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Unemployment Benefits and Incentives in Hungary: New Evidence

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Abstract:

This paper analyses the Hungarian unemployment insurance (UI) benefit reform of 1993, which led to the creation of a less generous benefit system. We investigate a sample of unemployment spells which are drawn from the Hungarian UI register directly before and after the reform came into force. The focus of the microeconometric duration analysis lies in the interaction of the exit behaviour of unemployed people with the UI system. As exit states we consider employment, subsidised employment, training and (early) retirement. Our results are in support of the hypothesis that the benefit reform increased, significantly, the transition rates from unemployment to all the labour market schemes. In particular, aged people responded to the benefit reform with a substantial rise in their transition rates to (early) retirement. A specific characteristic of our inflow cohort has to be taken into account for the analysis of the employment hazards. A large proportion of our sample represents workers who are likely to be on recall. Their job hazards are extremely sensible to seasonal labour demand fluctuations. We identify these workers and concentrate the analysis of the benefit reform on those workers who are not on recall. They represent workers who most likely were made redundant due to the restructuring of the Hungarian economy. The benefit reforms only slightly increased their speed of return to work. Furthermore, the transition rates to employment are found inelastic to the replacement rate with the exception of women who are below 30 years of age.

[1] I am extremely grateful to Prof. Gyula Nagy. This investigation would not have been possible without his support. He gained the access to the Hungarian unemployment register data and did all the work to set up this data for a duration analysis. I also would like to thank Prof. Janos Köllö and Gyula Nagy for providing a second data set on accepted wage offers. I am very grateful to my supervisor Prof. John Micklewright for his very helpful comments. All remaining errors are mine.
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Non-technical summary

The Hungarian labour market, from the start of transition, is characterised by outflow rates from unemployment that are low by western standards. For this reason long-term unemployment soon became a considerable problem. This paper evaluates whether one of the prime reasons for long unemployment spells in Hungary is the generosity of its unemployment insurance (UI) benefit system. We apply standard microeconometric duration methods in this analysis and rely on a 'quasi-natural experiment'. In Hungary such a situation was created by the UI benefit reform at the start of January 1993. The reason is that these new benefit rules did not apply to unemployed people who claimed their benefit before that date. This reform led to a considerably less generous UI system, since the most important change was a reduction of the length of all ten UI entitlement periods by one third. We compare the transition rates to jobs and to several labour market programs of unemployed people who made their UI claim in December 1992 with those of unemployed people who made this claim in January 1993. The main conclusion form this research is that the new and less generous UI system did nearly not increase the speed at which unemployed people return to work. A likely reason is that given the depressed labour market, unemployed people may not have much scope to increase the number of job offers by more intensive job search. However, we found that introducing a less generous benefit system directed unemployed people more rapidly into all kinds of labour market programs that are available.

1 Introduction

A very low turnover of the unemployment pool distinguishes many transition economies from the large western OECD countries [T. Boeri, 1994, 1996]. In Hungary, from the start of transition, the gross outflow rates from unemployment to jobs were low by western standards. The consequence is high long-term unemployment. There are a number of ways for economic policy to deal with this problem. This microeconometric study focuses on the Hungarian unemployment benefit system and the incentives it provides to individuals to leave unemployment. In particular, it addresses the question whether, during the transition process of the Hungarian economy, turning to a less generous benefit system is an appropriate policy to increase the outflow from unemployment.

An early discussion of the Hungarian unemployment insurance (UI) benefit system and incentives to leave (insured) unemployment has been carried out by J. Micklewright and Gy. Nagy in 1995. Their research focused on differences in the speed of exit from unemployment of two groups of people. One claimant group received UI according to the benefit rules of the year 1992, whereas the other group was entitled to UI according to the benefit rules of 1993, which are much less generous. Their Kaplan-Meier estimates of the transition rates to employment revealed meagre evidence for the hypothesis that less generosity of the UI system led to faster exit to employment. However, their analysis revealed that the introduction of the new benefit system in 1993 somewhat accelerated the transition rates to all other labour market states taken together. Our microeconometric analysis works with the same data set and extends this analysis.

One of the biggest problems that J. Micklewright and Gy. Nagy faced in their analysis, was a presumable presence of a large share of unemployed workers on recall, who return to work very quickly, in the sample. They considered this hypothesis as an explanation for the very large spikes that characterised their estimated employment hazards after an elapsed duration of unemployment of about three months. However, the share of the workers on recall among the 1992 claimant group and the 1993 claimant group was unknown. Therefore, the differences in the estimated employment hazards between the two claimant groups could just stem from a difference in these shares, but not from the distinct benefit rules. The following analysis attempts to identify individuals who are likely to be workers on recall. This allows us to distinguish between the benefit reforms impact on workers who lost their job for reasons of economic restructuring and on workers who lost their job only temporarily.

There is another issue that the analysis of Micklewright and Nagy did not address: In which way did the reform of the benefit system alter the exit behaviour to distinct active and passive labour market schemes. Thus, this research does not investigate composite transition rates to any exit state other than employment, but the transition rates to (early) retirement, subsidised employment and training. There is one reason why we expect that the benefit reform particularly altered the transition rates to these states. Typically, the demand side for labour in transition economies like Hungary is characterised by a low number of job offers relative to the number of job seekers, and vacancies are frequently taken by job-changers. So, there may not be much scope for an unemployed job searcher to raise the number of acceptable job offers by raising the search effort or lowering the reservation wage. Then, a policy that tightens the belt of UI recipients may raise, considerably, transitions to such labour market schemes rather than to jobs.

The paper is organised as follows: Section two introduces the Hungarian UI benefit system. It highlights its features before and after the benefit reform of January 1993. Additionally, a brief overview of the labour market programs which operate in the Hungarian labour market is given. In Section three we describe the microdata for our analysis. Section four introduces the econometric duration methods that we apply to our data. We start Section five by discussing the results of J. Micklewright and Gy. Nagy, and then extend their non-parametric work. Section six first sets up, briefly, the implications of search theory for our analysis. In this section, we also discuss problems with the specification of replacement rates, and of identification of the parameters of the benefit system in the econometric analysis that follows. The section finishes by presenting the results of Maximum Likelihood estimation of the piece-wise constant exponential model for transitions to employment, subsidised employment, and training. In Section seven we conclude.

2 The UI Benefit System and Labour Market Programs in Hungary

2.1 The Hungarian UI Benefit System and the 1993 Benefit Reform

In 1989, Hungary introduced a system of unemployment compensation. Two years later, in February 1991, an insurance-type unemployment system was established. The Solidarity Fund is the institution that is responsible for UI benefits. Its funds are provided by contributions from employers and employees. Its deficits are covered by the state budget. Reforms of this UI system made it more restrictive in the years 1992 and 1993. This occurred for reasons of high deficits of the Solidarity Fund[1]. Note that UI is „grandfathered"; i.e. people who receive UI are administered under the rules that were in place at the time of their benefit claim. Thus, new rules apply only to new claimants.

Let us now turn to the main features of the UI system during the period under review[2]. Its fundamental characteristics are the eligibility rules to UI benefits, the length of entitlement to UI benefits, and the level of benefits. Table 1 compares the benefit system of 1992 with that of 1993. There is no difference between the two regimes in their eligibility criteria. Unemployed people must have worked in contributory employment for at least 12 months during a four year period prior to the benefit claim[3]. Additionally, UI recipients are supposed to search actively for a job, to accept suitable jobs offered by the labour centre, and to co-operate with the labour centre. The employment record during the four years prior to the benefit claim determines the length of UI entitlement. In all, there are ten different entitlement periods. Under the 1992 benefit provisions they ranged from 135 days to 540 days. The 1993 benefit scheme reduced all ten entitlement periods by one third.

[1] The deficit was as high as 35 billion HUF in 1992; i.e. more than 50 percent of the Solidarity Fund's expenditure [E. Viszt and J. Ványai, 1994].
[2] Micklewright, J. and Gy. Nagy (1994a) provide a detailed discussion of the Hungarian UI system and its reforms.
[3] School-leavers were excluded from this requirement; in 1992 they were automatically eligible to six months of UI benefit.

The entitlement period is divided into two parts. A higher formal replacement rate applies to the first of these two parts. While this more generous first part accounted for two thirds of the total entitlement period under the 1992 benefit system, it accounts only for one quarter under the new benefit rules. The benefit levels are determined by an earnings-related benefit formula, where past earnings are defined as the average monthly earnings during the last four quarters before the benefit claim. Table 1 displays how the formal replacement rates vary over two phases of UI entitlement and over the two benefit schemes. The reform of 1993 raised the replacement rate of the first phase from 70 to 75 percent, and that of the second phase from 50 to 65 percent.

The benefit formula is not entirely earnings-related. Benefits are also subject to an upper and a flexible lower cap. The maximum benefit is set to twice the minimum wage under the 1992 benefit provisions. The same is true for the first phase of the entitlement period under the 1993 benefit regime though in the second phase the upper cap was set to only 1.7 times the minimum wage. Next, the minimum wage (9000 HUF in February 1993) was paid as a minimum benefit to people who made their claim before the end of 1992. After the reform, it was set to a somewhat lower level. However, this lower cap does not apply to UI recipients whose past earning are lower than the minimum benefit. For them the benefit equals their past earnings. Both the 1992 and 1993 benefit rules do not index benefits to inflation. Taken together, the modifications result in a much less generous UI system since January 1993.

People who exhaust their UI benefits, or who are recurrently unemployed and no longer eligible to UI, may apply for social benefits (SB). A means test requires an applicant's per capita household income not to exceed 80 percent of the minimum old-age pension, which is about two thirds of the minimum wage [J. Micklewright and Gy. Nagy, 1995b, p. 6]. The monthly social benefit is set at 80 percent of the minimum pension. So, even if somebody passes the means-test, social benefits should, in most cases, be much lower than UI benefits. J. Micklewright and Gy. Nagy (1996a) examined this difference for people who entered the UI register in early 1993. They found that the level of UI at the time of exhaustion is more than 40 percent higher than that of SB in more than 90 percent of the cases.

2.2 Labour Market Programs[4]

Unemployed people may be directed to labour market states other than employment, such as subsidised employment, training, and (early) retirement. If and how quickly they take this route out of unemployment depends on the supply, the eligibility criteria, and on the incentives provided by such programs. Changes of the UI benefit system can alter these incentives in a different manner than those for exiting to jobs. Hence, considering the overall escape rates from unemployment might not reveal the impact of the UI system on exits to specific states. This issue has been stressed by Atkinson and Micklewright (1991, p. 1713). Also Edin (1989, p. 639) emphasised that „assuming that all transitions from unemployment are governed by the same mechanism, may lead

[4] We focus the discussion of labour market programs to a number of features which may guide the decision and timing of an unemployed person to enter such a scheme. For a more detailed discussion see CCET (1995) or E. Viszt and J. Ványai (1994).

to state aggregation bias[4]. So, we deal with several exit states separately in order to consider aspects that are particular to these states. Let us describe the peculiarities of three labour market schemes which provide an alternative to exits to employment.

Subsidised Employment

In Hungary there exist three types of subsidised employment schemes which are available to unemployed job-seekers: subsidies for employment in the private sector, start-up allowances, and public works. In order to qualify for *subsidies for employment in the private sector*, an UI recipient has to be unemployed for at least six months. The employer is required not to have made redundant workers in a similar line of work so as to avoid displacement effects. The Employment Fund pays for up to half the wage cost for a year. Other programs for the private sector pay subsidies for part-time employment and job-creation investment. Let us now turn to *start-up allowances*. The Employment/Unemployment Act of 1991 introduced a scheme to start-up businesses[5]. It provides a lump-sum equal to six months' UI benefit and also a reimbursement of 50 percent of counselling services or training courses recommended by the labour offices in order to develop a sound business plan. *Public works* is directed to long-term unemployed people. It consists mainly of temporary tasks, such as renovating the infrastructure in less advantaged regions. The public works subsidy may account for up to 70 percent of wage bills.

Training

Training and Retraining options are open to people at risk of losing their job, as well as to all unemployed people. As an incentive to enter such schemes, the training support is set ten percent higher than the UI benefit during the first and more generous phase of the entitlement. Finance is provided by the Solidarity Fund and the Employment Fund. The latter provides the ten percent by which the training benefit exceeds UI.

Early Retirement

People who have been unemployed for at least six months and who are within three years of regular retirement are eligible for the *Pre-Pension Program*. The early retirement age of unemployed men is 57 years, that of women is 52 years. The scheme is financed by the Solidarity Fund. Not everyone who applies is guaranteed to receive benefits, as the funds are limited. The scheme provides a significant alternative to regular employment for aged workers. Like UI recipients, early retirees are still entitled to the full amount of their retirement benefit even when they are simultaneously achieving earnings that do not exceed the minimum wage.

[5] It was introduced in order to replace the Restart Loan Scheme, which was suspended in June 1990 since it had become too costly

3 The Sample of Hungarian UI Recipients

3.1 Strengths and Weaknesses of the UI Register Data

The data set that we study is the inflow to the UI register in December 1992 and January 1993. It was drawn from the Hungarian administrative records of the UI register. Neither spells of job quitters nor of people who received statutory severance pay prior to UI are included in this sample[6]. Also a few spells with a missing work-history record have been discarded beforehand. So, there remain 54911 male and 25200 female unemployment insurance spells. About 37.4 percent of men and 38.1 percent of women are administered by the 1993 benefit provisions[7]. The duration of these spells is measured in days, and we observe them at most until people ceased to be entitled to UI benefits.

The data set has major advantages: Duration is measured in days; systematic measurement errors of unemployment duration, as often found in retrospective surveys, cannot emerge; the size of the data set is considerable, and we observe people administered under two different UI schemes over roughly the same period. So they face the same macroeconomic conditions with respect to labour demand, wage and price inflation, and the wage offer distribution over time.

Nevertheless, there are a number of weaknesses of this UI register data: First, we observe spells only until the expiry of the UI receipt. Consequently, we can infer nothing about the exit rates after that date. Second, the data provides no information on marital status and the households in which people live. Household income and composition would be important to our analysis. On the one hand, they determine whether people receive social benefits after running out of UI. On the other hand, they determine an individual's sources of income that are additional to the UI benefits.

3.2 The Information Provided by the Hungarian UI Register

Exit States

The following duration analysis deals with four exit states: employment, subsidised employment, training and retraining, and (early) retirement. These, alongside duration, form the dependent variables in our study. We do not present any statistics on the duration of spells, as they would not be very indicative due to a high number of right-censored spells[8]. Table 2 displays the shares of completed UI spells by destination in order to distinguish their importance.

The first row of Table 2 shows that 48.4 percent of unemployed men and 33.7 percent of unemployed women exit to employment before their UI entitlement ends. Even when we

[6] The exact sample selection criteria is described in detail in J. Micklewright and Gy. Nagy (1995a).
[7] This composition may lead one to conclude that people who would have otherwise entered the register in January 1993 under the less generous benefit system made their claim in December 1992. J. Micklewright and Gy. Nagy (1995a) suggested a variety of reasons why this was not the case: Firstly, our sample consists of job-losers who cannot entirely control the date of their job losses. Secondly, the date of the of the introduction of the new benefit scheme was not clear until the 23rd of December 1992. Additionally the tax year coincides with the calendar year in Hungary, which may be a reason for enterprises shedding more staff in December than in January.
[8] Right-censored spells are spells that until the end of the observation period have not registered any exit. In our sample the bulk of right-censored spells are those that end by exhausting the UI benefit.

disaggregate by work-history groups, the male share is always higher than that of women. This may reflect the fact that women find a job at a slower rate than men. Unemployed people are much less frequently directed to subsidised employment or training. 4.5 percent of male sample and 3.3 percent of female sample end up in subsidised employment; only 1.3 percent of men compared with three percent of women exit to training. The share of (early) retirement exits in the fourth row is only displayed for people who are either close to the early retirement age or older[9]. For them, retirement is the dominant exit state: more then 65 percent of aged men and 52 percent of aged women retire. Finally, let us draw your attention to the importance of exhausting UI benefits. At 43.5 percent, it is the most frequent way to terminate an UI spell for women. The share of UI exhausters among men, at 35.1 percent, is also remarkable.

Observable Characteristics of our Sample

Table 3 shows a number of observable characteristics by benefit scheme and gender. These are mainly education, region, age and work-history. Our non-parametric analysis which compares the two benefit systems cannot control for all this observed heterogeneity. Hence, we would ideally wish a similar distribution of these characteristics over the two benefit schemes. Let us compare these distributions for the inflow cohorts of 1992 claimants with that of 1993 claimants:

Table 3 shows, that with respect to *educational categories*, the compositions of 1992 and 1993 UI claimants are strikingly similar. The second set of variables displays the *regional distribution* of our sample. We distinguish between the dominant regional activity alongside the regional unemployment rate in 1993[10]. The regional composition of the two claimant groups is apparently different. The share of the 1992 claimants who live in agricultural regions is about 10 percentage points higher than that of 1993 claimants. The opposite is true for diversified regions. The age-distribution over the two benefit schemes is very similar for men, though not for women. The share of women who are below 30 years is roughly eight percent higher for 1993 claimants compared with 1992 claimants.

Table 3 also displays the shares of the ten *work-history groups*. The bulk of spells, more than 40 percent, belong to the longest possible work-history of 48 months. The second largest group of the spells, 22.4 percent of men and 15.2 percent of women, are people with 44-47 months of work-history. The main difference between the two claimant groups emerges for four years of work-history. The 1992 claimants are much more frequently found to belong to this category than the 1993 claimants. The inflow cohort of 1992 claimants who entered unemployment in December 1992 is much larger than that of 1993 claimants who became unemployed in January 1993. One reason for this may be that, in Hungary, employers tend to make redundant those workers with the shortest job tenure. Thus we expect the share of spells with long work-histories to be lower the smaller the inflow cohort. To sum up, there are some observed characteristics, like region, age and

[9] Only 3.8 percent of men and 7.2 percent of women in the entire sample exit to retirement.

[10] We chose this division following an analysis S. Scarpetta (1995) and S. Scarpetta and P. Huber (1995): Their division into regional activity relies on an analysis of employment shares.

work-history, which are not distributed the same way over the inflow samples of 1992 and 1993 claimants.

Benefits and Earnings

We now turn to the benefit and earnings variables. The data set does not provide the UI benefit received by the unemployed people. Therefore the gross benefits are imputed according to the benefit rules: The minimum wage, which is important for this calculation, is set at the level of February 1993. Changes in nominal benefit levels over the spell are only allowed for when they are a consequence of the start of the second entitlement period[11]. Our data provides information on past and past indexed gross earnings. However, as earnings in the analysis that follows, we use a predicted wage. The appendix presents the corresponding equation for the change between post- versus pre-unemployment wages. It is estimated by OLS with data from a second data source on post- and pre-unemployment wages. Note that, due to the small number of explanatory variables that are available for both data sets, we cannot explain much of the variation of this wage gain. So the predicted wage is very close to past earnings indexed for wage inflation until January 1993.

Table 4 describes benefits, earnings, and the replacement rate for men younger than 55 years and women younger than 50 years[12]. The first row displays the benefit levels during the first part of the entitlement period. The mean benefit of women with 9411 HUF is close to the minimum wage (9000 HUF), while that of men is much higher. This reflects the fact that a larger percentage of women receive the minimum benefits[13], since their past earnings, on average, are lower than those of men. This is not so for those men and women whose spells reach the second part of the entitlement period. Both receive, on average, a benefit that is not far from the minimum wage. The replacement rates, as displayed in the last two rows of Table 4, reflect the benefit levels relative to the predicted post-unemployment wage. The mean replacement rate of men is 60.3 percent during the first entitlement period, and for those spells that reach the second entitlement period it is 55.1 percent. Women, in contrast, have higher replacement rates in both entitlement periods, which is again a consequence of gender earnings differences. Also the variation of the replacement rates is larger for women than for men.

4 Econometric Methods for Analysing Unemployment Duration

The statistical analysis of the random variable unemployment duration, T, relies on two concepts: the hazard rate and the survivor function. In the context of unemployment duration, hazard rates can be considered as a reduced form of the neoclassical job search model [Kiefer 1988, p. 648]. So, they are the product of the probability that a job offer arrives and the probability that the

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[11] The benefit level of 1992 claimants, who prior to unemployment achieved low earnings, would rise in line with the minimum wage. Since the minimum wage was raised in February 1993, the 1992 claimants' benefits before that date should be adjusted. So far they are not. Note that this adjustment is not necessary for 1993 claimants as the benefit rules are no longer related to the minimum wage [Micklewright/Nagy 1994b, p. 21].

[12] These variables become important when we turn to the semi-parametric model. By that time we will not analyse aged workers for reasons that later become clear. Note also that a small number of spells with implausible values of their entitlement periods have been excluded from the sample.

[13] About 65 percent of women receive no more than the minimum UI benefit against only 39 percent of men.

individual accepts it. As we distinguish between several exit states, we may refer to the hazard or transition rate as a measure of the speed of exit to a state s, where s has a finite state space.

The hazard or transition rate, $\theta^s(t)$, is the instantaneous probability that a spell terminates in the interval $[t,t+\Delta t)$ with an exit to state s provided that no exit has occurred before:

$$\theta^s(t) = \lim_{\Delta t \to 0, \Delta t > 0} \frac{1}{\Delta t} \cdot \Pr(t \le T < t + \Delta t | T \ge t) \tag{1}$$

$\theta^s(t) \cdot \Delta t$ may be interpreted as the conditional probability of an exit to state s in the interval $[t,t+\Delta t)$ [Blossfeld, H. J. et al., 1989].

The second important concept is the probability of survival, S(t). It represents the unconditional probability of no exit from the current state at T=t. Provided that there are S exit states, it is defined as:

$$S(t) = \Pr(T \ge t) = \exp\left[\sum_{s=1}^{S} -\int_0^t \theta^s(u) \cdot du\right] \tag{2}$$

Non-parametric estimation

The Kaplan-Meier estimator represents a non-parametric estimator of these expressions[14]. Its calculation is straightforward: Let d_{ts} be the number of spells that end in the duration interval $[t,t+\Delta t)$ with an exit to state s, and let N_t be the number of spells for which we observe a duration $T \ge t$. In other words, these are all spells that are observed to reach the start of this interval (the population at risk). The Kaplan-Meier estimator[15] of the transition rate and the probability of survival of this interval are

$$\theta^s(t) = \frac{d_{ts}}{N_t} \qquad (3)^{16}$$

$$S(t) = \prod_{u=t0}^{t} \left[1 - \frac{d_u}{N_u}\right] \qquad (4),$$

where $d_u = \sum_{s=1}^{S} d_{us}$.

[14] The non-parametric estimator of the hazard rate has a substantial advantage: It does not constrain the probability distribution of unemployment duration to a particular functional form. It is also an elegant instrument for displaying graphically how the hazards vary with the duration of spells. Therefore, it is well suited to compare hazards of distinct categories of unemployment spells like people who receive UI under the benefit rules of 1992 and 1993. However, we have to interpret the results of such an analysis with caution: The spell data may not be homogenous with respect to observable and unobservable characteristics of the individuals. For this reason we may find spurious differences.

[15] For a discussion of the Kaplan-Meier Estimator see Blossfeld, Hamerle, and Mayer (1989), Cox, and Oakes (1984), or Kalbfleisch, and Prentice (1980).

Semi-parametric estimation

Parametric estimators allow us to condition the transition rates on observed determinants of unemployment duration. Let us derive the likelihood function for a semi-parametric duration model: We specify the transition rates by an *exponential model with piece-wise constant terms*[17], for three exit states. In doing so we need to define a set of M time intervals. Let I_j denote a time interval where j=1,...,M, and I_j is:

$$I_j = \left[t \middle| v_j \leq t < v_{j+1} \right] \tag{5}$$

where $v_1,...,v_M$ are the points in time that represent the extremes of the M intervals. The transition rate of an individual i to state s is then specified in the following way:

$$\theta_i^s(t) = \exp(\gamma_j^s) \cdot \exp(\mathbf{x}_i^s(t) \cdot \boldsymbol{\beta}^s) \tag{6}$$

where t belongs to I_j . As exit states we consider employment, subsidised employment and training. $\mathbf{x}_i^s(t)$ represents a vector of explanatory variables. In an analysis of unemployment duration, these variables represent the determinants of the reservation wage, the job offer distribution and the arrival rate of job offers. β^s is an unknown parameter vector which will be estimated simultaneously with the γ_j^s (j=1,...,M). The latter represent the *piece-wise constant terms* that reflect the dependence of the hazards on the duration of a spell. As shift parameters they make the model very flexible with respect to dependence on duration against a model that specifies duration dependence by a parametric function.

Suppose there is one more covariate z that determines the hazard. Suppose also, that in contrast to the other covariates (**x**), the impact of z on the hazard varies over the spell length. One could then specify a hazard that allows for period-specific effects; i.e., it allows the coefficients of this covariate to vary with the spell length in the same way as the baseline hazard. Let δ_j (j=1,...,M) represent the parameters of this covariate, the hazard changes then to

$$\theta_i^s(t) = \exp(\gamma_j^s) \cdot \exp(z_i^s \cdot \delta_j^s + \mathbf{x}_i^s(t) \cdot \boldsymbol{\beta}^s) \tag{7}$$

Finally, we turn to the likelihood function. Let a = 1,...,A be the individuals in the sample with completed spells of unemployment insurance of length T_a who exit to state employment, b = 1,...,B those who exit to subsidised employment at length T_b, and c = 1,...,C those individuals who exit to training at spell length T_c. Further, let d = 1,...,D be the individuals whose spells are right-censored at length T_d. These are either spells for which no exit is reported since the individuals exhausted

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[16] This formula has to be altered when the interval length, Δt, is not infinitely small: The hazard in the midpoint of a large interval is defined as: $\dfrac{2 \cdot \theta^s(t)}{[2-\theta^s(t)] \cdot \Delta t}$, where $\theta^s(t)$ is defined according to equation (3).

[17] For a discussion of the exponential model piece-wise constant terms see Lancaster (1990).

their UI benefits, or spells with a transition to an exit state that is not modelled. The likelihood function of this sample is:

$$L = \prod_{a} \theta_i^a(T_a) \cdot S(T_a) \cdot \prod_{b} \theta_i^b(T_b) \cdot S(T_b) \cdot \prod_{c} \theta_i^c(T_c) \cdot S(T_c) \cdot \prod_{d} S(T_d) \tag{8}$$

The contribution to the likelihood function of completed spells, i.e. people who exit to one of the three exit states, is their state specific transition rate times the survivor function evaluated at their spell length. All right-censored spells are represented by the survivor function alone. This is the representation of a competing (independent) risk model. In practice we estimate the parameters of each of the three transition rates by the maximum likelihood of a single risk model; i.e. estimation proceeds for each exit state separately and treats spells that exit to another state as right-censored. Maximum likelihood estimation for exponential duration models with piece-wise constant terms is implemented in a software package called TDA, which has been developed by Götz Rohwer [TDA-Manual 5.7, 1994].

5 Transition Rates: Non-Parametric Analysis

5.1 The Analysis of Micklewright and Nagy in 1995

A non-parametric analysis of the Hungarian UI register data of UI spells has already been carried out by Micklewright and Nagy in 1995[18]. They aimed to reveal disincentives within the Hungarian unemployment benefit system. They examined differences in the speed at which people exit from the unemployment register when they claim benefits under the 1992 benefit scheme or the more restrictive 1993 benefit scheme. So they estimated their transition rates separately. The analysis distinguished between men and women, and four groups of work-history. It focused on transitions to employment (excluding subsidised employment) and composite transitions to all other exit states. Their main results were the following:

1. There is little or no evidence, for most employment history groups, that the more restrictive 1993 benefit scheme raised transition rates to employment.
2. A remarkable difference between the two benefit schemes occurred only for men, who experienced four years of work-history. The 1993 claimant group was found to have far higher transition rates to employment than the 1992 claimant group. This difference occurred in the form of some large spikes very early in their spell, at an elapsed duration of about three months. This elapsed duration coincides in calendar time with early spring 1993, during which the Hungarian labour market was characterised by a short and partly seasonal upturn. As a possible explanation, the authors suggested that the share of workers on recall is higher in the sample of 1993 claimants compared to that of 1992 claimants.
3. The authors found a small rise in the job hazards near the time when the benefits expire.
4. The composite hazards to other exit states of 1993 claimants exceed those of 1992 claimants. So there may be a link between generosity of UI and speed of exit to labour market schemes. Their results also point to a rise in these hazards just before benefits were exhausted.

[18] „Unemployment Insurance and Incentives in Hungary" in Newbery, D (ed.): Tax and Benefit Reform in Central and Eastern Europe and EUI Working Paper 95/7.

An extension of their research may reach stronger conclusions. First, we study the spikes in the employment hazards and thus the influence of the presence of workers on recall on the analysis. Second, we are interested in the impact of the benefit reform on exits to subsidised employment, training, and (early) retirement separately. Note, that for all the transition rates except for those to retirement we study the exit behaviour of men who are below 55 years and women who are younger than 50 years. The reason for this becomes clear when we analyse the transition rates to retirement.

5.2 Transitions to Employment

Workers on Recall and Spikes in the Employment Hazard

In Hungary, as a transition economy, the prime reason for lay-off is the destruction of jobs in the state sector. Hence, one may expect unemployed people to have lost their jobs permanently, so that a typical inflow cohort to unemployment is made up by a small share of workers on recall. Yet this argument is less persuasive for our specific inflow cohorts of December 1992 and January 1993. In these months, important reasons for labour-shedding are seasonal demand fluctuations. So our sample is possibly made up of a significant share of workers on recall, as argued by Micklewright and Nagy (1995). Let us first provide some arguments for this:

- Jobs in which workers are made redundant temporarily for reasons of seasonality are, by and large, male occupations. Consequently, we would expect our male sample to be characterised by a large share of unemployed workers on recall, but not the female one. That, in the previous analysis of Micklewright and Nagy, spikes in the transition rates to jobs are much larger for men than for women, corresponds to this expectation.
- Next, the spikes, that Micklewright and Nagy found for the early spring period of 1993, are most pronounced for unemployed workers with 44-47 months work-history prior to their unemployment spell. These are workers who were jobless for a brief period during the last four years, and this is a typical characteristic of workers who are made redundant temporarily for reasons of seasonal labour demand fluctuations.
- The hypothesis of the presence of a large share of workers on recall would be obsolete if we could repeat the finding of similar spikes for inflow cohorts from other calendar months. Micklewright and Nagy (1996b) estimated non-parametrically the employment hazards of an inflow sample of unemployment spells in Hungary that start in April/May 1994. These job hazards are mostly flat over the duration of unemployment. The only spikes they found arose when the spells reach the final month of UI receipt and the first week following UI exhaustion.

All this implies that our sample, and in particular, our male sample is likely to consist of two distinct types of unemployed workers: (1) Workers who are not on recall. They are workers who are made redundant because of the restructuring of the economy. For them we expect a relatively smooth hazard. (2) Workers on recall who lost their jobs for reasons of seasonality. Their employment hazards are affected strongly by seasonal labour demand fluctuations in early 1993. So, they return to work much faster than workers who are not on recall. This has a very important implication for our analysis: Suppose that the proportion of people laid-off for seasonal reasons is high in our sample. Next, assume it is substantially higher among the 1993 claimants compared with the 1992 claimants. A faster relative exit to jobs of the 1993 claimant group could reflect this fact, but not that the benefit system became more restrictive.

There is clearly need to shed some light on this issue. First of all, we have to prove the existence of a large share of workers on recall, which are not directly observable in our data. However, people who lost their job because of seasonal fluctuations in labour demand have a common characteristic. They are very likely to be unemployed for a short period each year or presumably every two years at around the same calendar time. This is a useful feature, as our data set provides information on how many days before the start of the current spell the last UI spell ended. This information, along with the calendar start of an unemployment spell, enables us to calculate when people ended their last UI spell in calendar time. Accordingly, we partition the sample into two categories of unemployed workers:

(1) unemployed people with *no prior UI spell or a prior UI spell that does not end in mid winter/early spring.* We suppose that they are *unlikely to be on recall.*

(2) unemployed people whose *last UI spell ended in mid winter/early spring* (i.e. between the last week of January and the end of April) either one or two years before 1993. They most likely found their last job during a seasonal upturn, and so we expect them *to be on recall.*

Table 5 shows the composition of our sample by these groups and work-history. It neglects workers with a four year contribution period, since for them nearly no previous UI spell is recorded. The proportion of men who ended their last UI spell in mid winter/early spring is much larger than for women. As expected, the share of these workers is particularly high in the work-history group 44-47 months: For men, 56.6 percent of 1992 claimants and 36.6 percent of 1993 claimants ended their last spell in mid winter/early spring. The corresponding numbers for women are 23.6 and 7.5 percent.

We first demonstrate that the hazards to jobs of workers who ended their last UI spell in late winter/early spring are very sensitive to seasonal labour demand fluctuations. We compare their job hazards to those of the other group. This is done in Figure 1a-b for males who belong to the work-history group 44-47 months and all work-history groups of a shorter employment record. It plots the (four-weekly) Kaplan-Meyer estimates of their job hazards against duration. There is one outstanding feature of this figure. We observe an immense temporary rise of the hazard at around three months of elapsed duration of the group of workers which are supposed to be on recall. The duration for which we observe this spike coincides with the early spring upturn of the Hungarian labour market. In contrast, the response of the job hazards of the other group to this upturn in the Hungarian labour market is negligible. Figure 2a-b shows the results of this analysis for women and we come to no different conclusion.

Let us briefly deal with a question that the previous two figures did not answer: How large is the share of workers on recall who are still jobless after the seasonal upturn in early 1993? Figures 3 and 4 plot for men and women the probability of survival, i.e. of remaining unemployed, against duration. The figures show that roughly 60 percent of workers who are not on recall are still unemployed after a spell length of half a year. By this time the seasonal upturn came to an end and workers on recall are far less likely to be still unemployed. The most distinct cases are male 1993 claimants on recall with 44-47 months of work-history. Less than 20 percent of them are unemployed half a year after their spell started.

To conlude, we illustrate that our sample consists of two types of unemployed workers with a very different elasticity to seasonal employment fluctuations. As suggested by J. Micklewright and Gy. Nagy (1995), a large share of workers on recall in our sample causes the spikes. Our most important concern with workers on recall is that they are not equally distributed over the two claimant groups. They are found more frequently in the sample of 1992 claimants than in that of 1993 claimants. Hence, over the early spring period, the hazards of the more generous benefit system are likely to exceed those of the less generous one substantially. So, we analyse the benefit reform's impact on the transition rates to employment for workers who are unlikely and likely to be on recall, separately. In doing so, we are constrained to dismiss the entire subsample of spells that are characterised by an employment record of four years during the last four years prior to their benefit claim. For them we cannot calculate our indicator variable, that identifies workers on recall, since, by definition, (nearly) no previous UI spell has been observed.

The UI Reform and Transition Rates to Jobs of Workers who are Unlikely to be on Recall

Figures 5 and 6 plot the estimated hazards of people who are not on recall for each benefit regime against duration. For men and women alike we distinguish between the work-history groups: (a) 44-47 months, (b) 28-43 months and (c) 12-27 months. The employment hazards are estimated as the midpoints of four-weekly intervals.

The transition rates to jobs of men are displayed in figure 5a-c. Nearly all of the estimated hazards of men are higher than 0.001 but below 0.003. The hazards of the two highest work-history groups are still characterised by some temporary rise of the hazards during the first three months after the spells start. For all three work-history groups 1993 claimant men exit more rapidly to employment than 1992 claimant men during the first three months of their spell. After an elapsed duration of about three months, it is the 1992 claimants that exit faster to employment.

Figures 6a-c shows the same comparison for women. Their estimated hazards are lower than those of men. Most of them are not higher than 0.0025 and not below 0.0005. 1993 claimant women with 44-47 months of work-history return to work somewhat faster compared with 1992 claimants during the first three months of their spells. Thereafter the hazards of both benefit schemes are about the same. The support for the claim that the benefit reform speeds up the female transitions to jobs is strongest for the work-history group 28-43 months: Most of the estimated transition rates to jobs of 1993 claimants exceed those of 1992 claimants. In contrast, we cannot find that the benefit reform raised the escape rates to jobs of women with 12-27 months of work-history, as shown in Figure 6c.

On the whole this set of figures did not reveal any strong disincentive effects. So, we cannot conclude that the introduction of the new benefit rules in January 1993 accelerated the speed of exit to employment when workers are not on recall. One common feature of the figures 5 and 6 is that employment hazards of 1993 claimants exceed those of 1992 claimants nearly always during the first three months of their spells. Over these months the hazards tend to rise. One may interpret this observation as evidence for the benefit reform exhibiting its main impact early during an unemployment spell. However, the observed difference is likely to be spurious. The spells of the 1993 claimants start roughly one calendar month later than those of the 1992 claimants. Therefore, this difference may only reflect that a general rise in labour demand during this period affects the hazards of 1992 claimants, in terms of elapsed duration, about one month later than those of 1993 claimants. Note also that none of these figures shows any considerable rise in the hazards just prior to UI benefit exhaustion. Thus, there is no evidence for an entitlement effect.

The UI Reform and Transition Rates to Jobs of Workers who are Likely to be on Recall

Workers who are made redundant for seasonal reasons are certainly not a group of workers whose exit behaviour from unemployment is important for the explanation of unemployment duration in transition economies. However, since a large share of our unemployment spells is made up by such workers on recall, we analyse the benefit reforms impact on their transition rates to employment. Their employment hazards are displayed by the benefit scheme in figure 7a-c for men and in figure 8a-c for women. Note, that the spikes in the employment hazard are highest for people with 44-47 months of work-history and become lower with work-history for both men and women. There are two competing interpretations of this feature. One is that our indicator variable identifies workers on recall better the higher the work-history. The other is that the longer workers have been jobless in the last four years the lower is the likelihood to return to work of unemployed workers who lost their job due to seasonal labour demand fluctuations.

For men the most distinctive hazards between the two claimant groups are found for the work-history group or 44-47 months. The hazards of unemployed men on recall of this work-history group administered according to the 1993 benefit rules exceed the hazards of 1992 claimant men by much during the seasonal upturn in the Hungarian labour market. This big difference, however, vanishes once the spells last about six months. For the two other work-history groups we do not repeat this observation. The hazards of both claimant groups are very similar for men with 28-43 months of work-history. In contrast, for men with only 12-27 months of work-history, it is rather the 1992 claimant group which exhibits higher exit rates to jobs, though only after four months of insured unemployment have gone past. To sum up, there is no clear distinction between the hazards of both claimant groups. We come to a similar conclusion when we analyse the female employment hazards in Figure 8a-c: There is some evidence that for women with 44-47 months of employment record the 1993 claimants speed of exit to jobs is somewhat faster compared with that of 1992 claimants. There is however no such evidence for the other two work-history groups as displayed in Figure 8b and 8c.

5.3 Labour Market Programs

We argued earlier that exits to one labour market scheme or another may be accelerated significantly as entitlement to UI becomes shorter. We no longer distinguish between workers who are likely or unlikely to be on recall, since this feature is of less importance when we consider the transition rates to labour market schemes. Here, we do not study this hypothesis for all work-history groups. The reason is that some of the eight work-history groups with less than 44 months of employment record are characterised by a very low number of exits to some schemes. As they are very heterogeneous in terms of the entitlement period, aggregating them to one group would not reveal any „entitlement effect". Therefore, we carry out a non-parametric analysis on exit to labour market schemes only for UI recipients with a work-history of 44-47 months and four years. They represent the bulk of spells of our inflow samples. We analyse them as one group, as their entitlement periods are quite similar. The Kaplan-Meier estimates of the transition rates are calculated for four-weekly intervals.

(Early) Retirement

Exits to (early) retirement are only relevant to men who, at the start of their spell, are at least 55 years old and women who are at least 50 years old. We expect these aged workers to speed up their exit to retirement when UI is paid for shorter periods as after the 1993 benefit reform group. As far as early retirement is concerned, this can only happen after the length of the unemployment spell exceeds the six months waiting period. Figure 9a plots the male hazards to retirement against duration. Not surprisingly, there is nearly no exit before the spells have lasted for at least half a year. After this spell length, men who receive UI benefits according to the 1992 benefit scheme exit at a rate of around 0.005 to retirement. In contrast, 1993 claimants exit much faster. Their hazards are usually more than 50 percent higher than those of the 1992 claimant group. Moreover, at the end of the entitlement period of 1993 claimants a large spike characterises their hazards. This is enough evidence for the hypothesis that shorter entitlement to UI benefits caused a much faster exit rate of the aged UI recipients to (early) retirement. Figure 9b confirms this hypothesis for women.

Subsidised Employment

Subsidised employment is an important alternative to regular employment, as people may gain new rights to claim UI benefits in the future. Thus they may consider this scheme to be particularly attractive when UI benefits are nearly exhausted. It may also be that labour centres offer entry to such schemes as Public Works more frequently to people whose UI benefit is about to run out. Both should result in a substantial rise of the hazards to subsidised employment when the remaining entitlement to UI becomes very short. And this effect should occur earlier, in terms of spell length, when the entitlement period is reduced.

Figure 10a shows the transition rates to subsidised employment for men with a work-history of 44-48 months for each claimant group. Most of these hazards are far lower than those to employment. For both benefit schemes, they rise strongly when UI is about to be exhausted. The 1993 claimant group does not always exhibit a more rapid exit to subsidised employment than 1992 claimants. Their hazards are about the same until 270 days have elapsed. After this period, 1993 claimants' hazards increase up to the end of their UI entitlement. We observe such a rise also by the end of the 1992 claimants' entitlement period. So we may argue that shorter entitlement periods lead to a faster exit to subsidised employment, because the rise in the hazards at around the date of UI expiry occurs earlier. There is less convincing evidence for women. Their transition rates to subsidised employment are displayed in Figure 10b. Though we do find female transition rates of both claimant groups increasing by the end of the entitlement period, this increase is much higher for the more generous 1992 scheme than for 1993 scheme.

Training

The last labour market scheme to consider is training. In contrast to subsidised employment, participation in training does not result in any new rights to UI entitlement. So we do not expect exit rates to training to become particularly high at the end of the UI entitlement period. An incentive to enter training is provided by the training benefit, which is ten percent higher than the UI benefit during the first part of the UI entitlement period. Thus, one may expect that UI recipients' exit rates to training to become higher when this first part of the UI entitlement ends.

We show the training transition rates of men in Figure 11a. The speed of exit to training is quite low. The figure shows that 1993 claimants exit faster to training than 1992 claimants. A small spike in the hazard of 1993 claimants appears in the fourth four-weekly interval just after 90 days. This is immediately after their second part of the entitlement period begins. However, in the next three intervals, their hazard falls off continuously to zero. In calendar time, this happens in the early summer; this may well be a time where no training courses start. By the end of the summer and the start of autumn, at around 270 days, the transition rates for both benefit schemes increase. This rise is clearly stronger for 1993 claimants. It is likely that in this period, a much larger number of training courses started than for the rest of the year. Unfortunately, this is also the time when the more generous first part of UI entitlement of 1992 claimants is about to end. So the figures cannot reveal whether this has any positive impact on the training hazards of 1992 claimants. The slope of training hazards of women in Figure 11b is very similar to that of men, though they are much higher for women. Another difference with men is that female 1993 claimants' hazards do not show a small rise just after 90 days, i.e. immediately after the first entitlement period ends. This is not surprising, since the share of people who actually experience a fall in their benefits is much lower in the female than in the male sample. This analysis suggests, for men and women alike, that exit speed to training accelerated due to the 1993 benefit reform.

6 Transition Rates: Semi-Parametric Analysis

6.1 Specification of the Benefit System

With the exception of transitions of aged workers to retirement, the non-parametric analysis was not sufficient to reach a general conclusion as to how the 1993 benefit reform changed the behaviour of UI recipients. Therefore, we extend this analysis by adopting an exponential hazard rate model with piece-wise constant terms for the baseline-hazard for the transition rates to jobs, subsidised employment and training.

Hazard rates, in our context, can be considered as the product of the probability of receiving a job offer and the probability of accepting it. This conditional probability is determined by the generosity of the UI benefit system, as well as personal characteristics, labour demand conditions, etc.. The models in this section will control for such characteristics as far as possible. In particular, we control for educational level, main regional activity and unemployment rate in 1993, age and squared age, previous unemployment spell, non-manual worker and entry to UI from employment. However, our focus lies on the impact of the benefit system and the 1993 benefit reform on the hazard rates. Therefore, we centre our discussion on the specification of variables related to the benefit system. Dynamic search theory [Mortensen, 1977] provides a guide for modelling the benefit system:

- A *disincentive effect* arises from the amounts of benefit paid, since the value of one more period of unemployment depends positively on the benefit level (relative to prospective earnings). Also, prolonging the potential duration of UI benefits positively affects the value of unemployment and is therefore supposed to diminish the employment hazard.
- A limited period of UI receipt leads to the *entitlement effect*: The value of a future job loss increases when unemployment spells come closer to the expiry date of the benefit. Therefore, their escape rates are higher the shorter their remaining entitlement to UI benefits. Once people run out of benefits, their employment hazards are stable if UI exhaustion is the only cause of non-stationary.

Typically disincentive effects of unemployment benefits are measured by the elasticity of the hazards to the replacement rate. We are including benefits and prospective earnings variables as separate covariates in order to test whether a replacement rate specification is appropriate. Next, the generosity of UI also depends on the benefit scheme under which people are administered. By the coefficients of a dummy that indicates whether people receive UI according to the less generous 1993 benefit rules, we measure how far the 1993 claimants hazards' deviate from those of 1992 claimants.

The econometric estimation results for the three distinct transition rates that we are going to present differ with respect to the treatment of a potential entitlement effect. The specification of the employment hazards does not deal with the impact of the remaining entitlement period. There are two reasons for this. The first is a result of the non-parametric analysis. It did not reveal any rise of the employment hazards when the spells come closer to the end of UI entitlement period. Second, we estimated a number of models in which we included dummies for the remaining entitlement period. Their coefficients were not found to be stable under different specifications of the UI benefit system[19]. However, we adopted a specification of the hazards to subsidised employment and training that takes an impact of the remaining entitlement period into account. At least for subsidised employment, the non-parametric analysis revealed a substantial rise in the transition rates during the last few months prior to end of the UI entitlement. Before turning to our results let us discuss two further issues that are important for the following analysis.

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[19] These estimation results are available on request.

Backward Versus Forward-Looking Replacement Rates

The replacement rate may be expressed either by a backward or forward-looking concept[20]. Whereas the backward-looking concept defines the replacement rate as benefit relative to past earnings, the forward looking concepts defines it as benefits relative to prospective earnings. The literature on incentives of unemployment benefits stresses that prospective wages should measure the *mean of the wage offer distribution* better than past wages do. Past earnings may be a misleading guide to the wages that are available in the market [St. Nickell (1979) and Atkinson et. al (1984)]. Nickell proxied the mean wage offer of an unemployed worker by wages received by employed workers in the same broad skill categories and with the same broad experience. The earnings of employed people may, however, be a bad guide to the wage offers available in the market. Accepted wages may be a better guide. Our earnings measure is based on the wage gain equation of unemployed people: We condition the difference between accepted wages and past (indexed) wages on a set of observable characteristics (see Appendix) and predict expected wages by this equation. However, the predicted wages are still close to the past indexed wage, as we can explain only a minor part of the variation of the post-unemployment wage gain.

Variation of the Explanatory Variables of the Benefit System

The second matter of concern is whether there is enough independent variation to identify the parameters of the benefit level and earnings. UI benefits are imputed by a benefit formula which, given the minimum and maximum benefit rules, introduces variation in the replacement rates. A source of further variation are the two phases of benefit receipt: For a part of the sample, the benefit level changes at different spell lengths during the unemployment spell. And, the two benefit systems, with different benefit formulas, introduce further variation in our benefit data. Despite all these sources of variation, the benefit levels remain highly correlated with the past and even with the prospective earnings measure.

Another matter of concern in the following analysis is the impact of *benefit regime*. We attempt to identify it by the coefficient of a dummy for the 1993 benefit scheme. However, our Kaplan-Meier estimators demonstrated that the slope of the employment hazards depends crucially on whether workers are on recall. In section five, we also discussed that workers on recall are more frequently found among the 1992 benefit claimants than the 1993 claimants. So, an econometric analysis of the employment hazard, that considers the entire sample, is likely to provide coefficients of a regime shift dummy that are downward biased. We avoid this problem by limiting the analysis of transition rates to employment to those workers who are unlikely to be on recall They presumably lost their jobs as a consequence of economic restructuring rather than because of seasonal demand fluctuations. In an economy in transition, they are the unemployed workers in whose exit behaviour we are most interested in. Since it is impossible to identify these workers

[20] There are a lot more unresolved problems with the concept of the replacement rate in our study. In particular, we emphasise that unemployed workers are allowed to work part-time while receiving the full amount of their UI benefit. Next, they may work in the shadow-economy, which is not negligible in Hungary. For these reasons, we would need a replacement rate that results from benefits augmented by additional earnings while unemployed relative to prospective earnings in a future job.

among the work-history group of four years, the modal group, we are constrained to throw away a large number of observations. However, this only applies to the employment hazard, since its slope depends crucially on the recall characteristic.

6.2 Transitions to Employment

We first turn to analysing the impacts of the benefit reform and the level of benefits on the job hazards. The subsample of workers who are unlikely to be on recall is made up by 23494 men and 11209 women. The reference individual of the following models did not experience an UI spell previously, and did not enter unemployment from employment. This person has a primary education. He/she also lives in an agricultural region with an unemployment rate of 12-18 percent in 1993. In addition, the reference individual claims benefits under the 1992 benefit system.

The following discussion focuses on the estimated coefficients of the variables that describe the UI benefit system. We discuss a variety of specifications in order to establish the stability of the effects of the benefit system. The estimated coefficients are displayed in Table 6a for men and Table6b for women[21][22].

Specification one

We start our discussion with the results of Maximum Likelihood estimation of the most simple semi-parametric model which we considered. The unemployment benefit system is represented by only three variables: The logarithms of UI benefits and expected earnings, and a dummy for the 1993 benefit rules. As displayed in Table 6a, we find the elasticity of the male employment hazard to benefits to be 0.17, and that to earnings to be close to zero. So there is no evidence for a disincentive effect of the benefit level, and a LR-test suggests that a specification in terms of the replacement rate would be inadequate. This is different for women. In their case, the estimated elasticity of the employment hazards to benefits is -0.18 and to earnings is 0.15 (Table6b). This result reveals a disincentive effect, though it is hardly significant. An LR-test cannot reject the hypothesis of a replacement rate specification, i.e. that the coefficient of benefits is of the same size and of the opposite sign to that of earnings. The estimated coefficients of the regime shift dummy for the 1993 benefit reform are highly significant and positive. They imply that the regime shift (holding the benefit levels constant) has increased the transition rates to jobs of men and women by around 18 percent.

Specification two

We extend our first specification by allowing the impact of the 1993 benefit rules to vary with duration. So we introduce period-specific effects of the benefit reform, and estimate a model that specifies the hazard as described in equation (7). One important reason for choosing this

[21] The baseline hazard is specified as piece-wise constant shift parameters as described in section 4. The number of observed exits to employment in our sample is high for short durations, but becomes much smaller if we regard spells that last longer than a year. The piece-wise constant terms are therefore estimated for relatively short interval lengths of 30 days, for durations that do not exceed 360 days. Thereafter, we work with an interval length of 45 days.

[22] The estimated coefficients of variables that are not related to the UI system are found in the Table B1 for the job hazard.

specification is that the non-parametric results suggested that the deviation of the employment hazards of 1993 and 1992 claimants varies over the spell length.

The results of Maximum Likelihood estimation of these models, for men and women alike, are displayed in Table 6a and 6b. We first note that, for men and women, an LR-Test rejects the hypothesis that our second specification may be simplified to the nested first specification at a 1 percent significance level. Let us first come to the benefit levels and earnings. The estimated elasticities of UI benefits and prospective earnings of the male model changed somewhat as compared to the first specification. Both are still found insignificant and the hypothesis of a replacement rate specification cannot be accepted. The impact of benefits on the male employment hazards is negligible. Not so for women, the estimated elasticity of benefits is -0.23 in this second specification. So, it is slightly higher in absolute terms as compared with specification one. The replacement rate specification is appropriate according to the corresponding LR-Test. However, the estimated elasticity of benefits is still not significant at a 5 percent significance level.

The period-specific effects for the 1993 benefit rules imply, that it had a relatively large positive effect on the transition rates to employment during the first 90 days of an unemployment spell. Over this period, the employment hazards of 1993 claimants men exceed those of 1992 claimants by 37.7-85.8 percent. The equivalent numbers for women are 12.5-92.5 percent. After an elapsed duration of 90 days, though, the estimated coefficients of the period specific effects of the 1993 benefit rules imply most of the time a less pronounced deviation between the employment hazards of the two claimant groups. Also the sign of this difference alters. Let us note here that these results may be misleading. We cannot interpret them as a relatively high impact of the 1993 benefit reform only early after the start of an UI spell. The reason is the spells of 1993 claimants start in calendar time, in principle, about one month after those of 1992 claimants. Therefore, the coefficients of the period specific effects may reflect the differences in labour demand over two subsequent calendar months. We will soon return to this issue.

Specification three

The first two specifications did not show any large impact of the UI benefits on the employment hazards for the entire sample. In our third specification we allow the coefficients benefit and earnings to vary with age. A reason for studying the age variation of benefit effects has been suggested by Narendranathan et al (1985): A change in the reservation wage has a larger impact on the speed of exit to jobs the less the variation in wage offers. This dispersion is supposed to be increasing in age. Therefore, the hazards to jobs of young people are expected to be most sensitive to benefit levels[23]. Thus our third specification of the employment hazard allows for an interaction

[23] W. Narendranathan et al. (1985) found evidence for this hypothesis by investigating a spell data set drawn from the UK Department of Health and Social Security (DHSS) Cohort Study of the Unemployed 1978/79. The elasticity of expected duration of unemployment with respect to benefit levels is 0.65 for teenage men, 0.47 for 20-24 year old men, 0.26 for the 25-44 year-old men and 0.08 for the over-45-year-old men. Also W. Arulampalam and M. Stewart (1996) found such evidence. They studied the unemployment duration of two different inflow cohorts. The first data set - spells that started in autumn 1978 - were drawn form the DHSS. The second data sets origin is the Department of Social Security (DSS) survey of Incomes In and Out of work, and contains unemployment spells that began in the four weeks starting March 16, 1987. They found a high negative elasticity of the employment hazards to benefit levels of teen-aged men during the first quarter of an unemployment spell for both cohorts. For 20-44 year-old

of benefits and earnings with age. The interaction terms are specified for two age-groups: below 30 years and above 29 years.

Let us first note, that an LR-test does not reject the hypothesis that the nested second specification of the male employment hazards is a valid reduction of the third one. Thus, the age interaction terms of benefits and earnings are irrelevant for determining the male employment hazard. In contrast, for the female employment hazard, the same LR-test rejects this hypothesis. For women who are below 30 years, the estimated elasticity of benefits is -0.47 and that of earnings 0.38. Both coefficients are significant. In contrast, when we consider these interaction terms of women who are above 29 years, both estimated elasticities are close to zero. This lends support to the hypothesis of Narendranathan et al.; only the employment hazards unemployed people who belong to the young age-cohorts respond considerably to variations in the benefit levels. As compared to the second specification, there is no change of the estimated period-specific effects for the 1993 benefit rules.

Specification four

The forth specification considers the work-history as a determinant of the employment hazards. Work-history and, thus, the period for which people paid contributions to UI together with the benefit scheme determines the potential duration of UI entitlement. On the one hand, a longer work-history leads to a higher value of one more period of unemployment. Therefore, one may expect the transition rate to employment to be decreasing in work-history. However, the previous employment record may also be a determinant of the willingness of an employer to take a person on [A. Atkinson and J. Micklewright (1991), p 1709]. On the other hand, one may expect that the employment hazards vary positively with work-history. We add to the model binary variables that indicate whether an individual belongs to one out of four work-history groups: 36-43 months, 28-35 months, 20-27 months, and 12-19 months. The base-group has a past employment record of 44-47 months.

The results of the Maximum Likelihood estimation of the coefficients of specification four are also displayed in Table 6a and 6b. For men we estimated the coefficients for benefits and earnings without age-interaction, since we found that age-interaction is relevant only for women. Again, when we regard the estimation results of the male employment hazard, an LR-test does not reject the hypothesis that this specification may be simplified to the second one. In other words, the coefficients of the work-history terms are jointly zero. So, the second model for men is our preferred specification.

As opposed to the male results, the fourth specification for the female employment hazard cannot be reduced to the nested third one. An LR-Test rejects this hypothesis. The coefficients of benefits and earnings, in absolute terms, are somewhat lower than in the third specification. The estimated elasticity for UI benefits of women who are below 30 years is now -0.36. The estimated

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men and over-44-year-old men they find this elasticity to be lower; in the case of the cohort of 1987 it was even insignificant. Additionally, they found that, for all these age-cohorts, after an elapsed duration of more than three months, the impact of benefit coefficients of the four terms for work-history are all positive and tend to be the higher, the lower the previous employment record, and so the longer the potential duration of the UI receipt. One may interpret this result by an incentive story. The lower the work-history and subsequently the length of UI entitlement, the higher is the employment hazard. However, we should be very cautious with such an interpretation. The three estimated coefficients of the work-history groups 36-43 months, 28-35 months and 20-27 months are roughly of the same magnitude (between 0.22 and 0.27). It is only the lowest work-history groups of 12-19 months for which we find a markedly higher coefficient of 0.41. However, according to this evidence, work-history is not a proxy for the willingness of an employer to employ a person.

The Period-Specific Effects of the 1993 Benefit Rules

The last three models specified the impact of the 1993 benefit rules as period-specific effects. That these coefficients vary with the spell length is partly due to calendar differences in labour demand, since the 1993 claimant group's UI spells start about one month later than those of 1992 claimants. Therefore, the impact of the benefit reform is evaluated best by comparing the hazard of elapsed duration of n months of a 1993 claimant with that of an elapsed duration of n+1 months of a 1992 claimant. We implicitly assume that for them the labour demand situation is the same. These relative hazards are displayed in Table 7. It displays them for 11 intervals of a length of 30 days. Thus, it displays them for those intervals for which we estimated period-specific effects of the 1993 benefit rules and for all three specifications that include these terms. Let us first note that these relative hazards are (with one exception) stable over the three specifications. During the first two months of an unemployment spell the 1993 benefit reform has a negative impact on the male job hazards and no impact on those of women. Thereafter, the hazards of 1993 claimants nearly most of the time exceed those of 1992 claimants.

In all, these results support the hypothesis that the benefit reform increased the incentives to return to work. This conclusion was not entirely clear after our non-parametric analysis. However, our results also tell us that the introduction of the 1993 benefit rules had only a moderate positive impact on the transition rates to employment.

6.3 Transitions to Subsidised Employment and Training

The following analysis includes workers who are likely to be on recall and thus also unemployed workers with four years of work-history. The reason is that being on recall is of much less importance for the slope of the subsidised employment or training hazards than for the job hazards. The specifications consider the remaining entitlement of UI as a determinant of the hazards. The coefficients of four dummy variables for a remaining entitlement of 180-120 days, 120-60 days, 60-30 days and 30-0 days, are assumed to capture this effect.

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levels on the employment hazards ceased to exist.

Subsidised Employment

We first turn to the transition rates to subsidised employment. The non-parametric analysis revealed that exit to subsidised employment speeds up, considerably, during the last few weeks prior to exhausting UI benefits. This entitlement effect does not come as a surprise: one of the important incentives to leave unemployment for these schemes is that they enable people to renew their eligibility to UI benefits after a future lay-off. However, there is a competing hypothesis: Employment offices offer jobs in such schemes more frequently to people who are about to run out of UI. Both suggest that we should concentrate the semi-parametric analysis on the remaining entitlement period. The benefit reform of 1993 reduced the entitlement period by one third. Therefore, it may increase the speed of exit to subsidised employment of 1993 claimants relative to 1992 claimants in two ways. Firstly, the shorter entitlement to UI reduces the value of a period of unemployment and so should increase the transition rates to employment in general. Secondly, if there is a considerable entitlement effect, then the 1993 benefit scheme brought the entitlement effect forward in terms of spell length.

These hypothesis have to take into account legal regulations, which allow only long-term unemployed people to exit to these schemes. Subsidised employment aggregates three different labour market schemes. One of them, subsidies for private sector employment, requires people to be unemployed for at least six months, and the bulk of exits to this scheme occurs after this waiting period. However, this formal eligibility constraint does not hold for the two other schemes: public works and start-up allowances. Given the six months of unemployment eligibility requirement of one of the schemes, we choose to analyse only a subsample of unemployment spells with a work-history of at least 28 months; this ensures that the entitlement period is long enough to reach the six months requirement under both benefit schemes. We only present one specification: The disincentive effects are represented by the logarithms of benefit and earnings and by the regime shift variable[24]. The entitlement effect is again measured by a set of binary variables for remaining entitlement.

The first column of men and women in Table 8 show the Maximum Likelihood estimates of these parameters of the variables which are related to the benefit system[25]. For both sexes, we cannot find any significant effect of benefit levels and earnings on the subsidised employment hazards. The estimated coefficients of the regime shift dummy imply that cutting the entitlement period by one third leads to a significant increase in the subsidised employment hazards: Those of men are 37.7 percent and those of women are 25.8 percent higher compared with those of the 1992 claimant group. Next, the coefficient of both men and women during the last 30 days before they exhaust their UI benefit implies a spike in the subsidised employment hazard: For both sexes, the exit rate is more than twice as high as before their spells reach a remaining entitlement to UI of 180 days. However, we find, only for women, that the subsidised employment hazards always increase

[24] As the number of exits to subsidised employment is much lower than to employment, we consider an interval length for the piece-wise constant terms in our baseline hazard to 45 days.

[25] The coefficients of any other covariate are shown in Table B2.

when the exhaustion date comes nearer. In all, we conclude that there is sufficient evidence for the hypothesis that the 1993 benefit reform raised the speed of exit to subsidised employment substantially[26].

Training

Training is not directed at people who have already been unemployed for a specific period, so we consider the entire sample[27]. The hazard rate model for training is specified in the same way as that for subsidised employment: The reasons for finding a disincentive effect are the same as before. A priori, there is not much reason to expect an entitlement effect: By exiting to training people cannot gain employment history. So they cannot raise the value of future unemployment directly. However, training benefits exceed the first period UI benefit by ten percent. For this reason, the UI recipients might regard an exit to training as a last resort.

Two specifications are discussed; the first excludes work-history as covariates, while the second controls for work-history. The second column of men and women in Table 8 shows the results of maximum likelihood estimation for the training hazard according to specification one. The elasticity of the training hazards with respect to benefits is found to be insignificant. Therefore, a disincentive effect associated with the benefit level does not emerge for either sex. The result for the coefficient of the 1993 benefit scheme dummy implies that the regime change led to a much higher exit to training. It increased the exit rates to training by 36 percent for men and by 49 percent for women. Also, an entitlement effect emerges: The coefficients of remaining entitlement dummies are all positive and jointly significant. They indicate a rise of the training hazards the closer the UI spell comes to the exhaustion date.

This last result, as well as the magnitude of the regime shift dummy, are much altered when we included dummies for five distinct work-history groups as covariates. The Wald tests in both the male and female models cannot reject the hypothesis that the coefficients of the dummies for remaining length of entitlement are jointly zero, but they do not reject this hypothesis for the coefficients of work-history. The coefficient of the regime shift dummy is much higher than in specification one. These results make sense: There is not much economic reason for an entitlement effect. But there is reason to believe that a disincentive effect arises from the length of the entitlement to UI. The shorter this entitlement length, and so the shorter the work-history and the less generous the benefit regime, the larger should be the training hazard. This is exactly what our results suggest.

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[26] A second specification that also takes work-history into account did not alter these results by much.

[27] Exit to training is also much less frequent than that to employment in our sample. Therefore, as for the subsidised employment model, we work with intervals for the piece-wise constant terms of 45 days length. We also distinguish between fewer educational categories and regions. Therefore, the characteristics of our reference person change slightly. He/she now either has primary education or incomplete primary education and lives in a region with an unemployment rate of 12-18 percent.

7 Summary and Conclusions

In this as well as in a previous study by J. Micklewright and Gy. Nagy, an attempt has been made to address one question: Has the reform of the Hungarian UI benefit system in 1993 raised the incentives of unemployed people to exit from unemployment? Our results shed some additional light on this topic:

Transitions to Employment

1. Nagy and Micklewright pointed out that the particular differences in the employment hazards of unemployed workers claiming their benefits under the rules of 1992 and 1993 are hard to explain by an incentive story. They suggested that they may be ascribed to a large share of workers on recall, that differs in size for each of the claimant groups. But how large is the share of workers on recall? To resolve this question we constructed a proxy that identifies two categories of workers with a very distinct elasticity of their job hazards to seasonal labour demand fluctuations. Thus, we were able to prove that the subsample with a work-history of less than four years indeed consists of large number of workers who have been laid off only temporarily. Their share is considerable among men, and significantly higher among the 1992 claimants than among 1993 claimants. A failure to reveal incentive effects of the benefit reform by comparing both claimant groups' employment hazards may emerge if we do not distinguish between these two groups of workers .
2. In a second step, we estimated non-parametrically the employment hazards of 1993 claimants and 1992 claimants for the two distinct groups of workers: workers who are unlikely and those who are likely to be on recall. This analysis is carried out for men and women separately, along three different work-history groups. The analysis yielded scant evidence for the benefit reform having fostered the speed of exit to jobs. This applies to men and women alike and workers that are unlikely and likely to be on recall. But this may still be the consequence of some observed, as well as unobserved, characteristics which determine the employment hazards and which are not distributed in the same way over the two claimant groups. There is no evidence that the job hazards increase while the spells approach the date of UI exhaustion.
3. The semi-parametric analysis of the transition rates to employment of workers who are unlikely to be on recall showed clearly that the overall impact of the UI reform is moderate. According to this analysis the benefit reform has raised the employment hazards of unemployed men and women who are not on recall by less than 20 percent. Note, however, that we achieved this result by assuming the same benefit level for 1992 and 1993 claimants, since benefits were treated as a separate covariate. For men this should not make any difference, since the coefficients of our preferred specification (two) are close to zero. However, this is not true for our preferred specification (four) of the transition rates to employment of women. This specification controls for age-interaction with earnings and benefits. It revealed an elasticity of the female employment hazards to the replacement rate of -0.37 for women who are below 30 years. In contrast, the replacement rate was found to have no impact on the transition rates to employment of women who are older.
4. Though we found some stable results for the impact of UI benefit levels, we need to cast some doubt on this. We cannot be very confident about our specification of the replacement rate by gross benefits and prospective earnings. Their net values would be preferable. What is more, information on other sources of income that are additional to UI benefits, like earnings from part-time employment or employment in the shadow economy, as well as household income, are not

available. So, we cannot consider alternative specifications of benefit and earnings variables that are better suited as guides to incentive effects.

Transitions to Labour Market Schemes

Overall, the analysis of the various labour market schemes and incentives yields results that are in line with economic intuition:

1. Aged UI recipients exit very rapidly to (early) retirement, after a spell length of six months. The benefit scheme seems to play an important role for the exit behaviour to retirement. Once their spells have lasted for the six month waiting period for early retirement, 1993 claimants exit at a rate that is most of the time 50 percent or more above that of 1992 claimants. This result is achieved for the two work-history groups with the longest entitlement periods.
2. The non-parametric analysis of the transition rates to subsidised employment and training did not yield strong conclusions. However, the semi-parametric hazard rate models showed that for both exit states, people in the 1993 benefit scheme exit considerably faster to such schemes than 1992 claimants. This is partly due to a strong entitlement effect. It emerges earlier in the spells of 1993 claimants, since the length of their entitlement to UI is only two thirds of the entitlement period of a comparable 1992 claimant. However, as far as we regard the training hazards, these results are not very stable when the work-history of the UI claimants is taken into account as a covariate: The strong entitlement effect vanishes for the training hazard. Still, one implication of the 1993 benefit reform is that people are driven, more rapidly, into all the kinds of labour market schemes that are available. One reason why the hazards to these schemes are more sensitive to the benefit scheme than the employment hazards may be the depressed labour market.

Appendix: Wage Gain Equation

In order to predict post-unemployment wages, we draw information from a second source of data: An outflow sample from the Hungarian UI register. This sample consists of workers who exited from the UI register by finding employment between March 20th and April 20th in 1994. These workers were interviewed in order to measure their post-unemployment wage.. Like the inflow data that we use for the duration analysis, this data set provides information about regions, education, age, and pre-unemployment earnings. J. Köllo and Gy. Nagy (1996) already carried out an analysis on post-unemployment wage change. We are interested only in a subsample of this data set: job-losers. This data is made up by 6768 men who are below 55 years, and 2577 women who are below 50 years.

Let us now describe the data. Table A1 provides descriptive statistics on education, county, age and duration of unemployment, measured in weeks. For the first two variables we present proportions. More than half the number of men are vocationally trained. With a share of 28.4 percent of the male sample, men with primary education represent the second largest educational category. The bulk of women belong to the categories primary education, vocational training or vocational secondary education. However, in contrast to men, the share of primary educated women, 37.8 percent, is much higher than the share of vocational trained women, which is 24.6

percent. The average age of men is 37.8 years and that of women 35.2. Average durations of joblessness in the outflow sample are 26.8 weeks for men and 33.9 weeks for women.

Table A2 presents pre- and post-unemployment earnings. Previous earnings are indexed to wage inflation, as we want to measure the real wage change. The mean of previous earnings of men is 21662 HUF, the median is about 2000 HUF lower. Post-unemployment wages are found to have an average that is roughly 900 HUF lower, and a median that is about 550 HUF lower than for pre-unemployment earnings. Measuring directly the post-unemployment wage relative to pre-unemployment wages, we observe a median of 94.3 percent for men and 97.1 percent for women. Table A3 describes the difference in logged wages by duration; i.e. since the time the last job ended. The table demonstrates that men and women experience the highest wage losses after very long durations of more than 51 weeks. However, wage gains occur when people experience very short spells of one to ten weeks, or when they are job-less for 21 to 30 weeks. We find a wage loss for those who accept a new job after 11 to 20 weeks. Note that this group of people accounts for roughly half the male sample, but only for a quarter of the female sample. What is not displayed in the table is that over this interval by far the highest wage loss occurred for people who accepted a job after 13 or 14 weeks of joblessness. The data does not provide us with any information about work-history and, thus, about the length of UI entitlement. Nevertheless, we know that for people whose work-history is 44-48 months, 13-14 weeks is the time just after the start of their second part of UI entitlement. So, people presumably accept a high wage loss because their UI benefits have declined. We will therefore control for this period in our wage change equation.

The advantage of this data set is that it certainly shares some observable characteristics with our inflow sample of UI spells. The great disadvantage is that this information is very limited, as we saw in Table A1. Education, county, age and duration (time since last job) of the spells are certainly insufficient to explain the variation of wage levels or the wage levels in logarithms. A large number of unobserved determinants are likely to lead to biased coefficients. Let us assume that these unobservables represent fixed effects. Now, we may interpret this wage data set as a panel, as for each individual we have an observation of the pre-unemployment and an observation of the post-unemployment wage. Thus, by taking the difference we would eliminate the fixed effects, and we could find unbiased estimates of the coefficients of all explanatory variables that vary over time from the first to the second wage observation. But we have only one variable, duration, which satisfies this criterion. Therefore, we assume that variables age, education and county, which are time-invariant, have different impacts on earnings before and after unemployment. Thus, we estimate an equation for wage change, considering this set of variables and the time of joblessness. The equation for the wage gain, the difference in the logged post- and pre-unemployment earnings, is estimated by OLS.

Table A4 presents the results on this equation, which has been estimated for men below 55 years and women below 50 years. Note, first the performance of this equation is poor. We can only explain a minor part of the variation in our data. The adjusted R^2s are 0.035 for men and only 0.024 for women. Furthermore, a Reset-test rejects the hypothesis of no omitted variables at a 5 percent significant level for women. According to the Cook-Weisberg-test, we also find evidence for heteroscedasticity.

Turning to our regressors, we find the same negative relation between logged wage differential and logged age for men and women; the elasticity of age is -0.151. Educational differences are not very important to determining the wage gain of women; all educational categories' coefficients are insignificant. Also, according to an F-test, they are not significantly different from zero. So, whether or not education is higher or lower than that of the base group, vocational trained people, does not help to explain the wage change of women. In the case of men, we cannot reject the joint hypothesis that all educational categories are different from zero. The coefficients for men of each educational category point to a higher wage loss as compared to vocationally trained men. It is most pronounced for those with primary education, whose wage loss is about five percent as compared to vocational trained workers.

The county variables turn out to be jointly significant for both sexes. Relative to the base group, people who live in Bekes, we find a number of counties who do better in terms of their wage gain. The male equation shows that this is the case in all counties. Men in Budapest, Pest, Tolna and Vas even achieve a wage gain that is more than ten percent higher than for men who live in Bekes. Also women in most of the other counties do better than women of Bekes. However, the differences relative to women in Bekes are not as pronounced as for men. E.g., only women who live in Budapest, Bacs-Kis and Tolna achieve a wage gain that is more than ten percent higher than the women in Bekes.

Duration has a negative impact on wage change. A one percent change in duration leads to a wage loss of 0.033 percent for men and 0.029 percent for women. However, we included a duration dummy for the 13th and 14th week of joblessness, which alters this result during this period. It turned out to be significant and highly negative for both men and women. Note that the results for duration have to be taken with much caution, as there is a problem of endogeneity. Expected post-unemployment wages for our inflow sample to the UI register are predicted by these two equations evaluated at a duration of one week.

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Table 1: The Hungarian UI Benefit System
Before and After the Reform in January 1993

1. Eligibility Criteria

- At least 12 months of contributory employment during last four years prior to UI claim.

2. Length of Entitlement

- Ten different entitlement lengths to UI, depending on the previous employment record.
- Each entitlement period is divided into two phases of which the first is characterised by a higher formal replacement rate.
- The reform in 1993 reduced all ten entitlement periods by one third. Also the balance between the two phases changed: In 1992 two thirds (first phase) and one third (second phase). In 1993 one quarter (first phase) and three quarters (second phase)

Work-history in the previous four years	December 1992 Length of UI Entitlement			January 1993 Length of UI Entitlement		
	First phase	Second phase	Total	First phase	Second phase	Total
(Months)		(Days)			(Days)	
12-15	90	45	135	23	67	90
16-19	120	60	180	30	90	120
20-23	150	75	225	38	112	150
24-27	180	90	270	45	135	180
28-31	210	105	315	53	157	210
32-35	240	120	360	60	180	240
36-39	270	135	405	68	202	270
40-43	300	150	450	75	225	300
44-47	330	165	495	83	247	330
48	360	180	540	90	270	360

3. Benefit Formula

- Benefits are earnings-related. The benefits are calculated according to the average monthly earnings during the last four quarters prior to the benefit claim.
- Maximum benefits are twice the minimum wage (minimum wage = 9000 HUF in Feb. 1993).
- The minimum benefits were equal to the minimum wage under the 1992 benefit provisions. The 1993 reform reduced the minimum benefits.

	December 1992		January 1993	
	First phase	Second phase	First phase	Second phase
(a) formal replacement rate (in percent)	70	50	75	65
(b) maximum benefit (in HUF)	18000	18000	18000	15000
(c) minimum benefit (in HUF)	9000	9000	8600	8600
(d) replacement rate if average past monthly earnings < minimum benefit (in percent)	100	100	100	100

4. Other

- Level of UI benefits remains unchanged, when UI claimant achieves earnings from part-time work that are not higher than the minimum wage.

Table 2: Share of Exit to Different Labour Market States (in percent)

Exits	Men			
	All spells	by work-history (in months)		
		48	44-47	12-43
	(54911 obs.)	(22242 obs.)	(12305 obs.)	(20364 obs.)
Exit States:				
Employment	48.4	51.0	59.4	38.9
Subsidised Employment	4.5	4.8	4.5	4.2
Training	1.3	1.2	1.1	1.4
(Early) Retirement[1)]	65.1 (2413 obs.)	73.8 (1659 obs.)	50.9 (409 obs.)	39.7 (345 obs.)
other exit states	3.0	2.5	2.7	3.8
Right-Censored:	39.0	33.8	31.9	50.5
of which: Exhaust UI	35.1	30.3	26.7	45.3

Exits	Women			
	All spells	by work-history (in months)		
		48	44-47	12-43
	(25200 obs.)	(11916 obs.)	(3832 obs.)	(9452 obs.)
Exit States:				
Employment	33.7	34.5	35.8	31.8
Subsidised Employment	3.3	3.9	3.4	2.4
Training	3.0	3.3	2.9	2.7
(Early) Retirement[1)]	52.7 (2993 obs.)	63.2 (1868 obs.)	47.1 (465 obs.)	26.8 (660 obs.)
other exit states	4.4	4.0	5.8	4.5
Right-Censored:	48.4	43.3	45.3	56.0
of which: Exhaust UI	43.5	38.4	39.7	51.3

1) only men older than 54 years, women older than 49 years

Table 3: Observable Characteristics by UI Benefit Scheme
(in Percent)

Observable Characteristics	Men			Women		
	92 scheme (34392 obs.)	93 scheme (20519 obs.)	Total (54911 obs.)	92 scheme (15609 obs.)	93 scheme (9591 obs.)	Total (25200 obs.)
Education						
inc. primary	6.3	5.8	6.1	5.0	5.6	5.2
primary	32.8	33.4	33.0	45.2	44.0	44.7
vocational	45.6	45.9	45.7	22.1	24.0	22.8
vocational sec.	9.8	9.2	9.5	13.8	12.8	13.4
gen.sec	2.7	3.1	2.9	11.2	10.8	11.1
college	1.8	1.7	1.7	1.8	2.1	1.9
university	1.1	1.0	1.0	0.8	0.9	0.8
Region						
Agricultural:	51.8	41.5	48.0	43.7	34.7	40.3
Unemployment Rate						
< 12%[1)]	2.4	4.2	3.1	2.9	4.4	3.5
12-18 %[2)]	37.3	30.0	34.6	31.0	25.0	28.7
> 18%[3)]	12.1	7.3	10.3	9.8	5.4	8.1
Industrial:	18.4	19.2	18.7	21.9	18.9	20.8
Unemployment Rate						
< 12%[4)]	6.7	7.3	6.9	7.7	8.3	7.9
12-18 %[5)]	2.6	2.5	2.6	3.6	2.8	3.3
> 18%[6)]	9.1	9.5	9.2	10.6	7.9	9.6
Diversified:	29.8	39.3	33.4	34.4	46.3	38.9
Unemployment Rate						
< 12%[7)]	22.3	31.3	25.7	26.5	38.2	30.9
12-18 %[8)]	7.5	8.0	7.7	7.9	8.1	8.0
Age						
younger than 30	30.6	33.5	31.7	26.7	34.4	29.7
30-54 (men) and 30-49 (women)	64.4	63.1	63.9	62.7	57.5	60:7
older than 54 (men) and 49 (women)	5.0	3.4	4.4	10.5	8.1	9.6
Work-history						
48	43.4	35.7	40.5	51.7	40.1	47.3
44-47	22.1	22.9	22.4	13.9	17.3	15.2
40-43	8.2	9.2	8.6	6.7	8.1	7.2
36-39	4.9	6.4	5.4	4.9	6.4	5.4
32-35	3.6	5.1	4.1	3.8	5.3	4.4
28-31	4.5	5.7	4.9	3.9	5.7	4.6
24-27	3.8	4.7	4.1	4.0	5.0	4.4
20-23	2.8	4.0	3.3	3.0	4.5	3.6
16-19	3.4	4.0	3.7	3.4	4.3	3.8
12-15	3.3	2.3	3.0	4.6	3.3	4.1
Other						
previous spell of UI	34.3	30.9	33.0	21.1	18.6	20.1
entry to UI from employment	94.6	91.1	93.3	91.3	86.3	89.4
non-manual	9.0	7.9	8.6	27.0	23.0	25.5

Regions: 1) Somogy; 2) Bacs-K., Bekes, Hajdu-B., Szolnok, Tolna;
3) Szabolcs; 4) Fejer, Veszprem; 5) Komarom; 6) Borsod, Nograd;
7) Budapest, Csongrad, Gyor-S., Pest, Vas, Zala; 8) Baranya, Heves

Table 4: UI Benefits and Earnings[1)]

Explanatory Variables UI system	Obs	Mean	Men Std.Dev.	Lower Decile	Upper Decile
Benefits/Wages (HUF)					
Benefit Period 1	52491	11089.7	3186.2	8396.0	16821.0
Benefit Period 2	30433	9514.5	2044.3	8000.0	12442.2
Indexed Wages	52491	18280.1	9391.0	10044.6	2892.2
Predicted Post-Unemployment Wage	52491	19789.5	10307.7	10884.7	30628.7
Replacement Rate					
in % Period 1	52491	60.3	11.3	48.7	76.4
in % Period 2	30433	55.1	15.0	38.3	77.6

Explanatory Variables UI system	Obs	Mean	Women Std.Dev.	Lower Decile	Upper Decile
Benefits/Wages (HUF)					
Benefit Period 1	22776	9410.6	2737.4	6997.0	12966.8
Benefit Period 2	15813	8681.8	1850.0	7000.0	9859.0
Indexed Wages (iw)	22776	14444.9	6987.0	8322.1	22416.2
Predicted Post-Unemployment Wage	22776	15906.9	7839.2	9096.8	24830.4
Replacement Rate					
in % Period 1	22776	63.1	12.5	49.4	80.5
in % Period 2	15813	61.3	15.3	40.1	80.4

1) only men younger than 55 years, women younger than 50 years

Table 5: Unemployment Spells by Calendar End of Previous UI Spell

Work-history	workers unlikely to be on recall1) obs.	share	workers likely to be on recall[2)] obs.	share	Total obs
Men, 1992 scheme					
12-27 months	4240	93.6	288	6.4	4528
28-43 months	5695	80.0	1428	20.0	7123
44-47 months	3186	43.4	4158	56.6	7344
Men, 1993 scheme					
12-27 months	2869	93.9	185	6.1	3054
28-43 months	4619	87.0	689	13.0	5308
44-47 months	2885	63.4	1667	36.6	4552
Women, 1992 scheme					
12-27 months	2171	96.5	79	3.5	2250
28-43 months	2490	88.4	327	11.6	2817
44-47 months	1501	76.4	463	23.6	1964
Women, 1993 scheme					
12-27 months	1492	95.6	69	4.4	1561
28-43 months	2153	92.8	167	7.2	2320
44-47 months	1402	92.5	114	7.5	1516

1) Either no previous UI spell or UI spell ended at other date than mid winter/early spring.

2) Last UI spell ended during mid winter/early spring one or two years before current spell

Table 6a: Maximum Likelihood Estimates of Transition Rate to Employment: Coefficients of Variables of the UI System

	Men							
	Spec. 1		Spec. 2		Spec. 3		Spec. 4	
Variable	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
(1) Benefits and Earnings								
log. of UI benefits	0.17	1.7	0.04	0.4			0.02	0.2
log. of earnings	0.04	0.7	0.12	1.9			0.13	2.0
LR-Test[1] (Ho:ß1+ß2=0, $Chi^2(1)$)	15.6	**	7.8	**			6.9	**
Age Interactions								
aged < 30 years								
log. of UI benefits					0.01	0.1		
log. of earnings					0.14	1.7		
LR-Test[1] (Ho:ß1+ß2=0, $Chi^2(1)$)					7.3	**		
aged >= 30 years								
log. of UI benefits					0.05	0.5		
log. of earnings					0.11	1.5		
LR-Test[1] (Ho:ß1+ß2=0, $Chi^2(1)$)					8.5	**		
(2) 1993 Benefit Rules								
(a) No Period Specific Effects	0.16	7.4						
(b) Period Specific Effects								
0-30 days			0.62	7.3	0.62	7.3	0.62	7.3
30-60 days			0.32	5.0	0.32	5.0	0.32	5.0
60-90 days			0.42	8.1	0.42	8.1	0.42	8.1
90-120 days			0.06	1.2	0.06	1.2	0.06	1.2
120-150 days			-0.04	-0.7	-0.04	-0.6	-0.04	-0.7
150-180 days			-0.05	-0.8	-0.05	-0.8	-0.06	-0.8
180-210 days			-0.25	-3.1	-0.25	-3.1	-0.25	-3.1
210-240 days			0.10	1.1	0.10	1.1	0.09	1.0
240-270 days			0.31	3.0	0.31	3.0	0.30	2.9
270-300 days			0.09	0.7	0.09	0.8	0.08	0.7
>300 days			0.11	0.7	0.11	0.7	0.12	0.8
(3) Work-history								
36-43 months							0.04	1.5
28-35 months							-0.02	-0.6
20-27 months							0.00	0.1
12-19 months							-0.01	-0.3
Log of the Likelihood	-65862.4		-65805.4		-65803.8		-65802.9	
Number of Spells	23494		23494		23494		23494	
LR-Tests[1)2)] (nested models)								
against Specification 1			114.07 ** $Chi^2(10)$		117.29 ** $Chi^2(12)$		119.13 ** $Chi^2(14)$	
against Specification 2					3.22 $Chi^2(2)$		5.06 $Chi^2(4)$	

1) Note, one asterisk "*" implies that the null hypothesis is rejected at a 5% significance level and two asterisks "" a 1% significance level.**

2) The Chi^2 statistic tests the null hypothesis that a specification may be reduced to a more parsimonious nested specification.

Table 6b: Maximum Likelihood Estimates of Transition Rate to Employment: Coefficients of Variables of the UI System

	Women							
	Spec. 1		Spec. 2		Spec. 3		Spec. 4	
Variable	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
(1) Benefits and Earnings								
log. of UI benefits	-0.18	-1.2	-0.23	-1.5				
log. of earnings	0.15	1.6	0.18	1.8				
LR-Test[1] (Ho:ß1+ß2=0, Chi²(1))	0.16		0.44					
Age Interactions								
aged < 30 years								
log. of UI benefits					-0.47	-2.7	-0.36	-2.0
log. of earnings					0.38	3.0	0.33	2.5
LR-Test[1] (Ho:ß1+ß2=0, Chi²(1))					1.3		0.2	
aged >= 30 years								
log. of UI benefits					-0.06	-0.3	0.02	0.1
log. of earnings					0.03	0.2	-0.01	-0.1
LR-Test[1] (Ho:ß1+ß2=0, Chi²(1))					0.1		0.0	
(2) 1993 Benefit Rules								
(a) No Period Specific Effects	0.17	4.8						
(b) Period Specific Effects								
0-30 days			0.66	5.9	0.66	5.9	0.67	6.0
30-60 days			0.16	1.6	0.16	1.6	0.17	1.8
60-90 days			0.12	1.3	0.12	1.3	0.14	1.5
90-120 days			0.24	2.6	0.25	2.7	0.28	3.0
120-150 days			-0.04	-0.4	-0.04	-0.3	0.01	0.1
150-180 days			0.10	1.0	0.11	1.0	0.14	1.3
180-210 days			-0.02	-0.1	-0.01	-0.1	0.01	0.1
210-240 days			0.32	2.3	0.32	2.3	0.36	2.6
240-270 days			0.12	0.8	0.13	0.8	0.18	1.2
270-300 days			-0.09	-0.5	-0.09	-0.4	-0.02	-0.1
>300 days			0.16	0.7	0.17	0.8	0.29	1.3
(3) Work-history								
36-43 months							0.22	4.5
28-35 months							0.27	4.9
20-27 months							0.25	4.1
12-19 months							0.41	5.9
Log of the Likelihood	-26800.1		-26785.0		-26770.4		-26748.6	
Number of Spells	11209		11209		11209		11209	
LR-Tests[1)2)] (nested models)								
against Specification 1			30.33 ** Chi²(10)		59.49 ** Chi²(12)		103.02 ** Chi²(16)	
against Specification 2					29.16 ** Chi²(2)		72.69 ** Chi²(6)	
against Specification 3							43.53 ** Chi²(4)	

1) Note, one asterisk "*" implies that the null hypothesis is rejected at a 5% significance level and two asterisks "" a 1% significance level.**

2) The Chi² statistic tests the null hypothesis that a specification may be reduced to a more parsimonious nested specification.

Table 7: Relative Hazards of 1993 Claimants versus 1992 Claimants[1)]

Interval (elapsed duration)	Men			Women		
	Spec. 2	Spec. 3	Spec. 4	Spec. 2	Spec. 3	Spec. 4
0-30 days	0.87	0.87	0.87	1.07	1.07	1.09
30-60 days	0.84	0.84	0.84	0.94	0.94	0.95
60-90 days	1.06	1.06	1.06	1.12	1.12	1.14
90-120 days	1.16	1.16	1.16	1.19	1.19	1.22
120-150 days	1.09	1.09	1.09	0.89	0.90	0.92
150-180 days	0.93	0.93	0.93	1.14	1.14	1.16
180-210 days	1.08	1.08	1.08	1.29	1.30	1.32
210-240 days	1.25	1.25	1.24	1.20	1.21	1.25
240-270 days	1.14	1.14	1.13	1.18	1.19	1.24
270-300 days	1.39	1.39	1.38	0.99	1.00	1.06
300-330 days	1.60	1.60	1.63	1.56	1.57	1.74

1) The relative hazards are calculated as the hazard of a 1993 claimant during the given interval of elapsed duration relative to that of a 1992 claimant during the subsequent interval of elapsed duration.
Both claimants are assumed to differ only by the benefit regime.

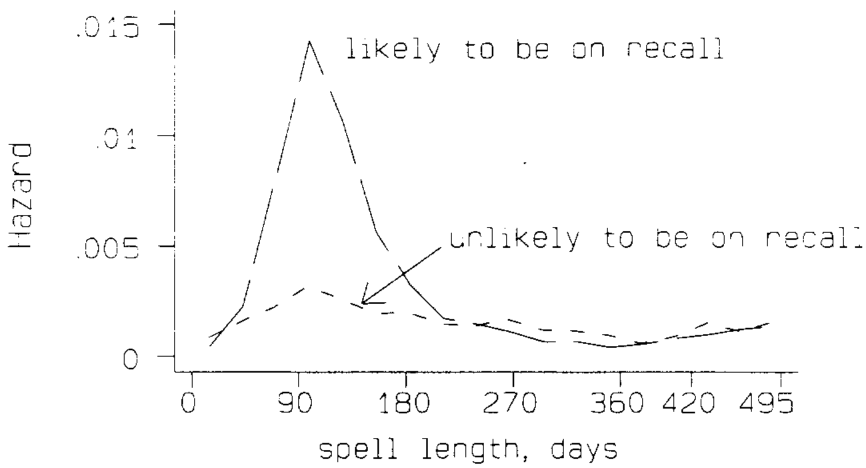
Table 8: Maximum Likelihood Estimates for Transition Rates to Subsidised Employment and Training: Coefficients of Variables of the UI System

	Men Transition Rates to						Women Transition Rates to					
	Subsidised Employment		Training				Subsidised Employment		Training			
			Spec. 1		Spec. 2				Spec. 1		Spec. 2	
Variable	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
(1) Benefits and Earnings												
log. of UI benefits	-0.13	-0.6	-0.12	-0.4	-0.07	-0.2	-0.36	-1.1	0.12	0.3	0.11	0.3
log. of earnings	-0.09	-0.7	0.55	3.2	0.55	3.2	0.07	0.3	0.43	2.2	0.45	2.2
Wald Test[1] (Ho:ß1+ß2=0, $Chi^2(1)$)	3.70		5.5*		6.5*		2.80		8.6**		8.8**	
(2) 1993 Benefit Scheme	0.32	4.7	0.30	3.2	0.52	4.8	0.23	2.0	0.40	4.3	0.55	5.1
(3) Remaining Entitlement to UI												
180-120 days	-0.19	-2.0	0.30	2.3	-0.04	-0.3	0.03	0.2	0.35	2.8	0.16	1.1
120-60 days	-0.10	-1.0	0.37	2.7	-0.07	-0.4	0.15	0.9	0.29	2.2	0.03	0.2
60-30 days	0.11	0.9	0.67	4.3	0.17	0.9	0.24	1.2	0.30	1.8	0.02	0.1
30-0 days	0.90	7.8	0.57	3.4	0.03	0.2	0.86	4.3	0.52	3.2	0.22	1.1
Wald Test[1] (joint signif., $Chi^2(4)$)	149.6**		21.0*		2.90		28.8**		13.3**		3.10	
(4) Work-history												
44-47 months					0.17	1.4					-0.12	-1.0
36-43 months					0.36	2.6					0.20	1.5
28-35 months					0.65	3.8					0.34	2.0
20-27 months					0.80	3.7					0.49	2.3
12-19 months					1.01	3.8					0.57	2.2
Log of the Likelihood	-19246.2		-7032.3		-7022.2		-7151.7		-6980.1		-6974.5	
Number of Observations	44909		52491		52491		18965		22776		22776	
LR-Tests[1,2] (nested models)												
against Specification 1					20.14 ** $Chi^2(5)$						11.24 * $Chi^2(5)$	

1) Note, one asterisk "*" implies that the null hypothesis is rejected at a 5% significance level and two asterisks "" a 1% significance level.**
2) The Chi^2 statistic tests the null hypothesis that a specification may be reduced to a more parsimonious nested specification.

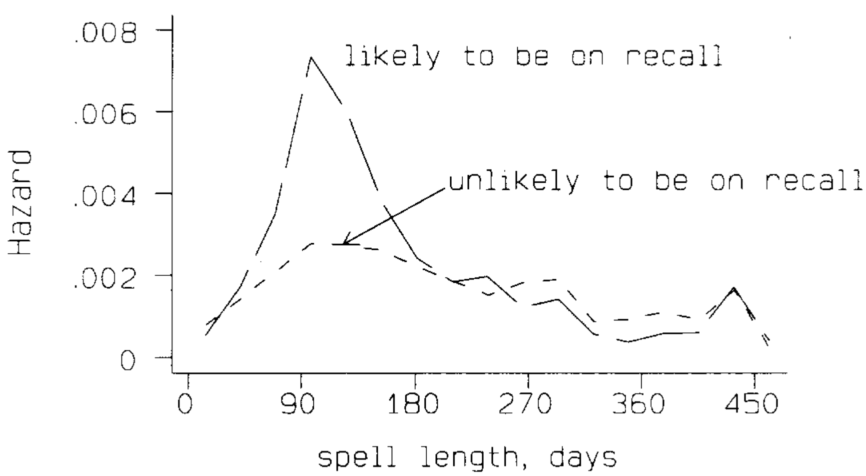
Figure 1 : Transition Rates to Employment of Men workers on recall versus workers not on recall

[a] 44-47 months of work-history



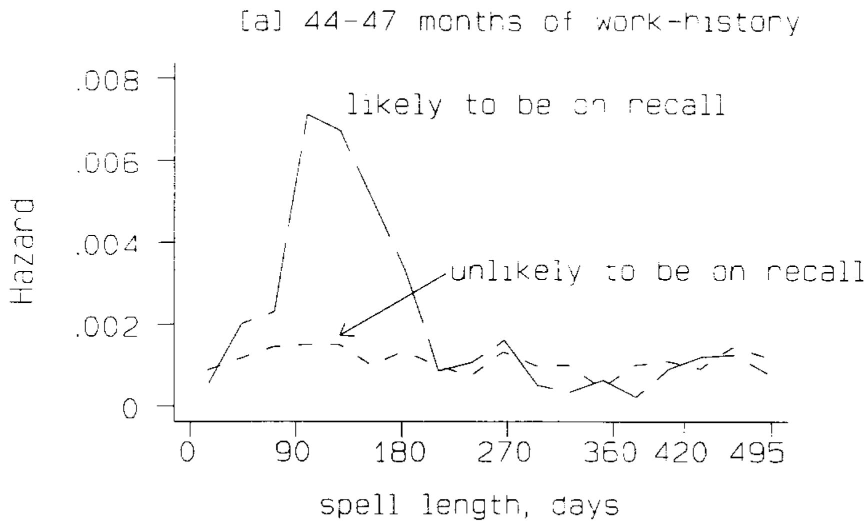
Note: sample sizes 6071 (unlikely to be on recall) and 5825 (likely to be on recall)

[b] 12-43 months of work-history

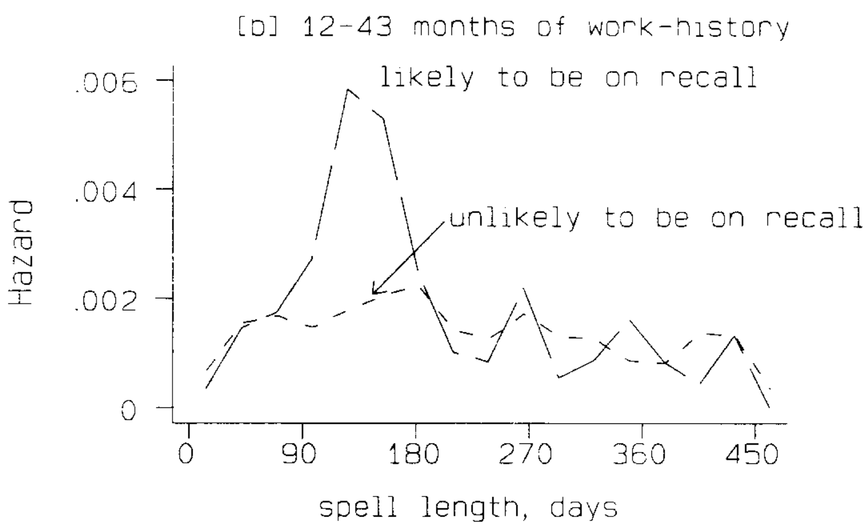


Note: sample sizes 17423 (unlikely to be on recall) and 2590 (likely to be on recall)

Figure 2 : Transition Rates to Employment of Women workers on recall versus workers not on recall



Note: sample sizes 2903 (unlikely to be on recall) and 577 (likely to be on recall)



Note: sample sizes 4661 (unlikely to be on recall and 406 (likely to be on recall)

Figure 3: Survivor Function of Men

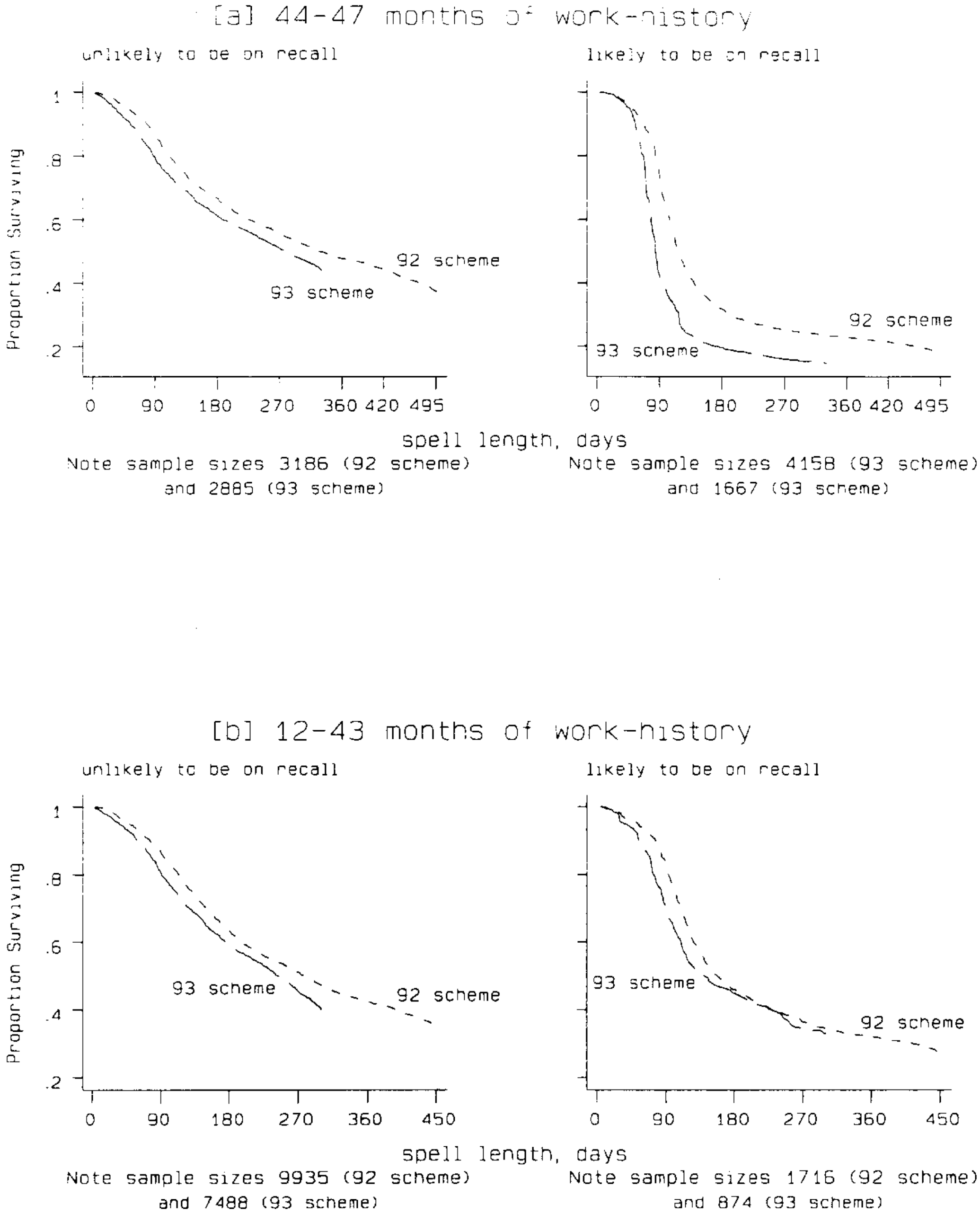


Figure 4: Survivor Function of Women

[a] 44-47 months of work-history

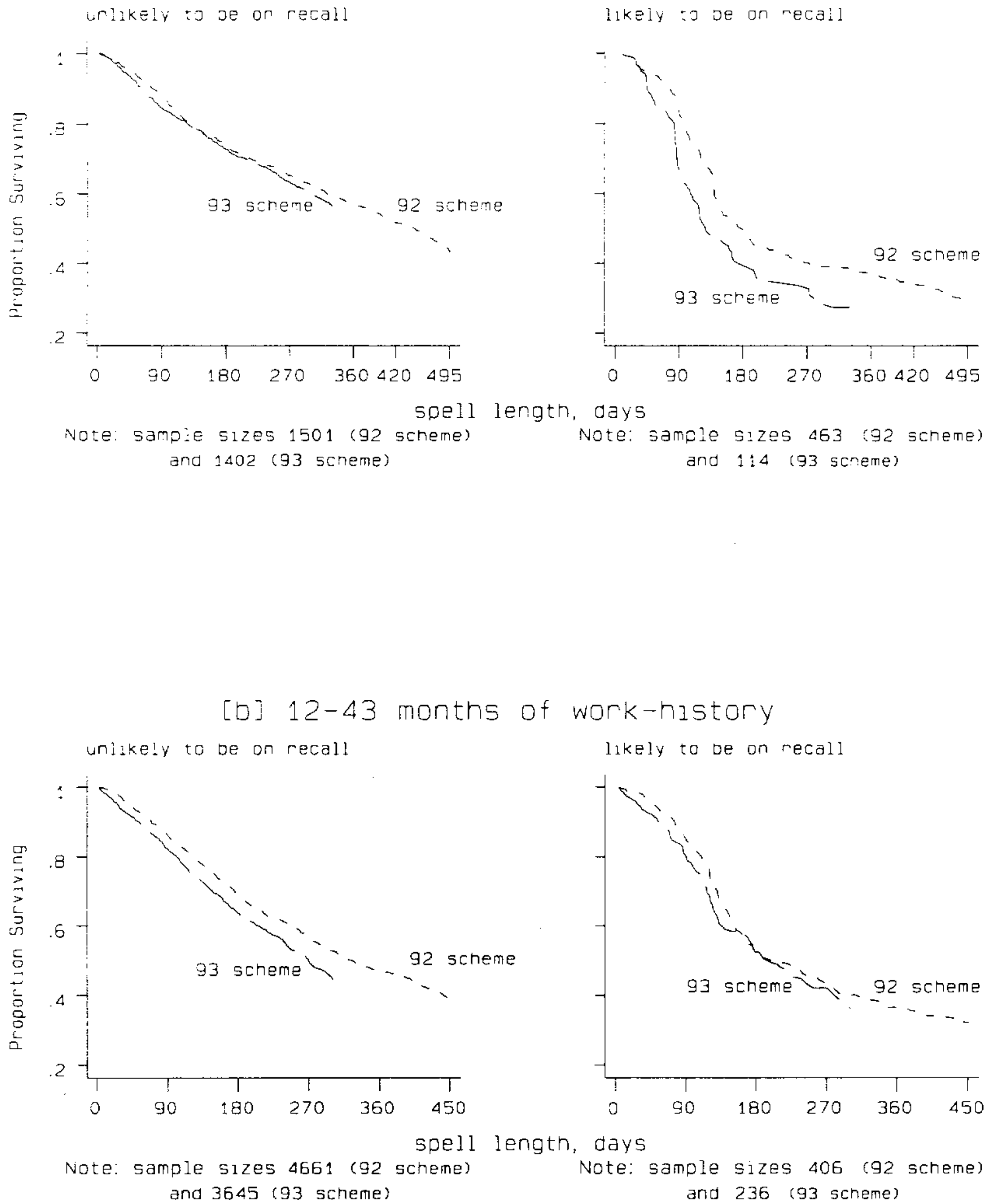
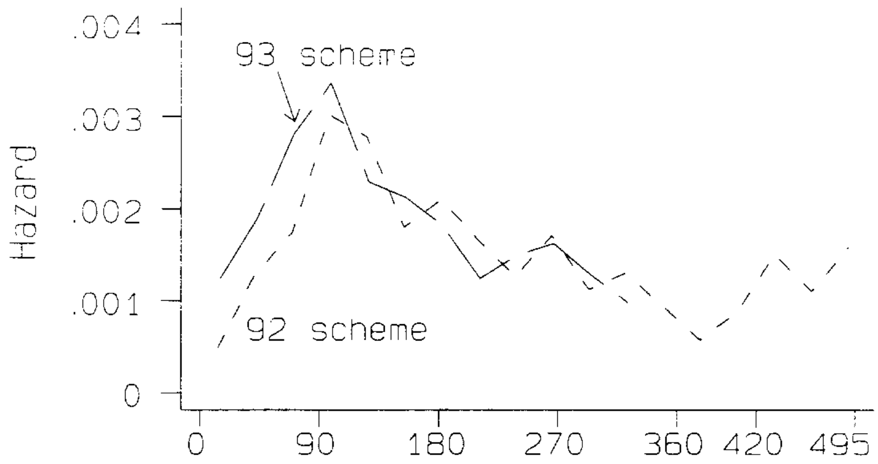


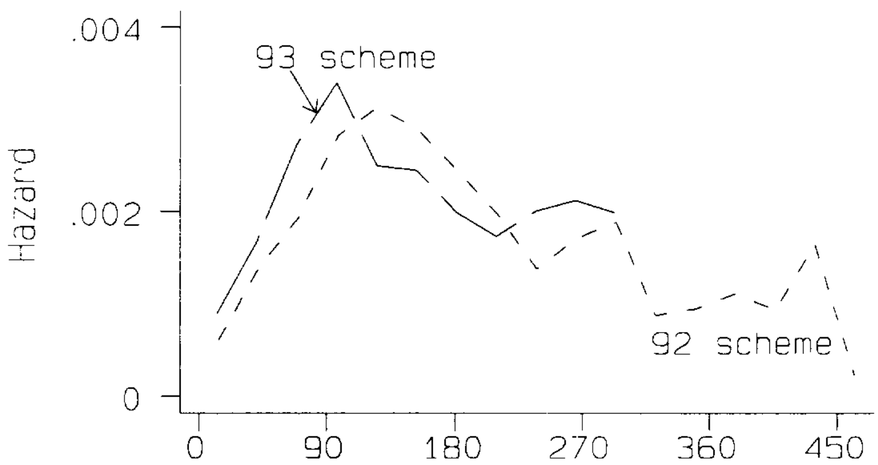
Figure 5 : Transition Rates to Employment of Men
unlikely to be on recall

[a] 44-47 months of work-history



Note: sample sizes 3186 (92 scheme) and 2885 (93 scheme)

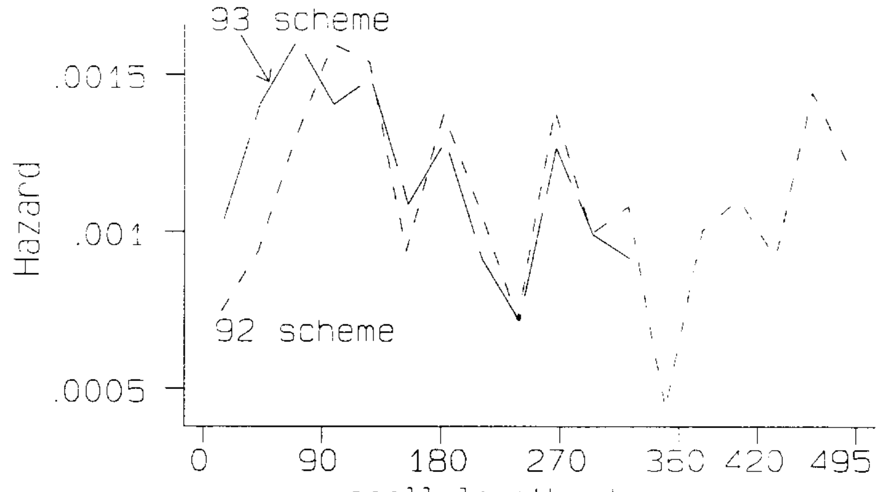
[b] 28-43 months of work-history



Note: sample sizes 5695 (92 scheme) and 4619 (93 scheme)

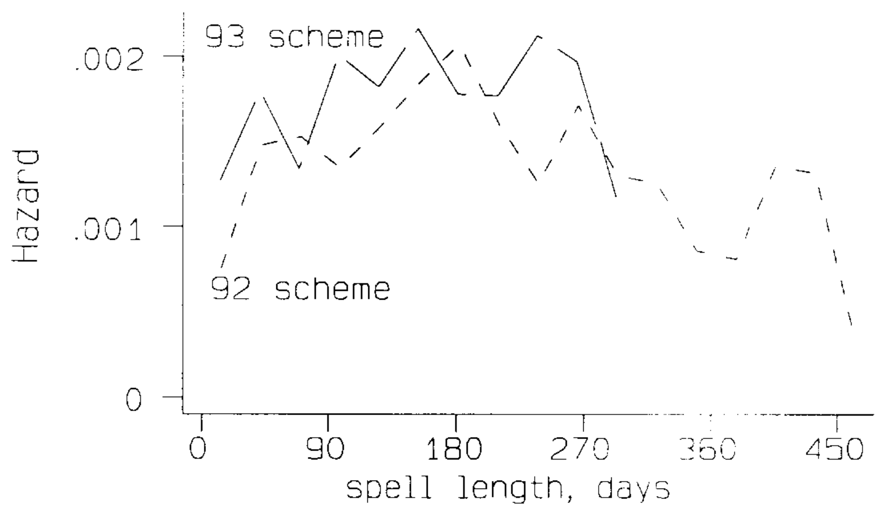
Figure 6 : Transition Rates to Employment of Women unlikely to be on recall

[a] 44-47 months of work-history



Note: sample sizes 1501 (92 scheme) and 1402 (93 scheme)

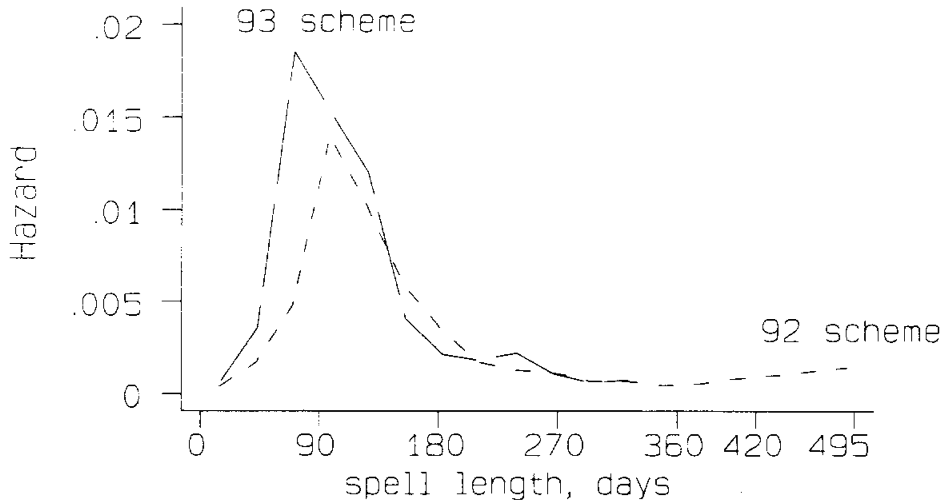
[b] 28-43 months of work-history



Note: sample sizes 2490 (92 scheme) and 2153 (93 scheme)

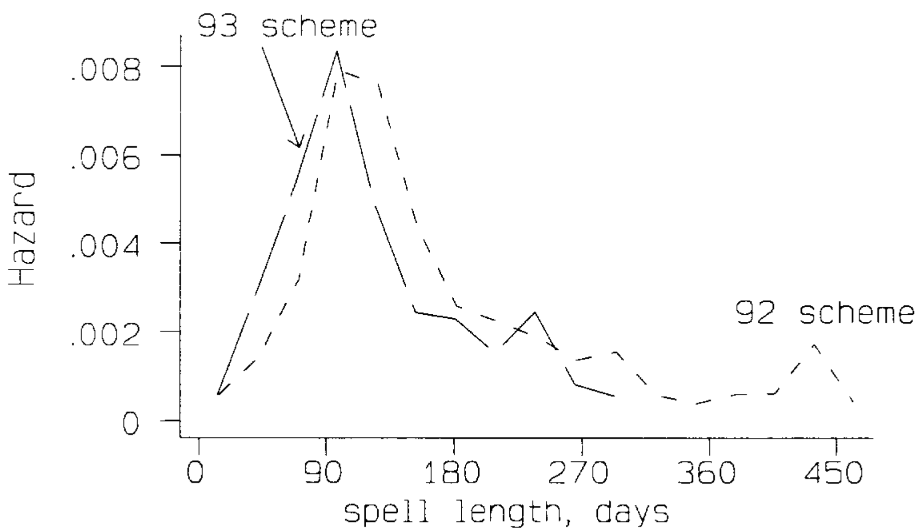
Figure 7 : Transition Rates to Employment of Men likely to be on recall

[a] 44-47 months of work-history



Note: sample sizes 4158 (92 scheme) and 1667 (93 scheme)

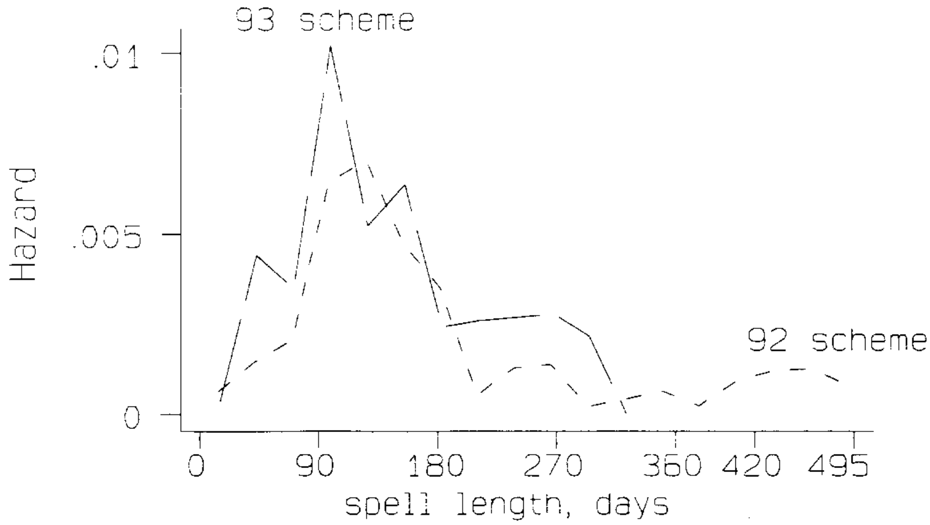
[b] 28-43 months of work-history



Note: sample sizes 1428 (92 scheme) and 689 (93 scheme)

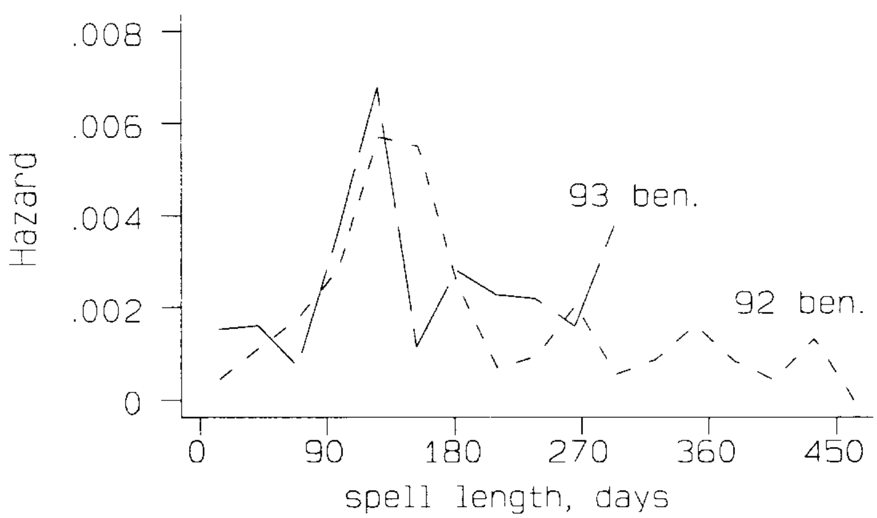
Figure 8 : Transition Rates to Employment of Women likely to be on recall

[a] 44-47 months of work-history



Note: sample sizes 463 (92 scheme) and 114 (93 scheme)

[b] 28-43 months of work-history

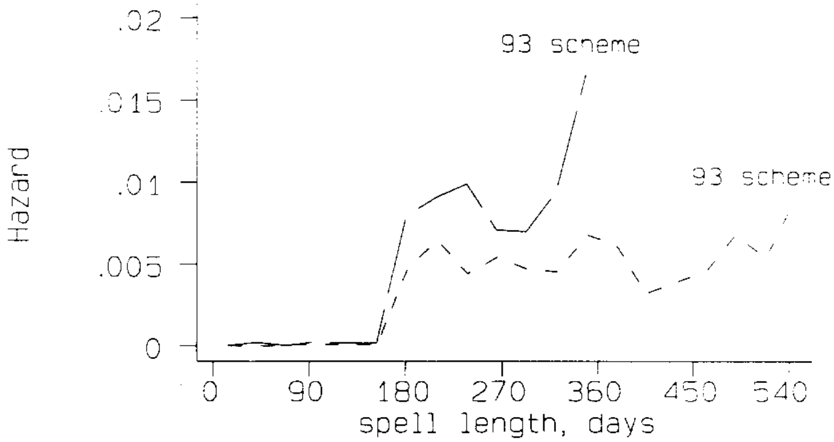


Note: sample sizes 327 (92 scheme) and 167 (93 scheme)

Figure 9 : Transition Rates to Retirement

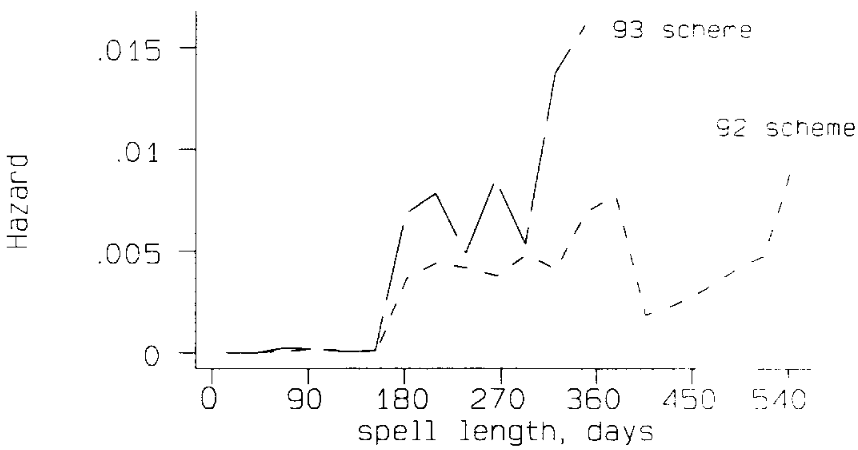
Work-history: 44-47 months or 4 years

(a) Men aged older than 54 years



Note: sample sizes 1497 (92 scheme) and 571 (93 scheme)

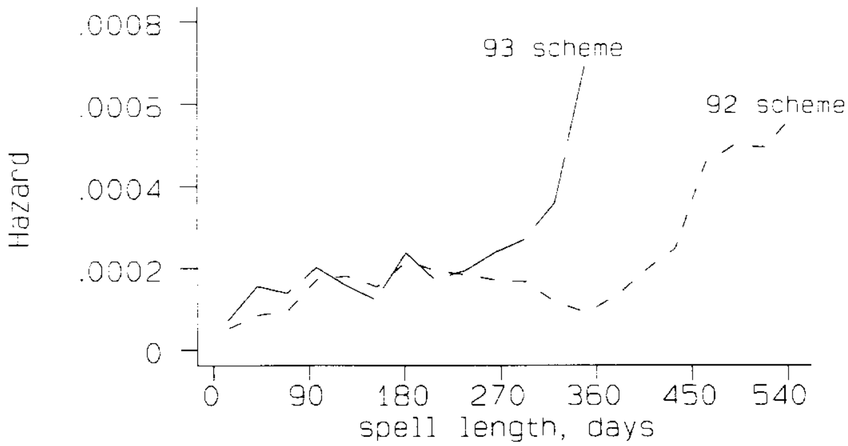
(b) Women aged older than 49 years



Note: sample sizes 1343 (92 scheme) and 575 (93 scheme)

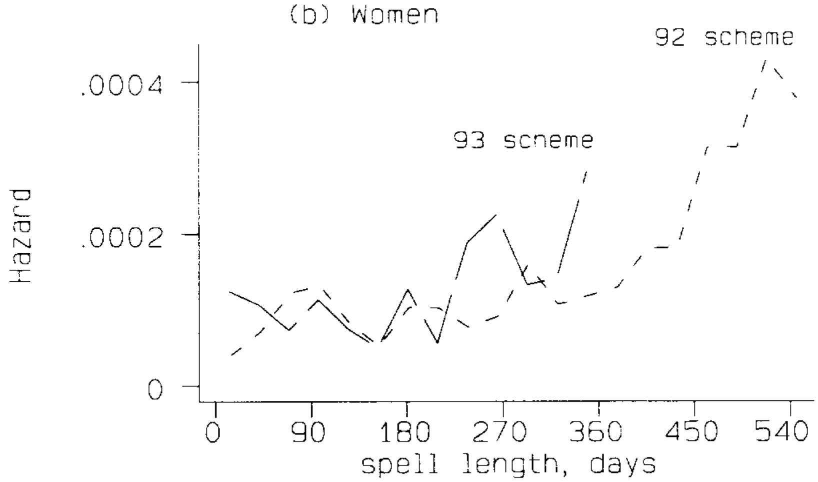
Figure 10 : Transition Rates to Subsidised Employment
Work-history: 44-47 months or 4 years

(a) Men



Note: sample sizes 21022 (92 scheme) and 11456 (93 scheme)

(b) Women



Note sample sizes 8896 (92 scheme) and 4832 (93 scheme)

Figure 11 : Transition Rates to Training
Work-history: 44-47 months or 4 years

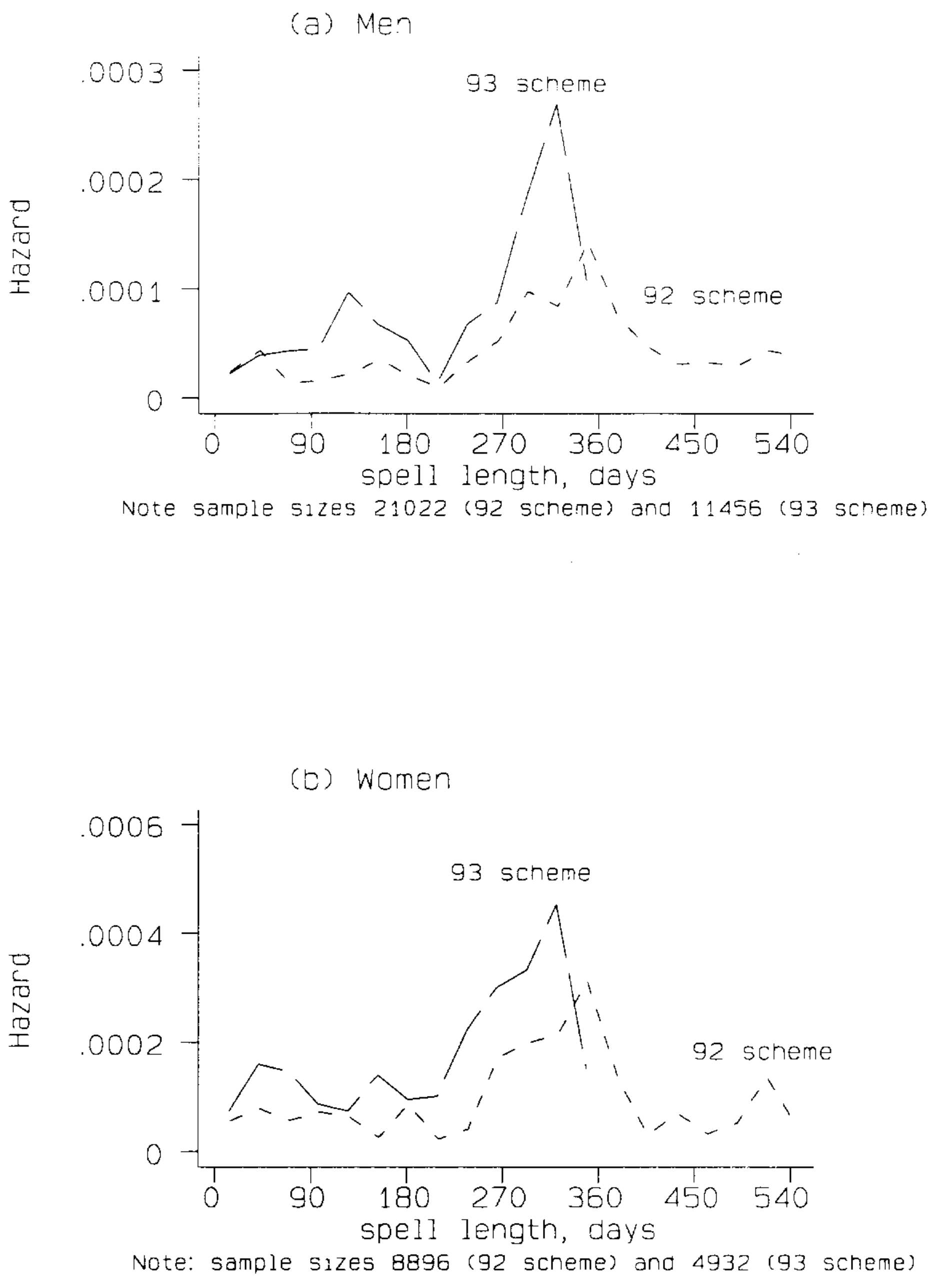


Table A1: Characteristics of the Sample: Proportions in Percent

	Men	Women
Observations	6768	2577
Education		
inc.prim	3.3	2.4
primary	28.4	37.8
vocation	53.2	24.6
voc.sec.	10.5	20.6
gen.sec	2.7	12.0
higher	1.9	2.4
County		
Budapest	3.9	7.8
Baranya	3.4	4.1
Bacs-Kis	4.8	4.9
Bekes	8.0	5.0
Borsod	8.3	5.7
Csongrad	3.2	3.5
Fejer	3.2	4.9
Gyor-Sop	4.9	4.7
Hajdu-Bi	8.4	5.6
Heves	4.5	3.8
Komarom	1.5	2.4
Nograd	3.0	2.8
Pest	5.7	7.8
Somogy	5.3	7.5
Szabolcs	8.8	3.5
Szolnok	5.5	5.6
Tolna	3.3	3.3
Vas	3.4	3.3
Veszprem	6.3	9.1
Zala	4.9	4.6
Age (average)	35.7	35.2
Duration of Unemployment (average in weeks)	26.8	33.9

Table A2: Pre- and Post-Unemployment Earnings

	Men		
	Mean	Median	Std.Dev.
Men			
Prev.Earnings (HUF) (indexed by March 1994)	21662.8	19556.8	10119.3
Post-Unemployment Earnings (HUF)	20733.1	19000.0	9691.2
Women			
Prev.Earnings (HUF) (indexed by March 1994)	16764.0	15133.0	7097.7
Post-Unemployment Earnings (HUF)	16304.0	15000.0	6889.3

Table A3: Difference in Log. Wages and Duration

	Logarithmic Wage Change		
Duration (weeks)	Mean	Std.Dev.	Obs.
Men			
1-10	0.0290	0.381	852
11-20	-0.0487	0.330	3245
21-30	0.0089	0.382	705
31-50	-0.0208	0.458	923
51-100	-0.0800	0.443	999
>100	-0.1664	0.495	44
Total	-0.0345	0.382	6768
Women			
1-10	0.0440	0.349	378
11-20	-0.0181	0.336	697
21-30	0.0157	0.389	394
31-50	-0.0217	0.388	519
51-100	-0.0593	0.386	546
>100	-0.1838	0.574	43
Total	-0.0161	0.374	2577

Table A4: OLS Estimates of Coefficients for Post-Unemployment Wage Gain

Regressors	Men			Women		
	Coefficient	t-value		Coefficient	t-value	
constant	0.203	6.473	**	0.217	3.931	**
ln(age/10)	-0.151	-8.938	**	-0.151	-4.956	**
Education						
inc.prim	-0.051	-1.927		0.007	0.151	
primary	-0.058	-5.362	**	-0.026	-1.36	
voc.sec.	-0.026	-1.674		0.006	0.274	
gen.sec	-0.041	-1.433		-0.023	-0.892	
higher	-0.015	-0.448		-0.004	-0.083	
County						
Budapest	0.175	6.147	**	0.138	3.276	**
Baranya	0.080	2.683	**	0.060	1.232	
Bacs-Kis	0.082	3.089	**	0.109	2.341	*
Borsod	0.077	3.393	**	0.045	1.008	
Csongrad	0.090	2.975	**	0.029	0.558	
Fejer	0.068	2.243	*	0.043	0.921	
Gyor-Sop	0.129	4.908	**	0.068	1.461	
Hajdu-Bi	0.099	4.375	**	0.038	0.842	
Heves	0.099	3.659	**	0.060	1.207	
Komarom	0.055	1.335		0.005	0.082	
Nograd	0.046	1.481		0.047	0.869	
Pest	0.134	5.375	**	0.091	2.177	*
Somogy	0.063	2.467		0.045	1.064	
Szabolcs	0.050	2.243	*	0.178	3.51	**
Szolnok	0.075	2.963	**	0.022	0.485	
Tolna	0.143	4.766	**	0.011	0.214	
Vas	0.135	4.554	**	-0.002	-0.038	
Veszprem	0.026	1.078		0.048	1.191	
Zala	0.070	2.67		0.101	2.137	*
Duration						
ln(duration)	-0.033	-5.54	**	-0.029	-3.492	**
duration of 13 or 14 weeks	-0.063	-4.779	**	-0.096	-2.061	*
Observations	6768			2577		
R^2	0.039			0.034		
adjusted R^2	0.035			0.024		
RESET-Test[1)]	F(3, 6737)=1.71			F(3, 2546)=2.85*		
Cook-Weisberg[2)] for heteroscedasticity	chi2(1)=110.6**			chi2(1)=9.99**		
Joint significance of						
Education	F(5,6740)=6.12**			F(5,2549)=0.75		
Counties	F(19,6740)=4.40**			F(19,2549)=1.87*		

* implies a 5 percent significance level, ** implies a one percent significance level

1) joint significance of coefficients of three powers of the precitions of wage change added to the specification

2) $H_0: \alpha = 0$ in $var(u) = \sigma^2 \exp(\alpha \hat{y})$

Table B1: Maximum Likelihood Estimates of Transition Rate to Employment: Coefficients of Variables not Related to the UI System

	Men								Women							
	Spec. 1		Spec. 2		Spec. 3		Spec. 4		Spec. 1		Spec. 2		Spec. 3		Spec. 4	
Variable	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
(1) Period Specific Effects (Duration Interval)																
0-30 days	-9.48	-19.6	-9.23	-19.1	-8.99	-17.9	-9.17	-18.3	-5.19	-7.3	-5.30	-7.5	-4.23	-5.7	-5.51	-7.1
30-60 days	-8.90	-18.4	-8.48	-17.6	-8.23	-16.5	-8.42	-16.9	-4.87	-6.9	-4.71	-6.7	-3.64	-4.9	-4.92	-6.3
60-90 days	-8.35	-17.4	-7.98	-16.6	-7.74	-15.5	-7.92	-15.9	-4.66	-6.6	-4.49	-6.4	-3.42	-4.6	-4.7	-6.1
90-120 days	-8.16	-17.1	-7.62	-15.9	-7.38	-14.8	-7.56	-15.2	-4.60	-6.5	-4.48	-6.4	-3.41	-4.6	-4.69	-6.1
120-150 days	-8.28	-17.3	-7.71	-16.1	-7.47	-15.0	-7.65	-15.4	-4.65	-6.6	-4.41	-6.3	-3.34	-4.5	-4.61	-6.0
150-180 days	-8.41	-17.6	-7.84	-16.3	-7.59	-15.2	-7.78	-15.6	-4.52	-6.4	-4.34	-6.1	-3.27	-4.4	-4.52	-5.8
180-210 days	-8.46	-17.7	-7.82	-16.3	-7.57	-15.2	-7.76	-15.6	-4.60	-6.5	-4.37	-6.2	-3.29	-4.4	-4.53	-5.9
210-240 days	-8.66	-18.1	-8.15	-17.0	-7.90	-15.8	-8.09	-16.3	-4.73	-6.7	-4.65	-6.6	-3.57	-4.8	-4.8	-6.2
240-270 days	-8.70	-18.2	-8.27	-17.2	-8.03	-16.0	-8.22	-16.5	-4.68	-6.6	-4.51	-6.4	-3.43	-4.6	-4.67	-6.0
270-300 days	-8.61	-18.0	-8.10	-16.8	-7.85	-15.7	-8.04	-16.2	-4.79	-6.7	-4.56	-6.4	-3.48	-4.7	-4.7	-6.1
300-330 days	-8.83	-18.4	-8.34	-17.3	-8.09	-16.2	-8.28	-16.7	-4.79	-6.7	-4.64	-6.5	-3.56	-4.8	-4.77	-6.1
330-360 days	-9.17	-19.0	-8.69	-18.1	-8.44	-16.9	-8.64	-17.4	-5.07	-7.1	-4.92	-6.9	-3.84	-5.1	-5.04	-6.5
360-405 days	-9.18	-19.1	-8.71	-18.2	-8.46	-16.9	-8.66	-17.5	-5.06	-7.1	-4.92	-6.9	-3.83	-5.1	-5.01	-6.5
405-450 days	-8.74	-18.3	-8.26	-17.3	-8.01	-16.1	-8.21	-16.7	-4.70	-6.6	-4.55	-6.4	-3.45	-4.6	-4.61	-6.0
> 450 days	-8.53	-17.7	-8.06	-16.7	-7.81	-15.6	-7.99	-16.1	-4.39	-6.1	-4.25	-5.9	-3.15	-4.2	-4.25	-5.5
(2) Education																
incomplete primary	-0.23	-3.6	-0.23	-3.6	-0.23	-3.6	-0.23	-3.6	-0.42	-4.0	-0.42	-4.0	-0.42	-3.9	-0.43	-4.1
vocational	0.35	13.9	0.35	13.7	0.35	13.7	0.34	13.6	0.13	3.0	0.13	3.0	0.13	3.2	0.133	3.2
vocational secondary	0.30	7.3	0.30	7.3	0.30	7.3	0.30	7.3	0.26	4.5	0.26	4.5	0.28	4.8	0.272	4.7
general secondary	0.19	3.0	0.19	2.9	0.19	3.0	0.19	2.9	0.11	1.8	0.11	1.8	0.13	2.1	0.118	1.9
college/university	0.58	7.4	0.58	7.4	0.58	7.4	0.57	7.3	0.44	4.2	0.44	4.2	0.48	4.5	0.472	4.5
(3) Regional Activity and Unempl. Rate																
agricult., <12 %	0.35	6.2	0.34	6.2	0.34	6.2	0.34	6.2	0.61	7.7	0.60	7.6	0.61	7.7	0.604	7.6
agrricult, >18 %	-0.26	-5.8	-0.26	-5.8	-0.26	-5.8	-0.26	-5.8	-0.28	-3.8	-0.28	-3.8	-0.29	-3.9	-0.28	-3.8
industry, <12 %	0.33	8.5	0.33	8.5	0.33	8.5	0.33	8.5	0.35	6.0	0.35	6.0	0.36	6.1	0.358	6.1
industry, 12-18 %	0.09	1.5	0.09	1.5	0.09	1.5	0.09	1.5	-0.15	-1.3	-0.14	-1.3	-0.14	-1.3	-0.14	-1.3
industry, >18 %	-0.13	-3.2	-0.13	-3.2	-0.13	-3.2	-0.13	-3.2	-0.36	-4.6	-0.36	-4.6	-0.35	-4.5	-0.35	-4.5
diversified, <12 %	0.06	2.2	0.06	2.0	0.06	2.0	0.06	2.0	0.01	0.3	0.01	0.3	0.02	0.5	0.028	0.6
diversified, 12-18 %	0.06	1.4	0.06	1.4	0.06	1.4	0.06	1.4	-0.10	-1.4	-0.10	-1.4	-0.10	-1.3	-0.1	-1.4
(4) Other																
age/10	-0.18	-2.1	-0.16	-1.9	-0.29	-2.6	-0.17	-2.0	-1.30	-8.7	-1.29	-8.7	-1.81	-10.0	-1.46	-7.7
$(age/10)^2$	0.00	0.2	0.00	0.1	0.01	1.0	0.00	0.1	0.17	7.5	0.17	7.4	0.22	8.8	0.178	6.9
previous UI spell	0.42	19.0	0.42	18.9	0.42	18.9	0.42	17.8	0.60	16.5	0.60	16.4	0.61	16.6	0.536	13.9
inflow from employment	0.44	10.3	0.47	10.9	0.46	10.9	0.47	10.7	0.46	7.5	0.47	7.7	0.46	7.5	0.535	8.6
non-manual	-0.26	-4.9	-0.25	-4.9	-0.25	-4.8	-0.25	-4.9	-0.06	-1.3	-0.07	-1.3	-0.06	-1.2	-0.05	-0.9

Table B2: Maximum Likelihood Estimates for Transition Rates to Subsidised Employment and Training: Coefficients of Variables not Related to the UI System

Variable	Men Transition Rates to						Women Transition Rates to					
	Subsidised Employment		Training				Subsidised Employment		Training			
			Spec. 1		Spec. 2				Spec. 1		Spec. 2	
	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
(1) Period Specific Effects (Duration Interval)												
0-45 days	-8.98	-8.4	-13.71	-8.1	-13.40	-7.4	-8.38	-5.1	-13.54	-7.9	-13.06	-7.1
45-90 days	-8.33	-7.9	-13.68	-8.1	-13.37	-7.4	-7.90	-4.9	-13.37	-7.8	-12.88	-7.0
90-135 days	-7.82	-7.4	-13.54	-8.0	-13.22	-7.3	-8.14	-5.0	-13.45	-7.9	-12.97	-7.1
135-180 days	-8.11	-7.7	-13.38	-7.9	-13.07	-7.2	-8.84	-5.4	-13.54	-7.9	-13.05	-7.1
180-210 days	-7.75	-7.3					-7.79	-4.8				
210-240 days	-8.01	-7.5					-8.36	-5.1				
240-270 days	-7.85	-7.4					-7.90	-4.8				
180-225 days			-14.39	-8.4	-14.07	-7.7			-13.92	-8.1	-13.42	-7.3
225-170 days			-13.47	-7.9	-13.15	-7.2			-13.13	-7.6	-12.64	-6.8
270-315 days	-7.88	-7.4	-12.64	-7.4	-12.32	-6.8	-7.74	-4.7	-12.32	-7.1	-11.83	-6.4
315-360 days	-7.99	-7.5	-12.60	-7.4	-12.28	-6.7	-7.99	-4.9	-12.32	-7.1	-11.83	-6.4
360-405 days	-7.82	-7.4	-13.26	-7.8	-12.94	-7.1	-7.75	-4.7	-13.13	-7.6	-12.64	-6.8
405-450 days	-7.40	-7.0	-13.75	-8.0	-13.42	-7.3	-7.51	-4.6	-13.83	-7.9	-13.33	-7.1
450-495 days	-7.19	-6.8	-14.16	-8.2	-13.83	-7.5	-7.16	-4.4	-14.01	-8.0	-13.52	-7.2
> 495 days	-7.00	-6.6	-13.54	-7.8	-13.21	-7.2	-7.00	-4.2	-13.10	-7.5	-12.61	-6.7
(2) Education												
vocational	0.34	6.3	0.09	0.8	0.09	0.8	0.46	4.5	0.02	0.1	0.01	0.1
vocational or general sec.	0.45	5.8	1.52	13.0	1.53	13.0	0.69	6.3	1.80	15.2	1.80	15.1
college/university	0.85	6.7	2.20	13.0	2.20	13.0	1.16	5.9	2.25	13.4	2.24	13.2
(3) Regional Activity and Unempl. Rate												
agricult., <12 %	0.10	0.7					-0.21	-0.9				
agrricult, >18 %	-0.33	-4.3					-0.14	-1.1				
industry, <12 %	-0.90	-6.4					-1.08	-4.9				
industry, 12-18 %	0.89	9.8					0.53	3.8				
industry, >18 %	0.07	0.9					-0.37	-2.9				
diversified, <12 %	-0.21	-3.5					-0.64	-6.1				
diversified, 12-18 %	-0.34	-3.5					-0.38	-2.6				
< 12%			0.38	4.4	0.38	4.5			0.58	6.5	0.57	6.4
>18%			-0.32	-2.6	-0.32	-2.6			0.06	0.5	0.06	0.5
(4) Other												
age/10	0.51	2.5	-0.49	-1.4	-0.70	-1.5	0.62	1.4	-1.29	-3.3	-1.48	-3.1
$(age/10)^2$	-0.08	-2.8	-0.01	-0.1	0.02	0.3	-0.08	-1.4	0.12	2.1	0.14	2.2
previous UI spell	0.59	11.5	0.14	1.4	0.14	1.5	0.30	2.7	0.12	1.1	0.12	1.1
inflow from employment	0.30	2.1	0.02	0.1	0.01	0.0	0.12	0.7	0.25	1.6	0.24	1.6
non-manual	0.34	4.1	0.13	1.2	0.13	1.2	0.22	2.2	0.15	1.6	0.15	1.6