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TIME-DEPENDENCE IN SEMI-PARAMETRIC MODELS OF UNEMPLOYMENT DURATION

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ABSTRACT: This paper examines the duration and calendar-dependence in the context of Cox's semi-parametric proportional hazards model. Duration-dependent unemployment benefits are used to study duration-dependent features of the Finnish unemployment insurance system and calendar-dependent covariates are used to study and the seasonal effects on the re-employment. Calendar-dependent covariates may not always be adequate to model the macroeconomic changes. Therefore the roles of duration and calendar time are changed in estimating the model. The underlying baseline hazards of the models based on duration and calendar time are presented.

1. Introduction

The proportional hazards model presented by Cox (1972) studies the effects of explanatory variables on the hazard rate without specifying the form of duration-dependence. The estimation of Cox's model leads to the partial conditional likelihood function, where the time-dependent part of the likelihood function is cancelled out, because it is identical for the individuals leaving unemployment and individuals in the risk set.

The environment of an unemployed person changes over the time. The changes may be duration-dependent. The Finnish unemployment insurance system is such that the rules concerning the eligibility of benefits vary over the duration of unemployment. On the other hand the changes may be calendar-dependent. Probably the most important calendar-dependent changes in the short run are the seasonal factors, which have an important effect on the re-employment probability. The interest of this study is in the calendar and duration-dependent effects on the re-employment probability.

Concerning any period the individuals experience two events, the entry τ^0 and exit τ^1, measured in calendar time. The calendar time is measured as the duration between the date and any fixed date before the events in question. The duration of unemployment is then $t = \tau^1 - \tau^0$. The hazard function of the proportional hazards model presented by Cox (1972) can be written

$$(1) \qquad h(t,x) = h_0(t)h_1(x;\beta),$$

where the first factor $h_0(t)$ is the unknown baseline hazard. These kinds of models are called semi-parametric, since one does not have to define the baseline hazard. The second factor, which is known up to a finite dimensional parameter vector β, usually takes the log linear form $h_1(x;\beta) = \exp(x\beta)$.

Let $t_1, t_2, \ldots, t_n$ denote the ordered durations of n individuals. The partial likelihood contribution can be written as follows [Cox (1975)]

$$(2) \qquad \ell_i(\beta) = \frac{\exp(x_i\beta)}{\sum_{j \in R(t_i)} \exp(x_j\beta)},$$

where $R(t_i)$ denotes the risk set, i.e. the observations with $t \geq t_i$. Multiplying the numerator and denominator by $h_0(t)dt$ it can be seen that the contribution of an observation i is just the probability that the duration ends in $[t_i, t_i+dt)$ given that some duration in the risk set ends in that interval.

Usually duration data includes censored observations. In estimation it is helpful to use an indicator to signify whether observations are complete durations or censoring times. Values of the censoring variable C and duration variable T are observed. The censoring indicator $c_i = 1$ if t_i is a complete spell, otherwise $c_i = 0$ and $t_i = \min(T_i, C_i)$. The risk set includes censored observations, which appear only in the numerator of the likelihood function. With censoring indicator the partial likelihood function can be written

$$(3) \qquad \ell(\beta) = \prod_{i=1}^{n} \ell_i(\beta)^{c_i} .$$

If $c_i = 0$, no contribution is made to the likelihood function.

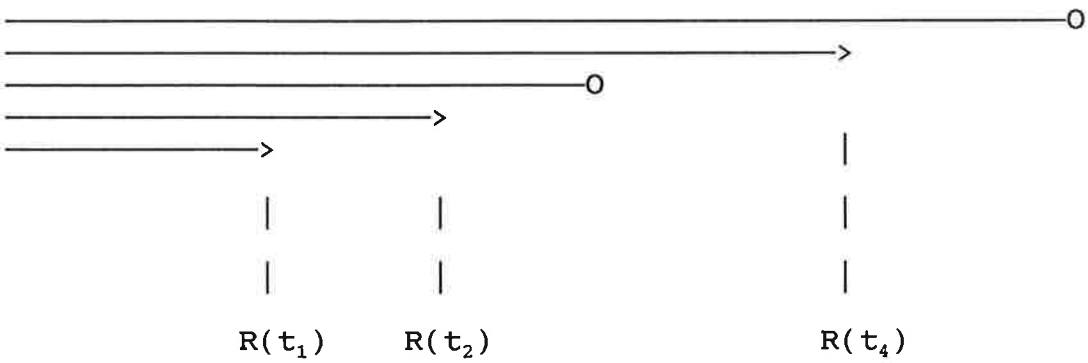
Usually duration data are to some degree grouped, i.e. there are spells, which end during the same unit of time. In this data the duration of unemployment is measured using dates. Therefore the grouping is rather mild. One way of taking the grouping into account is to write the partial likelihood function following Breslow (1974)

$$(4) \qquad \ell(\beta) = \prod_{i=1}^{d} \frac{\exp(\sum_{l \in T(t_i)} c_l x_l \beta)}{[\sum_{j \in R(t_i)} \exp(x_j \beta)]^{m_i}} ,$$

where d is the number of distinct exit times observed in the data and $T(t_i)$ denotes a set of m_i individuals who are observed to leave unemployment at t_i. Note that if $T(t_i)$ includes only censored observations then $m_i = 0$ and no contribution is made to the likelihood function.

In estimation of the unknown parameters β it is helpful to sort the durations and censoring times in descending order for the calculation of the risk set as shown in Figure 1. The calculation of the risk set for each distinct duration can then be done by adding $\exp(x_j\beta)$ terms of the individuals starting from the first observation. The partial likelihood function can then be easily maximized with respect to the parameters. The explanatory variables of the risk sets may vary even though the individuals are homogenous, since some of the explanatory variables may be calendar or duration-dependent. The effects of these variables are studied using Finnish data of unemployed workers.

Figure 1. Unemployment durations of five individuals (O = censored, > = complete spells)



Using parametric models of unemployment duration Kettunen (1990, 1991b, 1991c) has shown that the benefits have a negative effect on the re-employment probability and labour mobility. However, using parametric models the specification of duration distribution and duration dependence may be difficult. The Finnish unemployment insurance system is such that the rules concerning the eligibility of benefits vary over the duration of unemployment. Therefore the interest of this study is in the semi-parametric models and duration-dependent effects of the UI system on the re-employment and regional and occupational mobility of unemployed workers.

2. Time-Dependent Effects

2.1. Duration-Dependence

The values of explanatory variables may change over the duration of unemployment for the individuals. This section considers the inclusion of such variables in the semi-parametric proportional hazards model. Models with duration-dependent replacement ratios of unemployment benefits are estimated using different kinds of model specifications. With duration-dependent covariates the hazard function of the proportional hazards model can be written

$$(5) \qquad h(t,x,z) = h_0(t)e^{x\beta + z(t)\beta_z},$$

where x includes the covariates, which are constant in time and z(t) includes the duration-dependent covariates.

Models with duration-dependent benefit replacement ratios are estimated using the data on 2077 unemployed workers, who became unemployed during 1985 [see Kettunen (1991a)]. For comparison the first column of Table 1 includes the model, where the benefit replacement ratio is fixed at an average value over the unemployment spell. The second column includes a model, where the duration-dependent unemployment benefits were used 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 results show that the effect of unemployment benefits is lower with duration-dependent replacement ratios. The third possibility is to assume that the effects of duration-dependent variables vary over the time, remaining constant within predefined intervals. The reason for this kind of specification is that the explanatory variables may have different effects on the re-employment probability at different durations. With duration-dependent effects in the intervals $t_{j-1} < t \leq t_j$, $j = 1, 2$, the hazard function can be written

$$(6) \qquad h_j(t,x) = h_0(t)e^{x\beta + z(t)(\beta_z - \mu_j)}.$$

To avoid singularity it is defined that $\mu_1 = 0$. This approach has been followed by Moreau, O'Quigley and Mesbach (1985), Moreau, O'Quigley and Lellouch (1986) and O'Quigley and Pessione (1989).

The results of model (6) with duration-dependent replacement ratios are in the third column of Table 1. It can be seen that the unemployment benefits have a negative effect on the re-employment probability during the first three months, but after that period the effect turns positive. An obvious reason is that the eligibility rules of benefits become more strict. The unemployed persons have a risk of losing benefits after the first three months if they do not move to another region or change their occupations. Furthermore, after the 100*th* day of unemployment the earnings-related unemployment allowances decrease by 20 %. Because of these rules the incentive towards re-employment is higher for the persons with high benefits. These findings are confirmed by Table 2, which shows that the negative effect of benefits is higher for the non-members of labour unions even though their benefits are lower.

Since the rules concerning the labour mobility change during the unemployment, it is reasonable to model the probabilities of moving and changing occupations. The semi-parametric models of labour mobility are presented in Table 3. In the case of regional mobility the censoring indicator $c_i = 1$ if the person has moved to get a job, otherwise $c_i = 0$. In the case of occupational mobility the censoring indicator is defined in a similar way.

Table 1. Semi-parametric models of unemployment spell with duration-dependent replacement ratios

	Std.errors in parentheses		
Number of children	-0.002	-0.097	-0.086
	(0.048)	(0.049)	(0.049)
Married	0.143	0.171	0.159
	(0.067)	(0.070)	(0.069)
Sex	-0.014	-0.054	-0.056
	(0.063)	(0.061)	(0.061)
Age	-0.039	-0.037	-0.036
	(0.004)	(0.003)	(0.003)
Level of education	0.045	0.081	0.102
	(0.067)	(0.064)	(0.064)
Training for employment	0.183	0.206	0.202
	(0.074)	(0.079)	(0.079)
Member of UI fund	0.209	0.216	0.205
	(0.063)	(0.065)	(0.065)
Came from schooling	0.283	0.300	0.280
	(0.090)	(0.083)	(0.084)
Came from house work	-0.648	-0.655	-0.671
	(0.140)	(0.137)	(0.137)
Regional demand	0.114	0.248	0.131
	(0.256)	(0.252)	(0.252)
Occupational demand	0.551	0.656	0.547
	(0.586)	(0.622)	(0.590)
Taxable assets	0.783	0.770	0.682
	(0.994)	(1.112)	(1.112)
Replacement ratio, β_z	-1.232	-0.325	-1.375
	(0.132)	(0.138)	(0.205)
Replacement ratio, μ_2			2.127
			(0.271)
Log likelihood	-8415.6	-8453.6	-8422.1
Number of observations	2077	2077	2077

Table 2. Semi-parametric models of unemployment duration for the non-members and members of labour unions

(A) Non-members
(B) Members

	(A)	(B)
	Std.errors in parentheses	
Number of children	-0.019	-0.016
	(0.069)	(0.073)
Married	0.048	0.235
	(0.106)	(0.095)
Sex	-0.021	-0.047
	(0.083)	(0.096)
Age	-0.032	-0.048
	(0.005)	(0.005)
Level of education	0.154	-0.154
	(0.085)	(0.103)
Training for employment	0.181	0.179
	(0.119)	(0.108)
Came from schooling	0.300	0.200
	(0.097)	(0.188)
Came from house work	-0.729	-0.548
	(0.191)	(0.201)
Regional demand	-0.061	0.410
	(0.327)	(0.411)
Occupational demand	-0.387	1.276
	(0.925)	(0.886)
Taxable assets	0.120	1.494
	(2.015)	(1.193)
Replacement ratio	-1.725	-0.851
	(0.219)	(0.221)
Log likelihood	-4178.7	-3362.2
Number of observations	1212	865

Table 3. Semi-parametric duration models of labour mobility

(A) Regional mobility (B) Occupational mobility	(A)	(B)
	Std. errors in parentheses	
Age	-0.053	-0.035
	(0.022)	(0.010)
Level of education		0.305
		(0.182)
Training for employment		0.340
		(0.194)
Member of UI fund	-1.382	0.376
	(0.467)	(0.156)
Came from schooling		-0.034
		(0.248)
Came from house work		0.084
		(0.254)
Regional demand	-2.036	1.223
	(1.433)	(0.597)
Occupational demand	2.836	-3.544
	(3.179)	(1.637)
Taxable assets		-3.918
		(3.264)
Replacement ratio	-5.116	-0.446
	(1.005)	(0.334)
Log likelihood	-321.8	-1364.3
Number of observations	2077	2077

2.2. Calendar-Dependence

This section studies the seasonal effects on the re-employment using semi-parametric models of unemployment duration. With calendar-dependent covariates the hazard function of Cox's model can be written

$$(7) \qquad h(t,x,z) = h_0(t)e^{x\beta + z(\tau)\beta_z},$$

where x includes the covariates which are constant in time and $z(\tau)$ includes the calendar-dependent covariates. The seasonal variation and the effects of quarterly unemployment rates are examined in this study.

One way of introducing time-dependence is to use dummy variables to indicate periods in calendar time. Ridder (1987) allows for dependence on calendar-time by introducing dummies for two-year intervals. Quarterly dummy variables are in this study introduced to the model for the examination of seasonal effects. The variation across observations is used to estimate the effects of calendar-time, and the parameter is estimated for each dummy variable. The first column of Table 4 includes this model. The results show that the last quarter has the lowest re-employment probability, whereas the first two quarters have the highest probabilities. In great measure this result is due to the big changes of the Finnish weather.

It may be of interest to use a continuous calendar-dependent variable, e.g. the unemployment rate, and estimate its effect. This is done in the second model of

Table 4. Unemployment rate varies with calendar time, but in statistics it is calculated as an average value for intervals of calendar time, i.e. months, quarters and years. Quarterly data are used in this model. The variation across observations is used to take into account the calendar-time effects. The unemployment rate has a negative and statistically significant effect on the re-employment probability, as expected.

Table 4. Semi-parametric models of unemployment duration using calendar-dependent covariates

	Std.errors in parentheses	
Number of children	-0.017	-0.017
	(0.049)	(0.049)
Married	0.090	0.139
	(0.070)	(0.069)
Sex	-0.052	-0.019
	(0.061)	(0.061)
Age	-0.037	-0.031
	(0.003)	(0.003)
Level of education	0.085	0.079
	(0.064)	(0.064)
Training for employment	0.148	0.214
	(0.079)	(0.079)
Member of UI fund	0.193	0.200
	(0.065)	(0.065)
Came from schooling	0.324	0.180
	(0.084)	(0.084)
Came from house work	-0.609	-0.623
	(0.137)	(0.137)
Regional demand	0.179	-0.001
	(0.252)	(0.253)
Occupational demand	0.609	0.275
	(0.620)	(0.619)
Taxable assets	0.530	0.345
	(1.164)	(1.094)
Replacement ratio	-1.272	-1.129
	(0.156)	(0.154)
Quarter 1	0.763	
	(0.095)	
Quarter 2	0.765	
	(0.089)	
Quarter 3	0.362	
	(0.093)	
Unemployment rate		-0.684
		(0.068)
Log likelihood	-8368.3	-8364.1
Number of observations	2077	2077

2.3. Changing the Roles of Duration and Calendar Time

In this section the roles of the duration and calendar time are changed in order to study the seasonal effects and estimate and test the duration-dependence. Turning back to the seasonal effects it is obvious that the specification of the calendar-dependent part may be difficult, because the macro environment is changing continuously. If this process is not fully observed, one way to overcome the problem is to restrict the seasonal effects to a particular form. Using the models with calendar-dependent covariates the functional form of the calendar-dependent part has to be specified completely, but it may be difficult in practice. Here it is not preferred to specify the form of seasonal effects, but instead to ignore the specification of calendar-dependence by using calendar time as the basic time variable.

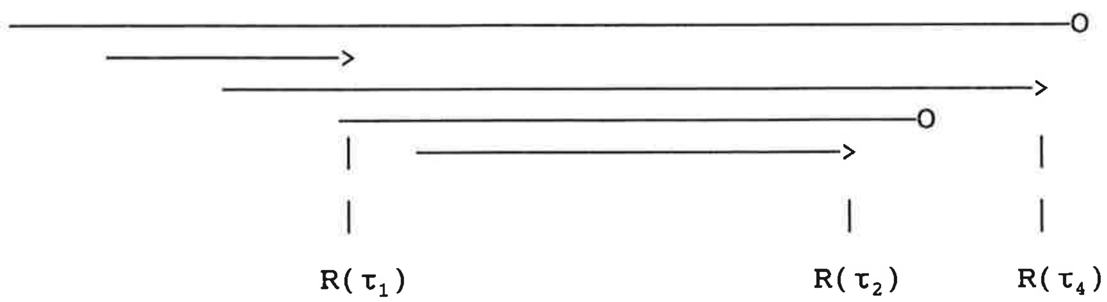
The idea of changing the roles of duration and calendar times is presented by Imbens (1990) and Ridder and Tunali (1990), even though they have not estimated that kind of models. Imbens suggests replacing the duration by the calendar time in order to eliminate common calendar-time related to macro shocks that affect all individuals in the same way. Ridder and Tunali study the child mortality and discuss the properties of these two observation plans.

After changing the roles of duration and calendar time the hazard function of the model can be written

$$(8) \qquad h(\tau,t,x) = h_0(\tau)h_1(t,x;\phi).$$

On any day of exit the risk set of a calendar time model consists of the individuals who are unemployed on that day. The estimation of the calendar time model is not possible using standard statistical packages. A scheme for estimating a calendar time model is following. Consider calendar time-dependent unemployment spells as shown in Figure 2. It is helpful to recode the dates of entry and exit as a difference between these dates and any date earlier than the first date of entry in the sample. The calculation of the risk set can then be performed easily starting from the latest date and moving towards the next latest date and so on. In the case of exit or censoring the observation is added to the risk set and in the case of entry it is subtracted from it. This procedure is continued until all the dates of exit have been proceeded in order to calculate the needed log likelihood function and the derivatives.

Figure 2. Unemployment durations of five individuals in a model based on calendar time (O = censored, > = complete)



The results of estimations of the calendar time models are in Table 5. Compared to the corresponding duration model in the first column of Table 1 and the models in Table 3, it can be seen that the parameter estimates of the calendar models do not differ substantially from the corresponding duration models. It can be concluded that the durations can be replaced by the calendar time in these cases. However, the loss of efficiency of the calendar model can be large, if there are a lot of durations that are not overlapping each other. Then some fraction of the persons does not contribute to the likelihood function.

The estimation of the calendar model is more time consuming than the model based on durations, since the calculation of the risk is based on comparisons of individual calendar times. However, the calendar model has an advantage. The baseline hazard can be expressed as a smoothly varying function of calendar time.

Table 5. Semi-parametric models of unemployment duration and labour mobility with calendar time

(A) Unemployment duration
(B) Regional mobility
(C) Occupational mobility

	(A)	(B)	(C)
	Std.errors in parentheses		
Number of children	-0.042		
	(0.049)		
Married	0.195		
	(0.070)		
Sex	0.018		
	(0.062)		
Age	-0.042	-0.050	-0.037
	(0.004)	(0.021)	(0.010)
Level of education	0.037		0.309
	(0.065)		(0.180)
Training for employment	0.214		0.381
	(0.079)		(0.193)
Member of UI fund	0.257	-1.234	0.425
	(0.064)	(0.466)	(0.156)
Came from schooling	0.237		-0.064
	(0.086)		(0.251)
Came from house work	-0.695		0.027
	(0.137)	-1.986	(0.252)
Regional demand	0.149	(1.463)	1.171
	(0.252)	2.561	(0.589)
Occupational demand	0.503	(3.214)	-3.733
	(0.629)		(1.644)
Taxable assets	0.756		-3.675
	(1.006)		(3.266)
Replacement ratio	-1.157	-4.841	-0.452
	(0.148)	(0.965)	(0.323)
Log likelihood	-7176.6	-272.8	-1167.9
Number of observations	2077	2077	2077

3. Baseline Hazard Functions

The estimated values of the parameters can be used to construct an estimator for the integrated baseline hazard, which has been proposed for Cox's model by Breslow (1972, 1974)

$$(9) \qquad I_0(t_i) = \sum_{t_j \le t_i} \frac{c_j}{\sum_{k \in R(t_j)} \exp(x_k \hat{\beta})}$$

The corresponding Breslow's estimate for the baseline hazard is based on the subdivisions of the time scale at those points where the event occurs

$$(10) \qquad \hat{h}_0(t_i) = \frac{c_i}{(t_i - t_{i-1}) \sum_{k \in R(t_i)} \exp(x_k \beta)} .$$

For the graphical presentation of the baseline hazard it is more natural to assume that $h_0(t)$ is a slowly varying function of t. In the calendar time model τ_i is substituted for t_i. If the estimates of the integrated baseline hazards are available, an estimate for the baseline hazard for each distinct duration can be rewritten as a difference quotient of the integrated baseline hazards

$$(11) \qquad \hat{h}_0(t_i) = \frac{\hat{I}_0(t_i) - \hat{I}_0(t_{i-v})}{t_i - t_{i-v}} ,$$

where $\nu = 1$. Then it can be seen that (10) is essentially equal to (11) for noncensored observations. It is suggested here that these smoother estimates are obtained by choosing ν for each t_i such that $\min(t_i - t_{i-\nu})$ is larger than a predefined constant ε. In this application ε was set to 5 weeks and the estimates of the baseline hazard function were centred at the midpoint of the intervals $(t_i, t_{i-\nu})$. An advantage of this kind of simple smoothing is that the baseline hazard can not get negative values, which is possible using the method suggested by Anderson and Senthilselvan (1980).

The estimates of the baseline hazard function of the duration and calendar time models are presented in Figure 3. Reluctant movers have a risk of losing benefits after the first three months. Figure 3 shows that the risk increases the re-employment probability. The elasticity of the hazard function with respect of the replacement ratio is the product of the replacement ratio and its parameter estimate. Therefore the effect of risk is larger for the members of labour unions, who are usually eligible for higher benefits.

Members of labour unions face a 20 % reduction in their benefits at the 20*th* week of unemployment. The reduction has a strong positive effect on the re-employment probability. Because of the reduction the hazard is just after it approximately 100 % higher than it would otherwise be indicating that in the case of decreasing benefits the elasticity of the hazard with respect to the replacement ratio is about -5.

The employment office has to offer a job to a person who has been unemployed for a year. Therefore the baseline hazard functions are increasing at the unemployment of a year. The low estimates of the baseline hazard function for the durations just less than a year are rather low. This is probably affected by the rules and practice of the employment office.

The calendar time model of unemployment duration is presented in Table 5 and the seasonal variation is illustrated by the baseline hazard function in Figure 3. The seasonal variation of the baseline hazard is rather similar during 1985 and 1986 except the baseline hazard is lower in 1986. Figure 3 shows that the re-employment probability is rather high in May and June, whereas it is low during the last quarter of the year.

The baseline hazard functions of the semi-parametric models of labour mobility are presented in Figure 4. The unemployed persons seem to be prone to move at the beginning of their unemployment spell and just after the three months of unemployment. There seems to be two peaks in moving. Unemployed persons move often in the beginning of June and August. However, one can not draw very strong conclusions about the regional mobility, since it is a rather rare phenomenon.

Occupational mobility is measured on the most accurate 5-digit level, which includes 1320 occupations. Unemployed persons change their occupations most often at the beginning of their unemployment spells and just after the three first months of unemployment. There is a peak in occupational mobility at the beginning of September. People change their occupations quite often also in the beginning of May, June and July, but rather seldom at the end of the year.

Figure 3. Baseline hazard functions of unemployment duration

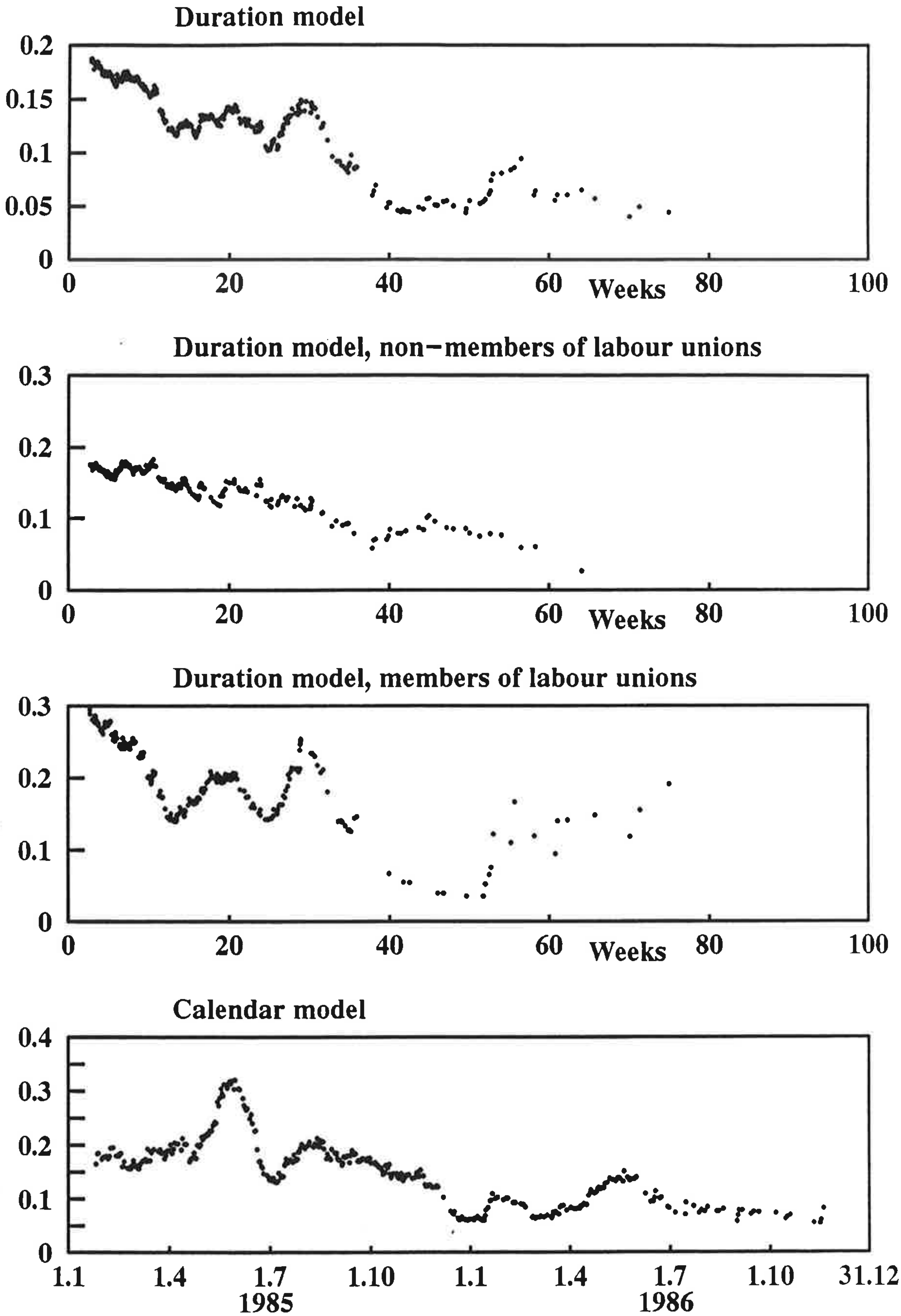
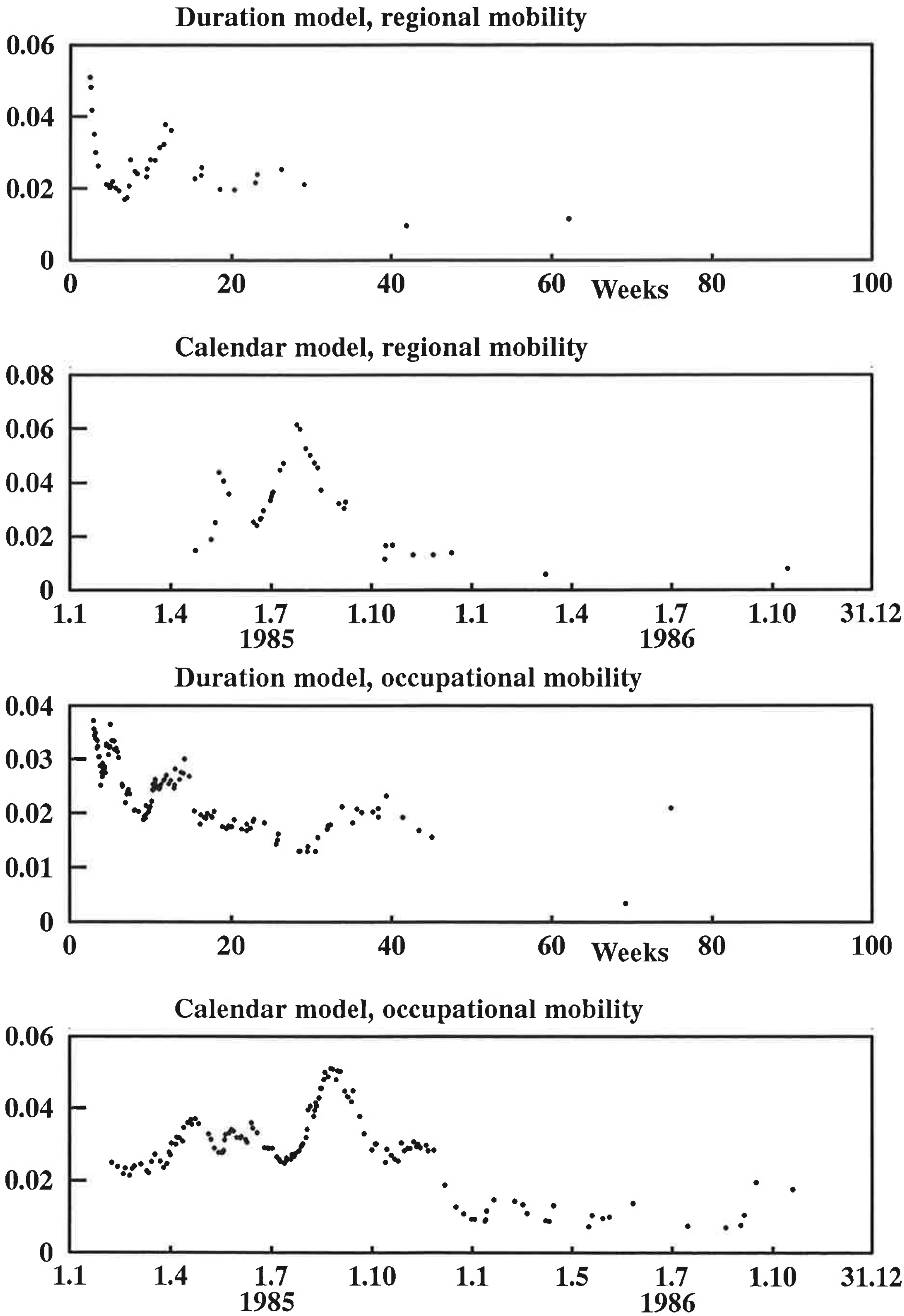


Figure 4. Baseline hazard functions of labour mobility



4. Conclusions

This paper studied the time-dependence in semi-parametric proportional hazards models of unemployment duration. Models with duration-dependent replacement ratios of unemployment benefits were estimated. The model with a duration-dependent replacement ratio gives a substantially lower parameter estimate of the replacement ratio than the model, where the benefit replacement ratio is fixed at an average value over the unemployment spell. Studying more carefully the duration-dependent effects, it was noted that the benefits have a negative effect on the re-employment probability during the unemployment of the first three months, but after that period the effect turns positive. One reason is that the unemployed persons may lose their benefits after the first three months if they do not move or change occupations. Another reason is that the reduction of benefits by 20 % doubles the re-employment probability just after the reduction. Using the graphs of the baseline hazard functions it was shown that these features of the UI system are of a great importance when looking at the probabilities of re-employment and regional and occupational mobility.

Specifications with calendar-dependent explanatory variables were studied and estimated using Finnish microeconomic data on unemployment spells. Statistically significant effects were found using calendar-dependent dummy variables and unemployment rates. The models were found useful in estimating the seasonal effects on the re-employment probability. However, using these kinds of models the functional form of the calendar-dependence has to be specified completely, which may be difficult in practice.

In order to allow a flexible form of the calendar-dependence the roles of duration and calendar time were changed following the suggestions by Imbens (1990) and Ridder and Tunali (1990). The baseline hazard functions of the calendar models were used to illustrate the seasonal variation of the re-employment and labour mobility. The baseline hazard of re-employment is rather high in May and June. Regional mobility has peaks in the beginning of June and August. Occupational mobility has a peak at the beginning of September, but the mobility is rather high also in the beginning of May, June and July. Labour mobility is rather low at the end of the year.

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