SEARCH BEHAVIOUR, TRANSITIONS TO NON-PARTICIPATION AND THE DURATION OF UNEMPLOYMENT*

Gerard J. van den Berg

In this paper we examine the estimation of a structural job search model using data on individual unemployment durations. The model allows for transitions from unemployment to non-participation. In an extended version of the model we deal with the influence of on-the-job search and prospective wage increases on search behaviour of the unemployed.

In empirical studies on unemployment duration the reduced-form approach, in which only hazards of the duration distribution are estimated (see, for example, Lancaster, 1979) seems to be replaced gradually by a structural approach. The latter way of modelling is characterised by the explicit use of the framework of job search theory in empirical analysis. The results from such analyses can be used for inferences about the behaviour of the unemployed. In particular a distinction can be made between choice and chance components of the transition rate into employment.

Several empirical studies using structural search models have been published (see, for example, Yoon, 1981; Lancaster and Chesher, 1983; Lynch, 1983; Narendranathan and Nickell, 1985; Ridder and Gorter, 1986; Wolpin, 1987) some of which use a very restricted model specification (notably the first three references). None of those studies uses a model that allows for transitions from unemployment to non-participation. In reality an individual who is unemployed and actively searching for a job may drop out of the labour force, at some point of time during unemployment. It may be that the papers referred to do not take account of transitions into non-participation because the data used are not rich enough to make the distinction between the states of unemployment and non-participation. This can be the case if the data collection is based on the receipt of unemployment benefits. Another cause for not taking account of such transitions may be that in the time those data were collected (typically the seventies) the occurrence of such transitions was less prominent. However, by now there is much evidence that a large portion of the flow out of unemployment consists of transitions into non-participation (for a survey of the literature, see Micklewright, 1988, who also forcefully argues that the state of non-participation should be incorporated in duration models of the labour market, especially if one is interested in the effects of benefits on

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unemployment duration). In the sample we use, almost 30% of all spells of unemployment ends up in a transition into non-participation. Therefore we estimate a structural job search model that allows for such transitions.

Further, up to now the structural models used in empirical analyses do not take into account that wage increases during employment may be expected. Wages can increase for several reasons such as accumulation of human capital or transitions from jobs with lower wages to jobs with higher wages without intervening spells of unemployment (on the job search, see, for example, Mortensen, 1986). The optimal strategy of an unemployed individual is likely to be dependent on changes of wages and jobs that occur after the acceptance of a job. We estimate an extended version of the model, which deals with these aspects.

In Section I we discuss the specification of the model. As a starting point we take a search model that resembles Narendranathan and Nickell's (1985) model. This model is extended to allow for transitions into non-participation. We outline how the model may be given an alternative interpretation which is more realistic with regards to the process of search. This interpretation allows for knowledge of the wage rate associated with a vacancy before one responds to that vacancy, i.e. before the job is actually offered. Section II contains a description of the data and a discussion of the empirical implementation of the model. The structural model is estimated by ML using the Newton–Raphson algorithm. The estimation method is analogous to that used by Narendranathan and Nickell (1985) and Wolpin (1987) in the sense that for every individual in the sample, for every iteration, the optimal search strategy, which follows from a dynamic programming problem, needs to be solved. Section III deals with the estimation of the wage offer distribution. Section IV gives the main results. We present estimates of the job offer arrival rate, the transition rate into non-participation and the utility function. For distinct age categories and levels of education we present sample averages of the main characteristics of the job search process. From a policy viewpoint it may be of interest to see whether a decrease in unemployment benefits has any influence on duration. If not, this may lead to a re-evaluation of benefits as a policy tool. Therefore we give special attention to the effects of changes in benefits on the reservation wage and the expected duration. In Section V the construction of the extended model is described and the results of the estimation of the extended model are discussed. Section VI concludes.

I. THE MODEL

I.1. Job Search Theory and Model Specification

We start by presenting the basic job search model for unemployed individuals who are searching sequentially for jobs until a suitable one has been found (for surveys on job search theory, see Mortensen, 1986 or McKenna, 1985). Job offers arrive randomly in time at the arrival rate \( \lambda \). Such job offers are random drawings from a wage offer distribution \( F(w) \). During unemployment a benefit \( b \) is received. The variables \( \lambda \), \( b \) and \( w \) are measured per unit time period.
Unemployed individuals aim at maximisation of their expected discounted lifetime utility (over an infinite horizon). To begin with, we also assume that once a job is accepted it will be held forever at the same wage.

The per-period utility function is a separable function of two arguments, income and state:

\[
\text{utility}(\text{income} = x, \text{state} = \text{employment}) = v^* u(x)
\]

\[
\text{utility}(\text{income} = x, \text{state} = \text{unemployment}) = vu(x).
\]

This utility function was first used by Nickell (see Lancaster and Cheshire, 1983). The function \( u \) is increasing in its argument and may take account of risk aversion. We normalise by setting \( v^* = 1 \). Somewhat loosely we call \( v \) the disutility of unemployment.

In the sequel only stationary job search models are considered. This means that we take \( \lambda, b, u, v \) and \( F(w) \) to be independent of unemployment duration and calendar time and independent of all events during unemployment. Obviously this is not very realistic. The level of unemployment benefits depends generally on the elapsed duration of unemployment. The job offer arrival rate may decrease during unemployment as a result of the stigma that the long-term unemployed may have. Further, \( \lambda, b \) and \( F(w) \) may change due to business cycle effects. The motivation for adopting the stationarity assumption is basically the same as it was in the other empirical studies using structural search models (see, for example, Lancaster and Cheshire, 1983 and Narendranathan and Nickell, 1985). That is, when estimating a non-stationary model the computational difficulties are likely to be even more burdensome, so it seems a good strategy to start off with a stationary model. (For an analysis of non-stationarity in job search theory, see van den Berg (1990).) In Section IV we turn to the effects that the presence of non-stationarity might have on the estimation results.

The optimal strategy of an unemployed individual in the model sketched above can be characterised by a fixed reservation wage \( \phi \). A job offer is accepted if its wage exceeds \( \phi \) while a wage that is smaller than \( \phi \) induces one to reject the offer and search for a better one. The transition rate from unemployment into employment \( \theta \) can be written as the product of the job offer arrival rate and the conditional probability of accepting a job offer.

\[
\theta = \lambda F(\phi), \quad F = 1 - F.
\] (1)

In reality an individual who is unemployed and actively searching for a job may drop out of the labour force, at some point in time during unemployment. This may be the result of a personal decision, such as deciding to dedicate all his time to household activities. It can also be a forced transition, for example, when he is conscripted or when he becomes disabled or when he retires. All these cases can be labelled as transitions out of unemployment into non-participation.

Flinn and Heckman (1982) present a three-state structural search model which could serve as a starting point for our model. In this three-state model
the distribution of returns of non-participants enters the equations that describe the behaviour of the unemployed. This implies that data on returns of non-participants are needed in order to estimate the model. Such data are not available. Therefore we adopt a reduced-form model of the transitions from unemployment into non-participation. Specifically, such transitions are assumed to occur according to a Poisson process with a parameterised transition rate $\zeta$.

The optimal strategy of an unemployed individual depends on the expected utility of becoming a non-participant. If the latter is high with respect to the expected utility of becoming employed then it is optimal to accept a job offer only if the wage corresponding to it is very high. Let $x$ denote the flow of income of a non-participant. We make the assumption

$$Eu(x) = u(b).$$

(2)

For a lot of cases the income flow after becoming a non-participant is close to the benefit level (for example, when an unemployed individual becomes disabled, when he retires, when he is conscripted, when he returns to school and applies for social assistance). If the dispersion of the distribution of $x$ is small, which we expect to be the case, then $E_x \approx b$ implies that $Eu(x) \approx u(b)$. To sum up, we do not assume anything about the distribution of the income flow $x$ in the state of non-participation except that equation (2) holds. In addition, we assume that the state of non-participation is absorbing and, for the moment, we assume that the non-pecuniary component of per-period utility in non-participation is the same as that in unemployment. As an additional condition for stationarity to hold we require that $\zeta$ is constant (though possibly different across individuals). Again this may not be very realistic. Individuals may enter non-participation at an increasing rate when they become discouraged about their chances on the labour market. This in turn may happen more frequently among the long-term unemployed.

Van den Berg (1988) proves that the reservation wage $\phi$, which characterises the optimal strategy in the model, satisfies the following equation, in which $\rho$ denotes the subjective rate of discount

$$u(\phi) = vu(b) + \frac{\lambda}{\rho + \zeta} \int_{\phi}^{\infty} [u(w) - u(\phi)] dF(w).$$

(3)

The exit rate out of unemployment is equal to the sum of $\theta$ and $\zeta$, with $\theta$ given by equation (1). Because $\theta$ and $\zeta$ do not depend on duration or on time or on events during unemployment this implies that the unemployment duration has an exponential distribution with parameter $\theta + \zeta$.

I.2. An Alternative Interpretation

It can be argued that the modelling of the search process so far is not very realistic. Generally one knows the wage rate associated with a vacancy before one responds to that vacancy, i.e. before the job is actually offered. Narendranathan and Nickell (1985) constructed a search model that deals with
this. Job vacancies arrive according to a Poisson process with arrival rate \( q_1 \). A vacancy is characterised by a random drawing from a distribution of wages associated with the flow of vacancies, \( G(w) \). The decision whether to apply or not is made with knowledge of the wage corresponding to the vacancy. If one does apply, then there is a (known) probability of \( q_2(w) \) that the job will actually be offered. The dependence of \( q_2 \) on \( w \) represents increased competition for vacancies with higher wages.

It is straightforward to show that the model developed in I.1 is equivalent to the model described here. To see this, set 

\[
\lambda = q_1 \int_0^\infty q_2(\omega) \, dG(\omega) \tag{4}
\]

and 

\[
F(w) = \frac{\int_0^w q_2(\omega) \, dG(\omega)}{\int_0^\infty q_2(\omega) \, dG(\omega)}. \tag{5}
\]

Consequently, the estimation results of the original model can be reinterpreted according to equations (4) and (5). Narendranathan and Nickell (1985) make the convenient assumption that 

\[
q_2(w) = q_3(w) \, q_4, \tag{6}
\]

in which \( q_3 \) depends on \( w \) only, while \( q_4 \) represents the dependence of \( q_2 \) on personal characteristics. If (6) holds then \( F(w) \) in (5) does not depend on \( q_4 \), i.e. does not depend on personal characteristics which influence the probability that the job is offered given application.

II. THE DATA

II.1. The Data Set

The data set used is constructed from the Netherlands Socio-Economic Panel, a survey conducted by the Netherlands Central Bureau of Statistics. Since April 1984 a random sample of about 12,000 individuals has been interviewed twice a year (in April and October). At every interview except the first one, respondents were asked to recall their labour market history for the past 6 months, that is, they were asked between which dates in the last 6 months they had a job, between which dates they were unemployed and searching for a job and between which dates they were doing something else. The latter category of activities includes being disabled, doing unpaid work in the household, being retired (the retirement age varies between (roughly) 55 and 70 years and centres on 65 years), being in full-time training, being conscripted and just doing nothing. At the first interview the observation period is extended to 12 months. Given present information we have labour market histories for 2·5 years, from May 1983 up to October 1985.

For our purposes we selected 223 men aged between 17 and 65, who reported
that at the moment of the first interview (April 1984) their main activity was being unemployed and searching for work. We determined for how long they were unemployed and searching for work at that moment, and (using subsequent waves) also for how long they would remain unemployed and searching for work after that moment. By analogy with the renewal theory literature we call these durations the backward and forward recurrence times, respectively. For 40 individuals we could not construct the forward recurrence time because they were not interviewed in subsequent waves. These are mainly young people leaving their parents’ home. Note that this might create a selection problem since these people might leave because they found a job elsewhere. We return to this issue in Section IV.

Of the backward and forward recurrence times, 64 and 39% are censored in the sense that it is only known that the realised time exceeds a certain value. Part of the 39% is due to respondents who drop out of the panel before October 1985. Of all 112 uncensored forward recurrence times 71% ended in a transition into employment. The other 29% became non-participants. This means that according to the labour market history as defined above there is a date such that the spell of unemployment ends on it while after that date the individual is doing something else than working in a paid job. Consequently, the state of non-participation covers the wide range of activities that was mentioned above as being included in the ‘third’ category of activities. The limited amount of observations in the sample prohibits a subdivision of the state of non-participation into different states.

By taking a closer look at the uncensored forward durations we observe a phenomenon that appears strange at first sight. Of the 112 uncensored forward recurrence times 54% seem to have ended at the day of an interview. That is, at wave $n$ ($n = 1, 2, 3$) the individual reports that he is unemployed whereas at wave $n + 1$ he reports that as of the date of the previous interview he has been in a different state. Clearly these people over-estimate the elapsed duration of the activities that they perform after leaving the state of unemployment. We have to account for these ‘memory problems’ when deriving the likelihood.

The data set provides a range of personal characteristics. We used the characteristics as reported in April 1984. Since we do not know the level of benefits that individuals obtained during spells of unemployment that started and finished between two successive waves of the panel, we decided to consider only those spells that contained the date of the first interview. Subsection II.3, in which the explanatory variables in the model are discussed, contains a table with sample characteristics.

The data on income variables all count for the survey. The unemployment-insurance benefit variable is not imputed but instead is measured directly by asking respondents who are unemployed in the first wave of the panel what their net (after-tax) unemployment income was at the date of the interview. Benefits need not be constant throughout the spell. In The Netherlands in the beginning of the eighties the benefits level during the first years of unemployment is related to the pre-unemployment wage while after about two years it is determined by the public assistance system. However, if the benefits
level related to the pre-unemployment wage is below the public assistance level, 
or if the individual did not have a job before becoming unemployed, then he 
obtains public assistance benefits from the beginning. Given the lack of 
information on pre-unemployment wages it is not possible to infer whether an 
individual in the sample faced decreasing benefits or not. In subsection IV.2 we 
examine the consequences that ignoring decreases in benefits (if present in 
reality) have on the estimation results.

As said before the data do not contain information on the return from being 
a non-participant. In some cases non-participants can receive unemployment 
insurance benefits, for example, if according to their own perception they do 
unpaid work in the household and do not search for a job while they are still 
officially registered as being unemployed. For most activities covered by the 
state of non-participation however the income level is not directly related to the 
unemployment insurance system. For instance an individual who is retired or 
disabled obtains a fixed amount of money every month, usually supplemented 
by a pension if he is retired.

The data on the income variables that are used to estimate \( F(w) \) will be 
discussed in Section III as that section is devoted entirely to the estimation of 
\( F(w) \).

II.2. Likelihood Function

In our stationary model the backward and forward recurrence time and the 
state of destination given exit from unemployment are stochastically in-
dependent (see, for example, Ridder, 1984). Because of this independence the 
individual log-likelihood contribution is simply the sum of three parts. The 
state of destination given exit from unemployment has a Bernoulli distribution 
with parameter \( \theta/(\theta + \zeta) \). The forward recurrence time has an exponential 
distribution with parameter \( \theta + \zeta \). By assuming that the individual entry rate 
into unemployment is constant before the moment of the first interview, the 
backward recurrence time follows this distribution as well. The forward and 
backward recurrence times are denoted as \( \tau \) and \( t \), respectively. The state of 
destination is denoted as \( \epsilon \) with \( \epsilon = 1 \) if the state is employment and \( \epsilon = 0 \) if the 
state is non-participation. The occurrence of censoring and the occurrence of 
the so-called memory problems are taken to be exogenous. If \( \tau \) is missing then 
this is taken to be exogenous as well.

First consider the state of destination. Let \( \epsilon_1 = 1 \) if \( \tau \) is censored and \( \epsilon_1 = 0 \) 
otherwise. Let \( \epsilon_2 = 1 \) if \( \tau \) is missing and \( \epsilon_2 = 0 \) otherwise. The part of the 
individual log-likelihood contribution \( L \) due to the state of destination is \( L_1 \),

\[
L_1 = (1 - \epsilon_2)(1 - \epsilon_1) [\epsilon \log \theta + (1 - \epsilon) \log \zeta - \log (\theta + \zeta)].
\] (7)

So if \( \tau \) is censored or missing then \( \epsilon \) is not observed and consequently \( L_1 = 0 \).

Next consider the backward recurrence time. Let \( \epsilon_3 = 1 \) if \( t \) is censored and 
\( \epsilon_3 = 0 \) otherwise. The part of \( L \) due to \( t \) is \( L_2 \),

\[
L_2 = (1 - \epsilon_3) \log (\theta + \zeta) - t(\theta + \zeta).
\] (8)
If no memory problems are present then the part of $L$ due to $\tau$ can be obtained by replacing in equation (8) $1 - c_2$ by $(1 - c_1)(1 - c_2)$ and $t$ by $(1 - c_2)\tau$. Recall that memory problems are present if the data suggest that the spell of unemployment ended on the day at which the individual was being interviewed for the first, second or third time. For such individuals it can only be inferred that the spell ended somewhere between two subsequent interviews, say the $n$th and the $(n + 1)$th ($n = 1, 2$ or $3$). By assumption it is ruled out that transitions can be forgotten. One is inclined to think that when the spell of unemployment ends shortly before the $(n + 1)$th interview the date of the transition will be reported more accurately than when the spell ends shortly after the $n$th interview. This is confirmed by the fact that most reported transitions between two subsequent interviews took place less than three months before the second interview. Therefore, if a memory problem is present in the sense that a spell seems to have ended at the date of the first, second or third interview, then this is interpreted as evidence that the spell has ended between that date and three months later. Later on it will be examined whether the results are sensitive with respect to the assumption that memory problems can only occur if the transition takes place in the three month period after each interview. Let $\tau_1$ denote the length of this three month period. Let $c_4 = 1$ if a memory problem is present and $c_4 = 0$ otherwise. The part of $L$ due to $\tau$ is $L_3$, 

$$L_3 = (1 - c_4) \left\{ (1 - c_1) \left[ (1 - c_4) \log (\theta + \zeta) - \tau (\theta + \zeta) \right] 
+ c_4 \left[ \log \left[ e^{-(\theta + \bar{\gamma}) - e^{-(\theta + \bar{\gamma} + \lambda)}} \right] + c_1 \left[ - \tau (\theta + \zeta) \right] \right] \right\} 
+ (1 - c_1) c_4 \log [1 - e^{-(\theta + \bar{\gamma} + \lambda)}].$$

(9)

It is likely that similar to the occurrence of memory problems in the reported values of $\tau$ there may be problems in the reported values of $t$. In the sample almost no transitions into unemployment are reported for the first three months after April 1983. We assume that whenever a transition into unemployment occurred before July 1983, individuals with a memory problem report at the date of the first interview that they have been unemployed for more than a year. Consequently in case the reported censored $t$ equals one year then this is interpreted as evidence that $t$ exceeds nine months. Let $t_4$ denote the length of that nine month period. Equation (8) has to be modified to

$$L_5 = (1 - c_4) \left[ (1 - c_1) \log (\theta + \zeta) - t (\theta + \zeta) \right] - c_4 t_4 (\theta + \zeta).$$

(10)

The log-likelihood contribution $L$ of an individual with known $c_1, c_2, c_3, c_4, t, \tau$ and $\epsilon$ is given by the sum of the right-hand sides of equations (7), (9) and (10). The structural parameters and functions of the job search model $(u, v, \rho, \lambda, F(u))$ enter the likelihood via $\theta$ (see equations (1) and (3)). The parameter $\zeta$ enters $L$ both directly and indirectly via $\theta$. 
II.3. *The Empirical Implementation*

Now that we have specified the structural model and described the data we examine in this subsection the functional forms of the exogenous variables and discuss parameterisations. However, the wage offer distribution will be examined in Section III because that section is devoted entirely to the estimation of $F(w)$.

The job offer arrival rate $\lambda$ and the transition rate into non-participation $\zeta$ are written as exponential functions of observable exogenous variables $x$ and $z$, respectively,

$$\lambda = \exp(x'\beta),$$

$$\zeta = \exp(z'\gamma).$$

The vector $x$ includes variables which are of interest to employers, for example, because they give an indication of the productivity of the job searcher. Examples are level of education (we distinguish between five levels: (1) no certificate after primary education, (2) lower secondary education, (3) secondary education, (4) higher vocational training, (5) university), age, nationality, whether the individual has had a job before (this was asked explicitly) and whether he is married. We include the local unemployment percentage as a (crude) indicator of labour market tightness. The vector $x$ also includes a variable that depends on the number of working individuals in the household. If this number is high then the unemployed individual may have easier access to employers.

The vector $z$ consists of variables which are important for the process of transiting into non-participation, either by chance or by choice. Obviously, age is important because young individuals may get drafted into the armed forces and older individuals retire or get disabled more often than younger ones. Furthermore, young unemployed individuals often return to school for additional training especially if they did not have any job before.

Table 1 contains some sample characteristics of the explanatory variables in

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean</th>
<th>Standard deviation</th>
</tr>
</thead>
<tbody>
<tr>
<td>Benefits (guilders/week)</td>
<td>305</td>
<td>114</td>
</tr>
<tr>
<td>Local % unemployment rate</td>
<td>18</td>
<td>2.7</td>
</tr>
<tr>
<td>Age</td>
<td>32</td>
<td>11</td>
</tr>
<tr>
<td>Level of education</td>
<td>2.0</td>
<td>0.2</td>
</tr>
<tr>
<td>No. working in household</td>
<td>0.37</td>
<td>0.62</td>
</tr>
<tr>
<td>Dutch</td>
<td>0.91</td>
<td>—</td>
</tr>
<tr>
<td>Head of household</td>
<td>0.71</td>
<td>—</td>
</tr>
<tr>
<td>Married</td>
<td>0.48</td>
<td>—</td>
</tr>
<tr>
<td>New entrant</td>
<td>0.11</td>
<td>—</td>
</tr>
</tbody>
</table>
the model. From the previous subsection it is clear that sample averages and standard deviations of the duration variables are not informative. Therefore they are not presented.

Similarly to Narendranathan and Nickell (1985) and Ridder and Gorter (1986) the utility function of income \( u \) is taken to be logarithmic. The subjective rate of discount \( \rho \) is fixed at 10% per year. In Section IV we examine the robustness of the results with respect to changes in the functional form of \( u \) and with respect to the numerical value of \( \rho \).

Non-wage income is not included in the model because figures on personal non-wage income components are not available in the first wave of the panel survey. A reduced form estimation of \( \theta \) with income of other household members included as a regressor in \( \log \theta \) showed that this variable has no influence at all on the transition from unemployment into employment. Therefore it was omitted in the structural model.

The estimation method we have employed was ML using the Newton–Raphson algorithm. Because of the assumptions that were made on the functional forms of \( F(w) \) (see Section III) and \( u \), it follows that equation (3) can be rewritten as an equation that can be solved numerically for \( \phi \) with a high level of precision. Via equation (1) the likelihood contributions can then be calculated as a function of the parameters.

### III. The Wage Offer Distribution

The most natural way to obtain information on \( F(w) \) in a structural job search model is to use data on post-unemployment wages, for these are drawings from \( F(w) \) truncated at \( \bar{\phi} \). Combining such data with duration data makes it possible to estimate \( F(w) \) jointly with the other parameters in the model, provided that \( F(w) \) satisfies Flinn and Heckman’s (1982) recoverability condition. However, as we saw in II.1, in our sample there are only 79 transitions from unemployment. Obviously we want to allow for different \( F(w) \) in different segments of the labour market. For some segments there are not enough post-unemployment wages available in order to be able to estimate \( F(w) \). For instance there are only two individuals with a university degree who provide such wages. Therefore we take a totally different route in estimating \( F(w) \). We estimate \( F(w) \) \textit{a priori} using data on individuals who were employed at the date of the first interview. Analogous to Narendranathan and Nickell (1985) the \textit{a priori} estimation results serve to predict individual wage offer distributions for the unemployed. These predictions are plugged in when estimating the structural model.

Wages of employed individuals are not random drawings from \( F(w) \). A working individual accepted his present job because its wage exceeded his reservation wage when he was unemployed. Consequently, observed wages are drawn from a truncated distribution. However, the point of truncation (the reservation wage before obtaining the job) is unknown and cannot be estimated, because the level of unemployment benefits received before obtaining the current job is not available in the data set. In order to deal with
this problem we use an ad hoc reduced-form wage model. The wage \( w \) is observed if and only if one is employed. Previous studies (for example, Kiefer and Neumann, 1979) assumed this to be equivalent to \( w \geq \phi \), that is, \( w \) is observed if and only if it exceeds the reservation wage prior to employment. However, this is only true in a discrete time model in which exactly one job offer arrives per period (see Flinn and Heckman, 1982) which is a very strong assumption because it neglects various sources of the dynamics and uncertainty in the process of search. Therefore we take a latent variable \( y^* \) as determining whether one is employed: \( w \) is observed if and only if \( y^* > 0 \). The wage offer distribution \( F(w) \) is assumed to be lognormal with parameters \( \mu \) and \( \sigma^2 \); \( \mu = \eta' x_1 \) with \( x_1 \) observed. The unobserved variable \( y^* \) is assumed to depend on a linear combination of observed exogenous variables \( x_2 \) and an additive error term. Obviously every factor that influences \( F(w) \) influences \( y^* \) as well. Therefore the variables in \( x_1 \) are included in the set of variables in \( x_2 \). In order to allow for different values of the parameters of the wage model in different segments of the labour market the wage model is estimated separately for each segment. Details of the estimation and the results are given in van den Berg (1988). The wage offer distribution of an unemployed individual with characteristics \( x_1 \) and parameters \( \eta \) and \( \sigma^2 \) associated with the segment he can be ascribed to, is predicted as being lognormal with parameters \( \eta' x_1 \) and \( \sigma^2 \). The predicted \( F(w) \) are plugged in when estimating the structural model. The sample averages of the estimated expectation and standard deviation of the wage offer distribution equal 470 and 99 guilders per week, respectively. (Sample standard errors of these estimated values equal 86 and 27 guilders per week, respectively.)

In terms of the alternative interpretation of the structural model (see I.2) the procedure described above does not give estimates of \( F(w) \) but instead it provides estimates of the individual distributions of vacancy wage offers corrected for wage competition (see equations (5) and (6)),

\[
\frac{\int_{0}^{w} q_3(w) \, dG(w)}{\int_{0}^{\infty} q_3(w) \, dG(w)}, \quad w \geq 0.
\]

A final thing to note is that for a variety of reasons the current wage rate of an employed individual may exceed the wage rate that he obtained directly after becoming employed. In Section V a model that deals with this issue is considered. Further it is outlined how the wage offer distribution can be estimated in the presence of such wage differences.

IV. RESULTS

IV.1. Parameter Estimates

The parameter estimates for the structural model described in I.1 and II.3 are presented in Table 2. The unit time period is one week. For the age and education dummies the reference categories are the age category 46–64 and the
level of education \( t \), respectively. Generally, the results seem to be in accordance with intuition. Education has a very significant influence on the job offer arrival rate. An individual having the highest level of education receives offers more than seven times as frequently as an individual with the lowest level of education. New entrants, having no experience, are offered jobs less often than experienced individuals. Being married is perceived by employers as a desirable property whereas being a head of a household is not. Single individuals are also defined as being head of a household, so it may be that what really matters for employers is not the sheer presence of a partner but the presence of a family which makes the employee feel responsible. The importance of the number of working household members may be because unemployed individuals, for which this number is high, have easier access to employers. However, it may also be a consequence of a positive correlation between unobserved characteristics of the unemployed individual and characteristics of other household members, as far as these characteristics are relevant for employers. The local unemployment rate has no significant influence on \( \lambda \). Other indicators of the tightness of the labour market like the local UV ratio performed even worse. Van Opstal and Theeuwes (1986) who estimated a reduced-form duration model using Dutch data from 1984, also report this lack of significance. Presumably, job search is not restricted to a region any more.

Table 2  
Parameter Estimates for the Search Model

<table>
<thead>
<tr>
<th>Variable/parameter</th>
<th>Coefficient</th>
<th>t ratio</th>
</tr>
</thead>
<tbody>
<tr>
<td>(i) Job offer arrival rate</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Constant</td>
<td>-6.08</td>
<td>6.4</td>
</tr>
<tr>
<td>Dutch</td>
<td>0.55</td>
<td>1.3</td>
</tr>
<tr>
<td>Education: level 2</td>
<td>0.91</td>
<td>3.3</td>
</tr>
<tr>
<td>Education: level 3</td>
<td>1.17</td>
<td>3.6</td>
</tr>
<tr>
<td>Education: level 4</td>
<td>1.74</td>
<td>2.8</td>
</tr>
<tr>
<td>Education: level 5</td>
<td>1.97</td>
<td>2.8</td>
</tr>
<tr>
<td>Age category 18–23</td>
<td>0.68</td>
<td>1.4</td>
</tr>
<tr>
<td>Age category 24–29</td>
<td>0.50</td>
<td>1.2</td>
</tr>
<tr>
<td>Age category 30–45</td>
<td>0.16</td>
<td>0.4</td>
</tr>
<tr>
<td>New entrant</td>
<td>-0.82</td>
<td>1.5</td>
</tr>
<tr>
<td>Head of household</td>
<td>-0.03</td>
<td>0.1</td>
</tr>
<tr>
<td>Married</td>
<td>0.78</td>
<td>2.5</td>
</tr>
<tr>
<td>log (1 + no. working in household)</td>
<td>1.03</td>
<td>3.0</td>
</tr>
<tr>
<td>Local % unemployment rate</td>
<td>-0.04</td>
<td>1.1</td>
</tr>
<tr>
<td>(ii) Rate of transition into non-participation</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Constant</td>
<td>-4.91</td>
<td>16.4</td>
</tr>
<tr>
<td>Age category 18–23</td>
<td>-0.41</td>
<td>0.8</td>
</tr>
<tr>
<td>Age category 24–29</td>
<td>-1.06</td>
<td>2.3</td>
</tr>
<tr>
<td>Age category 30–45</td>
<td>-1.39</td>
<td>2.9</td>
</tr>
<tr>
<td>New entrant</td>
<td>0.66</td>
<td>1.4</td>
</tr>
<tr>
<td>(iii) Disutility of unemployment</td>
<td>0.74</td>
<td>5.2</td>
</tr>
</tbody>
</table>

Log likelihood = -893.23.
Another explanation is that numbers on registered vacancies and unemployed individuals may not be accurate indicators of labour market tightness. Still, the estimate of $-0.04$ seems plausible: it implies that moving from the province with the highest rate of unemployment (24%) to the one with the lowest (15%) increases $\lambda$ with a factor of almost 1.5. The separate age coefficients in $\lambda$ are not significant. Replacement of the age dummy variables by log (age) and its squared value results in even less significant estimates. However, a Likelihood Ratio test of the hypothesis that all age dummy coefficients equal zero leads to a rejection at the 10% level. In Section II it was noted that in some cases censoring of the forward recurrence time of young individuals may arise because they leave their parents' home in order to start working elsewhere. If so, then the coefficient on the age category 18–23 in the job offer arrival rate is under-estimated.

In terms of the alternative interpretation of the model (see I.2) $\lambda$ is the product of the vacancy arrival rate $q_1$ and the term $q_4$ which captures the influence of non-wage variables on the acceptance probability conditional on application $q_4$. We expect the unemployment rate, experience in previous jobs, education and age to be linked to $q_1$ while nationality and household characteristics probably are linked to $q_4$. The signs of the coefficients seem to confirm these prior expectations.

Turning to the rate of transition into non-participation, we see that new entrants leave the labour market more often and that this is also true for individuals aged below 24 or over 45. The disutility of unemployment $v$ is smaller than one, implying that contrary to popular statements, being unemployed is regarded as unpleasant. From the standard error of 0.14 it follows that the hypothesis $v = 1$ is rejected by a Wald test at the 10% level but not at the 5% level. However, the Likelihood Ratio test statistic for this hypothesis equals $20.4 > \chi^2(0.95)$ so $v = 1$ is strongly rejected.

IV.2. *The Characteristics of the Search Process*

Given the parameter estimates, the main variables of the search process can be estimated and the influence of changes of the benefit level on these variables can be evaluated. We first present and interpret sample averages of the estimates. Subsequently the results are compared with other results in the literature and it is shown why our results differ in some respects. Table 3 presents sample averages of the estimates of $\lambda$, $\bar{F}(\phi)$ and $\zeta$ for different age categories and levels of education. The expected numbers of job offers and transitions into non-participation in a year can be obtained by multiplying the numbers in the $\lambda$ and $\zeta$ row by 52.1. What strikes most is that in most cases $\bar{F}(\phi)$ is nearly equal to one. In particular those who are aged under 24 or over 46, or who have a primary education only, accept virtually every job that is being offered. Still, even individuals with a university degree have a probability of 0.8 of accepting the first job offered. It means that the reservation wages are located in the left part of the left tail of the wage offer distribution. The reason for this is the combination of on the one hand a very small job offer arrival rate and on the other hand very low values of the utility function in unemployment.
### Table 3

**Probabilities and Expectations**

<table>
<thead>
<tr>
<th>(i) By age category</th>
<th>18–23</th>
<th>24–29</th>
<th>30–45</th>
<th>46–64</th>
<th>Average</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \lambda ) (expected number of offers in a week)</td>
<td>0.012</td>
<td>0.016</td>
<td>0.012</td>
<td>0.008</td>
<td>0.012</td>
</tr>
<tr>
<td>( \bar{F}(\phi) ) (proportion of offers acceptable)</td>
<td>0.999</td>
<td>0.994</td>
<td>0.966</td>
<td>1.080</td>
<td>0.977</td>
</tr>
<tr>
<td>( \xi ) (expected number of transitions into non-participation in a week)</td>
<td>0.007</td>
<td>0.003</td>
<td>0.002</td>
<td>0.007</td>
<td>0.004</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>(ii) By level of education</th>
<th>1</th>
<th>2</th>
<th>3</th>
<th>4</th>
<th>5</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \lambda )</td>
<td>0.004</td>
<td>0.014</td>
<td>0.018</td>
<td>0.024</td>
<td>0.033</td>
</tr>
<tr>
<td>( \bar{F}(\phi) )</td>
<td>1.000</td>
<td>0.989</td>
<td>0.954</td>
<td>0.899</td>
<td>0.882</td>
</tr>
<tr>
<td>( \xi )</td>
<td>0.004</td>
<td>0.004</td>
<td>0.004</td>
<td>0.004</td>
<td>0.003</td>
</tr>
</tbody>
</table>

\((w(b))\) relative to employment \((w(u))\). Rejection of an offer may well imply a waiting time of more than a year before the next offer arrives. In the meantime the only source of income is benefits, which appear to be rather low relative to wages: the sample average of \( \bar{F}(b) \) equals 0.99 and the ratio of the sample averages of \( b \) and \( E(u) \) equals 0.65. Moreover, because \( v < 1 \) there is a premium on being employed and one is willing to offer money for it by accepting lower-paid jobs. In fact, in our sample 79\% of the unemployed even accept jobs with wages below their benefit level, that is, for these individuals \( \phi < b \).

From Table 3 it can be inferred that for groups with a very low job offer arrival rate, almost 50\% of all spells of unemployment end in a transition into non-participation. In other words, without such transitions the durations of unemployment for such individuals would be approximately twice as long.

The results so far enable us to investigate a number of questions related to the effectiveness of policies aimed at a reduction of unemployment durations. Table 4 presents for different age categories and levels of education sample averages of the elasticities of the reservation wage, the transition rate from unemployment into employment \( \theta \), and the expected duration \( d \), with respect to the level of benefits. The results are unambiguous: a decrease in the level of benefits has virtually no effect on durations. Even for unemployed individuals with a university degree a 10\% drop in benefits causes only a 1\% drop in the expected duration. The individuals who suffer most from long spells (having primary education only, or aged under 24 or over 46) are completely insensitive to the benefits policy instrument. Note that elasticities refer only to infinitesimal changes. Still, even a large decrease in the level of benefits does not have much influence on duration. Individuals accept most jobs already, so a decrease in \( \phi \) forced by a large decrease in \( b \) does not help much. The expected duration is bounded from below by \( 1/(\lambda + \xi) \). From the results it is also clear that at an individual level additional educational training increases labour market opportunities.

The result that changing the benefits level has virtually no effect on the
transition rate into employment and on the expected duration is in contrast with most empirical literature on unemployment durations. Early studies by Lancaster (1979), Nickell (1979), Lancaster and Nickell (1980) and Lancaster and Chesher (1983) suggest values of around 0·6 to 1·0 for the elasticity of the expected duration with respect to benefits. More recent work by Atkinson et al. (1984), Narendranathan et al. (1985), Narendranathan and Nickell (1985) and Main and Shelly (1988) reports values of this elasticity that are typically ranging from about 0·1 to 0·3. The early studies use U.K. data from the beginning of the seventies while the later work uses more recent U.K. data. On the other hand, some studies that use Dutch data from the same observation period as we do (the mid-eighties) do not find any effect on duration of changing the benefits level. Van Opstal and Theeuwes (1986) and Groot and Huurne (1988) obtain zero estimates for the elasticity of expected duration with respect to benefits from the estimation of reduced-form duration models, using data from young individuals only. Vissers and Groot (1989) estimate a series of reduced-form duration models: they consider several model specifications and several ways of defining the unemployment benefit variable, and they use different Dutch datasets from the mid-eighties to estimate the model. Nevertheless their results are unambiguous in the sense that the elasticity of the transition rate into employment with respect to the benefits variable is insignificantly different from zero. Thus it seems that the results in Table 4 are not just an artefact of our particular sample but instead may be typical for the Netherlands in the mid-eighties.

### Table 4

<table>
<thead>
<tr>
<th>Age category...</th>
<th>18–23</th>
<th>24–29</th>
<th>30–45</th>
<th>46–64</th>
<th>Average</th>
</tr>
</thead>
<tbody>
<tr>
<td>𝛿log φ/𝛿log b (reservation wage)</td>
<td>0·36</td>
<td>0·24</td>
<td>0·25</td>
<td>0·46</td>
<td>0·30</td>
</tr>
<tr>
<td>𝛿log θ/𝛿log b (hazard)</td>
<td>0·01</td>
<td>0·05</td>
<td>0·04</td>
<td>0·00</td>
<td>0·03</td>
</tr>
<tr>
<td>𝛿log d/𝛿log b (expected duration)</td>
<td>0·01</td>
<td>0·05</td>
<td>0·03</td>
<td>0·00</td>
<td>0·03</td>
</tr>
</tbody>
</table>

(i) By age category

<table>
<thead>
<tr>
<th>Level of education...</th>
<th>1</th>
<th>2</th>
<th>3</th>
<th>4</th>
<th>5</th>
</tr>
</thead>
<tbody>
<tr>
<td>𝛿log φ/𝛿log b</td>
<td>0·44</td>
<td>0·24</td>
<td>0·23</td>
<td>0·19</td>
<td>0·16</td>
</tr>
<tr>
<td>𝛿log θ/𝛿log b</td>
<td>0·00</td>
<td>0·03</td>
<td>0·06</td>
<td>0·07</td>
<td>0·11</td>
</tr>
<tr>
<td>𝛿log d/𝛿log b</td>
<td>0·00</td>
<td>0·03</td>
<td>0·05</td>
<td>0·07</td>
<td>0·10</td>
</tr>
</tbody>
</table>

(ii) By level of education

\(d\) equals the expected duration of unemployment.

In order to shed more light on this issue we examine in some detail the expression for the elasticity \(e\) of \(\theta\) with respect to \(b\). For simplicity we set \(\zeta\) equal to zero. (Alternatively, \(\zeta\) is put into \(\rho\).) From equations (1) and (2) it follows,

\[
e = \frac{\delta \log \theta}{\delta \log b} = - \left[ \frac{\phi f(\phi)}{F(\phi)} \right] \frac{\rho \nu}{\rho + \theta},
\]  

(11)
with \( f \) being the derivative of \( F \). One sees immediately that \( e \) depends on all other variables in the model. In other words, according to search theory cross-effects play a role in the effect that changing the benefits level has on duration. Consequently it is hard to regard the elasticity as a parameter that is equal in different economic environments. (See Feldstein and Poterba, 1984 and Atkinson et al. 1984 for similar statements.) Now let us have a look at the way \( e \) varies with the other variables. The second part of the right-hand side of (11) is an increasing function of \( \phi \). At first sight it seems that the way the term between brackets \( k(\phi) \) varies with \( \phi \) depends crucially on the class of wage offer distributions under consideration. However, it can be shown that for virtually every class of distributions for \( F \), including the (truncated) Normal, Lognormal, Weibull, Gamma, Laplace, Uniform, Triangular, Beta, \( t \), Logistic and Log-logistic class of distributions, \( k(\phi) \) is a strictly increasing function on the interval in \([0, \infty)\) on which \( f \) is positive, with \( k(0) = 0 \). This implies that for almost every class of \( F \), including the class we adopted, it holds that the smaller the reservation wage is, the smaller the effect of changing the benefits level on the transition rate into employment is. (The Pareto class of distributions is an extreme case because then \( k(\phi) \) is a constant. This may explain the exceptionally high estimate of \( e \) in Ridder and Gorter (1986), who estimated a structural model in which \( F \) is a Pareto distribution.) Because the derivatives of \( \phi \) with respect to \( b \) and \( \nu \) are positive, it therefore follows that the smaller the benefits level and the smaller the non-pecuniary utility of unemployment are, the smaller the effect of changing \( b \) on \( \theta \) is. (This result is robust with respect to the functional form of \( \nu \); for example, it can be shown that it also holds if \( \nu \) is linear. See also Feldstein and Poterba, 1984 for some numerical examples in a simple model framework.) Further, numerical calculations show that in the neighbourhood of the estimates of Tables 2 and 3 it holds that the smaller the job offer arrival rate is, the smaller the effect of changing \( b \) on \( \theta \) is. In the Netherlands in the years around 1984 there was a very slack labour market. The level of unemployment was extremely high (the national unemployment percentage reached its peak in 1984 when it equalled 17.3\%\) while at the same time the number of vacancies was small (the \( V/U \) ratio was about 0.02 in 1984). Clearly labour market conditions were worse than the conditions that prevailed when the data for most of the previous studies mentioned before were collected. This shows up in the relatively low estimates of \( \lambda \) as reported in Table 3. Further, as shown before, in our dataset the level of benefits is generally very small as compared to most of the wage offers. As a result, the estimated reservation wages are located in the left part of the left tail of \( F(\nu) \) and the elasticities of \( \theta \) and the expected duration of unemployment with respect to \( b \) are almost zero. (Note that the latter elasticity is always smaller in absolute value than the former.)

A point that is related to the previous paragraphs concerns the influence of the sampling scheme on the sample averages in Tables 3 and 4. The dataset used to estimate the model is basically a sample from the stock of the unemployed (on the date of the first interview). The distribution of the observed explanatory variables in the stock of the unemployed differs from that
in the flow into unemployment, in the sense that values that are associated with a high expected duration are overrepresented in the stock (This can be inferred from the analytical results in Ridder, 1984.) As a result, a comparison of sample averages of, for example, the estimates of $\lambda$ and $\varepsilon$ obtained by using stock data (such as ours), with sample averages obtained by using flow data (such as that in Narendranathan and Nickell, 1985), is hampered by the fact that $\lambda$ and $\varepsilon$ depend on observed explanatory variables that are unequally distributed in the two sampling schemes. Further, it follows that the results in Table 4 only refer to the effect that changes in $b$ on average have on individuals in the stock of the unemployed. Note that in a reduced form model framework such problems with regard to $\varepsilon$ do not exist because in reduced form models $\varepsilon$ is a parameter that does not depend on explanatory variables.

In light of the previous paragraphs it is not surprising that the values of the elasticity of $\phi$ with respect to $b$ are somewhat different from those found in other studies. Lancaster and Chesher (1983) and Narendranathan and Nickell (1985) suggest values between $0.1$ and $0.2$ for this elasticity. Main and Shelly (1988) report values of about $0.32$ for youth training scheme participants and values of about $0.16$ for other unemployed youths. In our sample the job offer arrival rate is extremely small so for an unemployed individual $b$ is an important determinant of the expected discounted lifetime utility. Consequently the effect on $\phi$ of a change in $b$ is relatively large. As explained above this does not translate into a substantial change in $\theta$.

Obviously in the present context only micro effects of a cut in benefits can be investigated. On a macro level such a policy is likely to generate additional effects both on the inflow into unemployment and on the transition from unemployment into employment. (Narendranathan et al., 1985). Also, if there is an element of choice as to whether to become non-participant or not, then a cut in benefits may have an effect on $\zeta$. The sign of this effect depends among other things on the dependence of the distribution of income of non-participants on the level of benefits. If benefits are decreased whereas the incomes of non-participants like conscripts and disabled remain unchanged then equation (2) does not hold any more. Therefore an investigation of the relation between $b$ and $\zeta$ should be made in a wholly structural model setting and is beyond the scope of this paper. Inclusion of $\log (\text{benefits})$ as a regressor in $\log \zeta$ resulted in a highly insignificant parameter estimate of $-0.14$ ($t = 0.3$), all other things being almost identically equal.

Since the model does not allow for non-stationarity, it may be interesting to examine in what sense the results are affected by this omission. It is widely believed that the transition rate into employment $\theta$ is a decreasing function of duration. On the other hand, as we saw in II.1, $b$ may decrease during unemployment, and this makes $\theta$ ceteris paribus an increasing function of duration. One possible explanation for a decreasing $\theta$ is that the job offer arrival rate decreases sharply during unemployment, for example, as a consequence of a scar effect of being unemployed for a long time, and that this decrease of $\lambda$ offsets the increase in $F(\phi)$. If $\theta$ is a decreasing function of duration then the expected duration of the backward and forward recurrence
times exceeds the expected duration of completed durations of unemployment and a stock sample of unemployed individuals contains a relatively large number of long-term unemployed individuals. Further, if both \( b \) and \( \lambda \) decrease during unemployment then \( \phi \) also decreases. So if nonstationarity is present in reality in the sense that \( b, \lambda, \phi \) and \( \theta \) all decrease, then \( b \) in the sample is on average smaller than the benefits level for the short-term unemployed and \( \lambda, \phi \) and \( \theta \) are underestimated in the sense that shortly after the inflow into unemployment these variables are larger than estimated. From the discussion of equation (11) it then follows that in such cases the short-term unemployed are more sensitive with respect to changes in \( b \) than the results in Table 4 suggest, because in such cases \( \epsilon \) is more negative than reported shortly after the inflow into unemployment. However, short-term unemployed individuals may anticipate decreases of \( \lambda \) or \( b \) by modifying the reservation wage before these decreases take place. If such anticipations are strong then it is hard to elaborate on the effects that not allowing for nonstationarity may have on the estimates of the elasticity (van den Berg, 1990).

Another kind of nonstationarity is present if the transition rate into non-participation \( \zeta \) increases as a function of duration, as a result of a discouraged worker effect, for example. By analogy from the argument pointed out above it may be expected that in such a case \( \zeta \) is underestimated for individuals who are long-term unemployed.

IV.3. The Model Specification Revisited

This subsection examines whether the results are sensitive with respect to changes in some of the assumptions. Changes in the way jobs are characterised in the model (infinite duration, constant wages) are referred to in Section V where estimation results are presented for an extended model that deals with this.

In the structural model used for the empirical analysis \( v \) is the only exogenous variable which is estimated but not parameterised. It thus seems natural to extend the model by making \( v \) a function of observable individual characteristics. Also, one might ask why \( \rho \) is not estimated and why \( u \) is not parameterised, say, by assuming it to be a one-parameter CARA utility function. (CARA = constant absolute risk aversion; \( u(x) = -\exp(-cx) \) with \( c > 0 \)) Though such extensions do not raise identification problems in the statistical sense, it appears that there is not sufficient information in the data to be able to estimate such additional parameters. Apparently the likelihood is an almost completely constant function of such parameters in the neighbourhood of the optimum. This can be explained by recalling the results in Tables 3 and 4. First note that generally \( \phi \) is small with respect to most wage offers, which implies that \( f(\phi) \) is small so small changes in \( \phi \) given values of \( \lambda, \xi \) and \( F(w) \) do not affect the value of the likelihood function much. Secondly, \( u, \rho \) and \( v \) enter the likelihood only via \( \phi \). Therefore the correlation between estimates of parameters of \( u, v \) and \( \rho \) will be very high.

In the empirical model \( v \) is the only parameter that enters the likelihood via \( \phi \) only. The discussion in the previous paragraph suggests that \( \phi \) might be
biased if \( u \) is misspecified or if \( \rho \) has the wrong value. This is investigated by re-estimating the model with different \( u \) and \( \rho \). Throughout the range of acceptable values of \( \rho \) the estimation results for \( \lambda \) and \( \xi \) hardly differ from the original results (which are obtained by assuming \( \rho = 10\% \) per year). The differences in the value of \( \rho \) are absorbed by \( \delta \), higher values of \( \rho \) resulting in higher values of \( \delta \) thus holding \( \phi \) and therefore the fit of the model constant. For instance if \( \rho = 5\% \) then \( \delta = 0.67 \) (standard error: 0.19) while if \( \rho = 15\% \) then \( \delta = 0.78 \) (0.12). Still, \( \delta \) is always significantly smaller than 1 according to LR tests at the 1% level. Even in the limiting case of \( \rho = \infty \) the estimate of \( \delta \) is significantly smaller than 1 (\( \delta = 0.91 \)).

We also tried to re-estimate the model using a linear utility function \( u \) of income. This did not work. In the process of maximising the likelihood \( v \) tended to zero. This may be regarded as a justification for using a risk-averse specification of \( u \) because in that case the level of \( \phi \) for \( v = 0 \) is ceteris paribus lower than the corresponding level in the risk-neutral case.

In Section I we stated the assumptions that equation (2) holds and that the non-pecuniary utility of being a non-participant equals that of being unemployed. In what sense are the results affected if these assumptions are relaxed? Denote the non-pecuniary component of utility in non-participation by \( v_1 \) and the corresponding component in unemployment by \( v_2 \). It can be shown that if \( v_1 \neq v_2 \) or \( Eu(x) \neq u(b) \) then the parameter \( v \) in equation (3) has to be replaced by

\[
\frac{\xi v_1 \frac{Eu(x)}{u(b)} + \rho v_2}{\xi + \rho},
\]

in order to obtain the equation for the optimal reservation wage. So then \( \delta \) represents the estimate of expression (12). It follows that

\[ v_1 \frac{Eu(x)}{u(b)} > v_2 u(b) \iff \delta > v_2, \]

so if we believe that \( v_1 > v_2 \) or that \( Eu(x) > u(b) \) then the estimate of \( v \) implies that the estimate of the disutility of unemployment is even smaller than 0.74.

In Section II we discussed the so-called memory problems. There it was argued that values of 3 and 9 months for \( \tau_1 \) and \( t_1 \) respectively, were plausible. It appears that the parameter estimates are insensitive to changes of these values, though standard errors increase if \( \tau_1 \) increases or \( t_1 \) decreases.

When deriving the distribution of the backward recurrence time \( t \) we assumed that the rate of entry into unemployment is constant until May 1984. One may question whether this assumption holds true. According to Pissarides (1986) in the United Kingdom the entry rate was fairly constant between 1967 and 1983 apart from an increase in 1979–81. In the absence of reliable Dutch data we examine the sensitivity of the results with respect to the constant entry rate assumption by re-estimating the model with a time-varying entry rate. In particular we take as an alternative assumption that the entry rate \( q \) between January 1980 and January 1983 is twice as large as it is outside that time interval. In van den Berg (1988) the appropriate likelihood is derived. The
main effect of the alternative assumption about \( q \) on the estimation results is that the exit rate out of unemployment \( \theta + \zeta \) is estimated to be 13% larger. However, \( \theta \) and \( \zeta \) are still very small, and \( n, \overline{F} (\phi) \) and the elasticities are insensitive to the change in the assumption about \( q \). Thus, the main results and conclusions from IV.1 and IV.2 do not appear to be sensitive to changes in the assumptions about the time pattern of the entry rate into unemployment that are reasonable a priori.

One may question whether the estimation results are affected by a possible misspecification of the wage offer distribution which is estimated a priori. Obviously, \( F(w) \) plays a central role in the model because the trade-off between wages and benefits is a major determinant of search behaviour. We constructed \( F(w) \) which are lognormal and have same variances as before, but which have expectations that are shifted by 20% in comparison to the expectations derived in Section III. Re-estimation of the model using these alternative \( F(w) \) resulted in values that are almost identical to those presented in Tables 2-4. The shifts in \( E(w) \) are absorbed by \( \delta' \), a value of 1.2 times the original \( E(w) \) resulting in \( \delta' = 0.82 \) and a value of 0.8 times the original \( E(w) \) resulting in \( \delta' = 0.64 \). Consequently, the main conclusions are insensitive with respect to small misspecifications in the location of \( F(w) \).

When deriving the likelihood no account has been taken of unobserved heterogeneity in the sample. If unobserved heterogeneity is present in reality then the estimates may be inconsistent. However, estimating a structural model that allows for such heterogeneity is extremely complicated. Consider, for example, the case in which unobserved heterogeneity is present in \( \lambda \). We may rewrite \( \lambda \) as a product,

\[
\lambda = \nu \exp (x' \beta),
\]

in which the random term \( \nu \) represents unobserved heterogeneity in \( \lambda \) across the population. Substitution of equation (12) in equations (3) and (1) reveals that the hazard cannot be written explicitly as an analytical function of \( \nu \). Therefore, calculating the unconditional (on \( \nu \)) duration density by integrating the conditional density with respect to the density of \( \nu \) will be very complicated and is not pursued here.

V. AN EXTENDED MODEL

V.1. The Model

In reality the duration of employment is not infinite, nor are wages constant during employment. The prospective rate of wage increases and the distribution of the duration of employment affect the value of search of an unemployed individual. Therefore they should be incorporated in the model. In this section we deal with this.

We assume that the duration of employment has an exponential distribution with parameter \( s \) which is the layoff rate. During one period of employment one can hold several consecutive jobs without intervening spells of unemployment. It is assumed that one returns to the state of unemployment if a layoff occurs,
and that the duration of employment is stochastically independent of both the initial wage rate and the duration of unemployment that precedes employment.

During a spell of employment wages can increase for several reasons such as rising productivity or transitions from jobs with lower wages to jobs with higher wages without intervening spells of unemployment (on-the-job search). As a stylised description of this we assume that the wage pattern during employment is characterised by \( w(t) \) giving the wage rate as a function of the time \( t \) that one is employed conditional on the initial wage \( w(o) \),

\[
w(t) = w(o) e^{\alpha t}, \tag{14}
\]

in which \( \alpha \) does not depend on \( w(o) \) or \( t \) or on the duration of unemployment preceding employment. Though it is conceivable that mechanisms linking \( \alpha \), \( t \) and \( w(o) \) exist, the exploration of this is beyond the scope of the paper.

The extensions of the model do not affect the stationarity property of search behaviour of the unemployed. In van den Berg (1988) it is proven that the reservation wage \( \phi \) corresponding to the optimal strategy satisfies

\[
\log \phi = v \log b + \frac{\lambda}{\rho + \xi} \int_{\phi}^{\infty} (\log w - \log \phi) dF(w) - \frac{\alpha}{\rho + s}. \tag{15}
\]

\( F(w) \) is the distribution of initial-wage offers, which is the distribution from which the \( w(o) \) are drawn. Note that the derivative of \( \phi \) with respect to \( \alpha \) is negative. If \( \alpha \) is large then the value of search is high. However, this does not make the searcher more selective with regards to wage offers. It is profitable to give up more present income (a low \( w(o) \)) in order to obtain a higher income in the future.

The estimation of \( F(w) \) has to be reconsidered because in Section III we used a (cross-section) sample from the stock of the employed and therefore used data on current wages, that is, data on wages which are higher than the initial wages offered at the start of the current employment spell. We assume that the distribution of current wages is lognormal with parameters \( \mu \) and \( \sigma^2 \). Thus, in Section III these parameters are estimated. The distribution \( F(w) \) of initial-wage offers has to be recovered from the distribution of current wages. In van den Berg (1988) it is shown that \( F(w) \) can be approximated by a lognormal distribution with parameters \( (\mu + \log(s - \alpha) - \log s) \) and \( \sigma^2 \),

\[
F(w) \approx \text{LN} \left( \mu + \log \frac{s - \alpha}{s}, \sigma^2 \right). \tag{16}
\]

This requires \( s > \alpha \). The approximation is good for \( s >> \alpha \).

V.2. The Results

The approximation in equation (16) is used to obtain \textit{a priori} estimates of the individual distribution functions \( F(w) \). The results from Section III provide the individual values of \( \mu \) and \( \sigma^2 \). The parameter \( \alpha \) is fixed at 4% per year. We used the elapsed duration of employment of individuals who were employed in April 1984 to estimate \( s \). Since we assume that the entry rate into employment
is constant (the stationarity assumption) these incomplete durations have an
exponential distribution with parameter \( s \). In accordance with the treatment of
the memory problem in II.2, durations are censored at 9 months. The ML
estimate of \( s \) equals 14.4% per year (t ratio equals 14.2) which implies that the
expected duration of employment is almost seven years. This estimate may be
biased for a variety of reasons (such as neglected unobserved heterogeneity) but we
believe that for our purposes it is accurate enough.

From equation (16) it can be deduced that the expectation and the standard
deviation of \( F(w) \) are 100(a/s)\% = 28% smaller than those obtained in
Section III. The sample average of the probability that a random initial-wage
offer exceeds the benefit level is 0.61 as opposed to 0.91 when \( F(w) \) is estimated
as in Section III.

Table 5

\begin{tabular}{|c|c|c|}
\hline
Variable/parameter & Coefficient & Estimates for the basic model \\
\hline
\( v \) & 0.88 & 0.74 \\
\( \lambda \) & 0.012 & 0.012 \\
\( \zeta \) & 0.004 & 0.004 \\
\( \bar{F}(\phi) \) & 0.98 & 0.97 \\
\( \partial \log \phi / \partial \log b \) & 0.49 & 0.30 \\
\( \partial \log \theta / \partial \log b \) & -0.04 & -0.03 \\
\( \partial \log z / \partial \log b \) & 0.93 & 0.93 \\
\hline
\end{tabular}

The estimates and t ratios of the parameters of \( \lambda \) and \( \zeta \) hardly differ from
those presented in Table 2. Further, the general pattern of the results presented
in Tables 3 and 4 is preserved. Therefore only sample averages of the main
variables are presented for the extended model (see Table 5). \( \bar{F}(\phi) \), \( \lambda \) and \( \zeta \)
have almost the same sample averages as before. The parameter \( v \) is significantly
smaller than 1 according to a LR test (test-statistic value 36.0 > \( \lambda_1^2 \)(0.99)). The
job offer acceptance probability is large because of the combination of a small
job offer arrival rate and a low utility value attached to being in the state of
unemployment. The latter holds both because one dislikes being unemployed
for non-pecuniary reasons and because, in unemployment, income is constant
whereas one expects it to increase in employment. In the extended model \( b \) is
generally close to the median of \( F(w) \). So in this model it is the rate of income
increases rather than the level of income which makes employment preferable
from a material point of view. The elasticity of the expected duration with
respect to the level of benefits is very small. The reasons for this are similar to
those given in IV.2 to explain the results in Table 4.

In order to examine whether the estimates are sensitive to changes in the
assumed values of \( \rho \) and \( \alpha \), the model is re-estimated for alternative values of
\( \rho \) (5 and 15% per year) and \( \alpha \) (3 and 5% per year). It appears that the
differences in the values of \( \rho \) and \( \alpha \) are absorbed by \( \nu \) and that all other results
in Table 5 and the fit of the model are almost constant for the cases considered. The value of \( \rho \) ranges from 0.82 to 0.85 so the sensitivity of \( \rho \) to changes in the value of \( \rho \) is less than in the basic model.

In sum, the main conclusions from Section IV about the parameter estimates, about the relative magnitudes of the main variables for different age categories and levels of education, and about the effects of changes in the level of benefits, remain unaffected. The results in this section suggest that on-the-job search may be an important factor for search behaviour of the unemployed. Therefore a topic for further research would be to extend the model to include on-the-job search explicitly. Using data of employed and unemployed individuals simultaneously, the wage offer distribution could be estimated along with the other variables. Also, some of the rather rigid assumptions that were made in this section could be relaxed in such a model.

VI. CONCLUSIONS

In this paper we have extended the existing empirical literature on structural job search models by specifying and estimating a model that allows for transitions from unemployment into non-participation. Moreover, a version of the model deals with the influence of prospective wage increases during employment on the search behaviour of the unemployed. The model is estimated using Dutch data from 1983–5. The results indicate that almost every job offer is acceptable. The reason for this is the combination of a very small job offer arrival rate and low values of the utility function in unemployment relative to employment. If one turns down an offer then generally one has to wait for a very long time before the next offer arrives. In the meantime one is unemployed, which is disliked both for pecuniary and for non-pecuniary reasons. As for the pecuniary reasons, in the basic model these refer to the low level of benefits relative to wages. If account is taken of wage increases during employment then generally the estimated difference between benefits and initial-wage offers is much smaller. However, the prospect of wage increases causes the unemployed searcher to set a low reservation wage as well. The results imply that at an individual level a decrease in benefits is ineffective in reducing unemployment duration. The estimation results appear to be robust to varying certain assumptions.

University of Groningen

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