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The Hungarian Unemployment Insurance Benefit System and Incentives to Return to Work*

Joachim Wolff
Department of Economics
University of Munich
Germany
email: Joachim.Wolff@lrz.uni-muenchen.de

Abstract

This paper analyses the impact of the Hungarian unemployment insurance (UI) benefit system on the speed of exit from unemployment to regular employment. The duration analysis relies on unemployment spells from two inflow cohorts, which are administered under distinct UI rules. Thus, it exploits a natural experiment to identify disincentive effects. Kaplan-Meier estimates suggest that the benefit reform did not significantly change the transition rates. Moreover, a semi-parametric analysis cannot find remarkable disincentive effects but an entitlement effect. The hazards of men and women rise somewhat in the last two months before they run out of UI benefit.

JEL classification: J64,J65 and C23

Keywords: Unemployment Duration, Hazard Rate, Natural Experiment, Workers on Recall

1. Introduction

During the first years of the reform process, high unemployment and a very low turnover of the unemployment pool distinguished many transition economies from the large western OECD countries [Boeri, 1994, 1996]. The consequence of this low turnover was high long-

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term unemployment. Reducing the generosity of the unemployment benefit system is one possible policy to speed up the exit from unemployment. However, in economies with low labour demand such a policy may not alter outflow from unemployment considerably. For example Arulampalam and Stewart (1995) analysed for the United Kingdom the effects of unemployment benefits on the conditional probability of leaving unemployment. They estimated separately the effects of unemployment benefits for one sample of unemployment spells drawn during a period of low unemployment and one for a high unemployment period. They found the elasticity of the transition rates to the income while unemployed to be negative in both cases. However, in the high unemployment era this elasticity was insignificant and at a value of –0.11 less than a third of the value of the low unemployment era. Thus, in that period, even cutting benefit levels by 20 percent would have raised the transition rates by just by about two percent.

So for high unemployment countries, such as many transition economies, we should expect a low impact of unemployment benefits on the exit rates from unemployment. This paper studies this hypothesis for the case of Hungary. However, it is still important to know whether turning to a less generous benefit system is an appropriate policy to increase the outflow from unemployment when unemployment is already high.

An early discussion of the Hungarian unemployment insurance (UI) benefit system and incentives to leave (insured) unemployment has been carried out by Micklewright and Nagy (1995 a,b). This analysis of the duration of unemployment relied on a natural experiment and hence applied a treatment group approach: They focused on differences of the speed of exit from unemployment of two groups of people. The control group received UI according to the benefit rules of the year 1992, whereas the other group, the treatment 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. This microeconometric analysis works with the same data set and substantially extends the previous work of Micklewright and Nagy.

One of the biggest problems that Micklewright and Nagy faced in their non-parametric work emerged with the specific slope of their estimated job hazards. They were characterised by very large spikes at an elapsed duration of unemployment of about three months. The authors suggested that the cause could be a high share of workers on recall. However, the share of workers on recall among the 1992 claimants and the 1993 claimants was unknown. Hence, the hazards of the two claimant samples may differ due to a different share of workers on recall rather than due to the distinct benefit rules. In this paper, therefore, first unemployed people who are less and more likely to be on recall are identified. Then, by adopting a treatment group approach, I study whether the benefit reform of 1993 altered the speed of return for these two types of workers separately.

Apart from evaluating the UI reform by a treatment group approach, further insights can be reached about the benefit system's impact on the return to work. In particular, it is investigated how the UI benefit level relative to that of earnings as well as the remaining entitlement to UI benefits alter the individual transition rates to jobs. These are the most important characteristics of UI receipt, which explain unemployment duration. Therefore Micklewright and Nagy's analysis is extended by estimating a semi-parametric duration model, which controls for the impacts of the benefit system alongside other observed heterogeneity.

The rest of the paper is organised as follows: Section 2 presents the econometric continuous-time duration methods that are applied to the data. Section 3 introduces the Hungarian UI benefit system, highlighting its features before and after the benefit reform of January 1993. In Section 4, the microdata is presented. The advantages and disadvantages of this inflow sample of job-losers to the Hungarian unemployment register are discussed for the analysis that follows. Section 5, having discussed the results of Micklewright and Nagy, then deals with the issue of workers on recall. As they are not similarly distributed over the 1992 and 1993 claimant samples, the sample is divided into workers who are more likely to be on recall and those who are less likely to be on recall. Then the reform of 1993 is evaluated for those workers who are less likely to be on recall by comparing the Kaplan-Meier estimates of the job hazards of 1992 claimants with those of 1993 claimants. Section 6 analyses the effects of the UI benefit levels and time until exhaustion of UI on the job hazards within a semi-parametric framework. First, problems with the specification of these variables, and with the identification of their parameters in the econometric analysis that follows are discussed. The section finishes by presenting the results of Maximum Likelihood estimation of piece-wise constant exponential models for transitions into employment. Section 7 concludes.

2. Econometric Methods for Analysing Continuous-Time Duration Data

Let T be the continuous random variable duration of unemployment. 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 a labour force state s provided that no exit has occurred before:

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

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

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

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

Let me derive the likelihood function for a semi-parametric continuous duration model: the hazards are specified by an *exponential model with piece-wise constant terms*[1]. First, a set of M time intervals needs to be defined. Let I_j denote a time interval where I_j is

$$I_j = [t | v_j \le t < v_{j+1}], \qquad j=1,...,M \tag{3},$$

and $v_1,...,v_M$ 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:

$$\theta_i^s(t) = \exp(\alpha_j^s) \cdot \exp\left[\mathbf{x}_i^{s'}(t) \cdot \beta^s\right] \tag{4},$$

where t belongs to I_j. $\mathbf{x}_i^s(t)$ represents a vector of explanatory variables. $\boldsymbol{\beta}^s$ is an unknown parameter vector, which will be estimated simultaneously with the α_j (j=1,...,M). The latter are 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.

Let a = 1,...,A be the individuals in the sample with completed spells of unemployment insurance of length T_a who exit to state regular employment. Further, let b = 1,...,B be the individuals whose spells are right-censored at length T_b. These are either right-censored spells for which no exit is reported since the individuals exhausted their UI benefits, or spells with a transition to an exit state that is not modelled. The latter are treated as if they were right-censored. The likelihood function of this sample is:

$$L = \prod_a [\theta_i^a(T_a) \cdot S(T_a)] \cdot \prod_b S(T_b) \tag{5}$$

The contribution to the likelihood function of spells that finish with regular employment is thus their state specific transition rate times the survivor function evaluated at their spell end. All the other spells are represented by the survivor function alone. So the parameters of the employment transition rates are estimated by the Maximum Likelihood (ML) of a single risk model[2].

3. The Hungarian UI Benefit System and the UI Reform of 1993

In 1989 Hungary introduced a system of unemployment compensation and established an insurance-type unemployment benefit system in February 1991. The institution responsible for UI benefits is called the Solidarity Fund. Its funds are provided by contributions from employers and employees, while its deficits are covered by the state budget. Because of high

[1] For a discussion of the exponential model with piece-wise constant terms see Lancaster (1990).
[2] ML-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 6.1, 1997).

deficits it was reformed and made more restrictive in the years 1992 and 1993[3]. One of the most important features for the present analysis is that the Hungarian UI system is "grandfathered"; i.e. people receive UI according to the benefit rules that were in place at the time of their benefit claim. Thus their claims are unaffected by any subsequent changes in the rules, that affect new claims only.

Now turn to the main features of the UI system during the period under review[4]: these are the eligibility rules, the length of entitlement, and the UI benefits levels. To be eligible for UI, unemployed people must have worked in contributory employment for at least 12 months during the four years prior to the benefit claim[5]. Additionally, they are supposed to actively search 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 UI claim determines the length of entitlement. In all, there are ten different entitlement periods. Under the 1992 rules they ranged from 135 days to 540 days. Table 1 displays all potential entitlement periods and shows that the reform of 1993 reduced them by one third.

Each entitlement period is divided into two parts. A higher formal replacement rate applies to the first of these two parts. The more generous first part accounted for two thirds of the total length of entitlement under the 1992 benefit scheme but only for one quarter under the scheme of 1993. The UI benefit levels are determined by an earnings-related benefit formula. Earnings are defined as the average monthly earnings during the last four quarters before the UI claim. Table 1 shows 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 maximum and minimum benefit regulation. The maximum benefit was set to twice the minimum wage under the 1992 benefit provisions. The same is true for the first phase of UI entitlement under the 1993 benefit regime, although 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. Note that, the minimum benefit is not a lower cap. If past earnings were below the minimum benefit, a recipient's benefit equals these past earnings. Both the 1992 and 1993 benefit rules did not index benefits to inflation. Taken together, the modifications result in a much less generous UI system since January 1993. The main reason is, that according to the new rules UI benefit is paid for a much shorter period of time.

[3] The deficit was 35 billion HUF in 1992; i.e. more than 50 percent of the Solidarity Fund's expenditure (Viszt and Ványai, 1994).

[4] Micklewright and Nagy (1994) provide a detailed discussion of the Hungarian UI system and its reforms.

People, who exhaust their UI benefits, or who are recurrently unemployed and no longer eligible for 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 (Micklewright and Nagy, 1995b: 6). The monthly SB is set at 80 percent of the minimum pension. Thus, even if somebody passes the means test, the SB will, in most cases, be much lower than UI benefits. Micklewright and 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 was more than 40 percent higher than the SB in more than 90 percent of the cases.

4. Hungarian UI Register Data

As a consequence of the grandfathering of the UI benefit system, benefit reforms in Hungary give rise to quasi-experimental situations. Unemployment spells that start immediately before and after a reform are administered under two different benefit regimes, but face the same economic conditions. The reform that came into force on the first of January 1993 led to such a situation. Therefore, a data set is studied that represents the entire inflow to the UI register in December 1992 and January 1993. This data 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]. A few spells with a missing work-history record or implausible values of their entitlement periods have also been discarded. This leaves 54,901 male and 25,195 female UI spells. About 37.4 percent of men and 38.1 percent of women are administered by the 1993 benefit provisions[7].

There are various advantages to using this data apart from the fact that it allows a treatment groups approach in order to analyse the impact of the UI system. Duration of unemployment is measured very exactly, in days. Next, systematic measurement errors of this variable and the exit state, as often found in retrospective surveys, cannot emerge. The size of the data set is considerable; this is important since a high number of transitions is necessary to reach precise

[5] School-leavers were excluded from this requirement; in 1992 they were automatically eligible for six months of UI benefit.

[6] The exact sample selection criteria are described in detail in Micklewright and Nagy (1995a). They discarded 19,681 spells from the sample; this number corresponds to 19.6 percent of the entire inflow sample. These discarded spells are mainly UI claimants who quit their last job or received statutory severance pay after the job-loss. Their UI payment only starts after a waiting period. There is one important reason for discarding them: spells to which apply the 1992 benefit rules should not be included in the January 1993 sample.

[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. Micklewright and Nagy (1995a) suggested a variety of reasons why this was not the case: Firstly, the sample consists of job-losers who cannot entirely control the date of their job-losses. Secondly, the date 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.

estimates of the hazards and of the coefficients of their determinants. Finally, the data consists of spells from two different benefit regimes, so that it is rich in terms of variation of replacement rates and UI entitlement periods. This issue will be discussed later in relation to the subsample of spells with which the impacts of the UI benefit and remaining entitlement to UI are estimated.

However, there are a number of disadvantages to using this UI register data: First, spells are observed only until the expiry of the UI receipt. Thus, nothing be can inferred about the exit rates after that date. Second, there is no information on marital status, household income or composition, which determine whether people receive SB after running out of UI, and also determine an individual's sources of income that are additional to the UI benefits. How the UI receipt affects an unemployed person's job hazards is likely to depend on both. All that can be controlled for is age, education, region, time since the last UI spell ended, entry to UI from employment, whether a person is a non-manual worker and work-history. Next, the spells range from the end of 1992 and beginning of 1993 at most until June 1994. Therefore changes of the hazards over the spell length may reflect duration dependence as well as variations in labour demand over this specific period of time. Thus, seasonal labour demand fluctuations as well as the upturn of the Hungarian labour market in early 1993 characterise the baseline escape rates. It is also important to note, that the spells are observed during the period in which registered unemployment was highest. It peaked in spring 1993 at 14 percent and fell to 11 percent until June 1994. For this reason, the effects of the benefit system can be evaluated only for a situation of a very high, though decreasing, unemployment rate. Low labour demand may interact with the impact of the UI system on the speed of return to work. The reason is that there is not much scope for unemployed people to increase the arrival rate of job offers by searching more intensively or by reducing the reservation wages. This in turn would suggest that no large effect of the UI system on the job hazards may emerge.

The duration analysis that follows only deals with one exit state: employment. The exit state and duration form the dependent variables in this study. No statistics on the duration of spells are presented, as they would not be very indicative due to a high number of right-censored spells. Table 2 displays the shares of completed UI spells by destination as well as right-censored spells in order to distinguish their importance.

The third row of Table 2 shows that 48.4 percent of unemployed men and 33.7 percent of unemployed women find a regular job before their UI entitlement ends. Unemployed people are much less frequently directed to other labour force states. Taken together, 12.6 percent of men and nearly 19 percent of women exit to these states. What is not displayed in this table is the share of (early) retirement exits. Men who are at least 55 years old and women who are at least 50 years old [8] have reached or are about to reach the age limit for early retirement during

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

their UI receipt, which is 57 years for men and 52 years for women. For them, retirement is the dominant exit state: more than 60 percent of these men and women retire. Consequently, for aged people jobs are not an important exit state. Therefore these people are excluded from the analysis of the transitions into employment. Finally, the importance of exhausting UI benefits should not be ignored: at 43.5 percent, this is the most frequent way to terminate a UI spell for women. The share of UI exhausters among men, at 35.1 percent, is also remarkable.

At this stage it is important to stress that the data does not provide the UI benefit received by the UI claimants. As far as benefits are used, they are gross benefits that have been imputed according to the benefit formula: The minimum wage, which is important for this calculation, is set at the level of February 1993 (9000 HUF). Changes in nominal benefits over the spell are only allowed for when they are due to the start of the second entitlement period[9]. The data provides information on past gross wages and past indexed gross wages. The latter is indexed for wage inflation until January 1993.

5. Previous Results and Workers on Recall

5.1 The UI Reform of 1993 and Transitions into Employment: Previous Results

A non-parametric analysis of this Hungarian UI register data of UI spells has already been carried out by Micklewright and Nagy (1995a)[10]. Their aim was to reveal disincentives within the Hungarian UI 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. As far as the speed of return to work is considered, the following results emerged:

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 with four years of work-history. For a short period the 1993 claimants were found to have far higher job hazards than the 1992 claimants. This difference occurred in the form of some large spikes, at an elapsed duration of about three months. This spell length coincides in calendar time with early spring 1993. As a possible explanation for the spikes, the authors suggested that the share of workers on recall is higher in the sample of 1993 claimants than that of 1992 claimants. Note that their analysis discovered similar spikes for other work-history groups.
3. The authors found a small rise in the job hazards near the time when the UI benefits expire.

[9] 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. Note that this adjustment is not necessary for 1993 claimants as the benefit rules are no longer related to the minimum wage (Micklewright and Nagy, 1994b: 21).

[10] "Unemployment insurance and incentives in Hungary" in Newbery (ed.) (1995): Tax and benefit reform in Central.

5.2 Workers on Recall and their Consequences for the Analysis

The central problem that emerged in Micklewright and Nagy's analysis is the presumable presence of workers on recall. They return to work much faster than other unemployed workers. Suppose that the proportion of workers on recall is high in the 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. Under such conditions a treatment group approach for evaluating the UI reform would not be appropriate.

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. Then a typical inflow cohort to unemployment is made up by a small share of workers on recall. However, this may be different for the specific inflow cohorts of December 1992 and January 1993. In these months, important reasons for labour-shedding are seasonal demand fluctuations. By further inspection Micklewright and Nagy's analysis provide some evidence that the share of workers on recall in the sample is considerable:

- Labour shedding due to seasonal reasons in these inflow cohorts should lead to spikes at a spell length that coincides with a seasonal upturn, i.e. the late winter/early spring period in 1993. This is the only time during which Micklewright and Nagy found unemployed workers to return to work at a remarkable speed. A counter argument is that a general upturn in the Hungarian labour market started in this period. Yet, this argument cannot explain why the job hazards remained high only very briefly.
- The share of workers on recall should be higher for men than for women, because jobs in which workers are made redundant temporarily for seasonal reasons are, by and large, male occupations. That the spikes in jobs hazards are found 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 an attribute of workers who were made redundant temporarily for reasons of seasonal labour demand fluctuations. However, this could also be in line with other reasons for temporary breaks in the work-history.

All this suggests that the sample, and in particular the 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 I 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. Therefore, they return to work much faster than other workers.

Clearly some light needs to be shed on this issue. First of all, the existence of a large share of workers on recall, which are not directly observable in the data, has to be proved. Therefore an exogenous and available variable that is closely related to the recall characteristic is

needed. Let me argue that people who lost their job because of seasonal fluctuations in labour demand have one common characteristic: they are very frequently unemployed for a short period at the same calendar time of a year. This is a useful feature, as the data set provides information on the number of days that have gone past between the end of the last and the start of the current UI spell. 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, the sample is partitioned into two categories of unemployed workers:

(1) unemployed people with *no prior UI spell or a prior UI spell that did not end in mid winter/early spring.* This group of people should be characterised by a low proportion of workers on recall. They are here labelled as *less likely 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[11]. They most likely found their last job during a seasonal upturn. Therefore, they are labelled as *more likely to be on recall*.

First, it will be demonstrated that only 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. Their job hazards are compared to those of the first group. This is done in Figure 1 a and b for males who belong to the work-history group 44-47 months and all men with a shorter employment record. It plots the (four-weekly) Kaplan-Meier estimates of their job hazards against duration. There is one outstanding feature of this figure. An immense temporary rise of the hazard at around three months of elapsed duration of the second group of workers who are more likely to be on recall is observed. The duration, at which this spike occurs, 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 is negligible. The corresponding figures for women are not displayed here[12] but suggest no different conclusion. The evidence from these first two figures is strongly in favour of the hypothesis that the sample consists of two types of workers that differ extremely in the response of their hazards to changes in seasonal labour demand.

The proportion of workers in the sample who are more likely to be on recall is substantial. Table 3 shows the composition of the 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[13]. The subsample then consists of 31,909 men aged younger than 55 and 12,428

[11] Information about the days since the end of the last UI spell is only available over the last four years prior to the current UI spell. The results suggest that the hazards of workers whose last spell ended in mid winter/early spring during the previous two years are affected considerably by seasonal labour demand fluctuations. This is not so for workers who meet the same criteria for three or four years prior to the current spell.

[12] They are available on request.

[13] The information of the last spell end is only available for the four years prior to the start of the spells in our data. So, at first sight, it is puzzling that a past UI spell is recorded for any of the workers with an employment record of four years. An explanation may be that the exact definition of this

women below the age of 50. 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. For this work-history group as well for those with 28-43 months of employment record, the share of workers on recall is substantially higher among 1992 claimants than among 1993 claimants.

To sum up, it has been illustrated that this sample consists of two types of unemployed workers with a very different response of their job hazards to seasonal employment fluctuations. The 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. So, the benefit reform's impact on the job hazards for workers who are more and less likely to be on recall has to be studied separately. This leads to dismissing all spells with an employment record of four years. The indicator variable that identifies workers on recall cannot be calculated for these spells, since by definition, no information on previous UI spells during the four years prior to their current spell is available.

We will only analyse the impact of the benefit reform on those workers who are less likely to be on recall. They presumably lost their jobs as a consequence of economic restructuring and not because of seasonal demand fluctuations. In an economy in transition, they are the unemployed workers whose exit behaviour is of most interest. Since it is impossible to identify these workers among the work-history group of four years, the modal group, a large number of observations have been discarded.

Before turning to the analysis, one issue still needs to be discussed. There would ideally exist a similar distribution of observed characteristics over the two claimant groups. The reason for this is that I want to rule out as far as possible that their hazards differ due to some other heterogeneity. Table 4 shows the distribution of observed characteristics by benefit scheme and gender for the subsample workers who are less likely to be on recall. The attributes are mainly education, region, age and work-history. So let me discuss whether they are similarly distributed over the two claimant groups.

Table 4 shows, that with respect to *education, age* and *work-history* the compositions of 1992 and 1993 claimants are strikingly similar for the male and the female sample alike. Only with respect to *region* do some differences emerge. The table distinguishes between the

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work-history plays a role. Four years of work-history could be interpreted as an employment record that is longer than 47 months. In turn, a few people in this group may have experienced a very short UI spell during the last four years prior to their current spell.

dominant regional activities in 1993[14]. For both gender the share of the 1992 claimants who live in agricultural regions is about eight percentage points higher than that of 1993 claimants. While the proportion of people living in industrial regions is similar for both claimant groups, 1992 claimants less frequently come from diversified regions. However, the differences are not outstanding.

5.3 UI Benefits and Incentives to Return to Work: A Treatment Group Approach

Let Y_{it} and Y_{it}^{*} be two outcome variables, where the first is an outcome with treatment under the 1993 benefit rules and the second the corresponding outcome under the 1992 benefit rules. Suppose the outcome variable represents the log of the hazards of an individual (i) for a duration of length t. The mean impact of treatment on an individual assuming that it is a constant for all individuals is

$$\gamma_t = E\left(Y_{it} \middle| D_i = 1\right) - E\left(Y_{it}^{*} \middle| D_i = 1\right) \qquad (6),$$

where $D_i = 1$ represents a 1993 claimant, the treated, $D_i = 0$ will identify a 1992 claimant. An individual's mean impact cannot be measured this way since $Y_{it}^{*}|D_i = 1$ is unobserved. Let the expected outcome of 1992 claimants be $E(Y_{it}^{*}|D_i = 0)$. So, the difference between 1993 claimants' and 1992 claimants' expected outcomes is

$$E\left(Y_{it} \middle| D_i = 1\right) - E\left(Y_{it}^{*} \middle| D_i = 0\right) = \gamma_t + \left\{E\left(Y_{it}^{*} \middle| D_i = 1\right) - E\left(Y_{it}^{*} \middle| D_i = 0\right)\right\} \qquad (7),$$

where the expression in brackets on the right-hand side of the equation represents the selection-bias (Heckman and Hotz, 1989). In the case of random assignment into the two benefit regimes this bias equals zero. The assignment rule to the 1992 and 1993 benefit rules is due to the UI claim and so the start of the UI spell in December 1992 or January 1993. The critical assumption for the following analysis is that the observed selection variable, the start of the spell is not correlated with unobserved characteristics that determine the outcome. Since the sample represents people who were made redundant, the start of their UI spell should not be under their control. However, also keep in mind that the assumption of the treatment effect being the same for all individuals is unrealistic. This is particularity true because the benefit reform of 1993 did not change the rules in the same way for all individuals, at least as far as the UI benefit level is considered.

Let me turn to analysing the benefit reform's impact on the return to work by a treatment group approach[15]. I compare the Kaplan-Meier estimates of the hazards of the 1992 and 1993

[14] This division has been chosen following an analysis by Scarpetta (1995) and Scarpetta and Huber (1995): their division into regional activity relies on an analysis of employment shares.

[15] A critique of this treatment group approach may be that the treatment and control groups are still too heterogeneous and this heterogeneity may interact with the change of the benefit system. Because of this the effects of the benefit reform may not be identified properly. So, I also compared the job

claimant groups for workers who are less likely to be on recall. This analysis further distinguishes between men and women and the work-history groups: (a) 44-47 months, (b) 28-43 months and (c) 12-27 months. Each of the cases (b) and (c) aggregates three observable work-history groups. In both cases each claimant group aggregates spells with three distinct entitlement periods, as the work-history determines the length of UI receipt. The job hazards are estimated as the midpoints of four-weekly intervals.

Men's transition rates into jobs are displayed in Figure 2 a-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 during the first three months after the spells start. However, this rise is far smaller than for the group of workers that are more likely to be on recall as discussed earlier. For all three work-history groups 1993 claimant men exit more rapidly to employment than do 1992 claimant men during these first three months. After an elapsed duration of more than three months, the 1992 claimants tend to exit relatively faster to employment. 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. Hence, 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.

Figures 3 a-c shows the same comparison for women. Their estimated hazards are lower than those of men. Most of them are not larger than 0.0025 and not below 0.0005. 1993 claimants 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 return to work is strongest for the work-history group 28-43 months: Most of the estimated job hazards of 1993 claimants exceed those of 1992 claimants. Yet, the difference between their hazards is not large and according to the confidence bands also not significant. In contrast, this analysis did not find that the benefit reform raised the escape rates to jobs of women with 12-27 months of work-history, as shown in Figure 3c.

On the whole this set of figures did not reveal any strong disincentive effects and therefore it cannot be concluded that the introduction of the new benefit rules in January 1993 accelerated the speed of exit to work. Next, in a dynamic search theoretic framework, one would expect that a limited duration of benefit receipt leads to hazards, which rise the closer the unemployment spells get to the date of UI expiry. None of the Figures 2 and 3 shows any strong rise in the hazards just prior to UI benefit exhaustion. Thus, the evidence for an

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hazards of some relatively homogeneous groups of unemployed workers. This analysis, which is not presented in this paper, lead to no different conclusions. The results are available on request.

entitlement effect is weak. However, this may be due to labour demand fluctuations over time, since the spell length in these figures coincides with calendar time.

6. UI Benefits and Incentives to Return to Work: A Semi-Parametric Analysis

6.1 General Considerations and the Specification of the Benefit System

The non-parametric analysis showed that the benefit reform of 1993 did not substantially raise the speed of exit from unemployment to employment. Clearly, this treatment group approach is not enough to claim that the benefit system is not a source of disincentives to leave unemployment for work. Let me therefore consider a second approach, which studies the impact of the replacement rate and the remaining duration of UI benefits on the transition rates to jobs directly. In order to control for this and further heterogeneity, an exponential hazard rate model with piece-wise constant terms for the baseline-hazard is adopted. Again this analysis is applied only to those workers who are considered as less likely to be on recall.

Hazard rates, in this 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, they control for educational level, main regional activity and unemployment rate in 1993, age, previous unemployment spell, non-manual worker and entry to UI receipt from employment. However, the focus lies on the impact of the benefit system on the hazard rates. Therefore, the discussion centres 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: The value of one more period of unemployment depends positively on the benefit level (relative to prospective earnings).
- A limited period of UI receipt leads to the *entitlement effect*: The length of entitlement to UI is another parameter that determines the generosity of the UI system. It has a negative impact on the search intensity and a positive one on the reservation wage of an unemployed job-searcher. Since the length of entitlement decreases over an insured unemployment spell, search intensity rises and the reservation wage declines. For both reasons the escape rates are higher the shorter the 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-stationarity.

Typically disincentive effects of unemployment benefits are measured by the elasticity of the hazards to the replacement rate. Gross benefit and prospective earnings variables are included as separate covariates in order to test whether a replacement rate specification is appropriate. Note that their net values would be a more appropriate measure for disincentive effects. The size of the entitlement effect is captured by the coefficient of a variable that represents the remaining duration of UI receipt over an unemployment spell. This variable has

therefore to vary over an unemployment spell. Before the results are presented, some matters of concern with this approach of specifying the benefit system need to be discussed.

Backward versus Forward-Looking Replacement Rates

The replacement rate may be expressed either by a backward or forward-looking concept[16]. Whereas the backward-looking concept defines it as benefit relative to past earnings, the forward-looking concept 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 (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. Thus, the logarithmic prospective earnings of unemployed people are predicted as the logarithmic (indexed) pre-unemployment earnings augmented by an expected wage gain. The equation for this wage gain conditions the logarithmic difference between accepted and past (indexed) wages on a set of observable characteristics (see Appendix A). It is estimated by OLS using a second data source on pre- and post-unemployment wages in Hungary. However, these predicted wages are still close to the past indexed wage, as only a minor part of the variation of the post-unemployment wage gain can be explained.

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 separately. The reason for this concern is that the benefit levels depend on the previous wage rate. In all, there are four sources that introduce the independent variation to identify these parameters. First of all a predicted wage rate and not the past earnings of a UI recipient are used. Second, UI benefits are imputed by a benefit formula which, given the minimum and maximum benefit rules, leads to a variation in benefits that is independent from earnings. A third source of variation are the two phases of benefit receipt, which, for a part of the sample, imply that the UI benefit changes when a spell reaches the second phase of UI receipt. Since there are many entitlement periods, this happens at different spell lengths. Finally, the two benefit systems, with their distinct benefit formulas, introduce further variation into the benefit data that is independent from earnings.

[16] There are more unresolved problems with the concept of the replacement rate in this study. In particular, 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. Thus, one would need a replacement rate that results from benefits augmented by additional earnings while unemployed relative to prospective wages.

For all this, there is some independent variation between benefits and prospective wages. Table 5 describes the variation of benefits and prospective wages of workers that are less likely to be on recall. The coefficient of variation of the replacement rate is 0.22 for men and women alike. Next, the correlation between benefits and earnings is about 0.73 for men and 0.83 for women[17]. The difference between these two correlations mainly stems from one fact. Compared with men a much higher proportion of women earned such a low wage in the past that their benefits are below the minimum benefit[18]. For such individuals the UI benefit is not altered between the two phases of UI entitlement and their formal replacement rate is 100 percent. Thus their benefits are more closely related to their previous earnings.

Another identification problem may arise for the parameters of the remaining entitlement to UI and those of the baseline hazards. The latter directly measure the impact of duration on the hazards. As the sample consists of two inflow cohorts of two subsequent calendar months, December 1992 and January 1993, elapsed duration by and large coincides with calendar time. Thus, the baseline hazard picks up any unmeasured effect, such as variations in labour demand or wage inflation over the 16 months that follow. Suppose that all spells had the same length of UI entitlement. Then one could not identify the effect of duration and that of remaining duration of UI entitlement on the hazards, separately; their correlation would be -1.

This identification problem does not arise in this sample, as the benefit rules imply a large number of distinct entitlement lengths. First of all, the subsample consists of individuals from nine work-history groups. Each of these implies a different length of UI receipt. Next, the unemployment spells come from two different benefit regimes. The length of UI entitlement of 1993 claimants is exactly two thirds of that of 1992 claimants with the same work-history. Therefore, a total of 17 entitlement periods, which range from 90 to 495 days, is available.

6.2 Results of Maximum Likelihood Estimation

Now turn to analysing the impact of the benefit system on the job hazards. The subsample of workers who are less likely to be on recall consists of 23,494 men and 11,209 women. The reference individual of the following models neither experienced a UI spell before nor entered unemployment from employment. This person has a primary education and is older than 24 years but younger than 35 years. He/she lives in an agricultural region with an unemployment rate between 12 and 18 percent in 1993 and is a manual worker. The discussion focuses on the estimated coefficients of the variables that describe the UI system. These coefficients are

[17] The benefit variable involved in these calculations represents a weighted average of first and second phase UI benefits. The weights for each phase are defined as the time for which the people received the corresponding benefit divided by the total observed duration of the spell.
[18] In fact, more than 34 percent of women as compared with only 17 percent of men.

displayed in Table 6a for men and Table 6b for women[19] [20]. A variety of specifications is discussed in order to establish the stability of the effects of the benefit system.

The first and most simple specification represents the UI system by only two variables: the logarithms of UI benefits and of expected wages. Specification two extends the first model by including the remaining duration of UI entitlement in the set of covariates. It is specified by a sequence of five dummy variables that are set to one when a spell reaches the following times until exhausting UI: 270-180 days, 180-120 days, 120-60 days, 60-30 days and 30 days or less. In the base case time until exhaustion is longer than 270 days. Table 5 shows the proportion of spells that reach these intervals. The numbers in brackets represent the number of exits to employment during these intervals relative to all observed exits to jobs. In the case of men, no interval is characterised by a share of exits below seven percent. The corresponding number for women is nine percent. An entitlement effect would first imply that the coefficients of the exhaustion dummies are positive. Next, they must be the larger, the closer the interval of remaining entitlement is to the date when UI expires.

In the third specification I allow the impacts of benefit and earnings to vary with age. A reason for studying this age-interaction 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 rising in age. Thus, the hazards to jobs of young people are expected to be most sensitive to benefit levels[21]. For this reason, the third specification of the employment hazard allows for an interaction of benefits and earnings with age for two age-categories: below 30 years and above 29 years

Effects of Benefits and Earnings

[19] The baseline hazard is specified by piece-wise constant shift parameters as described in Section 2. The number of observed exits to jobs in the sample is high for short durations, but becomes much smaller after a duration of more than one year. The piece-wise constant terms are therefore estimated for relatively short interval lengths of 15 or 30 days, for durations that do not exceed 360 days. Thereafter, an interval length of 45 days is assumed.

[20] The mean values of variables that are not related to the UI system variables are displayed in Table B1. Their estimated coefficients are found in Table B2.

[21] 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. Arulampalam and Stewart (1996) also 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 from the DHSS. The second data sets originate from the Department of Social Security (DSS) survey of Incomes In and Out of Work, and contain unemployment spells that began in the four weeks starting March 16, 1987. They found a high negative elasticity of the job hazards to benefit levels of teen-aged men during the first quarter of an unemployment spell for both cohorts. For 20-44 year-old 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 levels on the job hazards ceased to exist.

Let me first consider the results for men with respect to benefit and earnings. They are displayed in Table 6a. For specification one to four the estimated elasticities of the hazards with respect to UI benefits range from 0.019 to 0.068 and thus are very close to zero. Next, at a 5 percent significance level they are never statistically significant. The estimated elasticities of the hazard with respect to earnings are positive and always close to 0.1. Only for the first specification though is the elasticity with respect to earnings statistically different from zero. In all, there is no evidence for a disincentive effect, and a Likelihood-Ratio (LR) test suggests that a specification in terms of the replacement rate is not adequate. In specification five, however, a surprising result emerges for men: the estimated elasticity of the hazards to the replacement rate is -0.2 and significant. This contrasts the findings of specification one to four on the benefit and earnings elasticity. A reason for this may be that for men the correlation between benefits and earnings is too high to identify their coefficients, separately. At 0.73 this correlation is high, however, it is not close to one. Whatsoever, according to the estimated coefficients for the elasticity of the hazards to the replacement rate a reduction of the replacement rate by one half would lead to a 10 percent increase in the transition rate to employment for men. In other words, replacement rates have to be reduced considerably to increase the outflow from unemployment by a relatively low percentage.

The results for women are shown in Table 6b. The effects of benefits and earnings for women differ from those of men. For specifications one to four the sign of the estimated elasticity of the hazard with respect to benefits is always negative while that of the elasticity with respect to earnings is always positive. This finding is in line with a disincentive effect. Next, according to an LR-test the hypothesis that these two elasticities sum up to zero is never rejected. So, a replacement rate specification is adequate. However, both elasticities are not stable over the specifications one to four. In specification one, where only benefits and earnings in logarithms are included, the elasticity of the hazard with respect to benefits is –0.29 and with respect to earnings it is 0.21. Adding time until exhaustion of UI as a covariate (specification two), decreases them by about one half. The estimated benefit and earnings elasticities of the second specification and the fourth specification, where dummies for work-history were added, are about the same. Finally, specification three provides some evidence that the effects of benefits and earnings vary with age. For those younger than 30 years, there is a significant disincentive effect. The benefit coefficient is –0.35 and the earnings coefficient is 0.31. For those older than 29 though, both coefficients are insignificant and very close to zero. So the evidence is in favour of the hypothesis of Narendranathan et al, that young people's hazards are more sensitive to UI benefits than those of older people. However, our preferred specification is specification five. The reason is that LR-tests suggest that a replacement rate specification is adequate and dummies for work-history should be included. The estimated elasticity of this specification suggests that reducing the replacement rate by one half would raise the transition rate of women by little more than six percent. So the impact of such a policy on the job hazard is very limited.

Effects of Remaining Entitlement.

Let me again first consider the results for men as displayed in Table 6a. In specification

two to five I estimated coefficients of dummies for time until exhausting UI entitlement. For men, the estimated coefficients for these dummies do not entirely fulfil the requirements for an entitlement effect. They do not imply that the hazards monotonously rise, the closer a spell is to the UI expiry date. A considerable rise only emerges in the last 30 days. In this period the hazard exceeds that of the base period of more than 270 days of remaining UI entitlement by 34.2 percent. This result remains unaltered in specification three, where age interaction of the benefit and earnings covariates has been considered. It is however not stable when dummies for work-history are introduced as covariates as is in specification four. In this specification, the coefficients of the remaining entitlement dummies at least by and large imply that the hazard is negatively related to time until exhaustion. The outstanding rise of the hazard happens in the last month of UI receipt. In this month the hazard exceeds that of the month before by 36 percent, and that of the base case, more than 270 days until UI expiry, by more than 50 percent. This is also my preferred specification for men, as a LR-test shows that the coefficients of the work-history dummies are jointly significant[22].

Now let us turn to women. The results in Table 6b are clearly in support of an entitlement effect. In specification two to four the coefficients for remaining UI entitlement imply that the hazards rise the nearer the date of UI expiry. The effect is statistically significant and it is quite stable over different specifications. Therefore, let us only discuss the size of the entitlement effect of the last and preferred specification[23]. Compared with the base case of more than 270 days before running out of UI, the job hazard during the last 270 to 180 days is 1.5 percent higher. During the next two intervals, it is 8 to 11.2 percent higher than in the base case. Next, for a remaining duration of 60 to 30 days, the corresponding value is already 22.8 percent, and during the last 30 days it even reaches 53.1 percent.

7. Summary and Conclusions

In this paper as well as in a previous study by Micklewright and Nagy, an attempt has been made to address one question: What is the role of the Hungarian UI system in guiding individual transitions from unemployment into jobs? The results of this paper shed some additional light on this topic. It was first shown that the sample is characterised by a large share of workers on recall. Their job hazards are, for a while, substantially higher than those of workers who are less likely to be on recall. They are not similarly distributed over the two claimant groups. Consequently, an analysis of the benefit reform of 1993 and its impact on transitions from unemployment back to work by a treatment group approach has to distinguish between workers who are more or less likely to be on recall.

[22] The estimated coefficients of the terms for work-history of the male hazards lend some support to the hypothesis that work-history is a screening device for employers. The lower the work-history, the lower the male hazards to jobs.

[23] This specification is the preferred one, since LR-tests suggest that the coefficients of remaining entitlement and work-history dummies are significant. Next, an LR-test cannot reject the hypothesis that a the replacement rate specification is adequate.

In a second step, the job hazards of 1993 claimants and 1992 claimants were estimated non-parametrically for each of these two distinct groups of workers. This analysis is carried out for men and women separately, along three different work-history groups. It yielded scant evidence for the UI reform having fostered the speed of exit to jobs. However, this may still be due to some observed, as well as unobserved, characteristics which determine the job hazards but which are not distributed in the same way over the two claimant groups. Moreover, there is little evidence that the job hazards rise while the spells come closer to the date of UI expiry. Sorting effects may be one reason why an entitlement effect is not observed. Variations of labour demand over calendar time may be another reason for this.

The semi-parametric analysis of the job hazards then focused the disincentive and entitlement effect of UI benefits. This part of the study attempts to exploit the variation in replacement rates and remaining duration of UI entitlement to identify both. The results do not support any substantial disincentive effect. The exception may be women who are younger than 30 years. The elasticity of their hazards with respect to UI benefits is -0.35 and with respect to wages 0.31. In contrast, the results are in support of an entitlement effect at least for women. Their hazards are negatively related to the time until expiry of UI and are particularly high during the last 60 to 30 days and the last 30 days of UI receipt. According to the preferred specification five, in these two intervals, the hazards exceed those of the base case of more than 270 days by 22.8 percent and 53.1 percent, respectively. This effect is stable over several specifications. The size of this effect is similar in the preferred male model. However, for men the effect is not stable over several specifications.

Some doubt may be cast on the results about the disincentives provided by the UI benefit system. It is not possible to be extremely confident about the 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 jobs or jobs in the shadow economy, as well as household income, are not available. Therefore, alternative specifications of benefit and earnings variables, that are better suited as guides to disincentive effects, cannot be considered.

The results are still, however, defensible. As gross benefits and earnings are certainly highly correlated with their net values, if there were a large disincentive effect from the UI system, it would have been identified. There is also a reasonable explanation for a low elasticity of the job hazards to UI benefits. At high levels of unemployment, one might expect their benefit elasticity to be smaller than during periods of low unemployment. Such a hypothesis was suggested by Arulampalam and Stewart (1996), who found some evidence for this in a study on unemployment duration and disincentive effects in Britain. The results of this paper are in line with these findings. In a depressed labour market such as the Hungarian one unemployed people are constrained from the labour demand side. Therefore, the UI system is not an extremely important determinant of the speed of return to work of

unemployed people.

Also the entitlement effect that has been identified does not imply that cutting the length of UI entitlement leads to a much higher escape rate from unemployment into employment. The reason is that it is only in the last two months of benefit receipt that this effect is substantial. This is only a short part of total UI entitlement over most of the sample. Next, in Hungary this effect is unlikely to continue for long once people run out of UI. Micklewright and Nagy (1996b, 1997) provided evidence for this. They found that a few weeks after the expiry of UI benefits, the hazards return to approximately the same level as during the two months prior to UI exhaustion. One reason for a small entitlement effect in the job hazard may be the availability of labour market programs (LMPs). Carling et al. (1994) argued that a sharp rise of the employment hazards just prior to UI exhaustion might not emerge if LMPs are present. These programs reduce the risk of income reductions, for people who are about to run out of their unemployment benefit. Carling et al. showed, in a dynamic search model with LMP exit, that the standard result that reservation wages fall the shorter the remaining entitlement to unemployment benefits may no longer hold. The strength of an entitlement effect in the hazards is then negatively related to the probability that a LMP is offered to unemployment benefit recipients.

The policy implication of the findings is straightforward. In Hungary, reforms of the benefit system were not a tool that made unemployed workers return to work faster during the transition process. However, this may change once the transition of the Hungarian economy leads to a considerable upturn of the labour market and is thus a question for further research. These conclusions, however, do not imply that in the long run a benefit reform could not lead to lower unemployment by fostering labour demand. It may reduce the Solidarity Fund's spending, making lower contributions of employers and employees possible. Thus, it may reduce wage costs and therefore increase labour demand. However, as the benefit system is grandfathered, this situation is only likely to happen after a considerable delay.

APPENDIX A: WAGE GAIN EQUATION

In order to predict expected wages, I draw information from a second source of data: An outflow sample from the Hungarian UI register, which consists of workers who left the UI register by finding a job between March 20th and April 20th in 1994. They were interviewed in order to measure their post-unemployment wage. Like the inflow data that is used for the duration analysis, this data provides information about regions, education, age, completed length of the unemployment spell (in weeks) and pre-unemployment earnings. Köllö and Nagy (1996) already carried out an analysis on post-unemployment wage change with this data. I am interested only in the subsample of job-losers. This sample consists of 6,768 men who are below 55 years, and 2,577 women who are below 50 years.

The expected (logarithmic) post-unemployment wages for the inflow sample to the UI register are computed in the following way: the logarithmic and indexed pre-unemployment wage is augmented by the logarithmic wage gain that results from these two equations (Table A) evaluated at a duration of one week.

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

Entitlement Length

		Work-history (months)									
		12-15	16-19	20-23	24-27	28-31	32-35	36-39	40-43	44-47	48
		Length of entitlement (days)									
December	Phase 1	90	120	150	180	210	240	270	300	330	360
1992	Phase 2	45	60	75	90	105	120	135	150	165	180
	Total	135	180	225	270	315	360	405	450	495	540
January	Phase 1	23	30	38	45	53	60	68	75	83	90
1993	Phase 2	67	90	112	135	157	180	202	225	247	270
	Total	90	120	150	180	210	240	270	300	330	360

Benefit Formula

	December 1992		January 1993	
	Phase 1	Phase 2	Phase 1	Phase 2
(a) formal replacement rate	70 percent	50 percent	75 percent	65 percent
(b) maximum benefit	18000 HUF	18000 HUF	18000 HUF	15000 HUF
(c) minimum benefit	9000 HUF	9000 HUF	8600 HUF	8600 HUF
(d) replacement rate if past wages < minimum benefit	100 percent	100 percent	100 percent	100 percent

Table 2: Share of Exit from UI receipt (in percent)

	Men	Women
Observations	54,901	25,195
Exit States		
Employment	48.4	33.7
Other	12.6	17.9
Right-Censored (total)	39.0	48.4
Right-Censored (exhausting UI)	35.1	43.5

Table 3: Unemployment Spells by Calendar End of Previous UI Spell

(a) Men

	1992 Benefit rules			1993 Benefit rules		
	Proportion of workers		Obs.	Proportion of workers		Obs.
	less likely to be on recall[1)]	more likely to be on recall[2)]		less likely to be on recall[1)]	more likely to be on recall[2)]	
Work-history						
12-27 months	93.6	6.4	4,528	93.9	6.1	3,054
28-43 months	80.0	20.0	7,123	87.0	13.0	5,308
44-47 months	43.4	56.6	7,344	63.4	36.6	4,552
Total	69.1	30.9	18,995	80.3	19.7	12,914

(b) Women

	1992 Benefit rules			1993 Benefit rules		
	Proportion of workers		Obs.	Proportion of workers		Obs.
	less likely to be on recall[1)]	more likely to be on recall[2)]		less likely to be on recall[1)]	more likely to be on recall[2)]	
Work-history						
12-27 months	96.5	3.5	2,250	95.6	4.4	1,561
28-43 months	88.4	11.6	2,817	92.8	7.2	2,320
44-47 months	76.4	23.6	1,964	92.5	7.5	1,516
Total	87.6	12.4	7,031	93.5	6.5	5,397

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 4: Distribution of Observable Characteristics over the two Claimant Groups

	Men		Women	
	Benefit rules		Benefit rules	
	1992	1993	1992	1993
Observations	13,121	10,373	6,162	5,047
Exits into employment	0.41	0.37	0.35	0.30
Education				
Incomplete primary	0.05	0.04	0.04	0.04
Primary	0.32	0.31	0.44	0.42
Vocational	0.49	0.49	0.26	0.26
Vocational secondary	0.09	0.09	0.12	0.13
General secondary	0.03	0.04	0.11	0.11
College	0.02	0.02	0.02	0.02
University	0.01	0.01	0.01	0.01
Region				
Agricultural[1)]	0.42	0.34	0.40	0.32
Industrial [2)]	0.21	0.21	0.20	0.18
Diversified[3)]	0.36	0.44	0.40	0.49
Age				
< 25 years	0.31	0.29	0.28	0.29
25-34 years	0.28	0.30	0.27	0.30
35-44 years	0.27	0.29	0.34	0.31
45-49 years			0.11	0.10
45-54 years	0.14	0.13		
Work-history				
12-27 months	0.32	0.28	0.35	0.30
28-43 months	0.43	0.45	0.40	0.43
44-47 months	0.24	0.28	0.24	0.28
Other				
Previous spell of UI	0.41	0.34	0.35	0.26
Entry to UI from employment	0.87	0.84	0.82	0.78
Non-manual	0.07	0.07	0.21	0.22

Regions:

1) Bacs-K., Bekes, Hajdu-B., Somogy, Szabolcs, Szolnok, Tolna;

2) Borsod, Fejer, Komarom, Nograd, Veszprem

3) Baranya, Budapest, Csongrad, Gyor-S.,Heves, Pest, Vas, Zala

Table 5: UI Benefits, Earnings and Remaining Entitlement

(a) Men[1)]

Explanatory variables related to the UI system	Mean	Std.Dev.	Lower Decile	Upper Decile
Benefits and wages (in HUF)[2)]				
UI benefit	10,186	2,668	8,000	16,127
Predicted post-unemployment wage	19,596	11,127	10,767	30,173
Replacement rate	0.569	0.126	0.426	0.751
Dummies for remaining entitlement to UI[3)]				
>270 days	0.57 (0.26)	.	.	.
270-180 days	0.65 (0.26)	.	.	.
180-120 days	0.68 (0.16)	.	.	.
120-60 days	0.65 (0.16)	.	.	.
60-30 days	0.55 (0.07)	.	.	.
< 30 days	0.51 (0.09)	.	.	.

(b) Women[1)]

Explanatory variables related to the UI system	Mean	Std.Dev.	Lower Decile	Upper Decile
Benefits and wages (in HUF)[2)]				
UI benefit	8,922	2,312	6,698	11,280
Predicted post-unemployment wage	16,030	7,795	9,119	24,804
Replacement rate	0.608	0.135	0.440	0.784
Dummies for remaining entitlement to UI[3)]				
>270 days	0.54 (0.23)	.	.	.
270-180 days	0.66 (0.21)	.	.	.
180-120 days	0.73 (0.18)	.	.	.
120-60 days	0.71 (0.18)	.	.	.
60-30 days	0.62 (0.09)	.	.	.
< 30 days	0.56 (0.11)	.	.	.

1) Only workers who are less likely to be on recall.

2) The statistics take into account that benefits vary over the observed spell length.

3) The numbers in brackets indicate the exits to employment during an interval of remaining entitlement to UI as a proportion of the total number of exits in our sample.

Table 6: Transition Rates into Employment, Maximum Likelihood Estimates of Coefficients of Covariates Related to the UI System

(a) Men

Covariates	Spec. 1 Coeff.	t	Spec. 2 Coeff.	t	Spec. 3 Coeff.	t	Spec. 4 Coeff.	t	Spec. 5 Coeff.	t
(1) Benefits and earnings										
Log. of UI benefits	0.019	0.19	0.041	0.39			0.068	0.66		
Log. of earnings	0.122	2.05	0.112	1.83			0.098	1.60		
LR-Test[1] (Ho:ß$_1$+ß$_2$=0, χ^2(1))	7.0 **		7.7 **				9.0 **			
Log. of replacement rate									-0.201	-3.95
Interacting with age										
(a) aged < 30 years										
Log. of UI benefits					0.036	0.31				
Log. of earnings					0.121	1.48				
LR-Test[1] (Ho:ß$_1$+ß$_2$=0, χ^2(1))					8.09 **					
(b) aged >= 30 years										
Log. of UI benefits					0.053	0.48				
Log. of earnings					0.100	1.38				
LR-Test[1] (Ho:ß$_1$+ß$_2$=0, χ^2(1))					7.77 **					
(2) Remaining entitlement to UI										
270-180 days			0.083	2.78	0.083	2.78	0.114	3.61	0.106	3.37
180-120 days			-0.030	-0.85	-0.030	-0.85	0.032	0.80	0.014	0.36
120-60 days			-0.028	-0.76	-0.028	-0.76	0.069	1.45	0.042	0.90
60-30 days			-0.003	-0.07	-0.003	-0.06	0.117	1.94	0.085	1.42
30-0 days			0.294	6.18	0.294	6.18	0.425	6.81	0.391	6.37
(3) Work-history										
36-43 months							0.023	0.81	0.024	0.82
28-35 months							-0.078	-2.09	-0.073	-1.97
20-27 months							-0.077	-1.63	-0.069	-1.47
12-19 months							-0.183	-2.95	-0.173	-2.79
Log. Likelihood	-65,872.2		-65,839.5		-65,838.7		-65,831.7		-65,836.2	
Number of spells	23,494		23,494		23,494		23,494		23,494	
LR-Tests[1)2)] (nested models)										
against specification 1			65.5**	χ^2(5)	67.2**	χ^2(7)	81.0**	χ^2(9)		
against specification 2					1.7	χ^2(2)	15.6**	χ^2(4)		

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

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

(b) Women

Covariates	Spec. 1 Coeff.	t	Spec. 2 Coeff.	t	Spec. 3 Coeff.	t	Spec. 4 Coeff.	t	Spec. 5 Coeff.	t
(1) Benefits and earnings										
Log. of UI benefits	-0.29	-1.89	-0.142	-0.92			-0.154	-0.99		
Log. of earnings	0.21	2.11	0.120	1.22			0.137	1.39		
LR-Test[1] (Ho:$ß_1+ß_2=0$, $\chi^2(1)$)	1.1		0.1				0.0			
Log. of replacement rate									-0.127	-1.49
Interacting with age										
(a) aged < 30 years										
Log. of UI benefits					-0.345	-1.95				
Log. of earnings					0.314	2.43				
LR-Test[1] (Ho:$ß_1+ß_2=0$, $\chi^2(1)$)					0.2					
(b) aged >= 30 years										
Log. of UI benefits					0.017	0.10				
Log. of earnings					-0.026	-0.22				
LR-Test[1] (Ho:$ß_1+ß_2=0$, $\chi^2(1)$)					0.0					
(2) Remaining entitlement to UI										
270-180 days			0.048	0.94	0.049	0.94	0.015	0.27	0.015	0.28
180-120 days			0.118	2.10	0.118	2.11	0.076	1.16	0.077	1.18
120-60 days			0.149	2.58	0.149	2.59	0.104	1.37	0.106	1.40
60-30 days			0.257	3.55	0.256	3.54	0.203	2.17	0.205	2.20
30-0 days			0.486	6.71	0.484	6.69	0.424	4.35	0.426	4.39
(3) Work-history										
36-43 months							0.191	3.91	0.191	3.91
28-35 months							0.195	3.20	0.195	3.20
20-27 months							0.113	1.48	0.113	1.48
12-19 months							0.151	1.59	0.151	1.58
Log. Likelihood	-26,803.7		-26,779.2		-26,776.6		-26,769.9		-26,769.9	
Number of spells	11,209		11,209		11,209		11,209		11,209	
LR-Tests[1)2)] (nested models)										
against specification 1			48.8**	$\chi^2(5)$	54.2**	$\chi^2(7)$	67.5**	$\chi^2(9)$		
against specification 2					5.3	$\chi^2(2)$	18.6**	$\chi^2(4)$		

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

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

Table A: OLS Estimates of Coefficients of Wage Gain Equation

	Men		Women	
Regressors	Coefficient	t-value	Coefficient	t-value
constant	0.211	6.71 **	0.210	3.82 **
ln(age/10)	-0.151	-8.95 **	-0.148	-4.86 **
Education				
Incomplete primary	-0.049	-1.86	0.008	0.17
Primary	-0.057	-5.32 **	-0.026	-1.35
Vocational secondary	-0.027	-1.72	0.007	0.32
General secondary	-0.042	-1.47	-0.023	-0.88
Higher	-0.018	-0.53	-0.003	-0.07
County				
Budapest	0.173	6.10 **	0.139	3.31 **
Baranya	0.079	2.65 **	0.059	1.22
Bacs-Kis	0.080	3.03 **	0.110	2.37 *
Borsod	0.077	3.40 **	0.047	1.04
Csongrad	0.089	2.95 **	0.029	0.57
Fejer	0.066	2.18 *	0.044	0.94
Gyor-Sop	0.129	4.90 **	0.067	1.44
Hajdu-Bi	0.098	4.35 **	0.040	0.88
Heves	0.098	3.63 **	0.062	1.24
Komarom	0.053	1.31	0.006	0.11
Nograd	0.044	1.41	0.048	0.88
Pest	0.133	5.32 **	0.093	2.21 *
Somogy	0.062	2.44 *	0.047	1.12
Szabolcs	0.051	2.27 *	0.179	3.52 **
Szolnok	0.074	2.92 **	0.023	0.52
Tolna	0.143	4.76 **	0.013	0.26
Vas	0.134	4.52 **	0.000	0.01
Veszprem	0.024	0.99	0.050	1.22
Zala	0.069	2.64 **	0.099	2.11 *
Duration				
ln(duration)	-0.035	-5.78 **	-0.029	-3.45 **
duration of 13-15 weeks	-0.072	-5.77 **		
duration of 14 weeks			-0.123	-2.07 *
Observations	6,768		2,577	
R^2 (adjusted R^2)	0.0403 (0.0365)		0.034 (0.0237)	
RESET-Test[1)]	F(3, 6737)=1.67		F(3, 2546)=3.38*	
Cook-Weisberg[2)] for heteroscedasticity	$\chi^2(1)$=112.49**		$\chi^2(1)$=8.36**	
Joint significance of				
Education	F(5,6740)=6.02**		F(5,2549)=0.77	
Counties	F(19,6740)=4.39**		F(19,2549)=1.85*	

* implies a five 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: Observable Characteristics (Proportions)

	Men	Women
Observations	23,494	11,209
Variable		
Exits into employment	0.39	0.32
Education		
Incomplete primary	0.05	0.04
Primary	0.31	0.43
Vocational	0.49	0.26
Vocational secondary	0.09	0.13
General secondary	0.03	0.11
College	0.02	0.02
University	0.01	0.01
Region		
Agricultural:		
Unemployment rate		
< 12%[1)]	0.04	0.04
12-18 %[2)]	0.27	0.25
> 18%[3)]	0.08	0.08
Industrial:		
Unemployment rate		
< 12%[4)]	0.08	0.09
12-18 %[5)]	0.03	0.03
> 18%[6)]	0.10	0.08
Diversified:		
Unemployment rate		
< 12%[7)]	0.32	0.37
12-18 %[8)]	0.08	0.07
Age		
< 20	0.07	0.09
20-24	0.23	0.19
25-29	0.15	0.14
30-34	0.14	0.14
35-39	0.16	0.18
40-44	0.12	0.15
45-49	0.08	0.11
50-54	0.06	
Work-history		
12-19	0.14	0.16
20-27	0.16	0.16
28-35	0.19	0.18
36-43	0.25	0.23
44-47	0.26	0.26
Other		
Previous spell of UI	0.38	0.31
Entry to UI	0.86	0.80
from employment		
Non-manual	0.07	0.22

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 B2: Transition Rate into Employment, Maximum Likelihood Estimates of Coefficients of Covariates not Related to the UI System

(a) Men

Covariates	Spec. 1		Spec. 2		Spec. 3		Spec. 4		Spec. 5	
	Coeff.	t	Coeff.	t	Coeff.	t	Coeff.	t	Coeff.	t
(1) Period Specific Effects (Duration Interval)										
0-30 days	-9.27	-19.7	-9.39	-19.3	-9.42	-19.4	-9.48	-19.4	-8.03	-104.7
30-45 days	-8.69	-18.5	-8.81	-18.1	-8.84	-18.2	-8.90	-18.2	-7.46	-94.9
45-60 days	-8.68	-18.5	-8.80	-18.1	-8.82	-18.1	-8.89	-18.2	-7.45	-94.2
60-75 days	-8.32	-17.8	-8.46	-17.4	-8.48	-17.5	-8.56	-17.6	-7.11	-93.7
75-90 days	-7.98	-17.1	-8.11	-16.8	-8.14	-16.8	-8.22	-16.9	-6.78	-91.8
90-105 days	-7.95	-17.1	-8.08	-16.7	-8.10	-16.8	-8.19	-16.9	-6.75	-90.7
105-120 days	-7.97	-17.1	-8.12	-16.8	-8.14	-16.8	-8.23	-17.0	-6.79	-90.5
120-150 days	-8.09	-17.4	-8.22	-17.0	-8.25	-17.1	-8.34	-17.2	-6.90	-96.0
150-180 days	-8.21	-17.7	-8.35	-17.3	-8.38	-17.3	-8.48	-17.5	-7.04	-94.7
180-210 days	-8.27	-17.8	-8.40	-17.3	-8.43	-17.4	-8.55	-17.6	-7.11	-91.9
210-240 days	-8.47	-18.2	-8.62	-17.7	-8.64	-17.8	-8.78	-18.0	-7.33	-88.8
240-270 days	-8.52	-18.3	-8.69	-17.9	-8.71	-17.9	-8.86	-18.1	-7.42	-85.6
270-300 days	-8.43	-18.1	-8.60	-17.7	-8.62	-17.7	-8.78	-17.9	-7.33	-81.3
300-330 days	-8.67	-18.6	-8.88	-18.2	-8.90	-18.2	-9.06	-18.4	-7.62	-75.4
330-360 days	-9.06	-19.3	-9.18	-18.7	-9.21	-18.8	-9.37	-19.0	-7.94	-62.0
360-405 days	-9.07	-19.4	-9.20	-18.8	-9.23	-18.8	-9.42	-19.1	-7.98	-65.6
405-450 days	-8.62	-18.5	-8.80	-18.0	-8.82	-18.0	-9.02	-18.3	-7.59	-63.9
> 450 days	-8.42	-18.0	-8.73	-17.7	-8.76	-17.7	-8.96	-18.0	-7.52	-55.2
(2) Education										
Incomplete primary	-0.24	-3.7	-0.24	-3.7	-0.24	-3.7	-0.24	-3.7	-0.24	-3.7
Vocational	0.35	13.7	0.35	13.8	0.35	13.8	0.35	13.8	0.35	13.6
Vocational secondary	0.30	7.2	0.30	7.2	0.30	7.2	0.30	7.2	0.30	7.3
General secondary	0.19	2.9	0.19	2.9	0.19	2.9	0.19	2.9	0.19	3.0
Higher	0.58	7.4	0.58	7.4	0.58	7.4	0.58	7.4	0.58	7.5
(3) Regional Activity and Unemployment Rate										
agriculture, <12 %	0.37	6.7	0.37	6.6	0.37	6.6	0.36	6.5	0.36	6.5
agriculture, >18 %	-0.26	-5.9	-0.26	-6.0	-0.26	-5.9	-0.26	-5.9	-0.26	-5.9
industry, <12 %	0.34	8.7	0.34	8.7	0.34	8.7	0.33	8.7	0.34	8.9
industry, 12-18 %	0.09	1.5	0.09	1.5	0.09	1.5	0.09	1.5	0.09	1.6
industry, >18 %	-0.12	-3.0	-0.12	-3.0	-0.12	-3.0	-0.12	-3.1	-0.12	-3.0
diversified, <12 %	0.07	2.5	0.07	2.5	0.07	2.5	0.07	2.3	0.07	2.3
diversified, 12-18 %	0.06	1.5	0.06	1.5	0.07	1.5	0.06	1.5	0.06	1.5
Age										
< 20	0.14	2.8	0.13	2.4	0.10	1.9	0.17	3.0	0.16	2.9
20-24	0.10	3.5	0.10	3.4	0.08	2.3	0.10	3.5	0.09	3.3
35-39	-0.11	-3.4	-0.11	-3.4	-0.08	-2.2	-0.11	-3.4	-0.11	-3.3
40-44	-0.24	-6.4	-0.23	-6.4	-0.21	-5.1	-0.24	-6.4	-0.23	-6.3
45-49	-0.22	-5.0	-0.22	-4.9	-0.19	-4.0	-0.22	-5.0	-0.21	-4.7
50-54	-0.39	-7.3	-0.39	-7.2	-0.36	-6.4	-0.39	-7.2	-0.38	-7.1
(4) Other										
Previous spell of UI	0.40	18.2	0.40	17.8	0.40	17.8	0.41	17.2	0.41	17.4
Entry to UI										
from employment	0.47	11.0	0.47	11.2	0.47	11.1	0.46	10.7	0.50	12.0
Non-manual	-0.26	-5.0	-0.26	-5.0	-0.26	-5.0	-0.26	-5.0	-0.25	-4.8

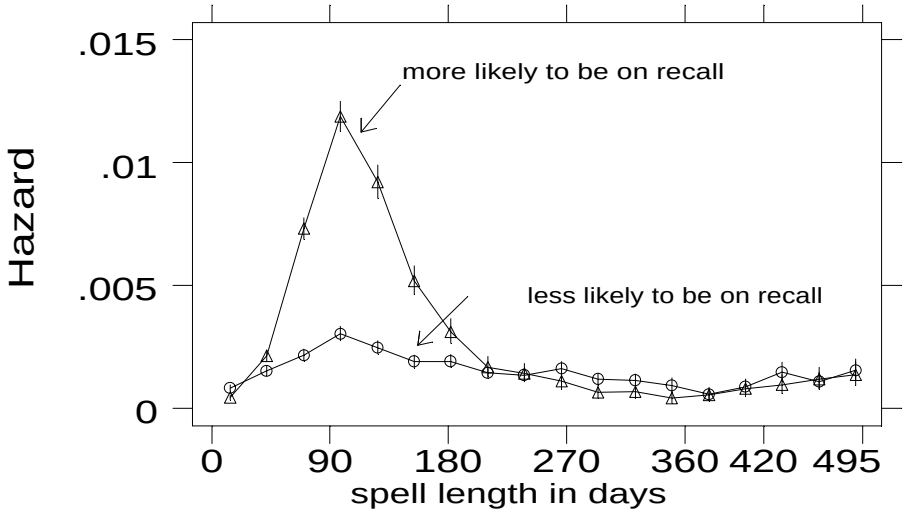
(b) Women

Covariates	Spec. 1 Coeff.	t	Spec. 2 Coeff.	t	Spec. 3 Coeff.	t	Spec. 4 Coeff.	t	Spec. 5 Coeff.	t
(1) Period Specific Effects (Duration Interval)										
0-30 days	-7.11	-10.6	-7.65	-11.1	-7.68	-11.1	-7.82	-11.3	-7.97	-72.7
30-45 days	-6.77	-10.1	-7.32	-10.6	-7.35	-10.6	-7.49	-10.8	-7.64	-65.7
45-60 days	-6.81	-10.2	-7.36	-10.7	-7.39	-10.7	-7.53	-10.9	-7.67	-65.1
60-75 days	-6.59	-9.9	-7.17	-10.4	-7.20	-10.4	-7.33	-10.6	-7.47	-64.9
75-90 days	-6.58	-9.9	-7.16	-10.4	-7.19	-10.4	-7.33	-10.6	-7.47	-64.4
90-105 days	-6.68	-10.0	-7.28	-10.6	-7.31	-10.6	-7.44	-10.8	-7.58	-63.4
105-120 days	-6.39	-9.6	-7.01	-10.2	-7.04	-10.2	-7.17	-10.4	-7.31	-63.5
120-150 days	-6.58	-9.9	-7.20	-10.5	-7.23	-10.5	-7.35	-10.7	-7.50	-68.1
150-180 days	-6.45	-9.7	-7.09	-10.3	-7.12	-10.3	-7.24	-10.5	-7.39	-66.4
180-210 days	-6.53	-9.8	-7.16	-10.4	-7.20	-10.4	-7.32	-10.6	-7.46	-63.7
210-240 days	-6.67	-10.0	-7.32	-10.6	-7.35	-10.6	-7.47	-10.8	-7.61	-61.2
240-270 days	-6.62	-9.9	-7.31	-10.6	-7.34	-10.6	-7.45	-10.7	-7.59	-58.8
270-300 days	-6.74	-10.1	-7.45	-10.7	-7.48	-10.8	-7.58	-10.9	-7.72	-55.1
300-330 days	-6.75	-10.1	-7.51	-10.8	-7.54	-10.8	-7.62	-10.9	-7.76	-52.3
330-360 days	-7.07	-10.5	-7.77	-11.1	-7.80	-11.1	-7.91	-11.2	-8.05	-44.2
360-405 days	-7.06	-10.5	-7.78	-11.1	-7.81	-11.2	-7.91	-11.2	-8.05	-46.3
405-450 days	-6.69	-10.0	-7.47	-10.7	-7.49	-10.7	-7.57	-10.8	-7.71	-43.9
> 450 days	-6.40	-9.5	-7.32	-10.4	-7.34	-10.4	-7.35	-10.4	-7.49	-39.0
(2) Education										
Incomplete primary	-0.41	-3.9	-0.42	-4.0	-0.42	-4.0	-0.43	-4.0	-0.43	-4.0
Vocational	0.13	3.1	0.13	3.1	0.14	3.3	0.13	3.1	0.13	3.1
Vocational secondary	0.28	4.8	0.27	4.7	0.28	4.7	0.27	4.6	0.27	4.6
General secondary	0.12	1.9	0.11	1.8	0.12	1.9	0.11	1.7	0.11	1.7
Higher	0.45	4.2	0.44	4.2	0.46	4.3	0.45	4.2	0.44	4.2
(3) Regional Activity and Unemployment Rate										
agriculture, <12 %	0.62	7.9	0.62	7.8	0.62	7.9	0.62	7.8	0.62	7.8
agriculture, >18 %	-0.32	-4.3	-0.29	-3.9	-0.29	-4.0	-0.29	-4.0	-0.29	-4.0
industry, <12 %	0.36	6.1	0.36	6.2	0.36	6.2	0.36	6.1	0.36	6.1
industry, 12-18 %	-0.14	-1.3	-0.14	-1.3	-0.14	-1.3	-0.14	-1.3	-0.14	-1.3
industry, >18 %	-0.35	-4.6	-0.35	-4.5	-0.35	-4.5	-0.35	-4.5	-0.35	-4.5
diversified, <12 %	0.03	0.7	0.03	0.8	0.04	0.8	0.04	0.8	0.04	0.8
diversified, 12-18 %	-0.10	-1.4	-0.10	-1.4	-0.10	-1.4	-0.10	-1.4	-0.10	-1.4
Age										
< 20	0.69	10.7	0.58	8.8	0.61	8.2	0.55	8.0	0.55	8.0
20-24	0.33	6.7	0.31	6.2	0.32	5.4	0.29	5.8	0.29	5.8
35-39	0.03	0.5	0.03	0.5	0.01	0.2	0.02	0.5	0.02	0.4
40-44	-0.08	-1.4	-0.07	-1.3	-0.09	-1.4	-0.07	-1.3	-0.07	-1.3
45-49	-0.04	-0.6	-0.03	-0.5	-0.05	-0.7	-0.03	-0.5	-0.03	-0.5
(4) Other										
Previous spell of UI	0.58	16.0	0.56	15.2	0.56	15.2	0.51	13.4	0.51	13.4
Entry to UI										
from employment	0.47	7.7	0.50	8.2	0.50	8.1	0.55	8.8	0.54	9.2
Non-manual	-0.07	-1.4	-0.05	-1.0	-0.05	-1.0	-0.05	-1.0	-0.05	-1.0

Figure 1: Transition Rate to Employment of Men:
more versus less likely to be on recall

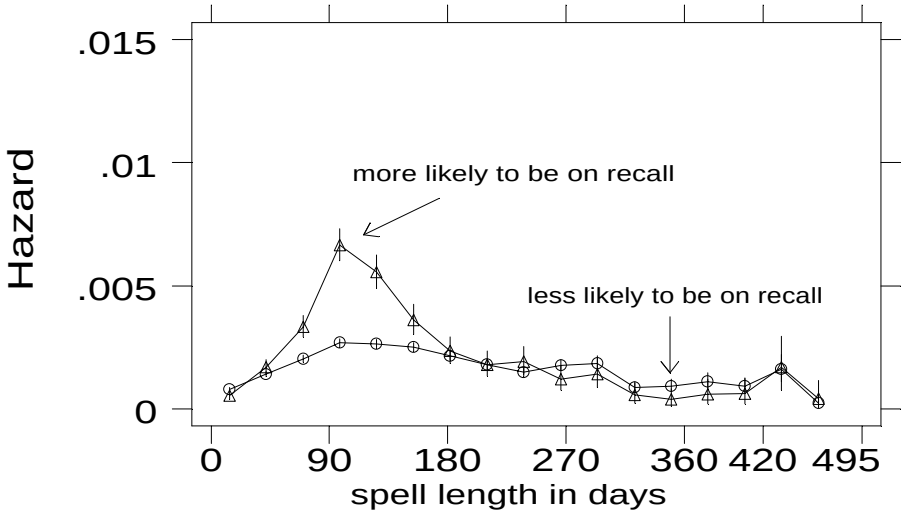
Kaplan-Meier Estimates together with 95 Percent Greenwood Confidence Bands

[a] 44-47 months of work-history



Note: sample sizes are 6,071 (less likely to be on recall) and 5,825 (more likely to be on recall)

[b] 12-43 months of work-history

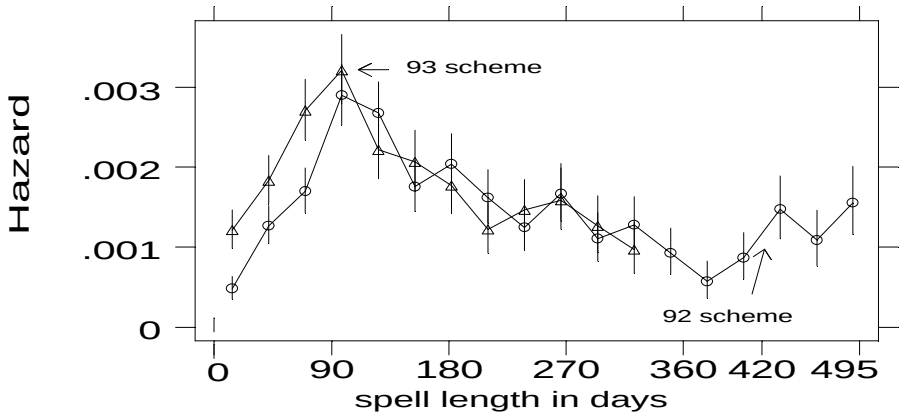


Note: sample sizes are 17,423 (less likely to be on recall) and 2,590 (more likely to be on recall)

Figure 2: Transition Rate into Employment of Men classified as less likely to be on recall

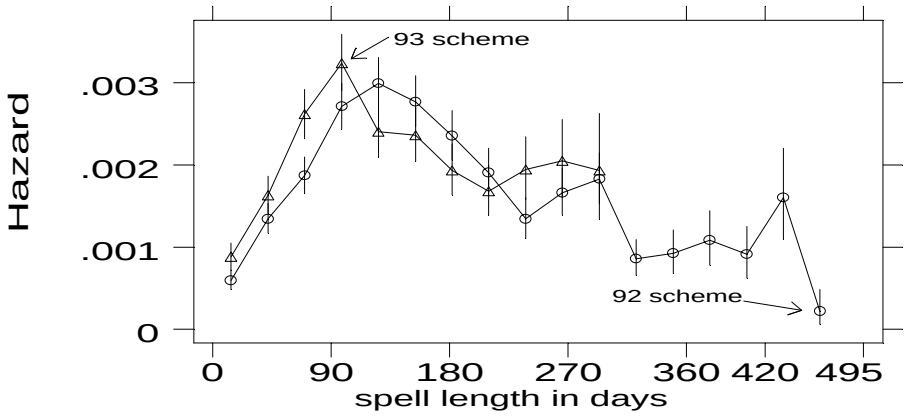
Kaplan-Meier Estimates together with 95 Percent Greenwood Confidence Bands

[a] 44-47 months of work-history



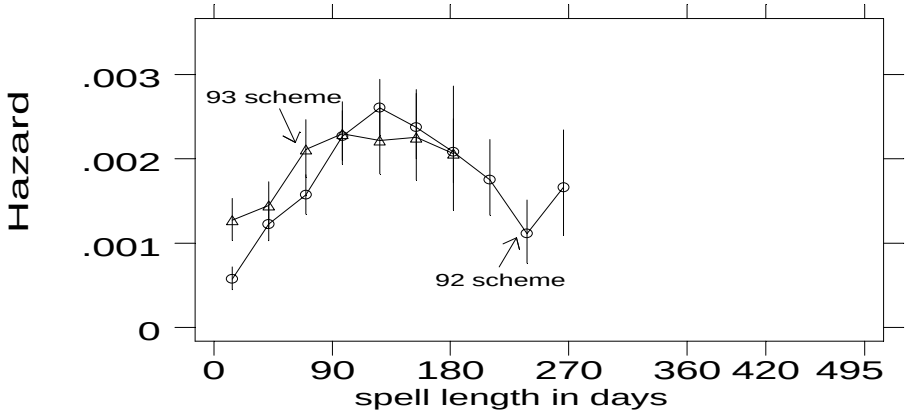
Note: sample sizes are 3,186 (92 scheme) and 2,885 (93 scheme)

[b] 28-43 months of work-history



Note: sample sizes are 5,695 (92 scheme) and 4,619 (93 scheme)

[c] 12-27 months of work-history

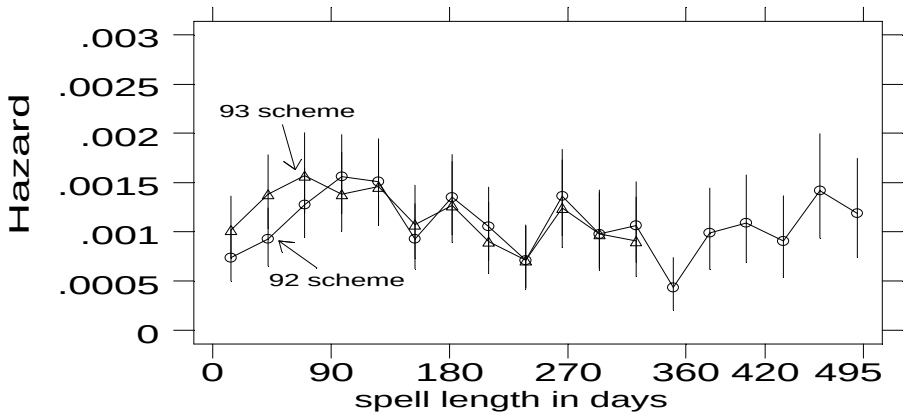


Note: sample sizes are 4,240 (92 scheme) and 2,869 (93 scheme)

Figure 3: Transition Rate into Employment of Women classified as less likely to be on recall

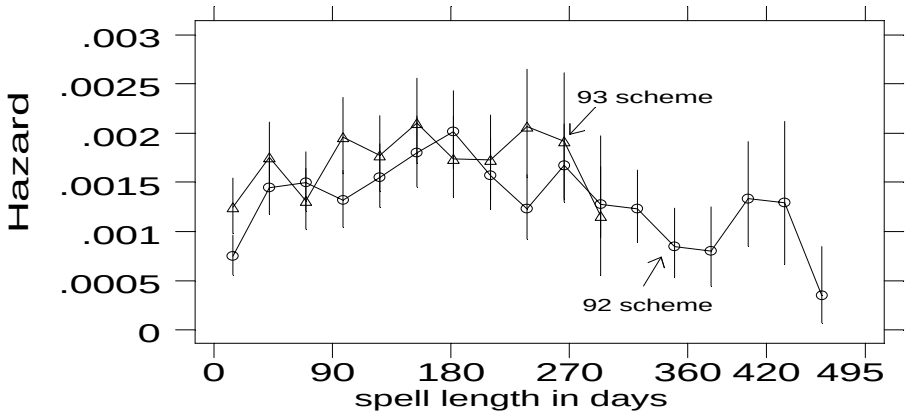
Kaplan-Meier Estimates together with 95 Percent Greenwood Confidence Bands

[a] 44-47 months of work-history



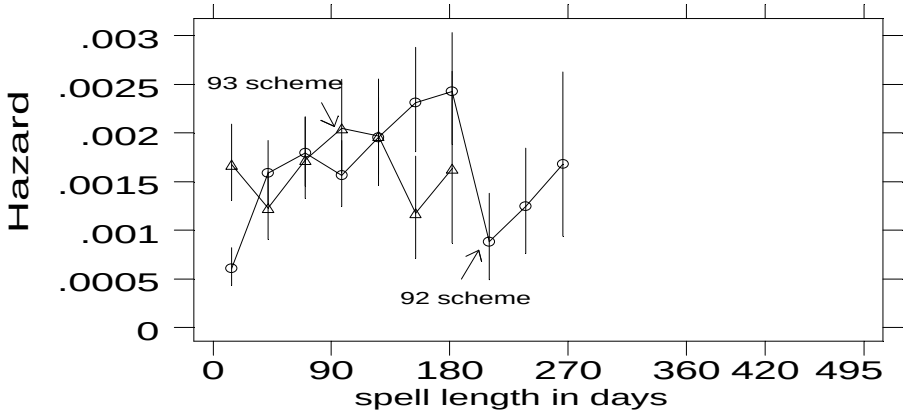
Note: sample sizes are 1,501 (92 scheme) and 1,402 (93 scheme)

[b] 28-43 months of work-history



Note: sample sizes are 2,490 (92 scheme) and 2,153 (93 scheme)

[c] 12-27 months of work-history



Note: sample sizes are 2,171 (92 scheme) and 1,492 (93 scheme)