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THE DIVERGING EFFECTS OF THE BUSINESS CYCLE ON THE EXPECTED DURATION OF JOB SEARCH

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Abstract

Efficiency wage theories of the labour market suggest that for a variety of reasons different groups of job seekers are not all affected in the same way by fluctuations in labour demand. Two testable predictions can be derived from those theories: first, the probability of finding a job will be affected more seriously for low qualified than for highly qualified job seekers, because the former cannot apply for lower ranked jobs. Secondly, job seekers of equal qualification will get better jobs in tight labour markets.

Both predictions are tested for the Dutch economy during the period 1981-1985. Individual transition rates from job search to employment are estimated. In a second model employment is classified by occupation to indicate the level of the job. Because explanatory variables on the aggregate and the individual level are combined, standard errors are normally underestimated. A method, similar to the two-error component model, is worked out to cope with this problem.

Estimation results show both phenomena to be relevant, with non-natives being the major victims of the increased selectivity of the market process.

1 INTRODUCTION

Efficiency wage theories of the labour market suggest that there is a discontinuity in wage formation.[1] As long as the market clearing wage is above a certain minimum level it is equal to the actual wage rate. In this case the wage rate can adjust more or less freely to shocks to the economy so as to maintain full employment. However, if market clearing requires a wage rate below this minimum, the economy will suffer unemployment. Shocks will be adjusted to by movements, not of relative prices, but of the level of employment. The wage level at which the economy switches over from price to quantity adjustment is the efficiency wage. It is the lowest wage profitable for employers to offer. Lower wages will lead to an increased wage bill due to a fall in labour productivity.[2]

If wages are restricted by the efficiency wage type of constraints then jobs are rationed: relative wages cannot perform their task of allocating workers to jobs completely. Employers can use criteria other than the minimum cost per efficiency unit of labour to determine which job seekers are to be hired.

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1 E.g. Weiss (1980).

2 For a review of efficiency wage type of models, see Yellen (1984).

For example, employers may dislike to employ non-natives. In a tight labour market with wages being market clearing, employers are kept from expressing this shared preference because the wages of non-natives will fall. Those employers whose discriminatory preferences are the weakest will employ non-natives and thereby gain a competitive advantage. If wages are prevented from falling by efficiency wages, then this mechanism breaks down. Employers can select the natives from the pool of job seekers without having to pay surplus wages. Discrimination is costless in the case of excess supply.[3)]

There are other mechanisms in efficiency wages models which lead to comparable results. Weiss[4)] suggests groups of job seekers having diverging efficiency wages per unit of effective labour. When labour demand falls, those with the highest efficiency wage are the first to be laid off. Akerlof and Yellen[5)] suggest that it is not the wage level but the difference between high and low pay within the same organisation which affects productivity. This imposes a minimum restriction only on the lowest wages. Therefore with excess supply, employment is rationed for low qualified workers alone.

These models are similar to Thurow's job competition model[6)] and to some aspects of the dual labour market hypothesis. There workers do not compete for the highest wage but for the best job. An increased rationing of high ranked jobs induces highly qualified workers to accept lower ranked jobs, thereby pushing lower qualified workers out of the market. The merit of efficiency wage theories compared to Thurow's model is that they offer explanations for wages being inflexible.

Concentrating on the process of job search two testable implications can be derived from these ideas. First, efficiency wage theories suggest that cyclical variations in market conditions do not affect the expected duration of search for all job seekers in the same way. Non-natives in particular will be affected more seriously, because employers can express their discriminatory feelings without cost during the downswing. But also the lower qualified will suffer more serious consequences of the cycle, because employers prefer higher qualified job seekers. With rationing of jobs this preference is costless, too.

Secondly, during the downswing job seekers with given characteristics like age, schooling, and labour market experience, will get lower ranked jobs than they get during the upswing of the cycle. The increased rationing blocks the entrance to jobs which were open to them when the labour market was more tight. Therefore, they are forced to accept lower ranked jobs.

The purpose of this paper is to establish both phenomena empirically. The deterioration of labour market conditions in the beginning of the eighties offers the opportunity to do this. In this paper the Dutch experience is examined. Unemployment in the Netherlands has risen from 9.2% in 1981 to 15.0% in 1983. This variation must be sufficient to identify both phenomena.

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3 Yellen (1984).

4 Weiss (1980).

5 Akerlof and Yellen (1988).

6 Thurow (1972).

I use micro data on job search activities between 1981 and 1985, which are obtained from a random sample of Dutch households. With these data the probabilities of finding a job can be estimated for various labour markets. These can be used to check the claim that the prospects of the least qualified job seekers are more seriously affected by the cycle than those of better qualified job seekers. Moreover, the data allow for a classification of jobs by occupation. These data can be used to analyse the allocation of jobs of different grades to job seekers. This enables me to search for the presence of the second phenomenon.

In Section 2, I sketch the analytical framework. This framework is based on notions from search theory and duration models. inference is based on individual transition rates from job search to employment under various labour market conditions.
In Section 3 the estimation procedure will be set out. Our data do not allow for the application of standard methods, so I have to devise a method to solve these problems. Furthermore, I shall argue that if explanatory variables on both the aggregate and the individual level are combined, standard errors of the variables on the aggregate level are underrated. I propose a specification and an estimation procedure to cope with this problem.
In Section 4 and 5 I discuss the empirical results: first the prospects of finding a job in general and secondly the probabilities by occupation are presented.

2 THE MODEL

The probability a job seeker will find a job during the next short interval of time Δt can be written as:

$$\Pr\left(t \le t^* \le t + \Delta t \mid t^* \ge t\right) = \lambda(t)\Delta t = \frac{f(t)\Delta t}{1 - F(t)} \tag{1}$$

where:

t = the time spent on searching up till the moment of observation (the incomplete duration of search);

t^* = the time spent on searching until the moment when the job seeker finds a job (the complete duration);

F(.),f(.) = respectively, the distribution and the density function of t^*.

$\lambda(t)$ is called the transition or hazard rate of job search to employment. The interval Δt is chosen so that the probability i finds two jobs within this interval can be ignored.
As is well-known, F(.) and f(.) are described by:

$$F(t) = 1 - \exp\left(-\int_0^t \lambda(s)ds\right) \tag{1a}$$

$$f(t) = \lambda(t)\exp\left(-\int_0^t \lambda(s)ds\right) \tag{1b}$$

The expected completed duration of search can be calculated from:

$$E(t^*) = \int_0^\infty t.f(t)\,dt = \int_0^\infty \exp\left(-\int_0^t \lambda(s)ds\right)dt \tag{2}$$

where the final step follows from integration by parts. In general $\lambda(t)$ will vary among individuals and during the elapsed duration of search t.[7] Because $\lambda(t)$ has to be positive by definition, it is convenient to use an exponential specification. Furthermore I assume the effects of t and of the characteristics of the individual to be multiplicatively separable. This model is known as the proportional hazard model:[8]

$$\lambda_i(t) = g(t)\exp\left(x_i'\beta\right) = \gamma\, t^{\gamma-1}\exp\left(x_i'\beta\right) \tag{3}$$

where:

x_i = a vector of characteristics of individual i

β = a vector of parameters describing the effect of each characteristic on the transition rate.

The final expression in (3) constitutes a convenient specification of g(.), which will be used throughout the paper. This is the basic framework, which has been used in many other papers on labour market dynamics. Two further elaborations have to be made in order to apply it to the problem in hand.

First, the way cyclical variations affect the transition rate has to be modelled. For expositional purposes I assume temporarily that the effects of personal characteristics and of the incomplete duration of search on the hazard rate can be neglected: $\lambda_i(t) = \lambda$.

Assume the search process on the labour market can be described with a Cobb-Douglas production function:

$$M_\tau d\tau = C.U_\tau^\varepsilon.V_\tau^\delta d\tau \tag{4}$$

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7 Since Lancaster's paper (1979) a lot of attention has been paid to the role of unobserved heterogeneity. Unobserved heterogeneity can be expected to bias the estimate of the effect of t mainly. As will be discussed in Section 3, I cannot even identify the effect of t with the available data. Therefore, I shall not consider the possibility of unobserved heterogeneity among individuals in the sequel.

8 Amemiya (1985), p. 449.

where:

U_τ = the stock of job seekers at time τ

V_τ = the stock of vacancies at time τ

M_τ = the rate of matches between job seekers and vacancies at time τ

and where τ denotes calendar time.

This concept is known as the match production function.[9] It reflects the fact that for a flow of matches to be realised, stocks of vacancies as well as job seekers are indispensable. The more one of the 'factors of production' is available, the more matches can be realised.
In the sequel the search process is assumed to be characterised by constant returns to scale: $\varepsilon + \delta = 1$.[10] Now we can ask what the transition rate of job search to employment will be in this labour market. For reason of our temporarily simplifying assumptions (homogeneous job seekers, no time dependence) the following relation holds:

$$M_\tau d\tau = U_\tau \lambda_\tau d\tau \tag{5}$$

After inserting (5) in (4) and some rearrangement one gets:

$$\lambda_\tau = C\left(\frac{V_\tau}{U_\tau}\right)^\delta \tag{6}$$

The transition rate appears to be a function of the number of vacancies divided by the number of job seekers. The simple form clearly depends on the choice of the Cobb-Douglas function, but for any production function exhibiting constant returns to scale the transition rate depends on the ratio of vacancies and unemployment.
Equation (6) is a justification for capturing the effect of cyclical variations on the transition rate by the ratio of vacancies and job seekers. This conclusion does not hinge on the assumption λ being independent of the incomplete duration of search. Using results discussed in section 3 it can be shown that under some stationarity assumptions this conclusion remains valid if this assumption is relaxed. However, if λ is allowed to depend on x_i, equation 6 is no longer valid. Not only does the number of job seekers matter, but also the composition of its stock. I leave all these complications for what they are. The discussion above is meant to give some idea of how this specification relates to a more general model of the labour market.

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9 Pissarides (1987), see Belderbos and Teulings (1990) for a further exposition on this subject.

10 Some evidence for this assumption can be found in Belderbos and Teulings (1990).

Now the purpose of the analysis is to test whether δ has the same value for all job seekers. Our hypothesis is that non-natives and lower qualified job seekers face a higher δ then others. Therefore I allow δ to be a linear function of some elements of x_i, say:

$$\delta = \underline{x}_i^{'}\alpha$$

where $\underline{x}_i$ is a vector with the elements being a subset of the elements of x_i and α is a parameter vector with corresponding dimension. To facilitate notation we drop the index for calendar time τ in the sequel. To identify δ we need variation of the ratio of the numbers of vacancies and job seekers. Therefore I introduce a subscript k denoting labour markets, each facing its own ratio of vacancies and job seekers. I assume a job seeker can only operate at one labour market in a time.
Combining (3) and (6) results into:

$$\lambda_{ik}(t) = g(t)\exp(x_i^{'}\beta)\left(\frac{V_k}{U_k}\right)^{\underline{x}_i^{'}\alpha} = g(t)\exp\left(x_i^{'}\beta + q_k\,\underline{x}_i^{'}\alpha\right) \tag{7}$$

where I write q_k in stead of ln (V_k/U_k).
It will be a central issue in this paper to establish elements of α other than the intercept being unequal to zero. Estimation results for equation (7) are presented in section 4.

The second elaboration concerns the heterogeneity of jobs within a labour market. Jobs can be categorized according to their occupation. Therefore, the job search process can be characterised not by just one but by more transition rates, one for every occupation. Assume the transition rate of individual i to occupation j after a spell of length t of unsuccessful search equals:

$$\lambda_{ijk}(t) = g(t)\exp\left(x_i^{'}\beta_j + q_k\alpha_j^0 + q_k\underline{x}_i^{'}\alpha\right) \tag{8}$$

The specification in (8) involves three assumptions.
First, g(.) is assumed to be independent of j implying that the duration dependence of the transition rates is the same for all occupations. This assumption enables an attractive decomposition of the model, which is discussed below. I have not investigated the consequences of this assumption any further. With the data at hand it seems hardly possible to derive testable implications from this assumption.
Secondly, q_o is not indexed for occupations. For V_k this seems to be counterfactual. The introduction of occupations in the model was meant to account for the heterogeneity of vacancies. Therefore it seems natural to distinguish vacancies by occupation. Data for the number of vacancies per occupation are indeed available, but inspection of these data reveals them to be not very realistic.

Earlier research has shown that using V_{kj} rather than V_k does not affect the transition rates by occupation significantly.[11] Therefore I have decided not to use these data here.
Thirdly, only α^o is allowed to vary over j. This is a crucial aspect of the model. I intend to show that, given x_i, the transition rates to higher ranked occupations are more sensitive to changing labour market conditions than the rates to lower ranked occupations: a job seeker who finds his opportunities to get a first rank job blocked by increased rationing, will revise his strategy by accepting second rank jobs too.
To capture this idea only the coefficient for product of q_k and the intercept has to vary over occupations. The coefficients for the product of other elements of $\underline{x}_i$ (apart from the intercept) and q_k are assumed to have same value for all occupations, for the data contain insufficient information to identify them all separately.
The total transition rate of job search to employment equals the sum of the transition rates by occupation:

$$\lambda_{ik}(t) = \sum_j \lambda_{ijk}(t) = g(t)\sum_j \exp\left(x_i'\beta_j + q_k\alpha_j^0 + q_k\,\underline{x}_i'\alpha\right) \tag{9}$$

In this case $\underline{x}_i$ does not contain a constant for reasons stated above.
The distribution of completed durations of search F(.) and thus the expected duration of search are solely determined by this total transition rate. Its decomposition by occupation does not affect the distribution of completed durations.
The probability that a job seeker who finds a job will find this job in occupation j equals:

$$\Pr(i \text{ finds a job in } j \text{ at } t \mid i \text{ finds a job at } t) = \frac{\lambda_{ijk}(t)}{\sum_\ell \lambda_{i\ell k}(t)} = \frac{\exp(x_i'\beta_j + q_k\alpha_j^0 + q_k\,\underline{x}_i'\alpha)}{\sum_\ell \exp(x_i'\beta_\ell + q_k\alpha_\ell^0 + q_k\,\underline{x}_i'\alpha)} \tag{10}$$

Here the advantage of g(.) being independent of j becomes clear: the probability described in (10) does not depend on the value of t. Further inspection shows this probability to be described by a multinomial logit model, which has been analysed thoroughly in the literature.[12]
Estimation results for equation (8) are presented in section 5.

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11 Teulings (1990), page 136.
12 e.g. Amemiya (1985), page 295.

3 ESTIMATION PROCEDURE

Normally, duration models are estimated using panel data. The ideal set-up involves data on a group of people all starting to search on the same date. After some period of time this group is interviewed again, which results in information about the completed duration of search of those who succeeded in finding a job. Using (1a) and (1b) a likelihood function can be easily written.

I postpone a general discussion of the data until the next section, but to understand the sequel it must be noted that no panel data were available. However, a retrospective question allows for identification of those persons who found a job during the year preceding the date of interview. [13] Furthermore it is also known if the respondent is searching for a job at the date of interview. This gives us information on two variables, namely the flow of people finding a job during the last year and the stock of people searching. In the sequel I show how to use this information to estimate the expected duration of job search or transition rates of search to employment.

The density of a job seeker who is presently searching to have started search at time - t is:[14]

$$\Pr(\text{started searching at } -t \mid \text{still searching at } 0) =$$

$$\frac{\Pr(t^* \geq t \mid \text{started searching at } -t)\Pr(\text{started searching at } -t)}{\int_0^\infty \Pr(t^* \geq s \mid \text{started searching at } -s)\Pr(\text{started searching at } -s)\,ds} \tag{11}$$

This expression can be greatly simplified if one is prepared to make two stationarity assumptions.

The first assumption is that the transition rate does not depend on the moment when the search activities start. This does not rule out the possibility of duration dependence (compare Section 2). This assumption leads $\Pr(t^* > t \mid \text{started searching at } -t)$ to be equal to $\Pr(t^* > t)$. So, the conditioning in the numerator and the denominator in (11) can be left out.

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13 This can be done by a comparison of the answer to the questions 'principal occupation?' (e.g. working, study, unemployed, disabled, working in own household) and 'principal occupation a year ago'.
One can ask why this information cannot be used to estimate transition rates by looking at which people, who were unemployed a year before, had found a job a year later. Basically there are two problems connected to this approach. First, analysis of the data shows unemployment being only partially reported in the retrospective question, probably because many unemployment spells last for only a short period. Secondly, being unemployed is not the same as searching for a job. The latter is the relevant variable.

14 Kooreman and Ridder (1983).

Secondly, I assume the probability of starting search to be constant over time. This causes the second part of the numerator and the denominator of (11) to be equal, so that they can be left out. Under these assumptions (11) simplifies to:

$$\Pr\text{ (started searching at } -t \mid \text{still searching at } 0) = \frac{\Pr(t^* \geq t)}{\int_0^\infty \Pr(t^* \geq s)\,ds} = \frac{1 - F(t)}{E(t^*)} \tag{12}$$

Do these assumptions hamper the validity of the estimation results seriously? In general the shorter the average spell of job search, the less serious are its consequences. This is because the stationarity assumptions need only to hold for the time needed to enable the majority of the job seekers to find a job. Most job search spells are rather short. The assumptions will have the most serious consequences for the estimation results for 1983: firstly, the average duration of search was very high in that year and secondly, in the period before labour market conditions were deteriorating rapidly. In general the assumptions will not affect the validity of the results seriously.[15)]

Now the rate of people finding a job M_t given the stock of job seekers U_t can be derived by multiplication of the probability that a job seeker has been searching for a period t and the corresponding transition rate:

$$M_\tau d\tau = U_\tau \int_0^\infty \lambda(s) \frac{1 - F(s)}{E(t^*)} d\tau\, ds = U_\tau d\tau \,/\, E(t^*) \tag{13}$$

After rearrangement equation (13) states that the expected completed duration of search is the number of job seekers divided by the rate of job seekers finding a job. This expression is the starting point for estimation.

To get better insight into what this procedure actually means, $E(t^*)$ has to be calculated using equation (2). First consider the case where no occupations are distinguished:

$$E(t^*) = \int_0^\infty \exp\left(- \int_0^t \gamma s^{\gamma-1} \exp(z_i' q)\, ds \right) dt = \int_0^\infty \exp(-t^\gamma \exp(z_i' q))\, dt =$$

$$= \int_0^\infty u^{1/\gamma - 1} \exp(-z_i' q)^{1/\gamma} \exp(-u)\, du = \exp\left(-\frac{1}{\gamma} z_i' q + A \right) \tag{14}$$

15 See Belderbos and Teulings (1990) for a more detailed analysis of these problems.

where:

$$A = \ln\{\tau(1/\gamma)/\gamma\}$$

$$z_i = \left[x_i^{'}; q_k \underline{x}_i^{'}\right]^{'}$$

$$\theta = \left[\beta^{'}; \alpha^{'}\right]^{'}$$

For the third step a change of variable is applied: $u = \exp(z_i'\theta)t^{\gamma}$. For convenience I suppress the subscript k in the notation of z_i.

Assume a sample including the stock of job seekers and the flow of people having found a job during the past period (i.e. the year preceding the date of observation). As a first order approximation we can use (13) for estimating this flow. The conditional probability that an individual in this sample belongs to the flow of people finding a job is:

$$\Pr(i \in M | i \in U \vee i \in M) = \frac{\exp\left(\frac{1}{\gamma} z_i^{'}\theta - A\right)}{1 + \exp\left(\frac{1}{\gamma} z_i^{'}\theta - A\right)} \qquad (15)$$

where U and M represent the sets of job seekers and those who found a job respectively. For convenience I drop the index of calendar time again.

Equation (15) is the well-known binomial logit model. So, the expected duration of search can be estimated by applying a logit model to the sample described before. The estimated coefficients measure θ/γ, except for the coefficient of the constant, which has to be corrected by adding A. By (14) these coefficients can be interpreted as minus the relative effect of the variables in x_i on the expected duration of job search. This procedure will be applied in section 4.

We now turn our attention to the case where jobs are distinguished by occupation. By a similar argument as for the case with homogeneous vacancies and by equation (10) the conditional probability of finding a job in occupation j can be shown to be:

$$\Pr\left(i \in M_j | i \in U \vee i \in M\right) = \frac{\left(\sum_{\ell} \exp\left(z_i^{'}\theta - \gamma A\right)\right)^{1/\gamma}}{1 + \left(\sum_{\ell} \exp\left(z_i^{'}\theta_{\ell} - \gamma A\right)\right)^{1/\gamma}} \frac{\exp\left(z_i^{'}\theta_j\right)}{\sum_{\ell} \exp\left(z_i^{'}\theta_{\ell}\right)} \qquad (16)$$

where M_j represents the set of those who found a job in occupation j.
Equation (16) is McFadden's nested logit model.[16] If $\gamma=1$ the model reduces to a simple multinomial logit model. Nested logit models can be estimated either by maximum likelihood or by using a two-stage method: first estimate the model described by equation (10) using maximum likelihood. This gives estimates for $\theta_j - \theta_1$ for all j except j =1. The first occupation is used as the benchmark. Next, impute these coefficients in a model describing $Pr(i \varepsilon M \mid i \varepsilon U \; \forall i \varepsilon M)$ and estimate the remaining coefficients with maximum likelihood.
As will be described in Section 5 I do not use a full set of occupations, so that procedure outlined above cannot be applied. Therefore γ is assumed to be equal to unity, so that equation (16) reduces to the multinomial logit model. Note that the coefficients θ_j can no longer be interpreted as minus the relative effects on the expected duration of job search, because this duration depends solely on the total transition rate. Therefore I have to resort to the interpretation of the coefficients as the relative effects on the transition rate by occupation.
Parameter estimates for this model are presented in section 5.

Before turning to the estimation results, one problem has to be discussed more deeply. In the models described by the equations (3) and (8) two kinds of effects are involved. The transition rates are affected firstly by personal characteristics and secondly by labour market conditions. The latter are identical for all job seekers who operate on the same market. So the job seekers in the dataset can be divided in a limited number of groups according to the labour market they belong to. Within a group labour market conditions are by definition the same.
The assumption in the present specification of the model is that q_k measures all the differences between labour markets: there are no other relevant variables, which are not included in the model. The stochastic nature of the model solely reflects the unpredictability of individual behaviour, not unobserved heterogeneity among labour markets.
A model that allows for this kind of unobserved heterogeneity would include a labour market specific error term, that is, an error term with the same value for all individuals within a market. This model is analogous to the two-error components models used in panel data analysis.[17] For every observation on individual i at time t there are two stochastic components: the first covering unobserved characteristics which vary over i but not over t, the second covering unobserved characteristics varying over i and t. Labour markets compare to individuals, and job seekers to moments of time in the two-error components model. Note however, that the assymptotic features are different here because normally the number of observation is large for the former and small for the latter, while here it is the other way around.
The omission of a labour market specific error term can be expected to induce underestimation of the variance of the parameters for the variables measuring market conditions. This can be seen by conside-

16 McFadden (1981).
17 Amemiya (1985), page 211.

ring an extreme example: suppose our dataset contains only two markets and suppose the probabilities of finding a job differ significantly in both markets, that is, a dummy for market 2 compared to market 1 proves to be significant. Now any variable with different values for both markets would produce exactly the same t-value (in absolute value), whether it be the number of vacancies per job seeker or cumulative rainfall.

The main problem is that the error terms for each individual are incorrectly perceived as independent, while within a market they are correlated due to the unobserved component of market conditions.[18)]

To solve this problem the model has to allow for this unobserved heterogeneity:

$$\lambda_{ik}(t) = g(t)\exp\left(\underline{x}_i'\beta + q_k\,\underline{x}_i'\alpha + \underline{x}_i'v_k\right) \tag{17}$$

where the index k indicates the labour market and where v_k is an error term which is distributed as $N(0,\Sigma)$. Note that model (17) reduces to equation (7) if Σ is equal to zero. This model can be estimated using a simple two-stage procedure. First, estimate the following model:

$$\lambda_{ik}(t) = g(t)\exp\left(x_i^{*\prime}\beta^{*} + \underline{x}_i'\alpha_k\right) \tag{18}$$

Equation (18) implies a separate vector α_k to be estimated for every labour market. To avoid the dummy trap, the variables included in $\underline{x}_i$ must be left out of x_i in this case. The symbol * for x_i and β is used to indicate this (so x_i^* is the complement of $\underline{x}_i$ with respect to x_i). By (15) this model is equivalent to a binomial logit model.

Now we can write down the following:

$$\hat{\alpha}_k = q_k\alpha + \underline{\beta} + v_k + w_k \tag{19}$$

where the symbol ^ above a parameter denotes its estimator, w_k represents the error involved in the estimation of α_k, and where $\underline{\beta}$ represents the elements of β that correspond to the variables in $\underline{x}_i$ ($\underline{\beta}$ being the complement of β^*).

Due to the consistency of the maximum likelihood estimates of a logit model the expectation of w_k is asymptotically equal to zero. An estimate of the covariance matrix of w_k results from the estimation of the logit model (18). Model (19) is a simple system of linear equations with the vector α and $\underline{\beta}$ as unknown parameters. Consistent estimates of α and β are obtained by applying FGLS to (19): the first equation (19) is estimated with OLS to obtain an estimate of Σ. This estimate can be used to calculate

18 This problem has more general importance. Any model using explanatory variables which have the same value for predefined subsets of the full set of observations suffer this problem e.g. Lancaster (1979).

the covariance matrix of $v_k + w_k$. Next we use this covariance matrix to estimate (19) with GLS. This procedure is described in more detail in the Appendix.
The method allows for the test of the restriction $\Sigma = 0$. If this restriction cannot be rejected, all relevant variation between labour markets is covered by q_k. In that case one can safely return to the more simple model (7). A practical approach is to note that model (7) is nested in model (18). The test $\Sigma = 0$ is equivalent to the test of model (7) against model (18); the former can be maintained if all v_k are equal to zero. For this purpose the likelihood ratio test can be used.
There are two limits to the applicability of this two-stage method. The first is the number of labour markets to be distinguished. Too many markets would increase the number of regressors in model (18) beyond what is manageable. The second is the number of observations per labour market. If there are too few observations the estimate of the covariance matrix of w_k is unreliable.

Mutatis mutandis this procedure can also be applied to the model with occupation specific transition rates. However, here the problem of the large number of regressors is acute, because it rises linearly with the number of occupations. To solve this problem I use the convenient feature of multinomial logit models which is that they can be decomposed into submodels, which are themselves logit models:[19)]

$$\Pr\left(i \in M_j \mid i \in U \vee i \in M_j\right) = \frac{\Pr\left(i \in M_j \mid i \in U \vee i \in M_j\right)}{\Pr\left(i \in U \vee i \in M_j \mid i \in U \vee i \in M\right)} = \frac{\exp\left(z_i'\theta_j\right)}{1 + \exp\left(z_i'\theta_j\right)} \tag{20}$$

However, the number of observations per labour market relating to finding a job in a specific occupation is in many cases rather small. As noted before this leads to problems in the estimation of Σ. Therefore, I apply this method only to a_j^0 and not to α.

A large number of observations is available. To reduce computation costs I do not use all of them for estimation. Keeping the total number of observations fixed, the parameter estimates of logit models are most reliable if the observations are divided more or less equally over the alternatives. During the period 1981-1985 job seekers outnumbered by far those who found a job. So to reduce the number of observations with the minimum loss of reliability of the parameter estimates the best one can do is to include only part of the job seekers. Here another practical feature of logit models can be used:[20)]

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19 Hausman and McFadden (1984).
20 Manski and Lerman (1977).

$$\Pr\left(i \in M \mid i \in U^* \vee i \in M\right) = \frac{\Pr(i \in M \mid i \in U \vee i \in M)\Pr\left(i \in U^* \vee i \in M \mid i \in U \vee i \in M\right)}{\Pr\left(i \in U^* \vee i \in M \mid i \in U \vee i \in M\right)} = $$

$$= \frac{\exp\left(z_i' q_j + \ln(p)\right)}{1 + \exp\left(z_i' q_j + \ln(p)\right)}$$

(21)

where U^* is the set of job seekers used for estimation and p is the probability-that a job seeker is included in U^*.

Equation (21) implies that if an alternative is under-represented in the sample all parameter estimates derived from maximum likelihood estimation remain consistent, except for the intercept. The intercept can be corrected easily by substracting ln(p). In our case p for job seekers is 50 %.

4 RESULTS: GENERAL MODEL

The dataset I use is the ArbeidsKrachten Telling AKT (Labour Market Survey) of the Dutch Central Bureau of Statistics. This survey is based on interviews of a 1-5 % sample of the total population of the country. The AKT was organised every two years up till 1985. Here the results for 1981, 1983 and 1985 are used. The editions before 1981 could not be used, because data on the number of vacancies were available only from 1980. These data are derived from a survey of enterprises also organised by the Central Bureau of Statistics. The Labour Exchange also produces data on the number of vacancies and it started to do so long before 1980, but these figures are known to be unreliable. Earlier research revealed that they do not fit the model because the registration rate of vacancies falls in the course of time.[21]

To analyse the effects of labour market conditions the data must cover as many markets as possible, in any case more than three (1981, 1983, and 1985). To increase variation markets are classified by region. Regional market conditions vary a lot in the Netherlands, the west experiencing a more tight labour market than the south and the north. For each year the market is divided into five regional markets, so together 15 markets are distinguished. To give some idea of the variation, the ratios of numbers of job seekers and vacancies are listed in Table 1. Earlier research has shown that these markets are almost perfectly separated, except for the higher vocational and the university level of education.[22] Therefore, these are left out of the analysis.

21 Belderbos and Teulings (1990).

22 Teulings and Koopmanschap (1989).

The analysis includes job-seekers having a job, but only if they are forced to look for another job because they expect to be dismissed in the near future. I call this involuntary job search, in contrast to voluntary search. Involuntary search can be distinguished from voluntary search by the questions 'reason for looking for a job?' (for job-seekers) or 'reason for leaving your former job?' (for those who found a job during the last year). The motivation to keep out voluntary search from the analysis is that it is not sure that the model can be applied to this. Many persons will have found their new job without any search at all, simply because they were asked to apply by their new employer. In that case the notion of a transition-rate from search to a new job does not make sense.

Table 1 The numbers of vacancies per job seeker by labour market

Region*	1981	1983	1985
North	0.061	0.019	0.029
East	0.079	0.019	0.050
North West	0.172	0.048	0.098
South West	0.203	0.033	0.086
South	0.082	0.023	0.047

* N: Groningen, Friesland, Drenthe
E: Overijssel, Gelderland
NW: Utrecht, Flevoland, North-Holland
SW: South-Holland, Zeeland
S: Brabant, Limburg

Furthermore, I leave out all persons finding or looking for a job of less then 15 hours or a job as self-employed. The labour market for these jobs operates in a substantially different way. Trying to start up your own firm can hardly be compared to trying to get hired by an employer. For small jobs turnover rates are high, which leads to substantial measurement errors.
Finally, the analysis is restricted to persons aged below 45 years. The reason for this is that schooling variables have to play an important role in the explanation of transition rates. The spread of secondary education among broader classes of society has increased tremendously since World War II. Moreover, nowadays schooling is of much greater importance for achieving an attractive position. Both factors make it hard to compare the degree of education of a 50 year old person with that of a 30 years younger entrant. Anyway, the chances of finding a job for people above the age of 45 were negligible during the period under consideration.

Table 2 contains the parameter estimates for the expected duration of job search before finding a job. As follows from equation (14) the parameters indicate the relative increase in the expected duration of job search due to a unit increase in the explanatory variable.

Table 2 Estimation results for the expected duration of job search

unobserved heterogeneity between markets:	without		with	
variable	coefficient	(st. err)	coefficient	(st. err)
intercept[1]	1.37	0.064	1.44	0.087[4]
level of education[2]	-0.37	0.031	-0.31	0.039[4]
direction of education (compared to general)				
- educational	0.01	0.18	-0.14	0.18
- agraricultural	-0.55	0.13	-0.49	0.14
- technical	-0.39	0.055	-0.32	0.060
- transport	-0.40	0.23	-0.40	0.24
- medical	-0.64	0.12	-0.80	0.12
- economic	-0.55	0.072	-0.57	0.073
- social/cultural	-0.34	0.24	-0.50	0.24
- care	-0.25	0.065	-0.21	0.067
age (except for second entry)[5]	0.33	0.040	0.30	0.040
age (second entry)[5]	-0.049	0.086	-0.034	0.087
labour market position (compared to dismissed)				
- entrant	-0.62	0.050	-0.65	0.050
- second entry	1.09	0.18	1.06	0.18
sex by composition of the household (compared to man, single)				
- man, non-single	-0.48	0.064	-0.51	0.064
- woman, non-single	-0.12	0.080	-0.12	0.081
- woman, single	0.15	0.055	0.16	0.056
non-native	0.87	0.13	0.91	0.22[4]
ln (vacancies/job seekers)[3]	-0.41	0.049	-0.39	0.096[4]
idem by level of education[2][3]	0.016	0.035	-0.009	0.095[4]
idem by non-native[3]	-0.51	0.15	-0.45	0.29[4]
log likelihood	-8154.60		-8117.23	
number of variables in logit model	21		60[6]	
number of observations	12807		12807	

1) corrected for endogenous sampling
2) value: 0 = basic; 1 = extended basic; 2 = intermediate
3) measured in deviations from their mean over labour markets (-2.915)
4) based on FGLS
5) (age minus 15)/10; age in years
6) in the logit model

As could be expected, the higher the education level attained, the quicker a job seeker finds a job. However, the coefficients show that a proper choice of the direction of education is even more important than reaching the highest possible level. In general vocational training is a better option than leaving school with only general education, except for training to become a teacher. This exception can be ascribed to demographic reasons. The high coefficient for medical education reflects the big shortage of nurses in the Netherlands. This problem is mainly due to the low pay for this occupation and it is still largely unsolved.

I distinguish between three types of labour market experience: those entering the labour market, job seekers who are dismissed or fear to be so in the near feature, and job seekers who re-enter the market after an interruption of a longer period. The latter are mainly women who have been taking care of their children and are now trying to get a job again. I have separated the effect of age for this group, because it can be expected that being older is not that much of a disadvantage for this group. Employers accept taking care of the children as a reasonable explanation for not being employed.
The estimation results confirm this idea: being aged is a serious disadvantage for entrants and dismissed persons, while for re-entrants age does not affect the expected duration of search that much. Note that the dummies for being an entrant or a re-entrant measure the effects of the labour market position at the age of fifteen. Because the prospects of laid-off persons to get a job deteriorate with age, the difference between this group and re-entrants gradually disappears. Taking into account the coefficients for age, it is better for a job seeker to be a re-entrant rather than to be laid off when he or she is older than 44.
Sex as such does not affect the expected duration of search that much. However, being a single male gives a substantially higher expected duration of search. Similar results can be found for income: single males earn less than their married colleagues.[23] Meyer and Wise attribute this result to the fact that in most cases married males have to earn the whole of the household income, while singles only have to take care of themselves. Probably a more likely explanation of this phenomenon is self-selection: success on the partner market and on the labour market are correlated. Our data give some weak evidence for this supposition: singles get lower ranked jobs than those having a partner. Moreover, their duration of search seems to be more seriously affected by the cycle.
We now turn our attention to the main point of interest: the variables reflecting labour market conditions. For transparancy I have measured the effect ln(vacancies per job seeker) in deviations from its mean over labour markets. This allows us to interpret the coefficients for non-nativeness and for the level of education as the effects in the 'average' labour market; this procedure does not affect the coefficients for the effect of labour market conditions.
First, the number of vacancies per job seeker proves to be highly significant. The value of the coefficient is somewhere between zero and unity as would be expected theoretically. For a 1 percent

23 Meyer and Wise (1983), Dickens and Lang (1985).

increase in the number of vacancies per job seeker, the expected duration of search falls by approx 0.4 percent.

As set out in the Introduction, the effect of market conditions on the duration of search is expected to be higher for non-natives and lower educated job seekers. The estimation results reveal that the effect of market conditions do not vary significantly with education level. However, market conditions do affect the relative position of non-natives. On average the prospects of non-natives finding a job are less than of those of natives. During the downswing this difference increases substantially: in our sample the coefficient for non-natives varies over labour markets from -0.20 (region SW, 1981) to -1.40 (region N and E, 1983).

These results are a partial confirmation of our first hypothesis that job seekers with lower qualifications suffer the downswing more seriously.

For comparison the results for the estimation method, without accounting for unobserved heterogeneity between labour markets, are listed in Table 2 next to the results when unobserved heterogeneity of market conditions is allowed for.

The likelihood ratio test shows that the restriction of no unobserved heterogeneity cannot be maintained.[24] Therefore, allowance for labour market specific error terms is necessary. Comparing the parameter estimates for both methods shows them to be more or less the same. The same is true for their estimated standard errors, except for the parameters α, which correspond to labour market specific variables, and their counterparts in β (i.e. $\underline{\beta}$). This had been expected: introducing labour market specific error terms accounts for the fact that the individuals within a labour market are not stochasticly independent. A comparison of the standard errors shows that this effect must not be neglected: the coefficient of the additional effect of labour market conditions for non-natives is shown to be only weakly significant in this case. The estimate for the covariance matrix of the labour market specific error terms (Σ) is presented in Table 2a.

Table 2a Estimate of covariance matrix of the labour market specific error terms (Σ)

	1	2	3
1. log (vacancies/job seekers)	0.040		
2. idem by education level	-0.028	0.032	
3. idem by non-native	-0.031	-0.041	0.398

24 The value of the test statistic is: 2 (8154.60 - 8117.22) = 74.76 This statistic is distributed χ^2 with 60 - 21 = 39 degrees of freedom. The critical value at 5 % significance level is 54.3.

5 RESULTS: MODEL WITH OCCUPATIONS

In this section I discuss the estimation results for a model with occupation specific transition rates. It is hardly possible to estimate such a model for the whole labour market simultaneously. Even with the restriction of persons with education above the intermediate level being left out of the analysis, a meaningful classification of jobs requires that at least 30 occupations are distinguished. The simultaneous estimation of 30 transition rates, where each rate involves something like 20 coefficients, is a hard job.

Therefore, only a part of the labour market is analysed that can be separated from the rest of the market rather easily. I chose the market for technical occupations. This choice is partly motivated by the degree of separation of the market and partly by the fine categorisation of jobs into occupations in manufacturing compared to services. This fine categorisation can be ascribed to Taylorist practices in manufacturing. In the model seven occupations are distinguished, mainly reflecting differences in the degree of schooling or experience needed to perform the tasks involved.[25)]

Restricting the analysis to technical occupations implies that the range of job seekers can be narrowed down too. Only male job seekers who have had training in general or technical direction are included. The motivation to include job seekers with general education is that at the basic level only general education exists. Women are excluded because they seldom find jobs in these occupations. This reduces the number of variables considerably, because (nearly) all re-entrants were female too.

The figures in Table 3 give an indication of the degree of separation between the selected occupations and job seekers and the rest of the labour market: 43% of the job seekers included in the model find a job in an occupation which is not include, while only 6% of the people not included find a job in an occupation which is included. The figures show the separation to be far from perfect.

Table 3 Job seekers distinguished by whether they are included in the model or not and by whether the occupation where they find a job is included or not (row percentages)

	the occupation where they find a job:		
people finding a job:	included	not included	total
included in the model	57	43	49
not included in the model	6	94	51

[25] This is shown by cluster analysis of the Labour Market Survey of 1981, see Teulings and Vriend (1988). The classification of occupations stems from this analysis.

The results are shown in Table 4. The ranking of occupations in the headings of the table is done more or less in accordance with the level of required qualifications. I used the average level of schooling of those holding a job in the occupation as a proxy for this.
In the table the coefficients of the transition rate are presented instead of those of the expected duration of search, as is done in Table 2. The former are the opposite of the latter. The reason for this is that in this case the interpretation in terms of expected duration of search is not valid, as this duration depends solely on the total transition rate and not on occupation specific rates.
In this analysis the level of education is considered to be an ordinal instead of a cardinal variable. This is done because in this model we have to allow for the possibility that in the transition rate to for example the occupation 'skilled production worker' (the middle in the ranking) the extended basic level of education has a positive effect compared to the basic as well as to the intermediate level of education.
The results stress the strong handicap of being non-native. In all occupations but one being a non-native reduces the relevant transition rate. Only for the one but lowest ranked occupation is the opposite true. So, according to the results in the former section, non-natives have a lower chance of finding a job in general and according to the results in this section, if they do find a job this is in a lower ranked occupation.
Entrants have higher transition rates to all occupations as have non-singles. Only for unskilled production workers does being single not affect the transition rate that much. This result supports to some extent the supposition that singles are less attractive to employers than their married colleagues.

Let us now turn to the variables describing labour market conditions. As in Section 4 labour market conditions are measured in deviation from their mean over labour markets.
First notice that there is a connection between the ranking of the occupation and the effect of the number of vacancies per job seeker: the effect increases with the required level of qualifications. This finding is in accordance with the second basic hypothesis of this paper: if labour market conditions deteriorate, job seekers will find less attractive jobs than in a tight labour market. The adverse ranking of the coefficients of truck drivers and unskilled production workers is probably the result of the inadequacy of the average education level as a measure of the ranking of occupations. The high coefficient for being non-native for unskilled production workers strongly suggests that working in this occupation is less attractive than being a truck driver.

Table 4 Estimation results for occupation-specific transition rates from job seeking to employment for the technical sub-market without unobserved heterogeneity (std. errors in parentheses)

	truck drivers	un-skill.prod. workers	constr. workers	skilled prod. workers	elec-tricians	superv. prod. workers	intermed. technical staff
average education level of those employed[4)]	0.77	0.87	0.90	1.30	1.57	1.67	2.07
intercept[1)]	-3.28	-3.53	-3.44	-3.36	-7.15	-8.66	-7.75
	(0.18)	(0.23)	(0.22)	(0.16)	(0.58)	(0.82)	(0.78)
level of education							
- extended basic	-0.67	-0.28	-0.69	-0.80	0.46	-1.41	-1.18
	(0.23)	(0.25)	(0.25)	(0.21)	(0.65)	(1.04)	(0.88)
- intermediate	-1.30	-0.32	-0.96	-0.49	1.14	1.13	2.24
	(0.25)	(0.26)	(0.27)	(0.21)	(0.64)	(0.93)	(0.80)
technical education	1.42	0.88	0.37	2.46	3.01	1.94	2.93
	(0.21)	(0.22)	(0.42)	(0.18)	(0.40)	(0.80)	(0.39)
age[3)]	-0.17	-0.57	-0.10	-0.31	-0.62	0.73	-0.26
	(0.30)	(0.17)	(0.13)	(0.087)	(0.23)	(0.26)	(0.23)
non-native	-0.74	0.53	-1.48	-1.30	-1.19	-0.89	$-\infty$
	(0.30)	(0.27)	(0.45)	(0.33)	(0.75)	(1.06)	
entrant	0.38	0.69	0.41	0.83	1.55	0.52	1.44
	(0.16)	(0.19)	(0.21)	(0.12)	(0.26)	(0.69)	(0.30)
non-single	0.78	0.11	0.72	0.53	0.78	1.33	0.32
	(0.15)	(0.22)	(0.20)	(0.12)	(0.27)	(0.53)	(0.29)
log (vacancies/ job seekers)[2)]	0.35	0.11	0.30	0.53	0.97	0.82	0.90
	(0.096)	(0.12)	(0.11)	(0.094)	(0.15)	(0.24)	(0.18)
generic coefficients							
log (vacancies/job seekers) by:							
- education level [2) 4)]	-0.11	(0.06)					
- idem by non-native [2)]	0.57	(0.23)					
log likelihood	-5383.08						
number of parameters	65						
number of observations	4828						
fraction of the alternative in the sample 22	0.040	0.041	0.076	0.154	0.008	0.029	0.0-

1) corrected for endogenous sampling
2) measured in deviations from their mean over labour markets (-2.915)
3) (age minus 15)/10; age in years
4) value: 0 = basic; 1 = extended basic; 2 = intermediate

Two variables measure the differential effect of labour market conditions on the transition rates of various groups. This effect is assumed to be equal for all occupations.[26] The sign of both coefficients is as would be expected. An increase in the number of vacancies per job seeker raises the transition rate of non-natives more than that of natives and reduces the transition rate of job seekers with an intermediate level of education compared to those with less education. This also confirms my second hypothesis. However, the latter effect is only weakly significant. To illustrate the implications of these results I have listed the estimated effects of labour market conditions on the transition rates for some groups of job seekers in Table 5.

Table 5 The coefficient of ln (vacancies/job seekers) per occupation for selected groups of job seekers[1)]

	basic education native	basic education non-native	intermediate education native
truck drivers	0.35	0.92	0.11
unskilled production workers	0.11	0.68	-0.13
construction workers	0.30	0.87	0.06
skilled production workers	0.53	1.10	0.29
electricians	0.97	1.54	0.73
supervising production workers	0.82	1.39	0.58
intermediate technical staff	0.90	1.47	0.66

1) based on the coefficients estimated without unobserved heterogeneity

For job seekers with the strongest position, and natives with an intermediate education level, the effects of labour market conditions are far less. For the transition rate to jobs such as unskilled production worker the effect is even negative. Although this effect is insignificant, it can be rationalised: if the prospects of getting a higher ranked job deteriorate sufficiently then job seekers react by accepting offers they would have rejected before. This effect on the transition rate can offset the effect of a reduced number of job offers.

A clear pattern is revealed. For job seekers with the weakest labour market position, non-natives with only a basic education level, a deterioration of labour market conditions greatly affects the transition rates, even to low ranked jobs such as unskilled production worker. For those who adhere to the dual labour market hypothesis, unskilled production workers probably belong to the secondary segment. Dickens and Lang stress that the important difference between the segments is that jobs are rationed in the primary segments only.[27] These results suggest that for non-natives even jobs in the secondary segment can be rationed.

26 A likelihood ratio test reveals that the implied restrictions on the coefficients cannot be rejected.

27 Dickens and Lang (1985).

For reasons stated before unobserved heterogeneity among labour markets is only allowed for labour market conditions in general and not for its cross-product with other variables. As is the case for the general model of Section 4, the coefficients and the standard errors of all coefficients, apart from those for the number of vacancies per job seeker and the intercept, are hardly affected by the introduction of market specific error terms. Therefore, only the latter coefficients with their standard errors are listed in Table 6. Again the estimated coefficients remain more or less unaffected, while their standard errors increase substantially. This increase is less than for estimation results reported in Section 4. Electricians and intermediate technical staff seem to be exceptions to this rule. However, these exceptions are due to the limited numbers finding a job in these occupations.

Table 6 Comparison of the estimation results with and without unobserved heterogeneity between markets

	with heterogeneity			without heterogeneity	
	intercept[1)]	ln(U/V)[2)]	σ	intercept[1)]	ln(U/V)[2)]
unskilled production workers[3)]	-3.51 (0.26)	-0.024 (0.23)	0.36	-3.53 (0.23)	0.11 (0.12)
construction workers	-3.42 (0.23)	0.36 (0.16)	0.16	-3.44 (0.22)	0.30 (0.11)
truck drivers	-3.34 (0.18)	0.43 (0.14)	0.095	-3.28 (0.18)	0.35 (0.096)
skilled production workers	-3.36 (0.17)	0.39 (0.13)	0.14	-3.36 (0.16)	0.53 (0.094)
supervising production workers	too few observations			-8.66 (0.82)	0.82 (0.24)
electricians	-7.39 (0.66)	1.11 (0.41)	0.32	-7.15 (0.58)	0.97 (0.15)
intermediate technical staff	-8.36 (1.14)	1.62 (0.89)	0.52	-7.75 (0.78)	0.97 (0.18)

1) corrected for endogenous sampling
2) compared to the mean value over labour markets (-2.915)
3) one labour market was kept out of the analysis due to nobody finding a job in that occupation in the sample.

6 CONCLUSION

The empirical evidence presented in Section 4 and 5 supports the idea that the transition rates to employment for various groups of job seekers are disproportionately affected by fluctuations in labour market conditions. Excess supply affects the transition rates of low qualified job seekers more than average. These results hold for the aggregate transition rate as well as for occupation specific transition rates, in the sense that the rates for higher ranked occupations are more seriously affected than those of lower ranked occupations. This implies that job seekers get better jobs during the upswing than during the downswing.

These findings fit into efficiency wage models where wages are sticky in the case of excess supply so that the labour market has to react to fluctuations in aggregate demand by adjusting quantities and not wages. These findings do not imply that there can be no relative wage adjustments at all. But as long as employers set wages within an occupation irrespective of the type of job seeker applying for the vacancy, excess supply will make employers more choosy. The only force which would stop them from doing so would be a rise in the wage differentials within an occupation, so that they have to pay for increasing their qualification standards.[28)]

The empirical results show that non-natives suffer the most from the increased selectivity during the downswing. This is in accordance with the efficiency wage model. Assume that employers for some reason or another dislike employing non-natives. In a competitive market there is a price for expressing these preferences: the wages of non-natives will tend to be below the wages of natives. Those employers whose discriminatory preferences are the weakest will employ non-natives, resulting in a lower wage bill.

If there is excess supply of labour, and wages are kept from falling by the existence of an efficiency wage, this mechanism breaks down. Jobs are rationed and employers can select their favourite candidates without having to pay surplus wages. In this case discrimination is costless.

This reasoning implies that the high unemployment among non-natives does not have to be caused by low productivity. On the contrary, the fact that the position of non-natives deteriorates relative to that of natives during the downswing of the cycle suggests that strong preferences against non-natives are only expressed when there are no costs involved.

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28 Compare the models of Mortensen (1970) and Lippman and McCall (1976), page 185.

Appendix

For convenience I rewrite equation (19) in matrix notation:

$$Y = X\Pi + V + W \tag{1}$$

where (dimension between brackets):

$Y = [\hat{\alpha}_1 \ldots\ldots \hat{\alpha}_K]'$ (*KxL*)

$x = [q_1, \ldots q_K]'$ (*Kx*1)

$X = [\iota_K, x]$ (*Kx*2)

$\Pi = [\underline{\beta}, \alpha]'$ (2*xK*)

$V = [v_1 \ldots v_K]'$ (*KxL*)

$E(V) = 0$

$\frac{1}{K} E(V'V) = \Sigma$ (*LxL*)

$W = [w_1 \ldots w_K]'$ (*KxL*)

$E(W) = 0$

$E[\text{vec}(W)\text{vec}(W)'] = Q$ (*KLxKL*)

with:

$Q = \text{var}(\text{vec}(\Pi) | X, V)$

K = the number of labour markets

L = the dimension of the vectors α_k, $\underline{\beta}$, and α

α_k, $\underline{\beta}$, α, v_k and w_k are defined before in the text.

The matrix Q is the covariance matrix of the maximum likelihood estimates of the elements of Y, which results from the estimation of model (18) in the text. So we have an estimate of Q.

In addition we make the following assumptions:

$$E(V'X) = 0 \tag{1a}$$

$$E\left(W'X\right) = 0 \tag{1b}$$

$$E\left(V'W\right) = 0 \tag{1c}$$

The assumptions (1a) and (1b) are standard in regression analysis. Assumption (1b) follows from the model specification: W is asymptotically a linear function of the stochastic component in individual behaviour conditional on the labour market specific error terms. The latter are covered by V.
Now let S be the L x L matrix defined by:

$$S = Y'\left[I - X\left(X'X\right)^{-1}X'\right]Y \overset{def}{=} Y'MY \tag{2}$$

The main diagonal of S contains the sum of squared residuals of the regression for each element of $\underline{y}_i$, the off diagonal elements reflect the correlation of the residuals for two different elements of $\underline{y}_i$. Substituting for (1) and taking expectations gives:

$$E(S) = (K - 2)\Sigma + J'(Q * \underline{M})J \tag{3}$$

where:

$\underline{M} = \left(\iota_K' \iota_K\right) \otimes M$

$J = I_L \otimes \iota_K$

and where A*B denotes the piecewise or Hadamard product of two matrices.
Alternatively, the i,j-th element of $J'(Q*\underline{M})J$ can be written as $tr(Q_{ij}M)$,

where:

Q_{ij} represents the i,j-th K x K block of Q.
From (3) a consistent estimate of Σ can be derived:

$$\hat{\Sigma} = \frac{1}{K - 2}\left[S - J'\left(\hat{Q} * \underline{M}\right)J\right] \tag{4}$$

where a ^ above a symbol denotes its estimator. Now we have estimates of Σ and Q so we can derive the variance matrix of V + W, which will be denoted A. This variance matrix can be used to estimate (1) by GLS.

$$vec\left(\hat{\Pi}\right) = \left(\underline{X}'A^{-1}\underline{X}\right)^{-1}\underline{X}'A^{-1}\, vec(Y) \tag{5}$$

$$var\left[vec\left(\hat{\Pi}\right)\right] = \left(\underline{X}'A^{-1}\underline{X}\right)^{-1} \tag{6}$$

where:

$$\underline{X} = I_L \otimes X$$

$$A = \hat{\Sigma} \otimes I_L + \hat{Q}$$

The estimator of Σ, see equation (4), is equal to the difference of two matrices, which are both positive definite. However, there is no guarantee that the difference between both matrices is positive definite too, although this is a requirement for Σ. Asymptotically there will be no problems, unless the model is mis-specified.

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