Job Separation in a Non-stationary Search Model: A Structural Estimation to Evaluate Alternative Unemployment Insurance Systems

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Abstract

This paper considers a job search model in which the environment is not constant throughout the unemployment spell and where jobs do not last forever. In this situation, reservation wages can be lower than they would be in a model without consideration of such separations, but they can initially be higher precisely because of the non-constant environment.

The model is estimated structurally by using Spanish data for the period 1985-1996. The main finding is that, after controlling for unobserved heterogeneity, the unemployment hazard rate is almost flat during the first six months. However, after this duration, the highly decreasing job offer arrival rate comes to be the only significant factor, given that acceptance probabilities become equal to one. The estimated parameters are used to evaluate different Unemployment Insurance designs. We conclude that a non-monotonic pattern in unemployment benefits, joint with a tax paid by workers and based on unemployment duration makes this duration to be 13.2% lower than it currently is in Spain.

JEL Classification: C41, J64

Key Words: Job Search, Nonstationarity, Unemployment, Separation probability, Structural estimation, Unemployment Insurance.

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1 Introduction

Recently, a great deal of research has been carried out on the job search behavior of unemployed workers. The analysis of unemployment duration has become an important way to better understand the issues behind the unemployment rate as an aggregate figure. Job search models study the problem of an unemployed worker searching for a new job in a dynamic and uncertain world. A traditional assumption in these models has been *stationarity* within the unemployment spell: parameters determining worker behavior are commonly assumed to be constant throughout the unemployment spell. But this assumption is often at variance with reality: reduced-form estimations of job search models usually show a manifest negative duration dependence of the re-employment probability, even when unobserved heterogeneity is considered (among others, see Meyer (1990) for US data, Narendranathan and Stewart (1993) for UK data, and Bover, Arellano and Bentolila (2002) for Spanish data). This time dependence turns the solution to the worker's problem into non-stationary throughout his unemployment spell. Such time dependence is supported by, for example, the lower number of offers received by long-term unemployed workers or the usual decreasing pattern in unemployment benefits.

The contribution of this paper is the introduction of a new element not considered in previous non-stationary job search models: an exogenous separation probability, which can represent both firing and quitting. In this situation the unemployed worker knows that once employed, he can leave or may be compelled to leave the job in the future and become unemployed once again. One of the most influential articles in the field of non-stationary job search, Van den Berg (1990), considers nonstationarity in a very general way but assumes that jobs are held forever, even though he admits the restrictions of such an assumption. The present paper incorporates the separation probability in a non-stationary search model. It is well-established that consideration of this fact makes the reservation wage more strongly time-dependent than it would be otherwise.[1] Moreover, in a stationary model reservation wages are always lower when the separation probability is greater.[2] The reason is that the future is discounted at a higher rate. However, when it is acknowledged that reservation wages change during the time the worker is unemployed, it is demostrated in this paper that this effect is not the only one possible: we may also see higher reservation wages with a higher separation rate, essentially at the beginning of the unemployment spell. This result means that, in some situations, the unemployed worker is more selective when accepting job offers because he knows that the probability of subsequently losing his job is not zero. But this will only happen at the beginning of the unemployment spell, when his situation is not so bad. As time passes, the worker will

[1] Ahn and García-Pérez (2002) show how reservation wages should be highly decreasing in Spain, given the pattern followed by self-reported wage aspirations in this country.

[2] In a model where search is also allowed once in the job, the effect of the separation probability depends on the relationship between the offer arrival rate while employed and unemployed.

increasingly want to be employed as soon as possible, because not only can he achieve better living conditions once in a job, but also because he realizes that, even after a possible future separation, his situation as newly unemployed will be better than the one he has now. This new result is of considerable importance because it offers an alternative explanation for long unemployment duration in countries like Spain, where there is a large separation rate: it is not only unemployment benefits which cause reservation wages to be high and hence the exit rate to be low at the beginning of the unemployment spell, but also the interaction of such benefits, and mainly their initial value, with the separation rate.[3]

This paper also has an empirical objective: to structurally estimate the non-stationary model by using Spanish data for the period 1985-1996. Given that we have to estimate the model without really taking much heterogeneity into account, the estimation procedure considers unobserved heterogeneity by means of a mixture technique inspired by Heckman and Singer (1984). Some simulations about the identification of this model with unobserved heterogeneity and duration dependence tell us that the structure of the model and the type of data used help to distinguish between these two elements.

The main empirical results are as follows. First, the predicted unemployment hazard rate is almost flat for the first six months in unemployment and decreases thereafter. The structural estimation indicates that during the first six months, decreasing reservation wages compensate the also decreasing arrival rate of offers. However, as reservation wages are highly decreasing with unemployment duration, after six months the acceptance probabilities are practically equal to one. Since then, the hazard rate is equal to the offer arrival rate, which also tends to decrease throughout the unemployment spell. The model predicts that those workers with access to unemployment benefits have a mean expected unemployment duration of more than five months whereas those without such benefits have an expected duration of 3.8 months.

Although this estimation procedure imposes some rigidity in the way unemployment benefits are considered, we have performed some policy simulations in order to evaluate the effect that different Unemployment Insurance designs may have on the expected unemployment duration. Moreover, we have analyzed whether the optimal Unemployment Insurance proposed in Hopenhayn and Nicolini (1997a) for the US is applicable to the Spanish situation with a much more serious problem of unemployment together with a high separation rate. Our results show that a design inspired by this proposal (decreasing replacement rate and a tax on workers depending on their unemployment duration which should be initially negative for relatively short unemployment durations) reduces the unemployment duration by 4.4% in Spain when compared to the current system. However, as suggested by Wang and Williamson (1996), the optimal way

[3] Empirical evidence from the European Community Household Panel shows that the correlation between self-declared reservation wages and unemployment duration is higher in countries with more presence of temporary contracts, that is, with more risk of job separation.

to provide incentives to the worker in a context of employment risk is a mix of non-monotonic replacement rates and a tax on re-employment wages depending on unemployment duration. Indeed, when we evaluate a design with a 40% replacement rate for the first three months of unemployment, 80% the following three months and a decreasing pattern after that, the expected duration in unemployment reduces by 9% (13.2% if we also use the optimal tax). Hence, given the high separation rate in our data, we conclude that providing incentives to the worker in order to quickly accept offers requires not only a tax on re-employment wages but also an initial penalty in the replacement rates at the beginning of the unemployment spell. This design together with the negative tax if the worker accepts early provides strong incentives to workers, which results in a substantial reduction in unemployment duration.

The structure of the paper is as follows. Section 2 deals with the non-stationary job search model with the separation probability, along with some simulation exercises which help us to better understand the results of the model. Section 3 describes the estimation procedure, the data used, and the main results, and Section 4 presents the main conclusions.

2 The model

I use a standard discrete-time search model (See, for example, Lippman and McCall, 1976 and Wolpin, 1987) where the parameters are allowed to vary in accordance with the unemployment duration and may depend also on certain observed characteristics, x_i.[4] The model is based on the continuous-time model of Van den Berg (1990), but I have modified it by introducing the probability of being separated from the job once employed.[5]

Consider a discrete-time economy where agents either work receiving a constant wage, w, or are unemployed and searching for a job in each period t. The following conditions are assumed:[6]

(**A1**) Wage offers at time t are random draws from a distribution function $F_i(w,t)$ where $w \in [0,\infty)$ and t is how long the agent has been unemployed.[7]

(**A2**) Job offers arrive with probability $\alpha_i(t)$ in each period t.

[4] We will have different regressors representing observed worker's heterogeneity in the empirical model to be estimated (basically age, qualification and whether or not the worker has access to Unemployment Benefits).

[5] The other modification with respect to Van den Berg (1990) is the discrete-time framework. The reason for departing from continuous time is not only the final objective of estimating the model using discrete data but also to better understand the effect of the separation rate.

[6] In order to have a compact notation we will represent the dependence of the parameters on the observed heterogeneity, x_i, with the subindex i in each parameter.

[7] Calendar time is assumed to start at the moment the individual becomes unemployed. Thus, t refers both to calendar time and to how long the individual remains unemployed.

(**A3**) During the spell of unemployment, the agent has an income $b_i(t) \in [0, \infty)$, net of search costs, which depends on unemployment duration and also on observed heterogeneity.[8]

(**A4**) When an offer is accepted, the agent works at the offered wage, w, but he can also be separated from that job at a rate $\delta_i \in [0, 1]$.[9]

(**A5**) The individual has a constant subjective discount rate $r \in [0, \infty)$.

(**A6**) There is also a period T from which all the parameters depending on unemployment duration are constant.

The expected present value of future net income for an unemployed worker who is searching for a job is defined as:

$$U_i(t) = b_i(t) + \frac{1}{1+r}\left[\alpha_i(t)E_{w,t+1}\max\left(W_i(w), U_i(t+1)\right) + \left(1-\alpha_i(t)\right)U_i(t+1)\right] \tag{1}$$

Thus, $U_i(t)$ is the value of unemployment time, $b_i(t)$, received at the beginning of period t, plus the expected and discounted value of the optimal decision to stop at $t+1$. This expected value is the maximum between the expected present value of accepting the offer, $W_i(w)$, in the event that an offer arrives in period t,[10] and continuing to search one more period, $U_i(t+1)$. If no offer arrives, then the worker will be unemployed again at period $t+1$, $U_i(t+1)$.

The expected present value of stopping a job search and beginning to work at wage w is:

$$W_i(w) = w + \frac{1}{1+r}\left[(1-\delta_i)W_i(w) + \delta_i U_i(0)\right] \tag{2}$$

That is, $W_i(w)$ is the value of the wage received in period t plus the expected present value of what can happen at period $t+1$: with probability $1-\delta_i$, the worker will continue employed and with the opposite probability the worker will leave the job and become unemployed again, with a duration of zero periods, $U_i(0)$.[11]

[8] This income can be interpreted as the value of time for the unemployed worker, and includes, among other things, unemployment benefits and income coming from other sources.

[9] That is, the job can be interrupted for any exogenous reason, such us firing or quitting. This probability can also be a function of accumulated tenure (See García-Perez, 1998). However, given we have no employment duration data, we will use the expression in that paper for δ_i, the "mean separation rate", which takes into consideration the evolution of the separation probability from the beginning of the employment spell to infinity.

[10] I assume that this offer is received at the end of the period so that we have to apply the time discount factor to its expected value.

[11] The value of unemployment time may depend on the previous wage of the worker. If this is the case, this value in the case of a future unemployment spell should depend on the accepted wage once the worker is employed. However, this would provoke the presence of two different sources of time dependence in the model, one with respect to the current spell and another one with respect to the following one. As this would complicate the model too much, I will not consider this in the analysis although it will be taken into account in the empirical approach.

In this context, as in all job search models, every time an offer arrives the decision is whether to accept or to reject it and search further. The individual will only be indifferent as to whether to work or search for one more period at a wage called the *reservation wage,* $w_i^*(t)$. This wage equates $U_i(t)$ and $W_i(w)$ and, hence, establishes that:

$$U_i(t) = \frac{(1+r)w_i^*(t)}{r} + \left(U_i(0) - \frac{(1+r)w_i^*(t)}{r}\right)\frac{\delta_i}{\delta_i + r} \tag{3}$$

Taking into account that $U_i(0) = \frac{(1+r)w_i^*(0)}{r}$, *i.e.* the value of $U_i(t)$ when $t = 0$, and substituting (3) in (1) we obtain the following difference equation for the reservation wage:[12]

$$w_i^*(t) = b_i(t) + \frac{\delta_i}{r}\left(b_i(t) - w_i^*(0)\right) + \frac{\alpha_i(t)}{r}\int_{w_i^*(t+1)}^{\infty}\left(w - w_i^*(t+1)\right)dF_i(w,t) + \frac{\Delta w_i^*(t)}{r} \tag{4}$$

where $\Delta w_i^*(t) = w_i^*(t+1) - w_i^*(t)$. It is straightforward to show that $w_i^*(0)$ satisfies this:

$$w_i^*(0) = b_i(0) + \frac{\alpha_i(0)}{\delta_i + r}\int_{w_i^*(1)}^{\infty}\left(w - w_i^*(1)\right)dF_i(w,0) + \frac{\Delta w_i^*(0)}{\delta_i + r} \tag{5}$$

From (4) we can distinguish four terms in the reservation wage: (*i*) the value of time for the unemployed worker, $b_i(t)$; (*ii*) the value associated with a future job separation, given by the difference between income in period t and the value of returning to period 0 of the following unemployment spell; (*iii*) the expected and discounted benefit associated with the arrival of a new offer; and, (*iv*) the appreciation or depreciation of the option represented by the reservation wage.

Given the expression of the reservation wage, equation (4), we obtain the *unemployment hazard rate,* $\phi_i(t)$, which is the probability for the individual i of exiting unemployment in t, conditional on not having exited before, and which is defined as:

$$\phi_i(t) = \alpha_i(t)\left[1 - F_i(w_i^*(t+1), t)\right] \tag{6}$$

that is, the rate at which offers arrive multiplied by the probability that a given offer is acceptable. Note that given (1), where the value of accepting a job offer arriving at the end of period t is compared with the expected present value of being unemployed at time $t+1$, the acceptance probability in period t is computed taking into account the reservation wage at time $t+1$.[13]

2.1 Nonstationarity of the reservation wage

The nonstationarity of the reservation wage is derived from the nonstationarity of the parameters of the model, which is established by the following alternative

[12] If the separation probability is taken to be equal to zero, this equation is the same as in Van den Berg (1990) but in discrete time.

[13] This is a consequence of discrete time. In continuous time (see García-Pérez, 1998), this acceptance probability would simply be $1 - F_i(w_i^*(t), t)$.

assumptions:[14]

(K1) $b_i(t) > b_i(t+1)$, $\forall t \in [0, T)$.

(K2) $\alpha_i(t) > \alpha_i(t+1)$, $\forall t \in [0, T)$.

(K3) $F_i(w,t)$ first order stochastically dominates $F_i(w, t+1)$, $\forall t \in [0, T)$, which implies that $1 - F_i(w,t) > 1 - F_i(w, t+1)$, $\forall w \in [0, \infty)$.

(K4) $F_i(w,t)$ is a mean preserving spread of $F_i(w, t+1), \forall t \in [0, T)$, that is, $E_i(w,t) = E_i(w, t+1)$, and $\forall x \in [0, \infty)$,

$$\int_0^x F_i(w,t)dw > \int_0^x F_i(w, t+1)dw.$$

The economic meaning of these assumptions is simple. For an unemployed worker the value of time decreases with unemployment duration because his income and unemployment benefits decline over time. The offer arrival rate and the wage offered may decrease as time proceeds as a result of the stigma effect that long-term unemployed workers may suffer from (see Berkovitch, 1990). The distribution of offers can be more concentrated around its mean for the long-term unemployed, because they may know more about this distribution as time passes (see Burdett and Viswanath, 1988). Also important is the assumption that people know how the parameters are related to the duration of unemployment.

The time dependence exhibited by the reservation wage is obtained in the following theorem, where it is helpful to use what I call a *stationary reservation wage, $w_i^0(t)$.* This wage is the optimal reservation wage at time t, for all $t \geq 0$, if the environment remains stationary after t, that is:

$$w_i^0(t) = b_i(t) + \frac{\delta_i}{r}\left(b_i(t) - w_i^0(0)\right) + \frac{\alpha_i(t)}{r}\int_{w_i^0(t)}^{\infty}\left(w - w_i^0(t)\right)dF_i(w,t) \qquad (7)$$

Theorem 1 *Let assumptions (A1) to (A6) be satisfied. Let one parameter satisfies assumptions (K1)-(K4) with strict inequality, while the remaining ones are constant over the time interval* $[0, \infty)$. *Then:*

(i) $w_i^*(t) < w_i^0(t)$,

(ii) $\Delta w_i^*(t) < 0$.

Proof : See Appendix A.

The meaning of this result is simple: any future decrease in the parameters of the model makes the value of a current job-search spell smaller than it would be if the parameters were constant. Accordingly, the unemployed worker, anticipating these future changes, sets a smaller reservation wage as his unemployment spell increases.

[14] The derivation of the nonstationarity of the reservation wage is similar to Van den Berg (1990) but in discrete time.

2.2 The effect of the separation probability

In stationary search models (see, for example, Devine and Kiefer, 1991) the effect of the separation rate upon reservation wages is negative. Given that the future is more uncertain, future opportunities are discounted at a higher rate and, thus, the reservation wage is lower. This is because the value of being employed is lower when jobs do not last forever. Given this, the value of being unemployed is also lower and the result is that the minimum acceptable wage for these workers is smaller.

However, in the present model, the nonstationarity of the search process introduces a new element: being separated from a job is not the same when considered by an unemployed worker at the beginning of the unemployment spell as it is after one year of unemployment, for example. This fact can be confirmed by analyzing equation (4). The effect of the separation probability is not only a direct one, through the presence of δ_i in the expression for $w_i^*(t)$, but also an indirect one because of its effect on $w_i^*(0)$. Hence, in order to obtain the complete effect of the separation probability we need a general expression for $w_i^*(t)$ as a function only of exogenous parameters.

This can be obtained by considering the fact that equation (4) determines a system of $T+1$ equations for reservation wages from period 0 to period T. If we work backwards in this system, we can obtain an expression for $w_i^*(0)$ and, after substituting in $w_i^*(t)$, obtain the following general expression for the reservation wage:

$$w_i^*(t) = (r+\delta_i)\mathrm{PV}(\mathrm{b}_i(\mathrm{t})) + \mathrm{PV}(\mathrm{E}_i(\mathrm{w,t})) - \frac{\delta_i \mathrm{D}_i(\mathrm{t})}{1+\delta_i \mathrm{D}_i(0)}\left[(r+\delta_i)\mathrm{PV}(\mathrm{b}_i(0)) + \mathrm{PV}(\mathrm{E}_i(\mathrm{w,0}))\right] \tag{8}$$

where,

$$PV(b_i(t)) = \sum_{k=t}^{T-1} \frac{b_i(k)}{1+r}\left(\prod_{j=t}^{k-1}\frac{1-\phi_i(j)}{1+r}\right) + \frac{b_i(T)}{r+\phi_i(T)}\left(\prod_{j=t}^{T-1}\frac{1-\phi_i(j)}{1+r}\right)$$

$$PV(E_i(w,t)) = \sum_{k=t}^{T-1} \frac{\alpha_i(k)\int_{w_i^*(k+1)}^{\infty} wdF_i(w,k)}{1+r}\left(\prod_{j=t}^{k-1}\frac{1-\phi_i(j)}{1+r}\right) + \frac{\alpha_i(T)\int_{w_i^*(T)}^{\infty} wdF_i(w,T)}{r+\phi_i(T)}\left(\prod_{j=t}^{T-1}\frac{1-\phi_i(j)}{1+r}\right)$$

$$D_i(t) = \sum_{k=t}^{T-1} \frac{1}{1+r}\left(\prod_{j=t}^{k-1}\frac{1-\phi_i(j)}{1+r}\right) + \frac{1}{r+\phi_i(T)}\left(\prod_{j=t}^{T-1}\frac{1-\phi_i(j)}{1+r}\right)$$

That is, the reservation wage at time t is the present discounted value, $PV(\cdot)$, of $(r+\delta_i)b_i(t)+\alpha_i(t)\int_{w_i^*(t+1)}^{\infty} wdF_i(w,t)$ from period t to period T, minus a fraction of this present discounted value but from period 0 to T. In these present values, the discount factor involves all the parameters of the model via the unemployment hazard rate, $\phi_i(t)$. Hence, this expression takes into account both a time discount, r, and a probability discount via the hazard rate. The latter regards whether the worker is unemployed or not in each of the periods considered.

Before discussing the derivative of $w_i^*(t)$ with respect to δ_i, we can observe the following useful result.

Lemma 2 *If $b_i(t)$ is decreasing, the derivative of $w_i^*(t)$ with respect to δ_i is also decreasing in t.*

Proof: The derivative of $w_i^*(t)$ with respect to δ_i, divided by $D_i(t)$, is:

$$\frac{dw_i^*(t)}{d\delta_i}\frac{1}{D_i(t)} = \frac{PV\big(b_i(t)\big)}{D_i(t)} - \frac{\delta_i D_i(0)}{1+\delta_i D_i(0)}\frac{PV\big(b_i(0)\big)}{D_i(0)} - \frac{D_i(0)}{\big(1+\delta_i D_i(0)\big)^2}\left[(r+\delta_i)\frac{PV\big(b_i(0)\big)}{D_i(0)} + \frac{PV\big(E_i(w,0)\big)}{D_i(0)}\right]$$

Given that $\frac{PV(b_i(t))}{D_i(t)}$ is a weighted average of the values of $b_i(t)$ from period t to T and $b_i(t)$ is decreasing in t, $\frac{PV(b_i(t))}{D_i(t)}$ is decreasing in t. But this means that $\frac{dw_i^*(t)}{d\delta_i}$ will also be decreasing in t because the other terms in this derivative are constant in t. Q.E.D.

Given this result, we can easily posit a general result for the sign of the derivative of $w_i^*(t)$ with respect to δ_i. This is established in the following proposition:

Proposition 3 *The effect of the separation rate on the reservation wage at period t, $\frac{dw_i^*(t)}{d\delta_i}$, will be negative if and only if*

$$\frac{PV\left(b_i(0)\right)}{D_i(0)} \leq \frac{PV\left(E_i(w,0)\right)}{1-rD_i(0)}$$

In the opposite case, the effect will be positive until t^* when

$$\frac{PV(b_i(t^*))}{D(t^*)} = \frac{\delta_i D_i(0)}{1+\delta_i D_i(0)}\frac{PV(b_i(0))}{D_i(0)} + \frac{D_i(0)}{(1+\delta_i D_i(0))^2}\left[(r+\delta_i)\frac{PV(b_i(0))}{D_i(0)} + \frac{PV(E_i(w,0))}{D_i(0)}\right]$$

Proof : Given the result of the previous Lemma, in order that $\frac{dw_i^*(t)}{d\delta_i} \leq 0$, it is sufficient that this derivative at time 0 is negative. This sufficient condition is satisfied if and only if, evaluating $\frac{dw_i^*(t)}{d\delta_i}$ at time 0

$$\frac{PV\big(b_i(0)\big)}{D_i(0)} - \frac{\delta_i D_i(0)}{1+\delta_i D_i(0)}\frac{PV\big(b_i(0)\big)}{D_i(0)} - \frac{D_i(0)}{\big(1+\delta_i D_i(0)\big)^2}\left[(r+\delta_i)\frac{PV\big(b_i(0)\big)}{D_i(0)} + \frac{PV\big(E_i(w,0)\big)}{D_i(0)}\right] \leq 0.$$

Rearranging the terms, we can establish that the condition to be satisfied is:

$$\frac{PV\left(b_i(0)\right)}{D_i(0)} \leq \frac{PV\left(E_i(w,0)\right)}{1-rD_i(0)}$$

In the opposite case, $\frac{dw_i^*(0)}{d\delta_i} \geq 0$ and it will continue to be positive until period t^*, where the second expression in the proposition is verified. For all periods after t^*, the derivative will be negative. Q.E.D.

This proposition tells us that in a non-stationary environment, the effect of the separation probability on reservation wages is not always negative. If the weighted average of $b_i(t)$ from period 0 to T, $\frac{PV\left(b_i(0)\right)}{D_i(0)}$, is high enough with respect to the present value of expected wages, we can find an initially positive effect which lasts for t^* periods. This is totally new and entirely different from the results found in a stationary environment: when the parameters of the model change with the time the worker is unemployed, a higher separation rate can cause the reservation wage to be higher instead of lower than without considering such separations. That is, the worker may be more selective at the beginning of his unemployment spell and this is the case when he enjoys a much better situation than he expects to have in a possible job. However, as time passes, the worker realizes that his income or his chances of a new offer will be lower. Nevertheless, he also knows that if he is hired, even in the case of a future separation, he will have access to greater values of all the parameters of the model. This means that the reservation wage decreases very rapidly as time passes.[15] That is, the separation probability also affects the time dependence of the reservation wage. The following proposition tells us that when the unemployed worker considers a future possibility of being unemployed, reservation wages will be even more negatively time-dependent.

Proposition 4 *If $b_i(t)$ is decreasing, a higher separation probability will make the negative time dependence of reservation wages even more negative.*

Proof : As $\Delta w_i^*(t) = w_i^*(t+1) - w_i^*(t)$, substituting each reservation wage by its own expression, we can establish

$$\frac{d\Delta w_i^*(t)}{d\delta_i} = \frac{\Delta b_i(t)}{1+r} + \frac{1}{1+r}\frac{d\Delta w_i^*(t+1)}{d\delta_i}.$$

[15] Of course, this effect is due to the assuption that the situation at the beginning of the unemployment spell is the same whatever the duration of the previous job. This is clearly at odds with the observed fact that, for example, unemployment benefits depend on the length of the previous job and on its associated wage. However, given the difficulty of accounting for these aspects, I have omitted them in the analysis.

Given that $\forall t \in [T, \infty)$ $\Delta w_i^*(t) = 0$, then $\frac{d\Delta w_i^*(T)}{d\delta_i} = 0$ and thus, we have $\frac{d\Delta w_i^*(T-1)}{d\delta_i} = \frac{\Delta b(T-1)}{1+r} < 0$ because $b_i(t)$ is decreasing. Therefore, a higher separation probability increases the negative time dependence of $w_i^*(t)$ provided that:

$$\frac{d\Delta w_i^*(t)}{d\delta_i} < \frac{d\Delta w_i^*(t+1)}{d\delta_i} < 0 \qquad \forall t \in [0, T).\ Q.E.D.$$

Hence, the worker's requirements for accepting job offers will decrease even more rapidly since he knows he could be unemployed again in the future. This means that acceptance probabilities will increase quickly during the unemployment spell.

I have carried out some simulations with the model in order to determine when we can expect to observe positive or negative effects of the separation probability on reservation wages. In these simulations I have combined five possible values for each of the parameters in the model to compute reservation wages and also to check the condition stated in Proposition 3. From this proposition we know that all the parameters in the model have an effect on whether we obtain either a positive or negative effect from the separation probability over reservation wages. We can affirm from these exercises that the most important factor to obtain an initially positive effect is the fact that the mean of the distribution of offered wages has to be sufficiently low with respect to the value of unemployment time. Moreover, we can establish that the reaction of the studied derivative to a 50% change on the baseline parameters (see the note to Figure 1) is more than proportional only for the mean of offered wages, $E_i(w)$, the value of unemployment time at period 0, $b_i(0)$, and for the coefficient of variation of offered wages, $CV_i(w)$.[16] In Figure 1 we find that the condition of Proposition 3, $\frac{PV\left(b_i(0)\right)}{D_i(0)} - \frac{PV\left(E_i(w,0)\right)}{1-rD_i(0)}$, is positive only for certain combinations of $b_i(t)$ and the parameters associated to the distribution of offered wages. Given the values used in the simulations, the level of $b_i(t)$ at period 0 has to be at least 80% larger than the mean offered wage in order to obtain an initially positive effect of the separation rate.[17] The coefficient of variation of offered wages is also very important. The larger this coefficient is, the higher $b_i(0)$ has to be in relation to the mean offered wage in order to observe a positive effect of the separation probability on reservation wages at time zero.

Hence, the effect of the separation probability on non-stationary reservation wages is quite important. Not only can we obtain a positive derivative of reserva-

[16] The elasticity of the condition stated in Proposition 3 to a 50% increase in each parameter (in absolute value) is -4.2 for $E_i(w)$, 2.5 for $b_i(0)$ and -2.4 for $CV_i(w)$. Hence, the higher the mean or the coefficient of variation of offered wages and the lower the value of unemployment time, the more difficult it is to obtain a positive effect of the separation probability on reservation wages. Other parameters show elasticities lower than 0.4.

[17] In a country like Spain, where severance payments are quite high, it is not unusual to have high values of unemployment time at the beginning of the unemployment spell.

tion wages with respect to the separation probability, but I have also found that the time dependence of the reservation wage will be even more negative when the probability of being separated from the job increases.

3 Structural estimation

There are not many papers that attempt to estimate dynamic programming models of individual behavior. In the context of job search models, some important references are Lancaster and Chesher (1983), Miller (1984), Narendranathan and Nickell (1985) and Frijters and Van der Klaauw (2003). However, Wolpin (1987) provides one of the most influential articles in this area. This paper develops a discrete-time search model, which is non-stationary because it assumes a finite search horizon. It is estimated by maximum likelihood, and uses data regarding duration, accepted wages and various individual characteristics.

The estimation in the present paper has been made using Spanish data derived from the Spanish Continuous Family Expenditure Survey (*Encuesta Continua de Presupuestos Familiares (ECPF)*) for the period 1985-1996. The *ECPF* is a rotating panel which interviews aproximately 3,200 households every quarter. One eighth of the sample is renewed quarterly and hence an individual can be monitored for a maximum of two consecutive years. This source gives information about unemployed workers during their spells of unemployment and also about their post-unemployment wages, in addition to information on income, consumption and other household characteristics.

The estimation sample consists of unemployed household heads, which is the only group within the *ECPF* whose educational level is reported. I have also restricted the sample to married men with non-working spouses in order to reduce heterogeneity since, given the estimation procedure, I am not able to consider many regressors.

The individuals in the sample are all entrants to unemployment. The observed spells can be either *complete*, if the worker subsequently exits from unemployment, or *censored*, if he does not. For the complete spells, the re-employment wage (for those who continue answering the survey two quarters after exiting from unemployment) is taken to be the labor income of the second quarter of employment.[18] As will be explained below, we will use information on previous wages to better identify the value of unemployment time, $b(t)$. As it is not possible to recover previous wages for all workers in the dataset, we also restrict the sample to those workers with an observed previous wage.

Table 1 shows that there are 390 completed spells of unemployment and 353 censored spells. Of the former, 228 have an observed re-employment wage. The histogram of re-employment wages (which are expressed in real terms of December 1996) can be seen in Figure 2.

[18] The reason for doing this is to reduce measurement error regarding the amount, which is simply the quarterly income.

Although the *ECPF* is a quarterly survey, it is possible to calculate monthly values of the variables. Monthly data is preferable because it better reflects the nonstationarity of the job search behavior, and the changing patterns of the parameters are likely to be estimated far more clearly.[19] In order to obtain monthly data, a few transformation rules have been applied. They are explained in Appendix B.

The model has been estimated structurally by using both the monthly data previously described and the usual assumption (see Van den Berg, 1990 or Wolpin, 1987) that wages are lognormal.[20] Moreover, we will also use data regarding household income to better identify the value of unemployment time. But the difficulties in the process of estimation make other simplifying assumptions necessary. This is because the estimation procedure involves calculating the reservation wages for each evaluation of the likelihood function. However, it is computationally very time-consuming to do this for each worker. The solution I have adopted is to restrict the heterogeneity of the sample and to divide workers into groups based on a few dichotomous variables.

In the results I present, there are five explanatory variables which, together with the also discretized previous wages, will be used to build the worker groups. The first variable is *Skill,* which is measured by the level of education: a skilled worker is one who has at least attended High School. *Age4565* is the second variable and is equal to one for workers older than 45 years and zero otherwise. Thirdly, there is a binary variable, *High Unemployment*, indicating the periods where the unemployment rate is higher than the mean in the whole period, 21.5%. I have also split the sample between observations before and after 1992 because the Unemployment Insurance System changed in Spain at the beginning of that year. Finally, I have also used a variable termed *Benefits* which indicates whether or not the individual has access to unemployment benefits.[21] Moreover, this variable will be different before and after 1992. Given that the design of Unemployment Insurance in Spain changed in 1992,[22] *Benefits* is de-

[19] Hence, the time period in our discrete-time model is one month. This length could represent quite a long time in some economies, which might create problems in the estimation of the offer arrival rate. However, this is not a problem for Spain where the duration in unemployment is generally long enough to imply monthly offer arrival rates lower than one.

[20] It is well-known that not all wage offer distribution functions satisfy the *recoverability condition* which is crucial for identifying the model (See Flinn and Heckman, 1982). One function which satisfies it is the lognormal one and this is the main reason for choosing it. Moreover, this function also works well in Wolpin (1987) and fits our empirical distribution of accepted wages (See Figure 2).

[21] This variable requires further comment. It indicates not only whether the unemployed worker actually receives unemployment benefits or not, but also whether he has received them or not. The basic idea behind this distinction is the following: in order to correctly estimate the effect of benefits on a structural estimation, we would need to know the complete sequence of benefit receipt over the spell of unemployment of each worker, both for workers with or without a complete spell. This requirement is not entirely satisfied by the data currently used. Accordingly, we have to follow an intermediate solution which will not fully depict the structural effect of unemployment benefits.

[22] Before 1992 the replacement ratio was 80% for the first six months of unemployment, 70%

fined to be equal to one when the replacement ratio was the highest in the whole period (80% before 1992 for the first six months of unemployment) and will take a relative value with respect to 80% in the rest of cases. Obviously when the worker has no unemployment benefits, this variable is set equal to zero. Finally, previous wages are also discretized into 5 values based on the four quartiles of their distribution.

In the estimation, a monthly discount rate of 0.3% (i.e. a 3.66% annual rate) was imposed and not estimated and T was set at 24 months so as to calculate the final condition for the reservation wage. Different discount rates have also been tried and the estimation results change only marginally. With respect to the separation probability, given that I have no data on employment duration, I cannot directly estimate this parameter. However, I will estimate it based on a different dataset, Social Security registers, for the same period. I will use these estimated values for each worker's group in the structural estimation. The results from this auxiliary estimation are shown in Appendix C. Hence, each worker type in the sample is associated with the individual estimation of his separation probability.

3.1 The likelihood function

The estimation technique presented here is clearly inspired by the one in Wolpin (1987) but it contains a new element: unobserved heterogeneity. Another difference is the use of more information (monthly income of the household) in order to better identify the value of unemployment time. As will be explained below, I have assumed household income to be equal to the value of unemployment time for the unemployed head, $b(t)$, multiplied by some measurement error.[23] I will first explain the likelihood function without taking into account unobserved heterogeneity, and afterwards, I will consider the modified function which controls its presence.

We have three types of individuals in the sample of unemployed workers: those with complete spells and observed re-employment wages, those with complete spells but without observed re-employment wages and finally, those with censored spells. Thus, the likelihood function will have three different components:[24]

between the seventh and the twelfth and 60% afterwards. After the 1992 reform the replacement ratio is 70% for the first six months of unemployment and 60% afterwards.

[23] I will assume this measurement error has zero mean and variance σ_y^2.

[24] v_i is an indicator variable equal to 1 if the re-employment wage of the worker i is observed and zero otherwise. c_i is an indicator of censoring: it is equal to 1 if the individual i has a complete spell and zero otherwise. T_i represents worker i's unemployment spell duration and W_i^o is his observed re-employment wage. $Y_{o_i}^t$ is the vector of observed values for household income, $Y_{o_{it}}$, in each of the t periods of unemployment. Finally, d_{it} is equal to one if the individual i has his last observation, t_i, at period t.

$$\begin{aligned}
\ln \pounds &= \sum_{i=1}^{N} L_i = \sum_{i=1}^{N} \left\{ c_i \left[v_i \ln \left(\Pr(T_i = t, W_{o_i}) \right) + (1 - v_i) \ln \left(\Pr\left(T_i = t\right) \right) \right] \right. \\
&\quad \left. + (1 - c_i) \ln \left(\Pr(T_i > t) \right) \right\} + \ln(\Pr(Y_{o_i}^{t})) \qquad (9) \\
&= \sum_{i=1}^{N} \sum_{j=1}^{t_i} \left\{ c_i d_{ij} \left[v_i \ln \left(f_{W_o}(W_{o_i} \mid j) \phi_i(j) \right) + (1 - v_i) \ln \left(\phi_i(j) \right) \right] \right. \\
&\quad \left. + (1 - c_i d_{ij}) \ln \left(1 - \phi_i(j) \right) \right\} + \ln(\Pr(Y_{o_{ij}}))
\end{aligned}$$

where $Y_{o_{ij}}$ is the observed income of individual i in period t which is equal to the value of unemployment time, $b_i(t)$, multiplied by a measurement error of variance σ_y^2. Given this likelihood function and taking into account the reservation wage, equation (4), we can estimate the parameters of the model, $\alpha(t)$, $b(t)$, δ, $\bar{W}$, σ_u, σ_ε and σ_y provided they are all identified.

The general idea behind identification is as follows: given that we have data on accepted wages, along with data on unemployment duration, the parameters of the wage offer distribution, $\bar{W}, \sigma_u$ and σ_ε will be clearly identified in the first component of the likelihood function. Furthermore, given data on income, the value of time for unemployed workers, $b(t)$, will also be identified. Finally, the separation probability, δ, and the offer arrival rate can be identified by making use of the system of reservation wages from 0 to T. However, without any data on previous employment spell durations or offer arrival, the distinction between these two parameters can be quite unclear. Hence, given that we can estimate the separation probability from a complementary data set, I will use this previous estimation and not estimate it directly in the structural estimation procedure.

3.1.1 The likelihood function with unobserved heterogeneity

The need to restrict the heterogeneity in our sample means that much of the sample heterogeneity will not be captured by the explanatory variables used. This problem, together with the fact that unobserved heterogeneity generates spurious negative duration dependence in the estimation, calls for the control of unobserved heterogeneity in the hazard rate.

Since I do not have multiple spells, unobserved heterogeneity cannot be controlled by using a fixed effect approach. However, I can apply a random effect technique such as the one used, for example, in Flinn and Heckman (1982). In order not to further restrict the estimation procedure, I will estimate the distribution of unobserved heterogeneity with a technique inspired by Heckman and Singer (1984). That is, with unobserved heterogeneity, η, the log-likelihood function takes the form:

$$\ln \pounds_h = \sum_{i=1}^{N} \ln \int L_i(\eta) dF(\eta) \qquad (10)$$

where $F(\eta)$ is the cumulative distribution function of η, which is a discrete function with two mass points, η_1 and η_2.[25] These mass points are selected in order to verify the assumption of $E(\eta) = 0$, which is necessary given the presence of a constant term in the parameters where unobserved heterogeneity is introduced. Hence, only one point will be estimated and the other one proceeds from this assumption.

The function $L_i(\eta)$ is the likelihood function described in the previous subsection, where its arguments are all functions of the unobserved heterogeneity variable, η.

I have carried out some simulations to understand to what extent we can jointly identify the effects of unobserved heterogeneity and duration dependence in the unemployment hazard rate. I have generated fifty random samples of 500 workers with a few binary variables similar to those present in our data set[26] and I have applied to them the estimation procedure previously described. For those exercises which deal with the presence of unobserved heterogeneity, one of the generated binary variables is dropped in different parameters of the model and assumed to be the unobserved heterogeneity term. I have tried different specifications with respect to this term. The results are shown in Table 2. It reflects this heterogeneity in both the mean of offered wages and the value of unemployment time, which is the final specification we use in our estimation.

The first conclusion from these simulations is that all the parameters are quite well identified given typical significance levels. With respect to the introduction of unobserved heterogeneity, I have found that the estimation procedure identifies its presence and its differential effect with respect to duration dependence. That is, when the time dependence of the offer arrival rate is -0.13, the estimation procedure results in an average duration dependence coefficient of -0.144 with a mean standard error of 0.006. The same is true for the time dependence of the value of unemployment time: its true value is -0.1 and its mean estimated value is -0.098 with a mean standard error of 0.001. The probability of the unobserved heterogeneity distribution, which is equal to 0.5 when generating the data, is estimated to be 0.498 on average with a mean standard error of 0.013. Finally, the level of the mass point representing unobserved heterogeneity is also estimated quite well: the mean estimated value is 0.247 (0.25 is the true value) and the standard error is 0.006. Hence, we can conclude that the data we have, together with the structure of the theoretical model, helps to clearly identify all the parameters in the model. Moreover, the use of household income to identify $b(t)$ is determinant given the fact that without using it the results are much

[25] I have only used two mass points in the distribution of the unobserved heterogeneity. It is known that the increase in the number of mass points could be a way of improving the control for unobserved heterogeneity. Moreover, a promising avenue for improving our analysis of unobserved heterogeneity in structural models would be the analysis of how the estimation results change as the number of mass points increases. This exercise could be complementary to that of Baker and Melino (2000) but it is left for future research.

[26] These variables are included, respectively, in the offer arrival rate, the mean and variance of offered wages and in the value of unemployment time.

poorer.[27]

The structural estimation uses the functional forms for the parameters of the model shown in Table 3. The offer arrival rate, $\alpha(t)$, is parameterized using the extreme value distribution function. The idea is to use a proportional assumption for the underlying continuous offer arrival rate. It is well known (see Meyer, 1990), that in discrete time, a continuous proportional hazard rate follows this distribution. The other parameters are assumed to be exponential because of the assumption of lognormal wages, in order to reduce their scale or to ensure they are positive. These parameters have some exclusion restrictions based on previous estimations which give us the best parameterization in terms of significance and likelihood values.

With respect to the parameterization of the offer arrival rate, I distinguish between people with and without access to unemployment benefits. It may be argued that access to unemployment benefits should not only alter the value of time for the unemployed worker, but also his search effort. Although not explored here, we would therefore expect the offer arrival rate of such an individual to be different to that of a worker without benefits. We have to remember that this variable does not represent unemployment benefits, but is merely an indicator of whether the worker has access to them or not, taking into account also the decreasing pattern of such benefits. Nevertheless, estimates without this indicator in the offer arrival rate are far less accurate than the one shown in terms of likelihood values and significance tests.

3.2 Results

The main results of the structural estimation can be seen in Tables 4 and 5. Table 4 shows the estimated coefficients of the model both when unobserved heterogeneity is and is not controlled for, under the previously estimated separation rates (its mean value is 4.9%). Table 5 shows the predicted values for the main elements in the model. It also presents the main predictions for the two estimated groups with respect to unobserved heterogeneity.

The first result shown in Table 4 is that the presence of unobserved heterogeneity in the data cannot be refuted (both estimated parameters are significant at 5% and even the test of the probability p being equal to 1 has a t-statistic of -20.71). Moreover, the likelihood ratio of a test for the absence of unobserved heterogeneity has a value of 1064.12 with a p-value lower than 10^{-4}. As is well known, these parameters are not identified under the hypothesis of no heterogeneity. In fact, the distribution of the LR test under this hypothesis is a mixture of the chi-square with one and with three degrees of freedom.[28] Hence, there

[27] The results from the simulation exercises that do not use data on income are available upon request.

[28] Instead of obtaining the empirical distribution of this statistic using a Monte Carlo test, we approximate this by taking the chi-square with three degrees of freedom as the approximate distribution of the LR test in order to reject the hypothesis of no heterogeneity. This can be

is unobserved heterogeneity in the data but we can also confirm from this table that the control of this element does not actually affect the duration dependence of the parameter where it is introduced although their estimated values change a bit. As the model which controls for unobserved heterogeneity is more general, we will focus on its results.

Duration dependence is estimated in the offer arrival rate, $\alpha(t)$, and in the value of time for unemployed workers, $b(t)$. We can see that there is a highly significant negative duration dependence in both parameters: a mean 1.26% monthly decrease in $b(t)$ for the average worker and a 9.23% mean monthly decrease in $\alpha(t)$. This is much higher than the 2.5% found in Wolpin (1987) for the offer arrival rate using US data. We allow for a different duration dependence of $b(t)$ for workers with and without unemployment benefits. As expected, I have found that the negative duration dependence is much more important for workers without such benefits.

As far as the skill variable is concerned, we can see that the variance of the logarithm of offered wages is considerably higher for these workers. This means, as can be seen in Table 5, that both the mean and the variance of offered wages are higher for skilled workers. The same result was found by Wolpin (1987).

The effect of having access to unemployment benefits is very strong. Not only is the value of time higher for the unemployed worker with such benefits, but their offer arrival rate is also much lower (see Table 5). These results might indicate a lower search effort among this type of workers, reflected in a lower offer arrival rate. Moreover, the known fact of lower hazard rates for workers with unemployment benefits can be interpreted much more accurately within this structural estimation.[29] Another interesting result is the fact that the effect of having such benefits is not so negative before 1992, that is, when the benefits had a more decreasing pattern throughout the unemployment spell. Compared with the case of having no benefits, both before and after the 1992 reform, the mean unemployment duration has increased, other things equal, by 9.70% after the reform implemented in 1992. Given that the rest of the factors which could influence such duration are controlled by this difference-in-difference calculation, we can conclude that this increase must be caused by the reduction in the decreasing pattern in the replacement rate.

The estimated values of $\alpha(t)$ and $E(w)$ are quite reasonable: the offer arrival rate at the sample mean values of both the regressors and the effect of unobserved heterogeneity begins at 22.13% in the first month of unemployment and has a value of only 5.72% fourteen months later. This value is higher for those without unemployment benefits: 36.87% in the first month and 10.26% in the fourteenth. The estimated mean offered wage, $E(w)$, at the sample mean of the regressors, is 662.71 euros, around 828.55 dollars per month (at the December 1996 exchange rate), which is 16.26% lower than the mean monthly accepted wage in the sam-

seen in Ridder (1987). I would like to thank this author for his assistance regarding this issue.

[29] We should not forget, however, that the estimation is not totally structural with respect to this variable.

ple (see Table 1). Finally, the value of the parameter $b(t)$ is estimated to be quite high in the first month of unemployment, 1029.75 euros, basically because of the significant effect of the zero-duration dummy in the estimated value of unemployment time. However, it decreases rapidly as the spell of unemployment lengthens. This high initial value could be on account of the severance payments received after being separated from the previous job. Moreover, this coefficient could also be affected by the influence of accepted wages over the value of unemployment time in future unemployment spells (See the reasoning behind this dummy variable in the theoretical model).

Table 5 shows the effect of unobserved heterogeneity on the parameters of the model. The estimated distribution of unobserved heterogeneity reveals the existence of two groups: There is a 54.8% probability of the worker having a lower value of unemployment time and a lower mean offered wage. This effect of unobserved heterogeneity means that this group of workers have comparatively lower reservation wages and hence, higher hazard rates. The remaining 45.2% of the sample have higher values of both parameters. Hence, their reservation wages are higher and their hazard rates are lower than for the first group of workers.

Estimated reservation wages and hazard rates can be obtained according to these estimated parameters. Reservation wages decrease very quickly with unemployment duration (See Table 5), just as the theoretical model predicts, and are higher for skilled unemployed workers (820.06 euros for skilled workers and 724.17 euros for the unskilled ones in the first month of unemployment). Still, it is noticeable that reservation wages start off at quite high levels, although their decreasing pattern is very important during the fourteen months analyzed. For sample mean values of the regressors, the reservation wage in the first month of unemployment is 10.2% higher than the mean of the distribution of wages but, after 14 months of unemployment, the reservation wage is almost half the mean of offered wages in view of its highly decreasing pattern. This results mainly from the effect of the separation rate which seems to have a strong effect on reservation wages. If we evaluate these results with a 10% higher separation probability, reservation wages are much more time decreasing. Furthermore, the initial positive effect of the separation probability predicted in the theoretical model is present in our results for unemployment spells after 1992. That is, if the separation rate was higher, reservation wages would be larger, at least for the beginning of the unemployment spell.

Such low reservation wages after some months in unemployment lead, as in Van den Berg (1990), to high *acceptance probabilities* (see Figure 3 where we distinguish between those with and those without unemployment benefits). However, the acceptance probability begins at a low level, 41.14% at the beginning of the spell (opposed to a mean value of 77% in Van den Berg, 1990) although it grows rapidly, reaching a value equal to one in 9 months for the larger unobserved heterogeneity group and in 12 months for the smaller group. This time pattern is quite different to that found previously in other papers and, given our theoretical model, we can conclude that it is mainly due to the consideration of the separation probability in the estimation process.

The final result of this model deals with the unemployment hazard rate. This rate is the product of the offer arrival rate and the acceptance probability. As shown in Figure 4, where the estimated hazard is compared with the Kaplan-Meier empirical one, we find a smoother pattern of the estimated hazard, almost flat in the first six months and slightly decreasing afterwards. We can better understand the difference between the observed and the estimated hazards by studying the hazards of each of the two estimated groups of workers with respect to unobserved heterogeneity. Thus, we obtain that the group with a low mean of offered wages and low values of unemployment time has a larger and strongly decreasing hazard rate which almost matches with the observed hazard for durations of less than 7 months. The other group shows a slightly increasing but much lower hazard rate which is the result of having much higher reservation wages. Hence, the structural estimation procedure is able to identify not only the reason of the duration dependence of the hazard rate but also the presence and effect of unobserved heterogeneity.

As shown in Figure 5, the hazard rates for workers with and without access to unemployment benefits are very different. The known stylized fact found in some reduced-form estimations (see García-Perez, 1997 and Bover *et al.,* 2002), is also present here: a worker without unemployment benefits has higher probabilities of exiting unemployment in every month of the spell. But since we have already undertaken a structural estimation, we can interpret this result and conclude that in the early stage of the spell, the main element at work is the acceptance probability, which is much larger for those without unemployment benefits. However, once this probability is high enough, the difference between the two groups of workers remains the same because the offer arrival rate is still comparatively high for those without such benefits.

Finally, we have also performed some exercises to evaluate the goodness of fit of our model. Given the characteristics of the workers in the sample, I have simulated, using the structurally estimated parameters, their behavior throughout their unemployment spells. That is, I have simulated when they will exit from unemployment for those who actually do. We can see in Table 6a that the structural estimated parameters generate durations slightly longer than the real ones: 4.1 months versus 4.8 months in the simulation. The main difference is that the model generates somewhat fewer durations of less than 3 months. However, the model predicts the re-employment wage quite well. Another exercise, shown in Table 6b compares the duration distribution, nonparametrically estimated by the Kaplan-Meier estimator, with a reduced-form estimation and with our structural estimation. As can be seen, the differences are minimal although we continue to obtain slightly longer durations using our structural estimated parameters. In fact, the mean predicted duration using the Kaplan-Meier estimates is 4.88 months while under the structural estimation, this mean duration is 5.01 months.

To conclude, the estimation of the search model shows that Spanish unemployed workers do not particularly differ from unemployed workers in other countries: their acceptance probabilities are very high except for the first nine to twelve months of unemployment (see Wolpin, 1987 for US data and Van den Berg, 1990 for the Netherlands). Thus, the main mechanism at play in the process of becoming employed is the arrival of offers from employers. The offer arrival rate, in spite of its initially high value, is very low for workers who are unemployed for more than 12 months, the so-called long-term unemployed. Thus, this group of unemployed workers, along with the unskilled ones, have serious problems in leaving unemployment in Spain.

3.3 Policy evaluation: the effect of different Unemployment Insurance systems

I have used the structurally estimated parameters of our model to evaluate the current Spanish Unemployment Insurance system and also to compare it with alternative designs in terms of different replacement rates. The basic question we want to answer is whether a more decreasing pattern in unemployment benefits results in a shorter unemployment duration. Moreover, we will analyse whether a monotonically decreasing replacement rate is always optimal or not. Table 7 shows the results of this exercise where, firstly, we can confirm that a decreasing pattern in replacement rates is not always better than a flat rate in terms of lower unemployment duration.[30] Moreover, we can see that an initially increasing replacement rate which is decreasing only after the six first months in unemployment is much better in terms of the reduction of unemployment duration for workers with unemployment benefits. The only way of getting these results with a monotonically decreasing replacement rate is by using a system very similar to the one in the U.S. (rapidly decreasing, initial value much lower than the current one in Spain and a duration of benefits of only 6 months).

Recently, there has been a lot of controversy regarding the issue of whether the Unemployment Insurance system is creating or not the correct incentives for the unemployed to exit early from unemployment. In a general equilibrium model with moral risk, Hopenhayn and Nicolini (1997a) show that the optimal system is the one which undergoes a not so rapidly decreasing replacement rate together with an increasing tax on workers, depending on their unemployment duration. I would like to evaluate the effects of this system using the structural estimates of the non-stationary search model estimated here.[31] In Table 7 we can see that the optimal tax they propose[32] together with their proposed replacement rate

[30] Indeed, we need much lower replacement rates than those currently used in Spain in order to obtain substantial reductions of unemployment duration.

[31] I will use the calibrated values the same authors have obtained for the Spanish economy (see Hopenhayn & Nicolini, 1997b) although they are derived under no turnover.

[32] Hopenhayn & Nicollini (1997b) proposed a subsidy to quick job acceptance, going from 2.7% if the worker exits from unemployment after just one month till 0.1% if he exits after seven months in unemployment. For those with more than seven months in unemployment, they propose a tax increasing with unemployment duration. The tax goes from 0.3% in the eighth month of unemployment to 3.2% after 15 months.

(going from 93.2% to 87.6% in 15 months) helps to decrease mean unemployment durations by only 1.7%. We interpret this result as the effect of not taking into account the separation rate in a non-stationary environment.[33] As the probability of being separated from the job once the worker accepts the offer is quite high in Spain, the main incentive for the worker to exit as soon as possible is to be faced with a highly decreasing flow of income when unemployed.

When we combine the optimal tax they propose with an alternative design for the replacement rate much more decreasing with unemployment duration the effect is more pronounced. With an always decreasing replacement rate (going from 70% to 30% in 15 months) the reduction in mean unemployment duration would be about 4.4%. However, a flat replacement rate of, say, 60%, together with the optimal tax on workers results in an average unemployment duration 8.4% lower than it actually is.

However, once the employment risk is taken into account, as obtained in Wang and Williamson (1996), the complete set of instruments to provide incentives to the unemployed worker includes not only a tax on re-employment wages, initially negative if the worker accepts early, but also a non-monotonic pattern of the replacement rate which includes an initially low value in the first quarter of unemployment. Indeed, if the replacement rate is 40% the first three months in unemployment, increasing to even more than the current value now in Spain, that is, to 80%, and decreasing afterwards, the predicted mean unemployment duration would fall by 13.2% in Spain. Hence, we can conclude that the complete set of tools to provide both reward and punishment to the unemployed worker must include this penalty early in the unemployment spell together with the reward of having a negative tax on re-employment wages if the worker accepts quickly. Hence, in a labor market with a high turnover rate, as is the case in Spain (see García-Pérez and Muñoz-Bullón, 2004), the incentives for workers to quickly accept job offers come not only from what will happen after accepting such offer but also from what is happening in the current unemployment spell.

4 Conclusions

This paper presents a non-stationary job search model where jobs do not last indefinitely. When the unemployed worker is looking for a new job, he takes into consideration that once employed he may be unemployed again in the future. This future risk should mean, in a normal situation, a reduction in reservation wages. However, given the nonstationarity of the process, at the beginning of the unemployment spell, the worker can be in quite a good situation with respect to his expectations for the future. Therefore, we can also observe higher reservation wages in the first stages of the unemployment spell when the separation rate is

[33] Although these two authors also considers the possibility of turnover in a more recent paper, Hopenhayn & Nicollini (2002), they have no calibrated figures for the optimal unemployment insurance system in this case.

higher. Hence, in an economy where there is a high turnover rate and a high value of unemployment time, we can find longer unemployment duration because of this new effect of the separation probability found in this non-stationary environment.

I have implemented a structural estimation of this search model for the Spanish economy using data observed at discrete time intervals. Furthermore, the estimation procedure takes into account the presence of unobserved heterogeneity by using the Heckman and Singer (1984) mixture technique.

One of the principal results of this estimation is the flat shape of the reemployment probability (the hazard rate) up to the sixth month of the unemployment spell, although it then decreases with unemployment duration. This result is still true even when we control the presence of unobserved heterogeneity. Given the decreasing pattern of reservation wages, we can conclude that in the first months of unemployment the most influential factor is the rapid increase of the acceptance probability. However as soon as these first months pass the only element influencing the hazard rate is the offer arrival rate, because acceptance probabilities are, in fact, very close to one.

As for the other results, we have found that the worker who receives or has received unemployment benefits has a much lower probability of exiting unemployment. The reason for this changes depending on whether we are considering the early stages of unemployment or later stages. In the former case, this is because those with unemployment benefits have higher reservation wages and thus, lower acceptance probabilities. From the sixth month of unemployment onwards, however, the main difference is in the offer arrival rates, which are much higher for those without unemployment benefits.

Given this structural estimation, we have evaluated the Unemployment Insurance system in Spain. We find that a non-monotonic pattern in the replacement rate would result in a 9% decrease of the mean unemployment duration. Moreover, we conclude that the optimal tax proposed in Hopenhayn and Nicolini (1997a) helps to reduce unemployment duration even more, from 23.4 to 20.3 weeks, 13.2%. Hence, we find that the optimal tax on workers they propose needs an initially low level in the replacement rate so that it can function in the Spanish economy.

Appendix A: Proof of Theorem 1

This proof consists of: firstly, proving the following Lemma which basically requires that $\Delta w_i^0(t) < 0$ for (i) and (ii) are true, and that $\Delta w_i^0(t) < 0$.

Lemma 5 *If assumptions (A1)-(A6) are satisfied and if, for every* $t \in [0, T)$, $\Delta w_i^0(t) < 0$*, then:*

(i) $w_i^*(t) < w_i^0(t)$,

(ii) $\Delta w_i^*(t) < 0$.

Proof: Suppose that at some $t \in [0, T)$ $w_i^*(t) \geq w_i^0(t)$ holds. Then, because of the relationship between $w_i^*(t)$ and $w_i^0(t)$ we will have that $\Delta w_i^*(t) > 0$. However, given that $w_i^*(t)$ and $w_i^0(t)$ are continuous functions and, by the Lemma's assumptions, $\Delta w_i^0(t) < 0$, it cannot be true that $w_i^*(T) = w_i^0(T)$, which must be verified at time T given the assumptions of the model. Thus, the opposite must hold: $w_i^*(t) < w_i^0(t)$ and as this implies, $\Delta w_i^*(t) < 0$. $Q.E.D.$

Now we have to prove that $w_i^0(t)$ is a decreasing function of t under all the assumptions (K1)-(K4). The proofs for each of them are quite similar so we will show only the proof under (K1), i.e. for $b_i(t)$:

Given (7) we will have that:

$$w_i^0(t) - w_i^0(t+1) - \frac{\alpha_i(t)}{\delta_i + r}\left(G_i(w_i^0(t), t) - G_i(w_i^0(t+1), t)\right) = b_i(t) - b_i(t+1)$$

where $G_i(w_i^0(t), t) = \int_{w_i^0(t)}^{\infty} \left(w - w_i^0(t)\right) dF_i(w, t)$.

If $b_i(t)$ is decreasing in t, the right-hand side of this expression will be positive and since the function $w_i^0(t) - \frac{\alpha_i(t)}{r} G_i(w_i^0(t), t)$ is increasing in $w_i^0(t)$, we will have that $w_i^0(t) > w_i^0(t+1)$, that is, $w_i^0(t)$ is decreasing in how long the worker is unemployed.

Appendix B

In order to obtain monthly wages, we have compared the labor income and the unemployment benefits declared in the corresponding quarter. If there are no unemployment benefits, the monthly wage is the declared labor income divided by three. If there are unemployment benefits, their amount is compared with the labor income: if the benefits are larger than 80% of the labor income (70% for periods after 1992), then the monthly wage is the total amount declared as labor income. If the benefits are less than 80% of the labor income, the monthly wage is the labor income divided by two. This rule is based on the characteristics of the unemployment benefit system in Spain before and after 1992.

Calculation of monthly duration data is also based on comparing the labor income of the first and the last quarters the worker is unemployed, if positive, with the unemployment benefits received that quarter or with the labor income of the following quarter. If there is no labor income in the first quarter and the individual answers that he is unemployed, it is considered that he is unemployed during all the quarter. If the reported labor income is relatively low a duration of two months is imputed in the corresponding quarter but if this income is sufficiently high, it is considered that the worker has been unemployed for only one month in that quarter.

Appendix C

The separation probabilities imposed in the structural estimation have been previously estimated by means of a sample of 114,177 employment spells. This information comes from Social Security registers, that is, administrative data. We have estimated these probabilities as a discrete-time hazard rate (see, for example, García-Pérez & Muñoz-Bullón (2004) for a similar approach) only for men and with the same variables we used in the structural estimation. The only difference is that skill is measured in this dataset as the required qualification for the job and hence we have considered as skilled those in the highest qualification category. The results from this auxiliary estimation (the baseline hazard is dropped simply to save space) is shown in Table C.1. The mean separation rates in our sample is 4.94%, which is much lower for skilled workers (3.11%)than for unskilled ones (5.08%). It is also much higher for periods of high unemployment (5.82%), in the case of young workers (6.12%) and after 1992 (6.09%) than for the opposite categories (4.12%, 4.76% and 3.97% respectively).

Table C.1
Results of the hazard rate for exiting employment

Parameter	Coef.	t-ratio
Constant	-2.490	-83.49
Skill	-0.645	-23.72
Skill × Duration	0.116	14.08
Skill × Age18-29	0.108	5.95
Skill × High Unempl.	-0.070	-3.60
Skill × Before92	-0.263	-13.58
Age18-29	0.204	7.95
Age18-29 × Duration	0.049	4.29
Age30-45	-0.067	-2.43
Age30-45 × Duration	0.045	3.70
High Unemployment	0.457	25.02
High Unempl. × Duration	-0.150	-22.56
Before92	0.500	28.86
Before92 × Duration	-0.305	-42.21
Before92 × High Unempl.	-0.023	-1.46

Note : The likelihood value is -329,658.3 with 1,514,379 monthly observations.

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Table 1
Distribution of unemployment duration and other variables in the sample

	Completed Spells				Censored Spells		
	Percentage	Accepted Wage	Previous Wage	Total Income	Percentage	Previous Wage	Total Income
Months							
0-1	9.74				12.78		
1-2	8.21				14.00		
2-3	17.18				16.63		
3-4	17.69				6.90		
4-5	10.51				5.48		
5-6	9.49				7.30		
6-7	6.92				2.43		
7-8	6.92				1.62		
8-9	3.59				4.67		
9-10	3.33				2.84		
10-11	2.05				1.22		
11-12	1.79				4.67		
12-13	1.54				2.43		
13-14	0.77				2.03		
14-15	0.26				15.01		
Age1845	61.44	808.10	835.47	707.70		965.41	760.81
Age4565	38.46	764.00	813.34	1037.53		910.97	1179.18
Skilled	6.15	988.47	1044.05	783.87		1336.59	1025.58
With benefits	69.49	823.82	832.43	840.18		975.93	986.82
Sample Mean		790.88	826.96	836.76		939.04	986.44
No. Spells	390				353		

Note : Mean accepted wages are in 1996 Euros (exchange rate: 1.25 dollar/euro).

Table 2

Identification with the estimation procedure: some simulations results

Coef.	a_0	a_1	a_2	e_0	e_1	v_0	v_1	b_0	b_1	b_2	r_0	p	η_1	y_0
True	-1	-0.13	0.4	11.75	0.3	-3.4	0.4	12.25	-0.1	-0.3	2	0.5	0.25	-2
Estim.	-1.11	-0.14	0.52	11.78	0.27	-3.43	0.25	12.246	-0.098	-0.298	1.74	0.498	0.247	-2.01
St. Error	0.076	0.006	0.076	0.020	0.021	0.186	0.138	0.012	0.001	0.006	0.394	0.013	0.006	0.016

Notes : The parameters take the following form:

$$\alpha(t) = 1 - \exp(-\exp(a_0 + a_1 \times t + a_2 \times var1)),$$
$$E(w) = e^{e_0 + e_2 \times var2 + \eta},\ b(t) = e^{b_0 + b_1 \times t + b_2 \times var3 + \eta},$$
$$Var(w) = e^{v_0},\ \rho^2 = \frac{Var(w)}{Var(w) + Var(\varepsilon)} = \frac{e^{r_0}}{1 + e^{r_0}}, Var(y) = e^{y_0}.$$

The distribution of the unobserved heterogeneity term is a discrete one with two mass points. The probability of $\eta = \eta_1$ is p.

Table 3

Functional forms of the estimated parameters

Job offers arrival rate:

$$\alpha(t) = 1 - \exp(-\exp(\beta_1 + \beta_2 \times dur + \beta_3 \times benefits + \beta_4 \times benefits \times before92))$$

Distribution of wages:

$$W_o = \bar{W}(\eta) e^u e^\varepsilon \textit{ with } u \sim N\left(0, \sigma_u^2\right)$$
$$\varepsilon \sim N\left(0, \sigma_\varepsilon^2\right)$$

$$\bar{W}(\eta) \exp(\beta_5 + \beta_6 \times age4565 + \beta_7 \times \eta)$$
$$\sigma_u^2 = \exp(\beta_8 + \beta_9 \times skill)$$
$$\sigma_\varepsilon^2 = \exp(\beta_{10})$$

Value of time for the unemployed worker:

$$Y_{o_t} = b(t) e^y \textit{ with } y \sim N\left(0, \sigma_y^2\right)$$

$$b(t, \eta) = w_{-1} \exp(\beta_{11} + \beta_{12} \times dur0 + \beta_{13} \times dur + \beta_{14} \times benefits + \beta_{15} \times benefits \times before92$$
$$+ \beta_{16} \times benefits \times dur + \beta_{17} \times \textit{high unempl.} + \beta_{18} \times age4565 + \eta)$$

$$\sigma_y^2 = \exp(\beta_{19})$$

Table 4
Main results of the structural estimation

	without Unob. Het.		with Unob. Het.	
Parameter	Coef.	t-ratio	Coef.	t-ratio
$\boldsymbol{\alpha}_i(\mathbf{t})$				
Constant	-0.726	-2.20	-0.777	-2.30
Duration	-0.094	-8.36	-0.103	-8.97
Benefits	-0.858	-3.15	-0.846	-2.90
Benefits × *Before92*	0.097	0.47	0.113	0.55
$\bar{\mathbf{W}}_i(\boldsymbol{\eta})$				
Constant	11.526	155.36	11.547	160.88
Age45-65	0.047	0.22	-0.002	-0.06
$\boldsymbol{\eta}$			0.112	1.56
$\boldsymbol{\sigma}^2_{\mathbf{u}}$				
Constant	-2.668	-7.17	-2.615	-7.29
Skill	0.761	1.25	0.776	1.94
$\boldsymbol{\sigma}^2_{\varepsilon}$				
Constant	-2.789	-17.12	-2.775	-14.94
$\boldsymbol{\sigma}^2_{y}$				
Constant	-1.263	-56.88	-1.994	-163.98
$\mathbf{b}_i(\mathbf{t}, \boldsymbol{\eta})$				
Constant	-0.251	-6.81	-0.279	-10.19
Duration	-0.064	-6.34	-0.051	-8.59
Duration = 0	0.112	4.23	0.115	5.77
Benefits	-0.031	-0.73	0.034	1.24
Benefits x Duration	0.070	5.46	0.059	8.05
Benefits x Before92	-0.067	-2.72	-0.202	-22.51
High Unemployment	-0.161	-7.83	-0.070	-9.30
Age 45-65	0.429	24.54	0.401	64.06
Unobs. Het.				
η_1			-0.356	-20.48
p			0.548	25.16
Log-likelihood	-50,351.44		-49,819.44	
No. of observ.	4,375		4,375	

Table 5

Predicted values for the main elements of the model

t	$\alpha(\cdot)$	$\phi(\cdot)$	$\bar{F}(w^*(\cdot))$	$w^*(\cdot)$	$b(\cdot)$
Mean values for all variables:					
0	22.13	9.10	41.14	730.42	1029.75
4	15.25	9.42	61.74	618.74	898.25
14	5.72	5.66	99.98	323.89	851.00
For the group with η_1 :					
0	22.13	13.07	59.08	600.26	667.38
4	15.25	12.44	81.54	501.83	582.16
14	5.72	5.72	99.98	245.56	551.53
For the group with η_2 :					
0	22.13	4.28	19.35	888.86	1470.07
4	15.25	5.75	37.68	760.79	1282.35
14	5.72	5.59	97.76	419.06	1214.89
With access to Unempl. Benefits (after 1992):					
0	19.70	7.40	37.57	768.86	1174.46
4	13.51	8.16	60.40	633.16	1050.23
14	5.58	5.58	99.94	259.51	950.45
Without access to Unempl. Benefits (after 1992):					
0	36.87	14.77	40.07	733.76	1010.04
4	26.23	18.54	70.67	585.67	734.17
14	10.26	10.26	100.00	160.58	440.87
Skilled workers:					
0	22.13	7.02	31.72	820.06	1186.48
4	15.25	6.75	44.24	714.59	1034.97
14	5.72	4.53	79.28	479.08	980.52
Unskilled workers:					
0	22.13	9.27	41.91	724.17	1016.45
4	15.25	9.57	62.79	614.25	886.66
14	5.72	5.67	99.15	322.32	840.02

Notes : $\bar{F}(w^*(\cdot)) = 1 - F(w^*(\cdot))$. The first three columns are percentages and the other two are expressed in 1996 euros. The predictions are carried out using the model with unobserved heterogeneity.

Table 6a: Goodness of fit tests:
simulated data for complete unemployment spells

	Sample Values	Simulation
Mean duration	4.06	4.80
Duration (Stand. dev.)	3.01	3.98
% dur. $\in [0, 3]$ months	52.82%	46.45%
% dur. $\in [4, 6]$ months	26.92%	22.15%
% dur. $\in [7, 14]$ months	20.26%	31.40%
Re-employment wage (mean)	790.89	753.40
Re-employment wage (stand.dev.)	259.49	190.60

Notes : The estimated parameters have been used for predicting the duration in unemployment and the re-employment wage of each worker in our sample. This table shows the predicted values once we have generated 500 different simulated durations and re-employment wages for each worker in the sample.

Table 6b: Goodness of fit tests:
comparing reduced-form and the structural estimations

Survival Prob. at duration:	Kaplan Meier estimation	Reduced form estimation	Structural estimation
0	100.00%	100.00%	100.00%
1	94.29%	97.71%	90.90%
2	90.29%	92.61%	82.47%
3	**80.00%**	**84.64%**	**74.73%**
4	68.33%	75.13%	67.68%
5	60.86%	65.73%	61.31%
6	**53.58%**	**57.47%**	**55.60%**
7	47.74%	50.68%	50.52%
8	41.60%	45.24%	46.03%
9	**38.30%**	**40.92%**	**42.10%**
10	34.93%	37.46%	38.66%
11	32.62%	34.71%	35.68%
12	**30.51%**	**32.53%**	**33.11%**
13	28.38%	30.87%	30.88%
14	27.16%	29.69%	28.97%

Note : The reduced form estimation has been carried out with the same variables used in the structural estimation and a polynomial on the logarithm of duration for capturing the duration-dependence in the hazard rate.

Table 7

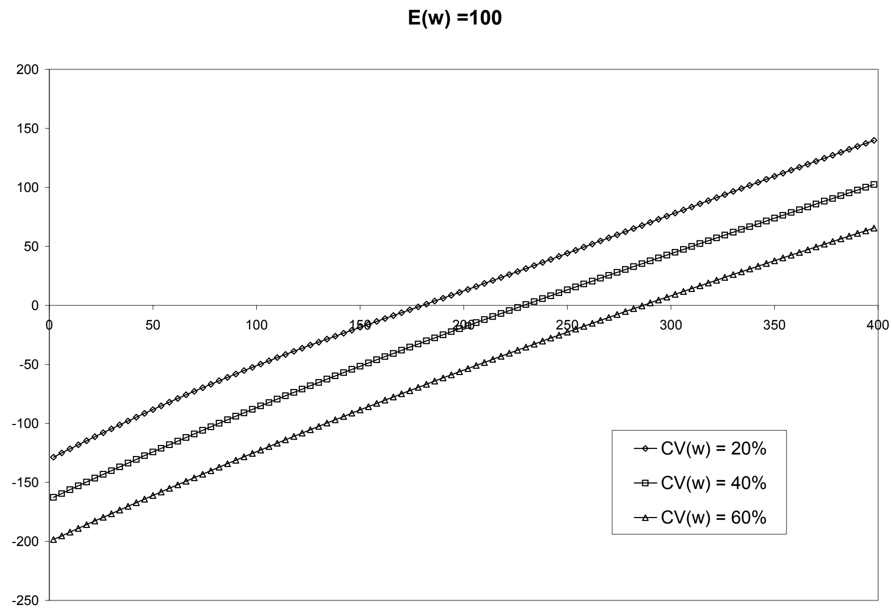
Simulations with different Unemployment Insurance regimes
Mean predicted durations

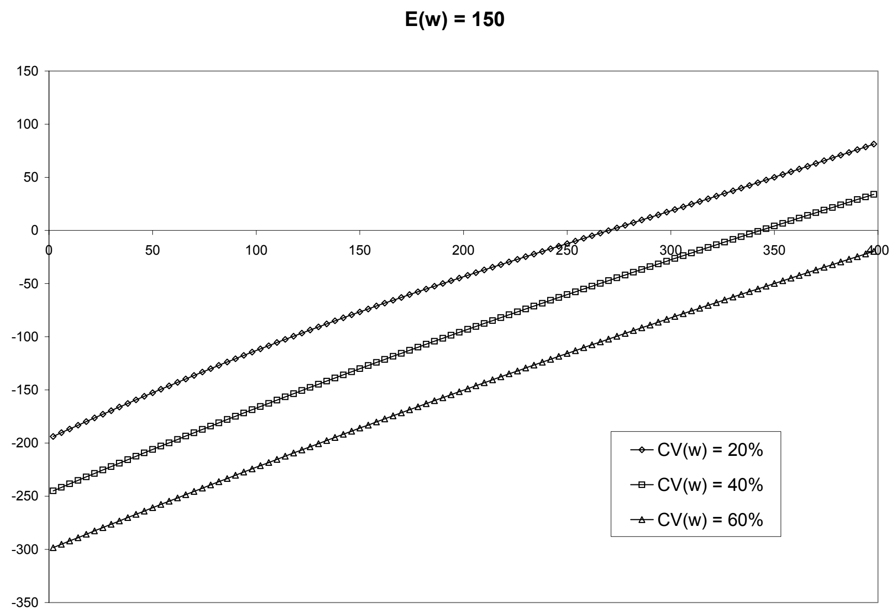
	Without optimal tax		With optimal tax	
	Weeks	Difference	Weeks	Difference
Without U.B.	**16.5**			
With actual U.B.	**23.4**		22.5	-3.9%
70% (always)	23.6	+0.6%	22.7	-3.3%
60% (always)	22.4	-4.5%	21.5	-8.4%
50% (always)	21.3	-9.1%	20.4	-13.0%
Hopenhayn & Nicollini (1997) proposal	23.9	+2.3%	23.0	-1.7%
70% - 60% - 50% - 40% - 30% (each 3 months)	23.3	-0.6%	22.4	-4.4%
60% - 50% - 40% - 30% - 20% - 10% - 0% (each month)	21.4	-8.7%	20.6	-12.1%
40% - 80% - 70% - 60% - 50% - 40% (each 3 months)	21.3	-9.0%	20.3	-13.2%

Note : The predicted durations are calculated under the mean values for all variables but the separation rate which is imposed to be the mean one in our sample from 1992 onwards (6.10%).

Figure 1: Simulation Results:

Effect of $b(0), E(w)$ and $CV(w)$ on the condition of Proposition 3





Note: The baseline parameters for these simulations are $b(t) = 200*\exp(-0.04t)$, $\alpha(t) = 0.3 * exp(-0.1t)$, $E(w) = 100$, $CV(w) = 40\%$, $r = 0.3\%$, $\delta = 5\%$. The parameter in the horizontal axis is $b(0)$.

Figure 2: Histogram of the re-employment wages

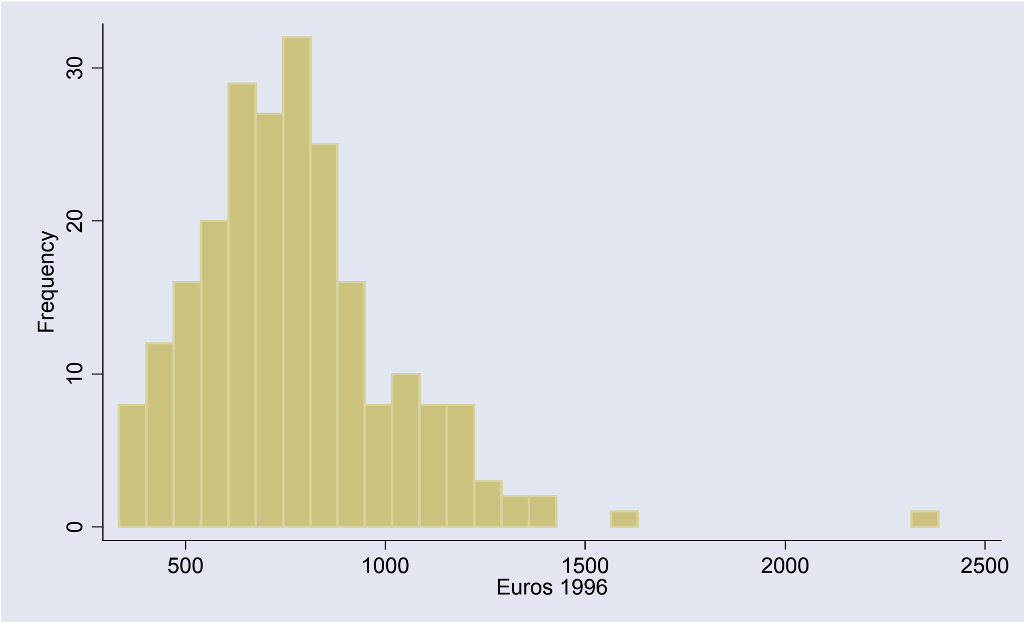


Figure 3: Estimated acceptance probabilities: the effect of unemployment benefits

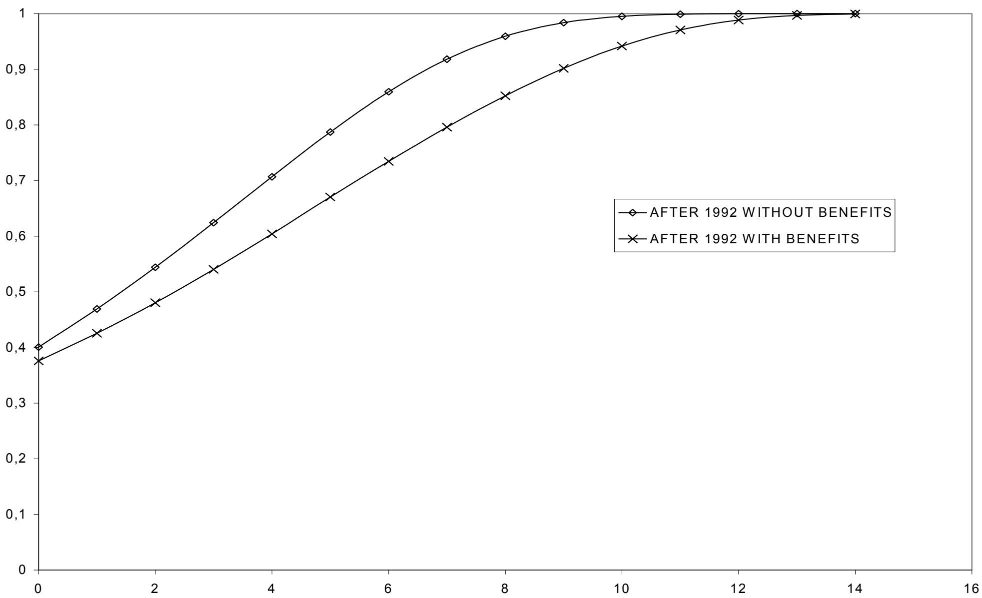


Figure 4: Kaplan-Meier empirical hazards and estimated hazard rates: the effect of unobserved heterogeneity

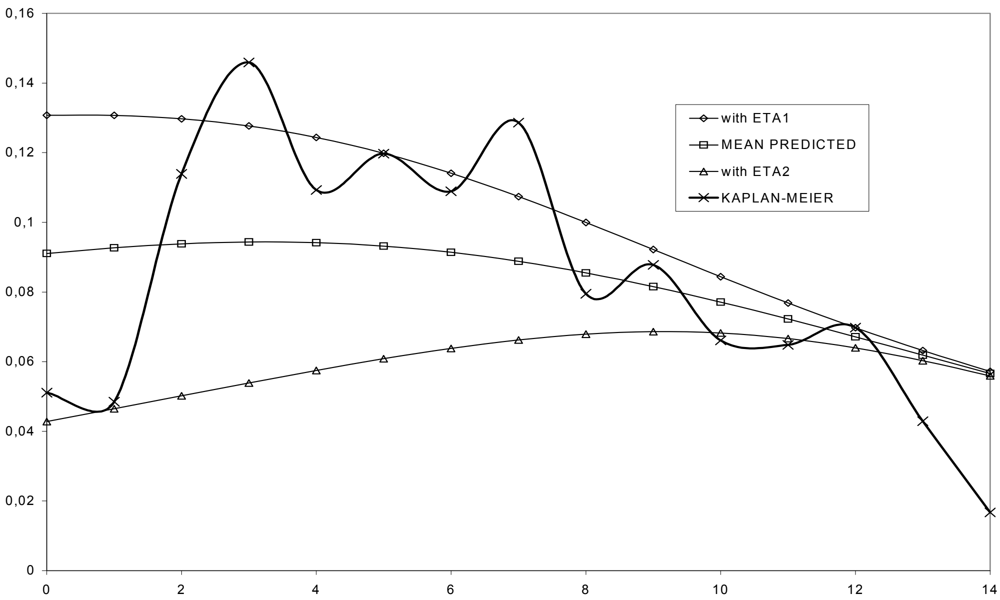


Figure 5: Estimated hazard rates: the effect of unemployment benefits