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THE DURATION OF UNEMPLOYMENT AND UNEXPECTED INFLATION - AN EMPIRICAL ANALYSIS

by

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I. Introduction*

A distinguishing message of the theory of search unemployment is that short-run unemployment fluctuations are explainable by inflationary surprises. Unemployment is basically viewed as productive investment in job search, chosen by employees in order 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.

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. A rising 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 address ourselves to the question of the empirical importance of the two competitive explanations. 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 our study lies in its ability to provide information about the relative importance of unexpected inflation and job opportunities as explanations of the duration of unemployment.[1] Another interesting feature of our paper is its comparative perspective; we apply the same model to both Swedish and U.S. data, thereby being able to reveal certain important differences between the labor markets in the two countries. We find e.g., perhaps somewhat surprisingly, that the U.S. unemployment duration is more or less unaffected by unexpected inflation, whereas the results for Sweden, on the other hand, give some support for the information-lag hypothesis. A third novelty of our study is the disaggregated data used (for Sweden only). By focussing the analysis on transition probabilities for workers with different lengths of (incompleted) spells, some interesting behavioral differences are observed; one finding is that the simple information-lag story is more valid for the short-term unemployed.

The paper is organized as follows: Section II below introduces the basic theoretical framework that guides our empirical estimation procedures; the latter are described in section III. Section IV presents the data employed and section V the empirical results. Some interpretations of our findings are discussed in the final section.

II. Optimal Search Policies and the Duration of Unemployment

Microeconomic explanations of unemployment have been focussing on the behaviour of the household, whereas the demand side generally has been considered as exogenous. We will follow that partial equilibrium approach, using a simple job search model as our theoretical framework.

Consider the behaviour of an unemployed worker according to the 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 left-ward 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 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 (μ) into two components, the job offer probability (θ) and the acceptance probability (P) we have

$$(1) \qquad \mu = \theta P = \theta[1-F(a)] \qquad \theta \leq 1$$

where a 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

$$D = 1/\mu = 1/\theta[1-F(a)] \tag{2}$$

Which are then the characteristics of an optimal search policy? In the simple case of infinite time horizon and no discounting, the optimal policy implies a certain time invariant reservation wage obtained as the solution to

$$C = \theta P[E(w|w>a)-a] = \theta\int_a^\infty (w-a)f(w)dw \tag{3}$$

where C is the (constant) marginal search cost and $f(\cdot)$ the known density function of wage offers.[2] Eq. (3) implies that the reservation wage declines as the job offer probability θ 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:

1. <u>The pure availability effect</u>: An increasing number of vacancies means a higher job offer probability, thereby reducing the dura-

tion of unemployment.

2. 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 ambigous a priori. Feinburg (1977) has, however, demonstrated that the availability effect will outweigh the supply effect under certain reasonable assumptions.

3. 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:

$$\mu = \underset{+}{\theta(V)}\,P(\underset{-}{V},\underset{+}{w/w^*}) = g(V,w/w^*) \qquad (4)$$

where V is the number of vacancies, w the actual average wage and w* the expected average wage.

We would argue that Eq. (4) 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.[3] We are suppressing other plausible determinants of unemployment duration, e.g. variations in unemployment compensation and the discount rate. These simplifications should not be too severe, since the cyclical fluctuations are dominating in the data. We have also excluded changes in the price level from consideration, perhaps a more questionable simplification. Unexpected price inflation does affect unemployment in some models within the microfoundations literature, although it is absent in the standard search model. The interpretation of this candidate regressor is, however, quite different in e.g. the Lucas-Rapping model compared to the Siven model (misperception of future prices versus misperception of current prices) and the theoretical predictions are completely opposite; a higher rate of unexpected price inflation will increase unemployment in Siven's model and decrease unemployment in the Lucas-Rapping model.[4] It is interesting to note that Seater's "unified model" represents a middle way; variations in unemployment are totally unaffected by how workers perceive changes in the price level.[5] We decided to exclude the price inflation variable from the regressions, thereby avoiding troublesome problems of interpretation.

III. Empirical analysis

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, i.e. to represent Eq. (4) above by a suitable functional form. The basic specification used will be:

$$(5) \qquad \ln\mu_t = \alpha_1+\alpha_2\ln V_t+\alpha_3\ln(w_t/w_t^*)$$

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.[6]

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. The expanding literature about the formation of expectations give several alternatives which all are quite plausible. However, no model which can be made operational can be considered "correct" in all respects. 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.

A. Adaptive expectations

The first model used is a type of adaptive expectations. These expectations are formed according to a finite distributed lag of past wage changes, i.e., with quarterly data (which is used for Sweden):

$$(6a) \qquad \left(\frac{w_t^*}{w_{t-4}}\right) = \sum_{i=1}^{4} \ell_i \left(\frac{w_{t-i}}{w_{t-4-i}}\right)$$

where

(6b) $$\sum_{i=1}^{4} \ell_i = \frac{1}{10} \sum_{i=1}^{4} (5-i) = 1$$

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

(7a) $$\left(\frac{w_t^*}{w_{t-12}}\right) = \sum_{i=1}^{12} \ell_i \left(\frac{w_{t-i}}{w_{t-12-i}}\right)$$

where

(7b) $$\sum_{i=1}^{12} \ell_i = \frac{1}{78} \sum_{i=1}^{12} (13-i) = 1$$

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.)

B. Expectations from an ARMA-process

Even though the simplicity of the simple adaptive model is appealing - since it might be argued that workers form their expectations in a simple and cheap way - it could also 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. Therefore we have proceeded as follows:

The process has been reestimated each period and reidentified each fourth period (with quarterly data) and each twelfth period (with monthly data).[7] For Sweden the character of the process changed over time; when observations from 1960 onwards were used the appropriate process changed from an AR(1) to an AR(1)MA(2), back again to an AR(1) and finally - during the past two years (1976-1977) - an MA(10) on the first differences of the variable (i.e. the process was non-stationary). All the time autoregressive seasonal terms had to be used.

For the U.S. the process was stationary when data from 1960 until 1969 were used.- AR(1) with first a seasonal autoregressive term and then a seasonal moving average term. From then on the process became non-stationary with an MA(1) term and a seasonal moving average term on the first differences.

C. Expectations from an estimated wage-equation

It could, finally, 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 estimated wage-equations (quarterly data) like:

$$(8)\ WCH = \beta_1 + \beta_2 \cdot (V_{t-1} + V_{t-2} + V_{t-3} + V_{t-4}) + \beta_3 WCH$$

where

$$WCH = (w_t - w_{t-4})/w_{t-4}.$$

Again the model was reestimated each fourth period and with data from the last five years. On the whole the estimated equations performed reasonably well for Sweden according to standard statistical criteria.

This approach was less successful for U.S.; the available vacancy indicators turned out to be bad predictors of wage inflation. We decided to exclude this expectations-formation scheme for the U.S. regressions.

IV. 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 survery are interviewed again three months later, whereas different individuals are interviewed in two subsequent months. In order to improve the estimates we decided 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:

$$(9) \qquad G_t^{1,14} = f \sum_{i=0}^{13} (1-\mu_1)^i$$

$$(10) \qquad G_{t+13}^{14,27} = G_t^{1,13}[1-\mu_2]^{13}$$

$$(11) \qquad G_{t+26}^{27,39} = G_{t+13}^{14,26}[1-\mu_3]^{13}$$

Three transition probabilities are obtained - μ_1, μ_2 and μ_3 - which can be regarded as conditional upon the length of the spell of unemployment. By using available data on f, $G_t^{1,14}$, $G_{t+13}^{14,27}$ etc. we obtain μ_1 from

$$(12) \qquad \frac{(1-\mu_1)^{13}-1}{(1-\mu_1)-1} = \frac{G_t^{1,14}}{f}$$

whereas μ_2 and μ_3 are calculated as

$$(13) \qquad \mu_2 = 1 - \left[\frac{G_t^{14,27}}{G_t^{1,13}}\right]^{1/13}$$

$$(14) \qquad \mu_3 = 1 - \left[\frac{G_{t+26}^{27,39}}{G_{t+13}^{14,26}}\right]^{1/13}$$

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 National Bureau of Statistics. All data used refer to manufacturing industry.

The U.S. transition probabilities refer to the labor market as a whole. They were computed by using the method proposed by Barron (1975). The essential idea is 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.

The U.S. wage data are average hourly earnings in manufacturing industry, reported in Employment and Earnings.[8] As vacancy data for the period 1965-1975 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.

V. Empirical results

The results from alternative estimations are presented in Tables 1 and 2 below. The estimation method is weighted-least-squares and the appropriate weights are derived in an appendix.

Let us first have a look at the results obtained for Sweden. We observe, in the first place, that the detection-lag variable is significant both for the short-term unemployed (1-13 weeks)

Table 1. Transition probability equations for Sweden.

Quarterly data 1968.1-1977.3

Adaptive expectations	V	w/w*	$\bar{R}^2$	DW
Short-term unemployed (μ_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 (μ_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 (μ_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
ARMA expectations				
Short-term unemployed (μ_1)				
(10)	0.98 (4.94)	7.47 (2.59)	0.50	1.43
(11)	-	11.00 (3.08)	0.18	0.79

cont.

ARMA expectations	V	w/w*	$\bar{R}^2$	DW
Medium-term unemployed (μ_2)				
(12)	0.36 (3.56)	2.21 (1.64)	0.33	2.19
(13)	-	3.56 (2.41)	0.11	1.72
Long-term unemployed (μ_3)				
(14)	0.36 (1.94)	-2.16 (-0.81)	0.05	2.27
Expectations from wage-equations				
Short-term unemployed (μ_1)				
(15)	1.13 (5.93)	8.43 (2.79)	0.51	1.66
(16)	-	7.76 (1.85)	0.06	0.65
Medium-term unemployed (μ_2)				
(17)	0.40 (4.23)	3.10 (2.38)	0.37	2.27
(18)	-	3.37 (2.14)	0.09	1.63
Long-term unemployed (μ_3)				
(19)	0.30 (1.72)	-3.20 (-1.31)	0.07	2.20

Note: $\bar{R}^2$ is the fraction of the weighted variance of the dependent variable explained by the weighted independent variables, adjusted for degrees of freedom. The $\bar{R}^2$ obtained when regressing μ_1 on $\hat{\mu}_1$ from Eq. (1) was 0.62.

Table 2. Transition probability equation for the U.S.

Monthly data 1969.4-1973.10 and 1965.2-1975.12, respectively

Adaptive expectations	HWA	Vm	w/w*	TIME	$\bar{R}^2$	DW	ρ
1969.4-1973.10							
1	-	0.23 (11.27)	1.62 (1.59)	-0.0008 (-1.61)	0.73	1.19	-
2	-	0.21 (8.57)	0.91 (1.11)	-0.0002 (-0.41)	n.a.	2.02	0.29
3	-	0.24 (11.81)	1.42 (1.38)	-	0.72	1.13	-
4	-	0.21 (8.71)	0.87 (1.08)	-	n.a.	2.03	0.30
5	-	-	0.14 (0.08)	-0.0021 (-2.38)	0.07	0.37	-
6	0.50	-	1.36 (1.33)	-0.0031 (-6.39)	0.72	1.18	-
7	0.44 (8.21)	-	0.71 (0.86)	-0.0022 (-3.84)	n.a.	1.97	0.31
8	0.45 (7.64)	-	0.21 (0.15)	-	0.51	0.67	-
1965.2-1975.12							
9	0.52 (16.81)	-	0.70 (1.28)	-0.0025 (-19.74)	0.83	1.34	-
10	0.53 (11.45)	-	0.47 (0.81)	-0.0025 (-13.26)	n.a.	2.03	0.34
11	0.49 (7.93)	-	1.55 (1.43)	-	0.33	0.34	-
12	-	-	0.71 (0.73)	-0.0024 (-10.62)	0.47	0.44	-

cont.

Adaptive expectations	HWA	Vm	w/w*	TIME	$\bar{R}^2$	DW	ρ
1969.4-1973.10							
13	-	0.23 (11.27)	2.57 (2.07)	-0.0007 (-1.54)	0.74	1.15	-
14	-	0.20 (8.63)	1.96 (1.98)	-0.0002 (-0.40)	n.a.	2.03	0.30
15	-	0.24 (11.92)	2.48 (1.98)	-	0.73	1.10	-
16	-	0.20 (8.81)	1.93 (1.98)	-	n.a.	2.04	0.31
17	-	0.24 (11.64)	-	-	0.71	1.12	-
18	-	0.20 (8.68)	-	-	n.a.	2.04	0.30
19	-	-	3.03 (1.31)	-0.0021 (-2.49)	0.10	0.36	-
20	0.49 (11.23)	-	2.48 (2.00)	-0.0030 (-6.44)	0.73	1.17	-
21	0.44 (8.36)	-	1.81 (1.80)	-0.0022 (-3.96)	n.a.	1.99	0.30
22	0.44 (7.71)	-	2.23 (1.34)	-	0.53	0.65	-
1965.2-1975.12							
23	0.52 (16.7)	-	-0.37 (-0.50)	-0.0026 (-19.19)	0.83	1.33	-
24	0.53 (11.37)	-	-0.07 (-0.10)	-0.0025 (-13.01)	n.a.	2.04	0.35
25	0.49 (8.03)	-	3.45 (2.48)	-	0.35	0.41	-
26	-	-	-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 approach.

and for the medium-term unemployed (14-26 weeks). These results hold for all models of expectations.[9] For the long-term unemployed, on the other hand, no significant detection-lag effect is revealed; the coefficient has even a wrong sign.

The job availability variable (V) is significantly positive in all regressions, even for the long-term unemployed. Dropping this variable produces in most cases a marked decrease in the DW-value, indicating the presence of specification errors.

Which are then 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".[10] 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 favour 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 finite search horizon will, under some stationary conditions, fall with the duration of unemployment, a theoretical prediction which has been given empirical support.[11] Eventually the reservation wage will coincide 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 find out 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 size of the parameters as well as the variation of the independent variables into account. The question might be illuminated by comparing the predicted transition probabilities using estimates from regressions in the table

$$\hat{\mu}_t = \alpha_1 \cdot V_t^{\alpha_2} \cdot \left(\frac{w_t}{w_t^*}\right)^{\alpha_3} \tag{15}$$

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

(16) $$\tilde{\mu}_t = \alpha_1 \cdot V_t^{\alpha_2}$$

Using the results from the adaptive model Figure 2 below demonstrates the relative unimportance of the detection-lag effect for the medium-term unemployed. Inflationary surprises produce, on the other hand, quite important unemployment effects for the short-term unemployed during the peak years 1969-70 and 1974-75. (Figure 1.) The main part of the variation is, however, attributable to the vacancy-variable.

Turning now to the U.S. regressions, 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 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.[12] 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,

Figure 1. The effects of unexpected inflation - short-term unemployed in Sweden

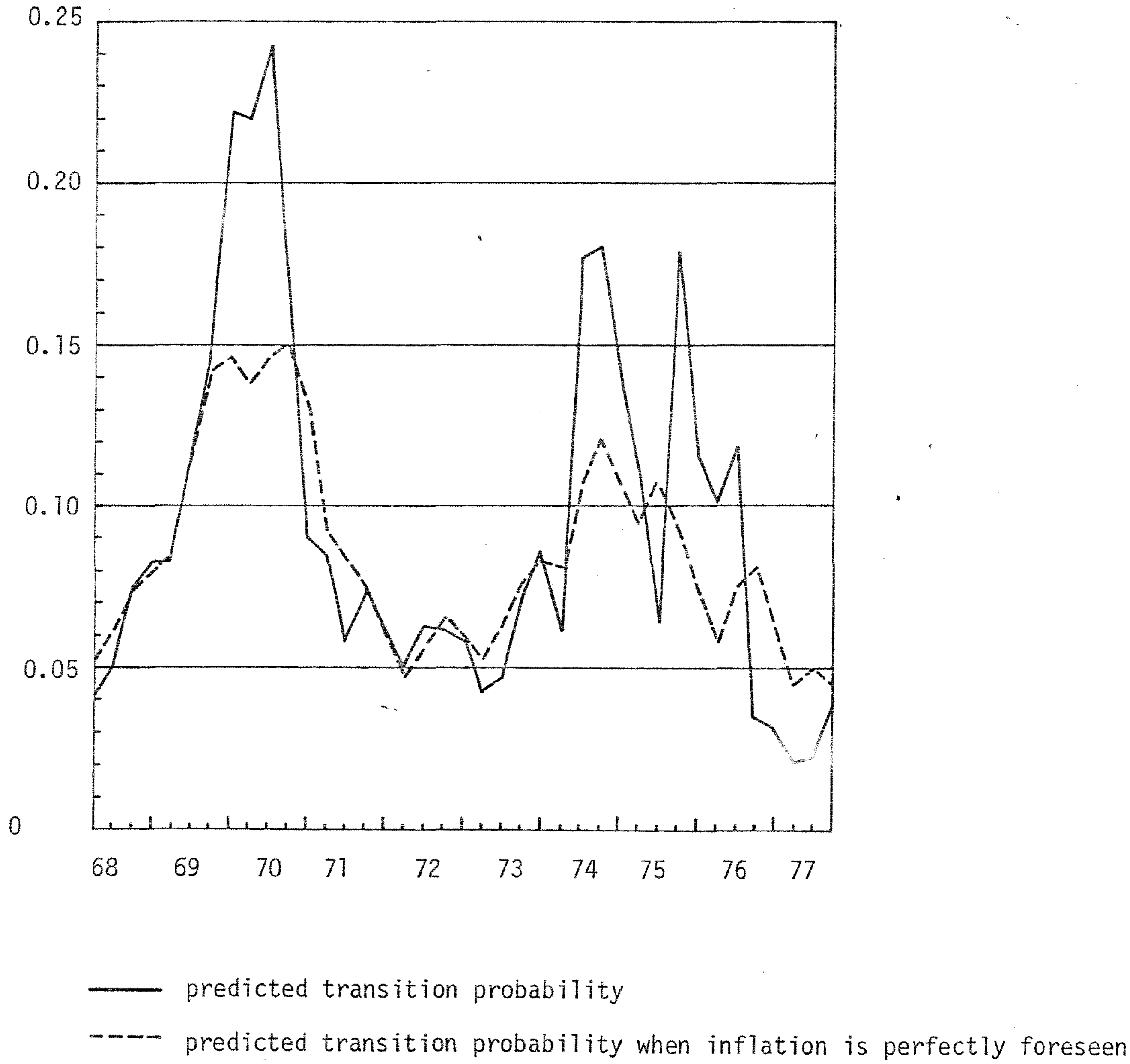
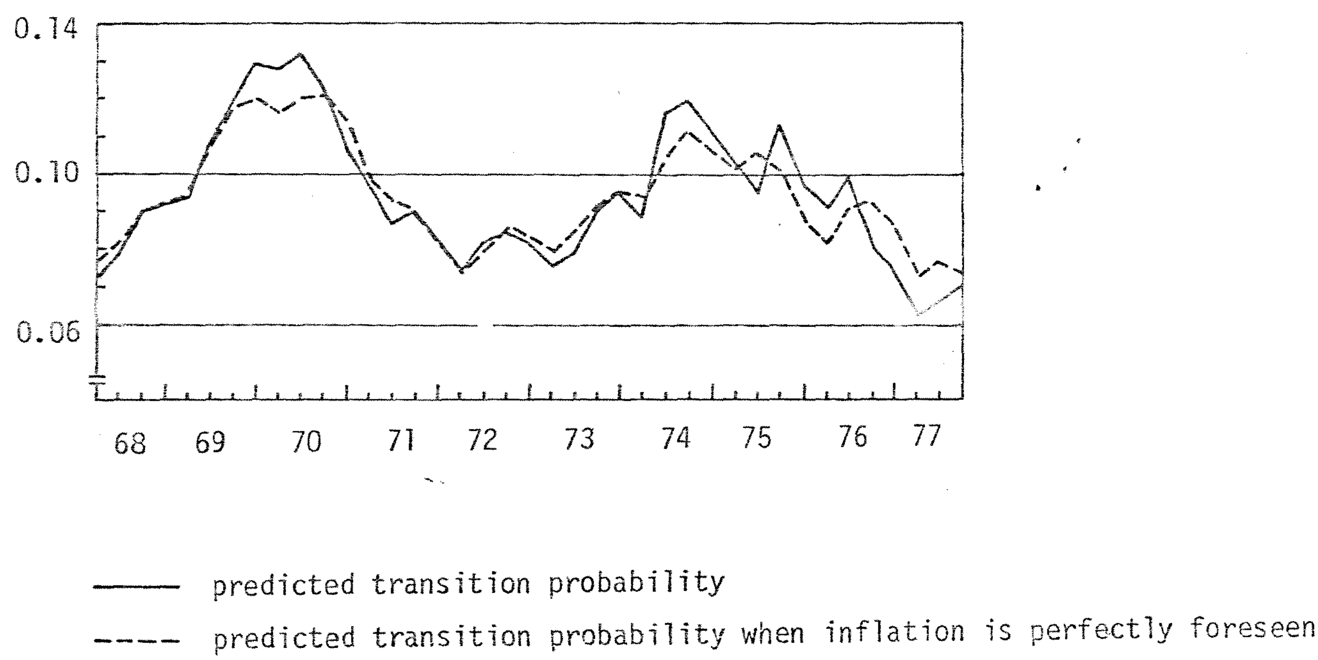


Figure 2. The effects of unexpected inflation - medium-term unemployed in Sweden



however, rule out the possibility of some detection-lag effects in operation, at least during certain time-periods and - especially - if the expectations are formed according to an ARMA-process rather than adaptively.

VI. 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. Since the flow into unemployment is fairly stable over the cycle, our results imply, moreover, that cyclical changes in the unemployment rate are only slightly affected by inflationary surprises.

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.

Let us, finally, offer some comments to the observed differences between the Swe[illegible] ly 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 the 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'[13] 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 U.S. - the importance of temporary layoffs. Temporary layoffs constitute - as Martin Feldstein has pointed out[14] - 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 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.[15] Feldstein's view of those laid off as "waiting" rather than "searching" has been questioned on empirical grounds.[16] 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 then a non-seeking unemployed worker on layoff respond to unexpected general wage inflation? He would, most likely, be less inclined to search, thereby reacting similar to his employed fellows; a familiar 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 would obviously be less suitable when applied to the middle state.

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FOOTNOTES

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[1] The question has earlier been addressed by Barron (1975) and Axelsson and Löfgren (1977). Their methods differ from ours.

[2] For a proof of (3), see e.g. Lippman and McCall (1976).

[3] The worker in Siven's and Seater's models is maximizing his lifetime utility by using search in the labor market as one important choice variable. Siven also considers search in the goods market but assumes leisure to be fixed; maximization of the utility functional is therefore equivalent to maximization of lifetime earnings. Seater, on the other hand, takes account of variable leisure but ignores search on the goods market.

[4] Unexpected price inflation implies in the Siven-model a re-allocation of time from search in the labor market to search in the goods market thereby causing a decline of the job offer probability. The reservation wage will also increase, reinforcing the effect on unemployment duration. The Lucas-Rapping model is hardly suitable for analyzing the length of spells of unemployment since it disregards job search and considers unemployment as pure leisure, resulting as a difference between actual and normal employment. Darby (1976) and Kesselman-Savin (1978) have

run unemployment regressions for the U.S. including unanticipated price increases as an explanatory variable. The results turn out to be unsatisfactory; the coefficients are as a rule insignificantly different from zero and the signs are unstable across different regressions.

[5] Seater (1978).

[6] The crucial trick in Barron's approach - followed by Axelsson and Löfgren - is to construct a model which gives an explicit specification of the relationship between the number of vacancies (V) and the job offer probability (θ). Given such a relationship, $\Theta = f(V)$, the acceptance probability is obtained as $P = \mu/f(V)$. The procedure is interesting since it can validate a pro-cyclical reservation wage pattern (i.e. P and V are inversely correlated). The approach requires, however, some fairly restrictive assumptions regarding the relationship between θ and V; Barron assumes that $\theta = k \cdot V$, implying that the elasticity $\partial \ell n\theta/\partial \ell nV$ equals one, an implication from the assumption that each firm has only one vacancy in each occupation. It can be shown that less restrictive assumptions produce an elasticity lower than one. Barron's procedure is, moreover, unable to separate the supply effect from the detection-lag effect. Our approach, on the other hand, can quantify the detection-lag effect but captures only the net availability effect.

[7] A Box-Jenkins-program called T-series available at the Stockholm School of Economics has been used. For identification criteria, see Nelson (1973).

[8] In some regressions we also tried average hourly earnings for the total private non-agricultural sector. The results were basically the same.

[9] We have also tried logit-specifications in some cases, as well as adaptive expectations with shorter lags. The results turned out to be fairly robust with respect to these changes.

[10] Santomero & Seater (1978) p. 525.

[11] See articles by Gronau (1971), Kasper (1967) and Kiefer & Neumann (1979).

[12] The coefficient of w/w* is significant in Eq.(25) but the DW-value indicates that the t-ratio should not be taken seriously.

[13] Phelps (1971) pp. 6-7.

[14] See Feldstein (1975).

[15] 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).

[16] See the paper by Bradshaw & Scholl (1976) and the following discussion in the Brookings Paper.

APPENDIX

An estimation problem arises because the dependent variable is an estimate of the "true" transition probability. This estimate of the transition probability is subject to sampling variation and this variation obviously enters in the regression equation as stochastic disturbances. Since this variation is not constant the assumptions of ordinary least squares are violated.

Theil (1967) has derived the following variance of the disturbances for the logit model:

$$\text{(A.1)} \qquad \text{Var}(\varepsilon_{1t}) = \frac{1}{N_t \hat{\mu}_t (1-\hat{\mu}_t)}$$

where $\hat{\mu}_1$ is the estimate of the transition probability and N_t the number of observations. By using the same procedure as Theil the following variance for the log-linear model can be derived (see below)

$$\text{(A.2)} \qquad \text{Var}(\varepsilon_{2t}) = \frac{1-\hat{\mu}_t}{N_t \hat{\mu}_t}$$

The appropriate weights are given by $1/\text{Var}(\varepsilon_{2t})$.

The derivation of (A.2) proceeds as follows: Consider the basic relation between the "true" transition probability for individual i at time t and the explanatory variable X_t

(A.3) $$\ln \mu_{it} = \ln \alpha + \beta \ln X_t$$

Since the explanatory variables have the same values for all individuals in these applications index i has been omitted.

When the estimate $\hat{\mu}_t$ of the death-risk is inserted into the equation (A.3) instead of μ_t the sampling variation necessitates the inclusion of a distrubance ε_t

(A.4) $$\ln \hat{\mu}_t = \ln \alpha + \beta \ln X_t + \varepsilon_t$$

Now, the problem is to express the variance of ε_t in terms of the observable $\hat{\mu}_t$.

The average of (A.3) is

(A.5) $$\frac{1}{N_t} \sum_i \ln \mu_{it} = \ln \alpha + \beta \ell n X_t$$

By subtracting (A.5) from (A.4) we obtain:

(A.6) $$\varepsilon_t = \frac{1}{N_t} \sum_i (\ln \hat{\mu}_t - \ln \mu_{it})$$

The expression in parentheses can be simplified to:

(A.7) $$\ln \hat{\mu}_t - \ln \mu_{it} = \ln \frac{\hat{\mu}_t}{\mu_{it}} = \ln\left(\frac{\mu_{it}}{\mu_{it}} - \frac{\mu_{it} - \hat{\mu}_t}{\mu_{it}}\right) = \ln\left(1 - \frac{\mu_{it} - \hat{\mu}_t}{\mu_{it}}\right)$$

The last expression can be simplified to

$$(A.8)\qquad \frac{\hat{\mu}_t-\mu_{it}}{\mu_{it}}$$

if $\hat{\mu}_t$ and μ_{it} are close to each other.

If (A.8) is inserted into (A.6) we have

$$(A.9)\qquad \varepsilon_t = \frac{\hat{\mu}-\bar{\mu}}{\bar{\mu}_t} \quad \text{where } \bar{\mu}_t = \frac{1}{N_t}\sum_i^N \mu_{it}$$

The variance of ε_t now becomes:

$$(A.10)\qquad \frac{E(\hat{\mu}_t-\bar{\mu}_t)^2}{\bar{\mu}_t^2} = \frac{\bar{\mu}_t(1-\bar{\mu}_t) - \text{var}(\mu_{it})}{N_t\bar{\mu}_t^2}$$

If we, as Theil, disregard $\text{var}(\mu_{it})$ and approximate $\dot{\mu}$ by $\hat{\mu}$ we obtain:

$$(A.11)\qquad \text{Var}\ (\varepsilon_t) = \frac{1-\hat{\mu}_t}{N_t\hat{\mu}_t}$$