Housing Tenure and Labour Mobility: A Comparison across European Countries

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This draft: November 2001
Preliminary and incomplete
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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's labour mobility is studied using a discrete unemployment duration model with two exits to employment, depending on whether they are associated with a residential change or not. Ownership is found to affect negatively geographical mobility. Second, the results are robust to potential endogeneity of the ownership status and institutional differences across countries. Third, post-unemployment wages are studied: the effects of unemployment spell length, geographical mobility and the existence of a self-selection bias.

Keywords: Labour mobility, Housing tenure, Duration models, Self-selection bias, Wage equation.
JEL classification: J61, R20, C41, C24, J31.

*I wish to thank Manuel Arellano for his comments and excellent supervision, as well as Samuel Bentolila, Olympia Bover, Michael Manove and Pedro Mira for their helpful comments and suggestions, and all my Ph. D. colleagues, specially Pedro Albarrán, Jesús Carro and Francisco Peñaranda. All errors are my responsability.
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1 Introduction

The purpose of this paper consists in showing empirical evidence for the impact of housing tenure on the unemployed's decision of geographical mobility in five European countries. That will be made by taking into account the institutional differences among them. To study both questions, housing tenure and labour mobility, I use a sample of unemployed family heads coming from the *European Community Household Panel Survey* (hereafter, ECHP) for the available period 1994-96.

A lot of empirical works have addressed the topics of mobility and housing tenure. In the case of UK, Hughes and McCormick (1981) and Henley (1998) have studied how housing tenure affects mobility; they find a negative impact of social rented housing and ownership, mainly the negative housing equity, on migration. These studies as well as others as Hughes and McCormick (1994) and Pissarides and Wadsworth (1989) have found that geographical mobility does not respond to the regional economic conditions as high unemployment rates in the expected way.

Housing and geographical mobility are also two important issues in Spain, given the high Spanish unemployment and ownership rates, 22.7% and 85.9% in 1995, the highest in the European Community. Empirical work studying the response of mobility to differentials in the regional labour markets in Spain can be found in Bentolila and Dolado (1990), Bentolila (1997), Antolín and Bover (1997), and Bentolila and Jimeno (1998).

The interest to analyse housing tenure and labour decisions using the ECHP is that this survey is like a merge of the *Labour Force Survey* and the *Family Expenditure Survey*; so, it provides wealthier information for addressing both topics than any of both surveys separately. The ECHP is also the first fixed panel carried out in Spain, at least, and it allows to study jointly housing, residential and labour mobility, since it continues interviewing migrant households.

Moreover, the implementation of this survey in other European countries allows to make a study joint with several countries. That 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. That is the main contribution of this paper to the studies of housing and labour markets. For that purpose, I will focus on the unemployed, since geographical mobility responds to different factors among the employed, that is highly related to their current job characteristics.

The choice of housing tenure is predetermined to a large extent by the institutional characteristics prevailing in each country. As housing is a basic good, policy makers have generally tried to make easier the access of a dwelling through providing social rented housing, subsidies and allowing tax reliefs if individuals rent a house. Home ownership has also been supported, for instance, through the interest relief in income taxes, the relief on maintenance of the house and through the absence of taxes on imputed rents and on capital gains derived from the sale of their own house. Often, these policies have benefited a housing tenure more than the other. This produces that in most European countries a given housing tenure dominates the other.

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 a house. This fact may help to explain up to which extent housing tenure may be influenced by the institutional characteristics of a country and how this structure can explain the degree of labour mobility observed in a country.

These five countries can be sorted in two groups depending on the composition of the housing stock according to the housing tenure, ownership or rental[1]. A first set is formed by Germany and France in which a high percentage of households lives in rental, private or social accommodation, 62% and 38% of the housing stock, respectively. A second group includes Spain, Italy and the United Kingdom in which ownership is the clearly preferred alternative; 78% of the housing stock in Spain is occupied through ownership, and the corresponding proportions in Italy and in the UK are 68% and 67%, respectively. In these countries, the inter-regional mobility as a percentage of population in 1993 is the following: 1.23% and 1.07% for Germany and France, 0.5% and 0.56% in Italy and Spain, and finally 1.58% in the UK. Thus, it seems to exist a negative relationship between ownership and inter-regional mobility rates in these countries except for the UK. Regarding the unemployment rate in 1993, there is also a negative relationship between unemployment and inter-regional mobility rates and another positive between ownership and unemployment rates, since Germany and Spain have the lowest and the highest unemployment rate in 1993, 7.60% and 22.22%. Meanwhile, this rate in France, UK and Italy reaches the values of 11.71%, 10.34% and 10.17%, respectively[2].

The empirical model of housing tenure and labour mobility that I use is a discrete unemployment duration model with two exits to employment, depending on whether they are associated with a residential change or not. These transitions to employment are

[1]Maclennan *et al.* (1998) collect information on the financial and housing markets in the European countries during the 1990s.

[2]The information on the unemployment rate in 1993 comes from "*OECD Economic Outlook*", No. 69.

assumed to follow a multinomial logit specification. Using this sample of unemployment spells, I will also investigate the main determinants of the housing tenure status by taking into account the institutional differences across countries. This ownership status equation will be used to make it endogenous to labour mobility controlling for the presence of unobserved heterogeneity correlated with each other.

Finally, I will be also interested in the wage level that these individuals obtain when they leave their unemployment spell. In particular, I wish to 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 that one they would have obtained in their local area. It is not clear in which direction the wage level moves when the unemployed migrate due to job reasons. That depends on whether the reservation wage of accepting a job in other region is higher than in their local area or not, which is related to their future expectations. This question will be explained in detail in following sections. First, I will estimate the average of wage distribution by controlling for self-selection bias; and second, I will estimate both transitions to employment and the average of the wage distribution jointly using the generalised method of moments, in order to obtain efficient estimates under the existence of self-selection bias. This 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 result is that owners are more reluctant to move than renters; this result is reinforced if I control for the presence of unobserved heterogeneity. As expected, in the local labour market, owners behave in a similar way to renters.

The rest of paper is organised as follows. Section 2 describes the empirical duration model of housing tenure and labour mobility and the results I obtain. Section 3 addresses the potential endogeneity of ownership status by controlling for unobserved heterogeneity in both transitions to employment. Section 4 studies the wage attained by the individuals when they left unemployment by one of these two alternatives: exit to a job associated with a residential change or not. The former alternative is identified as a case of geographical mobility in the empirical model. Finally, section 5 summaries 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' decision of geographical mobility, that is, their incentives to accept a job in other 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. Recently, theoretical studies analysing both decisions of housing tenure and labour mobility can be found in Haavio and Kauppi (2000) and Dohmen (2000).

The empirical study will consist in testing the two main conclusions about how housing tenure affects the individuals' labour decisions derived in Barceló (2001) for a sample of unemployed family heads coming from five European countries. For that purpose, differences in the institutional characteristics across countries will be captured by two different ways: first, I will include country dummies in order to find out whether these differences can identify through the sample and then, I will replace these dummies by indicators of relevant characteristics of housing and labour markets affecting the decision of labour mobility.

Barceló (2001) develops a two-region model of job search in which the individuals take three decisions jointly in 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 due to two reasons. First, individuals are assumed to obtain the consumption of housing services cheaper 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 to have a higher reservation wage than renters for accepting a job in other region. Thus, the probability of geographical mobility will be more reduced for this tenure regime. However, the performance in the labour market in the local area in which individuals live will be identical for both housing regimes. That is, housing tenure will not affect the decisions of acceptance or rejection of job offers which do not involve a residential change; thus, their reservation wage will be identical.

Both conclusions, lower mobility among owners and identical decision rules in local labour market, will be tested using the empirical model described in the following subsections.

Subsection 2.1 describes the data used in the estimates, Subsection 2.2 explains the empirical model and the estimation method implemented, and finally, Subsection 2.3

present the results obtained.

2.1 Data characteristics

The data used in the estimates come from the *European Community Household Panel Survey* (ECHP); currently, there are only available four waves covering period 1994-97. The ECHP consists in a fixed panel with an annual frequency; nevertheless, it also provides information on main economic activity in each month of previous year for all household's members aged over 16 years old. Thus, we can construct monthly durations 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 the Spanish *Labour Force Survey* (LFS) and the *Family Expenditure Survey* (FES); that is, it contains wealthy information on individuals' labour status and on housing characteristics at the same time. Separately, neither of both allows to study housing and labour questions 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, although they change of residence; meanwhile, the rotatory panels, as LFS and FES, stop interviewing those households that change of address, they are left out of the sample.

Therefore, we need a fixed panel as the ECHP to study labour and residential mobility, and this is the first fixed panel carried out in Spain. 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's member, but housing tenure and mobility decisions are taken by the whole household, mainly by the family head or by the person responsible for the accommodation. So, the sample is formed by household's heads. In order to obtain an homogeneous sample of individuals whose link to the labour market is stable, I restrict to family heads aged from 25 to 64 years old with previous experience.

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

The sample is formed by entrants into unemployment, that is, by those family heads who entered unemployment from January 1993 to December 1996.[3] I consider a residential change occurs when a change of address is produced in some point of the un-

[3]The information on 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 1997, the last wave of the survey.

employment spell until two months after the family head finds a job. That typically happens near the end of the unemployment spell.

An important limitation to study geographical mobility through the ECHP is the reduced geographical division of a residential change, 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. That means we cannot distinguish intra-town from town-to-town residential changes. Given the limited inter-regional mobility observed in Spain and in the other European countries and given the great importance of intra-regional movements nowadays, I decide to consider a case of 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'm aware that this measure can be potentially contaminated with residential changes due to other personal or housing reasons, non related to job. This is one of the reasons to estimate the decision of labour mobility and the housing tenure status jointly, to improve the estimates given that I use an imperfect measure of geographical mobility. An additional motivation to make housing tenure status endogenous to geographical mobility is to allow for the presence of unobserved heterogeneity correlated in both decisions.

Before describing the empirical approach, I will comment some characteristics of the sample used in the estimates. In Table A.1, again we can distinguish two sets of countries according to the housing tenure observed in the sample of unemployment spells: a first group formed by Germany and France, in which rental is the predominant tenure in 64.56% and 62.75%, respectively, of the households whose family head is unemployed. A second group includes Spain and Italy in which 81.56% and 67.13% of the households live in an owned house, whereas almost a half of the households lives in rental in the UK. Column 3 shows the size of geographical mobility 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; 44.27% of the observations corresponds to Spain, which reflects that the Spanish unemployment rate is much higher than in the rest of countries.

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, 75% against 21.43% and 3.57% of owners, respectively, depending on they have outstanding mortgage or not. 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's composition, single individuals not cohabiting in a relation are more mobile, since the 28.57% of them exit to a job spell after a residential change has produced against the 21.03% that found a job in the 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 less than 7 years old, households seem to be more mobile. Regarding the characteristics of previous job, people having worked in services and industry sectors are also more likely to move, whereas those having previous experience in the agriculture and in the construction are almost immobile. Finally, the family head's gender seems not to affect the decision of geographical mobility, and individuals having job tenure longer than two years at previous job are the least likely to move.

2.2 Econometric model of housing tenure and labour mobility

The empirical approach consists in estimating a discrete unemployment duration model with two exits to employment, depending on whether they are associated with a residential change or not. The sample consists in N multiple unemployment spells provided by I family heads, entrants into unemployment for several repeated times. Each spell and its length are denoted by the subscript i and T_i, respectively. As they are fresh spells, I avoid the problem of left-censoring. This 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 as retirement, education, etc.

The econometric method is similar to that used in Bover and Gómez (1999) and Barceló (2001). Let D_i be the indicator taking the value of 1 whether the exit is associated with a residential change, and the value of 0 whether the exit is not joined with a residential change. The transition intensity to employment with the alternative k, 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 alternative k (with residential change or not, $k = 1,\, 0$, respectively) 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 logarithm of duration, including a constant, in order to capture the dependence of duration on transitions[4], household's and family head's characteristics, aggregate economic variables and country dummies reflecting the differences in institutional characteristics across countries.

The estimation method consists in maximising the joint log-likelihood function defined for both transitions[5]. 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 that complete spells contribute to the likelihood in the probability of exit at t months with one of the two alternatives, say k, as follows:

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

Censored spells contribute to the likelihood in the probability of finding a job after having stayed unemployed for over t months:

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

Let c_i be an indicator of lack of censoring of spell i with duration T_i, the joint

[4]Due to the small size of the sample, mainly of exits associated with a residential change, I decide to specify the duration dependence as a second-degree polynomial in the logarithm of duration instead of estimating it semi-parametrically as the addition of duration dummies, as in Meyer (1990).

[5]As analysed in Bover and Gómez (1999), when the transition intensities follow a multinomial logit specification, this estimation method is equivalent to estimate a *competing-risk 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.

log-likelihood function can be expressed as:

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

Bover *et al.* (2001) and Bover and Gómez (1999) explain how to rewrite the log-likelihood function as the concatenation of the log-likelihood functions defined for the survival subsample in each duration. In this way, the estimation of the parameters of the transition intensities is made in a easier way.

Explanatory variables The explanatory variables included in the vector $X_i(t)$ are the following: first, a set of country dummies are introduced in the estimates except for Germany, so that we can include a constant in the polynomial of duration dependence. Later, these institutional differences will be reflected through some indicators of the relevant housing and labour market characteristics affecting the decision of labour mobility. Particularly, I use the information extracted from Maclennan *et al.* (1998) about the percentage of transaction taxes on the 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)[6]. I use this index to classify countries according to their low, medium and high degree of strictness, as shown in Appendix A. The strictness in the employment protection legislation refers to the degree of protection of regular and temporary employment and of collective dismissals. Concerning the regular employment, this index reflects regular procedural inconveniences to the employer for the dismissal as the delay to start of notice 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 reasons why

[6]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 of collective dismissals. As the rest of indices does not vary greatly from one to other date, I have used the latter.

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, 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). About the regulation of collective dismissals, the index of strictness evaluates its 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). Using this index, these five European countries can be classified in the following way: Germany and France are those countries with a medium degree of strictness, Italy and Spain have a great strictness in the employment protection and finally, the United Kingdom shows a very low degree of employment protection.

Second, the only aggregate economic variables considered are the quarterly national unemployment rate. I have not introduced a variable of real house prices due to the difficulty in finding it for all countries. Moreover, I have not included rent consumer price indices due to the fact that these indices refer to different national monetary currencies and quantities, although they are expressed in the same base year. Thus, they cannot be used to make comparisons across countries.

A third set of variables collects information on the family head's previous job. First, there are indicators of the economic sector, only for industry and services sectors; previous jobs in agriculture and in construction are considered as omitted category, given the shortage of observations of exits associated with a residential change for these sectors. Second, I include an indicator of working time, taking the value of 1 whether full-time job, the logarithm of monthly experience at previous job, and its interactions with the highest level of education and 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 also they will capture the effect of entitlement to receive unemployment benefits, since this is mainly determined by the time elapsed at previous job. In this sense, the experience variable will also reflect that: the greater the experience is, the higher the level of unemployment benefits is and the longer the spell of its receipt. Thus, the greater the experience, the higher the reservation wage is; that is why the sign of the coefficient of these indicators are not determined.

I have not included income nor the amount received related to benefits, although this information exists in the ECHP. That 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 of the first panel wave. 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 great measurement errors, and it would be highly endogenous.

Nevertheless, I constructed a dummy indicating whether the individual received any amount due to unemployment benefits during the years in which the unemployment rate is located. The coefficient estimate in exits associated with a residential change was insignificant, but it was positive and significant in exits in the local area. That means that this variable is highly endogenous or that it captures another effect, the stable link of the family heads to the labour market. For that reason, I preferred to remove it.

Finally, a fourth set of variables refers to the personal characteristics such as the level of education, the family head's gender, and the logarithm of family head's age. As indicators of the household's size and composition, I have included a dummy of whether the individual lives with a partner (married or not), an indicator of whether the spouse or the partner is employed, and a dummy pointing out the presence of children in the household interacted with the logarithm of the number of children aged 12 to 18 years old. The latter children variable was introduced under the belief that the presence of children at this range of age would make the household almost immobile, so that the probability of exit to a job spell associated with a residential change would lower significantly.

After having controlled for observed heterogeneity by the previous characteristics, I introduce an indicator taking the value of 1 if the individual is an owner, and the value of 0 if he is a renter, in order to find out whether housing tenure affects the willingness to geographical mobility. Moreover, I include a dummy variable revealing the existence of outstanding loans or mortgages among owners.

As Hughes and McCormick (1981), Henley (1998) and other studies in the case of UK state, renters living in social housing have as low incentives to move as owners, since they will lose their rent control if they move. I am exploring to consider them as an additional housing tenure other than renters; however, that introduces difficult econometric problems in the rest of the paper, mainly in the estimates by controlling for unobserved heterogeneity.

2.3 Estimation results

The parameter estimates of the transition intensities are shown in Table 1. Columns under the heading (i) present estimates 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 associated to each one of both transition intensities to employment, that is, associated with a residential change, $\theta_1 [t \mid X(t)]$, or not, $\theta_0 [t \mid X(t)]$. The explanatory variables, common in both specifications, have their coefficient and t-ratio values very similar to each other, except for the constant and the unemployment rate. For that reason, I will only comment the estimates of specification (i), except for those estimates varying significantly.

First, we observe that the ownership indicator is significant at 1% to explain exits associated with a residential change, $\theta_1 [t \mid X(t)]$; it has a coefficient estimate of -1.43, which means that owners are more reluctant to move to other area to work, and the probability of this exit still falls more for owners having outstanding loans or mortgages. In contrast, in exits to employment in the local area, $\theta_0 [t \mid X(t)]$, the ownership coefficient estimate is near zero with an insignificant value of 0.083, so that it seems to show evidence for the hypothesis that owners are less willing to migrate for job reasons, and they behave in a similar way to renters in the local labour market. Next, this statistical significance of housing tenure must be corroborated by its economic significance in the probabilities predicted by the model, presented in Table 2. In contrast, in the local labour market, unemployed owners having outstanding mortgage leave unemployment with a higher probability than both renters and the rest of owners at 1% of significance, maybe due to the fact that they have to repay debts.

Concerning the country dummies, only Italy and Spain are significant at 5% and 10%, respectively. Both coefficient estimates, -29.838 and -21.278, reflect that the institutional characteristics of their housing and labour markets make individuals much less willing to move than in the rest of European countries. However, in the local labour markets, country dummies only reflect important institutional differences affecting negatively on transitions to a job spell in Italy. Thus, the lower probability of finding a job in Italy, independently of the alternative chosen, may greatly respond to the rigidity of the Italian labour market.

Regarding the economic variables, the exits to a job occurs during the period 1995-96, not covering a complete business cycle for each country. However, the quarterly unemployment rate will not only capture the cross-section variation across countries, but also the business cycle if countries are located in different economic stages. In this sense, this rate is significant to explain exits associated with a residential change at 1%, with a coefficient estimate of -3.542. That reflects that geographical mobility is more likely to occur when the economic conditions are better. The positive and significant coefficient estimate of its interaction with the country dummy of Spain may collect the fact that the level of Spanish unemployment rate is very high in comparison with the rest of countries. That is, the Spanish unemployed will not respond in the same way to the rise or fall of 1% in the unemployment rate. The interaction with Italy is very significant and positive in both exits, producing that the total effect of unemployment rates on exits will be small or positive, it seems that the probability of finding a job in the local labour market increases in economic recessions.

It is expected that, when regional differences in unemployment rates and in other economic variables are bigger within a country, inter-regional migration for job reasons will be observed from depressed to richer regions more often. 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 British and in the Spanish economies using aggregate regional data. I tried to introduce a variable of dispersion in unemployment rates (a proxy of the economic conditions) across provinces, but the results were not satisfactory, mainly for the reason that residential mobility in the sample may occur between shorter distances, not having available data to capture this effect.

Regarding the characteristics at previous job, individuals having worked in the services sector have a lower probability of leaving unemployment than in the rest of sectors, whereas individuals who previously worked in the industry are the most likely to move. The experience attained at previous job is not significant to explain exits associated with a residential change, however, this affects negatively on exits to a job in the local area. The reason is that the predominant effect of these variables is the entitlement to receive unemployment benefits, which raises the reservation wage and decreases the probability of accepting a job. As the unemployment duration is longer, this negative impact on exits to a job spell disappear, indicating the approach to the end of the unemployment benefits receipt. This result has also been found in Bover and Gómez (1999).

The effect of the experience in the labour market is captured through the logarithm of age. The impact on geographical mobility cannot be estimated robustly due to small sample size of these events. 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 to find another job, due to the fact that their knowledge becomes more obsolete, and firms prefer to hire younger people. Finally, the effect of education on the transitions associated with a residential change cannot be estimated accurately as other variables due to the small sample size. Nevertheless, in the local labour market, individuals having completed the third level of education leave unemployment with a higher probability, although this positive influence disappears as the unemployment spells is longer. This latter effect is also captured in transitions associated with a residential change at 10% of significance level.

Concerning the working time, individuals who previously worked full-time have a higher probability of exit from unemployment in their local area. That 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 its job search intensity may be lower.

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 0.359 and it is significant at 1%. That may capture several effects, among them: women may have a higher reservation wage (but not very different since both are family heads) and the fact that women may be discriminated in the sense of receiving job offers at a lower rate or with worse wage conditions. About the household composition, the indicators of living with a partner and the spouse or partner's labour state are not significant to explain any transition, and the variable of children has the expected sign and it is significant.

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

Finally, in specification (ii), the country indicators have been replaced by some indices evaluating characteristics of housing and labour markets. Concerning the degree of strictness in the employment protection legislation, we see that the higher the strictness is in a country, the lower the probability of leaving unemployment. This feature is more stressed in countries with a high protection as the Southern of Europe, Spain and Italy. Another result is, in those countries with high and medium degree of strictness, the probability of exit to employment is less sensitive to the unemployment rate, since the interactions of the unemployment rate with the indicators of medium and high degree of rigidity are both positive and significant.

The logarithm of the percentage that the transaction taxes constitutes on the house price plays the role of proxy of the transaction costs that individuals incur when they buy a house. This measure does not behave as expected, since it has a positive coefficient estimate significant at 1% in exits associated with a residential change. I tried to interact this variable with the ownership indicator in order to capture a different behaviour of owners and renters, but any was significant. In addition, I tried another measure evaluating the transaction costs contained in Maclennan *et al.* (1998), but the results were identical. Perhaps, these transaction costs variables are correlated with other institutional characteristics as the generosity of the unemployment benefit system. For that reason, this latter specification is not attractive, and further work is needed to search another reliable indicators of transaction costs and of social protection system.

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 displays the predicted transition intensity 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 those characteristics. The first row shows these probabilities for the reference person, and the following rows display the probability for an individual with the same characteristics as the reference person's except for the characteristic quoted. 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 rented house in Spain; he is single, aged 30 years old, he 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, and he has previously worked full-time in the industry for 6 months. The aggregate variables are evaluated at their average level in 1995. For the reference person, the probability of transition to employment associated with a residential change is 0.424%, whereas that to a job in the local labour market is 9.63%. Thus, his predicted exit rate is 10.05%, the sum of both probabilities.

Two features are emphasised in Table 2. First, the difference in size of the probabilities displayed in both columns makes obvious that, in all the European countries analysed, the unemployed prefer to search and to accept a job in their local area rather than move to other place to work. Geographical mobility is rarely observed in all countries, although this fact is more stressed in the UK, Germany and Spain, in which the predicted probability for the reference person are 0.143%, 0.293% and 0.424%, respectively. The most mobile countries are France and Italy, whose predicted probabilities for the reference person are 0.815% and 0.564%. The predicted probabilities for Germany are extremely low, mainly in column 2; that is due to the small sample size of the German unemployed family heads (it accounts for 8.96% of the total unemployment spells, see Table A.1) and due to the high proportion of censored spells in it, 72.78% of the German sample, 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 tenant, since this probability falls from 0.424% for a renter to 0.101% in the case of Spain. However, the exit to employment in the local area is not affected by housing tenure, since the predicted probabilities do not vary significantly; 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

Last section considered ownership status as a predetermined variable in the two possible transition intensities to a job spell. However, the analysis carried out in previous section about the effect of housing tenure on the unemployed's decision of labour mobility 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 that purpose, Subsection 3.1 first investigates the main determinants of housing tenure status in the sample of unemployed family heads used in Section 2, regarding differences in the institutional characteristics of the housing and labour markets across European countries. Subsection 3.2 carries out the same analysis as previous section considering 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 hold by the individual i, which takes the value of 1 if he is an owner, and the value of 0 if he is a renter.

The probability of being an owner conditional on a vector of explanatory variables, Z_i, is assumed to follow a logit specification:

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

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 renter.

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

Before showing the estimates of tenure status equation, I will comment the densities of the logarithm of household's income and family head's age, for each type of housing tenure and country, estimated using the sample of unemployment spells. These estimated densities consist in Epanechnikov kernels, in which the bandwidth is chosen optimally; they are displayed in Figures 1 and 2. In Figure 1, age shows us how tenure status depends on the life-cycle. Among the younger, the probability of living in a rented house is the highest. This probability is increasing until the age band of $[30, 33]$, except for Italy, in which this probability continues increasing until the age of 39 years old; then, the probability of renting begins to decrease. Except for Germany, the largest proportion of owners is concentrated on the age range from 35 to 51, when individuals have accumulated enough wealth to invest in an owned house. In Germany, it happens later, about the age of 55 years old. At older ages, the proportion of owners starts to decrease. It is France the country in which the proportion of owners overcomes that of renters at the youngest age, 34 years old, followed by United Kingdom and Spain at the ages of 38 and 39 years old. Italy and Germany are those in which this transition happens much later at the age of 45 and 46 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 more stressed in Germany, France and in the UK, in which 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 in 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 Equation (3.1), 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 of Germany is omitted in order to include a constant in the estimates. In specification (i), France, Germany and the United Kingdom seem to be more similar in their institutional characteristics and in their housing market, since their country dummies are not significant. On the other hand, Spain and Italy seem to form another group of countries in which home ownership is more strongly supported, since their coefficient estimates take the values of 11.866 and of 13.460, respectively, and they are significant at 1%.

The logarithm of age is also significant at 1% to explain the status; its coefficient estimate is 2.048, which points out that the older the family head is, the more likely the household lives in an owned house. This fact is more stressed in France, whose interaction with age is positive and the only significant at 5%, indicating that given an age, the probability of being an owner is higher than in the rest of countries.

In addition to the logarithm of income, its interactions with the country dummies have been included in order to take into account different income distributions across countries. As we expect, the income variable has a positive coefficient estimate, significant at 1%; the higher the level of income, the higher the probability of living in an owned house is. This feature is more stressed in Germany, France and in the United Kingdom, since their interaction with income is not significant, whereas those with Spain and with Italy are significant at 1%, having negative coefficient estimates. Therefore, as we saw in the estimated densities in Figure 2, ownership does not seem to depend on income so strongly in Spain and in Italy, as if other characteristics make the access to a dwelling easier in these countries.

Regarding the household's composition, individuals being married or living with a partner are more likely to live in an owned house. The household's size, measured by the logarithm of the number of children aged 18 years old or less in the household, zero otherwise, affects negatively the probability of living in an owned house. It seems that the larger the household's size is, the smaller the part of income devoted to saving is, that is, the accumulated wealth by the household will be smaller. The level of education could also play the role of a proxy of wealth; thus, we would expect that the higher this level is, the larger the accumulated wealth and the higher the probability of living in an owned house are. As another proxy of wealth, I included an indicator of whether the household owned a second house, but I removed it from the estimates because of its insignificance.

In specification (ii), the country dummies have been replaced by other variables reflecting different characteristics of policies supporting for home ownership in these countries. These three variables have been constructed through the information provided in Maclennan *et al.* (1998). First, a variable consists in the logarithm of the stamp duty incurred in the purchase of a house as a percentage on the house price; a second one is the logarithm of the ratio of social to private rented accommodation prevailing in each country. The last is an indicator taking the value of 1 whether interest tax reliefs are allowed in the income tax due to the repayment of the outstanding mortgage. The weight of stamp duty on house price and the ratio of social to private rented accommodation are expected to affect negatively on the choice of living in an owned house; on the contrary, the interest tax relief will affect positively. This ratio of social to private rented housing is an indicator of the policy maker's encouragement of rental against the promotion the access to an owned dwelling.

In the estimates, all these variables are significant at 1%,with the expected signs. The interaction of income with the logarithm of the ratio social to private rented accommodation is also significant at 1% with a coefficient estimate of 0.253, indicating that owners have a higher level of income in average in those countries in which the social rented accommodation is more strongly supported.

3.2 Unobserved heterogeneity in transitions to employment and in ownership status

This subsection estimates specifications of this ownership 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 the ownership status.

Ownership status may be positively correlated with unobserved human capital and ability that increase the individuals' wealth accumulation. In addition, human capital contributes to increase the probability of transitions to employment, not only in the local labour market, but also in other geographical places. Thus, the coefficient of ownership status in exits associated with a residential change may be biased upwards spuriously if owners are higher skilled than renters in 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, in average, have a less favourable unobserved ability to abandon unemployment than owners. That will cause that it is 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.

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. Further research is needed to find out whether adding additional mass points improves the coefficient estimates and the log-likelihood.

Given 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).[7] As I have repeated spell durations for a non-negligible proportion of individuals (24.4%), allowing for unobserved heterogeneity becomes more advisable.

The unobservable are assumed to be independent of the rest of explanatory variables, X_i and Z_i, other than owner status, h_i. For that reason, I will redefine the vector of characteristics relevant in the duration model as $(X_i,\ h_i)$ in order to consider separately the owner status from the rest. Let η_i be the permanent unobserved effect on both transitions to employment and owner status equation for the individual i. Thus, this unobserved effect takes a value of the support $\{m_1,\ m_2\}$ with a probability 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 unobservable is reduced to estimate a mass point, say $m_1 = m$, and its associated probability, $p_1 = p$, $0 < p < 1$.[8]

The transition intensity to employment with the alternative k, $k = 0, 1$, and the owner status equation conditional on the unobserved heterogeneity and the observable

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

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

are specified, respectively, as follows:

$$\theta_k \left[t \mid X_i(t),\, h_i,\, \eta_i \right] \;=\; \frac{\exp(X_i(t)'\beta_k + h_i\beta_{hk} + \alpha_k\eta_i)}{1 + \sum_{j=0}^{1} \exp(X_i(t)'\beta_j + h_i\beta_{hj} + \alpha_j\eta_i)}, \quad k = 0, 1 \tag{3.2}$$

$$\Pr\left(h_i = 1 \mid Z_i, \eta_i\right) \;=\; \Lambda\left(Z_i'\delta + \alpha\eta_i\right) = \frac{\exp\left(Z_i'\delta + \alpha\eta_i\right)}{1 + \exp\left(Z_i'\delta + \alpha\eta_i\right)} \tag{3.3}$$

The unobserved heterogeneity in this model follows a one-factor structure. So, the permanent individual effect in each one of the three equations is 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. Thus, the parameters related to the unobservable that have to be estimated are $m,\ p,\ \alpha_0$ and α, if α_1 is normalised to 1.

The unobserved individual effect η is invariant in all the unemployment spells and housing tenure status 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 as in Section 2. Under the absence of unobserved heterogeneity, the duration model and the housing tenure status can be estimated separately by maximum likelihood. However, when controlling for unobserved heterogeneity, the joint log-likelihood cannot be split in the addition of both log-likelihoods defined separately.

Now, the sample is regrouped in $i = 1,\ 2, \ldots, I$ individuals who have been unemployed for $n_1,\ n_2, \ldots,\ n_I$ times, respectively each one of them; 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 one of the n_i unemployment spells in which the individual i has stayed. 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 by each individual i as follows: $X_i = \left[X_i\left(t_{i1}\right),\ X_i\left(t_{i2}\right), \ldots,\ X_i(t_{in_i})\right]'$, $Z_i = \left(Z_{i1},\ Z_{i2}, \ldots,\ Z_{in_i}\right)'$, $D_i = \left(D_{i1},\ D_{i2}, \ldots,\ D_{in_i}\right)'$ and $h_i = \left(h_{i1},\ h_{i2}, \ldots,\ h_{in_i}\right)'$.

Then, conditional on the unobserved heterogeneity and observable, the individual i contributes to the likelihood in the probability of observing that sequence of unemployment spell durations, exits to employment and housing tenure status in each one of the spells. Using the definition of transition intensities and Equations (3.2) and (3.3), this contribution can be factorised as follows :

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

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

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

$$\cdot \prod_{s=1}^{t_{ij}-1} \left(1 - \theta\left(s \mid X_i(s), \eta_i\right)\right) \Lambda\left(Z'_{ij}\delta + \alpha\eta_i\right)^{h_{ij}} \left[1 - \Lambda\left(Z'_{ij}\delta + \alpha\eta_i\right)\right]^{h_{ij}} \tag{3.4}$$

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:

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

$$= \sum_{i=1}^{I} \log \left\{ \sum_{l=1}^{2} \left(\prod_{j=1}^{n_i} \Pr\left(t_{ij}, D_{ij}, c_{ij} \mid h_{ij}, X_i(t_{ij}), Z_{ij}, \eta_i = m_l\right) \Pr\left(h_{ij} \mid Z_{ij}, \eta_i = m_l\right) \right) p_l \right\} \tag{3.5}$$

A simplified version of the duration model in Section 2 and the ownership status equation has been estimated controlling for unobserved heterogeneity. The results are found in Table 4. Most of the significant variables do not change a lot in their coefficient size nor significance when we control or not for unobserved heterogeneity. However, the estimated coefficient of the ownership status in exits associated with a residential change becomes more negative, from -1.388 to -2.404, giving evidence for the hypothesis that unobserved human capital variables could bias upwards this coefficient. However, the same variable in exits to a job in the local area becomes significant with a coefficient estimate of 0.734, indicating that this specification of heterogeneity does not capture all the unobservable, since owners could enjoy higher arrival rates of a job offer or better-paid job offers due to their human capital.

Concerning on age, after controlling for unobserved heterogeneity, older family heads have more difficulties in leaving unemployment. In the same way, the ownership status becomes much more dependent of the family head's age and of the household's total income, since their coefficient estimate have almost doubled.

Moreover, when unobserved heterogeneity is allowed for, the ownership status responds more intensively to housing policies which encourage home ownership as the existence of interest tax relief in income taxes, and the decrease in the stamp duties to pay for the purchase of a house. However, it becomes less sensitive to the ratio of social to private rented accommodation prevailing in the country.

Finally, after controlling for unobserved heterogeneity, the level of education becomes irrelevant to explain the ownership status.

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 concerns in the main determinants of the wage level obtained by individuals just after leaving unemployment, as in other studies as Addison and Portugal (1989). In the first place, I want to study whether this level depends on the length of the unemployment spell, thus 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 obsoleteness of the unemployed worker's knowledge, and second, the firms' reluctance to hire individuals having been unemployed for a long time, since firms think that long unemployment spells are a bad signal about their productivity (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 knowing whether those individuals who find a job associated with a residential change obtain a higher wage in average than those who exit from unemployment in their local area. The economic theory does not provide a clear prediction about the gain in wage 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 their origin region has, their reservation wage of moving will be comparatively lower than that for accepting a job in their local area. That will happen if moving costs are enough low. Individuals will have a higher reservation wage of 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 impact 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 been 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 average of the potential wages distribution. This distribution will depend on the regional characteristics as well as others related to job, 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. That means that a problem of self-selection bias arises, since individuals decide to remain unemployed and not to accept the wage offers possibly received 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[9].

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

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

equations (2.1) and (2.3); thus, 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 next specification, 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, in spite of the fact that this includes variables related to the unemployment duration.

The individual i will leave unemployment with the alternative k, $D_i = k,\ k = 0,\ 1$ if this provides him with the maximum level of 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, $\overline{w}_{ki} \geq 0$:

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

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

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

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 can consist in some characteristics as age or education, common to those in X_i. The distribution assumptions of the disturbance, u_i, are as follows:

$$\begin{gathered} u_i = \rho_0\varepsilon_{0i} + \rho_1\varepsilon_{1i} + v_i, \\ E\left[v_i \mid D_i,\ Z_i,\ T_i,\ X_i\right] = 0, \\ cov(v_i,\ \varepsilon_{ki}) = 0,\ k = 0, 1 \end{gathered} \tag{4.4}$$

The disturbance can be broken down as the sum of two terms: a first term, v_i, is uncorrelated with the explanatory variables of 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 wage offer comparison equations with the reservation wage.

In order to obtain an expression of the self-selection bias arising from Equation (4.3) when we estimate post-unemployment wages using only the subsample of individuals leaving unemployment, it is useful to define indicators A_i, A_{1i}, and A_{0i} for each 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 value of 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: on the one hand, 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 conditioned only on the subsample of individuals leaving unemployment, that is, on those individuals whose wage is observed, $A_i = 1$.

To study the average of potential wages, we can carry out a two-step procedure similar to the Heckman's two-step estimator[10], although 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:

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

By using the definitions in Equation (4.5) and the distribution function described in

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

Equation (4.1), we obtain that the probability terms in Equation (4.7) have the following form:

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

The average 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 by operating Equations (4.1), (4.4) and (4.5):

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

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:

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

In conclusion, in order to control for self-selection using an estimator similar to the Heckman's two-step estimator, the procedure will consist in estimating the duration model described in Section 2 in a first stage. In a second stage, I will have to 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, since they are treated as if they were the true values. Thus, 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 that 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 wages distribution conditional on the subsample of unemployed individuals entering a job spell, that is, conditional on $A_i = 1$ 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, over 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 are $2p + r + 4$, 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, in the following criterion function is irrelevant:

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

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

$$\xi_i = w_i - \alpha D_i - Z_i'\delta - \pi\log T_i - \rho_1\mu_{1i}(X_i;\,\beta) - \rho_0\mu_{0i}(X_i;\,\beta)$$
$$A = I_{(2p+r+4)}$$

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 in Arellano and Honoré (2000) 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 the new job and personal characteristics. First, country indicators are introduced except for Germany. As relevant characteristics of the new job, I have used an indicator whether the working time is full-time, and dummies of the firm size, considering a small firm if the number of workers is lower than 100, a medium firm if this number ranges from 100 to 499 workers, and a large firm if the number of workers is greater than 500. I have also included a dummy variable taking the value of 1 whether the firm belongs to the private sector and indicators of firm's economic activity according to NACE 2 digits, in which the omitted category is A+B+C+E, corresponding to the primary sector, mining and quarrying, electricity, gas and water supply.

The personal characteristics used are: the logarithm of duration in the unemployment spell, the logarithm of age, the indicator of whether the individual is male, and dummies pointing out the level of education, using levels lower than the second stage of secondary education as omitted category. Finally, I have included a dummy indicating whether the exit to employment is associated with a residential change, defined in Section 2.

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

The results shown in column (ii) correspond with the two-step estimator implied by the unemployment duration model with two exits to employment. Finally, column (iii) reports the results of the wage equation, jointly estimated with the duration model by GMM in order to obtain more efficient estimates than the two-step estimator.

After applying a series of filters to data, constructing the unemployment duration variable and assigning the explanatory variables to each duration, I have obtained a number of 1764 spells for these five countries, in which the proportion of complete spells is 56.18%. Thus, I should have a total of 991 wage observations from individuals having exited to employment. However, I have only been able to assign a wage to 612 individuals or spells, who represent the 61.76% of the complete unemployment spells.

The estimates of the three specifications are very similar in the size of the coefficients and in the significance of the explanatory variables. It seems not to exist big wage differences across countries after having allowed for job and personal characteristics. In specification (iii), the logarithm of duration has an estimated coefficient of -0.013, with the expected sign, but insignificant. That may be due to two reasons: first, the reduced sample size and the high proportion of right-censored spells make that this dependence cannot be estimated accurately. Second, this sample only consists of entrants into unemployment and they are followed in a short time period, hardly a year for most of them (for the 79.74% of the observations); thus, the sample has very short durations without enough variation to capture this dependence on wages.

Regarding the indicator of whether the exit to employment is associated with a residential change, in all specifications, its coefficient estimate is negative, but insignificant. That may be caused by the small sample size and the tiny frequency of residential mobility. In addition, there can exist opposite effects making this coefficient be bias to zero, as I have commented at the beginning of this section.

In conclusion, it is necessary to have available future waves of the ECHP for estimating the effects of both unemployment duration and exits associated with a residential change.

Concerning the personal and job characteristics, most of them are significant with the expected effects. The logarithm of age has a positive effect on wages that individuals obtained just after leaving unemployment, this variable captures the impact of the experience in the labour market on wages. Its coefficient estimate is higher in specifications controlling for self-selection bias. A possible explanation may be that: if older family heads have more difficulties in leaving unemployment due to the obsoleteness of their knowledge, this may influence negatively on wages that they obtained after leaving unemployment. Then, the impact of experience on the labour market will be biased downwards in OLS estimates, due to the fact that we do not take into account that the probability of leaving unemployment and finding better-paid offers may be positively correlated. I tried to introduce an interaction of age with unemployment duration, but it was not significant.

Other expected results are that individuals working full-time earn in average a higher monthly wage that those working part-time; however, wage differences are not observed between individuals working in private or public sector. With respect to the firm size, individuals working in medium firms are paid higher wages than in small firms, since their coefficient estimate in specification (iii) is 0.249 and significant at 1%. However, the indicator of large firm size is near zero and not significant in all specifications. I have tried to include interactions of firm size with level of education in order to capture different ways of remuneration of firms to the level of education according to their size. That is, individuals having a low level of education will not differ in their earnings so much if they work in a small or large firm; however, high skilled individuals are expected to be better-paid in large than in small size firms. Nevertheless, these interactions were not significant and I decided to remove them. As the sample consists in unemployed individuals, and unemployment is expected to affect more seriously to low skilled individuals, the composition of the sample may make difficult to identify this effect. The lowest skilled workers represent the 70.10% of the sample.

About the firm's or unit's economic activity, sectors as transport, storage and communication (I), hotels and restaurants (H), education (M), manufacture of metal products, machinery and equipment (DJ+DK), and the public administration and defense (L) seem to pay relatively more than the omitted category, the primary sector, mining, electricity, gas and water supply (A+B+C+E). Appendix A details the correspondence of this code with the economic activity.

The indicators of education level are significant at 1% and 10% to explain the wage level achieved just after leaving unemployment, and they have a positive effect on wages.

Finally, the coefficient estimates of the terms controlling for the self-selection bias, ρ_1 and ρ_0, are -0.010 and 0.139, respectively, but they are not significant. Both coefficient signs would be consistent with the idea that the unemployed have not found a job in their local due to the fact that their reservation wages are high, so that the average of the potential wage distribution is over that of accepted wages. However, transitions to a job spell associated with a residential change may be caused by high arrival rate of job offers and ability.

5 Conclusions

The purpose of this paper consists in studying three questions: first, how housing tenure, ownership or rental, influences the unemployed's decision of geographical mobility using a discrete model of unemployment durations with two 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 potential endogeneity of the ownership status in the transition intensities. For that purpose, I investigate the main determinants of housing tenure status, taking into account differences in policies supporting for the access to a dwelling across countries, since this ownership status equation will be used to estimate the unemployment duration model controlling for unobserved heterogeneity Third, I study the main factors influencing the wage level that individuals obtain just after leaving unemployment. In particular, I try to find out whether the unemployment duration affects negatively wage levels, and whether the unemployed 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 that purpose, I use individual data coming from the *European Community Household Panel* (ECHP) for the four currently available waves covering the period 1994-97. As the surveys carried out in each country are homogeneous, I can make comparisons among them.

The sample I use in all estimates consists in the spells of the unemployed family heads, aged from 25 to 64 years old with previous experience and coming from five European countries: Spain, Germany, France, Italy and the United Kingdom. An important limitation to study geographical mobility using the ECHP is the impossibility of differentiating the cases of geographical mobility from the other of residential mobility perfectly. Only I can know that these residential changes produce within the same province, from other province within the same country and from other country. Given the limited number of inter-regional movements observed in all European countries, I identify a case of geographical mobility as a residential change having taken place in some point of the unemployment spell and having been followed by an exit to employment (or preceded by an exit to employment until two months before the residential change). In spite of this problem, I have decided to make this study using the ECHP, since this is the only fixed panel available in Spain to study questions of housing tenure, residential mobility and labour mobility.

I have obtained the following results: concerning the effect of housing tenure on labour mobility, I found that home owners are more reluctant to accept a job associated with a residential change than renters, since the ownership indicator has a very negative coefficient estimate, and it is significant at 1%. On the contrary, housing tenure does not affect the unemployed's performance in the local labour market; that is, the reservation wage for job offers coming from the local area in which they live does not depend on housing tenure. However, owners having outstanding mortgage leave unemployment with a higher probability in exits not associated with a residential change; this type of owners seems to search a job more intensively, since they have to face debts. As observed in other studies, geographical mobility seems not to respond appropriately to regional differences in unemployment rates. Finally, concerning the probabilities predicted by the model, we can see that the unemployed's geographical mobility in all these countries is very small.

Concerning the main determinants of housing tenure status, living in an owned house depends on age in the sense predicted by the life-cycle theory: when individuals are young, 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 to home ownership and finally, when individuals are elder, the proportion of individuals living in an owned house starts to fall. The level of household's total income affects positively the probability of being an owner. This dependence on income is more stressed in Germany, France and United Kingdom; whereas the access to a dwelling does not depend on income so much in Spain and in Italy, as if other institutional characteristics make this access easier. When the country indicators are replaced by other indicators evaluating the policy of support for home ownership in the estimates, we observe that ownership status responds negatively to the weight of stamp duties incurred in the purchase of a dwelling, and to the ratio social to private rented accommodation, and positively to the existence of interest-tax relief on income tax.

When we make the ownership status endogenous to the transition to employment due to the presence of unobserved human capital, the results are reinforced. The ownership coefficient in exits associated with a residential change becomes more negative and significant. Further research is needed to introduce more mass points in the specification of the unobserved heterogeneity as well as to consider an additional housing tenure, renters living in a social housing, in order to reinforce the results about the reluctance of owners and renters having rent control to move.

Finally, the wage level that individuals obtained just after leaving unemployment depends on personal and job characteristics in the way we would expect, except for the indicator of large firm size and that of whether the firm operates in the private sector, since they are not significant. In order to estimate the average of the potential wage distribution, we have to control for the possible presence of a self-selection bias. This arises from the reason that a subsample of individuals decide to remain unemployed and reject all the possible offers received being lower than their reservation wage. The corrections are similar to the Heckman's two-step estimator, but they have been derived from a duration model with multinomial logit transitions. As this two-step estimator does not take into account the inclusion of predicted regressors, I estimate the post-unemployment wage equation jointly with the unemployment duration model with multiples exits by GMM in order to obtain more efficient estimates. However, given the small sample size of post-unemployment wages, the effect of the length of the unemployment spell and the impact of geographical mobility are not captured accurately. Future waves of the ECHP are crucial for studying them.

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A Construction of variables

A.1 Personal and household's characteristics

The data come from the *European Community Household Panel* (ECHP) for the period 1994-97, provided by *Eurostat.*

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

- Employed, whether full-time or part-time, paid apprenticeship or training under special schemes relative to employment, self-employment, unpaid work in a family enterprise, in education or training, retired, in community or military service and other economically inactivity different from doing housework, looking after children or other persons, without previous experience in a job, and having rent-free accommodation.
- Missing observations in the following variables: age, sex, cohabitation in a relation, lack of personal information on spouse or partner, education, household composition, year of move to the address, housing tenure and where individuals move from.

Duration: the sample consists in entrants into unemployment. Durations are constructed using the indicators of the individuals' main economic activity carried out in each month of the year prior to the survey. Thus, these durations start from January 1993 to December 1996.

Experience: constructed by the addition of the monthly indicators of the main economic activity referred to employed individuals from January 1993 to December 1996. For individuals being employed in January 1993, the experience variable is completed by multiplying 12 times the number of years worked at the current job in January 1993. This number is obtained through the variable of year of start at current job.

Economic sector and working time at previous job, working time, firm's size, firm's business sector (whether private or public) and economic activity according to NACE 2 digits in the job obtained just after leaving unemployment: constructed by matching the information available for current and last jobs in each wave of the survey according to the year and month of current job and the year and month of last job in the four waves.

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 household's total income received at previous year after having been made homogeneous using the purchasing power parity (PPP).

Residential change during the unemployment spell: constructed using information on year and on month from which the individual lives in the current address and using information on where the previous dwelling was located.

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

Economic activity according to NACE 2 digits: A+B+C+E: 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 off 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.

A.2 Aggregate variables

Purchasing power parity: provided by *Eurostat* for the period covered by the survey.

Quarterly national unemployment rate: coming from *"Main Economic Indicators"* from OECD.

Indices: the index of strictness in the employment protection legislation (EPL) is extracted from *"OECD Employment Outlook"*, June (1999). Taxes paid as % of the house price, % of stamp duties on house price, existence of interest tax relief, and the ratio social to private rented housing extracted from Maclennan *et al. (1998).*

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

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

EPL strictness: this index takes a value of the range [0,6], and the higher it is, the stricter the employment protection legislation is. It is aggregated according to three levels of strictness: low degree if the index of EPL strictness is smaller than 1.5, medium degree if the index takes a value of the interval [1.5,3] and high degree of strictness in EPL if this index is greater than 3.

Ln(stamp duty): constructed as

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

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

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

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

	Housing tenure		Geographical	Percentage of
	Ownership	Rental	mobility	unemployment spells
Germany	35.44	64.56	2.53	8.96
France	37.25	62.75	3.02	16.89
United Kingdom	57.14	42.86	0.60	9.52
Italy	67.13	32.87	1.39	20.35
Spain	81.56	18.44	1.15	44.27

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

	Non geographical mobility	Geographical mobility
Housing tenure		
Owner without outstanding loan	42.05	21.43
Owner with outstanding mortgage	23.27	3.57
Renter	34.68	75.00
Education		
Third level	8.70	7.14
2nd. level secondary	24.48	35.71
Less than 2nd level secondary	66.82	57.14
Living with a partner		
No	21.03	28.57
Yes:		
Spouse/partner employed	29.21	21.43
Spouse/partner not employed	49.77	50.00
Gender		
Male	85.02	82.14
Female	14.98	17.86
Children aged		
$[0, 6]$ years old	29.21	42.86
$[7, 11]$ years old	25.35	21.43
$[12, 18]$ years old	34.50	39.29
Economic sector		
Agriculture	11.06	3.57
Industry	19.35	25.00
Services	47.75	64.29
Construction	21.83	7.14
Experience at previous job		
≤ 12 months	58.93	60.71
$[13, 48]$ months	22.35	28.57
> 48 months	18.72	10.71

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

	Total	Germany	France	United Kingdom	Italy	Spain
Censored	43.82	72.78	35.23	61.31	35.38	41.36
Completed	56.18	27.22	64.77	38.69	64.62	58.64
Number of spells	1764	158.00	298	168	359	781

Duration of the unemployment spell	Percentage	Duration of the unemployment spell	Percentage
1 month	14.68	23 months	0.51
2 months	13.83	24 months	2.49
3 months	10.77	25 months	0.85
4 months	6.58	26 months	0.40
5 months	5.39	27 months	0.85
6 months	6.52	28 months	0.57
7 months	3.85	29 months	0.28
8 months	3.40	30 months	0.40
9 months	3.00	31 months	0.28
10 months	1.87	32 months	0.28
11 months	1.36	33 months	0.23
12 months	7.77	34 months	0.40
13 months	1.42	35 months	0.23
14 months	1.70	36 months	0.74
15 months	1.42	37 months	0.17
16 months	1.42	38 months	0.17
17 months	1.19	39 months	0.11
18 months	0.74	40 months	0.11
19 months	1.02	42 months	0.11
20 months	0.57	43 months	0.11
21 months	0.85	45 months	0.17
22 months	1.13	46 months	0.06

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)				(ii)			
	$\theta_1(t \mid X_i(t))$		$\theta_0(t \mid X_i(t))$		$\theta_1(t \mid X_i(t))$		$\theta_0(t \mid X_i(t))$	
	Coefficient	t-ratio	Coefficient	t-ratio	Coefficient	t-ratio	Coefficient	t-ratio
Country dummies								
France	-8.838	-0.49	4.080	0.70				
United Kingdom	-0.358	-0.03	9.657	1.72				
Italy	-29.838	-1.88	-25.349	-4.12				
Spain	-21.278	-1.32	1.107	0.21				
Indices								
Ln(taxation tax (%))					3.884	2.59	2.179	6.19
EPL strictness:								
Medium degree					-12.810	-1.39	-9.428	-3.17
High degree					-36.088	-8.81	-18.075	-7.61
Economic variables								
Unemployment rate	-3.542	-2.75	-0.810	-1.51	-4.234	-10.42	-2.083	-7.12
Un. rate*France	1.515	0.87	-0.089	-0.15				
Un. rate*UK	-0.702	-0.50	-1.277	-2.08				
Un. rate*Italy	3.292	2.08	2.433	3.95				
Un. rate*Spain	3.017	2.18	0.454	0.84				
Unemp. rate*medium EPL					1.823	1.68	1.089	2.99
Unemp. rate*high EPL					4.084	9.91	1.980	6.76
Ownership	-1.430	-2.83	0.083	0.86	-1.424	-2.82	0.082	0.86
Outstanding mortgage	-1.199	-1.09	0.224	2.34	-1.185	-1.09	0.227	2.38
Previous job characteristics								
Economic sector								
Industry	1.335	1.79	-0.118	-1.17	1.350	1.82	-0.146	-1.46
Services	0.857	1.27	-0.375	-4.26	0.845	1.25	-0.394	-4.52
Working time								
Full-time job	-0.040	-0.07	0.207	1.59	-0.032	-0.06	0.218	1.67

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

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

	(i)				(ii)			
	$\theta_1(t \mid X_i(t))$		$\theta_0(t \mid X_i(t))$		$\theta_1(t \mid X_i(t))$		$\theta_0(t \mid X_i(t))$	
	Coefficient	t-ratio	Coefficient	t-ratio	Coefficient	t-ratio	Coefficient	t-ratio
Experience								
Ln(experience)	0.029	0.08	-0.165	-3.75	0.019	0.05	-0.168	-3.78
Ln(experience)*ln(duration)	-0.027	-0.16	0.041	1.56	-0.021	-0.13	0.046	1.73
Personal characteristics								
Education								
Third level	1.302	1.18	0.588	2.64	1.295	1.20	0.614	2.75
Third level*ln(duration)	-0.713	-1.56	-0.230	-1.60	-0.716	-1.58	-0.238	-1.65
2nd stage secondary	0.437	0.90	0.078	0.84	0.443	0.91	0.078	0.85
Logarithm of age	-0.445	-0.59	-0.843	-5.42	-0.460	-0.62	-0.856	-5.53
Male family head	0.350	0.57	0.359	2.73	0.342	0.57	0.370	2.80
Living with a partner	0.367	0.71	0.038	0.33	0.370	0.71	0.047	0.40
Spouse/partner employed	-0.382	-0.75	-0.022	-0.27	-0.385	-0.76	-0.036	-0.43
Ln(Number of children aged $[0, 18]$*)*	0.041	0.11	0.156	2.11	0.047	0.12	0.144	1.97
Duration dependence								
Ln(duration)	0.034	0.04	0.031	0.24	-0.025	-0.03	-0.004	-0.03
Ln^2(duration)	0.078	0.38	-0.194	-5.06	0.092	0.44	-0.182	-4.74
Constant	27.400	2.14	7.627	1.49	27.058	5.60	17.300	7.22
Log-likelihood	-3489.32				-3488.94			
Number of spells	1764				1764			

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 taxation taxes in the purchase as % of the house price and the strictness of the employment protection legislation (EPL).

Table 2: Predicted probabilities (%) of multinomial logit transitions to employment.

Specification (i): Country dummies.

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

	Transitions to a job spell associated with	
	Residential change (θ_1)	Non residential change (θ_0)
Reference person	0.424	9.63
Countries		
Germany	0.293	4.81
France	0.815	15.70
United Kingdom	0.143	9.19
Italy	0.564	14.01
Ownership		
Non outstanding loans	0.101	10.41
Outstanding mortgage	0.030	12.70
Previous job characteristics		
Economic sector		
Agriculture/Construction	0.110	10.74
Services	0.269	7.63
Working time		
Part-time job	0.449	7.97
Experience attained		
12 months	0.424	8.99
24 months	0.424	8.39
51 months	0.425	7.78
Personal characteristics		
Education		
Third level	0.562	12.22
2nd stage secondary	0.649	10.31
Female family head	0.308	6.94
Age		
25 years old	0.452	11.05
45 years old	0.364	7.04
55 years old	0.337	6.01
Living with a partner		
Spouse/partner not employed	0.608	9.95
Spouse/partner employed	0.417	9.77
No. of children aged $[0, 18]$ *years old*		
1 child	0.424	9.63
2 children	0.431	10.61
3 children	0.435	11.22

Notes: The reference person is a male family head living in a rented house in Spain. He is single, aged 30 years old, he 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, and he has previously worked full-time in the industry for 6 months. The aggregate variables are evaluated at their average in 1995.

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

$$\Pr(h_i = 1 \mid Z_i) = \frac{\exp(Z_i'\delta)}{1+\exp(Z_i'\delta)}$$

	(i)		(ii)	
	Coefficient	t-ratio	Coefficient	t-ratio
Countries				
France	-0.851	-0.15		
United Kingdom	5.309	0.93		
Italy	13.460	2.87		
Spain	11.866	2.63		
Support for home ownership				
Ln(stamp duty)			-0.762	-6.62
Interest tax relief			1.929	8.16
Ln(ratio social to private rented housing)			-3.263	-4.14
Household's income (PPP)				
Ln(income)	1.496	3.93	0.676	4.66
Ln(income)*France	-0.613	-1.28		
Ln(income)*UK	-0.699	-1.45		
Ln(income)*Italy	-1.354	-3.36		
Ln(income)*Spain	-1.330	-3.37		
Ln(income)*ln(ratio social to private rented housing)			0.253	3.05
Family head's age				
Ln(age)	2.048	2.42	3.030	11.20
Ln(age)*France	2.045	1.91		
Ln(age)*UK	0.799	0.71		
Ln(age)*Italy	0.390	0.39		
Ln(age)*Spain	1.016	1.05		
Male family head	0.016	0.09	-0.032	-0.18
Ln(No. children aged $[0, 18]$ *years old)*	-0.391	-2.85	-0.372	-2.73
Living with a partner	0.603	3.41	0.691	4.02
Family head's education				
Third level	0.062	0.27	0.072	0.32
2nd stage secondary	0.478	2.99	0.443	2.85
Constant	-23.601	-5.73	-19.039	-10.76
Log-likelihood	-889.38		-897.30	
Number of spells	1764		1764	

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 four indicators evaluating the support for home ownership.

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

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

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

	No unobserved heterogeneity		Unobserved heterogeneity	
	Coefficient	t-ratio	Coefficient	t-ratio
Exit to a job associated with a residential change:				
France	0.529	0.63	0.598	0.70
United Kingdom	-0.926	-0.15	-0.708	-0.12
Italy	0.657	0.78	1.278	1.16
Spain	0.538	0.68	1.278	1.08
Ownership	-1.388	-1.95	-2.404	-2.05
Outstanding mortgage	-1.176	-0.19	-1.016	-0.17
Ln(age)	-0.487	-0.39	0.021	0.02
3rd level education	1.172	0.42	1.240	0.38
3rd level *ln(duration)	-0.717	-0.36	-0.724	-0.33
2nd stage secondary	0.346	0.53	0.399	0.58
Ln(experience)	-0.013	-0.04	-0.018	-0.05
Ln(experience)*ln(duration)	-0.015	-0.10	-0.006	-0.03
Industry	1.282	1.18	1.240	1.09
Services	0.771	0.74	0.770	0.69
Ln(duration)	-0.012	-0.01	-0.064	-0.06
Ln^2(duration)	0.085	0.35	0.083	0.33
Constant	-5.209	-1.14	-6.976	-1.50
Exit to a job not associated with a residential change:				
France	1.091	5.60	1.039	5.05
United Kingdom	0.417	1.84	0.308	1.31
Italy	1.192	6.03	0.973	4.54
Spain	0.930	4.77	0.667	3.11
Ownership	0.084	0.92	0.754	3.21
Outstanding mortgage	0.273	3.17	0.207	2.39
Ln(age)	-0.872	-5.87	-1.095	-6.90
3rd level education	0.536	2.42	0.481	2.12
3rd level *ln(duration)	-0.204	-1.39	-0.189	-1.26
2nd stage secondary	0.070	0.78	0.051	0.56
Ln(experience)	-0.165	-4.27	-0.162	-4.15
Ln(experience)*ln(duration)	0.049	1.97	0.047	1.88

Table 4: Joint estimates of multinomial transitions to employment with the ownership status equation. (Contd.)

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

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

	No unobserved heterogeneity		Unobserved heterogeneity	
	Coefficient	t-ratio	Coefficient	t-ratio
Industry	-0.206	-2.33	-0.205	-2.29
Services	-0.473	-5.91	-0.465	-5.71
Ln(duration)	-0.031	-0.26	-0.018	-0.16
Ln^2(duration)	-0.178	-4.87	-0.175	-4.81
Constant	0.639	1.07	1.191	1.93
Ownership status equation:				
Interest tax relief	1.936	8.56	4.843	6.56
Ln(stamp duty)	-0.762	-7.26	-2.191	-5.58
Ln(ratio social to private rented accommodation)	-3.290	-4.56	-2.945	-1.63
Ln(income)	0.685	5.57	1.319	4.77
Ln(income)*ln(ratio soc./priv.)	0.255	3.35	-0.049	-0.23
Ln(age)	2.977	13.57	5.446	7.84
No. children aged $[0, 18]$	-0.161	-3.93	-0.119	-1.00
Living with a partner	0.717	5.49	0.994	2.99
3rd level education	0.039	0.20	0.130	0.27
2nd stage secondary	0.434	3.33	0.243	0.85
Constant	-18.903	-12.51	-36.239	-8.47
Unobserved heterogeneity parameters:				
α_0			-0.663	-0.92
α			8.514	0.99
m			0.310	0.95
p			0.736	38.19
Number of individuals	1297		1297	
Number of spells	1764		1764	
Log-likelihood	-4435.88		-4274.50	

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

$$w_i = \alpha D_i + Z_i'\delta + \pi \log T_i + \rho_0\mu_0(X_i;\beta) + \rho_1\mu_1(X_i;\beta) + \xi_i \quad \text{for } w_i > 0$$

	(i) OLS estimates	(ii) Two-step estimates	(iii) GMM joint estimates
Countries			
France	−0.030 (−0.24)	(−0.081) (−0.52)	(−0.065) (−0.44)
United Kingdom	−0.077 (−0.39)	(−0.079) (−0.39)	(−0.064) (−0.33)
Italy	−0.165 (−1.13)	(−0.219) (−1.18)	(−0.190) (−1.07)
Spain	−0.053 (−0.41)	(−0.081) (−0.52)	(−0.071) (−0.47)
Job characteristics			
Economic activity (NACE 2)			
DA	0.153 (0.82)	(0.157) (0.84)	(0.119) (0.66)
DB+DC	0.228 (1.65)	(0.213) (1.54)	(0.178) (1.34)
DD+DE	−0.155 (−0.64)	(−0.153) (−0.64)	(−0.186) (−0.80)
DF-DI	0.248 (1.05)	(0.261) (1.10)	(0.229) (1.00)
DJ+DK	0.318 (2.05)	(0.358) (2.20)	(0.326) (2.06)
DL-DN	−0.247 (−0.81)	(−0.226) (−0.73)	(−0.260) (−0.87)
F	(0.169) (1.64)	(0.175) (1.69)	(0.142) (1.48)
G	(0.292) (2.23)	(0.332) (2.29)	(0.294) (2.11)
H	(0.513) (4.27)	(0.540) (4.21)	(0.504) (4.18)
I	(0.569) (4.55)	(0.590) (4.57)	(0.555) (4.54)
J	(0.312) (0.89)	(0.346) (1.02)	(0.312) (0.95)
K	(0.243) (1.65)	(0.271) (1.71)	(0.235) (1.57)
L	(0.292) (1.71)	(0.308) (1.72)	(0.282) (1.65)
M	(0.494) (1.83)	(0.522) (1.90)	(0.495) (1.86)
N	(0.009) (0.03)	(0.035) (0.12)	(0.005) (0.02)
O-Q	(0.342) (2.71)	(0.336) (2.53)	(0.301) (2.40)

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

$$w_i = \alpha D_i + Z_i'\delta + \pi \log T_i + \rho_0\mu_0(X_i;\beta) + \rho_1\mu_1(X_i;\beta) + \xi_i \quad \text{for } w_i > 0$$

	(i) OLS estimates	(ii) Two-step estimates	(iii) GMM joint estimates
Working time			
Full-time job	0.441 (4.13)	(0.436) (4.06)	(0.447) (4.29)
Firm size			
Medium	0.249 (2.33)	(0.253) (2.39)	(0.249) (2.42)
Large	0.014 (0.08)	(0.003) (0.02)	(0.005) (0.03)
Private sector	0.037 (0.33)	(0.031) (0.27)	(0.044) (0.40)
Personal characteristics			
Logarithm of age	(0.274) (2.22)	(0.375) (2.44)	(0.353) (2.23)
Education			
Third level	(0.283) (2.77)	(0.261) (2.42)	(0.264) (2.49)
2nd. stage secondary	(0.142) (1.59)	(0.122) (1.36)	(0.119) (1.34)
Male	(0.183) (2.09)	(0.137) (1.43)	(0.143) (1.37)
Ln(duration)	−0.034 (0.03)	(−0.008) (−0.15)	(−0.013) (−0.25)
Residential change	(−0.096) (−0.43)	(−0.131) (−0.57)	(−0.142) (−0.65)
Constant	(4.865) (8.93)	(4.802) (8.40)	(4.834) (8.54)
ρ_0		(−0.116) (−1.14)	(−0.110) (−0.98)
ρ_1		(0.114) (0.79)	(0.139) (0.76)
R^2	0.13		
Number of observations	612	1764	1764

Notes: 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.

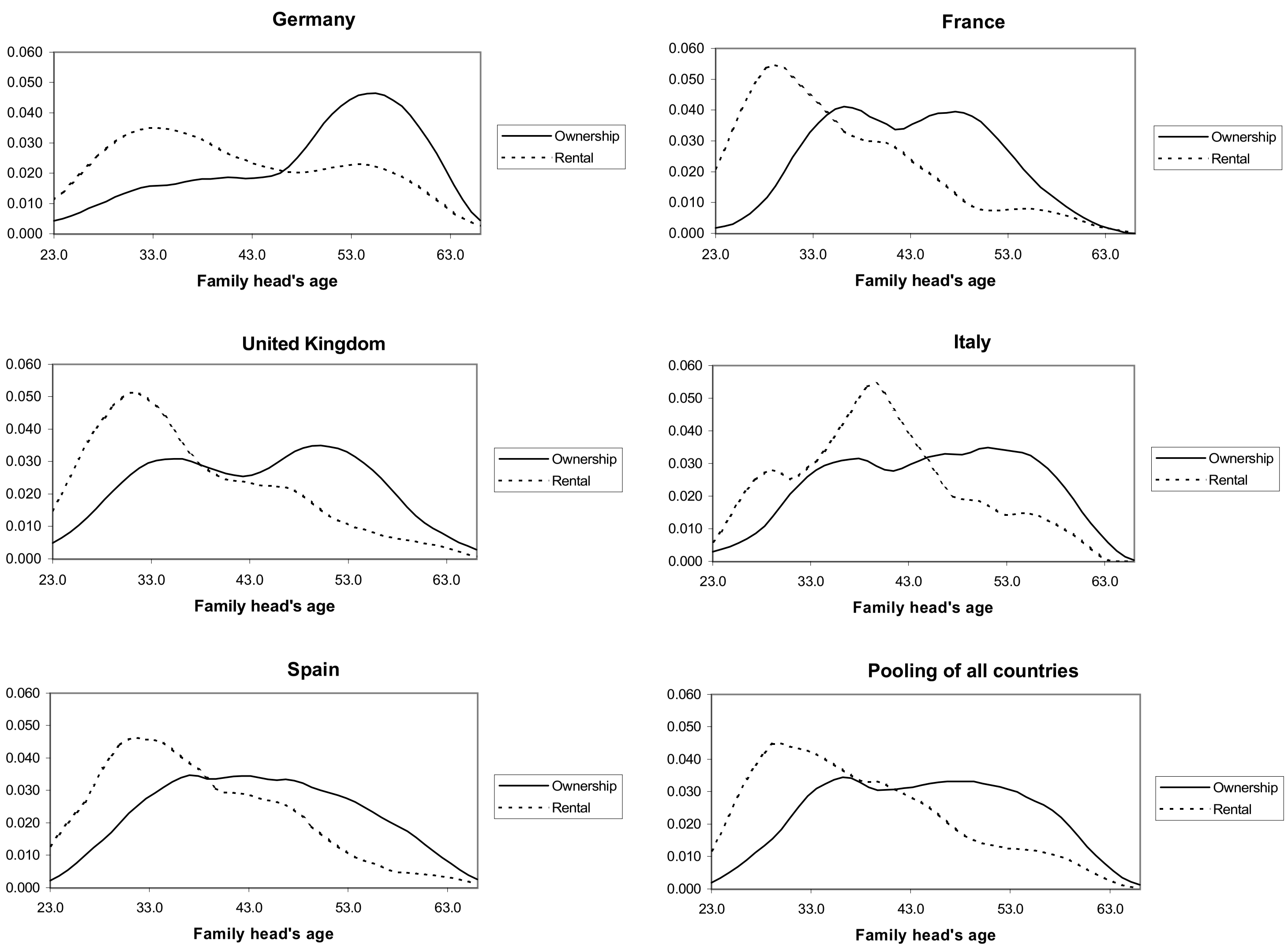

Figure 2: Kernel density estimates of the logarithm of the household's total income according the housing tenure.