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WORKING PAPER

Housing Tenure and Geographical Mobility in Belgium

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Abstract

Housing tenure is a key determinant of geographical mobility. We estimate several probit models to explain the probability that households move, using Belgian longitudinal PSBH and EU-SILC datasets which together cover the period 1994-2009. We confirm the general conclusion in previous literature, that homeowners are, ceteris paribus, less mobile than tenants. Within the first category, having a mortgage further hampers mobility. Earlier results for Belgium did not find a significant difference between outright owners and mortgagees. Furthermore, we make progress on the existing literature by paying particular attention to (and dealing with) methodological issues such as unobserved heterogeneity and state dependence. However, we also obtain some indications that the strict exogeneity assumption may be violated, implying that we cannot exclude the possibility of some bias in our estimated coefficients.

JEL classification: J61, R21, R23
Keywords: Housing tenure, geographical mobility, Belgian households, panel data

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1. Introduction

The most cited study in the literature regarding residential mobility, is assumably the work of Rossi (1955). This book is believed to have had the greatest influence on later empirical work since it was the first to extensively analyze geographical mobility at the micro level. Before, academic research mainly concentrated on aggregate mobility flows between regions. Rossi (1955) already concluded that housing tenure is one of the major determinants of mobility. Although he conducted his research from a primarily sociological point of view, his micro focus was rapidly adopted by economists. Many studies elaborated further on this early work, often focusing on diverse subfields (e.g. housing satisfaction by Diaz-Serrano & Stoyanova (2010); quality of the neighbourhood by Rabe and Taylor (2010a); public policy implications by Caldera Sánchez and Andrews (2011b); social capital by Kan (2007) and house prices and housing supply by Ferreira et al. (2008) and Rabe and Taylor (2010b)).

In economics, a major motivation for analyzing residential mobility is assuredly the link between geographical mobility and the labour market. As Blanchard and Katz (1992) demonstrate, mobility is a key instrument to resolve mismatches in the labour market. Mobile workers will move from regions hit by adverse labour demand shocks to regions with a higher labour demand or with specific requests for specialized skills. In the early research by Rossi (1955) and later by Speare et al. (1975), Hughes and McCormick (1981, 1987) and Clark and Dieleman (1996), the consensus holds that tenants are more mobile than homeowners. The latter experience higher costs when buying and selling a dwelling[1]. The research topic gained increasing attention since the influential work of Oswald (1996, 1997), in which these arguments where used to explain the harmful effects of homeownership on labour market outcomes. In particular, Oswald concludes that countries/states with a higher fraction of homeownership suffer from higher unemployment rates. In this paper we investigate whether owners are indeed less geographically mobile in Belgium. Of course, finding lower mobility would not directly prove Oswald's theory to be valid. It should be rather considered as a necessity and not a sufficiency for the validity of Oswald's hypothesis.

[1] These might include search costs, all kinds of transfer taxes, real estate agency fees, solicitor fees and refinancing costs.

The existing empirical literature can be categorized along several axes. A first one is the type of dataset that is used. First, a few studies (e.g. Gardner et al. (2001), Ermisch and Di Salvo (1996)) were able to use data that run over a very long period, even multiple decades. This enables the researchers to observe individual's complete moving history and therefore adopt a duration analysis. Although it is a powerful tool, important drawbacks are the high rate of attrition and the loss of quality when retrospective data are collected with large intervals. Second, many studies use cross-sectional data (e.g. Hughes & McCormick (1981, 1987), Helderman et al. (2004), Caldera Sánchez & Andrews (2011b)). The main shortcoming of these studies is that they fail to account for unobserved heterogeneity. Third, longitudinal panel data that cover a limited time period became increasingly popular over the last few decades. The advantages over the cross-sectional datasets are clear. The main drawback is that they are only able to analyze a snapshot of the total moving history. In the empirical section, we explain how we attempt to resolve this matter.

A second way to categorize is in the distinction between different types of housing tenure that have been introduced by researchers, and in the differential effects each type may have. First, tenants can be alternatively defined as private tenants or social tenants. Hughes and McCormick (1981, 1987) find that the higher mobility of social tenants compared to owners only holds over short distances. When only considering moves over a long distance, this group is even less mobile than owners. The reason for this is that public sector tenants could risk losing their benefits if they migrate. Often, the demand for social accommodation is higher than the supply, so people could end up on waiting lists or lose their chance of public housing entirely (Champion et al., 1998). Hence, the group of tenants cannot be considered as a homogeneous group. Second, a more recent distinction has been made between homeownership with a mortgage and outright homeownership. There is much less of a consensus about this topic. From a theoretical point of view, Böheim and Taylor (2002) address negative equity as a potential obstruction to mobility, in particular for mortgagees. If house prices drop, homeowners are restrained to move since a transaction will transpose their virtual loss into actual loss. Although Belgian average house prices did not decrease since the first half of the 1980s, the negative equity effect can still emerge for a different reason. According to Catte et al. (2004) and Van Ommeren and van Leuvensteijn (2005), the above described housing transaction costs are about 23% of the total house price in Belgium, directly causing a lock-in effect for homeowners[2]. We deduce two reasons why mortgage holders are constrained the most in this framework. A first one is that transaction costs are higher due to bank costs and extra solicitor fees. As a second one, mortgage holders are assumed to be more prone to the negative equity trap than outright owners, due to the financial constraints in case debt exceeds the market value of their current residence. Caldera Sánchez and Andrews (2011a) raise one counterargument to this: the mortgagees' greater incentive to remain employed or regain employment more rapidly. According to their theory, owners with a mortgage will have a larger urge for mobility to preserve the ability to repay their mortgage.

Empirically, the panel data analyses of Böheim and Taylor (2002) and Rabe and Taylor (2010a) conclude that outright owners are significantly more mobile than owners with a mortgage. However, Caldera Sánchez and Andrews (2011a,b) find the opposite result in almost all OECD countries, using a cross-section of EU-SILC data. They confirm the lower mobility of mortgagees only for Israel, Luxemburg, Norway, the UK and the US. Belgium is one of the exceptions where no significant difference between the two groups of homeowners is found.

In this paper, we investigate the impact of housing tenure on residential mobility using the PSBH and more recent EU-SILC datasets. We control for differences between both types of tenants and further investigate the disputed effect of having a mortgage for owner-occupiers. We derive the model specification from the abovementioned literature and adopt the methodology that has been employed by Tatsiramos (2009), Rabe and Taylor (2010a) and Diaz-Serrano and Stoyanova (2010). We handle unobserved heterogeneity and contribute to these models by treating the initial conditions problem. To our knowledge, this is the first extensive study for Belgium. Only Caldera Sánchez and Andrews (2011b) report estimates for Belgium within an international comparison, based on cross-sectional data. For various reasons, Belgium is a very interesting case to study. Since it is a very densely populated country, the social costs of commuting are considerable. Therefore, a high degree of residential mobility is desirable. Also, the number of homeowners in Belgium is very high compared to most Western countries, representing about 70% of Belgian households

[2] Since 2002, Flanders altered its system of stamp duties, permitting a discount in stamp duties subject to the stamp duties paid for a previous purchase. This policy measure mitigates the lock-in effect to some extent.

according to our sample. If homeowners are indeed less mobile, this may have a very large impact on the Belgian labour market. Finally, the above mentioned Belgian transaction costs are the highest among OECD countries. Consequently, residential mobility of homeowners, especially with a mortgage, is expected to be very low. The paper confirms that homeowners are indeed less mobile than tenants. If the owner-occupier has a mortgage, mobility further decreases. Social tenants have a lower propensity of moving than private tenants. Although we tackle a number of methodological issues by employing more advanced estimators, caution is still required when interpreting the results because of a potentially remaining endogeneity bias in the estimates.

The paper is organized as follows. Section 2 presents the methodological background of the different estimators that we use. In Section 3 we discuss the different determinants that are included in the model specification. Furthermore, some basic descriptive statistics are shown of both the explanatory variables and the dependent variable, geographical mobility, in order to provide the reader with some first insights. In Section 4 we present the estimation results. Section 5 concludes.

2. Methodology

Both data sources, PSBH and EU-SILC, offer the advantage of longitudinal data. First of all, this allows us to analyse expressed rather than revealed preferences of mobility. Second, panel data make it possible to take unobserved heterogeneity into account. Recent studies using only cross-sectional data such as Calderea Sánchez and Andrews (2011b) and Lux and Sunega (2012) explicitly acknowledge it as a shortcoming when a time dimension is lacking. In this section, we elaborate on the different econometric models that are used by Tatsiramos (2009), Rabe and Taylor (2010a) and Diaz-Serrano and Stoyanova (2010). We draw particular attention to the issue of unobserved heterogeneity. We make progress on these studies by also taking state-dependence and the strict exogeneity assumption into account.

2.1 The pooled Probit model

Whether a household moves or not over the subsequent year, is a binary choice. The observed variable y_{it} is equal to 1 in case of mobility between t and t+1, and 0 otherwise. To indicate the panel structure of our data, the subscripts i and t are used to denote the household and year respectively. The probability to move is assumed to depend on a continuous unobserved latent variable y_{it}^* (with $y_{it} = 1$ if $y_{it}^* > 0$ and $y_{it} = 0$ otherwise). We adopt a Probit model. This specifies the probability that $y_{it}^* > 0$ as a cumulative standard normal distribution. The latent propensity y_{it}^* can then be written as:

$$y_{it}^* = \beta_0 + \beta_H H_{it}^{'} + \beta_Z Z_{it}^{'} + \beta_R R_{it}^{'} + v_{it}, \tag{1}$$

with $H_{it}^{'}$: vector of housing tenure dummies

$Z_{it}^{'}$: vector of household characteristics

$R_{it}^{'}$: vector of area characteristics

v_{it}: error term

The exact elements out of which the different vectors consist, are described in the data section. Equation (1) shows a model which ignores unobserved heterogeneity. This is known as the Pooled Probit model and is very similar to the cross-sectional Probit model.

2.2 Dealing with unobserved time-invariant heterogeneity

Having repeated observations for each household enables us to deal with unobserved time-invariant heterogeneity. We decompose the error term from equation (1) into:

$$v_{it} = \mu_i + \varepsilon_{it}, \tag{2}$$

with μ_i denoting the time-invariant term and ε_{it} denoting the time-variant term. Several estimation options are possible to treat unobserved heterogeneity. If the time-invariant household effect μ_i is independent from the explanatory variables, the *Random Effects Probit estimator* is suitable and efficient. However, many scenarios can be thought of in which this assumption is violated. Some unobserved household-specific characteristics may affect both mobility and the explanatory variables. For example, households that have an intrinsically larger urge for stability, will have a lower probability of moving and might also have a higher propensity of homeownership. When these suspicions are justified, the Random Effects estimator produces inconsistent results. In this case, many refer naturally to the Fixed Effects estimator as a valid alternative. However, unlike linear models, non-linear models do not allow to estimate the fixed effects as dummies along with the rest of the equation through MLE when T does not go to infinity. This problem yields inconsistent estimates and is commonly known as the incidental parameters problem (see e.g. Wooldridge, 2010). To overcome this obstacle, it is common to use the Conditional Logit model (Chamberlain, 1980). Diaz-Serrano and Stoyanova (2010) see three important drawbacks to this method. First, observations with only positive or only negative outcomes of mobility are excluded from the sample. This is critical because non-movers in our dataset are a numerous group and particularly relevant to our research question. Second, the effect of time-invariant explanatory variables cannot be estimated. Third, explanatory variables with limited variation over time are weakly estimated because much of the variation is absorbed by the fixed effects. This is an unattractive feature since housing tenure varies rather little over time. Therefore, a more appropriate model is required.

In an attempt to preserve the convenient properties of the Random Effects model without conceding to the rather unrealistic assumption of the covariates being uncorrelated with the error term, Tatsiramos (2009), Rabe and Taylor (2010a) and Diaz-Serrano and Stoyanova (2010) suggest using an alternative specification. Following Mundlak (1978), the assumption is relaxed by allowing that μ_i is indeed correlated. The *Mundlak approach* assumes that the regression function of μ_i is linear in the within-individual mean values of the time-variant explanatory variables (denoted by $\bar{H}_i'$ and $\bar{R}_i'$). For our model, this boils down to:

$$\mu_i = \alpha_0 + \alpha_H \bar{H}_i' + \alpha_R \bar{R}_i' + e_i \text{,} \qquad (3)$$

where e_i is the individual effect with $e_i \sim N(0, \sigma_e^2)$ and not correlated with $H_{it}', Z_{it}', R_{it}'$ and v_{it}. α_0 is absorbed by the constant term β_0 in equation (1). Intuitively, this approach can easily be rationalized. If some unobserved fixed household-specific characteristic affects

H'_{it}, Z'_{it} or R'_{it}, then it will be reflected in the individual-specific mean values of these variables, denoted by $\bar{H}'_i$, $\bar{Z}'_i$ and $\bar{R}'_i$. By including these individual-specific mean values in the regression, any relevant influence of the underlying unobserved fixed characteristic is taken out of the error term. Correlation between included explanatory variables and the error term is then no longer possible, at least not for this reason[3]. Note that when we adopt this procedure in Section 4, the vector of within-means of household characteristics $\bar{Z}'_i$ is not included among the Mundlak terms. The reason is that we keep these characteristics constant at their initial value in the first period of observation (see Section 2.4. below).

2.3 State-dependence and the Wooldridge approach

Yet another issue is the possible appearance of state-dependence. In the introduction, we discussed the dichotomy between studies using duration models and research based on shorter longitudinal intervals, using discrete choice models. The latter attempts to evaluate a flow sample using a static model, also known as stock sampling. Many households probably have rarely or never moved while others have moved several times before the start of the survey. This information is ignored in the previously described specifications and is apparently not a point at issue in the aforementioned literature. Past mobility however may influence the probability of current mobility. One might be less tied to the neighbourhood or the house. Also, moving requires a large effort with respect to search and logistic costs. We believe that repeated moves lower these costs through learning effects. Hence, households that moved more recently may be more likely to move in a later period compared to an otherwise identical household. This situation is called state-dependence. Because the period of observation does not coincide with the whole mobility history of the household, the initial condition problem has to be considered. A bias could result from the possible correlation between unobserved heterogeneity and lagged values of the dependent variable. In the aforementioned literature this issue has not yet been tackled.

[3] The effect of unobserved fixed characteristics that are not correlated with the explanatory variables cannot be removed from the error term by this procedure. But that is not a problem. When there is no correlation with the included explanatory variables, the error term will not be correlated with these variables either.

Wooldridge (2005) suggests a computational inexpensive solution to this problem. The *Wooldridge approach* adds the initial condition y_{i0} as a supplementary regressor to the model. Alternatively, we use the number of years that passed (T_0) since the last move at the start of the observed period, as a proxy for the initial condition. This variable controls for the mobility prior to the observed period. Next, we add a yearly changing counterpart (T_t) which captures the duration effect throughout the observed period. The revised specification of the unobserved heterogeneity is denoted by[4]:

$$\mu_i = \alpha_0 + \alpha_H \bar{H}_i' + \alpha_R \bar{R}_i' + \alpha_{T0} T_0 + \alpha_{Tt} T_t + e_i \, , \qquad (4)$$

2.4 The strict exogeneity assumption

One of the requirements for using a panel model with unobserved heterogeneity is strict exogeneity. This issue has been generally neglected in the empirical mobility literature. Specifically, the error term ε_{it} must be uncorrelated with the explanatory variables in all time periods. A well-known example to clarify the issue is family expansion. The household may move to a more suitable residence because it has fertility plans. This may clearly bias the estimated coefficients. To overcome this issue, we keep most of the variables constant at the initial value in the first period of the observation. In particular we do this for the large vector of household characteristics Z_{it}'. The vector of area characteristics R_{it}' only contains variables at the aggregate level, so for these variables strict exogeneity is assured by definition. We do not adopt this procedure for the housing tenure dummies H_{it}'. The reason is that, when values are fixed to the first period, it is no longer possible to structurally interpret the coefficients. So, for these housing variables we cannot exclude the possibility of some endogeneity. We will test whether or not the strict exogeneity assumption is violated by including the lead variable of housing tenure in the regression. A second drawback of fixing variables at the initial value, is unavoidably that some of the information in our dataset is neglected. If for example the household composition changes during the observed period, this will not be taken into account when analysing the effect on mobility. In this scenario, an omitted variable bias can occur. These two disadvantages probably explain why the previous

[4] Again using a Random Effects estimator.

literature did not adopt this approach. Nevertheless, we believe that neglecting the apparent absence of strict exogeneity for many of the variables, may induce even stronger limitations.

After having discussed the data in Section 3, Section 4 provides the results and evaluates the four models that are introduced in this section. To sum up, the Pooled Probit model neglects unobserved heterogeneity, while the Random Effects Probit estimator requires that it is uncorrelated with the regressors. This assumption is weakened by the Mundlak approach. The Wooldridge approach tackles the potential initial condition problem. Last, we explicitly test whether the assumption of strict exogeneity holds.

3. Data

We use data from two different surveys of Belgian households. Together they cover the period 1994-2009 with a temporary break in 2003. For the period 1994-2002 we use the PSBH survey (Panel Study for Belgian Households). This survey ran from 1992 to 2002. However, since it was only normalized and integrated into the European Community Household Panel (ECHP) in 1994, the first two waves are withdrawn from our panel. The PSBH contains a wide range of socio-economic variables, both at the household level and at the level of its members[5]. It was built to serve as a longitudinal database, making it possible to analyse various social issues over time. From the 9 waves of the panel we can extract 8 time periods (T=8) and a total of 18,262 household-years after omitting the observations with missing values.

Second, we study the period 2004-2009 with the EU Statistics on Income and Living Conditions (EU-SILC), which is the more widespread successor of ECHP[6]. It comprises similar variables as its predecessor, although a direct comparison between both is not appropriate. The main difference is the set-up of the dataset. The EU-SILC dataset is also longitudinal but it is constructed as a rotating panel of 4 subsamples, each year replacing one fourth of the total sample. In this way, a household can stay in the panel for a maximum of 4 years. Each

[5] More information and metadata: http://www.psbh.be

[6] More information and metadata: http://epp.eurostat.ec.europa.eu/portal/page/portal/microdata/eu_silc

subsample contains initially 1,500 households. Our total panel counts 6 waves (so T=5) and a total of 13,434 household-years.

A first choice in constructing a feasible dataset is whether to work with households or individuals as sample units. To resolve this issue, we follow the arguments of Hughes and McCormick (1981) and van Ham et al. (2010). The event of a residential move is a household decision and the probability of moving depends to a large extent on household characteristics. Alternatively, Hughes and McCormick (1987) use the "heads of households" as sample unit, but this might seem somewhat arbitrary. Of course, sometimes households are forced to move when a split occurs as a result of a divorce or the end of co-housing. We consider these events as random and not in the scope of our subject. Therefore, these households are eliminated from the sample.

A second refinement of our dataset is to drop households without at least one person of the age group 25 to 64. The lower boundary (age 0-24) largely filters out first time movers and students while the upper boundary (age 65-…) filters out the so-called pension mobility. As noticed in the introduction, we are particularly interested in mobility benefiting the labour market and therefore restrict our sample to the population at working age (but older than 24). For analyses of the mobility of other groups than the population of working age, we refer to a wide range of existing literature (e.g. Andrew (2004), Angelini and Laferrère (2010)).

Before going through the different covariates that are included in the model, we first discuss the dependent variable, geographical mobility. More specifically, the dependent variable equals 1 if the household moved within the year after period t, conditional on the explanatory variables in period t. In an international perspective, residential mobility is rather low in Belgium. Caldera Sánchez and Andrews (2011b) constructed a comparative study of OECD countries. They show that in 2007, almost 12% of Belgian households had moved over the last 2 years. This is comparable to neighbouring countries as Luxembourg, the Netherlands and Germany but much lower than France and the UK. The Nordic and Anglo-Saxon countries experience typically higher mobility, while mobility is typically low in eastern and southern European countries.

Table 1 contributes more detailed information about the observed mobility in both datasets. Slightly more than 5% of all observed household-years show residential mobility. When making a distinction between short and long distance mobility, we see that the vast majority of moves are local. According to the PSBH dataset, only 0.91% of household-years resulted in a move from one district to another. This equals 15.6% of the total number of moves. EU-SILC does not contain information at the district level but it does on the larger scale of provinces. From the total number of observations, 0.47% concern a move between provinces, which corresponds to 8.8% of the total number of moves. These numbers are clearly lower than the statistics based on the British Household Panel Survey (BHPS) that Böheim and Taylor (2002) report. According to their findings, 17.0% of the total number of moves proceeded between regions (comparable to the Belgian provinces) and 33.1% between Local Authority districts (comparable to Belgian districts). In the PSBH, the households that moved were asked about the main reason for moving house (Table 1, n°4). Only 7.05% of households that responded to this question, stated that the main motivation for mobility was work-related.

Table 1: Some descriptive statistics of mobility

	Move (%)	Move (Abs.)	No move (%)	No move (Abs.)	Total (Abs.)
1. Overall fraction of residential mobility					
- PSBH	5.86%	1,039	94.14%	16,703	17,742
- EU-SILC	5.08%	683	94.92%	12,751	13,434
2. Overall fraction of long distance mobility					
- PSBH (between districts)	0.91%	162	99.09%	17,580	17,742
- EU-SILC (between provinces)	0.47%	60	99.53%	13,371	13,434
3. Overall fraction of short distance mobility					
- PSBH (within districts)	4.94%	877	95.06%	16,865	17,742
- EU-SILC (within provinces)	4.64%	623	95.36%	12,811	13,434
4. From the households that moved, which fraction said the main reason for mobility was (PSBH):					
A. Work	7.05%	59	-	-	-
B. The housing itself	69.41%	581	-	-	-
C. Personal reasons	23.54%	197	-	-	-

Source: own calculations; Panel Study on Belgian Households (1992-2002), Universiteit Antwerpen, Université de Liège and EU-SILC (2004-2009), FOD Economie, K.M.O., Middenstand en Energie

To summarize, the descriptive statistics in Table 1 show that only a small fraction of moves is non-local and that the observed mobility is only for a small part motivated by work related circumstances. Four our empirical analysis, the question can be asked which definition of mobility to use. For instance, some studies confined the dependent variable to long distance mobility (e.g. Hughes and McCormick (1985) and Böheim and Taylor (2002)) or mobility for job reasons only (Gardner et al. (2001)). Unfortunately, our two datasets do not allow using a more restricted definition for mobility. First of all, the number of non-local moves is not only small in relative numbers but also in absolute numbers. Only 162 non-local moves are observed in PSBH and 60 in the EU-SILC data. In order to properly estimate an extensive model, the sample size should be much larger. Futhermore, incorporating the reason of mobility would also introduce an important deficiency. It can be deduced from Table 1 that only 4 out of 5 movers responded to this question. The main reason is that in case of attrition in t+1, mobility can be observed for the most part but the reason remains unknown. The attrition can possibly bias the results. Although these limitations might seem unfortunate, we argue that both suggested refinements of the mobility definition are not particularly desirable. Our main research question is more general. We want to investigate the impact of housing tenure on residential mobility. The reason for this mobility or the distance is less important.

The selection of explanatory variables in the model is based on a numerous amount of earlier studies[7]. The applied regressors appear repeatedly in the literature, but sometimes in altered form. We discuss them concisely one by one. Table 2 for PSBH and Table 3 for EU-SILC show the composition for categorical variables and the mean and standard deviation for continuous variables. A distinction is made between the group of movers, non-movers and the total sample in order to provide some first insights into the determinants of mobility. We divide the explanatory variables into 3 broad categories: housing tenure, household characteristics and area characteristics.

As we clarified in the introduction, in this paper *tenure choice* is subdivided into 4 subcategories. We make the distinction between outright owners, owners with a mortgage, tenants paying rent at market rate and tenants paying rent at a reduced rate. In the total population the group of tenants paying rent at market rate is a little over 20%. When only

[7] E.g. Bartel (1979), Hughes and McCormick (1981, 1987), Gardner et al. (2001), Böheim and Taylor (2002), Helderman et al. (2004), Taylor (2007), Tatsiramos (2008), Diaz-Serrano and Stoyanova (2010), Rabe and Taylor (2010a) and Caldera Sánchez and Andrews (2011b),

taking the movers into account, the private tenants count for two thirds of the households-years.

Table 2: PSBH Descriptive Statistics of Explanatory Variables

	No move		Move		Overall	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Tenure choice						
Owner with mortgage	0.4752		0.1627		0.4569	
Outright owner	0.2746		0.0674		0.2624	
Private tenant	0.1876		0.6833		0.2167	
Reduced rent	0.0626		0.0866		0.0640	
Household characteristics						
Age	43.25	10.91	36.96	10.15	42.88	10.97
Family structure						
Single person	0.1881		0.2878		0.1939	
Single parent	0.0589		0.1261		0.0625	
Childless couple	0.4200		0.3012		0.4134	
Couple with children	0.3330		0.2849		0.3302	
Education						
Tertiary	0.4399		0.4398		0.4399	
Secondary	0.3129		0.3147		0.3130	
Less than secondary	0.2471		0.2454		0.2470	
Other						
Foreign nationality	0.1203		0.1578		0.1225	
Monthly income	2,109	1,090	1,866	997	2,094	1,086
Rooms per household member	1.90	1.28	1.83	1.13	1.89	1.28
Years since installation	10.97	7.23	5.88	5.03	10.67	7.22
Area characteristics						
Prov. unemployment rate	8.90	4.33	9.76	4.41	8.95	4.34
Housing trans./cap (t-1)	0.0070	0.0017	0.0069	0.0019	0.0070	0.0017
GVA Construction/cap (t-1)	0.997	0.201	0.985	0.188	0.996	0.200
Population density	971	1738	1354	2135	994	1766
Regional dummies						
Brussels	0.1094		0.1790		0.1135	
Flanders	0.4982		0.4042		0.4927	
Wallonia	0.3924		0.4167		0.3938	

Source: Own calculations; Panel Study on Belgian Households (1992-2002), Universiteit Antwerpen, Université de Liège and EU-SILC (2004-2009), FOD Economie, K.M.O., Middenstand en Energie

Table 3: EU-SILC Descriptive Statistics of Explanatory Variables

	No move		Move		Overall	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Tenure choice						
Mortgagee	0.4191		0.1483		0.4053	
Outright owner	0.3059		0.1057		0.2957	
Private tenant	0.2033		0.6622		0.2266	
Reduced rent	0.0717		0.0837		0.0723	
Household characteristics						
Age	46.73	10.86	39.70	10.96	46.37	10.97
Family structure						
Single person	0.1959		0.3098		0.2017	
Single parent	0.1045		0.1292		0.1057	
Childless couple	0.4428		0.3422		0.4377	
Couple with children	0.2568		0.2188		0.2549	
Education						
Tertiary	0.4579		0.4640		0.4582	
Secondary	0.3671		0.3363		0.3656	
Less than secondary	0.1750		0.1997		0.1762	
Other						
Foreign nationality	0.1615		0.2775		0.1674	
Monthly income	34,304	65,830	28,271	27,681	33,998	64,452
Rooms per household member	2.33	1.36	2.14	1.32	2.32	1.36
Years since installation	14.47	11.31	6.75	7.31	14.08	11.27
Area characteristics						
Prov. unemployment rate	12.25	6.29	13.18	6.57	12.30	6.31
Housing trans./cap (t-1)	0.0065	0.0017	0.0061	0.0019	0.0065	0.0017
GVA Construction/cap (t-1)	1.128	0.277	1.115	0.259	1.128	0.276
Population density	1133	1945	1733	2457	1164	1978
Regional dummies						
Brussels	0.1274		0.2305		0.1327	
Flanders	0.5470		0.4860		0.5439	
Wallonia	0.3256		0.2834		0.3234	

Source: Own calculations; Panel Study on Belgian Households (1992-2002), Universiteit Antwerpen, Université de Liège and EU-SILC (2004-2009), FOD Economie, K.M.O., Middenstand en Energie

Second, we incorporate a wide range of variables that capture *household characteristics*. In order to meet the strict exogeneity assumption, these variables are fixed at the value of the household's first observation in the sample as has been more elaborately discussed in the previous section. We introduce age and its square form to take life cycle effects into account. We also include its square because the relationship is not expected to be linear. We expect a decrease of mobility with age because of declining present discounted wage benefits of mobility while the cost of moving house does not decline (Schwartz, 1973 and Sjaastad, 1962). Because we need one observation per household, we use the age of the oldest member who is not yet 65. Next, a categorical variable is introduced that relates to family structure. The expected effect from having children is ambiguous. We would expect that it becomes more costly and complex to move with a larger household. On the other hand, households might need to move to a larger home to satisfy the household's needs. To take this last consideration into account, we follow the example of Helderman et al. (2004) and Böheim and Taylor (2002) and introduce a proxy for the so-called room-stress. We include the ratio of the number of rooms to the number of household members. If a household enjoys more space, it is expected to be less likely to move. By incorporating this variable, the expected estimate of having children on mobility is unambiguously negative. Cohabiting is expected to have a negative effect on mobility. Following the arguments of Helderman et al. (2004), we combine the properties of cohabiting and whether the household contains children. It is conceivable that having children will have a different impact on singles than on couples. Next, the level of education is included. To incorporate this we take into account the highest acquired degree of one of the household members. We make a distinction between tertiary education, higher secondary education and not having fulfilled secondary education. Last, we control for income and nationality. As to the latter, a dummy equals 1 if at least one household member has a foreign nationality. In the literature, the estimates for income are rather inconclusive, while being foreign is generally found to have a positive effect on mobility.

Finally, the third category contains *area characteristics*, based on aggregate data derived from Cambridge Econometrics data and data from the 'FOD Economie', Belgian Federal Government. We include the provincial unemployment rate in year t as an additional determinant. The expectation is that households have a higher propensity to move when the labour market is depressed. Next, we add a number of proxies to account for housing market conditions: the provincial per capita number of housing market transactions in the year t-1, provincial per capita real gross value added (GVA) of the construction sector in t-1 and last, the population density in the province. These proxies should capture housing supply as a determinant for mobility. In a more liquid housing market, mobility is expected to be more accessible. Last, to control for spatial disparities, we add dummies for the different regions. These capture a range of legislative, cultural and demographic circumstances that could influence the mobility of a household.

4. Results

Table 4 and Table 5 show the results for PSBH and EU-SILC respectively. In contrast to linear models, the estimated coefficients cannot be interpreted in a straightforward way. The size of the partial effects is subject to the selected values of the other regressors. Besides, when we use the panel dimension of the data, the partial effect also depends on the value of the unobserved heterogeneity, μ_i. We follow the suggestion in Wooldridge (2005) and calculate Average Partial Effects (APE's)[8]. This results in one single interpretable estimate for each determinant. For discrete variables, the partial effect equals the difference in probability when the dummy changes from 0 to 1. Accordingly, in case of categorical variables, the APE reveals the difference in moving probability compared to the reference category.

The columns in each table represent the alternative estimation methods as discussed in the previous section: (1) is the Pooled estimator; (2) is the Random Effects estimator; (3) is the so-called Mundlak-approach in which the means of the time-varying regressors are included and (4) is the estimation in which the time since last mobility is added as a supplementary control variable. From a first glance, we can see that the results are quite similar, irrespective of the method. However, the size of the APE's of interest changes rather considerably between the columns. We now discuss the suitability of the different models. At the bottom of the table, Rho indicates the variation that is captured by the unobserved household specific term. Using the RE estimator, this fraction is very low in the PSBH dataset, representing only 0.48% of variation. It is not significantly different from zero. Diaz-Serrano and Stoyanova (2010) obtain a similar result. For PSBH, we can confirm their conclusion that household specific effects are irrelevant and hence the pooled probit model is a more suitable framework compared to the Random Effects estimator. Table 5 shows that Rho equals 11.51% in the EU-SILC case and is significantly different from zero at the 5% level. To compare, Böheim and Taylor (2002) obtained 10% and Rabe and Taylor (2010a) 22%. In contrast to Table 4, unobserved heterogeneity should be accounted for.

[8] The calculation of APE's is very staightforward in the version of Stata®12.1 using the -margins- command.

Table 4: Probit coefficients and selected average partial effects based on PSBH data

	(1) Pooled	(2) RE	(3) Mundlak	(4) Wooldridge
Selected average partial effects				
Tenure choice				
Owner with mortgage	-0.005	-0.005	-0.053(***)	-0.063(***)
	(0.003)	(0.003)	(0.013)	(0.013)
Private tenant	0.131(***)	0.131(***)	0.133(***)	0.118(***)
	(0.007)	*(0.007)*	*(0.014)*	*(0.014)*
Reduced rent	0.048(***)	0.048(***)	0.069(***)	0.054(***)
	(0.008)	(0.009)	(0.019)	(0.018)
Area characteristics				
Prov. unemployment rate	-0.001	-0.001	-0.002	-0.004
	(0.001)	(0.001)	(0.003)	(0.003)
Housing trans./cap (t-1)	1.435	1.434	2.668	5.038
	(2.235)	(2.235)	(7.131)	(7.128)
GVA Construction/cap (t-1)	0.021	0.021	-0.001	0.019
	(0.015)	(0.015)	(0.048)	(0.049)
Population density	0.00007(***)	0.00007(***)	0.00009(***)	0.00009(***)
	(0.00002)	(0.00002)	(0.00002)	(0.00002)
Brussels	-0.135(***)	-0.135(***)	-0.382(***)	-0.336(***)
	(0.009)	*(0.009)*	*(0.118)*	*(0.119)*
Wallonia	0.032(***)	0.032(***)	0.037(***)	0.035(***)
	(0.008)	(0.008)	(0.011)	(0.011)
Probit coefficients				
Age	-0.050(***)	-0.050(***)	-0.055(***)	-0.036(**)
	(0.014)	(0.013)	(0.014)	(0.015)
Age squared	0.0004(**)	0.0004(**)	0.0005(***)	0.0003(*)
	(0.0002)	*(0.0002)*	*(0.0002)*	*(0.0002)*
FAM: Single person	0.124(**)	0.124(**)	0.194(***)	0.164(***)
	(0.059)	(0.059)	(0.060)	(0.062)
FAM: Single parent	0.201(***)	0.202(***)	0.246(***)	0.198(***)
	(0.064)	*(0.065)*	*(0.065)*	*(0.067)*
FAM: Couple with children	0.027	0.027	0.051	0.032
	(0.047)	(0.047)	(0.048)	(0.050)
EDU: Less than secondary	-0.002	-0.002	0.098	0.092
	(0.046)	(0.047)	(0.106)	(0.109)
EDU: Tertiary	-0.056	-0.056	-0.077	-0.101
	(0.041)	*(0.041)*	*(0.129)*	*(0.133)*
Foreign nationality	0.052	0.052	0.064	0.054
	(0.049)	(0.049)	(0.050)	(0.051)
Ln(Income)	0.057	0.057	0.011	0.0001
	(0.044)	*(0.044)*	*(0.045)*	*(0.047)*
Rooms per household member	-0.022	-0.021	-0.027	-0.030(*)
	(0.016)	(0.017)	(0.017)	(0.018)
Years since installation (constant)	n/a	n/a	n/a	0.026(**)
				(0.013)
Years since installation (continuous)	n/a	n/a	n/a	-0.049(***)
				(0.012)
Constant	-1.625(***)	-1.631(***)	-1.360(***)	-1.559(***)
	(0.476)	(0.479)	(0.486)	(0.503)
Time dummies	yes	yes	yes	yes
Means of time-varying covariates	no	no	yes	yes
Rho	n/a	0.0048	0.000002	0.000001
Log likelihood	-3,402.82	-3,402.80	-3,330.33	-3,149.63
Number of observations	18,262	18,262	18,262	17,728

Source: own calculations; Panel Study on Belgian Households (1992-2002), Universiteit Antwerpen, Université de Liège.

Note: * (**) (***) indicates statistical significance at 10% (5%) (1%). Between brackets are estimated standard errors. The reference category represents: outright owner, employee, age 35-54, childless couple, higher secondary education, no foreign nationality, in the Region Flanders.

Table 5: Probit coefficients and selected average partial effects based on EU-SILC data

	(1) Pooled	(2) RE	(3) Mundlak	(4) Wooldridge
Selected average partial effects				
Tenure choice				
Owner with mortgage	-0.010(***)	-0.009(***)	-0.044**)	-0.052(***)
	(0.004)	(0.003)	(0.019)	(0.019)
Private tenant	0.096(***)	0.090(***)	0.181(***)	0.181(***)
	(0.008)	*(0.008)*	*(0.025)*	*(0.024)*
Reduced rent	0.031(***)	0.028(***)	0.162(***)	0.161(***)
	(0.009)	(0.008)	(0.028)	(0.027)
Area characteristics				
Prov. unemployment rate	-0.0006	-0.0006	0.0003	-0.0029
	(0.0006)	(0.0006)	(0.0040)	(0.0044)
Housing trans./cap (t-1)	-0.427	-0.300	6.280	7.383
	(1.786)	(1.673)	(4.864)	(5.173)
GVA Construction/cap (t-1)	-0.006	-0.006	-0.016	0.101
	(0.010)	(0.009)	(0.111)	(0.128)
Population density	0.000001	0.000001	0.00005(***)	0.00005(***)
	(0.000002)	(0.000001)	(0.00001)	(0.00001)
Brussels	-0.003	-0.003	-0.003	-0.003
	(0.005)	*(0.005)*	*(0.005)*	*(0.006)*
Wallonia	-0.001	-0.001	-0.001	-0.002
	(0.004)	(0.004)	(0.004)	(0.004)
Probit coefficients				
Age	-0.045(***)	-0.049(***)	-0.048(***)	-0.044(***)
	(0.015)	(0.017)	(0.016)	(0.016)
Age squared	0.0003(*)	0.0003(*)	0.0003(*)	0.0003(*)
	(0.0002)	*(0.0002)*	*(0.0002)*	*(0.0002)*
FAM: Single person	0.057	0.056	0.068	0.053
	(0.060)	(0.065)	(0.063)	(0.061)
FAM: Single parent	-0.060	-0.069	-0.047	-0.061
	(0.073)	*(0.079)*	*(0.077)*	*(0.073)*
FAM: Couple with children	0.013	0.012	0.009	0.001
	(0.053)	(0.058)	(0.056)	(0.054)
EDU: Less than secondary	0.077	0.081	-0.016	-0.017
	(0.058)	(0.063)	(0.142)	(0.140)
EDU: Tertiary	0.113(**)	0.121(**)	0.108	0.107
	(0.047)	*(0.051)*	*(0.161)*	*(0.159)*
Foreign nationality	0.143(***)	0.148(***)	0.156(***)	0.124(**)
	(0.052)	(0.056)	(0.055)	(0.053)
Ln(Income)	-0.020	-0.020	-0.036	-0.027
	(0.038)	*(0.041)*	*(0.040)*	*(0.039)*
Rooms per household member	-0.016	-0.016	-0.021	-0.020
	(0.018)	(0.019)	(0.019)	(0.018)
Years since installation (constant)	n/a	n/a	n/a	0.083(**)
				(0.042)
Years since installation (continuous)	n/a	n/a	n/a	-0.099(**)
				(0.042)
Constant	-0.057	-0.054	0.135	0.205
	(0.514)	(0.556)	(0.544)	(0.523)
Time dummies	yes	yes	yes	yes
Means of time-varying covariates	no	no	yes	yes
Rho	n/a	0.1151(**)	0.0693(**)	0.00001
Log likelihood	-2,275.65	-2,273.42	-2,237.48	-2,221.00
Number of observations	13,434	13,434	13,434	13,431

Source: own calculations; EU-SILC (2004-2009), FOD Economie, K.M.O., Middenstand en Energie.

Note: * (**) (***) indicates statistical significance at 10% (5%) (1%). Between brackets are estimated standard errors. The reference category represents: outright owner, employee, age 35-54, childless couple, higher secondary education, no foreign nationality, in the Region Flanders.

Both tables show that the main dissimilarities appear when the so-called Mundlak terms are added to the Random Effects Probit model. As argued in the previous section, the results shown in column 2 are inconsistent if the unobserved heterogeneity term is correlated with the dependent variables. Because the extension of the Mundlak approach alters the estimates considerably and the Mundlak terms are jointly significant (not shown in the table), we can confirm our suspicions about the occurrence of unobserved heterogeneity as described in Section 2. The fourth column shows that conditioning the model on the duration spent in the residence, alters the magnitude of the estimated APE's rather strongly. The highly significant coefficient of "years since installation (constant)" proves that controlling for time spent in the residence, helps explaining the model. Its continuous counterpart reveals a significant negative coefficient. The longer a household remains in the same house, the more restrained it is for future mobility. This corresponds with the expectations as discussed in Section 2. We consider this last specification to be the most suitable model.

The results are very much in line with the expectations and similar in both datasets. The upper half of the tables shows the APE's of the key explanatory variables. Consistent with the earlier research that we described in the introduction, the results show that private tenants are the most likely to move, followed by tenants paying reduced rent. Unlike the results of Caldera Sánchez and Andrews (2011b) for Belgium, we do find a significant difference between outright owners and mortgagees. Outright homeowners appear to have a higher propensity to move (of 6.3% for PSBH, 5.2% for EU-SILC) which is in line with the estimates of Böheim and Taylor (2002) and Rabe and Taylor (2010a). The area characteristics have only limited explanatory power. The results do not provide evidence that households are more mobile if the aggregate unemployment rate is high, but neither did the aforementioned literature. From the proxies that we introduced to capture housing market conditions, only population density renders a significant coefficient. Households living in more densely populated areas experience higher mobility. A possible reason is the higher liquidity of the housing market. Finally, the regional dummies have significant estimates, but only in the PSBH dataset. The reason for the divergent outcome of both datasets is impossible to deduce. One would rather expect the opposite because Belgian housing policy was mainly decentralized in 2002. Of course, these regional dummies capture much more than this, so it is uncertain to what extent these APE's demonstrate policy effects.

The lower halves of Table 4 and 5 show the probit coefficients of the control variables. As clarified before, the estimates of these variables cannot be interpreted structurally because the values are fixed to the first observation of each household. Therefore, showing the APE's is otiose. For both datasets, the age categories have high explanatory power. Next we observe that the degree of statistical significance of the household characteristics differs substantially between both datasets. Whereas for the PSBH data family structure and room stress help to determine mobility, nationality is the most effective control in case of the EU-SILC data. Income seems to have no influence in both cases, which is consistent with Kan (2007) and Diaz-Serrano and Stoyanova (2010).

As announced in Section 2, we explicitly test whether or not the strict exogeneity assumption holds for the housing tenure dummies. When added to the Wooldridge regressions, we observe that the leads of some of these dummies do show significant coefficients. This result suggests that we fail to meet the assumption and caution is required when interpreting the estimated APE's. These are possibly driven by the correlation between 'shocks' to mobility and contemporaneous or future values of the housing tenure choice.

5. Conclusions

In this paper, we analyse the determinants of residential mobility in a large panel of Belgian households in 1994-2009. Like most papers in this literature, we find that - ceteris paribus - tenants are more mobile than owners. Neither of the two groups are homogeneous, however. Homeowners with a mortgage are less mobile than outright owners. Among tenants, those paying a reduced rate are significantly less mobile than tenants on the private market. The magnitude of the estimated average partial effects reveals that housing tenure is an economically significant determinant of mobility. Comparing the most and the least mobile groups (private tenants and mortgagees), we observe a difference of 18 to 23%-points in the probability per year to move. The hampered mobility of homeowners (especially with a mortgage) may have a large unfavourable effect in a country with a severely high homeownership rate.

Our estimation methods build on the Mundlak approach as applied in the previous literature (e.g. Tatsiramos (2009), Rabe and Taylor (2010a) and Diaz-Serrano and Stoyanova (2010)). Using the Wooldridge approach, we extend this estimation method to control for state-dependence. Both model specifications require strict exogeneity of the explanatory variables. Although the previous studies do not acknowledge this condition to obtain unbiased results, we have tried to avoid endogeneity by fixing as many potentially endogenous variables as possible at their initial value in the first period of the observation. Since this procedure implies, however, that a structural interpretation of the estimated coefficients is no longer possible, we could not impose it on the housing tenure variables. Tests reveal that in the end the strict exogeneity assumption may still be violated, implying that we cannot exclude the possibility of some bias in our estimated coefficients. The resulting bias probably manifests itself in all previous papers that analyse the effect of housing on mobility. A solution to this problem is not straightforward but progress may recently have been made by Biewen (2009). He developed a dynamic model (analysing poverty status) that explicitly allows for feedback from the dependent variable to future values of the explanatory variables. This might be a promising starting point for further research in order to avoid the remaining bias.

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