SIX WAYS TO LEAVE UNEMPLOYMENT

Pedro Portugal* and John T. Addison**

* Banco de Portugal and Faculdade de Economia da Universidade Nova de Lisboa
** Department of Economics, School of Management, Queen's University Belfast, and IZA Bonn

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ABSTRACT

This paper uses a unique Portuguese dataset to examine the effect of access to unemployment benefits and their maximum potential duration on escape rates from unemployment. In examining the time profile of transitions out of unemployment, the principal contributions of the paper are twofold. First, it provides a detailed state space of potential outcomes: open ended employment, fixed term contracts, part- time work, government-provided jobs, self employment, and labour force withdrawal. Second, it is able to exploit major exogenous discontinuities in the maximum duration of unemployment benefits to identify (two sets of) disincentive effects. While confirming strong disincentive effects, it is shown that use of an aggregate hazard function regression model compounds very different and even contradictory effects of the determinants of unemployment.

I INTRODUCTION

This paper offers an examination of the impact of access to unemployment benefits on unemployment duration in Portugal. The effect of subsidization of the search process on jobless duration is of course familiar territory, and so the present treatment seeks to extend the conventional analysis in two main ways. First, it allows for time-varying effects of unemployment insurance benefits on jobless duration, as suggested by both the labour-leisure and job search models. For reasons given below, our concern here is with the more potent benefit duration argument than with (changes in) replacement rates. Now the effects of potential benefit duration in influencing the pattern of exit rates from unemployment – a reduction in escape rates with extended benefits and a sharp increase in escape rates at benefit exhaustion – have been recognized in a growing literature, the newer variants of which focus on endogenous policy bias since all such studies exploit changes in unemployment rules over time or across jurisdictions (e.g. Card and Levine, 2000; Lalive and Zweimüller, 2004, Lalive, van Ours, and Zweimüller, 2006; van Ours and Vodopivec, 2006).[1] Here the distinctive feature of our analysis is that we are able to exploit large *and* exogenous discontinuities in maximum potential benefit duration reflecting rules on entitlement that are exclusively a function of age.

The second and more original contribution of the present paper is that it allows for an array of exit options open to the unemployed individual. First of all, the role of destination state is less often encountered in duration analysis. Second of all, although the sparse literature has recognized that unemployment and inactivity are behaviourally distinct states (see Flinn and Heckman, 1983; Jones and Riddell, 1999; Addison and Portugal, 2003), it has generally failed to draw distinctions between different types of employment. The principal exception concerns the distinction between full-time and part-time jobs (Narendranathan and Stewart, 1993b; McCall, 1996). To our knowledge there has been no attempt to investigate whether or not access to benefits in practice serves to mediate between these different routes to reemployment.

From a European perspective, it is germane to distinguish between open-ended or regular contracts of employment and fixed-term contracts. This is because the latter have tended to be the main means of circumventing ambitious employment protection rules. The classic example is of course Spain where fixed-term contracts constitute 35 percent of all dependent employment and around 90 percent of all new contracts (Bover et al., 2000). Fixed-term contracts perform the function of a labour buffer stock and, to complicate matters, also serve as a screening mechanism for inducting workers into open-ended employment (Varejão and Portugal, 2005, 2007). Another distinction worth pursuing in a European context (especially relevant in Mediterranean or southern European nations) is the option of self employment, again for reasons having to do with differential (i.e. more sweeping) employment protection legislation (see OECD, 1999). Finally, jobs provided through the public employment service are a more important exit option in Europe than in, say, the United States because of the greater emphasis placed on active labour market policies in the former region. And, although fixed term in nature, publicly-provided jobs should not be uncritically lumped together with the generality of fixed-term contracts. For all these reasons, a more realistic characterization of the European experience implies the identification of multiple destination states (see Bonnal et al., 1997; Røed and Raaum, 2006). In addition to labour market withdrawal and part-time employment, therefore, we will also consider open-ended employment, fixed-term contracts, government-provided jobs, and self employment.

If individuals do indeed choose between a number of reemployment options (and inactivity), there are several sources of aggregation bias attaching to estimations based on an aggregate hazard function regression model. First, and most obviously, if individuals attach different utilities to the various alternatives to unemployment, regression effects may differ markedly across destinations. Thus, to take unemployment insurance as a case in point, access to benefits can materially influence the choice of destination state because the variable will enter as a negative (and possibly time-varying) cost in the individual utility function. More concretely, if the individual is drawing benefits in a regime that does not allow benefits to be paid in conjunction with part-time employment, it is unlikely that we will observe transitions into such employment prior to the point of benefit exhaustion. Second, the underlying cause-specific baseline hazard functions themselves may differ materially across destination states, thereby yielding differences in the timing of transitions out of unemployment for observationally-equivalent unemployed persons. For instance, if individuals place a higher value on permanent job offers than shorter-term employment opportunities they may be expected (initially at least) to search more intensively over the former type of vacancies. And, to take another example, transitions into part-time employment may only be observed after all hope of obtaining a full-time job is extinguished, at which point hazard rates would spike sharply. In short, use of an aggregate hazard function (and associated regression effects, including access to and duration of benefits) may be expected to compound distinct and possibly even contradictory influences. The disaggregated, competing risks treatment pursued here is designed to account for such differences with the overall objective of offering an improved understanding of unemployment transitions and the role of unemployment insurance in this regard.

Finally, the relevance of the Portuguese case is that it is broadly representative of continental Europe in terms its joblessness and institutional framework. At the same time, the stringency of its employment protection regime (Blanchard and Portugal, 2001) offers an interesting context in which to examine transitions into *atypical* work.

II DATA

Our data are taken from the nationally representative Portuguese quarterly employment surveys *Inquérito ao Emprego* for the period 1992(2)-1997(4), conducted by the *Instituto Nacional de*

Estatistica (INE). The choice of period is dictated by changes in the methodology of the employment survey after the first quarter of 1992. The changes made included new sampling procedures and revisions to the definition of employment, unemployment, and inactivity. (Changes made to the *Inquérito* after 1997 include further, restrictive refinements to the definition of unemployment.)

Each quarter, the INE inquires of a random sample of individuals their current labour market status and past labour history. In this sense, just like the U.S. Current Population Survey, the Portuguese employment survey samples the population of members of a state at a given time and observes their elapsed durations. This sampling plan is referred to as stock sampling, and the elapsed (necessarily incomplete) durations are known as backward recurrence times. Familiarly, the distribution of elapsed durations of a stock (of, say, the unemployed) gives a distorted image of the distribution of complete durations of a flow of entrants (into the unemployment state). This is because the sampling plan over-samples long durations (so-called 'length biased sampling') and contains information only on spells currently in progress. As a result, mean unemployment duration is both over- and under-stated.

Such problems can be partially overcome, however, by a joint modelling of the elapsed duration distribution, the probability of being sampled, and the history of flows into a state (e.g. Flinn, 1986). Yet this procedure may still impose too much structure on the data and require information on entrant flows that is typically unavailable to the researcher. A feasible and much simpler alternative procedure – and one that is followed here – is available if the members of a state at a given time are observed over a fixed time interval. In these circumstances, we can obtain information on the remaining duration (or forward recurrence time) that, conditional on elapsed duration, is distributed as the entrant conditional density function. In essence, one conditions on elapsed duration and weights the likelihood function by the survival function evaluated at the observed elapsed duration at the beginning of the interval (see Lancaster, 1990, p. 183). In the presence of unobserved individual heterogeneity, as will be discussed below, matters become much more complicated.

The quarterly employment survey has a quasi-longitudinal capacity. One sixth of the sample rotate out of the sample each quarter, so that we can track transitions from unemployment for up to five quarters, and hence pursue the conditional approach. Transition rates are then obtained simply by identifying those unemployed individuals in the survey, and their elapsed duration in a given quarter, who move out of unemployment over the subsequent quarter.[2] The destination states of previously unemployed workers can also be identified. For the present purposes, we shall distinguish between six such states: open-ended employment (i.e. permanent jobs), fixed-term contracts, part-time employment, jobs provided by the public employment agency, self employment, and economic inactivity (i.e. withdrawal from the labour force). We note parenthetically that publicly-provided jobs are at subsidized wages in the municipal sector and are fixed-term in nature.

Focusing for the moment on unemployment, each survey contains information on the length of the current unemployment spell in months and the unemployment benefit status of the worker – as either a recipient or nonrecipient of benefits. 'Recipiency' may reflect either receipt of unemployment insurance proper or a lower order of unemployment benefits, termed unemployment assistance. We cannot with precision disentangle the two categories. Under Portuguese law, individuals have to have been employed for at least 18 months during the two years prior to the unemployment event to draw UI benefits proper. Individuals who do not fulfil these requirements can draw unemployment assistance if they have more than six months insured employment in the year preceding unemployment. In addition, workers who have exhausted UI benefits can claim unemployment assistance. In both cases, access to unemployment assistance hinges on per capita family income; only those whose per capita income is less than 80 percent of the minimum wage qualify for unemployment assistance.

For much of our analysis we will not distinguish between types of unemployment benefit recipient. Nevertheless, we will offer necessarily very tentative results for a measure of eligibility for the two types of unemployment benefit. The distinction is based solely on recipient status and tenure on the job that immediately preceded the unemployment event (our data do not contain information on per capita family income). In this exercise, persons recorded as collecting unemployment benefits who had at least 18 months of tenure on the last job are classified as eligible for UI benefits (ELIG). Those individuals drawing benefits with between 6 and 18 months of tenure on the last job are identified as recipients of unemployment assistance (ASSISTANCE). To repeat, this procedure is imprecise; in particular, those classified as recipients of unemployment assistance may in fact be receiving UI benefits if they had built up the necessary service requirement in (unobserved) jobs preceding that immediately prior to unemployment. That being said, the eligibility measure will assist us in going beyond the more aggregative results based on a simple benefits recipiency binary variable.

Note that the UI replacement rate in Portugal is in general 65 percent of the previous wage. That said, lower and upper and lower bounds apply. If the 65 percent value is in practice less than the minimum wage, then that minimum wage substitutes for it; and where this value exceeds three times the minimum wage, the latter amount applies. And in the separate case of unemployment assistance a lower replacement rate obtains, amounting to between 70 to 100 percent of the minimum wage. Unfortunately, the dataset does not contain information on the previous wage that would allow us to calculate replacement rates for these groups.

In addition to modelling the effects of recipiency (and, to a lesser extent, eligibility), we are also concerned to assess the impact of unemployment benefit duration on escape rates from unemployment and transitions to the various destination states. Under Portuguese law, duration of unemployment insurance benefits is exclusively a function of age. The maximum duration of benefits is 10 months for those aged less than 25 years and 12 months for those aged between 25 and 29 years. It then rises in 3-month intervals for each incremental 5 years of age, up to a maximum of 30 months at age 55.[3]

As a practical matter, however, in calculating maximum duration we will assume, first, that all individuals recorded as collecting unemployment benefits are entitled to UI benefits and, second, that they do not go on to receive unemployment assistance. (Observe that even if such individuals did proceed to collect unemployment assistance, the reduced amount of benefits then payable would also produce a spike in the transition rate out of unemployment.) Maximum duration is of interest because it allows us to determine the individual's time to benefit exhaustion on the basis of his or her elapsed jobless duration. Again, maximum potential duration is derived from the unemployment insurance rules, so that any individual going on to collect unemployment assistance is nevertheless assigned a zero time to exhaustion at the exhaustion of regular benefits. This is, then, a conservative approach.[4] Time-varying effects of unemployment benefits can be accommodated with information on the beneficiary's elapsed unemployment duration, either using the same intervals as employed for the baseline hazard or aggregating over certain of those intervals. In addition, nonlinearities can be introduced into time to exhaustion of benefits. Both approaches will be deployed, our favoured approach being the latter.

In sum, from the information in the survey we develop a number of variables to capture the effects of the unemployment benefit system. These are, first, a dummy variable denoting recipient status (compounding UI and unemployment assistance); second, a crude tenure-determined measure of eligibility for each type of benefit; and, third, remaining weeks of benefit entitlement (maximum duration of regular benefits less elapsed jobless duration). As noted, two methods of allowing for time-varying effects are also introduced.

Before turning to note the other arguments, it is important to recognize that the identification of the effect of unemployment benefits on unemployment durations derives in essence from the unemployment insurance rules. The distinction between UI recipients and nonrecipients arises in the first place from the eligibility rules noted earlier; that is, individuals must have been working for at least some time in order to collect benefits. In the second place, some unemployed may have exhausted their benefits and, for this reason, be classified as nonrecipients. This is a different source of identification, and one that may indeed be seen as not strictly exogenous. We would argue in this case, however, that if anything the bias would go against the moral hazard effect of unemployment benefits as nonrecipients would have longer durations and may have depreciated their human capital more profoundly. The third source of identification comes from the definition of maximum potential duration of benefits, which as we have seen is purely age related. These are the rules that we use to define our time to exhaustion variable. For all these reasons, we view the sources of identification of the unemployment benefit effect as basically exogenous. This is not of course to argue that the assignment to benefit status is random. Were that to be the case, we would have a much cleaner experimental environment in which to assess the effects of unemployment benefits.

The employment survey contains in addition to unemployment duration, destination status, and unemployment benefit status, information on the individual's age, marital status, level of schooling, tenure on the lost job, number of jobs held (and whether or not the individual is a new entrant to the labour market), broad occupational status, reason for job loss, and region of residence, inter al. Descriptive information on these and other variables is provided in Appendix Table 1.

The main restrictions placed on the data were that the individual be unemployed at the time of the survey, aged between 16 and 64 years, and resident in mainland Portugal. Further, given well-known gender differences in supply behaviour, we also excluded females. Finally, in recognition of potential sample attrition, we ensured that individuals appearing in contiguous surveys with the same identifier were in fact the same individual. The resulting sample size is 9,451 individuals.

III METHODOLOGY

A useful concept in statistical analysis of duration is the notion of a hazard function. In the study of unemployment duration, the hazard function gives the instantaneous probability of exiting unemployment at *t*, given that the individual stayed unemployed until *t*

$$h(t) = \lim_{\Delta t \to 0} \frac{P(t \leq T < t + \Delta t \mid T \geq t)}{\Delta t} = \frac{f(t)}{1 - F(t)} = \frac{f(t)}{S(t)}, \tag{1}$$

where *f(t)* is the probability density function, *F*(*t*) is the distribution function, *S(t)* is the survival function. A noteworthy relational function is the integrated hazard function

$$\Lambda(t) = \int_0^t h(u)\,du, \tag{2}$$

which relates to the survivor function simply by

$$S(t) = e^{-\Lambda(t)}. \tag{3}$$

In this paper, we consider a flexible form for the hazard function, namely, the piecewise-constant hazard function

$$h(t) = \begin{cases} e^{\lambda_1} & if \quad 0 \le t < c_1 \\ e^{\lambda_2} & if \quad c_1 \le t < c_2 \\ e^{\lambda_3} & if \quad c_2 \le t < c_3 \\ . & \\ . & \\ e^{\lambda_K} & if \quad c_{K-1} \le t \end{cases}, \tag{4}$$

where the time axis is divided into K intervals by points $c_1, c_2 \ldots, c_{K-1}$. In specifying the baseline hazard function, we use eleven intervals. The first six intervals correspond to calendar months, the next two intervals are three months each, while the ninth and tenth intervals are of six months length. The final (open-ended) interval thus covers elapsed durations of twenty-five months or more. In other words, the knot points are 1, 2, 3, 4, 5, 6, 9, 12, 18, and 24.

We shall also distinguish between six exit modes out of unemployment: full-time fixed-term contracts, full-time open-ended contracts, part-time employment, self-employment, public employment, and inactivity. Hence, we define the cause-specific hazard functions to destination j as

$$h_j(t) = \lim_{\Delta t \to 0} \frac{P(t \le T < t + \Delta t \mid T \ge t, J = j)}{\Delta t}, \qquad j=1, 2, \ldots, 6 \tag{5}$$

yielding the aggregate hazard function

$$h(t) = \sum_{j=1}^{6} h_j(t), \tag{6}$$

and the survivor function

$$S(t) = \prod_{j=1}^{6} S^j(t), \tag{7}$$

where $S^j(t) = e^{-\Lambda^j(t)}$, and $\Lambda^j(t) = \int_0^t h_j(u)\, du$.

The model has a conventional competing risks interpretation (see van den Berg, 2001). In this framework, a latent duration *(T_j)* of unemployment attaches to each exit mode. We only observe the minimum of each latent variable. If risks are assumed to be independent, with continuous duration this model simplifies to six separate single-cause hazard models. This conventional assumption of uncorrelated errors or independent risks is highly restrictive. It is made, in our case, in the interest of parsimony and computational tractability.[5] In fact, with six destination states, the parameters from the covariance matrix allowing for arbitrarily correlated risks proved too difficult to estimate using standard maximum likelihood methods.

A popular way to accommodate the presence of observed individual heterogeneity is to specify a proportional hazards model

$$h_j(t;x) = h_{0j}(t)\, e^{x'\beta_j}\,, \tag{8}$$

where $h_{0j}(t)$ denotes the baseline specific hazard function, that is, the hazard function corresponding to null values for the covariates x. In this case, the covariates affect the hazard function proportionally (i.e. $\frac{\partial h_j(x)}{\partial x_k} = \beta_k\, h_j(x)$). An implication of this assumption is that the impact of the covariates does not change (in relative terms) with the progression of the spell of unemployment.

Our information on elapsed duration of unemployment is grouped into monthly intervals (while transitions can only be observed over a fixed interval of three months).[6] Let m denote the elapsed duration (in months) at the time of the survey. The probability that an event occurs over the three-month window after being interviewed, and that such an exit is to destination r, will be given by

$$L_r = \left[S^r(m) - S^r(m+3)\right]\prod_{j\neq r} S^j(m+3) = \frac{\left[S^r(m) - S^r(m+3)\right]}{S^r(m+3)} S(m+3) = f^r(m+3) \tag{9}$$

where we neglect *x* for the sake of parsimony. This expression is simply the product of conditional and marginal probabilities. It corresponds to the product of the probability that an exit occurs and the probability that the exit route is *r*, given that a transition occurred. To keep things simple, we are assuming that exits can solely occur at the boundaries of the interval.

A censored observation (namely, a spell of unemployment that is still in progress after the three-month window) occurs with probability

$$L_C = S(m+3) \; = \prod_{j=1}^{6} S^j(m+3)\,, \tag{10}$$

which is simply the product of the specific survivor functions.

Apart from the discrete nature of the unemployment duration data, we need to pay attention to the type of sampling plan being used in order to avoid the length bias sampling problems induced by stock sampling (Flinn, 1986). Recall that in our sample the stock of unemployed individuals is observed over a fixed interval of three months. In other words, at the time of the first survey the elapsed duration of unemployment is recorded. With this sampling plan, we need to condition on elapsed duration at the time of the first interview in order to recover the entrant density function.

The likelihood contribution for a single individual who exits into destination r over the interval *[m; m+3]*, is given by

$$L_r^{'} = \frac{f^r(m+3)}{S(m)} \quad , \tag{11}$$

and the contribution from a censored observation is

$$L_C^{'} = \frac{S^r(m+3)}{S(m)}\,. \tag{12}$$

We have yet to incorporate the presence of unobserved individual heterogeneity.

Familiarly, this is achieved by assuming a multiplicative error term associated with each specific hazard function, as follows:

$$h_j(t;x) = h_{0j}(t)\ e^{x'\beta_j} v_j\,. \tag{13}$$

In these circumstances, the likelihood contribution for a single individual who exits into destination r over the interval *[m; m+3]*, is will now be given by

$$L_r^{"} = \frac{\int_{v_1}\int_{v_2}...\int_{v_6} f^r(m+3\,|\,v_1,v_2,...v_6)\,dG(v_1)\,dG(v_2)...dG(v_6)}{\int_{v_1}\int_{v_2}...\int_{v_6} S(m+3\,|\,v_1,v_2,...v_6)\,dG(v_1)\,dG(v_2)...dG(v_6)}, \tag{14}$$

and the contribution from a censored observation is

$$L_C^{"} = \frac{\int_{v_1}\int_{v_2}...\int_{v_6} S(m+3\,|\,v_1,v_2,...v_6)\,dG(v_1)\ dG(v_2)...dG(v_6)}{\int_{v_1}\int_{v_2}...\int_{v_6} S(m\,|\,v_1,v_2,...v_6)\ dG(v_1)\ dG(v_2)...dG(v_6)}\,. \tag{15}$$

In modelling unobserved individual heterogeneity in this framework, the econometrician is confronted with the need to make assumptions about either (a) the distribution of the unobserved heterogeneity in the population or (b) the distribution of unobserved heterogeneity in the sample. As here, the population of interest is in general the actual population, not the (length-biased) sample. In this case, it is entirely appropriate to separately integrate the conditional density function (in the numerator) and the weighting function (in the denominator) (see Santos Silva, 2003, pp. 47-48).[7]

Finally, we further assume that the errors v_j are gamma distributed with mean one and variance σ_j^2 and are uncorrelated. We then proceed, after Lancaster (1990, p.66), by redefining the specific survivor function using the well-known result for gamma mixtures, $\bar{S}^j(m) = \left[1+\sigma_j^2\Lambda^j(m)\right]^{-1/\sigma_j^2}$. After this transformation, the likelihood function is derived as

$$L=\left[\prod_{m=1}^{M}\frac{\prod_{j=1}^{6}\left[1+\sigma_j^2\Lambda^j(m)\right]^{-1/\sigma_j^2}-\left[1+\sigma_j^2\Lambda^j(m+3)\right]^{-1/\sigma_j^2}}{\prod_{j=1}^{6}\left[1+\sigma_j^2\Lambda^j(m)\right]^{-1/\sigma_j^2}}\right]^{\delta_{mj}}\left[\prod_{m=1}^{M}\frac{\prod_{j=1}^{6}\left[\left[1+\sigma_j^2\Lambda^j(m+3)\right]^{-1/\sigma_j^2}\right]^{\delta_m-\delta_{mj}}}{\prod_{j=1}^{6}\left[1+\sigma_j^2\Lambda^j(m)\right]^{-1/\sigma_j^2}}\right]\left[\prod_{m=1}^{M}\frac{\prod_{j=1}^{6}\left[1+\sigma_j^2\Lambda^j(m+3)\right]^{-1/\sigma_j^2}}{\prod_{j=1}^{6}\left[1+\sigma_j^2\Lambda^j(m)\right]^{-1/\sigma_j^2}}\right]^{1-\delta_m}$$

where δ_{mj} is a binary indicator function of an exit to j in the *[m;m+3]* interval and $\delta_m=\sum_{j=1}^{6}\delta_{mj}$ simply an indicator of an exit. The likelihood function retains its original structure. The first element refers to the probability of exiting into destination *j*. The second component corresponds to the product of survival probabilities for all destinations other than *r*, that is, the probability of not exiting into any other exit route over the same interval. And the third part accounts for a censored observation. Observations lost due **to** sample attrition, which is very light in our data (less than 2 percent), are treated as exogenous right-censored. In all cases, the conditional probabilities are weighted by the aggregate survival function up to elapsed duration *m*.

IV FINDINGS

Over our sample period, Portuguese unemployment rose by almost two-thirds – from 4.1.to 6.7 percent – and the mean (elapsed) duration of unemployment increased every year from 12.2 months in 1992 to 16.5 months in 1997. Not surprisingly, the distribution of unemployment has changed fairly profoundly; in particular, the share of long-term unemployment (12 months or more) rose by almost 75 percent, such that by the end of the sample period a little over two in five workers had been out of work for more than a year. But the proportion of workers covered by the unemployment benefit system has not changed since 1993. Also, the maximum duration of benefits and the replacement rate (at 65 percent, see above) have remained unchanged. Accordingly, it is the sharp increase in the number of unemployed individuals, and their jobless duration, that explain the near three-fold increase in nominal outlays on unemployment benefits between 1992 and 1997.

(Figure 1 near here)

Against this backdrop, we first consider the probability of escaping unemployment at the most general level (i.e. without distinguishing between destination states). An initial indication of the effects of unemployment benefits on escape rates is provided by the empirical hazard functions in Figure 1. Despite the narrowing in the difference between the escape rates of recipients and nonrecipients through time, there is prima facie evidence of both marked and persistent disincentive effects of access to benefits.

(Table 1 near here)

Results for the basic duration model are given in Table 1. As was noted earlier, the baseline hazard function is specified as an eleven-segment piecewise-constant function. The coefficient estimates in the table show the effects of the regressors in proportionally shifting the baseline hazard up or down. The coefficient estimate of the variable of principal interest indicates that receipt of unemployment benefits (the UB dummy) decreases the chance of exiting unemployment by 42 percent. The assumption that this disincentive effect is constant through time will of course subsequently be relaxed.

The effects of the other covariates can be very briefly described. First, note that we use seven age dummies – the omitted category is individuals aged less than 25 years – to coincide with the age-determined nature of duration entitlement. Absent this specification, it could be argued that the unemployment benefit effect is picking up the effects of aging on jobless duration. As can be seen, this is not the case because the decline in escape rates with age is near monotonic. Second, the effects of the TENURE, DISABLED, SCHOOLING, and MARRIED covariates are thoroughly conventional, with the first two arguments serving to reduce escape rates and the last two being associated with higher escape rates. Third, greater labour market experience/knowledge, indexed by the JOBS variable (and also negatively by FIRST JOB), seems to translate into reduced joblessness. Note in particular that new entrants are 21 percent less likely to exit from unemployment than other job seekers. On the other hand, whether or not a worker lost his job by reason of a mass layoff or through the termination of a fixed-term contract (respectively, LAYOFF and END FIXED) seems to play no part in influencing escape rates. Finally, and as expected, the current unemployment rate is a powerful determinant of escape rates, while the pattern of regional dummies picks up the persistence of unemployment rate differences across broad areas of the country (the high hazard rates of the Centre region being notable in this regard).

(Table 2 near here)

We next consider summary results from alternative characterizations of the effect of unemployment benefits on escape rates. The entry in the first column of Table 2 simply carries over the unemployment benefit (UB) coefficient estimate from Table 1. Specification (2) uses the alternative benefits measure TIMEEX, namely, time to exhaustion of benefits. It will be recalled that this measure pertains to the exhaustion of UI benefits and does not allow for any subsequent receipt of unemployment assistance. (When we reran the regression assuming that all those who received UI benefits proper went on to collect the maximum (again, age-determined) duration of unemployment assistance, the benefits coefficient estimate was somewhat reduced in absolute magnitude but remained statistically and economically significant. Full results of this exercise are available from the authors upon request.) It is apparent that escape rates decline substantially, the further is the insured unemployed worker from benefit expiration; specifically, the hazard rate declines by 4.1 percent for each remaining month of entitlement.

Specification (3) substitutes two imputed benefit measures for one, namely, ELIG and ASSISTANCE. The former variable proxies eligibility for UI benefits, while the latter measures an entitlement to lower-tier benefits in the form of unemployment assistance. As we have seen, each is defined on the basis of the unemployed individual's length of service in the job immediately preceding the unemployment event, given recipiency. Of the two measures, imputed receipt of regular benefits has the stronger effect. The relevant comparison is with the UB coefficient estimate in specification (1). Since the effect of actual receipt of benefits compounds the two effects, it follows that replacement rates drive the result that imputed receipt of UI is stronger than actual benefit receipt. Recall that the replacement rate is lower for UI proper than for unemployment assistance.

The balance of the material in Table 2 allows for time-varying effects in UB receipt (specifications (4) and (5)), as well as nonlinearities in the TIMEEX measure (specification (6)). As far as actual benefit receipt is concerned, specification (4) identifies time-varying effects by using the same intervals as the baseline hazard, whereas specification (5) offers a more parsimonious characterization by aggregating over those intervals. In the former case, it can be seen that the negative effects of benefits on escape rates last for up to two years. In the latter case, the use of a smaller number of intervals confirms the persistence of the benefits effect but perhaps makes more transparent the result that this influence is non-monotonic over the spell of joblessness.

For its part, the introduction of nonlinearities in the effects of TIMEEX provides evidence of rather dramatic disincentive effects, the longer the interval to benefit exhaustion. For example, with two or more years of remaining entitlement, the recipient is 53 percent less likely to escape from unemployment than his uninsured counterpart. At one year the difference is still 48 percent, falling very modestly to 47 percent at six months, and then more steeply to 35 percent at 3 months and to 14 percent at one month.

(Table 3 near here)

We next consider the issue of destination state. Sample means of jobless duration and time to exhaustion of benefits (both in months), as well as unemployment benefit status, are given in Table 3. Comparing the still unemployed (in the next quarter) with individuals entering the six destination states, it can be seen that their elapsed unemployment duration is much longer. The proportion of unemployment benefit recipients is also much greater among the remaining unemployed, with the obvious exception of those securing public employment. Individuals on government-sponsored manpower programs typically draw unemployment benefits prior to enrolment.

From the base of Table 3, it can be seen that the most common form of transition is to fixed-term contracts rather than open-ended employment.[8] In terms of elapsed duration, however, open-ended employment has the shortest associated joblessness. As implied earlier, part-time employment is associated with the most protracted unemployment, although we caution that the number of transitions in this case is rather small. Finally, *vis-à-vis* the remaining destination states, self-employed persons and those entering into fixed-term contracts use up most of their benefits.

(Table 4 near here)

The disaggregated version of the piecewise-constant hazards regression (first presented in Table 1) is given in Table 4.[9] The estimates correct for unobserved individual heterogeneity. It is immediately apparent that the regression coefficients vary widely from destination state to destination state. Abstracting from differences in the effects of unemployment benefits – which will be examined in detail below – there are a number of other interesting results. Thus, for example, the probability of finding employment in open-ended employment and fixed-term contracts is declining in age. But these effects of age are confined to full-time employment. (Not unexpectedly, similar results but of opposing sign are reported for married individuals.). These findings caution against uncritical aggregation by destination state. Another interesting result is that disability is associated with a sharply reduced likelihood of entering into open-ended employment. Although the same is true of labour market inexperience, those who are looking for their first job are also more likely to become inactive which is patently not the case for disabled individuals. Further, the probability of escaping into permanent jobs is negatively associated with the unemployment rate. For other destination states no such statistically significant relation is evident, with the one exception of inactivity. Interestingly, there are indications that labour market withdrawal rises in recovery and falls in recessions.

Perhaps the most interesting differences revealed by Table 4, however, pertain to open-ended employment versus fixed-term contracts. There are material differences in most of the coefficient estimates other than benefit receipt. The main result is that workers exiting into fixed-term contracts are typically high turnover groups. That is, such workers are more likely to be labour market entrants, to have been employed under fixed-term contracts in the past, to have held a larger number of jobs, and to have lower tenure. On the other hand, there is also the seemingly awkward result that workers with greater schooling are also more likely to transition into fixed-term contracts and less likely to find open-ended employment. The likely reconciliation is that there are two rather different processes underlying full-time employment that operate in tandem. The dominant story remains the high turnover one: fixed-term contracts are used as a buffer labour stock by employers, with high-turnover groups locating and taking such employment. But firms also seem to use fixed-term contracts as a screen and to deploy the screen more frequently in the case of educated individuals. Research using other datasets indicates that it is palatable for more educated workers to take employment under fixed-term contracts because such individuals have better prospects of subsequently exiting into open-ended employment (see Varejão and Portugal, 2005).

The sigma values at the foot of Table 4 suggest the presence of unobserved individual heterogeneity for five out of six destination states. But we refrain from placing too much emphasis on the interpretation of these variance parameters of the distribution of unobserved individual heterogeneity for a number of reasons. For the proportional hazards model, Lancaster (1990) shows that when observed and unobserved heterogeneity are independent in the population they cannot be independent in the stock sample. Conversely, independence between observed and unobserved variables in a stock sample implies non-independence in the flow. Further, Cook and Lawless (2007) show that, in general, initial conditions are very hard to tackle with left censoring (meaning stock sampling) and random errors (meaning unobserved individual heterogeneity). With multiple spells one can overcome the problem. In the absence of such data, these authors suggest dropping the error term altogether, and/or including elapsed duration as a regressor. What we can say here, while recognizing the seriousness of the problem, is that our results do not change materially if we do not control for unobserved individual heterogeneity or when we include elapsed duration. We would speculate that the seeming robustness of our regression results to these different specifications may largely be due to the flexibility of the baseline hazard specification, as has been argued by Ridder (1987).

(Table 5 near here)

The rest of our analysis is devoted exclusively to the effects of access to and duration of unemployment benefits on the probability of entering a particular destination state. Panel (a) of Table 5 carries over the UB findings from Table 4 and supplements them with summary results from a specification that substitutes TIMEEX for the binary unemployment benefits measure, UB. The general opening observation is that, again with the obvious exception of publicly-provided employment, there are strong disincentive effects of unemployment benefits across all destination states. Beginning with the UB variable, perhaps the most striking result is the absolute magnitude of the disincentive effect in the case of part-time employment – and, to a somewhat lesser extent, for self employment. Unemployment benefit recipients are respectively about seven times and two and one-half times less likely than non-recipients to enter these states. Neither result is surprising: insured workers have reservation wages that typically exceed the part-time wage, while for self employment the outcome presumably reflects optimal timing considerations (see below). There is no indication that unemployment benefits facilitate entry into stable jobs – compare the very similar point estimates for UB in open-ended employment and fixed-term contracts – but we have already commented on the possibility that for some individuals fixed-term contracts are a means of subsequently accessing open-ended employment.

The time to benefit exhaustion measure, TIMEEX, provides some additional information on the role of unemployment benefits. Disincentive effects for other than those entering public employment are indicated throughout and in each case these monthly values parallel those obtained for the UB binary measure.

(Figure 2 near here)

Panel (b) of Table 5 provides results for nonlinearities in the effects of TIMEEX. To facilitate interpretation, the relationships are also graphed in Figure 2. The figure denotes the percentage changes in transition rates of insured recipients over the entitlement period, the benchmark being nonrecipients. (Part timers and public employment are omitted because of the small number of transitions into these destination states and the large standard errors of the coefficient estimates.) The pattern is roughly consistent across destination states. That is, there is the suggestion that escape rates rise – albeit at different rates – as the benefit period shrinks. Open-ended employment is something of an exception in having relatively high escape rates for the two most remote intervals.

(Figure 3 near here)

Baseline hazard functions, corresponding to the specification in panel (a) of Table 5, are given in Figure 3. The functions apply to an individual with sample average characteristics in respect of the continuous variables SCHOOLING, TENURE, JOBS, and UNEMPLOYMENT RATE, but who is assigned a zero value – that is, the reference categories – for all the dichotomous variables including of course recipient status. Beginning with the main destination states of open-ended employment and fixed-term contracts, it is apparent that the former baseline hazard is characterized by declining escape rates as spell length progresses. The fall is precipitous over the first four months of unemployment although it is much reduced thereafter. For fixed-term contracts, the decline in escape rates is much less evident. Indeed, generally higher hazard rates characterize this destination state. Taken in conjunction, however, the two baseline hazards perhaps contain the suggestion that some unemployed job seekers initially looking for open-ended employment switch to sampling fixed-term contracts after a period of unsuccessful search.

Evidence of possible switching behaviour is more clearly revealed by the U-shaped pattern of the baseline hazard for transitions into part-time work. This configuration suggests that there are likely two types of unemployed individuals transitioning into part-time work: those wanting such jobs *ab initio* (who locate them quickly), and those who ultimately settle on part-time employment after unproductive search for full-time work. For its part, the pattern of the baseline hazard for transitions into government-provided jobs provides only weak evidence that the public authorities target the long-term unemployed.

As far as labour force withdrawal and self employment are concerned, there is evidence of positive duration dependence, with rising escape rates over time. This is most obviously the case for inactivity where the rising profile of escape rates is now more indicative of discouragement. Finally, escape rates into self employment display a pattern reminiscent of the part-time destination state. That is, there is every indication that those who value self-employment find such work readily with others tending to drift into self employment thereafter.

Finally, we need to address a potential problem of our competing risks framework that was flagged in the preceding discussion of the baseline hazards in Figure 3. The issue is whether at any given point in time there is competition amongst each of the six destination states, or whether some exit strategies – perhaps most obviously part-time employment – are not always an option. There are two ways of tackling this potential problem. The first would be formally to model the possibility that some risks are 'defective.' Unfortunately, the large number of destination states considered here would impose too much structure on the model. A second and more direct approach is to exploit the responses of survey participants to questions about their search activity. In particular, our dataset contains three questions indicating restrictions placed by the unemployed on their own search behaviour. The first is *wants a full-time job only*, identifying those individuals claiming solely to be looking for a full-time job (12 percent of unemployed individuals said this was the case). The second question is *wants dependent employment only*, identifying the much larger proportion (78 percent) of unemployed individuals desirous of working for anyone other than themselves. The final question identifies those who *want an open-ended contract only*, namely, individuals (some 5 percent) exclusively seeking permanent employment. Appendix Table 2 presents the frequency of response by destination status.

(Table 6 near here)

Table 6 provides summary results from adding these three arguments to our detailed controls for the specification contained in Table 4 (full results are available from the authors upon request). Our opening remark must be the more general point that the disincentive effects of unemployment benefits (to exits other than public employment) are preserved intact; that is to say, the point estimates for the UB variable are virtually unchanged (cf. the first rows of Tables 4 and 5). Moreover, we would add that the (U-) shape of the baseline hazard is uninfluenced by this potential source of individual heterogeneity.

As far as the performance of the new controls is concerned, this is decidedly mixed. For example, despite claiming to want full-time employment alone, workers responding in this manner are in fact more likely to transition into part-time unemployment than are other unemployed individuals. Again, individuals seeking open-ended employment are no more likely to enter regular employment – the sign of the coefficient estimate is even perverse – and somewhat more likely to enter into a fixed-term employment arrangement than their counterparts. (By the same token, they are also considerably less likely than their other unemployed counterparts to take part-time work and publicly-provided employment – both of which job assignments can be construed as temporary, especially the latter.) In neither case, however, are these results entirely unexpected. Thus, those individuals seeking open-ended employment may need initially to take a fixed-term job that has the potential to become open-ended employment after some trial period, while those wanting full-time jobs may end up with part-time employment as a means of maintaining their human capital intact and preserving their employment options longer term. Only in the case of those individuals claiming to want dependent employment do we find unequivocal confirmation of the indicated restriction on search behaviour: such individuals are more than four and one-quarter times less likely than other unemployed workers to enter into self employment.

V CONCLUSIONS

This paper has used a unique dataset to investigate the effects of unemployment benefits on jobless duration. Apart from its representative nature and useful human capital/demographic content, the dataset contains information on unemployment benefit recipiency, has a quasi-longitudinal capacity, and permits identification of a larger number of destination states than has hitherto been considered in the duration literature. Unlike administrative data, however, it does not contain information on benefit duration (which has to be imputed) or on the amount of benefits received. The former limitation is mitigated by the exclusively age-related nature of benefit duration in Portugal: as we know elapsed duration, simple subtraction permits an estimate of time to exhaustion of benefits. The latter limitation is of decidedly secondary importance because of the general uniformity of replacement rates.

The strengths of the paper are twofold. The major contribution resides in its detailed state space of potential outcomes. Six separate destination states representing behaviourally distinct choices on the part of unemployed job seekers are identified. Our results confirmed that one cannot assume common regression coefficients across destination states. Specifically, the use of an aggregate approach was shown to compound distinct and even contradictory effects of the covariates. The second contribution reflects the major discontinuities in maximum potential benefit duration, in turn yielding large discontinuities in economic incentives. This advantage is furthermore not contaminated by policy endogeneity.

In investigating the effects of the various unemployment benefit measures, large disincentive effects were observed across all destinations with the single (and fully anticipated) exception of publicly-provided employment. Some more specific findings included the huge disincentive effect of unemployment benefits on transitions into part-time employment, muted discouragement effects in respect of transitions into self employment and inactivity, and the similarity in the effect of unemployment benefits as between open-ended employment and fixed-term contracts. As far as the last result is concerned, however, one should resist the interpretation that access to unemployment benefits does not help workers obtain stable jobs. This is because there are seemingly two very different mechanisms at work here. On the one hand, high-turnover workers flow into and out of fixed-term contracts as employers take advantage of the unemployment insurance system. On the other hand, fixed-term contracts also seem to be used as a screening device that eventually leads to permanent jobs.

One result that does surprise is the similarity in the severity of the observed disincentive effects of unemployment benefits across the main ways of exiting unemployment (open-ended employment, temporary employment, self-employment, and inactivity). Seemingly, the moral hazard aspect of subsidization appears to play a vary important role vis-à-vis the intended insurance function of the unemployment benefit system irrespective of the exit options available.

There is one obvious policy implication of the present exercise: part-time employees should be allowed to draw benefits for some period after they transition into part-time employment. Indeed, and reassuringly, Portuguese law has recently been revised to allow this very option. The more general policy implication that maximum duration of benefits should be reduced would not follow if additional duration were found to lead to sufficiently higher wages or better jobs, although here the U.S. experience is not reassuring (see Addison and Blackburn, 2000).[10] And, since we have not modelled the behaviour of *employers*, any inference that the Portuguese unemployment insurance system would benefit from experience rating (so as to discourage the use of excessive buffer stocks of fixed-term contract workers by employers) is speculative. But perhaps a more ecumenical reading would see our results as supporting removal of the age criterion and its substation by one based on previous job attachment.

Endnotes

[1] For the early literature, see Meyer (1990), Katz and Meyer (1990), Fallick (1991), Narendranathan and Stewart (1993a), and Belzil (1995).

[2] Multiple events within a quarter are possible but unlikely and unobserved here. They are unlikely because of the documented sluggishness of the Portuguese labour market, where flows across states are very thin (see, inter al., Blanchard and Portugal, 2001; Addison and Portugal, 2003; Varejão and Portugal, 2007).

[3] For those insured recipients who have exhausted their regular benefits, the maximum duration of unemployment assistance is one-half that due under UI proper. Thus, for a 24-year old, the maximum duration of unemployment assistance would be another five months of income support. For unemployment assistance recipients proper, the maximum duration of these (reduced) benefits is exactly the same as applies for UI benefits, and is again age determined.

[4] Summary results from relaxing the assumption that insured individuals do not subsequently draw unemployment assistance for the relevant age-related period will be noted in passing below. Full results are available from the authors on request.

[5] Allowing for a single (perfectly correlated) error term across destinations imposes even more structure in the model. We note, however, that the estimation results are not perturbed if one actually imposes a common error term for each exit mode.

[6] At this stage a choice has to be made between a discrete-time and continuous-time model. Despite the simplicity of the standard discrete model, which can be estimated by, say, a multinomial logit specification, we chose to work with a continuous time model for three reasons. First, the continuous-time treatment of grouped duration data allows a conventional proportional hazards interpretation of the regression coefficients. Second, it leads to a neat factorization of the likelihood function under flow sampling. And, finally, it lends itself to a more straightforward manipulation in circumstances of stock-sampling and in the presence of unobserved individual heterogeneity. Those advantages are admittedly achieved at the cost of making the somewhat awkward assumption that exists occur solely at the boundaries of the intervals.

[7] This does not, however, fully solve the problem since, with left-censored observations, the distributions of observed and unobserved characteristics depend on the history of whole joint stochastic processes that are necessarily unobserved. Here, we make the strong assumption that observed and unobserved characteristics are uncorrelated in our (snapshot) sample. We note that

Narendranathan and Stewart (1993a) make a similar assumption.

[8] Although the number of fixed-term transitions is high, it is the case that their frequency is markedly lower than in neighbouring Spain. This difference reflects the relatively stricter rules governing fixed-term contracts in Portugal (see Bover et al., 2000).

[9]To save space, age is now redefined to be a continuous variable rather than a categorical variable as formerly**.** The use of age dummies, however, does not change the main thrust of the results. Details are available from the authors upon request.

[10]Shortening potential duration has been found by Van Ours and Vodopivec (2006) not to adversely affect the quality of the subsequent job in a study of changes in Slovenia's UI law that dramatically reduced maximum duration.

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Figure 1: Empirical Hazard Functions by Unemployment Benefit Recipiency Status

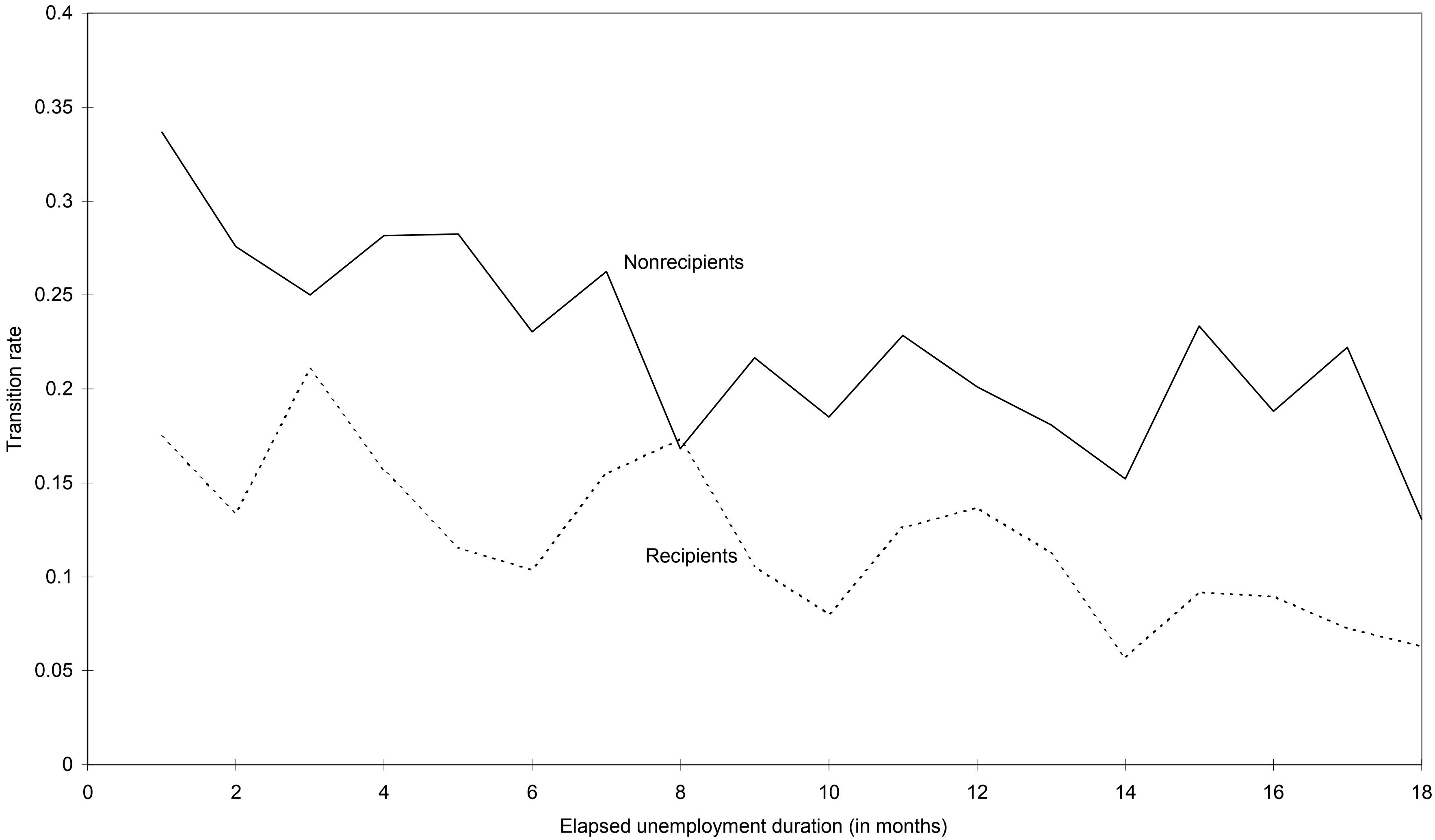


Figure 2: The Relative Effects of Time to Exhaustion on Escape Rates by Grouped Intervals

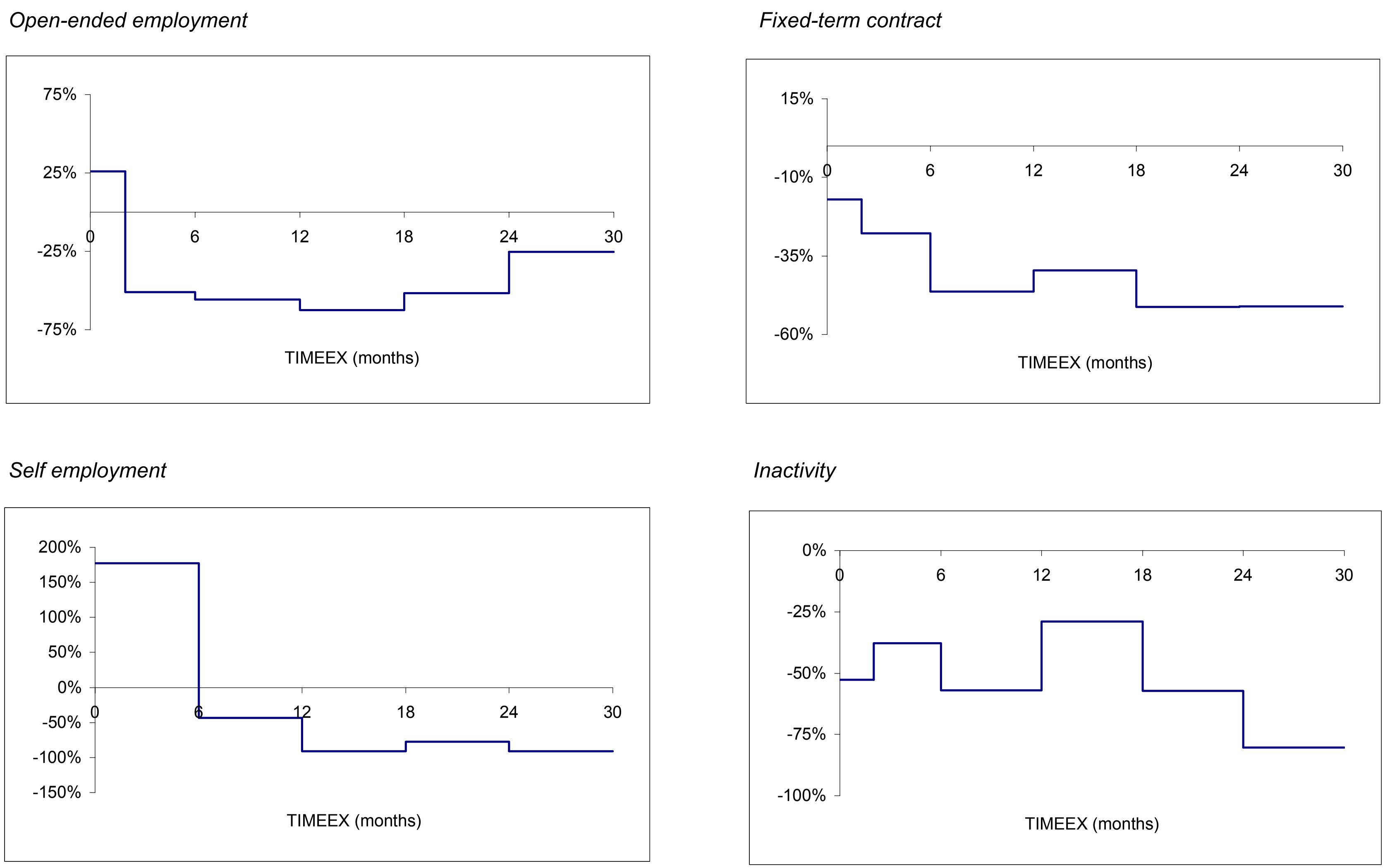
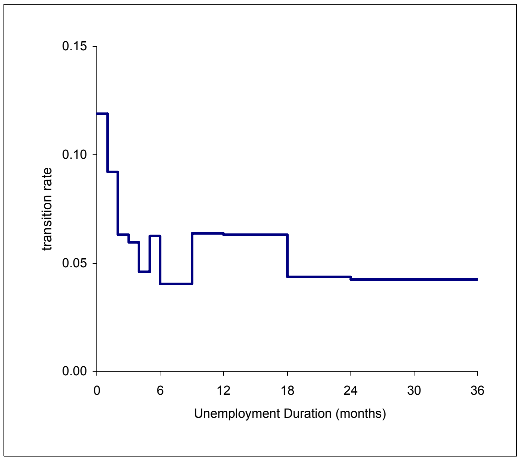
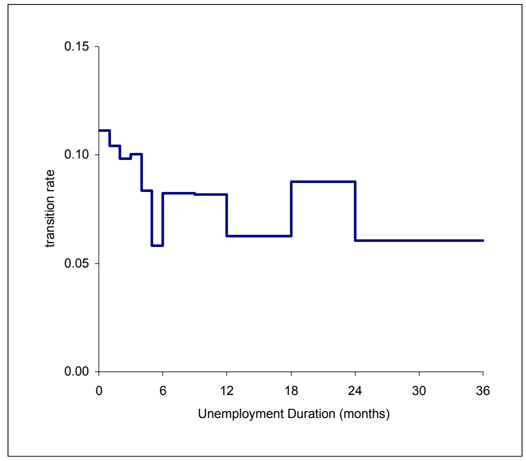


Figure 3: Baseline Hazard Functions by Destination State

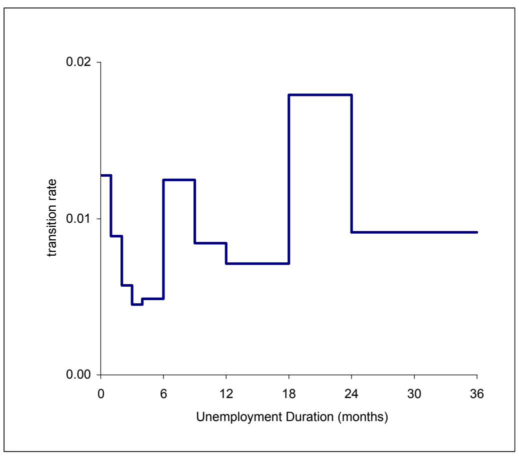
Open-ended employment



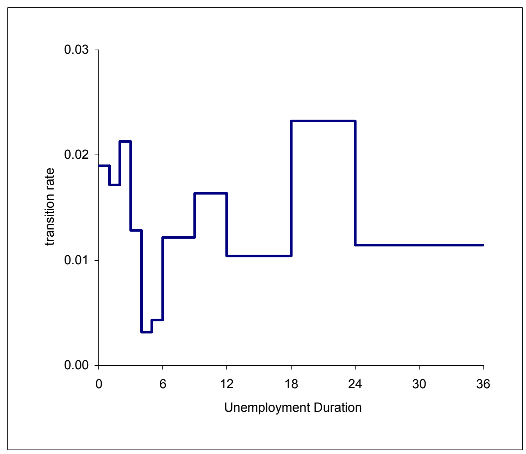
Fixed-term contract



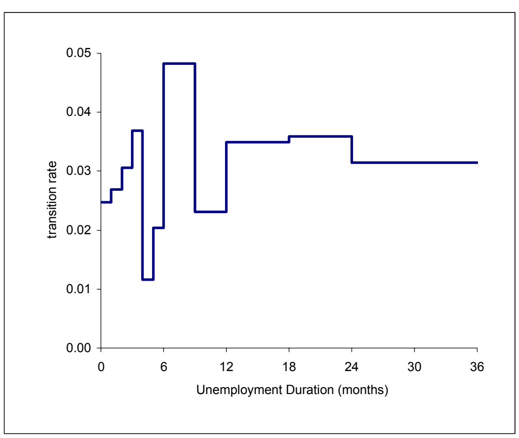
Part-time employment



Public employment



Self employment



Inactivity

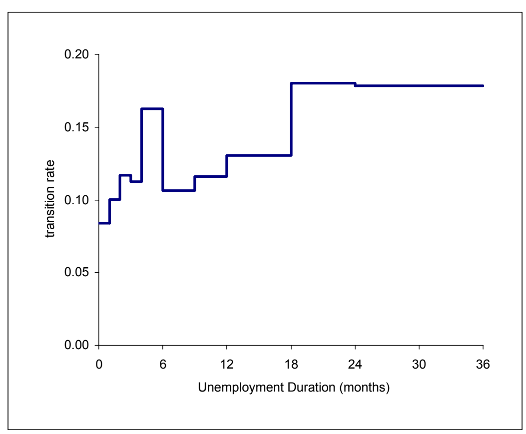


Table 1: Estimated Piecewise-Constant Hazards Regression, Aggregate Model (n=9,451)

Variable	Coefficient Estimate
UB	-0.550
	(0.064)
AGE GROUP	
25-29	-0.030
	(0.075)
30-34	-0.190
	(0.092)
35-39	-0.298
	(0.112)
40-44	-0.143
	(0.110)
45-49	-0.296
	(0.125)
50-54	-0.377
	(0.134)
55+	-0.738
	(0.137)
SCHOOLING	0.015
	(0.008)
TENURE	-0.010
	(0.004)
JOBS	0.021
	(0.007)
WHITE COLLAR	-0.109
	(0.074)
MARRIED	0.244
	(0.071)
DISABILITY	-0.665
	(0.239)
FIRSTJOB	-0.237
	(0.085)
LAYOFF	-0.087
	(0.084)
END FIXED	0.079
	(0.060)
UNEMPLOYMENT RATE	-0.056
	(0.026)
REGIONAL DUMMIES	
NORTH	-0.236
	(0.080)
CENTRE	0.059
	(0.098)
LISBOA	-0.229
	(0.078)
ALGARVE	-0.256
	(0.107)
Log-likelihood	-4361.755

Asymptotic standard errors in parenthesis

Table 2: Summary Results of the Effect of Unemployment Benefits on Transitions Out of Unemployment (n=9,451)

	Specification					
Variable	(1)	(2)	(3)	(4)	(5)	(6)
UB	-0.550 (0.064)					
TIMEEX		-0.042 (0.005)				
ELIG			-0.625 (0.083)			
ASSISTANCE			-0.483 (0.128)			
Recipient Elapsed Duration						
1 month				-0.704 (0.165)		
2 months				-0.738 (0.190)		
3 months				-0.180 (0.174)		
4 months				-0.698 (0.230)		
5 months				-0.917 (0.252)		
6 months				-0.820 (0.268)		
7-9 months				-0.323 (0.159)		
10-12 months				-0.478 (0.168)		
13-18 months				-0.758 (0.193)		
19-24 months				-0.603 (0.244)		
25 months or more				-0.135 (0.197)		
Recipient Elapsed Duration						
1-6 months					-0.638 (0.085)	
7-12 months					-0.396 (0.117)	
13-18 months					-0.756 (0.193)	
19 months or more					-0.333 (0.239)	
Recipient Time to Exhaustion						
1-2 months						-0.147 (0.229)
3-5 months						-0.434 (0.175)
6-11 months						-0.631 (0.099)
12-17 months						-0.647 (0.125)
18-23 months						-0.691 (0.150)
24 months or more						-0.752 (0.228)
Log-likelihood	-4361.8	-4363.1	-4365.2	-4352.9	-4358.7	-4356.6

Note: The full set of covariates are given in Table 1

Table 3: Mean Values of Elapsed Duration and Unemployment Benefit Status by Destination State

	Destination state[a]						
Variable	Open-ended employment	Fixed-term contract	Part time	Public employment	Self employment	Inactivity	Unemployed
DURATION	8.991	9.297	13.865	10.606	11.685	13.905	14.670
UB	0.223	0.206	0.077	0.451	0.260	0.210	0.362
TIMEEX	10.569	9.353	12.200	12.125	9.212	9.781	10.573
Number of events	457	743	104	71	119	305	7652

Note: [a]Individuals that exit unemployment into any of the six categories in the subsequent quarter

Table 4: Estimated Piecewise-Constant Hazards Regression with Gamma Heterogeneity by Destination State (n=9,451)

	Transition to:					
Variable	Open-ended employment	Fixed-term contract	Part time	Self employment	Public employment	Inactivity
UB	-0.652	-0.564	-1.910	-0.962	0.621	-0.605
	(0.142)	(0.113)	(0.513)	(0.314)	(0.324)	(0.215)
AGE	-0.028	-0.033	-0.007	0.008	-0.020	0.008
	(0.008)	(0.006)	(0.018)	(0.017)	(0.018)	(0.009)
SCHOOLING	-0.026	0.031	0.038	0.103	-0.012	0.004
	(0.018)	(0.014)	(0.041)	(0.041)	(0.063)	(0.027)
TENURE	-0.010	-0.026	0.003	0.004	-0.042	0.022
	(0.010)	(0.010)	(0.022)	(0.017)	(0.034)	(0.012)
JOBS	0.005	0.040	-0.020	0.038	0.018	-0.014
	(0.019)	(0.011)	(0.068)	(0.034)	(0.034)	(0.036)
WHITE COLLAR	-0.181	-0.310	-0.110	0.324	-0.253	0.233
	(0.176)	(0.131)	(0.366)	(0.351)	(0.462)	(0.230)
MARRIED	0.416	0.456	0.471	0.611	0.381	-0.368
	(0.157)	(0.124)	(0.352)	(0.362)	(0.374)	(0.258)
DISABILITY	-1.247	-1.071	0.137	-0.991	0.627	-1.677
	(0.520)	(0.459)	(0.851)	(1.249)	(0.692)	(0.744)
FIRSTJOB	-0.772	-0.328	-0.260	-1.289	-0.659	0.761
	(0.213)	(0.144)	(0.464)	(0.615)	(0.633)	(0.250)
LAYOFF	-0.016	-0.112	-0.103	-0.137	-0.427	-0.858
	(0.166)	(0.151)	(0.477)	(0.412)	(0.530)	(0.287)
END FIXED	-0.112	0.341	0.316	-0.203	-0.030	-0.344
	(0.127)	(0.101)	(0.297)	(0.303)	(0.353)	(0.207)
UNEMPLOYMENT RATE	-0.122	0.031	0.061	0.022	-0.041	-0.149
	(0.055)	(0.044)	(0.136)	(0.142)	(0.157)	(0.081)
NORTH	-0.075	-0.405	0.398	-0.173	-1.298	-0.775
	(0.174)	(0.191)	(0.508)	(0.450)	(0.440)	(0.261)
CENTRE	0.036	-0.008	1.260	..391	-1.346	-0.063
	(0.220)	(0.173)	(0.616)	(0.689)	(0.725)	(0.302)
LISBOA	-0.348	-0.170	0.425	0.108	-1.340	-0.674
	(0.185)	(0.136)	(0.509)	(0.455)	(0.493)	(0.249)
ALGARVE	-0.375	-0.405	0.095	0.621	-0.082	-0.877
	(0.247)	(0.191)	(0.694)	(0.545)	(0.489)	(0.360)
sigma	0.429	0.338	0.495	1.201	a	0.734
	(0.273)	(0.141)	(1.425)	(0.521)		(0.200)
Log-likelihooc			-6740.33			

Asymptotic standard errors in parenthesis

Note : [a] gamma variance parameter converged to zero

Table 5: Summary Results of the Effect of Unemployment Benefits on Transitions Out of Unemployment by Destination State (n=9,451)

	Transition to:					
Variable	Open-ended employment	Fixed-term contract	Part time	Self employment	Public employment	Inactivity
Panel (a)						
UB	-0.652	-0.564	-1.910	-0.962	0.621	-0.605
	(0.142)	(0.113)	(0.513)	(0.314)	(0.324)	(0.215)
sigma	0.429	0.338	0.495	1.201	a	0.734
	(0.273)	(0.141)	(1.425)	(0.521)		(0.200)
Log-likelihood			-6740.33			
TIMEEX	-0.044	-0.045	-0.086	-0.102	0.044	-0.048
	(0.010)	(0.008)	(0.029)	(0.024)	(0.027)	(0.014)
sigma	a	0.178	0.826	1.79	a	0.418
		(0.244)	(1.264)	(0.454)		(0.245)
Log-likelihood			-6744.5			
Panel (b)						
Recipient Time to Exhaustion						
1-2 months	0.330	-0.222				-0.677
	(0.481)	(0.410)				(0.823)
3-5 months	-0.746	-0.395				-0.471
	(0.444)	(0.282)				(0.514)
1-5 months			-1.526	1.630	-0.957	
			(1.257)	(1.503)	(1.319)	
6-11 months	-0.828	-0.689	-2.788	-0.575	0.809	-0.858
	(0.212)	(0.163)	(1.302)	(0.625)	(0.417)	(0.313)
12-17 months	-0.910	-0.562	-1.256	-2.557	0.125	-0.397
	(0.282)	(0.205)	(0.802)	(0.889)	(0.718)	(0.295)
18-23 months	-0.757	-0.729		-1.657	1.183	-0.932
	(0.304)	(0.278)		(0.671)	(0.685)	(0.431)
18 months or more			-1.282			
			(0.898)			
24 months or more	-0.292	-0.953		-2.641	0.524	-1.694
	(0.372)	(0.440)		(0.887)	(1.000)	(0.769)
sigma	0.299	0.270	1.02	1.819	a	0.406
	(0.401)	(0.175)	(1.046)	(0.430)		(0.254)
Log-likelihood			-6722.9			

Asymptotic standard errors in parenthesis

Note : [a] gamma variance parameter converged to zero

Table 6: Estimated Piecewise-Constant Hazards Regression with Gamma Heterogeneity by Destination State (n=9,451)

	Transition to:					
Variable[a]	Open-ended employment	Fixed-term contract	Part time	Self employment	Public employment	Inactivity
UB	-0.651	-0.559	-1.990	-0.912	0.609	-0.616
	(0.141)	(0.111)	(0.480)	(0.320)	(0.355)	(0.208)
Wants a full-time job only	0.108	0.019	0.744	0.348	-0.231	-0.854
	(0.191)	(0.145)	(0.390)	(0.425)	(0.505)	(0.394)
Wants dependent employment only	0.295	0.407	0.361	-1.457	0.466	0.325
	(0.138)	(0.117)	(0.353)	(0.344)	(0.507)	(0.239)
Wants an open-ended contract only	-0.211	0.237	-1.507	0.531	-1.048	-0.006
	(0.273)	(0.184)	(0.744)	(0.637)	(1.185)	(0.438)
sigma	0.347	0.275	[b]	1.249	[b]	0.750
	(0.333)	(0.158)		(0.359)		(0.194)
Log-likelihood			-6699.37			

Asymptotic standard errors in parenthesis

Note : [a] The model includes as regressors the folllowing variables: AGE, SCHOOLING, TENURE, JOBS, WHITE COLLAR, MARRIED, DISABILITY, LAYOFF, END FIXED, UNEMPLOYMENT RATE, NORTH, CENTRE, LISBOA and ALGARVE.

[b] gamma variance parameter converged to zero.

Appendix Table 1: Definition of Variables and Descriptive Statistics by Unemployment Benefit Recipiency

	Recipient		Nonrecipient	
Variable	mean	s.d.	mean	s.d.
UNOUT transition out of unemployment =1, 0 otherwise	0.125		0.223	
DURATION elapsed unemployment in months	11.828	11.240	14.888	18.764
AGE age in years	41.816	12.704	30.826	12.381
SCHOOLING years of schooling completed	5.771	3.430	7.117	3.782
TENURE years of tenure on previous job	9.785	10.408	3.900	7.710
JOBS number of previous jobs	3.483	3.925	2.508	3.444
WHITE COLLAR =1 if white-collar employee, 0 otherwise	0.232		0.168	
MARRIED =1 if married, 0 otherwise	0.741		0.341	
DISABILITY =1 if disabled, 0 otherwise	0.015		0.021	
FIRSTJOB =1 if looking for first job, 0 otherwise			0.231	
LAYOFF =1 if job lost by reason of mass layoff, 0 otherwise	0.305		0.090	
END FIXED =1 if job lost through termination of a fixed-term contract, 0 otherwise	0.258		0.258	
UNEMPLOYMENT RATE quarterly unemployment rate	6.640	0.839	6.540	0.940
NORTH =1 for the North region, 0 otherwise	0.400		0.351	
CENTRE =1 for the Centre region, 0 otherwise	0.076		0.087	
LISBOA =1 for the Lisboa and Vale do Tejo region, 0 otherwise	0.347		0.358	
ALGARVE =1 for the Algarve region, 0 otherwise	0.070		0.090	
n	3164		6287	

Appendix Table 2: Transitions by Destination According to Stated Restrictions on Search Behaviour

	Transition to:						
Variable	Open-ended employment	Fixed-term contract	Part time	Self employment	Public employment	Inactivity	Unemployment
Wants a full-time job only	59	95	16	23	5	20	931
Wants dependent employment only	383	645	84	70	62	248	5977
Wants an open-ended contract only	26	59	3	8	1	13	401
n	457	743	104	127	71	305	7644