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Abstract

This paper studies housing tenure and labour mobility using individual data from the ECHP for five European countries. First, the effect of housing tenure on the unemployed workers' labour mobility is studied using a discrete unemployment duration model with two alternative exits to employment, depending on whether they are associated with a residential change or not. Ownership is found to affect geographical mobility negatively. Second, the results are robust to potential endogeneity of the ownership status and institutional differences across countries. Third, post-unemployment wages are studied. We do not find any effects of the unemployment spell duration and the geographical mobility on wages after controlling for the self-selection bias.

Keywords: Labour mobility, Housing tenure, Duration models, Self-selection bias, Wage equation.

JEL classification: J61, R20, J31.

1 Introduction

The purpose of this paper is to provide empirical evidence on the impact of housing tenure on the unemployed workers' geographical mobility decision in five European countries, taking into account their institutional differences. To study the relationship between housing tenure and labour mobility, I use a sample of unemployed family heads from the European Community Household Panel (hereafter, ECHP) for the period, 1994-98.

Much empirical work has addressed the topics of mobility and housing tenure. In the case of the UK, Hughes and McCormick (1981) and Henley (1998) find a negative impact of social rented housing and ownership on migration, mainly from negative housing equity. These studies as well as others, like Hughes and McCormick (1994) and Pissarides and Wadsworth (1989), have found that geographical mobility does not respond to regional economic conditions, such as high unemployment rates.

Unlike papers mentioned above, this paper addresses the issue of how the owner-occupied housing may affect exits from unemployment negatively by making the home-owners more reluctant to accept a job implying a residential change. The analysis is carried out using microdata and only focuses on the unemployed workers' geographical mobility decision, so the more natural way of implementing this study is to use unemployment duration models with multiple transitions from unemployment. I focus on the unemployed, since employed workers' geographical mobility responds to different factors, possibly related to their current job characteristics.

Other microdata studies like van Leuvensteijn and Koning (2004) and Böheim and Taylor (1999) focus on other related issue of how home ownership and commuting affect the employed workers' job mobility decision in the Netherlands and the UK, respectively. In small countries where regions are connected by high-speed trains and other very good communications, commuting time becomes an important factor deterring from job mobility and the negative effect of home ownership disappears. However, in large countries like Spain and Italy, where regions are not well-communicated with each other, the structure of housing markets becomes an important factor explaining the geographical mobility decision.

Housing and geographical mobility are also two important issues in Spain, given the high unemployment and ownership rates, 15.2% and 83.4% in 1998 respectively,[1] which are among the highest in the European Community. Empirical work studying the response of mobility to differentials in regional labour markets in Spain can be found in Bentolila and Dolado (1990), Bentolila (1997), Antolín and Bover (1997), and Bentolila and Jimeno (1998). However, these authors do not study the effect of housing tenure on the unemployed workers' geographical mobility decision using microdata, possibly due to the absence of a fixed panel that allows to study both topics in Spain until the construction of the ECHP.

The interest in analysing housing tenure and labour decisions using the ECHP is that this survey is close to a merge of a Labour Force Survey and a Family Expenditure Survey; it provides more information for addressing both topics than any of either survey separately. This allows the joint study of housing, residential and labour mobility, since the survey keeps track of migrant households.

Moreover, the implementation of this survey in other European countries allows all countries to be studied jointly. This introduces exogenous variation in the analysis of housing and mobility, since the way in which individuals choose their housing tenure depends greatly on the institutional characteristics prevailing in each country. This is the main contribution of this paper to the studies of housing and labour markets.

The choice of housing tenure is predetermined to a large extent by the institutional characteristics. As housing is a basic good, policy makers have generally tried to make the access to a dwelling easier, by providing social rented housing and subsidies, and

[1]The harmonised unemployment and home ownership rates are constructed by Eurostat, using the European Labour Force Survey (EU-LFS) and the European Community Household Panel (ECHP), respectively.

allowing tax reliefs if individuals rent a house. Home ownership has also been supported, for instance through the interest relief in income taxes, the tax relief on maintenance of the house, and the absence of taxes on imputed rents and on capital gains derived from the sale of houses. Often, these policies have benefited one housing tenure more than the other, resulting in the dominance of one housing tenure over the other in most European countries.

The empirical study will focus on France, Germany, Italy, Spain, and the United Kingdom. This choice is based on the variation that these countries provide in the way that their institutions promote the access to housing. That variation helps to explain the link between home ownership and labour mobility observed in a country.

Using the information on the financial and housing markets in European countries during the 1990s that Maclennan *et al.* (1998) collect, those five countries can be sorted into two groups according to the relative importance of the two main types of housing tenure: ownership and rental. A first set is formed by Germany and France, in which a high percentage of households lives in rental housing (private or social accommodation), 62% and 38% of the housing stock, respectively. A second group includes Spain, Italy and the United Kingdom, where ownership is clearly the preferred alternative; 78% of the housing stock in Spain is occupied through ownership, and 68% and 67% in Italy and in the UK, respectively. In these countries, the yearly inter-regional mobility rates as a percentage of population are: 1.23% and 1.07% in Germany and France, 0.5% and 0.56% in Italy and Spain, and finally 1.58% in the UK (1993 data). Thus, it seems that in the aggregates there exists a negative relationship between ownership and inter-regional mobility rates in these countries, except for the UK. Regarding the unemployment rate, there is also a negative relationship between unemployment and inter-regional mobility rates and a positive one between ownership and unemployment rates, since Germany and Spain have the lowest and the highest unemployment rates in 1993, 7.6% and 22.2%, respectively. Meanwhile, the unemployment rates in France, the

UK and Italy are 11.7%, 10.3% and 10.2%, respectively.[2]

The empirical model of housing tenure and labour mobility that I use is a discrete unemployment duration model with two alternative exits to employment, depending on whether they are associated with a residential change or not. These transitions to employment are assumed to follow a multinomial logit specification. Using this sample of unemployment spells, I also investigate the main determinants of the housing tenure status by taking into account the institutional differences across countries. The ownership status and labour mobility are assumed to be both correlated through unobserved heterogeneity.

Finally, I also analyse the reemployment wage that individuals obtain when they leave their unemployment spell. Policy-makers should foster housing policies that encourage geographical mobility if this helps the unemployed workers to obtain a job with better work conditions. In particular, I study two aspects: first, whether the duration of the unemployment spell influences the attained wage level and, second, whether geographical mobility helps individuals to obtain a higher wage after leaving unemployment than the one they would have obtained in their local area.[3] First, I estimate the mean of the wage distribution by controlling for self-selection bias; and second, I estimate both transitions to employment and the mean of the wage distribution jointly using the generalised method of moments, in order to obtain consistent and more efficient estimates under the existence of self-selection bias. The bias arises due to the fact that a subsample of individuals decide to remain unemployed when the wage offers they receive are lower than their reservation wage.

The main empirical result is that owners are more reluctant to move than renters; and this result is reinforced when I allow for the presence of unobserved heterogeneity.

[2]Source: *OECD Economic Outlook*, No. 69, 2001.

[3]It is not clear in which direction the wage level should move when the unemployed migrate due to job-related reasons. It depends on whether the reservation wage for accepting a job in other region is higher than that in their local area or not, which is itself related to their expectations. This issue will be explained in detail in the following sections.

So, for fostering geographical mobility, policy-makers should encourage housing policies making private rental more attractive. As expected, in the local labour market owners behave in a similar way to renters. We do not find any effects of the unemployment spell durations and the geographical mobility on reemployment wages.

The rest of paper is organised as follows. Section 2 describes the empirical duration model of housing tenure and labour mobility and the results. Section 3 addresses the potential endogeneity of the ownership status by allowing for unobserved heterogeneity in both transitions to employment. Section 4 studies the wage attained by individuals when they leave unemployment through one of these two alternatives: exit to a job associated with a residential change and exit to a job without residential change. The former alternative is identified as a case of geographical mobility in the empirical model. Finally, Section 5 summarises the conclusions.

2 An empirical model of housing tenure and labour mobility

In this section, I try to find out whether housing tenure affects the unemployed individuals' geographical mobility decision, in particular, their incentives to accept a job in an area that implies a residential change. It is thought that owners are more reluctant to move to other region to work than renters, since they have higher moving costs.

The empirical study is to test the two main predictions about how housing tenure affects the individuals' labour decisions derived in Barceló (2005), for a sample of unemployed family heads in five European countries. For this purpose, differences in the institutional characteristics across countries are captured in two different ways: first, I include country dummies in order to find out whether these differences can be identified in the sample and afterwards, I replace these dummies by indicators of relevant characteristics of housing and labour markets affecting the labour mobility decision.

Barceló (2005) develops a two-region model of job search in which individuals take three decisions jointly each period: the choice of housing tenure, the region in which to live and the acceptance or rejection of the received wage offers. In this model, owners are less willing to accept a job in other region for two reasons. First, individuals are assumed to obtain housing services more cheaply if they live in an owned house than in a rented house. Second, individuals who become owners by buying a dwelling have to incur some transaction costs at the moment that this purchase takes place (transaction tax, stamp duty, and so on). These two assumptions make owners have a higher reservation wage than renters for accepting a job in another region. Thus, the probability of geographical mobility will be more reduced for this housing tenure type. However, the behaviour in the labour market in the local area in which individuals live will be identical for individuals in both tenure regimes. That is, housing tenure regime will not affect the decisions of acceptance or rejection of job offers not involving a residential change; the reservation wage will be identical.

Both conclusions, lower mobility among owners and identical decision rules in local labour market, are tested using the empirical model described in the following subsections. Subsection 2.1 describes the data used in the estimation, Subsection 2.2 explains the empirical model and the estimation method implemented, and finally Subsection 2.3 presents the results.

2.1 Data characteristics

The data used in the estimates come from the European Community Household Panel (ECHP); five waves of the survey are used, covering the period 1994-98. The ECHP consists of a fixed panel with an annual frequency; nevertheless, it also provides information on the main economic activity in each month of the previous year for all household members aged 16 years old or more. Thus, we can construct monthly duration of the unemployment spells for each individual.

The choice of the individual data coming from this survey is due to three reasons. First, the ECHP is like a merge of a Labour Force Survey (LFS) and a Family Expenditure Survey (FES); that is, it contains rich information on individuals' labour status and on housing characteristics at the same time. Separately, neither survey allows the study of housing and labour issues jointly. Second, the LFS does not include any information on income. Third, the most important reason is that the ECHP is a fixed panel that follows and interviews all the households belonging to the sample, even if they change residence. Meanwhile, rotating panels, such as the LFS and the FES, do not interview those households that change address.

Therefore, we need a fixed panel like the ECHP to study labour and residential mobility. In addition, it helps us to make comparisons across European countries, since this survey has been implemented in the same way and at the same time in all countries.

Labour market decisions are made individually by each household member, but housing tenure and mobility decisions are taken by the whole household, mainly by the family head. So, the sample is formed by household heads. In order to obtain a homogeneous sample of individuals whose link to the labour market is stable, I restrict the sample to family heads aged 25 to 64 years old with previous labour market experience.

The empirical approach that I use to study how housing tenure affects labour mobility is a discrete unemployment duration model with two alternative exits to employment, depending on whether they are associated with a residential change or not.

The stock sample is formed by those family heads who entered unemployment from January 1994 to December 1997.[4] I consider that a residential change occurs when a change of address is produced at some point of the unemployment spell or up to two months after the family head finds a job. This typically happens near the end of the unemployment spell.

An important limitation of using the ECHP in order to study geographical mobility is its reduced geographical breakdown of residential changes, since we can only know whether it took place within the same province, from outside the province but within the same country or from other country. This means that we cannot distinguish intra-town from town-to-town residential changes. Given the limited inter-regional mobility observed in Spain and in other European countries, and given the great importance of intra-regional movements nowadays, I consider that there is geographical mobility when an unemployed individual finds a job and this can be associated with a residential change in the way explained above. Nevertheless, I am aware that this measure can be potentially contaminated with residential changes due to other personal or housing reasons, non job-related. This is one of the reasons for estimating the labour mobility and the housing tenure status decisions jointly, so as to improve the estimates given that I use an imperfect measure of geographical mobility. An additional motivation for making housing tenure status endogenous to geographical mobility is to allow for the presence of unobserved heterogeneity, which is potentially correlated across decisions.

[4] The information on the main economic activity in each month of the year refers to the year prior to the survey, so that I do not have this information for 1998, the last wave of the survey included in the sample.

Before describing the empirical approach, I will comment on some of the characteristics of the sample used in the estimation. In Table A.1 of the Appendix we can distinguish two sets of countries according to the housing tenure regime observed in the sample of unemployment spells: a first group formed by Germany and France, where rental is the predominant tenure regime in 64% and 63.4%, respectively, of the households whose family head is unemployed. A second group includes Spain and Italy in which 82.6% and 68.9% of the households live in an owned house, whereas almost a half of the households lives in rental housing in the UK. Column 3 shows the geographical mobility rate in each country, here identified as an exit to employment associated with a residential change. Again, we can notice the existence of a negative relationship between ownership and geographical mobility. Finally, column 4 gives the proportion of observations coming from each country in the sample; 46.7% of the observations correspond to Spain, which reflects that the Spanish unemployment rate is much higher than in the remaining countries. Note that only three waves of the ECHP survey are available for Germany and the UK (from 1994 to 1996).

Table A.2 shows the main individual characteristics in the sample of unemployment spells, distinguishing whether the exit is associated with a residential change or not. First, we can observe that a high percentage of geographical mobility happens among renters, 73.69% against 26.31% among owners; moreover, 60.53% of migrants are tenants living in a private rented house. Second, geographical mobility is likely to occur among individuals with higher levels of education, mainly among those having completed the second stage of secondary education. Third, concerning household composition, single individuals not cohabiting in a relation are more mobile, since 26.32% of the individuals exiting to a job spell after changing residence are single, against the 19.79% of the individuals that found a job in their local labour market. Among those living with a partner, individuals whose spouse or partner is working are less likely to move. When children are younger, aged 6 years old or less, households seem to be more mobile. Regarding the characteristics of the previous job, people having worked in the service sector are also more likely to move, whereas those having previous experience in agriculture are almost immobile. Finally, the family head's gender seems not to affect the geographical mobility decision, and individuals having job tenure longer than one year in the previous job are the least likely to move.

2.2 Econometric model of housing tenure and labour mobility

The empirical approach consists of estimating a discrete unemployment duration model with two alternative exits to employment, depending on whether they are associated with a residential change or not. The stock sample is formed by N multiple unemployment spells provided by I unemployed family heads, who have entered an unemployment state possibly for several times since January 1994. The baseline of our sample is January 1995, the starting point at which the exits to employment with each of the two alternatives are observed to occur. Each spell and its length are denoted by the subscript i and by T_i, respectively. The elapsed duration, T_i, can be broken down into $\tau_i + r_i$; τ_i denotes the time spent in the unemployment state between the moment at which the individual becomes unemployed and the moment at which the exits to employment start to be observed, that is, January 1995; and r_i denotes the time spent in the unemployment state between January 1995 and the moment at which the exit to employment occurs. As we can observe entrants to unemployment, I avoid the problem of left-censoring, which happens due to the lack of knowledge of the exact date of the beginning of the unemployment spell. Spells can only be complete or right-censored; the latter case happens when the individual stops being interviewed by the ECHP before he finds a job or when he enters a non-employment spell like retirement, education, etc.

The econometric method is similar to that used in Bover and Gómez (2004) and Barceló (2005). Let D_i be the indicator taking the value of 1 if the exit is associated with a residential change, and the value of 0 if the exit is not associated with a residential change. The transition intensity to employment with the alternative k (with residential change or not, $k = 1,\, 0$, respectively), denoted as $\theta_k\left[t \mid X_i(t)\right]$, conditional on a vector of individual characteristics, $X_i(t)$, is defined as the probability of leaving unemployment at t months with the exit k given that the individual has been unemployed for at least t months:

$$\theta_k\left[t \mid X_i(t)\right] = \Pr\left(T_i = t,\, D_i = k \mid T_i \geq t,\, X_i(t)\right), \quad k = 0,\, 1 \tag{2.1}$$

The hazard function, $\theta\left[t \mid X_i(t)\right]$, or exit rate from unemployment conditional on the characteristics vector, $X_i(t)$, can be obtained as the sum of both transition intensities:

$$\theta\left[t \mid X_i(t)\right] = \Pr\left(T_i = t \mid T_i \geq t,\, X_i(t)\right) = \sum_{j=0}^{1} \theta_j\left[t \mid X_i(t)\right] \tag{2.2}$$

These transition intensities are assumed to follow a multinomial logit specification as follows:

$$\theta_k\left[t \mid X_i(t)\right] = \frac{\exp(X_i(t)'\beta_k)}{1 + \sum_{j=0}^{1} \exp(X_i(t)'\beta_j)}, \quad k = 0, 1 \tag{2.3}$$

The vector of characteristics, $X_i(t)$, includes a second-degree polynomial in the logarithm of duration as well as a constant, in order to capture duration dependence,[5] household and family head characteristics, aggregate economic variables and country dummies reflecting the differences in institutional characteristics across countries.

The estimation method consists of maximising the joint log-likelihood function defined for both transitions.[6] Using the relationships between distribution and density functions conditional on $X_i(t)$, the transition intensities of each alternative and the exit rate described in equations (2.1) and (2.2), we obtain the contributions of each un-

[5]Due to the small size of the sample, which consists mainly of exits not associated with a residential change, I specify duration dependence as a second-degree polynomial in the logarithm of duration instead of estimating it semi-parametrically with duration dummies, as in Meyer (1990).

[6]As analysed in Bover and Gómez (2004), when the transition intensities follow a multinomial logit specification, this estimation method is equivalent to estimating a *competing-risks model* for each exit separately. Thus, both provide consistent estimates of the same parameters; however, the first method produces more efficient parameters estimates, since they are obtained jointly. Given the small size of the sample, I have chosen this method.

employment spell to the likelihood function. Thus, as we have a stock sample, complete spells contribute to the likelihood of the probability of exiting at t months, $t = \tau_i + r_i$, with one of the two alternatives, k, conditional on τ_i, the time spent in the unemployment spell prior to the baseline (that is, before January 1995), as follows:[7]

$$\Pr\left(T_i = \tau_i + r_i,\ D_i = k \mid T_i > \tau_i,\ X_i(t)\right) = \theta_k\left[\tau_i + r_i \mid X_i(t)\right] \prod_{s=\tau_i+1}^{\tau_i+r_i-1} \left(1 - \theta\left[s \mid X_i(s)\right]\right),\ k = 0,\ 1 \tag{2.4}$$

Censored spells contribute to the likelihood of the probability of finding a job after having stayed unemployed for more than t months, conditional on the time spent in the unemployment spell prior to the baseline:

$$\Pr\left(T_i > \tau_i + r_i \mid T_i > \tau_i,\ X_i(t)\right) = \prod_{s=\tau_i+1}^{\tau_i+r_i} \left(1 - \theta\left[s \mid X_i(s)\right]\right) \tag{2.5}$$

Let c_i be the indicator of lack of censoring of spell i with duration t_i, $t_i = \tau_i + r_i$, then the joint log-likelihood function can be expressed as:

$$\begin{aligned} L(\beta) = \sum_{i=1}^{N} \{ & (1 - D_i)\, c_i \log \theta_0(t_i \mid X_i(t_i)) + D_i c_i \log \theta_1(t_i \mid X_i(t_i)) + \\ & + c_i \sum_{s=\tau_i+1}^{t_i-1} \log\left[1 - \theta_0(s \mid X_i(s)) - \theta_1(s \mid X_i(s))\right] + \\ & + (1 - c_i) \sum_{s=\tau_i+1}^{t_i} \log\left[1 - \theta_0(s \mid X_i(s)) - \theta_1(s \mid X_i(s))\right] \} \end{aligned} \tag{2.6}$$

Bover *et al.* (2002) and Bover and Gómez (2004) explain how to rewrite the log-likelihood function as the concatenation of the log-likelihood functions defined for the survival subsample in each duration, when the sample is formed by flow data. In this way, the estimation of the parameters of the transition intensities is made easier. In case of a stock sample, the concatenation is done for the survival subsamples in each duration of the spell that happens after the baseline, that is, we append the survival subsamples of durations from $\tau_i + 1$ to $\tau_i + r_i$ months.

[7]Unemployment spells starting after January 1995 are treated as flow data, since the unemployment spell duration elapsed before the baseline is zero months, that is, $\tau_i = 0$ and $r_i = t$ (see Lancaster, 1990).

Explanatory variables The explanatory variables included in the vector $X_i(t)$ are the following. First, the institutional differences will be captured by two different ways: by introducing country dummies in the estimates except for Spain (omitted category) and by including some indicators of relevant housing and labour market characteristics affecting the labour mobility decision. In particular, I use the information extracted from Maclennan *et al.* (1998) about the percentage of transaction costs on house prices, estimated for each country. This variable constitutes a proxy of the transaction costs incurred in the purchase of a house, and it is introduced in logarithms in the estimates.

I also use an index measuring the strictness of the employment protection legislation (hereafter, EPL) constructed by the OECD (1999).[8] This index is also entered in logarithms. The strictness in the employment protection legislation refers to the degree of protection of regular and temporary employment and against collective dismissals. See the Appendix for a detailed explanation of the different aspects of the employment protection legislation covered by the EPL index. Using this index, these five European countries could be classified in the following way: Germany and France are countries with a medium degree of strictness, Italy and Spain have great strictness in employment protection, and finally the United Kingdom shows a very low degree of employment protection.

Finally, the last institutional characteristic considered is an overall average of unemployment benefit replacement rates (%) that vary by unemployment duration and family circumstances (single, with dependent spouse or with spouse in work). This index is extracted from OECD (1994) and its logarithm is introduced in the estimation.

Second, the only aggregate economic variable considered is the quarterly national unemployment rate in logarithms, plus seasonal indicators, using the fourth quarter as the omitted category. I have not introduced a measure of real house prices due to the difficulty in finding it for all countries. Moreover, I have not included any Consumer

[8]This indicator is an average of different indices evaluating the employment protection across countries, constructed for the late 1980s and 1990s. For the most recent date, this indicator also collects information on the protection against collective dismissals. As the remaining indices do not vary greatly from the first to the second period, I have used the latter.

Price Indices for Rent and Rental Equivalence, due to the fact that these indices refer to different national monetary currencies and quantities, although they are expressed in the same base year. In addition, they cannot be used for making comparisons across countries, due to other reasons commented on by Deaton (1998).

A third set of variables collects information on the household head's previous job. There are indicators of the economic sector, only for industry and services; previous jobs in agriculture and construction are considered as the omitted category, given the shortage of exits associated with a residential change for these sectors. I include an indicator of working time, taking the value of 1 for a full-time job, the logarithm of monthly experience at the previous job, and an interaction of experience with the logarithm of unemployment duration.

Experience variables will not only pick up the positive effect of experience on the exit rate to employment, but they will also capture the effect of the entitlement to receive unemployment benefits, since this is mainly determined by tenure at previous job. In this sense, the experience variables will also reflect that: the greater the experience is, the higher the level of unemployment benefits and the longer the duration of its receipt. Thus, the greater the experience, the higher the reservation wage is; that is why the sign of its coefficient is not determined.

I have not included income nor the amount received related to benefits, although this information exists in the ECHP. This is due to the fact that the information on income and on main economic activity in each month is contemporary, so that the inclusion of their lags would imply the loss of a great number of observations. Given the reduced size of the sample, I prefer not to introduce them in the estimates. In addition, the amount of benefits can only be assigned to each unemployment duration in an imperfect way: I do not know the accurate dates in which they were received within the same year, so that this variable would suffer from large measurement errors.

However, I construct a dummy indicating whether the individual received any amount of unemployment benefits during the years in which the unemployment spell happens.

Finally, a fourth set of variables refers to personal characteristics such as the level of education (including an interaction of the highest level with the logarithm of unemployment duration), the family head's gender, and the logarithm of family head's age. As indicators of the household's composition and size, I have included a dummy for whether the individual lives with a partner (married or not), an indicator of whether the spouse or the partner is employed, and the number of children aged 18 years old or less in the household and its square.

After having controlled for such observed heterogeneity, I introduce an indicator taking the value of 1 if the individual is a homeowner, and the value of 0 if he is a renter, in order to find out whether housing tenure affects the geographical mobility decision. Moreover, I include a dummy variable revealing the existence of outstanding mortgages among owners.[9]

As Hughes and McCormick (1981), Henley (1998) and others have stated, in the case of the UK, renters living in social housing have as low incentives to move as owners, since they will lose their rent control if they move. In order to take this into account, an indicator of whether renters are living in social housing is introduced in the estimation.

2.3 Estimation results

The parameter estimates of the transition intensities are shown in Table 1. Columns under the heading (i) present estimates for a specification in which country dummies capture the institutional differences, while, in specification (ii), these dummies are replaced by indices evaluating several characteristics of the housing and labour markets in these countries.

Table 1 shows the coefficient estimates and t-ratios of each transition intensity to employment, whether associated with a residential change, $\theta_1 [t \mid X(t)]$, or not, $\theta_0 [t \mid X(t)]$.

[9] Although the sample size is very small and we control for a large number of explanatory variables, the model is not saturated and is estimated parametrically. I need to take into account a large number of explanatory variables that affect both the geographical mobility and the housing tenure regime. In this way, the estimates of the housing tenure coefficients are not biased due to a problem of omitted variables. Nevertheless, I have also estimated the model considering smaller sets of explanatory variables, and the estimation results are the same.

The explanatory variables, common to both specifications, have coefficient and t-ratio values which are very similar to each other, except for the constant and the unemployment rate. For this reason, I will only comment on the estimates of specification (i), except for those estimates diverging significantly.

First, we observe that the ownership indicator is significant at the 1% level in explaining exits associated with a residential change, $\theta_1 [t \mid X(t)]$; it has a coefficient estimate of -1.805, which means that owners are more reluctant to move to other area to work. The probability of this exit still falls more for owners having outstanding loans or mortgages. Similarly, social housing also affects negatively exits to a job spell associated with a residential change, since its coefficient estimate has a value of -1.177, which is significant at the 5% level. In contrast, in exits to employment in the local area, $\theta_0 [t \mid X(t)]$, the ownership and social housing coefficient estimates are near zero, 0.111 and 0.122 respectively, and not significant at the 5% level. This seems to provide evidence for the hypothesis that owners and renters living in social housing are less willing to migrate for job reasons, and they behave in a similar way to private renters in the local labour market. Next, the statistical significance of housing tenure must be corroborated by its economic significance in the probabilities predicted by the model, which are presented in Table 2. In contrast, in the local labour market, unemployed owners having an outstanding mortgage leave unemployment with a higher probability than both types of renters and the rest of owners at the 1% significance level, maybe due to the fact that they have to repay debts.

Concerning the country dummies, most of their coefficient estimates are significant in explaining both types of exits to a job spell, which reflects institutional differences in housing and labour market across countries. The intensity of mobility across countries is not only captured by the country dummies, but it is also reflected in the coefficient of the unemployment rate. For this reason, it is necessary to compute the predicted probabilities implied by the model, since it is not possible to observe this feature directly from the estimates.

Regarding the economic variables, the unemployment rate has the expected sign in both exits to a job spell, although it is only significant at the 1% level in exits associated with a residential change. However, exits to a job occur during the period 1995-97, which does not cover a complete business cycle for each country. As seasonal indicators have been introduced in the specifications, the negative sign of the quarterly unemployment rate will not only capture the effect of the business cycle, but also the cross-section variation across countries. This idea is corroborated in the estimates of specification (ii) in which the unemployment rate does not have the expected sign.

It is expected that, when regional differences in unemployment rates and in other economic variables are large within a country, inter-regional migration due to job reasons will be observed from more depressed to richer regions. However, McCormick (1997) and Bentolila (1997) have observed that geographical mobility seems not to respond to differentials in the regional economic conditions significantly in the UK and in Spain, respectively, using aggregate regional data. I tried to introduce a variable measuring the dispersion in unemployment rates (a proxy of economic conditions) across regions, but the results were not satisfactory, mainly for the reason that residential mobility in the sample may occur between shorter distances, but there is no available data to capture this effect.

Regarding the characteristics at the previous job, individuals having worked in the service and industry sectors have a lower probability of leaving unemployment than in the other sectors, whereas the economic sector does not seem to explain mobility. The experience attained at the previous job is not a significant influence on exits associated with a residential change; however, it affects negatively exits to a job in the local area. The reason for this is that the predominant effect is the entitlement to receive unemployment benefits, which raises the reservation wage and decreases the probability of accepting a job. As the unemployment duration increases, this negative impact on exits to a job spell disappears, indicating the approach to the end of the unemployment benefit receipt. This result has also been found in Bover and Gómez (2004).

The effect of experience in the labour market is captured through the logarithm of age. Its impact on geographical mobility cannot be estimated robustly due to the small sample size. However, it has a negative impact on exits to a job spell in the local labour market, indicating that as aged family heads become unemployed, it is more difficult for them to find another job, maybe due to the fact that their knowledge becomes more obsolete, and firms prefer to hire younger people. Finally, individuals having completed the second stage of secondary education seem to be more mobile; the coefficient estimate for this exit is 0.936, significant at the 5% level. In the local labour market, individuals having completed tertiary education leave unemployment with a higher probability, although this positive influence disappears as the unemployment spell lengthens.

Concerning working time, individuals who previously worked full-time have a higher probability of exiting from unemployment in their local area. This can be a consequence of their more stable and stronger link to the labour market, whereas the activity of those working part-time may be more sporadic, and their job search intensity may be lower. Nevertheless, this effect is not statistically significant.

Moreover, male family heads have a higher probability of leaving unemployment with respect to their female counterparts, since the gender indicator has a coefficient estimate of 0.326 and it is significant at the 1% level. This may capture several effects, among them the following: women may have a higher reservation wage than men (but not very different since both are family heads) and women may be discriminated in the sense of receiving job offers at a lower rate or with worse wage conditions than men. Concerning household composition, the indicator of being married or living with a partner is not significant in explaining any transition, although the fact that the spouse or partner is employed affects exits associated with a residential change negatively, as expected. Children variables have the expected sign, but they are not significant at the 5% level.

The receipt of unemployment benefits has the expected effects on both types of exits, although it is only significant in explaining exits to a job spell in local areas.

The negative effect of this entitlement disappears as the unemployment spell increases and the end of the receipt becomes closer.

Regarding duration dependence, it cannot be determined robustly in transitions to employment associated with a residential change given the small number of observations to this type of exit. On the contrary, dependence is negative for transitions in the local labour market: the longer the unemployment spell is, the more difficult the exit becomes, due to the obsolescence of the worker's knowledge, stigma effects or the unemployed worker's discouragement.

Finally, in specification (ii), the country indicators have been replaced by some indices measuring characteristics of housing and labour markets. Concerning the degree of strictness in employment protection legislation, we see that the higher the strictness is in a country, the lower the probability of leaving unemployment in the local area. This feature is stronger in countries with high protection, as Southern Europe, namely Spain and Italy. However, this result is not found in exits associated with a residential change. As mentioned above, the unemployment rate may be capturing cross-country variation rather than the business cycle, since its coefficient's sign is reverted, and it is significant in exits to a job in the local labour market.

The logarithm of the percentage that transaction taxes represent of the house price plays the role of a proxy of the transaction costs that individuals incur when they buy a house. This measure does not behave as expected, since it is not significant, although it has a negative coefficient estimate in exits associated with a residential change. I also interacted this variable with the ownership indicator in order to capture a different behaviour by owners and renters, but it was insignificant. In addition, I tried another measure evaluating the transaction costs contained in Maclennan *et al.* (1998), but the results were identical. Finally, the unemployment benefit replacement rate is expected to reflect the effect of unemployment protection legislation on the unemployed workers' geographical mobility. Its coefficient estimate is positive, but it is insignificant due to the small sample size.

As the index of transaction costs does not behave as expected and the sign and size of the coefficient estimates do not vary much except for the unemployment rate, the specification (i) will be used in the remaining sections.

Predicted probabilities In order to evaluate the size of the effects of explanatory variables on both transition intensities, Table 2 contains the probabilities predicted by the estimates of model (i) in Table 1. Column 1 shows the predicted probability of transition to a job spell associated with a residential change (θ_1) and column 2 to a job in the local labour market (θ_0). In each row, the sum of the percentages shown in both columns gives us the exit rate to employment (equation (2.2)) predicted for an individual with the relevant characteristic. The first row shows these probabilities for the reference person, and the following rows for an individual with the same characteristics as the reference person except for the characteristic indicated. Thus, the economic impact of each explanatory variable is analysed by comparing the probabilities shown in the corresponding row with those of the reference person.

The reference person is a male family head living in a private rented house in Spain; he is single, aged 30 years old, does not have any children, and his level of education is lower than the second level of secondary; he has been unemployed for 4 months, has previously worked full-time in industry for 6 months, and does not receive any unemployment benefits. The unemployment rate is evaluated at its average level in 1995, and the seasonal indicators are set at the fourth quarter. For the reference person, the probability of transition to employment associated with a residential change is 0.64%, whereas that to a job in the local labour market is 11.80%. Thus, his predicted exit rate is equal to 12.44%, the sum of both probabilities.

Two features stand out in Table 2. First, the difference in size of the probabilities displayed in both columns makes obvious that, in all European countries analysed, the unemployed prefer to search and to accept a job in their local area rather than moving to another place to work. Geographical mobility is rarely observed in all countries, although this fact is more stressed in the UK, Spain and Germany, in which the predicted probabilities for the reference person are 0.53%, 0.64% and 0.89%, respectively. The most mobile countries are France and Italy, whose predicted probabilities for the reference person are 1.77% and 0.94%, respectively. The predicted probabilities for Germany and the UK are extremely low, mainly in column 2; this is due to their small sample size of unemployed family heads (they account for 6.98% and 7.21%, respectively, of the total unemployment spells, see Table A.1) and due to the high proportion of censored spells, 70.67% and 59.35% of German and British samples, respectively, as we can observe in Table A.3 in the Appendix.

Second, when housing tenure is taken into account, the probability of geographical mobility is considerably smaller for an owner than for a private tenant, since this probability falls from 0.64% for a private renter to 0.10%, in the case of Spain. Social housing also reduces the incentives to geographical mobility, since this probability falls to 0.20%. However, the exit to employment in the local area is not affected by housing tenure, since the predicted probabilities do not vary significantly, except for the homeowners having outstanding mortgages; that is, owners and renters behave in a similar way in this market. Therefore, there is evidence for the hypothesis explained at the beginning of this section.

3 Addressing potential unobserved heterogeneity in labour mobility and ownership status

The previous section considered ownership status as a predetermined variable in the two possible transition intensities to a job spell. However, the analysis carried out about the effect of housing tenure on the unemployed workers' mobility decision may be contaminated by the presence of unobserved heterogeneity in these transitions, possibly correlated with housing tenure status. This section studies if those results are biased by the disregard of the unobserved heterogeneity. For this purpose, Subsection 3.1 first investigates the main determinants of housing tenure status in the sample of unemployed family heads used in Section 2, using differences in the institutional characteristics of the housing markets across European countries. Subsection 3.2 carries out the same analysis as in the previous section allowing for the presence of unobserved heterogeneity correlated with the ownership status equation considered below.

3.1 Determinants of the housing tenure status

Let h_i be the indicator of the housing tenure status of the individual i, which takes the values of 0 if he is a private renter, of 1 if he is an owner and of 2 if he lives in a social rented house.

The decision of housing tenure is modelled in a similar way as in King (1980) and Duce (1995). The preference of individual i between ownership and private rental is denoted by the indicator h_i^* that takes the value of 1 if he chooses ownership and the value of 0 if he chooses private rental. This choice is assumed to follow a logit specification as follows:

$$\Pr\left(h_i^* = 1 \mid Z_i\right) = \frac{\exp(Z_i'\delta)}{1 + \exp(Z_i'\delta)} \tag{3.1}$$

Each component of the parameter vector, δ, reflects how an increase in the corresponding explanatory variable affects the utility of being an owner in comparison with that of being a private renter.

The value of this indicator is unobserved for individuals living in a social rented house. These individuals may be unable to access either an owned-occupied house due to loan rationing reasons or a private rented house. However, the access to social housing is also restricted, since its low housing cost may cause a large demand. Thus, the admission to social housing is often based on some criteria, such as low income levels, number of children and family head's age among others. Let q_i be the indicator of this admission, that is, it takes the value of 1 if individual i obtains a subsidised house, and the value of 0 otherwise. The admission to social housing is also assumed to follow a logit specification as follows:

$$\Pr\left(q_i = 1 \mid Z_{2i}\right) = \frac{\exp(Z_{2i}'\gamma)}{1 + \exp(Z_{2i}'\gamma)} \tag{3.2}$$

The vector of explanatory variables, Z_{2i}, is a subset of household characteristics vector Z_i and it contains the criteria on which the admission is based. Particularly, I consider income, the number of children aged 18 years old or less in the household, the logarithm of the family head's age and country dummies to allow for different housing policies concerning social housing. I have excluded the family head's gender due to its non-significance in the estimates.

Thus, the probability of observing a private renter or an owner is as follows:

$$\Pr\left(h_i \mid Z_i\right) = \Pr\left(h_i^* \mid Z_i\right)\left[1 - \Pr\left(q_i = 1 \mid Z_{2i}\right)\right], \quad h_i = h_i^*,\ h_i^* = 0, 1 \tag{3.3}$$

That is, we observe both housing tenures when individual i is not admitted to obtain subsidised rental. Finally, the probability of an individual living in a social rented house is equal to the probability of being admitted:

$$\Pr\left(h_i = 2 \mid Z_i\right) = \Pr\left(q_i = 1 \mid Z_{2i}\right) \tag{3.4}$$

The main determinants of the housing tenure status can be studied using equations (3.1)-(3.4). The explanatory variables included in the vector Z_i are: the logarithm of the household's total income in the previous period, normalised by the purchasing power parity (hereafter, PPP), the indicator of living with a partner, the number of children aged 18 years old or less and its square, and the gender, the logarithm of age and the level of education of the family head.

Before showing the estimates of the tenure status equation, I will comment on the densities of the family head's age and the logarithm of household's income, for each type of housing tenure (ownership or rental) and country, estimated using the sample of unemployment spells. These estimated densities consist of Epanechnikov kernels, in which the bandwidth is chosen optimally; they are displayed in Figures 1 and 2. In Figure 1, the age profile shows how tenure status varies along the life-cycle. Among the younger workers, the probability of living in a rented house is the highest. This probability is increasing until the age band of 30 to 33 years old, except for Italy, where this probability continues increasing until the age of 39 years old; then, the probability of renting begins to decrease. The largest proportion of owners is concentrated in the age range from 45 to 55 years old, varying across countries, when individuals have accumulated enough wealth to invest in an owned house. At older ages, the proportion of owners starts to decrease. France is the country in which the proportion of owners overcomes that of renters at the youngest age, 35 years old, followed by the United Kingdom and Spain at the ages of 37 and 40 years old, respectively. Germany and Italy are the countries in which this transition happens much later, at the age of 46 and 47 years old, respectively.

Regarding the estimated density of income, Figure 2 shows that ownership is the predominant tenure only at high levels of income. This feature is stronger in Germany, France and the UK, where the density of the logarithm of income for the owners is shifted towards the right, far from that of the renters. On the contrary, in Spain and Italy, housing tenure status does not seem to depend on income so much, since both densities are more similar to each other.

Table 3 shows the estimates of the main determinants of tenure status equation, described in equations (3.3) and (3.4), for the pooling of subsamples of unemployed family heads' spells in all countries used to estimate the duration model in Section 2. Again, the indicator for Spain is omitted in order to include a constant in the estimates. The total number of spells, $2,150$, is used to estimate the equation of admission to subsidised rental; however, the parameters of the ownership status equation are identified only using $1,883$ spells corresponding with households that do not live in a social rented house. Specification (i) shows the existence of differences in housing policies supporting home ownership, since the country indicators except for the UK are significant in explaining home ownership status at the 1% level. In comparison with Germany and France, Spain, the United Kingdom and Italy are the countries in which home ownership is made more attractive than private rental by the specific characteristics of their housing market.

The logarithm of age is also significant at the 1% level in explaining the ownership status; its coefficient estimate is 3.102, which indicates that the older the family head is, the more likely it is that the household lives in an owned house. As expected, the income variable has a positive coefficient estimate, significant at the 1% level; the higher the level of income, the higher the probability of living in an owned house is. As we saw in Figure 2, the country indicators may capture, among other things, the movement of the density function of being an owner to the right in the case of Germany and France for a given level of income; however, the shape of the probability of living in an owned house by the level of income seems to be the same in all countries, and any interactions of income with country dummies are not included in the estimates.

Regarding the household's composition, individuals who are married or living with a partner are more likely to live in an owned house. Household size, measured by the number of children aged 18 years old or less in the household and by its square, affects positively the probability of living in an owned house, but at a lower rate as household size increases. This effect is found if we consider social housing as another housing tenure different from private rental. Otherwise, the effect of household size on the probability of being an owner is negative, indicating less ability to accumulate wealth as size increases.

The level of education could also play the role of a proxy for wealth; thus, we would expect that the higher this level is, the larger are the accumulated wealth and the higher the probability of living in an owned house. Nevertheless, education does not capture this effect; the highest level of education has an insignificant and wrong-signed coefficient estimate, and the second stage of secondary education has a positive coefficient value, but it is not significant at the 10% level. As another proxy for wealth, I included an indicator of whether the household owned a second house, but I finally removed it from the estimates because of its insignificance.

In specification (ii), country dummies have been replaced by other variables reflecting different characteristics of policies supporting home ownership across countries. Three variables have been constructed based on the information provided in Maclennan *et al.* (1998). First, I use the logarithm of the stamp duty incurred in the purchase of a house as a percentage of the house price; and also the logarithm of the ratio of social to private rented accommodation prevailing in each country. The last variable is an indicator taking the value of 1 if interest tax reliefs are allowed in income taxes due to the repayment of the outstanding mortgage. The first two variables are expected to affect negatively the choice of living in an owned house; on the contrary, the interest tax relief should affect it positively. In the estimates, all three variables are significant at the 1% level, with the expected signs.

Concerning the equation of admission to social housing, all country dummies have positive coefficient estimates, significant at the 1% level, indicating that Spain is the country in which subsidised rental is least promoted by the authorities. Meanwhile, the access to a social house can be more easily obtained in the UK, France and Germany. The rest of explanatory variables are significant with the expected signs, lower income levels, younger family head and bigger household size increase the probability of obtaining a subsidised rented house.

3.2 Unobserved heterogeneity in transitions to employment and in ownership status

In this subsection, I present estimates of the housing tenure status equation jointly with the unemployment duration model with multiple exits, in order to control for unobserved heterogeneity in transitions to a job spell correlated with ownership status.

Ownership status may be positively correlated with unobserved human capital and ability, since they increase the individuals' wealth accumulation. In addition, human capital contributes to raise the probability of transitions to employment, not only in the local labour market, but also in other geographical areas. Thus, the coefficient of ownership status in exits associated with a residential change may be biased upwards spuriously if owners are more skilled than renters on average, due to unobserved human capital variables that increase their probability of geographical mobility.

Moreover, the composition of the sample of unemployment spells according to housing tenure may be altered by a higher proportion of renters that, on average, have a less favourable unobserved ability to abandon unemployment than owners. This will cause it to be less likely to observe geographical mobility among renters, since this decision is greatly influenced by human capital and skill. Thus, the estimate of the owner status coefficient in exits associated with a residential change will be also increased spuriously for this motive if the presence of unobserved heterogeneity is not allowed for.

In addition, unobserved heterogeneity may also capture other factors making home ownership endogenous, such as the idea of rental as a temporary housing tenure regime for those individuals who expect to move in the near future.

Following Heckman and Singer (1984), the unobserved heterogeneity is specified as a discrete variable with a finite support, in this case, of two mass points. Given that the sample is formed by I individuals having multiple unemployment spells, the log-likelihood function has been constructed as in Ham and LaLonde (1996) and Meghir and Whitehouse (1997).[10] As I have repeated spell durations for a non-negligible

[10]Ham and LaLonde (1996) study the effects of a training program on subsequent employment and unemployment spells, and Meghir and Whitehouse (1997) analyse the labour history of individuals

proportion of individuals (27.2%) accounting for 50.6% of spells, allowing for unobserved heterogeneity becomes more advisable.

The unobservables are assumed to be independent of the explanatory variables, X_i and Z_i, other than ownership status, h_i^*. The admission to social housing is assumed to be based only on observed characteristics; so, unobserved heterogeneity will not take part in equation (3.4). For simplicity of exposition, I will redefine the vector of characteristics relevant in the duration model as $(X_i,\ h_i)$ in order to consider separately ownership status from the rest. Let η_i be the permanent unobserved effect on both transitions to employment and the ownership status equation for individual i. The unobserved effect is assumed to take a value on the support $\{m_1,\ m_2\}$ with probabilities of p_1 and p_2, respectively. Imposing that the expectation of the unobserved heterogeneity is null, $E[\eta_i] = 0$, and that the probabilities, p_1 and p_2, add up to one, the estimation of the unobservables is reduced to only a mass point, say $m_1 = m$, and its associated probability, $p_1 = p$, $0 < p < 1$.[11]

The transition intensity to employment with the alternative k, $k = 0, 1$, and the ownership status equation conditional on the unobserved heterogeneity and the observables are specified, respectively, as follows:

$$\theta_k\left[t \mid X_i(t),\ h_i,\ \eta_i\right] = \frac{\exp(X_i(t)'\beta_k + h_i\beta_{hk} + \alpha_k\eta_i)}{1 + \sum_{j=0}^{1}\exp(X_i(t)'\beta_j + h_i\beta_{hj} + \alpha_j\eta_i)}, \quad k = 0, 1 \tag{3.5}$$

$$\Pr\left(h_i^* = 1 \mid Z_i,\ \eta_i\right) = \Lambda\left(Z_i'\delta + \alpha\eta_i\right) = \frac{\exp\left(Z_i'\delta + \alpha\eta_i\right)}{1 + \exp\left(Z_i'\delta + \alpha\eta_i\right)} \tag{3.6}$$

$$\Pr\left(h_i \mid Z_i,\ \eta_i\right) = \Pr\left(h_i^* \mid Z_i,\ \eta_i\right)\left[1 - \Pr\left(q_i = 1 \mid Z_{2i}\right)\right], \quad h_i = h_i^*,\ h_i^* = 0,\ 1$$

$$\Pr\left(h_i = 2 \mid Z_i,\ \eta_i\right) = \Pr\left(q_i = 1 \mid Z_{2i}\right)$$

near the retirement age, facing a problem of initial conditions caused by the dependence of subsequent employment and non-employment spells. They solve this problem by assuming a different distribution for the first spell observed for each individual. Although my sample consists of multiple spells, this problem does not appear because I do not study entire labour histories, but only unemployment spells.

[11] This implies that $p_2 = 1 - p$ and $m_2 = -\frac{pm}{(1-p)}$.

Unobserved heterogeneity in this model follows a one-factor structure. Thus, the permanent individual effects in equations (3.5) and (3.6) are perfectly correlated with one another, although they can be inversely correlated and they can have a different impact in each equation according to the sign and size of the modifiers α_0, α_1, and α. As a result, the parameters related to the unobservables that have to be estimated are m, p, α_0 and α, if α_1 is normalised to 1.

The unobserved individual effect η is invariant across all unemployment spells and housing tenure regimes observed for each individual. Thus, the unemployment spells of each individual cannot be treated as independent observations in contrast with the homogeneous case of the duration model in Section 2. Under the absence of unobserved heterogeneity, the duration model and housing tenure status can be estimated separately by maximum likelihood. However, when allowing for unobserved heterogeneity, the joint log-likelihood cannot be split into the sum of two log-likelihoods defined separately.

The sample is regrouped in $i = 1,\ 2, \ldots, I$ individuals who have been unemployed for $n_i = n_1,\ n_2, \ldots,\ n_I$ times, respectively; that is, the individual i provides n_i unemployment spells to the sample, and $\sum_{i=1}^{I} n_i = N$, the sample size of unemployment spells in Section 2. Let $(T_{i1},\ T_{i2}, \ldots T_{in_i})$ be the sequence of the duration of each of the n_i unemployment spells in which the individual i has stayed, with the following realisation of spell durations, $(\tau_{i1} + r_{i1},\ t_{i2}, \ldots,\ t_{in_i})$, distinguishing the duration of the spell prior to the baseline, τ_{i1}, only for the first unemployment spell, $t_{i1} = \tau_{i1} + r_{i1}$, since the remaining spells are fresh for the individual i. In the same way, $(c_{i1},\ c_{i2}, \ldots, c_{in_i})$ denotes the sequence of indicators of lack of censoring. The explanatory variables relevant to each unemployment spell are regrouped for each individual i as follows: $X_i = [X_i(T_{i1}),\ X_i(T_{i2}), \ldots,\ X_i(T_{in_i})]$, $Z_i = (Z_{i1},\ Z_{i2}, \ldots,\ Z_{in_i})$, $D_i = (D_{i1},\ D_{i2}, \ldots,\ D_{in_i})'$ and $h_i = (h_{i1},\ h_{i2}, \ldots,\ h_{in_i})'$.

Then, conditional on the unobserved heterogeneity and observables, individual i contributes to the likelihood of the probability of observing that sequence of unemployment spell durations, exits to employment and housing tenure status in each of the spells.

Using the definition of transition intensities and equations (3.5) and (3.6), this contribution can be factorised as follows :

$$\Pr\left(\{t_{ij}, D_{ij}, c_{ij}, h_{ij}\}_{j=1}^{n_i} \mid T_{i1} > \tau_{i1}, X_i, Z_i, \eta_i\right) = \prod_{j=2}^{n_i} \Pr\left(t_{ij}, D_{ij}, c_{ij} \mid X_i(t_{ij}), h_{ij}, \eta_i\right) \cdot$$

$$\Pr\left(\tau_{i1} + r_{i1}, D_{i1}, c_{i1} \mid T_{i1} > \tau_{i1}, X_i(t_{i1}), h_{i1}, \eta_i\right) \prod_{j=1}^{n_i} \Pr\left(h_{ij} \mid Z_{ij}, \eta_i\right) =$$

$$= \prod_{j=1}^{n_i} \Big(\left[\theta_1\left(t_{ij} \mid h_{ij}, X_i(t_{ij}), \eta_i\right)\right]^{D_{ij} c_{ij}} \left[\theta_0\left(t_{ij} \mid h_{ij}, X_i(t_{ij}), \eta_i\right)\right]^{(1-D_{ij}) c_{ij}} \cdot$$

$$\left[1 - \theta\left(t_{ij} \mid h_{ij}, X_i(t_{ij}), \eta_i\right)\right]^{(1-c_{ij})} \cdot \Pr\left(h_{ij} \mid Z_{ij}, \eta_i\right) \Big) \prod_{s=\tau_{i1}+1}^{t_{i1}-1} \left(1 - \theta\left(s \mid h_{i1}, X_i(s), \eta_i\right)\right) \cdot$$

$$\prod_{j=2}^{n_i} \left(\prod_{s=1}^{t_{ij}-1} \left(1 - \theta\left(s \mid h_{ij}, X_i(s), \eta_i\right)\right) \right) \tag{3.7}$$

Finally, to obtain the contribution to the likelihood, $\mathcal{L}_i$, the unobserved heterogeneity has to be integrated out, and the log-likelihood function has the following form:

$$L = \sum_{i=1}^{I} \log \mathcal{L}_i =$$

$$= \sum_{i=1}^{I} \log \left\{ \sum_{l=1}^{2} \left[\Pr\left(\{t_{ij}, D_{ij}, c_{ij}, h_{ij}\}_{j=1}^{n_i} \mid T_{i1} > \tau_{i1}, X_i, Z_i, \eta_i = m_l\right) \; p_l \right] \right\} \tag{3.8}$$

The results are found in Table 4. Most of the significant variables do not change much in their coefficient size or significance when we control for unobserved heterogeneity. However, the estimated coefficient of the ownership status in exits associated with a residential change becomes more negative, from -1.805 to -2.066, giving evidence favourable to the hypothesis that unobserved human capital variables could bias upwards this coefficient. Moreover, the same variable in exits to a job in the local area becomes significant with a coefficient estimate of 0.600, indicating that this specification of heterogeneity does not capture all the unobservables, since owners could enjoy higher arrival rates of a job offer or better-paid job offers due to their human capital. These findings are corroborated by analysing the predicted probabilities of both models. About the indicator of living in social rented housing, the unconditional predicted probabilities of the model with unobserved heterogeneity are very similar to those of the model not controlling for it, although its coefficient estimates now become significant in both exits to employment.

Concerning age, after controlling for unobserved heterogeneity, older family heads have more difficulty in leaving unemployment. In the same way, ownership status becomes much more dependent of the family head's age and total income, since their estimated elasticities become larger. Moreover, when unobserved heterogeneity is allowed for, ownership status responds a bit less intensively to housing policies, such as the existence of interest tax relief in income taxes, and the ratio of social to private rented accommodation prevailing in each country. However, it becomes more sensitive to the decrease in the stamp duties to pay for the purchase of a house. Regarding education, after controlling for unobserved heterogeneity, individuals having completed the third level of education seem to prefer to live in a private rented house rather than in an owned house. Finally, the value of the log-likelihood function does not increase so much when the unobserved heterogeneity specification has three mass points.

4 Effects of mobility and unemployment duration on post-unemployment wages

This section is divided in two parts: subsection 4.1 describes the empirical model, and subsection 4.2 presents the estimation results.

4.1 Estimation method of the wage equation

This section focuses on the main determinants of the wage level obtained by individuals just after leaving unemployment, as in other studies like Addison and Portugal (1989). In the first place, I want to study whether this level depends on the length of the unemployment spell, so I introduce the logarithm of duration, $\log T_i$, as other determinant in the wage equation. This explanatory variable will capture two effects: first, the obsolescence of the unemployed worker's knowledge, and second, the firms' reluctance to hire individuals having been unemployed for a long time, since employers may think that long unemployment spells are a bad signal about their productivity (the stigma effect). Thus, this variable is expected to have a negative coefficient estimate.

In the second place, as commented in Section 1, I am also interested in finding whether those individuals who accept a job associated with a residential change obtain a higher wage on average than those who exit from unemployment in their local area. If so, policy-makers should promote housing policies encouraging geographical mobility among the unemployed workers. Economic theory does not provide a clear prediction about the gain in wages when geographical mobility occurs. In a partial job search model, when regions are heterogeneous in their offer arrival rates and in their wage distribution functions, individuals living in depressed regions are more likely to receive higher wage offers if they move to wealthier regions. In this case, geographical mobility has a positive impact on the wage level that individuals achieve just after leaving unemployment.

However, in a multi-period model of job search, individuals take into account that their current decisions influence their expected future utility, and they will accept a wage offer in other region only if this is greater than the reservation wage. As the reservation wage is determined by expectations, the wages accepted by migrants in the destination region might not be higher than those they would have obtained in their origin region. That is, if the destination region has a lower unemployment rate, a higher arrival rate of job offers to the employed, and a lower probability of being dismissed than the origin region does, their reservation wage of moving will be comparatively lower than that for accepting a job in their local area. This will happen if moving costs are low enough. Individuals will have a higher reservation wage for accepting a job in their depressed region, since the possibility of changing to another better-paid job will be lower. Therefore, in this sense, the effect of geographical mobility on the wage level obtained just after leaving unemployment may be negative in comparison with the wage level that they could have obtained in their local depressed region.

In order to capture the impact of geographical mobility on the wage level, I include the indicator, D_i, defined in Section 2, which describes the alternative chosen by the family head, whether a job associated with a residential change ($D_i = 1$) or a job in the local labour market ($D_i = 0$).

A problem arises when the wage equation is estimated using only the subsample of those unemployed individuals who exit from unemployment. In this way, this estimate does not correspond with the mean of the potential wage distribution. This distribution will depend on the regional characteristics as well as job-related ones, but it is not affected by the individuals' decision rules. On the contrary, the distribution of accepted wages is influenced by these rules, since all accepted wages are higher than the individuals' reservation wages. This means that a problem of self-selection bias arises, since individuals decide to remain unemployed and not to accept wage offers, if these are not greater than their reservation wage.

This bias can be avoided by computing the equivalent to the inverse Mill's ratio in the model of labour mobility described in Section 2. The inverse Mill's ratio or Heckman's lambda is typical of models in which the participation equation follows a probit specification, and in which the disturbances of the wage and participation equations are correlated with each other and distributed jointly as a Normal multivariate.[12]

In this paper, two participation equations exist: whether to accept a job associated with a residential change ($D_i = 1$), and whether to accept a job in the local labour market ($D_i = 0$). These two equations are given by the transition intensities defined in equations (2.1) and (2.3); so, they follow a multinomial logit specification.

Let $\overline{w}_{ki}(t),\ k = 0,\,1$, be the result of the comparison of the wage received with the corresponding reservation wage. These comparisons are not observed by the econometrician, and they are assumed to follow the specification below, in accordance with the model described in Section 2:

$$\overline{w}_{ki}(t) = X_i(t)'\beta_k + \varepsilon_{ki}, \quad k = 0,\,1;$$

$$F\left(\varepsilon_{0i},\, \varepsilon_{1i} \mid X_i(t)\right) = \frac{1}{1 + \exp(-\varepsilon_{0i}) + \exp(-\varepsilon_{1i})}; \tag{4.1}$$

where β_k is the parameter vector displayed in equation (2.3).

The vector of characteristics, $X_i(t)$, is the same as in Section 2. In order to simplify the notation, I will leave the index t out of the characteristics vector X_i, although it includes time-varying variables.

Individual i will leave unemployment with the alternative k, $D_i = k,\ k = 0,\,1$, if this provides him with the maximum level of the wage over the reservation wage of each alternative, $\max\{\overline{w}_{1i}, \overline{w}_{0i}\} = \overline{w}_{ki}$, and if this is greater than the reservation wage, $\overline{w}_{ki} \geq 0$:

$$D_i = k \quad \text{if } \max\{\overline{w}_{1i}, \overline{w}_{0i}\} = \overline{w}_{ki} \text{ and } \overline{w}_{ki} \geq 0,\ k = 0,\,1 \tag{4.2}$$

[12]Mroz (1987) estimated a model of female labour supply in which distributional assumptions different from those involved in the Heckman's lambda were also considered in the participation equation, in particular, a logit participation model.

Let w_i be the logarithm of the wage level that individual i obtains just after leaving unemployment, then this wage is specified as follows:

$$\begin{aligned} w_i^* &= \alpha D_i + Z_i'\delta + \pi \log T_i + u_i, \quad i = 1, 2, \dots N \\ w_i &= \begin{cases} w_i^* & \text{if } c_i = 1 \\ 0 & \text{if } c_i = 0 \end{cases} \end{aligned} \tag{4.3}$$

That is, the wage is only observed if the individual leaves unemployment with one of the two possible exits to employment. The variables in Z_i are related to new job characteristics, and they consist of some characteristics such as age or education, common to those in X_i. The distributional assumptions on the disturbance, u_i, are as follows:

$$u_i = \rho_0 \varepsilon_{0i} + \rho_1 \varepsilon_{1i} + v_i, \tag{4.4}$$

$$E\left[v_i \mid D_i,\, Z_i,\, T_i,\, X_i\right] = 0,$$

$$cov(v_i,\, \varepsilon_{ki} \mid D_i,\, Z_i,\, T_i,\, X_i) = 0,\ k = 0, 1$$

The disturbance can be broken down as the sum of two terms: a first term, v_i, is uncorrelated with the explanatory variables of the wage equation and with those of transition intensities, and a second term, $\rho_0\varepsilon_{0i}+\rho_1\varepsilon_{1i}$, is correlated with the disturbances, ε_{0i} and ε_{1i}, of the equations for the comparison of the wage offers with the reservation wage.

In order to obtain an expression of the self-selection bias arising from equation (4.3), it is useful to define indicators A_i, A_{1i}, and A_{0i} for individual i as follows:

$$\begin{aligned} A_{0i} &= 1\left(\max\left\{\overline{w}_{1i}, \overline{w}_{0i}\right\} = \overline{w}_{0i} \text{ and } \overline{w}_{0i} \geq 0\right) \\ A_{1i} &= 1\left(\max\left\{\overline{w}_{1i}, \overline{w}_{0i}\right\} = \overline{w}_{1i} \text{ and } \overline{w}_{1i} \geq 0\right) \\ A_i &= A_{0i} + A_{1i} \end{aligned} \tag{4.5}$$

Note that A_i can only take the values 0 or 1, since A_{0i} and A_{1i} cannot take the value of 1 at the same time. The dummy variable, A_i, indicates that the individual exits to a job spell: when he finds a job in his local area, $A_{i0} = 1$, or when he moves, $A_{i1} = 1$. These indicators have been defined in order to make the notation and the exposition of the wage equation model clearer.

If we estimate equation (4.3) by ordinary least squares (OLS) robust to heteroskedasticity, a problem of self-selection bias arises because we do not condition only on the subsample of individuals leaving unemployment, that is, on those individuals whose wage is observed, $A_i = 1$.

To study the mean of potential wages, we can carry out a two-step procedure similar to the Heckman's two-step estimator,[13] but applied to the duration model described in Sections 2 and 4. First, we have to know the functional form of this bias; for that purpose, we condition equation (4.3) on all the explanatory variables and on $A_i = 1$ as follows:

$$E\left[w_i \mid D_i,\, Z_i,\, T_i,\, X_i,\, A_i = 1\right] = \alpha D_i + Z_i'\delta + \pi \log T_i + E\left[u_i \mid D_i,\, Z_i,\, T_i,\, X_i,\, A_i = 1\right] \tag{4.6}$$

By developing the expectation of the disturbance in equation (4.6), we arrive at:

$$\begin{gathered} E\left[u_i \mid D_i,\, Z_i,\, X_i,\, T_i,\, A_i = 1\right] = \\ = \sum_{k=0}^{1} E\left[u_i \mid D_i,\, Z_i,\, X_i,\, T_i,\, A_{ki} = 1\right] \Pr\left(A_{ki} = 1 \mid A_i = 1,\, D_i,\, Z_i,\, T_i,\, X_i\right) \end{gathered} \tag{4.7}$$

By using the definitions in equation (4.5) and the distribution function described in equation (4.1), we obtain that the probability terms in equation (4.7) have the following form:

$$\begin{gathered} \Pr\left(A_{ki} = 1 \mid A_i = 1,\, D_i,\, Z_i,\, T_i,\, X_i\right) = \\ = \Pr(A_{ki} = 1 \mid A_i = 1,\, T_i,\, X_i) = \frac{\exp(X_i'\beta_k)}{\exp(X_i'\beta_0) + \exp(X_i'\beta_1)} \end{gathered} \tag{4.8}$$

[13] See Heckman (1976) and Chap. 10 in Amemiya (1985).

The mean of the disturbance u_i, conditional on all the explanatory variables and on the indicator A_{ki} taking the value of 1, has the following structure using equations (4.1), (4.4) and (4.5):

$$E\left[u_i \mid D_i,\, Z_i,\, T_i,\, X_i,\, A_{ki}=1\right] = \tag{4.9}$$
$$= \sum_{j=0}^{1} \rho_j E\left[\varepsilon_{ji} \mid X_i,\, T_i,\, \varepsilon_{ki} - \varepsilon_{(1-k)i} \geq X_i'\left(\beta_{(1-k)} - \beta_k\right),\, \varepsilon_{ki} \geq -X_i'\beta_k\right]$$

By developing both terms included in equation (4.9) and replacing equations (4.8) and (4.9) into equation (4.7), we obtain the following expression of the self-selection bias:

$$E\left[u_i \mid D_i,\, Z_i,\, T_i,\, X_i,\, A_i=1\right] =$$
$$= \rho_0 \frac{\left(1+\sum_{k=0}^{1}\exp\left(X_i'\beta_k\right)\right)}{\left(1+\exp\left(X_i'\beta_1\right)\right)\left[\sum_{k=0}^{1}\exp\left(X_i'\beta_k\right)\right]}\left\{\ln\left[1+\exp\left(X_i'\beta_0\right)+\exp\left(X_i'\beta_1\right)\right] - \right.$$
$$\left. - \frac{X_i'\beta_0\exp\left(X_i'\beta_0\right)}{1+\sum_{k=0}^{1}\exp\left(X_i'\beta_k\right)}\right\} +$$
$$+\rho_1 \frac{\left(1+\sum_{k=0}^{1}\exp\left(X_i'\beta_k\right)\right)}{\left(1+\exp\left(X_i'\beta_0\right)\right)\left[\sum_{k=0}^{1}\exp\left(X_i'\beta_k\right)\right]}\left\{\ln\left[1+\exp\left(X_i'\beta_0\right)+\exp\left(X_i'\beta_1\right)\right] - \right.$$
$$\left. - \frac{X_i'\beta_1\exp\left(X_i'\beta_1\right)}{1+\sum_{k=0}^{1}\exp\left(X_i'\beta_k\right)}\right\} \tag{4.10}$$

Note that this bias expression takes into account that the logarithm of the unemployment duration is endogeneous to reemployment wages. In order to control for self-selection using an estimator similar to Heckman's two-step estimator, the first stage will consist of estimating the duration model described in Section 2. In a second stage, I will compute estimates of those terms whose coefficients are ρ_0 and ρ_1 in the bias expression shown in equation (4.10), replacing the unknown parameters β_k by their estimates in the first stage; these terms are introduced in equation (4.6) and then this is estimated by OLS robust to heteroskedasticity using only the subsample of individuals entering an employment spell.

Although this two-step estimator is consistent, this procedure does not provide efficient estimates due to the fact that it does not take into account that predicted values have been introduced in the estimates. In order to obtain more efficient estimates, I estimate the unemployment duration model and the wage equation jointly using the generalised method of moments (GMM).

For this purpose, I use two sets of orthogonality conditions: the former is formed by the first-order conditions derived from the maximisation of the log-likelihood function defined for the duration model, shown in equation (4.11):

$$E\left[\frac{\partial L_i(\beta)}{\partial \beta_k}\right] = 0, \quad k = 1, 0 \tag{4.11}$$

The latter set of orthogonality conditions arises from the first moment of the potential wage distribution conditional on the subsample of unemployed individuals entering a job spell, that is, conditional on $A_i = 1$ as in equation (4.6). This equation can be rewritten as follows:

$$w_i = \alpha D_i + Z_i'\delta + \pi \log T_i + \rho_1 \mu_{1i}(X_i;\, \beta) + \rho_0 \mu_{0i}(X_i;\, \beta) + \xi_i$$

The expression $\rho_1 \mu_{1i}(X_i;\, \beta) + \rho_0 \mu_{0i}(X_i;\, \beta)$ denotes the self-selection bias displayed in equation (4.10), which depends on the parameter vector, β, coming from the transition intensities to employment. The noise ξ_i is an error term coming from the regression of the logarithm of wage, w_i, on the explanatory variables D_i, Z_i, $\log T_i$, $\mu_{1i}(X_i;\, \beta)$ and $\mu_{0i}(X_i;\, \beta)$ from the subsample of individuals entering a job spell. Thus, this prediction error is mean independent of the regressors:

$$E\left[\xi_i \mid D_i, Z_i, \log T_i, \mu_{1i}(X_i;\, \beta), \mu_{0i}(X_i;\, \beta)\right] = 0$$

This kind of independence implies the second set of orthogonality conditions shown in equation (4.12):

$$E\left[\begin{pmatrix} D_i \\ Z_i \\ (\log T_i) \\ \mu_{1i}(X_i;\,\beta) \\ \mu_{0i}(X_i;\,\beta) \end{pmatrix} \xi_i\right] = 0 \tag{4.12}$$

Let X_i and Z_i be matrices of dimension $p \times 1$ and $r \times 1$, respectively. The number of parameters to be estimated from both the duration model and the wage equation is $2p+r+4$, corresponding to $\beta_1,\ \beta_0,\ \delta,\ \alpha,\ \pi,\ \rho_0$ and ρ_1. As the number of parameters is equal to the number of orthogonality conditions, the parameters are exactly identified. Therefore, the choice of the weighting matrix, A_N, becomes irrelevant in the following criterion function:

$$\min_{\{\beta,\delta,\alpha,\pi,\rho_1,\rho_0\}} b_N\left(\beta,\delta,\alpha,\pi,\rho_1,\rho_0\right)' A_N b_N\left(\beta,\delta,\alpha,\pi,\rho_1,\rho_0\right)$$

$$b_N\left(\beta,\delta,\alpha,\pi,\rho_1,\rho_0\right) = \left\{\begin{array}{c} \frac{1}{N}\sum_{i=1}^{N} \frac{\partial L_i(\beta)}{\partial \beta} \\ \frac{1}{N}\sum_{i=1}^{N} \begin{pmatrix} D_i \\ Z_i \\ (\log T_i) \\ \mu_{1i}(X_i;\,\beta) \\ \mu_{0i}(X_i;\,\beta) \end{pmatrix} \xi_i \end{array}\right\}$$

$$\xi_i = w_i - \alpha D_i - Z_i'\delta - \pi \log T_i - \rho_1 \mu_{1i}(X_i;\,\beta) - \rho_0 \mu_{0i}(X_i;\,\beta)$$

$$A_N = I_{(2p+r+4)}$$

This optimisation problem has been solved using the Newton-Rhapson method and taking the two-step estimates as initial values.

Finally, the wage equation estimates have not been corrected using the panel data techniques described by Arellano and Honoré (2001), in order to take into account the presence of unobserved heterogeneity correlated with the explanatory variables. The reason is that I cannot apply first differences to the wage equation to remove the fixed effect, since most explanatory variables are time invariant, and their coefficients will not be identified.

4.2 Estimation results

The vector of explanatory variables, Z_i, includes new-job and personal characteristics. First, country indicators are introduced except for Spain. Concerning new-job characteristics, I use an indicator of whether the working time is full-time, and firm size indicators, in which a firm is small if the number of workers is lower than 100, medium if this number ranges from 100 to 499 workers, and large if the number of workers is larger than 500. I also include a dummy variable taking the value of 1 if the firm belongs to the private sector, and indicators of the firm's economic activity according to the two-digit NACE, in which the omitted category is A+B+C+E, corresponding to the primary sector, mining and quarrying, electricity, gas and water supply.

I consider the following personal characteristics: the logarithms of the unemployment spell duration and the age, the indicator of whether the individual is male, and dummy variables indicating the level of education, using the level lower than the second stage of secondary education as the omitted category. Finally, I include the indicator of whether the exit to employment is associated with a residential change, as defined in Section 2.

Table A.5 shows some descriptive statistics of the average monthly earnings (adjusted by the Purchasing Power Parities, PPP) that the household heads obtain after leaving unemployment, according to their job and personal characteristics. We can see that the median of the average monthly earnings are higher among household heads having found a full-time job and working in the public sector. The higher the education level and the larger the firm size are, the higher the median of the average monthly earnings that individuals obtain after leaving unemployment and the larger the wage dispersion (measured by the interquartile range) are. Moreover, both the younger and the older people are, the lower the earnings they can achieve. It also seems to exist a big difference in the level and dispersion of the earnings obtained by male and female household heads that enter a job spell. Furthermore, post-unemployment wages seem to decrease when the family heads have been unemployed for more than 18 months, and exits to a job spell related to geographical mobility are associated with both lower levels and greater dispersion of earnings.

Table 5 shows the estimation results. Specification (i) is estimated by OLS robust to heteroskedasticity, and the other two specifications take into account the possible presence of a self-selection bias by including the two terms, $\mu_{0i}(X_i;\beta)$ and $\mu_{1i}(X_i;\beta)$, which are associated with coefficients ρ_0 and ρ_1, respectively, in equation (4.10).

Column (ii) shows the results of the two-step estimator implied by the unemployment duration model with two alternative exits to employment, and column (iii) the GMM joint estimates of the wage equation and the duration model.

After applying a series of filters to the data, constructing the unemployment duration variable and assigning the explanatory variables to each duration, I obtain a number of $2,150$ spells for these five countries, in which the proportion of complete spells is 64.65%. Thus, I should have a total of $1,390$ wage observations of individuals exiting to employment. However, I can only assign a wage to 800 spells, which represent the 57.55% of the complete unemployment spells.[14]

The estimates of the three specifications are very similar in the coefficients size and the significance of the explanatory variables. It seems not to exist big wage differences

[14]We cannot assign a reemployment wage to those individuals that have worked in two or more different jobs in the same year (in more than two jobs as employees or in more than two jobs self-employed), since we use both the annual total wage and salary earnings and self-employment income for obtaining a monthly average reemployment wage. We cannot identify the reemployment wage, as we cannot distiguish how much income was obtained in each job. See the Appendix for further details.

across countries after having allowed for job and personal characteristics. In specification (iii), the logarithm of the duration has an estimated coefficient of -0.032, with the expected sign, but it is insignificant. This may be due to two reasons: first, the reduced sample size and the high proportion of right-censored spells make duration dependence be not estimated accurately. Second, this sample only consists of entrants into unemployment that are followed for a short time period, for nearly a year for most of them (81.86% of the observations); so, the sample has very short durations without enough variation to capture dependence on wages.

Regarding geographical mobility, the coefficient estimate of the residential change indicator is negative in all specifications, but insignificant. This may be caused by the small sample size and the tiny frequency of residential mobility. In addition, the existence of opposite effects can make this coefficient be biased toward zero, as commented at the beginning of this section.

Concerning personal and job characteristics, most of them are significant and have the expected effects. The logarithm of age has a positive effect on post-unemployment wages; this variable captures the influence of the labour market experience on wages. The education indicators are also significant at the 1% level in explaining the post-unemployment wages as expected.

Another expected result is that individuals working full-time earn on average a higher monthly wage than those working part-time; however, wage differences are not observed among individuals working in the private and public sectors. With respect to the firm size, individuals working in medium firms are paid higher wages than those in small firms, since its coefficient estimate in specification (iii) is 0.277, significant at the 1% level. Although the effect of large firm size is positive, it is not significant. I tried to include interactions of the firm size with the level of education in order to capture different ways of firms' remuneration for the level of education according to their size. That is, individuals having a low level of education may earn similar wages independently of the firm size, small or large; however, high-skilled individuals are expected to be better-paid in large firms than in small firms. Nevertheless, these interactions are not significant, and they are finally removed from the estimates. As unemployment is expected to affect low-skilled individuals more seriously, the composition of the sample of post-unemployment wages makes this education effect difficult to identify. The lowest-educated workers account for the 69.88% of the sample.

Concerning firm's or unit's economic activity, sectors like education (M); transport, storage and communication (I); hotels and restaurants (H); real state, renting and business activities (K); and wholesale and retail trade (G) pay relatively more than the omitted category does, (A+B+C+E). Appendix A details the correspondence of this code with the economic activity.

Finally, the coefficient estimates of the terms allowing for the self-selection bias, ρ_1 and ρ_0, are 0.050 and 0.045, respectively, but they are not significant. Both values would be consistent with the idea that wages and exits to a job spell are positively correlated. The unemployed workers may refuse wage offers that are very low, under their reservation wage. So, the mean of the accepted wage distribution seems to be higher than that of the potential wage distribution.[15]

[15] The two terms that controls for the self-selection bias, $\mu_{0i}(X_i;\beta)$ and $\mu_{1i}(X_i;\beta)$, always take positive values in the sample of wages.

5 Conclusions

The aim of this paper is to study three questions: first, how housing tenure, ownership or rental, influences the unemployed workers' geographical mobility decision using a discrete model of unemployment duration with two alternative exits to employment, depending on whether they are associated with a residential change or not; these transition intensities follow a multinomial logit specification. Second, I address the potential endogeneity of the ownership status to the transition intensities. I investigate the main determinants of the housing tenure status, taking into account the differences in the policies supporting the access to a dwelling across countries. The ownership status equation is used to estimate the unemployment duration model allowing for unobserved heterogeneity. Third, I study the main characteristics affecting the wage that individuals obtain just after leaving unemployment. Particularly, I try to find out whether the unemployment duration affects wages negatively, and whether the unemployed workers who exit to a job spell associated with a residential change have higher wages than those who find a job in their local area.

For this purpose, I use individual data from the European Community Household Panel (ECHP) for five waves covering the period 1994-98. As the surveys are homogeneous across countries, I use a pooling of countries to make comparisons among them.

The sample in all estimates consists of unemployment spells of family heads aged 25 to 64 years old, having previous labour market experience, and coming from one of the following European countries: France, Germany, Italy, Spain and the United Kingdom. An important limitation to study geographical mobility using the ECHP is the impossibility of differentiating geographical mobility from residential mobility perfectly. I can only know that the residential changes occur within the same province, from other province within the same country and from other country. Due to the limited number of inter-regional movements observed in all European countries, I identify a case of geographical mobility as a residential change that occurred at some point of the unemployment spell or that was preceded by an exit to employment two months ago, at most. In spite of this problem, I use the ECHP survey, since this is the only fixed panel available in some of these countries for studying issues of housing tenure, residential mobility and labour mobility. I have obtained the following results: concerning the effect of housing tenure on labour mobility, I find that homeowners and renters living in social housing are more reluctant to accept a job associated with a residential change than private renters, since they have large negative coefficient estimates, and they are significant at the 1% and 5% levels, respectively. On the contrary, housing tenure does not affect the unemployed workers' behaviour in the local labour market; that is, the reservation wage for accepting job offers coming from their local area is not altered by the type of housing tenure. However, owners having outstanding mortgages leave unemployment in their local area with a higher probability; this type of owners seems to search for a job more intensively, since they have to face outstanding debts. Finally, concerning the probabilities predicted by the model, we can see that the unemployed workers' geographical mobility is very reduced in all countries.

With respect to the main determinants of the housing tenure status, the age profile of the home ownership is similar to that predicted by the life-cycle theory: when individuals are younger, they have a higher probability of living in a rented house; when they have accumulated enough wealth to invest in an owned house, they switch from rental to home ownership, and finally when they are older, the proportion of individuals living in an owned house starts to fall. The household income affects the probability of being an owner positively. Housing markets in Spain, the United Kingdom, and Italy make home ownership more attractive than private rental, in comparison with those in Germany and France. When country indicators are replaced by other indicators evaluating housing policies that encourage home ownership, this status is observed to respond negatively to the importance of the stamp duties incurred in the purchase of a dwelling, and the ratio social to private rented accommodation, and positively to the existence of interest tax-reliefs on income taxes.

When we make the ownership status endogenous to transitions to employment due to the presence of unobserved human capital or other factors, such as the idea of rental as a temporary housing tenure regime for those individuals who expect to move in the near future, the results are reinforced. The ownership coefficient in exits associated with a residential change becomes more negative and significant. The value of the log-likelihood function does not increase so much when unobserved heterogeneity is specified as a discrete variable with a finite support of more than two mass points. Finally, the average monthly wage that individuals obtain just after leaving unemployment depends on personal and job characteristics in the way expected, except for the indicators of large firm size and the private sector, since neither are significant. In order to estimate the mean of the potential wage distribution, we have to control for the possible presence of a self-selection bias. This arises due to the fact that a subsample of individuals decides to remain unemployed; they reject all wage offers being lower than their reservation wage. The corrections are similar to Heckman's two-step estimator, but they are derived from the duration model with multinomial logit transitions to employment. As this two-step estimator does not take into account the inclusion of predicted regressors, I estimate the wage equation jointly with the unemployment duration model of multiple employment exits by GMM, in order to obtain more efficient estimates. However, due to the small sample size of post-unemployment wages, the effect of the length of the unemployment spell and the geographical mobility are not captured accurately.

Appendix: Database description

A Personal and household characteristics

The data come from the European Community Household Panel (ECHP) for the period 1994-98, and they are provided by Eurostat.

Sample: is formed by family heads aged 25 to 64 years old, who do not satisfy any of the following characteristics:

- Full-time or part-time employed, self-employed, retired, staying in paid apprenticeship or training under special schemes relative to employment, unpaid work in a family enterprise, education or training, community or military service, and other economically inactivity different from doing housework and looking after children or other persons; not having previous experience in a job, and living in rent-free accommodation.
- Missing observations in the following variables: age, gender, cohabitation in a relation, lack of personal information on the spouse or the partner, education, household composition, year of the move to the address, housing tenure and where individuals move from.

Duration: the sample consists of entrants into unemployment. The durations are constructed using the indicators of the individuals' main economic activity carried out in each month of the year prior to the survey. These durations can start in any month from January 1994 to December 1997.

Experience: is constructed by using the indicators of the main economic activity from January 1993 to December 1997 and summing the months for which individuals are employed. When individuals are employed in January 1993, the experience variable is completed by multiplying 12 by the number of years worked at the current job in January 1993. This number is obtained through the variable of year of the start of the current job.

Previous job economic sector and working time; post-unemployment job working time, firm's size, firm's business sector (whether private or public) and two-digit NACE economic activity: are constructed by matching the information available for the current and last jobs in each wave of the survey, according to the year and the month in which the individual starts and finishes to work in both the job previous to an unemployment spell and the post-unemployment job.

Education: is the highest level of education completed, broken down by three categories: recognised third level of education (ISCED 5-7), second stage of secondary level of education (ISCED 3), and less than second stage of secondary education (ISCED 0-2).

Income: is the logarithm of the household income received in the previous year. Income is made homogeneous across countries using the Purchasing Power Parity (PPP).

Indicator of a residential change during the unemployment spell: is constructed using information on the year and the month in which the individual starts to live at the current address, and using information on where the previous dwelling was located.

Housing tenure: Three types of housing tenure are considered: home ownership, private rental and subsidised rental (the accommodation is rented from a public, municipal, voluntary, or non-profit agency or from a household member's employer). Rent-free accommodation is left out of the sample.

Average monthly wage: is constructed using information on the total net income from work, earned in the year prior to the survey (wage and salary earnings or self-employment income), and the indicators of the main economic activity done in each month of the year prior to the survey.

Two-digit NACE economic activity: A+B+C+E denotes primary sector, mining and quarrying, electricity, gas and water supply; DA manufacture of food products, beverages and tobacco; DB+DC manufacture of textiles, clothing and leather products; DD+DE manufacture of wood and paper products, publishing and printing; DF-DI manufacture of coke, refined petroleum/chemicals/rubber and plastic products, etc.; DJ-DK manufacture of metal products, machinery and equipment n.e.c.; DL-DN other manufacturing; F construction; G wholesale and retail trade, repair of motor vehicles, motorcycles and personal/household goods; H hotels and restaurants; I transport, storage and communication; J financial intermediation; K real state, renting and business activities; L public administration and defense, compulsory social security; M education; N health and social work; O-Q other community, social and personal service activities, private households with employed persons, extra-territorial organisations and bodies.

B Aggregate variables

Purchasing Power Parity: is provided by Eurostat for the period covered by the survey.

Quarterly national unemployment rates: come from the *Main Economic Indicators* from the OECD.

Economic indices: the index of the strictness of the employment protection legislation (EPL) is extracted from the *OECD Employment Outlook*, June (1999). The percentages of transaction costs and the stamp duties incurred in the purchase of a house on the house price, the indicator of the existence of an interest tax-relief on income taxes, and the ratio social to private rented accommodation are extracted from Maclennan *et al.* (1998).

Ln(transaction tax(%)): is constructed as

$$\ln(\text{transaction tax }(\%)) = \ln\left(\frac{\text{transaction taxes}}{\text{house price}}\cdot 100\right)$$

EPL strictness: this index takes a value of the range from 0 to 6, and the higher it is, the stricter the employment protection legislation. The logarithm of this index is introduced in the estimation.

Ln(stamp duty): is defined as

$$\ln(\text{stamp duty}) = \ln\left(\frac{\text{stamp duty}}{\text{house price}} \cdot 100\right)$$

Ln(ratio social to private rented housing): is constructed as

$$\ln(\text{ratio social to private rented housing}) = \ln\left(\frac{\text{social rented housing stock}}{\text{private rented housing stock}}\right)$$

Concerning regular employment, the index of EPL strictness reflects regular procedural inconveniences to the employer for the dismissal, such as the notice period and other procedures; severance pay for no-fault individual dismissals; difficulty in the dismissal, such as the definition of unfair dismissal; trial period before eligibility arises; unfair dismissal compensation; etc. With respect to temporary employment, this index captures the strictness in the conditions under which fixed-term contracts may be used, such as specific projects, seasonal work, replacement of temporarily absent permanent workers (sickness, maternity leave) and exceptional workload; the maximum number of successive contracts; the maximum cumulated duration; and the regulation of temporary work agencies (TWAs) (types of work for which TWA employment is legal, restrictions on the number of renewals and maximum cumulated duration of temporary work contracts). Regarding the regulation of collective dismissals, the index of strictness evaluates their definition, the additional notification requirements to employee representatives and to government authorities, additional delays involved and other special costs to employers (severance pay and social compensation plans).

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Table A.1: Percentage of housing tenure and geographical mobility in the sample of unemployment spells.

	Housing tenure		Geographical	Percentage of	Sample
	Ownership	Rental	mobility rate	unemployment spells	period
Germany	36.00	64.00	2.00	6.98	1994-96
France	36.59	63.41	2.51	16.65	1994-98
United Kingdom	57.42	42.58	0.65	7.21	1994-96
Italy	68.94	31.06	1.24	22.47	1994-98
Spain	82.57	17.43	1.89	46.70	1994-98

Table A.2: Individual characteristics in the sample of unemployment spells.

	No geographical mobility	Geographical mobility
Housing tenure		
Ownership without outstanding mortgage	43.94	23.68
Ownership with outstanding mortgage	23.58	2.63
Private rental	20.08	60.53
Social rental	12.41	13.16
Education		
Third level	8.29	5.26
2^{nd} level secondary	23.58	39.47
Less than 2^{nd} level secondary	68.13	55.26
Living with a partner		
No	19.79	26.32
Spouse/partner employed	29.36	15.79
Spouse/partner not employed	50.85	57.89
Gender		
Male	85.84	86.84
Female	14.16	13.16
Children aged		
$[0, 6]$ years old	29.69	42.11
$[7, 11]$ years old	25.24	23.68
$[12, 18]$ years old	33.95	34.21
Economic sector		
Agriculture	12.78	5.26
Industry	19.70	18.42
Services	44.84	55.26
Construction	22.68	21.05
Experience at previous job		
≤ 12 months	58.10	71.05
$[13, 24]$ months	13.64	10.53
$[25, 36]$ months	5.59	2.63
> 36 months	22.68	15.79

Table A.3: Duration frequencies in entrants into unemployment.

	Total	Germany	France	United Kingdom	Italy	Spain
Censored	35.35	70.67	33.52	59.35	27.95	30.58
Completed	64.65	29.33	66.48	40.65	72.05	69.42
No. of spells	2,150	150	358	155	483	1,004

Duration of the unemployment spell	Percentage	Duration of the unemployment spell	Percentage
1 month	15.53	25 months	0.37
2 months	15.02	26 months	0.37
3 months	11.35	27 months	0.56
4 months	7.02	28 months	0.37
5 months	5.86	29 months	0.28
6 months	6.65	30 months	0.23
7 months	3.35	31 months	0.37
8 months	3.63	32 months	0.09
9 months	3.44	33 months	0.05
10 months	1.77	34 months	0.09
11 months	1.40	35 months	0.05
12 months	6.84	36 months	0.56
13 months	0.98	37 months	0.23
14 months	1.63	38 months	0.14
15 months	1.26	39 months	0.14
16 months	1.40	40 months	0.05
17 months	1.35	41 months	0.09
18 months	0.70	42 months	0.05
19 months	0.84	43 months	0.05
20 months	0.47	44 months	0.05
21 months	0.74	45 months	0.05
22 months	0.88	46 months	0.05
23 months	0.60	47 months	0.09
24 months	2.56	48 months	0.37

Table A.4: Percentage of individuals exiting to a job spell broken down by the receipt of unemployment benefits and the duration of the unemployment spell.

	Completed spells (%)			Completed spells (%)	
Duration of the unemployment spell	Unemployment benefits	No benefits	Duration of the unemployment spell	Unemployment benefits	No benefits
1 month	11.09	14.58	25 months	0.00	3.28
2 months	13.34	12.05	26 months	4.88	0.00
3 months	11.21	10.57	27 months	8.57	5.88
4 months	9.00	9.35	28 months	0.00	2.17
5 months	8.55	8.08	29 months	4.00	2.44
6 months	12.90	8.81	30 months	0.00	10.00
7 months	8.09	4.52	31 months	5.26	8.33
8 months	8.58	7.22	32 months	6.67	3.13
9 months	6.34	6.69	33 months	0.00	0.00
10 months	4.95	4.15	34 months	0.00	3.33
11 months	4.18	2.59	35 months	0.00	0.00
12 months	10.06	10.04	36 months	0.00	3.33
13 months	3.08	3.07	37 months	0.00	4.55
14 months	5.74	3.13	38 months	20.00	5.26
15 months	2.72	1.33	39 months	0.00	5.56
16 months	4.88	3.50	40 months	0.00	0.00
17 months	6.90	3.03	41 months	0.00	7.14
18 months	1.60	2.44	42 months	0.00	8.33
19 months	3.51	2.52	43 months	0.00	0.00
20 months	3.92	0.00	44 months	0.00	0.00
21 months	5.32	1.80	45 months	0.00	0.00
22 months	4.94	3.70	46 months	0.00	0.00
23 months	5.33	1.05	47 months	0.00	0.00
24 months	10.94	9.68	48 months	0.00	0.00

Table A.5: Average monthly earnings converted to equivalent units using the Purchasing Power Parities (PPP).

	Median	Interquartile Range	25% percentile	75% percentile
Countries:				
Germany	945.65	760.40	620.73	1,381.13
France	820.34	420.85	582.44	1,003.30
United Kingdom	1,095.00	880.09	693.26	1,573.35
Italy	942.84	577.49	589.28	1,166.77
Spain	874.62	444.77	652.34	1,097.11
Job characteristics				
Working time:				
Full-time job	900.90	465.49	675.68	1,141.16
Part-time job	555.97	362.73	406.50	769.23
Firm size:				
Large	1,076.37	732.64	745.54	1,478.18
Medium	999.41	715.97	820.34	1,536.31
Small	866.43	493.53	591.72	1,085.25
Business sector:				
Private	883.91	469.78	638.14	1,107.92
Public	909.17	644.11	558.66	1,202.77
Personal characteristics				
Age bands:				
[25, 29]	778.53	431.90	589.28	1,021.17
[30, 44]	894.05	501.46	658.13	1,159.59
[45, 64]	885.71	505.77	603.70	1,109.47
Education:				
Third level	1,067.46	808.65	748.06	1,556.71
2nd stage secondary	974.38	654.86	687.38	1,342.24
Less than 2nd stage secondary	843.02	470.33	580.72	1,051.05
Sex:				
Male	897.79	457.89	668.24	1,126.13
Female	690.94	557.98	443.79	1,001.77
Geographical mobility	817.59	702.50	596.63	1,299.13
Local labour market	885.80	480.60	630.10	1,110.70
Unemployment spell duration:				
[1, 6] months	885.16	508.52	638.14	1,146.66
[7, 12] months	885.71	421.70	638.99	1,060.70
[13, 18] months	889.55	485.55	517.75	1,003.30
> 18 months	838.32	486.35	512.13	998.49

Table 1: Estimates of multinomial logit transition intensities to employment associated with a residential change (θ_1) or not (θ_0).

$$\theta_k(t \mid X_i(t)) = \frac{\exp(X_i(t)'\beta_k)}{1+\exp(X_i(t)'\beta_0)+\exp(X_i(t)'\beta_1)}; \quad k = 0,\, 1$$

	(i) Country dummies				(ii) Economic variables			
	$\theta_1(t \mid X_i(t))$		$\theta_0(t \mid X_i(t))$		$\theta_1(t \mid X_i(t))$		$\theta_0(t \mid X_i(t))$	
	Coefficient	t-ratio	Coefficient	t-ratio	Coefficient	t-ratio	Coefficient	t-ratio
Country dummies								
Germany	−14.126	−3.42	−1.728	−2.37				
France	−9.956	−3.40	−0.625	−1.22				
United Kingdom	−17.474	−3.46	−1.604	−1.85				
Italy	−10.538	−3.50	−0.492	−0.93				
Economic variables								
Ln(transaction costs (%))					−0.299	−0.25	1.284	4.24
Ln(EPL strictness)					1.010	0.47	−1.651	−3.32
Ln(benefit replacement rate(%))					0.260	0.58	−0.515	−4.68
Ln(unemployment rate)	−16.607	−3.29	−1.209	−1.41	0.194	0.22	0.630	3.03
Housing characteristics								
Housing tenure:								
Ownership	−1.805	−4.03	0.111	1.25	−1.617	−3.65	0.128	1.44
Social housing rental	−1.177	−2.14	0.122	1.08	−1.092	−2.00	0.128	1.14
Outstanding mortgage	−1.364	−1.28	0.343	4.26	−1.384	−1.30	0.343	4.26
Previous job characteristics								
Economic sector:								
Industry	0.220	0.42	−0.149	−1.74	0.305	0.60	−0.148	−1.73
Services	0.057	0.13	−0.242	−3.39	0.028	0.07	−0.245	−3.44
Working time:								
Full-time job	0.228	0.36	0.155	1.38	0.230	0.36	0.156	1.39
Experience:								
Ln(experience)	0.042	0.16	−0.166	−4.75	0.019	0.07	−0.169	−4.83
Ln(experience)*ln(duration)	−0.138	−1.14	0.016	0.76	−0.144	−1.19	0.017	0.82

Table 1: Estimates of multinomial logit transition intensities to employment associated with a residential change (θ_1) or not (θ_0). (Cont.)

$$\theta_k(t \mid X_i(t)) = \frac{\exp(X_i(t)'\beta_k)}{1+\exp(X_i(t)'\beta_0)+\exp(X_i(t)'\beta_1)}; \quad k = 0,\, 1$$

	(i) Country dummies				(ii) Economic variables			
	$\theta_1(t \mid X_i(t))$		$\theta_0(t \mid X_i(t))$		$\theta_1(t \mid X_i(t))$		$\theta_0(t \mid X_i(t))$	
	Coefficient	t-ratio	Coefficient	t-ratio	Coefficient	t-ratio	Coefficient	t-ratio
Personal characteristics								
Education:								
Third level	1.122	0.90	0.508	2.68	1.045	0.86	0.506	2.67
Third level*ln(duration)	−0.604	−0.77	−0.147	−1.28	−0.599	−0.80	−0.151	−1.32
2^{nd} stage secondary	0.936	2.31	0.097	1.21	0.861	2.15	0.090	1.11
Ln(age)	−0.716	−0.90	−0.791	−5.66	−0.688	−0.88	−0.784	−5.62
Male family head	0.555	0.89	0.326	2.95	0.680	1.10	0.329	2.98
Living with a partner	−0.104	−0.18	0.045	0.44	−0.127	−0.22	0.048	0.47
Spouse/partner is employed	−0.770	−1.58	−0.024	−0.33	−0.630	−1.31	−0.025	−0.34
No. of children aged $[0, 18]$	0.596	1.40	0.030	0.45	0.634	1.51	0.033	0.51
Squared no. of children aged $[0, 18]$	−0.086	−0.83	−0.001	−0.05	−0.102	−1.00	−0.003	−0.15
Unemployment benefits receipt	−0.136	−0.18	−0.157	−1.44	−0.171	−0.23	−0.170	−1.55
Benefits receipt*ln(duration)	0.024	0.07	0.209	3.27	0.004	0.01	0.212	3.32
Unemployment duration dependence								
Ln(duration)	0.389	0.54	−0.079	−0.70	0.348	0.49	−0.087	−0.76
Ln^2(duration)	0.013	0.08	−0.177	−5.49	0.027	0.16	−0.176	−5.45
Constant	47.925	3.06	4.827	1.79	−4.848	−1.37	−0.174	−0.26
Seasonal indicators								
First quarter	−0.942	−1.91	−0.328	−3.84	−1.245	−2.61	−0.373	−4.50
Second quarter	−0.747	−1.62	−0.052	−0.63	−1.015	−2.28	−0.082	−1.01
Third quarter	−1.211	−2.40	0.011	0.14	−1.350	−2.68	−0.020	−0.25
Log-likelihood	-4,476.45				-4,484.91			
Number of spells	2,150				2,150			

Notes: Specification (i) includes country dummies in the estimates, and specification (ii) replaces these dummies by indices evaluating characteristics of the housing and labour markets, particularly, the transaction costs in the purchase as % of the house price, the strictness of the employment protection legislation (EPL) and the average of the unemployment benefit replacement rates by unemployment duration and family circumstances across country.

Table 2: Predicted probabilities (%) of multinomial logit transitions to employment. Specification (i): Country dummies.

$$\theta_k(t \mid X_i(t)) = \frac{\exp\left(X_i(t)'\beta_k\right)}{1 + \exp\left(X_i(t)'\beta_0\right) + \exp\left(X_i(t)'\beta_1\right)}; \quad k = 0,\, 1$$

	Transitions to a job spell associated with	
	Residential change (θ_1)	No residential change (θ_0)
Reference person	0.64	11.80
Countries		
Germany	0.89	6.33
France	1.77	13.60
United Kingdom	0.53	8.64
Italy	0.94	15.33
Ownership		
Non outstanding loans	0.10	13.07
Outstanding mortgage	0.03	17.50
Social housing rental	0.20	13.18
Previous job characteristics		
Economic sector		
Agriculture/Construction	0.30	10.11
Services	0.33	8.11
Working time		
Part-time job	0.52	10.29
Experience attained		
12 months	0.59	10.81
18 months	0.55	10.26
24 months	0.53	9.89
51 months	0.48	8.97
Personal characteristics		
Education		
Third level	0.82	15.33
2nd stage secondary	1.60	12.72
Female family head	0.38	8.83
Age		
25 years old	0.72	13.37
45 years old	0.50	8.86
55 years old	0.44	7.67
Living with a partner		
Spouse/partner not employed	0.58	12.28
Spouse/partner employed	0.27	12.07
No. of children aged $[0, 18]$ *years old*		
1 child	1.06	12.05
2 children	1.47	12.29
3 children	1.73	12.53
Unemployment benefits receipt	0.57	13.25

Notes: The reference person is a male family head living in a private rented house in Spain. He is single, aged 30 years old, does not have any children and his level of education is lower than the second level of secondary. He has been unemployed for 4 months, has previously worked full-time in the industry for 6 months, and does not receive any unemployment benefits. The unemployment rate is evaluated at its average level in 1995, and the seasonal indicators are set at the fourth quarter.

Table 3: Logit estimates of determinants of the ownership status in the sample of unemployment spells.

$$\Pr(h_i \mid Z_i) = \Pr(h_i^* \mid Z_i)\,[1 - \Pr(q_i = 1 \mid Z_{2i})]\,,\; h_i = h_i^*,\; h_i = 0,\, 1;$$
$$\Pr(h_i = 2 \mid Z_i) = \Pr(q_i = 1 \mid Z_{2i});$$

	(i)Country dummies		(ii) Economic variables	
$\Pr(h_i^* = 1 \mid Z_i)$	Coefficient	t-ratio	Coefficient	t-ratio
Countries				
Germany	−2.344	−9.27		
France	−1.814	−9.99		
United Kingdom	−0.450	−1.57		
Italy	−0.668	−4.22		
Support for home ownership				
Ln(stamp duty)			−0.975	−6.64
Interest tax relief			2.076	7.63
Ln(ratio social to private rented housing)			−0.697	−9.31
Household characteristics				
Ln(total income (PPP))	0.371	4.41	0.359	4.34
Ln(family head's age)	3.102	11.01	3.116	11.07
Male family head	0.061	0.32	0.053	0.27
Living with a partner	0.609	3.29	0.627	3.41
No. children aged $[0, 18]$ years old	0.241	1.68	0.236	1.64
Squared no. children aged $[0, 18]$	−0.072	−1.91	−0.071	−1.88
Family head's education:				
Third level	−0.090	−0.39	−0.111	−0.49
2nd stage secondary	0.230	1.45	0.222	1.40
Constant	−13.707	−10.98	−15.556	−11.75
Log-likelihood	-837.90		-838.25	
Number of spells	1883		1883	
$\Pr(q_i = 1 \mid Z_{2i})$				
Countries				
Germany	2.106	7.56		
France	2.291	10.67		
United Kingdom	2.473	9.68		
Italy	1.136	5.05		
Household characteristics				
Ln(total income (PPP))	−0.247	−2.42		
Ln(family head's age)	−1.651	−5.04		
No. children aged $[0, 18]$ years old	0.381	2.68		
Squared no. children aged $[0, 18]$	0.001	0.02		
Constant	4.573	3.20		
Log-likelihood	-664.41			
Number of spells	2,150			

Notes: In specification (i), the tenure status equation is estimated by controlling the institutional differences across countries through the inclusion of country dummies, and, in specification (ii), through three indicators evaluating the support for home ownership.

Table 4: Joint estimates of multinomial transitions to employment with the ownership status equation controlling for unobserved heterogeneity.

$$\theta_k\left[t \mid X_i(t),\, h_i,\, \eta_i\right] = \frac{\exp(X_i(t)'\beta_k + h_i\beta_{hk} + \alpha_k\eta_i)}{1 + \sum_{j=0}^{1}\exp(X_i(t)'\beta_j + h_i\beta_{hj} + \alpha_j\eta_i)}, \quad k = 0, 1;$$

	$\theta_1(t \mid X_i(t))$		$\theta_0(t \mid X_i(t))$	
	Coefficient	t-ratio	Coefficient	t-ratio
Country dummies				
Germany	−14.406	−3.34	−1.543	−2.09
France	−10.145	−3.33	−0.498	−0.96
United Kingdom	−17.662	−3.43	−1.529	−1.76
Italy	−10.664	−3.46	−0.430	−0.81
Economic variables				
Ln(unemployment rate)	−16.787	−3.26	−1.126	−1.31
Housing characteristics				
Housing tenure:				
Ownership	−2.066	−2.79	0.600	2.21
Social housing rental	−1.318	−2.04	0.426	2.18
Outstanding mortgage	−1.353	−1.27	0.327	4.01
Previous job characteristics				
Economic sector:				
Industry	0.207	0.40	−0.150	−1.74
Services	0.060	0.14	−0.245	−3.40
Working time:				
Full-time job	0.245	0.39	0.141	1.23
Experience:				
Ln(experience)	0.037	0.14	−0.164	−4.67
Ln(experience)*ln(duration)	−0.134	−1.10	0.013	0.64
Personal characteristics				
Education:				
Third level	1.075	0.86	0.547	2.85
Third level*ln(duration)	−0.602	−0.77	−0.153	−1.33
2^{nd} stage secondary	0.934	2.30	0.097	1.19
Ln(age)	−0.577	−0.69	−0.902	−5.93
Male family head	0.578	0.92	0.312	2.79
Living with a partner	−0.112	−0.20	0.045	0.43
Spouse/partner is employed	−0.739	−1.50	−0.045	−0.60
No. of children aged $[0, 18]$	0.623	1.46	0.009	0.13
Squared no. of children aged $[0, 18]$	−0.091	−0.87	0.004	0.22
Unemployment benefits receipt	−0.111	−0.15	−0.172	−1.56
Benefits receipt*ln(duration)	0.013	0.04	0.212	3.30
Unemployment duration dependence				
Ln(duration)	0.372	0.52	−0.063	−0.55
Ln^2(duration)	0.015	0.09	−0.179	−5.51
Constant	48.127	3.05	4.631	1.71
Seasonal indicators				
First quarter	−0.937	−1.90	−0.332	−3.87
Second quarter	−0.744	−1.61	−0.054	−0.65
Third quarter	−1.211	−2.40	0.007	0.08

Table 4: Joint estimates of multinomial transitions to employment with the ownership status equation controlling for unobserved heterogeneity. (Cont.)

$$\Pr(h_i \mid Z_i,\ \eta_i) = \Pr(h_i^* \mid Z_i,\ \eta_i)\left[1 - \Pr(q_i = 1 \mid Z_{2i})\right],\ h_i = h_i^*,\ h_i = 0,\,1;$$

$$\Pr(h_i = 2 \mid Z_i,\ \eta_i) = \Pr(q_i = 1 \mid Z_{2i});$$

Housing tenure status:	Coefficient	t-ratio
$\Pr(h_i^* = 1 \mid Z_i,\ \eta_i)$		
Support for home ownership		
Ln(stamp duty)	−3.232	−6.60
Interest tax relief	5.916	8.02
Ln(ratio social to private rented housing)	−1.455	−6.04
Household characteristics		
Ln(total income (PPP))	1.184	6.28
Ln(family head's age)	6.022	7.25
Male family head	0.750	1.88
Living with a partner	0.859	2.08
No. children aged $[0, 18]$ years old	0.570	1.67
Squared no. children aged $[0, 18]$	−0.167	−1.98
Family head's education:		
Third level	−1.325	−2.55
2^{nd} stage secondary	−0.174	−0.45
Constant	−34.130	−8.99
$\Pr(q_i = 1 \mid Z_{2i})$		
Countries		
Germany	2.106	7.56
France	2.291	10.67
United Kingdom	2.473	9.68
Italy	1.136	5.05
Household characteristics		
Ln(total income (PPP))	−0.247	−2.42
Ln(family head's age)	−1.651	−5.04
No. children aged $[0, 18]$ years old	0.381	2.68
Squared no. children aged $[0, 18]$	0.001	0.02
Constant	4.573	3.20
Unobserved heterogeneity parameters:		
α_0	−1.944	−0.42
α_1	1.000	
α	27.029	0.43
m_1	0.059	0.43
m_2	−0.247	
p_1	0.807	47.21
Number of individuals	1,459	
Number of spells	2,150	
Log-likelihood	-5,806.10	

Table 5: Estimates of a wage equation from a sample of individuals just after leaving unemployment.

$$w_i = \alpha D_i + Z_i'\delta + \pi \log T_i + \rho_0\mu_0(X_i;\beta) + \rho_1\mu_1(X_i;\beta) + \xi_i \quad \text{for } c_i = 1$$

	(i) OLS estimates		(ii) Two-step estimates		(iii) GMM joint estimates	
Countries						
Germany	0.005	(0.05)	−0.077	(-0.58)	−0.042	(-0.31)
France	−0.080	(-1.24)	−0.074	(-1.15)	−0.080	(-1.25)
United Kingdom	−0.127	(-0.70)	−0.165	(-0.89)	−0.148	(-0.80)
Italy	−0.031	(-0.49)	−0.023	(-0.36)	−0.028	(-0.44)
Job characteristics						
Working time:						
Full-time job	0.457	(5.22)	0.466	(5.24)	0.462	(5.32)
Firm size:						
Medium	0.272	(3.98)	0.282	(4.08)	0.277	(4.07)
Large	0.209	(1.55)	0.205	(1.52)	0.207	(1.57)
Private sector	−0.102	(-1.22)	−0.125	(-1.46)	−0.113	(-1.35)
Personal characteristics						
Logarithm of age	0.241	(2.32)	0.158	(1.28)	0.206	(1.72)
Education:						
Third level	0.309	(3.75)	0.332	(3.90)	0.321	(3.82)
2^{nd} stage secondary	0.229	(3.56)	0.232	(3.55)	0.230	(3.54)
Male	0.229	(3.23)	0.267	(3.71)	0.249	(3.47)
Ln(unemployment duration)	0.017	(0.66)	−0.028	(-0.74)	−0.003	(-0.08)
Residential change	−0.019	(-0.11)	−0.032	(-0.18)	−0.028	(-0.16)
Constant	4.913	(11.20)	4.875	(10.34)	4.871	(10.41)
ρ_0			0.100	(1.35)	0.045	(0.62)
ρ_1			0.088	(0.63)	0.050	(0.32)
R^2	0.16					
No. of observations	800		2,150		2,150	

Notes: These specifications include industry dummies NACE-2 classification. See Appendix A.
The specification (i) is estimated by OLS robust to heteroskedasticity.
The specification (ii) is also estimated by OLS robust to heteroskedasticity, and it controls for the presence of self-selection bias using a two-step estimator similar to Heckman's lambda.
In columns (i) and (ii), both participation equations arise from specification (i) of the duration model with two multinomial logit transitions to employment, shown in Table 1.
Column (iii) shows joint estimates using the generalised method of moments (GMM).

Figure 1: Kernel density estimates of the family head's age according to the housing tenure.

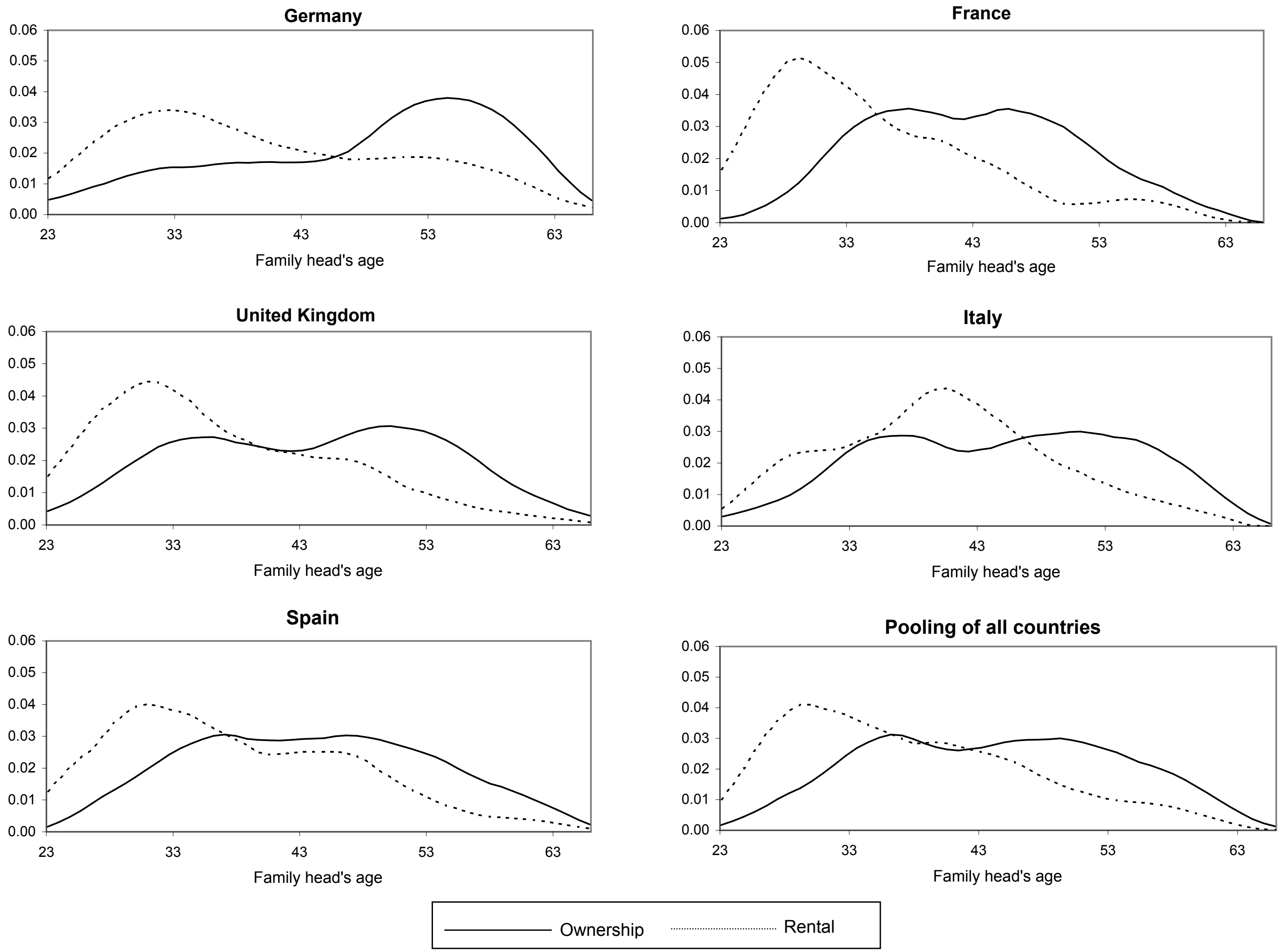
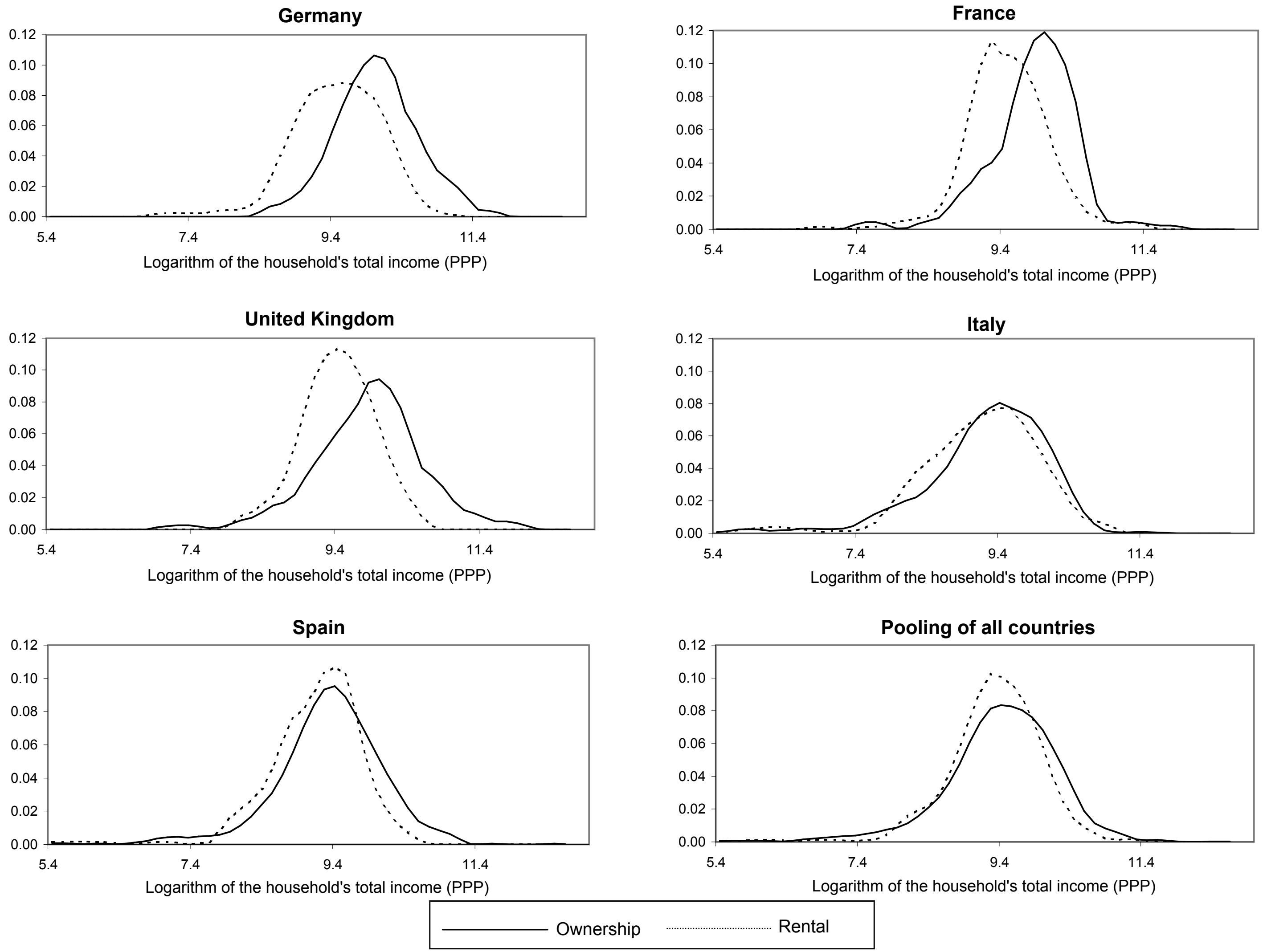


Figure 2: Kernel density estimates of the logarithm of the household's total income according to the housing tenure.



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