Search behaviour, transitions to nonparticipation and the duration of unemployment

van den Berg, G.

Publication date:
1988

Document Version
Publisher's PDF, also known as Version of record

...SITY
...8
...HOLIEKE UNIVERSITEIT BRABANT
POSTBOX 90153
5000 LE TILBURG
THE NETHERLANDS

DEPARTMENT OF ECONOMICS

RESEARCH MEMORANDUM

[redacted]

SEARCH BEHAVIOUR, TRANSITIONS TO NONPARTICIPATION AND THE DURATION OF UNEMPLOYMENT

Gerard J. van den Berg

FEW 338

SEARCH BEHAVIOUR, TRANSITIONS TO NONPARTICIPATION AND THE DURATION OF UNEMPLOYMENT

Gerard J. van den Berg
Department of Economics
Tilburg University

Using longitudinal micro data on unemployed individuals for 1983-1985 a structural job search model is estimated. The model allows for transitions from unemployment to nonparticipation. An extended version of the model deals with the influence of on-the-job search and prospective wage increases on search behaviour of the unemployed. The empirical results show that the probability of accepting a job offer is almost one for most unemployed individuals. A large portion of unemployment spells ends in a transition out of the labour force. The effects of changes in benefits on duration appear to be extremely small.

I am grateful to Arie Kapteyn, Wiji Narendranathan, Andrew Chesher, Geert Ridder, Stephen Nickell, Peter Kooreman, Mark Stewart and Maarten Lindeboom for their helpful comments. Financial support from the Netherlands Organization for the Advancement of Pure Research (ZWO) is acknowledged. This research is part of a research project included in the Specific European Community Action to Combat Poverty. The Netherlands Central Bureau of Statistics (CBS) provided the data.

keywords: job search theory, unemployment duration, nonparticipation, wage increases.

JEL classification: 210,810.

1. Introduction

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 nonparticipation. 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 appraoch, in which only hazards of the duration distribution are estimated (see e.g. Lancaster (1979)) seems to be replaced gradually by a structural approach. The latter way of modeling is characterized 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.

Up to now a small number of empirical studies using structural search models have been published (Yoon (1981), Lancaster & Chesher (1983), Lynch (1983), Narendranathan & Nickell (1985), Ridder & Gorter (1986), Wolpin (1987)) some of which use a very restricted model specification (notably the first three references). None of the papers referred to uses a model that allows for transitions from unemployment to nonparticipation. 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 accounts of transitions into nonparticipation because the data used are not rich enough to make the distinction between the states of unemployment and nonparticipation. 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 evidende that a large portion of the flow out of unemployment consists of transitions into nonparticipation (for a survey of the literature, see Micklewright (1988) who also forecefully argues that the state of nonparticipation should be incorporated in duration models of the labour market, especially if one is interested in the effects of benefits on unemployment duration). In the sample we use, almost 30% of all spells of unemployment ends up in a transition into nonparticipation. 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 e.g. 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 2 we discuss the specification of the basic search model. 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 3 contains a description of the data, and a discussion of the empirical implementation of the model. Section 4 deals with the estimation of the wage offer distribution. Section 5 gives the main results. We present estimates of the job offer arrival rate, the transition rate into nonparticipation, 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. White's Information Matrix test is used in order to check whether unobserved heterogeneity is present in the structural parameters.

In section 6 the construction of the extended model is described and the results of the estimation of the extended model are discussed. Section 7 concludes.

2. The model

2.1. *Job search theory and model specification*

Job search theory describes the behaviour of unemployed individuals who are searching sequentially for jobs until a suitable one has been found (for surveys, see Mortensen (1986) or McKenna (1985)). Job offers arrive randomly in time at the arrival rate λ. Such job offers are random drawings (without recall) from a wage offer distribution $F(w)$. During unemployment a benefit b is received. The variables λ, b en w are measured per unit time period. Unemployed individuals aim at maximization of their expected discounted lifetime utility (over an infinite horizon). For now 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:

$$\text{utility (income} = x, \text{ state} = \text{employment)} \quad = \quad v^{*}.u(x)$$

$$\text{utility (income} = x, \text{ state} = \text{unemployment)} \quad = \quad v.u(x)$$

The function u is increasing in its argument and may take account of risk aversion. We normalize 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 λ, b, u, v and $F(w)$ to be independent of unemployment duration and calender 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, λ, 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 e.g. Lancaster & Chesher (1983) and Narendranathan & Nickell (1985)). That is, when estimating a nonstationary 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 nonstationarity in job search theory, see van den Berg (1987)) In section 4 we return to the effects that the presence of nonstationarity might have on the estimation results.

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

$$(2.1) \qquad \theta = \lambda\bar{F}(\varphi) \qquad\qquad \bar{F} = 1-F$$

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. This may be the result of a personal decision of that individual e.g. if he decides to dedicate all his time to household activities. It can also be a forced transition, e.g. when he is conscripted or when he becomes disabled or when he retires. All these cases can be labeled as transitions out of unemployment into nonparticipation.

Flinn & 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 nonparticipants enters the equations that describe the behaviour of the unemployed. This implies that data on returns of nonparticipants are needed in order to estimate the model. Such data are not available. Therefore we adopt a reduced-form modeling of the transitions from unemployment into nonparticipation. Specifically, such transitions are assumed to occur according to a Poisson process with a parametrized transition rate ζ.

The optimal strategy of an unemployed individual depends on the expected utility of becoming a nonparticipant. 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 nonparticipant. We make the assumption

(2.2) $\quad Eu(x) = u(b)$

For a lot of cases the income flow after becoming a nonparticipant is close to the benefit level (e.g. 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 $Ex \approx b$ 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 nonparticipation except that equation (2.2) holds. In addition, we assume that the state of nonparticipation is absorbing and, for the moment, we assume that the nonpecuniary component of per-period utility in nonparticipation is the same as that in unemployment. As an additional condition for stationarity to hold we require that ζ is constant (though possibly different across individuals). Again this may not be very realistic. Individuals may enter nonparticipation 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.

In appendix 1 we prove that the reservation wage φ which characterizes the optimal strategy in the model satisfies the following equation

(2.3) $$u(\varphi) = v.u(b) + \frac{\lambda}{\rho+\zeta}. \int_{\varphi}^{\infty} (u(w)-u(\varphi))dF(w)$$

The exit rate out of unemployment is equal to the sum of θ and ζ, with θ given by equation (2.1). Because θ and ζ 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 $\theta + \zeta$.

2.2. *An alternative interpretation*

It can be argued that the modeling 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 & Nickell (1985) constructued a search model that deals with this. Job vacancies arrive according to a Poisson process with arrival rate q_1. A vacancy is characterized 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 subsection 2.1 is equivalent to the model described here. To see this, equate

$$(2.4) \qquad \lambda = q_1 \int_0^\infty q_2(\omega)dG(\omega)$$

$$(2.5) \qquad F(w) = \frac{\int_0^w q_2(\omega)dG(\omega)}{\int_0^\infty q_2(\omega)dG(\omega)}$$

Consequently, the estimation results of the original model can be reinterpreted according to equations (2.4) and (2.5). Narendranathan & Nickell (1985) make the convenient assumption that

$$(2.6) \qquad q_2(w) = q_3(w).q_4$$

in which q_3 depends on w only, while q_4 represents the dependence of q_2 on personal characteristics. If (2.6) holds then $F(w)$ in (2.5) 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.

3. The data

3.1. *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. As of April 1984 a random sample of about 12000 individuals is being interviewed twice a year (in April and October). At every interview except the first one, respondents are asked to recall their labour market history for the past 6 months. At the first interview this 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 of 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 5.

Of the backward and forward recurrence times, 64% and 39% are censored in the sense that it is only known that the realized 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 nonparticipants. This means that according to their own perception they were not unemployed and searching for a job anymore though they weren't employed either. The state of nonparticipation covers a wide range of activities like being conscripted, being disabled, being retired, doing unpaid work in the household, being in full-time training and just doing nothing. The limited amount of observations in the sample prohibits a subdivision of the state of nonparticipation into different states. Note that in some cases nonparticipants can receive unemployment insurance benefits.

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.

3.2. *Likelihood function*

In our stationary model the backward and forward recurrence time and the state of destination given exit from unemployment are stochastically independent (see e.g. 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 $\vartheta/(\vartheta+\zeta)$. The forward recurrence time has an exponential distribution with parameter $\vartheta + \zeta$. 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 τ and t, respectively. The state of destination is denoted as ε with $\varepsilon = 1$ if the state is employment and $\varepsilon = 0$ if the state is nonparticipation. The occurrence of censoring and the occurrence of the so-called memory problems are taken to be exogenous. If τ is missing then this is taken to be exogenous as well.

First consider the state of destination. Let $c_1 = 1$ if τ is censored and $c_1 = 0$ otherwise. Let $c_2 = 1$ if τ is missing and $c_2 = 0$ otherwise. The part of the individual log-likelihood contribution L due to the state of destination is L_1.

$$(3.1) \qquad L_1 = (1-c_2)(1-c_1)(\varepsilon \log \vartheta+(1-\varepsilon)\log \zeta-\log(\vartheta+\zeta))$$

So if τ is censored or missing then ε is not observed and consequently $L_1 = 0$.

Next consider the backward recurrence time. Let $c_3 = 1$ if t is censored and $c_3 = 0$ otherwise. The part of L due to t is L_2,

$$(3.2) \qquad L_2 = (1-c_3).\log(\vartheta+\zeta) - t.(\vartheta+\zeta)$$

If no memory problems are present then the part of L due to τ can be obtained by replacing in equation (3.2) $1 - c_3$ 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)st (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 some weeks before the (n+1)st interview that date of the transition will be reported more accurately than when the spell ends some weeks 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 latest of both interviews. 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, than 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 τ_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 τ is L_3,

$$\begin{aligned} L_3 = {} & (1-c_2)[(1-c_1).\{(1-c_4)(\log(\vartheta+\zeta)-\tau.(\vartheta+\zeta)) \\ & + c_4.(\log(e^{-(\vartheta+\zeta)\tau} - e^{-(\vartheta+\zeta)(\tau+\tau_1)}))\} \\ (3.3) \qquad & + c_1.\{-\tau(\vartheta+\zeta)\}] \\ = {} & (1-c_2)[-\tau(\vartheta+\zeta) + (1-c_1)(1-c_4)\log(\vartheta+\zeta) \end{aligned}$$

$$+ (1-c_1).c_4.\log(1-e^{-(\vartheta+\zeta).\tau_1})]$$

It is likely that similar to the occurrence of memory problems in the reported values of τ 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_1 denote the length of that nine month period. Equation (3.2) has to be modified to

$$(3.4) \quad L_2 = (1-c_3)(\log(\vartheta+\zeta) - t.(\vartheta+\zeta)) - c_3.t_1.(\vartheta+\zeta)$$

The log-likelihood contribution L of an individual with known c_1, c_2, c_3, c_4, t, τ and ε is given by the sum of the right-hand sides of equations (3.1), (3.3) and (3.4). The structural parameters and functions of the job search model $(u,v,\rho,\lambda,F(w))$ enter the likelihood via ϑ (see equations (2.1) and (2.3)). The parameter ζ enters L both directly and indirectly via ϑ.

3.3. *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 parametrizations. As for the wage offer distribution however this will be done in section 4 because that section is devoted entirely to the estimation of F(w).

The job offer arrival rate λ and the transition rate into nonparticipation ζ 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 e.g. because they give an indication of the productivity of the job searcher. Examples are level of education (we distinguish between five levels), age, nationality, whether the individual has had a job before (this was being 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 nonparticipation, 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.

Similarly to Narendranathan & Nickell (1985) and Ridder & Gorter (1986) the utility function of income u is taken to be logarithmic. The subjective rate of discount ρ is fixed at 10% per year. In section 5 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 ρ.

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 θ with income of other household members included as a regressor in log θ 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-Rapson algorithm. Because of the assumptions that were made on the functional forms of F(w) (see section 4) and u, it follows that equation (2.3) can be rewritten as an equation that can be solved numerically for φ with a high level of precision. Via equation (2.1) the likelihood contributions can then be calculated as a function of the parameters.

4. The wage offer distribution

4.1. *Estimation strategy*

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 φ (Flinn & Heckman (1982)). Combining such data with duration data makes it possible to estimate F(w) jointly with the other parameters in the model. However, as we saw in subsection 3.1, in our sample there are only 79 transitions from unemployment into employment. 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) a priori using data on individuals who were employed at the date of the first interview. Analogous to Narendranathan & Nickell (1985) the 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 (e.g. Kiefer & Neumann (1979)) assumed this to be equivalent to $w \geq \varphi$, 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 & 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 μ and σ^2; $\mu = x_1'\eta$ with x_1 observed. The unobserved variable y^* is assumed to depend on a linear combination of observed exogenous variables x_2. This gives the wage model

(4.1) $\log w = x_1'\eta + \varepsilon_1$ $\qquad\qquad$ w observed $\Leftrightarrow y^* > 0$

(4.2) $y^* = x_2'\beta^* + \varepsilon_2$

$$\begin{bmatrix}\varepsilon_1\\ \varepsilon_2\end{bmatrix} \sim N\left[0, \begin{bmatrix}\sigma^2 & \sigma_{12}\\ & \sigma_2^2\end{bmatrix}\right]$$

Equation (4.2) can be interpreted as a reduced form description of the way factors x_2 influence the probability of being employed. Obviously every factor that influences w, influences y^* as well. Therefore the variables in x_1 are included in the set of variables in x_2. We are only interested in η and σ^2, so the identifying restriction $\sigma_2 = 1$ is harmless.

4.2. *Empirical implementation and results*

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. Consider the way such segments can be defined. For the purpose of predicting wage offer distributions it is obvious that the explanatory variables appearing in the wage model must be observed both for employed respondents who provide data on observed wages, and for unemployed respondents. The same holds for variables defining the segments. For unemployed individuals there is no information on their previous job, and new entrants have no previous job at all. Therefore, segments are defined using data on the level and type of education.

Five levels of educations labelled 1 to 5 (from low to high) are distinguished. For every level a distinction is made between two types of education: technically (including economic) and socially oriented education. Within segments μ is made dependent on age. This is very much the best that can be done given the limited availability of productivity indicators that are observed for both unemployed and employed individuals in the data set. Also, a more detailed classification into segments results in very small numbers of wage observations from some segments.

In order to facilitate the estimation of the wage model (in particular equation (4.2)) data on unemployed individuals ($y^* \leq 0$) have been used in addition to data on employed individuals ($y^* > 0$). For reasons of simplicity the possibility of transiting into nonparticipation is disregarded in this section. Analogous to the estimation of the main model attention is restricted to data on male individuals aged between 17 and 65. Wages are net weekly wages.

The wage model has been estimated by ML using the BHHH algorithm, for every segment. Tests show that in accordance with prior beliefs there is no difference between the estimates of equation (4.1) for different types of education given that the level of education equals 2 or 3. For the lowest level the data set does not provide information on the type of education. Therefore the technical and social segments were aggregated when estimating the wage model for the levels 1, 2 and 3. Tests also show that for the levels 4 and 5 the parameter σ^2 does not depend on the type of education, so we imposed this as a restriction. For every segment the covariance σ_{12} turned out to be insignificantly different from zero at the 10% level. This means that the events as captured by the latent variable y^* have no significant influence on the wage level. (This is a result which is frequently encountered in the literature, see e.g. Van Opstal & Theeuwes (1986) and Narendranathan & Nickell (1985)) Therefore equation (4.2) is dropped and F(w) is estimated by OLS on equation (4.1) using data on employed individuals only.

Table 1 presents the estimation results. Figures 1 and 2 show the estimated mean wage offers as a function of age. For every segment this is a concave function. Since we are dealing with cross-sectional data this is to be interpreted as a cohort effect rather than a life-cycle effect. For most ages the mean wage offer is increasing in the level of education. Further the technical type of education has always larger mean wage offers than the social type has. The variance of log wage offers is increasing with the level of education. For the segment with level = 5 and type = social there are only data available on middle-aged employed individuals. The extrapolation of $\hat{\mu}$ to low and high ages results in very low values of $\hat{\mu}$ for these ages.

Table 1. Parameters estimates for the wage offer distribution

level of education	estimates
1	$\mu = \underset{(0.1)}{-0.34} + \underset{(2.7)}{3.84x} - \underset{(2.6)}{0.46x^2}$
	$\sigma = 0.19$, $n = 171$
2	$\mu = \underset{(2.2)}{-3.37} + \underset{(5.9)}{5.10x} - \underset{(5.5)}{0.68x^2}$
	$\sigma = 0.20$, $n = 258$
3	$\mu = \underset{(1.6)}{-2.50} + \underset{(5.1)}{4.57x} - \underset{(4.8)}{0.59x^2}$
	$\sigma = 0.23$, $n = 646$
4	$\mu = \underset{(0.3)}{-2.59} + \delta(\underset{(0.5)}{4.91} + \underset{(0.5)}{1.74x} - \underset{(0.4)}{0.16x^2}) +$
	$+(1-\delta).(\underset{(1.1)}{4.71x} - \underset{(1.1)}{0.61x^2})$
	$\sigma = 0.26$, $n = 203$
5	$\mu = -29.07 + \delta(18.48 + 8.76x - 1.10x^2)$
	$+(1-\delta).(\underset{(1.6)}{18.79x} - \underset{(1.6)}{2.47x^2})$
	$\sigma = 0.26$, $n = 85$

$x = \log$ (age)

n = number of individuals

t-ratios in parentheses

$\delta = 1$ if type = technical and $\delta = 0$ otherwise

However, the unemployed individuals in this segment are all middle-aged as well. Therefore we have confidence that for these individuals the prediction of $F(w)$ is reliable. The wage offer distribution of an unemployed individual with characteristics x and parameters η and σ^2 associated with the segment he can be ascribed to, is predicted as being lognormal with parameters $\hat{\eta}'x$ and $\hat{\sigma}^2$. The predicted $F(w)$ are plugged in when estimating the structural model.

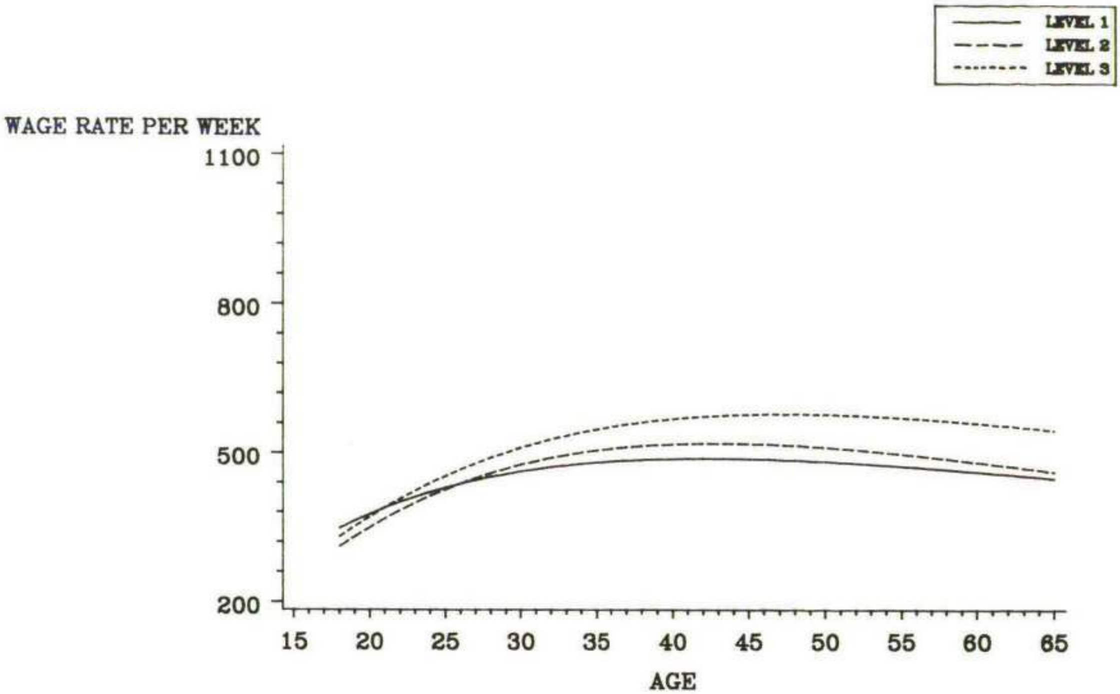

FIGURE 1

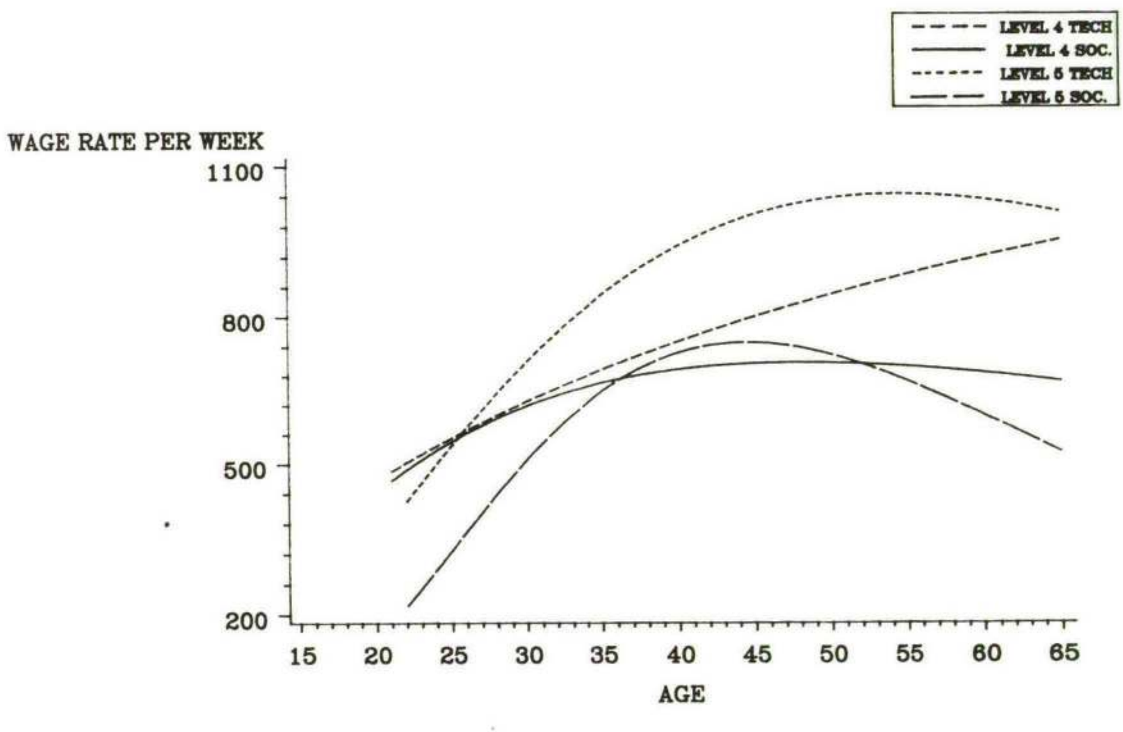

FIGURE 2

In terms of the alternative interpretation of subsection 2.2 the estimation of equation (4.1) 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 (2.5) and (2.6)),

$$(4.3) \quad \frac{\int_0^w q_3(\omega)dG(\omega)}{\int_0^\infty q_3(\omega)dG(\omega)} \qquad w \geq 0$$

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 6 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.

5. Results

5.1. *Parameter estimates*

The parameter estimates for the structural model described in subsections 2.1 and 3.3, 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 1, 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 due to the fact that such unemployed individuals 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 λ. Other indicators of the tightness of the labour market like the local UV ratio performed even worse. Van Opstal & 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 anymore. 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 λ with a factor of almost 1.5. The separate age coefficients in λ are not significant. However, a Likelihood Ratio test of the hypothesis that all age coefficients equal zero leads to a rejection at the 10% level. In section 3 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 subsection 2.2) λ 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_2. 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 nonparticipation, 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 be 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 $20.4 >> \chi_1^2(0.95)$ so $v = 1$ is strongly rejected.

Table 2. Parameter estimates for the search model

variable/parameter	coefficient	(t-ratio)
(i) *job offer arrival rate*		
constant	-6.08	(6.4)
Dutch	0.55	(1.3)
education: level 2	0.91	(3.3)
education: level 3	1.17	(3.6)
education: level 4	1.74	(2.8)
education: level 5	1.97	(2.8)
age category 18-23	0.68	(1.4)
age category 24-29	0.50	(1.2)
age category 30-45	0.16	(0.4)
new entrant	-0.82	(1.5)
head of household	-0.03	(0.1)
married	0.78	(2.5)
log (1 + # working in household)	1.03	(3.0)
local % unemployment rate	-0.04	(1.1)
(ii) *rate of transition into nonparticipation*		
constant	-4.91	(16.4)
age category 18-23	-0.41	(0.8)
age category 24-29	-1.06	(2.3)
age category 30-45	-1.39	(2.9)
new entrant	0.66	(1.4)
(iii) *disutility of unemployment*		
v	0.74	(5.2)
Log likelihood = -898.23		

5.2. *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. Table 3 presents sample averages of the estimates of λ, $\bar{F}(\varphi)$ and ζ for different age categories and levels of education. The expected numbers of job offers and transitions into nonparticipation in a year can be obtained by multiplying the numbers in the λ and ζ row by 52.1. What strikes most is that in most cases $\bar{F}(\varphi)$ 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 (v.u.(b)) relative to employment (u(w)). 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 0.9. 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 $\varphi < b$.

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 nonparticipation. In other words, without such transitions the durations of unemployment for such individuals would be approximately twice as long. See also figures 3 and 4 in which the escape rate out of unemployment is split in its two parts θ and ζ.

Since the model does not allow for nonstationarity, 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 θ is a decreasing function of duration. On the other hand, benefits generally decrease during unemployment, which ceteris paribus makes θ an increasing function of duration. One possible explanation for a decreasing θ is that the job offer arrival rate decreases sharply during unemployment e.g. as a consequence of a scar effect of being unemployed for a long time, and that this decrease of λ offsets the increase in $\bar{F}(\varphi)$. If θ 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 amount of long-term unemployed individuals. Further, if both λ and b decrease during unemployment then φ also decreases. So if nonstationarity is present in reality in the sense that b, λ, φ and θ all decrease, then λ, φ and θ are under-estimated in the sense that shortly after the inflow into unemployment these variables are larger than estimated. Another kind of nonstationarity is present if the transition rate into nonparticipation increases as a function of duration e.g. as a result of a discouraged worker effect. By analogy of the argument pointed out above it may be expected that in such a case ζ is over-estimated for individuals who are short-term unemployed.

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 θ, 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 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. The reasons are clear from table 3. A small decrease in benefits decreases reservation wages, but reservation wages are generally located in the left part of the left tail of F(w), so $\bar{F}(\varphi)$ is almost constant on a small interval around φ. In other words the decrease in benefits does not increase the proportion of acceptable jobs substantially. Consequently, the transition rate from unemployment to employment does not increase very much. 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 φ forced by a large decrease in b does not help much. The expected duration is bounded from below by $1/(\lambda+\zeta)$. 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, Nickell & Stern (1985)). Also, if there is an element of choice as to whether to become a nonparticipant or not, then a cut in benefits may have an effect on ζ. The sign of this effect depends among other things on the dependence of the distribution of income of nonparticipants on the level of benefits. If benefits are decreased whereas the incomes of nonparticipants like conscripts and disabled remain unchanged then equation (2.2) does not hold anymore. Therefore an investigation of the relation between b and ζ should be made in a wholly structural model setting and is beyond the scope of this paper. Inclusion of log (benefits) as a regressor in log ζ resulted in a highly insignificant parameter estimate of -0.14 ($t = 0.3$), all other things being almost identically equal.

From the results it is also clear that at an individual level additional educational training increases labour market opportunities.

5.3. *The model specification revisited*

In this subsection it is examined whether the results are sensitive with respect to changes in some of the assumptions made. As for changes in the way jobs are characterized in the model (infinite duration, constant wages) we refer to section 6 in which estimation results are presented for an extended model that deals with this.

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 e.g. the case in which unobserved heterogeneity is present in λ. We may rewrite λ as a product

Table 3. Probabilities and expectations

	18-23	24-29	30-45	46-64	average
(i) *by age category*					
age category	18-23	24-29	30-45	46-64	average
λ (expected number of offers in a week)	0.012	0.016	0.012	0.008	0.012
$\bar{F}(\varphi)$ (proportion of offers acceptable)	0.99	0.94	0.96	1.00	0.97
ζ (expected number of transitions into nonparticipation in a week)	0.007	0.003	0.002	0.007	0.004
(ii) *by level of education*					
level of education	1	2	3	4	5
λ	0.004	0.014	0.018	0.024	0.033
$\bar{F}(\varphi)$	1.00	0.98	0.94	0.89	0.82
ζ	0.004	0.004	0.004	0.004	0.003

$$(5.1) \quad \lambda = \nu . \exp(x'\beta)$$

in which the random term ν represents unobserved heterogeneity in λ across the population. Substitution of equation (5.1) in equations (2.3) and (2.1) reveals that the hazard cannot be written explicitly as an analytical function of ν. Therefore, calculating the unconditional (on ν) duration density by integrating the conditional density with respect to the density of ν will be very complicated. There is however a way of testing unobserved heterogeneity which does not require specifying a more general class of models. According to e.g. Chesher & Spady (1988) White's Information Matrix (IM) test which is generally considered to be an omnibus test for misspecification can be interpreted as a test for unobserved heterogeneity across individuals in a model of the behaviour of individual agents. In order to detect in which parts of the model unobserved heterogeneity

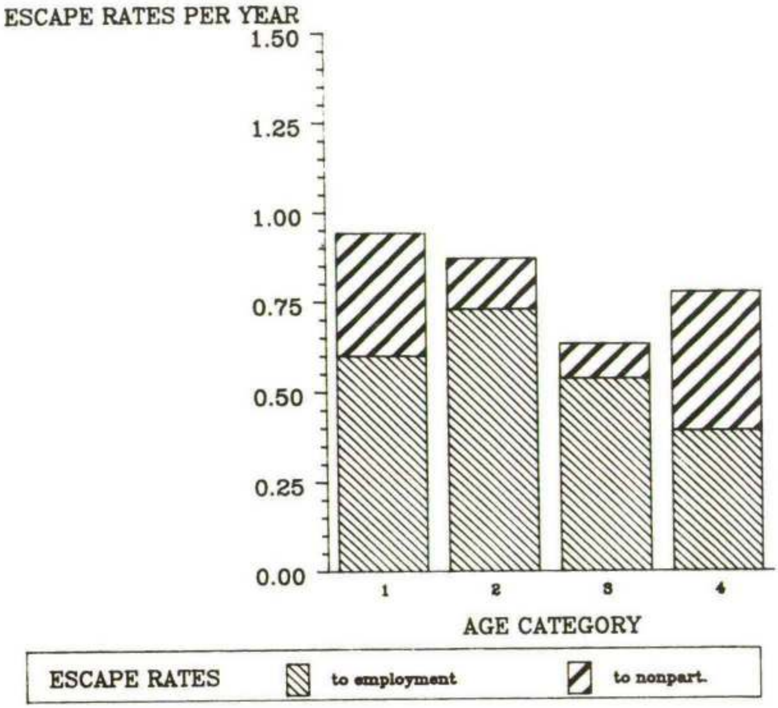

FIGURE 3

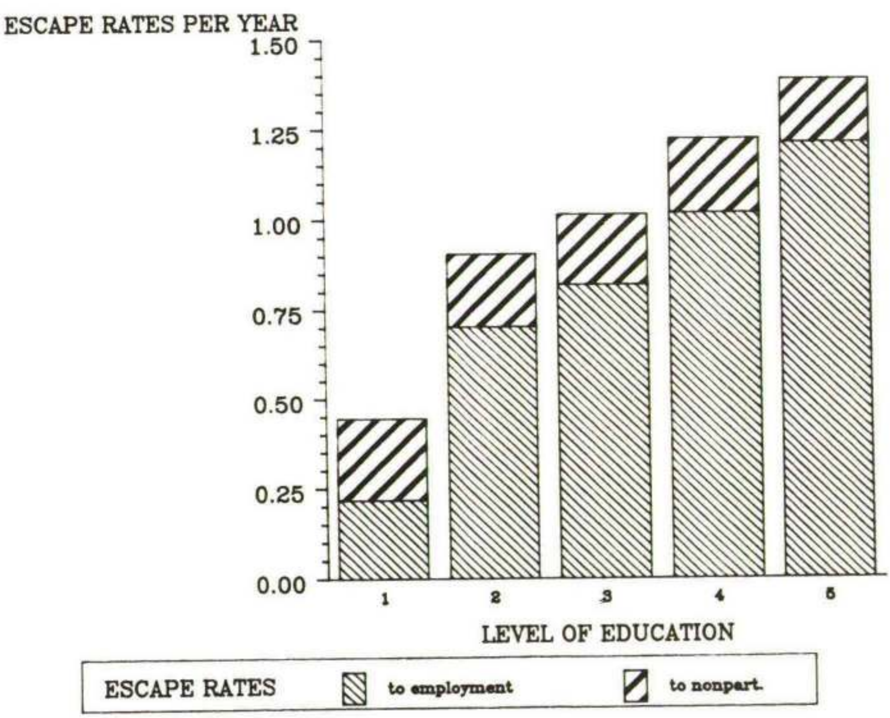

FIGURE 4

Table 4. Elasticities with respect to benefits

(i) *by age category*					
age category	18-23	24-29	30-45	46-64	average
$\frac{\partial \log \varphi}{\partial \log b}$ (reservation wage)	0.36	0.24	0.25	0.46	0.30
$\frac{\partial \log \vartheta}{\partial \log b}$ (hazard)	-0.01	-0.05	-0.04	-0.00	-0.03
$\frac{\partial \log d}{\partial \log b}$ (expected duration)	0.01	0.05	0.03	0.00	0.03
(ii) *by level of education*					
level of education	1	2	3	4	5
$\frac{\partial \log \varphi}{\partial \log b}$	0.44	0.24	0.23	0.19	0.16
$\frac{\partial \log \vartheta}{\partial \log b}$	0.00	-0.03	-0.06	-0.07	-0.11
$\frac{\partial \log d}{\partial \log b}$	0.00	0.03	0.05	0.07	0.10

d equals the expected duration of unemployment.

may be present we performed IM tests on the set of parameters which constitutes λ, on the set which constitutes ζ and on v. Because the degrees of freedom get very large if all elements of the IM are used for the test

Table 5. Information Matrix tests

parameters	test statistic	degree of freedom	critical level
λ	18.9	14	23.7
ζ	6.9	5	11.1
v	5.8	1	3.8

we restricted ourselves to diagonal elements of the IM. Table 5 reports the test statistics along with the 5% critical levels of the corresponding limiting chi-square distribution. The results show that the hypothesis of no unobserved heterogeneity cannot be rejected for λ and ζ. The test statistic for v indicates that individuals are heterogeneous with respect to the disutility of unemployment. According tot Chesher & Spady (1988) the IM test based on the chi-square distribution generally has excessive size even in quite large samples. However, the test result on v is plausible in the sense that v is the only estimated exogenous variable which is not parametrized. It thus seems natural to extend the model by making v a function of observable individual characteristics. Also, one might ask why ρ is not estimated and why u is not parametrized e.g. by assuming it to be a one-parameter CARA utility function. Though such extensions do not raise identification problems in the statistical sense, it appeared 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 φ is small with respect to most wage offers, which implies that $f(\varphi)$ is small so small changes in φ given values of λ, ζ and $F(w)$ do not affect the value of the likelihood function much. Secondly, u, ρ and v enter the likelihood only via φ. Therefore the correlation between estimates of parameters of u, v and ρ will be very high.

In the empirical model v is the only parameter that enters the likelihood via φ only. The discussion in the previous paragraph suggests

Table 6. Alternative values of ρ

rho (per year)	log-likelihood value	$\hat{v}$
5%	-898.23	0.67
10%	-898.23	0.74
15%	-898.26	0.78

that $\hat{v}$ might be biased if u is misspecified or if ρ has the wrong value. This is investigated by re-estimating the model with different u and ρ. Table 6 presents some results for alternative values of ρ. The estimates for λ and ζ hardly differ from the original results. The differences in the value of ρ are absorbed by $\hat{v}$, higher values of ρ resulting in higher values of $\hat{v}$ thus holding φ and therefore the fit of the model constant. Still, throughout the range of acceptable values of ρ, $\hat{v}$ is significantly smaller than 1 according to LR tests at the 1% level. Even in the limiting case of $\rho = \infty$ the estimate of v is significantly smaller than 1 ($\hat{v}$=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 maximizing 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 φ for $v = 0$ is ceteris paribus lower than the corresponding level in the risk-neutral case.

In section 2 we stated the assumptions that equation (2.2) holds and that the non-pecuniary utility of being a nonparticipant 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 nonparticipation 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 (2.3) has to be replaced by

$$(5.2) \qquad \frac{\zeta \cdot v_1 \cdot \dfrac{Eu(x)}{u(b)} + \rho v_2}{\zeta + \rho}$$

in order to obtain the equation for the optimal rservation wage. So then $\hat{v}$ represents the estimate of expression (5.2). It follows that

$$v_1 . Eu(x) > v_2 . u(b) \quad \Leftrightarrow \quad \hat{v} > v_2$$

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. Note that in the more general model in which it is allowed that $v_1 \neq v_2$ or $Eu(x) \neq u(b)$, the IM test result on v can be interpreted as evidence that individuals are heterogeneous with respect to expression

(5.2). This result is not surprising because individuals differ with respect to the values of b and ζ.

In section 3 we discussed the so-called memory problems. There it was argued that values of 3 and 9 months for τ_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 τ_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 U.K. the entry rate was fairly constant between 1967 and 1983 apart from an increase in 1979-1981. In the absence of reliable Dutch data we examine the sensitivy 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 appendix 2 the appropriate likelihood is derived. The main effect of the alternative assumption on q on the estimation results is that the exit rate out of unemployment $\vartheta + \zeta$ is estimated to be 13% larger. However, ϑ and ζ are still very small, and v, $\bar{F}(\varphi)$ and the elasticities are insensitive to the change in the assumption on q. Thus, the main results and conclusions from subsections 5.1 and 5.2 do not appear to be sensitive to a priori reasonable changes in the assumptions about the time pattern of the entry rate into unemployment.

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 4. 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 $\hat{v}$, a value of 1.2 times the original E(w) resulting in $\hat{v} = -0.20$ and a value of 0.8 times the original E(w) resulting in $\hat{v} = -0.45$. Consequently, the main conclusions are insensitive with respect to small misspecifications in the location of $F(w)$.

6. An extended model

6.1. *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 preceeds 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 wager without intervening spells of unemployment (on-the-job search). As a stylized description of this we assume that the wage pattern during employment is characterized by $w(t)$ giving the wage rate as a function of the time t that one is employed conditional on the initial wage $w(0)$.

$$(6.1) \quad w(t) = w(0).e^{\alpha t}$$

in which α does not depend on $w(0)$ or t or on the duration of unemployment preceeding employment. Through it is conceivable that mechanisms linking α, t and $w(0)$ 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 appendix 3 it is proven that the reservation wage φ corresponding to the optimal strategy satisfies

(6.2) $$\log \varphi = v.\log b + \frac{\lambda}{\rho+\zeta}.\frac{\rho}{\rho+s}.\int_{\varphi}^{\infty}(\log w - \log \varphi)dF(w) - \frac{\alpha}{\rho+s}$$

$F(w)$ is the distribution of initial-wage offers, which is the distribution from which the $w(0)$ are drawn. Note that the derivative of φ with respect to α is negative. If α 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 $w(0)$) in order to obtain a higher income in the future.

The estimation of $F(w)$ has to be reconsidered because in section 4 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 μ and σ^2. Thus, table 1 gives estimates of these parameters. The distribution $F(w)$ of initial-wage offers has to be recovered from the distribution of current wages. In appendix 4 it is shown that $F(w)$ can be approximated by a lognormal distribution with parameters $(\mu + \log(s-\alpha) - \log s)$ and σ^2,

(6.3) $$F(w) \approx LN(\mu + \log \frac{s-\alpha}{s}, \sigma^2)$$

This requires $s > \alpha$. The approximation is good for $s \gg \alpha$.

6.2. *The results*

The approximation in equation (6.3) is used to obtain a priori estimates of the individual distribution functions $F(w)$. The results in table 1 provide the individual values of μ and σ^2. The parameter α 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 subsection 3.2 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 (e.g. because of neglected unobserved heterogeinity) but we believe that for our purposes it is accurate enough.

From equation (6.3) it can be deduced that the expectation and the standard deviation of F(w) are $100.(\alpha/s)\% = 28\%$ smaller than those obtained in section 4. The sample average of the probability that a random initial-wage offer exceeds the benefit level is 0.61 as opposed to 0.91 in case F(w) is estimated like in section 4.

Table 7. Estimates for the extended search model

variable/parameter	coefficient	estimates for the basic model
v	0.83	0.74
λ	0.012	0.012
ζ	0.004	0.004
$\bar{F}(\varphi)$	0.98	0.97
$\partial \log \varphi/\partial \log b$	0.49	0.30
$\partial \log \theta/\partial \log b$	-0.04	-0.03
$\partial \log d/\partial \log b$	0.03	0.03

Table 8. Alternative values of ρ and α

rho (per year)	alpha (per year)	log-likelihood value	$\hat{v}$
10%	3%	-898.50	0.84
10%	4%	-898.50	0.83
10%	5%	-898.49	0.82
5%	4%	-898.58	0.85
15%	4%	-898.50	0.83

The estimates and t-ratios of the parameters of λ and ζ differ hardly 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 7). $\bar{F}(\varphi)$, λ and ζ 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 >> $\chi_1^2(0.99)$). 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. This is basically a consequence of the large value of $\bar{F}(\varphi)$.

From table 8 we infer that the results are insensitive to changing the assumptions on the values of ρ and α. The fit of the model is almost constant for the cases considered. Note that the sensitivity of $\hat{v}$ to changes in the value of ρ is less than in the basic model.

In sum, the main conclusions from section 5 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.

7. 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 nonparticipation. 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 1983-1985. 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.

Appendix 1

Derivation of equation (2.3)

Basically, the derivation proceeds along the lines of Lancaster & Chesher (1983)'s derivation of the reservation wage equation in a standard model with income maximization and $\zeta = 0$. First, consider a moment t at which an offer is pending. Let I_e denote the value at time t of following the optimal strategy. An acceptance policy can be characterized by a function p mapping $[0,\infty\rangle$ onto $[0,1]$ and giving for every w the probability that a wage offer w will be accepted. R is defined to be the return of rejecting the offer and behaving optimally afterwards. Because of the stationarity assumption I_e, p and R do not depend on t. Thus, at every moment at which an offer is pending, I_e denotes the present value of following the optimal strategy.

$$\text{(A1.1)} \qquad I_e = \sup_p \int_0^\infty \left[p(w)\,\frac{u(w)}{\rho} + (1-p(w)).R \right] dF(w)$$

It follows that the optimal acceptance policy p^* is given by

$$\text{(A1.2)} \qquad \begin{aligned} p^*(w) &= 1 \qquad && \text{if } u(w) \geq \rho.R \\ p^*(w) &= 0 && \text{otherwise} \end{aligned}$$

so p^* can be characterized by a reservation wage φ, satisfying

$$\text{(A1.3)} \qquad u(\varphi) = \rho.R$$

Thus (A1.1) can be written as

$$\text{(A1.4)} \qquad I_e = R + \frac{1}{\rho}.\int_\varphi^\infty (u(w)-u(\varphi))\,dF(w)$$

Let I_n denote the expected return at a moment at which a transition into nonparticipation occurs. From the assumptions about expected utility during nonparticipation it follows that

$$\text{(A1.5)} \qquad I_n = \int_0^\infty \int_0^\infty e^{-\rho t}.v.u(x)dt\ dH(x) = \frac{v.u(b)}{\rho}$$

in which $H(x)$ is the c.d.f. of income flows of nonparticipants. Let $k(\tau)$ denote the p.d.f. of the distribution of the waiting time at t until the next event (job offer or transition into nonparticipation) occurs. Because of the stationarity assumption $k(\tau)$, does not depend on t and is distributed exponentially with parameter $\lambda + \zeta$. If an event occurs, the probability that this event is a job offer is equal to $\lambda/(\lambda+\zeta)$. Now R can be written as

$$\text{(A1.6)} \qquad R = \int_0^\infty k(\tau)\left[\int_0^\tau v.u(b)e^{-\rho s}ds + e^{-\rho\tau}\left\{\frac{\lambda}{\lambda+\zeta}.I_e + \frac{\zeta}{\lambda+\zeta}.I_n\right\}\right]d\tau$$

which reduces to

$$\text{(A1.7)} \qquad R = \frac{v.u(b)+\lambda I_e+\zeta I_n}{\rho+\lambda+\zeta}$$

Substitution of (A1.3), (A1.4) and (A1.5) in (A1.7) gives the desired result. Note that for equation (2.3) to hold it is not necessary that the distribution of income flows of nonparticipants and the per-period utility function of nonparticipants, are independent of the time spent in the state of nonparticipation. What is essential is that the expected discounted lifetime utility at the moment that one becomes a nonparticipant I_n equals $v.u(b)/\rho$. Therefore equation (A1.5) can be replaced by

$$\text{(A1.8)} \qquad I_n = \int_0^\infty e^{-\rho t} \int_0^\infty v(t)u(x;t)dH(x|t)dt = \frac{v.u(b)}{\rho}$$

in which t denotes the duration in the state of nonparticipation; the definitions of $v(t)$, $u(x;t)$ and $H(x|t)$ are obvious.

The model can also be extended in another direction without changing the outcomes. From the examples of transitions from unemployment into nonparticipation it is clear that one can also expect transitions from employment into nonparticipation to be present in reality. If so then the unemployed individual is assumed to take account of this when determining his optimal strategy. Let transitions from employment into nonparticipation arrive according to a Poisson process with arrival rate ω. We assume that the expected discounted lifetime utility at the moment that one becomes a nonparticipant is independent of the origin state and is denoted by I_n. It can be proven that, instead of equation (2.3), the reservation wage satisfies

$$\text{(A1.9)} \quad u(\varphi) = \frac{1}{\rho+\zeta}\,(\rho.I_n(\zeta-\omega)+v.u(b).(\rho+\omega)) + \frac{\lambda}{\rho+\zeta}.\int_{\varphi}^{\infty}(u(w)-u(\varphi))dF(w)$$

If we impose that $\omega = \zeta$, that is, if we assume that the transition rate into nonparticipation is the same for employed and unemployed individuals then equation (A1.9) reduces to equation (2.3). This result holds regardless of the value of I_n as long as it is fixed. For our purposes it is even more interesting that if equation (A1.8) is substituted in equation (A1.9) this equation again reduces to equation (2.3). That is, the reservation wage does not depend on ω if (A1.8) holds.

Appendix 2

Likelihood function in case of a time-varying entry rate.

If the entry rate into unemployment is dependent on time then the backward recurrence time t no longer has an exponential distribution. Consequently the likelihood contribution L_2 (see equation (3.4)) has to be modified. From Ridder (1984), the density function $h(t|x)$ of t given time-independent personal characteristics x is given by

$$\text{(A2.1)} \quad h(t|x) = \frac{q(-t|x).e^{-\omega t}}{\int_0^{\infty} q(-s|x).e^{-\omega s}ds} \quad t \geq 0$$

in which $\omega \equiv \omega(x) \equiv \theta(x) + \zeta(x)$ and in which $q(-t|x)$ is the entry rate at t units of time before April 1984. In subsection 5.3 it is assumed that

$$\text{(A2.2)} \quad \begin{aligned} q(-t|x) &= q(0|x) && 0 \le t < t_2,\ t \ge t_3 \\ q(-t|x) &= 2.q(0|x) && t_2 \le t < t_3 \end{aligned}$$

with t_2 and t_3 equal to 16 and 52 months, respectively. The variable t is censored at t_1 (see subsection 3.2). By substituting (A2.2) in (A2.1), taking account of the censoring, and by taking the logarithm, the modified L_2 is obtained. This expression does not depend on $q(0|x)$.

Appendix 3

Proof of equation (6.2)

The line of argument and the notation of appendix 1 are followed. Equations (A1.7) and (A1.8) remain valid.
Equation (A1.1) is replaced by

$$\text{(A3.1)} \quad I_e = \sup_p \int_0^\infty \left[p(w).\left\{E_t\left[\int_0^t e^{-\rho\omega}u(e^{\alpha\omega}.w)d\omega + e^{-\rho t}.R\right]\right\} + (1-p(w)).R\right]dF(w)$$

The expectation E_t is taken w.r.t. the duration of employment. The reservation wage φ characterizes the optimal strategy,

$$\text{(A3.2)} \quad E_t \int_0^t e^{-\rho\omega}.u(e^{\alpha\omega}.\varphi)d\omega = E_t(1-e^{-\rho t}).R$$

Substitution in (A3.1) gives, noting that u is the logarithmic function,

$$\text{(A3.3)} \quad I_e = R + \frac{1}{\rho}E_t(1-e^{-\rho t}).\int_\varphi^\infty (\log w - \log\varphi)dF(w)$$

Equation (A3.2) can be simplified to

$$\text{(A3.4)} \quad E_t\left[\int_0^t \alpha\omega e^{-\rho\omega}d\omega\right] = E_t(1-e^{-\rho t}).\left[R - \frac{\log\varphi}{\rho}\right]$$

Because $t \sim$ exponential (s) it holds that

$$E_t(1-e^{-\rho t}) = \frac{\rho}{s+\rho}$$

$$E_t\left[\int_0^t \alpha\omega e^{-\rho\omega}d\omega\right] = \frac{\alpha}{(s+\rho)^2}$$

which gives

$$\text{(A3.5)} \quad \log \varphi = \rho.R - \frac{\alpha}{\rho+s}$$

$$\text{(A3.6)} \quad I_e = R + \frac{1}{\rho+s}.\int_{\varphi}^{\infty}(\log w - \log \varphi)dF(w)$$

Substitution of (A3.5), (A3.6) and (A1.8) in (A1.7) gives the desired result.

Appendix 4

Approximation of the distribution of initial-wage offers.

In order to avoid confusion between initial wages and current wages the latter is denoted by y and the former by w. The distribution over the population of completed durations of employment is exponential with parameter s. We observe a (cross-section) sample from the stock of the employed, which means that the durations of employment t are incomplete. However the entry rate into employment is time-independent due to the stationarity assumption. Therefore such incomplete durations have an exponential distribution with parameter s as well.

An observed (current) wage y is the product of two unobserved stochastic terms

$$\text{(A4.1)} \quad y = e^{\alpha t}.w$$

in which t denotes the incomplete duration of employment. Since t and w are independent the moments of y are easily expressed in terms of the moments of w, (A4.2) $\quad E(y) = E(e^{\alpha t}).E(w) = \frac{s}{s-\alpha}.E(w)$

(A4.3) $\quad \mathrm{var}(y) = (\frac{s}{s-\alpha})^2.\mathrm{var}(w) + (\frac{\alpha}{s-\alpha})^2.\frac{s}{s-2\alpha}\,E(w^2)$

Define $\xi = \alpha/s$. Equations (A4.2) and (A4.3) can be rewritten as

(A4.4) $\quad E(w) = (1-\xi).E(y)$

(A4.5) $\quad \mathrm{var}(w) = (1-\xi)^2.\mathrm{var}(y) + O(\xi^2)$

Consequently, if we use

(A4.6) $\quad w = (1-\xi).y$

in order to recover $F(w)$ from the distribution of y then the first moment of the distribution thus obtained is correct while the second central moment is correct up to the second order of α/s. For α small as compared to s the distribution of w based on equation (A4.6) is a good approximation of the true $F(w)$ though the variance of $F(w)$ is somewhat overstated. It is assumed that $y \sim LN(\mu,\sigma^2)$ so

(A4.7) $\quad (1-\xi).y \sim LN(\mu+\log(1-\xi),\sigma^2)$

Therefore $F(w)$ is approximated according to equations (A4.6) and (A4.7). From equation (A4.1) and from the assumptions on the parametric forms of the distributions of y and t the true $F(w)$ can be deduced. However this gives rather problematic results for the parametric forms that were chosen. Rather than modifying these choices we prefer to approximate $F(w)$ as set out in the previous paragraph.

References

Berg, G.J. van den (1987), 'Nonstationarity in job search theory.', Research Memorandum, Tilburg University.

Burdett, K., Kiefer, N.M., Mortensen, D.T. and Neumann, G.R. (1984), 'Earnings, unemployment, and the allocation of time over time.' *Review of Economic Studies*, vol. 51, pp. 559-578.

Chesher, A. and Spady, R. (1988), 'Asymptotic expansions of the Information Matrix test statistic.', Discussion Paper, Bristol University.

Coleman, T.S. (1983), 'A dynamic model of labor supply under uncertainty.', Mimeo, University of Chicago.

Flinn, C. and Heckman, J. (1982), 'New methods for analyzing structural models of labor force dynamics.' *Journal of Econometrics*, vol. 18, pp. 115-168.

Kiefer, N.M. and Neumann, G.R. (1979), 'Estimation of wage offer distributions and reservation wages.' In *Studies in the economics of search* (ed. S. Lippman and J. McCall), Amsterdam: North Holland.

Lancaster, T. (1979), 'Econometric methods for the duration of unemployment.' *Econometrica*, vol. 47, pp. 939-956.

Lancaster, T. and Chesher, A.D. (1983), 'An econometric analysis of reservation wages.' *Econometrica*, vol. 51, pp. 1661-1676.

Lynch, L.M. (1983), 'Job search and youth unemployment.' In *The causes of unemployment* (ed. C.Greenhalgh, R. Layard and A. Oswald), Oxford: Clarendon Press.

McKenna, C.J. (1985), *'Uncertainty and the labour market.'* Brighton: Wheatsheaf Books.

Micklewright, J. (1988), 'Unemployment duration: some modelling and policy issues for European countries.', Mimeo, University of London.

Mortensen, D.T. (1986), 'Job search and labor market analysis.' In *Handbook of labor economics* (ed. O. Ashenfelter and R. Layard), Amsterdam: North Holland.

Narendranathan, W. and Nickell, S. (1985), 'Modelling the proces of job search.' *Journal of Econometrics*, vol. 28, pp. 29-49.

Narendranathan, W., Nickell, S. and Stern, J. (1985), 'Unemployment benefits revisited.' *Economic Journal*, vol. 95, pp. 307-329.

Opstal, R. van and Theeuwes, J. (1986), 'Duration of unemployment in the Dutch youth labour market.' *De Economist*, vol. 134, pp. 351-367.

Pissarides, C. (1986), 'Unemployment and vacancies in Britain.' *Economic Policy*, October 1986.

Ridder, G. (1984), 'The distribution of single-spell duration data.' In *Studies in labor market dynamics* (ed. G. Neumann and N. Westergård-Nielsen), Heidelberg: Springer-Verlag.

Ridder, G. and Gorter, K. (1986), 'Unemployment benefits and search behaviour, an empirical investigation.', Report, University of Amsterdam.

Wolpin, K.I. (1987), 'Estimating a structural job search model: the transition from school to work.' *Econometrica*, vol. 55, pp. 801-818.

Yoon, B.J. (1981), 'A model of unemployment duration with variable search intensity.' *Review of Economics and Statistics*, vol. 63, pp. 599-609.

IN 1987 REEDS VERSCHENEN

242 Gerard van den Berg
Nonstationarity in job search theory

243 Annie Cuyt, Brigitte Verdonk
Block-tridiagonal linear systems and branched continued fractions

244 J.C. de Vos, W. Vervaat
Local Times of Bernoulli Walk

245 Arie Kapteyn, Peter Kooreman, Rob Willemse
Some methodological issues in the implementation of subjective poverty definitions

246 J.P.C. Kleijnen, J. Kriens, M.C.H.M. Lafleur, J.H.F. Pardoel
Sampling for Quality Inspection and Correction: AOQL Performance Criteria

247 D.B.J. Schouten
Algemene theorie van de internationale conjuncturele en strukturele afhankelijkheden

248 F.C. Bussemaker, W.H. Haemers, J.J. Seidel, E. Spence
On (v,k,λ) graphs and designs with trivial automorphism group

249 Peter M. Kort
The Influence of a Stochastic Environment on the Firm's Optimal Dynamic Investment Policy

250 R.H.J.M. Gradus
Preliminary version
The reaction of the firm on governmental policy: a game-theoretical approach

251 J.G. de Gooijer, R.M.J. Heuts
Higher order moments of bilinear time series processes with symmetrically distributed errors

252 P.H. Stevers, P.A.M. Versteijne
Evaluatie van marketing-activiteiten

253 H.P.A. Mulders, A.J. van Reeken
DATAAL - een hulpmiddel voor onderhoud van gegevensverzamelingen

254 P. Kooreman, A. Kapteyn
On the identifiability of household production functions with joint products: A comment

255 B. van Riel
Was er een profit-squeeze in de Nederlandse industrie?

256 R.P. Gilles
Economies with coalitional structures and core-like equilibrium concepts

274 Raymond H.J.M. Gradus
The net present value of governmental policy: a possible way to find the Stackelberg solutions

275 Jack P.C. Kleijnen
A DSS for production planning: a case study including simulation and optimization

276 A.M.H. Gerards
A short proof of Tutte's characterization of totally unimodular matrices

277 Th. van de Klundert and F. van der Ploeg
Wage rigidity and capital mobility in an optimizing model of a small open economy

278 Peter M. Kort
The net present value in dynamic models of the firm

279 Th. van de Klundert
A Macroeconomic Two-Country Model with Price-Discriminating Monopolists

280 Arnoud Boot and Anjan V. Thakor
Dynamic equilibrium in a competitive credit market: intertemporal contracting as insurance against rationing

281 Arnoud Boot and Anjan V. Thakor
Appendix: "Dynamic equilibrium in a competitive credit market: intertemporal contracting as insurance against rationing

282 Arnoud Boot, Anjan V. Thakor and Gregory F. Udell
Credible commitments, contract enforcement problems and banks: intermediation as credibility assurance

283 Eduard Ponds
Wage bargaining and business cycles a Goodwin-Nash model

284 Prof.Dr. hab. Stefan Mynarski
The mechanism of restoring equilibrium and stability in polish market

285 P. Meulendijks
An exercise in welfare economics (II)

286 S. Jørgensen, P.M. Kort, G.J.C.Th. van Schijndel
Optimal investment, financing and dividends: a Stackelberg differential game

287 E. Nijssen, W. Reijnders
Privatisering en commercialisering; een oriëntatie ten aanzien van verzelfstandiging

288 C.B. Mulder
Inefficiency of automatically linking unemployment benefits to private sector wage rates

289 M.H.C. Paardekooper
A Quadratically convergent parallel Jacobi process for almost diagonal matrices with distinct eigenvalues

290 Pieter H.M. Ruys
Industries with private and public enterprises

291 J.J.A. Moors & J.C. van Houwelingen
Estimation of linear models with inequality restrictions

292 Arthur van Soest, Peter Kooreman
Vakantiebestemming en -bestedingen

293 Rob Alessie, Raymond Gradus, Bertrand Melenberg
The problem of not observing small expenditures in a consumer expenditure survey

294 F. Boekema, L. Oerlemans, A.J. Hendriks
Kansrijkheid en economische potentie: Top-down en bottom-up analyses

295 Rob Alessie, Bertrand Melenberg, Guglielmo Weber
Consumption, Leisure and Earnings-Related Liquidity Constraints: A Note

296 Arthur van Soest, Peter Kooreman
Estimation of the indirect translog demand system with binding non-negativity constraints

IN 1988 REEDS VERSCHENEN

297 Bert Bettonvil
Factor screening by sequential bifurcation

298 Robert P. Gilles
On perfect competition in an economy with a coalitional structure

299 Willem Selen, Ruud M. Heuts
Capacitated Lot-Size Production Planning in Process Industry

300 J. Kriens, J.Th. van Lieshout
Notes on the Markowitz portfolio selection method

301 Bert Bettonvil, Jack P.C. Kleijnen
Measurement scales and resolution IV designs: a note

302 Theo Nijman, Marno Verbeek
Estimation of time dependent parameters in lineair models
using cross sections, panels or both

303 Raymond H.J.M. Gradus
A differential game between government and firms: a non-cooperative approach

304 Leo W.G. Strijbosch, Ronald J.M.M. Does
Comparison of bias-reducing methods for estimating the parameter in dilution series

305 Drs. W.J. Reijnders, Drs. W.F. Verstappen
Strategische bespiegelingen betreffende het Nederlandse kwaliteits-concept

306 J.P.C. Kleijnen, J. Kriens, H. Timmermans and H. Van den Wildenberg
Regression sampling in statistical auditing

307 Isolde Woittiez, Arie Kapteyn
A Model of Job Choice, Labour Supply and Wages

308 Jack P.C. Kleijnen
Simulation and optimization in production planning: A case study

309 Robert P. Gilles and Pieter H.M. Ruys
Relational constraints in coalition formation

310 Drs. H. Leo Theuns
Determinanten van de vraag naar vakantiereizen: een verkenning van materiële en immateriële factoren

311 Peter M. Kort
Dynamic Firm Behaviour within an Uncertain Environment

312 J.P.C. Blanc
A numerical approach to cyclic-service queueing models

313 Drs. N.J. de Beer, Drs. A.M. van Nunen, Drs. M.O. Nijkamp
Does Morkmon Matter?

314 Th. van de Klundert
Wage differentials and employment in a two-sector model with a dual labour market

315 Aart de Zeeuw, Fons Groot, Cees Withagen
On Credible Optimal Tax Rate Policies

316 Christian B. Mulder
Wage moderating effects of corporatism
Decentralized versus centralized wage setting in a union, firm, government context

317 Jörg Glombowski, Michael Krüger
A short-period Goodwin growth cycle

318 Theo Nijman, Marno Verbeek, Arthur van Soest
The optimal design of rotating panels in a simple analysis of variance model

319 Drs. S.V. Hannema, Drs. P.A.M. Versteijne
De toepassing en toekomst van public private partnership's bij de grote en middelgrote Nederlandse gemeenten

320 Th. van de Klundert
Wage Rigidity, Capital Accumulation and Unemployment in a Small Open Economy

321 M.H.C. Paardekooper
An upper and a lower bound for the distance of a manifold to a nearby point

322 Th. ten Raa, F. van der Ploeg
A statistical approach to the problem of negatives in input-output analysis

323 P. Kooreman
Household Labor Force Participation as a Cooperative Game; an Empirical Model

324 A.B.T.M. van Schaik
Persistent Unemployment and Long Run Growth

325 Dr. F.W.M. Boekema, Drs. L.A.G. Oerlemans
De lokale produktiestructuur doorgelicht.
Bedrijfstakverkenningen ten behoeve van regionaal-economisch onderzoek

326 J.P.C. Kleijnen, J. Kriens, M.C.H.M. Lafleur, J.H.F. Pardoel
Sampling for quality inspection and correction: AOQL performance criteria