INVESTIGACIONES ECONÓMICAS. vol. XXVIII (2), 2004, 285-314

ANOTHER LOOK AT UNEMPLOYMENT DURATION: EXIT TO A PERMANENT *VS.* A TEMPORARY JOB

OLYMPIA BOVER
Banco de España
RAMÓN GÓMEZ
European Central Bank

We investigate the determinants of exit rates from unemployment to permanent and temporary jobs. First, we present a theoretical model to discuss the effects of reservation wages, unemployment benefits and job offers on the exit probabilities to permanent and temporary jobs. Then, using micro data from the Spanish Labour Force Survey we estimate a multinomial duration model, including unemployment benefits, the cycle and personal characteristics. Important differential effects are unmasked by distinguishing by type of employment. The negative impact of receiving benefits dominates the combined effect of business cycle variables in exits to temporary employment but not to permanent jobs.

Keywords: Unemployment duration, temporary vs. permanent job, unemployment benefits, business cycle.

(*JEL* J64, J65, E32)

1. Introduction

Aside from a well-known high unemployment rate, the Spanish labour market has another distinctive characteristic: an extensive use of temporary employment. At the end of 1984, new fixed-term contracts (with lower firing costs than the traditional permanent contracts) were introduced in an attempt to ease strong employment protection and foster net job creation[1]. Since then, temporary employment has seen

We would like to thank Pilar García-Perea, the Bank of Spain seminar audience, the editor and a referee for their comments, and Manuel Arellano for very helpful discussions.

[1] For a detailed description of employment protection in Spain, see for example Bover, García-Perea, and Portugal (2000).

unprecedented growth reaching over 30% of total employees in recent years.

The empirical evidence on exits from unemployment to employment usually does not distinguish by type of employment found. This is justified for many countries but not necessarily for those with a high proportion of temporary employment. In Spain, since 1984, the probability of receiving a permanent job offer is much lower than the one of receiving a temporary one, other things equal. In contrast, the influence of benefits on the reservation wage may be less negative if the offered job is permanent rather than temporary, given the likely higher utility attached to higher job stability. Therefore, the results of the papers that do not distinguish by type of employment (notably Bover, Arellano, and Bentolila (2002) for Spain) may compound two very different effects. Indeed, we show in this paper that this is the case for the estimated impacts of some important economic variables.

The goal of this paper is to study the determinants of the exit rates for the unemployed to permanent and temporary employment as distinct alternatives. This will allow us to characterize possible differences and to learn about some relevant issues. For example it seems important to know whether the effect of unemployment benefits on exits to permanent jobs and to temporary ones differ substantially or not[2]. To this end we will first present a theoretical framework and discuss its predictions for the two types of contracts to help us understand the effects we estimate in our empirical analysis.

The data we use are the individual records from the rotating panel of the Labor Force Survey, between 1987Q2 and 1994Q3. This sample period covers a full cycle of the Spanish economy. Importantly, our data are particularly suited to study the effect of receiving unemployment benefits on unemployment duration. Namely, in contrast with data used in other studies in the literature, our data provide exogenous variation across workers in the receipt of unemployment benefits. In 1984, the labor market reform in Spain allowed for the possibility to extensively use short-term contracts with low firing costs. As a result of the reform the unemployed can be thought to be almost randomly assigned to the benefit/non-benefit categories. Therefore workers without benefits coexist with similar workers entitled to benefits. This natural experiment feature of the data has been used in

[2] For details on the unemployment benefit system in Spain see Bover, Arellano, and Bentolila (2002).

Bover, Arellano, and Bentolila (2002) to study the effect of unemployment benefits on unemployment durations by comparing the exit rates of workers with and without benefits given unemployment duration, holding demographics and other variables constant. The present paper can be seen as a follow up that focuses on the potentially different effects of benefits on exits to permanent vs. temporary jobs[3].

Our dataset, however, is different from that of Bover, Arellano, and Bentolila (2002). They used a sample of entrants whereas we use both entrants and the stock of unemployed workers at the time of interview. By focusing on entrants, one avoids the use of, possibly less reliable, retrospective information, but at the cost of having relatively few long durations and no spells longer than 18 months. Here we felt that for distinguishing permanent from transitory exits, a larger sample size and longer durations may be needed. Moreover, we found that the enlarged dataset that included the stock of the unemployed essentially reproduced the single-exit results obtained by Bover, Arellano, and Bentolila (2002) with the sample of entrants.

The paper is structured as follows. Section 2 presents our theoretical framework. Section 3 describes the database used. Section 4 discusses the alternative empirical models and estimation methods. Section 5 presents the estimation results. Finally, Section 6 summarises the conclusions.

2. Theoretical framework

2.1 A reservation wage model

Each period, unemployed workers receive at most one temporary and one permanent job offer with probabilities λ_τ and λ_p, respectively. Offers are indexed by wages W_τ and W_p whose logs have a joint distribution $F(.,.)$, and marginal distributions $F_\tau(.)$ and $F_p(.)$.

Suppose that the unemployed respond to job offers as follows:

1) A permanent job offer with log wage w_p is accepted if w_p exceeds a (log) reservation wage $q : w_p > q$.

2) A temporary job offer with log wage w_τ is accepted if $w_\tau > q + \gamma$, where $\gamma \geq 0$ i.e. the salary offered in a temporary job has to be $\gamma\times$ 100 per cent higher than that offered in a permanent job to be acceptable.

[3] Other papers looking at unemployment exits in Spain are Alba (1996), Antolín (1995), Cebrián *et al.* (1995), and García Brosa (1996).

3) If in the same period two offers are received (one of each kind) such that $w_p \geq w_\tau$, the permanent offer is accepted as long as $w_p > q$.

4) In the event of two offers with $w_\tau - w_p \geq 0$ there are two possibilities

4a) If $w_\tau - w_p \geq \gamma$, the temporary job is accepted provided $w_p > q$.

4b) If $w_\tau - w_p < \gamma$, the permanent job is accepted provided $w_p > q$.

This is a restrictive but convenient model that allows us to investigate the effects of arrival rates and reservation wages on the exit rates to permanent and temporary jobs. We are not concerned with the issue of whether this model can be derived from utility maximization under certain assumptions.

Let T denote unemployment duration, and let D_τ, D_p be indicators of exit to temporary and permanent employment, respectively. The hazard rate is given by

$$\phi_H(t) = Pr(T = t \mid T \geq t)$$

and it can be decomposed as follows

$$\begin{aligned} \phi_H(t) &= \phi_\tau(t) + \phi_p(t) = Pr(T = t, D_\tau = 1 \mid T \geq t) \\ &\quad + Pr(T = t, D_p = 1 \mid T \geq t). \end{aligned}$$

Given the previous model, abstracting from t for the time being, we have

$$\begin{aligned} \phi_\tau &= (1-\lambda_p)\lambda_\tau[1 - F_\tau(q+\gamma)] \\ &+ \lambda_p\lambda_\tau \int_q^\infty Pr(w_\tau > \gamma + z \mid w_p = z)\, dF_p(z) \\ &+ \lambda_p\lambda_\tau Pr(w_p \leq q, w_\tau > \gamma + q) \end{aligned} \quad [1]$$

$$\begin{aligned} \phi_p &= \lambda_p(1-\lambda_\tau)[1 - F_p(q)] \\ &+ \lambda_p\lambda_\tau \int_q^\infty Pr(w_\tau \leq \gamma + z \mid w_p = z)\, dF_p(z) \end{aligned} \quad [2]$$

On the other hand, the probability of remaining unemployed is

$$\begin{aligned} 1 - \phi_H &= (1-\lambda_p)(1-\lambda_\tau) + (1-\lambda_p)\lambda_\tau F_\tau(q+\gamma) + \lambda_p(1-\lambda_\tau)F_p(q) \\ &+ \lambda_p\lambda_\tau Pr(w_p \leq q, w_\tau \leq \gamma + q) \end{aligned} \quad [3]$$

2.2 The effect of unemployment benefits

It is reasonable to assume that, other things equal, benefit earners will have a higher reservation wage q than non-earners. Benefit entitlement may also affect the arrival rates of job offers through differential search intensity, or even the preference for permanent jobs over temporary ones as captured by γ. However, one would expect the leading benefit effects to occur through changes in q.

First, from expression [3] it can be seen that the effect of q on exit rates is unambiguously negative since

$$\frac{\partial}{\partial q}\left(1-\phi_H\right) > 0$$

This is so because $F_\tau\left(q+\gamma\right)$, $F_p\left(q\right)$, and $Pr\left(w_p \leq q, w_\tau \leq \gamma + q\right)$ are all non-decreasing in q.

Next, from expression [2] we can see that the effect of q on the exit intensity to permanent jobs is also negative:

$$\frac{\partial \phi_p}{\partial q} < 0$$

This is so because $[1-F_p(q)]$ is decreasing in q and the integral in the second term of [2] shrinks as q increases.

Finally, let us consider the effect on the exit intensity to temporary jobs. The derivatives with respect to q of the first two terms of [1] are negative, but the sign on the derivative of the third term is ambiguous. When q increases the probability of $w_p \leq q$ increases and the probability of $w_\tau > \gamma + q$ decreases, but the change in the probability of the intersection of the two events $Pr\left(w_p \leq q, w_\tau > \gamma + q\right)$ can be positive or negative.

2.3 A simplified model

We simplify the model by considering the case where $\lambda_\tau = 1$, and (w_p, w_τ) are independently distributed. The assumption of independence of the temporary and permanent wage offer distributions is plausible conditionally on workers characteristics. Moreover, permanent and transitory offers will often be made by different firms. The assumption $\lambda_\tau = 1$ captures the idea that temporary jobs are much more easily available than permanent jobs, which arrive with a (perhaps small) probability λ_p.

In this situation we have:

$$\phi_\tau = (1-\lambda_p)\left[1-F_\tau(q+\gamma)\right] + \lambda_p \int_q^\infty \left[1-F_\tau(\lambda+z)\right] dF_p(z)$$
$$+\lambda_p \left[1-F_\tau(\gamma+q)\right] F_p(q) \qquad [4]$$

$$\phi_p = \lambda_p \int_q^\infty F_\tau(\gamma+z)\, dF_p(z) \qquad [5]$$

Moreover

$$\phi_\tau + \phi_p = 1 - \left\{1-\lambda_p\left[1-F_p(q)\right]\right\} F_\tau(q+\gamma)$$

The main thrust of the formulae can be seen when $\lambda_p = 1$ also.

$$\phi_\tau = \int_q^\infty \left[1-F_\tau(\gamma+z)\right] dF_p(z) + \left[1-F_\tau(\gamma+q)\right] F_p(q)$$
$$\phi_p = \int_q^\infty F_\tau(\gamma+z)\, dF_p(z)$$
$$\phi_\tau + \phi_p = 1 - F_\tau(\gamma+q)\, F_p(q)$$

In words, ϕ_p is the probability that the temporary wage does not exceed the permanent wage in more than γ per cent, and the latter is not smaller than the reservation wage. The transitory intensity ϕ_τ is the sum of two probabilities. The first one is the probability that the temporary wage exceeds the permanent wage in more than γ per cent, and the latter is not smaller than the reservation wage. The second is the probability that the permanent wage is smaller than q but the transitory wage is larger than $q+\gamma$.

2.4 Log normal wage offers

Here we assume that offered wages are log normally distributed:

$$\log\ W_\tau \sim N\left(\mu_\tau, \sigma_\tau^2\right), \qquad \log W_p \sim N\left(\mu_p, \sigma_p^2\right)$$

so that $F_\tau(r) = \Phi\left[(r-\mu_\tau)/\sigma_\tau\right]$ and $F_p(r) = \Phi\left[\left(r-\mu_p\right)/\sigma_p\right]$. We make this assumption to be able to obtain an explicit expression for the integral

$$I(q,\gamma) = \int_q^\infty F_\tau(\gamma+z)\, dF_p(z),$$

which appears in the formulae for ϕ_τ and ϕ_p given in [4] and [5]. In this way we can study in greater detail the effects of changing q and

γ on the transitory and permanent intensities, and be able to make numerical calculations.

Using the formulae for nomal probabilities in the Appendix we obtain the following result:

$$I\left(q,\gamma\right)=\Phi\left(\frac{\gamma+\mu_p-\mu_\tau}{\sqrt{\sigma_p^2+\sigma_\tau^2}}\right)-\Phi_2\left(\frac{q-\mu_p}{\sigma_p},\frac{\gamma+\mu_p-\mu_\tau}{\sqrt{\sigma_p^2+\sigma_\tau^2}};\frac{-\sigma_p}{\sqrt{\sigma_p^2+\sigma_\tau^2}}\right), \quad [6]$$

where Φ_2 (.,.; ρ) is the bivariate standard normal probability with correlation coefficient ρ.

If $\sigma_p^2=\sigma_\tau^2$, we have

$$\frac{-\sigma_p}{\sqrt{\sigma_p^2+\sigma_\tau^2}}=-\frac{1}{\sqrt{2}}\cong -0.7$$

If in addition $\mu_p=\mu_\tau$ we have

$$I\left(q,\gamma\right)=\Phi\left(\frac{\gamma}{\sqrt{2}\sigma_p}\right)-\Phi_2\left(\frac{q-\mu_p}{\sigma_p},\frac{\gamma}{\sqrt{2}\sigma_p};-0.7\right) \quad [7]$$

Recall that $I(q,\gamma)$ is the probability that the temporary wage does not exceed the permanent wage in more than γ per cent and the latter is not smaller than the reservation wage. Clearly,

$$\frac{\partial I\left(q,\gamma\right)}{\partial q}<0.$$

The behaviour of $\partial I\left(q,\gamma\right)/\partial\gamma$ is more complicated because it is the result of two off-setting effects. We have

$$\frac{\partial}{\partial\gamma}\Phi\left(\frac{\gamma}{\sqrt{2}\sigma_p}\right)=\frac{1}{\sqrt{2}\sigma_p}\phi\left(\frac{\gamma}{\sqrt{2}\sigma_p}\right)>0$$

where $\phi(.)$ is the $N(0,1)$ density.

Also, using (A.3):

$$\Phi_2\left(\frac{q-\mu_p}{\sigma_p},\frac{\gamma}{\sqrt{2}\sigma_p};-0.7\right)=\int_{-\infty}^{(q-\mu_p)/\sigma_p}\Phi\left(\frac{\gamma}{\sigma_p}+s\right)\phi\left(s\right)ds$$

Thus,

$$\frac{\partial}{\partial\gamma}\Phi_2\left(\frac{q-\mu_p}{\sigma_p},\frac{\gamma}{\sqrt{2}\sigma_p};-0.7\right)=\frac{1}{\sigma_p}\int_{-\infty}^{(q-\mu_p)/\sigma_p}\phi\left(\frac{\gamma}{\sigma_p}+s\right)\phi\left(s\right)ds>0$$

Therefore, $\partial I\left(q,\gamma\right)/\partial\gamma$ can be positive or negative depending on the value of q.

2.5 Explicit expressions of the exit intensities

The exit intensity to a permanent job in the simplified model is given by

$$\phi_p=\lambda_p I\left(q,\gamma\right) \quad [8]$$

where $I\left(q,\gamma\right)$ is given in [6], or [7] when the distributions of w_p and w_τ are equal.

The exit intensity to a temporary job can be calculated as the difference between ϕ_H and ϕ_p:

$$\phi_\tau=\left(1-\lambda_p\right)\left[1-F_\tau\left(q+\gamma\right)\right]+\lambda_p\left[1-F_\tau\left(\gamma+q\right)F_p\left(q\right)\right]-\lambda_p I\left(q,\gamma\right) \quad [9]$$

Alternatively, using again Lemmas 1 and 2, we can find an expression for the integral

$$I^*\left(q,\gamma\right)=\int_q^\infty\left[1-F_\tau\left(\gamma+z\right)\right]dF_p\left(z\right),$$

which is[4]

$$I^*\left(q,\gamma\right)=\Phi\left[\frac{-\left(\gamma+\mu_p-\mu_\tau\right)}{\sqrt{\sigma_p^2+\sigma_\tau^2}}\right]$$
$$-\Phi_2\left[\frac{q-\mu_p}{\sigma_p},\frac{-\left(\gamma+\mu_p-\mu_\tau\right)}{\sqrt{\sigma_p^2+\sigma_\tau^2}};\frac{\sigma_p}{\sqrt{\sigma_p^2+\sigma_\tau^2}}\right] \quad [10]$$

Thus, we can also calculate ϕ_τ as follows:

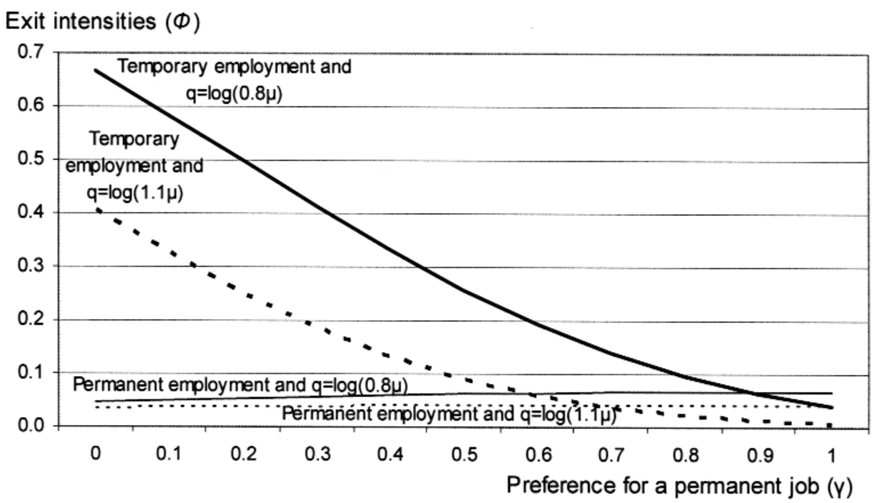
$$\phi_\tau=\left(1-\lambda_p\right)\left[1-F_\tau\left(q+\gamma\right)\right]+\lambda_p I^*\left(q,\gamma\right)+\lambda_p\left[1-F_\tau\left(\gamma+q\right)\right]F_p\left(q\right). \quad [11]$$

[4] The two integrals are related by
$I(q,\gamma)+I^*(q,\gamma)=1-F_p(q)$

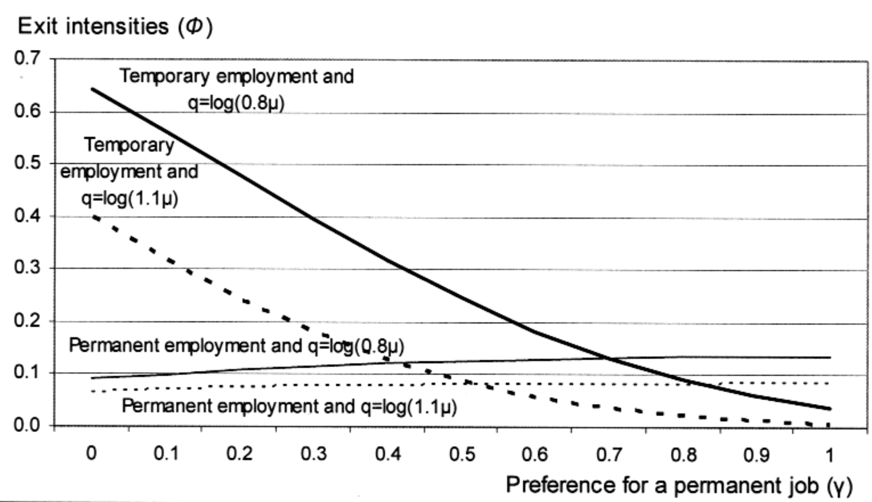
2.6 Some numerical calculations

In Figure 1 we show how the exit intensities to a permanent job ϕ_p (given by [8]) and to a temporary job ϕ_τ (given by [11]) vary as γ increases, for certain values of the model parameters.

FIGURE 1
Numerical exit intensities
(i) $\lambda_p = 0.1$ (permanent offer arrival rate)



(ii) $\lambda_p = 0.2$



In particular, μ and σ^2 are based on the results in Bover, Bentolila, and Arellano (2002) ($\mu = \log$ (90592) and $\sigma^2 = 0.452$)[5]. For the log reservation wage q we consider two values ($q = \log$ (1.1μ) and $q = \log$ (0.8μ)) that could be thought to correspond to the (log) reservation wage of people with and without benefits, respectively. These values have been chosen to roughly match the three-month empirical exit rates for workers with and without benefits.

It is clear from the graphs that in our model the effect of q is negative but more so for the exits to temporary employment. Furthermore, the higher is γ the closer ϕ_p and ϕ_τ become, for a given λ_p. Moreover, the higher λ_p the sooner ϕ_p overtakes ϕ_τ.

3. Description of the data

Our sample comes from the individual data of the Labour Force Survey rotating panel, for the period 1987Q2 to 1994Q3[6]. In this survey households are interviewed for a maximum of six quarters, and each quarter one-sixth of the sample is replaced. Given the large size of the original sample, in Bover, Arellano, and Bentolila (2002) only individuals that started an unemployment spell on one of the dates on which they were interviewed were used. It was considered that the information provided by these workers on the beginning of their period of unemployment would be more reliable than that from those who were already unemployed for over three months at the time of the first interview. However, only periods of unemployment with a maximum duration of eighteen months can be so studied. Given the high level of long-term unemployment in Spain, we thought useful to include in the study unemployment durations over eighteen months as well. Accordingly, after certain preliminary filters, our analysis is based on unemployment durations obtained from the answers of individuals to the questions about how long they have been unemployed and not on the time they are actually observed to be unemployed in the successive interviews. In any case, to avoid stock sample biases, our analysis will always condition on elapsed unemployment duration.

[5] We evaluate σ from the quantiles information using the standard formula
$s = [quantile(i) - quantile(j)]/[\Phi^{-1}(i) - \Phi^{-1}(j)]$
where $\Phi^{-1}(.)$ is the inverse of the $N(0,1)$ probability.

[6] Note that the sample period ends before the 1994 reform, which increased in theory the possibilities of a dismissal to be ruled fair due to economic reasons, and that of 1997 which reduced severance payments for new permanent contracts.

The data available do not reveal information about the other members of the family, so we omit women from the study, the characteristics of their household (number of children and their ages, household income, etc.) being fundamental in their decision whether to seek employment and their incentives to search. Also omitted are the group aged 16-19 years, due to the evident instability in its activity rate, the group aged over 65 due to its high retirement rate, and the unemployed without previous work experience, given our interest in studying the effect of sectoral economic conditions[7].

Even for workers under 65, exit to inactivity is a potential problem. Bover and Gómez (1999) considered separate exits to employment and inactivity for the long term unemployed, but found that separate consideration of exits to inactivity made little difference to measured determinants of exit to employment (compare the results in their Table 3 and Table A.2, column 2). Hence we do not consider an explicit exit to inactivity here.

From this sample of unemployed the transitions to permanent employment or temporary employment as against staying unemployed or inactive are constructed. This is, as in Bover, Arellano, and Bentolila (2002), a wide definition of unemployment, which seems justified in a sample of only men. When defining the unemployment benefit receipt variable, we also consider as recipients those individuals who, at very short durations, do not still receive benefits (owing to administrative delays), but do eventually receive them, in line with Bover, Arellano, and Bentolila (2002). The size of our final sample including unemployed from one to thirty-six months is 110,233 spells.

In this paper we specify single and multiple discrete duration models. As we shall see below, those models can be regarded as a sequence of discrete choice models (with cross-equation restrictions) defined for the population which remains unemployed at each duration (cf. Kiefer (1987), Narendranathan and Stewart (1993a), Sueyoshi (1995), and Jenkins (1995)).

[7] Detailed description about the data and the sample can be found in an Appendix in Bover and Gómez (1999).

4. Empirical specification and econometric methods

4.1 Single-exit discrete duration models

Consider first a single-exit model that includes both entrants into unemployment and the stock of the unemployed workers at the time of interview. For this model, the log-likelihood function is given by the sum of the contributions of the N individuals (i)

$$L(\beta) = \sum_{i=1}^{N} \left\{ (1-c_i) \sum_{t=q_i}^{T_i^0} \log\left[1-\phi_{Hi}(t)\right] \right.$$
$$\left. + c_i \left(\sum_{t=q_i}^{T_i^0-1} \log\left[1-\phi_{Hi}(t)\right] + \log \phi_{Hi}\left(T_i^0\right) \right) \right\}$$

where, following the notation used by Bover, Arellano, and Bentolila (2002): β is the parameter vector to be estimated; c_i is an indicator which takes the value 1 if the end of the period of unemployment is observed, and 0 if it is not; $\left(T_i^0\right)$ is observed duration; q_i is the number of months in unemployment at the time of the first interview (for an entrant $q_i = 1$), $\phi_{Hi}(t)$ represents the hazard rate at $t : \phi_{Hi}(t) =$ Prob $(T = t \mid T \geq t, x(t)) = F\left[\beta_0(t) + \beta_1(t)' x(t)\right]$; $F(.)$ is a cumulative probability function (we use a logistic specification in this paper); $x(t)$ is a vector of individual, sectoral and aggregate characteristics that can vary over time, including an indicator of benefit receipt and interactions between the explanatory variables; $\beta_0(t)$ is a specific parameter of each duration t to capture in a flexible way additive duration dependence, and $\beta_1(t)$ is a vector of polynomials in $\log(t)$ which are introduced to capture interactive effects between the duration and the explanatory variables.

4.2 Multiple-exit discrete duration models

Consider now a model in which there is more than one possible exit from unemployment. Specifically, we are going to distinguish between exits to a permanent job and exits to a temporary job for persons who have been unemployed for between 1 and 36 months.

Specifically, if we have a discrete duration variable T and two alternatives represented by the indicators D_1 and D_2, we can define the following intensities of transition to each of the states:

$$\begin{aligned}\phi_1(t) &= \Pr(T = t, D_1 = 1 \mid T \geq t) \\ \phi_2(t) &= \Pr(T = t, D_2 = 1 \mid T \geq t)\end{aligned}$$

such that the hazard rate from unemployment is given by:

$$\phi_H(t) = \phi_1(t) + \phi_2(t)$$

Likewise, in order to see the discrete duration models as discrete choice models, it is useful to introduce sequences of exit indicators at t to a given alternative:

$Y_{1t} = 1(T = t, D_1 = 1), Y_{2t} = 1(T = t, D_2 = 1)$ for $t = 1, 2, 3\ldots$ According to this notation, $\phi_1(t) = \Pr(Y_{1t} = 1 \mid T \geq t)$ and $\phi_2(t) = \Pr(Y_{2t} = 1 \mid T \geq t)$[8].

Alternatively, we can define exit rates to each of the states conditional upon not exiting to the alternative state:

$$\begin{aligned}h_1(t) &= \Pr(Y_{1t} = 1 \mid T \geq t, Y_{2t} = 0) \\ h_2(t) &= \Pr(Y_{2t} = 1 \mid T \geq t, Y_{1t} = 0)\end{aligned}$$

The relationship with the previous transition intensities is given by:

$$h_1(t) = \frac{\Pr(Y_{1t} = 1 \mid T \geq t)}{\Pr(Y_{2t} = 0 \mid T \geq t)} = \frac{\phi_1(t)}{1 - \phi_2(t)}$$

and similarly

$$h_2(t) = \frac{\phi_2(t)}{1 - \phi_1(t)}$$

Thus, unlike the continuous case, in the context of discrete duration variables and multiple alternatives, we can choose between modeling the intensities $\phi_j(t)$ or the conditional hazard rates $h_j(t)$. For example, if T represents the duration of unemployment and exits 1 and 2 are permanent employment and temporary employment, respectively, $\phi_1(t)$ is the probability of exiting to permanent employment at $T = t$ among those who remain unemployed for at least $T \geq t$ periods. For

[8] Exits 1 and 2 correspond to those labeled p and τ in our theoretical discussion in Section 2. Here we prefer to use numeral indices for simplicity and to facilitate the connection with related formulae in the multiple-exit duration literature.

its part, $h_1(t)$ is the probability of exiting to permanent employment at $T = t$ among those who remain unemployed for at least $T \geq t$ and do not exit to temporary employment at $T = t$.

A specification commonly used in multiple choice problems which we shall also use here is the multinomial logit model. Accordingly, the dependence of $\phi_1(t)$ and $\phi_2(t)$ on the explanatory variables x is specified by

$$\begin{aligned}\phi_1(t) &= \frac{e^{x'\beta_1}}{1 + e^{x'\beta_1} + e^{x'\beta_2}} \\ \phi_2(t) &= \frac{e^{x'\beta_2}}{1 + e^{x'\beta_1} + e^{x'\beta_2}}\end{aligned}$$

Note that, in accordance with the relationships given above, this specification for $\phi_1(t)$ and $\phi_2(t)$ implies that

$$\begin{aligned}h_1(t) &= \frac{e^{x'\beta_1}}{1 + e^{x'\beta_1}} \\ h_2(t) &= \frac{e^{x'\beta_2}}{1 + e^{x'\beta_2}}\end{aligned}$$

That is, if the transition intensities are multinomial logit, the conditional exit rates are binary logit with the same parameters. As a result, the use of the logistic specification leads to the same model in both cases.

Nevertheless, having obtained estimates of the parameters β_1, β_2, we can obtain two different measurements of the effect of the explanatory variables on the probabilities of exit to a specific alternative depending on whether changes in the $\phi_j(t)$ or changes in the $h_j(t)$ are used. Specifically, for a continuous variable and for the exit to alternative 1 we can use

$$\varepsilon_{\phi_1 tk} = \frac{\partial \phi_1(t)}{\partial x_k} \cdot \frac{x_k}{\phi_1(t)}$$

or else

$$\varepsilon_{h_1 tk} = \frac{\partial h_1(t)}{\partial x_k} \cdot \frac{x_k}{h_1(t)}$$

It can be easily shown that the relationship between the two elasticities is given by

$$\begin{aligned}\varepsilon_{h_1 tk} &= \varepsilon_{\phi_1 tk} + \frac{\phi_2(t)}{1 - \phi_2(t)} \varepsilon_{\phi_2 tk} \\ \varepsilon_{h_2 tk} &= \varepsilon_{\phi_2 tk} + \frac{\phi_1(t)}{1 - \phi_1(t)} \varepsilon_{\phi_1 tk}\end{aligned}$$

In addition, in the logistic case:

$$\varepsilon_{h_1 k} = \beta_{1k}\left(1 - h_1\right) x_k$$

where β_{1k} denotes the k^{th} coefficient of the vector β_1.

The differences between the two types of elasticity may be greater when the temporal aggregation of the durations is large. In the description of the results given below we have chosen to assess the effects of the explanatory variables using the changes in the transition intensities $\phi_j(t)$ since this fits more closely with our theoretical framework, although we also provide in certain cases the conditional probabilities $h_j(t)$ to see the extent to which the conclusions may vary in our case.

The models for the conditional probabilities $h_1(t)$, $h_2(t)$ are usually called *competing risk models.* This name derives from the fact that if we consider the existence of two latent duration variables T_1^* and T_2^*, such that the observed duration is $T = \min\left(T_1^*, T_2^*\right)$ and T_1^*, T_2^* are independent, then the conditional exit rates can be interpreted as exit rates for the latent durations:

$$\begin{aligned} h_1(t) &= \Pr\left(T_1^* = t \mid T_1^* \geq t\right) \\ h_2(t) &= \Pr\left(T_2^* = t \mid T_2^* \geq t\right) \end{aligned}$$

That is, to analyse exits to alternative 1 we take the exits to alternative 2 as censored observations, and vice versa.

Note that irrespective of whether T_1^*, T_2^* correspond to well defined concepts (and in the case of exits to permanent or temporary contract it is difficult to imagine that they do), $h_1(t), h_2(t)$ generally represent useful descriptive characteristics for the durations and exits observed[9].

We now turn to consider the estimation of the parameters (β_1, β_2) of the logistic specification. We are going to consider two different methods for estimating the same model. The first consists in the joint estimation of β_1 and β_2 by maximum likelihood, while the second consists in separate estimation of β_1 and β_2 by conditional maximum likelihood. Both methods provide consistent and asymptotically normal estimates of the parameters, although the first estimator is in general asymptotically more efficient than the second. The advantage of the second is basically that its computation is faster. Moreover, separate

[9] Certainly, T_1^*, T_2^*, do not have any meaningful interpretation in the context of our theoretical model.

estimators of the parameters corresponding to one of the alternatives are robust to specification errors in the regression index for the other alternative.

The sample in this paper is very large so that the relative inefficiency of separate estimation of β_1 and β_2 is of little importance. For the same reason, the time to calculate the joint estimation is significantly greater than that of separate estimation, which would hamper the specification searches. Accordingly, we have chosen to use mainly the conditional maximum likelihood estimators, although we also present some joint estimates, which show that the differences between them are very small in our case.

The joint log-likelihood function is given by:

$$L(\beta_1,\beta_2) = \sum_{i=1}^{N} \left\{ (1-c_i) \sum_{t=q_i}^{T_i^0} \log\left[1-\phi_{1i}(t)-\phi_{2i}(t)\right] \right.$$
$$\left. +c_i \left(\sum_{t=q_i}^{T_i^0-1} \log\left[1-\phi_{1i}(t)-\phi_{2i}(t)\right] + D_{1i}\log\phi_{1i}\left(T_i^0\right) + D_{2i}\log\phi_{2i}\left(T_i^0\right) \right) \right\}$$

Likewise, using the sequences of indicators defined above we can express $L(\beta_1,\beta_2)$ as

$$L(\beta_1,\beta_2) = \sum_{t=1}^{\max\left(T_i^0\right)} L_t$$

where

$$\begin{aligned} L_t &= \sum_{i=1}^{N} 1\left(T_i^0 \geq t \geq q_i\right) \{c_i Y_{1ti}\log\phi_{1i}(t) + c_i Y_{2ti}\log\phi_{2i}(t) \\ &\quad + (1-c_iY_{1ti}-c_iY_{2ti})\log\left[1-\phi_{1i}(t)-\phi_{2i}(t)\right]\} \end{aligned}$$

which shows that $L(\beta_1,\beta_2)$ can be regarded as the log-likelihood of a multinomial logit model defined on the basis of the concatenation of the samples surviving at each duration. The joint maximum likelihood estimators $\left(\widehat{\beta}_1,\widehat{\beta}_2\right)$ are defined as the values which maximise $L(\beta_1,\beta_2)$.

In addition, the conditional log-likelihood function for exit 1 is given by:

$$L_{c1}\left(\beta_1\right)=\sum_{i=1}^{N}\left\{c_i\left(D_{1i}\log h_{1i}\left(T_i^0\right)+D_{1i}\sum_{t=q_i}^{T_i^0-1}\log\left[1-h_{1i}\left(t\right)\right]\right)\right.$$
$$\left.+\left[D_{2i}+\left(1-c_i\right)\right]\sum_{t=q_i}^{T_i^0}\log\left[1-h_{1i}\left(t\right)\right]\right\}$$

with a similar expression for the likelihood corresponding to exit 2, $L_{c2}\left(\beta_2\right)$. Note that in $L_{c1}\left(\beta_1\right)$ the exits to alternative 2 are treated as censored observations, so that formally it is a function with exactly the same form as the likelihood with a single exit of the previous section. The implication is that the conditional maximum likelihood estimators, $\left(\widetilde{\beta}_1,\widetilde{\beta}_2\right)$ defined as the maximisers of $L_{c1}\left(\beta_1\right)$ and $L_{c2}\left(\beta_2\right)$, respectively, can be obtained as separate maximum-likelihood estimates of two binary logit models.

Both types of estimators have been used in various papers in the literature, but generally without relating the alternative models and estimation methods to each other. Narendranathan and Stewart (1993b) and Carrasco (1998) obtain estimates by conditional maximum likelihood, while Portugal and Addison (1997) and Alba (1998) obtain estimates by joint maximum likelihood. In the first type of studies the analysis focuses on the conditional exit rates $h_j(t)$ ("competing risk models") while in the second the intensities of transition to the states are studied. Nonetheless, given the logistic specification, both cases use the same model.

5. Results

The analysis of the results is based on Table 1, which provides the sign and significance of the explanatory variables, and on Table 2 and some figures which help to assess the quantitative importance of the effect of the main variables in terms of the transition intensity[10,11].

[10] The calculations in Table 2 and the figures correspond to a set of individual characteristics considered representative.

[11] Some additional comparisons can be found in Bover and Gómez (1999).

TABLE 1
Estimates of logistic hazards of leaving unemployment[1]:permanent and fixed-term employment

Individual Characteristics:	Maximum Likelihood	Conditional Max. Likelihood (ML)				Joint ML	
	Employment 1	Permanent 2a	Fixed-term 2b	Permanent 3a	Fixed-term 3b	Permanent 4a	Fixed-term 4b
Benefits	-0.791	-0.838	-0.792	-0.820	-0.785	-0.796	-0.782
	(13.77)	(6.05)	(13.09)	(5.93)	(12.97)	(5.80)	(12.92)
Benefits x log Dur	0.266	0.199	0.273	0.195	0.274	0.191	0.272
	(9.44)	(2.83)	(9.19)	(2.77)	(9.21)	(2.70)	(9.16)
Benefits x Tenure in previous job	-0.128	-0.083	-0.137	-0.086	-0.139	-0.094	-0.137
	(6.44)	(2.49)	(5.93)	(2.59)	(6.00)	(2.85)	(5.99)
Benefits x Tenure in prev. job x log Dur	0.035	0.0005	0.041	0.002	0.042	0.006	0.042
	(3.47)	(0.03)	(3.66)	(0.09)	(3.74)	(0.30)	(3.73)
Benefits x Tenure in previous job^2	0.003	0.002	0.004	0.002	0.004	0.002	0.004
	(4.48)	(1.41)	(4.23)	(1.54)	(4.36)	(1.74)	(4.32)
Benefits x Tenure in prev. job^2 x log Dur	-0.001	0.001	-0.001	0.001	-0.001	0.0004	-0.001
	(2.09)	(0.76)	(2.57)	(0.67)	(2.68)	(0.53)	(2.62)
Benefits x Age 20-24	0.017	0.073	0.010	0.071	0.009	0.072	0.008
	(0.33)	(0.57)	(0.18)	(0.55)	(0.17)	(0.55)	(0.15)
Benefits x Age 30-44	-0.280	-0.378	-0.264	-0.382	-0.269	-0.388	-0.267
	(3.56)	(2.09)	(3.20)	(2.11)	(3.26)	(2.16)	(3.24)
Benefits x Age 30-44 x log Dur	0.177	0.318	0.161	0.318	0.160	0.318	0.158
	(4.39)	(3.25)	(3.80)	(3.25)	(3.76)	(3.25)	(3.73)
Benefits x Age 45-64	-0.502	-0.628	-0.487	-0.628	-0.490	-0.660	-0.486
	(5.51)	(2.95)	(5.07)	(2.95)	(5.11)	(3.14)	(5.07)
Benefits x Age 45-64 x log Dur	0.173	0.266	0.168	0.266	0.167	0.270	0.164
	(3.56)	(2.30)	(3.25)	(2.30)	(3.24)	(2.35)	(3.20)
Age 20-24	0.046	-0.023	0.059	-0.021	0.055	-0.018	0.054
	(1.42)	(0.29)	(1.74)	(0.27)	(1.61)	(0.23)	(1.59)
Age 30-44	0.009	0.352	0.158	0.351	-0.034	0.196	-0.030
	(0.12)	(2.93)	(2.60)	(2.92)	(0.44)	(1.11)	(0.39)
Age 30-44 x log Dur	-0.167	-0.309	-0.155	-0.306	-0.150	-0.279	-0.151
	(5.74)	(4.90)	(5.06)	(4.86)	(4.90)	(4.10)	(4.93)
Age 45-64	-0.266	0.010	0.008	-0.296	-0.270	-0.389	-0.266
	(2.84)	(0.07)	(0.11)	(1.47)	(2.74)	(1.80)	(2.71)
Age 45-64 x log Dur	-0.235	-0.245	-0.250	-0.242	-0.244	-0.211	-0.244
	-(5.94)	(3.11)	(5.94)	(3.07)	(5.80)	(2.41)	(5.82)
Tenure in previous job	0.035	0.141	0.001	0.144	0.004	0.151	0.001
	(2.31)	(6.05)	(0.05)	(6.16)	(0.20)	(6.54)	(0.06)
Tenure in previous job x log Dur	-0.022	-0.030	-0.015	-0.031	-0.016	-0.034	-0.016
	(2.78)	(2.05)	(1.67)	(2.13)	(1.86)	(2.35)	(1.80)
Tenure in previous job^2	-0.001	-0.003	-0.001	-0.003	-0.001	-0.003	-0.001
	(1.90)	(3.06)	(1.22)	(3.18)	(1.33)	(3.46)	(1.16)
Tenure in previous job^2 x log Dur	0.0002	-0.0002	0.0003	-0.0001	0.0003	-0.0001	0.0003
	(0.74)	(0.30)	(0.87)	(0.22)	(1.01)	(0.08)	(0.89)
Secondary Education	0.001	-0.016	0.001	-0.020	0.002	-0.016	0.004
	(0.03)	(0.32)	(0.04)	(0.39)	(0.11)	(0.32)	(0.20)
University Education	-0.073	0.182	-0.128	0.178	-0.130	0.202	-0.128
	(1.53)	(1.81)	(2.47)	(1.77)	(2.51)	(2.01)	(2.47)
Head of household	0.447	0.453	0.448	0.449	0.441	0.519	0.435
	(10.73)	(8.21)	(10.23)	(8.13)	(10.06)	(5.55)	(9.94)
Head of household x log Dur	-0.072	--	-0.078	--	-0.077	-0.051	-0.074
	(3.18)		(3.28)		(3.23)	(0.96)	(3.11)

TABLE 1
Estimates of logistic hazards of leaving unemployment *(cont.)*

Sectoral and Time Dummies, and Economic Variables	Maximum Likelihood	Conditional Max. Likelihood (ML)				Joint ML	
	Employment 1	Permanent 2a	Fixed-term 2b	Permanent 3a	Fixed-term 3b	Permanent 4a	Fixed-term 4b
GDP growth	0.069			0.108	0.061	0.115	0.061
	(9.94)			(5.91)	(8.41)	(6.27)	(8.39)
Sectoral unemployment rate	-0.018			-0.049	-0.016	-0.045	-0.016
	(3.95)			(5.05)	(3.47)	(3.51)	(3.46)
Sectoral unemployment rate x log Dur	-0.004			--	-0.005	-0.003	-0.005
	(3.00)				(3.03)	(0.72)	(3.06)
Sectoral unemployment rate x	0.012			--	0.013	0.009	0.013
Age 30-44	(3.59)				(3.82)	(0.91)	(3.78)
Sectoral unemployment rate x	0.018			0.022	0.018	0.027	0.018
Age 45-64	(4.59)			(2.21)	(4.46)	(2.48)	(4.45)
Change in the sectoral unemployment rate	0.004			0.0005	0.003	-0.002	0.003
	(0.80)			(0.04)	(0.61)	(0.15)	(0.67)
Sectoral temporary employment ratio	0.002			-0.022	0.007	-0.021	0.007
	(0.96)			(3.60)	(2.67)	(3.33)	(2.63)
Industry	-0.329	-0.250	-0.370	-1.000	-0.201	-0.972	-0.205
	(4.20)	(3.74)	(11.85)	(5.17)	(2.43)	(5.04)	(2.48)
Construction	-0.157	-0.321	-0.260	-0.099	-0.156	-0.142	-0.154
	(4.67)	(5.31)	(9.85)	(1.18)	(4.42)	(1.69)	(4.38)
Services	-0.483	-0.346	-0.498	-1.096	-0.370	-1.058	-0.373
	(6.57)	(5.66)	(18.01)	(5.87)	(4.77)	(5.68)	(4.83)
1988		-0.101	0.087				
		(1.43)	(2.25)				
1989		-0.179	0.167				
		(2.47)	(4.34)				
1990		-0.312	0.194				
		(4.09)	(5.00)				
1991		-0.552	0.125				
		(6.95)	(3.25)				
1992		-1.109	-0.195				
		(13.03)	(5.12)				
1993		-1.264	-0.292				
		(15.28)	(7.90)				
1994		-1.318	-0.137				
		(12.40)	(3.17)				
Second quarter	0.108	0.170	0.104	0.173	0.094	0.173	0.097
	(4.73)	(3.03)	(4.29)	(3.10)	(3.91)	(3.10)	(4.05)
Third quarter	-0.007	0.024	-0.027	0.073	-0.026	0.078	-0.022
	(0.27)	(0.39)	(0.99)	(1.21)	(0.96)	(1.29)	(0.82)
Fourth quarter	-0.115	-0.304	-0.114	-0.196	-0.106	-0.192	-0.104
	(4.66)	(4.83)	(4.29)	(3.18)	(4.12)	(3.12)	(4.03)
Number of spells	86,660	69,528	84,018	69,528	84,018	86,660	
Log likelihood	-43,255	-10,330	-39,524	-10,332	-39,547	-50,682	

[1] Notes:
1 - *t*-ratios in parentheses.
2 - In all the specifications we include monthly duration dummies variables for spells up to 24 months and quarterly duration dummies for 25 to 36 month spells.

TABLE 2
Transition intensities and conditional hazard rates[1]

	Unemployment duration (in months)									
	1		3		7		12		24	
	Transition	Conditional	Transition	Conditional	Transition	Conditional	Transition	Conditional	Transition	Conditional
Exit to a permanent job										
Receiving benefits	0.028	0.037	0.026	0.035	0.020	0.026	0.019	0.024	0.015	0.019
Not receiving benefits	0.050	0.093	0.043	0.072	0.032	0.046	0.029	0.038	0.022	0.026
Not receiving benefits										
Age 20-24	0.047	0.091	0.041	0.070	0.030	0.046	0.028	0.037	0.021	0.026
Age 25-29	0.050	0.093	0.043	0.072	0.032	0.046	0.029		0.022	0.026
Age 30-44	0.064	0.127	0.044	0.073	0.026	0.037	0.021	0.026	0.012	0.014
Age 45-64	0.051	0.095	0.038	0.057	0.023	0.030	0.019	0.022	0.012	0.013
6 months tenure in prev.job	0.041	0.077	0.037	0.062	0.028	0.042	0.027	0.035	0.020	0.025
2 years tenure in prev.job	0.050	0.093	0.043	0.072	0.032	0.046	0.029	0.038	0.022	0.026
High GDP growth	0.065	0.132	0.057	0.104	0.043	0.068	0.041	0.056	0.031	0.039
Mean GDP growth	0.050	0.093	0.043	0.072	0.032	0.046	0.029	0.038	0.022	0.026
Low GDP growth	0.036	0.062	0.031	0.048	0.022	0.030	0.020	0.025	0.015	0.017
High sec.unemployment rate	0.025	0.043	0.022	0.033	0.016	0.021	0.014	0.017	0.010	0.012
Mean sec.unemployment rate	0.050	0.093	0.043	0.072	0.032	0.046	0.029	0.038	0.022	0.026
Low sec.unemployment rate	0.064	0.124	0.055	0.097	0.041	0.063	0.039	0.052	0.029	0.036
High sec.temporary emp.ratio	0.028	0.058	0.025	0.045	0.018	0.028	0.017	0.023	0.013	0.016
Mean sec.temporary emp.ratio	0.050	0.093	0.043	0.072	0.032	0.046	0.029	0.038	0.022	0.026
Low sec.temporary emp.ratio	0.092	0.154	0.079	0.121	0.058	0.080	0.053	0.066	0.039	0.046
Exit to a fixed-term job										
Receiving benefits	0.244	0.251	0.265	0.272	0.249	0.254	0.204	0.208	0.187	0.190
Not receiving benefits	0.464	0.488	0.401	0.419	0.318	0.328	0.230	0.237	0.174	0.178
Not receiving benefits										
Age 20-24	0.478	0.502	0.414	0.432	0.330	0.340	0.240	0.247	0.182	0.186
Age 25-29	0.464	0.488	0.401	0.419	0.318	0.328	0.230	0.237	0.174	0.178
Age 30-44	0.494	0.527	0.399	0.417	0.291	0.299	0.196	0.200	0.134	0.136
Age 45-64	0.462	0.487	0.341	0.354	0.227	0.232	0.142	0.144	0.089	0.090
6 months tenure in prev.job	0.468	0.488	0.409	0.424	0.328	0.337	0.241	0.247	0.185	0.189
2 years tenure in prev.job	0.464	0.488	0.401	0.419	0.318	0.328	0.230	0.237	0.174	0.178
High GDP growth	0.509	0.545	0.447	0.475	0.363	0.380	0.269	0.281	0.207	0.214
Mean GDP growth	0.464	0.488	0.401	0.419	0.318	0.328	0.230	0.237	0.174	0.178
Low GDP growth	0.411	0.427	0.349	0.360	0.270	0.276	0.191	0.195	0.143	0.145
High sec.unemployment rate	0.411	0.422	0.327	0.334	0.237	0.241	0.159	0.162	0.111	0.113
Mean sec.unemployment rate	0.464	0.488	0.401	0.419	0.318	0.328	0.230	0.237	0.174	0.178
Low sec.unemployment rate	0.482	0.515	0.429	0.454	0.352	0.367	0.263	0.273	0.206	0.212
High sec.temporary emp.ratio	0.514	0.528	0.447	0.458	0.358	0.365	0.263	0.267	0.200	0.203
Mean sec.temporary emp.ratio	0.464	0.488	0.401	0.419	0.318	0.328	0.230	0.237	0.174	0.178
Low sec.temporary emp.ratio	0.402	0.442	0.345	0.375	0.272	0.289	0.195	0.205	0.147	0.153

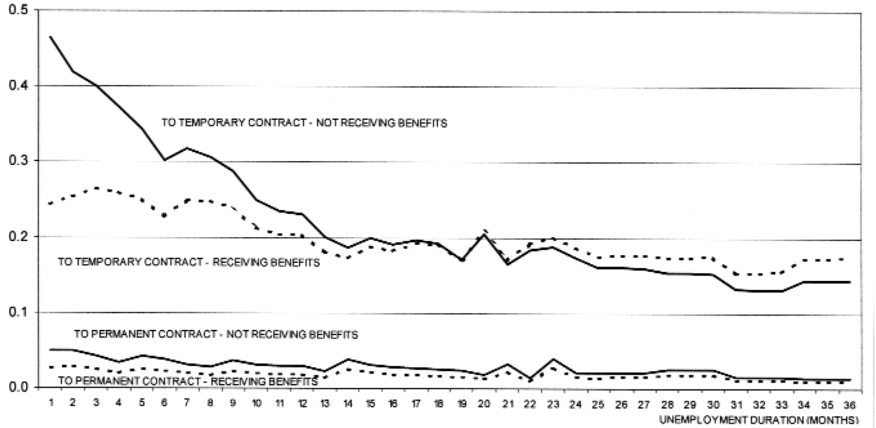
[1] Reference characteristics: 25-29 years old, secondary education, 2 years tenure in previous job, previous job in industry, economic variables at their sample average levels, 2nd quarter.

In the first column of Table 1 we present an estimation of the hazard rates obtained from our sample but without distinguishing by type of employment found. These estimates will allow to see the differences when distinguishing between temporary and permanent employment.

Columns 2 and 3 of Table 1, with a breakdown between the exits to temporary and to permanent employment, present the results of the estimation by conditional maximum likelihood. Aside from personal characteristics, in column 2 sectoral and annual dummy variables appear as regressors. In column 3 economic variables are included without excluding the sectoral dummies to avoid any of the sectoral economic variables capturing only permanent unobservable differences between sectors. The latter is the specification chosen. The last column replicates the third but estimating by joint maximum likelihood.

Before analysing the effect of the different characteristics, it is necessary to point out the important difference in the magnitude of the hazard rates to the two types of employment (see for example Figure 2). These large differences have to be taken into account when assessing the importance of the effects of the different variables and, therefore, such effects must be evaluated not only in absolute but also in relative terms.

FIGURE 2
Predicted hazards and benefits:
Temporary and permanent employment (1)



(1) Men with experience in industry, head of household, secondary education, 2 years tenure in previous job, second quarter, 25-29 years old. Average level of economic variables.

The dependence of the hazard rate on the time spent unemployed is captured through duration dummies, and through the interaction of the explanatory variables with log duration. As expected, the longer the time spent unemployed the lower the hazard rate to a job, whether temporary or permanent. In both cases, the negative relationship is especially significant during the first year, the decline being much smoother thereafter.

Our results show that in order to correctly assess the effect of age it is relevant to distinguish between exits to temporary and to permanent jobs. Indeed, when this distinction is not made, the estimated hazards of mature and young workers are practically identical (see Bover, Arellano, and Bentolila (2002) and the Appendix in Bover and Gómez (1999) for a directly comparable single-exit model for all durations). However, that masks a significantly higher probability of permanent employment for the 30-44 age group for the first few months. After the fourth month, the hazard rates to both types of employment are practically inversely related to age.

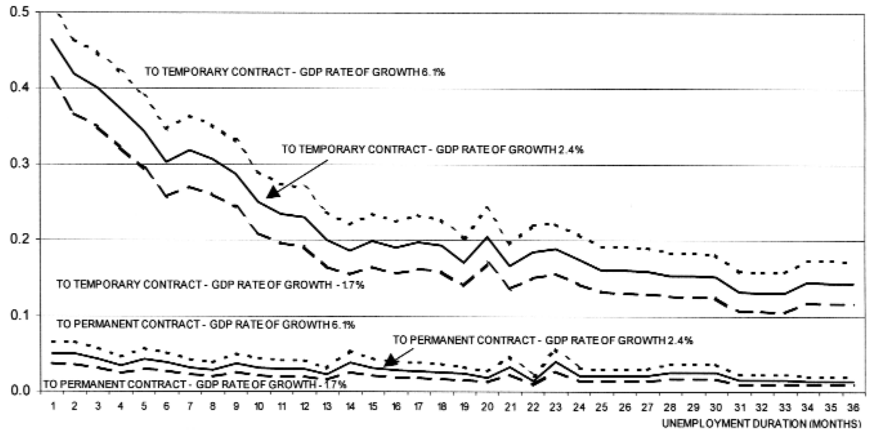
Splitting by type of employment found helps explain the puzzling negative or non-significant effect of university education on the hazard rate to employment in general (see Bover, Arellano, and Bentolila (2002) and the Appendix in Bover and Gómez (1999)). Specifically, having a university degree reduces the hazard rate to a temporary job and increases the one to a to a permanent one.

Receiving benefits affects the hazard rate to both types of jobs negatively, although its effect lessens over time. The effect of benefit is greater, in absolute terms, on exits to a temporary job than to a permanent one; the reduction in the first month is 22 and 2.2 percentage points, respectively. However, in relative terms, the hazard rates fall by approximately one-half in both cases. This reduction induced by the receipt of benefit widens when combined with two variables. The first is the time spent in the previous job which is the fundamental factor when determining the duration of the benefits in Spain. The second is whether a person is over the age of 45, a factor which affects the amount of the benefit entitlement, although there may be other underlying factors, such as the closeness of the transition to retirement.

The economic cycle can be captured by means of annual and quarterly dummy variables along with sectoral dummies or, alternatively, using macroeconomic variables. Specifically, we use the GDP growth rate, which reflects the state of economic activity as a whole, and the unemployment rate and the ratio of temporary to total employees across sectors, as indicators of the situation of the sector in which the person has worked previously.

The results with annual dummies (column 2a and 2b of Table 1) indicate that the hazard rate to temporary employment has a procyclical nature, with increases in the years of net job creation (1998-1991) and decreases in those of net job destruction (1992 onwards). In contrast, for permanent employment the hazard rate diminishes in all years in relation to the baseline year (1987), probably reflecting the trend decline in permanent employment over these years, albeit more so in years of decline in employment as a whole[12].

FIGURE 3
Predicted hazards and GDP growth:
Temporary and permanent employment
Not receiving benefits (1)



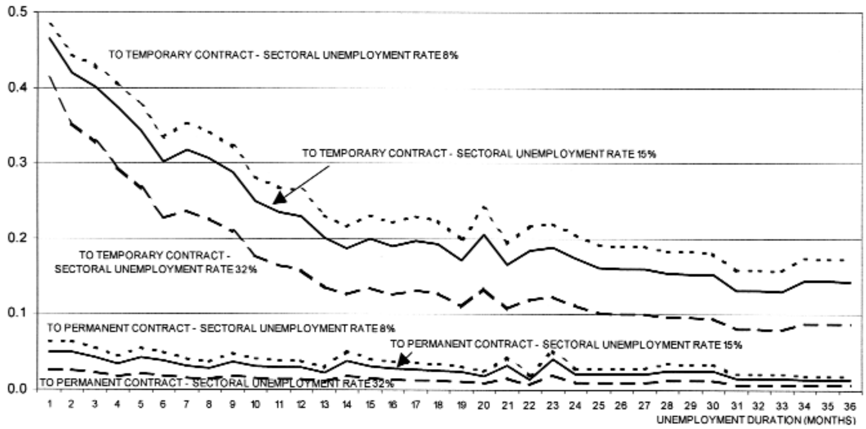
(1) Men with experience in industry, head of household, secondary education, 2 years tenure in previous job, second quarter, 25-29 years old. Average level of economic variables.

Using economic variables (column 3 of Table 1), the positive response of exits to permanent and temporary employment to GDP growth, shown in Figure 3, is clear. Moreover, the sectoral unemployment rate adversely affects the hazard rate to the two types of employment (see Figure 4) but, in the case of temporary employment, especially those younger than 30, and, in permanent employment, those younger than 45. Further, for temporary employment the effect becomes more

[12] For an analysis of workers flows and the cycle, distinguishing between temporary and permanent jobs, see Estrada, García-Perea, and Izquierdo (2002).

negative with duration. Lastly, the sectoral temporary employment ratio proves significant in the two hazard rates although, as expected, it positively affects temporary employment and bears negatively on permanent employment. In contrast, when exits to permanent and temporary employment are not separately considered, the ratio of sectoral temporary employment is not found to have a significant effect on the hazard rate to employment.

FIGURE 4
Predicted hazards and sectoral unemployment:
Temporary and permanent employment
Not receiving benefits (1)



(1) Men with experience in industry, head of household, secondary education, 2 years tenure in previous job, second quarter, 25-29 years old. Average level of economic variables.

It is relevant to compare the effect of unemployment benefits and that of the economic variables, for each type of employment. Consideration is given to the effect of GDP growth, of the unemployment rate and of both jointly, i.e. the unemployment rate observed with the lowest GDP growth (for detailed calculations see Table A.III.1 in Bover and Gómez (1999)). In exits to permanent employment, the effects of receiving benefits and of a high sectoral unemployment rate are very similar and larger than the effect of low GDP growth. However, viewed jointly, the effect of benefits is dominated by the combined effect of GDP and unemployment. By contrast, in exits to temporary employment the effect of benefit exceeds the others to a greater extent; yet, as unemployment duration increases, it is the economic variables which have a larger effect. The importance of benefit relative to the effects of the business cycle variables is, therefore, greater in exits to a temporary job than to a permanent one and this differential impact seems relevant for policy considerations.

The entire analysis of the effects of the various variables on hazard rates from unemployment has been conducted in terms of the transition intensities to each of the states (called $\phi_j(t)$ in Section 4). Table 2 shows the effects of the variables in terms of the conditional hazard rates, $h_j(t)$ in the notation of Section 4, comparing them with the corresponding $\phi_j(t)$. As is to be expected, the hazards of exit conditional upon not exiting to the alternative state are higher, but the conclusions on the effect of the various variables on the hazards of leaving unemployment in each duration do not vary.

Finally note that in our case, as anticipated in Section 4, the results we obtain using joint maximum likelihood (column 4) do not practically differ from the ones obtained by conditional maximum likelihood.

6. Conclusions

In this paper we have studied the determinants of exits to employment distinguishing between exits to permanent and temporary jobs, using data from the Spanish Labour Force Survey (1987Q2 to 1994Q3). Firstly we develop a theoretical framework to inform the empirical discussion and then we estimate multinomial logit duration models with a flexible specification of the duration dependence.

Exit rates to temporary jobs are ten times larger than exit rates to permanent jobs. However, the reduction in exit rates as unemployment duration increases is larger for exits to temporary jobs than to permanent ones. The effect of receiving unemployment benefits is to halve the exit rates in both cases, although these strong effects slowly die out. In contrast, the effect of GDP growth or sectoral unemployment is smaller but longer lasting in both cases. When comparing the relative magnitude of the effects of receiving benefits and the cycle, the negative impact of receiving benefits dominates the combined effect of business cycle variables during the first six months in exits to temporary employment but this is not the case for exits to permanent jobs.

Other differential effects that are unmasked by distinguishing by type of employment found are those of age and university education. Regarding age, the 30 to 44 group have a higher probability to exit into a permanent job at the beginning of the unemployment spell, but after the fourth month, the exit rates to both types of employment are inversely related to age. When not distinguishing by type of employment, the estimated hazards of mature and young workers have been found to be practically identical. University education is found to increase the probability of exit to a permanent job while reducing the exit probability to a temporary one. This explains the counter-intuitive negative or non-significant effect of having a university degree on the overall probability of exit to employment found when no distinction by type of job is made.

Appendix: Exit rates under log normal wages

We first collect two results in normal probabilities that are used in the calculation of the intensities when the wage offer distributions are assumed to be log normal.

LEMMA 1 *Let $X \sim N(\mu, \sigma^2)$ and let $\Phi(.)$ be the standard normal cdf. Let a and b be arbitrary coefficients. Then*

$$E[\Phi(a+bX)] = \Phi\left(\frac{a+b\mu}{\sqrt{1+b^2\sigma^2}}\right) \qquad [A.1]$$

PROOF. Define $V \sim N(0,1)$ independent of X, and form the variable $W = V - a - bX$. Then

$$W \sim N\left(-a-b\mu, 1+b^2\sigma^2\right)$$

so that

$$\Pr(W \le 0) = \Phi\left(\frac{a+b\mu}{\sqrt{1+b^2\sigma^2}}\right),$$

but also

$$\begin{aligned} \Pr(W \le 0) &= \Pr(V \le a+bX) \\ &= E[\Pr(V \le a+bX \mid X)] = E[\Phi(a+bX)], \end{aligned}$$

which proves the result.

LEMMA 2 *Let $\Phi_2(.,.;\rho)$ be the bivariate standard normal cdf with correlation coefficient ρ, and let $\Phi(.)$ and $\phi(.)$ be the standard normal cdf and pdf, respectively. Let m, a, and b be arbitrary coefficients. Then we have*

$$\int_{-\infty}^{m} \Phi(a+bs)\,\phi(s)\,ds = \Phi_2\left(m, \frac{a}{\sqrt{1+b^2}}; \frac{-b}{\sqrt{1+b^2}}\right) \qquad [A.2]$$

PROOF. Let (Z_1, Z_2) be random variables with joint *cdf* $\Phi_2(.,.;\rho)$. That is,

$$\Pr(Z_1 \le r_1, Z_2 \le r_2) = \Phi_2(r_1, r_2; \rho) \\ = \int_{-\infty}^{r_1} \Pr(Z_2 \le r_2 \mid Z_1 = s)\,\phi(s)\,ds$$

Since $(Z_2 \mid Z_1 = s) \sim N(\rho s, 1-\rho^2)$, we have

$$\Phi_2(r_1, r_2; \rho) = \int_{-\infty}^{r_1} \Phi\left(\frac{r_2 - \rho s}{\sqrt{1-\rho^2}}\right)\phi(s)\,ds \qquad \text{[A.3]}$$

Now we can obtain

$$\Phi_2\left(m, \frac{a}{\sqrt{1+b^2}}; \frac{-b}{\sqrt{1+b^2}}\right) = \int_{-\infty}^{m} \Phi\left(\frac{\frac{a}{\sqrt{1+b^2}} + \frac{b}{\sqrt{1+b^2}}s}{\sqrt{1-\frac{b^2}{1+b^2}}}\right)\phi(s)\,ds = \int_{-\infty}^{m} \Phi(a+bs)\,\phi(s)\,ds,$$

which proves the result.

Next, we wish to prove that assuming $F_\tau(r) = \Phi\left(\frac{r-\mu_\tau}{\sigma_\tau}\right)$ and $F_p(r) = \Phi\left(\frac{r-\mu_p}{\sigma_p}\right)$, the integral $I(q,\gamma) = \int_q^\infty F_\tau(\gamma+z)\,f_p(z)\,dz$ is given by

$$I(q,\gamma) = \Phi\left(\frac{\gamma+\mu_p-\mu_\tau}{\sqrt{\sigma_p^2+\sigma_\tau^2}}\right) - \Phi_2\left(\frac{q-\mu_p}{\sigma_p}, \frac{\gamma+\mu_p-\mu_\tau}{\sqrt{\sigma_p^2+\sigma_\tau^2}}; \frac{-\sigma_p}{\sqrt{\sigma_p^2+\sigma_\tau^2}}\right)$$

To be able to use Lemmas 1 and 2, we rewrite $I(q,\gamma)$ as

$$I(q,\gamma) = \int_{-\infty}^{\infty} F_\tau(\gamma+z)\,f_p(z)\,dz - \int_{-\infty}^{q} F_\tau(\gamma+z)\,f_p(z)\,dz \qquad \text{[A.4]}$$

where now

$$F_\tau(\gamma+z) = \Phi\left(\frac{z+\gamma-\mu_\tau}{\sigma_\tau}\right)$$
$$f_p(z) = \frac{1}{\sigma_p}\phi\left(\frac{z-\mu_p}{\sigma_p}\right)$$

Using Lemma 1, the first term of the right-hand side of [A.4] equals

$$\int_{-\infty}^{\infty} F_\tau(\gamma+z)\,f_p(z)\,dz = \Phi\left(\frac{\frac{\gamma-\mu_\tau}{\sigma_\tau} + \frac{\mu_p}{\sigma_\tau}}{\sqrt{1+\frac{1}{\sigma_\tau^2}\sigma_p^2}}\right) = \Phi\left(\frac{\gamma+\mu_p-\mu_\tau}{\sqrt{\sigma_p^2+\sigma_\tau^2}}\right) \qquad \text{[A.5]}$$

Next, we introduce a change of variable in order to evaluate the second term:

$$v = \frac{z - \mu_p}{\sigma_p},$$

so that $z = \mu_p + \sigma_p v$ and $dz = \sigma_p dv$. Thus, changing the index of the integral and using Lemma 2:

$$\int_{-\infty}^{q} F_\tau (\gamma + z) f_p (z) dz = \int_{-\infty}^{(q-\mu_p)/\sigma_p} \Phi \left(\frac{\gamma + \mu_p - \mu_\tau}{\sigma_\tau} + \frac{\sigma_p}{\sigma_\tau} v \right) \phi (v) dv$$

$$= \Phi_2 \left(\frac{q - \mu_p}{\sigma_p}, \frac{\frac{\gamma+\mu_p-\mu_\tau}{\sigma_\tau}}{\sqrt{1 + \frac{\sigma_p^2}{\sigma_\tau^2}}}; \frac{-\frac{\sigma_p}{\sigma_\tau}}{\sqrt{1 + \frac{\sigma_p^2}{\sigma_\tau^2}}} \right)$$

$$= \Phi_2 \left(\frac{q - \mu_p}{\sigma_p}, \frac{\gamma + \mu_p - \mu_\tau}{\sqrt{\sigma_p^2 + \sigma_\tau^2}}; \frac{-\sigma_p}{\sqrt{\sigma_p^2 + \sigma_\tau^2}} \right) \qquad [A.6]$$

Finally, subtracting [A.5] from [A.6] the result in [6] follows.

References

Alba, A. (1996): "Explaining the transitions out of unemployment in Spain: the effect of unemployment insurance", Working Paper 96-71, Economic Series 29, Universidad Carlos III de Madrid.

Alba, A. (1998): "Re-employment probabilities of young workers in Spain", *Investigaciones Económicas* 22, pp. 201-24.

Antolín, P. (1995): "Transitions probabilities to employment and non-participation", Working Paper EC 95-20, Instituto Valenciano de Investigaciones Económicas.

Bover, O., M. Arellano, and S. Bentolila (2002): "Unemployment duration, benefit duration, and the business cycle", *The Economic Journal* 112, pp. 223-65.

Bover, O., S. Bentolila, and M. Arellano (2002): "The distribution of earnings in Spain during the 1980s: the effect of skill, unemployment, and union power", in D. Cohen, T. Piketty, and G. Saint-Paul (eds) *The Economics of Rising Inequalities,* CEPR and Oxford University Press, pp. 3-53.

Bover, O. and R. Gómez (1999): "Another look at unemployment duration: long-term unemployment and exit to a permanent job", Documento de Trabajo 9903, Banco de España.

Bover, O., P. García-Perea, and P.Portugal (2000): "Labour market outliers: lessons from Portugal and Spain", *Economic Policy* 31, pp. 381-428.

Carrasco, R. (1999): "Transitions to and from self-employment in Spain: an empirical analysis", *Oxford Bulletin of Economics and Statistics* 61, pp. 315-41.

Cebrián, I., C. García, J. Muro, L. Toharia, and E. Villagómez (1995): "Prestaciones por desempleo, duración y recurrencia del paro", in J. Dolado y J. Jimeno (comp.). *Estudios sobre el funcionamiento del mercado de trabajo español,* Fundación de Estudios de Economía Aplicada, Madrid.

Estrada, A., P. García-Perea and M. Izquierdo (2002): "Los flujos de trabajadores en España: el impacto del empleo temporal", Documento de Trabajo 0206, Banco de España.

García Brosa, G. (1996): "Prestaciones por desempleo y duración del paro", Colección Estudios, Consejo Económico y Social.

Imbens, G. and L. Lynch (1994): "Re-employment probabilities over the business cycle", mimeo, Harvard University.

Jenkins, S. (1995): "Easy estimation methods for discrete-time duration models", *Oxford Bulletin of Economic and Statistics* 57, pp. 120-138.

Kiefer, N. (1987): "Analysis of grouped duration data", Working Paper 87-12, Cornell CAE.

Narendranathan, W. and M. Stewart (1993a): "How does the benefit effect vary as unemployment spells lengthen?", *Journal of Applied Econometrics* 8, pp. 361-81.

Narendranathan, W. and M. Stewart (1993b): "Modelling the probability of leaving unemployment: competing risks models with flexible base-line hazards", *Applied Statistics* 42, pp. 63-83.

Portugal, P. and J. Addison (1998): "Unemployment insurance and joblessness: a discrete duration model with multiple destinations", Discussion Paper 99-03, Mannheim, Centre for European Economic Research.

Sueyoshi, G. (1995): "A class of binary response models for grouped duration data", *Journal of Applied Econometrics* 10, pp. 411-31.

Resumen

Estudiamos los determinantes de las tasas de salida del paro a un empleo fijo o temporal. Primero presentamos un modelo teórico de los efectos de los salarios de reserva, la prestación por desempleo y las ofertas laborales sobre las tasas de salida a empleos permanentes y temporales. A continuación, usando datos individuales de la Encuesta de Población Activa, estimamos un modelo de duración multinomial incluyendo las prestaciones, el ciclo y características personales. Distinguiendo por tipo de empleo aparecen importantes diferencias. El impacto negativo de recibir prestaciones domina el efecto del ciclo en las salidas a empleos temporales pero no a empleos fijos.

Palabras clave: Duración del desempleo, empleo temporal vs. fijo, prestaciones por desempleo, ciclo económico.

Recepción del original, junio de 2002
Versión final, abril de 2003