ON THE EFFECTS OF DURATION ON FINNISH UNEMPLOYMENT

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The overall exit rate from unemployment conditional on duration of unemployment is often shown to be downward sloping. This can be result of a) individual heterogeneity, i.e. some individuals have initially fewer chances finding of job than others or b) negative state-dependence, i.e. unemployment experience worsens individual exit probability from unemployment. This paper provides a test based on time series behaviour to exclude the pure heterogeneity hypothesis. We also apply the test on Finnish data and find support for rejecting pure heterogeneity in favour of some negative state-dependence effects.

1. Introduction

The persistence of the higher unemployment rate in Finland (as well as in most other European countries) since the mid-seventies has given rise to an extensive discussion among economists. Despite the efforts devoted to finding explanations of the events and hence the appropriate measures for the cure, there are still many aspects of unemployment which remain unsettled.

Ordinary newspaper discussion about the unemployment problem simply recognizes the higher level of unemployment rate, but for a more thorough analysis of the problem we have to know: 1) how and why do individuals become unemployed, 2) what is it that the stock of unemployment consists of, 3) how long do people stay unemployed and 4) how and when do they leave unemployment. We also need to take account of the individual experience of unemployment, or the structure of unemployment duration at individual level in order to make clear how severe actual unemployment is. Here we will concentrate on the individual exit probability from unemployment, and, in particular, we aim to analyse whether the long-term unemployed suffer more from their unemployment than the short-term unemployed in terms of lower probability of escaping from unemployment.

It is a well known fact that there are substantial variations in the length of unemployment spells, thus it has become a commonplace to talk about the short-term unemployed and the long-term unemployed. There is also evidence of great differences between these groups when it comes to the exit probability from unemployment. It has been shown that the proportion leaving unemployment is much higher among the short-term unemployed than among those with longer unemployment spells. The crucial question is, does this mainly

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* A preliminary draft of this paper was prepared for VIII Yrjö Jahnsson Graduate Course, 1986—87, and the work was completed during a visit to Nuffield College, University of Oxford, in 1987—88. Comments and helpful suggestions by Tor Eriksson, Richard Layard, Stephen Nickell, Tuire Santamäki, Pekka Sauramo, the participants of the VIII Yrjö Jahnsson Graduate Course and two anonymous referees are most gratefully acknowledged. My thanks are also due to Anthony Johnson, who checked the language. The remaining errors, of course, are totally my responsibility.
depend on the heterogeneity of unemployed people, e.g. that individuals differ in their chances of getting a job or is there some type of negative state-dependence behind the longer spells? By negative state-dependence we mean that unemployed people either get demoralized during longer unemployment spells, so that their search activity in the labour market declines the longer the duration of unemployment, or that long-term unemployment is seen by employers as an adverse sign of the worker’s productivity. This stigmatization of the long-term unemployed might be due to actual depreciation of individual human capital or simply the prejudice of employers.¹ In either case, the situation of the long-term unemployed is much more severe than in the case of heterogeneity.

Knowledge about which one of these two possible explanations of declining exit rates is closer to the actual situation of long-term unemployed people is of great importance for practical labour market policy. If long-term unemployment is due to state-dependence effects, there is a greater variety of measures which can be introduced in order to combat long-term unemployment. In the case of heterogeneity it is not so easy to envisage measures which would make the long-term unemployed more attractive for employers.²

Discrimination between the two hypotheses is, however, no simple matter. Basically, the problem is due to lack of suitable and reliable data on unemployment spells. As Jackman and Layard (1988) put it, there is always the possibility of unobservable differences between individuals when cross-section data is utilized.³ The alternative way of shedding some light on the issue is to predict aggregate time-series behaviour from assumptions about certain types of micro behaviour. Below, a test based on aggregate time-series, which would lead to a distinction between the two hypotheses, will be applied. Clearly, these two types of approach may be complementary.

We shall proceed as follows. In section 2, we describe the main features of the development of the Finnish labour market since 1968. We concentrate on some special aspects, which are particularly relevant. The exit probability conditional on the duration of unemployment spell is analysed in section 3. In section 4, the connection between long-term unemployment and the U-V-relationship is discussed. In section 5, we make use of time-series data to estimate the steady-state U-V-curve in order to test the proposition that declining exit rates by duration might have caused a shift in the U-V-curve. A discussion and a summary of the results are given in section 6.

2. The main features of the rise in the Finnish unemployment

The rise in the unemployment rate in mid-seventies was accompanied by a decrease in vacancy rate, and both have stayed at their new levels since then. In the eighties both unemployment and vacancies have been quite stable, but the greater difference between the two has persisted (see figures 1.a and 1.b). In terms of the standard U-V-curve we can recognize an outward shift of the relationship, (figures 1.c and 1.d) in 1972. A visual inspection of figure 1.d reveals that there is a possible new shift of the U-V-curve in the late 1970s and at the start of the 1980s.

1972 was the year of enforcement of the Employment Act, which extended the number of people eligible for state unemployment assistance (mostly young people and females). At the same time the coverage of unemployment insurance and assistance was gradually increased. The replacement ratio, though, has been declining since 1972 until 1985. ¹ Erikkson (1986) has calculated that the unemployment assistance in average was about 76 % of an average after-tax wage in 1972 and that it has since then declined until 1984 to 44 %.


¹ Especially, in case some of the longer durations are due to heterogeneity, the whole group of long-term unemployed easily becomes stigmatized.
² Jackman and Layard (1988).
Unemployment pensions were introduced already in the beginning of the 70s, but became more widely used first since end of the 70s. Because of the complexity of the unemployment income security system, the effects that these changes in the system have on unemployment are not straightforward.

The duration structure of Finnish unemployment has recently been studied by Eriksson (1985). The aggregate picture of duration structure is quite clear: there has been a positive trend in the proportion of long-term unemployed up to 1979, after which the proportion has stayed at about 27% of total unemployment, (see figure 2).6 According to Eriksson the average duration of unemployment spells is rather short, but a substantial fraction of unemployment is concentrated in long spells that have occurred since the mid-seventies. This disproportionate number of long spells indicates that every individual cannot have the same chance of obtaining a job, i.e. exit probabilities from unemployment vary. Eriksson also finds that the picture of unemployment being concentrated in long spells remains unchanged when more homog-

* Naturally, the concept of the long-term unemployment is a relative matter; in the Finnish case we would characterize a spell over a year as long-term (compared with e.g. the British practice of long-term unemployment consisting of those, who have been unemployed for a year or more).
Figure 1.c. The U-V-relationship, quarterly data, 1968—86.

Figure 1.d. The U-V-relationship, annual averages, 1968—86.

Enous populations, such as age groups or occupations, are considered.

During the same time the inflow into unemployment has not shown any increasing trend (figure 3.c)\(^7\), which indicates that the increase in unemployment was mainly due to the longer duration of spells (figure 3.b), and not due to an increase in the number of people becoming unemployed.

\(^7\) The inflow rate varied with unemployment rate, but not as much (figure 3.a and figure 3.c.).

Eriksson also studied state-dependence in unemployment using a micro data set, which contained information about each individual's labour market history. His approach, though, differs from ours below: instead of directly studying state-dependence in exit rates, he examined whether unemployment experience creates future unemployment or so-called occurrence dependence, i.e. there is a causal link between unemployment at different times. The result obtained was that this type of state dependence — in which people
who have been unemployed before have difficulties in obtaining a permanent job — cannot, indeed, be ruled out. This can, of course, be caused by the employers’ attitude towards persons who have been unemployed or by individuals’ loss of productivity through interruptions in work experience.

3. The exit probability from unemployment

In the case of Finland the issue of state-dependence in exit probabilities has not been explored. This is not surprising since there are obvious problems in making inferences about state-dependence effects from cross-section data and only few ideas of revealing it through examining aggregate time-series. This is also due to a lack of adequate data on exit rates conditional on duration for a longer period, a problem we will examine below. There are, however, some results reported by Eriksson (1985) on weekly exit probabilities, but these refer to populations grouped by age, occupation and region and duration-specific exit rates are not reported. He finds that the overall exit rates for all age groups, regions and types of occupations declined in the mid-seventies, implying that the change in the duration structure of Finnish unemployment has been a rather «democratic» one, since every group has suffered from it. Also Kettunen (1988) finds in his study based on cross-section data pooled with observations on each individual from 1985 to 1986 that hazard functions are declining in almost all duration categories. These results are consistent with our findings of the decline in duration-specific outflow rates for the long-term unemployed.

The duration-specific exit rates (i.e. the conditional probability of leaving unemployment) can be calculated only if there is data on outflows from unemployment for categories of equal length over time. This is the major problem in Finnish data on outflows. We have monthly data on outflows of up to 3 months, increased since 1976, but this is mainly because of introducing unemployment pensions and other forms of early retirements.

* As mentioned earlier, cross-section studies are always subject for biased estimates because of possible unobserved heterogeneity, which might result in downward bias of estimated exit probability. In Kettunen’s study this possibility is reduced by introduction of duration-dependent explanatory variables, in that study, i.e. benefit ratios. The question of state-dependence in exit rates is not, however, explicitly discussed in that study.

* The aged are an exception. Their exit rates actually
Figure 3.a. The unemployment rate, annual averages, 1968—85.

Figure 3.b. The average expected length of completed spells.

Figure 3.c. The weekly inflow rate, 1968—85.
and we are able to construct quarterly outflow rates of up to 6 months. For durations over 6 months (reported by Labour Ministry) we have data on quarterly outflows only from 1984 onwards. One way of dealing with this problem is to assume equal proportions leaving unemployment during a period of a half year and then calculate the quarterly exit rate. Let \( \pi_j \) be the quarterly exit rate for a certain duration category \( j \), \( S_{jt} \) the number of unemployed of duration category \( j \) staying on the register at time \( t \), \( S_{jt+\tau} \) the number of unemployed of duration cohort \( j \) still staying on the register after \( \tau \) quarters. Then the quarterly exit rate can be calculated as:

\[
\pi_j = 1 - \left( \frac{S_{jt+\tau}}{S_{jt}} \right)^{1/\tau}
\]

(1)

As we can see from table 1, there have not been large changes in the duration-specific exit rates in 1980s. However, this is not surprising considering the stable unemployment and vacancy situation during the period. Unfortunately, this stability of the labour market in '80s rules out a direct test of some state-de-

Table 1. Outflow from unemployment.

Percentage of unemployed in January leaving unemployment in next 3 months by duration in January.

<table>
<thead>
<tr>
<th>Year</th>
<th>0—3 months</th>
<th>3—6 months</th>
<th>6—9 months</th>
<th>over 9 months1</th>
</tr>
</thead>
<tbody>
<tr>
<td>1981</td>
<td>54.1</td>
<td>54.1</td>
<td>38.1</td>
<td>38.1</td>
</tr>
<tr>
<td>1982</td>
<td>54.1</td>
<td>54.1</td>
<td>35.6</td>
<td>35.6</td>
</tr>
<tr>
<td>1983</td>
<td>52.9</td>
<td>52.9</td>
<td>40.1</td>
<td>40.1</td>
</tr>
<tr>
<td>1984a</td>
<td>54.2</td>
<td>54.2</td>
<td>44.4</td>
<td>44.4</td>
</tr>
<tr>
<td>1985a</td>
<td>52.0</td>
<td>52.0</td>
<td>46.1</td>
<td>46.1</td>
</tr>
<tr>
<td>1986a</td>
<td>47.5</td>
<td>47.5</td>
<td>42.7</td>
<td>42.7</td>
</tr>
<tr>
<td>1984b</td>
<td>49.8</td>
<td>47.5</td>
<td>30.1</td>
<td>8.3</td>
</tr>
<tr>
<td>1985b</td>
<td>49.0</td>
<td>48.2</td>
<td>31.8</td>
<td>15.8</td>
</tr>
<tr>
<td>1986b</td>
<td>44.6</td>
<td>46.5</td>
<td>28.7</td>
<td>15.2</td>
</tr>
</tbody>
</table>

1 Compared with those who are still in the register after 6 months
a Computed using formula (1), as also 1981—1983, which leads to an assumption of constant outflow rate during a half year
b Computed using the actual outflows per quarter

dependence vs. pure heterogeneity proposed by Jackman & Layard (1988), since the test presupposes a comparison of two different states of the economy. In the case of Finland we do not have sufficiently large change in the state of the economy in 1980s to make this test reliable in any sense.

In any case we may notice that exit rates are clearly time-dependent; the longer the duration, the lower the exit rate. It is also remarkable that there is a sharp drop in the exit rate when we exceed a duration of six months: the decline is almost 20 percentage points compared with those whose spell of unemployment has lasted less than 6 months.\(^\text{10}\) This seems to be a good reason for defining duration longer than half a year as long-term unemployment. The clear drop of exit rates at 6 months duration is followed by a further sharp decline of escape probability for durations longer than 9 months. These latter escape probabilities, when defined as outflow from unemployment into employment, should be corrected for the large outflow of long term unemployed into the out-of-labour-force state because of the increasing number of unemployment pensioners since the early 1980’s.\(^\text{11}\) Unfortunately, such a correction cannot be done properly since it is difficult to estimate how long these persons, who have been earlier retired, would have stayed unemployed were there no unemployment pension arrangements. It is clear, though, that the reported exit probabilities in the group over 9 months duration would be smaller, if the correction could be done.\(^\text{12}\) On the other hand, among durations between 6 and 9 months, we might have persons waiting to qualify for unemployment pension, who are not efficiently searching employment. This increases the number of long-term unemployed (when defined in durations longer than 6 months) and is likely to cause a downward bias in the calculated exit rates for those being unemployed for over 6 months, but less than 9 months.

10 Unfortunately, this can be shown only for the last 3 years, since formula (1) smooths over the differences when it is applied for the previous years.
11 Because it is a prerequisite for receiving unemployment pension that prospective pensioner has been unemployed for over 200 days, it is clear that the exclusion of unemployment pensioners from the unemployment statistics reduces the number of long-term unemployed.
4. The U-V-relationship and long-term unemployment

In the Finnish case there is an obvious difference between the short-term and long-term unemployed based on the differences in the exit probability: becoming a long-term unemployed individual seems to make it more difficult to escape from unemployment. The underlying reason for this is not easy to reveal. As our data on duration-specific exit rates show that there is a major difference between short-term and long-term unemployment based on the low probability of exit when unemployment spells have lasted 6 months, we try to utilize this fact to accomplish a connection between movements of the U-V-relationship and long-term unemployment. This approach has been successfully applied by Budd, Levine & Smith (1986a, 1986b)\(^\text{13}\), who have developed a model for the pattern of unemployment for the case when duration-specific exit rates differ, and have shown that a rise in the share of long-term unemployed will shift the U-V-curve to the right if the differences in exit probabilities are due to negative state-dependence. They also show in a multi-country study that this model fits well for the labour markets in the UK and Germany. In the cases of the Netherlands and the United States this relationship cannot, however, be confirmed.

Let us outline the analysis the U-V-relationship and the role of long-term unemployment connected with it. In the steady state we find that the change in unemployment \(\Delta U = 0\), which implies that inflow into unemployment \(I\) equals outflow from unemployment \(O\):

\[
I = O
\]

Using (2) we can write the ratio \(U/N\), (where \(U\) is unemployment and \(N\) employment) as a ratio of overall inflow rate \(s\) and overall exit rate \(\pi\):

\[
\frac{U}{N} = \frac{I}{U} = \frac{S}{\pi}
\]

Assuming a constant inflow rate into unemployment, let us investigate how the overall exit rate \(\pi\) is generated under the two competing hypothesis of individual exit probability.

The pure state dependence hypothesis suggests that each individual becoming unemployed has initially the same probability of obtaining re-employment, say \(\pi_0\), but that unemployment experience itself affects the exit probability, so that the individual exit probability becomes a function of duration, say, that exit probability is given by \(\pi_t h(t)\). In case \(h'(t) > 0\), we have positive state-dependence, and \(h'(t) < 0\) implies negative state-dependence. The latter case in more interesting in from our point of view, since then we would also observe declining duration specific exit probabilities from aggregate time series data.

The pure heterogeneity hypothesis implies that persons entering unemployment have individual exit probabilities from unemployment, say \(\pi_{it}\) that will stay constant during an unemployment spell. We will, however, observe declining duration-specific exit probabilities from an aggregate time series data, due to the initial heterogeneity of entrants. This is because the proportion of individuals with lower exit probabilities among unemployed increases with time, as the individuals with high exit probability leave unemployment faster, i.e. as time passes the »quality« of the stock of unemployed in each duration category decreases as persons with higher individual exit probability leave unemployment, and consequently, the duration-specific overall exit rates are declining.\(^\text{14}\)

Obviously, it can also be that we have a mixed case, with different initial individual exit probabilities combined with some duration effects, i.e. the individual exit probability is given by \(\pi_t h(t)\). If duration effects are negative, the two factors reinforce each other in producing a declining overall exit rate.

Let us now assume, quite reasonably, that the individual exit rates are affected by labour market conditions, say by \(\lambda\), which can be thought as an index describing the state of the economy. Then, in case of pure state-dependence the individual exit rates are given by \(\lambda \pi_t h(t)\) and in case of pure heterogeneity by \(\lambda \pi_{it}\). It can be shown (see Appendix\(^\text{15}\)), that the overall exit rate \(\pi\) is given by the following for the two cases respectively:

\(^{13}\) See also Franz (1987).


\(^{15}\) Also, see e.g. Kiefer (1988), who gives a thorough survey of the theoretical foundations of the analysis of duration data and exit probability.
(4.1) \( \pi = \lambda \Phi(R) \)

in the case of state-dependence, where \( R \) is the proportion of long-term unemployed of total unemployed and

(4.2) \( \pi = \lambda \tilde{\pi} \)

in the pure heterogeneity case, where \( \tilde{\pi} \) is a weighted mean of individual exit probabilities, which is a constant and independent of duration-structure.

Assume now that \( \lambda \) is a function of relevant labour market variables, say \( \lambda = \lambda(U/L, V/N) \), where \( V \) is vacancies, and substitute this in (4.1) and (4.2) and use (3) to give:

\[
(5.1) \quad \frac{U}{N} \lambda \left( \frac{U}{L}, \frac{V}{N} \right) = \frac{s}{\Phi(R)}
\]

\[
(5.2) \quad \frac{U}{N} \lambda \left( \frac{U}{L}, \frac{V}{N} \right) = \frac{s}{\tilde{\pi}}
\]

or in terms of unemployment and vacancy rates \( (u = U/L, \nu = V/N) \):

\[
(6.1) \quad \frac{u}{1-u} \lambda (u, \nu) = \frac{s}{\Phi(R)}
\]

\[
(6.2) \quad \frac{u}{1-u} \lambda (u, \nu) = \frac{s}{\tilde{\pi}}
\]

These determine the U-V-relationship, which in case of pure heterogeneity (6.2) is unaffected by the duration-structure \( (s/\tilde{\pi}) \) is constant. In the case of state-dependence (6.1), the U-V-curve shifts, if the proportion of long-term unemployment changes. As shown in the appendix, \( \Phi'(R) < 0 \), if there is negative state-dependence, i.e. the U-V-curve shifts outwards, if \( R \) increases (and vice versa for positive state-dependence). So the basic idea for the test of state-dependence is quite simple: if it can be shown that the share of long-term unemployed is a relevant variable in a U-V-relationship, there must be state-dependence effects in unemployment. If the share of long-term unemployed causes an outward shift of the U-V-curve, this implies negative state-dependence.

The result of negative state-dependence can be given different intuitive interpretations. First, the long-term unemployed might get discouraged by their long unemployment experience (e.g. after several unsuccessful job applications) and this can reduce their further search activity. If employers use long-term unemployment as a screening device, then the long-term unemployed are rejected more often than the short-term unemployed. In both of these cases we would expect that a given number of vacancies is associated with higher unemployment if the share of long-term unemployment increases. Furthermore, in case individual exit rates vary with time, a temporary rise in unemployment inflow will induce and increase of steady state long-term unemployment after a time period.\(^{16}\) So, an increase in long-term unemployment can be simply explained by the increase of unemployment itself.

5. The evidence on movements of the Finnish U-V-curve

Our next step is to estimate a U-V-curve equation of quarterly data from 1968 q2 until 1986 q3. The static version of the basic relationship to be estimated in log linear form is:

\[
(7) \quad \log u = \alpha_0 + \alpha_1 \log \nu + \alpha_2 \log R
\]

where \( u \) is the unemployment rate \((= U/L)\) and \( \nu \) is the vacancy rate \((= V/N)\), \( L = N + U \) is the labour force. \( R \) is the ratio of long-term unemployed against total unemployment.\(^{17}\) Our data on persons who are unemployed is quite reliable, a minor change in reporting occurred in 1980, when those who had been temporarily laid off, were partly excluded from the unemployment statistics (i.e. duration data does not include temporary layoffs).\(^{18}\) There was also another change in reporting in 1980, which might have had a larger effect on our estimates, namely the exclusion of persons receiving unemployment

\(^{16}\) Budd, Levine and Smith (1986a) p. 17—18, 22; see also footnote 27 below.

\(^{17}\) The ratio of long-term unemployed is the ratio of those who have been unemployed for 6 months or more to the total number of unemployed.

\(^{18}\) This exclusion of temporary layoffs does not prevent us from calculating the exact total number of unemployed, but in order to correct our data concerning long-term unemployed, we have to make the assumption that all temporary layoffs are short-term. As temporary layoffs seldom exceed 3 months, this does not affect our results.
pension. This means that the reported proportion of long-term unemployment in the 1970s is overestimated compared with the end of the estimation period. We try to capture this change in reporting by introducing a dummy variable in connection to the long-term unemployment ratio. 19

Earlier results 20 indicate that the enforcement of the Employment Act in 1972 had considerable effects on the basic U-V-relationship. We are going to investigate, if these results hold in our approach also, by including a shift dummy in 1972 in our basic equation.

The vacancies used here are those which have been reported by the Employment Service; this data covers approximately 25—35 % of all vacancies. 21 There is an apparent source for measurement error in using the reported vacancy data, if the proportion of reported vacancies varies over the estimation period as has been suggested by some authors. 22 Unfortunately, any corrections suggested in the literature 23 to deal with varying notification of vacancies cannot be applied in the Finnish case. This is because we have no data on engagements in the public sector or in jobs, which are short term and under LEL-pension system, 24 and, especially, because we do not have flow data on notified vacancies before the 1980s. There are also changes in the public sector norms of notification of vacancies at the Employment Service (one in 1975 and another one in 1978), 25 which might have affected the notification rate. This study, as all the previous studies in Finland, therefore suffers from the uncertainty about the development of true vacancies in Finland.

The method of estimation applied is an instrumental variables estimation 26 for the equation above. The instruments used are three lags of the vacancy rate and the proportion of firms operating at full capacity level (FFC) as an exogenous demand shock variable. To check for the choice of instruments an ordinary least squares estimation is applied, which shows that 94 % of variation in log v, is explained by the chosen instruments. In all estimated equations quarterly seasonal dummies (Q1, Q2 and Q3) are included (see table 2, where estimation results are given).

Equation 1, reported in the first column, shows the basic U-V-relationship, when the ratio of long-term unemployed is not included in the estimated equation. We find that there is a dynamic structure in our equation which includes the dependent variable lagged three times. The relationship between unemployment and vacancies does not seem to be very well determined. The steady state elasticities for the independent variables are reported at the bottom of the table.

After some experimentation we found that we have two alternative ways to represent the changes in the relationship over time. If we include a time trend in the equation, this improves the estimation result considerably (see equation 2). The trend is significant and positive and the relationship between unemployment and vacancies is now significantly negative. On the other hand, if we replace the trend with a simple step dummy variable D72 (D71 equal to 1 as from 1972), the resulting equation 3 gives almost as good fit as equation 2. As can be seen from equation 3, the dynamic structure of our model is modified: only one lag of the dependent variable are found to be significant in the presence of D72.

We tried also include the time trend and the dummy variable for 1972 in the same equation, which results in a highly significant positive coefficient for the dummy and the time trend loses its significance (this equation is not reported). As this result confirms the earlier

19 A more satisfactory way to deal with people who are registered as unemployed but are actually waiting to qualify for unemployment pension, would be to exclude them from the analysis. (This group of people is now included in the number of long-term unemployed for, at least, the duration category of 6—9 months, since only persons with an unemployment spell exceeding 200 days are entitled to early retirement.) It is, however, impossible to identify this group because we would need flow data for people retiring on the grounds of unemployment, and such data, unfortunately, is not readily available.

20 Eriksson (1985) and Sauramo and Solttila (1983), which, however, are based on different specification of the U-V-curve.


24 Moreover, the data on the number of engagements under TEL-pension system in the private sector is only available in annual form.


26 We have made use of the software package »PC-GIVE» in our estimations.
Table 2. The estimated equations for the U-V-curve.

Instrumental variables estimation
Estimation period: 1968 q2—1986 q3
Dependent variable: log \( u_l \)
Instruments: log \( v_{-1} \), log \( r_{-2} \), log \( v_{-3} \), FFC
t-ratios: in parentheses below the estimated coefficients

<table>
<thead>
<tr>
<th>Independent variable</th>
<th>EQ1</th>
<th>EQ2</th>
<th>EQ3</th>
<th>EQ4</th>
<th>EQ5</th>
<th>EQ6</th>
</tr>
</thead>
<tbody>
<tr>
<td>log ( v_l )</td>
<td>-0.0718</td>
<td>-0.3097</td>
<td>-0.2866</td>
<td>-0.2468</td>
<td>-0.3836</td>
<td>-0.4889</td>
</tr>
<tr>
<td></td>
<td>(1.40)</td>
<td>(5.91)</td>
<td>(3.90)</td>
<td>(4.15)</td>
<td>(5.82)</td>
<td>(6.86)</td>
</tr>
<tr>
<td>log ( u_{l-1} )</td>
<td>1.2736</td>
<td>0.8034</td>
<td>0.7746</td>
<td>1.0823</td>
<td>0.7705</td>
<td>0.5916</td>
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<tr>
<td></td>
<td>(10.04)</td>
<td>(6.75)</td>
<td>(5.28)</td>
<td>(9.89)</td>
<td>(5.90)</td>
<td>(4.32)</td>
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<tr>
<td>log ( u_{l-2} )</td>
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<td>-0.4432</td>
<td>-0.2300</td>
<td>-0.1588</td>
</tr>
<tr>
<td></td>
<td>(3.82)</td>
<td>(3.71)</td>
<td>(0.67)</td>
<td>(4.25)</td>
<td>(2.11)</td>
<td>(1.51)</td>
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<tr>
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<td></td>
<td>(3.10)</td>
<td>(3.21)</td>
<td>(5.97)</td>
<td>(5.97)</td>
<td>(6.60)</td>
<td>(6.60)</td>
</tr>
<tr>
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<td>0.2722</td>
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<td>0.2987</td>
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<td></td>
<td>(4.47)</td>
<td>(4.47)</td>
<td>(5.74)</td>
<td>(5.74)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>log ( R_{l-2} )</td>
<td>0.2452</td>
<td>0.2452</td>
<td>0.1695</td>
<td>0.1325</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(5.02)</td>
<td>(5.02)</td>
<td>(3.59)</td>
<td>(3.59)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>D80 * log ( R_{l-2} )</td>
<td>0.0257</td>
<td>0.0257</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(2.74)</td>
<td>(2.74)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

| Intercept            | 0.1560 | 0.1560 | 0.2286 | 0.1939 |
|                      | (2.09)  | (2.09)  | (1.85)  | (1.85)  |

s.e.       0.1151 | 0.0871 | 0.0986 | 0.0998 | 0.0880 | 0.0824 |
\( R^{2} \) 0.9707 | 0.9745 | 0.9596 | 0.9641 | 0.9668 | 0.9702 |
DW        1.623 | 1.667 | 1.680 | 1.768 | 1.625 | 1.691 |
LM(1) ~ \( \chi^{2} \) 16.10 | 6.65 | 12.69 | 2.95 | 8.41 | 3.91 |
LM(4) ~ \( \chi^{2} \) 12.67 | 4.38 | 9.05 | 4.59 | 7.95 | 4.61 |
\( z_1 - \chi^{2} \) 8.60 | 1.98 | 1.76 | 2.51 | 1.33 | 1.73 |
\( z_2 - \chi^{2} \) 1.08 | 2.03 | 2.92 | 4.37 | 2.79 | 9.98 |
FORE(19) 0.90 | 4.86 | 2.32 | 0.91 | 2.17 | 1.75 |
CHOW(19) 0.53 | 0.79 | 1.73 | 0.55 | 1.33 | 0.92 |

Steady state elasticities:

<table>
<thead>
<tr>
<th>EQ1</th>
<th>EQ2</th>
<th>EQ3</th>
<th>EQ4</th>
<th>EQ5</th>
<th>EQ6</th>
</tr>
</thead>
<tbody>
<tr>
<td>log ( v )</td>
<td>-1.2088</td>
<td>-0.7234</td>
<td>-0.9640</td>
<td>-0.6856</td>
<td>-0.8348</td>
</tr>
<tr>
<td>time</td>
<td>0.0418</td>
<td>0.0003</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( D72 )</td>
<td>0.9156</td>
<td>0.6794</td>
<td>0.6794</td>
<td>0.4662</td>
<td>0.5267</td>
</tr>
<tr>
<td>log ( R_{\Delta-10} )</td>
<td>0.3689</td>
<td>0.3689</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>log ( R_{\Delta-16} )</td>
<td>0.2790</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

\( a \) This refers to the estimated reduced form of equation.
\( b \) Test for the validity of instruments.
\( c \) Test for the normality assumption.
results mentioned above, and we have, indeed, a very good motivation for including the dummy variable in our equation, we choose to continue on this line.

We now proceed by including the ratio of long-term unemployed in the U-V-equation, which appears with two lags in our relationship. Again, the dynamic structure of the relationship is affected: in the presence of log $R_{t-2}$ (equations 4 and 5) two lags of the dependent variable are found to be significant.

The two equations (4 and 5) are reported in order to make clear the significance of long-term unemployment ratio in this equation. In equation 4, the long-term unemployment ratio is the only additional regressor, and it turns out to be highly significant. When we include the dummy variable for 1972 in the equation, both variable coefficients are significant and positive, but the effect of long-term unemployment is now reduced. Naturally, the increase in long-term unemployment might have at least partly coincided with the change in the unemployment benefit system in 1972. Introducing the dummy does, however, improve the explanatory power of the equation, so that equation 5 is to be preferred to equation 4.

But we can do even better than this by trying to capture the possible change in the reported long-term unemployment ratio in 1980, due to expansive use of unemployment pensions. The long-term unemployment ratio multiplied with a dummy ($D80 = 1$ as from 1980) turns out to be significant in equation 6. The effect of long-term unemployment of

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27 In the presence of a time trend, the long-term unemployment ratio, however, turns out to be insignificant in the equation. Note, however, that in addition to the secular increase there are also considerable fluctuations in long-term unemployment. We have also estimated an equation for the share of long-term unemployed, which confirms the positive time trend. The estimated equation is ($\hat{R}$-ratios in parentheses after the estimated coefficient): $R_t = -0.60 (0.62) + 0.50 R_{t-1} (7.34) - 1.75 u_t (2.94) + 3.22 u_{t-1} (4.97) + 4.24 D72 (4.26) + \text{seasonals}$

$R^2 = 0.954, \text{s.e.} = 2.17, LM(1) - \chi(1) = 0.146$

The fluctuations of the proportion of long-term unemployed can be traced back to the fluctuations of unemployment rate (unemployment rate at time $t$ and the first lag of it are significant in the equation). This equation makes good sense in the light of earlier discussion: as unemployment increases (inflow increases) we observe decline in the ratio of long-term unemployed, but in the longer run, long-term unemployment increases with total unemployment.

---

the U-V-relationship is now decomposed in two periods, 1968—79 and 1980—86. We can see that the long-term unemployment ratio has shifted the U-V-curve even more outwards in the 1980s. This result might imply that the share of long-term unemployed has increased partly because of the higher proportion of persons waiting to qualify for unemployment pension in the 1980s.28

All the reported equations are stable over time, as the reported FORE(19) and Chow-test statistics based on splitting the sample at the end of 1981 show clearly. In addition to Durbin-Watson test against autoregressive errors (which in the presence of lagged dependent variable is biased), two Lagrange Multiplier test statistics are reported. The first one tests the null hypothesis of zero first order autocorrelation (LM(1)) and the second one tests no autocorrelation hypothesis up to fourth order (LM(4)). Only in the case of equation 1, the first order autocorrelation of residuals can be detected. In all other cases, the null hypothesis of no autocorrelation cannot be rejected.

We find that the steady state properties of all reported equations are quite similar to each other. In the steady state the U-V-relationship is less than proportional in all cases (except equation 1). The steady state elasticity of log $R$ varies along the line we would expect from the estimated equations: from 0.6794 in equation 4 to 0.2336 in 1968—79 and 0.2790 in 1980—86 in equation 6. Figure 4 shows the steady state U-V-curves drawn at different times for different mean levels of long-term unemployment based on the estimated equation 6, which is our preferred equation.29 This shows that the estimated equation fits quite well to the observed pattern of U-V-relationship both in the 1970s and in the 1980s.

Our estimates show clearly that the U-V-curve has shifted outwards during the estimation period. We find support for the hypothesis that the proportion of long-term unemployed, which has increased since the early 70s, has been a major factor behind these shifts. Our findings correspond well to results presented by Budd, et al. (1986a), and their

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conclusions of the ratio of long-term unemployed being a significant labour market variable, which affects the U-V-curve through the special characteristics of the long-term unemployed, i.e. the long-term unemployed have lower search intensity and/or they are discriminated by employers, as a result of the unemployment experience. These results can be interpreted in the light of the theory presented above as rejecting the pure heterogeneity hypothesis, i.e. that there are at least some state-dependence effects in unemployment.

6. Conclusions

In this paper we have discussed long-term unemployment in Finland since the late 1960s. The major aspect of interest has been the hypothesis of negative state-dependence, i.e. whether unemployment experience creates more unemployment. Unfortunately, there is no straightforward way of testing this hypothesis, mainly because of the lack of suitable data. Earlier results based on cross-section data suffer from the »problem of unobservables» and consequently, the interpretation of such tests can always be questioned.\(^\text{30}\)

Recently, tests of state-dependence, which rely on the aggregate time series behaviour, have been suggested. One test proposed by Jackman & Layard (1988) concentrates on the observed overall exit rate under the hypothesis of pure heterogeneity of unemployed contrasted with a hypothesis of some state-dependence. To be able to use their test we have to have access to a sufficiently long time-series on outflows from unemployment at different states of economy. In the Finnish case we failed to create such a long series, which would include distinct different states of the economy. This makes it impossible to utilize this test.

The test, originally proposed by Budd, et al. (1986a) that we have used, is based on the notion that if there are differences in exit rates between the short-term and the long-term unemployed, and if an increase of long-term un-

employment shifts the U-V-curve outwards, this must be due to some negative state-dependence. We tried to establish a connection between the ratio of long-term unemployed and movements of the Finnish U-V-curve. The results are promising. In particular, we find that our estimated equations, which include the share of long-term unemployed as an explanatory variable, fit well into the actual behaviour of the U-V-relationship during the estimation period.

There is an alternative way of explaining the shifts of the U-V-relationship by including a positive time trend in the equation, but we prefer the equations which include the ratio of long-term unemployed because of the economic interpretation. (To explain the shifts of the U-V-curve simply with a positive time trend is to say: the U-V-curve has shifted because of shifts over time.)

In this study we have not considered a number of other possible variables (in addition to the ratio of long-term unemployed and two dummy variables), which may have contributed to the shift of U-V-relationship in Finland during the estimation period. Such variables that might be included in the estimated U-V-function are, for instance: some measure of structural mismatch between the unemployed and vacancies; replacement ratio; share of young or old in the labour force; and the separation rate from employment. However, we do not think these factors have changed sufficiently to lead to profound changes in our results. A closer analysis of other factors influencing the steady state U-V-relationship is, however, left for further research.

Also, on the basis of this type of U-V-curve analysis, it is not possible to distinguish between different reasons for state-dependence in unemployment, i.e. is it so that unemployed persons are demoralized or discouraged by the unemployment experience, so that their search activity is reduced, or is it so that employers discriminate against job applicants with long-term unemployment experience. Unfortunately, there is no straightforward solution to this question as long as we do not have any data on firms' hiring and/or firing behaviour.

Our results indicate that the increasing long-term unemployment has been a major factor in shifting the U-V-relationship outwards, along with a major shift of the relationship in 1972. This, together with the fact that the exit probability of long-term unemployed is clearly lower than the short-term unemployed exit rate, would then indicate that there are, indeed, negative state-dependence effects in unemployment.

Data

Unemployment U
The number of unemployed in the Finnish Employment Service unemployment registers

Vacancies V
Vacancies registered by the Employment Service

Employment N
Total number of employed

Labour force L = N + U

Long-term unemployment ratio R
The number of people who have been unemployed for more than 6 months as a proportion of the total number of unemployed

FFC
The proportion of firms operating at full capacity level

D72
Dummy-variable 0, if ≤1972; 1, if >1972

D80
Dummy-variable 0, if ≤1980; 1, if >1980

Appendix

Deriving the overall exit rate and the U-V-curves in the two cases

A. The overall exit rate: pure heterogeneity case

Assume that we have two different types of individuals entering unemployment with two different exit probabilities \( \pi_1 \) and \( \pi_2 \) according to pure heterogeneity hypotheses.

Let \( \lambda \) be an index of the state of economy which affects the individual exit rates. Then the overall exit rate \( \pi \) can be expressed in terms of these two category exit probabilities \( \lambda \pi_1 \) and \( \lambda \pi_2 = \lambda \theta \pi_1 \). (The distribution of \( \pi_1 \) is assumed to be constant over time.)

\(^{31}\) c.f. Budd, Levine and Smith (1986a).
Suppose that proportion \( \gamma \) of total \( M \) entrants have exit probability \( \pi_1 \). We then have in the steady state, which is characterized by \( I = O \), i.e. inflow to unemployment equals outflow from unemployment:

\[
(1.A) \quad \pi = \frac{O}{U} = \frac{I}{U} = \frac{M}{\lambda \pi_1 + \frac{1 - \gamma}{\theta}} \pi_1 = \lambda \pi
\]

where \( \pi \) is a weighted mean of \( \pi_n \), which is a constant and does not depend on the duration-structure of unemployment.

**B. The overall exit rate: pure state-dependence case**

Define a survival function \( S(t) \) for unemployed as:

\[
(1.B) \quad S(t) = 1 - G(t) = \int_0^t g(\tau)d\tau
\]

where \( G(t) \) is the distribution of completed unemployment spells.

Then the time dependent hazard function (the hazard or probability of leaving unemployment) is given by:

\[
(2.B) \quad h(t) = \frac{g(t)}{S(t)}
\]

where \( g(t) = -S'(t) \), which implies:

\[
(3.B) \quad h(t) = \frac{-S'(t)}{S(t)} = \frac{d \log S(t)}{dt}
\]

Integration of both sides of (3.B) gives:

\[
(4.B) \quad \int_0^t h(\tau)d\tau = - \int_0^t D \log S(\tau)d\tau
\]

\[\Rightarrow \log S(t) = - \int_0^t h(\tau)d\tau\]

\[\Rightarrow S(t) = \exp \left[ - \int_0^t h(\tau)d\tau \right]\]

which gives the connection between survival function and hazard function.

The expected mean duration \( \bar{d} \) of unemployment spells is given by:

\[
\bar{d} = \int_0^\infty \int_0^t g(t)d\tau dt
\]

but since

\[
\int_0^\infty S(t)dt = \int_0^\infty (1 - G(t))dt = \int_0^\infty t(1 - G(t))dt + \int_0^\infty t g(t)dt
\]

so long as \( t(1 - G(t)) \to 0 \), as \( t \to \infty \), we can write:

\[
(5.B) \quad \bar{d} = \int_0^\infty S(t)dt
\]

Let us now introduce a simple step hazard function in accordance to the state-dependence hypothesis:

\[
(6.B) \quad \lambda h(t) = \begin{cases} \lambda h_1, & \text{if } t < T \\ \lambda h_2, & \text{if } t \geq T \end{cases}
\]

which indicates a) positive state-dependence if \( h_1 < h_2 \); b) negative state-dependence if \( h_1 > h_2 \); and c) constant exit probability if \( h_1 = h_2 \).

Then the survival function is:

\[
(7.B) \quad S(t) = \exp \left[ - \int_0^t \lambda h(\tau)d\tau \right]
\]

\[
= \begin{cases} \exp \left[ - \int_0^t \lambda h(\tau)d\tau \right], & \text{if } t < T \\ \exp \left[ - \int_0^T \lambda h(\tau)d\tau - \int_0^t \lambda h(\tau)d\tau \right], & \text{if } t \geq T \end{cases}
\]

\[
= \begin{cases} \exp \left[ -\lambda h_1 t \right], & \text{if } t < T \\ \exp \left[ -\lambda h_1 T - \lambda h_2 (t - T) \right], & \text{if } t \geq T \end{cases}
\]

and the mean duration of unemployment is:

\[
(8.B) \quad \bar{d} = \int_0^\infty S(t)dt = \int_0^T S(t)dt + \int_T^\infty S(t)dt
\]

\[= \frac{1}{\lambda h_1} (1 - e^{-\lambda h_1 T}) + \frac{1}{\lambda h_2} (e^{-\lambda h_2 T}) \]
The overall exit rate $\pi$ is by definition:

\[
\pi = \frac{1}{d} = \lambda \left[ \frac{1}{h_1} (1 - e^{-\lambda h_1 T}) + \frac{1}{h_2} (e^{-\lambda h_2 T}) \right]^{-1}
\]

where

\[
\begin{align*}
\omega'(\lambda) &> 0 \quad \text{if } h_1 < h_2 \\
\omega'(\lambda) &< 0 \quad \text{if } h_1 > h_2 \\
\omega'(\lambda) &= 0 \quad \text{if } h_1 = h_2
\end{align*}
\]

Let us now characterize all durations longer than $T$ as long-term unemployment. Then, in the steady state the proportion of long-term unemployed of total unemployment $R$ is given by:

\[
R = \frac{\int_{0}^{\infty} S(t) dt}{\int_{0}^{\infty} S(t) dt} = \frac{1}{h_2} \left( e^{-\lambda h_1 T} \right)
\]

\[
= \frac{1}{h_1} (1 - e^{-\lambda h_1 T}) + \frac{1}{h_2} (e^{-\lambda h_2 T})
\]

\[= f(\lambda), f'(\lambda) < 0 \]

Then the inverse function of $f(\lambda)$ gives us:

\[
\lambda = \Psi(R), \Psi'(R) < 0
\]

by the inverse function rule for derivatives. Substituting (11.B) into (9.B) gives:

\[
\pi = \lambda \omega(\Psi(R)) = \lambda \Phi(R)
\]

where

\[
\begin{align*}
\Phi'(R) &> 0 \quad \text{if } h_1 < h_2 \\
\Phi'(R) &< 0 \quad \text{if } h_1 > h_2 \\
\Phi'(R) &= 0 \quad \text{if } h_1 = h_2
\end{align*}
\]

C. The steady state $U$-$V$-curve in the two cases

Assume that $\lambda$ is a function of relevant labour market variables, say $\lambda = \lambda \left( \frac{U}{L}, \frac{V}{N} \right)$, where $U = \text{unemployment}$, $V = \text{vacancies}$, $N = \text{employment}$ and $L = N + U$ is labour force.

Steady state in the labour market is characterized by $I = O$, i.e. inflow into unemployment equals outflow from unemployment, which enables us to write:

\[
\frac{U}{N} = \frac{\frac{1}{N}}{\frac{U}{\lambda}} = \frac{s}{\pi}
\]

where $s$ is the overall inflow rate and $\pi$ is the overall outflow rate. Assume now a constant inflow rate.

Then in the case of pure heterogeneity using (1.A) we can write:

\[
\frac{U}{N} \lambda \left( \frac{U}{L}, \frac{V}{N} \right) = \frac{s}{\pi}
\]

or in terms of unemployment rate ($u = \frac{U}{L}$) and vacancy rate ($v = \frac{V}{N}$):

\[
\frac{u}{1-u} \lambda(u, v) = \frac{s}{\pi}
\]

(2.C) determines an $U$-$V$-relationship, which is independent of duration-structure, i.e. changes in the duration-structure do not shift the $U$-$V$-curve.

In the case of pure state-dependence, the $U$-$V$-curve is given by (using (12.B) and (1.C)):

\[
\frac{U}{N} \lambda \left( \frac{U}{L}, \frac{V}{N} \right) = \frac{s}{\Phi(R)}
\]

or in terms of unemployment and vacancy rates:

\[
\frac{u}{1-u} \lambda(u, v) = \frac{s}{\Phi(R)}
\]

(3.C) This relationship depends on the ratio of long-term unemployed ($R$). In case of positive state-dependence, $\Phi'(R) > 0$, the $U$-$V$-curve shifts inwards, if $R$ increases. If $\Phi'(R) < 0$, i.e. individual exit probabilities exhibit negative state-dependence, an increase in the share of long-term unemployed shifts the $U$-$V$-curve outwards.
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