**Residential Mobility: Wealth, Demographic and Housing Market Effects**

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Abstract

The paper presents a very simple model in which housing equity can influence mobility, and then estimates parameters that gauge the impact of housing equity, local house prices and other variables associated with household structure and change on residential movement within the UK. The data come from the British Household Panel Study (BHPS) over 1992-2008, which allow us to use within person variation to identify the parameters. The parameter estimates indicate that estimates based on cross-section variation are seriously biased in our analysis. We check the robustness of our results to errors in measuring equity by using an instrumental variable estimator. Our main finding is that an increase in a household’s housing equity encourages residential mobility substantially, and a decline discourages it.

Introduction

Internal migration is clearly one of the key demographic processes, and migration over longer distances provides an important channel for labour market adjustment and flexibility. While most residential moves are local, primarily for adjustment of housing consumption, factors that discourage residential mobility also tend to discourage migration over longer distances. In this paper our focus is on the level of housing equity as a potential constraint on mobility. We lay out a very simple theoretical model in which housing equity can influence mobility, and then add measures of equity and local house prices to a standard empirical model of the major demographic influences on the likelihood of a move within the UK. Our interest is in whether, as has often been argued, having small amounts of equity in one’s home reduces the ability to realize a desired move (e.g. Ferraria et al. 2010).

The classic view of residential mobility holds that moves come about when changes in the household environment or its composition give rise to dissatisfaction with the current housing situation (Rossi, 1955; Brown and Moore, 1970). The life course approach utilizes the notion of a ‘housing career’, one that interacts with careers in the occupational, household and other domains. This approach has been the basis for a large body of empirical work emphasizing the ‘triggering’ role of life course events on mobility such as leaving the parental home, union formation and dissolution and childbearing (Clark and Dieleman, 1996; Mulder, 1996). Transitions to life stages involving higher levels of commitment – such as marriage and parenthood – lead to requirements for ‘long-stay housing’ (single family dwellings and owner-occupied housing) and to changing preferences over dwelling and neighbourhood attributes (Feijten and Mulder, 2002; Rabe and Taylor, 2010). Work in the economics literature has also focused on the investment motive as a reason behind moves to owner-occupied housing (e.g. Kiel, 1994). Desires for mobility may be prompted by a wide range of life events, but desires cannot always be realized. The literature has also focused on the constraints that determine the set of feasible alternatives to the current dwelling, and ultimately whether the household prefers to stay where it is or to move. Income and wealth constraints, transactions costs that vary with tenure, social ties, the supply of dwellings and the functioning of the housing and mortgage markets all affect whether a desired move will in fact be realized (Belot and Ermisch, 2009; Linneman and Wachter, 1989; Helderman et al., 2004; Stein, 1995; Venti and Wise, 1984; Wheaton, 1990).

The degree of residential mobility and the constraints that discourage it are important for the operation of the housing and mortgage markets. For instance, in Wheaton’s (1990) matching model, in which imperfect information makes it necessary for households to search for a dwelling that meets their needs, the rate at which households change their desired house affects key housing market variables like the proportion of households searching, expected time to sell and house prices. In his model, there is a tendency for a higher rate of demand change to increase the rate at which houses are sold and to raise house prices. Figure 1 illustrates a broad positive correlation (*r*=0.57) between the number of property transactions[[1]](#footnote-1) and the real house price of existing houses (i.e. deflated by the retail price index (RPI))[[2]](#footnote-2), as Wheaton’s theory would suggest.[[3]](#footnote-3) The size of the impact of housing equity on mobility also has implications for the formulation of regulations on mortgage lending. Large falls in house prices, such as those in the early 1990s in the UK and in many parts of the USA in recent years, are alleged to reduce residential mobility because they wipe out housing equity. In the extreme case the household may end up with negative equity —owing more on the mortgage than the house is worth. High loan-to-value ratios, if permitted, make it more likely that negative equity emerges.

In this paper we estimate parameters that gauge the impact of housing equity and local house prices, alongside other more standard mobility triggers, using data from the British Household Panel Study (BHPS). These data permit us to allow for persistent unobservable effects on mobility in our econometric analysis by using within person variation to identify the parameters. We find that, after accounting for the major demographic determinants of mobility and for unobserved individual characteristics, a rise in housing equity increases the ability of the household to realize a move substantially.

A Simple Model of Financial Influences on Residential Movement

Our model abstracts from imperfect information and search considerations, but it provides a simple framework for our empirical analysis. Suppose there are two types of house—a ‘smaller or lower quality’ one, denoted as H0, and a ‘larger or higher quality’ house, denoted as HN. We measure ‘housing services’ such that HN=λH0, where λ>1, and normalise so that H0=1. As we focus on younger couples, we can think of households contemplating a move to a larger or higher quality house.

Let W=the household’s financial wealth (i.e. non-housing), yielding a return of r, and y its annual earnings. Let M0 be the mortgage stock in the present house, and rM is the mortgage rate of interest. The household’s current consumption is c0=y+rW-rMM0 and its current utility, derived from consumption (c) and housing, is given by

U(c0, H0)=U(y+rW-rMM0, H0) (1)

If the household wishes to buy the ‘larger house’ and sell its current one, then its wealth available for house purchase is W+PH0-M0=W+qP, where P is the price of housing services in household’s housing market relative to other consumption, and q=(PH0-M0)/PH0=(P-M0)/P is the ‘housing equity ratio’. If the household buys the larger house, the household must expend PHN-MN =Pλ-MN, where MN is the mortgage on it.

*Down-payment constraint*

Previous analyses of the relationship between housing equity and residential mobility have often focused on the ‘down-payment constraint’ (Stein 1995; Englehardt, 2003; Ferraria et al. 2010). Its foundation is a maximum loan-to-value ratio for purchase of the new house, denoted here as k (0≤k≤1), and so if the household cannot borrow against future earnings without a house as collateral it faces a constraint that

W+qP≥(1-k)Pλ (2)

If they do not satisfy this down-payment constraint (e.g. when they have negative equity and W=0), then they cannot move to the new house. We can rewrite (2) as

W/P+ q≥(1-k)λ (3)

Clearly the chances of satisfying this constraint increase with the housing equity ratio, the maximum loan-to-value ratio and financial wealth; they decline with higher P and higher λ. Households headed by younger people are likely to have low values of W.

*Wealth effects of housing equity*

Suppose the down-payment constraint is not binding. Housing equity will still affect the mobility decision via the size of the mortgage payments required on the new house, and hence the sacrifice in consumption following a move. The household’s consumption in future periods after purchase of the new house is cN =y+r(W+ qP− PHN)+(r-rM)M­N. For simplicity, we shall assume r=rM in what follows. Then consumption is y+r{W+ (q-λ)P}and the household’s utility in the larger house is

U(cN, HN)=U(y+rW+ r(q-λ)P, λ) (3)

The household moves if

U(y+rW+ r(q-λ)P, λ) > U(y+rW-rM0, 1) (4)

Because q≤1, q<λ. The more ‘upmarket’ the household aims to move, the higher is λ. While the ‘larger’ house increases the household’s utility if it moves, moving comes at a sacrifice of other consumption if (λ-q)P>M0; note that the sale of the present house releases some wealth from housing, P-M0. Because (q-λ)P only affects utility in the new house, it is clear that the chances of moving increase when (q-λ)P is higher. Depending on preferences, some households are more willing to sacrifice other consumption for housing than others. A change in the household circumstances (e.g. the birth of a child, so that HN would be the ‘more appropriate’ house) might alter preferences to put more weight on housing, making them more willing to make the trade-off for the larger house[[4]](#footnote-4).

*Predictions*

Obviously there are many influences on residential movement. This very simple model does however suggest that a higher housing equity ratio would increase the propensity to move, either through relaxing the down-payment constraint or by increasing fungible wealth. Distinguishing these two reasons empirically is difficult, but we would expect the down-payment constraint effect to operate only amongst those households with very low levels of equity.

For a given q, the level of relative house prices in the market (P) exerts a negative impact on residential mobility by making it more likely that the down-payment constraint is binding, and the impact of house prices on mobility among the unconstrained is also negative because q<λ. The latter relationship reflects the fact that it is more expensive to move up-market (more other consumption must be foregone) in a housing market in which house prices are higher. Changes in nominal house prices over time affect q and may affect P. If nominal house prices increase purely because of general inflation, q rises because mortgages are written in nominal terms, but the relative price of housing (P) remains constant. An increase in relative house prices over time raises q and increases P, producing offsetting effects on mobility.[[5]](#footnote-5) The model also suggests that the size of the positive impact of the equity ratio on consumption is increasing in the level of house prices.

The analysis above has not considered the stimulating impact of mortgage default on mobility, which would be more likely for households with low values of equity. Default has been relatively low in Britain, but it is possible that for some households it is default that causes movement. Their existence tends to understate the impact of equity through the two other channels discussed above.

Empirical analysis

Our data is from the British Household Panel Study (BHPS), which interviewed people annually from 1991-2008 (see www.iser.essex.ac.uk/bhps/documentation). In addition to the housing equity and relative house price variables, these data provide other mobility relevant characteristics, such as the person’s age, number of children of particular ages in the household in the previous year, whether or not the person has a partner, real monthly household income in the previous year, change in household income, change in household size and (importantly) tenure in current residence. All of these variables are measured in year t-1 for the analysis of moves between year t-1 and year t. The ‘change variables’ are meant to capture possible ‘triggering events’ (allowing for asymmetric effects). We do not include contemporaneous changes in family status, such as divorce and childbearing, because we believe these to be endogenous processes, and these may indeed by influenced by housing equity. Unfortunately we do not have a measure for the household’s financial wealth on an annual basis, but we check for correlation between changes in net financial wealth and change in the equity ratio between two years in which we have wealth data.

We also allow for a fixed effect on residential mobility, which captures persistent unobserved influences on mobility, including long-term indicators of the level of wealth, income and preferences. That is, we use within-person variation to estimate the impacts of the explanatory variables on mobility. We can identify these parameters from multiple moves by some people. This is an advance on previous studies, which rely on cross-person variation, and we show below that allowing for fixed effects is important; in particular, ignoring them strongly biases the estimated impacts of housing equity and the relative house price toward zero.

*Measurement of ‘housing equity’*

The household’s equity in their house is measured as the homeowner’s estimated current value of their property (wHSVAL) minus the total mortgage on all property (wMGTOT) (names in parentheses are BHPS variable names). Because the former applies to their residence and the latter to all property, we confine the calculation and the analysis to homeowners with a mortgage who only own one property (wHS2OWND==2) and we exclude people living with their parents. The equity ratio (q in the theoretical model) is the ratio of this measure of equity to wHSVAL. We also know the time of purchase, the value of the house and the mortgage value at that time (information collected in the first interview after the purchase for those who purchased during the panel), as well as the type of dwelling (e.g. detached, terraced house, etc.).

There may be measurement errors in self-reported home values. We use the sales volume in the household’s local authority area and the original mortgage value as instruments for self-reported current equity, following in the footsteps of Ferraria et al. (2010).

*House prices*

Another key variable is the level of the relative house price (per unit of ‘housing services’) in the relevant market (P). We use an indicator of average market prices, drawn from Land Registry data at the local authority level (available from January 1995).[[6]](#footnote-6) On the one hand, this level of aggregation may be considered too small for the housing market relevant to mobility decisions, despite the fact that most moves are local. But these data have many advantages: (1) they are comprehensive (e.g. compared to data from particular large building societies, such as Nationwide or Halifax) ; (2) they are available monthly, allowing a close match to interview dates and flexibility in defining the time period for measuring house prices; and (3) they can be disaggregated by house type (e.g. detached, semi-detached etc.). Details on the data and how we match them to the BHPS data are supplied in Appendix A.

*Econometric analysis*

We take women as the ‘household marker person’ and confine the analysis to women aged under 45 not living with their parents, who are owners of a single home and have a mortgage, over the period 1992-2008.[[7]](#footnote-7) There are not substantive differences if men are chosen as the marker person. For instance, the pattern shown in Table 1 is very similar and we indicate below that estimates of the key parameters do not differ if men are taken as marker.

The annual residential movement rate is 8.6 percent (N=number of woman-years=9638). On the basis of changes in real house value, 87% of the moves are ‘upmarket’, in line with the assumption in our theoretical model that λ>1. Among the movers, one-fifth cease being a homeowner after the move, mainly moving into private rental housing. Some other analysts confine their analysis to ‘own-to-own’ moves (e.g. Englehardt 2003). In the British context, such censoring is less appropriate because renting in the private sector for long durations is much less common after a person has become an owner. The observed moves out of owner-occupation into private tenancy are likely to be mainly transitional moves on the way to buying another house. Of course, a small number of these may also reflect mortgage default. Our main analysis is on all moves among owners with a mortgage, but we also consider own-to-own moves separately.

There are some measurement problems for housing equity, which are very visible at the low end. For instance, there are 55 observations that have an equity ratio less than -0.5. These cases involve 52 individuals, indicating lack of persistence in these values. The large negative equity comes mainly from high reported mortgage debt: a median value of £200,000 for these cases compared with £45,000 for the rest of the sample (means of £444,000 cf. £57,000). Those with extreme values have a mean residential of tenure one year shorter than the rest. We exclude these observations.

The quartiles of the distribution of the equity ratio in the previous year are 0.25, 0.45 and 0.63, with a mean ratio of 0.44. Table 1 shows how residential movement varies with the equity ratio in the previous year.

Table 1: Residential movement by Equity Ratio (q) in previous year

|  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- |
| Move | **q>0.20** | **0.10<q≤0.20** | **0.05<q≤0.10** | **0.05<q≤0** | **q<0** |
| % move | 7.93 | 11.76 | 8.78 | 9.96 | 20.86 |
| N (100%) | 7879 | 986 | 353 | 281 | 139 |

It appears that the probability of movement is lowest for households with considerable housing equity (q>0.20). But this is likely to reflect other characteristics of women with high levels of housing equity (e.g. they are older, more likely to have a partner and less likely to have children aged under 12).

Our econometric model takes the following form:

Mit\*= γqit-1+ Xt-1β + αi + uit (1)

Mit\* is a latent variable for the propensity to move for person i between years t-1 and t, with movement taking place when Mit\*>0; qit-1 is the equity ratio in year t-1; the vector Xt-1 contains other relevant variables affecting the propensity to move, including our measure of relative house price; αi is person-specific fixed effect (αi may be correlated with qit-1 and Xt-1); and uit captures unmeasured residual influences on movement, where uit is assumed to have a logistic distribution. While the impact of the equity ratio is taken to be linear in equation (1), we consider other formulations, including categorical variables and linear splines.

The parameters are identified from a sample of women with at least one move among which some have at least two moves. This means that an individual must be observed at risk of a move in at least two spells with different values of the explantory variables. Among women who have at least one move, one-half of the observations are from women with 2 or more moves in the panel (35% have only two and the remaining 15% have more than 2 moves). Appendix B shows descriptive statistics. Table B1 provides these for the observations that are used in estimating the conditional (fixed effects) logit model, and Table B2 shows the statistics for the full sample, including persons who never move and those who move in the only pair(s) of years that we observe them. As must be the case, the estimation sample has a much higher movement rate, 17.9% compared with 8.4%, and of course a lower average residential tenure, 4.1 years compared with 6.4 years.

The first model in Table 2 shows estimates of the impact of the same categories of the equity ratio (from a conditional logit model) as in Table 1. In contrast to Table 1, it indicates a tendency for those with equity ratios below 0.10 to be less likely to move. As suggested earlier, we might expect the equity ratio to have a larger effect on movement for those with low levels of equity than those with higher equity because of the operation of the down-payment constraint. Model 2 examines this hypothesis by using a linear spline for the four categories of equity ratio.

Model 2 shows imprecisely estimated effects of the equity ratio, statistically indistinguishable from zero, for levels of housing equity below 20 percent, but strong positive effects of increases in equity above that level.[[8]](#footnote-8) The imprecision reflects the relatively small samples for equity ratios below 0.20 (see Table 1). Model 3 estimates a two-part spline for equity ratios below and above 0.20, and finds a similar impact of the equity ratio for the two groups. We cannot reject the hypothesis that the equity ratio has the same effects at both levels (p-value=0.97), and we constrain this to be so in Model 4[[9]](#footnote-9). When using men as markers, the estimate of the equity ratio parameter from Model 4 is 1.49 (SE=0.54) compared with 1.62 in Table 2.

**Table 2: Fixed Effect Logit Estimates of Impact the Equity Ratio (q) and house prices in the previous year on the log odds of residential movement\***

|  |  |  |  |  |
| --- | --- | --- | --- | --- |
|  | **Model 1** | **Model 2** | **Model 3** | **Model 4** |
| **0.10<q≤0.20** | -0.13  (0.21) |  |  |  |
| **0.05<q≤0.10** | -0.59  (0.34) |  |  |  |
| **0.05<q≤0** | -0.65  (0.39) |  |  |  |
| **q<0** | -0.53  (0.57) |  |  |  |
| **Linear spline:** |  |  |  |  |
| **q≤0.05** |  | -0.57  (3.16) |  |  |
| **0.05<q≤0.10** |  | 14.61  (8.86) |  |  |
| **0.10<q≤0.20** |  | -1.69  (3.25) |  |  |
| **q>0.20** |  | 1.71  (0.54) | 1.61  (0.53) |  |
| **q≤0.20** |  |  | 1.67  (1.37) |  |
| **q (linear)** |  |  |  | 1.62  (0.46) |
| **Log(real LA house pricet-1)** | -1.04  (0.43) | -1.32  (0.45) | -1.32  (0.45) | -1.32  (0.45) |

\*Number of women=427; number of woman years=3059. Standard errors are in parentheses. Other covariates are the woman’s age, age squared, number of children of particular ages in the household, whether or not the woman has a partner, real monthly household income, change in household income (positive and negative changes separately), change in household size (positive and negative changes separately), tenure in current residence and its square, and the local unemployment rate. All of these variables are measured in year t-1 for the analysis of moves between year t-1 and year t.

As noted earlier, previous analyses have used cross-section variation to estimate the parameters, implicitly assuming that any person-specific persistent influences on mobility are not correlated with the covariates in the model. We test the null hypothesis that the fixed effects (αi) are not correlated with qit-1 and Xt-1 using a Hausman test, which compares the conditional logit estimates of Model 4 in Table 2 with the corresponding random effect (RE) estimates, which use the cross-section variation in estimating the parameters as well as within variation. The test soundly rejects the hypothesis (p-value=0.0000).[[10]](#footnote-10) The test result suggests that previous estimates using cross-section variation are biased. Table 3 compares the conditional logit (FE) and RE estimates. It shows that the RE estimate of the impact of the equity ratio is virtually zero, implying that the correlation between the equity ratio (qit-1) and the person’s underlying mobility propensity (αi) is negative and large.[[11]](#footnote-11) Also, the RE estimate of the impact of the relative house price is one-fourth the size of the conditional logit estimate, which indicates a tendency for those living in housing markets with higher prices to have higher mobility propensities. The RE estimates also underestimate the size of the impacts of residential tenure and the unemployment rate.

**Table 3: Comparison of Fixed Effect and Random Effect Estimates on the log odds of residential movement**

|  |  |  |  |  |
| --- | --- | --- | --- | --- |
| Variable | **FE est.** | **RE est.** | **Difference\*** | **SE diff.** |
| qt-1 | 1.62 | -0.03 | 1.66\* | 0.39 |
| age | -0.56 | -0.29 | -0.27 | 0.15 |
| Age-squared/100 | 0.48 | 0.27 | 0.20 | 0.20 |
| Partnert-1 | -0.84 | -0.79 | -0.05 | 0.24 |
| Number of children aged: |  |  |  |  |
| 0-2 | -0.23 | -0.04 | -0.19 | 0.12 |
| 3-4 | -0.18 | 0.04 | -0.22 | 0.13 |
| 5-11 | -0.25 | -0.18 | -0.08 | 0.13 |
| 12-15 | 0.03 | -0.01 | 0.04 | 0.20 |
| 16-18 | -0.04 | -0.12 | 0.07 | 0.32 |
| Mo. household incomet-1 000s | 0.10 | 0.14 | -0.04 | 0.08 |
| Positive change in inct-1 000s | -0.01 | -0.05 | 0.04 | 0.09 |
| Negative change in inct-1  000s | -0.04 | -0.05 | 0.01 | 0.07 |
| Positive change in hh sizet-1 | -0.08 | 0.09 | -0.17 | 0.07 |
| Negative change in hh sizet-1 | -0.48 | -0.28 | -0.20\* | 0.10 |
| Res.tenuret-1 yrs. | 0.57 | 0.15 | 0.42\* | 0.06 |
| Res.ten.squared | 0.02 | -0.79 | 0.82 | 0.45 |
| Log(real LA house pricet-1) | -1.32 | -0.23 | -1.09\* | 0.42 |
| LA unemployment ratet-1 | -0.17 | -0.04 | -0.13\* | 0.06 |

\*Statistically significant difference at the 0.05 level.

Hausman test: chi-square (18 df)= 295.34 (p=0.0000).

*Impacts of housing equity and house prices*

Because each woman has her own underlying mobility rate (her fixed effect), average marginal effects cannot be calculated. But Model 4 suggests that for a woman with an underlying annual movement rate around the mean of 8.6%, an increase in their equity ratio of 0.10 would increase the probability of moving by 0.013 (0.086\*0.914\*1.62\*0.10), which is quite substantial.[[12]](#footnote-12)

Moves that involve a move out of owner-occupation to renting may be affected in a different way by the equity ratio and house prices than own-to-own moves. For instance, such moves could reflect mortgage defaults, or they may be discouraged less by low equity, provided that the household has some positive equity. We re-estimated Model 4 of Table 2 for own-to-own moves only, censoring other moves. The estimated impact of housing equity on the log-odds of own-to-own residential mobility is indeed larger than that for all moves: 1.89 (std. error=0.51), although the increase is only about one-half of the standard error.[[13]](#footnote-13)

With respect to the effect of the market price of housing services (P in the theoretical model), Table 2 indicates a negative effect of the average house price (relative to the retail price index) in the local authority area in the previous year in every model, as predicted from the theoretical model.[[14]](#footnote-14) We do not find evidence of a significant interaction effect between the house price and the equity ratio (p-value=0.49). In the model, an increase in the relative house price has two influences: it raises the equity ratio (e.g. given a 10% increase in house prices, an initial equity ratio of 20% would rise to 27%), which tends to increase mobility, but the house price increase also reduces mobility directly, because the household aims to improve its housing (λ>1). The parameter estimates in Model 4 of Table 2 suggest that these two effects are offsetting (0.086\*0.914\*1.62\*0.07=0.009 and -0.086\*0.914 \*1.32\* 0.10 = -0.010, respectively). But a rise in nominal house prices (matched by a rise in other prices) only affects mobility through the equity ratio effect, thereby raising it (e.g. a rise in the equity ratio from 20% to 27% arising from a general 10% increase prices would raise the mobility rate by 0.009).

There may be concern that house price effects may be partially reflecting local labour market effects. To address this, we have included the average local unemployment rate in the local area in the previous year in the mobility equation. A one percentage point higher unemployment rate is found to reduce the log odds of movement by 0.17 (std. error=0.07).

*Addressing Measurement error*

In order to address measurement error in self-assessed current home values, we instrument the equity ratio using the sales volume in the household’s local authority and in the previous year and the original mortgage value at purchase. As the under-identification test at the bottom of Table 4 shows, these are very strong instruments.

The instrumental variable (IV) estimates in column 2 of Table 4 are obtained from a fixed effects linear probability model, which can be interpreted as an estimate of the conditional expectation function for the movement variable. We use a GMM estimation procedure, which produces estimates that are efficient for arbitrary hetero-skedasticity and standard errors robust to heteroskedasticity. For comparison purposes, estimates in the first column of Table 4 show parameter estimates from an ordinary fixed effects linear probability model.[[15]](#footnote-15) The effects of the equity ratio and relative house price are roughly the same as the marginal effects in the conditional logit model when calculated around the mean movement rate. The IV estimate of the impact of the equity ratio in column 2 of Table 4 is more than six times as large as that in column 1, implying an increase of 0.066 in the mobility rate for a 0.10 increase in the equity ratio. The test of the over-identifying restrictions cannot reject the hypothesis that these are valid instruments (i.e. uncorrelated with the residual error term of the mobility equation). [[16]](#footnote-16)

Columns 3 and 4 of Table 4 repeat the exercise using the sample excluding non-movers and always-movers, which is similar to that in Table 2 (the sample is smaller than in Table 2 because of missing values on some instruments). While apparently very large, the IV estimates suggest that the estimates in Table 2 are likely to understate the true impact of the equity ratio, as we would expect with classical measurement error. Using a sample of male markers produces a pattern of estimation results similar to those in Table 4.

**Table 4: Fixed effects Linear Probability Estimates of the Impact of the Equity Ratio (q) in previous year on the probability of residential movement\***

|  |  |  |  |  |
| --- | --- | --- | --- | --- |
|  | **FE**  **Full Samp.a** | **FE-IVb**  **Full Samp. a** | **FE**  **Tb.2 Samp.c** | **FE-IVb**  **Tb.2 Samp.c** |
| **q (linear)** | 0.094  (0.032) | 0.628  (0.177) | 0.147  (0.059) | 0.735  (0.200) |
| **Log(real LA house pricet-1)** | -0.079  (0.028) | -0.227  (0.060) | -0.131  (0.052) | -0.249  (0.069) |

\*Robust standard errors are in parentheses. For other covariates, see Table 2.

aFull sample: number of women=1229; number of woman years=6553.

**b**Excluded instruments are the sales volume in the household’s local authority and original mortgage value in the previous year. GMM estimation.

Under-identification (rank) test: Kleibergen-Paap rk LM statistic (chi-square 2 d.f.) = 138.62 (p=0.0000).

Over-identification test of all instruments: Hansen J statistic (chi-square 1 d.f.) = 0.035 (p=0.852)

cTable 2 Sample, excluding those with missing values on the instruments: number of women=411; number of woman years=2939.

Under-identification (rank) test: Kleibergen-Paap rk LM statistic (chi-square 2 d.f.) = 134.99 (p=0.0000).

Over-identification test of all instruments: Hansen J statistic (chi-square 1 d.f.) = 0.022 (p=0.883)

As noted earlier, we do not have annual measures of financial wealth. As a consequence, we would overstate the impact of the equity ratio if it is positively correlated with financial wealth over time within persons. We check whether this might be a problem using the information from the BHPS on net financial wealth in 1995 and 2000. The correlation between change in a person’s net real financial wealth and their housing equity ratio between these two dates is only 0.02. Thus, we do not think that the absence of a financial wealth measure biases our estimates. Note that this because we only use within-person variation to estimate the parameters; it would be more likely to be a problem if we used cross-section variation because the level of net financial wealth and the equity ratio are correlated (e.g *r*=0.20 in the 2000 BHPS).

It might also be the case that people with a shorter expected residential duration after a particular move may choose different initial levels of equity when they purchase. To the extent that this expectation is persistent over time it should be captured in the fixed effect, but it is possible that some moves are perceived to be more temporary than others.

Some of the other variables have an important influence on mobility (see Table B3 in Appendix B). The older the woman, the less likely she will move, but the longer that she have resided in her current house, the more likely she will move. Women with a partner are much less likely to move, and the presence of children aged 5-11 also discourages movement, as does experience of a decline in household size in the previous year.

Conclusions

Our main finding is that an increase in a household’s housing equity encourages residential mobility, and a decline discourages it. Matching models of housing markets, such as Wheaton (1990), would then suggest that exogenous shocks to market house prices can be amplified because of the influence of equity on residential mobility. A boom will encourage mobility of existing owners, and the additional market turnover will raise house prices further. Correspondingly, a house price bust will be larger because of reduced residential mobility. We find that relative house price shocks do not have these amplification effects because the effects of equity and relative house prices given equity have offsetting effects.

We believe that our econometric estimates are more robust than earlier ones because we do not rely on cross-section variation in housing equity across households, but rather within-person variation and multiple moves by some persons. Indeed, our analysis rejects random effect or pooled estimates of the parameters, indicating that these bias the estimated impact of housing equity and relative house prices on residential mobility downwards substantially. Our estimates suggest that, given the relative house price, the impact of equity on movement is relatively large, with an increase of 0.10 in the equity ratio increasing the probability of moving house by about 0.01 for households with an average underlying mobility rate of about 0.09.

Appendix A

**Local authority house price data**

1. Official BHPS datasets mapping the 4-digit local authority district code to BHPS (wave-specific) household IDs (HID) are available under Conditional Access conditions (see www.esds.ac.uk). Files doing this are named alad.dta to rlad.dta, within which the LA id variable is called wLAD.
2. The Land Registry dataset contains monthly series from Jan 1995 onwards (www. www.landreg.gov.uk/house-prices). There are 172 localities at the lowest (least disaggregated) level, giving a total of 172 areas \* ((16 years \* 12 months)+2 months from 2011) = 33,338 observations. In addition, series are provided for 18 more aggregated areas: England and Wales as a whole, Greater London, 9 regions/unitary authorities, and 7 metropolitan counties (a further 3492 observations). The series provided are:
   1. House Price Index (HPI) with Jan 1995=100.
   2. Average price (£)
   3. Monthly change (%)
   4. Annual change (%)
   5. Sales volume (all sales)
   6. Average price of detached houses (£)
   7. Average price of semidetached houses (£)
   8. Average price of terraced houses (£)
   9. Average price of maisonettes/flats (£)
3. *Matching the data.* In total there are 376 local authorities in England and Wales at the lowest level, potentially represented in the BHPS. The Land Registry provide data for 172 local areas, the difference being that individual districts within non-metropolitan counties are not disaggregated. (For example, Preston is a district within Lancashire. The BHPS indicates a household is located in Preston via a 4-digit code, but house price data is only available for Lancashire as a whole.)We’ve created a look-up file that matches all 376 districts and their codes to the 172 areas in the Land Registry data. This matching was conducted using the information provided at <http://en.wikipedia.org/wiki/ONS_coding_system>.
4. The house price data are all in current prices. To deflate to real values we’ve downloaded monthly RPI data from the ONS website.
5. The real house price variable used in this study is the average for the local authority in question over the interview month t-1 and the preceding 11 months.

Appendix B: Descriptive Statistics

Table B1: Estimation sample (N=3059)

|  |  |  |  |  |
| --- | --- | --- | --- | --- |
| Variable | Mean | Std. Dev. | Min | Max |
| Move | 0.179 | 0.384 | 0 | 1 |
| qt-1 | 0.431 | 0.232 | -0.386 | 1.000 |
| Age | 33.968 | 5.792 | 19 | 44 |
| Partnert-1 | 0.827 | 0.378 | 0 | 1 |
| Number of children aged: |  |  |  |  |
| 0-2 | 0.174 | 0.399 | 0 | 2 |
| 3-4 | 0.166 | 0.390 | 0 | 2 |
| 5-11 | 0.477 | 0.755 | 0 | 4 |
| 12-15 | 0.168 | 0.454 | 0 | 3 |
| 16-18 | 0.026 | 0.163 | 0 | 2 |
| Mo. household incomet-1 000s  1991 prices | 2.407 | 1.397 | 0 | 22.207 |
| Positive change in inct-1  000s | 0.323 | 0.861 | 0 | 17.990 |
| Negative change in inct-1  000s | 0.272 | 0.825 | 0 | 17.076 |
| Positive change in hh sizet-1 | 0.121 | 0.350 | 0 | 5 |
| Negative change in hh sizet-1 | 0.095 | 0.411 | 0 | 6 |
| Res.tenuret-1 yrs. | 4.078 | 3.767 | 0 | 26.292 |
| Log(real LA house pricet-1) | 11.465 | 0.441 | 10.465 | 12.788 |
| LA unemployment ratet-1 | 2.837 | 1.789 | 0.417 | 13.750 |

Table B2: Sample including non-movers and those only observed moving (N=6912)

|  |  |  |  |  |
| --- | --- | --- | --- | --- |
| Variable | Mean | Std. Dev. | Min | Max |
| Move | 0.084 | 0.278 | 0 | 1 |
| qt-1 | 0.461 | 0.238 | -0.481 | 1.000 |
| Age | 35.528 | 5.814 | 18 | 44 |
| Partnert-1 | 0.853 | 0.355 | 0 | 1 |
| Number of children aged: |  |  |  |  |
| 0-2 | 0.137 | 0.360 | 0 | 2 |
| 3-4 | 0.148 | 0.369 | 0 | 2 |
| 5-11 | 0.567 | 0.790 | 0 | 4 |
| 12-15 | 0.263 | 0.547 | 0 | 3 |
| 16-18 | 0.041 | 0.209 | 0 | 2 |
| Mo. household incomet-1 000s  1991 prices | 2.334 | 1.349 | 0.001 | 22.207 |
| Positive change in inct-1 | 0.306 | 0.810 | 0 | 17.990 |
| Negative change in inct-1 | 0.259 | 0.852 | 0 | 24.516 |
| Positive change in hh sizet-1 | 0.095 | 0.314 | 0 | 5 |
| Negative change in hh sizet-1 | 0.093 | 0.406 | 0 | 7 |
| Res.tenuret-1 yrs. | 6.363 | 5.278 | 0 | 36.292 |
| Log(real LA house pricet-1) | 11.423 | 0.434 | 10.453 | 12.830 |
| LA unemployment ratet-1 | 3.096 | 1.948 | 0.417 | 14.433 |

Table B3: Full set of parameter estimates for Model 4 of Table 2\*

|  |  |  |  |  |
| --- | --- | --- | --- | --- |
|  | **Coeff.** | **Std, Error** | **z** | **P>|z|** |
| qt-1 | 1.62 | 0.46 | 3.52 | 0 |
| Age | -0.56 | 0.17 | -3.25 | 0.001 |
| Age-squared/100 | 0.48 | 0.24 | 1.96 | 0.05 |
| Partnert-1 | -0.84 | 0.27 | -3.13 | 0.002 |
| Number of children aged: |  |  |  |  |
| 0-2 | -0.23 | 0.18 | -1.27 | 0.205 |
| 3-4 | -0.18 | 0.18 | -1.04 | 0.299 |
| 5-11 | -0.25 | 0.15 | -1.72 | 0.085 |
| 12-15 | 0.03 | 0.23 | 0.15 | 0.88 |
| 16-18 | -0.04 | 0.44 | -0.1 | 0.921 |
| Mo. household incomet-1 000s | 0.10 | 0.09 | 1.04 | 0.298 |
| Positive change in inct-1 | -0.01 | 0.11 | -0.05 | 0.958 |
| Negative change in inct-1 | -0.04 | 0.10 | -0.39 | 0.698 |
| Positive change in hh sizet-1 | -0.08 | 0.17 | -0.48 | 0.634 |
| Negative change in hh sizet-1 | -0.48 | 0.16 | -3.01 | 0.003 |
| Res.tenuret-1 | 0.57 | 0.07 | 8.29 | 0 |
| Resten.squared | 0.02 | 0.50 | 0.05 | 0.96 |
| Log(real LA house pricet-1) | -1.32 | 0.44 | -2.97 | 0.003 |
| LA unemployment ratet-1 | -0.17 | 0.07 | -2.4 | 0.016 |

\*Conditional fixed-effects logistic regression:

Number of obs =3059

Number of groups = 427

Obs per group: min = 2

avg = 7.2

max = 13

LR chi2(18) = 456.24

Log likelihood = -687.12104 Prob > chi2 = 0.0000



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1. Property transactions are not exclusively related to residences, but the vast majority are (e.g. 1.4 million out of 1.6 million in 2002. [↑](#footnote-ref-1)
2. The real price of existing houses for 1959-1983 comes from Holmans (1988), Table A1; and the nominal price of existing houses for 1983-2008 is from the Halifax price index for existing homes, which is deflated by the RPI, excluding housing costs. [↑](#footnote-ref-2)
3. A positive correlation between the real house price and property transactions is also consistent with Stein’s (1995) theoretical model, which emphasises a down payment requirement and the existence of liquidity constraints for some households. [↑](#footnote-ref-3)
4. Our focus in this model is on the effects of wealth constraints on mobility as captured by the equity ratio. Household members also potentially face income-related constraints associated with debt-financing of residential property, such as maximum mortgage payment-to-income ratios (Linneman and Wachter, 1989). We included controls for the level of and recent changes in household income in our empirical models and the effects were not significant, providing suggestive evidence that income-related constraints are not important in our sample. [↑](#footnote-ref-4)
5. There are other ways of parameterising the model, making movement a function of the mortgage amount and the relative house price. Other studies, such as Englehardt (2003) and Ferraria et al. (2010) exclude the relative house price from their mobility equations, but this ignores the impact of the trade-off between housing and other consumption when moving ‘upmarket’ in housing. [↑](#footnote-ref-5)
6. An unfortunate consequence of the lack of Land Registry house price data for 1991-94 is that this was a period when nominal house prices were falling in most areas, which produced important within-person variation in housing equity. [↑](#footnote-ref-6)
7. The marker person approach is necessary to avoid the ‘double-counting’ of moves made jointly by individuals in a couple. Sample size is larger using women, in part because of more attrition among young men: 7887 man-year observations compared with 9638 woman-years. In the fixed effect analysis, there are 361 men compared with 427 women. When using men the overall residential mobility rate is 9.1 percent. [↑](#footnote-ref-7)
8. Indeed we cannot reject the hypothesis that the equity ratio has the same effects at every level (p-value=0.54). [↑](#footnote-ref-8)
9. If we exclude the relative house price and the local unemployment rate from the mobility equation, the impact of the equity ratio on the log-odds of movement is slightly smaller: 1.34 (SE=0.45). [↑](#footnote-ref-9)
10. Similar results emerge when comparing the conditional logit estimates with ordinary logit estimates, a cross-section estimator which has been used in previous studies [↑](#footnote-ref-10)
11. When using male markers, the RE estimate of the impact of the equity ratio is again significantly smaller: 0.33 (SE=0.27) compared with 1.49 for the conditional logit estimate, [↑](#footnote-ref-11)
12. The average marginal effect is calculated using the formula P\*(1-P)\*βj\*ΔXj where P is the expected probability of success, βj is the coefficient of interest on the variable Xj and ΔXj is the magnitude of the change in Xj whose effect we wish to evaluate.  [↑](#footnote-ref-12)
13. The size of the negative impact of house prices on own-to own moves is smaller than that for all moves: -0.86 (0.47), although not significantly so in the light of the standard error of the estimate. The smaller impact could reflect a large impact of high rents in housing markets with higher house prices, causing house prices to also discourage movement involving a change in tenure. [↑](#footnote-ref-13)
14. The estimated parameter (-1.39) is very similar when we use the average house price of the type of dwelling the household lived in, and the standard error is lower (0.27). [↑](#footnote-ref-14)
15. The larger reported ‘full samples’ in Table 4 than Table 2 reflects the fact that the linear fixed effect estimates do not drop the non-movers and those who are only observed moving. These women contribute little to the identification of the parameters. [↑](#footnote-ref-15)
16. The GMM estimate of the impact of the local unemployment rate is virtually zero. [↑](#footnote-ref-16)