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Studies in Labor Market Behavior: Sweden and the United States

Proceedings of a Symposium at IUI, Stockholm, July 10-11, 1979

THE INDUSTRIAL INSTITUTE FOR
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STUDIES IN LABOR MARKET BEHAVIOR: SWEDEN AND THE UNITED STATES

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FOREWORD

Labor market studies have long been a central theme of IUI research, covering such topics as labor mobility, education and productivity on the one hand, and wage and salary formation on the other. This area has become even more important during the last decade as new legislation affecting the labor market is proposed and enacted and as new demands for rapid structural adjustment are imposed on Swedish industry by changing international market conditions.

Since comparative research from different countries is especially valuable in addressing these issues, IUI organized a small conference in July, 1979, around some of them. The papers given at the conference are presented in this volume. They cover a number of important areas, and several of them compare labor markets in Sweden and the U.S. The topics are closely related to the economic policy debate of today in both countries: youth unemployment, the determinants of labor supply and the relationships between profits and wage change. Both theoretical propositions as well as the effects of labor market intervention policies have been analyzed.

The Institute wants to thank all outside participants for their contributions. We hope that this volume will find interested readers and act as a stimulus to future research in labor economics.

Stockholm, November 1981

Gunnar Eliasson

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Labor Market Behavior in Sweden and the U.S.

An Introduction

Gunnar Eliasson, Bertil Holmlund
Frank P. Stafford

The functioning of the labor market is of central importance for the possibility to realize several traditional objectives for economic policy. Macroeconomic *stabilization* policies attempting to affect inflation and unemployment will require quantified knowledge about how wages are formed and unemployment determined. Likewise, a policy intended to improve resource *allocation* and to foster economic *growth* must be based on information on how workers respond to changes in the available compensation packages and how on-the-job training is acquired. Finally, it is obvious that the application of an *income distribution* policy should be based on a firm understanding of how the initial wage and income inequalities were generated.

The IUI-symposium on labor economics in July 1979 offered a unique opportunity to discuss some of the above mentioned topics in a comparative U.S.–Swedish setting. Emphasis was on unemployment and unemployment policies, labor supply and wage determination. Several papers included comparisons between Sweden and the U.S.—two countries with supposedly very different labor markets.

Sweden vs the U.S.

Swedish employers and employees are operating in a labor market setting that exhibits several unique features compared to the situation in other countries and especially in the U.S. One characteristic feature of the Swedish post-war economy is the extensive application of various *selective labor market programs*. During the deep recession 1976–1978—with an unemployment rate around 2%—these programs employed almost 4% of the labor force. In addition to traditional labor market programs (manpower training, temporary public jobs), several institutional changes have taken place during the 70s. Of special importance is the employment security legislation from 1974, which tends to make employer-initiated job separations very costly.

Other policy parameters with implications for labor market behavior are defined by the Swedish *tax system*. Total tax yields, as a percentage of GNP, have increased from 21% to 53% during the period 1950 to 1977; Sweden has now the highest tax rates in the OECD-area. Of special relevance is that Swedish tax scales are highly progressive, with marginal tax rates around 60–70% for most of full-time working employees.

A third noteworthy feature of the Swedish labor market setting is the role played by *centralized collective bargaining*. An explicit objective for the Trade Union Confederation has been to enforce a more "equitable" wage structure, in practice regularly interpreted as requiring a reduction of prevailing before tax wage differentials. This wage policy has presumably been at variance with the wage structure that a market "solution" would imply.

The policies pursued in Sweden have had substantial effects on labor market behavior. The labor market programs have been able to reduce individual welfare losses associated with high unemployment. To illustrate, there has been a substantial reduction in the layoff rate subsequent to the 1974 employment security legislation. On the other hand, there might have been adverse allocation and growth effects. The employment preserving measures as well as the progressive tax system have most likely weakened the incentives for labor to move in response to offered wage increases and obstructed productivity-augmenting structural changes in the economy. This is partly reflected by the decline in the new hire rate subsequent to the 1974 legislation and many believe the burden of the reduced new hire rate to have been disproportionately borne by youth and other job market entrants.

Compared to Swedish conditions, the U.S. labor market may seem relatively free from government regulation. It is, however, important to note that in the U.S. "market solutions" to an important extent are also influenced by collective bargaining processes. The outcome of collective bargaining in the U.S. occasionally displays striking similarities with some recent Swedish labor market legislation. The emphasis on *tenure* is a case in point here; U.S. practice as well as Swedish legislation apply "last in—first out" as the basic layoff rule.

Whereas firm-specific employment security has been emphasized in Swedish labor market policies of the 1970s, this has not been a universal element of what may be called the Swedish labor market model. Most social benefits and pension schemes in Sweden are unrelated to tenure in a particular firm, in contrast to the extensive application of various firm-specific fringe benefit schemes in the U.S. Seniority rules and firm-specific benefits will, of course, not necessarily have adverse productivity effects. The weakened mobility incentives may be offset by other effects conducive to on-the-job learning and work efficiency. Experience from the Japanese labor market underlines the relevance of such explanations of productivity growth. Suffice it here to say that the incentive structures facing Swedish and U.S. workers have different content—but not necessarily different effects. A Swedish employee contemplating job mobility will find that the major part of his prospective income increase will be absorbed by taxes. The U.S. worker, in contrast, can expect to keep the major fraction of an income increase; on the other hand, he or she must consider the possibility of firm-specific fringe benefit losses associated with job mobility. Simple views of institutional and incentive differences between Sweden and the U.S. can be quite wrong at places.

One might expect the revealed inclination for social innovations on the part of Swedish authorities to have a mirror image in the form of a substantial ongoing program of policy evaluation and research. Some policy-oriented research has been undertaken, but in the area of our chosen topic—labor market behavior—it is not difficult to find numerous areas which have been analyzed only briefly so far.

For instance, practically no research has been devoted to the macro economic effects of the system for temporary jobs (relief works), despite its pivotal role in Swedish labor market policy; relief workers accounted for about 1% of the labor force during the last recession of the 1970's.

Empirical evidence on incentive effects of the Swedish income tax system is also sparse or close to non-existent. Conventional wisdom suggests substantial, negative responses on labor supply from a steeply progressive tax system. In the U.S. the *joint* income of married couples is taxed. In Sweden married persons are taxed at rates independent of the earnings of the spouse. Hence one should expect differences in the contribution of married women to labor supply between the U.S. and Sweden. Numerous policy differences of this kind should give rise to differences in the supply composition of the U.S. and Swedish labor markets but predictions made have been left more or less untested so far.

Our knowledge of how centralized negotiations affect the process of wage changes is also very limited. To what extent has union policy been able to influence the structure of wages and wage changes over time?

These examples represent only a few of the areas which have been left relatively unexplored by empirical research. The first part of this volume discusses some policy issues of importance and tests certain fundamental hypotheses about labor market behavior. What is the role of inflationary surprises for unemployment fluctuations and how valid is the human capital interpretation of returns to education?

Unemployment and Unemployment Policies

The first set of studies considers the role of *labor market policy* in Sweden. *Stafford's* introductory paper primarily reviews the major public policies designed to influence unemployment in Sweden and the U.S. and comments on their effects. There are many conceptual problems associated with measuring unemployment. *First*, there is the problem of deciding what it should stand for: Do we want a measure of individual welfare costs (psychological sufferings, social hardship), or a measure of the output losses associated with an underutilized labor force? We need at least two, probably several, measures to capture these two different concepts. *Second*, one has to decide what existing measures in fact represent and to what extent comparability can be obtained. Stafford's paper shows that the countries are in fact very similar in terms of structural variables that explain intercountry differences in unemployment rates. He finds that the fraction of 18 year olds enrolled in school shows a strong positive association with unemployment indicating that the larger the youth cohorts entering the labor market the larger is youth unemployment. The basic difference between the labor markets in the U.S. and Sweden turns out to be the much larger emphasis placed on selective labor market policies in Sweden. This difference presumably explains a substantial part of the observed unemployment rate differentials.

Johannesson explores the composition and development of Swedish labor market policy in detail. He shows that expenditures on labor market policy have increased gradually during the post-war period, with marked increases during recessions and with negligible decreases during boom years. Policy-priorities have changed in interesting ways. The relative importance of supply side measures in Sweden—heavily emphasized during the 1960's—have been reduced in favor of demand-oriented measures during the 1970's. These demand-oriented measures have, furthermore, been oriented primarily towards preventing unemployment through layoffs in the private sector.

Holmlund's paper investigates the effects of different labor market policy programs in Sweden. The policies considered include traditional programs—such as temporary jobs and manpower training—as well as recent policy innovations and the employment security legislation in particular.

Holmlund uses a longitudinal data set in order to estimate how participation in labor market programs affect future unemployment risks. The results obtained indicate, *inter alia*, a marked autocorrelation in individual unemployment probabilities; workers with previous unemployment experiences are facing much higher current unemployment risks even after controlling for various personal characteristics. However, these adverse effects of previous unemployment appear to be mitigated by participation in labor market programs.

Another interesting observation is that the higher separation costs introduced through the employment security legislation have decreased layoffs but have also induced more careful screening procedures on the part of firms. The net effect so far, however, turns out to be a substantial reduction in measured unemployment.

The paper by *Gramlich and Ysander* evaluates the role played by Swedish relief work programs for the local governments' employment demand. To what extent are relief workers performing jobs "normally" done by regular employees, thereby reducing normal labor demand? Given the great emphasis placed on relief works in Sweden, the possibility of such displacement effects should clearly be taken account of and Gramlich's and Ysander's paper represents an interesting econometric method to do so. The aggregate data used hardly allow any far-reaching conclusions; the results do, however, indicate considerable displacement effects in the public road work while no such effect is observed for health and welfare relief workers.

Unemployment Fluctuations and Inflationary Surprises

The next theme in the volume deals with the determinants of short-run unemployment fluctuations. *Burdett's* study approaches this issue from a search-theoretic viewpoint. The problem addressed is how changes in labor demand affect the outcome of a job search process. A shift in labor demand conditions will affect the worker's job offer probability as well as his reservation wage. The outcome depends strongly on the extent to which changes in conditions are fully predicted by the worker.

Burdett uses a search model where expected discounted lifetime income is maximized. *Unpredicted* improvements in labor market conditions will decrease the expected duration of unemployment because they do not affect reservation wages. Burdett, however, also demonstrates that *fully predicted* changes in labor market conditions may have unambiguous implications for the outcome of the search process. In particular, Burdett derives the sufficient restrictions that must be placed on the wage offer distribution in order to obtain determinate results.

The paper by *Björklund and Holmlund* investigates the extent to which fluctuations in actual unemployment duration are explained by short-run deviations between actual and expected wages, as predicted by search theory, and the extent to which they are caused by fluctuations in job offers. The paper demonstrates that unexpected inflation can explain some of the short-term unemployment in Sweden. But it also turns out that inflationary surprises can explain only a small part of actual fluctuations in unemployment duration. Changes in job availability are found to be the most important determinant.

In addition to these results the study by Björklund and Holmlund offers a comparison of unemployment patterns in Sweden and the U.S. It is interesting to see that cyclical unemployment fluctuations in Sweden are almost exclusively due to changes in unemployment duration whereas unemployment inflow is a significant additional source for aggregate unemployment fluctuations in the U.S.

The Supply of Labor

A third group of studies in this volume is focused on issues related to *labor supply*. Among these is the paper by *Axelsson, Jacobsson and Löfgren*, which includes estimates of neo-classical labor supply functions on Swedish household data. The results are well in conformity with earlier results from U.S. studies as far as males are concerned: men's labor supply appears to be relatively insensitive to changes in the wage rate. Male labor supply is furthermore shown to be unaffected by the presence of children in the household. The Swedish female labor supply functions have significant negative slopes, contrary to most results obtained in U.S. studies. However, when local tax rates are introduced in the labor supply equations, negative tax elasticities are arrived at. The simultaneous prevalence of negative (gross) wage elasticities and negative tax rate coefficients is somewhat puzzling and possible to interpret only with some difficulty. As the authors point out, measurement errors in reported hours will imply a negative bias in the wage rate coefficient.

The question of labor supply responses to changes in tax rates is also in focus in the paper by *Jakobsson and Normann*. Their procedure, however, is quite different from the traditional econometric approach pursued in the aforementioned paper. Jakobsson and Normann start off with an explicit utility function of the individual with (net) income and leisure as arguments. Utility maximization yields labor supply as a function of an exogenous wage rate and exogenous tax parameters. This "micro-model" is embedded in a simulation model of the Swedish system for personal income taxation and the labor supply responses to certain policy changes are investigated. The results are also evaluated by means of an explicit social welfare function, that takes account of individual utility levels as well as the dispersion of individual utilities.

The simulations reveal the existence of *perverse government revenue effects*. In other words, increases in marginal tax rates will actually *decrease* government tax revenues because labor supply diminishes as a consequence of the increased marginal tax rate. This result is, however, not consistent with the estimated tax rate elasticities obtained by Axelsson, Jacobsson and Löfgren and clearly indicates the need for further research on labor supply effects of the highly progressive Swedish tax system. To Jakobsson and Normann the Swedish income tax shedules differ greatly from those that would be prescribed by the theory of optimal taxation. They also conclude that more lump sum transfers should increase social welfare and that which policy to choose is quite independent of the importance attached to a more or less even income distribution in the social utility function.

Applications of Human Capital and Signalling Theories

The paper by *Gustafsson* deals with the labor supply issue from a different viewpoint. The focus here is not how wage rates affect labor supply behavior but instead how previous labor supply decisions influence current wages. Of primary interest is the extent to which male-female wage differentials can be explained by differences in education and work histories between the sexes. Standardizing for these human capital related variables, it turns out that Swedish women earn about 20% less than Swedish men; the corresponding U.S. earnings differential appears to be somewhat larger.

The human capital interpretation of the return to education is challenged in *Albrecht's* paper. The basic objective is to develop an econometric procedure to test the *signalling* model and the question is whether employers use education for purely informational purposes in their hiring decisions. The role of education is decomposed into a pure "productivity" component and a pure "information" component. The basic idea explored is that employers will be forced to rely more heavily on education when considering those applicants about whom they have the least information. The hypothesis is tested—but not supported—by means of a data set that includes information on hired job applicants as well as refused applicants.

The Determinants of Wage Changes

The remaining two papers in the volume are both studies on the process of *wage changes*. *Schager's* paper focuses on wage drift in Sweden, i.e., the difference between total wage increases and centrally negotiated increases. Several earlier studies have documented a close (Phillips-) relationship between wage drift and unemployment for post-war years. During the 1970's, however, this relationship seems to have vanished. Schager's theoretical framework of the wage drift process is focusing on firms' active recruitment behavior and suggests that the *duration* of vacancies should be the variable most closely related to the tightness of the labor market. The empirical results strongly confirm this hypothesis. Another noteworthy result in Schager's paper is that also profits appear to be a significant factor behind wage drift. A remarkable finding is that inflationary expectations—measured by changes in consumer prices—play a negligible role for wage drift.

The determinants of wage increases are also the topic for *Jonsson's and Klevmarken's* paper. Their approach represents an interesting attempt to integrate the human capital wage theory with Phillips curve oriented views of wage changes responding to market disequilibria. The analysis—performed on pooled cross-sectional data for salaried employees—shows that both market changes and the outcome of central negotiations are important to explain age-earnings profiles. The study clearly indicates that downward wage rigidity is a characteristic feature of the labor market; salaries are much more sensitive to excess demand situations than they are to excess supply. Another interesting finding is that central negotiations have a substantial net effect on salary growth when the market is characterized by excess supply but no significant effect when excess demand prevails.

Conclusions

As this brief overview indicates, the conference dealt with a variety of important issues. We believe that the volume has filled some gaps in our knowledge of labor market behavior and that it will stimulate further research as well. Unexplored and important research areas abound.

For instance, the desirability of careful *evaluation research* could hardly be underestimated in a country like Sweden which devotes 2–3% of GNP to labor market programs. The study of displacement effects included in this volume outlines a methodology that can be applied to other—and hopefully richer—data sets. Evaluation of labor market programs should also provide some information on individual welfare effects of program participation as compared to unemployment.

The quantitative importance of labor market programs makes it necessary to address the displacement issues from a broad allocational perspective as well. To what extent is subsidized employment in sheltered or semi-sheltered workplaces crowding out private sector employment? Our knowledge here is extremely scarce.

Another issue that should be further analyzed concerns incentive effects of the Swedish system for personal taxation. The somewhat conflicting results from two studies in this volume are cases in point here. But the tax system has, of course, implications for a broader set of labor market phenomena, e.g., the mobility of labor between firms, regions and labor force states.

It is also important to understand the extent to which the different agents in the labor market are involved in active search activities. Traditional search models emphasize job search among workers, whereas active search on part of the firms have been largely ignored in the literature. A better understanding of unemployment and mobility patterns in the labor market will require a more careful analysis of firms' search and recruitment behavior.

Finally, much discussion today centers around the allocation effects and the macroeconomic consequences of alternative labor market policies. The papers presented in this volume have dealt with the allocation theme only in passing, but the papers still represent a wealth of evidence on the matter. Combined with other evidence it should eventually be possible to shed some coherent light on a number of very pressing and complex policy problems of today that so far appear to be resolved without recourse to the necessary background knowledge.

Part I

Unemployment and Unemployment Policies

Unemployment and Labor Market Policy in Sweden and the United States

Frank P. Stafford

1. INTRODUCTION*

Unemployment rates are far lower in Sweden than in the United States. During the most recent recession in Sweden the unemployment rate was 2.2 percent (1977-78), while during the most recent recession in the United States unemployment averaged 8.5 percent (1975). These differences cannot be explained by simple differences in methods of measurement since both countries rely on a household survey with similar question sequences and definitions of unemployment.[1] The large difference in measured unemployment raises the question of the extent to which the unemployment rates differ because of differences in public policy and to what extent they differ because of structural differences in the labor markets, such as proportion of teenagers, duration of spells of older workers, or rate of growth of the labor force.

In this paper we review the major public policies designed to influence unemployment in Sweden and

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[1] International Comparisons of Unemployment, U.S. Bureau of Labor Statistics, 1978, p.14.

* I would like to thank Gunnar Eliasson, Ned Gramlich and Bertil Holmlund for helpful suggestions.

the United States and comment on their effects on unemployment. A cost-benefit analysis of the policies is not developed, and it is therefore possible that some of the policies have costs exceeding their benefits, even though they appear to pass the direct test of lowering unemployment rates. Cost-benefit analysis in this area is seldom applied and one of the major difficulties is defining the benefits realized by foregoing unemployment.

If unemployment is reduced through a policy which shows long duration spells, the benefits could be approximated by the reduced loss of market output less the value (positive or negative) of unplanned or involuntary leisure. On the other hand, some short duration spells may truly reflect job search at the intensive margin, and eliminating these spells may represent a cost rather than a benefit. If the policy itself had other implementation costs, then in this latter case the analysis would involve determining the sum of the two cost components.

This paper provides a comparison of major differences in labor market policy between the two countries and offers comment on the structural similarities and differences in the labor markets. Most of the differences arise from differences in the demographic composition of the workforce and from differences in education and tax policies. The more ambitious task of cost-benefit analysis of the various labor market programs is not attempted.

2. OVERALL DIFFERENCES IN UNEMPLOYMENT AND LABOR MARKET POLICY

In both Sweden and the United States there has been increased emphasis on countercyclical labor market policy in the last 10 years. Both countries placed greater reliance on supply oriented policies in the 1960's and have come to place more emphasis on demand oriented policies in the 1970's, though the overall level of resources devoted to labor market policy, defined as expenditures for training, employment and cash assistance to the unemployed by the Federal government, is higher in Sweden than in the United States. This can be seen in Table 1. About 1 percent of Swedish GNP was devoted to labor market policy in the mid 1960's but by the late 1970's (1976 and 1977) about 3 percent of Swedish GNP was devoted to labor market policy. For the U.S. the comparable figures are 0.5 percent in the mid 1960's and 1.25 percent during 1975-77. The share of total federal labor market resources in the form of cash assistance to the unemployed is far higher in the United States ranging from as high as 85 percent in 1965 to an (estimated) low of 47 percent in 1979. For Sweden the largest share of labor market resources in the form of cash assistance to the unemployed was 13 percent during 1978.

Comparing the two countries in terms of share of GNP devoted to training and employment measures, the last line of Table 1 indicates that between 1965 and 1975 Sweden was spending about 8 times more than the United States. Since 1975 the U.S. has increased resources to training and employ-

Table 1. Expenditure and labor market policy measures in relation to GNP and labor force, Sweden and the United States, 1963-1980

	1963	1964	1965	1966	1967	1968	1969	1970	1971
SWEDEN									
Training and employment expenditure[a]	663.9	-	891.4	978.3	1,461.7	1,778.1		2,112.9	3,382.3
Unemployment compensation[a]	58.1	-	76.0	99.1	172.1	242.4		257.9	462.6
Total[a]	722.0	-	967.4	1,077.4	1,633.8	2,020.5		2,370.8	3,844.9
GNP[a]	81,765[e]	91,000[e]	98,300	105,700	109,000	116,200	136,000	151,400	159,300
Total as percentage of GNP	0.9	-	1.0	1.0	1.5	1.7		1.6	2.4
Total per person in labor force(Skr)	194	-	259	284	433	529	-	606	971
U-rate	1.7	1.6	1.2	1.6	2.2	2.3	1.9	1.5	2.5
Average duration[g]	9.0	9.6	8.3	8.4	10.3	10.8	12.4	12.0	13.5
Recession year(1-0)	1	0	0	0	1	1	0	0	1
UNITED STATES									
Training and employment expenditure[h]	209	299	534	989	1,236	1,587	1,560	1,602	1,952
Unemployment compensation[h]	-	-	2.980	2.440	2.552	2.412	2.583	3.364	6.168
Total	-	-	3,514	3,429	3,788	3,999	4,143	4,966	8,120
GNP[h]	516,300	616,200	657,100	721,100	774,400	829,900	903,700	959,000	1,019,300
Total as percentage of GNP	-	-	.53	.48	.49	.48	.46	.52	.80
Total per person in labor force ($)	-	-	45	43	47	49	49	58	93
U-rate[i]	5.7	5.2	4.5	3.8	3.8	3.6	3.5	4.9	5.9
Average duration[j]	14.0	13.5	12.0	10.5	8.8	8.5	8.0	8.8	11.4
Recession year(1-0)	0	0	0	0	0	0	0	0	1
Percent GNP on training and employment in Sweden	-	-	$\frac{.907}{.081}=11.2$	$\frac{.925}{.137}=6.8$	$\frac{1.341}{.1596}=8.4$	$\frac{1.530}{.191}=8.0$	-	$\frac{1.39}{.167}=8.3$	$\frac{2.123}{.191}=11.1$

	1972	1973	1974	1975	1976	1977	1978	1979	1980
SWEDEN									
Training and employment expenditure[a]	3,802.3	3,845.7	3,831.1	4,736.3	7,316.7	9,824.1	7,025.3[c]	8,025.6[d]	-
Unemployment compensation[a]	435.2	507.0	449.5	347.4	462.8	794.1	1,082.3[c]	1,090.9[d]	
Total[a]	4,237.5	4,352.7	4,280.6	5,083.7	7,779.5	10,618.2	8,107.6[c]	9,1165[d]	-
GNP[a]	172,200	190,400	219,000	251,700	284,300	308,500	347,300[f]	389,000[f]	
Total as percentage of GNP	2.5	2.3	2.0	2.0	2.7	3.4	2.3	2.3[f]	-
Total per person in labor force(Skr)	1,067	1,094	1,059	1,231	1,872	2,544	1,926	2,145[f]	-
U-rate	2.7	2.5	1.9	1.6	1.6	1.8	2.2	1.7	-
Average duration[g]	16.2	16.7	15.8	15.6	15.2	15.3	16.2	16.5[f]	
Recession year(1-0)	1	0	0	0	1	1	1	0	-
UNITED STATES									
Training and employment expenditure[h]	2,894	3,283	2,910	4,063	6,288	6,877	10,784	11,729[f]	11,002[f]
Unemployment compensation[h]	7,076	5,356	6,065	13,459	19,452	15,258	11,769	10,295[f]	12,410[f]
Total	9,970	8,639	8,975	17,522	25,740	22,135	23,337	22,025[f]	23,412[f]
GNP[h]	1,110,500	1,237,500	1,359,200	1,457,300	1,621,700	1,834,400	2,043,400	2,289,400[f]	2,505,700[f]
Total as percentage of GNP	.72	.70	.66	1.20	1.59	1.21	1.14	.96	.93
Total per person in labor force ($)	112	95	96	186	266	222	228	214[f]	223[f]
U-rate[i]	5.6	4.9	5.6	8.5	7.7	7.0	6.0	5.9	-
Average duration[j]	12.1	10.0	9.7	14.1	15.8*	14.3	11.8	10.7[k]	-
Recession year (1-0)	1	0	0	1	1	0	0	0	-
Percent GNP on training and employment in Sweden	$\frac{2.203}{.26}=8.5$	$\frac{2.02}{.265}=7.6$	$\frac{1.749}{.714}=8.2$	$\frac{1.88}{.299}=6.7$	$\frac{2.57}{.38}=6.6$	$\frac{3.184}{.575}=7.1$	$\frac{2.022}{.527}=3.8$	$\frac{2.06}{.512}=4.0$[f]	-

[a] Millions of Swedish Crowns in current prices. Source: Johannesson, Jan, "Swedish Labour Market Policy during the 1960's and the 1970's", Stockholm, April 1979.

[b] From 1967 on data are for 1967-68, 1968-69,...,for Sweden.

[c] Allocation as reported in Swedish Employment Policy, 1977/78, p.51.

[d] Requested for 1979/80, Swedish Employment Policy 1977/78, p.51.

[e] Statistiska Meddelanden, National Accounts 1963-64, Swedish Central Bureau of Statistics 1975:98, p.52 (estimated).

[f] Estimated.

[g] Statistiska Meddelanden, The Labour Force Surveys 1963-75, National Central Bureau of Statistics 1978:32, p.101.

[h] Budget of the United States, 1963-79.

[i] Economic Report of the President 1978.

[j] Employment and Earnings, 1963-79, Table A37.

[k] January.

Source: Handbook of Labor Statistics 1977. U.S. Pol. Res. Lab. Stat. Bulletin 1966.

ment, particularly youth programs which are estimated at about $ 3.5 billion of the $ 11.7 billion for training and employment in 1979. Estimates for 1978 and 1979 indicate that Sweden now spends only 4 times more than the United States on training and employment measures as a share of GNP.

In this regard the U.S. is more similar to Germany in emphasizing what Johannesson and Schmid refer to as a compensatory policy rather than selective demand policy. In comparison to Germany, Sweden spends five times more for selective demand measures relative to the Gross National Product.[1]

Generally, Sweden has a far more active labor market policy[2] and has placed more emphasis on a greater diversity of demand side labor market programs. The U.S. has utilized public sector employment and subsidies for employment of welfare recipients (the Work Incentive Program), while Sweden has utilized public sector employment, in-plant manpower training to avoid layoffs, wage subsidies, sheltered and semisheltered employment for the handicapped, subsidies for inventory accumulation, and advance warnings of layoffs.[3] The chang-

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[1] Johannesson, Jan and Schmid, Günther, "The Development of Labor Market Policy in Sweden and in Germany: Competing or Convergent Models to Combat Unemployment?", International Institute of Management, Berlin, May 1979, p.iii.

[2] As distinct from aggregate monetary and fiscal policy.

[3] <u>Swedish Employment Policy 1977/78</u>, National Labor Market Board, 1978, pp.1-12, 35, 39, 41, 46-48.

ing patterns in Swedish labor market policy are well summarized by Jan Johannesson.[1] The public policy in Sweden does appear to have averted a significant rise in unemployment rates during the 1976-78 recession. As can be seen in Table 1 the unemployment rate rose from 1.6 percent to a high of 2.2 percent in 1978.

Because several new policies were operating simultaneously along with increased levels of previously established policies, the relative effectiveness and cost of the different policies are less clear.

The most important countercyclical labor market policies would appear to be the inventory subsidy program, government orders to firms which have given notice of reduced operation or closure, advance notification to workers of layoffs, the program for in-plant training to avoid layoffs, and the temporary jobs program. Of these programs perhaps the most distinctive is the advance notification of layoffs. Since July 1974 there is a statutory requirement that the Labor Market Administration shall be informed in advance of layoffs and closure planned by firms. A layoff notice is required for employees who have been with the firm for six consecutive months or for a total of more than 12 months during the past two years and the duration of lead time varies from one month for employees under 25 years of age to six months if

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[1] Johannesson, Jan, "Swedish Labor Market Policy in the 1960's and 1970's", Expertgruppen för utredningsverksamhet i arbetsmarknadsfrågor (EFA), Stockholm, April 1979.

the employee is over 45.[1]

Firms may not lay off workers who have not been given advance notice. Such a policy could have a substantial impact on layoffs since a firm will presumably bear costs if workers, disgruntled by knowledge of impending layoffs, are less productive. On the other hand there are no formal penalties for failing to lay off someone previously notified, and firms may face a net assymetric loss function for errors of over- and underpredicting layoffs. With smaller losses to overstating planned layoffs, warning notices should exceed subsequent layoffs, but the policy still may have had a significant impact in keeping the unemployment rate low during the recent recession.

Besides a wider range of countercyclical labor market policies in Sweden than in the United States there are also more resources devoted to policies which are not heavily countercyclical. Perhaps the most notable difference here is between the Swedish and U.S. employment services. In Sweden 70 percent of unemployed persons and some 30 percent of employed persons wanting a change in their employment situation use the employment service.[2] In contrast 27.5 percent of the unemployed jobseekers in the U.S. made use of a public employment agency during 1977.[3] Research evaluating com-

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[1] Swedish Laws on Security of Employment, Status of Shop Stewards and Mitigation in Labour Disputes, Ministry of Labor, Stockholm, Sweden, May 1977, pp.11-12.

[2] Att utvärdera arbetsmarknadspolitik, Statens Offentliga Utredningar (SOU) 1974:29, p.466.

[3] Employment and Training Report of the President 1978, U.S. Department of Labor, p.221.

pulsory notification of vacancies for jobs lasting longer than 10 days in selected Swedish counties (Kristianstad, Blekinge and Malmöhus) suggests that the impact is favorable, particularly for employed persons seeking new employment.[1]

Overall, Swedish labor market policy is more comprehensive than that of the U.S. or any other industrialized economy, and the recent policy applications appear to have averted a substantial rise in the unemployment rate during the 1976-78 recession. The questions which the different labor market policy commitments raise are:

1. Which of the various labor market policies in Sweden and the U.S. have been the most effective in influencing unemployment rates and other labor market outcomes, such as unemployment durations or wage growth for the less skilled? What have been the costs of the different policies? Were some policies redundant in the sense that their marginal payoff was reduced by the presence of others?

2. What are the non-policy or structural differences in the labor markets of Sweden and the U.S. which give rise to the large disparity in unemployment rates?

The purpose of this paper is to give an overview of the similarities and differences between Sweden and the U.S. in recent trends in employment and unemployment and the extent to which specific policies oriented to the labor market may (rather

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[1] <u>Labour Market Policy in Transition</u>, The Expert Group for Labor Market Research, Stockholm, 1978, pp.36-39.

than overall monetary and fiscal policy) have influenced the level of unemployment during both recessionary and normal periods, assuming one can imagine the latter as at least a hypothetical state. Further, there have been specific policies which have not been concerned with unemployment. For example, in the U.S. there has been a rapid growth of what is referred to as "work testing" of welfare payments such as the Food Stamp program. Benefit recipients without current employment are required to establish that they were looking for work as a condition of eligibility. This may encourage a substantial rise in nominal unemployment[1]: persons who would otherwise have been out of the labor force are conditioned to report to both the Employment Service and the interviewer for the Current Population Survey that they are looking for work.[2]

The focus of this paper is to exploit major differences between the countries in labor market com-

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[1] Kenneth W. Clarkson and R.E. Meiners estimate a 2.1 percentage point increase in 1976 arising from food stamp and AFDC work registration requirements. See their paper "Government Statistics as a Guide to Economic Policy: Food Stamps and the Spurious Increase in Unemployment Rates", Policy Review, Summer 1979. Richard Devins, using Current Population Survey data on AFDC and food stamp participants registering at the public employment service concludes that a 4 percentage point increase in the unemployment rate in 1976 would be a likely upper bound on the rise in nominal unemployment owing to worktesting of welfare recipients. See his paper, "Unemployment Among Recipients of Food Stamps and AFDC", Monthly Labor Review, March 1979, pp.47-52.

[2] This also raises a policy conflict for the Employment Service which is asked to function as a potential adversary to applicants who are seeking its services for job placement.

position and explicit and implicit labor market policies. These intercountry differences will be used to draw some inferences about the potential role of policy and labor force composition on employment and unemployment in industrialized economies. In terms of scientific method many of the inferences will be admittedly crude, and in a sense a goal of this paper is to identify questions which may be addressed by more traditional research methods. The belief is that international comparisons can yield important insights despite the potential and actual difficulties in data comparability.[1]

In the next section of the paper an analysis of long run differences in unemployment in Sweden, the United States, and other industrial countries will be attempted. After this rather global view of unemployment rate differences and possible explanations, a more detailed look at U.S.-Swedish differences in the structure of employment and unemployment will be considered in Section 4.

3. INTER COUNTRY DIFFERENCES IN LONG RUN RATES OF UNEMPLOYMENT

The demand side labor market programs in Sweden are unique not only in terms of range and magni-

[1] An illustration of this point is that most observers assume that labor force participation increases by married women are a normal pattern for growing industrial economies. Yet data on the Japanese labor force from the Annual Report of the Labor Force Survey (1974, 1970, 1965) show that participation rates of women have declined between 1965 and 1974, with the largest declines registered for women under the age of 45.

tude, but in terms of the tie-in of participation with actual or imminent unemployment. The study by the U.S. Bureau of Labor Statistics provided international comparisons of unemployment rates using a uniform definition of unemployment, and offered an adjusted unemployment rate series for Sweden over the period 1965 to 1976. In their adjusted series (Table 2) they included in unemployment a monthly average of persons in training for labor market reasons, work training programs, public relief works, archive and relief work for musicians, and sheltered and semisheltered workshops.

The adjusted unemployment rate in Sweden in column (4) is an answer to the question, "What has been the impact of labor market policy on the unemployment rate?", but it is an answer which needs qualification. First, some of those in the labor market programs in column (3) may have dropped out of the labor force in the absence of the programs, and if so (4) is an overestimate of the unemployment rate which would have prevailed in the absence of the policies. Further, as pointed out by Björklund[1], whether one wants use unemployment rates to indicate social hardship or labor market tightness or reserves would influence one's willingness to use the adjusted unemployment rate.

For the purpose of the analysis in this section, which is to examine the role of broad structural differences in labor markets on long run unemployment rates, it does seem appropriate to use the

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[1] Björklund, Anders, "The Measurement of Unemployment in Sweden", International Institute of Management, Berlin 1979.

Table 2. Alternative unemployment rates for Sweden 1965-1978

Year	(1) Official rate	(2) Rate adjusted to U.S. definitions	(3) Number in labor market programs[a] (1000)	(4) Unemployment rate based on (2) and (3)
1963	1.2	-	29[b]	2.5[b]
1965	1.2	1.2	33	2.1
1967	2.2	2.1	48	3.4
1968	2.3	2.2	63	3.9
1969	1.9	1.9	65	4.1
1970	1.5	1.5	70	3.3
1971	2.5	2.6	83	4.6
1972	2.7	2.7	103	5.3
1973	2.5	2.5	112	5.3
1974	1.9	2.0	102	4.5
1975	1.6	1.6	94	3.)
1976	1.6	1.6	112	4.3
1977[b]	1.8	-	138[b]	5.0
1978[b]	2.2	-	154[b]	5.8

[a] See text for the programs included. Annual averages.

[b] From Johannesson, Jan, "Swedish Labor Market Policy in the 1960's and 1970's", op.cit., p.7.

Source: International Comparisons of Unemployment Rates, U.S. Bureau of Labor Statistics, Washington, D.C., 1978, p.33.

adjusted figure. Use of the adjusted figures has its limitations, the major of which depends on whether other countries, notably the United States, should have their unemployment rates subject to analogous kinds of adjustments so that reasonable comparisons can be made. My tentative answer to this question is no. The first reason is because the U.S. and other industrial countries to be covered in this discussion have nowhere near as comprehensive a policy for labor markets as Sweden. Second, in the United States, although demand side labor market policies have grown, particularly for youth, there has not been a particularly strong tie between an individual's unemployment state and program eligibility. Although recent program implementation has emphasized low-income, long-term unemployed individuals,[1] analysis of the general experience in the 1970's indicates that much of the demand oriented policies have resulted in an implicit increase in unrestricted revenue sharing grants. In other words, there can be fiscal substitution; that is, with the "PSE (public service employment) approach it is quite possible that the S and L's (State and Local governments) will use their PSE subsidies to pay for incumbent employees, or, more subtly, for workers who would have been hired even in the absence of the program".[2] The evidence on the extent of fiscal substitution is not unanimously interpreted, but most researchers agree that there

[1] The Budget for Fiscal Year 1979, U.S. Government Printing Office, Washington, D.C., 1979, p.171.

[2] Johnson, George E. and Tomola, Tomas D., "The Fiscal Substitution Effect of Alternative Approaches to Public Service Employment", Journal of Human Resources, Winter 1977, pp.3-27.

exists a substantial gap between nominal transfers and effective net expenditures on S and L employment.[1] Moreover, increased employment is not necessarily an indicator of reduced unemployment since the new hires may come about through increased labor market participation. For several reasons then, it seems that major revisions to the U.S. unemployment rate are unwarranted if one wishes to compare Sweden and the United States.

What are some of the broad structural factors which might explain intercountry differences in unemployment rates? Here we will consider six major factors which may account for the large differences among industrial countries.

1. Aggregate demand pressure as indicated by price inflation. In a simple Phillips curve view one might expect demand conditions, as indexed by the rate of price inflation, to influence short run changes in unemployment.

   In terms of the "natural rate theory"[2] one would expect there to be a more modest relation between decade averages of inflation and unemployment. This is because suppliers of labor may be unable to distinguish average prices (overall inflation) from relative prices (particularly their wage rate relative

[1] Johnson and Tomola put it at 100 percent after 5 quarters for direct employment grants with administrative restrictions, but at 56 percent for wage subsidies. Ibid, p.17.

[2] Lucas, Robert E., "Some International Evidence on Output Inflation Tradeoffs", American Economic Review, June 1973, pp.326-34.

to the prices of goods and services) in the short run, but consistently large average price increases will be adjusted for in assessing relative prices. As a result there will not be erroneous acceptance of high nominal wage rates and a corresponding reduction in unemployment rates with routinely high inflation year after year.[1]

For countries with high decade averages of inflation there should still be somewhat greater levels of unanticipated inflation than for countries with low decade averages of inflation. If this is so, one would expect some relation between unemployment and inflation under the "natural rate hypothesis" as well as under a traditional interpretation of the Phillips curve. Friedman has argued that in transitional periods inflation can influence output adversely because relative price shifts are more difficult to assess in such an environment.[2] If so, high rates of inflation should be positively related to unemployment in a structural sense, though not necessarily in a single equation approach to predicting unemployment rates.

2. The share of employment in jobs characterized by formal, market employment relations rather than informal, family employment relations as

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[1] Particularly if the inflation is "pure" inflation with all prices rising at the same rate rather than "mixed" inflation with a dispersion in rates of price increase across commodities.

[2] Friedman, Milton, "Nobel Lecture: Inflation and Unemployment", Journal of Political Economy, June 1977, pp.451-72.

in traditional agriculture. This hypothesis has been advanced in the previously mentioned BLS volume. If a labor market is characterized by extensive division of labor and market signals to allocate labor among firms, one would expect greater use of formal wage and salary contracts. Changes in output demand among firms will give rise to reallocation of labor across firms and this may be thought of as causing frictional unemployment even during periods of normal aggregate demand.

3. Rate of growth of output of the economy. A view that there are jobs created by economic growth can be given a somewhat more economic interpretation if there is substantial on-the-job training. An unexpectedly slow rate of economic growth means that there is excess job capital in the current labor force. Firms will therefore have less incentives to hire and train new workers and this reduced demand for new hires will be borne disproportionately by youth. Even with wage flexibility the lower wage rates offered to young people will make employment less attractive relative to non-market activity and will induce periods of intermittant employment, unemployment and non-participation.

4. The size of entering cohorts. The arguments for the effect of changing cohort sizes are similar to growth of output, but the lower wage is initiated by supply side shifts. Even if the wage structure across demographic groups is not affected, a large influx of young job market entrants will imply greater

turnover directly and indirectly as firms alter their workforces to utilize the new entrants.

5. Related to 4 is the school-to-work transition. Schooling of older youth (age 16-19) often involves intermittant labor market participation and the extent to which older youth continues in school varies substantially across industrialized countries.

6. Labor force participation of women. Increased labor force participation rate by women increases the supply of part-time workers. This in turn lowers the wages for youth and makes the desirability of employment more ambiguous for both youth and women. As a result there is more time spent in transition states between employment and non-market activity.

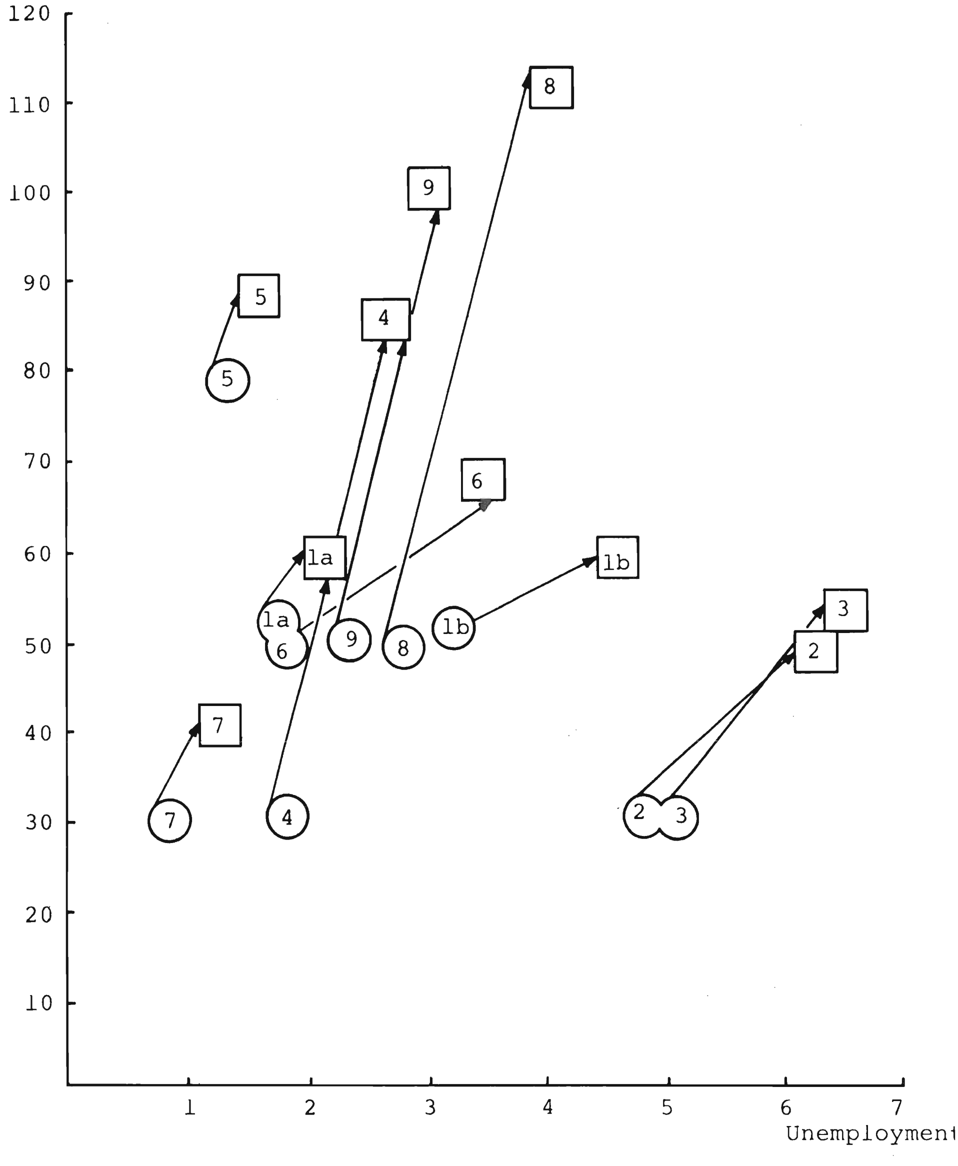
7. Policy measures. Various labor market policies may influence long run unemployment rates. These include both the training and employment measures of various kinds discussed in Section 2 and unemployment compensation, which has been alleged to extend the duration of unemployment spells and to increase the probability of spells.

Data on unemployment rates consistent with the definitions used by the Bureau of Labor Statistics, by decade, for 9 industrialized countries are given in column (A) of Table 3. Data on percentage change in CPI (ΔCPI) by decade are given in column (B). The share of the labor force in wage and salary employment (PCTWGSAL) is given in column (C). Data were not available for the 1960's and in subsequent regression analysis it was assumed that the 1974 fraction applied in the 1960's as well. The rates of growth of GNP were not consistently available in the OECD statistics, and it was decided that growth of industrial production (ΔINDPROD) which was consistently available for all countries in both decades could be used instead [column (D)]. Data on change in the composition of the labor force (ΔTEEN, ΔWOMEN) and percent of 18 years olds enrolled in school (PCTSCHOOL) were available in the BLS report [columns (E)-(G)].

Except for the Swedish unemployment rates, which have been adjusted to include those in labor market programs targeted to the unemployed, the data in Table 3 do not include policy variables since these are less readily quantifiable. For the purpose of analysis the approach will be to examine the relation of inflation and non-policy variables to long run levels of unemployment. Against this more general setting we will turn to a discussion of major developments in employment and unemployment in Sweden and the United States.

Turning first to the simplest Phillips curve relation, in Figure 1 are plotted data on decade averages of unemployment and inflation. The data points with encircled numbers indicate the 1960's, and the data points with numbers in boxes indicate the 1970's. The order of the numbers is that of the countries listed in Table 3. The data suggest a weak inverse relationship within decade between price inflation and unemployment but with all countries reporting higher levels of both inflation and unemployment as one more across decades as indicated by the arrows.

Figure 1. Unemployment rates and inflation for selected industrial economies by decade, 1960-70 and 1970-76



Source: Unemployment rates: U.S. Bureau of Labor Statistics, International Comparisons of Unemployment Rates. Price inflation: Economic Report of the President, 1978, p.378.

Table 3. Unemployment, inflation and selected structural variables for nine industrial economies, 1960-69 and 1970-76

Country	(A) Unemployment rate[a]	(B) ΔCPI[b] (%)	(C) Pct. labor force[c] in wage and sal. empl.	(D) ΔIndustrial[d] production (%)	(E) Change in share[c] of LF teenagers	(F) Pct 18 year[c] olds in school	(G) Change in share of women in LF
	(a) (b)						
(1) Sweden							
1960-69	1.7 (3.2)	50		70.9	-3		+5
1970-76	2.1 (4.6)	59	91.0	21.3	0	40.7 (1972)	+3
(2) United States							
1960-69	4.8	30		67.8	+2		+5
1970-76	6.2	47	90.4	16.8	+1	58.1 (1970)	+3
(3) Canada							
1960-69	5.0	30		80.5	+1		+7
1970-76	6.1	53	88.7	28.6	+1	45.5 (1970)	+3
(4) Australia							
1960-69	1.8	28		60.0	-2		+3
1970-76	2.6	85	85.8	20.8	0	18.0 (1972)	+2
(5) Japan							
1960-69	1.3	76		210.2	-5		-1
1970-76	1.5	88	69.3	42.7	-2	29.5 (1970)	-2
(6) France							
1960-69	1.8	48		60.6	-2		+2
1970-76	1.2	67	80.6	30.7	0	30.6 (1970)	+1
(7) Germany							
1960-69	0.7	30		45.5	-2		-2
1970-76	1.2	41	83.9	21.1	+1	12.9 (1969)	+2
(8) Great Britain							
1960-69	2.7	49		31.3	-2		+3
1970-76	4.0	111	92.0	5.2	-1	17.4 (1970)	+2
(9) Italy							
1960-69	3.3	47		84.2	-4		-3
1970-76	2.8	71	71.5	29.9	-1	19.7 (1966)	+2

[a] U.S. Bureau of Labor Statistics, International Comparisons of Unemployment, p.19-21.

[b] Economic Report of the President 1978, p.378.

[c] U.S. Bureau of Labor Statistics, International Comparisons of Unemployment, p.51 (percent of employment in wage and salary employment, 1974), p.50 (share of labor force consisting of women and teenagers, p.64 (percent of 18 year olds in educational institutions).

[d] Economic Report of the President 1978, p.379 and Main Economic Indicators, Historical Statistics 1960-75, OECD, Paris 1976 and Main Economic Indicators, OECD, January 1978.

Regression of decade average unemployment rates on decade average inflation rates[1] and selected structural variables are given in Table 4. The relation between unemployment and inflation is very weak and requires disaggregation by decade before the expected inverse relationship between inflation and unemployment is observed. The coefficient for ΔCPI is statistically significant in none of the equations, lending support to the view that the Phillips curve is not a long run relationship. The point estimate of the ΔCPI coefficient suggests that a 100 percentage point increase in the decade rate of inflation would lower unemployment by 3.2 percentage points in the 1960's (equation (2)), but by 1.4 percentage points in the 1970's (equation (3)). The same information is expressed in the interactive form in equation (4).

By far the most important structural variable explaining differences in unemployment is percentage of 18 year olds enrolled in school (PCTSCHOOL) with a 10 percentage point increase associated with a .9 percentage point higher unemployment rate. As can be seen from the correlation matrix at the bottom of Table 4, PCTSCHOOL is rather highly correlated with rate of entrance of women and teenagers into the labor market. For equation (7) PCTSCHOOL is deleted and ΔTEEN, ΔWOMEN have larger expected values, but still have high standard errors. Change in CPI exhibits a weak, negative correlation with unemployment (-.3 within decade), but, generally, the structural variables have much stronger correlations with unemployment and in the expected directions.

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[1] Inflation rates for 1970-76 are adjusted to allow for the fact that 7 years rather than 10 years are covered. The variables in columns (D), (E) and (G) are similarly adjusted.

Table 4. Regressions of decade average unemployment rates on decade average inflation rates and selected structural variables

	Equation						
Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)
ΔCPI	.001 (.011)	-.032 (.034)	-.014 (.021)	-.032 (.038)	-.021 (.025)	.015 (.051)	.008 (.069)
1970's (1-0)	-	-	-	.87 (2.61)	-.59 (1.70)	.36 (1.92)	.93 (2.62)
1970•ΔCPI	-	-	-	.018 (.043)	.027 (.027)	-.008 (.048)	-.007 (.065)
PCTSCHOOL	-	-	-	-	.088 (.019)	.081 (.026)	-
PCTWGSAL	-	-	-	-	-	.021 (.071)	.059 (.095)
ΔINDPROD	-	-	-	-	-	-.009 (.015)	.009 (.019)
ΔTEEN	-	-	-	-	-	.041 (.369)	.306 (.494)
ΔWOMEN	-	-	-	-	-	.066 (.186)	.127 (.255)
INTERCEPT	3.06	4.12	4.99	4.12	1.00	-1.45	-2.96
R^2(ADJ)	-.06	-.01	-.07	-.04	.58	.53	.108
SAMPLE	ALL	1960's	1970's	ALL	ALL	ALL	ALL

Correlation among structural variables and unemployment

Pct School	1.00					
Pct WGSAL	0.33	1.00				
ΔINDPROD	0.10	-.55	1.00			
ΔTEEN	0.46	.60	-.58	1.00		
ΔWOMEN	0.50	.75	-.31	.68	1.00	
UNEMPL	0.76	.52	-.28	.61	.60	1.00

In light of the small sample and rather descriptive nature of our regression analysis, drawing strong inferences about the relative importance of the different structural factors on unemployment seems unwarranted. Yet, the strength of the relation between unemployment and share of youth in school merits some comment. It may be that protracted schooling of youth creates a pool of unmarried "near adults" who have an ambiguous attachment to the job market. Their intermittant experience in the job market has some spillover effects which may be important. As employers replace young people with adults and vice versa there may be an increase in adult unemployment as well. Countries with early school leaving ages (Germany, Australia, Britain) are more likely to have their young people fully settled into the job market, whereas in the U.S., Canada and Sweden young people are less vocationally trained and may experience job transitions as part of acquiring information on occupational preferences or on-the-job training.

In summary, Sweden and the United States have labor market settings which are quite similar and which would give rise to high unemployment rates in both countries were it not for the role played by an active labor market policy in Sweden. If we pursue the implication of our analysis, that protracted schooling appears to be related to high unemployment rates, it suggests that recent labor market policies for youth in Sweden and the United States should be given close scrutiny. In Sweden there has been a dramatic rise in the share of youth participation in public employment programs[1], and in the United States some innovative

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[1] See Johannesson, op.cit. pp.21-24.

policies have been instituted which tie youth job program eligibility to school enrollment.[1] The evidence set out in this section also presents a challenge for search theory because the current models do not spell out how increased search and turnover by a given group (here young people) may, through demand side interaction, influence the search and turnover of other groups.

4. U.S.-SWEDISH PATTERNS OF EMPLOYMENT AND UNEMPLOYMENT

Overall, Sweden has experienced greater increases in labor force participation of married women and greater decreases in the labor force participation of married men. In the last 10 years both Sweden and the United States have experienced a rapid growth in the labor force participation rates (LFPR's) of women with the LFPR of women aged 20-64 reaching 69.0 percent in Sweden and 58.5 percent in the United States (see Table 5). Particularly remarkable is the rise in the LFPR for married women with preschool children (under age 7) in Sweden, and this may be partly related to the provision of subsidized childcare.[2]

In both countries participation rates of men 20-64 have declined only modestly while declines in LFPR's for married men of all ages are larger. In

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[1] The Youth Employment and Demonstration Projects Act. See the Budget for the Fiscal Year 1979, *op. cit.*, p.173.

[2] See Gustafsson, Siv, "Cost Benefit Analysis of Early Childhood Care and Education", Manuscript, Industriens Utredningsinstitut, May 1978.

Table 5. Weekly hours in the labor market and participation Rates, Sweden and the United States 1965-1975

SWEDEN
SCB Estimates

	Normal work (Hours worked last week)			Participation rates		
	1965	1975	Percent change	1965	1975	Percent change
Married Men	-	-	-	90.2	83.9	-7.0
Married Women						
All	31.0	29.4	-5.2	43.9	59.3	35.1
With children under 7	26.7	26.6	0.4	34.0	58.8	72.9
Men 20-64	46.0	41.5	-9.8	92.4	90.9	-1.5
Women 20-64	34.2	31.5	-7.9	53.4	69.0	29.2
Men 55-74	44.2	40.5	-8.4	68.6	55.5	-19.1

Source: Arbetskraftsundersökningen, Tables 1 and 19, 1965 and 1975.

UNITED STATES
CPS Estimates

	Normal work (Hours worked last week)			Participation rates		
	1965[b]	1975[c]	Percent change	1965[d]	1975	Percent change
Married Men[a]	44.2	42.9	-2.9	95.5	92.2[e]	-3.5
Married Women[a]	34.5	34.0	-1.4	38.7	49.0[e]	26.6
Men 20-64	43.9	42.6	-3.0	94.6	92.2[f]	-2.5
Women 20-64	35.7	35.0	-2.0	51.7	58.5	13.2

TIME DIARY ESTIMATES FOR THOSE WORKING[g]

	Normal work			Travel to work		
	1965[h]	1975[h]	Percent change	1965[h]	1975[h]	Percent change
Married Men	44.7(448)	41.3(244)	-7.6	5.0	4.5	-10.0
Not Married Men	46.0(73)	35.2(78)	-23.5	3.9	4.4	+12.8
All Men	44.9(521)	39.9(322)	-11.1	4.8	4.2	-12.5
Married Women	34.3(190)	26.5(117)	-22.7	3.2	2.3	-28.1
Not Married Women	34.9(152)	35.6(102)	+2.0	3.6	3.7	+2.8
All Women	34.6(343)	30.8(219)	-11.0	3.4	2.9	-14.7

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[a] Married, spouse present

[b] November 1965

[c] November 1975

[d] November 1965

[e] May 1976

[f] April 1976

[g] Sample size in parenthesis

[h] Source: 1965-75/76 Time Use Comparison Tape, Survey Research Center, University of Michigan. Hours of normal work were defined to include regular work for pay outside the home or brought home, overtime, waiting, or interruption during worktime (e.g., machine breakdown), and coffee breaks. Data are weighted using day of the week as a stratification variable, and are available only for those reporting at least 10 hours per week in the labor market. For further information see Stafford, Frank P. and Duncan, Greg J., "The Use of Time and Technology by Households in the United States", forthcoming in R.G. Ehrenberg (ed.), Research in Labor Economics, Vol.3.

both countries the major declines in LFPR's for men have been for those in age groups traditionally considered as "preretirement" ages.[1] In the United States much of the change has been attributed to the Social Security system. In 1972 the system began providing an option for benefits at 80 percent of normal amounts for those electing retirement at age 62 rather than age 65. Since this provision has been in effect the LFPR of those at age 62-64 has declined sharply while the LFPR of those at age 60-61 has not.[2] It can also be shown that the likely impact of the system on labor supply is to increase hours prior to retirement because retirement benefits are based on preretirement earnings. As a consequence the Social Security system in the United States may have had much of its effect on labor supply through changing the timing of lifetime labor market hours rather than simply reducing them through earlier retirement.[3]

In Sweden the national retirement income system was modified in July 1972 to make it easier to receive a pension beginning at age 60 rather than the normal retirement age (67 prior to July of 1976 and 65 since that time). The number of persons receiving an early pension has risen at an annual rate of 5.9 percent between 1972 and 1976 in comparison to an annual rate of 6.6 percent

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[1] See Table 5. The LFPR of Swedish men age 55-74 has fallen by 19.1 percent between 1965 and 1975.

[2] See Economic Report of the President, 1976, Table 34, p.114.

[3] Burkhauser, Richard and Turner, John, "A Time Series Analysis of Social Security and Its Effect on the Market Wage of Men at Younger Ages", Journal of Political Economy, August 1978.

between 1966 and 1972.[1] If there has been a substantial impact of changes in early retirement on labor force participation rates of older workers in Sweden, it does not appear to relate directly to the changed early retirement provisions as it does in the U.S. While LFPR's of older Swedish men have declined in recent years, it would appear that the decline is a continuation of a trend present in the late 1960's.[2] Hence, the policy changes in both countries may simply be collective responses to preferences for earlier retirement induced by rising family income and other social changes.

Average hours at work of those working exhibit larger declines in Sweden than in the United States. Partly offsetting the rapid growth in labor market participation by Swedish women is a decline in the average workweek of 8 percent from 34.2 hours to 31.5 hours between 1965 and 1975 as reported in the official statistics. If the U.S. experience is any guide the official statistics are likely to understate the decline in hours in the job market in Sweden.[3]

By using time diaries (Table 5, last panel) one can obtain a chronology of activities of respond-

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[1] Björklund, Anders, "The Measurement of Unemployment in Sweden", International Institute of Management, Berlin 1979.

[2] Labour Market Policy in Transition, op. cit., p.20.

[3] For Japan the official statistics show virtually no change in hours of market work for those participating but time diaries show declines of 16 and 27 percent for adult men and adult women, respectively. See Stafford and Duncan, op.cit., Table 5.

ents over a 24 hour period beginning at midnight and add up time spent in market work. This is an alternative to directly questioning the respondent for an estimate of hours of work per week which is subject to bias. With direct questioning respondents overstate time spent in socially acceptable activities and fail to conform to the simple budget constraint of 24 hours per day. In contrast, time diaries are non-directed and provide an expensive but methodologically superior approach to measuring time use.[1] If part of the increased leisure in both Sweden and the U.S. has come in the form of "partial absenteeism" (late arrival and early departure from work and long lunch breaks) and if respondent reports of hours represent hours per week for which a person can be obligated, then the "official" statistics of both countries may be understating the decline in hours actually on-the-job.

For both married men and married women the time diary estimates show larger declines than the CPS estimates, and the discrepancy is the greatest for married women. The large declines in labor market hours of those in the job market in the U.S. are also consistent with changes in other time uses of the employed population which indicate a 47 per-

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[1] In previous methodological work time-diary estimates were compared to estimates from recorded time use at random intervals indicated by a signal from a beeper carried by the respondent. The mean beeper and diary estimates matched quite well except for time uses away from home. Presumably respondents were reluctant to make public explanations of the beeper signal. See Robinson, John P., "Methodological Studies", draft, Survey Research Center, University of Michigan, 1976.

cent rise in time spent watching TV.[1] Perhaps new TV programming, which has succeeded in attracting the more educated viewers in the U.S., accounts for more of the declining rate of productivity growth than does the pattern of capital accumulation and obsolescence!

The labor force participation rates of U.S. teenagers have been rising over the last 15 years with cyclical plateaus or declines occurring during the recessions of 1971-72 and 1975-76.[2] Because of sharp rises in teenage unemployment rates during recessions combined with discouraged worker effects on participation rates, the employment/population rates for teenagers is strongly procyclical in the United States.

In Sweden the LFPR of teenagers declined between 1965 and 1970, largely because of the expansion of the school system, and this can be seen in Table 7. Subsequent to the leveling off in school enrollment rates in 1975, there has been a decline in labor force participation of teenagers during the 1977-78 recession. This is despite the fact that those in temporary jobs are considered to be in the labor force and the share of young people in relief work programs has risen dramatically[3] as

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[1] Stafford and Duncan, op.cit., Table 8. The employed adults in the U.S. spend an average of 14 hours a week in TV viewing as a primary activity. Data for Japan also show TV viewing as the major growth sector in adult time use.

[2] See Employment and Earnings, November 1978, Chart 3, p.8.

[3] Johannesson, op.cit., p.22.

Table 6. Labor force participation rates of 16-19 year olds, U.S. and Sweden, selected years

	1965	1970	1975	1978
U.S.	45	50	54	58
Sweden	60	52	58	55

Source: (U.S.) Employment and Earnings.
(Sweden) Arbetskraftsundersökningen.

Table 7. Percent of given age groups enrolled in school in Sweden 1966, 1969 and 1975

Age	1966	1969	1975
16	59.8	70.5	74.6
17	45.1	55.2	62.6
18	35.4	42.3	40.0
19	21.6	21.4	20.7
20	12.0	12.9	15.6
21	7.6	9.3	13.9
22	5.9	7.2	11.5
23	4.3	5.3	9.6
24	3.1	3.8	8.4

Source: Arbetsmarknadspolitik i förändring, Ministry of Labor, Stockholm, 1978, p.139 (SOU 1975:60).

has unemployment.[1] An interpretation consistent with the finding in Section 3 of a strong association between school attendance of teenagers and higher unemployment rates is that schooling of older youth leads to intermittant labor force attachment and greater resulting transition unemployment. This may be further exacerbated by rapid increases in participation by married women, since many of them are competing in the same labor markets. Changes in the labor market activity of teenagers may exhibit mutual spillovers with changes in labor market activity of married women, particularly those seeking part time employment.[2]

In the U.S. teenagers and, to a lesser extent, women have LFPR's which vary directly with the cycle. As the job market deteriorates teenagers have become discouraged workers and their withdrawal from the labor market has kept unemployment rates from rising still further in the 1971-72 and 1975-76 recessions. As the economy begins its recovery teenagers, perceiving the improved job prospects, reenter the labor market. Because reentrance to the job market often involves a spell of unemployment these encouraged workers prolong the high rates of unemployment even when the economic recovery has been underway for some time.

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[1] Labour Market in Transition, p.22, 16-19 year old males had unemployment rates as follows: 1965, 2.0; 1970, 3.1; 1974, 5.0; and 16-19 year old females had unemployent rates as follows: 1965, 3.9; 1970, 5.3; 1974, 7.3.

[2] In the U.S. total voluntary part time employment has grown from about 5 million in 1960 to about 13 million in 1979. See Employment and Earnings.

Patterns of labor force activity of Swedish teenagers during the recent recession parallel those in the U.S. The LFPR of teenagers fell from 58 in 1975 and 1976 to 57 in 1977 and 55 in 1978. As of January 1979 the rate was 50.7 in comparison to 52.6 in January of 1978.[1] The question which arises is whether as the economy recovers the teenage LFPR will rise and slow the rate of decline in overall unemployment because of transitional spells of job search. Also, those in relief work programs may switch to the regular job market as part of the economic recovery and this too would slow the decline in official unemployment rates. If this pattern of reentering into the regular labor market is observed in Sweden as well as in the U.S. it would provide some market level interpretation of the strong association between school retention rates for teenagers and levels of unemployment in the different countries.

In the U.S. the group with the highest unemployment rate is black teenagers. The incidence is high but duration of spells is short for all teenagers and the major reason for unemployment is either entrance or reentrance to the labor market (about two-thirds of unemployed 16-19 year olds were new entrants or reentrants in early 1973). Very young blacks have had very high unemployment rates even prior to the 1975-76 recession when they reached the 40 percent level.

In early 1973 the rates were about 32 percent for teenage black females and about 28 percent for teenage black males. Although data on duration by

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[1] <u>Arbetskraftsundersökningen</u>, selected issues.

race for 16-19 year olds are not available for 1973 Employment and Earnings, my guess is that the length of spells for young blacks are not very different from the length of spells for young whites.[1]

Young persons are characterized by relatively high quit rates which one would expect to observe as they try alternative employment to learn where they have the most talent and interest. Because the category of "quit" is rather broad it is difficult to tell whether the quit was in some sense anticipated and part of a career plan or was motivated by unanticipated dissatisfaction with a given job and what job characteristic or characteristics were the source of the dissatisfaction (wage, learning opportunities, location of job, work hours or schedule).

The rate of unemployment for black teenagers has been in the 35-40 percent range during 1977 and 1978. For all blacks the rate of unemployment is about twice that for whites during non-recessionary times. For example, in the month of April 1978 the unemployment rate was 5.2 percent for whites and 11.8 percent for blacks. For black teenagers the unemployment rate was 35.3 percent. The reasons for the large black-white differences are numerous and the major reasons include direct dis-

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[1] See Barrett, Nancy and Morgenstern, Richard, "Why do Blacks and Women have High Unemployment Rates?", Journal of Human Resources, Fall, 1974, p.457. White males 16-19 have durations of 7.8 weeks and black males 16-19 have durations of 8.7 weeks. In contrast, spells of unemployment as estimated by "turnover" are 54 for white males 16-19 and 95 for black males ages 16-19.

crimination, poorer schooling and geographic segregation. Schooling attainment of blacks has increased, and there has been a federal policy supporting equal job opportunity since the 1960's. If these policies have been even modestly effective (and there is an extensive literature arguing this point) the importance of geographic segregation for the rising unemployment rates for young blacks may be greater.

The overall data show that jobs have been moving to suburban locations at a much faster rate than people[1] and there are several important reasons why such a pattern would have a substantial impact on job market prospects of central city blacks. If workers are defined by skill and location, then a standard economic model of production will usually imply lowered wages for a group which is an increased proportion of the labor force. In principle, a location-specific oversupply can be offset by reduced commuting costs (both time and money). It does imply that if blacks face housing discrimination outside the central city, and have poor transportation and are disproportionately in certain age and education categories there will be low wages. These low wages relative to income support programs or living at home will act to increase unemployment and non-participation.

The high black unemployment rates raise the overall average unemployment rate in the United

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[1] See Harrison, Bennett, Urban Economic Development, Urban Institute, 1974. His analysis (based on work by Charlotte Freeman) shows that while there has been growth in employment in a number of SMSA's, there has been a decline in jobs held by central city residents. See p.48-49.

States quite substantially. In October 1978 the unemployment rate for whites 16 years of age or older was 4.7 percent but was 10.6 percent for blacks, resulting in an overall rate of unemployment of 5.4 percent, not seasonally adjusted. Policy and structural changes which would lower the black unemployment rate in the United States would be important in lowering the overall unemployment rate, but even if black unemployment rates were reduced to those for whites the U.S. would still have far higher official rates of unemployment than Sweden.

One interesting aspect of the labor market for youth in Sweden and the U.S. relates to the size of the youth cohorts. As in the U.S. unemployment rates of teenagers have risen in Sweden, despite a declining relative cohort - in contrast to the U.S. where the teenage cohort has been rising, at least until recently for whites and continues to rise for blacks. This leads to the inference that relative cohort size may not be as important an explanation of the high teenage unemployment rates in the U.S. In the context of a search model[1] one can interpret a spell of search unemployment (which is more likely to apply to teenagers than to older workers) as increased in probability and duration by the expected market wage relative to some value of non-market activity - either the value of leisure or income support or the value (in terms of future income or current consumption) of time in school. If this is so and if labor hours of women, particularly of younger women and

[1] See, for example McCall, John J., "Economics of Information and Job Search", Quarterly Journal of Economics, February 1970, pp.113-126.

those working part time, are viewed as relatively good substitutes for teenage labor hours we may conclude that part of the rise in unemployment among teenagers in both U.S. and Sweden is attributable to increased labor force participation of women. This supply increase may depress the wage rates of teenagers relative to the value of non-market activities, including continued schooling.

A major difference between unemployment in Sweden and the United States lies in the duration versus incidence of spell. The extensive literature on the subject supports the conclusion that Sweden has moderately longer duration spells of unemployment whereas the U.S. has a far higher incidence or flow of individuals into unemploment.[1] Both countries have experienced a rise in average duration as measured by the household surveys. In the U.S. average, non-recessionary duration has risen from about 8 weeks in 1969 to 10 weeks in 1974 and about 11 weeks in 1978.[2] In Sweden average, non-recessionary duration has risen for all age groups rather than being a simple compositional shift in the labor force toward older workers, who are characterized by longer durations of unemployment.

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[1] Flanagan, Robert, "The U.S. Phillips Curve and International Unemployment Rate Differentials", American Economic Review, March 1973, pp.114-31; Barrett, Nancy, "The U.S. Phillips Curve and International Unemployment Rate Differentials; Comment", American Economic Review, March 1975, pp.213-21; Axelsson, Roger, Holmlund, Bertil and Löfgren, Karl-Gustaf, "On the Length of Spells of Unemployment in Sweden: Comment", American Economic Review, March 1977, pp.218-21; and Björklund and Holmlund, "The Structure and Dynamics of Unemployment - Sweden and the United States", this volume.

[2] Employment and Earnings, Table A37.

With higher duration spells for workers of a given age, the rise in duration would thus not appear to be primarily the consequence of reducing the growth of short duration spells through the growth of youth enrollment in public employment programs. Are there some other policy or labor market variables, which could provide a sensible explanation for rising duration in both countries? It has been hypothesized that more generous unemployment benefits can increase the duration of unemployment[1] as well as the incidence of unemployment.[2] Even if the effects of unemployment insurance (UI) are substantial and growing in both countries,[3] the effect on unemployment duration is therefore ambiguous because it would depend on whether duration or incidence was more strongly affected.

The source of the incentive effects of UI on unemployment rates is in the imperfect experience rating of claims against firms and in the tax status of benefits. In Sweden UI benefits are not experience rated at the firm level but since 1974 are taxed as personal income[4], while in the U.S. benefits are not taxed as personal income, but the firms pay taxes which are, for the most part,

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[1] Ehrenberg, Ronald, and Oaxaca, Ronald", Unemployment Insurance, Duration of Unemployment and Subsequent Wage Gain", American Economic Review, December 1976, pp.754-79.

[2] Feldstein, Martin S, "The Effect of Unemployment Insurance on Temporary Layoff Unemployment", American Economic Review, December 1978, pp.834-46.

[3] Table 1 indicates a growth in Swedish unemployment insurance outlays.

[4] Persson-Tanimura, Inga, "On the Costs of Unemployment in Sweden", manuscript, Stockholm, December 1978, p.16.

experience rated. There are some clear propositions about the impact of replacement rates for individuals and the tax rates for firms which could be tested by a comparative analysis of the two unemployment compensation systems.

One might think the greater rise in unemployment duration in Sweden than in the U.S. to be explained by the policy of layoff notification that began in 1974. The data in Table 8 indicate a sharp decline in the layoff rate with recessionary layoff rates falling from .36 per month in 1971 to .15 in 1977. Employers can respond to higher effective layoff costs by putting greater reliance on statistical discrimination[1] (attempts to predict productivity from formal characteristics in the work history of applicants). Trial hiring will be more costly and a decline in trial hiring implies fewer short duration spells of employment and unemployment. If so one offset to the benefits of fewer spells of unemployment for the advance layoff notification policy could be longer duration spells.[2]

On the positive side the advance layoff warning policy has reduced turnover into unemployment through reduced layoffs and does not appear to have induced a sharp upward rise in average duration of spells. The rise in duration of spell set

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[1] Aigner, Dennis I. and Cain, Glen, "Statistical Theories of Discrimination in Labor Markets, "Industrial and Labor Relations Review, January 1977, pp.175-87.

[2] Note, however, that the tests presented in Holmlund's paper (this volume) lend no support to this hypothesis.

in well in advance of 1974 and has continued steadily upwards - apart from recessionary periods which result in duration increases as recorded by the survey measure. How can this be when all the obvious components in Table 8 would seem to imply a sharp rise in duration from 1974 on? That is, layoff, quit, and new hire declines would suggest that spells of unemployment would be associated with only major dislocations likely to involve longer duration spells of unemployment. One interpretation is that on-the-job search of those with

Table 8. Monthly layoff, quit and new hire rates per 100 workers. Swedish mining and manufacturing, 1969-1977

	1969	1970	1971	1972	1973	1974	1975	1976	1977
Layoff	.24	.24	.36	.30	.22	.16	.15	.12	.15
Quit	2.94	2.92	2.09	1.86	1.94	2.20	1.95	1.80	1.51
New Hire	3.78	3.47	2.29	2.46	2.67	2.97	2.19	1.87	1.30
Recession year(1-0)	0	0	1	1	0	0	0	1	1
Persons affected[a] by advance layoff notice	-	-	-	-	-	1560[b]	2974	4268	8350

Source: Layoff, quit, and new hire rates are from SCB, unpublished data.

[a] Source: Swedish Employment Policy 1977-1978, National Labor Market Board, p.53. Quarterly average of workers and salaried staff notified because of planned closures or planned reductions in operations.

[b] Third and fourth quarters only.

advance notification but prior to layoff has increased. This means that total search duration may have increased substantially, but search duration while unemployed may not have increased much at all. On net, the advance layoff notification policy may have reduced the Swedish unemployment rate.[1]

5. CONCLUSION

Both Sweden and the United States would have high unemployment rates were it not for differences in labor market policy. We have seen that the countries are very similar in terms of long term variables that appear to relate to intercountry differences in unemployment rates: a high proportion of the labor force in wage and salary employment, a large proportion of youth in school at age 18, high rates of growth of female labor force participation and similar rates of growth of industrial production. The countries even have fairly similar decade rates of inflation in the 1970's, though long run inflation rates were not found to be important in explaining intercountry differences in unemployment rates.

Structural differences between Sweden and the United States include a higher LFPR for women and a smaller cohort of younger workers in Sweden. Despite this smaller cohort of younger workers there has been a rise in teenage unemployment

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[1] Average of other impacts of the legislation was projected via simulations in Eliasson, Gunnar, "Competition and Market Processes in a Simulation Model of the Swedish Economy", American Economic Review, February 1977, pp.277-81.

rates, and this suggests that a simple "cohort-size-crowding" explanation offered for high teenage unemployment rates in the U.S. may not be valid but that increased postponement of labor market entry because of greater formal schooling may explain much of the rise in teenage unemployment in both countries. Both countries have had a rise in duration of unemployment, but the rise in non-recessionary periods is greater in Sweden. This suggests that added research is needed to explain this because it does not appear to be the direct consequence of the policy of advance notification of layoffs that started in 1974.

Labor force hours of men have declined in both countries, and time use data for the U.S. indicate a decline in hours of married women who are in the job market. In fact, this decline is large enough to more than offset the increased LFPR of married women so that average market hours of all married women, both in and out of the labor market, have declined by 2 percent in the U.S. between 1965 and 1975. In the U.S. the decline in labor market hours is associated with a large rise in TV viewing with TV viewing growing most rapidly among the more educated.

On the Composition of Swedish Labor Market Policy

Jan Johannesson

1. INTRODUCTION

The main objective of this paper is to study the changes that have taken place in the total volume and composition of Swedish labor market policy during the 1960's and 1970's. Furthermore, a brief examination on some effects on unemployment will be given.[1]

Years with high capacity utilization will be used as the basis for a discussion about the volume and composition of labor market policy at the peak of the business cycle. In the same way years with low capacity utilization will be the basis for a discussion about labor market policy during recessions. The objectives of labor market policy during peak years are among other things to raise the level of employment, to reduce frictional-, seasonal- and structural unemployment and to counteract inflationary tendencies in the economy. During recessions labor market policy will have additional tasks of a stabilization character to

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[1] This paper is mainly built on materials from two reports from the Swedish expert group for labor market research (EFA); Björklund-Johannesson-Persson, "Labour market policy and labour market development in Sweden during the 1960's and 1970's", IIM, Berlin, 1979, and "Arbetsmarknadspolitik i förändring", SOU 1978:60, Stockholm 1978.

perform; how large will depend on the size of the changes in the business cycle and on the extent to which other economic policy measures are undertaken. The development of the volume and composition of the labor market policy should reflect these underlying objectives.

The operational definition of labor market policy used in the paper is measures financed out of the budget of the Labor Market Board (AMS).[1] However, there are measures to improve employment (e.g., industrial policy) that are financed outside AMS. Hence the operational definition is in some respects rather arbitrary.

2. THE DEVELOPMENT OF SWEDISH LABOR MARKET POLICY

2.1 Changes in Volume: Total Expenditure and Number of Persons in Labor Market Policy Measures

Total expenditure on labor market policy

The expansion of Swedish labor market policy during the last few decades is illustrated in Figure 1 which shows the growth of total expenditure on labor market policy at current and fixed prices, as a share of total government expenditure and as a share of Gross National Product.

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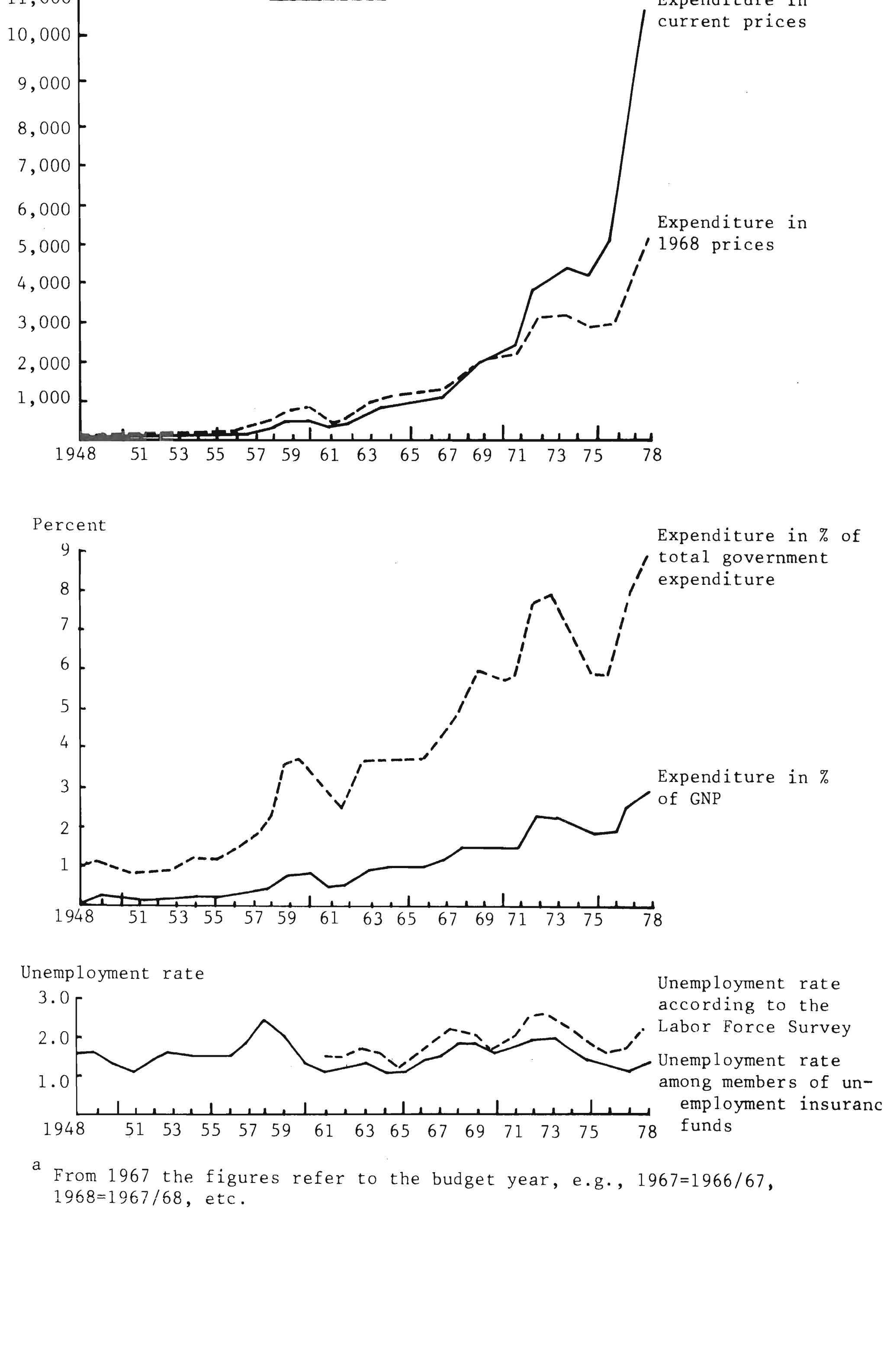
[1] The Appendix shows the size of these measures. Information about expenditure and labor market status have been collected from AMS' annual reports and from the labor force surveys (AKU).

Figure 1. Total expenditure of the National Labor Market Board (incl. of the expenditure for labor market training of the National Board of Education) 1948-1978



[a] From 1967 the figures refer to the budget year, e.g., 1967=1966/67, 1968=1967/68, etc.

Expenditure has grown in almost a stepwise fashion: with marked increases during the recessions and with nearly no decreases during booms.

During the recession of 1957-1959 when there was a marked expansion of labor market policy, total expenditure amounted to just over 500 million Skr. During the recession year 1977/78 total expenditure was instead approximately 11,000 million Skr (at current prices). In the former case labor market policy's share of total government expenditure was approximately 4 percent (1 percent of GNP), in the latter it had risen to about 9 percent (3 percent of GNP).

The expansion of labor market policy from 1948 onwards can be divided into three separate phases. Until about 1960 total expenditure on labor market policy followed a fairly distinct cyclical pattern, increasing its share of government expenditure and GNP during periods of recession and then declining again during periods of more rapid economic activity.

The second phase covers the period until 1969. During this period labor market policy's share of government expenditure rose continuously although the rate of increase was more rapid during recessionary periods. The 1970's represent a third phase where expenditure once again appeared to display a cyclical pattern; e.g., the improvement in economic activity during 1974/75 brought about a fall in expenditure. The most important factors behind this new cyclical pattern during the 1970's have been the variations in traditional manpower training and the introduction of new counter-recessionary measures such as financial assistance for stockpiling and in-plant manpower training to avoid temporary and permanent lay-offs.

Number of persons in labor market policy measures

The same picture of the development of the total volume of labor market policy emerges from an examination of the changes in the average number of persons taking part in different labor market policy measures. Table 1 shows that AMS has developed a considerable capacity for counteracting the effects of recessions and local structural disturbances on employment. The average number of persons in labor market policy measures was 0.9 percent of the labor force in the peak year 1965. In the peak years 1970 and 1974 it had risen to 1.8 and 2.5 percent, respectively. The average number of persons in different measures grew almost continuously during the whole period up to the 1970's. During the 1970's on the other hand there has been a marked cyclical pattern. After a decrease in 1974 and 1975 the number of persons involved has again approached and surpassed the previous peak level. (For more information about the average number of persons in different measures see Appendix B, Table B1).

2.2 Changes in the Composition of Labor Market Policy

The development of total expenditure on labor market policy conceals underlying changes in the "mix" of different policy measures. Some idea of

Table 1. The average number of persons participating in different labor market policy measures (thousands), 1963-1978

Measures	1963	1965	1968	1970	1972	1973	1974	1975	1976	1977	1978
Temporary jobs (relief works)	10.5	9.8	20.3	14.6	32.4	33.3	23.4	16.6	26.3	29.1	45.7
Measures to create employment for hard-to-place labor[a]	4.7	7.4	13.0	20.4	25.0	30.2	35.3	38.6	41.1	42.7	44.2
Rehabilitation and other measures on behalf of hard-to-place labor[b]	1.2	1.4	1.8	2.1	2.6	3.0	3.2	3.3	3.1	3.4	3.5
Training (excl. of in-plant training)	11.3	12.5	25.8	28.8	36.2	34.9	30.6	26.9	28.2	41.7	46.9
In-plant training (incl. of in-plant training to avoid lay-offs)[c]	1.2	3.0	3.8	4.9	6.4	10.4	9.6	8.0	8.0	21.5	13.3
Sum	28.9	34.1	64.7	70.8	102.6	111.8	102.1	93.4	106.7	138.4	153.6
In percent of LF	0.8	0.9	1.7	1.8	2.6	2.8	2.5	2.3	2.6	3.3	3.6
Unemployment rate	1.7	1.2	2.2	1.5	2.7	2.5	2.0	1.6	1.6	1.7	2.2

[a] Sheltered employment, work at home, semi-sheltered employment, archive work and assistance to musicians (see Appendix A).

[b] Courses for adjustment and training (in-plant training for the handicapped and elderly is included under the heading of in-plant training).

[c] The average number of persons receiving in-plant training to avoid lay-offs during 1974, 1975, 1976, 1977 and 1978 has been estimated to 100, 200, 1,600, 16,000 and 11,500 respectively (see Appendix B).

these changes can be gained from an examination of the trend of expenditure for separate policy programs as defined by the Labor Market Board (see below and Appendix A, Table A1).

The Contents of Different Labor Market Programs at the National Labor Market Board

Program	
1) Labor Market Information	
2) Vocational and Geographic Mobility	Labou market training Geographic mobility grants Purchase of owner-occupied houses
3) Measures for Creating Employment	Temporary jobs (Relief work) Orders to industry Detail planning grants Investment funds administration Stockpiling support Wage subsidies to different industries
4) Regional Development Assistance	Regional development grants Training subsidies Employment premiums Municipal employment measures Localization loans
5) Rehabilitation and Other Measures on behalf of Hard-to-place Labor	Assessment of capacity and work training Training of hard-to-place labor
6) Measures to Create Employment for Hard-to-place Labor	Occupational aids for the handicapped Industrial assistance Archive work Semi-sheltered employment Sheltered employment
7) Measures for Refugees	Transfer of refugees Care of refugees

8) Cash Assistance to Unemployed Persons	Grants to unemployment insurance funds Adjustment grants Cash labor market assistance
9) Total Defence Activities	Defence and preparedness planning Non combatant conscripts
10) Acquisition and Administration of Equipment	Accomodation facilities Machinery, vehicles

In terms of both fixed and current prices, there has been a substantial expansion of almost all types of policy measures between comparable periods of economic activity during the period under analysis (1957-1978). However the rate of expansion has differed between different programs and between different time periods.

An attempt has been made to categorize the various measures of labor market policy according to how they affect and modify the functioning of the labor market in order to reach the economic-political goals of stabilization, allocation, growth and income distribution. The labor market policy measures are oriented to affect the labor supply, the labor demand or the matching between supply and demand. The supply-oriented measures (programs) - influencing geographic and occupational mobility as well as providing rehabilitation for hard-to-place labor - can increase the possibilities of becoming employed. These measures will decrease structural unemployment by supporting the development generated in the market. The demand-oriented measures (programs) for temporary jobs, regional policy measures, employment creation for the handicapped, in-plant manpower training and wage subsidies will affect cyclical-, seasonal- and structural unemployment by maintaining or increasing labor demand and thus modify the development generated in the market. The measures intended to improve the matching process between supply and demand (the program for labor market information) will above all affect frictional unemployment.

By categorizing different measures in this way[1] it can be seen that a number of shifts have occurred in the general orientation of labor market policy. Table 2 shows that from the late 1950's up to 1970/71 there was a continuous increase in the share of expenditure going to supply-affecting measures at the expense of demand-affecting measures.

During the recession of 1971-1973 the share of supply-affecting measures decreased and was kept at an almost constant level until the boom in 1974/75. The recession that started in the autumn of 1975 resulted in a strong increase in the share of supply-affecting measures, and during the recession they reached almost the same level as in the beginning of the 1970's.

The demand-affecting measures show, except cyclical fluctuations, a downward trend during the 1960's. This trend seems to have been broken in the 1970's. In the latest recession their share of total expenditure has been almost the same as in the recession of 1967/68.

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[1] Every such categorization has to be rather arbitrary but in this case the tendencies in the development are (as will be shown) so clear that the arbitrariness is of minor importance.

Table 2. Total expenditure on labor market policy divided into different kinds of measures, 1957-1977/78. Percentage and in million Skr.

Measures	1957	1959	1960	1963	1965	1967/68	1968/69
To affect the supply of labor	3.0	5.0	6.6	21.7	20.5	26.3	27.8
To affect the demand of labor	70.8	86.4	85.3	68.8	70.6	65.4	64.5
To improve matching	26.2	8.2[a]	8.1	9.5	8.9	8.3	7.7
Sum	100.0[b]	100.0	100.0	100.0	100.0	100.0	100.0
in million Skr	113.4	400.4	463.3	663.9	891.4	1.461.7	1.778.1
Cash assistance	55.5	87.4	74.3	58.1	76.0	172.1	242.4
Others	27.6	16.2	15.5	12.8	26.6	28.0	39.0
Total in million Skr	196.5	504.0	553.1	734.8	994.0	1.661.8	2.059.5
In % of total government expenditure	1.6	3.6	3.7	3.6	3.7	5.0	5.9

[a] The first major expansion on labor market policy measures (relief work and labor market training) occurred in 1958 and 1959. While the expenditure for labor market training rose from 0.5 millon to 10 million Skr and expenditure for relief work from 75 million to 332 million between 1957-1959 expenditure for labor market information was kept almost constant at about 30 million Skr.

70/71	1971/72	1972/73	1973/74	1974/75	1975/76	1976/77	1977/78
31.6	24.6	24.0	25.5	25.8	28.0	28.2	29.6
59.6	68.8	69.6	66.7	64.8	62.2	64.5	63.6
8.7	6.6	6.4	7.8	9.4	9.8	7.3	6.8
100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0
112.9	3.382.3	3.802.3	3.845.7	3.831.1	4.736.3	7.316.7	9.824.1
257.9	462.6	435.2	507.0	449.5	347.4	462.8	794.1
58.1	52.9	45.1	44.7	63.6	72.6	148.0	128.6
428.9	3.897.7	4.282.6	4.397.4	4.344.2	5.156.3	7.927.5	10.746.8
5.8	7.7	7.9	6.9	5.9	5.9	7.7	9.0

[b] This sum comprises total expenditure for measures usually referred to as active labor market policy, i.e., measures to create employment and training possibilities and to facilitate the matching process. This corresponds to the first six labor market programs according to the list on page 73.

Measures intended to improve the matching process

Expenditures on policy measures that operate on the matching process have increased during the 1960's and 1970's although generally at a slower rate than policy measures intended to influence the demand and supply of labor. However, the 1970's have been a period when the Employment Service has been rapidly expanded and modernized. For example, ADB-produced vacancy lists covering the vacancies at all Employment offices in the country have been introduced.

The increase of the labor market policy has given the Employment Service a more central role in the whole economy since all labor market policy measures are administrated via the Employment Service offices. In addition the introduction of compulsory notification of vacancies has strengthened this tendency.

Supply-oriented measures

During the period from the late 1950's to the early 1970's supply-oriented measures were the most rapidly expanding of the three above mentioned policy categories. During this period demand-oriented measures (which have always taken the largest share of total expenditure on labor market policy) were thus supplemented by measures that sought to remove obstacles to geographic and occupational mobility. Hence the 1960's can be viewed as a period when emphasis was given to both geographic and occupational mobility; the period was, e.g., a developmental phase for manpower training. During the 1970's, the level of expenditure on manpower training has varied with the level of economic activity. It was expanded during the recession of 1971-1973 and in 1976-1978 while during the intervening period it remained at the level reached during the 1960's. The number of persons receiving geographic mobility grants was expanding during the 1960's, from approximately 8,000 in 1961 to 24,000 in 1970. However, during the 1970's, this figure has fallen to around 20,000. The decrease has been due to a reduction in the number of migrants from the forest counties in northern Sweden.

Demand-oriented measures

During the 1970's there has been a marked change in the composition of the demand-oriented measures due to the increased emphasis on firm-oriented[1] measures rather than on measures directed towards particular individuals.[2] (See Table 3.) These changes are reflected in the expansion of policy measures such as stock-piling support, orders to industry, in-plant training to avoid lay-offs, temporary employment assistance for older employees in textile and clothing industries and employees of companies that are of vital importance to a local labor market (the "75-% wage subsidy"). These measures have primarily been directed towards the avoidance of temporary and permanent lay-offs. The tendency to use wage subsidies in the form of training subsidies has been especially strong in the latest recession. However, some

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[1] Measures intended to maintain or increase employment at company level without any instructions from the labor market authorities as to which particular individuals are to receive assistance.

[2] Measures where the labor market authorities have specified the individuals that are to take part in the measures.

Table 3. Expenditure on demand-oriented measures, 1967/68-1977/78

Measures	1967/68	1970/71	1972/73	1974/75	1975/76	1976/77	1977/78
Directed to firms, %	17	25	17	27	29	37	34
Directed to individuals, %	83	75	83	72	71	63	66
Sum, %	100	100	100	100	100	100	100
Million Skr	957.1	1,260.1	2,645.2	2,479.0	2,946.2	4,725.3	6,249.6

changes in the composition of this policy could be found during the second half of 1978. The use of the subsidy for in-plant training to prevent lay-offs has been reduced markedly.

In addition a new type of hiring subsidy has been introduced in 1978. This subsidy can be said to be directed towards expanding rather than stagnating firms. Since the third quarter of 1978 there is a subsidy to companies if they make new recruitments during the second half of 1978 or the first half of 1979; at most the companies can get 12,000 Skr per person permanently added to the workforce.

There has also been a change in the type of measures directed to individuals during the 1970's. In 1972 55 percent of the persons engaged in temporary jobs[1] in the public sector were employed in building and construction work or in forestry work while in 1977 the corresponding figure was 30 percent. What has happened is that temporary jobs

[1] The development of persons in temporary jobs is shown in Appendix B, Table B2

have become increasingly common in the public service sector and within private industry (private sector temporary jobs). Wage subsidies are available for private companies and public authorities which arrange temporary jobs for persons younger than 20 years who are registered as unemployed. One of the reasons for this change has been the altered composition of the unemployed. The proportion of women and young persons involved in temporary jobs has thus increased.

Temporary job projects during the 1970's have also become more oriented towards labor intensive projects, i.e., projects where the wages for those directly involved represent the principal cost. In addition new types of in-plant training directed to particular individuals have been introduced. During 1972-1974 there were training subsidies to firms which, in addition to normal recruitment, employed and trained young people and women. Since 1974 there are wage subsidies to firms which employ women in traditionally male professions and vice versa.

Among the demand-oriented measures there has also, during the entire period, been a rapid expansion of employment-creating measures for the handicapped. This increase applies to sheltered and semi-sheltered employment as well as to archive work. The average number of persons covered by employment-creating measures for the hard-to-place has thus increased from approximately 4,000 in the early 1960's to about 40,000 or about 1 percent of the labor force in 1978.

Thus it seems as if major changes have occurred in the orientation of the demand-oriented measures during the 1970's. They have to a higher degree than before become firm-oriented to make it possible to protect employment.

3. A SUMMING UP OF THE DEVELOPMENT OF THE LABOR MARKET POLICY AND ITS OUTCOME

The composition of labor market policy has shown continuous change and adaptation. During the 1960's there was a development phase for supply-oriented measures such as labor market training and geographic mobility grants. Special emphasis was put on the capability of labor market policy to ensure employment by increasing mobility in the labor market. During the 1970's there was an adjustment in the composition of labor market policy towards an expansion and change of the demand-oriented measures. These measures have since the middle of the 1970's to a higher degree than before become firm-oriented and oriented towards preventing unemployment by avoiding lay-offs in the private sector.

The period after 1974 is interesting in several respects. In spite of the recession open unemployment was kept at roughly the level normally associated with boom periods in the Swedish economy (see Table 1). This was achieved by a combination of general demand policies, labor market policies, industrial policies[1] and job security provisions.

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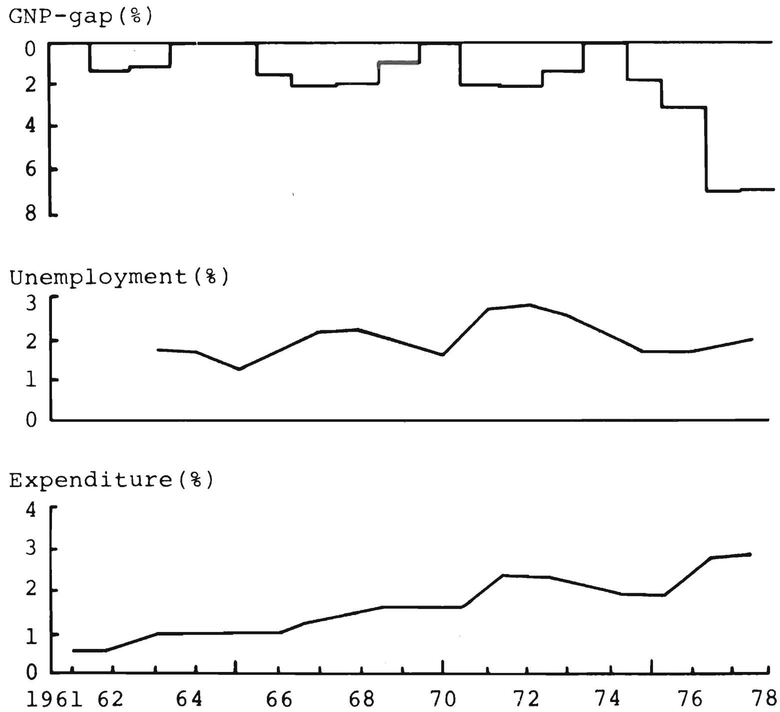
[1] The labor market policy has since 1976 to a great extent been complemented with selective measures in form of industrial policy for the maintenance of employment which are not financed via the budget of AMS. There has been extensive assistance to above all steel-, shipbuilding-, textile- and forest industries. The extent of these measures in terms of the number of people directly affected is not easy to fix.

But one of the most notable changes was that the increase in labor market policies during these years consisted to a large part of measures aimed at preserving already existing jobs. The low unemployment rate can be mainly explained by the fact that the number of unemployed due to lay-offs and close-downs have been kept down during 1975-1978; during this recession they amounted to an average of about 15,000 as compared to the recession 1971-1973 when they amounted to about 33,000. It has not been possible to protect new entrants and re-entrants into the labor market to the same extent. The number of unemployed entrants and re-entrants into the labor market shows only a minor decrease during 1975-1978 (when they amounted to an average of about 26,000) as compared to 1971-1973 (when they amounted to 30,000).

The decrease in the total number of unemployed persons (from an average of 103,000 during 1971-1973 to about 75,000 during 1975-1978) means that <u>the share</u> of entrants and re-entrants (mainly younger age groups) in the total number of unemployed has risen. Thus the labor market policy measures have been insufficient to prevent a marked increase in youth (16-24 years) unemployment during 1975-1978, although labor market policy measures have to a greater extent been directed toward young people. But the expanding share of young people in different labor market policy measures may to a certain degree explain the shorter average duration of unemployment for the young compared with other groups.

One motive for labor market policy is to reduce output losses associated with cyclical unemployment. Figure 2 displays estimated GNP-gap. We can notice that there have been four recessions during the 1960's and 1970's, the one with the largest GNP-gap being the one after 1974. During the other three recessions the GNP-gap has been fairly small.

Figure 2. The GNP-gap, the rate of unemployment and the expenditure on labor market policy for the Swedish economy during the period 1961-1978



Note: The figure shows the GNP-gap as a percentage of potential GNP, estimated by a peak-to-peak method.

Source: Persson-Tanimura, Inga, On the Cost of Unemployment, Berlin 1979.

Furthermore, the cyclical pattern of Swedish unemployment seems to have changed. From the early 1960's until 1975 there was a close relationship between changes in capacity utilization (as measured by the GNP-gap) and changes in the level of open unemployment. The GNP-gaps during 1966-1969 and 1971-1973 were, for example, reflected in marked increases in the unemployment rate and the number of unemployed. This, however, was not the case during 1975-1977 when open unemployment remained at low levels despite large GNP-gaps. The estimated loss (GNP-gap) for 1978 in current prices was about 30,000 million Skr, almost 4,000 Skr per capita or about 7.5 percent of GNP.

Appendix A. ON VARIOUS LABOR MARKET POLICY MEASURES AND THE EXPENDITURE FOR VARIOUS LABOR MARKET PROGRAMS AT THE NATIONAL LABOR MARKET BOARD

Labor market training is to be specially designed for those who are, or run the risk of becoming, unemployed or whose position in the labor market is weak in other respects. A large part of labor market training is arranged by the National Board of Education at the request of the National Labor Market Board. This training is undertaken in specially designed courses in about fifty training centers. Labor market training at firms is arranged in certain cases for unemployed persons who are referred to employment/training by the Public Employment Service and in other cases to persons who are already employed at the firm arranging the training.

On several occasions during the 1970's one has used temporary subsidies to firms arranging training in connection with new recruitment of personnel. In 1972-1974 subsidies were granted firms which in addition to their normal recruitment employed and trained women and young people. Since 1974 there is a training subsidy to encourage the employment of women in occupations traditionally held by men or the employment of men in "female" occupations. Training subsidies for existing personnel have been available since 1976 on condition that those in training will be replaced by new recruitments.

The most important change in the system of labor market training has been the use of various kinds of in-plant training in order to avoid temporary or permanent lay-offs during recessions. During the first six months of 1972 a training subsidy to avoid temporary lay-offs was available for firms which could not use stockpiling support. During the oil crisis in 1973 there was a special training subsidy and since 1974 there has been a subsidy to in-plant training to avoid lay-offs.

Relief work (temporary jobs) provides unemployed persons with temporary employment during seasonal and cyclical recessions. Special relief work is arranged in order to provide employment for persons who are elderly, have local ties or a handicap. Relief work can take the form of national, local government or private projects.

Wage subsidies to different industries

In the spring 1977 temporary employment subsidies were given to firms with long-run supernumerary staff if they postponed lay-offs. Thus temporary employment subsidies are given for older employees in the clothing and textile industries and for employees in companies of vital importance to a local labor market ("the 75-% wage subsidy"). These measures have primarily been directed towards the prevention of unemployment and the avoidance of lay-offs. Since the third quarter of 1978 there is a subsidy to companies if they make new recruitments during the second half of 1978 or the first half of 1979.

Sheltered employment, i.e., employment in sheltered workshops can be arranged for persons with occupational handicap who cannot obtain a job in the open market.

Table A1. Expenditure on labor market policy programs (total expenditure of the National Labor Market Board, AMS, inclusive of the expenditure for labor market training of the National Board of Education, SÖ) divided into different programs 1965-1977/78. Million Skr (current prices)

Programs	1965	1966	1966/ 67	1967/ 68	1968/ 69	1969/ 70	1970/ 71	1971/ 72	1972/ 73	1973/ 74	1974/ 75	1975/ 76	1976/ 77	1977/ 78
(1) Labor market information	80.0	95.6	103.9	120.7	136.6	205.4	184.2	221.8	244.7	298.2	359.3	465.8	531.2	670.8
(2) Occupational and geographic mobility	173.4	197.9	269.7	372.0	469.4	559.0	625.3	783.4	871.8	948.7	921.0	1237.5	2029.7	3252.7
thereof Geographic mobility grants	27.3	22.1	22.4	32.0	56.0	63.0	55.2	56.2	55.0	65.1	60.3	64.7	72.6	104.5
In-plant training to avoid lay-offs								0.5[a]	-	0.9	1.1	19.0	90.4	482.8
Other in-plant training									7.8[b]	27.8	13.5	1.8	5.8	9.7
(3) Employment creation measures	548.0	484.6	558.3	693.0	833.5	706.4	657.7	1568.1	1929.0	1560.9	1108.7	1486.9	2955.1	3839.4
thereof Temporary jobs (relief work)	538.9	469.8	542.2	670.0	792.1	689.3	624.0	1482.5	1740.7	1488.6	1067.0	1193.4	1796.9	3141.9
Orders to industry	5.5	1.2	3.2	4.1	1.3	2.0	5.2	42.7	49.4	1.9	0.5	43.9	174.3	299.3
Stock-piling support									113.0	30.2	1.1	236.6	954.5	295.3
(4) Regional development assistance	22.5	113.6	129.8	162.3	165.6	250.1	337.9	436.4	339.7	527.2	732.6	615.3	624.1	682.5
thereof Training subsidies							32.5	41.0	58.5	88.1	74.3	52.6	42.3	28.9
Employment subsidies							4.8	10.3	13.7	16.3	29.3	31.3	28.0	18.0

Programs	1965	1966	1966/67	1967/68	1968/69	1969/70	1970/71	1971/72	1972/73	1973/74	1974/75	1975/76	1976/77	1977/78
(5) Rehabilitation and other measures on behalf of hard-to-place labor	9.0	10.7	9.6	13.1	25.8	35.0	43.3	49.7	48.2	60.8	87.4	108.4	141.9	157.4
thereof In-plant training for hard-to-place labor											1.0[d]	1.7	2.2	2.0
Training for sheltered work													13.0[e]	11.9
(6) Measures to create employment for hard-to-place labor	58.5	68.5	80.1	100.6	147.2	184.0	264.5	322.8	368.9	449.9	622.1	822.4	1034.7	1221.3
(7) Measures for refugees	6.0	8.4	6.3	7.1	17.9	34.0	19.3	16.8	15.2	23.6	31.8	42.5	102.0	79.8
(8) Cash assistance to the unemployed	76.0	99.1	108.9	172.1	242.4	238.5	257.9	462.6	435.2	507.0	449.5	347.4	462.8	794.1
(9) Others (defence activities and acquisition and administration of equipment)	20.6	18.6	18.5	21.0	21.1	29.9	38.8	36.1	29.9	21.1	31.8	30.1	46.0	48.8
Total expenditure	994.0	1096.0	1285.0	1661.8	2059.5	2242.3	2428.9	3897.7	4282.6	4397.4	4344.2	5156.3	7927.5	10746.8
In % of Government Expenditure	3.7	3.7	4.2	5.0	5.9	5.7	5.8	7.7	7.9	6.9	5.9	5.9	7.7	9.0
In % of GNP	1.0	1.0	1.2	1.5	1.6	1.6	1.6	2.4	2.4	2.1	1.8	1.9	2.7	3.0

[a] Cannot be separated until 1971/72.
[b] Cannot be separated until 1972/73.
[c] Cannot be separated until 1970/71.
[d] Cannot be separated until 1974/75.
[e] Cannot be separated until 1976/77.

Archive work is arranged for unemployed persons who cannot obtain work in the open market and whose age, state of health or other special personal circumstances make them unsuitable for the different forms of relief work or sheltered employment. The work places can be arranged as state, local government or private archive work.

Semi-sheltered employment can be provided for handicapped and elderly who cannot obtain work in the open market under ordinary conditions. Such activities carry a state grant to firms and to local government authorities or institutions as well as to national public enterprises.

Appendix B. NUMBER OF PERSONS IN VARIOUS MEASURES

Table 1 (p.72) contains estimated figures of the number of persons receiving in-plant training to avoid lay-offs. For this kind of training the firms can get a training subsidy to maintain employment (without any instructions from the labor market authorities as to which employed individuals are to receive assistance). This training subsidy was introduced in July 1, 1974 and has been reformed several times. During 1978 the training subsidy consisted of 25 Skr/hour for up to 480 hours' occupational training or for up to 160 hours' general education. For training thereabove the firms could get 15 Skr per person and hour.

The National Labor Market Board (AMS) publishes figures about this form of labor market training in the middle of the month. Because this training is usually undertaken on a part-time basis it is difficult to interpret the number of persons in this kind of measure. The average number of persons in in-plant training according to AMS' statistics during the period February 1-December 31, 1977 was about 44,700. During this period the authorities had granted training of this kind for 30.4 million hours, that is full-time (40 hours per week) training for 190,000 persons during one month.

For the period February-December this would be equivalent to 17,300 persons in full-time in-plant training to avoid lay-offs. This figure corresponds to about 40 percent of the figure in the official AMS-statistics. If the official figure for this kind of training in January 1977 is transformed in the same way the average number of persons 1977 in fulltime in-plant training to avoid lay-offs should be 16,000. This figure can be said to correspond to the other figures in Table 1. This means that on average 3.3 percent of the number of persons in the labor force took part in labor market policy measures in 1977. The same estimation method applied to the years 1974, 1975, 1976 and 1978 gives 100, 200, 1,600 and 11,500 persons in full-time in-plant training to avoid lay-offs.

Table B1. Number of persons in different labor market policy programs, 1959-1978

Annual averages, thousands.

Programs	1959	1960	1961	1962	1963	1964	1965	1966	1967	1968
In-plant training	0.4	0.4	0.6	0.8	1.2	1.8	3.0	3.4	3.2	3.8
thereof in-plant training to avoid lay-offs	-	-	-	-	-	-	-	-	-	-
Other labor market training	4.2	6.2	7.5	9.3	11.3	11.9	12.5	15.4	20.3	25.8
% of women	-	-	-	-	30	34	36	35	33	34
Sum of labor market training	4.6	6.6	8.2	10.1	12.5	13.7	15.9	18.8	23.5	29.6
Temporary jobs (relief work)		6.1	3.6	5.7	10.5	9.6	9.8	9.2	13.7	20.3
Sheltered work and homework		1.8	1.9	2.2	2.6	3.6	4.7	5.7	6.8	7.8
Semi-sheltered work										0.3
Archive work	1.3	1.5	1.6	1.8	2.1	2.4	2.7	3.1	3.7	4.9
Assessment of capacity and work training	-	-	-	-	1.2	1.3	1.4	1.5	1.7	1.8
Sum of measures on behalf of hard-to-place labor	1.3	3.3	3.5	4.0	5.9	7.3	8.8	10.3	12.2	14.8
Total	5.9	16.0	15.3	19.8	28.9	30.6	34.5	38.3	49.4	64.7
in % of LF			0.4	0.5	0.7	0.8	0.9	1.0	1.3	1.6
In labor force			3.699	3.746	3.720	3.718	3.742	3.792	3.775	3.822
Unemployed			55	58	65	60	45	61	82	86
Rate of unemployment			1.5	1.5	1.7	1.6	1.2	1.6	2.2	2.3
Employed			3.644	3.688	3.655	3.658	3.697	3.731	3.693	3.736

Cont.

Programs	1969	1970	1971	1972	1973	1974	1975	1976	1977	1978
In-plant training	4.0	4.9	5.8	6.4	10.5	9.6	8.0	8.0	21.5	13.3
thereof in-plant training to avoid lay-offs	-	-	-	-	-	0.1	0.2	1.6	16.0	11.5
Other labor market training	26.6	28.8	33.2	36.2	34.9	30.5	26.9	28.2	41.7	46.9
% of women	48	48	46	44	47	51	54	52	51	52
Sum of labor market training	30.6	33.7	39.0	42.6	45.4	40.1	34.9	36.2	63.2	60.2
Temporary jobs (relief work)	15.6	14.6	19.0	32.4	33.3	23.4	16.6	26.3	29.1	45.7
Sheltered work and home work	9.3	10.8	11.9	12.8	13.8	15.1	16.0	16.4	16.8	17.2
Semi-sheltered work	0.7	1.3	1.7	2.1	4.0	6.3	8.3	9.8	10.8	11.7
Archive work	6.3	8.3	9.6	10.1	12.4	13.9	14.3	15.0	15.1	15.3
Assessment of capacity and work training	1.8	2.1	2.4	2.6	3.0	3.2	3.3	3.1	3.4	3.5
Sum of measures on behalf of hard-to-place labor	18.1	22.5	25.6	27.6	33.2	38.5	41.9	44.3	46.1	47.7
Total	64.3	70.8	83.6	102.6	111.9	102.0	93.4	106.8	138.4	153.6
in % of LF	1.6	1.8	2.1	2.5	2.8	2.5	2.2	2.5	3.3	3.6
In labor force	3.841	3.913	3.961	3.970	3.977	4.042	4.129	4.155	4.174	4.209
Unemployed	74	59	101	107	98	77	67	66	75	94
Rate of unemployment	1.9	1.5	2.5	2.7	2.5	1.9	1.6	1.6	1.8	2.2
Employed	3.767	3.854	3.860	3.863	3.879	3.965	4.062	4.089	4.099	4.115

Table B2. Number of persons in temporary jobs (relief work) 1965-1978. Thousands.

Year	Number of persons in temporary jobs[a]	Percentage women	Percentage of the women that are under 25 years old	Percentage of the total number (men and women) that are under 25 years old	Total number of persons in temporary jobs (incl. of craftsmen and foremen)	
					thousands	% of LF
1965	8.2				9.8	0.3
1966	7.6				9.2	0.2
1967	11.2				13.7	0.4
1968	16.0				20.3	0.5
1969	12.6				15.6	0.4
1970	9.6	4	2	4	14.6	0.4
1971	14.0	5	14	8	19.0	0.5
1972	23.1	9	30	16	32.4	0.8
1973	25.2	16	43	22	33.3	0.8
1974	17.8	19	42	20	23.4	0.6
1975	12.9	18	39	18	16.6	0.4
1976	21.5	31	73	47	26.3	0.6
1977	24.9	35	80	57	29.1	0.7
1978	41.7	39	84	66	41.7	1.1

[a] The number of persons put into relief works by AMS, due to employment problems. This figure does not contain those craftsmen and foremen who have no labour market problems but are employed by the authorities to hold key positions in the relief work.

Determinants and Characteristics of Unemployment in Sweden: The Role of Labor Market Policy

Bertil Holmlund

1. INTRODUCTION[1]

The rate of unemployment has in the late seventies displayed a substantial increase in a number of western countries, Sweden being one striking exception. Despite a pronounced fall in aggregate demand Swedish unemployment stayed at lower levels 1976-1978 than it did during the more normal recession 1971-72. The rate of unemployment reached a decade high in 1972 - 2.7 percent - but was on average only 2.2 percent in 1978.

This paper is focusing on the relationships between labor market policy and labor market behavior in Sweden. The policies taken account of include traditional programs - such as relief works and manpower training - as well as recent policy innovations such as (i) the employment security legislation, (ii) the inventory subsidy scheme and (iii) the wage subsidy program.

The plan of the paper is as follows: First, in Section 2, a brief description of the institutional setting and the policies undertaken are given. Section 3 contains a rough attempt to establish

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[1] I have benefited from Anders Björklund's constructive comments on an earlier version of this paper.

certain microeconomic effects of two important labor market programs. The basic question is how enrollment in relief works or retraining affects the individual's future unemployment probability. Are trainees rewarded by lower unemployment risks compared to non-participants? This analysis, which applies standard econometric methodology, as a by-product gives microinformation on the determinants and characteristics of unemployment.

Section 4, in turn, approaches the policy issues from a different perspective: How is the probability of leaving unemployment affected by the availability of relief works and training programs? This analysis, therefore, investigates how the contracyclical objectives of these programs are translated into practice. The section includes, moreover, analyses of possible unemployment effects of the employment security legislation. Have the rising firing costs - implicit in this legislation - increased the duration of unemployment via more careful screening practices on part of the firms?

In Section 5 of the paper, finally, we embed a structural unemployment relationship in a macrosimulation model of the labor market.[1] This model - consisting of ten stochastic equations estimated on monthly data - describes the manufacturing sector's decisions on production, vacancies, recruitments, layoffs and hours per worker. The behavior of the household sector is taken account of in quit and unemployment equations. The model is

[1] The model is described in some detail in Holmlund (1980).

driven as a selfcontained system by exogenous output demand and provides a suitable framework for dealing with the indirect effects and interactions that occur throughout the labor market if one part of it is directly changed. In particular, we present a policy-simulation where the actual development during the seventies is compared to an alternative case, produced by a "passive" labor market policy. Implications for employment, unemployment, labor turnover and labor productivity are revealed by this simulation.

2. THE INSTITUTIONAL SETTING

Swedish post-war labor market policy has to a large extent been fostering mobility and structural change in the labor market. The well-known Rehn-Meidner policy-proposals for selective policy measures had given priority to job security in the entire labor market rather than job security in a particular firm.[1]

A change of the labor market policy appears to have taken place during the seventies. The policies have now to a greater extent emphasized employment protection rather than employment creation.[2] Various layoff-discouraging measures have been undertaken in order to reduce the flow into unemployment whereas less priority has been given efforts aimed at influencing the duration of unem-

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[1] For an overview of the Swedish discussions on stabilization issues during the 40's and the 50's, see Lindbeck (1974).

[2] See Johannesson, this volume, for more details about the Swedish labor market policy.

ployment. These changes are reflected in the expansion of various policy measures, including economic "carrots" as well as legal "sticks". Three of the most important policies are the inventory subsidy schemes 1972 and 1975-77), the employment security legislation (in operation from July 1974) and the employment subsidy program (most actively undertaken during the recession 1977-78).

The inventory subsidies are - roughly speaking - paid to firms conditional upon unchanged volume of employment; the firms under consideration are those where employment reductions are likely in the absence of the policy. The employment subsidies, in turn, are intended to affect temporary and permanent layoffs; firms with high expected layoff probabilities are subsidized, given that they provide their excess labor with a certain degree of training during a specified period of time. The employment security legislation (LAS),[1] finally, is intended to protect employment for tenured workers. The employers' free right to dismiss is replaced by legislation which demand objective cause for dismissal. "Shortage of work" is considered to be one such objective cause. Firms have, furthermore, to give employed workers up to six months advance notice before anyone can be laid off.

The two subsidy programs are temporary contracyclical measures whereas the LAS-scheme represents a permanent institutional change. The employment subsidy scheme will mainly affect unemployment via longer durations of existing jobs. The employment security legislation has employment pre-

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[1] LAS = Lagen om anställningsskydd.

serving aims and strong layoff-discouraging rules and is therefore likely to reduce the flow into unemployment. But the recruitment behavior of firms will, presumably, also be influenced, since the laws make it difficult to revise past hiring decisions based on erroneous information. The result is likely to be longer vacancy durations (falling recruitment probabilities) as well as increasing duration of unemployment. The inventory subsidy program, finally, is - at least for the period 1975-77 - focusing upon net changes in employment as a requirement for subsidies; hence we expect that hiring decisions as well as layoffs will be affected.

Along with these new policies, a continuous expansion of more traditional measures has taken place. Of special importance here are relief works and labor market training programs. Relief works are intended to counteract seasonal and cyclical unemployment as well as unemployment due to personal characteristics (e.g., among elderly or handicapped workers) or adverse characteristics of certain regional labor markets. Relief works can take the form of national, local government or private projects. Labor market training, in turn, is arranged for persons who either are or run the risk of becoming unemployed or who are difficult to place in the labor market. Relief works and labor market training engaged each on average about 50,000 persons in 1978 (slightly above one percent of the labor force for each type of program).

3. UNEMPLOYMENT, PERSONAL CHARACTERISTICS AND POLICY EXPERIENCES

The human capital model views the pecuniary remuneration of labor services as the basic indicator of returns to accumulated investments in knowledge. In empirical human capital studies, therefore, the typical dependent variables are wage rates or earnings.

It is fairly straightforward to pursue the human capital approach beyond these confinements. Acquired skills will not only affect the rewards per hour worked but may also affect the individual's incentives and capacity to find and keep jobs. Consider a simple job search model of the type

$$P = \Theta \cdot [1-F(w^*)] \tag{1}$$

where

P = the probability of leaving unemployment (the transition probability)

Θ = the probability of finding a job offer

$[1-F(w^*)]$ = the probability of accepting a job offer (w^* is the reservation wage and $F(\cdot)$ is the distribution of wage offers).

The probability of receiving job offers will, for obvious reasons, be related to the seeker's productivity, as perceived by the employers. Information on the seeker's characteristics in terms of age, education, previous labor market experience, etc. will be useful guides for employers' hiring decisions. This holds true irrespective of whether education is regarded as a productivity augmenting activity or merely as a signalling device.

The job-seeker's stock of human capital will, of course, also affect the reservation wage, in general making the net effect on the transition probability indeterminate *a priori*.

The probability of job separations will also be affected by the individual's human endowments. In particular, we expect that acquired firm specific knowledge will reduce the likelihood of unemployment inflow via lower quit and layoff probabilities. It is also reasonable to hypothesize that workers with more years of schooling will face lower layoff risks at given length of tenure.

There are well-known difficulties associated with efforts to evaluate labor market programs, such as manpower training. A basic problem is related to the choice of an adequate comparison group.[1] The standard procedure applies multiple regression analysis in order to control for pre-training differences between program-participants and non-participants. There is, however, likely that the standardizing variables are imperfect controls and that left-out variables, affecting *both* the outcome measures and the program-enrollment probability, will produce biased estimates of the program coefficients. An approach which may reduce such risks utilizes the longitudinal properties of some data sets. If we, e.g., have information on *past*

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[1] ....."the programs are not usually offered to conform the standards of experimental design; there is not a randomly chosen control and experimental group and more usually those who enroll do so on the basis of some only partially known self-section criteria". (Stafford, 1979, p.10). The ideal design is, however, seldom available in practice, requiring the use of non-random comparison groups.

unemployment experiences (or past wage rates), the effects of leaving out variables affecting current unemployment (current wage levels) will be less serious.[1] This kind of approach will be adhered to in this paper.

Our empirical analysis will focus on one outcome measure, the unemployment probability. In particular, we propose to explain the probability of being unemployed at least once during 1973, conditional upon labor force participation. The microeconomic data source is based on the Level of Living Survey (Levnadsnivåundersökningen), conducted by the Institute for Social Research, Stockholm, in 1968 and 1974.[2] The major fraction of the respondents in 1968 were also interviewed in 1974. Our analysis will focus on individuals that were participating in the labor force in 1967 and in 1973. The criteria for being a participant in 1973 (1967) were that the respondent was unemployed or employed at least once during 1973 (1967).

The functional form adopted is the logistic:

$$Pr(U73) = 1/(1+e^{-X\beta}) \qquad (2)$$

where Pr(U73) is the probability of being unemployed and X is a vector of explanatory variables. The latter includes the following background factors:

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[1] For studies along these lines, see papers by Ashenfelter, Kiefer and Goodfellow, included in Bloch (1979).

[2] See Vuksanovic (1979).

EXP = years of previous labor market experience
S = years of schooling
MAR = dummy for marital status (MAR=1 if the respondent was married, zero elsewhere)
WOM = dummy for female respondents
U67 = dummy for previous unemployment experiences (U67=1 for respondents with at least one spell of unemployment during 1967)
EXCH = dummy for respondents with experiences of employment exchange contacts during 1968-72.

The unemployment equation is a straightforward application of a standard human capital earnings function. The individual's unemployment experiences are related to past labor market experience (reflecting on-the-job training) as well as formal education. In addition, controls are included for marital status, sex and past unemployment and job search.[1]

Uncertainty is associated with the proper specification of the policy variables. We have represented them as follows:[2]

(i) DRT = dummy for retraining at least once during 1968-72

(ii) DRT•U67 = interaction between retraining and previous unemployment

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[1] Regressions were also run with age and age squared included. The results were basically the same; the age-coefficients were insignificant, probably due to strong multicollinearity between age and experience.

[2] The retraining variable also includes those engaged in vocational rehabilitation programs.

(iii) DRW = dummy for enrollment in a relief work program at least once during 1968-72

(iv) DRW•U67 = interaction between relief work enrollment and previous unemployment.

A description of sample characteristics is set out in Table 1. We observe, among other things, that 25 percent of the individuals with unemployment experiences in 1973 also were unemployed some time during 1967; the corresponding 1967-figure for all labor force participants in 1973 is 5 percent.

The results of alternative maximum likelihood estimates are set out in Table 2. The basic human capital approach appears to be consistent with the findings. The variables representing education and labor market experience have the expected signs. The education coefficient implies that one additional year of schooling reduces the unemployment rate for representative individuals by about 0.5 percentage points.[1] The highly autoregressive nature of unemployment is noteworthy; individuals with past unemployment experiences are also facing much higher current unemployment risks. The estimates imply, furthermore, a standardized unemployment rate differential between married and non-married persons of about 3-4 percentage points.

Turning next to the policy variables, the results show that the effects of program enrollment depend

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[1] The derivative

$\frac{\partial P}{\partial S} = -0.11 \cdot P(1-P)$ is calculated at the mean of P where P=Pr(U73).

Note that the unemployment concept used above differs from the conventional measure since one year - rather than one week - is the observation period.

Table 1. Summary of sample characteristics
(Means)

Variable	All individuals	Individuals unemployed at least once 1973
U73	0.051	1.00
EXP	22.0	16.8
S	9.6	9.3
MAR	0.776	0.608
WOM	0.381	0.358
EXCH	0.202	0.642
U67	0.052	0.258
DRT6872	0.027	0.083
DRW6872	0.013	0.058
Sample size	2399	123

on whether the individuals had experienced previous unemployment or not (i.e., unemployment in 1967 in these data). Put differently, we may say that the adverse effects of past unemployment on future unemployment tend to be mitigated by participation in labor market programs.

The quantitative effects are displayed in Table 3, showing predicted unemployment probabilities for individuals with different experiences of previous unemployment and labor market training. If persons without previous unemployment and training enrollment are taken as a reference group [row (i)], we observe that previous unemployment is increasing the current unemployment probability by almost 0.10 [row (ii)]. However, if those with previous unemployment experiences are affected by labor market training their unemployment probabilities are substantially reduced [row (iv)]; in fact, the adverse effects of previous unemployment appear to be completely offset by participation in training programs.

Table 2. Unemployment probabilities and labor market programs. Logit estimates (t-ratios in parentheses).

	(1)	(2)	(3)	(4)
INTCPT	0.905 (-1.457)	-1.000 (-1.597)	-0.991 (-1.587)	-1.003 (-1.602)
S	-0.119 (-2.988)	-0.112 (-2.800)	-0.115 (-2.879)	-0.113 (-2.840)
EXP	-0.084 (-2.672)	-0.081 (-2.578)	-0.082 (-2.602)	-0.082 (-2.618)
EXPSQ	0.001 (1.669)	0.001 (1.590)	0.001 (1.561)	0.001 (1.625)
MAR	-0.778 (-3.851)	-0.818 (-3.919)	-0.765 (-3.667)	-0.790 (-3.790)
WOM	-0.151 (-0.712)	-0.143 (-0.669)	-0.122 (-0.568)	-0.141 (-0.661)
EXCH	1.603 (7.438)	1.529 (6.895)	1.552 (7.112)	1.488 (6.637)
U67	1.381 (5.188)	1.608 (5.793)	1.457 (5.280)	1.637 (5.705)
DRT		0.824 (1.988)		
DRT•U67		-2.565 (-2.214)		
DRW			1.261 (2.130)	
DRW•U67			-1.407 (-1.505)	
DRTDRW				0.934 (2.589)
(DRTDRW)•U67				-1.649 (-2.320)
Log likelihood	-397.8	-394.1	-395.8	-394.0

Note: $EXPSQ=EXP^2$.
DRTDRW=1 implies participation in retraining programs, relief work programs or both.

Table 3. Predicted unemployment probabilities from Eq.(2)

	Background characteristics	Predicted value
(i)	U67=0,DRT=0	0.027
(ii)	U67=1,DRT=0	0.122
(iii)	U67=0,DRT=1	0.060
(iv)	U67=1,DRT=1	0.024
	Mean	0.051

Note: All variables, except U67 and DRT, are set to their mean values.

The results spelled out for retraining programs are not equally well established as far as relief works are concerned; the coefficient for the interaction term is not significant at conventional levels. However, when the two program categories are lumped together, the basic conclusions remain intact [cf. Eq. (4)].

These exercises should, needless to say, be taken with some care. There are few persons with policy experiences in the sample and it is also likely that the standardizing variables are imperfect controls for differences between those enrolled in the programs and the non-participants. The estimates are, however, significant in most cases and fairly robust with respect to changes in the specifications.[1]

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[1] The Swedish labor market training programs have been evaluated by Dahlberg (1972) and Dahlström (1974). The studies show income increases for the trainees but no significant effects on wage rates. The favorable employment effects revealed by our estimations are consistent with those earlier findings.

To what extent is previous enrollment in labor market programs associated with adverse signalling effects? The positive coefficients for DRT and DRW are at least consistent with a signalling interpretation. The coefficients for these program-dummies are, however, smaller in magnitude than the estimated coefficients for U67, i.e., previous unemployment.[1] (Cf. row (ii) and row (iii) in Table 3). We are inclined to interpret these observations, together with the negative coefficients for the interaction dummies, as indicating that the labor market programs are able to improve the participants' future labor market performance.[2]

4. UNEMPLOYMENT FLOWS AND UNEMPLOYMENT POLICIES

4.1 Unemployment Outflow

The theoretic framework already introduced is a useful point of departure for analyses of the duration of unemployment. The worker's probability of leaving unemployment and finding employment equals the product of two terms, the acceptance probability and the job-offer probability. The job-offer probability, first, depends on (i) the likelihood of encountering a job opening and (ii) the probability of obtaining a wage offer from a located vacancy. The first of these probabilities is related to the number of vacancies available and to the worker's search intensity. An unem-

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1 Unfortunately, the data include no satisfactory information on unemployment experiences during 1968-72.

2 Of course, such a conclusion does not necessarily imply that the programs will pass a cost-benefit test.

ployed job searcher has, however, some opportunity to choose between "regular vacancies", and "vacancies" created by the labor market board, i.e., relief works and labor market training. Although no direct measures of these kinds of vacancies are available, some proxies will be tried. The question of interest here is: How have two important elements of the Swedish labor market policy - relief works and retraining - affected the duration of unemployment?

It should be noted that these labor market programs may affect unemployment duration via the job-offer probability as well as the labor force exit probability. An unemployed worker leaving unemployment for training is by definition leaving the labor force. Another circumstance to be kept in mind is that individual unemployment in most cases is a precondition for participation in relief works or retraining. The possibility of induced voluntary unemployment inflow should be recognized; the training programs may be attractive educational alternatives for individuals who not initially are unemployed.

The likelihood of receiving offers from located vacancies will presumably be affected by the searchers' productivities, as perceived by the employers, and by actual firing costs. If past hiring decisions based on erroneous information are costly to revise, this will most likely induce careful screening procedures on part of the firms. The Swedish employment security legislation may be interpreted along these lines; the hypothesis, therefore, is that the introduction of LAS has decreased, *ceteris paribus*, the probability of leaving unemployment and increased the duration of unemployment.

The variables used for explaining unemployment outflow are:

(i) VAC_t = the total number of vacancies (unfilled vacancies in the beginning of the month plus the flow of new vacancies during the month).[1]

(ii) DLAS = dummy for the employment security legislation (DLAS=1 from July 1974, zero otherwise).

(iii) "vacancies" associated with relief works and retraining. As proxy-variables we use, first, the number of persons engaged in relief works or training (LMB_t) and, second, the change in this number.[2] Of relevance for an unemployed job searcher is, of course, not the number of relief works or training positions per se but rather the availability of unfilled employment or training opportunities. It is assumed that the latter are closely related to the two variables mentioned above.

Data and Results

The transition probabilities used are calculated by applying a method suggested by Barron (1975).

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[1] Adjusted to changes associated with compulsory registering in the late seventies according to investigations undertaken by the National Labor Market Board.

[2] LMB = Labor Market Board.

This method assumes that the transition probability is constant during the whole spell of unemployment. The idea is to compare the number of people in one week who have been unemployed less than, say, five weeks with the number of people four weeks later who have been unemployed five to eight weeks (or five to twelve weeks). The difference consists in principle of people who have left the pool of unemployed.[1] The transition rates obtained include the job-finding probability as well as the labor force exit probability. The expectation is that more training opportunities will increase the exit probability.

Estimation results for three different model specifications are set out in Table 2. The specifications adopted are (i) a linear equation, (ii) a log-linear equation and (iii) a logit-equation; the latter restricts the predicted transition probability to values in the interval between zero and one.

$$P_t = \alpha_0 + \alpha_1 VAC_t + \alpha_2 DLAS + \alpha_3 LMB_t + \alpha_4 \Delta LMB_t \tag{3}$$

$$\ell n P_t = \beta_0 + \beta_1 \ell n VAC_t + \beta_2 DLAS + \beta_3 \ell n LMB_t + \beta_4 \Delta LMB_t \tag{4}$$

$$\ell n\left(\frac{P_t}{1-P_t}\right) = \gamma_0 + \gamma_1 VAC_t + \gamma_2 DLAS + \gamma_3 LMB_t + \gamma_4 \Delta LMB_t \tag{5}$$

Two noteworthy results are displayed in the table. First, the coefficient for DLAS is insignificant

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[1] For details, se Barron (1975) or Axelsson and Löfgren (1977). Transition probabilities in this paper have been calculated by using the intervals 1-4 weeks and 5-12 weeks according to the monthly Swedish labor force surveys (AKU).

and has, moreover, an unexpected sign. Thus no adverse effects on unemployment duration are revealed, which is contrary to the basic hypothesis. Second, only <u>transitory</u> effects of relief works and training are found; the coefficient for ΔLMB is significant whereas the stock variable LMB appears to play a negligible role.

The results obtained should be taken with some care. The analysis here is, e.g., focusing on the short-term unemployed, whose behavior may differ from groups with longer spells of unemployment.[1] It may be the case that relief works and labor market training are of special importance for the long-term unemployed, whose reservation wages are likely to be lower. A somewhat speculative story about job-search during unemployment may read as follows: The worker initially contemplates only "regular jobs" and refuses less desirable opportunities such as relief works and retraining. During a period of unsuccesful search, the relative attractiveness of the latter alternatives will increase; the worker's reservation wage - or reservation utility - will fall with the duration of unemployment according to well-known results from search theory.[2]

There is another interpretation to be recognized. The availability of relief works may decrease the probability of transitions from unemployment to the not-in-the-labor force state. It has been shown that these exit rates vary procyclically, at

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[1] Cf. Björklund and Holmlund, this volume.

[2] See, e.g., Gronau (1971).

Table 4. Unemployment outflow and labor market policy. Monthly data 1970.2-1979.5

	$\ell n\ P_t$	$\ell n\ P_t$	$\ell n\ P_t$	$\ell n\ P_t$	P_t	$\ell n(P_t/1-P$
INTCPT	-10.03 (-7.887)	-10.32 (-7.928)	-11.01 (-8.373)	-11.40 (-6.211)	-0.002 (-0.073)	-2.856 (-10.89)
$\ell nVAC_t$	0.706 (6.255)	0.730 (6.338)	0.785 (6.785)	0.816 (5.274)		
VAC_t					0.14E-05 (6.390)	0.10E-((5.882)
DLAS		0.042 (1.038)	0.043 (1.096)	0.042 (1.042)	0.85E-03 (0.141)	0.038 (0.827)
LMB_t				0.46E-06 (0.302)	0.85E-07 (0.390)	0.35E-((0.209)
ΔLMB_t			0.73E-05 (2.198)	0.75E-05 (2.206)	0.99E-06 (1.970)	0.68E-((1.781)
DW	1.23	1.25	1.30	1.31	1.44	1.40
R^2	0.58	0.58	0.60	0.60	0.59	0.62

Note: $LMB_t = LMB_{t+1} - LMB_{t-1}$. All regressions include eleven seasonal dummies.

least in the U.S.[1] Relief works may, however, affect exit behavior in a way different from regular vacancies. The reason for this possibility is that workers contemplating relief work employment have to be registered at the employment exchange offices, clearly an indicator of active job search according to the labor force surveys.

The absence of significant LAS-effects on unemployment duration may be the result of aggregation phenomena. The groups most likely to be hurt by higher recruitment standards on part of the firms

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[1] See Björklund and Holmlund, this volume, for results and a discussion of these findings.

are workers with weak or unstable previous employment experiences or other "undesirable" signalling characteristics. Adverse effects on these groups are perhaps not easily detected by using aggregate data.[1]

Another interesting possibility should, however, also be observed. The advance warning rules included in LAS clearly contribute to on-the-job search among workers threatened by layoffs. This effect implies that the expected number of search periods in unemployment may decrease, even if the total number of search periods increase.

4.2 The Rate of Change of Unemployment

The preceding sections have exclusively been focusing on unemployment outflow whereas the inflow component has been completely disregarded. In this section the relationship for the flow into the unemployment pool will be combined with an outflow equation. The aim is to derive an explicit unemployment equation. The point of departure is the approximation

$$U_t = U_{t-1} + IU_t - P_t U_{t-1} \qquad (6)$$

where IU_t is unemployment inflow and P_t the transition probability, i.e., the probability of leaving the pool of unemployed.[2] Rearranging (6) gives

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[1] Disaggregation by age, however, does not change the basic results. See Holmlund (1978).

[2] (6) is an approximation since those who are entering as well as leaving unemployment within the same month are disregarded.

$$UCHR_t = IU_t/U_{t-1} - P_t, \tag{7}$$

where $UCHR = (U_t - U_{t-1})/U_{t-1}$ is the relative rate of change in unemployment.

The next step is to specify explicit relationships for the inflow and for the transition probability, thereafter substituting these functions into (7) in order to obtain an equation suitable for estimation purposes. Consider first the inflow relation. As a pure identity, the flow into unemployment consists of quits (Q), layoffs (Y) and labor force entries (E):

$$IU_t = \mu_1 Q_t + \mu_2 Y_t + \mu_3 E_t, \tag{8}$$

where μ_i (i=1,2,3) is "unemployment propensities" associated with the three flows. Empirical evidence - at least for the U.S. labor market - indicates that μ_1 is substantially below unity; most job changes take place without intervening unemployment spells. We postulate the following relationship for the unemployment inflow:

$$IU_t = \tau_0 + \tau_1 Y_t + \tau_2 DLAS \cdot Y_t + \tau_3 VAC_t \tag{9}$$

The inflow consists of an autonomous component, τ_0, unaffected by the state of the labor market. Its magnitude is also affected by the number of layoffs. The contributions by firms - via layoffs - to unemployment is, however, assumed to have been decreased by the introduction of LAS. The advance warning rules, giving workers up to six months' notice before layoff, will most likely reduce the "unemployment propensity" associated with layoffs. The specification takes account of that effect by the slope dummy DLAS • Y.

The unemployment inflow is, finally, affected by quits and labor force entries via the number of vacancies. The hypothesis is that quits into unemployment as well as labor force entries will increase with a tighter labor market.[1]

Consider, next, the transition probability, P_t. The basic search theoretic framework is retained, explaining P_t by variables associated with the likelihood of finding vacancies, receiving wage offers and accepting offers. Recalling the arguments already given and approximating the theoretically preferable multiplicative model with a linear function, the transition probability equation is written as

$$P_t = \tau_4 + \tau_5 \ (VAC_t/U_{t-1}) \tag{10}$$

Substituting (9) and (10) into (7) yields

$$UCHR_t = -\tau_4 + \tau_0\left(\frac{1}{U_{t-1}}\right) + \tau_1\left(\frac{Y_t}{U_{t-1}}\right) + \tau_2\left(\frac{DLAS.Y}{U_{t-1}}\right) + (\tau_3 - \tau_5)\left(\frac{VAC_t}{U_{t-1}}\right) \tag{11}$$

as estimating equation.

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[1] There is theoretical arguments - as well as U.S. empirical evidence - indicating that quits into unemployment will rise with the number of vacancies. See Barron and McCafferty (1977).

The regression equation set out in (11) has been estimated on monthly data with results as follows:[1]

$$UCHR_t = \underset{}{-0.020} + \underset{(3.279)}{3043.63}(1/U_{t-1}) + \underset{(2.505)}{0.743}(Y_t/U_{t-1}) - \underset{(-2.027)}{0.686}(DLAS \cdot Y/U_{t-1}) - \underset{(-2.843)}{0.037}(VAC_t/U_{t-1}) + SD \qquad (12)$$

Period: 1969.2-1979.5

R^2=0.62 DW=1.93

The reported equation is satisfactory in terms of statistical performance and reasonable estimates. The "autonomous" component of the unemployment inflow equals about 3000 workers each month, or about 0.5 % of the number of employed workers. The cyclical part of the inflow is captured by the three remaining terms. The LAS-reform seems to have contributed to a marked decrease in lay-off-associated unemployment inflow via the advance warning rules; the slope dummy coefficient is negative and significant. The final term, VAC_t/U_{t-1}, has an estimated coefficient with negative sign, suggesting that the unemployment reducing effect (via the transition probability) outweighs the unemployment creating effect (via quits and new entrances into unemployment).

The estimated constant term, averaged over all months, is -0.13. Recalling equations (10) and (11) above, we have the transition probability as

$$P_t = 0.13+0.037(VAC_t/U_{t-1})+\tau_3(VAC_t/U_{t-1}) \qquad (13)$$

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[1] Data refer in this case to the industrial sector (bluecollar workers). The unemployment rate for workers in manufacturing occupations has been applied to the number of employed workers in industry (according to the establishment statistics). SD is short for eleven seasonal dummies.

By setting $\tau_3 = 0$, thereby completely disregarding voluntary search unemployment due to demand fluctuations, a minimum estimate of the transition probability is obtained for different values of VAC_t/U_{t-1}. If the mean value of the latter ratio is inserted, we obtain P_t (min) = 0.2, which is approximately equivalent to a maximum estimate of the average duration of unemployment of about five months. Available extraneous information suggests that the average duration has been about 14 weeks.[1] The difference between five months (21 weeks) and 14 weeks is due to the cyclical sensitivity of voluntary unemployment inflow.

The estimated unemployment equation may be rewritten as a first order difference equation

$$U_t=0.87U_{t-1}+3043+(0.74-0.69DLAS)\cdot Y_t-0.037VAC_t \tag{14}$$

where the coefficient 0.87 comes from the deseasonalized intercept (0.87=1-0.13). A "quasi-reduced" form of this equation is obtained by fixing DLAS and Y_t and focusing on the unemployment-vacancy relationship. It is easily seen that cyclical variations in VAC will produce clockwise loops in the (U,VAC)-space, thus being consistent with one element in Edmund Phelp's theory of the Phillips curve.[2] We would accordingly, at certain phases of the business cycle, expect to observe simultaneous increases in vacancies and unemployment.

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[1] Björklund (1978).

[2] Phelps (1971).

5. LABOR MARKET POLICY IN A SIMULATION MODEL OF THE LABOR MARKET

5.1 Model Description

The previous analyses have been completely partial, ignoring all possible feed-backs among, e.g., unemployment, vacancies, hirings and layoffs. Clearly, such feed-backs are present in the real world. A high level of unemployment is likely to speed up hirings and reduce the stock of unfilled vacancies via shorter vacancy durations. The layoff-behavior among firms may also be affected by the level of unemployment; firms will presumably be more careful in firing workers if they feel that an eventual unexpected sudden upturn will entail recruitment difficulties due to low unemployment.

In order to deal with some of these interactions, a simulation model of the labor market for industrial workers has been developed and estimated. The system includes the unemployment equation estimated in Section 4 as one relationship. The theoretical framework underlying the model has been described in some detail elsewhere[1] and will not be fully presented here. The model is basically demand-oriented, although both employment and unemployment are partially affected by job search and job acceptance decisions taken by the household sector.[2] The quit and unemployment specifications

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[1] Holmlund (1980).

[2] The driving exogenous variables are sales and unfilled orders. Potential productivity is "explained" by time and estimated on a small number of peak observations.

are focusing on job availability effects rather than inflationary surprises, an approach which has empirical underpinning.[1]

As workers are searching for jobs, employers are searching for employees to fill job vacancies. Employers attempted recruitments are distinguished from their actual hirings, the latter being affected by the state of the labor market.

Changes of employment via hirings or layoffs are likely to be affected by significant adjustment costs. An optimal policy would therefore most likely imply employment smoothing with respect to demand fluctuations. The principal short-run mean of adjusting labor input will be variations in hours per worker. We describe the firms' decisions on men and manhours as being recursive in nature; the basic determinant of hours per worker is the difference between the actual and the desired number of workers on hand.

A condensed description of the model is given in an appendix. We will here give a brief account of how the various policies are taken account of. Inventory subsidies, in the first place, are affecting the system via their effects on firms' production decisions. These, in turn, are dictated by expected demand, represented by sales and unfilled orders. The inventory subsidies are introduced as dummies for the subsidy periods. When planned output increases, so will the size of the desired workforce, thereby indirectly affecting vacancies, hirings, layoffs, which - in turn - influence quits and unemployment, etc.

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[1] Parsons (1973), Barron (1975), Axelsson and Löfgren (1977), Björklund and Holmlund (1981).

The employment subsidy program is directly affecting layoffs, and indirectly therefore unemployment, employment, etc. Technically, the policy is represented by a variable NSUB, the fraction of non-subsidized excess labor.

The third policy taken account of, the employment security legislation (LAS), is transmitted via the equations for hirings, layoffs and unemployment changes. As expected, we find that the higher firing costs - implicit in LAS - have had significant layoff-discouraging effects. Another finding is that these layoff difficulties appear to have produced more careful screening procedures among firms; vacancy durations are significantly higher, ceteris paribus, after the introduction of LAS. A ten percent higher level of unemployment will, furthermore, increase hirings by one percent.

Possible LAS-effects on unemployment duration have already been investigated in Section 4 above. Given the observed effects on layoffs and vacancy durations, it is surprising to note the absence of significant adverse effects on unemployment duration. Possible interpretations of this result have already been discussed in Section 4.

5.2 Simulation Experiments

5.2.1 The Predictive Accuracy of the Model

In order to investigate the tracking ability of the model some historical simulations have been performed. The model takes initial values of exogenous and endogenous variables for the months preceding 1970.7 and is allowed to run on its own from there and 102 months ahead, ending in December 1978. The exogenous variables are sales, unfilled orders, the policy parameters, season and time. The historical simulations reveal that the model stays roughly on track over time. Actual and simulated values for employment and unemployment are displayed in Figures 1 and 2 (yearly averages). The model ends the 102-months simulation with the simulated unemployment rate equal to the actual rate for the second halfyear of 1978.

The predictive accuracy of the model is summarized in Table 5, where the mean absolute percentage errors between actual and simulated values are set out. One simulation was run with output as exogenous driving variable. A comparison between the two columns in the table reveals that errors in the production equation are of minor importance for the predictive performance of the model as a whole.[1]

The prediction errors are smallest for the stock of employed workers (0.9 percent). This result is unsurprising since the monthly flows are small relative to the stock. The errors are considerably higher (about 23 percent) for the stock of unfilled vacancies, where the monthly flows are of the same magnitude as the size of the stock. Layoffs turn out to be the most difficult variable to predict with precision, a result which partly may be due to relatively large measurement errors in the variable itself.

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[1] The performance of the model turns out to be slightly better with output endogenous, which may be due to large measurement errors in the output data itself.

Figure 1. Employed workers, actual and simulated numbers

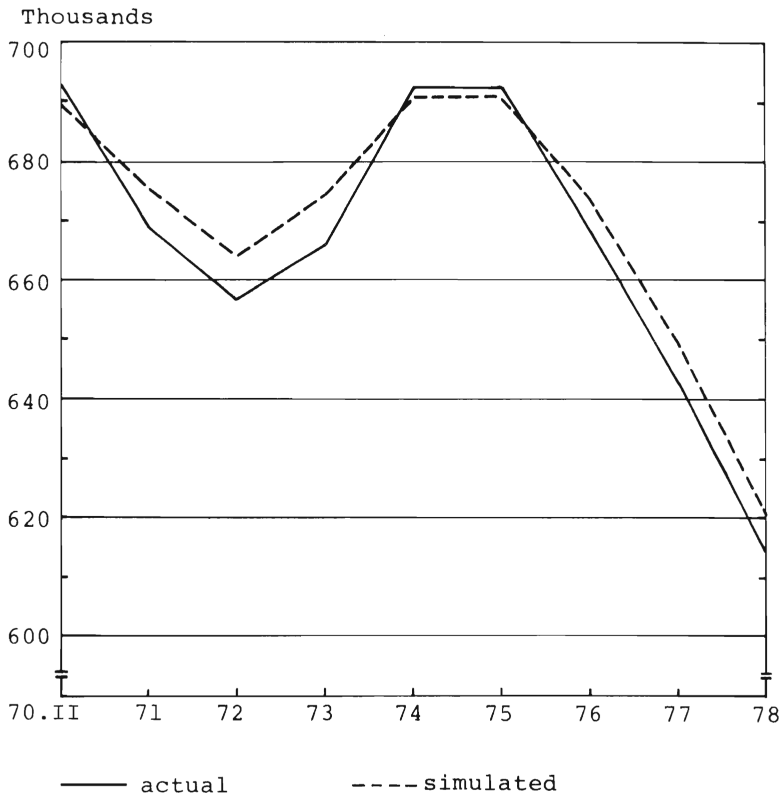

Figure 2. Actual and simulated unemployment rate

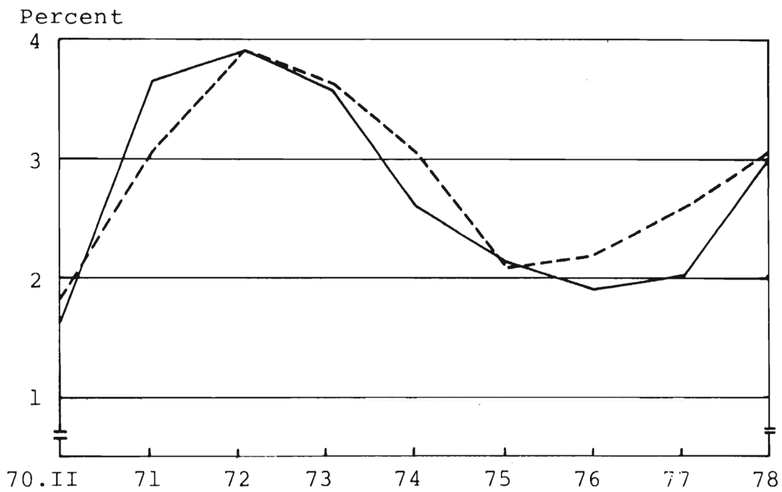

Table 5. Historical simulations - mean absolute errors

	Exogenous sales and orders	Exogenous output
Output	2.2	
New vacancies	13.6	15.8
Unfilled vacancies	22.8	24.8
Hirings	14.7	15.2
Layoffs	34.5	35.1
Quits	15.6	15.5
Hours per worker	4.2	4.1
Total hours	4.2	4.1
Employment	0.9	1.0
Unemployment	15.2	15.5

5.2.2 A policy experiment - Sweden with passive labor market policy

The level of Swedish industrial production fell with about 10 percent between 1974 and 1978. Despite this fairly dramatic fall employment stayed at lower levels than it did during the more normal recession 1971-72. The model developed is a suitable mean for analyzing this development. We will compare the "actual" development, as predicted by a full historical simulation of the model (the reference case), with an alternative case, simulated without employment and inventory subsidies and without the employment security legislation. The alternative case depicts a development produced by a "passive" policy, in contrast to the "active" policy adhered to in practice. Table 6 shows the results as average halfyear differences

Table 6. Sweden's active labor market policy - results of a policy simulation

	N_A/N_R	U_A/U_R	RU_A-RU_R	H_A/H_R	Q_A/Q_R	Y_A/Y_R	HRS_A/HRS_R	$OHRS_A/OHRS_R$
1974.II	0.998	1.169	0.40	1.053	1.020	2.254	0.999	1.001
1975.I	0.992	1.529	1.27	0.942	0.898	2.305	0.991	0.998
1975.II	0.986	1.934	1.66	0.873	0.812	3.165	0.987	1.002
1976.I	0.981	2.012	2.38	0.975	0.819	2.794	0.983	1.006
1976.II	0.981	1.991	2.00	0.980	0.862	2.778	0.983	1.006
1977.I	0.974	1.984	2.68	1.104	0.941	3.052	0.977	1.012
1977.II	0.965	2.012	2.62	1.343	1.029	2.806	0.972	1.028
1978.I	0.954	1.976	3.38	1.775	1.331	2.621	0.964	1.037
1978.II	0.957	1.782	2.34	1.857	1.489	2.567	0.967	1.034

Note: The variables are as follows:
N = The number of employed workers
U = The number of unemployed workers
RU = the rate of unemployment (%)
H = new hires
Q = quits
Y = layoffs
HRS = paid-for hours worked
OHRS = output per paid manhour

for a set of variables. The effects are presented in ratio terms, giving variable values under the alternative development (subscript A) divided by values under the reference case (subscript R). The absolute difference for the unemployment rate (RU) is also set out.

The simulation reveals that the policies undertaken have had important effects on employment, unemployment, labor turnover and labor productivity. The results indicate that employment would have been about five percent lower 1978 with a passive policy, corresponding to nearly 30,000 workers. The level of unemployment would have been approximately twice as high, implying a rise in the unemployment rate of about 2-3 percentage points. The "price" paid by this employment preserving policy is lower productivity - more labor hoarding - as displayed by the last column.

The policy effects on labor turnover are somewhat ambiguous. The most significant change concerns layoff behavior; a passive policy would have more than doubled the number of layoffs and the yearly layoff-rate (in relation to employment) would have been approximately 7 per cent in 1977. The number of quits, on the other hand, have been - due to the policy adhered to - higher in the beginning of the recession, responding to the lower unemployment, the higher vacancy stocks (due to larger vacancy durations) and the higher vacancy flow associated with inventory subsidies. A passive policy, on the other hand, would have induced more quits during the upturn of the business cycle. This, in turn, depends on the larger number of new vacancies and new hires that a less employment preserving policy would have required during the recovery.

6. CONCLUDING REMARKS

The aim of this paper has been to investigate certain unemployment effects of labor market policy programs undertaken in Sweden. This objective has been addressed via different econometric models, intended to elucidate (i) future unemployment experiences for those enrolled in relief works or training programs and (ii) macroeconomic effects of recent policies, in particular the consequences for the rate and duration of unemployment.

The microeconomic results obtained indicate, inter alia, a marked autocorrelation in individual unemployment probabilities; workers with unemployment experiences in 1967 are facing substantially higher unemployment risks in 1973 even after controlling for various personal characteristics. However, these adverse effects of previous unemployment appear to be mitigated by participation in labor market programs. The probability of being unemployed in the future shows a marked decline for unemployed workers in training programs.

Ambiguity is revealed to some extent when we explore the contra-cyclical role played by relief works and retraining programs. We are thus only able to establish transitory unemployment duration effects of these programs; changes in the (aggregate) magnitude of relief works and retraining are associated with changes in unemployment duration, but no effects from the magnitude per se are discernible. The results clearly underline the need for further and more elaborate evaluations of these programs.

Sweden's experiences during the deep recession of the late 70's are fairly remarkable. Despite a dramatic fall in aggregate demand it was possible to avoid rapid increases in unemployment. The policies have caused a marked reduction in layoffs. The employment security legislation may also have reduced turnover into unemployment via a significant "on-the-job-search effect"; the estimated unemployment equation indicates that a given number of layoffs have smaller unemployment effects during the late seventies than during earlier periods. It is notable that the higher firing costs - implicit in the legislative changes - so far appear to have caused no or neglible effects on the duration of unemployment.

DESCRIPTION OF THE SIMULATION MODEL

Variables explained by stochastic equations

FCV_t = filled and canceled vacancies (employment exchange statistics)

H_t = new hires (establishment data)

$HRSN_t$ = paid-for hours per worker

IV_t = the flow of new vacancies

O_t = index of industrial production in real terms

$\widehat{OHRS}_t$ = potential productivity

Q_t = quits

$SHRSN_t$ = standard paid-for hours per worker

$UCHR_t t$ = the relative rate of change of unemployment

Y_t = layoffs

Exogenous variables

B = index of backlog of orders in real terms

DINV1 = dummy for the inventory subsidy scheme 1972

DINV2 = dummy for the inventory subsidy scheme 1975-77

DLAS = dummy for the employment security legislation (DLAS = 1 from 1974.7 on, zero otherwise)

D2, D3... D12 = seasonal dummies

NSUB = the fraction of non-subsidized excess labor

S = index of sales in real terms

T = time

Definitions that close the model

$HRS_t = HRSN_t \cdot N_{t-1}$	(the total number of paid-for hours during month t)
$\widehat{HRS}_t = O_t/\widehat{OHRS}_t$	(the total number of productive hours during month t)
$\widehat{N^*_{t-1}} = \widehat{HRS}_t/SHRSN_t$	(desired employment at the beginning of month t, desired excess labor excluded)
$N_t=N_{t-1}+H_t-O_t-Y_t$	(actual employment at the end of month t)
$U_t = (1+UCHR_t)U_{t-1}$	(the number of unemployed at the end of month t)
$RU_t = 100 \cdot U_t/(U_t+N_t)$	(the unemployment rate at the end of month t)
$VAC_t = V_{t-1} + IV_t$	(the total number of vacancies during month t)
$V_t = V_{t-1} \; IV_t-FCV_t$	(the stock of unfilled vacancies at the end of month t)
$PFCV_t = FCV_t/VAC_t$	(filled and canceled vacancies as a fraction of the total number of vacancies)

Table A1. Estimation results (t-ratios in parentheses)[a]

Potential productivity

$$\ell n\widehat{OHRS}_t = \underset{}{0.0656} + \underset{(5.182)}{[0.502E{-}02]} \cdot T$$

Periods: 1969.1–1969.3, 1970.4–1970.6, 1974.4–1974.6

$R^2 = 0.79$ DW = 3.58

Standard hours[b]

$$\ell nHRSN_t = -1.846 - \underset{(-11.289)}{[0.171E{-}02]} \cdot T + SD$$

Period: 1969.1–1978.6

$R^2 = 0.95$ DW = 2.74

Industrial production

$$\ell nO_t = 4.861 + \underset{(4.966)}{0.122}\ \ell nB_{t-1} + \underset{(1.528)}{0.018}DINV1 + \underset{(1.336)}{0.011}DINV2 + \underset{(10.64)}{1.325} \sum_{i=0}^{5} w_i \ell nS_{t-i} + SD$$

Period: 1970.8–1978.12

$R^2 = 0.98$ DW = 1.36

New vacancies

$$\ell nIV_t = 6.008 + \underset{(4.868)}{0.453}\ \ell nQ_t + \underset{(3.152)}{2.206}\ ln(N^*_t \cdot N^*_{t+1} \cdot N^*_{t+2}/3N_{t-1}) + \underset{(5.741)}{5.492} \sum_{i=1}^{6} w_i \ell n(N^*_{t-i}/N_{t-i}) + SD$$

Period: 1969.2–1977.10

$R^2 = 0.95$ DW = 2.07

$\rho = 0.60$

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[a] Seasonals are not presented but indicated by SD if they appear in an equation. The intercept refers to January.

[b] Note that predicted values from this equation are interpreted as (log) standard hours.

Table A1 (cont.)

Hirings

$$\ell n H_t = 1.380 + \underset{(19.25)}{0.706}\,\ell n VAC_t + \underset{(2.080)}{0.103}\,\ln U_{t-1} - \underset{(-2.273)}{0.077} DLAS + \underset{(4.236)}{1.755}\,\ell n(N^*_t \cdot N^*_{t+1} \cdot N^*_{t+2}/3N_{t-1}) + SD$$

Period: 1969.2-1978.12
$R^2 = 0.97$ DW = 1.43

Layoffs

$$\ell n Y_t = 4.273 + \underset{(2.252)}{0.261}\,\ell n U_{t-1} - \underset{(-2.683)}{0.452} DLAS + \underset{(2.637)}{1.205}\,\ell n NSUB_t - \underset{(-2.506)}{1.859}\,\ell n(N^*_t \cdot N^*_{t+1} \cdot N^*_{t+2}/3N_{t-1}) - \underset{(5.097)}{6.216}\sum_{i=1}^{6} w_i\,\ell n(N^*_{t-i}/N_{t-i})$$

Period: 1969.2-1978.12
$R^2 = 0.55$ DW = 1.66

Hours per worker

$$\ln(HRSN_t/HRSN_{t-1}) = 0.032 + \underset{(54.97)}{1.058}\,\ln(SHRSN_t/HRSN_{t-1}) + \underset{(3.920)}{0.352}\,\ln(N^*_{t-1}/N_{t-1}) + \underset{(1.647)}{0.043}\,\ln(\widehat{HRS}_{t+1} \cdot \widehat{HRS}_{t+2} \cdot \widehat{HRS}_{t+3}/3\widehat{HRS}_t)$$

Period: 1969.2-1978.12
R^2 0.98 DW = 2.55

Table A1. (cont.)

Quits

$$\ell nQ_t = \underset{}{2.683} + \underset{(2.369)}{0.291}\ \ell nVAC_{t-1} - \underset{(-1.866)}{0.093}\ \ell nU_{t-2} + \underset{(3.682)}{0.486} \sum_{i=0}^{8} w_i \ell nH_{t-i} + SD$$

Period: 1969.2-1977.10

$R^2 = 0.94$ DW = 1.42

Unemployment

$$UCHR_t = -0.020 + \underset{(3.279)}{3043.63}(1/U_{t-1}) + \underset{(2.505)}{0.743}(Y_t/U_{t-1}) - \underset{(-2.027)}{0.686}(DLAS \cdot Y_t/U_{t-1})$$

$$- \underset{(-2.843)}{0.037}(VAC_t/U_{t-1}) + SD$$

Period: 1969.2-1979.5

$R^2 = 0.62$ DW = 1.93

Filled and canceled vacancies

$$\ell n\left(\frac{1-PFCV_t}{PFCV_t}\right) = -6.479 + \underset{(14.16)}{0.774}\ \ell nVAC_t - \underset{(-3.146)}{0.195}\ \ell nU_{t-1} + \underset{(2.924)}{0.178}DLAS +$$

$$+ \underset{(4.521)}{[0.466E-02]} \cdot T + SD$$

Period: 1969.2-1978.12

$R^2 = 0.86$ DW = 1.85

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Relief Work and Grant Displacement in Sweden

Edward M. Gramlich and Bengt-Christer Ysander

INTRODUCTION*

Public relief work has been a main-stay of the very active Swedish labor market policy. As early as 1965, when expenditures on labor market policy were still less than one percent of GNP, 10 thousand workers were employed under the relief work program. By 1979, when labor market policy expenditures had risen to more than three percent of GNP, 48 thousand workers were employed. Under the very strong assumption that all or most of these workers would have been unemployed without the program, the relief work program could have single-handedly reduced the overall unemployment rate for Sweden by one and a half percentage points.

The program could have had an even stronger impact on unemployment rates for certain subgroups of the labor force. In 1979, 30 thousand youths were employed by the program - implying that the pro-

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* We have benefited from discussions with Richard Murray, from the help in finding and interpreting data of Charles Öberg and Stefan Goés of the Labor Board. Erik Mellander has provided first class research assistance both in data collection and computation. Our FIML-estimations were made possible by the generous collaboration of Leif Jansson, the originator of the program used.

gram could have lowered youth unemployment rates by as much as three percentage points. Regionally, 19 thousand jobs were located in the seven forest countries, lowering the unemployment rate by as much as two and a half percentage points in those regions.[1]

These calculations assume that the relief workers would have been unemployed without the relief work jobs. There are several reasons why such an assumption could overstate, perhaps dramatically, the unemployment impact of the program:

a) relief workers may be performing jobs "normally" done by regular state or local government employees, hence reducing normal public sector employment demands. This is the possibility commonly known as the grant displacement effect.

b) relief workers may be performing jobs that would "normally" have been done by regular government employees at some later time, say as part of an effort to induce a counter-cyclical timing of government expenditure and employment. We could call this special kind of grant displacement "intertemporal" displacement.[2] It might be offset by

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[1] Handicapped workers also used to be heavily overrepresented in relief work. In recent years, however, the relative importance of relief work for this group has diminished, partly due to the setting up more permanent workshops, specially adjusted to their needs.

[2] To effect this kind of displacement has indeed been the explicit aim of Swedish labor market policy in the post-war period. Expansion of public capacities in times of recession should be matched by moderation during the up-swing, relieving some of inflationary pressure in the private labor market.

the fact that the presence of relief workers in a recession creates pressure to make the positions permanent later on.

c) the relief work program could push up lower wage rates in the private market, forcing private employers to lay off some low wage workers.

d) the relief work program could, by pushing up low wage rates and/or providing employment guarantees, expand the supply of labor and not generate a one-for-one reduction in unemployment. The latter two possibilities might be termed wage displacement effects.

Should any of these displacement effects occur, one does not necessarily become less enamored of relief work as a labor market policy option. The supposed benefits of the program as a way to reduce overall unemployment are a good deal less than might be imagined, however, and the evaluation of the program becomes a good deal more complex.

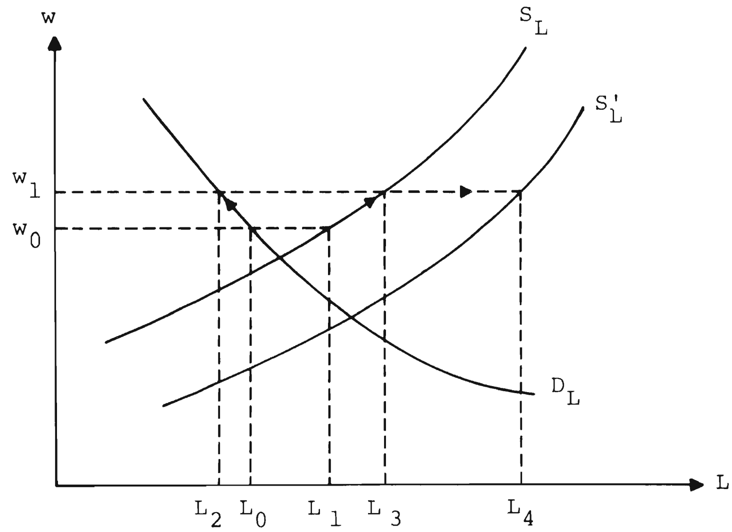
The same displacement questions - although often perhaps more conveniently couched in terms of expenditure - arise also with other kinds of grants. Measuring displacement is one way of indicating the allocative effectiveness of a grant policy. The effectiveness of grants to local governments - the receiving group we have here chosen to study - is indeed a question of strategic importance for Swedish economic development today. Real expenditures by local governments have during the last two decades increased twice as fast as GNP, raising their share to a quarter. During the same period the part financed by state grants has increased from about a fifth to a fourth with a growing emphasis on categorical grants. In the perspective of an expected slow growth of the total economy in the eighties much interest is focused on the possibilities of controlling the volume and pattern of local government expenditures by grant policy.

In this paper we try to examine the displacement effects for the Swedish relief work (beredskapsarbeten) program and to compare its impact on local government expenditure and employment with that of ordinary state grants. We begin with a simple theoretical demonstration of how the displacement processes might work and ways in which their presence or absence can be identified. We then proceed to outline a model by which the strength of the grant displacement process can be estimated, a model which draws on both of our own previous work. (Gramlich and Galper, 1973; Ysander, 1981). We go on to describe how the Swedish program fits the model, and try to estimate the model using Swedish data for the period 1964-77. Although such empirical examinations have been conducted a few times in the United States (Johnson and Tomola, 1977; Borus and Hamermesh, 1978) to our knowledge nobody has yet done any econometric estimations of grant displacement for Sweden.

WAGE DISPLACEMENT

Taking up first the wage displacement phenomenon, the process can be depicted by the elementary demand and supply diagram shown in Figure 1. The

Figure 1. Public employment and wage displacement



curves D_L and S_L represent the normally-sloped private demand and supply curves for low wage labor. For any of a number of well-known and much-discussed reasons (turn-over unemployment, minimum wages, etc), we assume that this labor market does not clear initially, leaving a wage of w_0 and initial unemployment of the amount L_0L_1. Policy-makers respond by passing a public employment program which offers, say, a higher wage of w_1 to all who are willing to work. Private employers must also pay w_1 or risk losing all of their work force: when they do so, private employment falls by L_0L_2. The higher wage brings L_1L_3 workers into public employment. In addition some "discouraged workers" will be induced to supply labor by the job guarantee - this shifts the supply curve to the right and adds to public employment by L_3L_4.

The total public employment work force of L_4L_2 is then composed of some workers who were formerly unemployed (L_1L_0), some who were formerly employed privately (L_2L_0), and some who were formerly out of the labor force (L_1L_4).

It should be pointed out immediately that even though only a share of the public employment job gains actually reduce unemployment, such an outcome is not necessarily socially undesirable. For one thing, all low wage workers gain wage income, and if wages are at least somewhat correlated with family income, a public employment program operating in this manner will redistribute income. Secondly, the expansion of supply may consist of workers who really should have been counted as unemployed already - they just were not in the labor force because of perceived difficulties in finding a job - and so the social gain represented by the increased wage income may be almost as great as that for unemployed workers.

In the United States, where public employment wage levels are set as part of the program and are normally slightly above prevailing minimum wages, the wage displacement issue is a very important one - many more likely entrants to a public employment program come from existing low wage private employment or supply expansion than from existing unemployment. (See Betson, Greenberg and Kasten, 1980; or Gramlich and Wolkoff, 1979.) For Sweden, generalizations are risky but the phenomenon does not appear to be so important. One indication that it may not be so important lies in the newly-emerging studies of the Swedish labor market; supply wage elasticities appear, e.g., to be very small

Sweden.[1] If these results hold at low wages, wage displacement for the relief work program should be much more modest. Moreover, the Swedish program appears to be much more confined to workers likely to be unemployed initially - youth, handicapped, and workers in unemployment pockets in the forest counties - than the U.S. program.

But even though wage displacement may not be as large as in the U.S. program, it may not be entirely absent either. The relief work program is supposed to pay "going wages" for a particular task, but these wages are probably above what relief workers could have commanded from the private sector (hence a quality-corrected w_1 would exceed w_0). It has also been asserted that inexperienced youths prefer relief work to ordinary employment at the same wage, and indeed will shun private sector vacancies to take the relief work jobs. If true, this phenomenon could either have the same implications as the outward shift in supply described by Figure 1 or exert a certain upward pressure on private wages, hence implying that there may be some wage displacement.

GRANT DISPLACEMENT

The other type of displacement is that working through the grant system. Both in the U.S. and Sweden, public employment programs are actually carried out through grants to recipient agencies

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[1] See the paper by Axelsson, Jacobsson, and Löfgren in this volume for the supply results. A precise interpretation is that the amount represented by L_1L_3 is negative.

to hire eligible workers. In the U.S. these grants go from the Labor Department to local governments and a few quasi-governmental bodies called Community Based Organizations. In Sweden they go from the Labor Board to some state agencies, local governments, and a small number of private employers. While it may be necessary to conduct public employment through grants to agencies that are producing normal government services, this does lead to the possibility of grant displacement. Estimates of the phenomenon are very large for the U.S. To suspect that the possibility is relevant for Sweden, one need go no further than the law itself. For public investment activities it is required that relief workers be used to build projects already in the long term plans of state and local governments.[1]

In studying the employment effect of grants to local governments, the first thing one has to do is to differentiate between types of grants.[2] Three types can be distinguished:

1) Close-ended noncategorical grants, which shift upwards the budget-line of the receiver, acting like a straight income subsidy.

2) Open-ended categorical grants, which reduce the relative prices facing the receiver at the margin.

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[1] The details of both the relief work system and other Swedish categorized grants are summarized in three recent Royal Commission reports (SOU 1973:45, SOU 1975:39, SOU 1977:78). These reports contain a good deal of discussion on grant displacement in general but no attempts at measurement.

[2] For a more thorough discussion of the various types of grants see Gramlich (1977).

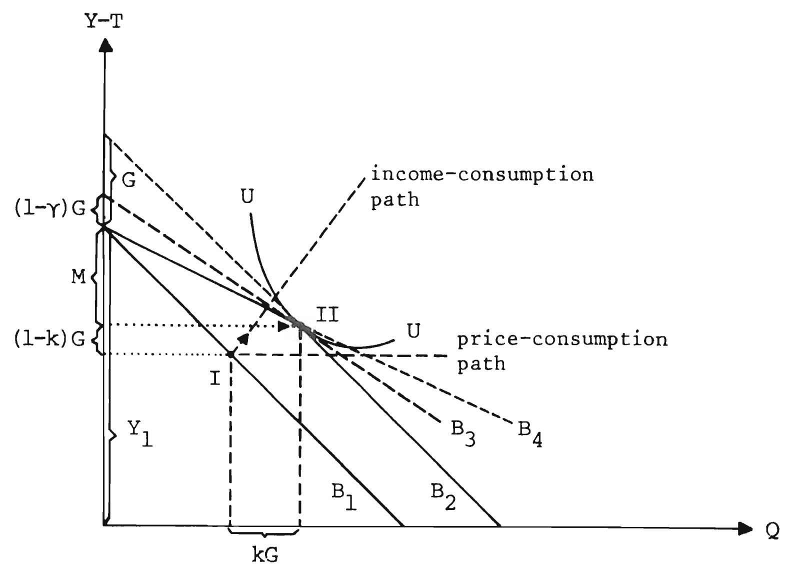
3) Close-ended categorical grants, which reduce prices but only up to a point determined by central government restrictions on the size of the overall grant.

In the first two cases we know directly how the receiver's budget constraint will be affected. From estimated price and income elasticities we can then predict the grant effect on expenditure and employment for a utility-maximizing receiver.

In Sweden, however, most grants are not of either of these types. Most are categorical and all have a limit on the total amount, making them type 3 grants in the above classification. For these hybrid grants, the proper treatment becomes more involved.

The various possibilities can be represented by the indifference curve diagram shown in Figure 2.

Figure 2. <u>Displacement of close-ended grants - the general case</u>



As applied to governmental decision making, the indifference diagram is supposed to reflect the behavior of some vague amalgam called "the recipient government decision-making body". In certain carefully specified but probably not very realistic models of governmental behavior, that body is the famous median-voter; in less precise but perhaps more realistic models it is not entirely clear who is behaving - some weighted average of voters and bureaucrat-politicians, where the latter have a higher per capita weight.

In the diagram the government is choosing between expenditures on public services, Q, and private goods and other income uses, which we call after tax income (Y-T). To simplify the exposition, we assume that initial prices are equal, so that the slope of the initial budget line B_1 is -1. The initial allocation is then at point I.

Taking first type I grants (close-ended noncategorical), these would shift the budget line to B_2, parallel to B_1, and move the community along the income-consumption path. Not all of the grant would go into public expenditures in this case unless the income elasticity of demand for private goods were zero, or the income-consumption path horizontal.

Type 2 grants would be treated as a straight-forward reduction in the price of public services and move the community along the price-consumption path (drawn horizontally to correspond to the case where the price elasticity of demand is unity). Type 3 grants then kink the budget line at point II and move the recipient unit to this kink point.

As the funds limit gets tighter and tighter, or the kink point moves leftward, the impact of these grants approaches that of type I grants; as the limit becomes less stringent, the impact approaches that of type 2 grants.

There are two ways of dealing with type 3 grants in empirical work. One approach, used by Gramlich and Galper (1973), is to find point II directly. This is done by assuming the government receives the entire amount of the grant, G, spends its legally-mandated matching amount, M, and then <u>reduces</u> spending on other public goods that are substitutable with the grant-supported goods. The latter reduction is called the grant displacement impact, and it obviously can imply that the overall spending effects of grants are a good deal less than those mandated by law.

A second approach was first used by McGuire (1978). It converts a type 3 grant into an income term (like type I grants) and a price term (like type 2 grants), and finds the appropriate shares. Diagrammatically, this amounts to finding the budget line that passes through point II tangent to the indifference curve, so that γG of the grant works like a price subsidy and $(1-\gamma)G$ like an income subsidy. Such a budget line is described by B_3 on Figure 2. Were the kink point to the right of the intersection of budget line B_4 and the price-consumption path, the limit is ineffective, the grant is entirely of type 2, the tangent to the indifference curve is also line B_4, and $\gamma = 1$. Were the kink point to the left of the intersection of B_4 and the income-consumption path, the tangent to the indifference curve is parallel to B_2, the grant is entirely of type 1, and $\gamma = 0$.

Of more interest than γ is a parameter k, by which we denote the impact of the grant on public spending. This parameter can be shown to be related to γ by the identities

$$\frac{\partial Q}{\partial G} = k = -\gamma e_{\pi} + (1-\gamma) e_{y} \frac{Q}{(Y-T)} \tag{1}$$

where e_{π} and e_{y} denote price and income elasticities of demand respectively. In the diagram kG, the increased public spending due to the grant, is the horizontal distance between points I and II; and (1-k)G, the increased private spending, is the vertical distance. Returning to our original issue, (1-k)G is also called the displacement amount because it shows how much of the grant "trickles down" to private spending.

In our empirical work we have chosen to use both approaches. For non-relief work categorical grants the central government share of total expenditures $g = G/(M+G)$ is relatively small and there is a good deal of negotiating its exact size and other grant conditions between the granting authority and recipient governments. Given this bargaining, and the large number of such grants, we have found it more convenient to deal with these grants using the McGuire approach and simply estimate γ from the data.

For relief work grants the situation is much different. The central government share, g, is very high and indeed very close to one. It is constant across communities, relief work grants, and employment, is a small share of normal employment in the projects the relief workers participate in. In view of this, it makes sense to constrain $\gamma=0$, assuming that relief work grants have no marginal impact on relative prices, and estimate the displacement coefficient (1-k) directly.

As with wage displacement, it should be pointed out that a strong degree of grant displacement does not necessarily imply that the relief work program is failing as a device to raise the demand for certain types of labor. Indeed, if the conditions of the grant are enforced adequately, the relief work program will stimulate demand for unemployed or otherwise disadvantaged workers. In this regard, data supplied by the Labor Board do indicate that the overwhelming majority of workers have been referred from the labor exchanges, and hence were not employed before entering the program. The fact of grant displacement then, could merely imply less demand for higher wage regular public sector workers, perhaps not an unfavorable result if these regular workers can get other jobs. Hence, if grant displacement is of this employment-switching type, the program is altering the composition of overall labor demand in favor of disadvantaged workers or workers in regional pockets of unemployment and simultaneously allowing the work of the public sector to be performed at lower wage rates - no mean feat.

AN EMPIRICAL MODEL OF GRANT DISPLACEMENT

An empirical model of grant displacement can be developed from orthodox utility maximization principles. The procedure, as applied in the following to local government, involves assuming that our governmental decision-making body is motivated by a utility function consisting of argument in:

a) currently produced governmental goods and services,
b) private consumption,
c) government capital stock,
d) the change in the net value of assets, reflecting fiscal independence and flexibility of the community.

Mathematically, the utility function - assumed to be quadratic - can be expressed as:

$$U = U(Q,\ Y\text{-}T,\ K,\ F), \tag{2}$$

where Q stands for public services, Y-T is income (Y) less local taxes and charges (T), a measure of private consumption, K is the measure of capital stock and F is the change of net value of assets, with all variables being defined in real terms.[1] All arguments are assumed to have positive first and negative second derivatives. This utility function involves directly flows of current income and expenditure, Q, Y-T and F, and one stock, K. This yields a stock adjustment behavior in the government's response to income or price changes - a rise, say, in income will raise public consumption and private consumption directly, and also

[1] One could argue for including either the level or the change in F in the utility function. In some sense the community is better off the higher is the level of F, regardless of how much this level has grown recently. In another sense, however, local governments face a legal constraint on F - it cannot go below a certain level (penalizing governments more the closer is F to this legal constraint).

generate a desire for increased capacity or stocks of capital. Once construction and internal saving have increased these stocks to the proper level, no further accumulation is necessary and public and private consumption can rise yet again.[1]

The Q variable needs however to be further specified, due to the interpretation we have chosen to give relief work grants. We are assuming that there is a constant value, p, attached to relief work output compared to that of regular employment. The utility function can therefore be written as:

$$U = U(Q'+p\,\frac{R}{g_r},\ Y-T,\ K,\ F) \tag{3}$$

where R is the real value of relief work grants, g_r denotes the grant share of total relief work costs, R/g_r is the total real cost of relief work, and Q' stands for the regularly produced services.

The utility function is then maximized subject to the government's budget constraint. This yields public goods demand functions of the form:

$$Q' = a_0+a_1\Theta+a_2\pi-a_2(\gamma g\pi)+a_3(Y-T)+a_4\,\frac{R}{g_r} \tag{4}$$

where Θ is some socio-economic shift variable representing service needs, π is the relative price of regular services, $\gamma g\pi$ is the effective reduc-

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[1] The precise details of all this are worked out in Ysander (1981). The Gramlich-Galper model (1973) deviates slightly in using stocks of financial assets directly in the utility function.

tion in this relative price due to regular categorical grants, with central government matching share g and with the γ value estimated from the data. The terms represented by the parameter a_2 together give the effective marginal price for public services, while the a_3 coefficient shows the marginal public spending propensity as community income changes. The parameter a_4 measures relief work displacement; if a_4 is close to zero, there is little displacement; if it is close to minus one, a great deal.[1]

It is also possible to estimate (5) for public employment by making use of the following approximations:

$Q' \rightarrow E$ = regular employment
$R/g_r \rightarrow E_R$ = relief work employment
$\pi \rightarrow w$ = real wage (gross of subsidies but net of the cost percentage paid by user charges).

Expression (5) then becomes:

$$E = a_0 + a_1\Theta + a_2 w - a_2\gamma g w + a_3(Y-T) + a_4 E_R \qquad (5)$$

The model was meant to be used for the study of grants to local governments. There are several reasons for not trying to use the same explanatory

[1] Note that a_4 = -p from the utility function (3). When a_4 = 0, the lack of displacement results from the fact that R and Q' are not substitutes in the utility function.

framework for state agencies, the other main recipient of relief work grants from the Labor Board. For one thing, state agencies do not usually think of themselves as allocating resources between private and public users the way local governments do. Also in Sweden projects suitable for relief work are usually earmarked years in advance within the agencies' revolving five-year plans. Total displacement is then virtually guaranteed. Defense is, however, a notable exemption, since up till recently, relief work grants - although used for purposes laid out in long-term plans - were not included in the financial four-year (usually) expenditure limits set out by parliament. In this case we would therefore expect little displacement in terms of production, although there could still be a considerable employment displacement because of the changing product mix within the financial limits.

THE DATA

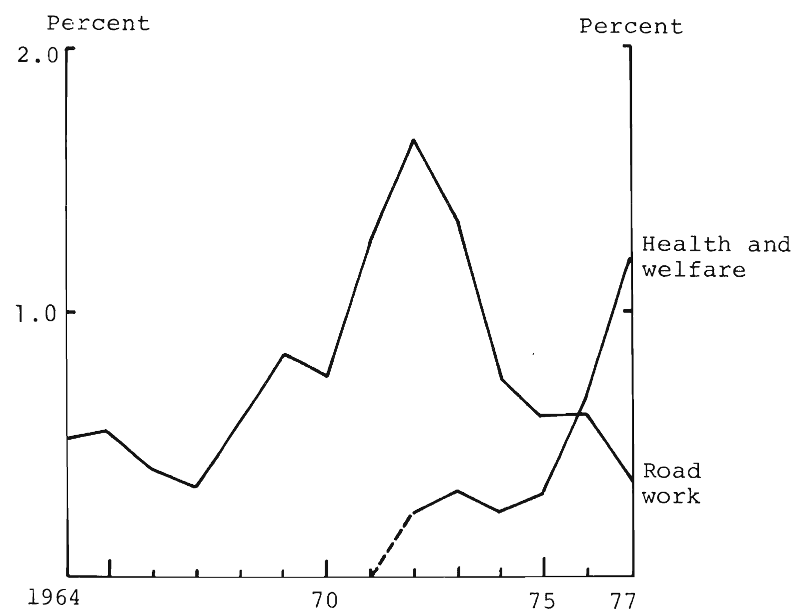
Within local governments in Sweden relief work is very unevenly distributed between different categories of service. The distribution has also changed dramatically in the seventies, with the traditional construction work being more and more superseded by jobs within health and welfare. Relief work grants to local governments have tended to increase in relative importance over the last two decades and at the end of the seventies paid for around one percent of all service expenditure. It is hard to estimate reliable relief work displacement coefficients for health and welfare, however, because the program only got started in the early seventies and its expansion coincided with the expansion (for other reasons) of overall health and welfare spending.

For our empirical study of grant impact we have chosen two categories, health and welfare, and road work. The official relief work statistics are such that a further separation within the category health and welfare cannot be made with any confidence. We have picked road work, although it by no means dominates the traditional relief construction work, because the classification of relief work statistics and other financial statistics here agrees better than with other construction works. In defining road work, we have lumped together maintenance and new construction to avoid being misled by the possible shifting of regular employees into new construction occasioned by an increased relief work on maintenance.

Within the local government sector one could fit the model described here either to time series or cross section data. Cross section data for local governments are in general very good in Sweden, with numerous observations and a reasonable amount of variance in most independent variables (Murray, 1977, 1981). These general advantages may, however, not hold unreservedly when dealing with relief work grants. Since provisions of the grants are essentially constant across counties, there is very little variation in these critical independent variables. In any case, for this paper we have not had a chance to use cross-section data, but have confined our attention to the more limited time-series data. Hence we study annual time-series data on employment and expenditures for the period 1964-77.

Figure 3 shows the percentage of total work within the categories that has been done as relief work in the period studied. As noted above, the share of total expenditures comprised by relief work is very small, one or two percent. This is why we constrain $\gamma = 0$ in our estimation. The figure also high-lights the fact that the recession in the early seventies was the last time road work was used as a major form of relief work, while later relief work endeavors have tended to be more and more directed towards the health and welfare area.

Figure 3. <u>The share of relief work in total expenditure, $(R/g_r)/(Q'+R/g_r)$</u>



ESTIMATION RESULTS

In fitting the equations (5) and (6) above we were trying to estimate the price-subsidy effect of ordinary grants and the displacement of relief work respectively. From these estimates we can derive and compare the net impact on expenditures of these two kinds of grants.

The main econometric difficulty involves the parameter γ, determining the price and income components of non-relief work categorical grants. There is no reason why (5) could not be estimated directly with OLS, γ being then determined by comparing a_2 and $-a_2\gamma$, the coefficients of π and πg respectively. However this approach would not give a standard error for γ, and to fill that gap we have therefore estimated the model with a FIML program.

On a more practical level, as a shift or needs variable for health and welfare we have used a population index, where the various age groups are weighted by their earlier relative per capita share of total expenditure in this area. The corresponding shift variable for road work is an index of the number of heavy trucks in traffic, meant to reflect the changing demands made by heavy road transport.

HEALTH AND WELFARE

The results for the health and welfare category are shown in Table 1. All equations explain normal expenditures, Q', because the results for employment were not reliable. In the first equation the

Table 1. Equations explaining local government regular expenditures for Health and Welfare, Q', 1964-77

Estimated in index form; 1975=100
Absolute T value below coefficients

Equation	Dep. variable	a_0	a_1	a_2	γ_I	a_3	a_4	R^2	DW
(1)	Q'	-374.4	5.68	-1.23[a]	0.22[b]	0[c]	0.015[d]	0.98	1.39
			(6.2)	(2.6)	(1.4)		(1.3)		
(2)	Q'	-383.8	5.87	-1.06[a]	1.0[b]	0[c]	0.017[d]	0.98	1.35
			(6.6)	(2.4)			(1.3)		
(3)	Q'	-383.1	5.60	-1.13[a]	1.0[b]	0.33[c]	0.017[d]	0.98	1.68
			(6.4)	(3.0)		(1.3)	(1.4)		

[a] Implies price elasticities of -1.36, -1.17, and -1.25 respectively.

[b] In equation (1) the γ_I value estimated in index form implies a value of γ in absolute terms of 1.32. In equations (2) and (3) γ_I is constrained to equal 1.0.

[c] In equations (1) and (2) a_3 is constrained to equal zero. In equation (3) the implied income elasticity is 0.41.

[d] Implies relatively large negative values of p in text equation (3), or that relief work services and regular services are complements.

income term is omitted because of its collinearity with relative prices, and γ is estimated to be 1.32, outside of its theoretical band. Hence in equations (2) and (3) γ_I is just set at one - implying that categorical grants are treated like open-ended price subsidies - and the equation re-estimated with and without the income term. The price elasticity is computed to be slightly in excess of one in all equations, and the income elasticity is 0.41 in the one equation where it could be estimated. In comparison with other studies, these estimated price elasticities are on the high side, but the estimated income elasticity is standard.

Regarding displacement, all equations showed relief work to have a positive effect on normal employment. Relief work employment is not a substitute but a complement to normal employment in the health and welfare area. The coefficients are not statistically significant, but are nevertheless fairly large. The only explanation for the result we can see is that relief work employment necessitates more regular employees in supervisory positions. We are inclined to view our precise estimates skeptically, but we should stress that there is no evidence of displacement as far as health and welfare spending go. Indeed, if anything grant displacement is negative.

ROAD WORK

Table 2 shows the results for road work, this time estimated both for normal expenditures, Q', and regular employment, E. As before the initial estimate of γ in (1) was high and we constrained $\gamma=1$

Table 2. <u>Equations explaining local government regular expenditures for Road work, Q' and E, 1964-77</u>

Estimated in index form; 1975=100
Absolute T value below coefficients

Equation	Dep. variable	a_0	a_1	a_2	γ_I	a_3	a_4	R^2	DW
(1)	Q'	-97.9	1.27	-0.41[a]	0.41[b]	0.98[c]	-0.0098[d]	0.95	1.86
			(9.1)	(2.1)	(1.5)	(4.0)	(0.9)		
(2)	Q'	-71.6	1.18	-0.41[a]	1.0[b]	0.98[c]	-0.0078[d]	0.95	2.09
			(7.4)	(2.1)		(4.0)	(4.0)		
(3)	E	-11.3		0.56[a]	1.0[b]	0.56[c]	-0.048[d]	0.89	1.87
				(3.2)		(1.4)	(2.1)		

[a] Implies price elasticities of -0.37, -0.37 and -0.51 respectively.

[b] The index estimate implies for (1) a value of γ in absolute terms of 2.20.
In equations (2) and (3) γ is constrained to 1.0.

[c] Implies income elasticities of 0.88, 0.88, and 0.50 respectively.

[d] Implies values of p around 1 in equations (1) and (2) and around 5 in equation (3).

in (2) and (3), again indicating that grants appear to be treated mainly as price subsidies. This time both the price and income elasticities are less than one in absolute value, as is usually found for public expenditure functions. But this time the estimates indicate relatively complete displacement for the two normal expenditures equations, and more than complete displacement in the employment variant. Whether we should believe the precise estimates is again questionable, but the evidence suggests pretty strongly that there is a great deal of displacement in this area.

Although the estimate is statistically insignificant an inspection of the time series shows that the high figure is no mere trick played by multicollinearity, etc. When, e.g., the relief work multiplied during the recession in the early seventies, the stagnation of regular road expenditure turned into an outright fall, which was even more accentuated in terms of employment. The aggregate figures seem to suggest that total local road work during the period has tended to move with the business cycles. The effort to comply at the same time with the requirement of concentrating road investment to periods of high unemployment has resulted in a downturn in the labor intensity of regular road work during these periods.

While complying with all formal requirements the local governments thus seemed to have managed to make a negative total contribution to the labor market efforts. This is probably what shows up in the large displacement coefficients estimated for relief work. Whether the percentage-wise rather small number of relief workers affect this situation significantly is, however, difficult to know and impossible to ascertain from the aggregated series available.

In many ways these estimates leave a lot to be desired, but at least within functional categories the results are reasonably consistent on the displacement issue - there is not much for health and welfare, and there is a great deal for road work. To go beyond these conclusions one would appear to need a more detailed analysis - perhaps utility functions elaborated to allow complements as well as substitutes, certainly longer time series, and perhaps more use of cross section data.

There is a further statistical distinction that would be interesting to pursue. Above all it would be interesting to see whether the omission of handicapped workers would alter the estimated grant effects. The share of positions for handicapped workers - defined as positions that are tailor-made for the needs of people with physical or psychic handicaps or locally-tied elderly workers - has fluctuated from about one-third in the early days to two-thirds in the early seventies and back to one-fourth recently. One might expect displacement to be less for these workers.

CONCLUSIONS

The estimation results unfortunately do not permit any far-going or general conclusions to be drawn. As for ordinary grants, the results would seem to indicate that grant policy within the categories studied is a rather effective way of controlling local government expenditure. The estimations suggest the existence of considerable price effects and do not make it possible to reject the hypotheses that all categorical grants, which cannot *a priori* be viewed as bloc grants, work as if they were open-ended.

Displacement of relief work could only be identified in one of the categories - road work. There, the aggregate data do undoubtedly indicate a very considerable displacement effect - the regular work-force becoming reduced by more than the number of relief workers. But the evidence is just as strong that there is no displacement in the other category - health and welfare. The explanation appears to be related to the fact that in Sweden health and welfare relief workers are complementary with normal workers, and hence the employment-inducing impact of grants in this area is very strong.

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Part II

Unemployment Fluctuations and Inflationary Expectations

A Useful Restriction on the Offer Distribution in Job Search Models

Kenneth Burdett*

1. INTRODUCTION

In recent years a large amount of work has been accomplished on the empirical implementation of job search theory (see, for example, Barron, 1975, and Kiefer and Neumann, 1979). Much of this work, however, has been hindered by the relatively well known fact that there are, in general, indeterminant consequences on the outcome of the job search process when there is a change in labor demand conditions.[1] The purpose of this short study is to specify a reasonably general restriction which implies that specific predictions can be made about the consequences of changes in the demand for labor conditions. Further, the restriction used, which is placed on the distribution of wage offers faced by the unemployed worker, adds some new insight into the nature of job search models.

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[1] See Barron(1975) for details.

*Thanks are due to A.S. Goldberger for helpful discussions. This work was supported in part by National Science Foundation grant. Further work on material in this paper is currently underway with Jan Ondrich.

A shift in the demand for labor conditions faced by an unemployed worker will involve at least one of the following types of changes.

Type 1: A change in the probability (per period) the worker finds a vacancy, given a particular search intensity.

Type 2: A change in the wage offer distribution.

The objective of the study is to investigate the consequences of both types of changes on (a) the expected duration of a completed spell of unemployment, D, and (b) the expected post-unemployment wage of a worker, $\bar{W}$. For the majority of the study it will be assumed that an unemployed worker fully predicts the changes in labor market conditions. Nevertheless, the case where the worker is unaware of such changes will also be investigated. It has been noted that a type 1 change in labor market conditions may increase or decrease D. Further, it can be shown that a type 2 change may increase or decrease $\bar{W}$. Later in this study it will be demonstrated that if a particular restriction is placed on the cumulative distribution function (cdf) describing the wage offers, then a fully predicted change of either type will lead to a reduction in D and an increase in $\bar{W}$.

Recently, several authors have studied the consequences of a type 1 change in labor market conditions (Barron, 1975, Feinberg, 1977, Axelsson and Löfgren, 1977, and Björklund and Holmlund, 1981). Within the context of a job search model, Barron has shown that a type 1 improvement (when fully predicted) will increase an unemployed worker's reservation wage. Thus, a type 1 improvement implies the worker is more likely to receive offers (Barron's "Effect 1"), but less likely to accept an offer (Barron's "Effect 2"). Barron concludes that the net effect is, in general, indeterminate. He conjectures, however, that it would depend on such things as the cost of search and the unemployment compensation received by the worker. Feinberg subsequently demonstrated that if the cdf of wage offer is Uniform, D will always decline when there is a fully predicted type 1 improvement.[1] One objective of the present study is to generalize Feinberg's result.

To illustrate the restriction used to obtain our results let

$$V(z,G) = E\{w \mid w > z \text{ and } w \text{ is distributed according to cdf } G\}. \qquad (1)$$

$V(\cdot, G)$ is termed the truncated mean function (tmf) of G. It is straightforward to show that the tmf of any particular cdf is a unique transformation of it. It will be shown that if the tmf has a slope less than 1, then D will always decline with a fully predicted type 1 improvement in labor market conditions. It should be noted that many well known types of distribution functions have a tmf which satisfy this restriction. Nevertheless, there are some exceptions.[2] When the restriction

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[1] Feinberg (1977) conjectures, and goes far in establishing, that if the cdf of wage offers is Normal, then a type 1 improvement lowers the expected duration of unemployment.

[2] Goldberger (1980) specifies the tmf of several well known types cdfs. He also shows which have a slope less than 1.

does not hold, it will be shown that a fully predicted type 1 improvement will increase D if the discount factor is low enough. The expected post-unemployment wage, $\bar{W}$, will always increase with a predicted type 1 improvement.

Both the Barron study and the Björklund and Holmlund study consider a type 2 change in the demand for labor conditions. However, both these studies assume that such a change in demand for labor is not fully predicted by unemployed workers (Barron's "Effect 3"). Suppose for the moment that such a change is fully predicted. Within the context of a job search model, a type 2 change may be interpreted as a change in the cdf of wage offers. There are many ways in which a cdf of wage offers can be said to improve. For example, suppose the wage offer distribution is a Normal cdf. In this case a type 2 improvement may take the form of an increase in the mean and the variance of this distribution. In the present study such improvements will be ruled out by assumption. Specifically, it will be assumed that any change in the cdf of wage offers takes the form of a translation of location (in the case of a Normal cdf this would imply that the mean can change but not the variance).[1] In this case a fully predicted type 2 change in labor market conditions has an indeterminate effect on $\bar{W}$. Nevertheless, it will be shown that a predicted type 2 improvement will always increase $\bar{W}$ if the tmf generated from the wage offer distribution has a slope less than 1. If this restriction is not satisfied, and if the

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[1] There is little evidence about actual distribution of wages in labor markets and how they change through time.

discount factor is large enough, a type 2 improvement will lower $\bar{W}$. The expected duration of unemployment always decreases with a type 2 improvement when the discount factor is strictly positive.

Suppose now that a change in the demand for labor conditions is not predicted by the worker. In this case the worker will not alter his or her reservation wage when a change occurs. This implies a type 1 improvement will reduce the expected duration of unemployment and leave the expected post-unemployment wage unchanged. A type 2 improvement that is not predicted will reduce the expected duration of unemployment, but has, in general, an indeterminate effect on the expected post-unemployment wage. However, if the tmf of the distribution of wage offers has a slope less than 1, then such an improvement will increase $\bar{W}$. These results have important implications for the derivation of the short-run Phillips curve.

The restriction that the tmf of the wage offer distribution has a slope less than 1 has important implications for empirical work on job search as well as for the derivation of a short-run Phillips curve. Only if this restriction is satisfied will changes in the demand for labor conditions have predictable consequences on the outcome of a worker's job search process. Further, if this restriction is satisfied it is straightforward to construct downward sloping short-run Phillips curves, whereas the short-run Phillips curve may be upward sloping if the wage offer distribution does not satisfy this restriction. Before describing the model used some properties of truncated mean functions will be presented.

2. TRUNCATED MEAN FUNCTIONS OF A CLASS OF DISTRIBUTIONS

Let $\{F^{\mu}(\cdot)\}_{\mu\varepsilon R^1}$ denote a class of cdfs identical up to a translation of location, i.e., for given z, $F^{\mu}(z) = F^{\mu'}(z-(\mu-\mu'))$ for any μ and μ' in R^1. For notational convenience let $H(z) = F^{\mu}(z)$ for any z when $\mu = 0$. In this case for any particular z we have

$$F^{\mu}(z) = H(z-\mu) \tag{2}$$

for any $\mu\varepsilon R^1$. For given $\mu\varepsilon R^1$, let $V(\cdot,\mu)$ denote the tmf when the cdf if $F^{\mu}(\cdot)$. From (1) and (2) it follows that

$$V(z,\mu) = \frac{\int_z^{\infty} wdF^{\mu}(w)}{\left(1 - F^{\mu}(z)\right)} = \frac{\int_{z-\mu}^{\infty} wdH(w)}{\left(1 - H(z-\mu)\right)} + \mu \tag{3}$$

$$= \mu + V(z-\mu,0).$$

Clearly, $V(\cdot,\mu)$ is a unique transformation of $F^{\mu}(\cdot)$ for any given $\mu\varepsilon R^1$. This fact and (3) imply that the restrictions placed on $H(\cdot)$ alone will determine on the properties of $V(\cdot,\mu)$ for all $\mu\varepsilon R^1$. To simplify the exposition it will be assumed that $H(\cdot)$ is strictly increasing and differentiable on the real line. Taking the partial derivative of (3) with respect to z yields

$$V_1(z,\mu) = V_1(z-\mu,0) = \frac{h(z-\mu)}{\left(1-H(z-\mu)\right)} - \left[V(z-\mu,0)-(z-\mu)\right] \tag{4}$$

where $h(z-\mu) = H'(z-\mu)$. Since $H(\cdot)$ is strictly increasing, $V(z-\mu,0) > z-\mu$, and thus $V_1(z,\mu)$ for any z and $\mu\varepsilon R^1$, i.e., and increase in the truncation-point, z, increases the conditional expectation. Further, from (2), (3), and (4) we obtain

$$V_2(z,\mu) = 1 - V(z,\mu) = 1 - V_1(z-\mu,0), \tag{5}$$

where $V_2(\cdot,\mu) = \partial V(z,\mu)/\partial\mu$. Thus, a small shift to the right in the location of the cdf will increase the conditional expectation if and only if $V_1(z,\mu) < 1$. If $H(\cdot)$ is assumed to be a particular type of cdf, then it is relatively easy to check if the slope of its tmf is less than 1 everywhere. For example, if $H(\cdot)$ is assumed to be a Normal cdf or a Logistic cdf, then the implied $V_1(x,0) < 1$ for all x. However, if $H(\cdot)$ is assumed to be a Student cdf, then $V_1(x,0) > 1$ for some x.

Recently, Goldberger (1980) has demonstrated the following claim.

<u>Claim 1</u>: If $h(\cdot)$ is strictly logconcave, i.e., $\log H'(\cdot)$ is strictly concave, then the implied $V_1(x,0) < 1$ for all $x \in R^1$.

The above result provides a simple-to-check sufficient condition for the restriction used. If the probability density function of a particular cdf is not logconcave, the slope of the tmf has to be calculated to determine if it is less than 1 everywhere. The economic model will now be presented.

3. <u>THE MODEL</u>

In the first part of this section a simple infinite life job search model is briefly outlined. Consider an unemployed worker looking for a job in a labor market. Each period the worker visits a firm and enquires if there is a job. Let λ denote the probability any firm contacted has a vacancy. If the worker obtains a job offer, assume the wage rate offered is a random draw from cdf $F(\cdot)$ for some fixed $\mu\varepsilon R^1$. Hence, $\lambda(1-F^{\mu}(z))$ indicates the probability the worker obtains an offer with a wage rate at least as great as z in a period. The worker's expected discounted lifetime income given an offer with wage rate w' is accepted is assumed to be

$$w'\sum_{i=1}^{\infty} (1+r)^{-i} = w'/r,$$

where r denotes the discount factor.

Let $\psi(z,\lambda,\mu)$ denote the unemployed worker's expected discounted lifetime income (net of job search costs) when the worker accepts the first job offer with a wage rate at least as great as z, i.e., if z is used as the reservation wage. It follows that

$$\psi(z,\lambda,\mu) = \frac{1}{(1+r)} \Big[u-c+\lambda\int_z^{\infty} (w/r)dF^{\mu}(w) + \{1-\lambda(1-F^{\mu}(z))\}\psi(z,\lambda,\mu)\Big], \tag{6}$$

where u-c indicates the unemployment compensation per period minus the cost of visiting a firm. Given the worker desires to maximize expected discounted income it is well known that reservation wage R will be used where $\psi(R,\lambda,\mu) = R/r$ (see Lippman and McCall, 1976). This fact, (2) and (6) imply

$$R = u-c + (\lambda/r)\int_{R-\mu}^{\infty}(w+\mu-R)dH(w). \tag{7}$$

R will be termed the optimal reservation wage conditional on the parameters λ and μ.

Within the context of the model developed, a type 1 change in the demand for labor conditions will be reflected in a change in λ, whereas a type 2 change will be reflected by a change in μ. Using (4) and (7) we have

$$\frac{dR}{d\lambda} = \frac{1}{[r+\lambda(1-H(R-\mu))]} \int_{R-\mu}(w+\mu-R)dH(w) \tag{8}$$

$$= \frac{(1 - H(R-\mu))^2}{h(R-\mu)[r+\lambda(1-H(R-\mu))]} V_1(R-\mu,0) > 0$$

and

$$\frac{dR}{d\mu} = \frac{\lambda(1 - H(r-\mu))}{[r+\lambda(1-H(R-\mu))]} > 0. \tag{9}$$

Thus, an improvement in the demand conditions of either type will increase the worker's optimal reservation wage when it is fully predicted. Note that (9) implies that $dR/d\mu < 1$, if the discount rate is strictly positive. This fact, in turn, implies that if all wage offers in the market are increased by 1, then the worker's optimal reservation wage will increase by less than 1 when $r > 0$. Although the above results are of theoretical interest they are not of much use in themselves for empirical work as workers' reservation wages are usually not directly observable. Nevertheless, researchers can often ascertain how long a worker was unemployed and what was the worker's post-unemployment wage. Consequently, the effect of a change in labor market conditions on D and $\bar{W}$ will be analyzed.

To simplify the exposition the escape from unemployment probability, E, will be considered instead of the expected duration of unemployment, D. This is possible as $D = 1/E$, and hence

$$dD/dX \gtrless 0 \text{ as } dE/dX \gtrless 0 \qquad (10)$$

for any variable X. The probability of finding an acceptable wage offer in a period can be written as

$$E = \lambda\left(1 - H(R-\mu)\right) \qquad (11)$$

where R satisfies (7). From (3) and (7) it follows that the expected post-unemployment wage can be written as

$$\bar{W} = \mu + V(R-\mu,0), \qquad (12)$$

where R satisfies (7). Using (11) and (12) yields

$$\frac{dE}{di} = \begin{cases} \left(1-H(R-\mu)\right) - \lambda h(R-\mu)dR/d\lambda, & \text{if } i = \lambda \\ \lambda h(R-\mu)\left[1-dR/d\mu\right], & \text{if } i = \mu. \end{cases} \qquad (13)$$

$$\frac{d\bar{W}}{di} = \begin{cases} V_1(R-\mu-0)dR/d\lambda, & \text{if } i = \lambda \\ 1-V_1(R-\mu,0)\left[1-dR/d\mu\right], & \text{if } i = \mu. \end{cases} \qquad (14)$$

If there is an unpredicted change in the demand for labor conditions the worker will not change his or her reservation wage. This fact, (10), (13), and (14) allow us to make the following claim.

Claim 2: (i) An unpredicted type 1 improvement in labor market conditions implies (a) the expected duration of unemployment will decrease, and (b) the expected post-unemployment wage will remain unchanged.

(ii) An unpredicted type 2 improvement in labor market conditions implies (a) the expected duration of unemployment will decrease, and (b) the expected post-unemployment wage will increase (decrease) if $V_1(R-\mu,0)<1$ (if $V_1(R-\mu,0) > 1$).

As the above claims follow immediately from inspection of (10), (13), and (14) no proof will be presented.

In this final part of the study it will be assumed that any change in labor market conditions is fully predicted by unemployed workers. First, suppose there is a change in λ when all other parameters are held constant. From (8), (9), (13), and (14) we have[1]

$$\frac{dE}{d\lambda} = \frac{\lambda(1-H(R-\mu))^2}{[r+\lambda(1-H(R-\mu))]}\left[\frac{r}{\lambda(1-H(R-\mu))} + (1-V_1(R-\mu,0))\right] \tag{15}$$

and

$$\frac{d\bar{W}}{d\lambda} = V_1(R-\mu,0)\,\frac{\lambda(1-H(R-\mu))}{[r+\lambda(1-H(R-\mu))]}. \tag{16}$$

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[1] In Barron's model the discount factor is assumed to be zero. Thus (15) reduces to $dE/d\lambda = (1-H(R-\mu))(V_1(R-\mu,0)-1)$ when r=0. In this special case $dD/d\lambda < 0$, if and only if $V_1(R-\mu,0) < 1$.

Inspection of (10), (15), and (16) allows us to state the following claims without any formal proof.

Claim 3: (a) $d\bar{W}/d\lambda > 0$

(b) $dD/d\lambda < 0$, if $V_1(R-\mu,0) < 1$

(c) $dD/d\lambda > 0$ for some $r > 0$, if $V_1(R-\mu,0) < 1$

Claim 3 implies that the consequences of a type 1 improvement can be predicted fully only if the distribution of wage offers is such that the tmf has a slope less than 1. In this case a type 1 improvement which is fully predicted, will reduce the expected duration of unemployment and increase the expected post-unemployment wage. If the distribution of wage offers is such that the tmf has a slope greater than 1 in places, then a type 1 improvement will always increase the expected post-unemployment wage, but may increase the expected duration of unemployment if the discount factor is small enough.

Finally, suppose there is a shift in the cdf of wage offers that is fully anticipated by the unemployed workers, i.e., a predicted type 2 change in labor market conditions. From (9), (13), and (14) it follows that

$$\frac{dE}{d\mu} = \lambda h(R-\mu) \frac{r}{[r+\lambda(1-H(R-\mu))]} \tag{17}$$

and

$$\frac{d\bar{W}}{d\mu} = 1-V_1(R-\mu,0) \frac{r}{[r+\lambda(1-H(R-\mu))]}. \tag{18}$$

Using (10), (17), and (18) the following claims can be stated without formal proof.

<u>Claim 4</u>: (a) $dD/d\mu < 0$.

(b) $d\bar{W}/d\mu > 0$, if $V_1(R-\mu,0) < 1$

(c) $d\bar{W}/d\mu < 0$, for large enough r, if $V_1(R-\mu,0) > 1$

Claim 4 establishes that if the tmf of the distribution of wage offers has a slope less than one everywhere, then a type 2 improvement will reduce the expected duration of unemployment and increase the expected post-unemployment wage. Although the expected duration of unemployment will always fall with a fully anticipated type 2 improvement, the expected post-unemployment wage cannot be guaranteed to increase with such an improvement if $V_1(R-\mu,0) < 1$. If $V_1(R-\mu,0) > 1$, then a type 2 improvement can reduce the expected post-unemployment wage if the discount factor r is large enough.

From the above analysis it can be concluded that if the distribution of wage offers is such that its tmf has a slope less than 1 everywhere, then there are predictable consequences to a type 1 or type 2 change in labor market conditions independent of the discount factor used by the unemployed workers. When the distribution of wage offers is such that its tmf has a slope greater than 1 in places sign predictions cannot be made without knowledge of the discount rate.

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The Structure and Dynamics of Unemployment: Sweden and the United States

Anders Björklund and Bertil Holmlund

1. INTRODUCTION*

A characteristic feature of modern theories of labor market behavior is their focus on the dynamic nature of unemployment. Various studies have documented the importance of high turnover between labor force states and emphasized the need to explain these transitions in order to understand unemployment patterns.

The characteristics and determinants of unemployment inflow and outflow have received special attention in this context. One interesting hypothesis - suggested by job-search theory - claims that unemployment fluctuations are explainable by inflationary surprises. Unemployment then is viewed as a productive investment in job search, chosen by employees to enhance their lifetime earnings. An increase in aggregate demand will imply a temporary fall in unemployment due to short-run deviations between actual and expected wages; workers are fooled into accepting more employment.

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* This paper includes as a sub-set an abbreviated version of another paper of ours, Björklund and Holmlund (1981). We are indebted to Ned Gramlich and Mats Persson for helpful comments on earlier versions.

This information-lag interpretation of changes in unemployment might be compared to an alternative view, where the quantity-rationing rules of the labor market are emphasized. An increasing flow of labor from unemployment to employment is, according to this theory, caused by the relaxation of job-rationing constraints rather than unanticipated inflation.

In this paper we adopt the search-turnover framework as a vehicle for exploring unemployment patterns in Sweden and the U.S. The study has two basic objectives. The first, essentially descriptive, is to characterize the structure of unemployment in Sweden and the U.S. Section 2, therefore, involves a decomposition of the unemployment rate into components such as (i) the average duration of unemployment spells, (ii) the relative number of individuals experiencing unemployment spells and (iii) the average number of spells per unemployed worker. We also investigate possible sources of cyclical unemployment fluctuations in the two countries. Are these primarily associated with variations in the duration component or with fluctuations in the number of spells? The calculations included in Section 2 reveal certain significant U.S.-Swedish differences with respect to structural characteristics as well as sources of unemployment cycles.

In Sections 3 and 4 we present econometric tests of the two competitive explanations of unemployment duration fluctuations, i.e., the detection-lag hypothesis and the job-rationing view. The two stories are, of course, not mutually exclusive; we try, via a fairly simple specification, to capture both views in one equation. The principal contribution of the analysis lies in its ability to provide information about the relative importance of unexpected inflation and job opportunities as explanations of the duration of unemployment. We apply the same model to both Swedish and U.S. data, thereby revealing important differences between the labor markets in the two countries. We find, perhaps somewhat surprisingly, that the U.S. unemployment duration is more or less unaffected by unexpected inflation, whereas the results for Sweden give some support for the information-lag hypothesis. Another finding is that the simple information-lag story is more valid for the short-term unemployed.

The analysis of U.S. transition probabilities is to some extent performed on data disaggregated by demographic groups and including information on hiring probabilities (transitions from unemployment to employment) as well as labor force exit rates (mobility between unemployment and the not-in-the-labor force state). A noteworthy result is that the exit probabilities display strikingly <u>pro-cyclical</u> behavior.

2. THE STRUCTURE AND CYCLICAL FLUCTUATIONS OF UNEMPLOYMENT IN SWEDEN AND THE U.S.

A decomposition of the unemployment rate according to the familiar formula

$$u = f \cdot D \tag{1}$$

is a natural starting point to discuss the cyclical fluctuations in unemployment. This relation -

where u is the unemployment rate, f the rate of inflow to unemployment and D the average duration of completed spells of unemployment - holds exactly only in a stationary state with constant inflow and "survival rates" of the unemployed. The bias will, however, be of minor importance when yearly data are considered.[1]

It is likely that the two components represent different economic mechanisms. If the cyclical fluctuations can be attributed to the inflow component, a theory of quits, layoffs and labor force entrances would be called for. Such a theory may be quite different from a theory of cyclical fluctuations of the duration of unemployment.

The basic patterns for Sweden and the U.S. are presented in Figures 1 and 2.[2] It appears that

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[1] For Swedish data this is documented in Björklund (1979) Appendix 1.

[2] The Swedish data (ages 16-74 years) have been obtained from a measure of the weekly inflow to unemployment (Björklund (1979), Appendix 2). The U.S. data (ages above 16) are calculated from measurements of monthly probabilities of leaving unemployment, obtained from unpublished CPS-data. We are grateful to Ralph Smith and Jean Vanski at the Urban Institute for making available these series. The duration has been computed from the relation D = 1/P, where P is the probability of leaving unemployment. This D has been multipled by 4.3 to convert the measure from months to weeks. According to Clark and Summers (1978), footnote 5, this measure gives results which are consistent with measures based on inflow data. The probability of an upward bias should be recognized since those who leave unemployment are doing this sometimes during the month and not at the end of the month. A natural approximation would be to assume that the average person leave unemployment at the middle of the month and consequently subtract 2.15 from the measure. Relying on Clark and Summer's observations this change has <u>not</u> been undertaken.

Figure 1. The unemployment rate (left axis) and the unemployment inflow rate (right axis); Sweden and the U.S. Percent.
—— unemployment rate
- - - - inflow rate

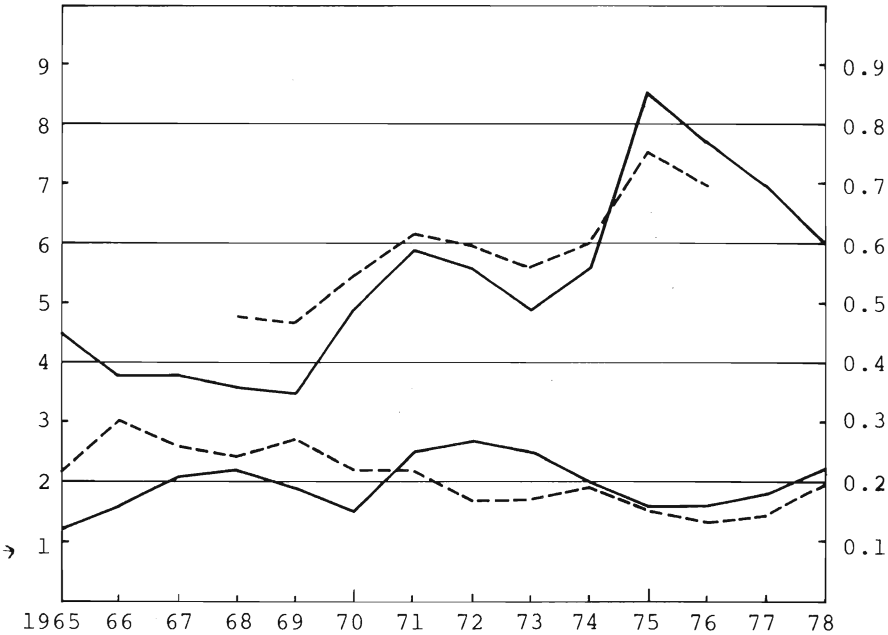
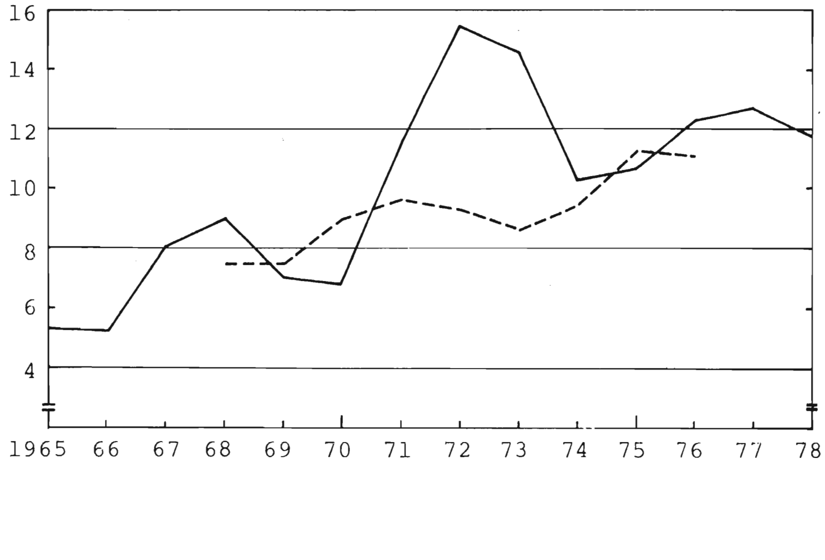


Figure 2. The average duration of the spells of unemployment (weeks): Sweden (——) and the U.S. (- - -).



most of the cyclical fluctuations of Swedish unemployment can be attributed to the duration-component. For the U.S., on the other hand, unemployment inflow as well as unemployment duration displays a marked cyclical pattern.

The figures also show some other interesting differences between the two countries. Although the unemployment rate is almost three times higher in the U.S. than in Sweden, the average duration of a completed spell of unemployment is lower in the U.S. The U.S. inflow rate is consequently much higher. Furthermore, a trend towards lower inflow rate can be found in Sweden whereas the U.S. inflow has increased over time. In both countries duration has increased according to trend.

As is well known, the duration of unemployment equals the inverse of the probability of leaving unemployment. This transition probability, in turn, equals the sum of (i) the job-finding probability (Pue) and (ii) the labor force exit probability (Pun). Data are available for the U.S. to make such a decomposition possible. These are presented in Figure 3.[1]

The figure reveals that the probability of leaving the labor force - conditional upon being unemployed - is only slightly lower than the probability of entering employment, which no doubt indicates that a comprehensive explanation of the duration of unemployment should take account of participation behavior. The development of the two curves is, however, strikingly similar. This is

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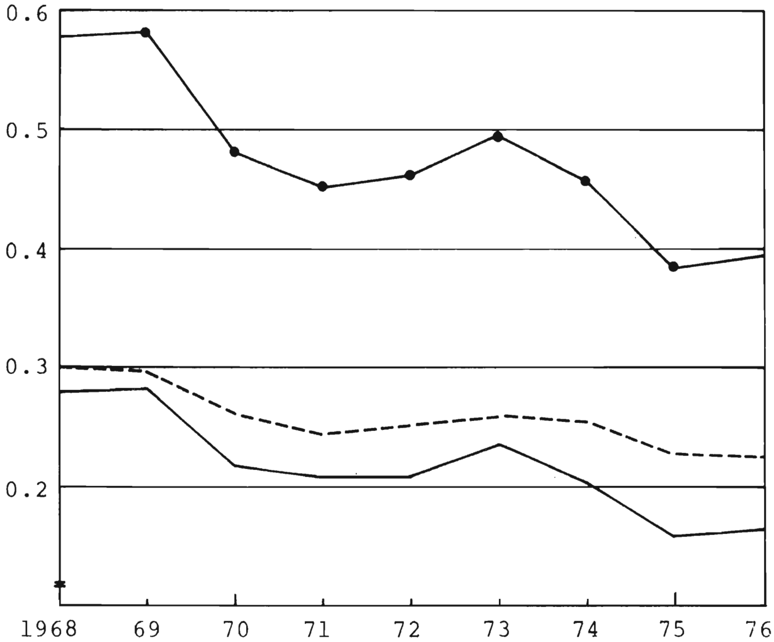
[1] Unpublished CPS data mentioned in footnote above.

Figure 3. <u>Components of U.S. unemployment outflow.</u> Yearly averages of monthly transition rates.



●——● P = the probability of leaving unemployment

---- Pue = the probability of transitions from unemployment (u) to employment (e)

—— Pun = the prob ability of transitions from unemployment (u) to the not-in-the-labor force state (n) somewhat surprising since it could be expected that the probability of leaving the labor force would be high when the probability of entering employment is low.

The decomposition of the unemployment rate into an inflow and a duration component is, of course, not the only one which is possible. It might be argued that a more welfare-relevant decomposition should take account of the number of individuals who are hit by unemployment and the total length of unemployment per unemployed individual. Such a decomposition is possible by using information from the retrospective labor force surveys about the number of individuals who have been unemployed any time during a year.[1]

The following relationship holds:

$$52 \cdot U = N \cdot W \tag{2}$$

where

U = the average stock of unemployed during a year,
N = the number of people who have been unemployed any time during a year,
W = the average number of unemployment weeks per unemployed individual and year.

What expression (2) says is that 52·U weeks of unemployment during a year are distributed among N persons who are unemployed W weeks each on average.

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[1] These retrospective surveys are performed in the beginning of each year in both countries. Swedish data are published by the National Labor Market Board and the U.S. figures are from various issues of Monthly Labor Review.

The latter component, in turn, can be decomposed further, since the W weeks of unemployment is the result of a certain number of unemployment spells (S) and the average duration of the spells (D). Making this decomposition and dividing U and N by the labor force, L, we obtain:

$$52 \cdot u = \frac{N}{L} \cdot W = \frac{N}{L} \cdot S \cdot D \qquad (3)$$

These components are presented in Figures 4 and 5, respectively.[1] In the Swedish case it now appears that most of the cyclical fluctuations of unemployment can be attributed to the number of persons who are hit by unemployment (N/L). The fluctuations of the total length of unemployment (W) are much smaller since S and D counteract each other. In the U.S. case, the data reveal that both components (N/L and W) have a marked cyclical pattern.

Summarizing this empirical analysis the characteristic country-differences should be observed. The basic message includes the following points.

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[1] Here we combine two types of unemployment data, namely the retrospective data and the regular monthly survey data. These might be inconsistent. However, the inconsistency is almost negligible in the Swedish case (Björklund (1979), p.19). A crude test by Morgenstern and Barrett (1974) indicates that the inconsistency might be more important for the U.S. However, as long as the inconsistency over time is constant only the levels of the components are affected whereas the development is unaffected. The relation between (1) and (3) is clear by noting that S equals (52•F/N) where F is the weekly inflow of spells of unemployment. Actually, S can be lower than one since 52•F is the number of spells which have been started during the year whereas some of the N persons might have become unemployed during the preceding year (see Björklund, 1979, Appendix 3, for a more detailed analysis).

Figure 4. The percentage of workers with unemployment experiences during a year (N/L) and rates of unemployment (u); Sweden and the U.S.

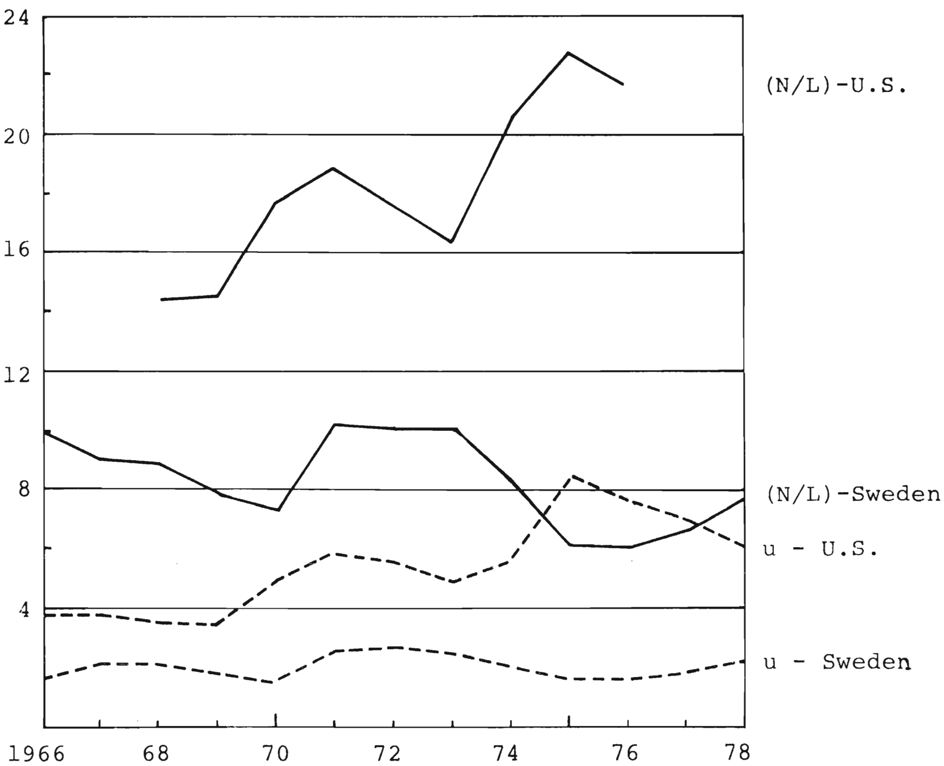


Figure 5A. The average number of yearly unemployment weeks per worker with unemployment experiences during a year (W); Sweden and the U.S.

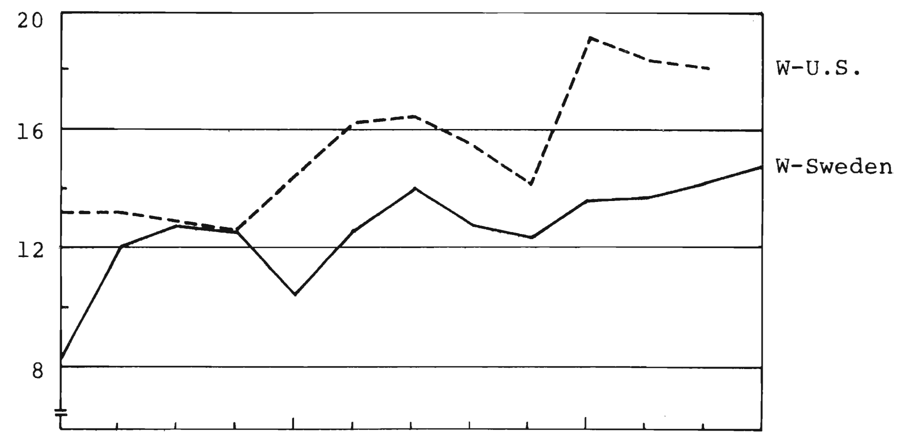
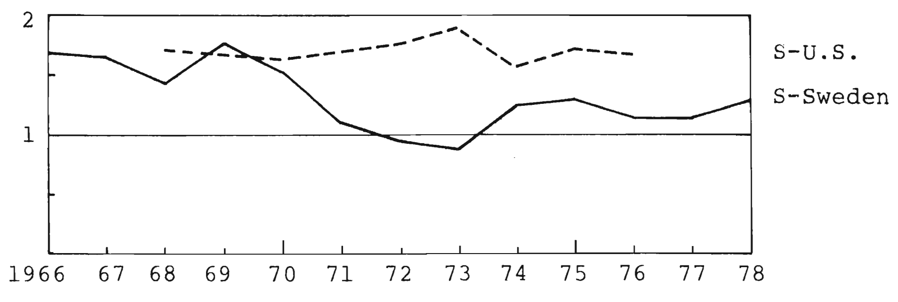


Figure 5B. The average number of unemployment spells per worker with unemployment experiences during a year (S); Sweden and the U.S.



(i) The much higher U.S. unemployment rate cannot be attributed to longer duration of unemployment spells. In fact, we find that U.S.-Swedish duration-differentials are fairly small (less than 2 weeks on average for the period 1968-76).[1]

(ii) Taking account of the frequency of unemployment spells (for unemployed workers) we observe, furthermore, a larger number of (yearly) unemployment weeks for an average U.S. unemployed. The U.S. "unemployment burden", so defined, has amounted to 15-16 weeks for the period 1966-77 whereas the corresponding Swedish figures have been 12-13 weeks.

(iii) The higher U.S. unemployment rate is to an important degree associated with higher U.S. probabilities of being hit by unemployment. The relative number of persons with unemployment experiences is approximately twice as high in the U.S.

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[1] The well-informed reader may wonder how our results are related to an earlier controversy on these issues. (See the discussion in AER by Flanagan, 1973, Barrett, 1975 and 1977, and Axelsson, Holmlund and Löfgren, 1977). A basic difference between our computations and those presented in the AER-papers mentioned is that we use the best data sources available, data which, moreover, are comparable (see footnote 2, p.186). Our analysis is also more detailed w.r.t. decompositions of the unemployment rate. The results presented here are qualitatively consistent with Barrett's observation concerning duration differentials, although they basically confirm the point made by Flanagan and AHL: these differences are fairly negligible. This conclusion is given additional strength by an analysis presented in Björklund (1979, Appendix 2), indicating that Swedish unemployment durations in 1972 and 1973 might be lower than the figures obtained from the method used here.

An additional interesting observation, finally, regards the sources of cyclical unemployment fluctuations. We have seen that Swedish unemployment duration can account for most of these variations; the unemployment inflow rate displays a significant negative trend but only irregular cyclical patterns. The U.S. inflow, on the other hand, is closely related to fluctuations in the unemployment rate; variations in the latter are reinforced by similar developments of inflow and unemployment duration.

It is far beyond the scope of this paper to assess the welfare implications of the inter-country differentials w.r.t. the structure and fluctuations of unemployment. The high U.S. inflow rate is, probably, to an important degree attributable to the frequency of layoffs. This will also briefly be touched upon in the concluding section of the paper. Next, however, we turn to possible explanations of the cyclical fluctuations of unemployment. This analysis will focus on the duration of <u>spells</u> of unemployment and a simple job search model will be used as our theoretical framework. We are of course aware of the fact that other characteristics of unemployment, such as fluctuations in unemployment inflow and the number of spells per unemployed, are not easily explained by available versions of job search theory.

3. DETERMINANTS OF UNEMPLOYMENT DURATION FLUCTUATIONS: THE SEARCH-THEORETIC VIEW

Consider the behavior of an unemployed worker according to the standard search model. His problem is to choose an acceptance wage which assures him an income greater than what he might have received by continued search. The decision is affected by the perceived location of the wage offer distribution. If a monetary contraction produces a leftward shift of the wage offer distribution - or a lower rate of wage inflation - this change in general market conditions is assumed to be imperfectly detected by job seekers, who mistakenly blame local circumstances rather than changes in aggregate demand. Unemployed workers will search for a longer time causing the length of spells of unemployment to rise.

A common assumption in standard search models is that the number of job offers received per period equals one. The probability of leaving unemployment - the transition probability - is then solely determined by the job seeker's offer-acceptance probability. The simplifying job offer assumption is, however, not inherent in search theory *per se*; by a modest generalization the case with a random number of job offers is easily incorporated into the basic search theoretic framework. Consider the job-seeker's transition probability, which - in the absence of labor force exits - equals the hiring probability. Decomposing the transition probability into two components, the *job offer probability* and the *acceptance probability*[1] we have

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[1] Note that the job offer probability is the product of the probability of finding a vacancy and the probability of being offered a job, conditional upon having found one.

$$P = \Theta[1-F(w^*)] \qquad 0 \leq \Theta \leq 1 \tag{4}$$

where w* is the reservation wage and $F(\cdot)$ the distribution function of wage offers. If the transition probability is constant during search, the expected duration of unemployment (D) is

$$D = 1/P = 1/\Theta[1-F(w^*)] \tag{5}$$

Which are then the characteristics of an optimal search policy? In the simple case of infinite time horizon and discount rate r, the optimal policy implies a certain time invariant reservation wage obtained as the solution to

$$c + w^* = \frac{\Theta}{r} \int_{w^*}^{\infty} (w-w^*)f(w)dw \tag{6}$$

where c is the (constant) marginal search cost and $f(\cdot)$ the known density function of wage offers. Eq. (6) implies that the reservation wage declines as the job offer probability decreases. Likewise, a known leftward shift of the wage offer distribution will also reduce the reservation wage.

We have so far briefly outlined the basic search story, strictly valid only in a stationary world. Now consider the possibility of fluctuations in aggregate demand, influencing the job-seeker's transition probability via the job offer probability (more vacancies) and/or via imperfect reservation wage adjustments. Three different effects may be identified:

(i) The pure availability effect: An increasing number of vacancies means a higher job offer probability, thereby reducing the duration of unemployment.

(ii) The supply effect: A permanent increase of the job offer probability will increase the expected returns from search, thus increasing the worker's reservation wage. It follows that the unemployment effect of a rising number of vacancies is ambiguous *a priori*.

(iii) The detection-lag effect: Changes in aggregate demand will affect the location of the wage offer distribution. Assuming a lag in the discernment of a rising rate of inflation, reservation wages will be unaffected in the short run, implying a rising flow of new hires from the pool of unemployed.

Summarizing these three effects we have:

$$P = \underset{+}{\Theta(V)}A(\underset{-}{V},\underset{+}{w/w^e}) = g(V,w/w^e) \tag{7}$$

where V is the number of vacancies, A the acceptance probability, w the actual average wage and w^e the expected average wage.

We would argue that Eq.(7) represents the kernel of the search theory of cyclical unemployment. The standard search model outlined does rely on some very restrictive assumptions, e.g., a stationary wage offer distribution, fixed leisure time and a constant job offer probability. More complex search models, e.g., those of Siven (1979) and Seater (1977, 1978, 1979) are, however, fairly consistent with the simple search model in their emphasis on unexpected inflation and vacancy contacts. We are suppressing other plausible determinants of unemployment duration, e.g., variations in unemployment compensation and the discount rate.

4. EMPIRICAL ANALYSIS

4.1 The Basic Model

A straightforward method of investigating the validity of the detection-lag hypothesis is to specify explicit transition probability equations with vacancies and unexpected wage increases as explanatory variables,[1] i.e., represent Eq.(7) above by a suitable functional form.[2] The basic specification used will be:[3]

$$\ell nP_t = \alpha_1 + \alpha_2 \ell nV_t + \alpha_3 \ell n(w_t/w_t^e) \qquad (8)$$

The obtained α_2-estimate reflects the net result of the positive availability effect and the negative supply effect; intuition and some theoretical predictions suggest that α_2 (the net availability effect) will have a positive sign.[4]

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[1] We have also excluded price changes from consideration. For a discussion of this issue see Björklund and Holmlund (1981).

[2] For a different approach, see Barron (1975), and Axelsson and Löfgren (1977).

[3] Logit models have also been tried with qualitatively similar results.

[4] Feinberg (1977).

The main problem with the approach chosen is, of course, that it requires an analysis of perceived as well as actual wages. Since no direct data about expected wages or wage-changes are available some model of the formation of expectations must be used. Our approach has been to try three different models in order to investigate how robust the information-lag hypothesis is with respect to the different specifications. Two of the applied forecasting functions are consistent with the idea that workers learn from past errors, reestimating the parameters of their forecasting equations when more information is obtained.

The first model postulates <u>adaptive expectations</u> and assumes that expectations are formed according to a finite distributed lag of past wage changes, i.e., with quarterly data (which is used for Sweden):

$$\frac{w_t^e}{w_{t-4}} = \sum_{i=1}^{4} \ell_i \frac{w_{t-i}}{w_{t-4-i}} \tag{9a}$$

where

$$\sum_{i=1}^{4} \ell_i = \frac{1}{10} \sum_{i=1} (5-i) = 1 \tag{9b}$$

and with monthly data (which is used for the U.S.)

$$\left(\frac{w_t^e}{w_{t-12}}\right) = \sum_{i=1}^{12} \ell_i \left(\frac{w_{t-i}}{w_{t-12-i}}\right) \tag{10a}$$

where

$$\sum_{i=1}^{12} \ell_i = \frac{1}{78} \sum_{i=1}^{12} (13-i) = 1 \tag{10b}$$

Models like these - where the sum of the weights has been constrained to one - are often used in empirical work even though it has been pointed out that the theoretical basis is quite weak (see, e.g., Persson (1979), where it is shown that the sum should equal one only in very special cases if the forecast is to be optimal).

Even though the simplicity of the adaptive model is appealing it could be argued that individuals have some knowledge about historical regularities of wage changes, and that they use this information when forming their expectations. One possible way to represent these regularities is to apply a time-series approach. The assumption is that people have in their mind an auto-regressive moving average-process (ARMA) which is generating forecasts from period to period. Both the specification and the parameters of this process are, however, likely to be revised when people receive more information about wage-changes. The process was therefore reestimated each period and reidentified each fourth period (with quarterly data) and each twelfth period (with monthly data).[1]

Finally, it could be argued that workers are still more rational than using information only from an ARMA process of wage changes. They might even have in mind an empirical model incorporating different economic variables. An unemployed worker forming his expectations may, e.g., use a wage-equation of the Phillips curve type. We have therefore also estimated such wage-equations, reestimating them each fourth period and with data from the last

[1] More detailed information on these excercises are given in Björklund and Holmlund (1981).

five years. On the whole the estimated equations performed reasonably well for Sweden according to standard statistical criteria, but were less successful for U.S.

In the following sections we report results only for adaptive expectations (Sweden) and ARMA expectations (U.S). More detailed results are given in Björklund and Holmlund (1981).

4.2 The Data

Swedish transition probabilities have been estimated as follows. The rotating system of the Swedish Labor Force Surveys is constructed so that almost 90% of those who are interviewed in one survey are interviewed again three months later, whereas different individuals are interviewed in two subsequent months. This property of the surveys can be used to compute quarterly transition probabilities.

Denoting the number of unemployed for at least a weeks but less than b weeks at time t by $G_t^{a,b}$ and the weekly inflow into unemployment by f we can describe the estimates as follows.

$$G_t^{1,14} = f \sum_{i=0}^{12} (1-P_1)^i \tag{11}$$

$$G_{t+13}^{14,27} = G_t^{1,14}(1-P_2)^{13} \tag{12}$$

$$G_{t+13}^{27,39} = G_t^{14,27}(1-P_3)^{13} \tag{13}$$

Three transition probabilities are obtained - P_1, P_2 and P_3 - which can be regarded as conditional upon the length of the spell of unemployment. By using available data on f and G we obtain P_1 from

$$\frac{(1-P_1)^{13}-1}{(1-P_1)-1} = \frac{G_t^{1,14}}{f} \tag{14}$$

whereas P_2 and P_3 are calculated as

$$P_2 = 1 - \left(\frac{G_{t+13}^{14,27}}{G_t^{1,14}}\right)^{1/13} \tag{15}$$

$$P_3 = 1 - \left(\frac{G_{t+13}^{27,39}}{G_t^{14,27}}\right)^{1/13} \tag{16}$$

The Swedish vacancy statistics are from labor market statistics, published by the National Labor Market Board. Quarterly wage data are obtained from the labor market issues of Statistical Reports, published by the Swedish Bureau of Statistics. All data used refer to manufacturing industry.

The U.S transition probabilities are of two kinds. In the first place, we used the method proposed by Barron (1975). The essential idea is the same as above, namely to compare the number of people in one week who have been unemployed less than five weeks with the number of people four weeks later who have been unemployed five to eight weeks. The difference consists of people who have left the pool of unemployed. The duration data reported in

Employment and Earnings are grouped in the classes 1-4 weeks, 5-14 weeks, etc., which requires a slight modification of the method outlined above; for details, see Barron (1975). The other type of data used is described below (Section 4.5).

The U.S wage data are average hourly earnings in manufacturing industry, reported in Employment and Earnings. As vacancy data for the period 1965-75 we used the Help-wanted advertising index (HWA) published in Main Economic Indicators (OECD). For the period 1969.4-1973.10 manufacturing vacancies (Vm) according to establishment data were also tried (Employment and Earnings); the latter series are available only for (approximately) this period.

4.3 Empirical Results - Sweden

The results from the estimations on Swedish data are presented in Table 1 below. The estimation method is weighted-least-squares.[1] We observe, in the first place, that the detection-lag variable is significant both for the short-term unemployed (1-13 weeks) and for the medium-term unemployed (14-26 weeks). These results hold for all models of expectations. For the long-term unemployed, on the other hand, no significant detection-lag effect is revealed; the coefficient even has the wrong sign.

The job availability variable (V) is significantly positive in all regressions, even for the long-

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[1] Derivations of the weights are available on request.

Table 1. Transition probability equations for Sweden

Quarterly data 1968.1-1977.3

Adaptive expectations	V	w/w^e	R^2	DW
Short-term unemployed (P_1)				
(1)	0.81 (4.29)	10.30 (4.10)	0.60	1.75
(2)	1.11 (5.36)	-	0.42	1.57
(3)	-	14.51 (5.18)	0.41	1.05
Medium-term unemployed (P_2)				
(4)	0.34 (3.24)	1.97 (1.69)	0.33	2.27
(5)	0.41 (4.11)	-	0.30	2.29
(6)	-	3.42 (2.83)	0.16	1.84
Long-term unemployed (P_3)				
(7)	0.39 (2.19)	-3.35 (-1.57)	0.09	2.16
(8)	0.31 (1.78)	-	0.05	2.28
(9)	-	-1.99 (-0.93)	0.004	2.03

term unemployed. Dropping this variable produces in most cases a marked decrease in the DW value, indicating the presence of specification errors.

What are the economic interpretations of the different results for the three groups of unemployed? No straightforward answer is available, partly because the "hypothesis-testing includes a joint test of the underlying model and the expectations-generating mechanism."[1] The absence of any significant detection-lag effect for the long-term unemployed may have at least two explanations: the expectations model might be inappropriate and/or the variable reservation wage hypothesis could be erroneous. There are arguments in favor of both these interpretations. In the first place, it makes sense to hypothesize that the long-term unemployed (more than six months in our data) are better informed about the actual wage offer distribution, simply because they have experienced a longer period of "learning" through full time job search. This argument implies that the parameters of the forecasting function might differ across workers with different unemployment histories.

The second interpretation stated above (the possible unrealism of the variable reservation wage hypothesis) may be elucidated by recalling some familiar results from search theory. The reservation wage of a job-seeker with a finite search horizon will, under some stationary conditions, fall with the duration of unemployment, a theoretical prediction which has been given empirical support.[2] Eventually the reservation wage will coin-

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[1] Santomero and Seater (1978), p.525.

[2] See Gronau (1971), Kasper (1967), and Kiefer and Neumann (1979).

cide with the minimum value of the wage offer distribution, implying an acceptance probability equal to one. In that extreme case all job offers are accepted and there is no detection-lag effect.

Both of the hypotheses outlined are consistent with the results obtained. Intuition would suggest that both of the mechanisms are in operation to some extent, reinforcing each other and thereby producing the observed results.

Since both the (net) availability effect and the detection-lag effect are significant, it is important to discover the relative importance of these variables as determinants of the cyclical variations of the duration of unemployment. To find out this we must take the <u>size</u> of the parameters as well as the <u>variation</u> of the independent variables into account. The question might be illuminated by comparing the predicted transition probabilities using estimates from regressions in the table

$$\hat{P}_t = \alpha_1 \cdot V_t^{\alpha_2} \cdot \left(\frac{w_t}{w_t^e}\right)^{\alpha_3} \tag{17}$$

with the transition probabilities obtained when inflation is perfectly foreseen ($w_t = w_t^e$)

$$\tilde{P}_t = \alpha_1 \cdot V_t^{\alpha_2} \tag{18}$$

Using the results from the adaptive model, Figure 7 below demonstrates the relative unimportance of the detection-lag effect for the medium-term unemployed. Inflationary surprises produce, on the

Figure 6. The effects of unexpected inflation - short-term unemployed in Sweden

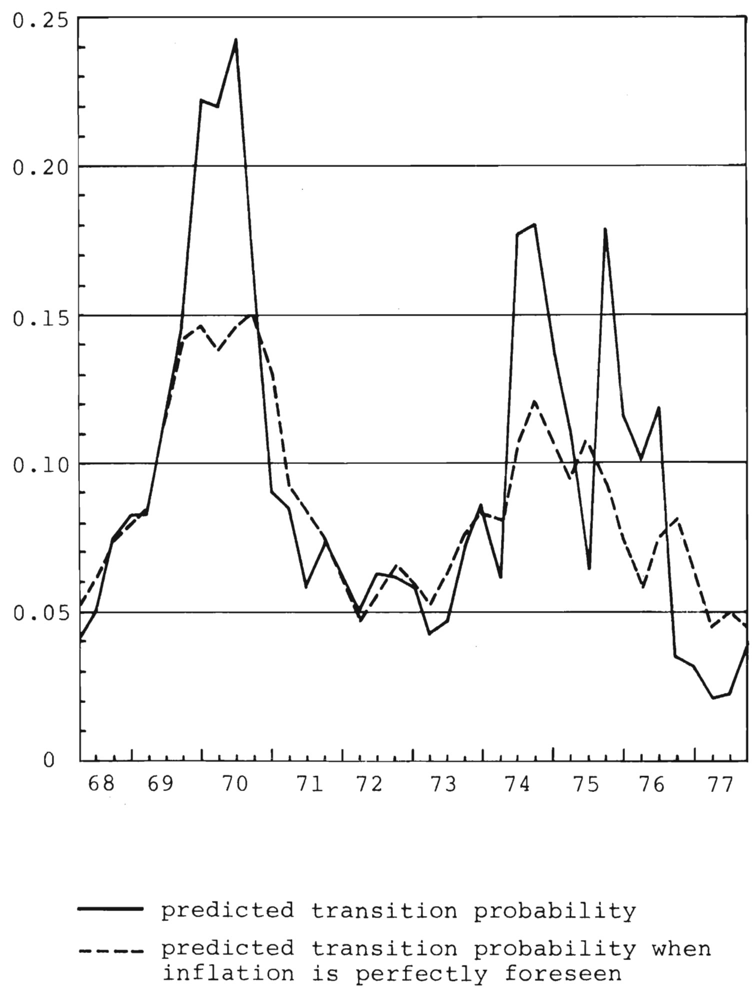

Figure 7. The effects of unexpected inflation - medium-term unemployed in Sweden

——— predicted transition probability

- - - - predicted transition probability when inflation is perfectly foreseen

other hand, quite important unemployment effects for the short-term unemployed during the peak years 1969-70 and 1974-75 (Figure 6). The main part of the variation is, however, attributable to the vacancy variable.

4.4 Empirical Results - United States

Turning now to the U.S. regressions (Table 2), the dominant availability effect is even more pronounced than in the Swedish case. The vacancy variables used are highly significant in all regressions whereas the detection-lag coefficient is fairly sensitive with respect to the choice of expectations model and estimation period. A significant detection-lag effect is obtained only by applying an ARMA-expectations-generating mechanism for the period 1969.4- 1973.10. These results

Table 2. Transition probability equations for the U.S.
Monthly data 1969.4-1973.10 and 1965.2-1975.12, respectively.

ARMA expectations	HWA	Vm	w/w^e	TIME	R^2	DW	ρ
1969.4-1973.10							
1	-	0.23 (11.27)	2.57 (2.07)	-0.0007 (-1.54)	0.74	1.15	-
2	-	0.20 (8.63)	1.96 (1.98)	-0.0002 (-0.40)	n.a.	2.03	0.30
3	-	0.24 (11.92)	2.48 (1.98)	-	0.73	1.10	-
4	-	0.20 (8.81)	1.93 (1.98)	-	n.a.	2.04	0.31
5	-	0.24 (11.64)	-	-	0.71	1.12	-
6	-	0.20 (8.68)	-	-	n.a.	2.04	0.30
7	-	-	3.03 (1.31)	-0.0021 (-2.49)	0.10	0.36	-
8	0.49 (11.23)	-	2.48 (2.00)	-0.0030 (-6.44)	0.73	1.17	-
9	0.44 (8.36)	-	1.81 (1.80)	-0.0022 (-3.96)	n.a.	1.99	0.30
10	0.44 (7.71)	-	2.23 (1.34)	-	0.53	0.65	-
1965.2-1975.12							
11	0.52 (16.7)	-	-0.37 (-0.50)	-0.0026 (-19.19)	0.83	1.33	-
12	0.53 (11.37)	-	-0.07 (-0.10)	-0.0025 (-13.01)	n.a.	2.04	0.35
13	0.49 (8.03)	-	3.45 (2.48)	-	0.35	0.41	-
14	-	-	-0.09 (-0.07)	-0.0025 (-10.31)	0.46	0.44	-

Note: ρ is the first-order autocorrelation coefficient obtained by using the Cochrane-Orcutt's method.

are independent of the choice of vacancy variable. Exclusion of the latter also gives rise to a strong decline in the DW statistic, indicating specification errors. When the estimation period is extended (1965.2-1975.12), the significance of unexpected inflation disappears.[1] It should also be noted that a negative and significant trend-coefficient is obtained when HWA is used as vacancy variable.

The main conclusion from these excercises on U.S. data is that the job availability variables are the dominant determinants of the cyclical fluctuations of unemployment duration. We cannot, however, rule out the possibility of some detection-lag effects in operation, at least during certain time periods and - especially - if expectations are formed according to an ARMA process rather than adaptively.

4.5 Hiring Probabilities Versus Labor Force Exit Probabilities

In all estimations presented so far we have used the probability of leaving unemployment as the dependent variable. Clearly, the detection-lag hypothesis is relevant for the job finding probability but has little to say about labor force exits. More decisive tests of the hypothesis should therefore focus explicitly on transitions from unemployment to employment. Fortunately, such data are available for the U.S. The monthly current popula-

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[1] The coefficient for w/w^e is significant in Eq. (13) but the DW-value indicates that the t-ratio should not be taken seriously.

tion survey (CPS) uses a rotation system, implying that the major part of the sample is common from one month to the next. Hence, it is possible to compare labor force status month t with that in month t-1 and calculate various transition rates of interest.[1] Pue, e.g., would then represent the proportion of unemployed workers last month who were employed this month. Estimation results for the group 16-59 years are set out in Table 3;[2] ARMA expectations have been postulated and the functional form is log-linear.

Consider first the hiring probability, Pue. The basic finding is that vacancies and time explain most of the variation, whereas unexpected inflation plays a negligible role. These results are, obviously, perfectly consistent with our earlier findings for the U.S. Consider, next, the exit probability, Pun. A remarkable result is the strong positive vacancy elasticity obtained; the pro-cyclical behavior of exit probabilities is even more pronounced than the corresponding pattern for hiring probabilities![3] The negative trend-coefficients in the equations for Pue and Pun, respectively, indicate that the duration of unemployment have been increasing over the sample period.

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[1] The idea is the same as the one presented in Section 4.2 but these transition probabilities have been computed from data on identical individuals.

[2] Estimates of the intercept and 11 seasonals are not presented.

[3] Exit-behavior of this type is also observed by Ralph Smith (1977).

The striking pro-cyclical exit rates are also observed for demographic groups, as revealed by estimation results in Appendix 1 and the summary in Table 4. Two interesting age-patterns can be observed: First, hiring probabilities for the prime-age group (25-59) display weaker cyclical sensitivity than hiring probabilities for young workers. Vacancy elasticities for exit probabilities, on the other hand, are increasing with age, being especially large for prime-age males.

Can the pro-cyclical exit pattern be given intelligible explanations? Two alternative interpretations might be formulated in terms of (i) aggregation effects and (ii) expected utility-maximizing job search behavior. The aggregation hypothesis, suggested by Ralph Smith (1977), assumes systematic cyclical patterns of the composition of the unemployed, even within the demographic groups. The idea is that those unemployed in a tight labor market are more likely to be workers with some particularly adverse characteristics, possibly associated with higher propensity to drop out. Hence, the group exit probability may display a pro-cyclical pattern even if there is no such individual behavior.

The other hypothesis suggested requires a slight modification of the basic job search model - or the basic theory of labor supply. Consider a non-employed worker who is contemplating job search and who attempts to maximize expected utility, where utility depends positively on consumption and leisure. Assume, furthermore, that the worker can affect his job offer probability via his search intensity, e.g., the fraction of available

Table 3. Determinants of the probability of transitions to employment and not-in-the-labor force, respectively. Monthly data 1967.7-1975.12.

	HWA	w/w^e	TIME	R^2	DW
Dependent variable					
Pue	0.389 (6.272)			0.64	0.84
Pue	0.384 (6.180)	1.832 (1.079)		0.65	0.88
Pue	0.284 (7.574)	0.562 (0.557)	-0.003 (-12.82)	0.88	2.34
Pun	0.775 (9.109)			0.54	0.28
Pun	0.772 (9.007)	1.264 (0.540)		0.54	0.29
Pun	0.617 (16.76)		-0.004 (-20.28)	0.92	1.51
Pun	0.618 (16.73)	-0.702 (-0.707)	-0.004 (-20.19)	0.92	1.51
Pue+Pun	0.438 (16.73)	-0.057 (-0.081)	-0.003 (-22.97)	0.93	2.04

Table 4. Transition probability elasticities w.r.t. job vacancies - results for demographic groups

	White males		White females		Non-white males		Non-white females	
	Pue	Pun	Pue	Pun	Pue	Pun	Pue	Pun
-19	0.48 (7.44)	0.30 3.96)	0.18 (1.85)	0.21 (2.85)	0.86 (4.75)	-0.09 (-0.60)	-0.37 (-1.58)	-0.22 (-1.70)
-24	0.73 (8.23)	0.35 (2.65)	0.91 (9.63)	0.36 (3.80)	0.52 (2.06)	0.21 (0.69)	0.54 (2.23)	0.48 (3.29)
-59	0.18 (3.01)	1.57 (13.41)	-0.002 (-0.03)	0.67 (11.95)	0.05 (0.39)	1.22 (4.68)	-0.02 (-0.10)	0.45 (4.19)

time devoted to job search. A worker in this setting will face an important "investment problem", since by sacrificing current leisure (and perhaps paying monetary search costs) he can enhance his future utility.

How will changed labor market conditions then affect the worker's behavior? An increase in the number of vacancies may be interpreted in terms of a shift of the worker's job-offer production function, simultaneously raising marginal and average productivity. More vacancies, therefore, will increase the marginal return to job search and increase the job offer probability at a given search intensity.

Will the worker increase his search in response to the more favorable labor market conditions? Not necessarily, since there is an "income effect" in operation; the worker is better off and will most likely increase his demand for leisure, assuming the latter to be a normal good. This effect may decrease (planned) hours worked and/or job search. Stated differently, the worker can decrease job search but expect to receive the same wage rate. This effect may offset the "own-substitution effect", associated with the higher return to job search at the margin. The net effect on search is ambiguous *a priori*.

What relevance has all this for the observed procyclical exit-behavior? Our point is the following: If job search among non-employed workers is decreasing when the labor market tightens - which is quite possible, as outlined above - this will probably in a number of cases be registered as labor force exits. The reason is that CPS requires people to search actively in order to classify as unemployed. The borderlines between unemployment and not-in-the-labor force are not very sharp and we may also consider methods of job search not captured by CPS.

The suggested search-hypothesis is admittedly a tentative one. There is, however, some interesting additional evidence available. Given that job search decreases with a tighter labor market, this would also imply a lower probability of unemployment entry from the not-in-the labor force state; the transition probability Pnu would display a contra-cyclical behavior, opposite to the exit probability Pun. When job openings are more readily available, jobless workers - registered as being outside the labor force - may reduce their search, knowing that they probably will find job offers anyhow.

In fact, this hypothesis appears to be consistent with observed behavior. The probability of unemployment entry, conditional on being outside the labor force last month, displays a contra-cyclical pattern in the U.S. The transition probability equations estimated by Ralph Smith for the Urban Institute simulation model[1] reveal (i) that the probability of labor force entry is unaffected by market tightness whereas (ii) the probability of "successful entry", Pne, is varying pro-cyclically. Clearly, these findings imply that "unsuccessful entries", Pnu, are displaying contra-cyclical variations.

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[1] Smith (1977).

5. CONCLUDING REMARKS

In job search literature there has been a tendency to overlook the importance of vacancy contacts as determinants of the duration of unemployment, the emphasis instead being placed on inflationary surprises. This (mis)use of the search story does not necessarily follow from the logic of the theory; most search models do recognize the significance of the stream of job offers. The popularity of the detection-lag view is, probably, its ability to provide a reasonable interpretation of the short-run Phillips curve. The transmission mechanism of aggregate demand policies is explicated in a fairly simple way: an increase in the money growth rate will increase inflation thereby fooling the acceptance decisions of job seekers.

In this paper we have demonstrated that this view has some empirical validity, at least for the short-term unemployed and for a labor market like Sweden's. But we have also shown that unexpected inflation can explain only a small part of the actual fluctuations in unemployment duration.

The elementary search model - where variations in the job offer probability are disregarded - is then clearly inadequate as an explanation of the short-run Phillips curve. Our results also rule out one of the mechanisms which imply a vertical long-run Phillips curve; the natural rate theory must of course be valid if the detection-lag hypothesis is a sufficient explanation of cyclical changes in unemployment. The results are thus more in accordance with the "mainline" view of inflation and unemployment, stressing that aggregate demand influences employment and unemployment via the relaxation of job rationing constraints rather than via misperceptions of relative wages. It is possible that unanticipated price inflation may be of some importance even within the latter framework - as a determinant of the flow of vacancies into the labor market. We are, however, unaware of solid theoretical work on that issue.

Finally, let us offer some comments on the observed differences between the Swedish and the U.S. labor market. Sweden has a highly unionized labor market and wage bargaining at the national level gives rise to relatively uniform and long-term wage contracts. One would be inclined to expect that this institutional setting would produce fast dissemination of information about wages in general, thus reducing the importance of information-lag effects. The less unionized U.S. labor market is probably more resembling the familiar Phelpsian "island parable"[1] than the Swedish is and the scope for temporary wage misperceptions would therefore be greater. In fact, we find the opposite. Why? Let us focus on one additional significant difference between labor market functioning in Sweden and the U.S. - the importance of temporary layoffs. Temporary layoffs constitute - as Martin Feldstein has pointed out[2] - an important source of U.S. unemployment. The U.S. manufacturing layoff rate has varied between 10 and 20 percent (of the number of employed workers) per year whereas the corresponding Swedish figures are 2-4 percent. The major part (60-70 percent) of the

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[1] Phelps (1971).

[2] See Feldstein (1975).

U.S. layoffs are temporary, implying that most workers are ultimately rehired by the same employer. Temporary layoffs in Sweden are, on the other hand, very unusual. Unemployed workers on temporary layoff accounted for 2-3 percent of Swedish unemployment during the period 1975-1978. The corresponding U.S. figures seem to have fluctuated between 10 and 20 percent.[1] Feldstein's view of those laid off as "waiting" rather than "searching" has been questioned on empirical grounds.[2] The Feldstein hypothesis might, however, be considered as modestly corroborated by our results; one interesting interpretation of our revealed U.S.-Sweden differences would be that the extent and intensity of job search among the unemployed is lower in the U.S. If unemployed workers on layoff act as if they will be recalled - and therefore abstain from search - there is little scope for detection-lag effects of the traditional type.

A laid off worker "has a job" in some sense; he is attached to a particular firm and expects to be recalled by his employer. He is probably also well informed about wage changes in his firm. How would a non-seeking unemployed worker on layoff respond to unexpected general wage inflation? He would, most likely, be less inclined to search, thereby reacting similarly to his employed fellows; a fa-

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[1] For Sweden, see the labor force survey (AKU). Feldstein's (1975) figures imply that 18 percent of those unemployed in March 1974 were on temporary layoff. The corresponding figure for March 1978 is 11 percent (Employment and Earnings). Note also that temporary layoffs in general are very short in duration. Thus their fraction of spells of unemployment is even higher.

[2] See the paper by Bradshaw and Scholl (1976) and the following discussion in the Brooking Paper.

miliar implication of search theory is that quits will decrease - via lower propensity to search - as a response to unexpected wage increases. Clearly, temporary layoffs represent a middle state between employment and unemployment. Economic theories designed to explain individual behavior in the polar cases are less suitable when applied to the middle state.

The structural characteristics of the labor markets of the two countries are also - probably to an important degree - related to the frequency of layoffs. We have found that the average duration of completed spells of unemployment is somewhat lower in the U.S., whereas the average number of yearly unemployment weeks per unemployment-experienced worker is higher in the U.S. The most important source of observed unemployment differentials is, however, the different probabilities of being hit by unemployment. The fraction of U.S. workers with unemployment experiences during a year has been 2-3 times as high as the corresponding Swedish figures. The frequent U.S. layoffs may also to some extent account for the pro-cyclical U.S. unemployment inflow.

The results presented in this paper also underline the importance of participation behavior for unemployment duration. Clearly, the marked pro-cyclical fluctuations of exit probabilities are reinforcing cyclical unemployment variations. Further research on these issues should include an analysis of the dynamics of participation behavior at the micro level. Such investigations may also be important for understanding the limitations inherent in current labor force classification schemes.

APPENDIX 1

Table A. Transition probability equations for demographic groups. Monthly U.S. data 1967.7-1975.12.

Dependent variable: Probability of transition from unemployment to employment (Pue).

	HWA	w/w^e	TIME	R^2	DW
WM 1619	0.477 (7.436)	-0.729 (-0.423)	-0.002 (-4.956)	0.82	1.97
WM 2024	0.726 (8.225)	1.039 (0.438)	-0.005 (-9.758)	0.77	1.70
WM 2559	0.175 (3.009)	-0.793 (-0.507)	-0.003 (-8.185)	0.75	2.22
WF 1619	0.176 (1.851)	1.700 (0.667)	-0.001 (-1.001)	0.59	2.09
WF 2024	0.914 (9.635)	-0.216 (-0.085)	-0.007 (-12.20)	0.81	2.02
WF 2559	-0.002 (-0.035)	2.929 (1.615)	-0.001 (-2.314)	0.64	2.07
NM 1619	0.862 (4.752)	3.904 (0.801)	-0.005 (-4.459)	0.60	1.94
NM 2024	0.519 (2.055)	-11.33 (-1.670)	-0.006 (-3.938)	0.37	1.94
NM 2559	0.048 (0.385)	0.263 (0.079)	-0.004 (-5.968)	0.53	1.83
NF 1619	-0.365 (-1.578)	4.480 (0.720)	0.000 (0.105)	0.40	2.30
NF 2024	0.542 (2.226)	2.412 (0.369)	-0.009 (-6.813)	0.49	2.59
NF 2559	-0.015 (-0.097)	2.081 (0.488)	-0.005 (-5.550)	0.46	2.06

Note: WM 1619 = white males, age 16-19. WF = white females, NM = non-white males, NF = non-white females.

Table B. Transition probability equations for demographic groups. Monthly U.S. data 1967.7-1975.12.

Dependent variable: Probability of transition from unemployment to not-in-the-labor force (Pun).

	HWA	TIME	R^2	DW
WM 1619	0.300 (3.960)	-0.005 (-11.22)	0.80	1.53
WM 2024	0.349 (2.645)	-0.003 (-3.563)	0.30	1.95
WM 2559	1.566 (13.41)	-0.005 (-7.146)	0.79	1.74
WF 1619	0.206 (2.851)	-0.003 (-8.180)	0.68	1.79
WF 2024	0.362 (3.799)	-0.004 (-7.650)	0.58	1.89
WF 2559	0.671 (11.95)	-0.006 (-18.07)	0.88	1.47
NM 1619	-0.088 (-0.597)	0.000 (0.100)	0.34	1.63
NM 2024	0.215 (0.691)	-0.002 (0.968)	0.10	2.25
NM 2559	1.216 (4.684)	-0.004 (-2.752)	0.39	2.11
NF 1619	-0.224 (-1.700)	0.003 (3.434)	0.30	1.79
NF 2024	0.480 (2.290)	-0.003 (-4.084)	0.38	2.38
NF 2559	0.453 (4.194)	-0.002 (-3.327)	0.35	1.75

APPENDIX 2

BASIC DATA

	Sweden					U.S.				
	u	f	D	N/L	W	u	f	D	N/L	W
1965	1.2	0.22	5.4	-	-	4.5	-	-	16.2	14.4
1966	1.6	0.30	5.3	10.0	8.3	3.8	-	-	15.0	13.2
1967	2.1	0.26	8.0	9.1	12.0	3.8	-	-	15.0	13.2
1968	2.2	0.24	9.0	8.9	12.8	3.6	0.48	7.5	14.4	13.0
1969	1.9	0.27	7.1	7.8	12.6	3.5	0.47	7.5	14.5	12.6
1970	1.5	0.22	6.9	7.4	10.5	4.9	0.55	9.0	17.6	14.5
1971	2.5	0.22	11.6	10.2	12.7	5.9	0.61	9.6	18.8	16.3
1972	2.7	0.17	15.5	10.0	14.1	5.6	0.60	9.4	17.7	16.5
1973	2.5	0.17	14.6	10.0	12.9	4.9	0.56	8.8	16.3	15.6
1974	2.0	0.19	10.3	8.3	12.5	5.6	0.59	9.4	20.4	14.3
1975	1.6	0.15	10.7	6.1	13.7	8.5	0.75	11.3	22.8	19.3
1976	1.6	0.13	12.3	6.0	13.8	7.7	0.69	11.1	21.6	18.5
1977	1.8	0.14	12.7	6.5	14.3	7.0	-	-	19.9	18.3
1978	2.2	0.19	11.8	7.7	14.9	6.0	-	-	-	-

u = the unemployment rate (%)

f = the unemployment inflow rate (%)

D = the average duration of unemployment spells (weeks)

N/L = the fraction of the labor force with unemployment experiences during the year (%)

W = the number of yearly unemployment weeks per unemployment-experienced individual.

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The Structure and Dynamics of Unemployment: Sweden and the United States

Comment
Edward M. Gramlich

This paper was extremely interesting, a nice combination of some skillful use of modern techniques for analyzing unemployment and very provocative results.

Björklund and Holmlund (B&H) focus on the duration of unemployment, or its inverse the exit probability, for the United States and Sweden. They first show that:

a) Duration levels are about the same in both countries, even though unemployment rates are higher in the United States.

b) Duration provides all the cyclical variation for unemployment rates in Sweden and only some of it in the United States, with the flow into unemployment also varying cyclically.

Then they get down to business and analyze the duration figures more carefully. Using a generalized search model, they can derive equations of the form

$$-\ell n\ D = \ell n\ P = \alpha_1 + \alpha_2\ \ell n\ V + \alpha_3\ \ell n\ (W/W^e)$$

where D is unemployment duration
P is the exit (from unemployment) probability
V is job vacancies
W/W^e is the ratio of actual to expected wages, with the latter defined as an autoregressive distributed lag of various types.

Basically, what happens is that α_3 comes out to be close to zero in most cases (excepting only short term unemployment in Sweden, where α is not zero but still explains relatively little of the variation in P).

B&H don't give the full implication of this finding for Phillips Curve research, so I will try to. Most economists agree that Phillips Curves (relating money wage growth to unemployment) are downward sloping in the short run and vertical in the long run. The traditional explanation, Keynesian modified to include adaptive inflationary expectations (MK), runs something like the following. Rates of unemployment serve as a proxy for excess aggregate labor demand, and show how varying levels of excess demand cause varying rates of wage change, along a downward sloping curve. Expected rates of wage and price inflation are held constant along this short run curve, but as they adjust to actual rates given endogenously by he short run curve, the long run curve becomes vertical if the coefficient of adaptive expectations is unity. The key feature here is that unemployment is the causal variable.

This view of Phillips Curves has been attacked vigorously by the rational expectations (RE) school, which has built on earlier micro foundations and search theory work to explain the same facts. They view movements in unemployment around the natural rate as being only transitory, caused by misperceptions of true real wages. When $W/W^e > 1$, workers will think they have been moved along their labor supply curve, they will supply more labor, and unemployment will be temporarily low, with the situation continuing until they get wise. As before, the unemployment cycle lasts until wage observations and expectations coincide, but now the causal variable is wage perceptions, not unemployment.

The two theories can have different policy implications and are certainly radically different ways of modelling labor market behavior. Hence it is surprising there has been so little testing of their differing implications. One can imagine many such tests:

a) Within a Phillips Curve framework, one might use new techniques of identifying lagging and leading variables to find the true causal force. I'm not aware that this has been done, though the RE devotees are now engaged in a systematic attempt to do this for all equations in the macro system, so perhaps it soon will be. I would like to see the tests also made by people who are neutral in the dispute, however (if there is anybody left who is still neutral).

b) Also within a Phillips Curve framework, one might disaggregate the wage lag process. The wage lag term in a Phillips Curve exists in the RE model only because of misperceptions: prices and wages are supposed to clear quickly once information lags have been exhausted. But in the MK model there is also the possibility that wages will be slow to adjust because of institutional rigidities - long term wage contracts and so forth. It does not seem impossible to try to distinguish these lags and make a discriminating test.

c) Within a RE framework, one might test independently to see if wage misperceptions do indeed raise cyclical unemployment. This is the unique contribution of B&H, and it seems to me they pretty decisively reject the RE hypothesis. Wage misperceptions might still affect flows into unemployment - quits and layoffs - but apart from the short term Swedish unemployment - they consistently do not affect duration in either country. On the other hand, the job vacancy variable, which can be fit into a search theory, RE type model but is more characteristic of an MK explanation for unemployment, works extremely well. I would have to give round one to the MK fighter.

Let me first make one technical comment regarding this important result. B&H do actually present one U.S. equation, for the 1969-1973 period, where wage misperceptions have a very slight, though insignificant, positive effect. I would tend to disregard this period because for most of it, the U.S. had in effect devices for legally-controlling wage movements. It is extremely tricky to model expectations in such a period, but also extremely clear that the usual ARMA techniques are inadequate. Hence, if I were they, I would omit these observations, not run regressions for only those observations. This comment, by the way, shows why we can only award round one to the MK fighter: if it were easy to model expectations, economists wouldn't be so divided and confused about their effects.

I would like also to comment on an important subsidiary result found by B&H. They use the identity

$$P = P_{ue} + P_{un} \tag{2}$$

to disaggregate unemployment exit probabilities into one term explaining exit into employment (P_{ue}) and another explaining exit from the labor force (P_{un}). They then use U.S. data to run model (1) with both P_{ue} and P_{un} as dependent variables. As we might now anticipate, the wage growth terms are still either insignificant or negative, again providing evidence against the RE Phillips Curve hypothesis. Also the vacancy terms again work well, but now too well. As they should be, they are strong explainers of P_{ue}. But as they shouldn't they also explain P_{un}. Subsidiary tests in the Appendix show that this result comes through across the board for all white age and sex groups, and only doesn't work for non-white teenagers. That it doesn't work for non-white teenagers is reassuring - non-whites should enter, not leave, the labor force when there are more jobs. The puzzle is why the effect does work for all other groups. Why does a growth in vacancies cause unemployed workers to drop out of the labor force?

I should point out first that this finding by B&H is not necessarily inconsistent with the discouraged worker hypothesis - that the labor force rises when V rises - if movements from employment to not in the labor force exceed those from unemployment to not in the labor force.

But even though the discouraged worker hypothesis may still be alive and well, the result is nevertheless very surprising. B&H advance two tentative explanations for it, both of which I find totally unconvincing. The first is an aggregation bias explanation - that P_{un} rises with vacancies because the only unemployed left are those who are exogenously inclined to drop out. The reason this is unconvincing is that B&H have estimated their equation with data from high and low vacancy periods. Their argument might explain why the impact of V on P_{un} could be close to zero when V is high, but it would certainly not explain a positive coefficient when V is low. They could test this explanation either by splitting the sample into high and low V periods, or by using a longitudinal data bank to see if P_{un} varies with V for the same individuals. Such tests ought to be made, and I presume they would reject the explanation. If not, that is if the impact on P_{un} is an aggregation phenomenon, then the whole technique of running these sorts of regressions is open to question, for one can generate almost any kind of counter-intuitive explanation and justify it by aggregation bias.

The second explanation is similarly unconvincing, focusing on the income effect in a search model. If there is an income effect, high levels of V make unemployed workers better off, and they will use this gain in welfare to buy more leisure time, or stop searching. While some of these workers might search slightly fewer hours per worker, just enough to be dropped from the labor force by the CPS, for the most part this income effect makes no sense when talking about transition probabilities. What "rational search" model would have workers stop searching altogether, and drop out of the labor force, as B&H find? To use a homely analogy, I know fishermen who will come in sooner when the fish are biting, because they catch their limit. I don't know any fishermen who will <u>not</u> go fishing when the fish are biting, because in their hearts they know they are better off.

Of course the harder job is to be constructive. As a discussant, it is easy to tear down somebody else's tentative explanation: what is my own for the finding? The only one I can come up with uses the recent Clark-Summers (1979) finding that some funny things happen when we try to analyze the unemployment-not in the labor force relationship too carefully. One of their main points is that unemployed workers float into and out of the labor force for little apparent reason, and it may simply be that even though the CPS classifies these workers as P_{un} workers, they are really nothing more than workers who have not yet exited the unemployment status. Or maybe they were truly not in the labor force all along, and now are afraid to make even minimal attempts to look for a job for fear they might find one. Whatever the case, my own bet would be that the key to the puzzle lies in a more thorough understanding of the CPS and its questioning procedures, not in a more complete development of search models to include income effects.

I should close by saying that even though I did not appreciate the B&H tentative explanations for some bizarre findings, I still congratulate the authors for an important contribution in testing different explanations for the short run Phillips Curve.

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Part III

Labor Supply, Human Capital and Labor Force Participation

Male–Female Lifetime Earnings Differentials and Labor Force History

Siv Gustafsson

1. INTRODUCTION

Sweden has an international reputation for relatively high standards of equity between the sexes. This reputation is justified formally by our family legislation. But is it justified as well by achieved smaller earnings differentials? This paper analyzes the effects of work experience and other human capital variables on earnings for men and women. Using these results, sex earnings differentials are calculated. Comparisons are then made between the results for Sweden and results from earlier studies for the U.S.

The human capital approach to income determination has been implemented by means of an earnings function which was developed by Jacob Mincer (1974). Studies on sex differentials in earnings using this approach are, e.g., Oaxaca (1973), and Malkiel and Malkiel (1973) and for Swedish data Gustafsson (1976). The earnings function was developed for the case of an interrupted career in Mincer and Polachek (1974) and has already been applied to different sets of data with information on work histories. (See Sandell and Shapiro, 1978; Mincer and Polachek, 1978; Corcoran and Duncan, 1979; and Mincer and Ofek, 1980.)

2. EFFECTS OF LABOR FORCE INTERRUPTIONS ON FUTURE EARNINGS

Schooling and on the job training may be seen as productive investments in human capital that increase the earnings capacity of the individual. The general form of the lifetime earnings profile for a continuous work history may be approximated by the well known log-linear earnings function developed by Mincer (1974):

$$\ln(y) = \beta_0 + \beta_1 S + \beta_2 e + \beta_3 e^2 \qquad (1)$$

where the logarithm of earnings (y) is regressed on schooling (s), experience (e) and a squared experience term (e^2).

Not all individuals work continuously in the labor market but some drop out for various reasons, e.g., unemployment, health or child care.

There are three reasons why we may expect lower earnings if the individual experiences labor force interruptions. First, the opportunity to make post-school investments in earnings capacity by on-the-job-training is sacrificed. Second, there may be depreciation of human capital due to nonuse because a person may forget what was once learned or because the accumulated human capital becomes obsolete. Third, expectations of lower labor force participation imply fewer periods in which to collect returns. This may lead to low investment in the preinterruption period.

The work history of the individual is segmented into periods of market activity e_j and periods of non-market activity h_i:

$$h_1 + e_1 + h_2 + e_2 \ldots + h_K + e_\ell = \sum_{i=1}^{K} h_i + \sum_{j=1}^{\ell} e_j \qquad (2)$$

if there are K periods of home time and ℓ periods of labor force experience.

A major determinant of the size of investment in any particular period is the potential number of years of working life in the future. If individuals who work continuously and individuals who made labor force interruptions have an equal number of years of potential remaining working life until retirement there is a reason to believe that the earnings profiles in the last continuous period of work experience before retirement should be parallel.[1] As will be shown empirically later this may be achieved by letting the slope of the earnings function depend on age and combining (1) and (2):

$$\ln(y) = \alpha+\beta s+\sum_{i=1}^{K} \gamma_i h_i + \sum_{j=1}^{\ell} \delta_j e_j+\varepsilon_1 a+\varepsilon_2 a^2 \qquad (3)$$

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[1] In Mincer and Ofek (1980) one hypothesis which may lead to more rapidly increasing earnings for returning individuals is that the human capital that was depreciated or became obsolete can be fairly easily repaired to restore its previous value.

3. DATA AND VARIABLE DEFINITIONS

The data consist of a one in ten random sample of white-collar workers in the private sector of Sweden. The sample includes 32,287 individuals of whom 23,366 are men and 8,921 women.

The set of data was arranged by matching salary statistics of 1974 on an individual basis with a one in ten random sample of pension funds data for the Swedish population. The pension funds data make a time series of pension points "ATP" for every year from 1960-74 (see Eriksen, 1973). The salary statistics are collected jointly by the Swedish Employers' Confederation (SAF) and the employee organizations in the private sector. Most white-collar workers in the Swedish private sector, industry, retail trade and services are covered by this set of data.

For each individual there is the time series of pension points "ATP" and cross sectional data for 1974 of monthly salary, education, hours of work per week, occupation, industry branch, size of the company and geographical location[1].

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[1] Occupation is recorded according to the "Position Classification System for Salaried Employees", worked out in cooperation by the Employer and Employee organizations. The classification system is a 4-digit system where the first 3 digits give an occupation. The first 3 digits thus give the horizontal dimension distinguishing about 60 occupations. By the first digit occupations are grouped together in ten occupational fields. The 4th digit contains a vertical classification of jobs into job levels of 7 different degrees of difficulty.

Education is given by a 3-digit code, the Swedish Educational Nomenclature (SUN). Although the SAF

cont.

Work histories have been calculated from the time series of ATP scores.[1] In the pension funds data ATP points are calculated from annual earnings of the individual and the "base amount" according to the formula:

$$ATP = \frac{y - b}{b} \tag{4}$$

where y = annual earned income

b = "base amount".

An individual is defined to be a labor force participant if:

ATP > 0

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cont.

types of education shorter than 11 years the code comprises 200 different values in this set of data. Education has been coded 000 in most cases where employees have compulsory schooling because SAF has not bothered to collect educational information according to the SUN code on these people.

[1] ATP means **A**llmänna **T**illäggs**p**ensioneringen. The scores of ATP-points are given by three digits in the data. The ATP are real incomes because they are deflated by the base amount (b) which is altered discretionarily by Swedish authorities to take account of price increases. The base amount is used not only for calculating the basis on which old age pensions are paid but also for the current income taxation and in the determination of minimum standards for welfare. The ATP are truncated from above at 6.50 since incomes higher than 7.5 times the base amount do not increase pensions further above this level. It is also clear by the definition of an ATP-point that it is truncated from below. The base amount, one b, in August 1974 was 8 500 Skr.

This definition of labor force participation is dependent on earnings, which means that a labor force interruption must be a whole year in order to be observed. One advantage of these data is that they are recorded rather than calculated retrospectively.

Years in which the individual is a labor force participant according to definition (5) are defined as years of experience (e). Years in which the individual is not a labor force participant can be either years of schooling (s) or years of home time (h).

Years of schooling are separated from years of experience by making use of two identities and information on type of education completed and age.[1]

The following identities are valid:

$$g + e + h \equiv a$$

and (5)

$$s \equiv g - 7,$$

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[1] The third digit in the Swedish Educational Nomenclature (SUN) gives number of years of schooling required to complete the education. This is given in seven steps which have been aggregated to compulsory schooling (1-2), secondary schooling (3-4) and university graduates (5-7). The quality of the educational coding in the salary statistics does not really admit distinguishing all the different steps. This is particularly true of the shorter educations. A majority of those not having a longer education are coded 0 instead of 1 or 2.

where g = age at graduation from school, e = number of years of experience, h = years of schooling.

We have assumed that:[1]:

$$g = 16 \text{ for compulsory schooling}$$
$$g \leq 19 \text{ for secondary schooling} \qquad (6)$$
$$g \leq 25 \text{ for university schooling}$$

If a person is young enough to have finished school not more than 15 years ago there is full knowledge of the labor force experience and nonexperience per year since the end of schooling. For older women we focus on the effects of the last 15 years of actual experience on earnings.

The time series of ATP makes possible the calculation not only of years of home time (h), but also of the individual movements into and out of the labor force during the 15 year period. A year (i) of labor force entrance occurs if:

$$ATP_{i-1} = 0 \text{ and } ATP_{i} > 0 \qquad (7)$$

and a year (j) of labor force exit occurs if:

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[1] The weak inequalities have been chosen in order to avoid counting as a year of schooling if it is actually known by the data that a person has been working. This may be observed for young persons who finished school not more than 15 years ago.

An earlier attempt to put g = 24 for young female university graduates resulted in a number of negative observations on h. Allowing g to vary resulted in a variation of s from a minimum of 9 years to a maximum of 18 years and a mean of 15.3 years for the young female university graduates.

$$ATP_{j-1} > 0 \text{ and } ATP_j = 0 \tag{8}$$

By computing the years of entrance by (7) and the years of exits from the labor force by (8) it is possible to calculate the number of years spent in and out of the labor force in different segments.[1] The first segment of home time (h_1) is calculated in the following way:

$$h_1 = i_1 - b - g \tag{9}$$

where i_1 = the year of the first entrance into the labor market

b = the year of birth of the individual

g = age of graduation.

If the person is old enough to have graduated more than 15 years ago h_1 is a truncated variable and is calculated:

$$h_1 = i_1 - 1960 + 1 \tag{10}$$

where i_1 = the calendar year (i) of the first labor force entrance in the observed period.

Because individuals in the sample are labor force participants in 1974 the last continuous period of labor force participation e_ℓ is calculated:

$$e_\ell = 1974 - i_\ell + 1 \tag{11}$$

where i_ℓ is the calendar year of the last intrance. For the individuals who made no labor force interruption:

$$i_1 = i_\ell \text{ and } e_1 = e_\ell. \tag{12}$$

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[1] Since there are 15 years, the maximum number of entrances and exits if the person works every second year are seven of each. Most of the individuals have much smaller mobility into and out of the labor force. Only 2 individuals had left the labor market for a fourth time and only one had left the labor market for a fifth time and entered again.

4. WORK HISTORIES OF SWEDISH WOMEN

Out of the 6,745 fulltime working women in the sample more than one fourth were in the labor force in 1960 and stayed in all the following 15 years without any labor force interruption. If all women are distributed according to the number of labor force interruptions they made during the 15 years period we find that 5,278 made no interruptions, 1,221 made one labor force interruption of varying length, 221 women made 2 labor force interruptions and 22 women made 3 labor force interruptions. Twenty-two per cent of the women made one or more labor force interruptions and the average number of years out of the labor market was 2.34. Half of those who made a labor force interruption were out only one year and only one fifth were out more than three years.

In Table 1 work histories are given. The largest proportions of total experience of the Swedish women occurs in the last period of uninterrupted labor force experience. The women spent 9.5 years in the labor market and 8.9 of these were spent in the last period. Years out of the labor market were most common in the form of delayed entrances rather than labor force interruptions.

Comparisons with work histories for American women are difficult because American studies do not cover the same period, have different definitions of labor force interruptions and are not restricted to private sector white collar workers. But some crude comparisons are possible.

There are two sources of American data to compare with. The Michigan panel data used by Corcoran and

Table 1. Work history 1960-74. Full time female workers in private sector of Sweden in 1974. Means of variables.

	n	h_1	e_1	h_2	e_{2+}	h_{3+}	e_ℓ	h	e
Age 16-64									
All educations	6745	1.83	0.55	0.44	0.07	0.07	8.89	2.34	9.51
Compulsory education	5110	1.94	0.57	0.44	0.07	0.07	9.22	2.45	9.86
Secondary education	1432	1.55	0.47	0.41	0.07	0.06	8.11	2.02	8.65
University education	203	1.06	0.61	0.57	0.05	0.04	6.12	1.67	6.78
Age									
16-19	321	1.56	0.00	0.00	0.00	0.00	1.69	1.56	1.69
20-24	1393	2.04	0.14	0.10	0.01	0.01	4.20	2.15	4.35
25-29	1422	2.17	0.60	0.51	0.07	0.06	7.48	2.74	8.15
30-34	854	2.45	1.24	0.94	0.19	0.18	10.19	3.57	11.62
35-39	548	1.58	0.91	0.99	0.19	0.17	11.18	2.74	12.28
40-44	600	1.99	0.58	0.57	0.08	0.10	11.68	2.66	12.34
45-49	555	1.48	0.55	0.26	0.06	0.05	12.59	1.79	13.20
50-54	546	1.05	0.38	0.25	0.04	0.06	13.22	1.36	13.64
55-59	354	0.60	0.29	0.19	0.01	0.00	13.87	0.79	14.17
60-	116	0.75	0.43	0.16	0.02	0.03	13.61	0.94	14.06

Definitions of variables:

n = number of observations,

h_1 = years of delayed entrance to the labor market after schooling,

e_1 = years of first labor force experience,

h_2 = years of hometime in the first labor force interruption,

e_{2+} = years of labor force experience between labor force interruptions,

h_{3+} = years of labor force interruption (hometime) for the second or subsequent time,

r years of labor force experience in the last continuous period up to e_ℓ 1974.

Duncan (1979) show that white American women who were working in 1975 had spent 5.8 years out of the labor force since finishing school and 10.4 years in the labor force, or 36% of the total period in non-market activities. Swedish women according to Table 1 spent 20% out of the labor force but then the labor force histories of older women are truncated.

Mincer and Polachek (1974), provide work histories from the NLS of women 30-44 years old broken down by three educational groups and by five years age intervals. By this comparison American women of the same age and approximately the same length of education seem to have spent more years out of the labor force than the Swedish women. American women with less than 12 years of education have spent more than half of the available number of years since leaving school out of the labor market. The proportion of years spent out of the labor force decreases with length of education but is still .4 for university graduates. Note that this is higher than .2 reported for Swedish women 30-34 years of age.

Differences in definition may be one reason for the small amount of time spent out of the labor force by Swedish women. Maternity leave with sick security payment for most of the covered period is six months.

Half a year of market earnings is enough to earn a positive ATP and to be considered a labor force participant for the year. However, the impression that Swedish women spend less time out of the labor force than American women is consistent with the fact that female labor force participation in Sweden is higher than in the U.S. (Gustafsson,1980).

The general impression of this comparison is that Swedish women work a larger proportion of the available number of years and are less inclined to make labor force interruptions and stay out for any extended period of time.

5. IS THERE DEPRECIATION FROM LABOR FORCE INTERRUPTIONS?

Mincer and Polachek (1974) found that for every additional year out of the labor force, wages diminished by 1.5 percent at return, i.e., a depreciation rate of -.015. Sandell and Shapiro argued that the rate of depreciation was overstated by Mincer and Polachek and that a more correct figure is -.005. Mincer and Polachek (1978) in their reply to Sandell and Shapiro claimed that the estimates of the two studies are not significantly different by standard statistical tests. Corcoran and Duncan (1979) report a depreciation rate for white women of -.005 which is statistically significant. None of the American studies have shown positive coefficients on home time.

In Table 2 estimates of earnings functions for Swedish women are reported. The most striking result is that the coefficient of home time is negative and significant only for the earnings function where women 30-44 years of age were subselected. In all other regressions the coefficient of home time is positive and significant implying

Table 2. Earnings functions for Swedish full time working women with different specifications of experience and for different age groups. Dependent variable: logarithm of monthly salaries.

	All aged 16-44		All aged 30-44	Young, labor force history fully known e = exp	Aged 16-64 with no labor force interruption h = 0
	(1)	(2)	(3)	(4)	(5)
Intercept	6.9191 (0.0152)	7.0155 (0.0150)	7.5412 (0.0578)	6.7780	6.9566 (0.0215)
s	0.0598 (0.0013)	0.0583 (0.0014)	0.0476 (0.0026)	0.0673 (0.0014)	0.0632 (0.0018)
h	0.0210 (0.0008)	0.0101 (0.0008)	-0.0106 (0.0020)	0.0315 (0.0012)	
e	0.0598 (0.0021)			0.0813 (0.0029)	
e^2	-0.0011 (0.0001)			-0.0025 (0.0002)	
exp		0.0398 (0.0007)	0.0081 (0.0041)		0.0443 (0.0009)
$(exp)^2$		-0.0006 (0.0000)	-0.0001 (0.0001)		-0.0007 (0.0000)
R^2	0.5786	0.5221	0.2519	0.6471	0.5644
n	6745	6745	2038	3656	3197

Note: The regressions are numbered so that reference can be made to them in the text. Numbers in parentheses are standard errors.

that reentering women have higher salaries than women with the same number of years of experience with continuous careers. In regression (1) only the last 15 years of experience are allowed to influence the results. In regression (2) women who finished school more than 15 years ago are assumed to have had an equal proportion of years out of the labor force in the unknown period as in the known period:

$$exp = (1 - \frac{h}{e+h})(a-g-1) \qquad (13)$$

Assumption (13) is also made for the subsample of women 30-44 in equation (3). For the young women whose labor force history is fully known the home time coefficient is likewise positive and significant.

The marginal effect of an additional year of schooling is about 5-6%. The estimates have t-values of close to 50 for schooling and experience. The schooling and experience variables explain a great deal of the variation in earnings, R^2:s are above 50% for the full sample, which is high for micro data.

The results may be explored further by segmenting the earnings function. Results for men and women according to the segmented earnings function controlling for age are given in Table 3. We consider the effects of only the last 15 years but it matters how old the individual is when the labor force interruption takes place. It also means that investment is affected by the remaining horizon since the slope of the earnings function is determined by age.

Table 3. Segmented labor force history.
Estimates for all men and all women.

	Women		Men	
	Regression estimate (6)	Mean of variable	Regression estimate (7)	Mean of variable
ln(y)		7.98		8.38
Intercept	6.6351* (0.0301)		6.5250* (0.0558)	
s	0.0534* (0.0014)	9.74	0.0657* (0.0010)	10.88
h_1	0.0129* (0.0012)	1.83	0.0179* (0.0022)	0.76
e_1	0.0189* (0.0019)	0.54	0.0229* (0.0029)	0.52
h_2	0.0107* (0.0021)	0.44	0.0107 (0.0038)	0.33
e_2	0.0080 (0.0054)	0.07	0.0195* (0.0084)	0.05
h_3	0.0178* (0.0059)	0.06	0.0107 (0.0119)	0.04
e_3	0.0211 (0.0331)	0.01	-0.0341 (0.0725)	0.00
h_{4+}	-0.0074 (0.0300)	0.00	0.0571 (0.0757)	0.00
e_ℓ	0.0299* (0.0013)	8.90	0.0376* (0.0018)	12.1
a	0.0274* (0.0023)	34.2	0.0300* (0.0035)	41.2
a^2	-0.0003 (0.0000)	1170	-0.0003 (0.0000)	1697
R^2	0.5853		0.5010	
n	6745		7048	

* Significant at the 5% level.

Not all the coefficients of this function are statistically significant probably because there are few observations of individuals who make several labor force interruptions. One of the results of Mincer and Polachek (1974) is that the coefficient of preinterruption experience periods is smaller than the coefficient of postinterruption experience periods. The interpretation is that individuals invest more in on-the-job-training during a period in which they expect their labor force participation to be continuous than is the case during a period which they expect to be terminated soon. The first experience period is about half a year on the average for men and women alike. The regression coefficient for this period (e_1) is significant. The regression coefficients for the last continuous period of work experience are highly significant for both men and women and show higher increases in earnings capacity during the last period than during the first experience period (e_ℓ is 8.9 years on the average for women and 12.1 years on the average for men). This result conforms to the Mincer and Polachek result.

Years of schooling would be more likely to be included in h_1 than in h_2-h_4. Table 3 shows, however, that the coefficients of all the home time segments are positive ruling out the possibility that fulltime investment activities explain the positive coefficients of hometime in Table 2.

Corcoran and Duncan (1979) ran segmented earnings functions for men and women and concluded that by segmenting the work history they get more similar regression coefficients across sex and race groups than are obtained when years of experience are measured as a single variable. In Table 3 the experience variables have very similar effects on earnings between the sexes.

In Table 4 earnings functions for separate educational groups are given. The segmented earnings function does not work for subsamples broken down by education because there are too few observations of people who had several entries and exits into the labor force. Instead traditional earnings functions are given. The coefficients of the experience variables are more similar between the sexes than is the case across the educational groups.

One result of Mincer and Polachek is that the depreciation of human capital increases with schooling. The coefficients of home time in Table 4 are positive but they show a decreasing pattern with length of schooling. Also the experience variables increase with length of education implying that there is a higher penalty to home time the higher education an individual has.

In all regressions the marginal effect of a year of experience is larger than the marginal effect of a year of home time. In regressions on Swedish earnings data the net effect on future earnings of a year of home time is positive rather than negative. This implies that investment during home time has taken place to such an extent as to exceed negative depreciation. Another explanation is that wages and salaries are to such a large extent determined by central negotiations that investments in skills do not entirely determine wages. (Gustafsson, 1976).

Table 4. Earnings functions for men and women of different educational groups

	Compulsory		Secondary		University	
	Women (8)	Men (9)	Women (10)	Men (11)	Women (12)	Men (13)
Inter-cept	7.5372	7.6947*	6.9596*	7.2358*	7.2237*	7.1510*
s	-	-	0.0600*	0.0434*	0.0386*	0.0514*
h	0.0129*	0.0219*	0.0028	0.0259*	-0.0185*	0.0150*
exp	0.0382*	0.0391*	0.0505*	0.0552*	0.0693*	0.0767*
$(exp)^2$	-0.0006*	-0.0006*	-0.0009*	-0.0009*	-0.0017*	-0.0016*
R^2	0.4947	0.1554	0.5591	0.4137	0.4191	0.5324
n	5 100	2 620	1 428	2 124	203	2 322
Means of variables:						
ln(y)	7.9500	8.2600	8.0537	8.4422	8.2175	8.6363
a	34	43	32	40	32	37
s	9	9	11.5	11.8	15.7	16.5
h	1.88	0.72	1.26	0.72	0.51	0.30
exp	17.3	27.3	13.1	21.3	9.3	13.7

* Significant at the 5% level.

$$exp = (1 - \frac{h}{e+h})(a-g-1)$$

where h = years of hometime

e = years of experience

a = age

g = age at graduation

6. PREDICTED LIFETIME EARNINGS FOR INDIVIDUALS WITH DIFFERENT LABOR FORCE HISTORIES

The fact that home time does not result in a net depreciation of human capital does not mean that it is costless to the individual to make labor force interruptions. There are costs of lost investment opportunities. In Table 5 predicted earnings at age 35 for women with secondary education are given, according to different specifications of the earnings function. All specifications show lower earnings after labor force interruptions than for continuous careers as predicted by theory.

In panel 1 predicted earnings for women with continuous careers are given. At age 35 earnings range between the logarithm of 8.21 and 8.37. When only the past 15 years of experience have an influence the estimated earnings are higher than when the older women are assumed to have worked in the market in the same proportion in the unknown period as in the known period.

In panel 2 earnings of women who made one labor force interruption of differing length 2, 5 or 10 years, are compared.

Only by the segmented earnings function of Table 3 it is possible to make the distinction between one or two spells of the same length. Five years of hometime in one spell reduces earnings less than five years in two spells (2+3 years), which is seen by comparing the third column of panels 2 and 3.

Table 5. Predicted log of salary at age 35 for a woman with secondary schooling (s=12)

Predicted from Tables 2-4. Equation No.	(1)	(2)	(6)	(10)
1. Continuous career	8.3354	8.2183	8.3757	8.2730
2. Interrupted career one spell				
Left at age 25 for:				
a) 2 years	8.3284	8.1973	8.2713	8.2352
b) 5 years	8.3009	8.1568	8.2137	8.1650
c) 10 years	8.2114	8.0653	8.1177	8.0120
3. Interrupted career two spells				
left at age 25 and 30 (2+3 years)	8.3009	8.1568	8.1693	8.1650
4. Marginal differential between a year of experience and a year of nonexperience				
a) $\partial \ln(y)/\partial e - \partial \ln(y)/\partial h_2$ at age 25	0.0256	0.0319	0.0192[a]	0.0397
b) $\partial \ln(y)/\partial e - \partial \ln(y)/\partial h_2$ at age 35	0.0224	0.0194	0.0192[a]	0.0199
5. Total differential between a continuous career and a career with 10 years of home time				
a) total differential	0.1240	0.1530	0.2580	0.2610
b) percent	11.7	14.2	22.8	23.0

[a] $\frac{\partial \ln(y)}{\partial e_\ell} - \frac{\partial \ln(y)}{\partial h_2}$.

The marginal differential between a year of experience and a year of nonexperience is calculated. In panel 4 the differential is constant according to equation (6) but varies over the life cycle according to the other equations because of the second order term of experience. From equation (6) foregone opportunity of investment in human capital results in an opportunity loss of about 2% per year (0.0192) for an individual chosing a year of hometime. In equation (10) the opportunity loss is almost 4% at age 25 but only 2% at age 35.

In the bottom panel the total earnings differential between the continuous career and a career with ten years of home time is calculated. The estimate varies between 11.7% and 23.0% lower earnings at age 35 if 10 years of potential working life has been spent out of the labor force.

In Figure 1 predicted earnings from regressions 6 and 7 of Table 3 are drawn. As seen in panel 4 of Table 5 the only specification of the earnings function that gives parallel logarithmic earnings functions between continuous and interrupted careers is the one where the slope of the earnings function depends on age rather than on experience. The other specifications give interrupted careers that approach the continuous career. The parallelity of the continuous with the interrupted careers for women is seen in Figure 1A. In Figure 1B the same age earnings curves are transformed into Skr. The earnings curves are not parallel when measured in absolute rather than relative terms.

Figure 1A. Predicted logarithmic earnings from regression estimates (6) and (7)

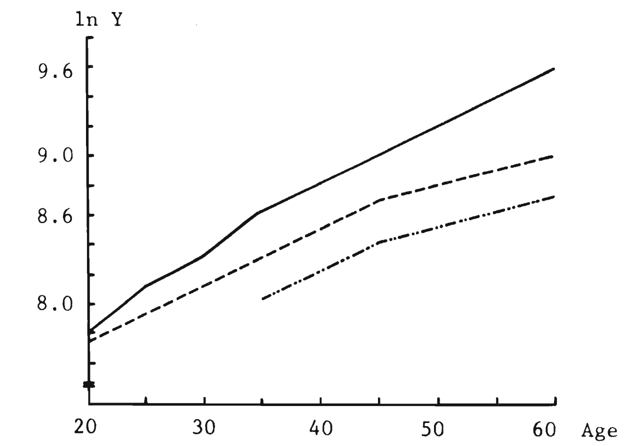
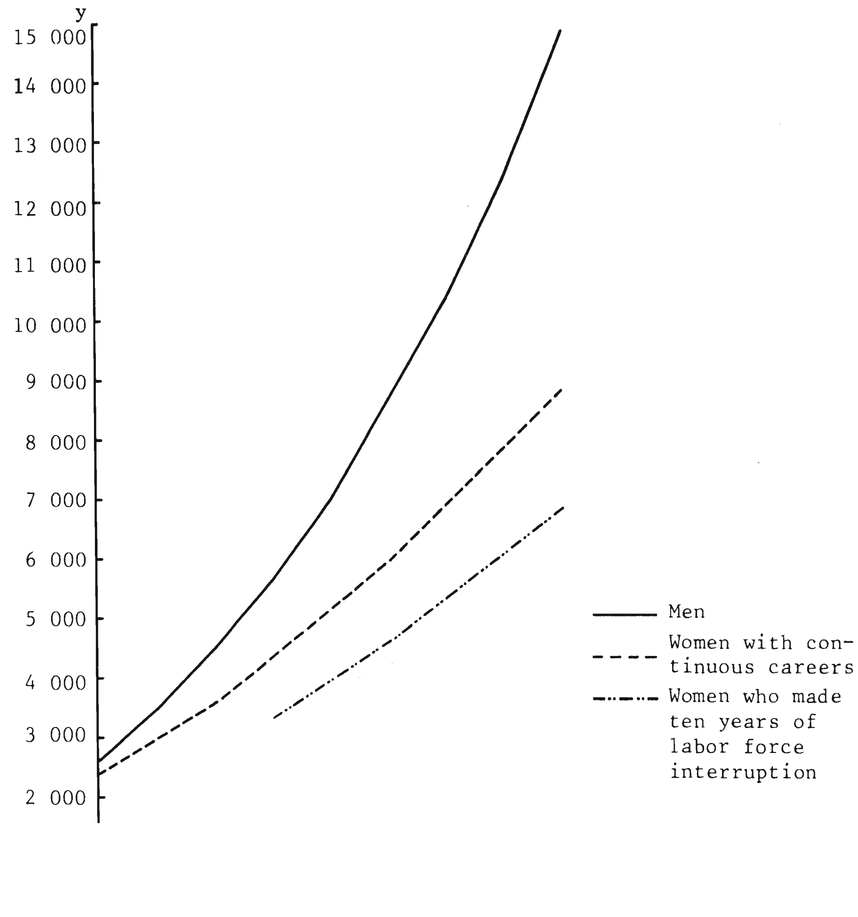


Figure 1B. Predicted earnings in Skr from regressions estimates (6) and (7)



7. MALE-FEMALE EARNINGS DIFFERENTIALS

One of the purposes of this paper is to determine whether earnings differentials between the sexes are smaller in Sweden than in the United States. In comparing the earnings differentials we want to assess what part of it is due to different coefficients of the earnings function and what part is due to different work histories.

A comparison of equations (6) and (7) of Table 3 shows that there is a great deal of similarity in the earnings functions for men and women. However, earnings of men rise at a steeper rate than earnings of women. The difference in steep is:

$$\left[\frac{\delta \ln y}{\delta a} + \frac{\delta \ln y}{\delta e_\ell}\right]^M - \left[\frac{\delta \ln y}{\delta a} + \frac{\delta \ln y}{\delta e_\ell}\right]^F = 0.0103 \qquad (14)$$

i.e., male earnings of the continuous career increase by 1% per year faster than female earnings of the continuous career. At age 60 women following a continuous career earn 59% of what men earn according to these estimates. If they made one labor force interruption of 10 years they earn 46% of what men following a continuous career earn at age 60. These estimates are in accordance with the general observed pattern of age-earnings curves (see Gustafsson, 1976). The difference in the increase in earnings power by a marginal year of experience between American white men and white women reported by Corcoran and Duncan (1979) is smaller than the differential between Swedish men and women reported here. Mincer and Polachek (1974 and 1978) and Sandell and Shapiro (1978) do not report male earnings functions.

The mean earnings differential may be decomposed according to:

$$\ln(\bar{y}^M) - \ln(\bar{y}^F) = (C^M - C^F) + (\bar{X}^M - \bar{X}^F)b^M + \bar{X}^F(b^M - b^F) \tag{15}$$

where the proportion explained by different characteristics is:

$$(\bar{X}^M - \bar{X}^F)b^M$$

The C:s are estimated intercepts, the b:s are vectors of estimated regression coefficients for men and women respectively, the X:s are vectors of mean characteristics of men and women respectively. The decomposition may also be done by comparing earnings standardized at the means of male work history data and using the female regression equation in the standardizing procedure:

$$\ln(\bar{y}^M) - \ln(\bar{y}^F) = (C^M - C^F) + (\bar{X}^M - \bar{X}^F)b^F + \bar{X}^M(b^M - b^F) \tag{16}$$

Earnings differentials between the sexes in Sweden and the U.S. can now be compared using the above formulas for the estimates of Table 3 equations (6) and (7) and for the data reported in Corcoran and Duncan.

The results are given in Table 6. Total earnings differentials between men and women are only slightly smaller in Sweden than in the U.S.

The regression estimates of Table 3 are less similar between the sexes than is the case for the American data. This results in different standardized earnings differentials when the standardiza-

Table 6. Earnings differentials between men and women in Sweden and the U.S.

	Sweden 1974 Regression estimates (6) and (7)		U.S. 1975 Corcoran and Duncan (1979)	
Total differential				
$\ln(y^M)-\ln(y^F)$	0.400	0.400	0.438	0.438
Explained differential				
Schooling:[a]				
$b^F(S^M-S^F)$	0.0587		0.009	
$b^M(S^M-S^F)$		0.0749		0.007
Work history:				
$b^F(X^M-X^F)$	0.0803		0.155	
$b^M(X^M-X^F)$		0.1027		0.157
Age:				
$b^F(A^F-A^M)$	0.0337			
$b^M(A^F-A^M)$		0.0519		
Residual standardized differential				
$X^M(b^M-b^F)+C^M-C^F$	0.2273		0.274	
$X^F(b^M-b^F)+C^M-C^F$		0.1725		0.2744
Female earnings in percent of male earnings				
a) total differential	67	67	65	65
b) standardized by schooling	71	72	65	65
c) standardized by schooling and work history	77	80	(76)	(76)
d) standardized by schooling, work history and age	80	84	76	76

[a] S and A are components of the vector of X:es in (15)
S = the average number of years of schooling
A = the average age.

tion is done at the means of male characteristics than from a standardization at the mean of female charasteristics. Differentials in schooling in Sweden reduces the earnings differentials whereas it has almost no effect for the U.S. The effect of age is sizable in the Swedish data. The female employees are considerably younger than the male employees.

The American data includes full work histories which means that a comparison with age included in the Swedish standardization is relevant.[1]

In the lower panel of Table 6 the results of standardization are given in percent. Swedish women earn 67% of what Swedish men earn and the corresponding proportion for the U.S. is 65%.[2] For the same age and work history female earnings as a percent of male earnings is 80-84% for Sweden and 76% for the U.S. The Swedish standardized earnings differential is rather an overestimate than an underestimate because individuals may have had hometime during the unobserved period that if taken account of would have decreased the standardized differential still more. Furthermore we are dealing exclusively with private sector white collar workers and the government sector shows smaller sex differentials in earnings (Gustafsson, 1976).

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[1] Since $\Sigma h + \Sigma e + s + 7 = a$ for individuals for whom we know the full work history.

[2] $e^{-0.4} = 0.67$, $e^{-0.438} = 0.645$.

8. HOW MUCH DOES LABOR FORCE HISTORY EXPLAIN?

The smaller accumulation of human capital on the part of women explains between 43 and 57% of the total male-female earnings differentials and labor force history variables make up the larger part of this. (See Table 6).[1]

In estimating the effect of labor force interruptions there is a simultaneity problem. Are lower earnings the result of labor force interruptions or do people that have low earnings drop out of the labor force more frequently? It is certainly less costly to drop out of the labor force if the alternative earnings foregone are smaller than if they are larger. One can thus read the earnings function with causation running the opposite way as a lifetime labor supply function.

One way of coping with the simultaneity problem is to employ two stage least squares (TSLS). The simultaneity problem and its implications are discussed and TSLS estimates are given in Mincer and Polachek (1974, 1978), and in Sandell and Shapiro (1978). The idea in TSLS is to find instrument variables that are uncorrelated with the stochastic terms. Mincer and Polachek use health, geographical area of residence, number of children, own education and husband's education as exogenous instrumental variables. They also remark that the occupational choice may be simultaneous. All of their TSLS estimates come out with more deprecia-

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[1] (0.0587 + 0.0803 + 0.0337)/0.4 = 0.1727/0.4 = 0.43175 = 43% or alternatively: (0.0749 + 0.1027 + 0.0519)/0.4 = 0.2295/0.4 = 0.5737 = 57%.

tion than OLS estimates do and consequently with a larger part of the male-female earnings differentials explained by the models.

In Table 7 instrumental equations for hometime and experience time are given with the second stage estimates of the earnings function. Most coefficients of the instrumental equations are significant and have expected signs. Schooling also reduces experience, a result probably due to the truncated work history data. The occupations do not have the expected effects. Being a general office worker significantly reduces hometime contrary to expectation. Total previous (MEANATP) earnings has the expected effects. Being a general office worker significantly reduces hometime and increases experience time. Age increases hometime which is consistent with the observation that older cohorts participate less in the labor force (Gustafsson, 1980) but it also increases experience probably because of the truncation of the work history data.

Living in a big city does not have any effect. Estimates not reported here have shown that the effect of CITY differs between the first hometime period of delayed entrance to the labor market (h_1) for which it is insignificant and for periods of withdrawal from the labor market after periods of experience (h_2, h_3, etc.), for which it is positively significant.

The structural earnings equation of the TSLS estimates differs sharply from earlier OLS results in that the coefficient of home time is negative, large in absolute value and strongly significant.

Table 7. TSLS-estimates of earnings functions for full-time working women

Structural equation		Instrumental equations			
Dependent variable	ln(y)	Dependent variable	h	e	ln(y)
Intercept	7.1125* (0.0288)	Intercept	-4.3760* (0.3761)	-12.148* (0.3991)	6.6502 (0.0166)
s	0.0522* (0.0025)	s	-0.0647* (0.0207)	-0.4738* (0.0230)	0.0104* (0.0009)
h	-0.0685* (0.0052)	OFFICE	-0.1869* (0.0731)	0.0900 (0.0776)	-0.0280* (0,0032)
e	0.0470* (0.0017)	SECR	-0.0046 (0.0890)	0.0326 (0.0945)	0.0175* (0.0004)
e^2	0.0003* (0.0000)	MEAN-ATP	-0.0067* (0.0004)	0.0074* (0.0005)	0.0017* (0.0000)
R^2	0.5919	AGE	0.4551* (0.0187)	1.1927* (0.0199)	0.0457* (0.0008)
n	6 745	AGESQ	-0.0056* (0.0002)	-0.0123* (0.0003)	-0.0005* (0.0000)
		CITY	0.1135 (0.0632)	-0.0880 (0.0670)	0.0270* (0.0028)
		R^2	0.0977	0.6877	0.8079
		n	6 745	6 745	6 745

*Significant at the 5% level.

MEANATP = $(\Sigma\ ATP_i)/15$, i = 1960, 1961,...,1974 for all years including zeroes.

OFFICE = dummy variable = 1 if the woman is coded "general office worker".

SECR = dummy variable = 1 if the woman is coded "secretary".

CITY = dummy variable = 1 if the woman lives in Stockholm, Gothenburg or Malmö with suburbs.

9. CONCLUDING REMARKS

Sweden has a very modern family legislation with separate taxation of husband's and wife's earnings and the right for fathers as well as mothers to take paid parental leaves for child care as important and internationally advanced features. This study has shown that when it comes to earnings differentials between men and women Sweden is only slightly more equal than the U.S. Average earnings of women are 65% of average earnings of men in the U.S. and the corresponding figure for Sweden is 67%.

If women can by their own choices affect their earnings this is less serious than if they can not. A reason for their lower earnings can be that they have chosen a smaller lifecycle labor force participation and correspondingly smaller investments in human capital. It turns out that in Sweden women earn 80-84% of what men earn if schooling and labor force history are held constant. American women earn 76% of what American men earn when schooling and labor force history are held constant.

Swedish women seem to have worked a larger proportion of the available time than American women.

Appendix 1. THE TIME SERIES OF ATP SCORES USED TO CALCULATE EARNINGS

In this paper ATP-scores have been used only to build variables on the work history. We have not used the information on earnings contained in this variable. Panel data on earnings helps in finding a better solution to the simultaneity problem: "If wage setbacks due to interruptions were attributable solely to the loss of work experience (foregone on the job investments) wages of returnees would be lower than wages of stayers but not lower than their own wages at exit." (Mincer and Ofek, 1980).

ATP is a truncated function of annual earnings but we do not know how many hours the individual has worked. The range of the ATP is 0 through 650 because earnings higher than 6.5 times the base amount are not included.

In Table A.1 the average ATP at the start of different work history intervals has been calculated in order to assess its relevance as an earnings variable. At the start of the first experience period the average ATP-score for women aged 16-64 is 99.9. At the end of the first work experience period, however, ATP is lower than at the start. If the earnings potential increased during the first work experience period, earnings per hour would be higher at the end of the first work experience period than is the case at the end of this period. However, it is very likely that the person who enters a period of home time finished working in the market sometime in the middle of the last year for which we observe earnings.

Earnings at the start of the second work history period are in most cases higher than earnings at the start of the first experience period. The conclusion of this exercise is that the time series earnings information in the ATP scores is useless as long as hours of work are not known.

Table A.1 Mean earnings according to work history for fulltime female workers in the private sector

	n	Mean earnings at the start of				
		e_1	h_2	e_2	h_2	e_3
Age 16-64						
All women	6 745	99.9	84.3	118.6	85.5	135.1
Compulsory education	5 110	96.4	81.6	112.0	82.0	124.3
Secondary education	1 432	106.0	92.5	135.8	99.3	168.4
University education	203	145.0	96.0	159.1	73.0	186.8
Age						
30-34	854	66.4	100.2	117.7	105.7	123.2
35-39	584	108.0	80.0	117.7	105.7	123.1
40-44	600	126.8	114.7	104.9	47.5	65.0

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On the Determinants of Labor Supply in Sweden

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1. INTRODUCTION

The purpose of this paper is to analyze how relative prices (wage rates), nonemployment income, and the tax structure affect the quantity of labor supplied in the Swedish labor market. The paper contains to our knowledge the first empirical analysis of the neoclassical labor supply function on Swedish household data.[1] This is a non-attractive orginality, which has its roots in, until recently, a lack of suitable empirical data.

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[1] For a survey in Swedish, compare Axelsson, Jacobsson and Löfgren (1979). Estimations of labor supply functions on individual data have been made by Blomqvist (1979).

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The structure of the paper is the following: In Section 2 we give a short summary of the history of the supply of labor in economic theory. Section 3 contains a summary of how the supply of labor is treated in contemporary economic analysis and how certain theoretically and empirically interesting problems are handled. In Section 4 we describe the contents of our data source and how the theoretical determinants of labor supply have been operationalized in our models. The empirical results are presented in Sections 5 and 6. The content of Section 6 is especially interesting as we are able to take explicit account of how an increased tax progression affects the labor supply. The paper ends with a few comments on how the analysis can be further developed.

2. BACKGROUND

The theoretical foundations of the supply of labor are of approximately the same age as the marginal analysis and utility theory. They can be found in almost contemporary shape in the work of Stanley Jevons (1879). Before that date the supply of labor was mostly referred to in terms of the number of workers as is the case in the works of Malthus (1793) and Ricardo (1821). Ricardo's "iron law of wages" is, e.g., based on the assumption that the number of workers is determined from Malthus principle of population.

The marginalists before Jevons are very brief in their treatment of the supply of labor. This is the case in Longfield (1834), Lloyd (1883) and

Jennings (1855). Jenkin (1870) introduced the labor supply function without attempting to base it on any fundamental economic assumptions.

In The Theory of Political Economy (1870) Jevons investigates the labor supply of an individual in utility-theoretical terms. He derives no explicit supply function, but the supply function is of course implicit in the marginal conditions (first-order conditions) for equilibrium.

Marshall introduces in his Principles of Economics (1890), a more explicit derivation of the labor supply function[1] and gives rather elaborate discussions of the determinants of the supply of labor. In his formal derivation he uses rather restrictive assumptions on the properties of the "utility function" and these enable him to derive a labor supply function that is a non-decreasing function of the wage rate.[2]

The possibilities of both a backward bending and a forward bending supply function is for the first-time brought up by Lionel Robbins (1930).[3] In his paper Robbins shows that a backward bending supply function is a theoretical possibility, but not a

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[1] For a more comprehensive discussion of Marshall's contribution to the supply of labor compare Walker (1975). In this connection it should be mentioned that Walras (1874) contains a discussion of the supply of labor, but this discussion is far from as thorough as the discussion in Marshall (1890).

[2] Marshall maximizes some sort of quasi-rent that equals income from work minus the disutility of work.

[3] The problem is also treated in Barone (1899).

fact as Frank Knight (1921) and Arthur C. Pigou (1920) seemed to believe. In his book Risk, Uncertainty and Profit Knight introduces arguments for why it would be rational for the individual to reduce the supply of labor when the wage rate increases, and in The Economics of Welfare Pigou indicates that higher (proportional) taxes imply an increased supply of labor.

As is well-known today, the problem can be transferred to the so-called Slutsky-decomposition of the effects of a price change. If the so-called income effect dominates the substitution effect the supply curve will be backward bending, while the opposite is the case if the substitution effect dominates the income effect. Robbins was the first to introduce this analysis into a labor market context, but the results are of course implicit already in the analysis of Slutsky (1915).

The theory that we have been concerned with up till now, tries to explain the supply decision of a single individual. In other words, one assumes implicitly that each family member has a utility function of his own, which is independent of the consumption and supply decisions of other family members.

Inspired of works by Lewis (1956), Mincer (1962) and Becker (1965), Marvin Kosters (1966) makes a "neighboring" development of the theory. He regards the family as a unit that possesses one utility function, which is maximized subject to a budget constraint that contains family income as revenue and family consumption as outlay. Variations on this theme have dominated theory since then, and our main approach is based on Kosters' original model.[1]

3. THE THEORY OF LABOR SUPPLY

The household is assumed to possess the utility function

$$U = U(x_1, x_2, x_3, \dots, x_n) \qquad (1)$$

which is assumed to be quasi-concave and twice continuously differentiable. We will use the convention that

$x_1 = T - \ell_1$ = the leisure of the husband in hours

$x_2 = T - \ell_2$ = the leisure of the wife in hours

where T = the total available number of hours

ℓ_1 = labor supplied by the husband in hours

ℓ_2 = labor supplied by the wife in hours

x_i = the consumption of other goods than leisure: $(i = 3, \dots, n)$

The utility function (1) is assumed to be maximized subject to the constraint

$$\Sigma p_i x_i = C \qquad (2)$$

where $C = p_1 T + p_2 T + \pi$

π = income of the family from nonemployment sources

p_i = the prices of consumer goods $(i=1, \dots, n)$

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[1] Compare, e.g., Ashenfelter and Heckman (1974), Wales and Woodland (1976, 1977, 1979), and Rosen (1976a, 1976b).

As we are not interested in the demand for commodities other than leisure we will reduce the number of such commodities from n-2 to one. Maximization of (1) subject to the constraint (2) results in the familiar first-order conditions

$$U_i - \lambda p_i = 0; \qquad i = 1,2,3 \tag{3}$$

$$C - \sum_{i=1}^{3} p_i x_i = 0$$

An interior solution of these equations will, due to the assumed quasiconcavity of the utility function, be optimal. The conditions (3) can be interpreted as follows: The marginal utility of the last dollar spent is independent of the commodity on which it is spent, and equals the Lagrange multiplier λ. This multiplier can be shown to be the marginal utility of additional nonemployment income.

The solution of (3) will give us demand functions of the form:

$$x_i = x_i(p, \pi) \qquad i = 1,2,3 \tag{4}$$

$$p = (p_1, p_2, p_3)$$

The supply of labor for husband and wife can therefore be written

$$\ell_i = T - x_i(p, \pi) \qquad i = 1,2 \tag{5}$$

Hence the supply of labor for both members of the household will depend on all prices and nonemployment income. It is also easily seen that the supply functions (demand for leisure) are homogeneous of degree zero in prices and nonemployment income, i.e., if these parameters are multiplied by a common constant this does not change the behavior of the household. The constant can be extracted from the budget constraint and the feasible choice set therefore remains unchanged. It follows that:

$$x_i(p,\pi) = x_i(kp,\ k\pi) \qquad \begin{array}{l} i = 1,2,3 \\ k > 0 \end{array} \tag{6}$$

If (6) is differentiable we can derive the following "aggregation" condition[1]

$$0 = \sum_{j=1}^{3} p_j \frac{\partial x_i}{\partial p_j} + \pi \frac{\partial x_i}{\partial \pi} \qquad i = 1,2,3 \tag{7}$$

If each term in (7) is divided by x_i we obtain after change of notation

$$\sum_{j=1}^{3} \eta_{ij} + \eta_{i\pi} \tag{8}$$

which can be interpreted in the following way: The sum of the own price elasticity, the cross price elasticities and the income elasticity of a commodity is zero.[2]

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[1] Differentiate both members with respect to k, when k=1.

[2] For commodities i=1,2 equation (7) is formulated in leisure terms, but it should be obvious that similar conditions hold in supply terms.

For more details on the properties of the demand and supply functions we differentiate (3) totally to obtain

$$\begin{bmatrix} 0 & -p_1 & -p_2 & -p_3 \\ -p_1 & U_{11} & U_{12} & U_{13} \\ -p_2 & U_{21} & U_{22} & U_{23} \\ -p_3 & U_{31} & U_{32} & U_{33} \end{bmatrix} \begin{bmatrix} d\lambda \\ dx_1 \\ dx_2 \\ dx_3 \end{bmatrix} = \begin{bmatrix} \sum_{i=1}^{2} (x_i - T)dp_i + x_3 dp_3 - d\pi \\ \lambda dp_1 \\ \lambda dp_2 \\ \lambda dp_3 \end{bmatrix} \tag{9}$$

If A denotes the matrix of the system we obtain when solving (9) for $dp_j = d\pi = dp_3 = 0, \quad j=1,2.$[1]

$$\frac{\partial x_i}{\partial p_i} = \lambda \frac{A_{ii}}{\det A} - \ell_i \frac{A_{0i}}{\det A} \qquad i = 1,2 \tag{10}$$

or more generally

$$\frac{\partial x_j}{\partial p_i} = \lambda \frac{A_{ij}}{\det A} - \ell_i \frac{A_{0j}}{\det A} \qquad i,j = 1,2,3 \tag{11}$$

If we investigate the solution for $dp_i=0$, $i=1,2,3$ we obtain:

$$\frac{\partial x_j}{\partial \pi} = - \frac{A_{0j}}{\det A} \qquad j = 1,2,3 \tag{12}$$

This derivative expresses how the demand for leisure or the commodity is changed when nonemployment income increases (ceteris paribus). Leisure

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[1] We assume that det $A \neq 0$. This does not follow from the assumptions made, det $A=0$ can be the case in a non dense subset of R^3. A_{ij} denotes the cofactor that is obtained by deleting from A the i:th row and the j:th column.

is said to be a normal good if the income effect is positive. This is usually considered to be the case and it means, e.g., that a transfer of nonemployment income to a household (<u>ceteris paribus</u>) decreases the supply of labor of both members of the household.

It is easily shown that the effect on the demand for a commodity of a change of a commodity price, provided that the household receives income compensation that enables it to stay on the initial utility level, is expressed by the derivative

$$\left(\frac{\partial x_j}{\partial p_i}\right)_{u=k} = \lambda \frac{A_{ij}}{\det A} \qquad i,j = 1,2,3 \tag{13}$$

This effect is usually called the substitution effect or the compensated price effect. The total effect of a change of the price of leisure can hence be decomposed into a substitution effect and an income effect. We can write

$$\frac{\partial x_j}{\partial p_i} = \left(\frac{\partial x_j}{\partial p_i}\right)_{u=k} + \ell_i \left(\frac{\partial x_j}{\partial \pi}\right)_{p=k} \qquad i,j = 1,2,3 \tag{14}$$

From the fact that the household is assumed to maximize utility it follows that [1]

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[1] A necessary condition for maximum is that the quadratic form

$$\sum_{i=1}^{3} \sum_{j=1}^{3} U_{ij} dx_i dx_j \leqslant 0 \quad \text{(negative semi-definite)}$$

given that dx_i, dx_j are chosen arbitrarily, but in such a way that the budget constraint is satisfied. A sufficient condition for this is that the principal minors of det A alter signs and (15) follows. In the strict sense, both det A and A_{jj} can be zero on a non dense subset of R^3, i.e., in single points.

$$\left(\frac{\partial x_j}{\partial p_j}\right)_{u=k} = \frac{A_{jj}}{\det A} < 0 \qquad j = 1,2,3 \tag{15}$$

which means that the own substitution effect of a price increase always is negative. Expressed in supply terms this means that the substitution effect of a wage increase increases the supply of labor (decreases the demand for leisure). If leisure is a normal good it follows by definition that a wage increase decreases the supply of labor via the income effect. To sum up: The theory provides no *a priori* information whether a wage increase increases or decreases the supply of labor. In other words the supply curve can be either negatively or positively sloped.

Up till now we have not explicitly considered taxes, but it is easily seen that p_1 and p_2 can be interpreted as net wages if the tax is proportional. It then immediately follows, that a tax increase decreases the supply of labor via the substitution effect, and increases the supply of labor via the income effect.[1] That is, a tax increase tends to increase the supply of labor if the supply curve is negatively sloped, and tends to decrease the supply of labor if the supply curve is positively sloped.[2] It can also be shown

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[1] It might, however, be argued that there will be no income effect if the tax rate change is perceived as transitory.

[2] Note that if the tax rate is changed this affects both net wage rates and we have to take account of two substitution effects (the own substitution effect and the cross substitution effect) and two income effects. Therefore the word "tends" in the statement above.

that a similar result holds under a progressive tax system.[1]

There are, however, more direct methods to analyze the effects on labor supply of changes in tax progression. If we describe the tax system by a parameterization Θ we will obtain labor supply functions that can be written

$$\ell_i = \ell_i(p, \pi, \Theta) \qquad i = 1,2 \tag{16}$$

and on a suitable empirical material it might be possible to estimate empirical specifications of these functions. The problem is usually that the households face the same tax system in a cross section material, and, hence, the same parameters Θ. In our data, however, the municipal tax differs between the households and it is theoretically possible to get a direct estimate of how changes in tax progression affect the supply of labor.

We have thus derived practically all empirically meaningful theorems embedded in conventional neo-classical theory. There remains, however, one contra-intuitive conclusion. It holds that an increase of the husband's wage rate affects the compensated supply of the wife in exactly the same way as a change of the wife's wage rate affects the compensated supply of the husband. This result follows from the symmetry of A. Formally

$$\frac{A_{ji}}{\det A} = \frac{A_{ij}}{\det A} \qquad i,j = 1,2,3 \tag{17}$$

or

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[3] See, e.g., Axelsson, Jacobsson and Löfgren (1979).

$$\left(\frac{\partial x_1}{\partial p_2}\right)_{u=k} = \left(\frac{\partial x_2}{\partial p_1}\right)_{u=k} = -\left(\frac{\partial \ell_1}{\partial p_2}\right)_{u=k} = -\left(\frac{\partial \ell_2}{\partial p_1}\right)_{u=k} \qquad (17a)$$

The sign of (17) is a priori ambiguous, but it is customary to classify commodities as substitutes, neutral or complementary according to the sign of the cross substitution effect. According to this classification scheme we say that the husband's and the wife's leisure are substitutes if

$$\left(\frac{\partial x_1}{\partial p_2}\right)_{u=k} = \left(\frac{\partial x_2}{\partial p_1}\right)_{u=k} > 0$$

complementary if (18)

$$\left(\frac{\partial x_1}{\partial p_2}\right)_{u=k} < 0$$

neutral if

$$\left(\frac{\partial x_1}{\partial p_2}\right)_{u=k} = 0$$

We are now ready to suggest an empirical specification of the supply function. As has been shown all prices (wages) and nonemployment income are the theoretical determinants of both the husband's and the wife's supply of labor. Disregarding commodity prices we obtain:

$$\ell_1 = f(p_1, p_2, \pi) + \varepsilon_1$$

$$\ell_2 = g(p_1, p_2, \pi) + \varepsilon_2 \qquad (19)$$

where ℓ_i is an error term. Assuming the functions

(19) to be linear[1] and with the tax system explicit we obtain the following supply functions:

$$\ell_1 = \alpha_0 + \alpha_1 p_1 + \alpha_2 p_2 + \alpha_3 \pi + \alpha_4 t + \varepsilon_1$$
$$\ell_2 = \beta_0 + \beta_1 p_1 + \beta_2 p_2 + \beta_3 \pi + \beta_4 t + \varepsilon_2 \qquad (19a)$$

where t = the municipal tax rate.

The theory provides no a priori hypothesis concerning the signs of the coefficients. According to equation (14) it should, however, hold that

$$\left(\frac{\partial \ell_1}{\partial p_1}\right)_{u=k} = \alpha_1 - \bar{\ell}_1 \alpha_3 > 0$$
$$\left(\frac{\partial \ell_2}{\partial p_2}\right)_{u=k} = \beta_2 - \bar{\ell}_2 \beta_3 > 0 \qquad (20)$$

where ℓ_i is an estimate of the mean in the population under consideration.[2] If leisure is a normal good it should in addition hold that:

$$\frac{\partial \ell_1}{\partial \pi} = \alpha_3 < 0; \qquad \frac{\partial \ell_2}{\partial \pi} = \beta_3 < 0 \qquad (21)$$

If the husband's and the wife's leisure are complementary goods it holds that:

$$\left(\frac{\partial \ell_1}{\partial p_2}\right)_{u=k} \approx \alpha_2 - \bar{\ell}_2 \alpha_3 > 0$$
$$\left(\frac{\partial \ell_2}{\partial p_1}\right)_{u=k} \approx \beta_1 - \bar{\ell}_1 \beta_3 > 0 \qquad (22)$$

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[1] The choice of a linear model is somewhat arbitrary. See, e.g., Dickinson (1975a) for a discussion of its implications for the underlying preference structure.

[2] Cf Dickinson (1975b) on problems in using $\bar{\ell}_i$ in the estimation of the substitution effect.

and

$$\left(\frac{\partial \ell_1}{\partial p_2}\right)_{u=k} = \left(\frac{\partial \ell_2}{\partial p_1}\right)_{u=k} \tag{23}$$

Equality (23) will hold only in the case of a single household utility function.

In the empirical estimations presented below we have used variations of equations (19). In addition to the variables indicated certain control variables are used to take account of differences in "tastes" and commodity prices.[1]

4. THE DATA

In this section we make a short presentation of the data source used, the criteria for selecting the sample, and finally we briefly describe the variables used.

The estimations presented in this paper are based on data from the annual Income Distribution Survey conducted by the Swedish Central Bureau of Statistics.

The survey from 1976, which we have used, covers a sample of roughly 10,000 households (or 28,000 individuals). The main part of this data source consists of income and other money variables collected from various government records. This information is supplemented with supply variables

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[1] "Taste" is, e.g., controlled by age dummies and commodity prices by region dummies.

(e.g., weeks in employment and average hours worked per week) and information on profession and place of work. These labor market variables are collected by means of a questionnaire.

The Income Distribution Survey provides information about all members of the household. Another advantage lies in the very detailed specification of income and transfer variables. Furthermore, the fact that these figures have been collected from government records ensures a high degree of accuracy (although the problem of tax evasion of course still remains). The main drawback of the data source used concerns the quality of the supply include errors. However, a great deal of effort has been made in checking the consistency of these data and some obvious errors (due to misunderstanding of the questionnaire) have been corrected.

The regression equations are estimated on households with two adults. All households receiving any entrepreneurial income (from farming, business or apartment houses) have been excluded.

Quality aspects on the supply and wage variables have necessitated the exclusion of households if the model has included the supply or wage variable of a household member who had an employment income less than 5,000 Skr per year, received parent or labor market assistance, educational grants or pensions. For similar reasons teachers, clergymen and home day care personnel have been excluded. This means that the models with both husband's and wife's wages are estimated on a smaller number of observations than models including, for example, only husband's wage. The equations are estimated on persons aged 20-64.

The dependent variable has been computed as the product of the number of weeks the individual received employment income in 1976 (including vacation, sick leave and other payed absence) and the average number of hours worked per week. This supply concept thus measures the number of hours employed per year and not the actual hours worked. As mentioned above this variable may be erroneously reported. People with varying working hours per week may have had problems in estimating a correct average.

The wage variable has been computed as the quotient between gross income from employment (including sickness payments) and the supply variable.

Nonemployment income is measured as the sum of income of capital and net income of owner occupied dwellings. In one model we have omitted the spouse's wage and instead added his/her employment income to nonemployment income.

The age dummies have been defined as follows for husband (m) and wife (k),

m_1, k_1	for ages	20 - 24
M_2, k_2	"	25 - 34
m_3, k_3	"	35 - 44
m_4, k_4	"	45 - 54
m_5, k_5	"	55 - 64

with the ages 55-64 as the reference category. Regression equations have also been estimated separately for the five age groups.

Presence of pre-school children has been controlled for with the dummies

c_1 for households with at least one child younger than three years,

c_2 for households with at least one child between three and six years old, but no child under three years,

c_3 for all other households.

The last category has been used as reference group.

The regional dummies are based on the so called "H-region" concept with the following groupings.

h_1 for the three largest cities (Stockholm, Göteborg and Malmö)

h_2 for other large cities

h_3 for other areas of Southern and Mid-Sweden

h_4 for other areas of Northern Sweden

The fourth category has been used as reference group.

5. EMPIRICAL RESULTS

In this section we report the regression results.[1] We begin with an estimation of equations (19a) on

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[1] The data have been weighted inversely to the sampling frequencies in different strata.

a sample that contains all households where both spouses perform work in the labor market. As control variables we use dummies for children in different age classes, regional dummies, and age dummies. Nonemployment income is measured as the sum of income from capital and net income of owner occupied dwellings.[1]

The estimated versions of equations (19a) are given below.[2]

$$\ell_1 = \underset{}{2104} - \underset{(3.191)}{0.0172\, p_1} - \underset{(0.586)}{0.0042\, p_2} - \underset{(5.000)}{0.0046\, \pi^T} + \underset{(1.228)}{25\, c_1} +$$

$$\underset{(2.914)}{55\, h_1} + \underset{(2.248)}{42\, h_2} + \underset{(1.767)}{35\, h_3} - \underset{(4.169)}{118\, m_1} - \underset{(4.516)}{59\, m_2} - \underset{(0.377)}{5\, m_4}$$

$$R^2 = 0.04 \qquad F = 6.184 \qquad n = 1414$$

$$(24)$$

$$\ell_2 = 1866 - \underset{(2.369)}{0.0367\, p_1} - \underset{(8.978)}{0.1837\, p_2} - \underset{(2.509)}{0.0070\, \pi^T} + \underset{(7.491)}{455\, c_1} -$$

$$\underset{(5.717)}{266\, c_2} + \underset{(3.697)}{136\, h_1} + \underset{(1.679)}{60\, h_2} + \underset{(6.408)}{404\, k_1} + \underset{(3.362)}{191\, k_2} +$$

$$\underset{(1.106)}{58\, k_3} + \underset{(1.691)}{88\, k_4}$$

$$R^2 = 0.13 \qquad F = 19.598 \qquad n = 1407$$

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[1] Income from owner-occupied dwellings has beeen defined as the schematically computed income used in tax filing less deductions for interest payments.

[2] Numbers in parentheses below coefficients are absolute values of the t-statistic.

where

ℓ_1 = Labor supplied by the husband in hours per year

ℓ_2 = Labor supplied by the wife in hours per year

p_1 = Gross wage rate of the husband in Swedish kronor/100 per hour

p_2 = Gross wage rate of the wife in Swedish kronor/100 per hour

π^T = Nonemployment income in Swedish kronor (capital income and net property income)

$m_1, \ldots, m_4$, $k_1, \ldots, k_4$ = Dummy variables controlling for age differences

c_1, c_2 = Dummy variables controlling for presence of pre-school children.

h_1, h_2, h_3 = Dummy variables controlling for geographic location of the household

The R^2-values are low, but all purely economic variables (p_1, p_2, π^T) are significant[1], except p_2 in the first regression. We can see from the first equation that c_2 (children of age 3-6 years) is not included, and that c_1 (children of age 0-2 years) is insignificant.[2] In other words, the presence of children does not affect the labor supply of the husbands.

From the age dummies it is obvious that young men (m_1 and m_2) work significantly less than the oldest generation (55-64), which serves as reference group. No other age group among men differs significantly from the reference group. This

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[1] When we speak of significance we mean significance at the 5-percent probability level, two-tailed test.

[2] The regression program used excludes variables with an F-statistic under a certain minimum level.

result is not inconsistent with the results from recent developments of life-cycle theory.[1]

The situation is a bit different among wives. The presence of children decreases the labor supply of "married" females. Both c-coefficients are negative, and significantly different from zero. This result is wholly in line with conventional wisdom, and results from previous studies.[2]

A bit surprising, however, is that young females work significantly more than the reference group (55-64).[3]

If we return to the purely economic variables we find that both supply curves are backward bending, i.e., the own price elasticities are negative. The elasticity for husbands is rather low, -0.025, which is clearly lower than the wage elasticities found by Kosters (1966). The wives are a lot more wage sensitive as can be seen from the value of the wage elasticity, -0.30.[4]

From the theory it follows that the substitution effects or the compensated wage rate effects are positive. A check reveals that this is not generally the case in our estimates. In Table 1 we have summarized the calculations.

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[1] Compare Blinder and Weiss (1976), and Ryder, Stafford and Stephan (1976).

[2] Compare, e.g., Gramm (1975).

[3] It should be noted that we here do not discuss participation rates, but only working time among those already in the labor force. We thus do not take account of a (probably) positive participation effect of higher wages.

[4] The elasticities are computed at the mean values.

Table 1. Results computed from equation (24)

Wage elasticity		Own substitution effect		Cross substitution effect	
Husband	Wife	Husband	Wife	Husband	Wife
-0.025	-0.30	7.85	-8.15	6.29	10.90

The compensated wage effect on the wife's supply of labor is negative and, hence, seems to refute the theory. We will, however, hint at a possible explanation to this anomaly in Section 6 below.

The cross substitution effects are positive, which indicates that the leisure of both spouses are complementary goods. According to theory the cross substitution effects should be equal. A t-test assuming independence reveals that the seemingly large difference between the cross substitution effects is not significantly different from zero.[1]

Finally, it is clear that leisure is a normal good; the nonemployment coefficient is negative and significant.

To make a comparison between different age categories regarding the sensitiveness of labor supply to the economic variables possible, we have estimated equations (19a) for each sex and five different age groups within each sex. The results are summarized in Table 2.

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[1] This indicates that the estimation could be made non-biased, and more efficient by using this theoretical relationship. Compare Ashenfelter and Heckman (1974), and Abbott and Ashenfelter (1976).

Table 2. Regression results for different age groups

	Constant	p_1	p_2	π^T	c_1	c_2	h_3	h	h	R^2	F	n	Substitution effect
Husband													
20–24	2202	0.0208	−0.0737	−0.0293	−51	−414	−121	−120	391	0.26	1.159	35	60.39
		(0.136)	(0.512)	(1.076)	(0.411)	(2.490)	(0.447)	(0.486)	(1.308)				(1.07)
25–34	2200	−0.0839	−0.0113	−0.0066	24	18	94	110	51	0.12	5.702	331	5.17
		(5.774)	(0.633)	(3.385)	(0.761)	(0.646)	(2.047)	(2.409)	(1.059)				(1.21)
35–44	2024	−0.0038	−0.0068	−0.0064	15	22	80	85	84	0.04	2.252	422	13.11
		(0.313)	(0.526)	(3.282)	(0.292)	(0.678)	(2.150)	(2.360)	(2.095)				(3.06)
45–54	2122	−0.0054	−0.0050	−0.0012	−28	45	8	−24	−21	0.01	0.665	413	1.96
		(0.751)	(0.398)	(0.769)	(0.294)	(1.200)	(0.275)	(0.826)	(0.713)				(0.59)
55–64	2100	0.0058	−0.0103	0.0006		−193	6	−17		0.12	4.504	213	−0.67
		(0.924)	(1.076)	(0.034)		(4.191)	(0.355)	(0.993)					(0.18)
Wife													
20–24	2109		−0.1902	−0.0247	−709	−287	323	345	−66	0.59	11.588	65	−61.65
			(1.720)	(1.980)	(6.413)	(1.940)	(1.793)	(1.832)	(0.336)				(2.55)
25–34	1978	0.0237	−0.2070	−0.0122	−432	−408	105	62	−97	0.24	15.412	405	−3.11
		(0.691)	(5.948)	(2.792)	(6.058)	(6.744)	(1.035)	(0.698)	(0.904)				(0.43)
35–44	1988	−0.0573	−0.1696	0.0017	−403	−75	88	61	−10	0.08	4.880	442	−19.41
		(1.939)	(4.658)	(0.324)	(1.716)	(0.825)	(0.966)	(0.713)	(0.101)				(2.35)
45–54	2224	−0.0669	−0.2384	−0.0107		569	144	−156		0.14	9.890	368	−8.09
		(2.637)	(5.671)	(1.829)		(2.939)	(2.200)	(2.288)					(0.84)
55–64	1855	−0.0331	−0.1193	−0.0214			24	−121	−29	0.06	1.209	127	18.58
		(0.716)	(1.579)	(1.589)			(0.117)	(0.564)	(0.128)				(0.90)

It is worth noting that the purely economic variables have small coefficients in the regressions concerning husbands. The coefficient for the direct wage effect is insignificant in all cases except one. The cross wage effect is insignificant in all regressions. The nonemployment income coefficient is significant in two out of five cases. Of the dummies for children only c_2 is significant, and this is the case for the youngest and the oldest age categories. The substitution effect has the correct sign in four out of five cases. (Assuming independence, it is significantly greater than zero in one case).

The purely economic variables have larger (negative) coefficients in the regressions for wives. The coefficient for the direct wage effect is strongly significant in three out of five cases. The coefficient for the cross wage effect is significant in two out of five cases, and the same holds for the nonemployment coefficient. The dummies for presence of children work in the expected age groups - 20-24 and 25-34 - and in the expected direction. The substitution effect has the expected sign in only one case, and is significantly negative for two groups.

As can be expected the regressions show larger coefficients for women. The reason is probably that institutional factors contribute to dampen the variations in the dependent variable for husbands.

Finally, it is worth noting that the supply curves are most often backward bending and leisure seems to be a normal good.

We will now introduce two regression equations from a somewhat different approach. If one assumes that each member of the household possesses a utility function, and that he or she regards (part of) the income of the other members of the household as his nonemployment income, one gets a theoretical model that can be empirically specified as follows.

$$\ell_i = \alpha_0 + \alpha_i p_i + \alpha_3 \Pi_i + \text{dummies} + \varepsilon_i \qquad i=1,2 \qquad (25)$$

where

Π_i = the nonemployment income of the household plus the income(s) of the other member(s) of the household

p_i = the wage for the household member under consideration

The data include in this case individuals that are members of households, where at least one member participates in the labor market. The results for males and females are presented in equations (26) below.

$$\begin{aligned}
\ell_1 = 2136 &- \underset{(7.286)}{0.0247}\, p_1 - \underset{(3.682)}{0.0008}\, \Pi_1 + \underset{(1.279)}{18}\, c_1 + \underset{(1.030)}{14}\, c_2 + \underset{(3.845)}{58}\, h_1 + \\
&\underset{(2.756)}{41}\, h_2 + \underset{(2.025)}{32}\, h_3 - \underset{(6.694)}{165}\, m_1 - \underset{(2.366)}{35}\, m_2 + \underset{(1.983)}{27}\, m_3 + \underset{(1.503)}{20} m_4
\end{aligned}$$

$R^2 = 0.05 \qquad F = 12.920 \qquad n = 2823$

$$(26)$$

$$\begin{aligned}
\ell_2 = 1778 &- \underset{(9.121)}{0.1682}\, p_2 - \underset{(3.636)}{0.0020}\, \Pi_2 - \underset{(8.189)}{443}\, c_1 - \underset{(7.197)}{310}\, c_2 + \underset{(3.000)}{144}\, h_1 + \\
&\underset{(0.796)}{38}\, h_2 - \underset{(0.160)}{8}\, h_3 + \underset{(8.185)}{495}\, k_1 + \underset{(6.670)}{305}\, k_2 + \underset{(3.673)}{157}\, k_3 + \underset{(4.261)}{181}\, k_4
\end{aligned}$$

$R^2 = 0.14 \qquad F = 28.877 \qquad n = 1745$

The coefficients of the purely economic variables are strongly significant in both equations. The supply curves are negatively sloped and the "nonemployment" coefficients have the correct sign. The age dummies in the equations for males render support for the life cycle theories, and the corresponding age dummies for females show that all age groups work significantly more than the oldest age category. Moreover, the presence of children affect the women's supply of labor in the expected direction.

6. TAXES AND LABOR SUPPLY

A comparison of our results with the corresponding results from econometric studies on US material show a good correspondence with respect to the effect of the wage variable on the labor supply of males, i.e., their labor supply is relatively insensitive to changes in the wage rate. The picture is quite different for females. Our estimates show a significantly negatively sloped supply curve, while most of the U.S. studies indicate a positively sloped supply curve.

As was pointed out in Section 3 above a negatively sloped supply curve indicates that an increased marginal tax rate will increase the supply of labor. This implication intuitively seems to be a bit unreasonable due to the high marginal tax rates in Sweden compared to the U.S., and our knowledge of the behavior of relatives and neighbors. Intuition cannot, however, be used to refute scientific hypotheses, but it should force the researcher to think twice.

There is one obvious reason why the slopes of the supply curves could have a negative bias. Our method of calculating the hourly wage rate means that every misspecification of the true number of hours worked contributes to this bias. If an individual reports a too low working time this implies a too high hourly wage rate and vice versa.[1]

This possible bias could explain the negative substitution effects obtained in Section 5 above. A negative bias in the wage coefficient means that the substitution effect is underestimated and if this underestimation is sufficiently large, it will generate a negative substitution effect.

One way to solve this problem would be to use a (up till now unknown) data source, where the wage variable is collected directly. There is also the possibility of using an instrumental wage variable. A more direct way to analyse the effect on labor supply of changes in tax progression would be to introduce a variable measuring the tax progression. We have previously given a possible empirical specification of the labor supply function, where we use the fact that the municipal tax rate differs between the different households in our sample. We will now introduce estimation results of a variation on equations (24). The differences are two: the municipal tax rate has been introduced and the region dummies have been omitted to avoid possible multicollinearity.

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[1] The suspected bias due to negative correlation between the supply and wage variables may, however, to some extent be offset if part-time jobs are less payed per hour.

$$\ell_1 = \underset{}{2360} - \underset{(3.037)}{0.0164\, p_1} - \underset{(0.639)}{0.0046\, p_2} - \underset{(4.763)}{0.0046\, \pi^T} + \underset{(1.140)}{23\, c_1} -$$

$$\underset{(4.027)}{121\, m_1} - \underset{(3.396)}{58\, m_2} - \underset{(0.123)}{2\, m_3} - \underset{(0.281)}{4\, m_4} - \underset{(1.948)}{8.20\, t}$$

$$R^2 = 0.04 \qquad F = 6.294 \qquad n = 1414$$

$$(27)$$

$$\ell_2 = \underset{}{2773} - \underset{(2.118)}{0.0324\, p_1} - \underset{(9.113)}{0.1859\, p_2} - \underset{(2.536)}{0.0071\, \pi^T} + \underset{(7.612)}{459\, c_1} -$$

$$\underset{(5.803)}{268\, c_2} + \underset{(4.986)}{366\, k_1} + \underset{(2.747)}{161\, k_2} + \underset{(0.296)}{16\, k_3} + \underset{(1.073)}{58\, k_4} - \underset{(2.560)}{30.95\, t}$$

$$R^2 = 0.13 \qquad F = 21.435 \qquad n = 1407$$

where

t = the municipal tax rate.

We observe no major differences from the results in the previous estimations. The new variable (the tax rate) has negative sign for both men and women, however slightly insignificant for men. The negative tax rate coefficient means that a tax increase would decrease the supply of labor. This finding is inconsistent with our observations of a backward bending supply curve (if we can rely on our theory). Our presumption concerning bias in the wage variable is thus strengthened. In the case that we have a backward bending supply curve (seemingly inconsistent with the sign of the tax rate coefficient) but a positive substitution effect (consistent with theory) there remains, however, an additional possibility that the wage rate coefficients are unbiased: if the household perceives that the tax payments are redistributed to the household we can forget about income effects. An increased tax progression will only have a (negative) substitution effect.[1]

The tax rate elasticities are for men -0.10 and for women -0.56. Thus, an increase of the tax rate would not, it seems, as other authors[2] believe, diminish tax revenues.

We will finally in equations (28) present results from estimations of a variation on equation (25) where we have introduced the municipal tax rate variable in the same way as above.

$$
\begin{aligned}
\ell_1 = {} & 2364 - \underset{(7.120)}{0.0240\, p_1} - \underset{(3.361)}{0.0007\, \Pi_1} + \underset{(1.147)}{16\, c_1} + \underset{(1.026)}{14\, c_2} - \\
& \underset{(6.713)}{166\, m_1} - \underset{(2.298)}{34\, m_2} - \underset{(1.939)}{26\, m_3} - \underset{(1.609)}{22\, m_4} - \underset{(2.070)}{7.27\, t} \\
& R^2 = 0.04 \quad F = 14.526 \quad n = 2823 \\
\ell_2 = {} & 2644 - \underset{(9.048)}{0.1667 p_2} - \underset{(3.111)}{0.0017\, \Pi_2} - \underset{(8.202)}{445 c_1} - \underset{(7.270)}{313 c_2} + \\
& \underset{(7.815)}{477\, k_1} + \underset{(6.400)}{298\, k_2} + \underset{(3.173)}{137\, k_3} + \underset{(3.990)}{173\, k_4} - \underset{(2.812)}{30.97\, t} \\
& R^2 = 0.14 \quad F = 30.290 \quad n = 1745
\end{aligned}
\tag{28}
$$

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[1] This possible explanation has been suggested to us by Ingemar Hansson and Charles Stuart, University of Lund, Sweden.

[2] See, e.g., Jakobsson and Normann, this volume, and Stuart (1979). Our conclusion of course needs further qualification. We investigate a rather special sample with a partial equilibrium approach.

The tax rate coefficients have the same sign and magnitude as in equations (27). All other coefficients correspond very well with those obtained in equations (26). We should of course note that we face the same possibility of bias in equations (28) as discussed above.

6. CONCLUDING REMARKS

The models estimated in the previous sections are just a first step in several approaches that have to be investigated.

The bias problem demands an inquiry into the possibilities of using alternative data sources, where the gross wage variable is obtained independently of the number of ours worked.

A well-known problem with the gross wage rate variable lies in the possibility of the wage rate not being independent of the hours worked (e.g., over time premiums). The wage rate variable being endogenous implies correlation between the dependent variable and the error term.

A similar problem arises in the estimation of models with a marginal net wage rate variable. This problem, however, does not disqualify a net wage approach. Estimation techniques have been developed to cope with such simultaneity problems.[1]

The data source we use also allows an intertemporal approach. The income distribution survey has a panel, where half the sample is substituted every

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[1] See, e.g., Wales and Woodland (1979).

year. This enables us to study, for example, the determinants of entrance into and exits from the labor market.

The Swedish tax system with a proportional and locally differentiated municipal tax rate has enabled us to study the supply effects of changes in income tax progression. Another approach could be to pool samples from different years and more fully parameterize the tax system for the years studied. This could give a more complete picture of how changes in the tax system affect labor supply.

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On the Determinants of Labor Supply in Sweden

Comment

Frank P. Stafford*

As Axelsson, Jacobsson and Löfgren (AJL) note, their paper has the distinction of being the first econometric labor supply study in Sweden. This is remarkable since Sweden is the developed economy with the highest adult labor force participation rate and is at the same time the most interesting from the point of view of the potential impact of the tax structure on labor supply and work effort. Tax rates are high on average and at the margin and, for couples filing according to the married column in the tax tables since 1971, the wife and husband both face independent tax schedules; the earnings of one spouse do not affect the marginal tax rate of the other.

If one is inclined to believe that people do respond to taxes in choosing family labor supply, then, in comparison to the U.S., married women in Sweden should have greater incentives to work some, but not long, market hours. This is because in the U.S. married persons are taxed at their spouse's marginal tax rate on the first dollar of labor income[1] in contrast to a zero marginal tax rate in Sweden. If you like, while the U.S. system is "unfair" to married couples of equal earning capacity, the Swedish system is

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[1] If they file jointly. If they file separately they face a different tax schedule and this "pays" only if both spouses have high incomes. The Swedish tax system had similar provisions over the period 1966-71.

* I would like to thank Ned Gramlich for helpful comments on these comments.

"unfair" to married couples of unequal earning capacity and taxes the well-known division of labor within the household. My guess is that the U.S. system may evolve in the direction of the Swedish system, but for now major structural differences exist and invite comparative research.

As a consequence of the two tax systems, the comparative static model of labor supply predicts that in Sweden the high (average and marginal) tax rates on the spouse will make for smaller income effects from husband's earnings which in turn say to married women "work in the market". If the wife's earnings are low she will face low marginal tax rates, hence there are wage (substitution) effects for the first hours of work[1] but not for long hours of work because the progressive tax rates will create a reduced role for substitution effects so as to discourage long hours.[2] Tax rates are sufficiently progressive in Sweden as to likely offset the well-known wage premium which most labor markets generate for longer hours.

Assuming labor-leisure preferences to be the same in both countries, the prediction from a simple labor supply model is that married Swedish women should have higher labor force participation rates (LFPR's) and should be more likely to work short hours. The gross evidence is consistent with these predictions. In Sweden LFPR's of married women are higher than in the U.S. (63.2 versus 49.5 percent) but reported hours of work per week are less (28.9 versus 34.2 hours). Of course the gross evidence is very limited in value and simple time trends indicate an only moderately higher rate of growth in LFPR's of married Swedish women

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[1] Apart from income conditioned transfers and an apparent discontinuity with tax rates rising from .02 to .30 at 7,000 Skr per year.

[2] Marginal tax rates equal 50 percent at annual income levels of 53,000 Skr.

after 1971 than between 1966 and 1971. For married women age 35-44, the LFPR rose by 11.4 percentage points or (11.4/43.7)/5 = 5.2 percent per year relative to the nonparticipants in 1966-1971. From 1971 to 1975 it rose by 7.7 percentage points or (7.7/32.3)/4 = 6.0 percent per year relative to nonparticipants. This modest acceleration in LFPR's of married women is despite the fact that the major change allowing each spouse to face independent marginal tax rate schedules occurred in 1971. Because of the limitations of such a global overview, studies of the type presented by AJL are essential for public analysis. Their paper represents early results of a larger project, and my comments should be regarded as much suggestions for future research as a comment on their results to date.

Besides the usual static labor supply model developed in Section 1 of their paper, their future work could include models designed to illuminate intertemporal aspects of labor supply. With high marginal tax rates there are incentives for income averaging by keeping earnings constant from one year to the next. However, for the lowest wage groups income conditioned transfer payments can result in incentives for cycling between market work and non-participation with receipt of maximum transfer benefits in the non-work years. As a point in theory suppose some groups of individuals receive transient wage opportunities where wage rates may be high one year and low the next. In an intertemporal context this will encourage a larger apparent cross-sectional wage elasticity because the income effects of the temporarily higher wage will be spread out over a longer planning period.[1]

If the job market for persons in poverty or near poverty is characterized by highly variable wage rates for the same

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[1] See Metcalf, Charles, "Making Inferences from Controlled Income Maintenance Experiments", American Economic Review, June 1973, pp.478-85.

person through time, lower wage workers can appear to have a more elastic supply of labor when in fact their preferences may not differ from the non-poor. This theoretical possibility could explain the greater apparent labor supply elasticities of women and minorities which are often observed in the United States, and it is also consistent with higher unemployment rates because in some periods the optimal reduction in labor supply may be a corner solution of zero hours.

Another advantage of intertemporal models is that they can be adapted to consider endogenous retirement and the impact of tax and pension policies on lifetime labor supply. The sharp decline in LFPR's of older men in Sweden and the U.S. suggests this is worth considering. Further, models of this sort will generally indicate that the net wage from current labor market activity should be adjusted to include the discounted value of added future pension benefits.[1] As a result people may work long hours despite a low current marginal after tax income benefit. Finally, intertemporal models can be used to consider the effects of the tax system on human capital accumulation. Some of the theoretical work in this area indicates that people can have alternative life cycle paths of equal utility which differ dramatically in terms of skill accumulation.[2] Sharply progressive income taxes, which reduce the returns to extensive formal and on-the-job training will certainly encourage a choice toward the lower skill life cycle path.

On the demand side it is pretty well-established (or at least agreed upon) by labor economists that there are wage

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[1] Burkhauser, Richard and Turner, John, "A Time Series Analysis on Social Security and Its Effect on the Market Wage of Men at Younger Ages", Journal of Political Economy, August 1978.

[2] Ryder, Harl, Stafford, Frank and Stephan, Paula, "Labor, Leisure and Training Over the Life Cycle", International Economic Review, October 1976, pp. 651-74.

premiums for longer hours of work. A person choosing between a part-time job and a full-time job will usually get a higher hourly wage rate on the latter. The reasons for this are not absolutely clear, but can be partly explained by the fact that longer hours or at least 8 hours per day makes it easier to utilize more fully certain kinds of physical capital through the use of work shifts.[1] From the point of theory, marginal wage rates which rise with hours of work make the model less tractable because it imparts a (pretax) convexity of the opportunity set.

However, there is a Say's Law of Search. As my colleague Paul Courant has pointed out, supply creates its own demand in that more widely available kinds of units, be they housing or labor hours, are easier for searchers to find and this will lower expected costs. If only a few persons want part-time hours Say's Law of Search says that they may have lower wage rates which is in contrast to a usual factor demand argument which would imply higher wage rates to a scarce factor. In the U.S. many firms now plan for and actively recruit part-time workers and this must be the case in Sweden as well, but there is still probably a wage premium for full-time hours.

Combining the rising marginal wage on the demand side and the fact that progressive marginal tax rates will make for falling after tax wage rates is something of an analytical mess, but research indicates that, at least for the U.S., it is probably important in evaluating the effects of taxes on labor supply.[2]

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[1] Deardorff, Alan and Stafford, Frank, "Compensation of Cooperating Factors", Econometrica, July 1976, pp. 671-84.

[2] Rosen, Harvey, "Taxes in a Labor Supply Model with Joint Wage-Hours Determination", Econometrica, May 1976, pp. 485-507.

The current AJL work does not cope with a number of econometric problems, and it may be worthwhile to consider the following in future work:

1. Because they obtain wage rate measures by dividing labor income (before tax) by respondent reports of hours of work per week times weeks paid for per year, if there is a response error understating hours it will yield an overstated wage rate and conversely for overstated hours. These probable measurement errors will yield a spurious increase in the probability of observing a backward bending supply curve of labor and a false conclusion that work hours are increased by a tax rate. One method for coping with this is to reduce the negative error covariance between wage and hours by obtaining independent measures of hours or wage rates. One could use time diary estimates as the dependent variable or respondent reports of wage rates as the independent variable. There are still errors-in-variables problems (the wage elasticity will be biased toward zero), but using independent hours and wage estimates would likely be an improvement over the current methodology.

2. If panel data are available one could cope with the spurious correlation between non-labor income and labor supply: given observationally identical individuals in a cross-section the one with a history of more market work will have built up greater non-labor income. If so the coefficients in equation (23) will understate the income elasticity of demand for leisure.

3. Panel data or some sort of averaging would also permit the authors to deal with the econometric problems associated with non-participa tion. In their paper the analysis is restricted to participants, but it would seem

that one of the major aspects of the Swedish tax system is to encourage participation. Panel data would permit a substantial reduction in the number of zero hours cases since hours would be averaged over a number of years.

4. If an averaging procedure is not used one could develop a reservation wage model, but the models developed to date represent the case of a constant hourly wage rate[1], a condition unlikely to be met if there are very progressive taxes or a market wage premium for long hours.

The future and current work faces difficult measurement issues. To introduce tax rates requires not only examining the tax schedule but also factoring in the effect of income-conditioned benefits such as child care and the deductability of interest payments. The latter may be non-trivial and leads me to an Archipelago theory of labor supply. If your labor earnings place you in the 82 percent marginal tax bracket (114,000 Skr per year or more) and if you were considering cutting back your work effort because of this, the lure of sailing in the Archipelago could encourage you to keep working longer hours, at least in the winter. If the boat you buy is financed either directly or indirectly, by increasing your home mortgage, the boat could be bought with after tax crowns approaching 20 öre.

Another measurement issue centers on hours of work. Methodological work on the validity of respondent reports of work hours is, to my knowledge, non-existent, despite the use of

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[1] Heckman, James, "Shadow Prices, Market Wages and Labor Supply", Econometrica, July 1974, pp.679-94 or Devaney, Barbara "The Labor Supply of Married Women: An Analysis of the Allocation of Time to Market and Non-Market Activities", unpublished Ph D dissertation, Department of Economics, University of Michigan, 1977.

Table 2. Comparisons of alternative hours at work measures

Reported hours per week in an average week	Time diary reports of mins., worked per Week				
	Under 1770	1770-2369	2370-2969	2970 or more	No diary time
Under 30	13	10	2	1	2
30-39	12	12	10	1	2
40-49	45	35	105	36	3
50 or more	6	10	23	27	0

such data in most industrialized economies. The data in Table 2 demonstrate that time diary estimates of hours at work during a twelve month period are only weakly correlated with respondent reports of hours of work per week in an average week, and respondent reports appear to miss secular trends toward fewer hours. Further, elapsed hours at work do not measure work effort nor time spent at work not working, all of which may, over the long run, be influenced by tax policy.

So much for preliminary remarks.

What do we know right now about labor supply in Sweden in light of the AJL work? If we take their results literally the income and substitution effects tells us that men in the U.S. and Sweden are fairly similar in their *preferences* in that neither respond very much to economic variables. However, women in Sweden and the United States are *different* in their labor supply *preferences*. I emphasize preferences because according to the methodology used by AJL, with wage rates independent of hours of work and "exogenous" non-labor income, we are pretending to "identify" the opportunity set by *a priori* zero and non-zero restrictions and are using this to estimate parameters describing the peference side. Women in Sweden have, according to AJL, negative wage elasticities (the resultant effect of underlying substitution and income elasticities) while in the U.S. they have moderate, positive wage elasticities.[1]

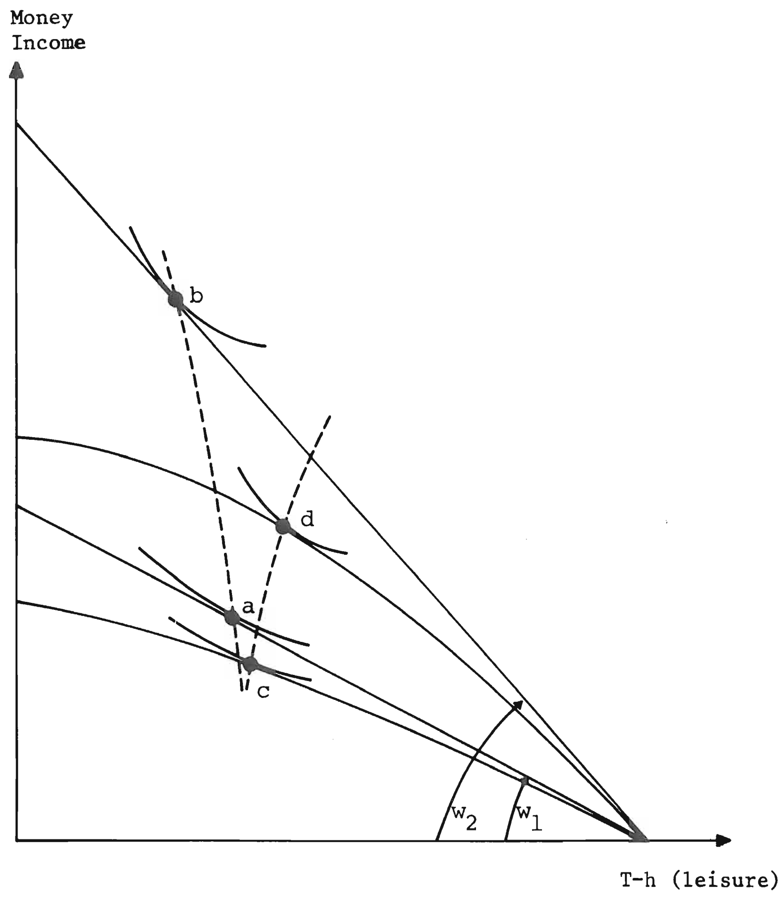
An alternative to the interpretation that women's preferences differ in the two countries is that the negative wage elasticities reported in Table 1 of their paper may reflect the fact that at higher before tax wage rates and given participation, the disincentives to longer hours are relatively greater in Sweden than in the U.S. In light of the comments on theoretical, econometric and measurement problems, this is assuredly a tentative inference. For example, annual work hours are defined to include weeks of compensation rather than weeks at work, and if vacations are longer for higher wage women this would yield an overestimate of work hours and should impart an upward algebraic bias to the wage elasticity estimated by AJL.

Not including progressive tax rates in the model can be viewed as an erroneous prior restriction to identify preference parameters. To visualize this, consider a situation where all persons participate and face pretax wages of either w_1 or w_2 and after tax marginal wage rates of Θ_1, Θ_2. In the case of progressive taxes there is a relatively larger income effect and this induces shorter work hours - in fact it is the converse of overtime wage rates (say double time for hours beyond the "standard" work day) which act to increase the role for substitution effects relative to income effects. To AJL what appears to be a wage increase from w_1 to w_2 is really a marginal equilibrium wage rate increase (wage rates are endogenous) from Θ_1 to Θ_2 and the

[1] Rees, Albert, "Summary of Results on Negative Income Tax", Journal of Human Resources, Spring 1974.

latter "induces" a reduction in labor supply. See Figure 1 where cd represents the price and consumption curve observed with progressive taxes and when ab is the price consumption curve without taxes or with proportional taxes. As a first step forward in their current research the authors could utilize something pretty close to the current model, but use marginal wage rates.

Figure 1. Labor supply and progressive taxation



In conclusion it should be noted that the tax systems in Sweden and the U.S. can also have significant impacts on marriage, fertility and divorce. In a broader analysis family composition is an endogenous variable and studies of labor supply by different demographic groups can easily overlook the fact that how many people are in which groups can, in the long run, be related to labor supply and therefore be subject to influences of tax policy. In the U.S. divorce rates have doubled within a decade[1] and this is not unrelated to greater lifetime career commitments to the labor market by younger women. More generally, tax rates can affect a whole array of non-market behavior. To take a specific case, in the U.S. a variety of market and non-market work activities have declined substantially between 1965 and 1975. However, time spent in household repairs and projects appears to have exhibited a substantial percentage increase.[2].Time spent in such repairs is tax exempt, but the rapid increase in housing prices which motivated greater activity of this sort may also be the result of efforts to respond to personal income taxes.

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[1] Michael, Robert T., "Why Has The U.S. Divorce Rate Doubled Within The Decade?" working paper, National Bureau of Economic Research, Stanford California, September 1977.

[2] See Stafford, Frank P. and Duncan, Greg J., "The Use of Time and Technology by Households in the United States", forthcoming in R.G. Ehrenberg (ed.), Research in Labor Economics.

Welfare Effects of Changes in Income Tax Progression in Sweden

Ulf Jakobsson and Göran Normann

1. INTRODUCTION

In this paper we will investigate the implications of the theory of optimal income taxation for the graduation of the income tax schedule in Sweden.

The optimal income tax problem deals with the trade-off between equality and efficiency, when deciding on the progressiveness of the income tax schedule. The trade-off problems considered in the literature are of two kinds:

(i) between equity and efficiency losses due to distortions of labor-leisure choice. (See, e.g., Diamond, 1968; Mirrlees, 1971, and Phelps, 1973.)

(ii) between equity and distortions of the incentives to invest in human capital. (See, e.g., Atkinson, 1973, Phelps, 1973, and Sheshinski, 1976).

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So far there are few works where these trade-offs have been studied in connection with an actual tax system.[1]

We will investigate the first mentioned trade-off problem in connection with the Swedish system of personal income taxation. Even though we cast the problem into an optimal taxation mould, we do not intend to find the optimal tax schedule. Instead we will look for welfare improving tax reforms.

The instrument used in this analysis is an extended version of the model for simulating the Swedish system of personal income taxation first presented in Jakobsson and Normann (1972). The original simulation model belongs to a class of models with explicit public parameters that by now is quite common.[2] This article might be seen as an attempt to indicate how these models can be extended to include behavioral relations, which opens up the possibility of using them for a broader range of problems than today.

The next section of the article is a description of the model used. We start by presenting the original simulation model by which tax revenues at the individual and aggregate levels can be computed. The original model provides us with one of the essential features of the optimal tax problem, namely a tax function defined on individual

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[1] The only examples we know of are Bruno and Habib (1976) and Rosen (1976). None of these works however did primarily investigate the rate structure of the tax system.

[2] For early examples or models of this type see Pechman (1970) and Rechtenwald (1972)

income. This model is then extended to encompass the other main ingredients of the optimal income tax problem as posed by Mirrlees (1971). These are individual utility functions defined on consumption and leisure, a skill distribution, a social welfare function defined on individual utilities, and a production relation. We give a fairly detailed description of how this extension is made in the last part of Section 2.

To find the optimal tax system, the social welfare function is maximized subject to two constraints. The first is that the individual maximizes his utility subject to his income constraint. The second is that the total labor supplied can produce the total quantity of goods demanded. Welfare improving tax reforms will analogously be tax changes that improve social welfare subject to these two constraints. Sections 3 and 4 of the article are devoted to finding that kind of tax changes, where the present Swedish tax system is the initial state. This is done by simulation in the extended tax model.

We find that under the assumptions usually made in the literature on optimal income taxation progression in the Swedish income tax should be decreased. The most striking result is that all statutory marginal tax rates should be diminished in brackets above Skr 30,000 (ca $7,500) which was well below the median income in 1975, the year covered in the study.[1] The main explanation for this turns out to be a "perverse revenue" effect. Revenues will actually be increased when marginal tax rates are diminished. The extra revenues could

[1] Skr 30,000 in 1975 correspond to Skr 43,000 or $10,000 in 1979 prices.

be used for introducing a lump sum transfer. This combination of parameter changes will obviously increase the utility for everybody. Therefore the specification of the social welfare function is not important for the result mentioned, as long as we restrict ourselves to Paretian functions.

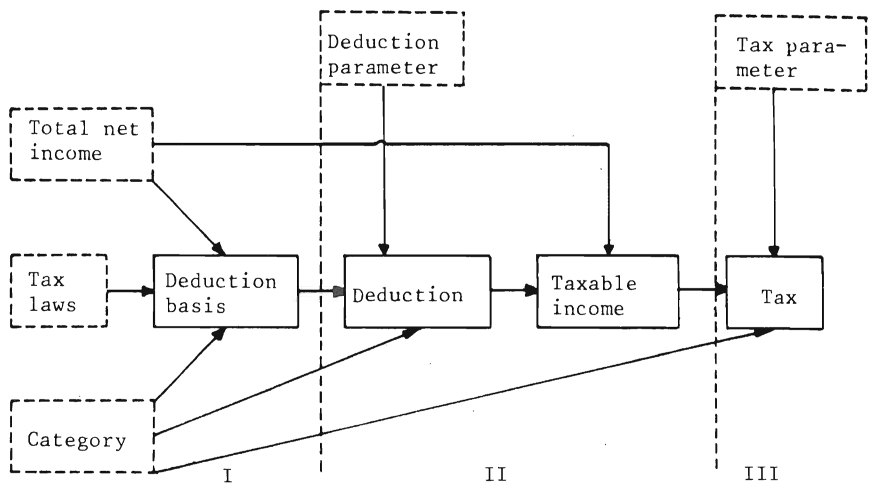
What is important, however, is the labor supply response to a change in marginal tax rates, since this response obviously is crucial for the "perverse revenue effect." In Section 5 we investigate how sensitive this effect is to different assumptions on the elasticity of substitution (σ) between consumption and leisure in the individual utility function. It is found that this effect appears in most rate brackets for $\sigma > 0.4$.

In the last section we briefly discuss what kind of conclusions can be drawn from our results.

2. MODEL DESCRIPTION

The original model consists of two parts namely a micro part and an aggregative part. The former part is constructed to compute the tax for a random individual. The individuals were partitioned in ten categories such that all individuals in a category are treated at least approximately equal by the tax laws. The categories are of the type single persons (age 17-66) without children, married men (age 17-66) and so on. An individual is characterized in the model not only by the category he belongs to but also by the level of his income before tax. Thus the micro model is an algorithm that for a given set of public parameters computes the tax for an individual on the

Figure 1. Chart of the micro-model



basis of two pieces of information of him, namely: (1) the individual's level of income before tax; (2) the category the individual belongs to.

As can be seen from Figure 1 the micro model is the place where the public parameters are introduced. Jakobsson and Normann (1972) give a short description[1] of how the tax laws were formalized and to some extent simplified so that they could be integrated in the model.

If we consider a specific category a condensed description of the micro model is given by:

$$t = F(y; P) \tag{1}$$

---

[1] For a full description, see Jakobsson and Normann (1974).

where t = individual tax payments
y = individual income before tax[1]
P = set of deduction and tax parameters.

To get from (1) to a macro-relation between income and taxes an aggregation procedure is introduced. The one we have used relies on knowledge of the income distributions in different categories. Still considering a specific category the total tax (T), paid by the category is given by

$$T = N \cdot \int_{y_{min}}^{y_{max}} F(y;P) \cdot \psi'(y)dy \qquad (2)$$

where N = number of persons in the category,
$\psi'(y)$ = density function of incomes in the category.

In this simulation model it is possible to distinguish and compare the effects on, e.g., revenues and income distribution after tax of different specified changes in the parameter set. The level and distribution of income before taxes also appear explicitly so the built-in-flexibility of the tax system can be investigated. An important limitation of the model, however, is that income before tax is exogenous. By introducing, in the micro-model, utility maximizing choice between labor and leisure on part of indviduals, this assumption is relaxed in the present version of the model.

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[1] The income before tax concept used here is total net income (sammanräknad nettoinkomst). Our choice of this concept that is defined by the tax law has been dictated by the existing data on income distribution.

2.1. Individual Behavior

The assumptions on individual behavior made here are those of standard labor leisure analysis. We will thus assume that individuals have identical preferences and try to maximize their utility. It is also assumed that consumptions of goods (C) and consumption of leisure (L) enter a utility function U(C;L). Each individual makes his (C;L) choice under his budget constraint;

$$C = f(wH;P) \equiv wH - F(wH;P) \tag{3}$$

where H = hours worked (H=Q-L; Q=hours available)
w = wage rate
f represents the function from income before tax to income after tax.

The formulation of the budget constraint implies two assumptions, both common in the optimal tax literature:

(i) Savings are ignored.

(ii) Other income than wage income is ignored, i.e., y = wH.

In order to make a quantitative analysis it is necessary to be more specific on the form of the individual utility function. We have here chosen the standard assumption that the utility function is of the Cobb-Douglas type. In a special section we will discuss how sensitive our basic results are to this assumption.

On the assumption that the individual tries to maximize his utility, he will face the following optimum problem:

$$\text{Max } U = C^{\alpha}(Q-H)^{1-\alpha} \quad \text{subject to } C=f(wH;P) \tag{4}$$

The optimal labor supply of the individual will be

$$H = \frac{Q \cdot e \cdot \alpha}{1 - \alpha + e \cdot \alpha} \qquad (5)^{1}$$

where

$$e = \frac{f_1(wH;P) \cdot wH}{f(wH;P)} = \text{residual progression.}^{2}$$

If we suppose that the wage rate (w) for each individual is given exogenously then (5) in principle can be solved for H, provided that f is completely specified. Furthermore it is clear that to each specific set of public parameters (P) we get a related solution for H. So (5) defines a function from (w;P) to H or

$$H = g^1(w;P) \tag{6}$$

By (6), the budget-restriction (3), and the utility function we get

$$U = g^2(w;P) \tag{7}$$

---

[1] Q stands for maximal labor supply. Supposing that there is a limit at 16 hours per day every day, we get for a full year Q=5,840. To get realistic values on labor supply we have chosen α=0.33. Experimentation with different values on α indicates that our results are not sensitive to changes in α.

[2] For a discussion of this concept see Jakobsson (1976).

Since we are assuming that $y = wH$, we also get by (6) and (1) individual tax payments

$$t = g^3(w;P) \tag{8}$$

2.2 Aggregation over Wage Rates

A basic difference between the micro-model defined by (1) and that defined by the preceding equations is that the wage rate is exogenous in the latter while income is exogenous in the original model. From the empirical point of view this represents a difficulty since the only information we have got on individuals is the distribution of income. In order to aggregate the model (6)-(8) it is therefore necessary to relate individual income in the initial position to wage rates. This is done by (5). At the existing tax system we can observe the income distribution before tax. Formula (5) then relates each income to a specific value of H. Since $y = wH$, we also get a specific wage rate associated with each income level in the initial stage. From the observed income distribution we can then derive a distribution of wage rates that is exogenously given in the model and constant throughout the experiments carried out here. For a specific category of income earners aggregate tax payments can be obtained as

$$T = N \cdot \int_{w_{min}}^{w_{max}} g^3(w;P)\ \phi'(w)dw \tag{9}$$

where $\phi'(w)$ is the "derived" distribution of wages. We will assume that this distribution is equivalent to the skill distribution in the optimal income tax problem. Concerning production we adopt the assumption that the production of each worker equals his wage.

2.3 The Social Welfare Function

A central element for the whole concept of an optimal tax schedule is an interpersonal comparison of utilities. The valuation of utilities for different persons is made by a social welfare function. The proper specification of this function is of course a very difficult problem. We have, however, chosen the form most commonly used in the literature on optimal taxation, namely addition of individual utilities raised to the power of $1-\varepsilon$, where ε could be interpreted as social inequality aversion (Atkinson, 1970) $(U^{1-\varepsilon}/(1-\varepsilon);\ \varepsilon > 0;\ \varepsilon \neq 1)$. By this function we have social welfare

$$W = \frac{1}{1-\varepsilon} N \cdot \int_{w_{min}}^{w_{max}} [g^2(w;P)]^{(1-\varepsilon)} \phi'(w)dw \qquad (10)$$

Restricted as this form might seem it still allows for a wide range of social preference orderings. Included are the strictly utilitarian approach ($\varepsilon=0$) and the Rawlsian welfare function, max-min, ($\varepsilon \to \infty$). This illustrates the well-known fact that the sensitivity of the function W to changes in different parts of the distribution is affected by the valaue of the parameter ε. The higher the value of ε the larger is the weight given to changes in the lower part of the distribution. A higher value does also increase the general sensitivity for inequality.

By (10) our extended simulation model is complete and it will now be used to investigate what effects we get when public parameters are changed. By simulations with the model we compute partial derivatives of H, U, t (individual level), W and T (aggregate level) with respect to specific public parameters P_j.

3. SIMULATION RESULTS

All simulations are restricted to the category married men in active ages with wives having no income. Important for our analysis is that in this category a very high fraction of total income is wage income. Table 1 gives for this category average pre-tax income in each income class (1975) and corresponding average and marginal effective tax rates in the 1975 tax system.

The policy instruments we are going to consider are the statutory marginal tax rates at national taxation, the local tax rates and the basic tax deduction. In addition to these existing parameters we consider the effects of the introduction of a lump-sum transfer equal to all persons in the distributions.

3.1. Effects on The Individual

On the individual level we can, according to (6)-(8), compute $\partial H/\partial P_i$; $\partial U/\partial P_i$; $\partial t/\partial P_i$, etc., for each specific wage rate. Before we report on the results of these computations we shall indicate

Table 1. Tax rates and income distribution for married men (wife not assessed) in 1975

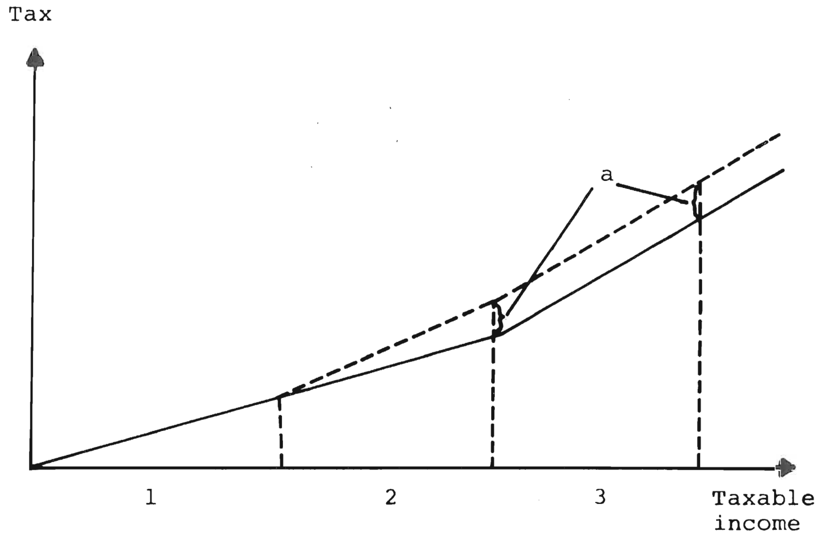
Income class	Relative frequency of tax-payers %	Pre-tax mean income Skr	Average tax rate %	Marginal tax rate %	Residual pro-gression[a]
1	2.2	118	0	0	1.00
2	0.1	2 801	0	0	1.00
3	2.0	9 076	0	31	0.69
4	4.1	14 411	9	31	0.76
5	5.7	20 259	16	36	0.76
6	8.2	25 598	20	41	0.74
7	13.6	31 416	24	46	0.71
8	17.6	36 634	28	52	0.66
9	14.6	42 373	31	52	0.69
10	15.1	49 323	35	57	0.65
11	6.6	61 274	40	62	0.63
12	5.7	74 882	44	72	0.50
13	2.1	98 865	51	72	0.57
14	2.3	161 158	61	80	0.51

[a] Elasticity of income after tax with respect to income before tax.

the nature of the different parameter changes and the kind of individual response we might expect under the assumptions made.

The effect on individual labor supply from a tax change can be divided in an income effect and a substitution effect. The income effect is positive, which in this context means that an isolated increase in the average tax rate will increase labor supply. The negative substitution effect implies that an isolated increase in the marginal tax rate will lead to a diminished labor supply. For a given tax schedule a specific revenue is collected from the individual.

Figure 2. Increase of the statutory marginal tax rate within a specific bracket



The tax schedule in the Swedish tax system can be described as an increasing step-wise linear function from income to tax payments. The general shape of the function is determined by the statutory marginal tax rates at national taxation and the so called basic tax deduction. Figure 2 illustrates an increase of the statutory marginal tax rate within a specific bracket (bracket 2 in the figure). Obviously, people below this bracket will not be affected by the change. Everybody in bracket 2 and above will have their utility levels diminished. An individual within the bracket gets his marginal tax rate as well as his average tax rate increased, so the effect on labor supply is in principle undetermined and so is the revenue effect. If the effect on labor supply is positive, the revenue effect will of course also be positive. A negative supply effect might, however, diminish the tax-base enough to offset the effect on revenue from the upward shift in the tax schedule.

As the tax increase in bracket 3 and above is of the same nature as an additional lump-sum tax labor supply in these brackets will be greater than before and so will revenues collected. Utility levels, however, will of course be diminished.

If we now go to the local tax it could mainly be seen as a linear tax with constant marginal tax rate which is equal to the local tax rate. It is clear that for the whole range of income an increase in this tax rate will give rise to exactly the same effects as we met within bracket 2 in the preceding paragraph.

The qualitative effects of changes in the other two instruments (basic tax deduction, lump-sum transfer) are obvious since they do not affect marginal tax rates and therefore only give rise to income effects.

Results on the micro level for changes in the statutory marginal tax rates in brackets Skr 0-10,000 and Skr 30,000-40,000, can be seen in Table 2. Each of these parameters has been increased by one percentage unit. In the table the resulting changes in percent of initial values are given for tax payments, hours worked and individual utilities at different income levels. To pick an example we can in row 8, column 9, read the value of $(\partial t/\partial P_j)/(t)\cdot 100$ at income level $\approx$36,600, where P_j stands for the marginal tax rate in the bracket Skr 30,000-40,000.

Table 2. Effects of parameter changes on the individual at different income levels

Income class	Pre-tax mean income before tax change (1)	Increase of statutory marginal tax rate in taxable income bracket: 0-10,000 Skr				30,000-40,000 Skr			
		Marginal tax rate (2)[a]	Work effort (3)[b]	Utility (4)[b]	Tax payment (5)[b]	Marginal tax rate (6)[a]	Work effort (7)[b]	Utility (8)[b]	Tax payment (9)[b]
1	118	0	0	0	0	0	0	0	0
2	2 801	0	0	0	0	0	0	0	0
3	9 076	+1	-1.1	-0.1	0.0	0	0	0	0
4	14 411	+1	-0.6	-0.2	4.8	0	0	0	0
5	20 259	0	0.5	-0.2	5.5	0	0	0	0
6	25 598	0	0.4	-0.2	2.3	0	0	0	0
7	31 416	0	0.4	-0.1	2.1	0	0	0	0
8	36 634	0	0.4	-0.1	1.6	+1	-2.1	-0.0	-3.7
9	42 373	0	0.3	-0.1	1.3	+1	-1.8	-0.1	-2.5
10	49 323	0	0.3	-0.1	1.1	0	0.3	-0.1	1.1
11	61 274	0	0.3	-0.1	0.8	0	0.3	-0.1	0.8
12	74 882	0	0.3	-0.1	0.8	0	0.3	-0.1	0.8
13	98 865	0	0.3	-0.1	0.4	0	0.2	-0.1	0.4
14	161 158	0	0.2	-0.1	0.3	0	0.2	-0.1	0.3

[a] Change given in percentage units.

[b] Change given in percent of initial value.

As could be expected, utilities are decreased for all individuals affected by the tax increase. Furthermore, those individuals that get their tax rates increased with unchanged marginal tax rates will increase their hours worked. The amount of tax collected from these people will, of course, also increase. These results do not depend on our specific choice of utility function for the individual. The Cobb-Douglas assumption is, however, important in the brackets where marginal tax rates are increased. Here we get a decrease in labor supply. For individuals with taxable income in the bracket Skr 30,000-40,000 this effect is strong enough to produce a negative overall effect on their tax payments.

This negative effect is essential for the results we will give later on. Some readers might find it so extreme that it would rule out any form of the individual utility function producing this effect. However, as soon as any incentive effects at all are admitted, a perverse revenue effect does not seem to be too far fetched which should be clear from the following example.

Consider a full time worker supplying 2,000 hours/year at a wage rate of 22.5 Skr/hour. This gives a yearly wage of Skr 45,000 and a taxable income of approximately Skr 40,000. Tax payments are roughly Skr 12,000. Now let the marginal tax rate in the brackets above the taxable income Skr 30,000 be increased by one percentage unit. At a taxable income of 40,000 this gives an initial tax increase on Skr 100 or 0.8 percent of taxes paid. By how much must hours worked be diminished in order to offset this positive revenue effect? Since the elasticity of tax payments with respect to income in this bracket is roughly equal to 2, an adjustment in hours worked by 0.4%, or 8 hours per year, would be sufficient to give a zero revenue effect. Higher adjustments than 8 hours per year will consequently give negative revenue effects.

3.2 Aggregate Effects

From the aggregative part of he model (e.g., (9)-(10)) we can investigate the effects of specific parameter changes on tax revenues and the social welfare function. Table 3 gives computed values of $\partial T/\partial P_i$ and $\partial W/\partial P_i$ for different parameters.

The most striking result of the table is that the perverse revenue effects we could observe at the micro-level in certain cases give rise to similar effects at the macro-level. Take, e.g., the bracket Skr 30-40,000. From the table we can see that a rise of the marginal tax-rate in this bracket by 1 percentage unit will decrease the aggregate tax revenues by Skr 19 million. From the micro-simulations (Table 2) it is clear that this figure is the net effect of diminished revenues from people within the bracket getting their marginal tax rate increased and revenue increases from people above the bracket, where the average tax rate is increased while the marginal tax rate is unchanged.

The interpretation of the perverse revenue effects for certain brackets is that the tax schedule in these brackets is not Pareto-optimal under the assumptions on individual behavior made here. Lowered marginal tax rates would increase utilities for the persons affected at the same time as total revenues would be increased.

Table 3. Aggregate effects of parameter changes on social welfare and tax revenue

	Parameters		Change of parameter	Effect on			
	National income tax schedule			Tax revenue mill. Skr	Social welfare[a]		
	Taxable income bracket. Thousands of Skr	Initial statutory marginal tax rate, %		$\partial T/\partial P_i$	$\partial W/\partial P_i$		
					ε=0.8	ε=3.0	ε=6.0
P1	0-15	7	+1 p.u[b]	99	-4.21 10^{-1}	-8.9 10^{-3}	3.3 10^{-5}
P2	15-20	12	"	27	-1.41 "	-2.7 "	-.8 "
P3	20-25	17	"	22	-1.27 "	-2.3 "	-.6 "
P4	25-30	22	"	10	-1.03 "	-1.7 "	-.4 "
P5	30-40	28	"	-19	-.89 "	-1.3 "	-.3 "
P6	40-45	33	"	-30	-.25 "	-.3 "	-.1 "
P7	45-65	38	"	-0	-.44 "	-.4 "	-.0 "
P8	65-100[c]	43	"	-14	-.46 "	-.39"	-.0 "
P9	100-[d]	52	"	-33	-.68 "	-.49"	-.0 "
P10	Lump sum transfer[e]		+100 Skr	-71	1.0	1.0	1.0
P11	Basic tax deduction[f]		+100 Skr	-35	1.73 10^{-1}	3.4 10^{-3}	1.2 10^{-5}
P12	Local income tax[g]		+1 p.u	33	-6.40 "	-12.9 "	-4.5 "

[a] These effects are normalized so that the effect of the introduction of a lump-sum transfer by 100 Skr is equal to one.

[b] Percentage unit.

[c] Two brackets put together. The statutory marginal tax rate is 48% in the subbracket 70,000-100,000 Skr.

[d] Cf. c). The statutory marginal tax rate is 56% in the subbracket 150,000-.

[e] This parameter does not exist in the actual tax system.

[f] Presently 4,500 Skr allowed to all income earners subject to the restriction that taxable income should not become negative.

[g] Flat rate of approximately 26% applied to taxable income.

We can also observe that the effect on social welfare of introducing a lump-sum transfer, with one exception is much greater than any other welfare effect. The exception is the rate of the regressive local tax. For $\varepsilon=0.8$ it would not increase social welfare to finance an increased lump-sum transfer with an increase in the local tax rate.

For higher values of ε the welfare effect of other parameter changes become almost negligible compared to the welfare effect of a change in the lump-sum transfer.

4. WELFARE IMPROVING POLICIES UNDER A FIXED BUDGET-CONSTRAINT

We are now equipped to answer the question of which parameter changes to choose in order to increase social welfare. As we do not consider other branches of public policy than personal income taxation it is natural to restrict the changes in the tax schedule to leave total net revenues constant. Under the assumptions made here this restriction is equivalent to the restriction that changes in consumption shall be equal to changes in production (see Stern, 1976). By the help of Table 3 it is easy to design policies, i.e., combinations of parameter changes that improve social welfare keeping total revenues constant.

In terms of our previous notation our task is to find combinations of parameter changes dP_k; dP_c such that

$$dW = \frac{\partial W}{\partial P_k} \cdot dP_k + \frac{\partial W}{\partial P_c} \cdot dP_c > 0$$

$$dT = \frac{\partial T}{\partial P_k} \cdot dP_k + \frac{\partial T}{\partial P_c} \cdot dP_c = 0 \qquad (11)$$

In Table 4 we give a selection of combined parameter changes that fulfills (11). The results are in accordance with those reached by Mirrlees (1971) and Phelps (1973). Both authors present results indicating that the optimal marginal tax rates should be falling at higher income levels. Here it is clear that marginal tax rates in brackets above 30,000 should be lowered. In Table 4, II and III are examples of such policies. It should also be mentioned that these two policies are of special interest since they as well as policy VI represent Pareto improvements.

We have introduced the possibility of a lump-sum transfer in the tax system. Our results strongly indicate that such an element should be included in the actual tax system. This is of course also in accordance with the results reached in theoretical literature.

In our analysis this result can be explained by the heavy weight attached to income in the lowest part of the distribution, already by the utilitarian sum of utilities. This tendency is reinforced by the social welfare function. It should also be pointed out that the financing of such policies is comparatively easy in the category married men since it has few persons in the lower end of the income spectrum (see Table 1).

Table 4. Combination of parameter changes improving social welfare under a fixed revenue constraint

	I		II		III	
Parameters involved	P1[a] marginal tax rate bracket 0'-15' Skr	P10[b] lump-sum transfer	P6[a] marginal tax rate bracket 40'-45' Skr	P10[b] lump-sum transfer	P6[a] marginal tax rate bracket 40'-45' Skr	P1[a] marginal tax rate bracket 0'-15' Skr
Parameter changes	+0.71	+1	-2.3	+1	-3.3	-1

	IV		V*		VI	
Parameters involved	P1[a] marginal tax rate bracket 0'-15' Skr	P3[a] marginal tax rate bracket 20'25,	P12[a] local tax rate	P10[b] lump-sum transfer	P12[a] local racket	P1[a] marginal tax rate bracket 0'-15' Skr
Parameter changes	-1	+4.5	+2.2	+1	-3	+1

[a] Change given in percentage units.

[b] Change given in hundreds of Skr.

* For ε=0.8 the indicated combination of changes in local tax rate (P12) and lump-sum transfer (P10) leads to a decreased value of the social welfare function.

Another general conclusion from the results is that the valuation of different policies do not change much with the value of ε. For the piecemeal policy analysis done here, it is in most cases indifferent if ε is equal to zero (the strictly utilitarian approach) or if we let ε tend to infinity (the Rawlsian criterion). A related point is that utility changes in the higher income classes mostly could be neglected. What is important here is the revenue effect. Therefore the assumptions made on disincentives in these classes are important for the results we will get.

From Table 3 it is seen that an increase in the local tax rate combines a low revenue effect with a high welfare loss. Policies V and VI in Table 4 are both encompassing a change in the local tax rate (P12). When it is used to finance an increased lump-sum transfer we get a welfare increase only when ε is greater than 0.8. This increase is much less than the one we get when the local tax rate is lowered in combination with an increase in the marginal tax rate in the lowest bracket (policy VI).

5. DISINCENTIVES AND THE REVENUE EFFECT

A clear-cut result of our previous analysis is that, under the assumptions made, marginal tax rates should be decreased in all brackets above Skr 30,000. This result depends crucially on the fact that in these brackets a decreased marginal tax rate leads to an increase in aggregate tax revenues (T).

It is important to check how sensitive this result is to changes in the elasticity of substitution between consumption of goods and consumption of leisure. We have done this by letting the individual's labor supply be governed by a utility function of the CES-type.[1] By simulating the response of hours worked and revenues for different values of σ for a change in the marginal tax rates in each one of the brackets above 30,000 Skr we get an indication of the range of σ where the disincentive effect is strong enough to create a perverse revenue effect.

From Table 5 it is seen that in the two highest brackets there is quite a wide range of values on σ that will give a perverse aggregate revenue effect. For the lower brackets, however, we get a picture that is a bit more mixed. Still, the Cobb-Douglas assumption does not seem to be essential for our results. An interesting result in this connection is provided by Stern (1976) who calculated implied elasticities of substitution from supply curves estimated by Ashenfelter and Heckman. This calculation gives $\sigma = .4$ which indicates that the range of σ in Table 5 for most brackets contains realistic values.

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[1] $U = \left[\alpha C^{-\mu}(1-\alpha)(T-H)^{-\mu}\right]^{1/\mu}$ $\left(\sigma = \frac{1}{1+\mu}\right)$. If U is maximized subject to the budget constraint the number of hours worked will be determined implicitly by the following equation

$$\frac{C}{T-H}^{-(\mu+1)} = \frac{1-\alpha}{\alpha w f_1'(wH;P)}$$

In lack of data on hourly wage rates we have computed values on w from yearly incomes on the assumption that everybody <u>initially</u> is working 2,000 hours/year.

Table 5. Least value on σ in the CES-function where an increased tax rate produces diminished aggregate tax revenues

	P 5	P 6	P 7	P 8	P 9
Bracket of the tax schedule (thousands of Skr)	30-40	40-45	45-65	65-100	100-
Revenues will be diminished for $\sigma \geq$	0.8	0.4	1.0	0.4	0.3

6. CONCLUDING REMARKS

A clear-cut conclusion of our analysis is that the graduation of the Swedish income tax schedule differs greatly from what would be prescribed by the theory of optimum income taxation with its usual assumptions. One may then take either the position that the tax system should be changed or the position that the assumptions in the theory of optimal income taxation need re-examination.

Certainly one would like to have more empirical evidence on individual behavior before using our results for policy prescriptions. The analysis made has highlighted the crucial importance of the labor supply response to tax changes. Therefore one objection against the results reached might be that the assumptions on disincentives have little empirical support. Econometric work in this area indicates that labor force participation and average hours of adult men are affected relatively little by changes in tax rates. As we could see in Section 5, calculations made by Stern (1976) indicate that the elasticity of substitution between labor and leisure among adult men still is high enough to produce the "perverse revenue effect" in a wide range of tax brackets. A more important fact, however, is that there is a downward bias in the estimates of these studies since they only are concerned with one dimension of labor supply, namely hours of work, while more important dimensions are left out, like work effort, choice of job, demand for education.

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A Procedure for Testing the Signalling Hypothesis

James W. Albrecht

1. Introduction*

The signalling model of the returns to education as developed by Spence (1974), Arrow (1973) and Stiglitz (1975) represents an important theoretical contribution to the economics of information, but whether this contribution is of significant empirical consequence is an open question. This paper develops and applies a general method for addressing this question.

The signalling interpretation of the returns to education depends upon employers' lack of information about job applicants. Workers (applicants) are assumed to have a good idea about their productivities, but, *a priori*, employers are not. If the less productive cannot be induced to admit to that fact, then the employer considering job applicants will be forced to "estimate" applicants' productivities.

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* This paper has gone through several versions and two data sets. The earliest version was presented at the 1974 Econometric Society meetings in San Francisco. The guidance and encouragement of Roy Radner on the early versions is gratefully acknowledged. This version has also been published in Journal of Public Economics, February 1981.

It is suggested that educational background may serve as an ideal observable trait for the employer to use to infer other, unobservable traits related to productivity. That education can be so used depends upon the assumption that the cost of education varies inversely with productivity. Under this assumption only the inherently more productive will find extra education worthwhile. Employers' initial beliefs that the educated are more productive will be self-fulfilling.

This basic objection to the signalling hypothesis is that the educational screen is a costly one. Ought not there exist less expensive alternative mechanisms to elicit information about productivities from applicants? It is sometimes asserted, for example, that any signalling component to the rewards to education would be eroded by the establishment of "testing firms". Alternatively, firms may be able to structure their promotion policies in such a way as to deter applicants from misstating their qualifications; that is, applicants may be induced to self-select into the proper job slots. These arguments, however, lack any empirical basis.

My approach to the signalling hypothesis will be to examine directly the question of whether employers reward education for purely informational purposes in the hiring decision. The role of education in the hiring decision will be decomposed into a pure "productivity component" and a pure "information component". This is most naturally done within a 2-way analysis of covariance framework with interactions between education and "information".

In the next section I develop the statistical procedure for testing the signalling hypothesis. Then, in the third section, I present an application of this method to a recruitment by the Swedish auto manufacturer Volvo. The results of this application provide both an illustration of the procedure and some substantive evidence about the signalling hypothesis. Finally, in a concluding section, I summarize the method and relate my procedure to another approach presented in Riley (1979).

2. A GENERAL PROCEDURE

Suppose an employer is considering applicants for a position who can be characterized by their educational background and by their "information level", i.e., the amount of *a priori* information the employer has about them. According to the signalling hypothesis, employers need to use education as a source of information about applicant productivities, i.e., applicants cannot be induced to properly selfselect by some cheaper means. Therefore, if the signalling hypothesis is valid, employers will be forced to rely more heavily on education when considering those applicants about whom they have the least information. The test procedure presented below is an exploitation of this simple idea.

Typically "information level" will be a qualitative variable, and often educational attainment will be as well. Let $i = 1,\ldots,I$ index educational categories, and let $j = 1,\ldots,J$ index informational categories. The k^{th} individual in the $(i,j)^{th}$

cell has observable characteristics X_{ijk}. Assume that the (lifetime, discounted, etc.) marginal product ($= Z_{ijk}$) of this applicant as perceived by the prospective employer can be expressed as a linear combination of these characteristics plus a $N(0,\sigma^2)$ error term. That is,

$$Z_{ijk} = X_{ijk}\eta + U_{ijk}, \tag{1}$$

where U_{ijk} is $N(0,\sigma^2)$.

The employer's decision problem can be modelled as one of accepting only those applicants whose perceived marginal product exceeds a critical value w. Then, the probability that the k^{th} applicant in the $(i,j)^{th}$ cell will be accepted can be written as

$$P_{ijk} = \Pr(Z_{ijk} > w) = \Pr(X_{ijk}\eta + U_{ijk} > w) =$$

$$= \Pr(U_{ijk} < X_{ijk}\eta - w)$$

$$= \int_{-\infty}^{(X_{ijk}\eta - w)/\sigma} (2\pi)^{-1/2} e^{-z^2/2} dz$$

$$= \Phi(X_{ijk}\eta^*), \tag{2}$$

where η^* is the standardized parameter vector and $\Phi(\cdot)$ is the distribution function of the standardized normal random variable.

To pursue the 2-way analysis of covariance approach, assume

$$X_{ijk}\eta^* = \mu + \alpha_i + \beta_j + \lambda_{ij} + \sum_{h=1}^{H} \delta_h Y_{hk}, \tag{3}$$

where

$$\sum_i \alpha_i = \sum_j \beta_j = \sum_i \lambda_{ij} = \sum_j \lambda_{ij} = 0.$$

The interpretation of the parameters is as follows:

μ = mean (standardized, perceived) productivity

α_i = main effect on productivity of being in educational category i

β_j = main effect on productivity of being in informational category j

λ_{ij} = interaction effect on productivity of being jointly in educational category i and informational category j

δ_h = effect of the h^{th} concommitant variable on productivity.

The main effects of education are the effects of educational categories averaged across all informational categories, and likewise for the main effects of information. The interaction effect in the $(i,j)^{th}$ cell is the effect of the i^{th} level of education on the employer's perception of applicant productivity specific to the j^{th} informational category; that is, it is the effect of the i^{th} level of education above and beyond the main effect, α_i.[1]

It is the interaction effects which are of principal interest. To see this it is useful to consider a simple "2×2" example. Imagine an applicant pool

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[1] The concepts of main effects and interaction effects in 2-way analysis of variance models are lucidly discussed in Scheffé (1959).

differentiated according to high versus low education level and high versus low information level.

If employers are forced to use education for information, then the interaction effects can be expected to take on the sign pattern indicated below:

		Information	
		High	Low
Education	High	−	+
	Low	+	−

We expect education to receive a positive overall weight in the employer's assessment procedure. If part of this positive overall weight can be ascribed to an informational component, then the positive effect of education ought to be decreased in the presence of alternative information; i.e., we expect the interaction effect for high education together with high information to be negative. Analogous arguments can be made to sign the other interaction terms, but these are redundant since there is only one independent interaction parameter in this 2x2 case. Alternatively, if the employer is not forced to use education as a source of information, then the effect of education should be constant across all information levels. Thus, a test of the hypothesis that the employer does not use education for informational purposes may be expressed as

$$H: \lambda_{ij} = 0; \quad i=1,\ldots,I \quad j=1,\ldots,J.$$

It is to be emphasized that the hypothesis of zero interaction effects is <u>not</u> the hypothesis that the employer is indifferent about the educational attainment of applicants, nor is it the hypothesis that the employer is indifferent about the amount of *a priori* information available about prospective employees. These hypotheses instead translate into hypotheses about the main effects.

Nor does the hypothesis of zero interaction effects imply that an employer's preference for applicants about whom more information is available need solely reflect a preference for more information. There may be differences in average productivity across information classes, but these differences ought to be reflected in the main effects of information, rather than in differential rewards to education. However, one must be on guard for other mechanisms that might introduce an interaction between education and information, and such alternative mechanisms are easier to imagine when information is not "neutral". The point, of course, is that one must be careful in specifying "information classes".

3. AN APPLICATION

The data used in this application come from records of applicants for entry-level blue collar positions at Volvo's Torslanda auto works for the month of June 1978.[1] Excluding those applying for

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[1] These data were kindly made available to me by Göte Bernhardsson and Anne-Marie Qvarfort of the Employment Commission in the Swedish Ministry of Labor (sysselsättningsutredningen). Their report on Volvo's recruitment practices is available in mimeo as "Personalrekryteringen till Volvo-Torslandaverken, Juni 1978", *Sysselsättningsutredningen*, October 1978.

part-time work, a total of 515 applicants were considered and of these 291, or 56.5%, were hired. Data on the educational attainment and on the recruitment source of each applicant are available from these records. Educational attainment is a dichotomous variable with "low education" identified with attainment of less than the gymnasium level. The gymnasium is normally attended for 3 years in Sweden between the ages of 16-19 and roughly corresponds to the last years of senior high school plus parts of junior college in the U.S. Today the completion rate in the gymnasium is quite high, but this is a very recent phenommenom, and in this sample 42% of the applicants have not completed the gymnasium.

The information class of the applicant is identified with the source of his or her recruitment. The first recruitment source - and this is the source to be identified with greater prior information - is recommendation by a current Volvo employee; that is, the applicant has given the name of a Volvo employee who has informed him of the job opening and from whom the personnel department can solicit an evaluation. Of course, such an evaluation may not be unbiased, but it seems reasonable that the company can take the caliber of the reference into account. The other two recruitment sources are identified with less prior information. The first of these relatively low information sources is the Swedish Labor Market Board (AF). This refers to job seekers who have searched AF's position announcements and have then come to Volvo with a notification from that Board. No active placement on AF's part is implied. Secondly, there are those who have simply applied in response to newspaper advertisements (plus a small group from "miscellaneous" sources). In principle, those who come via AF and those who come via advertisement are in an equally low information category. However, there is the possibility of more active placement on the part of the AF for some candidates. This potentially has both the implication of more information and the implication of a decrease in the probability of hire for those candidates since AF is more likely to make an active effort on behalf of those who are "difficult to employ". These two low information categories have been combined in the empirical results presented below.[1]

Besides the information about education and recruitment source, data are available on the age, the nationality, the residence and the sex of each applicant. These data are presented in Table 1. Ignoring any covariation between these variables for the moment, Table 1 indicates a preference for (1) more highly educated applicants, (2) applicants in the high information category, (3) younger applicants, (4) Swedish and Finnish nationals, (5) non-Gothenburg residents and (6) males. The only surprise in the data is the preference given to those living out of the greater Gothenburg region where the plant is located. However, the relatively low number of non-Gothenburg residents

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[1] In fact, Bernhardsson and Qvarfort conjecture that some applicants recorded as recruited via advertisement may also have searched the AF position announcements. There are 2 bases for this suspicion: (i) some applicants may feel that any identification with AF hurts their chances and (ii) the fraction of applicants coming from AF seems "abnormally low".

Table 1 The basic data

		Appli-cants	Hired	Relative frequency
	Total	515	291	.565
Education:	Low	215	101	.470
	High	300	190	.633
Information:	Rec	180	110	.611
	AF	115	58	.504
	Ad	220	123	.559
Age:	$\leqslant 20$	202	129	.639
	21-27	186	104	.559
	$\geqslant 28$	127	58	.457
Nationality:	Swedish	298	182	.611
	Finnish	122	76	.623
	Other	95	33	.347
Residence:	Gothenburg	415	228	.549
	Other	100	63	.630
Sex:	Male	455	270	.593
	Female	60	21	.350

Source: Unpublished data from the Employment Commission in the Swedish Ministry of Labor.

(and the even lower number of females) among the applicants should be noted.

The model that has been estimated inverts equation (2) to express $\Phi^{-1}(p)$ as a constant plus a sum of main effects for education, information, age, nationality, residence and sex plus an education-information interaction. The parameters have been estimated using maximum likelihood (probit), and test statistics for assessing the significance of the main and interaction effects have been computed as -2 times the logarithm of the appropriate likelihood ratio. The test statistics are asymptotically χ^2 with degrees of freedom equal to the number of independent restrictions implied by the null hypothesis. These parameter estimates and test statistics are presented in Table 2.

The parameter estimates may be interpreted with the aid of a simple example. An applicant who (1) has a low level of education, (2) falls in the high information category, (3) is between the ages of 21-27, (4) is Swedish, (5) is a Gothenburg resident and (6) is male would be hired with an estimated probability of $\Phi(0.305) = 0.620$. An applicant with a high level of education but otherwise identical attributes would be hired with an estimatted probability of $\Phi(0.569) = 0.715$ with the change ascribable to the increase via the main effect of education (from -0.191 to +0.191) and to the decrease via the education-information interaction (from +0.059 to -0.059).

The pattern of main effects in Table 2 is in basic accord with that suggested by the raw data in Table 1. Completion of the gymnasium, Swedish or Finnish nationality and being male strongly increase the chance of getting hired, and these main effects are significant at the 1% level. Having a Volvo employee to use as a reference also increases the hire probability, but not as strongly; and the factors of age and residence, while retaining the same pattern as in the raw data, become much less important. In fact, the anomalous apparent preference for non-Gothenburg residents essentially becomes zero when the covariation between residence and other variables is taken into account. The significance probabilities for the main effects of information, age and residence (0.15,

Table 2 Probit estimates and significance tests

Estimate	Maximum likelihood (χ^2)	Test statistic
Mean	-0.145	-
Main effects		
Education		9.48*
High	0.191	
Low	-0.191	
Information		2.18
High	0.115	
Low	-0.115	
Age		1.98
⩽20	0.095	
21-27	-0.022	
⩾28	-0.073	
Interaction effect		0.54
High Ed. × High Inf.	-0.059	
High Ed. × Low Inf.	0.059	
Low Ed. × High Inf.	0.059	
Low Ed. × Low Inf.	-0.059	
Nationality		14.78*
Swedish	0.185	
Finnish	0.240	
Other	-0.425	
Residence		0.34
Gothenburg	-0.016	
Other	0.016	
Sex		10.98*
Male	0.320	
Female	-0.320	

* Significant at 1 percent level.

0.35 and 0.65, respectively) are above conventionally accepted levels.

The interaction effects take on the sign pattern suggested by the signalling hypothesis, i.e., the positive effects of extra education are decreased in the presence of extra information, but these effects are quite small in magnitude. The significance probability for the education-information interactions is only slightly less than 0.5. The hypothesis of zero interaction effects clearly cannot be rejected; that is, the hypothesis that there is no purely informational component to the preference exhibited for the more educated applicants cannot be rejected. Volvo's hiring behavior gives no support to the signalling hypothesis in this instance.

Finally, it should be noted that the results are insensitive to re-parameterization of the basic model. Alternative models have been estimated with (1) age as a continuous variable, (2) 3 information categories instead of 2, (3) interactions between education and nationality and information and nationality and (4) residence and sex supressed as separate variables. In addition the model with residence and sex supressed has been re-estimated by the alternative technique of "minimum normit chi-square", i.e., weighted least squares based on the cell relative frequencies, as developed by Berkson (1955). The basic conclusions remain the same.

4. DISCUSSION

This paper has presented a procedure for testing the signalling hypothesis based on a decomposition of the role of education in the hiring decision into a pure "productivity" component and a pure "information" component. The procedure was applied to a recruitment of auto workers by Volvo, and in this instance Volvo's hiring behavior indicates no support for the signalling hypothesis. Volvo prefers applicants with more education and (weakly) prefers applicants about whom more information is available, but in the absence of that extra information no significantly different premium is attached to extra education. That is to say, Volvo does not appear to rely on education for purely informational purposes in the hiring process.

Of course, this same procedure could be applied to different sets of data, and one aim of this paper is to motivate the collection of richer data sets for replication. As explained above, and as illustrated in the Volvo application, the trick is to define the concept of "information level" in a suitable way.

The procedure developed in this paper is very "micro" in the sense that it focuses on the significance of signalling at the level of the individual job and at the level of the individual employer. More "macro" approaches are also possible, and such approaches can be considered complementary to the method advocated here. In my opinion, the best of these macro approaches is presented in Riley (1979).[1] Riley's method is based on an idea

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[1] Some other empirical papers on signalling are Layard and Psacharopoulos (1974), Taubman and Wales (1973), and Wolpin (1977). Riley gives a good discussion and critique of these papers.

similar to that of information levels. However, instead of differentiating among applicants for a particular job according to the amount of available prior information, he divides occupations into those for which productivity may be easily ascertained versus those for which signalling might conceivably be important. A test of the signalling hypothesis is then based on a comparison of lifetime earnings functions at each level of education for those in the "screened" sector versus those in the "unscreened" sector. Using this test, Riley concludes that signalling is a significant phenomenon.

However, as Riley points out, there is no obvious best method for classifying occupations as screened or unscreened. In fact, he is forced to use *ex post* data analysis to perform the classification. Nor is there any way to ensure the differences in earning profiles between the screened and unscreened sectors for a given education level can be solely ascribed to the screening function of education. But these practical problems are analogous to those which make the application of this paper's procedure difficult; namely, suitably defining what one means by "information" and ensuring that specious interactions between education and information are controlled.

To summarize, empirical analysis of the significance of signalling appears to have reached the point where well-founded techniques are becoming available. However, the data requirements imposed by these techniques have proved to be rather stringent. One advantage of the procedure and application presented in this paper is that these data requirements have been clarified, and one can hope that further applications will be possible.

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Part IV
Determinants of Wage Increases

Disequilibrium and Non-Neutral Market Effects on Age-Earnings Profiles

Anita Jonsson and Anders Klevmarken *

1. INTRODUCTION

The development of individual earnings over time reflects current and past individual decisions as well as changes in the economic environment. In the human capital theory the rate of increase in earnings is explained by investments in the earnings capacity of each individual. However, the curvature of age-earnings profiles does not only depend on investments in human capital but also on the general growth of the economy and the conditions on the particular sector of the labor market to which these individuals belong. This study attempts an integration of studies of the income distribution in the human capital tradition and studies of wage increases related to the Phillips curve.

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This paper was also presented at the symposium "Statics and Dynamics of Income", September 20-21, 1979 at the University of Gothenburg, and it will appear in a conference volume with the same title edited by A. Klevmarken and J. Lybeck.

Previous studies of age-earnings profiles have, with few exceptions (cf. Hanushek and Quigley, 1978), assumed neutral market and growth effects on earnings. This implies parallel cross-sectional (log) earnings profiles. Growth and market are thus assumed to induce parallel shifts of the cross-sectional profiles but not to influence their shape. Our study differs from previous studies in the human capital approach in that our model explicitly includes the effects of short-run changes in demand and supply of labor on the earnings profiles. It also attempts to estimate an effect of negotiations net of the market effect. The model also allows these effects to differ in magnitude for initial earnings and for the earnings of those who already are in the market, i.e., non-neutral effects.

Our study also differs from previous studies of the Phillips curve and related studies of wage-drift, etc., in several aspects. It uses micro-data and not aggregate time-series data and it analyzes the changes in earnings of salaried employees and not wage changes of "blue collar" workers. The emphasis partly on distributional aspects explains why our model seeks to explain real earnings rather than nominal earnings. We also treat the problem of wage drift versus negotiated wage increases somewhat differently from the approaches commonly used in Phillips curve studies. We do not assume that the total increase in earnings is the sum of two components - wage drift and negotiated wage increases - with two independent explanations, nor do we assume that the two components are governed by the same forces, but we take the more general approach that the negotiated increase contributes to the total increase in earnings by a factor which might differ from one or zero.

In most studies of the Phillips curve the tightness of the market is measured by the unemployment rate, the number of unfilled vacancies or the difference between vacancies and unemployment. An excess demand (supply) variable is used in this study too but in addition there are shift variables for the demand and supply functions. The rationale for this specification is that total changes in earnings have an equilibrium component as well as a disequilibrium component; the latter is proportional to the excess demand (supply) variable.

This paper continues and extends earlier studies of the earnings profiles for engineers in Sweden by the present authors (Klevmarken, 1974, and Jonsson and Klevmarken, 1978). In these studies the sample covered the period 1961-1970. Additional data for 1973-1976 now make possible a more powerful test of the model. The market for engineers in Sweden gradually changed during this entire period from excess demand to excess supply and at the same time the relative importance of the negotiations between the employers and the unions increased. It will be shown below that the model used in Jonsson and Klevmarken (1978) cannot fully predict the observed earnings profiles for 1973-1976 which suggests a respecification of the model.

The model, first suggested in Klevmarken (1974) and then also used in Jonsson and Klevmarken

(1978), is reproduced with only minor modifications in Section 2 below. Empirical results from the sample period 1961-1970 as well as predictions for 1973-1976 based on these results are given in Section 3. Section 4 gives the estimates of the same model for the entire sample period 1961-1976 while the model is respecified and reestimated in Section 5.

2. A MODEL FOR THE EARNINGS PROFILES OF ENGINEERS WITH NON-NEUTRAL MARKET EFFECTS

An individual earnings profile is formally represented by the following two relations:

$$\ell n\ y_{bbi} = \alpha_0 + \sum_{t=1}^{b} \beta_t + \delta_{bi} \tag{1a}$$

$$\ell n\ y_{bTi} = \ell n\ y_{bbi} + \sum_{t=b+1}^{T} \gamma_{bti} \tag{1b}$$

where

y_{bbi}	is the initial earnings of an individual i who started his career in year b,
α_0	is the average initial log-earnings in a base period 0,
β_t	is the increment in average initial log-earnings from year t-1 to year t,
δ_{bi}	is the individual deviation from the average initial log-earnings in year b,
y_{bti}	is the earnings in year T of an individual i who started his career in year b,
γ_{bti}	is the increment in log-earnings from year t-1 to year t of an individual i who started his career in year b.

Different assumptions about the annual increments β_t and γ_{bti} in this formal representation result in different models for age-earnings profiles. A simple model which has been extensively used in the human capital approach (cf. Mincer, 1974) is the following

Model I:

$$\beta_t = \beta_0 + \varepsilon_t^\beta \tag{2a}$$

$$\gamma_{bti} = \gamma(t-b) + \varepsilon_{bti}^\gamma \tag{2b}$$

The annual increments of the average initial earnings are assumed constant except for a random disturbance. For individuals in the career the increase in earnings is a result of investments in human capital accounted for by years of experience (t-b).

Model I was used as reference model and compared with the following more general model which specifies earnings as a function of years of experience and calendar-time specific factors.

Model II

$$\beta_t = \beta(x_t) + \varepsilon_t^\beta \tag{3a}$$

$$\gamma_{bti} = \gamma(t-b,\ x_t) + \varepsilon_{bti}^\gamma \tag{3b}$$

In this model increases in earnings do not only depend on investments in human capital but also for example on the general growth in the economy, supply and demand conditions in the labor market, and the outcome of negotiations. The vector x_t, which will be specified in more detail below, includes these time-specific factors. In order to complete the specification of equations (3a) and (3b) we have to analyze the particular sector of the labor market to which the engineers belong.

The market for new graduates was analyzed in Klevmarken (1974) by a model of market disequilibrium with five structural relations, namely functions for demand and supply of new graduates, a definition of an equilibrium salary, a salary reaction function and a function for negotiated salary increases.

The demand and supply functions were respectively,

$$D_t = d_0 + d_1 W_t + d_2 X_{1t} + \varepsilon_{dt}; \tag{4}$$

$$S_t = s_0 + s_1 W_t + s_2 X_{2t} + \varepsilon_{st}; \tag{5}$$

W_t is log of earnings measured as the stipulated monthly salary. X_{1t} is the log of an index of industrial production, which was assumed to catch shifts in demand for engineers. The shift variable in the supply function X_{2t} is the log of the number of new graduates.[1] ε_{dt} and ε_{st} are random errors with zero expectation and d_0, $d_1, \ldots, s_1$, s_2 are parameters.

The logarithmic equilibrium salary $\bar{W}_t$ was defined by,

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[1] A more detailed account for the data is given in Klevmarken (1974), Appendix A.

$$D_t=S_t \rightarrow \bar{W}_t=\frac{1}{d_1-s_1}\,(s_0-d_0-d_2X_{1t}+s_2X_{2t}-\varepsilon_{dt}+\varepsilon_{st}), \quad (6)$$

which differs from the observed salary as long as the market is not in balance, i.e., from eqs.(4), (5) and (6)

$$W_t = \frac{1}{d_1-s_1}\,(D_t-S_t) + \bar{W}_t. \quad (7)$$

We will now assume that there are constraints on the adjustment mechanism of the market which prevents a balance to be reached in each period. Earnings are assumed to adjust according to the following relation,

$$W_t-W_{t-1} = w_0+w_1(\bar{W}_t-\bar{W}_{t-1}) + w_2(\bar{W}_{t-1}-W_{t-1}) + w_3\Delta A_{t-1} + \varepsilon_{wt-1}. \quad (8)$$

The relative change in earnings thus depends on the change in the equilibrium salary, the size of the disequilibrium and the negotiated rate of increase ΔA_{t-1}.[1] This model includes as a special case, if $w_1=w_2$, that the rate of change in earnings depends only on the whole difference between the equilibrium salary at t and the observed salary at t-1. If, however, it takes time for the market to transmit information about shifts in the supply and demand curves to employers and employees, they might know the magnitude of last years

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[1] Since W_t-W_{t-1} is the log-change in real earnings ΔA_{t-1} is the negotiated relative change less the relative change in the consumers price index.

disequilibrium salary. The disequilibrium component might thus have a stronger influence on changes in earnings than the equilibrium component.

The special case $w_3=1$ is equivalent to a model which explains wage-drift by market changes, while the negotiated increase is exogenous or explained independently of the wage-drift. $w_3=0$ is the case of complete market determination of salary increases. Since the model is applied to annual data it is reasonable to allow intermediate cases between these two extremes. We do not necessarily have to treat ΔA_t as an exogenous variable. The result of central negotiations presumably depends on the tightness of the labor market, the general economic outlook which employers have to face, increases in consumer prices, etc., as well as the particular policy of the unions. Fortunately we do not have to specify these relations in detail.

Substitution of eqs.(6) and (7) into eq.(8) gives,

$$W_t-W_{t-1}= w_0+ \frac{w_1 d_2}{d_1-s_1}(X_{1t}-X_{1t-1})+ \frac{w_1 s_2}{d_1-s_1}(X_{2t}-X_{2t-1}) +$$

$$\frac{w_2}{d_1-s_1} (D_{t-1}-S_{t-1})+ \frac{w_3}{d_1-s_1} \Delta A_{t-1}+u_{t-1}; \qquad (9)$$

where u_{t-1} is a composite random error with zero expectation. In order to simplify and obtain conformity with the notation of eq.(3a), eq.(9) is rewritten in the following way

$$\beta_t = \beta_0+\beta_{11}x_{1t}+\beta_{12}x_{2t}+\beta_{13}x_{3t-1}+\beta_{14}x_{4t}+\varepsilon_t^{\beta}; \qquad (10a)$$

where

$$\beta_t = W_t - W_{t-1}$$

$$x_{it} = X_{it} - X_{it-1} \; ; \quad i = 1,2$$

$$x_{3t} = D_t - S_t$$

$$x_{4t} = \Delta A_{t-1}$$

Changes in earnings for those who already are in the market are explained by the same model but with one modification. The supply of new graduates is assumed not to influence these changes. The introduction of non-neutral growth and market effects makes less straightforward the interpretation of years of experience as an indicator of investments in human capital. Changes in the shape of profiles and in the life-time earnings of alternative careers will certainly influence the incentives to invest. The supply of labor with various kinds of training will change and subsequently influence the earnings profiles. The view taken in this study is that investment activities only change gradually and that there is a "normal" or long-run return to on-the-job training, which can be separated from short-run or medium term fluctuations due to market imperfections. The effect of the years of experience variable is thus assumed to be sufficiently stable to justify the following model,

$$\gamma_{bti} = \gamma_0 + \gamma_1 (t-b) + \gamma_{21} x_{1t} + \gamma_{23} x_{3t-1} + \gamma_{24} x_{4t} + \varepsilon^{\gamma}_{bti} . \quad (10b)$$

The disequilibrium variable x_{3t} was measured as

$$x_{3t} = {}^{e}\log(TDN_t/0.285TS_t),$$

where

TDN_t = number of job openings for engineers and technicians advertised in Dagens Nyheter in year t

TS_t = number of engineers employed in Swedish industry in year t.

The constant 0.285 is an estimate of the ratio of the number of jobs advertised to the number of engineers when the market is in balance. It is an average for the two years 1966 and 1969. The status of the market was determined from Labor Market Tendency Surveys.[1] A discussion of this and other measures is deferred to later. The observed series are drawn in Figures 1-4.

During the 1950's and the first half of the 1960's there was an excess demand for engineers in Sweden. A large increase in the output of new graduates and a decreased additional demand resulted in a gradual decrease in the excess demand; by 1966 the market was in balance; in 1967 and 1968 there was a minor excess supply and in 1969 the market was in balance again. Since then the supply of new graduates did not show a trendwise change and except for the years 1973-1974 demand did not increase much either. The result became an excess supply of engineers.

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[1] Forecasting Information (IPF), Swedish Central Bureau of Statistics, Stockholm.

In Klevmarken (1974) it was found - although with a slightly different model - that the negotiated increases did not contribute to the explanation of the rate of increase in initial earnings. For this reason β_{14} and γ_{24} was *a priori* set to zero in Jonsson and Klevmarken (1978). The model obtained with this constraint is below labelled Model II:1.[1] In other model versions we will use the more general formulation.

The age-earnings profiles which are generated from the different models are obtained by inserting the expressions of average increments and in the formal relations (1 a) and (1 b). Using the expressions (2a) and (2b) in Model I and assuming $\gamma(t-b)$ to be a linear function of (t-b) the earnings profile becomes

$$\ell n\ y_{bTi} = \alpha_0 + \beta_0 b + (\gamma_0 + \frac{\gamma_1}{2})\ (T-b) + \frac{\gamma_1}{2}\ (T-b)^2 + u_{bTi}; \tag{11a}$$

where

$$u_{bTi} = \delta_{bi} + \sum_{t=1}^{b} \varepsilon_t^{\beta} + \sum_{t=b+1}^{T} \varepsilon_{bTi}^{\gamma}. \tag{11b}$$

The systematic part of equation (11a), identical in form to the specification commonly used in the human capital literature, gives logarithmic earnings as a second degree polynomial in years of experience. The individual earnings profile im-

[1] Model IIa in Jonsson and Klevmarken (1978).

plied by the specifications (10a) and (10b) of Model II with β_{14} and γ_{24} both set to zero is

$$\ell n\ y_{bTi} = \alpha_0 + \beta_0 b + \beta_{11} \sum_{t=1}^{b} x_{1t} + \beta_{12} \sum_{t=1}^{b} x_{2t} +$$

$$\beta_{13} \sum_{t=1}^{b} x_{3(t-1)} + (\gamma_0 + \frac{\gamma_1}{2})(T-b) + \frac{\gamma_1}{2} (T-b)^2 +$$

$$\gamma_{21} \sum_{t=b+1}^{T} x_{1t} + \gamma_{23} \sum_{t=b+1}^{T} x_{3(t-1)} + u_{bTi} \qquad (12)$$

where

u_{bTi} is given by (11b)

The corresponding cross-sectional profile for a given year T_0 is obtained by holding $T=T_0$ constant in (12) and varying b and i. As discussed and illustrated in greater detail in Jonsson and Klevmarken (1978), the curvature of this cross-sectional profile will depend on the market conditions between the year when the oldest employee entered the market and the year T_0. If the effects of changing market conditions are neutral with respect to years of experience ($\beta_{1j}=\gamma_{2j}$ for all j) the curvature of the cross-sectional profile is the same as that of the profile derived from Model I. If, however, these effects are non-neutral, the estimates of for example the returns to investments in human capital based on Model I become biased.

The parameters in Model I and Model II:1 were estimated from pooled cross-sectional earnings data on Swedish graduate engineers specialized in electrical engineering. The earnings data were obtained from the Swedish Association of Graduate Engineers for the years 1961-1970. These individuals were grouped by physical age and active age (the number of years since graduation) and the mean logarithmic monthly salary were given for each combination. Before estimation mean earnings were discounted to 1961 Swedish kronor. Because data for our market indicators were not available before 1954, those who entered the market before this year were eliminated from the sample. After this reduction, our data included 987 observations of mean earnings, grouped by physical and active age.

Efficient estimation of the parameters in the models requires realistic assumptions about the stochastic properties of the residuals in the estimated equations. Since mean earnings are given, information about the properties of the mean $u_{bT}.$ of the residuals u_{bTi} in (11a) and (12) is needed. According to (11b) these residuals are built up by sums of the residuals in the equations (1a), (2a) and (2b). Partly for convenience we made the assumptions that δ_{bi}, ε_t^β and ε_{bti}^γ have zero expectation and constant variance σ_δ^2, σ_β^2 and σ_γ^2 respectively for all values of b and t. Furthermore, we assumed that all covariances are zero, that $\sigma_\delta^2=\sigma_\gamma^2=\sigma^2$ and that σ_β^2 is much smaller than σ^2. These simple assumptions imply that the variance of $u_{bT}.$ is a function of b and T according to the following approximate expression

$$\text{Var}\ (u_{bT}.) \approx \sigma^2 \frac{(T-b+1)}{n_{bT}} \tag{13}$$

The property that the variance increases with active age (T-b) agrees with previous experience from earnings data (cf. Klevmarken, 1972). The estimation method was weighted least squares with the weights implicitly given by (13).

In order to investigate the properties of our estimates we have to scrutinize the estimated equation carefully. One condition that must be fulfilled, if the least squares estimates are to be consistent, is that all the independent variables in the equation are uncorrelated with the residual. In this study the shift variables in the demand and supply functions (x_1 and x_2) are exogenous non-stochastic variables and thus uncorrelated with the residual. The excess demand (supply) variable (x_3) is an endogenous stochastic variable. It is thus possible that the presence of this variable makes our estimates inconsistent and biased. For instance, it may be argued that the deviation from equilibrium on the market in a certain year t_0 depends on the increase in ear nings in the previous years. Since the error components ε and ε according to eqs.(10a) and (10b) contribute to the disequilibrium for $t<t_0$ and since they also are part of the residual in the estimated relation, there would be a correlation between the residual u_{bTi} and the disequilibrium variable $\Sigma x_{3(t-1)}$ in eq.(12). However, this correlation is probably small since earnings data refer to electrical engineers, while the variable x_3 refers to a much larger sector of the labor market, namely to all engineers and technicians in Swedish industry.

If x_4, the negotiated rate of increase, is included there is also another potential source of bias. If x_4 is an endogenous stochastic variable, and if this variable depends on earnings, then simultaneity makes x_4 correlated with the residual in the estimated relation. However, as the indicator variable x_4 refers to negotiated salary increases for all male employees and the variable x_3 is a measure of disequilibrium on a much smaller sector of the labor market, namely the market for engineers and technicians, the correlation between these variables will be small and the conditions for consistent least-squares estimates will be satisfied in practice.

The efficiency of our estimation method depends on the agreement between the assumed and the true error structure. A plausible alternative would be to assume that the errors ε_{bi} and ε_{bti} in the equations (1a) and (2b) include a random individual component. This implies a non-zero correlation between the residuals in the individual earnings profiles, i.e., between u_{bti} and u_{bTi} for $t \neq T$. Although our data are not longitudinal, many individuals have been observed for several years and we would thus expect a minor correlation between the residuals in the estimated equation, i.e., between $u_{bt.}$ and $u_{bT.}$ for $t \neq T$. However, since we are not able to trace an individual from one year to another, it is not possible to estimate and account for the variance of this individual component in the estimation procedure. A consequence of disregarding the non-zero correlation between the residuals is inefficient estimates, but this is counter-weighted by a relatively large sample. Another consequence of not using the true error structure is biased estimates of the residual variance (σ^2) in expression (13). This will influence both the significance level of the tests used in the following sections and the length of the prediction intervals given in Section 3. These possible short-comings will most certainly not invalidate the main conclusions drawn in this study.

3. EMPIRICAL RESULTS FROM THE SAMPLE PERIOD 1961-1970 AND PREDICTIONS FOR 1973-1976

Estimates of Model I and two versions of Model II are presented in Table 1 for the sample period 1961-1970. Model II:1 does not include the negotiated rate of increase while Model II:2 does. Standard errors of the estimates are given in parenthesis. The values of the adjusted coefficients of determination ($\bar{R}^2$) show that the closest fit is obtained for Model II.[1] According to Model I, the average yearly increase in initial real earnings, obtained from the estimate of β_0 is 1.1%. For those already in the market, the predicted increase, obtained as the sum of γ_0 and γ_1, is 9% during the first year of experience. After this year, the rate of increase is a declining function of active age.

The specification of Model II:1 allows yearly increases in earnings to vary over time depending on labor market conditions. Information about the market is thus needed in order to obtain predictions from this model. The predicted increases for

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[1] Note that the models are fitted to data on mean earnings and not to individual data which partly explains the high $\bar{R}^2$.

a certain year during the period 1955-1970 can be obtained from the relations (10a) and (10b) with the parameter estimates in the second column of Table 1 and the values of the indicator variables (x_1, x_2 and x_3) inserted. For instance, the predicted rates of increase in initial earnings range from -0.6% in 1969 to 3.9% in 1957.

The estimated partial effects of changes in demand, supply and in the market disequilibrium all conform to the expected signs. The β-estimates show a higher absolute value than the corresponding γ-estimates and we thus find that the effects of changing labor market conditions on initial earnings are larger than the effects on earnings for those who are already in the market. Contrary to previous results the variable for negotiated increases contributes significantly. This variable also lacks in neutrality. As mentioned in the previous section, the implication of non-neutral time specific effects is unstable cross-sectional profiles. There is thus no simple model of years of experience only from which the effect of changing experience can be estimated.

A good test of a model is to use it for predictions. From Model I it is possible to predict age-earnings profiles over an arbitrary range of active age. Profiles from Model II which includes the effects of changes in labor market conditions can also be obtained over a large range provided that these conditions are known for a long time period. Because the market indicators are not known before 1954 we can not in this study predict earnings for engineers with more than 22 years of experience. The test of the models must therefore rely on comparisons between observed and predicted profiles over a rather short period of experience. The forecasting period is 1973-1976. Also for this period earnings data are given as average monthly salaries, grouped by age and the number of years since graduation. The observed mean of log earnings for each group of active age is the weighted average of the log geometrical means in each physical age group. The observed and predicted cross-sectional profiles for the years 1973 and 1976 are shown in Figures 5 and 6. Except for differences in levels, the profiles for 1974 give almost the same picture as the 1973 profiles, while the 1975 and the 1976 profiles are similar.

The predicted profiles were adjusted by the consumer price index to make them comparable to the observed profiles in current prices. As the figures were drawn in a logarithmic scale this adjustment only results in a difference in level but not in the curvature of the profiles. To investigate the influence of random variation on the predicted profiles these have been supplemented with 95% prediction intervals.[1] For Model I and Model II:1

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[1] The half length of the intervals for the mean of earnings in an age group is calculated from the expression

$$t_{\nu,0.975}\left[\underline{x}_0'\,\underline{V}(\hat{\beta})\underline{x}_0 + \hat{\sigma}^2\,\frac{T-b+1}{n_{bT}}\right]^{1/2}$$

where $\underline{x}_0$ is the prediction vector, $\underline{V}(\hat{\beta})$ is the covariance matrix of the parameter estimates and $\hat{\sigma}^2$ is the estimate of the residual variance based on ν degrees of freedom. The length thus depends on both the variance of the estimate of the expected mean, active age (T-b) and the number of individuals in the age group (n_{bT}).

the intervals are displayed in Figures 5 and 6 for selected years of age.

Let us first compare the predicted and the observed profiles for 1973 in Figure 5. While the simple Model I overestimates the observed profile, Model II:1 which includes time specific effects gives good prediction in all age intervals. Model I overestimates the earnings profile also for the following years. For 1975 and 1976 Model II:1 also gives large deviations between observed and predicted profiles. It underestimates initial earnings and the earnings for those who only have a few years of experience. The extent of this underestimation is augmented for each year while the difference in the curvature between observed and predicted profiles increases. As observed mean earnings for several years of active age lie outside the corresponding prediction interval, prediction errors cannot be the result of random variation. The predictions from Model II:2 are better than those from Model II:1, but the prediction errors show the same systematic pattern. These results thus show that there is also a need for modification of Model II. Before we enter into a respecification of the model it might be useful to investigate if only a reestimation is sufficient to bridge the gap between predicted and observed profiles.

4. EMPIRICAL RESULTS FROM THE SAMPLE PERIOD 1961- 1976

By adding earnings data from the years 1973- 1976, our sample is extended to 14 cross-sections ranging from 1961 to 1976. The number of observed mean-earnings is increased from 987 to 1393. Although earnings data for the years 1971-1972 are missing it is possible to estimate the model from the extended sample. Information about the labor market is however required also for these two years. The new estimates of Model 1, Model II:1 and Model II:2 are given in Table 2. They were obtained by weighted least squares in the same way as previously. All parameter estimates in Model I are lower than those estimated for the sample period 1961-1970 (compare the results in Table 1 and Table 2). For instance, instead of the previously predicted growth rate of 1.1% in real initial earnings the model now predicts a decrease of 0.4%. This instability is easily explained, since Model I does not account for the changes in the market between the two sample periods. These findings obviously show that the assumption of a constant rate of increase in initial earnings, except for a random error, is not an adequate description of the growth of earnings.

We now turn to the results for Model II:1 in the second column of Table 1 and Table 2. As this model includes calendar year specific effects we expect small differences in the parameter estimates. However, also in this model there are differences between the estimates from the two periods. Moreover, a few parameter estimates are difficult to explain. For instance, the effect of an increase in demand would be negative! Model II:2 gives the same results.

Although these findings confirm that the model should include time specific effects they also cast doubts on the model specification. One possible explanation to these somewhat discouraging results is that the labor market indicators used do not measure what they are supposed to measure or do not do so persistently. These problems will be analyzed in more detail in the next section.

5. ALTERNATIVE MODEL SPECIFICATION

Model II assumes that shifts in the demand curve is adequately accounted for by an index of industrial production. Demand for highly educated manpower like graduate engineers might however be relatively invariant to short-run changes in output, at least as compared to "blue collar" workers. We therefore smoothed the time-series of the rate of changes in industrial production by a moving average and reestimated the model. The results, however, did not improve and they were relatively insensitive to the span of the moving average.

We experienced more success when the aggregate output measure for the entire industry[1] was replaced by the corresponding index for Manufacture of Electrical Machinery, Apparatus and Supplies[2]. This latter index is probably a better proxy for shifts in demand for electrical engineers, since most electrical engineers will be found in this sub-industry. An even better proxy might have been an index with weights proportional to the number

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[1] Industri 2 and 3 according to Swedish Standard Industrial Classification of all Economic Activities (SNI).

[2] Group 383 according to Swedish Standard Industrial Classification of all Economic Activities (SNI).

of electrical engineers in each sub-industry, but this would involve the compilation of a completely new index. Figure 1 shows that output from Manufacture of Electrical Machinery, Apparatus and Supplies increased more than the total industrial output and in particular so for the years 1974-1975. The index for the whole industry was 131 in 1976 (1968=100) while it was as much as 175 for Manufacture of Electrical Machinery, etc.

All estimates with these new data have the expected signs and all standard errors are satisfactory small (Table 3). The fit has improved a little compared to the results in Table 2. A comparison of the two model versions in Table 3 shows that there is a net effect of negotiated increases on total increases and that it is close to 1, i.e., a negotiated real increase of one percent will increase total real earnings by approximately the same amount. The net effect on initial earnings is a little stronger than the effect on earnings for those who have a job. Like the estimates of previous model versions these new estimates show also that the market effects are non-neutral. The point estimate of β_{13} is almost twice the estimate of γ_{23} and the standard error for the difference is small.

The excess demand variable is another proxy which might not have quite the same properties as its theoretical correspondence. It could be argued that the measure of excess demand should reflect both the flow of new vacancies and the duration of vacancies. The number of job openings advertised will first of all depend on the flow of new vacancies, but if recruitment difficulties result in more than one ad per unfilled vacancy it will also to some extent depend on the duration of vacancies. Time-series for the number of new vacancies and number of unfilled vacancies for technicians are available in the Labor Market Statistics since 1969.[1] A comparison for the years 1969 - 1975 shows a very close agreement between the number of ads (TDN) and these two series. The correlation with the number of new vacancies is 0.93 and with the number of unfilled vacancies 0.97. These comparisons thus indicate that the number of ads do reflect both the flow of unfilled vacancies and their duration, but seven years are admittedly a short period for a comparison and more data would be needed for a more definitive conclusion. However, on the bases of these results it is not justified to argue that our disequilibrium variable systematically overestimates the size of the excess supply for the last five years of the sample period.

Even if there is no systematic measurement error in the disequilibrium variable Model II might still react too strongly on an excess supply of engineers.

With the present model specification an excess supply and an excess demand of equal magnitude would imply a decrease and an increase respectively in real earnings of the same size. It is well-known that it is virtually impossible to decrease nominal salaries which suggests that it would also be more difficult to decrease real salaries as a result of an excess supply than to increase them

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[1] Labor Market Statistics, The National Labor Market Board (AMS), monthly issues 1969-.

as a result of excess demand. To investigate this hypothesis the effect of the disequilibrium variable was made dependent on the sign of the variable.

It might also be argued that the net effect of the negotiations are not the same in periods of excess demand as in periods of excess supply. If there is an excess supply of engineers there would be no incentives from the market to increase salaries. The net effect of negotiated increases would then be 1. In periods of excess demand the net effect would depend on the relative magnitudes of the negotiated increases and the market induced increase. If the negotiated increase would be higher than the market induced increase, the observed increase would be determined by the negotiated increases. If the market would set higher salaries than negotiated an additional small increase in the negotiated rate of increase would have no effect on the observed salaries. The market induced increase is however the result of competition for labour and a negotiated general increase might not effectively reduce demand but make some employers compete at a higher salary level. The result would be that an increase in the negotiated rate of increase would have some net effect on the observed salaries even if the negotiated increase is relatively low.

These respecifications result in a new model version, Model II:3,

$$\beta_t = \beta_0+\beta_{11}x_{1t}+\beta_{12}x_{2t}+\beta_{13(1)}x_{3t-1(1)}+\beta_{13(2)}x_{3t-1(2)}+$$

$$\beta_{14(1)}x_{4t(1)}+\beta_{14(2)}x_{4t(2)}+\varepsilon_t^{\beta} \qquad (14a)$$

$$\gamma_{bti}=\gamma_0+\gamma_1(t-b)+\gamma_{21}x_{1t}+\gamma_{23(1)}x_{3t-1(1)}+\gamma_{23(2)}x_{3t-1(2)}+$$
$$\gamma_{24(1)}x_{4t(1)}+\gamma_{24(2)}x_{4t(2)}+\varepsilon^{\gamma}_{bti} \tag{14b}$$

where

$$x_{3t(1)} = \begin{array}{ll} x_{3t} & \text{if } x_{3t} < 0 \\ 0 & \text{otherwise} \end{array}$$

$$x_{3t(2)} = \begin{array}{ll} x_{3t} & \text{if } x_{3t} > 0 \\ 0 & \text{otherwise} \end{array}$$

$$x_{4t(1)} = \begin{array}{ll} x_{4t} & \text{if } x_{3t} < 0 \\ 0 & \text{otherwise} \end{array}$$

$$x_{4t(2)} = \begin{array}{ll} x_{4t} & \text{if } x_{3t} > 0 \\ 0 & \text{otherwise} \end{array}$$

The estimates of Model II:3 are given in Table 4. These results conform well with the hypothesis suggested above. The effect of an excess demand is approximately three times as strong as the effect of an excess supply. The net effect of negotiated increases is not significantly different from 1 in periods of excess supply and from 0 in periods of excess demand. Also this model version shows that initial earnings are more sensitive to market changes and the outcome of negotiations than the earnings of those who already have a job. The over all fit of the model is good and the prediction of the cross-sectional profile is very close to the observed profile as shown in Figure 7. These favorable prediction results do not however give the same strong support to the model as predictions for a period outside the sample period.

In order to check the results obtained for electrical engineers Model II:2 and Model II:3 were also fitted to data for engineers specialized in mechanical engineering. We used an index of the output from "Manufacture of Fabricated Metal Products, Machinery and Equipment"[1] as shift variable in the demand function, which is probably not ideal. All other variables were the same. The results from this new data set agree with those obtained for electrical engineers, except that the estimates of the effect of shifts in the demand function - γ_{21} in Model II:2 and β_{11} and γ_{21} in Model II:3 - have the wrong sign. In Model II:3 $\hat{\gamma}_{21}$ is significantly different from zero, which indicates that a better measure of shifts in demand is desirable.

6. CONCLUSIONS

A simple human capital approach which explains age-earnings profiles by a function of years of experience only was extended by time specific effects on the growth of earnings. The consequences of omitting time specific effects were illustrated by applying alternative models to earnings data for engineers. Since we have only analyzed two homogeneous groups of engineers it was possible to relate these time specific effects to a certain market.

Our analysis shows that both market changes and the outcome of central negotiations are important to explain age-earnings profiles. The effect of

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[1] Division 38 according to Swedish Standard Industrial Classification of all Economic Activities (SNI).

these factors are not the same independently of years of experience, which implies that cross-sectional profiles of log-earnings will not be parallel but instead change in shape. One implication of this result is that the return to investments in human capital cannot be estimated from single cross-section.

The single most important explanation to changes in earnings for engineers is the lack of balance in the market, which shows how important it is that recruitment to higher education and the output from colleges and universities balance changes in demand. Earnings are more sensitive to an excess demand situation, which will generate a return to human capital above "normal" return, than to an excess supply situation, which will only have a minor depressing effect on earnings. Contrary to many previous studies we have also shown that the outcome of negotiations has a net effect on increases in earnings and this is so in particular when demand for engineers is low.

Table 1. Estimated earnings profiles for electrical engineers. Sample period: 1961-1970

	Model I		Model II:1		Model II:2	
β_0	0.0107	(0.0010)	-0.0072	(0.0109)	-0.0235	(0.0123)
β_{11}	--		0.0043	(0.0020)	0.0072	(0.0022)
β_{12}	--		-0.0016	(0.0006)	-0.0023	(0.0007)
β_{13}	--		0.0338	(0.0080)	0.0379	(0.0046)
β_{14}	--		--		0.0107	(0.0035)
γ_0	0.0892	(0.0029)	0.0753	(0.0093)	0.0656	(0.0111)
γ_1	-0.0030	(0.0004)	-0.0021	(0.0004)	-0.0025	(0.0005)
γ_{21}	--		0.0017	(0.0016)	0.0025	(0.0017)
γ_{23}	--		0.0148	(0.0066)	0.0261	(0.0050)
γ_{24}	--		--		0.0077	(0.0027)
$\bar{R}^2$	0.8418		0.8536		0.8572	

Table 2. Estimated earnings profiles for electrical engineers. Sample period: 1961-1976

	Model I		Model II:1		Model II:2	
β_0	-0.0036	(0.0005)	0.0283	(0.0109)	0.0086	(0.0079)
β_{11}	--		-0.0012	(0.0013)	0.0023	(0.0014)
β_{12}	--		-0.0020	(0.0003)	-0.0029	(0.0004)
β_{13}	--		0.0289	(0.0024)	0.0269	(0.0025)
β_{14}	--		--		0.0141	(0.0021)
γ_0	0.0817	(0.0019)	.1236	(0.0057)	0.1071	(0.0063)
γ_1	-0.0036	(0.0002)	-0.0033	(0.0002)	-0.0032	(0.0002)
γ_{21}	--		-0.0061	(0.0009)	-0.0038	(0.0010)
γ_{23}	--		0.0202	(0.0019)	0.0166	(0.0020)
γ_{24}	--		--		0.0063	(0.0013)
$\bar{R}^2$	0.8612		0.8917		0.8959	

Table 3. Estimated earnings profiles for electrical engineers and mechanical engineers
Sample period: 1961 - 1976

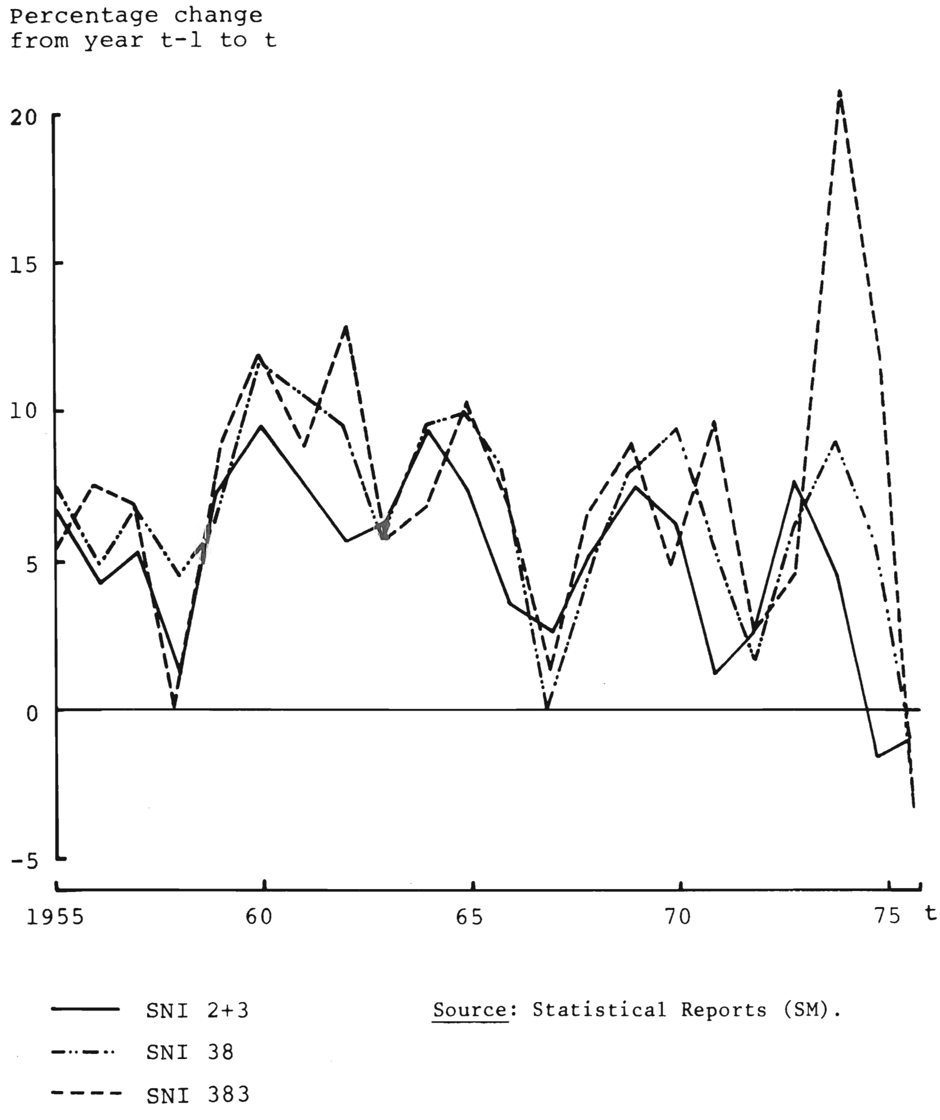
	Electrical engineers				Mechanical engineers	
	Model II:1		Model II:2		Model II:2	
β_0	-0.0170	(0.0056)	-0.0139	(0.0055)	-0.0131	(0.0069)
β_{11}	0.0037	(0.0007)	0.0027	(0.0007)	0.0027	(0.0011)
β_{12}	-0.0009	(0.0003)	-0.0011	(0.0003)	-0.0013	(0.0027)
β_{13}	0.0469	(0.0027)	0.0489	(0.0028)	0.0317	(0.0037)
β_{14}	--		0.0122	(0.0019)	0.0158	(0.0018)
γ_0	0.0772	(0.0037)	0.0792	(0.0037)	0.0904	(0.0052)
γ_1	-0.0033	(0.0002)	-0.0033	(0.0002)	-0.0033	(0.0002)
γ_{21}	0.0017	(0.0005)	0.0078	(0.0050)	-0.0013	(0.0008)
γ_{23}	0.0272	(0.0026)	0.0269	(0.0026)	0.0273	(0.0030)
γ_{24}	--		0.0080	(0.0012)	0.0124	(0.0011)
$\bar{R}^2$	0.8945		0.8982		0.9094	

Note: These estimates are obtained with the output measure for the entire industry replaced by an index for sub-industries (SNI 383 for electrical engineers and SNI 38 for mechanical engineers).

Table 4. Estimated earnings profiles for electrical engineers. Sample period: 1961 - 1976

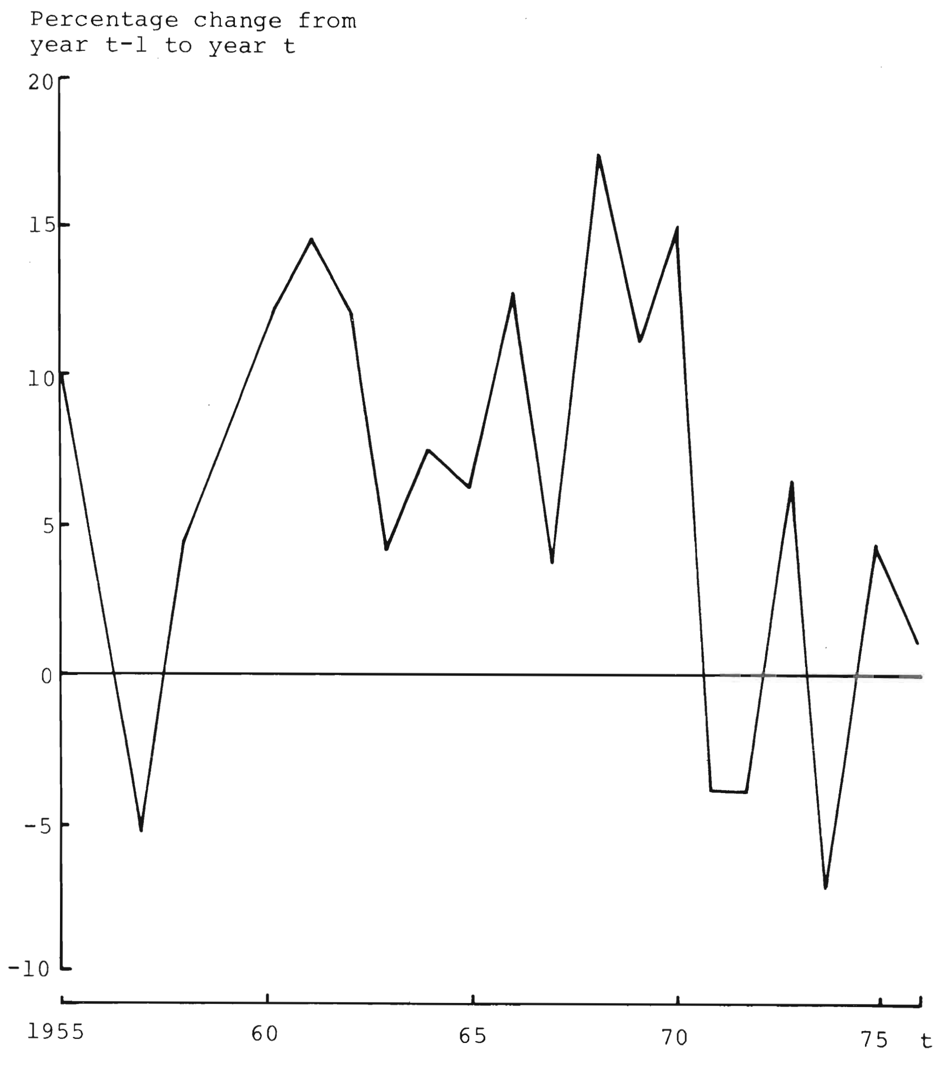
Model II:3					
β_0	-0.0192	(0.0065)	γ_1	-0.0033	(0.0002)
β_{11}	0.0014	(0.0011)	γ_{21}	0.0002	(0.0009)
β_{12}	-0.0008	(0.0004)	$\gamma_{23(1)}$	0.0166	(0.0148)
$\beta_{13(1)}$	0.0226	(0.0180)	$\gamma_{23(2)}$	0.0454	(0.0252)
$\beta_{13(2)}$	0.0827	(0.0273)	$\gamma_{24(1)}$	0.0089	(0.0017)
$\beta_{14(1)}$	0.0130	(0.0024)	$\gamma_{24(2)}$	0.0028	(0.0075)
$\beta_{14(2)}$	0.0050	(0.0081)	$\bar{R}^2$	0.8983	
γ_0	0.0782	(0.0047)			

Figure 1. Relative changes in the volume of industrial production



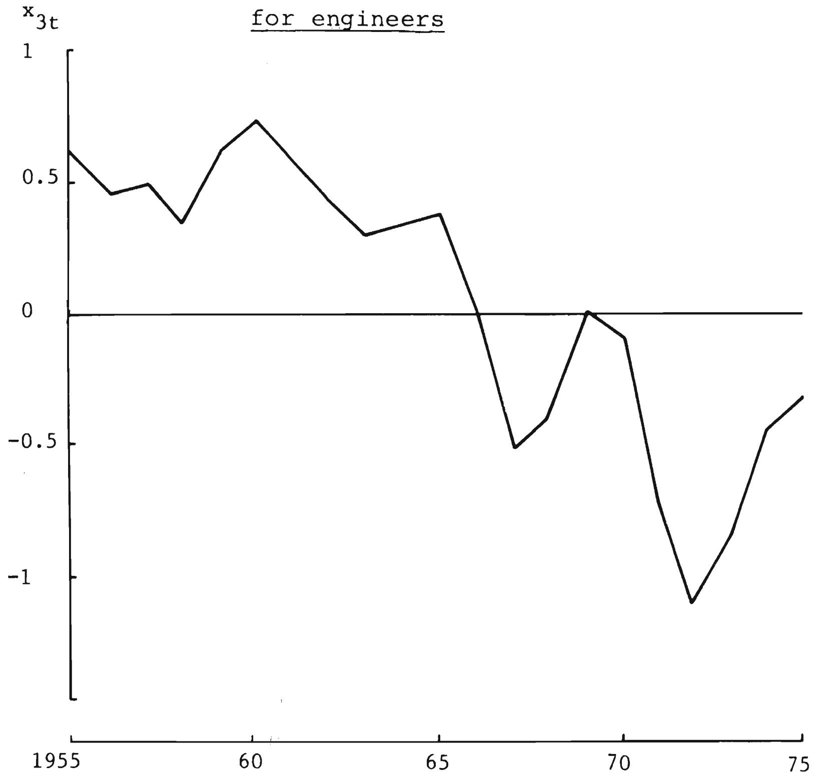
Source: Statistical Reports (SM).

Figure 2. Relative changes in the number of new graduated engineers



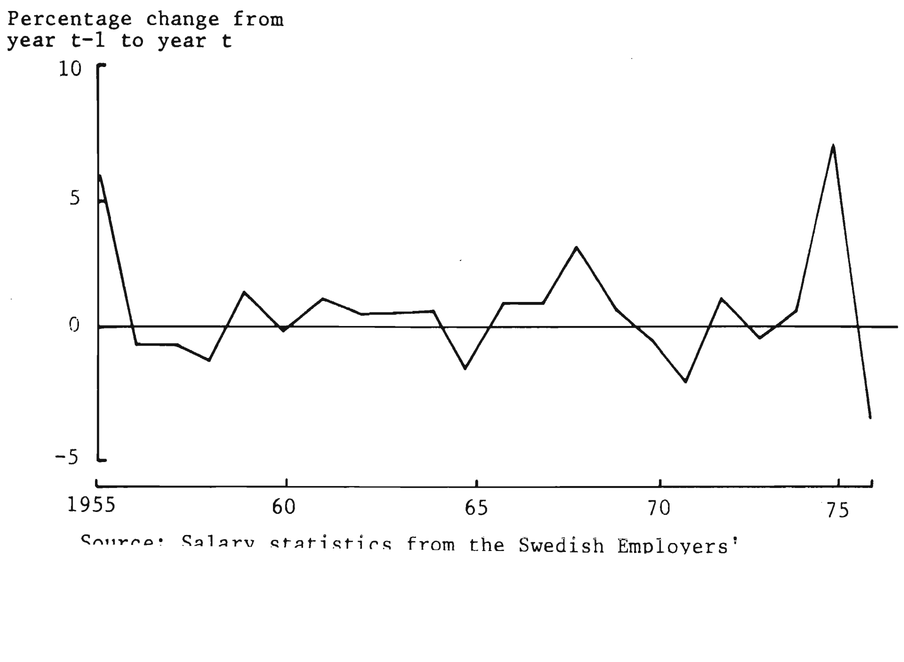
Source: Statistical Abstract of Sweden (SÅ).

Figure 3. The disequilibrium in the market for engineers



Source: See Section 2 for the definition of the disequilibrium measure (x_3) and sources.

Figure 4. Negotiated salary increase in constant prices for male employees



Source: Salary statistics from the Swedish Employers'

Figure 5. Observed and predicted earnings profiles for electrical engineers in 1973

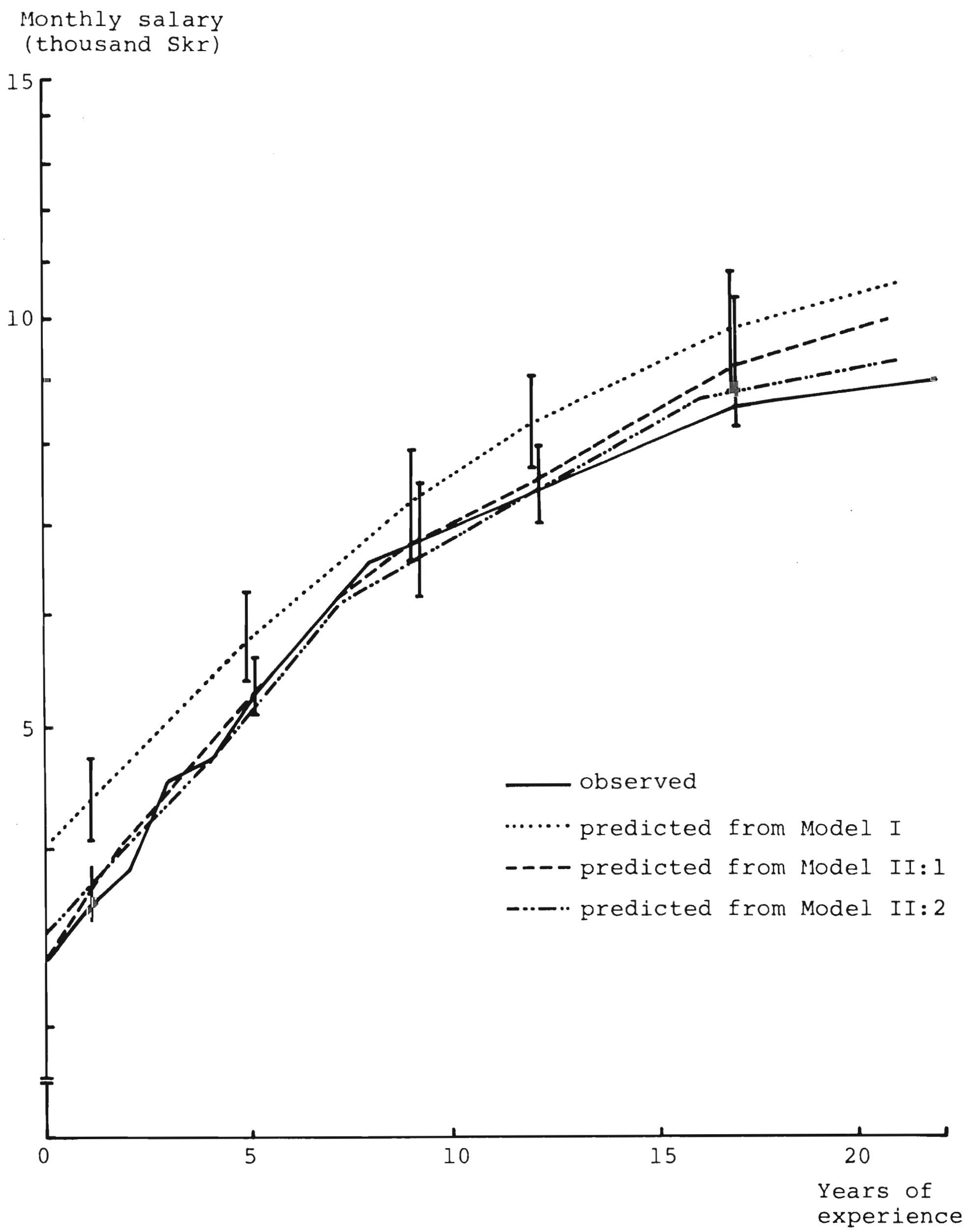


Figure 6. Observed and predicted earnings profiles for electrical engineers in 1976

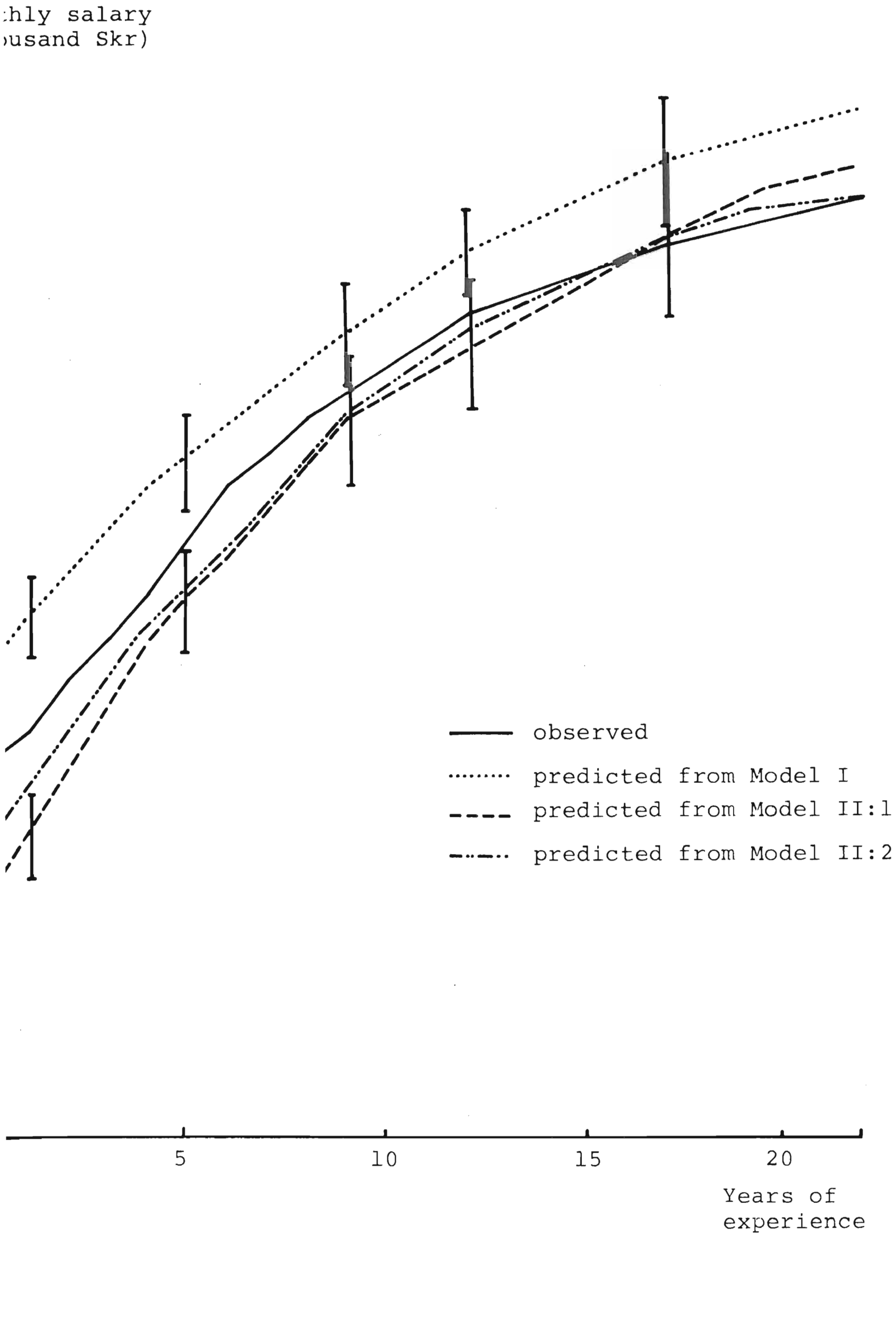
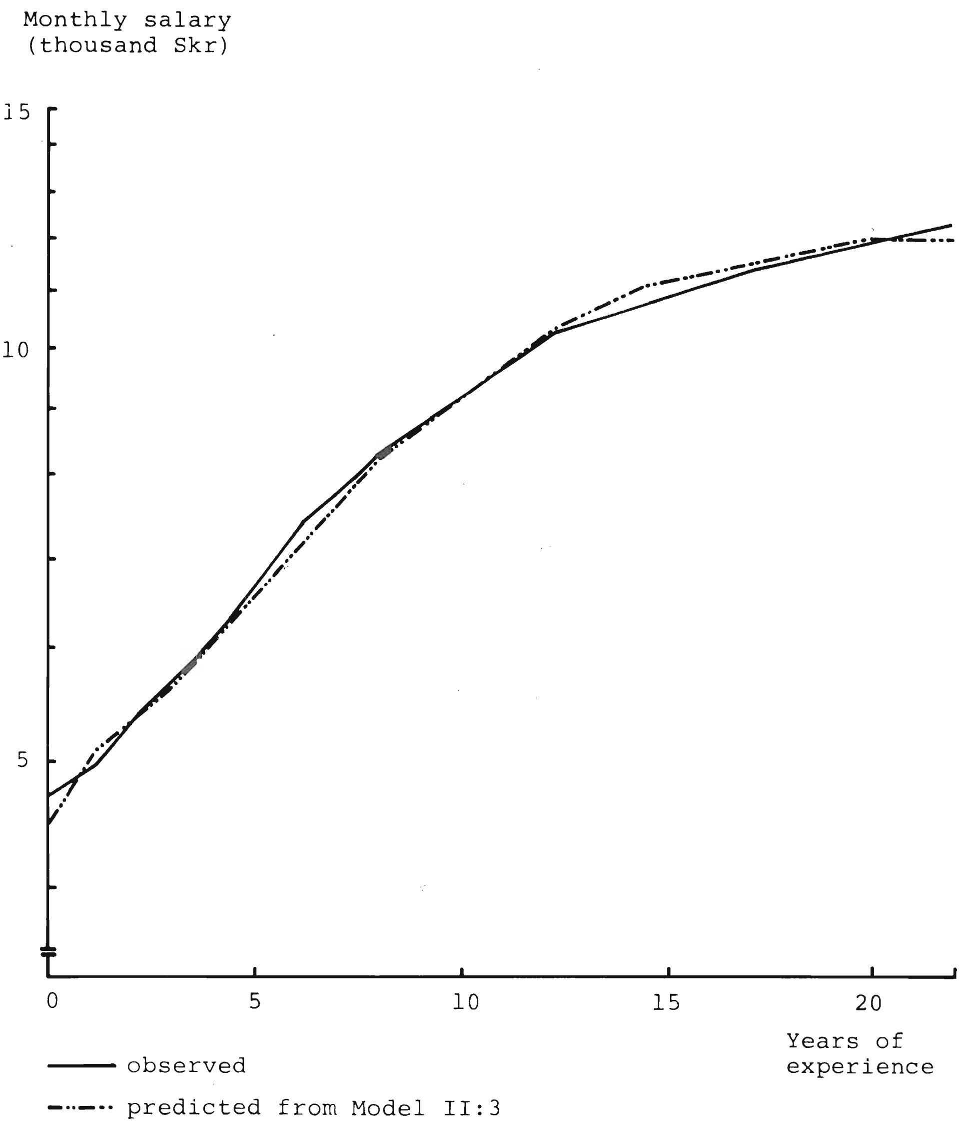


Figure 7. Observed and predicted earnings profiles for electrical engineers in 1976



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The Duration of Vacancies as a Measure of the State of Demand in the Labor Market. The Swedish Wage Drift Equation Reconsidered

Nils Henrik Schager

INTRODUCTION*

During the last twenty years a multitude of empirical studies has demonstrated the close relationship between the state of demand in the labor market and the rate of money wage increases in various countries. Following the pattern of A.W. Phillips' celebrated study most researchers have used the unemployment ratio as the indicator of

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the labor market situation.[1] Others have on theoretical grounds preferred to work with a combination of unemployment and vacancy data, where such figures have been available. While the relationship seemed to be stable up to the second half of the sixties, the wage equation based on these variables has not performed well during the last decade. The estimated wage increases have fallen short of the actual ones in most countries.

Sweden is no exception in this respect. Traditionally the Swedish "Phillips curve" includes both unemployment and vacancies (or a combination of them) as explanatory variables. Before 1969 the post-war data behaved in accordance with the original Phillips hypothesis. There was little evidence as to the question whether the number of unemployed alone, the number of unfilled vacancies alone or the difference between them contained the best explanatory power. However, from 1969 onwards the stability of the unemployment-based wage increase equation disappeared. The same is true for the vacancy-based equation as regards negotiated wage increases. Unfilled vacancies continued to show a close correlation with wage drift up to 1973.[2] Since then even this remnant of a "Phillips relation" vanished.

This paper is an attempt to restore the stability of the Swedish wage drift equation. My approach follows the mainstream of economic thinking on these matters in so far as the state of demand in

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[1] See Phillips (1958) and - for a survey of the later research - Trevithick-Mulvey (1975).

[2] See Calmfors-Lundberg (1974) and Holmlund (1976).

the labor market is supposed to exert a dominant influence on the amount of wage drift. However, I will argue that the traditional ways of measuring that state of demand suffer from serious shortcomings. Instead of the stocks of vacancies and of unemployed I will put forward the duration of vacancies as a preferable alternative. Such a duration measure can be constructed by combining figures on stocks and flows of vacancies.

Intuitively it should be obvious that for an individual firm the duration of its vacancies is a very tangible sign of the situation in the labor market. If it changes, there should be a need for a change in the firm's recruitment policy, especially in its wage policy. Nevertheless, at least to my knowledge there has been no study attempting to explain the variations in wage increases by relating them to the duration of vacancies. In other words, the literature seems to look at the unemployment and/or vacancy rates as the "true" indicators of the labor market situation, that is relevant to the wage formation process. The movements of the duration measures are regarded more as selfevident statistical phenomena without any real independent significance in this respect, although their covariation with the corresponding stock entities has been recognized in some contributions on the Phillips curve issue.[1]

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[1] See Holt-Martin (1966) and Hansen (1970) for a theoretical discussion on this point and Holmlund (1976) for an empirical investigation, based on Swedish data. The closest parallel to my approach should be those studies, following the approach applied in an article by Holt in Phelps (1970), where the 'reservation' wage of an unemployed job applicant is explicitly stated as a declining function of the length of the duration of unemployment.

The approach of this paper it the opposite one. I will argue that the time which is required for a firm to fill its vacancies is crucial for its decision on which wage policy to follow. It will be demonstrated that the duration of vacancies reflects the basic characteristics of the search activity in the labor market, when regarded as a stochastic process. This analysis strengthens the theoretical support for my hypothesis, which I will test by relating an aggregate measure of the duration of vacancies for the Swedish labor market as a whole to the amount of wage drift for workers employed in the private sector.

Before proceeding to the analysis it will be useful to shortly review earlier studies of wage drift behavior in Sweden.

EARLIER RESEARCH ON WAGE DRIFT IN SWEDEN

A path-breaking study on wage drift was published as early as in 1956 by Bent Hansen and Gösta Rehn.[1] They developed in a rigorous fashion a model for a labor market in disequilibrium within the framework of Hansen's earlier theoretical study on inflation.[2] Partly on the basis of this model, partly with the help of more ad-hoc reasoning, the authors postulated that the amount of wage drift was to be expected to depend on the state of demand in the labor market, on the profit margins and on the rate of increase in physical productivity within industry.

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[1] Hansen-Rehn (1956).

[2] Hansen (1951).

This hypothesis was tested on time-series data for eight branches of industry embracing the years 1947-1954. It was clearly established that the labor market variable - defined as the difference between vacancies and unemployment - was highly significant in almost all branches of industry. On the other hand the influence of profits was weak and that of productivity almost non-existent.

In the OECD-report "The Problem of Rising Prices", of which Bent Hansen was co-author, the study of Hansen and Rehn was followed up.[1] When wage drift within the manufacturing and mining sector as a whole was related to the number of unfilled vacancies (no data on unemployment were used!) the result was very convincing. A simple regression on data covering the period 1950-1960 yielded a coefficient of determination of 0.83.

The Hansen-Rehn study was replicated by Lars Jacobsson and Assar Lindbeck more than a decade later.[2] The approaches of the two studies were fairly similar, but Jacobsson and Lindbeck also made an attempt to explain the amount of centrally negotiated wage increases. Their wage data referred to workers in manufacturing as a whole.

Jacobsson and Lindbeck used the same theoretical explanatory variables as Hansen and Rehn. The state of demand in the labor market, profits and productivity increases were supposed to influence the amount of wage drift (as well as negotiated wage increases). The authors used alternative measures when identifying the demand situation in the

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[1] Marris(-Hansen) (1964).

[2] Jacobsson-Lindbeck (1969).

labor market: the unemployment figures according to the unemployment insurance statistics, the unfilled vacancy figures recorded at the local employment offices and the ratio between them.

The estimation period was 1955-1965, which gave eleven yearly observations on wage drift. When ordinary least squares multiple regression was carried out on these data, the results were unequivocal. The labor market variable (on any definition) was highly significant (on 1% level), while the remaining independent variables were insignificant (on 5% level). When simple regressions were run, unfilled vacancies gave the best goodness of fit ($R^2 = 0.86$).

As was mentioned Jacobsson and Lindbeck also tried to explain the amount of negotiated wage increases. Even in this case only the labor market situation could be found to exert a significant influence on wage development. The authors used the opportunity to study whether any systematic covariation between negotiated wage increases and wage drift could be traced. It is sometimes maintained that there exists an inverse relationship between these two forms of wage increases. When negotiated increases are "too low", this would lead to higher wage drift in subsequent periods. However, Jacobsson and Lindbeck could not find any pattern in their estimates substantiating such an hypothesis.[1]

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[1] Strictly speaking, Jacobsson-Lindbeck's "testing procedure" does not exclude the possibility that a centrally negotiated increase which is so low as to improve the competitiveness of Swedish industry would at a later stage lead to stronger labor demand and hence to higher wage drift.

They gave an interesting theoretical explanation of this lack of (negative) correlation. Wage drift, they argued, reflects to a high degree competition for labor between employers and in that case it is the deviation of levels of earnings from the average which is important. Negotiated wage increases affect (mainly) the average but should leave the competitive position of an individual employer unaltered, thus giving him no incentive to wage drift regardless of the size of the negotiated increase.

A few years later, Jacobsson and Lindbeck published another study within the area of wage formation.[1] From our point of view the most interesting feature of this second study is that the estimates on wage drift in manufacturing now include data from the second half of the sixties, a period when the earlier observable close correlation between movements in vacancy and unemployment figures ceased to exist.

The period of observation now ranges over 1950-1970. In a simple regression the coefficient of the labor market variable, here defined as unfilled vacancies minus unemployment, is again highly significant (on 1% level). Jacobsson and Lindbeck also ran a multiple regression with both vacancies and unemployment as explanatory variables. In this estimate the vacancy variable turned out to contain all explanatory power. It performed even better than the composite measure of vacancies minus unemployment.

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[1] Jacobsson-Lindbeck (1971).

As a conclusion we are able to maintain that wage drift for workers in Swedish manufacturing during the fifties and sixties was to a high degree influenced by the state of demand in the labor market. There is also evidence that the rate of unfilled vacancies for the labor market as a whole was the most reliable indicator of the labor market situation in this context. No other factor has been shown to have a significant effect on wage drift.

We will now turn to the wage development during the seventies, a period of more violent fluctuations both in the Swedish and in the world economy. We know from foreign studies that in many countries the Phillips relation has dissolved during this period. The same development seems to have occurred in Sweden.

In their comprehensive book on inflation and unemployment Lars Calmfors and Erik Lundberg noticed that the pure Phillips relation - in which unemployment is the only explanatory variable - lost its former stability around 1969.[1] During the period 1969-1973 the amount of wage drift lies on a considerably higher level compared to that which the unemployment figures would predict. When the measure vacancies minus unemployment is used, the upward shift is still clearly visible but is postponed to 1970. However, when unfilled vacancies alone are used as the explanatory variable there is no evidence of a shift in the wage drift equation during the whole period 1956 - 73.

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[1] Calmfors-Lundberg (1974).

Calmfors and Lundberg made no explicit comment on the fact that the number of unfilled vacancies seems to 'explain' wage drift much better than the alternative traditional measures of the labor market situation. This point was discussed, however, by Bertil Holmlund in his doctoral dissertation.[1] There he demonstrated the existence of a shift around 1969 in the relationship between the stock of unemployed and the stock of vacancies. This phenomenon he attributed to an increase in the average duration of unemployment, which in turn he interpreted as a sign of decreasing adaptability in the labor market.

Holmlund confirmed that the number of unfilled vacancies continued to show a high degree of correlation with the amount of wage drift for the period 1963-1973 (a simple regression yielded an R^2 of 0.87) while the unemployment figures failed to do so. From this outcome Holmlund drew the conclusion "that changes on the demand side of the labor market are more inflationary than changes on the supply side".

Holmlund's interpretation of his wage drift estimation results can be questioned and the unfilled vacancy variable is certainly not able to accomplish its task in explaining wage drift in the post-1973 period, as this paper will show. However, according to the approach used in the present study Holmlund touches upon some very crucial matters when he analyzes and interprets extensively the relationships between stocks, flows and durations of vacancies and unemployment.

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[1] Holmlund (1976).

THE LABOR MARKET VARIABLE IN THEORY

On the basis of established economic theory it is evident that there are good reasons to expect the state of demand in the labor market to exert a strong influence on the amount of wage increases. In theoretical discussions on dynamic wage behavior there seems to be a substantial agreement that a measure of the difference between vacancies and unemployment is the most satisfactory indicator of the state of demand in the labor market. However, even within such a purely theoretical setting the notions of vacancies and unemployment are not always clear. It has been pointed out that according to neoclassical theory vacancies and unemployment cannot exist simultaneously in one market, if jobs and job applicants are considered as homogeneous. Such a statement obviously refers to vacancies and unemployment as stock concepts. Flows of vacant jobs and of job applicants meet in the market and match immediately (or at least within the "transaction period") according to the neoclassical assumptions. As a result there should remain <u>either</u> unsatisfied labor demand (unfilled vacancies) <u>or</u> unsatisfied labor suppliers (unemployed) whenever the market does not clear. So the degree of disequilibrium might be described equivalently as <u>the difference between the flows</u> of vacancies and job applicants or as <u>the stock</u> of vacancies/unemployed (whichever is relevant). Either of these two stock entities is to be interpreted as the outcome of the interaction between demand and supply in the labor market. Theoretically they contain the same information. Only the flow entities can be said to reflect just one side of the market.

The observed simultaneous existence of stocks of vacancies and of unemployed causes considerable difficulties within a neo-classical framework. As long as the assumption of one homogeneous market, in which agents possess complete information, is upheld such a phenomenon does not make sense. As a consequence one has to postulate that the total vacancy and unemployment figures are the result of an aggregation of several submarkets, in which excess demand and excess supply prevail alternatively. While such a disaggregated approach may have some merits on its own, it demonstrates painfully clear how difficult it is to maintain the hypothesis of a stable aggregate wage increase equation. Several restrictive assumptions have to be made in order to derive, e.g., a stable Phillips relation.[1]

Most empirical research on the relationship between wage increases and the state of demand in the labor market has been based on notions from neoclassical theory. This is presumably the reason why stocks of vacancies and of unemployed have been regarded as containing sufficient information on the state of demand and supply. The necessary assumption of heterogeneities in the labor market motivates the use of a vacancies-minus-unemployment measure in the context. Recently, however, a new theory of the labor market has emerged, which emphasizes that the matching of vacant jobs and job applicants is a costly and time-consuming process.

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[1] See Hansen (1970) for a comprehensive analysis in the neoclassical tradition.

In a given period of time a certain amount of new vacancies are "thrown into the market" as a sign of temporarily unsatisfied demand for labor. In the market these vacancies meet job applicants (which may or may not be officially unemployed) and after a search process some of these vacancies are filled.

If the period is not too long (in comparison with the time required for searching) there will always remain some unfilled vacancies as well as some "unsatisfied" job applicants at the end of the period. This result holds even if we assume homogeneity in jobs and applicants in the market. Thus one of the merits of the modern labor market theory is that it reconciles the analytical implications of an aggregate model with the fact that unfilled vacancies and unemployment are indeed observed to exist at the same time.

However, the simple link between flow and stock concepts in a neoclassical model is not preserved. We might think of two situations where the flows of vacant jobs and of job applicants during a certain period are equal. In one case job-matching turns out to be very "efficient" and only small numbers of unfilled vacancies and unsatisfied applicants remain at the end of the period. In the other case large proportions of the flows remain as stocks, when the period expires.

Should we measure the degree of disequilibrium in the labor market as the difference between the two flows? That alternative cannot be appropriate, as it would imply that it does not matter what the agents experience in the market. Should we use the difference between the two stocks? Perhaps, but the crucial question is whether we should rely on stock concepts at all, not how the stock figures should be combined, because search models do predict that the flows and stocks interact in such a way as to establish a one-to-one correspondence between the stock of vacancies and the stock of unsatisfied applicants. In such a case it does not matter whether the degree of disequilibrium is measured by the one stock or the other or by any combination of them.[1]

As is nowadays recognized, there exists no reliable way to measure the number of job applicants. The unemployment figures have ceased to behave as a good indicator for a number of reasons.[2] However, the foregoing discussion has shown that as long as we have faith in aggregate models at all search theory models predict that stocks of vacancies contain the same information on the state of the labor market as stocks of job applicants. So it is not to deviate from established disequilibrium analysis to concentrate on the behavior of vacancies only, although it is certainly not in

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[1] Obviously such a conclusion does not hold if we apply the model to observations, where the flows and stocks of unemployed are substituted for those of job-applicants, a procedure used both in theoretical analysis and in empirical applications. Swedish and U.K. data contradict the hypothesis of a stable relationship over time between the number of unfilled vacancies and the stock of unemployed. (See Holmlund, 1976 and Trevithick-Mulvey, 1975).

[2] The reasons why the unemployment figures have lost their capacity of reflecting the availability of labor fall outside the main line of this paper. For such a discussion, see, e.g., the references in the preceding footnote as well as Taylor (1974), where some correction procedures are suggested.

accordance with established practice in empirical research.[1]

There is another reason why an analysis of wage increases based on the behavior of vacancies should be more adequate. Several years ago neo-classical disequilibrium theory was criticized because it failed to answer the basic question: who increases prices?[2] The modern theories of price adjustment in an environment of incomplete information do give an answer to that question. More specifically, search models of the labor market postulate that firms act as temporary monopsonists and set wages so as to regulate the time-paths of new hires (and of quits) in an optimal way.[3] What

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[1] The reliability of existing vacancy statistics can certainly not be taken for granted either. This issue will be discussed in the next section. But the general argument is that the vacancy figures, not being a "policy target" have run less risk of being distorted than the unemployment ones.

[2] Arrow (1959).

[3] Cf. the models presented in Phelps, ed. (1970), Siven (1979), Pissarides (1976). As should be clear, like these models we are here concerned with decentralized wage formation in the absence of union bargaining. It may be argued that the choice of a vacancy-based variable is accurate if one regards wage drift only as the outcome of a reaction from the employer to his recruitment situation. However, wage drift may also contain elements of labor action, e.g. collective actions at plant level. As to this question I am basically of the same opinion as was expressed by Jacobsson and Lindbeck that to a dominant extent wage drift reflects the competition for labour between employers. This may not hold in exceptional circumstances when the functional distribution of income is strongly unbalanced and to this qualification I will return at the end of this paper.

the firms observe is the rate at which their vacancies are filled. This process is certainly influenced by supply conditions in the labor market, but what happens to job applicants (or the unemployed) is not directly relevant to the wage setting agents.

Let us, however, return to the analysis of stocks and flows and their interaction. We can now reformulate our earlier question as 'does the stock of vacancies accurately indicate the degree of disequilibrium in the labor market?' It will simplify the discussion if we write down the following relationship:

$$v \cdot T = V \tag{I}$$

where v denotes the flow of vacancies during the period, T the duration of vacancies and V the stock of vacancies at the end of the period. Eq. (I) applies strictly only to a steady state situation, but the general form will soon be presented and the argument will not be affected by this simplification.

Equation (I) epitomizes the firms' recruitment process. The flow of new vacancies (v) represents the demand for labor. This flow interacts in a search process with the flow of job applicants, which represents the supply side of the labor market. The outcome of the search is reflected in T, the value of which shows how efficient the matching of vacant jobs and job applicants is. The stock of unfilled vacancies (V) is then the combined (multiplicative) effect of the volume of labor demand and the ability of the labor market to satisfy that demand. Thus it can be maintained that the unfilled vacancy figures do reflect not only the demand for labor but also the availability of labor to match that demand. This is certainly an appealing explanation of the apparent success in using unfilled vacancies alone in the wage drift equation.[1]

However, it is equally obvious that the stock of unfilled vacancies alone cannot tell the whole story. When the demand for labor (i.e., the flow of new vacancies) changes, the number of unfilled vacancies changes in the same direction, too, even if the labor market is equally apt to respond to this change, i.e., the average duration period is unchanged. To the extent that variations in the unfilled vacancy figures reflect pure demand changes they should not, according to my view, be interpreted as a sign of increased or decreased degree of disequilibrium in the labor market.

However, a case can be made for attributing this property to the duration measure T. When the demand for labor increases and new vacancies are opened up there will be an increase in new hires as well as in unfilled vacancies. Rational behavior on the part of a recruiting firm requires it to work with quantity signals alone (i.e., vacancies) as long as this is sufficient for hiring new labor and to postpone the costly price signals

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[1] So this success is not to be interpreted as if "demand factors" influence wage increases more heavily than "supply factors". However, it may be regarded as a support for the hypothesis that wage drift is initiated from the demand side of the labor market.

(i.e., wage increases) until the quantity signals fail to take effect.

When do they fail? Is it when the number of unfilled vacancies increases, just because the number of new vacancies increases, although the duration of vacancies is unaltered and the recruitment process is going on at the same speed as before? Or is it when the firm realizes that it has to wait for a longer time until a qualified applicant turns up and accepts the job? I would maintain that firms are more sensitive to the second situation, in which they face a real recruitment problem and might well contemplate speeding up the hiring process by raising wages.[1] In other words, I claim that the duration of vacancies is a more reliable measure of the degree of disequilibrium in the labor market.

We will now complete the general argument with a more rigorous analysis of the firms' recruitment process.[2] In the first place, this will offer us a possibility of finding alternative ways to interpret the concept of duration of vacancies in the context of a stochastic search process. Secondly, in the absence of any published figures on the length of duration itself we are bound to derive a

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[1] The same conclusion can be found in Pissarides (1976): "In our model of optimal firm behavior ... the number of periods the vacancy has been standing idle is an important factor in the determination of the wage offer, in that the 'reasonable' firm will be increasing its offer over time as workers search it and walk out again without taking the job." (p.202).

[2] Those uninterested in formal mathematical analysis may without loss of continuity leave out the remainder of the present section.

duration measure from the flow and stock figures. For that purpose the steady-state assumptions behind eq.(I) are not satisfactory; we must find a relation that can apply to short-term changes.

The last problem may be tackled by simply postulating that during a certain period every vacancy has the same probability of being filled at each point of time, i.e., every vacancy does have the same constant "deathrisk". If we denote the "death intensity" as Θ, it is a well known result that the probability function of the completed durations of the variable is

$$f(t) = \Theta \cdot e^{-\Theta t}$$

and that consequently the expected value of the duration is[1]

$$T = \int_0^\infty t \cdot \Theta \cdot e^{-\Theta t} dt = \frac{1}{\Theta} \tag{II}$$

Suppose now that for a certain time interval (t_0, t_1) we are able to observe the stock of unfilled vacancies at t_0 and t_1, respectively; we denote these variables V_0 and V_1. We also observe the total flow of new vacancies during (t_0, t_1) and denote it $v_1(t_1-t_0)$ (by using this notation we assume that the flow is uniformly distributed over (t_0, t_1)). The postulated stochastic process implies that every vacancy will have a "survivor function" $EXP(-\Theta_1 \cdot s)$, where s denotes the relevant

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[1] The results can be found in any comprehensive textbook on probability theory. I have consulted Blom (1970).

time interval. We can then formulate the following relationship:

$$V_0 \cdot e^{-\Theta_1(t_1-t_0)} + \int_{t_0}^{t_1} \cdot v_1 \cdot e^{-\Theta_1(t_1-t)} \, dt = V_1$$

or (III)

$$V_0 \cdot e^{-\Theta_1(t_1-t_0)} + \frac{v_1(t_1-t_0)}{\Theta_1(t_1-t_0)} \cdot \left(1-e^{-\Theta_1(t_1-t_0)}\right) = V_1$$

Eq.(III) can be solved for $\Theta(t_1-t_0)$, and by eq.(II) we directly obtain a value of T when the length of the time interval (t_0,t_1) is given.[1]

It is easily seen that when eqs. (III) and (II) are applied to a steady-state situation, where $V_0=V_1=V$, they reduce to eq.(I).

By the described procedure we are thus able to calculate a measure of the duration of vacancies, which can be interpreted as the inverse of the instantaneous probability that a vacancy should be filled.[2]

We may, however, go somewhat further than merely postulating that the durations of vacancies follow

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[1] An explicit analytical solution to eq.(III) is not easily found, but a numerical one can be approximated with any degree of precision.

[2] Given the exponential distribution of durations, it is not only the expected value T which shows this type of simple relation to the parameter $\Theta(t_1-t_0)$. The time it takes to fill any fractile of the vacancies is in fact T multiplied with a constant.

an exponential distribution. This result can be derived from more basic assumptions concerning the nature of the search process in the labor market.

Consider a firm which announces its need for more labor by opening up new vacancies. It is searched by prospective employees, who turn up according to a stochastic pattern. It has been demonstrated that this kind of search can well be envisaged as a Poisson process,[1] i.e., the number of contacts (j) within a time interval (t_0, t_1) between a firm and job-applicants follows a Poisson distribution

$$P(j) = \frac{[\lambda_1(t_1-t_0)]^j}{j!} \cdot e^{-\lambda_1(t_1-t_0)} \qquad j=0,1,2...$$

where $\lambda_1(t_1-t_0)$ is the parameter of the distribution.[2]

If each contact during the interval has a constant probability ρ_1 of leading to a hire, the number of hires will follow a distribution that is a combination of the Poisson and the binomial distribution. This compounded distribution is itself a Poisson distribution, the parameter of which is $\rho_1 \cdot \lambda_1 \cdot (t_1-t_0)$.[3]

The recruitment process may thus be envisaged as a Poisson process. It is well-known that the distribution of time intervals until the first event

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[1] See Siven (1979) and examples given in Blom (1970).

[2] For an analysis of the characteristics of the Poisson distribution, see Blom (1970) and Moran (1968).

[3] See Moran (1968), p.85.

happens is exponential, if the number of events follows a Poisson distribution.[1] In this case the probability function of the duration variable would be

$$f(t) = \rho_1 \cdot \lambda_1 \cdot e^{-\rho_1 \cdot \lambda_1 t}$$

which is nothing but the distribution of the vacancy durations.

Our parameter Θ can thus be interpreted as the product of λ and ρ, which parameters reflect the crucial characteristics of a firm's recruitment situation: the possibility to attract a qualified job applicant's attention and - at a second stage - to make the job offer attractive enough. The success or failure in this respect is directly reflected in our duration measure T. The argument for using it as explanatory variable in a wage increase equation is certainly not weakened by this interpretation.[2]

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[1] See Blom (1970), p.9/15.

[2] The fact that the exponential distribution of vacancy durations can be derived from the assumption that hiring occurs according to a Poisson process is of course no proof that the 'real' durations of vacancies have such a distribution. However, the exact time-paths of vacancies have, according to my knowledge, never been analyzed. Bartholomew (1973) makes an extensive review of applications of stochastic models to labor market processes and cites results which indicate that the assumption of 'time independence' underlying the Poisson process is often too restrictive. However, it is not evident that the same objections apply to vacancies. Be it as it may, as long as we lack data on the time-paths of vacancies, no further progress can be made on this issue.

THE LABOR MARKET VARIABLE IN PRACTICE

The vacancy data that we shall use are the monthly figures recorded at the public employment offices. The wage drift data, to which they are to be related, cover workers in the private sector and it would of course be preferable if we could use vacancy data of the same coverage. It is not quite impossible to make such adjustments as would approximate this end, making use of an existing classification by occupation system ("Nordisk yrkesklassificering, NYK") but in this study we will simply use vacancy data for the Swedish labor market as a whole.

Our primary interest is to consider the reliability of the recorded monthly flows and stocks of vacancies only to the extent that is relevant for a judgement of the transformed measure, the duration of vacancies. Being a "pure number", it is clear that the duration of vacancies should not be influenced by changes in the size of employment, while the absolute number of vacancies ought to be affected. However, there are other more important sources of biases. Our period of estimation extends from 1965/66 to 1978/79 (second quarter), during which there has been a change in legislation, requiring employers to report their vacancies to the employment offices. These new regulations have been introduced gradually in different regions, beginning in November 1976. This change affects our three last observations, which justifies that we run regressions with these observations excluded as an alternative.

Beyond doubt the introduction of compulsory vacancy reporting must have increased the number of recorded vacancies, *ceteris paribus*, as it is evident that many employers did not earlier use the service of the employment offices for announcing vacancies.[1] Whether and to what extent the reporting of all vacancies affects our duration measure cannot be established; certainly we are not entitled to maintain that it has remained unchanged. However, we might guess that it was the more easily recruited vacant jobs that were not earlier reported, so that the introduction of compulsory reporting should lower our duration measure, *ceteris paribus*.

On the other hand, the proportion of vacancies that refer to salaried job positions ought to have increased which affects the duration measure in the opposite direction. The total effect is uncertain.

Quite apart from the problem caused by the structural shift in recorded vacancies from 1976 and onwards, we have to consider whether the existence of underestimation in relation to the "true" number of vacancies, i.e., the demand for new labor, might seriously affect the reliability of our labor market measure. It cannot be taken for granted that the number of recorded and "true" vacancies move *pari passu*. For example, there seems to be a tendency for employers to make use of the employment offices to a larger extent when the labor market becomes tighter. This means that the long durations of vacancies are calculated on a broader basis than the short ones and that according to the hypothesis mentioned above vacancies situated at the "short" end of the duration scale are included in the record, which would give

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[1] See Wadensjö (1977).

the duration measure a downward bias (although it has generally become longer).

There is one statistical deficiency in the series on unfilled vacancies, which may also have adverse effects on our duration measure. It is obvious that the filling of announced vacancies is not always adequately followed up by the employment offices or reported on time by the employers.[1] This means that the amount of unfilled vacancies will be artificially high. Such a phenomenon is bound to introduce an upward bias in the duration measure. This situation is most likely to arise when demand for labor and the number of announced vacancies are high (compared with the resources of the labor market offices).

So far we may conclude that there are deficiencies in our duration measure, when calculated on the basis of recorded vacancies. The measure is likely to show an upward bias in periods of strong demand for labor, because the last-mentioned 'delay' effect should be expected to dominate over the first-mentioned 'broader-base' effect. However, the traditional measure of unfilled vacancies is at least as much affected in the same direction without any offsetting elements.

However, there is one change in the statistical series on recorded vacancies during our estimation period, which can be supposed to influence our duration measure but leave the number of unfilled vacancies unaffected. Before 1968 the employment offices kept a record of new hires and it did often happen that when a job applicant, registered at the local office, was hired by a firm this was

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[1] See Wadensjö (1977).

not only recorded as a new hire but also as a new announced vacancy, although it had never been reported as such by the firm.[1] This practice introduces a downward bias in our duration measure before 1968, because the vacancies, created in this way are formally treated as immediately filled!

The effect of such a statistical artifice cannot be established with any precision. No study has been made on the number of vacancies created in the described way. Fortunately, the number of observations that are affected is small and we will just keep the possibility of a bias in mind, when analyzing the regression results.

THE WAGE DRIFT DATA

As has been indicated this study is limited to the behavior of wage drift. This concept is defined as the difference between the calculated effects of the central wage agreements on average earnings and the actual increase in earnings for a specified period. It is important to note that the central negotiating parties of the private sector of the labor market in Sweden (the Swedish Employers' Confederation, SAF, and - for workers - the Swedish Trade Union Congress, LO) are fully informed of the development in earnings within their area, thanks to an elaborate system for collecting earnings statistics. On the basis of the criteria set down in the agreements and subsequent earnings statistics it is possible to see how earnings deviate from the negotiated wage "kitties" not only at national but also at branch and

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[1] Ibid.

plant level. Although there are certainly problems of measurement involved, the wage drift concept in Sweden has thus an accurate and well defined meaning, especially when measured at an aggregate level.[1]

Traditionally the earnings statistics of SAF and LO (covering about 800 000 workers) are gathered on the most comprehensive basis for the second quarter each year. This means that the wage drift figures represent the increase between two consecutive years in average earnings during the second quarter minus the calculated effects of negotiated wage increases, usually the ones taking effect in the beginning of the second year.

Earnings and hence wage drift can be calculated on the basis of different wage form concepts; time rates ("tidlön"), piece rates ("ackord"), including or excluding holiday payment and shift work payment respectively. Which of these concepts to use is not self-evident. The problem is related to that of the effects of structural shifts not only in the composition of wage forms but also in the distribution of sex, occupation, age and employment between firms and industries. Hansen and Rehn looked into these matters in their study and argued that structural shifts were not of decisive importance when measuring the amount of wage drift on an aggregate level.[2] It should be added that it

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[1] For a discussion on possible errors in measuring wage drift, see Edgren-Faxén-Odhner (1973).

[2] Edgren-Faxén-Odhner (1973) confirm that the effect of structural shifts on wage drift for the whole group of industrial workers amounts to only a few tenths of one per cent.

is by no means clear to what extent these "structural" shifts are exogenous to economic conditions reflected in the labor market situation, so it would be wise not to indulge too much in purging the wage drift data. In this study I have chosen to work with a wage drift concept based on both time and piece rates but excluding other forms of payment. These wage forms dominate the wage bill of private employers and the corresponding wage drift figures show a high degree of reliability.

Our wage drift data will not be confined to workers in manufacturing as has often been the case in other studies but are based on the whole SAF-LO area. This means that wage drift for workers in transportation, trade and other private services is included in our figures. There seems to be little reason for excluding these groups _per se_ and the widest possible coverage is in itself desirable because we will relate the wage drift data to the labor market as a whole.

The fact that we rely on wage data for the whole SAF-LO area puts a restriction on how long a time-series we can use. In 1965 SAF was joined by the Commercial Employers' Confederation ("Handelns Arbetsgivareorganisation"), by which the SAF-LO area increased by approximately 10 per cent. Within the commercial trades wage drift is considerably lower than within manufacturing industry, so the merger is to be expected to cause a shift in the level of wage drift, calculated for the whole SAF-LO area. As a consequence the estimation period begins with 1965/66 (second quarters).

As we are working with yearly data, based on an average for only one quarter, we are able to specify the influence of the labor market variable on wage drift more precisely than, e.g., Jacobsson-Lindbeck could. These authors had to work with changes in yearly averages with the consequence that a weighting system had to be applied to the labor market variable, comprising data for a two year period for each observation on wage drift.[1] Strictly speaking our corresponding weighting system should comprise 15 months, where the observations for April, May and June overlap (the weights would be 1/6-5/6, 1/2-1/2 and 5/6-1/6 for these months in the first and the second of two consecutive years, respectively). However, taking into account the probability of some reaction lag, we will simply work with an average lag of two months. This means that we will relate the wage drift between two consecutive second quarters to the labor market situation during the period May-April, measured as an average of monthly figures.

PRESENTATION OF THE ESTIMATION RESULTS

We will first present the estimated wage drift equations for the observation period 1965/66-75/76 where the labor market variable is identified as unfilled vacancies (UV) and the duration of vacancies (DV), respectively. The exact definitions of the variables are found in the preceding text. More precisely the duration measure is given as

[1] See Jacobsson-Lindbeck (1969), Appendix II. The rationale for this procedure has nothing to do with any supposed time-lag in the influence of the labor market situation on wage drift. The reason is the simple statistical fact that when changes are measured in terms of averages, calculated over a period of time (as opposed to changes in levels calculated at a point of time), the measured amount of change depends on when the change occurs within the period.

the value of T according to eqs.(II) and (III)[1]. (The values of the variables are reproduced in Appendix II). Both linear and loglinear specifications are tried.

The estimation method is ordinary least squares. t-values are given below coefficients. F- and Durbin-Watson-(DW-) statistics are also presented as well as the standard error of the regression (SE).

The estimated equations turn out to be

(1 a) $WD = 0.02 + 1.02 \cdot 10^{-4}\, UV$ $\quad\bar{R}^2=0.40$; $F=7.73$;
$\qquad\qquad(0.01)\ (2.78)$ $\quad DW=0.78$; $SE=1.19$

(1 b) $\log WD = -9.47 + 1.02 \cdot \log UV$ $\quad\bar{R}^2=0.41$; $F=7.96$;
$\qquad\qquad(-2.44)\ (2.82)$ $\quad DW=0.76$; $SE=0.26$

(2 a) $WD = -0.92 + 8.28 \cdot DV$ $\quad\bar{R}^2=0.75$; $F=31.08$
$\qquad\qquad(-0.91)\ (5.57)$ $\quad DW=2.30$; $SE=0.77$

(2 b) $\log WD = 2.01 + 1.23 \cdot \log DV$ $\quad\bar{R}^2=0.77$; $F=34.77$
$\qquad\qquad(19.11)\ (5.90)$ $\quad DW=2.10$; $SE=0.16$

The estimation results are rather striking. The traditional measure, unfilled vacancies, balances on the edge of significance on 1% level. However, the significance of the duration of vacancies is clearly much higher, its t-value being twice as high as that of unfilled vacancies. The $\bar{R}^2$'s and the F-values tell the same story.

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[1] The numerical value of $\Theta(=1/T)$ in eq.(III) is determined with a degree of exactness corresponding to an error in the functional value of less than /10/.

The choice of functional specifications does not seem to influence the results very much. Indeed, the two alternative estimates on unfilled vacancies are consistent as far as they indicate a linear relationship without an intercept. With regard to the duration of vacancies there is a slight indication that the loglinear specification gives a better fit.

At this point it must furthermore be observed that the Durbin-Watson test indicates a high degree of positive first order serial correlation in the residuals of the equations (1). The coefficient of autocorrelation shows a value in the neighborhood of 0.5, when estimated by the Cochrane-Orcutt procedure. On the other hand, no high degree of serial correlation seems to disturb the estimates when the duration of vacancies is used. As is well known the presence of serial correlation may indicate the existence of a specification error in the equation at the same time as it puts the reliability of the ordinary least square estimators in doubt; especially the t-statistics may be biased, although the small number of observations makes any conclusion uncertain.[1,2]

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[1] See Maddala (1977).

[2] An inspection of the residuals of equations (1) shows that they exhibit a very spectacular positive time trend. As a matter of fact the introduction of a simple trend variable in eq.(1) does improve the values of the various test statistics considerably. However, a trend variable offers no "explanation", if it cannot be shown that it performs as a proxy for another variable that is theoretically relevant (e.g., increases in the price level). In Appendix III several alternative specifications of the wage drift equation are tried. The estimations show that such an interpretation of the time trend cannot be sustained.

So far we are able to maintain that our hypothesis has not been refuted that the duration of vacancies is a satisfactory indicator of the state of demand in the labor market and that it exhibits a close relationship to the amount of wage drift in the private sector. Furthermore, it can be claimed that the duration of vacancies performs better than the number of unfilled vacancies in this respect, which is also in accordance with our hypothesis.

However, I must admit that I am surprised that the difference in explanatory power between unfilled vacancies and the duration of vacancies should turn out to be so large. My astonishment concentrates upon the deterioration in the goodness of fit of the former measure, compared to the satisfactory results earlier reported on time-series extending into the seventies. A scrutiny of the residuals in the equations (1) gives a clue to the answer: although a positive time trend in the residuals is visible also for the first part of our estimation period it becomes much more pronounced when observations from the years 1973/74-1975/76 are added. It seems as if we have been aided in discriminating between the two labor market variables by a structural trend, in which the time paths of those variables have parted. The characteristic of that trend is that the duration of vacancies has increased while the number of unfilled vacancies has remained fairly stable.[1] This development strengthens the impression that

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[1] These phenomena are arithmetically reconciled by the fact that the number of announced new vacancies shows a downward trend. See the figures reproduced in Appendix II.

the labor market has become less apt to respond to changes in demand during the last decade. Our findings are consistent with the notion that this deterioration in flexibility has immediate consequences for the propensity to wage drift.

One may legitimately ask whether not the exaggeration of announced vacancies and the corresponding underestimation of our measure of duration before 1968 has by chance helped us to explain low wage drift figures. However, only three observations are affected and one of them only partly. Furthermore, two of the residuals are positive so the effect of a "correction" is not bound to impair the goodness of fit of eq.(2).

Let us now extend our estimation period to include data from 1976/77-78/79. The labor market variables during these three observation periods were to some extent influenced by the new legislation requiring employers to report their vacancies to the local employment office. The number of recorded unfilled vacancies must unequivocally have been augmented by this, *ceteris paribus*, causing an expected overestimation of wage drift. The effect on the measure of duration of vacancies is less certain, but as was mentioned earlier it is most likely that the expected result is an underestimation of wage drift, although the effect is probably rather weak.

If we first look at the estimated equations for the period 1965/66-78/79 as a whole, we immediately observe an equalization in the goodness of fit compared with the results given by (1) and (2).

The equations turn out to be

(1'a) $WD = -0.31 + 1.06 \cdot 10^{-4} UV$ $\bar{R}^2=0.42$; $F=10.44$
$(-0.21)\ (3.23)$ $DW=1.25$; $SE=1.10$

(1'b) $\log WD = -9.76 + 1.05 \cdot \log UV$ $\bar{R}^2=0.42$; $F=10.47$
$(-2.83)(3.24)$ $DW=1.17$; $SE=0.25$

(2'a) $WD = -0.47 + 7.08 \cdot DV$ $\bar{R}^2=0.50$; $F=13.93$
$(-0.36)(3.73)$ $DW=1.60$; $SE=1.03$

(2'b) $\log WD = 1.85 + 1.04 \cdot \log DV$ $\bar{R}^2=0.51$; $F=14.42$
$(14.16)\ (3.80)$ $DW=1.41$; $SE=0.23$

Obviously there is a marked deterioration in the goodness of fit when the duration measure is used. A slight increase in the values of the t- and F-statistics of the unfilled vacancies equations can be noticed. However, the performance of the duration of vacancies is still the best one. It might also be observed that the values of the DW-statistics are more inconclusive than in our estimations (1) and (2), but still the unfilled vacancies estimates seem to be more disturbed by serial correlation. As regards the importance of the functional specifications both (1') and (2') may be taken to indicate the existence of a linear relationship without an intercept.

The explanation of this relative change in goodness of fit and in the test statistics is evident when we observe that the amount of wage drift was low during 1976/77-78/79, and produces negative residuals in all the estimated equations (see Appendix I). The swing in residuals between 1975/76 and 1976/77 is of the same order of magnitude in the equations (1') and (2'). As the residual of

1975/76 is very strongly positive in equation (1') - a result of the earlier established positive time trend - the downward swing produces only moderate negative residuals for subsequent periods. On the other hand, in equation (2') the observation for 1975/76 is not very far off the regression line, so in this case the swing produces large negative residuals for the last observations.

The fact that the swing in the residuals is almost the same when unfilled vacancies and the duration of vacancies are used in the estimations makes it difficult to ascertain whether the introduction of compulsory vacancy reporting really has produced those expected biases in the estimates of wage drift, which we supposed earlier[1]. For our estimation results they do not seem to be of crucial importance, at least not if compared with other influences, which we will shortly consider.

As they stand, the estimation results given by (1') and (2') clearly weakens the case for the duration of vacancies as the superior measure of the degree of disequilibrium in the labor market when wage drift behavior is considered. However, I maintain that a more comprehensive residual analysis will restore its position and as well indicate what kind of explanatory factor is ignored in the present approach.

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[1] A comparison of the residuals for 1975/76 and for 1976/77-78/79 may indicate that the introduction of compulsory vacancy reporting has had opposite effects on the number of unfilled vacancies and on the magnitude of the duration of vacancies which is at least not contradictory to what we have supposed.

The residual pattern of the unfilled vacancies equations has already been outlined: a clearly visible positive time-trend throughout the estimation period up to 1975/76, abruptly broken by three concluding negative residuals (see Appendix I). It is not easy to derive an explanation of such a pattern from economic reasoning.[1]

The residuals of the duration of vacancies equations behave differently. Their pattern is strikingly similar to that of the Swedish business cycle during the period of estimation, not only with regard to timing but also to amplitude (see Appendix I).

It is well-known that the Swedish economy has experienced far more violent swings during the seventies than during the earlier part of the post-war period. I would argue that the large residuals of the duration equations coincide with periods when 'exceptional circumstances' have prevailed, in which wage drift is influenced by other forces besides the degree of disequilibrium in the labor market. I would guess that some measure of the functional distribution of income should be able to capture the effects of the forces operating in such situations.

Clearly, the formulation of a theory, which establishes the influence of such a measure on the wage drift process, requires as careful an investigation as that which has been accomplished with regard to the labor market influence. This task

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[1] See Appendix III, where some (unsuccessful) attempts are made to complement the unfilled vacancies variable with other explanatory variables.

cannot be carried through in this paper. I would like to point out, however, that there exists a theoretical basis for this kind of influence in the literature on bargaining theory in situations of bilateral monopoly in the labor market, where the possibility of collective labor action on local level as well as employer reaction is taken into account.[1]

A TENTATIVE TEST OF PROFIT INFLUENCE ON WAGE DRIFT

Consequently, we will make an attempt to include a measure of the functional distribution of income as a complementary explanatory variable in the estimated wage drift equation. As should be clear, this exercise is not to be interpreted as a test of a new elaborated hypothesis. Its value is chiefly derived from the fact that it may illustrate whether we have been successful in correctly identifying the degree of disequilibrium in the labor market, i.e., whether the 'unexplained' variations in wage drift depend on an incomplete specification of the estimated equation rather than on an incorrectly specified labor market variable.

There are several alternatives to choose between if we want to measure the share of income that goes to capital. I have selected the series on the percentage yield on equity capital in real terms that is calculated and published by Aktiv Placering AB.[2] The reason for this choice is twofold.

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[1] See de Menil (1971).

[2] The figures are published in Skandinaviska Enskilda Banken Quarterly Review, No.3-4 1979, where the matters of method are also presented.

First of all these profitability figures are calculated in a consistent way both over time and between companies. Secondly the figures include all business companies quoted at the Stockholm stock exchange (except banks) which means that, e.g., building industry and retail trade are represented.[1] Thus the correspondence between the profit data and the wage drift data are better than if we restricted ourselves to manufacturing industry. As a matter of fact to use profitability figures from manufacturing is otherwise the only remaining alternative, if we want data that are at all reliable.

As was mentioned the yield on equity capital is measured in real terms. This should be an advantage as it excludes the possibility that the estimated coefficient of the profitability variable partly reflects the impact of price increases on wage drift. Whether this profitability measure, calculated by external experts quite a long time after the relevant profit period, really corresponds to the profit situation as it was originally perceived by the bargaining parties at local level is of course questionable. However, this dilemma is not easily solved if we want the aggregation to be carried out on data that are internally consistent.

The postulated time pattern, according to which the profitability is supposed to influence wage drift, is indeed crude. I simply assume that the impact is immediate. As the profitability figures

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[1] In the present context, one obvious deficiency of the Aktiv Placering series is the inclusion of the performances of foreign subsidiaries.

are given on a calendar year basis, they are weighted so as to roughly correspond to the wage drift period (meaning that 2/3 of the profitability of the first year and 1/3 of the profitability of the second year are related to the corresponding second quarter wage drift figures).

There remains one relevant piece of information that ought to be taken into account before we run our final regressions. It is known that some wage drift regularly occurs when the centrally negotiated agreement is implemented at local level. For some years during our estimation period the central agreements were delayed so the implementations did not manage to affect the second quarter earnings statistics (1968/69, 1970/71, 1974/75 and 1976/77). As a consequence the subsequent periods were affected by two implementations. We would expect this to cause a shift in the timing of wage drift with regard to our second quarter period, leading to a tendency towards negative serial correlation in the residuals. We will handle this problem by introducing an "implementation" variable which takes on the values 0, 1 or 2, corresponding to the number of central agreement implementations during the period in question.[1]

Denoting the profitability variable P and the implementation variable D, we have the following

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[1] According to Edgren-Faxén-Odhner (1973) the statistical calculations might systematically underestimate the impact of the central agreements on actual earnings. Thus an alternative interpretation of the coefficient of the implementation variable is that it reflects an item of wage drift that is really to be regarded as a centrally negotiated wage increase. However, the basis for such a claim is somewhat arbitrary and it is safer just to talk of 'wage drift in connection with central agreement implementations'.

estimated linear equations for the period 1965/66-1978/79.

$$(3)\quad WD = -2.12 + 6.86 \cdot DV + 0.42 \cdot P \qquad \bar{R}^2=0.79;\ F=24.80$$
$$\qquad (-2.25)\ (5.52)\qquad (4.13)\qquad DW=3.02;\ SE=0.67$$

$$(4)\quad WD = -3.13 + 7.13 \cdot DV + 0.47 \cdot P + 0.53 \cdot D$$
$$\qquad (-3.71)\ (7.20)\qquad (5.67)\qquad (2.74)$$
$$\bar{R}^2=0.87\ ;\ F=28.79\ ;\ DW=1.65\ ;\ SE=0.53$$

Both equations perform very satisfactorily. Equation (4) reaches as high a degree of goodness of fit as one can hope for when working with real life data and its F-value is remarkably high, considering four parameters being estimated.

It is especially worth noting that when we include more explanatory variables as we go from equations (2') to (3) and (4) the significance of the duration variable does increase markedly. Likewise the inclusion of the implementation variable increases the significance of the profitability measure. As was expected it eliminates the negative serial correlation present in (3) according to the DW-test. So while the implementation variable is in itself barely significant at 1% level, its contribution is not to be overlooked.

On the whole the high values of the t-statistics indicate that the influence of the explanatory variables are easily separable (i.e., the degree of multicollinearity is low) and complement each other very well in contributing to a high total degree of goodness of fit.[1]

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[1] Anyone suspecting that several combinations of labor market, profit and price variables can reach the same result should consult Appendix III. They do not!

CONCLUDING REMARKS

On the basis of the preceding analysis and the presented estimation results we are able to maintain that the degree of disequilibrium in the labor market, that is relevant for the development of wage drift, is adequately measured by the length of duration of vacancies. In addition there seems to exist an independent impact on wage drift from profits.

If the above hypothesis constitutes an accurate model of reality, this offers an additional explanation why the number of unfilled vacancies has shown a close correlation with wage drift in earlier studies. Let us return to equation (I), i.e., $v \cdot T = V$. As was pointed out, the number of unfilled vacancies (V) reflects not only the length of the duration of vacancies (T) but also the number of new vacancies (v), i.e., the demand for new labor.

In a typical business cycle, the upswing is characterized by increases in utilization of fixed capacity, increased profits <u>and</u> increased labor demand. However, during this phase the labor market is not yet strained and T does not change very much. As a consequence changes in the value of V will reflect changes in v and so V is likely to be highly correlated with profits.

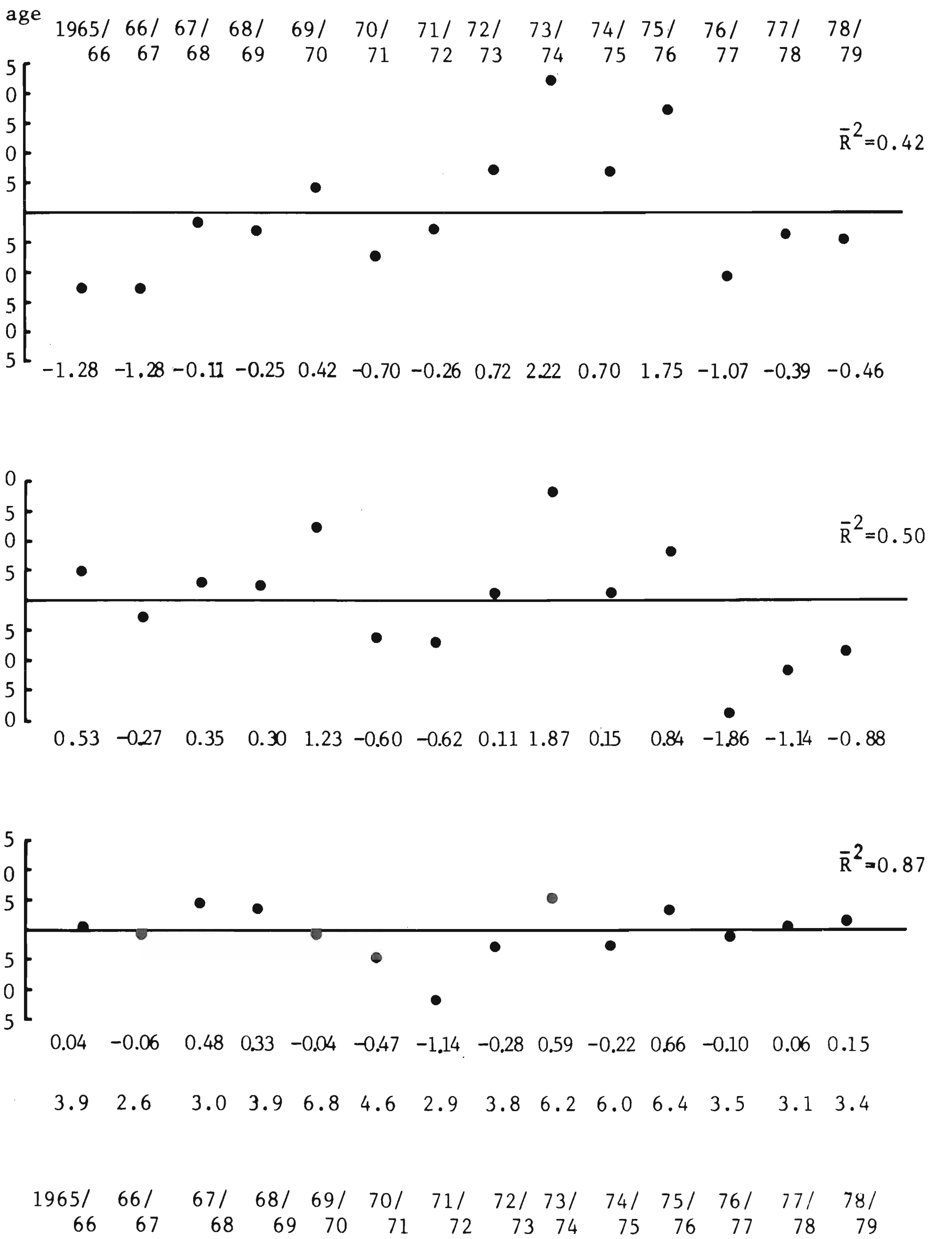
During a later phase of the cycle both profits and the demand for additional employees stagnate and start to decline, but the high level of total demand for labor relative to supply will lead to a sustained increase in T. So now changes in the value of V will mainly reflect the development of T.

Thus the time path of unfilled vacancies will follow those of both profits and the duration of vacancies during different phases of the traditional business cycle and the unfilled vacancies variable is able to perform their roles alternatively in an estimated wage drift equation. However, when structural changes disrupt the stability of the traditional relationships between the relevant variables during the cycle, the number of unfilled vacancies ceases to be a satisfactory proxy.[1]

[1] We have already discussed extensively the expected effects of the shifts that have taken place in the relations between various stocks and flows in the labor market in the seventies. There is another phenomenon, that may also contribute to the failure of unfilled vacancies to explain the amount of wage drift during this decade. Profits within Swedish industry have then been much more influenced by price changes than during the sixties and fifties (after the Korean boom). The number of vacancies is most likely to be highly correlated with profits, when these are influenced mainly by changes in volume of production and in capacity utilization. On the other hand, the vacancy figures are not to be expected to reflect profit increases that are purely price induced in situation where the utilization of fixed capacity has reached its upper limit.

APPENDIX I

RESIDUALS OF ESTIMATED EQUATIONS



APPENDIX II

PRESENTATION OF BASIC DATA

Table 1. Unfilled vacancies (UV)

	May	June	July	Aug	Sept	Oct	Nov	Dec	Jan	Feb	March	April	Average May/April
1965/66	65764	64397	60917	58648	53894	49622	46836	43099	40499	41643	45388	51967	51889
1966/67	56401	55380	51545	45929	40542	35767	33331	30650	27924	28950	31807	36894	39593
1967/68	41459	41561	38569	33762	28471	26194	27317	26663	25954	27190	30823	40269	32352
1968/69	46837	45594	42439	38949	35727	34590	35824	36728	37736	41154	49081	61480	42178
1969/70	70871	69692	65956	63285	59441	57594	58507	57514	55637	58299	65538	76844	63265
1970/71	81964	76687	70329	63338	55160	49185	45047	40240	34901	33859	38730	46470	52992
1971/72	48828	44599	38883	33107	27697	25554	26666	25781	24309	26436	31656	39463	32748
1972/73	44366	41528	36584	31912	26866	25018	26302	26259	25537	26825	32483	40741	32035
1973/74	44357	42576	39766	37077	33873	34119	35896	33440	32995	38571	50095	63625	40532
1974/75	68790	61139	53575	52100	44894	45156	44618	42167	42659	47648	55647	78339	53061
1975/76	79259	61467	49648	48808	37247	33865	35921	32875	33363	37282	49052	64004	46899
1976/77	69778	60550	48994	45700	39110	38346	37775	33188	31048	35857	49312	63441	46092
1977/78	59980	49276	37096	31957	23667	23090	27540	24262	23700	29204	42777	58303	35904
1978/79	53986	38700	29774	30547	24968	25423	30994	26939	30091	44979	60344	76482	39436

Source: Swedish Labour Market Board (AMS)

Table 2. <u>Announced vacancies</u>

	May	June	July	Aug	Sept	Oct	Nov	Dec	Jan	Feb	March	April
1965/66	119479	118560	82530	98635	91532	85153	88432	81365	80558	86231	100597	96261
1966/67	116453	116248	76803	91594	83835	76434	73716	64611	67147	71198	72042	86278
1967/68	99315	95980	66917	82083	73103	72156	71060	60794	69768	59654	64222	78807
1968/69	91146	85151	65870	76843	71535	71065	66668	56665	73256	65170	77144	90325
1969/70	88950	95671	68379	79367	78303	71564	69457	62491	70722	62523	66976	92803
1970/71	88517	88154	58835	69526	67398	57937	58611	50141	48301	45540	62924	70521
1971/72	73354	68269	48116	56375	53444	48312	54302	50641	43716	46117	56828	68670
1972/73	73041	67142	46552	59903	51548	51335	50644	41513	50776	43341	58923	60990
1973/74	76223	64601	47921	61545	53591	59264	58880	45048	56071	53961	69782	82857
1974/75	86278	67323	47041	60097	57350	61211	56092	41795	54708	54553	57291	84745
1975/76	68562	60396	37825	51999	49501	52503	55674	42182	41593	50328	65953	70664
1976/77	73554	61024	38273	55073	50869	50340	58466	42035	45302	50989	68542	70875
1977/78	61866	63781	33859	52526	44487	44028	52842	43966	45368	47567	63683	78482
1978/79	69741	61703	36003	57633	50321	49698	58626	46582	52807	62096	80590	79481

<u>Source</u>: Swedish Labour Market Board (AMS)

Table 3. Durations of vacancies (DV) (fractions of months)

	May	June	July	Aug	Sept	Oct	Nov	Dec	Jan	Feb	March	April	Average May/April
1965/66	.506	.541	.724	.589	.578	.572	.524	.558	.498	.485	.456	.481	.543
1966/67	.490	.475	.657	.492	.475	.460	.448	.469	.412	.408	.446	.434	.472
1967/68	.422	.433	.567	.406	.384	.361	.442	.436	.371	.458	.488	.533	.442
1968/69	.527	.533	.632	.500	.493	.484	.541	.652	.517	.646	.666	.727	.576
1969/70	.843	.724	.935	.784	.742	.794	.848	.912	.776	.956	1.046	.887	.854
1970/71	.957	.843	1.121	.865	.772	.807	.744	.767	.689	.736	.636	.693	.802
1971/72	.675	.637	.553	.566	.502	.521	.497	.509	.549	.583	.577	.601	.564
1972/73	.625	.608	.748	.519	.506	.482	.524	.632	.501	.627	.571	.712	.588
1973/74	.593	.651	.806	.592	.617	.577	.617	.724	.587	.752	.782	.839	.678
1974/75	.822	.907	1.012	.856	.739	.739	.792	.976	.784	.920	1.060	1.107	.893
1975/76	1.166	.879	1.101	.930	.682	.629	.655	.753	.807	.770	.818	1.036	.852
1976/77	.994	.916	1.081	.806	.725	.756	.644	.751	.672	.735	.798	1.013	.824
1977/78	.940	.717	.914	.587	.503	.522	.536	.538	.520	.645	.747	.827	.666
1978/79	.778	.574	.741	.532	.481	.512	.547	.560	.584	.826	.829	1.099	.672

Calculated from tables 1 and 2 according to eqs. (II) and (III).

Table 4. Wage drift (%) (WD)

1965/66	66/67	67/68	68/69	69/70	70/71	71/72	72/73	73/74	74/75	75/76	76/77	77/78	78/79
3.9	2.6	3.0	3.9	6.8	4.6	2.9	3.8	6.2	6.0	6.4	3.5	3.1	3.4

Source: Swedish Employers' Confederation (SAF)

Table 5. Profitability (P)

1965	66	67	68	69	70	71	72	73	74	75	76	77	78	79[a]
6	4	3.5	5	6	5.5	4.3	4.2	6.9	7.2	4.1	2.3	0.3	1.1	4

[a] Estimated

Source: Aktiv Placering AB

APPENDIX III

In this Appendix some alternative specifications of wage drift equations are estimated for the period 1965/66-78/79. Besides the variables presen-
t-
ed in the main text, the increase in consumer price index (CPI) and the unemployment ratio (U) according to the Labor Force Surveys are introduced as explanatory variables.

The unemployment ratio is supposed to influence the amount of wage drift according to the same pattern as unfilled vacancies and the duration of vacancies. The CPI variable is measured as the increase in percentage terms from December year i-1 to December year i when the wage drift period is second quarter year i to second quarter year i+1.

t-statistics						F-statistics	$\bar{R}^2$	DW
Variables								
U	UV	DV	CPI	P	D			
-0.88						0.77	-0.02	1.15
-0.80			0.20			0.37	-0.11	1.19
-0.86			0.15		0.47	0.30	-0.19	0.98
-1.23				2.48		3.62	0.29	1.15
-1.57				2.80	1.30	3.13	0.33	0.53
	3.10		0.37			4.91	0.38	1.36
	3.17		0.32		0.86	3.45	0.36	0.95
	2.64			1.74		7.62	0.50	1.32
	2.86			2.02	1.37	6.12	0.54	0.67
		4.51	-1.91(!)			10.32	0.59	1.82
		4.68	-2.05(!)		1.12	7.47	0.60	1.22

The only result worth commenting is the extremely weak performance of the CPI variable. When complementing the duration of vacancies variable it shows significance at 5% level but with negative sign!

Even if a postulated influence of consumer price increases on wage drift may be specified in more sophisticated ways, it does not seem likely that such a variable should contribute very much towards a deeper understanding of the wage drift process.

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