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TIME-DEPENDENT EFFECTS OF UNEMPLOYMENT BENEFITS

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ABSTRACT: This paper presents methods to estimate and test the effects of time-dependent covariates in parametric duration models. Using Finnish microeconomic data it is shown that unemployment insurance benefits have a negative effect on the probability of becoming employed during the first months, but after that period the effect vanishes. The reason is that in the Finnish system persons who are eligible for the benefits have a risk to lose them after the first three months. This result remains after correcting for omitted variables assuming gamma and mass point heterogeneity across unemployed perons.

KEY WORDS: Unemployment benefits, unemployment duration, time-dependent effects, heterogeneity.

1. Introduction

In this paper the effects of unemployment insurance (UI) on spells of unemployment are examined. The circumstances of unemployed persons do not usually stay constant over time. The purpose of this paper is to estimate the time-dependent effects of time-dependent benefits on the re-employment probability during the unemployment spell. A technique for estimating these effects is presented using Finnish microeconomic data.

Duration models based on the proportional hazards (PH) assumption imply constant effects of explanatory variables over time. A score test for testing the PH assumption is presented and a method for estimating the time-dependent coefficients in a Weibull model is developed. Often it may be preferable to avoid estimating an alternative non-proportional hazards model and therefore the focus is on a score test. A computationally convenient form of the test statistic and the appropriate connection with the pseudo-regression based on ordinary least squares is presented.

It is inevitable that econometric models do not include all the necessary explanatory variables, either because they are unmeasurable or because their importance is unsuspected. The neglected heterogeneity may bias the parameter estimates towards zero. To correct for unobservable variables gamma and mass point heterogeneity across individuals is introduced into the model.

The paper is organized as follows. Section 2 introduces the parametric duration models and time-dependent effects.

Furthermore, it provides a score test for the time-dependent effects. Section 3 analyzes the effects of omitted variables and introduces gamma and mass point heterogeneity into the model. The results of the estimations are presented in section 4 and section 5 concludes the study.

2. Time-dependent effects of UI benefits

Parametric duration models

Before presenting the model with time-dependent covariates a generic form of the likelihood function with the right censored duration data is presented. Let us consider independent pairs of independent random variables T and Z, where T is the duration variable of primary interest and Z is a censoring variable. A censoring time or a duration time are observed, $t = \min(T, Z)$, with the censoring indicator, $\bar{c} = 1$ if $T \geq Z$ and $\bar{c} = 0$ otherwise. An indicator of a completed spell of unemployment is defined as $c = 1 - \bar{c}$.

Econometric duration models are specified in terms of the hazard function $h(t)$, which is the conditional probability that the person leaves unemployment at t given that he still is unemployed. The probability of being still unemployed until the duration t is given by the survivor function

$$S(t) = e^{-I(t)}, \tag{1}$$

where $I(t)$ is the integrated hazard

$$I(t) = \int_0^t h(\tau)d\tau. \tag{2}$$

Using the rule of conditional probabilities, the unconditional probability (density function) that an individual leaves unemployment at t is a product of the hazard and survivor functions

$$f(t) = h(t)e^{-I(t)}. \tag{3}$$

The likelihood function for individuals can then be written using the indicator for complete spells of unemployment c as follows

$$\ell = h(t)^c e^{-I(t)}, \tag{4}$$

which is equal to f(t) if c = 1 and S(t) if c = 0. The distribution of unemployment spells needs to be parametrised, and maximizing the likelihood function ℓ over the unknown parameters θ may be accomplished by maximizing a concave functional $L(\theta) = \Sigma \log \ell(\theta)$.

A commonly applied specification is the PH model, where the hazard function $h(t) = h_0(t)h(x)$ factors into the product of a function of duration time, the base-line hazard, and function of the explanatory variables. The Weibull model is a versatile family of duration distributions in view of its interpretation and its flexibility for empirical fit. Therefore it has been widely used in applications of duration models to unemployment spells. In a Weibull model the base-line hazard can be parametrised as $h_0(t) = \alpha t^{\alpha-1}$, where α is the shape parameter. If $\alpha > 1$, the hazard function is increasing over the duration of the unemployment spell and it is said that there is *positive duration dependence*. If $\alpha = 1$, the hazard function is constant and the distribution of unemployment spells is exponential. If $\alpha < 1$, the hazard function is decreasing over time and it is said that there is *negative duration dependence*. The explanatory variables are introduced into the model in a log-linear form $h(x)=e^{x\beta}$. An advantage of this form is that it renders positive estimates. The integrated hazard is written as $I(t) = \int_0^t h(\tau)d\tau + C$. The constant C is chosen such that $I(0) = 0$. Then the integrated hazard in the Weibull case can be written simply as

$$I(t) = t^{\alpha}e^{x\beta}. \tag{5}$$

Consequently, the survivor, density and hazard functions of the Weibull distribution can be written as

$$S(t) = e^{-t^{\alpha}e^{x\beta}} \tag{6}$$

$$f(t) = \alpha t^{\alpha-1}e^{x\beta} - t^{\alpha}e^{x\beta} \tag{7}$$

$$h(t) = \alpha t^{\alpha-1}e^{x\beta}. \tag{8}$$

To estimate the unknown parameters $\theta = (\alpha, \beta)$ the hazard function (8) and integrated hazard (5) are substituted into the likelihood function (4), which is maximized with respect to the parameters.

Time-dependent effects

According to the Finnish UI system the circumstances of an unemployed person are different during the unemployment of the first three months. If there are not suitable jobs in the unemployed person's area of residence within the first three unemployment months, the person does not have to accept an offer outside his area of residence. Also during the first months the unemployed person does not have to accept an offer if the job is not suitable to him with respect to his education or previous work experience. This rule concerns persons with education and at least one year of job experience or alternatively persons without education and at least two years experience in their job. A person who after being unemployed for three months does not accept an offer may lose his benefits. If the effect of the time-dependent change is handled in a flexible manner, it should account for the higher hazard just after the first three months. This is allowed for letting the unemployment benefits and their parameters vary over time.

The time trended variables may be replaced with their within spell average or using beginning of spell values [Heckman, Singer (1984a)]. Usually in parametric models the variation in the explanatory variables across observations is used to take into account the time-dependent effects. The problem with these kinds of models is that the variation over time of the covariates may be absorbed by the base-line specification. Prentice and Gloeckler (1978) specify a semiparametric model and use the variation in the mean of the covariates, i.e. the variation in the covariates across observations, to estimate the base-line hazard and structural parameters. No assumptions are made about the base-line hazard. In that sense the Prentice and Gloeckler approach is similar to Cox's partial likelihood technique (Cox, 1972, 1975). Their method has been proposed also by Han and Hausman (1986) and used by Moffitt (1985). Recently Meyer (1990) divided the duration of the unemployment into intervals of one week and extended the Prentice and Gloeckler model by using time-dependent covariates.

The approach taken in this paper has several advantages. The method is more efficient than the Prentice and Gloeckler approach in the sense that the duration is continuous. It is not partitioned into intervals. The parameter estimates may be sensitive with respect to how the duration is classified into days or weeks. The approach avoids inconsistent estimation of covariate coefficients due to allowing for the time-dependent covariates and their parameters to vary over time. Furthermore, unobserved heterogeneity across observations is taken into account. Meyer found the computation of the discrete mixing distribution difficult. In this context the computation is rather easy.

The PH model assumes that the effect of an explanatory variable is constant during the duration of the unemployment. An alternative is to assume that the effect varies with the duration, remaining constant within predefined intervals. Such an alternative may be relevant in long-term studies, and in cases where the environment of an individual changes starting at a known point in time it may even be the more natural model to apply.

Consider q intervals of duration time $(t_0, t_1], \ldots, (t_{q-1}, t_q)$ with $t_0 = 0$ and $t_q = \infty$. The hazard function of the Weibull model with time-dependent effects is

$$h(t) = \alpha t^{\alpha-1} e^{x(\beta+\mu_j)}, \quad \text{for } t_{j-1} < t \le t_j,\ j=1,\ldots,q, \tag{9}$$

where $\beta = (\beta_1 \ \ldots \ \beta_p)$ and $\mu_j = (\mu_{1j} \ \ldots \ \mu_{pj})$ are 1+q vectors of p parameters. By definition $\mu_j = 0$, as $j = 1$. One reason for this kind of specification of time-dependent effects is that the integrated hazard has a closed-form expression. The integrated hazard is obtained by integrating the hazard functions by intervals, which leads to the expression

$$I(t) = \sum_{s=1}^{j-1} [I_s(t_s)-I_s(t_{s-1})] + [I_j(t)-I_j(t_{j-1})]. \tag{10}$$

In the Weibull case, e.g. in the third interval, it would be $I(t) = t_1^{\alpha} e^{x\beta} + (t_2^{\alpha} - t_1^{\alpha})e^{x(\beta+\mu_2)} + (t^{\alpha} - t_2^{\alpha})e^{x(\beta+\mu_3)}$. The likelihood contribution of an individual is written as

$$\ell = \prod_{j=1}^{q} [h(t)^c e^{-I(t)}]^{d_j}, \tag{11}$$

where d_j indicates the interval, i.e. $d_j = 1$ if $t_{j-1} < t \le t_j$ otherwise $d_j = 0$.

The explanatory variables may be time-dependent as well, i.e. the time-dependent variables may take different values in the intervals. In this paper the interest concerning the PH assumption is in a single time-dependent explanatory variable, the benefit replacement ratio. Its effect is tested in two intervals $(t_0,t_1]$ and $(t_1,t_2]$, where $t_0 = 0$, $t_1 = 3$ and $t_2 = 24$ months. The economic reason for estimating the change of the hazard function is that after the first three unemployment months the rules of the UI system are different. Thus, in our case $x(\beta+\mu_j)$ is written as $x\beta + x_{rj}(\beta_r+\mu_j)$, where x_{rj} is the benefit replacement ratio in the intervals $j=1,2$, β_r is its parameter and μ_j is its parameter in the interval j. The rest of the explanatory variables x are constant over time.

Testing the proportional hazards assumption

In this section a score test for the PH assumption is presented. Specification tests are particularly important for many econometric models estimated by maximum likelihood, such as parametric duration models, where few diagnostic tests are currently available. A chi-squared test for the PH assumption based on the difference between the number of failure times observed and its expected value in each category from a given partition of the time axis is suggested by Schoenfeld (1980). A Wald type of test for the PH assumption in a two-step regression model has been suggested by Anderson and Senthilselvan (1982). Moreau, O'Quigley and Mesbach (1985) presented a score test for checking the assumption. The test was extended by O'Quigley and Pessione (1989). All the tests have been developed in the context of Cox's model and they are not applicable to parametric duration models.

The alternative non-proportional hazards model assumes that the effect of a covariate varies as a step function. However, this model may be awkward to estimate. Therefore it seems preferable to avoid such an estimation and develop a score test. The null hypothesis for the PH model is H_0: $\mu_2 = \mu_3 = \ldots = \mu_q = 0$, which leads to the hazard function

$$h(t) = \alpha t^{\alpha-1} e^{x\beta+x_{rj}\beta_r}. \tag{12}$$

The score test is based on the statistic $S_j = n^{-1/2}L_{\mu_j}$, where $L_{\mu_j} = L/ \mu_j$ is evaluated under the null hypothesis. In the Weibull case it becomes

$$S_j = \begin{cases} 0, & t \leq t_{j-1} \\ n^{-1/2}\{c-\epsilon(t)[1-(t_{j-1}/t)^{\alpha}]\}x_{rj}, & t_{j-1} < t \leq t_j \\ n^{-1/2}\{ -\epsilon(t)[(t_j/t)^{\alpha}-(t_{j-1}/t)^{\alpha}]\}x_{rj}, & t_j < t, \end{cases} \tag{13}$$

which have been written using the generalized residuals $\epsilon(t) = t^{\hat{\alpha}} e^{x\hat{\beta}+x_{rj}\hat{\beta}_r}$, i.e. the integrated hazard of the fitted Weibull model. The definition of residuals is found in Cox and Snell (1968).

The information matrix $- E[\partial^2 L/\partial\theta\partial\theta']$ can be consistently estimated by

$$I = - n^{-1} \begin{bmatrix} L_{\mu\mu} & L_{\mu\beta} \\ L_{\beta\mu} & L_{\beta\beta} \end{bmatrix}. \tag{14}$$

It can be inverted using the method outlined by Theil (1983, p. 13). The top left hand block of the inverse of the information matrix is of relevance towards calculating the test statistic

$$V = -\ n^{-1}\ (L_{\mu\mu} - L_{\mu\beta}\ L_{\beta\beta}^{-1}\ L_{\beta\mu}'). \qquad (15)$$

Under the null hypothesis $\hat{S}$ is asymptotically $N_{p(q-1)}(0,V)$. The asymptotic null distribution of the score test is not affected if the required estimates of V are evaluated using any estimator which is consistent under H_0. The matrix V can be expressed consistently as the outer product form of the information matrix identity

$$V = n^{-1}\ (L_\mu'L_\mu - L_\mu'L_\beta(L_\beta'L_\beta)^{-1}\ L_\beta'L_\mu), \qquad (16)$$

which is a convenient form because it requires neither an expression for the Hessian of the log likelihood function nor analytic evaluation of the information matrix. The score test statistic is then of the form

$$T = \hat{S}'\hat{V}^{-1}\hat{S}. \qquad (17)$$

The statistic is based upon the result that the quadratic form $S'V^{-1}S$ manifests an asymptotic chi-squared distribution when the null hypothesis is true. The test statistic can then be calculated as nR^2 from a pseudo-regression based on ordinary least squares, where a vector of ones is regressed on L_α, L_β and L_μ. The procedure based on the pseudo-regressions is described by Chesher (1983) and Lancaster (1984).

3. The effects of omitted variables

Gamma heterogeneity

The first approach in this section combines the correction for gamma heterogeneity and a parametric duration model with time-dependent covariates. The usual method for incorporating heterogeneity is to assume a parametric functional form for the pattern of heterogeneity. The gamma mixing distribution has been chosen because it is analytically simple to use and it provides quite a flexible model for the distribution of the heterogeneity component. If unobserved characteristics are not adequately captured by explanatory variables, this may lead to biases in parameter estimates. Even if the omitted variables are uncorrelated with those which are included in the model, the parameters will be biased. It may be shown under fairly general conditions that the coefficients of explanatory variables are then biased towards zero (Lancaster and Nickell, 1980). Therefore, the parameters of the model may be expected to increase in absolute value when omitted variables are taken into account.

Suppose that the individuals in the sample differ to a certain degree with respect to some unobservable variable, say, motivation v. Each individual has his own v and hence his own hazard function h(t). Lancaster (1979) assumed that these hazards have a gamma distribution. The conditional hazard function in a Weibull model allowing for time-dependent effects and gamma heterogeneity is

$$h(t|v) = v\alpha t^{\alpha-1} e^{x(\beta+\mu_j)}, \tag{18}$$

where v has a gamma density. The expected value of v is normalized to one and its variance σ^2 is estimated. The marginal survivor function, not conditional on v, is obtained by integrating over the assumed mixing distribution. The density function is obtained from the survivor function by differentiating and the hazard function is obtained as a ratio of the density and survivor functions. The hazard function allowing for gamma heterogeneity is written as

$$h_g(t) = \alpha t^{\alpha-1} e^{x(\beta+\mu_j)} [1 + \sigma^2 I(t)]^{-1}, \tag{19}$$

and the corresponding integrated hazard is

$$I_g(t) = 1/\sigma^2 \log[1 + \sigma^2 I(t)], \tag{20}$$

where I(t) is the integrated hazard of the original model (10). In our case $x\beta + x_{rj}(\beta_r+\mu_j)$ is substituted for $x(\beta+\mu_j)$, because we are interested in the time-dependent effects of a single time-dependent variable. The hazard function (19) and integrated hazard (20) are substituted into the likelihood function (4) to estimate the unknown parameters.

Mass point heterogeneity

In this section a mass point approach to the incorporation of heterogeneity into duration models is described. The main method for incorporating heterogeneity has been to assume a parametric functional form for the pattern of heterogeneity. Heckman and Singer, who propose a discrete pattern of heterogeneity (1984 a,b), have shown that estimates of the structural parameters may be sensitive with respect to the parametric forms assumed for heterogeneity. Furthermore, there are a limited number of tractable forms for mixing distributions available.

The approach dispensing with the need to specify a parametric distribution for the heterogeneity component has its origins in the work of Kiefer and Wolfowitz (1956), who showed that a nonparametric characterization of the heterogeneity distribution ensures consistent estimation of simultaneously estimated structural parameters. Further work on the properties of mass point mixing distributions has been carried out by Simar (1976), Laird (1978), Lindsay (1983 a,b) and Heckman and Singer (1984 a,b). Applications of the mass point approach in the context of discrete choice models have been presented by Davies and Crouchley (1984), Dunn, Reader and Wrigley (1987), Davies (1987) and Card and Sullivan (1988). Applications to duration models have been presented by Brännäs (1986 a,b), Trussell and Richards (1987) and Ham and Rea (1987).

The idea of mass point models is that the constant parameter of the basic model β_0 is partitioned into m location parameters u_k and each of the location parameters is given a probability p_k. Thus the explanatory variables do not include the vector of ones. In the case where m = 1, when there is one location parameter, the parameter u_1 is equal to the constant of the basic Weibull model β_0. Consequently, the likelihood function of mass point models reduces to the likelihood function of the basic Weibull model, and the model with one mass point and the basic Weibull model coincide.

In the case of parametric duration models the mixing likelihood for an individual can be written as

$$f_Q = \prod_{j=1}^{q} \sum_{k=1}^{m} [p_k h_k(t)^c e^{-I_k(t)}]^{d_j}, \tag{21}$$

where $h_k(t) = \alpha t^{\alpha-1} e^{u_k + x(\beta+\mu_j)}$ and $I_k(t) = t^{\alpha} e^{u_k + x(\beta+\mu_j)}$ are the hazard function and integrated hazard for the group k in the Weibull case. In a model with one time-dependent explanatory variable $x\beta + x_{rj}(\beta_r+\mu_j)$ is substituted for $x(\beta+\mu_j)$.

To ensure that $p_k \in (0,1)$ and $\Sigma\, p_k = 1$, the probabilities associated with each location have been defined using a multinomial logit type of formula

$$p_k = \frac{e^{g_k}}{1 + \sum_{l=1}^{m-1} e^{g_l}}, \quad k = 1,\ldots,m-1, \tag{22}$$

where g_l, $l = 1,\ldots,m-1$ are parameters to be estimated. The probability of the last mass point is $p_m = 1-p_1-p_2-\ldots-p_{m-1}$. By definition $p_1 = 1$ when $m = 1$. The parameters g_1 work only as a device, and they do not have an interesting economic interpretation in this context.

The objective is to estimate the discrete mixing distribution Q consistently with the atomic densities, a maximizer of the mixture likelihood function $\ell(Q) = \pi\ f_Q$. Maximizing the likelihood function $\ell(Q)$ over Q may be accomplished by maximizing the log likelihood function $L(f) = \Sigma \log f_Q$. Following Lindsay (1983a) it can be seen that $L(f) = \Sigma \log f_Q$ is differentiable with the directional derivative of L at L_{Q_0} towards L_{Q_1} being

$$D(u;Q) = \lim_{p\to 0} \{L[(1-p)f_{Q_0} + pf_{Q_1}] - L(f_{Q_0})\}/p$$

$$= \Sigma\ [(f_{Q_1} - f_{Q_0})/f_{Q_0}]$$

$$= \Sigma\ f_{Q_1}/f_{Q_0} - n, \qquad (23)$$

where it will be understood that the summing is over observations. The procedure of estimating a discrete mixing distribution is to increase the number of points of support until $D(u;Q) \le 0$. Then the procedure is stopped and the semiparametric ML estimator is obtained. This procedure is suggested also by Brännäs and Rosenqvist (1988) in the context of count data models. Maximum likelihood algorithms are directly applicable to the constrained problem of maximization over discrete mixtures Q with a fixed number of support points. The Berndt, Hall, Hall and Hausman (1974) algorithm is used to estimate the unknown parameters. A simple first order check for a global maximum is to verify that $D''(u^*; Q) \leq 0$ at the support points of measure Q.

4. The results

Data on 2077 Finnish unemployed persons was collected for this study from various registers. The data set is more reliable than the data sets based on surveys. In order to guarantee that the sample would be randomly generated, every hundredth individual was picked from the flow into unemployment during the year 1985. The sample was taken from the unemployment register of the Ministry of Labour. The individuals were then followed until the end of their unemployment periods, but at most until the end of 1986. The income and asset information was compiled into the data from the tax register. The information on unemployment benefits was compiled from the registers of the bank Postipankki and the Social Insurance Institution. 40 % of the observations are right censored, i.e. the complete spells of unemployment were not observed. The appendix includes the description of the variables used in the study.

The results of the estimations are presented in Table 1. The first model is the basic Weibull model with the hazard function (8), where an average replacement ratio over the unemployment period is used. A Weibull model with hazard function (12) including the time-dependent replacement ratios is estimated in the two intervals, $(t_0, t_1]$ and $(t_1, t_2]$, where $t_0 = 0$, $t_1 = 3$ and $t_2 = 24$ months. The second column of Table 1 includes this model. The negative effect of the replacement ratio β_r decreases substantially when time-dependent replacement ratios are introduced into the model.

A score test for the PH assumption is made. The test statistics calculated under H_0 takes a value of 8.84, which exceeds the critical value $\chi^2_{1,0.95} = 3.84$. The conclusion is that the PH assumption is rejected for the replacement ratio. The test suggests estimating a model with time-dependent parameters.

A Weibull model with time-dependent effects are in the third model. The parameter estimate β_r takes a value -0.894, and after the first three months the additional parameter estimate μ_2 takes a positive value 0.871. Unemployed persons who are eligible for benefits face a risk of losing benefits after the first three months. The effect of risk increases the re-employment probability and it is captured by the parameter μ_2.

The PH assumption is not valid either after allowing for gamma heterogeneity in the fourth model, since the corresponding parameter estimates take the values -1.506 and 1.475. It is interesting to note that Nickell (1979) using the data from the U.K. and a different kind of model found similar effects of UI benefits. In both of the studies the effect is first negative and statistically significant but later on the effect vanishes.

Table 1

Time-dependent effects of UI benefits and gamma heterogeneity

	(1)	(2)	(3)	(4)
	Std.errors in parentheses			
Shape parameter	0.861	0.853	0.822	1.096
	(0.020)	(0.021)	(0.022)	(0.054)
Variance of heterogeneity				0.955
				(0.161)
Constant	-1.478	-1.576	-1.454	-1.183
	(0.136)	(0.137)	(0.139)	(0.207)
Children	-0.004	-0.088	-0.082	-0.134
	(0.050)	(0.050)	(0.049)	(0.074)
Married	0.170	0.207	0.203	0.198
	(0.065)	(0.066)	(0.065)	(0.099)
Sex	-0.007	-0.040	-0.041	-0.074
	(0.056)	(0.057)	(0.057)	(0.088)
Age	-0.042	-0.044	-0.044	-0.060
	(0.003)	(0.003)	(0.003)	(0.005)
Level of education	0.064	0.053	0.066	0.029
	(0.058)	(0.059)	(0.059)	(0.093)
Training for employment	0.176	0.187	0.181	0.331
	(0.072)	(0.074)	(0.073)	(0.116)
Member of UI fund	0.213	0.243	0.237	0.356
	(0.060)	(0.062)	(0.061)	(0.093)
Came from schooling	0.291	0.299	0.291	0.380
	(0.078)	(0.079)	(0.078)	(0.128)
Came from house work	-0.711	-0.716	-0.731	-0.895
	(0.124)	(0.125)	(0.124)	(0.171)
Regional demand	0.168	0.308	0.271	0.487
	(0.238)	(0.240)	(0.240)	(0.330)
Occupational demand	0.641	0.743	0.715	0.457
	(0.600)	(0.609)	(0.602)	(0.943)
Taxable assets	1.021	1.997	1.224	1.223
	(1.080)	(1.102)	(1.073)	(1.393)
Replacement ratio, β_r	-1.223	-0.376	-0.894	-1.506
	(0.150)	(0.120)	(0.191)	(0.264)
Replacement ratio, μ_2			0.871	1.475
			(0.208)	(0.267)
Log likelihood	-4962.5	-4993.4	-4985.4	-4950.4

The results of estimations of mass point models are in Table 2. The model with two mass points produces approximately constant hazard functions for the two groups which are not controlled for explanatory variables. The models with three or four mass points produce increasing hazard functions. An increasing hazard function is in concordance with standard search theories. The absolute values of statistically significant parameter estimates increase in most cases when more mass points are introduced into the model, as is to be expected.

Lindsay's criterion was used to decide the number of mass points. The values of the D function of the models with 2,3 and 4 mass points are 1.34, 9.12, and -3.63 respectively, showing that four mass points are enough to rectify the effect of omitted variables in this data set.

Many of the explanatory variables have significant effects on the re-employment probability. Unemployed persons who have children or persons who are not married have lower probabilities. Age is a very significant factor. Old people are more apt to incur problems in finding jobs. Training for further employment has a significant and positive effect on the re-employment probability. Members of the UI funds, i.e. members of the Finnish labour unions, are often skilled workers and therefore they become employed earlier than the non-members. The persons leaving school or the army usually have no great problems. They leave unemployment clearly earlier than the others. The persons who have come from house work find it very difficult to find a job.

In the model with four mass points the parameter estimate of the replacement ratio takes a value of -1.890 and the additional parameter μ_2 due to the risk of losing benefits takes a value of 1.752. The negative effect of the replacement ratio vanishes after the first three months and the PH assumption is not valid.

Table 2

Time-dependent effects of UI benefits and mass point heterogeneity

	Number of mass points		
	m=2	m=3	m=4
	Std.errors in parentheses		
Shape parameter	0.987	1.194	1.288
	(0.040)	(0.064)	(0.091)
Number of children	-0.109	-0.175	-0.184
	(0.065)	(0.079)	(0.089)
Married	0.201	0.216	0.199
	(0.087)	(0.103)	(0.115)
Sex	-0.060	-0.088	0.085
	(0.077)	(0.092)	(0.104)
Age	-0.055	-0.063	-0.069
	(0.004)	(0.006)	(0.007)
Level of education	0.029	0.066	0.040
	(0.080)	(0.098)	(0.109)
Training for employment	0.331	0.240	0.296
	(0.102)	(0.118)	(0.132)
Member of UI fund	0.333	0.368	0.411
	(0.083)	(0.097)	(0.110)
Came from schooling	0.341	0.408	0.441
	(0.110)	(0.131)	(0.148)
Came from house work	-0.819	-1.011	-1.114
	(0.153)	(0.187)	(0.206)
Regional demand	0.440	0.420	0.489
	(0.297)	(0.350)	(0.388)
Occupational demand	0.609	0.451	0.512
	(0.826)	(0.980)	(1.106)
Taxable assets	0.881	2.087	1.369
	(1.255)	(1.594)	(1.621)
Replacement ratio, β_r	-1.276	-1.741	-1.890
	(0.229)	(0.303)	(0.345)
Replacement ratio, μ_2	1.302	1.644	1.752
	(0.244)	(0.297)	(0.327)
u_1	-0.980	-0.004	0.459
	(0.193)	(0.247)	(0.366)
u_2	-2.738	-2.103	-1.360
	(0.238)	(0.259)	(0.478)
u_3		-4.865	-2.792
		(0.816)	(0.489)
u_4			-7.277
			(7.803)
g_1	0.610	0.890	0.992
	(0.262)	(0.387)	(0.346)
g_2		1.709	1.608
		(0.372)	(0.544)
g_3			1.752
			(0.353)
p_1	0.648	0.272	0.187
p_2	0.352	0.617	0.345
p_3		0.112	0.399
p_4			0.069
Log likelihood	-4956.5	-4945.9	-4943.6

5. Concluding remarks

In this paper the effects of unemployment benefits on the unemployment spells were examined. In the Finnish system the circumstances of an unemployed person changes starting at a known time point. In this paper it was allowed for the effects of unemployment benefits to vary with the duration of unemployment spell, remaining constant within predefined intervals. A technique for estimating the time-dependent effects of time-dependent explanatory variables on the re-employment probability was presented.

Often it may be preferable to avoid estimating an alternative non-proportional hazards model, because the estimations of these kinds of models may not be straightforward. Therefore the focus was at first on a score test. Tests for the PH assumption has been studied in the context of Cox's model by many authors. In this paper a score test for the PH assumption was extended to parametric duration models. The test shows that the effect of benefits do not stay constant during the unemployment spell.

If the average replacement ratio during the unemployment spell is used the effect of the replacement ratio on the re-employment probability is negative. However, the microeconomic data collected from various registers include the time-dependent replacement ratios. Alternative models with time-dependent effects of unemployment benefits were estimated. The replacement ratio has a negative effect on the re-employment probability during the first three months, but after that period the effect vanishes.

Even though the data are rich of explanatory variables and more reliable than the data from surveys, there is reason to assume that relevant variables have been omitted from the model. The influence of omitted variables was taken into account in estimation assuming that the effects have a gamma and discrete mass point distribution. When heterogeneity is introduced into the model the absolute values of parameter estimates increase, but the correction for omitted variables does not eliminate the result that the effect of UI benefits vanish after the first three months.

References

Andersen, J.A. and A. Senthilselvan, 1982, A Two-step Regression Model for Hazard Functions, Applied Statistics 31, 44-51.

Berndt, E.R., B.H. Hall, R.E. Hall and J.A. Hausman, 1974, Estimation and Inference in Nonlinear Structural Models, Annals of Economic and Social Measurement 3, 653-666.

Brännäs, K., 1986a, Small Sample Properties in a Heterogeneous Weibull Model, Economics Letters 21, 17-20.

Brännäs, K., 1986b, On Heterogeneity in Econometric Duration Models, Sankhya 48, 284-93.

Brännäs, Kurt and Gunnar Rosenqvist, 1988, Semiparametric Estimation of Heterogenous Count Data Models, Working Papers no. 187, Swedish School of Economics and Business Administration.

Card, D, and D. Sullivan, 1988, Measuring the Effect of Subsidized Training Programs on Movements in and out of Employment, Econometrica 56, 497-530.

Chesher Andrew, 1983, The Information Matrix Test, Simplified Calculation Via a Score Test Interpretation, Economic Letters 13, 45-48.

Cox, D.R., 1972, Regression Models and Life-Tables, Journal of the Royal Statistical Society B, 34, 187-202.

Cox, D.R., 1975, Partial Likelihood, Biometrika 62, 269-276.

Cox, D.R. and E.J. Snell, 1968, A General Definition of Residuals, Journal of the Royal Statistical Society, Series B, 30, 248-275.

Davies, R.B., 1987, Mass Point Methods for Dealing with Nuisance Parameters in Longitudinal Studies, in: R. Crouchley, ed., Longitudinal Data Analysis, Surrey Conference on Sociological Theory and Method 4, Gower Publishing Company Limited, Avebury, 1987.

Davies, R.B. and R. Crouchley, 1984, Calibrating Longitudinal Models of Residential Mobility and Migration, Regional Science and Urban Economics, 14, 231-47.

Dunn, R., S. Reader and N. Wrigley, 1987, A Nonparametric Approach to the Incorporation of Heterogeneity into Repeated Polytomous Choice Models of Urban Shopping Behaviour, Transportation Research 21A, 327-43.

Ham, J.C. and S.A. Rea Jr, 1987, Unemployment Insurance and Male Unemployment Duration in Canada, Journal of Labor Economics 5, 325-53.

Han, Aaron and Jerry A. Hausman, 1986, Semiparametric Estimation of Duration and Competing Risk Models, Working Paper, M.I.T.

Heckman, James J. and Burton Singer, 1984a, Econometric Duration Analysis, Journal of Econometrics 24, 63-132.

Heckman, James J. and Burton Singer, 1984b, A Method for Minimizing the Impact of Distributional Assumptions in Econometric Models for Duration Data, Econometrica 52, 271-320.

Kiefer, J. and J. Wolfowitz, 1956, Consistency of the Maximum Likelihood Estimator in the Presence of Infinitely Many Incidental Parameters, Annals of Mathematical Statistics 27, 887-906.

Laird, N., 1978 Nonparametric Maximum Likelihood Estimation of a Mixing Distribution, Journal of the American Statistical Association 73, 783-92.

Lancaster Tony, 1979, Econometric Methods for the Duration of Unemployment, Econometrica 47, 939-956.

Lancaster, Tony, 1984, The Covariance Matrix of the Information Matrix Test, Econometrica 52, 1051-1053.

Lancaster, Tony and Stephen Nickell, 1980, The Analysis of Re-Employment Probabilities for the Unemployed, Jouranal of the Royal Statistical Society, Series A, 143, 141-165.

Lindsay, B.G., 1983a, The Geometry of Mixture Likelihoods: A General Theory, The Annals of Statistics 11, 86-94.

Lindsay, B.G., 1983b, The Geometry of Mixture Likelihoods, Part II: The Exponential Family, The Annals of Statistics 11, 783-92.

Meyer, Bruce D., 1990, Unemployment Insurance and Unemployment Spells, Econometrica 58, 757-782.

Moffit, Robert, 1985, Unemployment Insurance and the Distribution of Unemployment Spells, Journal of Econometrics 28, 85-101.

Moreau T., J. O'Quigley and M. Mesbah, 1985, A Global Goodness-of-fit Statistic for the Proportional Hazards Model, Applied Statistics 34, 212-218.

Nickell Stephen, 1979, Estimating the Probability of Leaving Unemployment, Econometrica 47, 1249-1266.

O'Quigley, John, Fabienne Pessione, 1989, Score Tests for Homogeneity of Regression Effect in the Proportional Hazards Model, Biometrics 45, 135-144.

Prentice, Ross and L. Gloeckler, 1978, Regression Analysis of Grouped Survival Data with Application to Breast Cancer Data, Biometrics 38, 57-67.

Schoenfeld, D., 1980, Chi-squared Goodness-of-fit Tests for the Proportional Hazards Regression Model, Biometrika 67, 145-153.

Simar, L., 1976, Maximum Likelihood Estimation of a Compound Poisson Process, The Annals of Statistics 4, 1200-9.

Theil, Henri, 1983, Linear Algebra and Matrix Methods in Econometrics, in: Zvi Griliches and Michael D. Intrilligator, eds., Handbook of Economics, vol I, (North-Holland, Amsterdam).

Trussell, J. and T. Richards, 1985, Correcting for Unmeasured Heterogeneity in Hazard Models Using the Heckman-Singer Procedure, Sociological Methodology, 242-76.

Data appendix

Variables of the data

Duration of unemployment is calculated in weeks and it is the difference between the date of entry into unemployment and the date of returning back to work. Mean = 15.03.

Number of children is the number of unemployed person's children who are younger than 18 years. Mean = 0.23.

Married is a dummy variable, 1 = yes. Mean = 0.37.

Sex is a dummy variable, 1 = male. Mean = 0.54.

Age is measured in years. Mean = 31.2.

Level of education is a dummy variable, 1 = at least 12 years of education. The level of education is based on the education code of the Central Statistical Office of Finland. Mean = 0.45.

Training for employment is a dummy variable, 1 = The person has got training for further employment. Mean = 0.15.

Member of UI fund is a dummy variable, 1 = yes. Mean = 0.42.

Came from schooling is a dummy variable, 1 = The person has come from schooling or from the army. Mean = 0.13.

Came from house work is a dummy variable, 1 = The person has come from house work or elsewhere outside the labour force. Mean = 0.07.

Regional demand describes the regional rate of jobs available. It is the number of vacancies divided by the number of job seekers in the area. Mean = 0.10.

Occupational demand describes the occupational rate of jobs available in the whole country. It is the number of vacancies divided by the number of job seekers in the occupation group. Mean = 0.12.

Taxable assets has been compiled from the tax register and it is measured in millions of marks. Mean = 0.011.

Replacement ratios in the intervals, 1-3 and 3-24 months are unemployed person's average replacement ratios after tax in those intervals. Average weekly unemployment benefits after tax have been divided by the average weekly income in 1985 after tax. Means: 0.15 and 0.25.

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