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.

---

* 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

$$u = f \cdot D \qquad (1)$$

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

---

[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

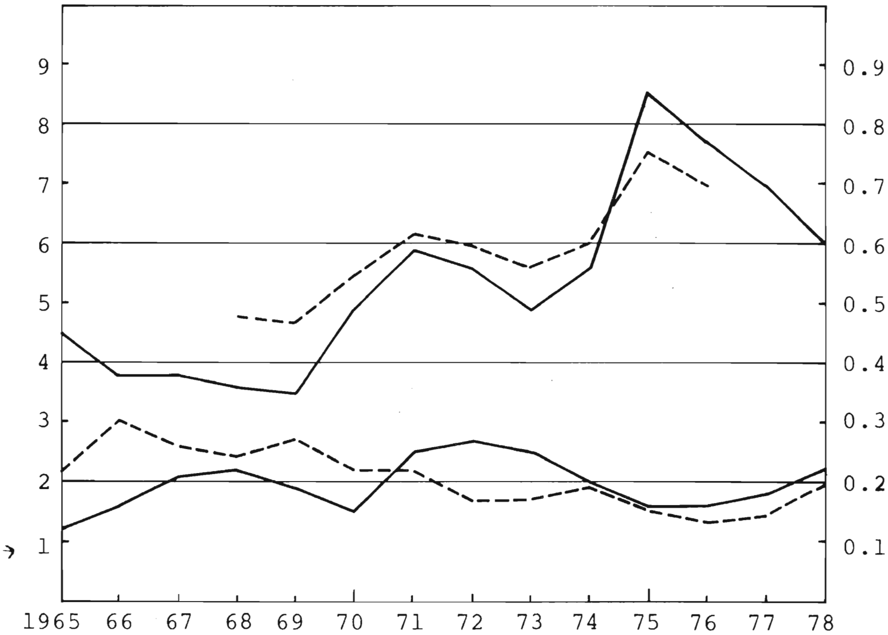
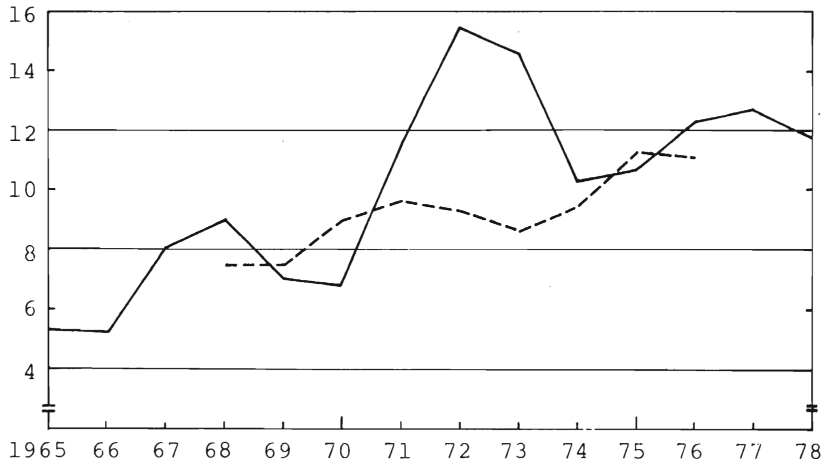


Figure 2. The average duration of the spells of unemployment (weeks): Sweden (———) and the U.S. (- - -).



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.

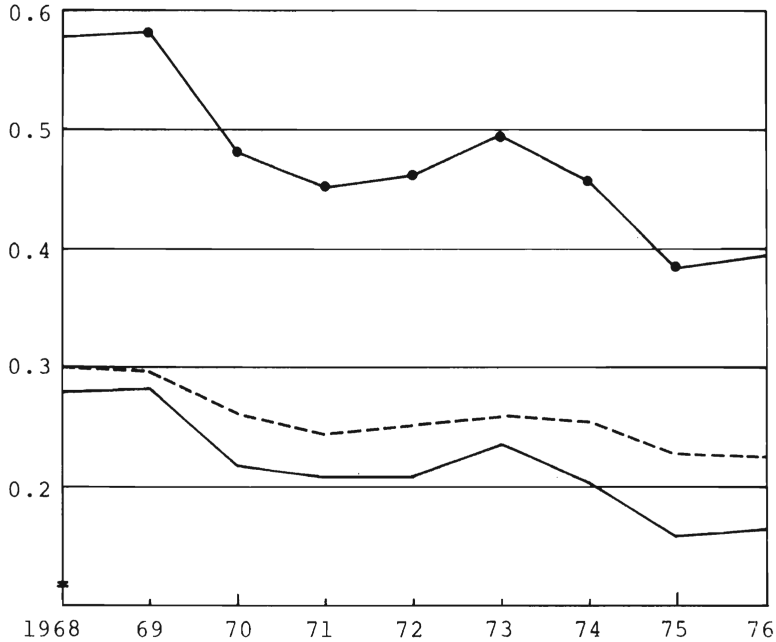
As is well known, the duration of unemployment equals the inverse of the probability of leaving unemployment. This transition probability, in turn, equals the sum of (i) the job-finding probability (Pue) and (ii) the labor force exit probability (Pun). Data are available for the U.S. to make such a decomposition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

---

[1] Unpublished CPS data mentioned in footnote above.

Figure 3. Components of U.S. unemployment outflow. 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.

---

[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.

---

[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.

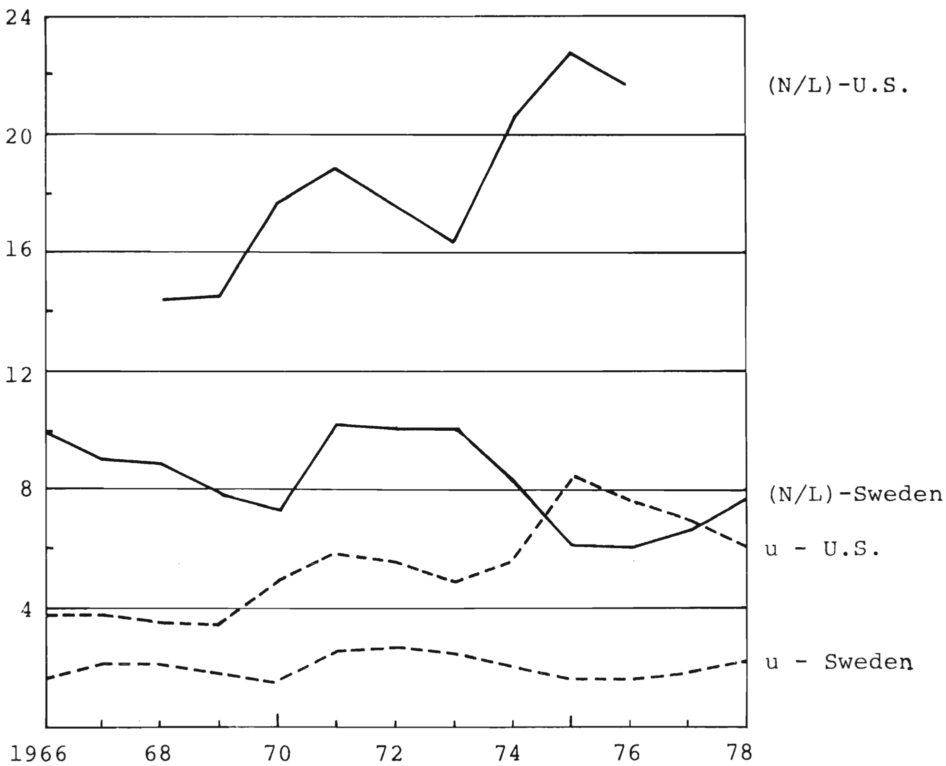


Figure 5A. The average number of yearly unemployment weeks per worker with unemployment experiences during a year (W); Sweden and the U.S.

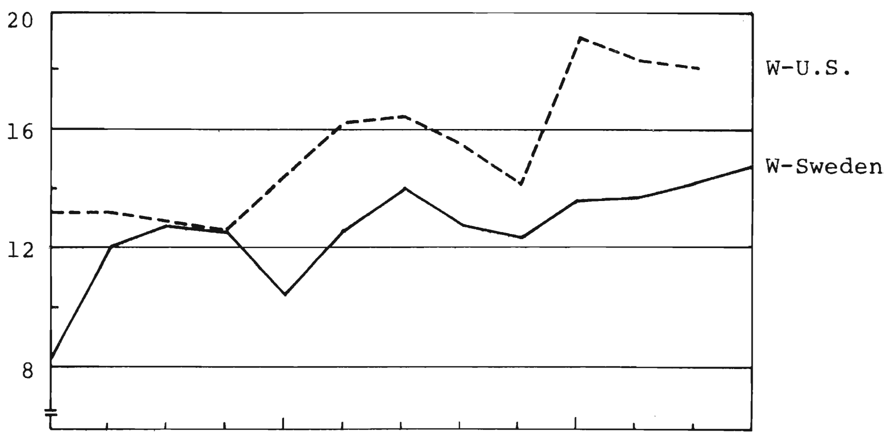
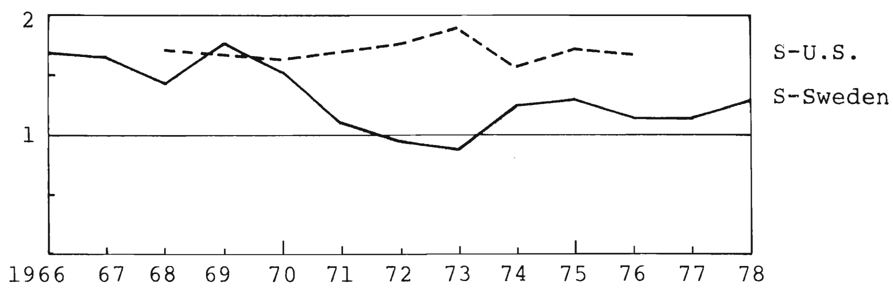


Figure 5B. The average number of unemployment spells per worker with unemployment experiences during a year (S); Sweden and the U.S.



(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.

---

[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 spells 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

---

[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:

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

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]

---

[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-

---

[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-

---

[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 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

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

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. <u>The effects of unexpected inflation - short-term unemployed in Sweden</u>

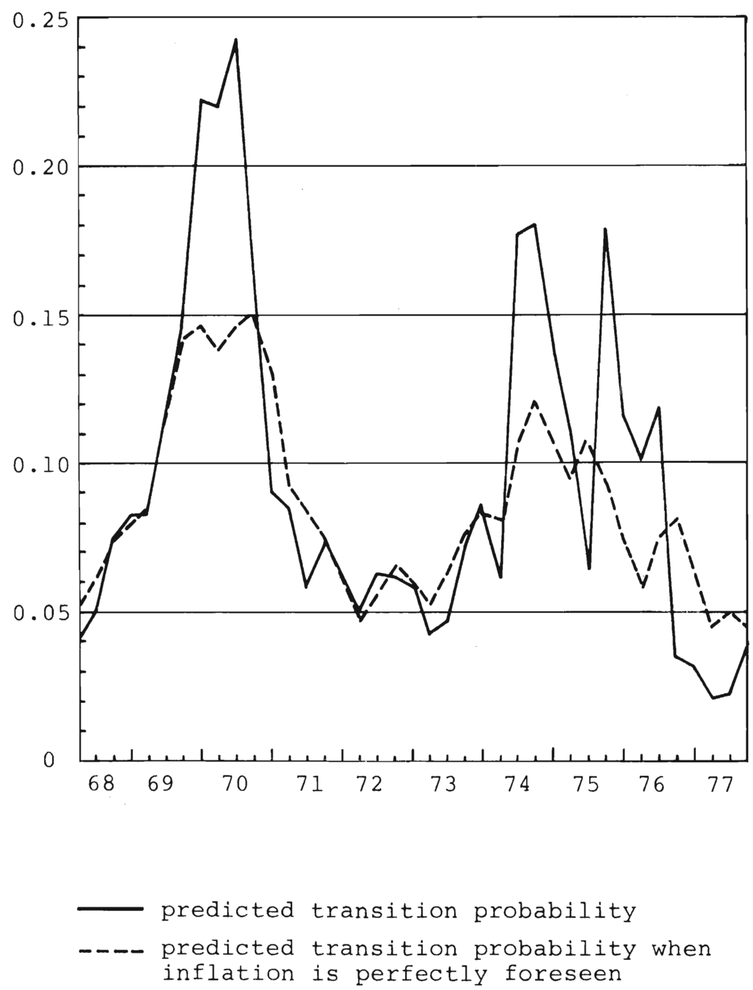
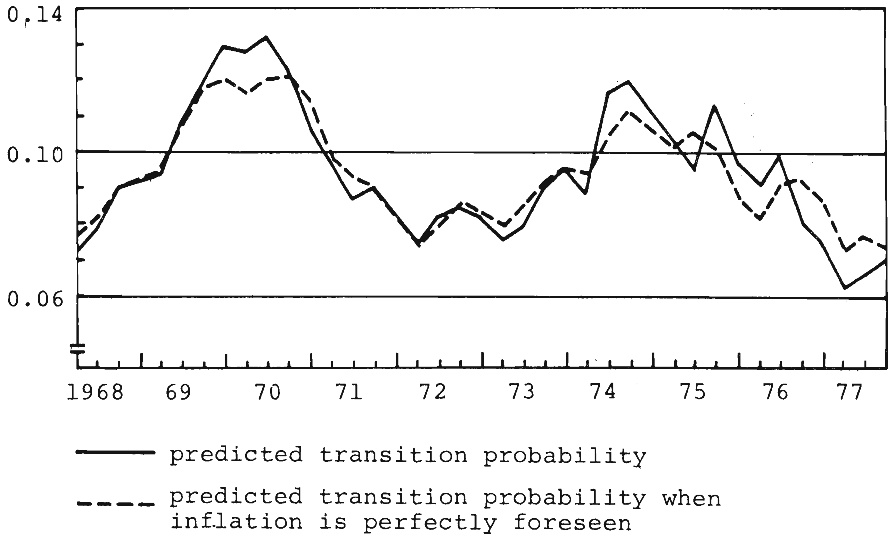


Figure 7. The effects of unexpected inflation - medium-term unemployed in Sweden



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-

---

[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 <u>positive</u> 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.

---

[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.

---

[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

---

[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-

---

[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.

REFERENCES

Axelsson, R. and Löfgren, K.G. (1977), "The Demand for Labor and Search Activity in the Swedish Labor Market", Europ. Econ. Rev., 9, 1977, 345-59.

Axelsson, R., Holmlund, B. and Löfgren, K.G. (1977), "On the Length of Spells of Unemployment in Sweden: Comment". Amer. Econ. Rev., March 1977, 67, 218-221.

Barrett, N.S. (1975), "The U.S. Phillips Curve and International Unemployment Rate Differentials: Comment", Amer. Econ. Rev., March 1975, 65, 213-21.

--- (1977), "On the Length of Spells of Unemployment in Sweden: Reply". Amer. Econ. Rev., March 1977, 67, 222-224.

Barron, J.M. (1975), "Search in the Labor Market and the Duration of Unemployment: Some Empirical Evidence", Amer. Econ. Rev., Dec. 1975, 65, 934-42.

Björklund, A. and Holmlund, B. (1981), "The Duration of Unemployment and Unexpected Inflation - An Empirical Analysis." Amer. Econ. Rev., March 1981.

Björklund, A. (1979), "Spells and Duration of Unemployment in Sweden 1965-1978." IIM-papers 79-17, International Institute of Management, Berlin. June 1979.

Bradshaw, T.F. and Scholl, J.L. (1976), "The Extent of Job Search during Layoff", Brookings Papers, Washington 1976, 2, 515-26.

Clark, K.B. and Summers, L.H. (1978), "Labor Force Transitions and Unemployment". NBER Working Paper, No. 277, August 1978.

Feinberg, R. (1977), "Search in the Labor Market and the Duration of Unemployment: Note", Amer. Econ. Rev., Dec. 1977, 67, 1011-13.

Feldstein, M. (1975), "The Importance of Temporary Layoffs: An Empirical Analysis", Brookings Papers, Washington 1975, 3, 725-44.

Flanagan, R.J. (1973), "The U.S. Phillips Curve and International Unemployment Rate Differentials", Amer. Econ. Rev., March 1973, 63, 114-31.

Gronau, R. (1971), "Information and Frictional Unemployment". Amer. Econ. Rev., June 1971, 61, 290-301.

Kasper, H. (1967), "The Asking Price of Labor and the Duration of Unemployment", Rev. Econ. and Statis., May 1967, 49, 165-72.

Kiefer, N.M. and Neumann, G.R. (1979), "An Empirical Job-Search Model with a Test of the Constant Reservation-Wage Hypothesis", J. Polit. Econ., Febr. 1979, 87, 89-107.

Morgenstern, R.D. and Barrett, N.S. (1974), "The Retrospective Bias in Unemployment Reporting by Sex, Race, and Age." J. Amer. Statist. Assn., June 1974, 69, 355-57.

Persson, M. (1979), Inflationary Expectations and the Natural Rate Hypothesis, Stockholm School of Economics (Diss.).

Phelps, E.S. (1971), "Introduction: The New Microeconomics in Employment and Inflation Theory", in E.S. Phelps et al. (1971),1-23.

Phelps, E.S. et al. (1971), "Microeconomic Foundations of Employment and Inflation Theory", Macmillan, London.

Santomero, A.M. and Seater, J.J. (1978) "The Inflation-Unemployment Trade-Off: A Critique of the Literature", J. Econ Lit., June 1978, 499-544.

Seater, "J.J. (1977), "A Unified Model of Consumption, Labor Supply, and Job Search", J. Econ. Theory, Apr. 1977, 14, 349-72.

--- (1978), "Utility Maximization, Aggregate Labor Force Behavior, and the Phillips Curve", J. Monet. Econ., Nov. 1978, 4, 687-713.

--- (1979), "Job Search and Vacancy Contacts", Amer. Econ. Rev., June 1979, 69, 411-19.

Siven, C.H. (1979), "A Study in the Theory of Inflation and Unemployment", North Holland, Amsterdam.

Smith, R. (1977), "A Simulation Model of the Demographic Composition of Employment, Unemployment and Labor Force Participation". In R.G. Ehrenberg (ed.), "Research in Labor Economics", Greenwich, Connecticut 1977.

Arbetsmarknadsstyrelsen (AMS), "Arbetsmarknadsstatistik", (The National Labor Market Board, "Labor Market Statistics"), various issues, Stockholm.

OECD, "Main Economic Indicators", 1960-1975, Paris 1976.

Statistiska Centralbyrån (SCB), "Arbetskraftsundersökningar" (The Swedish Central Bureau of Statistics, "Swedish Labor Force Surveys", AKU), yearly averages 1975-78.

U.S. Bureau of Labor Statistics (BLS), "Employment and Earnings", various issues, and "Monthly Labor Review", various issues.