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Labour Market Transitions and Economic Incentives

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No 06	By Nico Keilman: Demographic Translation: From Period to Cohort Perspective and Back. 17 p.
No 05	By Pål Longva and Oddbjørn Raaum: Earnings Assimilation of Immigrants in Norway – A Reappraisal. 28 p.

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25 April, 2000

Labour Market Transitions and Economic Incentives

By Knut Røed and Tao Zhang*

Abstract

Building on register data describing monthly labour market status for the whole Norwegian population 1992-95, we estimate grouped competing risk hazard rate models for transitions between *employment*, *unemployment* and *non-participation*. The models impose no parametric restrictions on either calendar time- or spell duration effects. There is a substantial business cycle element in all transition rates apart from transitions from employment to non-participation. Conditioned on observed covariates, there is a strong negative duration dependence in transitions from unemployment to employment. Economic incentives play an important role in transitions from unemployment to employment, as well as in transitions from employment to non-participation. A 10 per cent increase in the unemployment benefit replacement ratio reduces the transition rate from unemployment to employment with 4-7 per cent, while it reduces the transition rate from unemployment to non-participation with 10-13 per cent. Access to a particularly generous voluntary early retirement scheme increases the hazard rate from employment to non-participation with 40-80 per cent. More human capital entails higher job-to-job transition rates and lower transition rates to unemployment and non-participation. It is also associated with higher transition rates from unemployment to employment.

Keywords: Labour market transitions, unemployment duration.

JEL Classification: C41, J64.

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1 Introduction

The purpose of this paper is to identify determinants of individual employment and unemployment spells. Building on Norwegian register data covering the whole adult population 1992-1995, we first analyse transitions out of regular jobs, either to other jobs, to unemployment or out of the labour force. Those that enter the unemployment pool are then subject to an additional transition rate analysis, either back to work or out of the labour force. We consider how all these transitions depend on human capital, economic incentives, business cycles, and spell durations.

The determinants of individual unemployment durations have received considerable attention in the literature (see e.g. Devine and Kiefer, 1991, or Pedersen and Westergård-Nielsen, 1998, for recent surveys). The present paper contributes to this literature in a number of ways. First, we extend the standard unemployment duration analysis to a combined incidence/duration analysis, also focusing on transitions from employment to unemployment (as well as transitions between jobs and transitions out of the labour force). Second, we apply data and econometric methods that allow us to disentangle calendar time effects (business- and seasonal cycles) from spell duration effects (duration dependence and unobserved heterogeneity), without arbitrary parametric assumptions. Third, we calculate the unemployment benefit replacement ratio as a time varying covariate, making it possible to evaluate its incentive effects with the aid of cross sectional- as well as longitudinal variation. And finally, we take advantage of a quasi-natural experiment to asses how economic incentives embedded in the retirement system affect the flows from employment to non-participation and to unemployment.

2 The Data

The database is obtained by merging a number of administrative registers, and in the present form it gives an account of the main labour market activity for the whole Norwegian adult population at the end of each month during the period 1992-1995. We start out with the complete stock of employed persons (according to the employer/employee register) in January 1992. Then, for each subsequent month until

December 1995, we add persons entering jobs and record job spells that are terminated. A terminated spell is either identified as a transition to another job, to unemployment or an exit out of the labour force. Those that in the course of the four-year observation period become unemployed are tracked until a new transition takes place, or until the end of the observation period. A terminated unemployment spell is either identified as a transition back to employment or an exit out of the labour force. Table 1 provides summary statistics about the transitions and some selected covariates[1].

Table 1
Descriptive Statistics of the Data

Origin state	Employment		Unemployment	
	Men	Women	Men	Women
Total number of spells	122830	110077	109752	87460
Number of transitions to:				
Employment	32526	29373	84576	62075
Unemployment	6823	5432	-	-
Out of the labour force	11947	13561	10608	13786
Average duration of completed spells (measured in months)	60.7	54.0	5.4	4.9
Average duration of completed spells according to destination:				
Employment	51.5	47.0	5.1	4.3
Unemployment	40.7	39.9	-	-
Out of the labour force	97.2	75.0	8.3	7.8
Averages and fractions taken over the total number of months at risk:				
Average age	38.9	38.6	38.2	37.1
Average years of work experience	15.2	11.4	13.9	9.9
Educational attainment (per cent):				
Less than 10 years	13.8	14.8	23.4	22.4
10 years	17.6	27.8	28.6	36.0
11 to 12 years	36.6	25.7	33.5	29.5
13 to 16 years	21.0	26.8	8.1	8.1
Over 16 years	8.0	2.8	1.0	0.6
Unknown	2.9	2.1	5.4	3.4
Average unemployment benefit ratio	-	-	51.9	54.7
Per cent of employed in relevant age group satisfying individual eligibility rules for early retirement (AFP)	98.2	75.4		
Per cent of eligible working in AFP firm	63.9	74.0	-	-

The covariates include standard *demographic variables*, such as gender, age, county of residence, family situation, and nationality; and *human capital variables* such as educational attainment and work-experience. In order to improve the characterisation

[1] At the present stage, only a 10 per cent sample of employment spells have been used in the econometric analysis, while all unemployment spells have been used).

of individuals' human capital, we take advantage of income records (based on pension point accumulation) for the years back to 1967. The basic idea is that individual ability, conditioned on education and work-experience, is partly revealed through the actual income path. We use the following procedure to proxy the level of human capital embedded in individual ability: We first divide the whole population into 120 relatively homogenous groups with respect to gender, educational attainment and work experience, and retrieve for each person the maximum yearly income earned after the education was completed[2]. We then compute a set of dummy variables indicating the decile in the *within-group* maximum earnings distribution to which each person belongs. These variables reflect earnings capacity *conditioned* on educational attainment and work experience. And even though wage formation in Norway to a large extent is governed by egalitarian considerations (Moene and Wallerstein, 1995), wages are likely to be positively correlated to individual skills. In order to avoid arbitrary functional form relationships, we use these dummies directly in the econometric models explaining labour market transitions. In addition, we apply the more standard (unconditional) measure of average yearly income during the whole work career[3].

Economic incentives facing unemployed persons are represented in the dataset by a time varying measure of the replacement ratio, i.e. the amount of unemployment benefits received, relative to expected income. The latter is basically assumed to be equal to previous income, adjusted for general wage growth (see Røed and Zhang, 1999a, for details). The basic rule in Norway is that unemployment benefits amount to 62.4 per cent of previous income. Nevertheless, our replacement ratio embodies cross sectional- as well as longitudinal variation. The cross section variation stems from variation in eligibility, from an upper ceiling on benefit entitlements, and from a floor on benefit entitlements for the elderly. The longitudinal variation stems from incomplete wage indexation (for ongoing unemployment spells), from changes in payments associated with participation in labour market programs, and from the reduction in benefits associated with long term unemployment.

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[2] We believe maximum income is a better proxy for individual ability than e.g. average income, as many persons have periods with part-time work or work only in parts of the year. Moreover, average income measures are affected by previous unemployment periods.

[3] As there is a ceiling on pension point generating income, approximately 5 per cent of the incomes are censored.

Finally, we take the opportunity to investigate possible incentive effects associated with early retirement schemes for employed workers. A sort of natural experiment facilitates the identification of these effects. The reason is that the fairly generous voluntary early retirement scheme (AFP), recently introduced in Norway, encompasses only a part of the labour force, and eligibility depends on firm-specific criteria (see Hernæs et al, 2000, for details). Approximately 60 per cent of the workers between 64 (65 before October 1993) and 66 years had access to this particular exit path in our estimation period (the age limit is now further reduced to 62 years). Based on our access to the pension point accumulation files, we have been able to calculate each person's AFP replacement ratio quite accurately. The replacement ratios ranges from around 40 per cent (for the highest incomes) to 70 per cent (for the lowest incomes)[4]

3 The Econometric Model

Let $j=e,u$ be the two origin states of employment and unemployment respectively, let T_j be the duration of that state, and let $k=e,u,o$ be the three alternative destination states of employment, unemployment and out of the labour force (non-participation). Furthermore, let $q_{ijk}(t,t)$ be the hazard rate at which individual i transits from state j to state k at spell duration t and calendar time t, i.e.

$$q_{ijk}(t,t) = \lim_{\Delta t \to 0} \frac{P(t \le T_j \le t + \Delta t, K = k \mid T_j \ge t, t, i, j)}{\Delta t}. \tag{1}$$

As we observe labour market status by the end of each month only, we set up the econometric model in terms of discrete (grouped) hazard rates (Prentice and Gloeckler, 1978; Meyer, 1990; Narendranathan and Stewart, 1993). Let $\bar{t}_i$ be the calendar time at which individual i entered the present state j. The grouped composite hazard, i.e. the probability of exiting state j during duration month t, given that no exit occurred before that, is given as:

$$h_{ij}(\bar{t}_i + t, t) = 1 - \exp\left(-\sum_k \int_{t-1}^{t} q_{ijk}(\bar{t}_i + u, u)\,du\right). \tag{2}$$

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[4] Since pensions are subject to a lower tax rate than work income, the after tax replacement ratios are 10-15 per cent higher.

The grouped destination specific hazard, i.e. the probability of exiting to a *particular* destination state k during duration month t, given that no exit occurred before that, is given as:

$$h_{ijk}\left(\bar{t}_i + t, t\right) = \int_{t-1}^{t} q_{ijk}(\bar{t}_i + u, u)\exp\left(-\sum_k \int_{t-1}^{u} q_{ijk}(\bar{t}_i + s, s)ds\right)du\ . \tag{3}$$

Time has two dimensions; calendar time t and spell duration t. Assume that the hazard rates can be factorised into three terms, depending on individual characteristics, calendar time and spell duration, respectively. Let x_{it} be a vector of time varying individual characteristics, let s_{jkt} measure calendar time effects, and let l_{jkt} measure spell duration effects. Imposing exponential link functions between individual characteristics and the hazard rates, we have (Røed and Zhang, 1999a):

$$\int_{t-1}^{t} q_{ijk}(\bar{t}_i + u, u)du = \exp(x_{it}'b_{jk} + s_{jkt} + l_{jkt}),\quad t_i = t - \bar{t}_i. \tag{4}$$

Given the proportionality assumption, the calendar time- and spell duration coefficients (s_{jkt}, l_{jkt}) are simply the logs of the respective integrated continues time baseline hazard components. The proportionality assumptions may easily be relaxed by including in the x-vector interaction terms between e.g. individual characteristics and spell duration.

Let E_{92} be the set of employed persons in January 1992 and let t_{92} be the corresponding job spell durations at this point. The likelihood function is then given as:

$$l_{jk} = \prod_{i=1}^{N_j}\left[(h_{ijk}(\bar{t}_i + t_i, t_i))^{d_i}\prod_{s=\bar{t}_i}^{t_i^*-1}(1 - h_{ij}(\bar{t}_i + s, s))\right], \tag{5}$$

$$i = 1,2,\ldots,N_j,\quad j = e,u,\quad k = e,u,o,\quad t = 1992.3,\ldots,1995.10,\quad t = 1,2,\ldots,T_j,$$

where $d_i=1$ if there is a transition to state k and zero otherwise, $\bar{t}_i = t_{i92}$ if $i \in E_{92}$ and $\bar{t}_i = 1$ otherwise, $t_i^* = t_i$ if $d_i=1$ and $t_i^* = t_i + 1$ if $d_i=0$.

A problem with the likelihood function in (5) is that it cannot be factorised into destination specific parts, and that the contribution of the integral in $h_{ijk}(t, t_i)$ (equation 3) makes maximisation cumbersome. This problem is usually disregarded (Narendranathan and Stewart, 1993; Carling et al, 1999; Jensen et al, 1999; Tysse and Waage, 1999), and each transition rate is estimated *as if* competing transitions can only occur at integer points on the border of (and not within) the time intervals (Kiefer, 1990). This may not be a bad approximation for monthly unemployment data if the hazards are relatively stable within each duration month. For example, in the case of constant within-month hazard rates, we may use the following approximation:

$$\begin{aligned} h_{ijk}(t,t_i) &= h_{ij}(t,t_i)\frac{\exp(x_{it}'b_{jk}+s_{jkt_i}+l_{jkt_i})}{\sum_k \exp(x_{it}'b_{jk}+s_{jkt_i}+l_{jkt_i})} \\ &\approx 1-\exp\left(-\exp(x_{it}'b_{jk}+s_{jkt_i}+l_{jkt_i})\right). \end{aligned} \tag{6}$$

Inserting the approximation in (6) into (5) yields the standard decomposable complementary log-log model, which is straightforward to estimate. Results reported by Røed and Zhang (1999b) indicate that the approximation in (6) yields reliable estimates even when the assumption of constant within-month hazards does not hold, as long as the hazard rates are relatively smooth (e.g. Weibull hazards).

An advantage with the present modelling framework is that it can be used to decompose changes in the transition rates into two factors, i) changes in the composition of the group at risk and ii) changes in the economic environment faced by the group as a whole. The composition effect may in turn be separated into its observed- $(\exp(x_{it}'b_{jk}))$ and its unobserved component $(\exp(l_{jkt}))$. The unobserved component may reflect selection mechanisms (those with the most "transition-enhancing" unobserved characteristics transit first, ceteris paribus) as well as duration dependence[5] (individual hazard rates depend on elapsed duration).

Note that employment spells are analysed *conditional* on the prevailing stock distribution in January 1992, while unemployment spells are analysed unconditionally. Since a stock distribution overstates the number of long spells and understates the number of short spells, we do not identify the underlying true duration distribution of job spells.

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[5] Note that the standard result that unobserved heterogeneity always contribute to negative observed duration dependence, does not hold within a competing risk framework (Røed, et al, 1999). The reason is that unobserved characteristics that are favourable with respect to one transition may be unfavourable with respect to a competing transition,

4 Estimation Results

The model was estimated separately for men and women for each of the five possible transitions. The results are presented in Tables 2 and 3 and in Figures 1-6[6]. In the following subsections, we discuss the results in relation to relevant economic theory and in relation to previous empirical findings. We focus on the roles played by business cycles (Section 4.1), and spell durations (Section 4.2). We then turn to individual factors such as economic incentives (Section 4.3), human capital (Section 4.4), demographics (Section 4.5), and participation in labour market programs (Section 4.6).

4.1 Business Cycles

The calendar time parameters $(\exp(s_{jkt}))$ are proxies for the pure business- and seasonal cycle effects. The estimates of these parameters taken together thus provide an opportunity to evaluate the pattern with which changes in the economic cycle affect the various labour market flows. The “raw” calendar time parameters display a strong seasonal pattern, and also a relatively strong pattern of spikes at certain customary dates (particularly in job-to-job transitions and transitions out of the labour force). For example, there are typically few transitions in December and many transitions in January. In order to highlight the business cycle pattern, Figures 1 and 2 present smoothed calendar time effects for the various transitions, together with the development of actual exit rates (the “calendar effects” obtained when no covariates are entered apart from calendar time dummies). Figure 1 displays the estimated calendar effects for transitions out of jobs. They indicate a steady increase in job-to-job transitions and a steady decline in employment-to-unemployment transitions after the cyclical trough in the winter 1992/93. The out-of-labour-force hazards, on the other hand, seem to be almost unaffected by the business cycle. There are only small differences between the estimated calendar time effect and the actual transition rates, indicating that the composition of the labour force did not change substantially with respect to transition probabilities.

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[6] Only selected estimates are reported. The models also contain a number of control variables, such as region and industry. The total number of estimated coefficients was 128 for the out-of-job-hazards and 119 for the out-of-unemployment hazards.

Table 2
Transitions out of employment. Estimated effects of selected covariates
(standard errors in parentheses)

Destination:	Another job		Unemployment		Out of labour force	
	Men	Women	Men	Women	Men	Women
Educational attainment with maximum 5 years work experience (ref.:11-12)						
9 years or less	-0.1321	-0.2863	0.3463	-0.0009	0.4050	0.2944
	(0.0327)	(0.0288)	(0.0679)	(0.0630)	(0.0802)	(0.0505)
10 years	-0.0623	-0.1894	0.5658	0.4551	0.7122	0.4485
	(0.0303)	(0.0275)	(0.0539)	(0.0502)	(0.0612)	(0.0420)
13-16 years	0.0510	0.1024	-0.5667	-0.7286	-1.1797	-0.5704
	(0.0234)	(0.0200)	(0.0648)	(0.0606)	(0.1112)	(0.0476)
17 years or more	0.0619	0.2718	-1.0999	-0.7092	-1.4572	-0.5312
	(0.0466)	(0.0540)	(0-1712)	(0.1808)	(0.2705)	(0.1354)
Educational attainment with at least 6 years work experience (ref.:11-12)						
9 years or less	0.0034	-0.1054	0.3158	0.3724	0.3296	0.3794
	(0.0250)	(0.0312)	(0.0448)	(0.0576)	(0.0278)	(0.0321)
10 years	-0.0758	-0.0758	0.2757	0.1900	0.1797	0.1244
	(0.0219)	(0.0238)	(0.0404)	(0.0490)	(0.0271)	(0.0289)
13-16 years	0.1510	0.2086	-0.3415	-0.7963	-0.5212	-0.3791
	(0.0194)	(0.0216)	(0.0499)	(0.0649)	(0.0374)	(0.0337)
17 years or more	0.2933	0.5276	-1.1259	-0.4796	-0.6676	-0.5644
	(0.0266)	(0.0415)	(0.1081)	(0.1501)	(0.0625)	(0.0930)
Work experience (# years)	-0.0073	-0.0125	-0.0164	-0.0239	0.0287	0.0069
	(0.0023)	(0.0022)	(0.0050)	(0.0047)	(0.0042)	(0.0030)
Work experience squared	0.0001	0.0002	-0.0016	-0.0016	-0.0022	-0.0013
	(0.0002)	(0.0002)	(0.0004)	(0.0004)	(0.0004)	(0.0002)
Less than one year work experience	0.1850	0.1399	-0.1715	-0.0240	-0.4682	-0.0650
	(0.0273)	(0.0252)	(0.0600)	(0.0547)	(0.0659)	(0.0342)
Average income in work career (censored at the highest incomes)	-0.0305	-0.0534	-0.1577	-0.1203	-0.1161	0.0054
	(0.0061)	(0.0081)	(0.0144)	(0.0200)	(0.0111)	(0.0120)
Decile in conditional maximum wage distribution (ref.: 4-7):						
1 (the lowest decile)	-0.1696	-0.0463	0.1653	0.0714	0.1059	0.1039
	(0.0298)	(0.0340)	(0.0526)	(0.0720)	(0.0376)	(0.0367)
10 (the highest decile)	0.1864	0.2786	0.0984	0.0917	-0.0100	-0.1142
	(0.0213)	(0.0259)	(0.0522)	(0.0625)	(0.0391)	(0.0388)
Access to early retirement (AFP)#:		0.4185	-0.7106	-0.6032	0.2368	0.6566
		(0.1767)	(0.4122)	(0.2428)	(0.0654)	(0.0438)
AFP replacement ratio#		-0.1072	0.4196	-0.7200	0.5292	-0.1922
		(0.3485)	(0.7683)	(0.4691)	(0.1261)	(0.0852)
Immigrant from Non-OECD country	-0.0408	-0.0486	0.2700	0.2669	0.5414	0.5490
	(0.0508)	(0.0697)	(0.0876)	(0.1273)	(0.0891)	(0.0912)
Married	-0.0677	-0.1947	-0.4324	-0.2237	-0.1216	-0.0429
	(0.0152)	(0.0162)	(0.0324)	(0.0368)	(0.0231)	(0.0212)
Small children	0.0274	-0.1985	0.1233	0.0061	0.0852	0.8056
	(0.0270)	(0.0248)	(0.0605)	(0.0531)	(0.0492)	(0.0277)

These estimates are obtained from separate estimations based on the whole population above 56 years of age.

Table 3
Transitions out of the unemployment pool. Estimated effects of selected covariates (standard errors in parentheses)

Destination:	Job		Out of labour force	
	Men	Women	Men	Women
Inflow intensity (smoothed) at time of entry	0.9336	0.9204	-0.1587	-0.2445
	(0.0595)	(0.0750)	(0.1399)	(0.1229)
Educational attainment with maximum 5 years work experience (ref.:11-12 years):				
9 years or less	-0.3500	-0.2600	-0.0792	0.0164
	(0.0226)	(0.0217)	(0.0450)	(0.0359)
10 years	-0.2652	-0.2596	-0.1528	-0.0123
	(0.0163)	(0.0170)	(0.0373)	(0.0293)
13-16 years	0.1711	0.2329	0.5857	-0.0989
	(0.0231)	(0.0215)	(0.0532)	(0.0520)
17 years or more	0.1585	0.1449	0.2545	0.0000
	(0.0755)	(0.0785)	(0.2104)	(0.1870)
Educational attainment with at least 6 years work experience (ref.:11-12 years):				
9 years or less	-0.1134	-0.1209	-0.1054	0.0247
	(0.0113)	(0.0144)	(0.0359)	(0.0321)
10 years	-0.1473	-0.1037	-0.0667	-0.0206
	(0.0099)	(0.0122)	(0.0326)	(0.0280)
13-16 years	-0.0405	0.0850	-0.1118	-0.0050
	(0.0154)	(0.0177)	(0.0502)	(0.0452)
17 years or more	0.0592	0.1380	-0.0616	-0.3090
	(0.0372)	(0.0517)	(0.1211)	(0.1811)
Work experience (# years)	0.0161	-0.0019	-0.0324	-0.0224
	(0.0014)	(0.0012)	(0.0036)	(0.0028)
Less than one year work experience	-0.2340	-0.1662	0.3459	0.0754
	(0.0146)	(0.0140)	(0.0381)	(0.0297)
Average income in work career (censored at the highest incomes)	-0.0124	-0.0159	-0.1270	-0.1417
	(0.0039)	(0.0058)	(0.0111)	(0.0128)
Decile in conditional maximum wage distribution (ref.: 4-7):				
1 (the lowest decile)	-0.3116	-0.2419	0.1571	0.1167
	(0.0148)	(0.0189)	(0.0380)	(0.0368)
10 (the highest decile)	0.1210	0.1688	0.0122	-0.2794
	(0.0146)	(0.0166)	(0.0494)	(0.0495)
Replacement ratio	-0.7491	-0.4017	-1.3971	-1.0855
	(0.0151)	(0.0167)	(0.0404)	(0.0346)
Participation in labour market programs:				
Current course	-0.6759	-0.8749	-0.6065	-0.8577
	(0.0247)	(0.0317)	(0.0631)	(0.0547)
Current employment program	-0.1523	-0.4916	0.0303	-0.5082
	(0.0221)	(0.0294)	(0.0582)	(0.0552)
Previous course	0.1312	0.2162	-0.1168	-0.1623
	(0.0177)	(0.0230)	(0.0419)	(0.0362)
Previous employment program	0.2391	0.4569	-0.4051	-0.4366
	(0.0200)	(0.0266)	(0.0482)	(0.0432)
Immigrant from Non-OECD country	-0.2617	-0.3609	-0.3608	-0.1991
	(0.0269)	(0.0439)	(0.0547)	(0.0715)
Married	0.2012	0.0346	-0.0525	0.0874
	(0.0090)	(0.0097)	(0.0285)	(0.0212)
Small children	-0.0361	-0.2435	0.0456	0.4814
	(0.0168)	(0.0147)	(0.0523)	(0.0243)

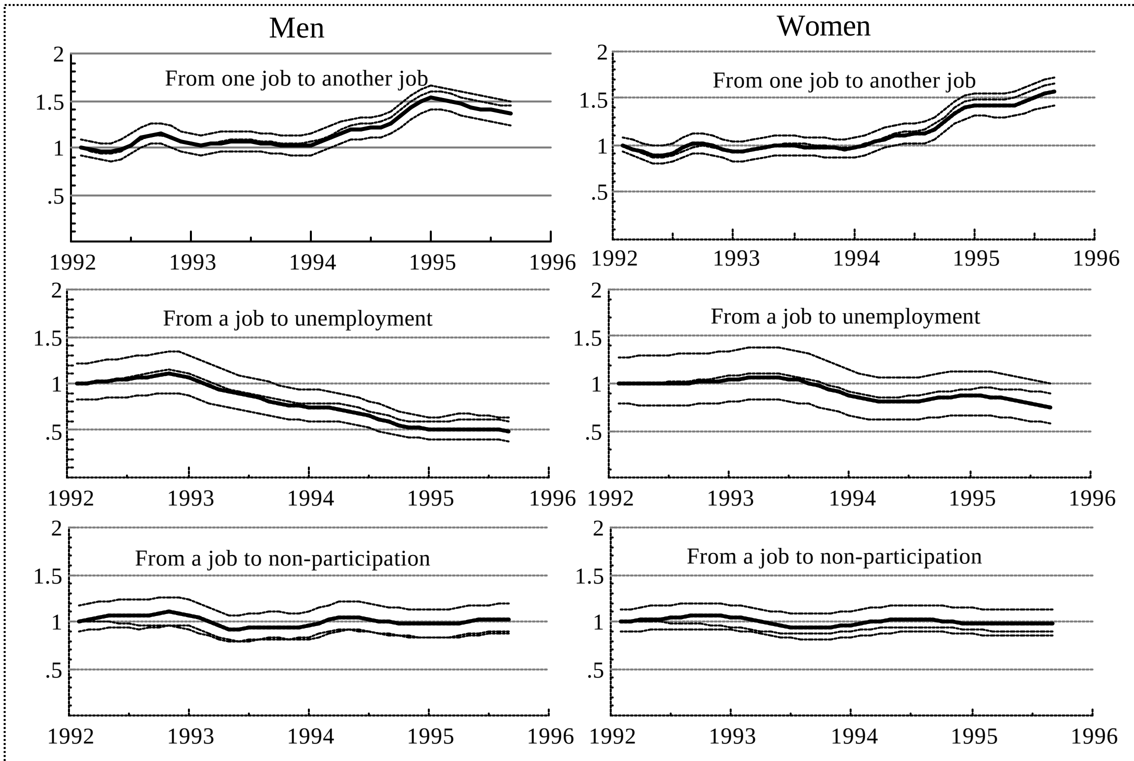

Figure 1. The out-of-job calendar time baselines. Smoothed estimates for calendar time effects (thick solid lines), together with 95 per cent confidence intervals (dashed lines) and actual transition rates (thin solid lines).

Note: The estimates are smoothed with the X-11 additive method (see Bureau of the Census, 1999, 70-84), and normalised on the first month (February 1992).

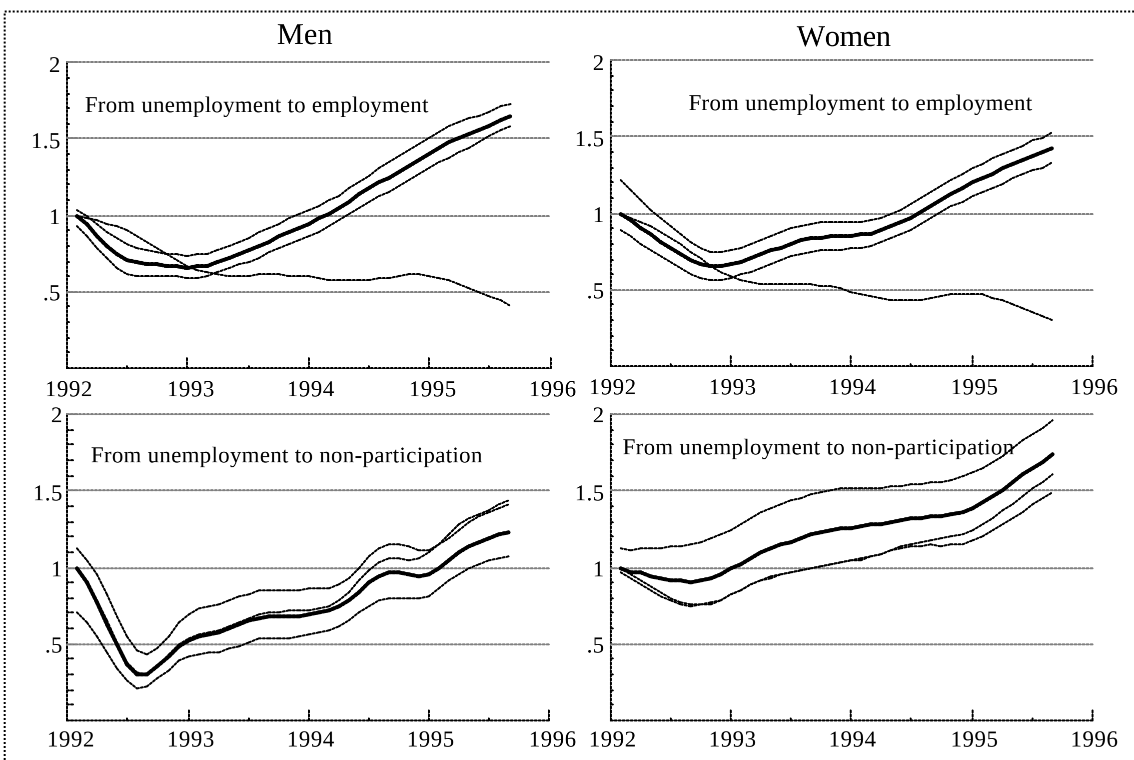

Figure 2. The out-of-unemployment calendar time baselines. Smoothed estimates for calendar time effects (thick solid lines), together with 95 per cent conficence intervals (dashed lines) and actual transition rates (thin solid lines).

Note: The estimates are smoothed with the X-11 additive method (see Bureau of the Census, 1999, 70-84), and normalised on the first month (February 1992).

Figure 2 displays the estimated calendar effects for transitions out of unemployment. They reveal a steady increase in transitions from unemployment to employment. More surprisingly, they also reveal a parallel increase in transitions out of the labour force. For the unemployment-employment transitions, the estimated calendar effects differ sharply from actual transition rates. This is a composition effect related to the flow structure of unemployment spells. In the beginning of the estimation period, all the unemployed in our dataset are newly unemployed. As time passes, the fraction of long term unemployment increases, and this produces a decline in actual transition rates that are unrelated to the business cycle pattern captured in the estimated calendar time effects.

There is a weak tendency for employment-unemployment transitions to lead unemployment-employment transitions. According to the smoothed series for men in Figures 1 and 2, the macroeconomic turning point occurred two months earlier for the employment-unemployment transitions (November 1992) than for the unemployment-employment transitions (January 1993). According to the raw point estimates, the employment-unemployment calendar effects are significantly negatively correlated (at the 5 per cent level) with the seven subsequent leads of the unemployment-employment calendar effects, but only with the three closest lags.

In order to investigate the extent to which the composition of the flow from employment to unemployment changes over the business cycle, we included the estimated (smoothed) calendar time effects from the male employment-unemployment transition as a regressor in the unemployment-employment hazard rate[7]. The result, reported in Table 3, indicates that persons becoming unemployed in a slump on average have much better job prospects than persons becoming unemployed in a boom. This is likely to reflect that the latter group is more selected according to individual characteristics than the former group. The estimated elasticity is close to one for both men and women, indicating that e.g. a 10 per cent higher inflow rate at the time of entry is associated with almost a 10 percent higher transition rate back to employment (ceteris paribus).

[7] We used the male calendar effect as the proxy for business cycle conditions at the time of inflow for both men and women, as its explanatory power outperformed that of the female calendar effect. This is likely to reflect that male jobs are more business cycle sensitive than female jobs.

4.2 Spell Durations

The spell duration parameters $(\exp(l_{jkt}))$ measures the extent to which the grouped hazard rates, conditioned on all the observed covariates, depend on elapsed spell duration. Figures 3 and 4 plot the estimated duration baselines for the various transition rates. The duration baselines for job spells are depicted in Figure 3. Job-to-job transitions occur most frequently in the beginning of job spells, but there are substantial spikes after exactly one and (to some extent) two years (reflecting customary contract periods). After that, there is a steady decline in the probability of switching to another job. There is a similar pattern in transitions to unemployment, i.e. a small increase during the first year, and a decline thereafter. Transitions out of the labour force are typically less likely, the longer the job has lasted.

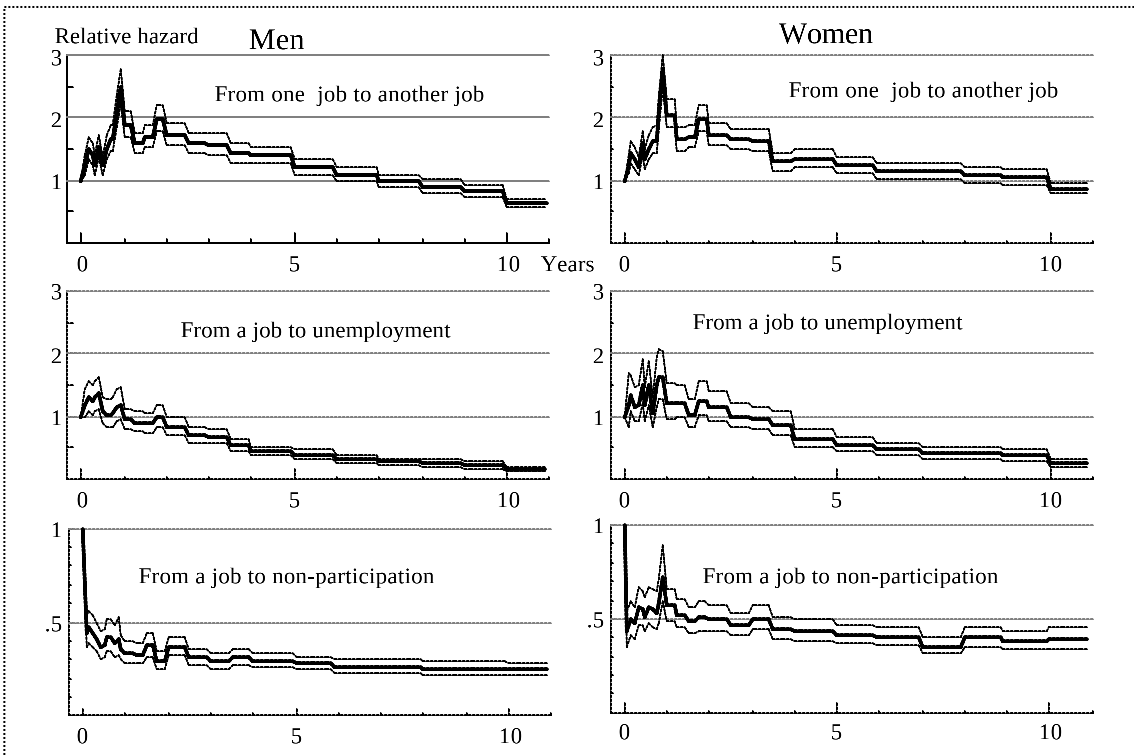

Figure 3. The out-of-job duration baselines. Point estimates (solid lines) together with 95 per cent point-wise confidence intervals (dashed lines).
Note: The hazard rates are normalised on the first duration month.

Figure 4 displays the duration baselines for unemployment spells. There is a strong, and very precisely determined, decline in the hazard rate from unemployment to employment as the spell is prolonged. This pattern may reflect negative duration dependence at the individual level, as well as selection according to unobserved characteristics. We cannot disentangle these two mechanisms without further assumptions about either the duration baseline or the distribution of unobserved heterogeneity, and this is not a topic of the present paper. The hazard rates to non-participation display a less precisely determined pattern of positive duration dependence.

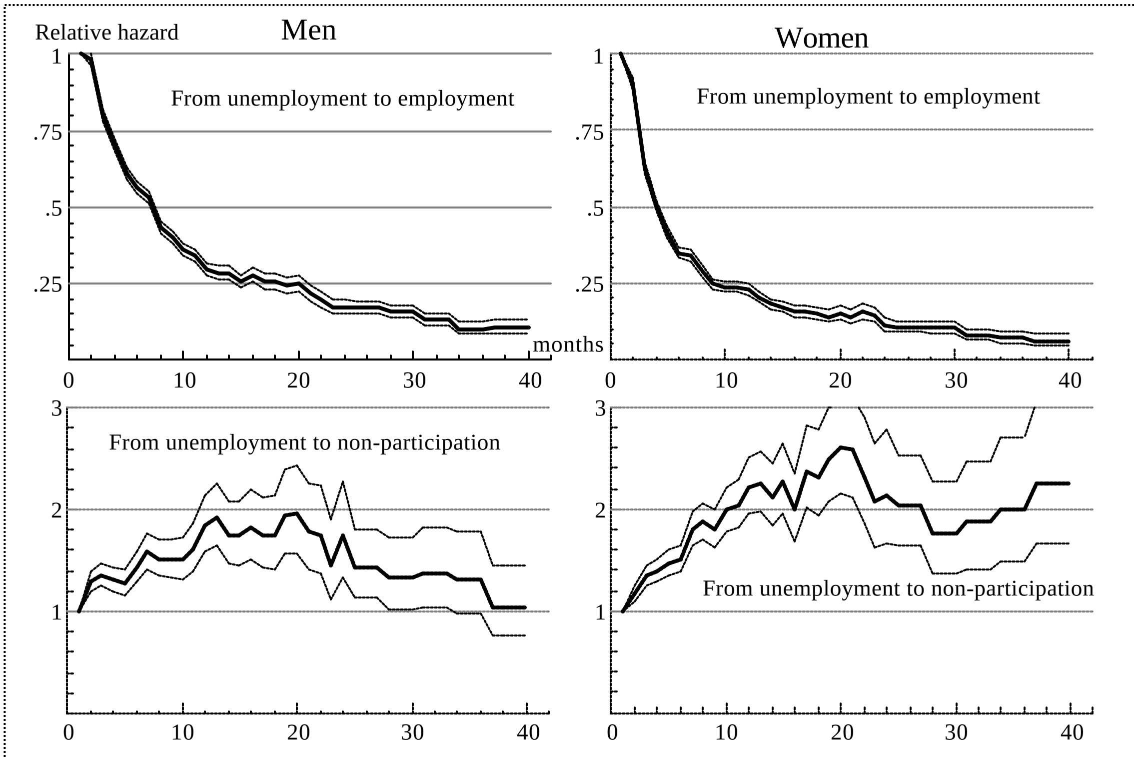

Figure 4. The out-of-unemployment duration baselines. Point estimates (solid lines) together with 95 per cent point-wise confidence intervals (dashed lines).
Note: The hazard rates are normalised on the first duration month.

4.3 Economic Incentives

The unemployment duration literature is strongly preoccupied with the incentive effects associated with unemployment benefits. Most of the results are obtained with the aid of cross sectional variation in benefit entitlements. This implies that the estimates may be biased due to unobserved characteristics that are correlated with benefit entitlements. Our approach to this problem is to use a rich set of control variables (that almost fully explain benefit entitlements), and to take advantage of longitudinal, as well as cross sectional variation. The results indicate that e.g. an increase in the replacement ratio with 10 percentage points reduces the transition rate from unemployment to employment with approximately 7 per cent for men and 4 per cent for women. Evaluated at the mean level of compensation, this is comparable to elasticities (with respect to the benefit level) ranging from –0.35 for men to –0.20 for women. These results are roughly in line with previous findings for the United Kingdom (Narendranathan et al, 1985; Narendranathan and Stewart, 1993; Arulampalam and Stewart, 1995). They indicate weaker responses than recently found in Sweden, based on quasi experimental data (Carling et al, 1999), but stronger responses than typically found in continental Europe (Hujer and Schneider, 1989; Groot, 1990; van den Berg, 1990; Steiner, 1990). However, there are even stronger effects in the transition rate from unemployment to non-participation, suggesting that pure unemployment duration analyses may overestimate the incentive effect regarding transitions from unemployment to employment. The point estimates indicate that an increase in the replacement ratio with 10 percentage points reduces the out-of-labour-force transition rate with 13 per cent for men and 10 per cent for women. Hence, the absolute value of the benefit elasticities are roughly twice as large for out-of-labour-force transitions as for employment transitions.

For many older workers, early retirement is an alternative exit route that may be affected by economic incentives. In 1989, a new voluntary early retirement scheme was introduced (AFP) with relatively generous after tax replacement ratios for workers that qualified. The aim was not to reduce the retirement age, but rather to give "worn-out" workers a more honourable exit from the labour market than the alternatives of disability pension or long term unemployment. Hence, the hope was that the overall early retirement hazard rate should not increase very much and that the unemployment probability should decline. As the eligibility rules were not only related to the workers themselves[8], but also to the affiliation of their employers, the reform constitutes a rather unique natural experiment. The results reported in Table 2 indicate that workers with access to AFP do have a significantly higher exit rate from employment to non-participation. For men, there is also evidence suggesting that the effect is stronger, the higher is the AFP replacement ratio. For example, at a 40 per cent replacement ratio, the point estimates indicate that exit rate to non-participation is 55 per cent higher than for a similar worker without access to AFP. At a 70 per cent replacement ratio, the difference increases to almost 85 per cent. For women, we also find that access to AFP increases transition rate to non-participation significantly.

[8] Individual eligibility rules require a relative firm attachment to the labour force over some time. Since qualification according to these rules may be correlated to retirement behaviour we also included a set of dummies indicating individual eligibility.

However, the replacement ratio effect is barely significant and wrongly signed. Access to AFP also reduces the transition rate to unemployment, although this effect is less precisely determined.

4.4 Human Capital

Human capital is proxied by educational attainment, work experience and previous absolute and education/experience-specific relative income (see Section 2). More human capital is associated with higher job-to job transition rates and higher transition rates from unemployment to employment, but lower transition rates to unemployment and out of the labour force. Higher unemployment incidence for persons with less human capital may reflect that low-skilled workers often perform tasks that involve less attachment to a particular firm. But the fact that they also have longer unemployment durations suggests that low-skilled workers are less active job seekers and/or less demanded by employers in general. The former may reflect that low-skilled workers have weaker work incentives, as their take-home market wages may be relatively small compared to the value of leisure. The latter may reflect that differences in wages do not fully reflect differences in productivity. As Norwegian wages are strongly related to education and work experience (particularly in the public sector), the failure of relative wages to reflect productivity differences is likely to be particularly noticeable in the case of “within-group” productivity differences. This is probably also the reason why a person’s position in the education- and experience-specific maximum income distribution seems to be important for labour market transitions.

4.5 Demographics

The age effects were estimated on the basis of a non-parametric approach (for five-year intervals up to the age of 60, and one-year intervals thereafter). The results are displayed in Figures 5 and 6. Unsurprisingly, the job-switching probability depicted in Figure 5 declines monotonically with age. The probability of transiting from employment to unemployment is relatively stable, with a weak upwards trend until the age of 60, where it rises sharply, particularly for men. The latter is most likely related to a special rule in the benefit entitlement system, offering workers above 61 years permanent benefits (until retirement age at 67). Also transitions out of the labour force rise substantially as the workers approach the retirement age. The transition rates from unemployment to employment (Figure 6) decline sharply and monotonically as a function of age, for both men and women. The transition rates from unemployment to non-participation are relatively high for young people (as they take up more education) and, for men, they are also relatively high for the middle aged. For both men and women, there is a sharp decline after the age of 60, again reflecting the strong economic incentives older persons have to register as unemployed.

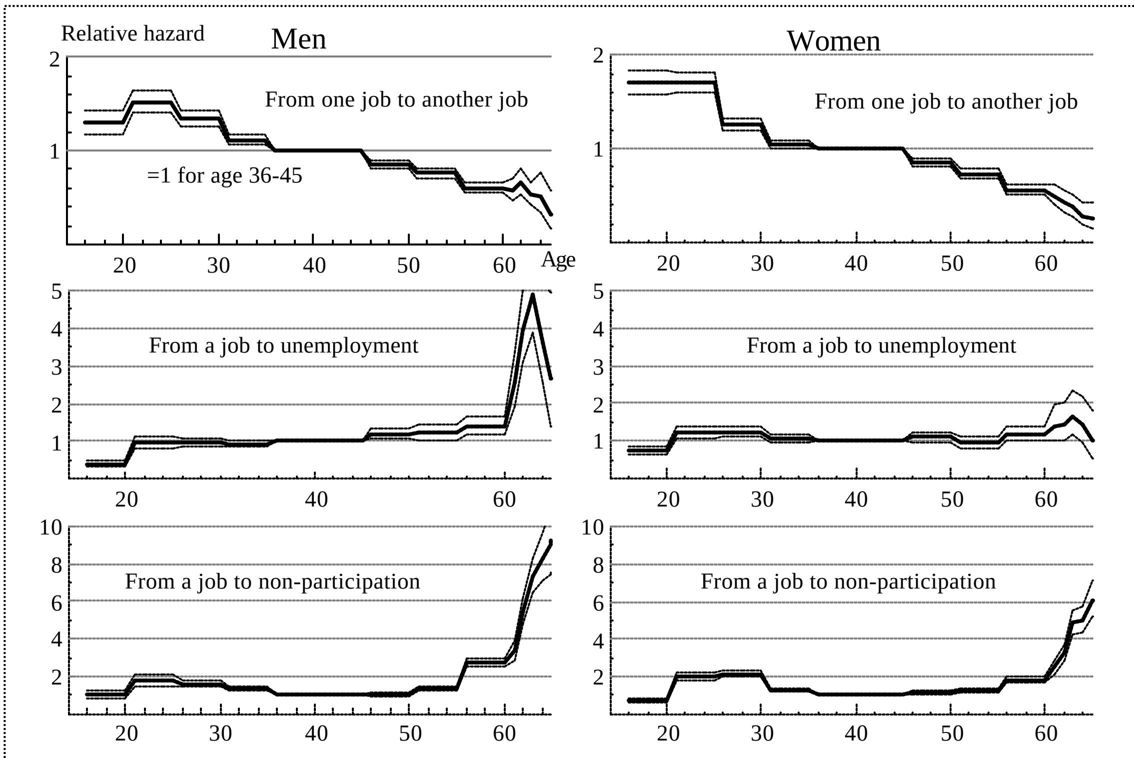

Figure 5. The estimated effects of age on transitions out of jobs. Point estimates (solid lines) together with 95 per cent point-wise confidence intervals (dashed lines).

Marital status and responsibility for small children affect men and women differently. For example, while a small child (below 5 years) is associated with an increase in the transition rate from employment to non-participation of 100 per cent for women, it leaves the male transition rate almost unaffected. Child responsibility also raises the transition rate from employment to unemployment (for men with 10-15 per cent), as well as reduce the subsequent transition rate back to work (for women with as much as 22 per cent). The higher degree of unemployment exposure for mothers may reflect that the responsibility for small children entail a strong preference for leisure (and hence reflect a moral hazard problem with respect to the benefit system), but it may also reflect that employers prefer employees without responsibility for small children.

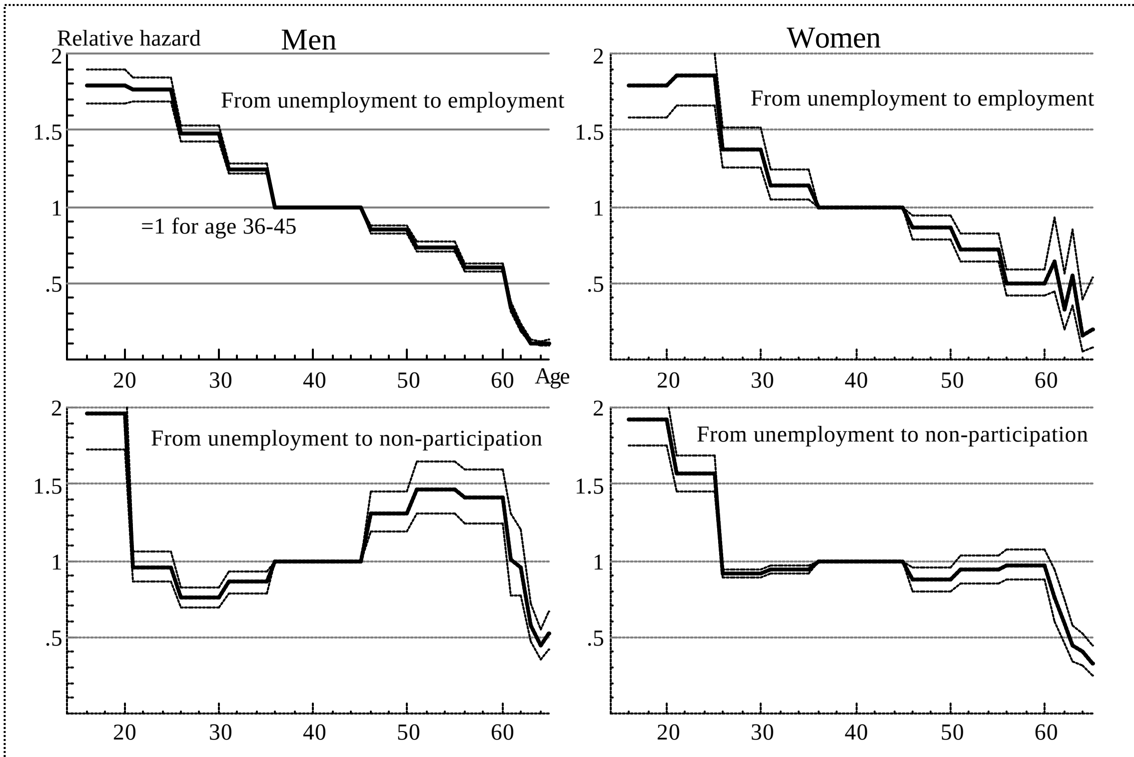

Figure 6. The estimated effects of age on transitions out of unemployment. Point estimates (solid lines) together with 95 per cent point-wise confidence intervals (dashed lines).

Immigrants from non-OECD countries are known to be much more unemployed than observationally equal natives. The estimates in Tables 2 and 3 indicate that this is attributed to both incidence and duration (in roughly equal portions). For example, according to the point estimates, male immigrants have roughly a 30 per cent higher transition rate from employment to unemployment, and a 23 per cent lower transition rate from unemployment to employment (ceteris paribus).

4.6 Participation in Labour Market Programs

As there are a number of selection processes going on in the allocation of labour market programs, the estimates of the program participation coefficients given in Table 3 cannot be interpreted directly as treatment effects. The results convey that, although most program participants are obliged to seek work actively, the predicted transition rates from unemployment to both employment and non-participation fall significantly (with roughly 50 per cent) during the course of the programs. This is particularly the cases for courses, which typically have a pre-determined duration. Previous program participation, however, seems to affect the two hazard rates in opposite directions. The have-completed-a-program-effect significantly positive for transitions to employment and significantly negative for transitions to non-participation. Previous participation in employment programs raises the predicted hazard rate to employment with 25-50 per cent, while it reduces the predicted hazard rate to non-participation with roughly 40 per cent.

5 Concluding Remarks

We have taken advantage of Norwegian register data 1992-1995 to estimate various competing risk transition rates in the labour market, without parametric restrictions on either calendar time- or duration effects. Our main results may be summarised as follows.

1. There is a substantial business cycle element in job-to-job transitions and in transitions between employment and unemployment (both ways). There is no visible business cycle pattern in transitions from employment to non-participation.

2. Conditioned on observed covariates, there is a strong negative and monotonic duration dependence in transitions from unemployment to employment, and a weak and non-monotonic positive duration dependence in transitions from unemployment to non-participation.

3. A 10 per cent increase in the unemployment benefit replacement ratio reduces the transition rate from unemployment to employment with 4-7 per cent, while it reduces the transition rate from unemployment to non-participation with 10-13 per cent.

4. Access to a particularly generous voluntary early retirement scheme increases the hazard rate from employment to non-participation with 40-80 per cent.

5. More human capital entails higher job-to-job transition rates and lower transition rates to unemployment and non-participation. It is also associated with higher transition rates from unemployment to employment.

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