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IFS Working Paper W14/25

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HOUSE PRICES, WEALTH EFFECTS AND LABOR SUPPLY

by

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Revised Version: September 2014

Abstract

We examine the impact of housing wealth on labor supply decisions using data on exogenous local variation in house prices merged into household panel data for Britain. Our estimates are conditioned on variations in local labor demand and income expectations as these may co-determine housing wealth and labor supply. We use renters as a control group and test for the potential endogeneity of tenure and location. We find significant housing wealth effects on labor supply among young married / co-habiting female owners and older male owners, consistent with leisure being a normal good. The size of these effects is economically important. Our estimates imply housing wealth effects have stronger effects then local labor market conditions upon participation decisions for these workers.

Key words

Labor supply; Wealth effects; House prices

JEL classification

D12 E21 J22

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Earlier versions of this paper were presented at the Society of Labor Economics Annual Conference 2014, the Royal Economic Society Conference 2014, at the Institute for Fiscal Studies, London, the National Institute of Economic and Social Research, London, and the Universities of Nottingham, St Andrews, Stirling and Sussex. Our thanks go to seminar participants and others for comments on earlier drafts, notably Richard Blundell, Chris Carroll, Carl Emmerson, Eric French, Mike Shields, Jim Spetzler and Gemma Tetlow. The data and tabulations used in this paper were made available through the ESRC Data Archive. The data were originally collected by the ESRC Research Centre on Micro-social Change at the University of Essex (now incorporated within the Institute for Social and Economic Research). Neither the original collectors of the data nor the Archive bear any responsibility for the analyses or interpretations presented here. All remaining errors are the responsibility of the authors.

HOUSE PRICES, WEALTH EFFECTS AND LABOR SUPPLY

1. Introduction

This paper estimates the size of housing wealth effects on labor supply for a panel of households in Britain. The motivation for this paper is the following. Several recent studies based on micro-data have shown that changes in housing wealth affect consumption spending and household indebtedness. A typical finding, however, is that the average response is small and that the effect is disproportionately concentrated among a minority of credit constrained households.[1] One reason for this finding might be that, for some types of households, housing wealth changes primarily affect labor supply – for example, the decision to retire. Hence, housing wealth gains cause some households to reduce income as they take more leisure, instead of increasing consumption for a fixed income and labor supply. Leisure, like consumption, is typically thought of as a normal good so we might expect housing wealth gains to increase leisure and decrease labor supply for some households, and *vice versa*, for housing wealth losses.

Motivated by this research question we consider whether household labor supply in Britain responds to housing wealth gains and losses and also whether there are heterogeneous responses across different types of households. Our results show that labor supply responses to changes in housing wealth are highly heterogeneous across household types. We find most households do not adjust their labor supply when house prices change, but for subsets of households there are large effects. The household types that show significant responses to housing wealth changes are those at the margins of labor supply: married women, at the margin of household labor supply; and men close to retirement at the inter-temporal margin

[1] Recent studies on the impact of house prices upon household consumption and saving include Campbell and Cocco (2007), Disney, Gathergood and Henley (2010), Attanasio, Leicester and Wakefield (2011), Carroll, Otsuka and Slacalek (2011), Browning, Gørtz, and Leth-Petersen (2013), Mian, Rao and Sufi (2013), Cooper (2013); on indebtedness see Hurst and Stafford (2004), Disney and Gathergood (2011) and Mian and Sufi (2011).

of lifetime labor supply. The prior literature shows these households are responsive to changes in marginal tax rates. We show wealth effects are also important for understanding the labor supply decisions of these households.

The effects we find are economically significant. We find that a 10% rise in local house prices relative to the national trends is associated with a reduction in the labor market participation rate among young married / co-habiting women of 1.7% and a reduction in the participation rate among older men of approximately 4.5%. Therefore, our results show that house price changes have distributional effects on labor supply (as well as consumption) which correlate with life-cycle characteristics. These may also arise due to collateral effects on labor supply choices among younger workers. Hence there is a life-cycle as well as an overall effect of house price changes on labor supply.

The paper proceeds as follows. Section 2 describes our broad modeling strategy. The British Household Panel Survey, which we use to estimate labor supply equations for various dimensions of labor supply, is described in Section 3. Section 4 describes our econometric model and our identification strategy. Section 5 describes our main results concerning participation and hours. Where we find that house price gains (losses) lead to reduced (increased) labor market participation, we then investigate the types of activities individuals undertake when they withdraw from the labor market. Section 6 tests some alternative specifications and applies a sensitivity analysis to our main findings. Section 7 provides a brief summary and conclusions.

2. Modeling strategy

Existing studies of wealth effects on labor supply based on exogenous wealth changes such as lottery wins (Imbens et al., 2001; Cesarini et al., 2013) and inheritances (Joulfaian and Wilhelm, 1994; Brown et al., 2010) in general confirm the intuition that labor supply falls when wealth increases. Moreover, studies on United States (US) data have shown that housing wealth changes impact on decisions that have implications for labor supply without estimating labor supply effects directly. Lovenheim (2011) shows that increases in housing wealth raise college enrolments while Lovenheim and Reynolds (2013) find that housing wealth gains increase the likelihood of enrollment at public flagship universities. Lovenheim and Mumford (2013) show that housing wealth gains also raise the likelihood of home owners choosing to have children. College enrollment and childbearing are both likely to affect household labor supply.

Prior studies also show movements in non-housing wealth, and therefore possibly housing wealth also, are important at the margin of retirement timing. Blundell *et al* (2013) for the UK, and French and Benson (2011) and Daly, Kwok and Hobijn (2009) for the US all argue that asset price declines may be one reason why labor supply in the post-2008 recession remained higher than in previous recessions due to delayed retirement. However Coile and Levine (2011), find evidence that labor market changes dominate non-housing asset (wealth) effects in explaining patterns of retirement over the business cycle[2]. One prior study on the role of housing by Farnham and Sevak (2007) using an earlier sample of US data from the Health and Retirement Study finds house price gains typically cause households to bring forward their intended retirement date.

An important issue in this context is that wealth effects on labor supply should be identified only off exogenous shocks. In the canonical life cycle model, consumption, wealth accumulation including housing wealth, and labor supply are simultaneously determined. Households may, for example, work more in order to acquire a more expensive house. They are also likely to anticipate that their existing stock of housing wealth may grow in value over time due to the overall relative growth in the price of housing, and understand that house

[2] For related studies based on UK data see Disney, Ratcliffe and Smith (2013) and Crawford (2013)

prices are broadly pro-cyclical in nature. Figure 1 illustrates the pro-cyclicality of house prices in the UK using de-trended data. House prices are strongly pro-cyclical and more variable than GDP. The correlation coefficient between house prices and GDP is 0.6 over many periods of business cycle fluctuations[3]. Hence it is reasonable to assume that households understand the trend and cyclicality of house prices. Modeling the 'exogenous' component of house price changes to households is therefore an important practical issue.[4]

It is not possible to randomly assign housing wealth. In our baseline model we utilize changes in local house price indices conditioned on time, household and neighborhood effects, as our measure of exogenous variation in house prices. We do not use self-reported housing wealth as this may be accumulated endogenously. We also control for tenure and locality choices. Therefore our primary source of identification arises from differential changes in house prices across localities relative to average house price changes controlling for neighborhood effects (such as local amenities which may affect house price levels in the area) and household preferences.

In taking this approach, we assume that households form a general expectation of broad house price trends (e.g. from discussion in the news media) and that the exogenous component of housing wealth changes arises from realised local variations in the rate of change of house prices relative to this national trend.[5] We believe that it is reasonable to assume that households can identify this local component *ex post* from posted prices by local realtors ('estate agents' in British parlance) and online property search engines that provide valuations of existing properties.

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[3] The figure plots the percentage deviation from trend for UK real house prices and real GDP. House prices are more volatile than GDP. The percentage standard deviation from trend in house prices expressed as a percentage of the percentage standard deviation in trend in GDP is 376%.

[4] And one pertinent for other measures of exogenous wealth shocks insofar as inheritances and even lottery wins may be anticipated – arguably it is only the *timing* of such events that is unknown.

[5] For further discussion of issues concerning the modeling of income and house price expectations, see Browning, Gørtz, and Leth-Petersen (2013), and Disney, Gathergood and Henley (2010).

A relaxation of our modeling strategy would be to assume that households do not anticipate average fluctuations in house prices over the business cycle. This is equivalent to removing time effects from the model. We see this as an unrealistic approach to how households form expectations of house price movements. Under this approach our results would then have an interpretation identical to that of the US studies cited earlier, and can be used to show how wealth effects moderated fluctuations in labor supply over the course of the economic cycle. We discuss the effects of such a change to our modeling strategy in Section 5.

To examine the effect of asset prices on labor supply, we must also control for variation in local labor demand, given the likely covariance of shocks to asset and labor markets which might co-determine local house prices and labor supply decisions. This provides an additional quasi-identification strategy for our modeling, insofar as local labor market conditions should have an effect on the labor supply of *renters* which is independent of house price changes. Hence, once we control for local labor market conditions, the labor supply of renters should be unaffected by any local movements in house prices. Indeed insofar as renters may include some would-be homeowners, we might expect that higher house prices, while inducing lower labor supply among owners, might induce *higher* labor supply among some renters. Inclusion of renters provides a test that our measured housing wealth effects for owners are not contaminated by local labor market effects.

A particular advantage of our UK panel is that it includes individual-level income expectations data. This is important as income expectations may explain a negative correlation between housing wealth and labor supply. In inter-temporal models of labor supply higher expected future income (arising, for example, from higher expected future wages) might induce workers to reduce current labor supply. Higher expected future income also increases current consumption and raises current housing demand (to smooth housing consumption), hence increasing house prices. Elsewhere, we show that failing to control for income expectations causes upward bias in the estimated housing-consumption wealth effect (Disney et al., 2010). Attanasio et al. (2011) come to a similar conclusion using a calibrated model. Individual-level income expectations data is not available in US household panels covering the working age population.[6]

Finally, our modeling strategy has to allow for potential endogeneity of housing tenure status and also that migration between localities may induce a potential bias into our estimates. Suppose that households move to localities where there are increased work opportunities. Given that housing supply is very inelastic, we would expect such localities also to exhibit faster rises in house prices relative to the national average. Hence, worker mobility may induce increased measured hours of work or participation probabilities which correlate with local house prices increasing above trend. This 'migration effect' will then bias the local 'house price effect' downwards. We discuss our strategy for dealing with this issue in due course.

3. Data Sources

We use United Kingdom (UK) data combining variation in house prices across geographic localities with household panel data to estimate exogenous housing wealth effects on labor supply - both for total hours and separately at the extensive margin (participation). We show results for total household hours, by household type, by age and whether single or couples, and for men and women separately to identify differential responses within couples. We also examine the impact of housing price shocks on routes in and out of the labor market – for example, into retirement and full-time child care.

[6] The US Health and Retirement Study now incorporates a wide-ranging module of questions on individual expectations but the sample is limited to older individuals.

Our primary data set is the British Household Panel Survey (BHPS). The BHPS is a high-quality source of panel data on work activity and is commonly used in studies on labor supply in the UK as in, for example, Blundell et al. (2008). The BHPS is an annual survey of each adult member (16 years of age and older) of a nationally representative sample of more than 5,000 households, comprising a total of approximately 10,000 individual interviews.

Major topics covered in the survey are household composition and demographics, participation in the labor market, income, wealth and housing. The same individuals have been re-interviewed in successive waves and, if they split-off from original households, all adult members of their new households have also been interviewed. Children are interviewed once they reach the age of 16. The sample is representative of the population of the United Kingdom. We use 18 waves of data that are available from 1991 to 2009.

The sample used here is the head of household and spouse or live-in partner only, aged 18-75. We limit the top age to 75 as 99% of BHPS respondents are retired by that age and our interest is in labor market participation and hours of work. We exclude the self-employed as the relationship between house prices and self-employment has been considered elsewhere (for example, Hurst and Lusardi, 2004, and Disney and Gathergood, 2009) and our focus is on participation in the labor market for non-self employed individuals.

The labor market status measure in the dataset is a question on the individual's current activity from which they choose one from the following menu of options: self-employed / in paid employment / unemployed / retired / family care / full time student / long-term sick or disabled / maternity leave / government training scheme / other status. Hours of work are measured in the data set as the sum of hours normally worked per week plus overtime hours for first and second jobs.[7] We define an individual as participating in the labor market is they

[7] Individuals who report they are suffering short-term sickness leave from work or are on vacation from work are classified by their regular labor market status (employed or self-employed).

report their labor market status as 'in paid employment' or 'unemployed'. This is our measure of labor supply at the extensive margin. We define hours of work as the sum of weekly hours plus 'overtime' hours for all jobs worked by the individual. This is our intensive margin labor supply measure.

The financial expectations measure included in the survey is an individual level answer to the question: 'Looking ahead, how do you think you yourself will be financially a year from now, will you be better than now / worse than now / about the same?' Although this question is asked only of a short time-frame, it captures something of changes in the household's income expectations which might cause changes in labor supply in the current period and is similar to those used in consumer confidence indices[8]. We take answers to this question and code two 1/0 dummy variables for 'positive financial expectations' and 'negative financial expectation' which we include in our econometric specification, allowing the labor supply responses of individuals to positive and negative expectations to differ in sign and magnitude.

Instead of using self-reported house values in the BHPS for our measure of housing wealth, we use county-level house price data from a separate source. This approach, which is similar to that used by Lovenheim and Reynolds (2013) and Farnham and Sevak (2007), has two purposes: first, it gives a measure of exogenous variation in house prices and second, it allows us to assign a proxy measure of the cost of housing for renters for our test of whether local house price changes proxy changes in local economic conditions. Hence we match into the BHPS survey data local level house price data derived from house price sales.

[8] For example, the question about future income expectations in the Michigan Survey of Consumer sentiment is 'During the next 12 months, do you expect your (family) income to be higher or lower than during the past year?'

Our house price data is the Halifax county-level house price index provided by Halifax Bank of Scotland (now part of the Lloyds banking group), the UK's largest mortgage lender.[9] The Halifax index comprises standardized house prices which reflect the sale price of a medium-sized family home in each county in each year.[10] Throughout we adjust all financial variables to 2000 prices using the Retail Prices Index. We also match into the BHPS two county level variables which capture local labor market conditions: first, registry unemployment data provided by the Office for National Statistics (ONS) and second, county level average earnings derived from the ONS Annual Survey of Hours and Earnings employer survey.[11]

Summary statistics for key variables appear in Table 1. All financial variables are adjusted to year 2000 prices. Our dataset comprises approximately 135,000 individual-year observations, 56% of which are for men and 77% of which are for married survey respondents. The average age of a respondent to the survey is 47.2 years. A little less than 60% of the individual-year observations are for workers in employment (this employment rate is lower than the 70% in the working age population as our sample includes individuals up to 75 years of age and in total 26% of our sample are retired at the point of interview). A little more than two-thirds of individual-year observations in our sample are for home owners with the average house value among owners at £133,000.

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[9] On average the population of county in the UK in 2012 is 880,000 individuals comparable to the population of a US Metropolitan Statistical Areas which average 700,000 individuals in 2012.

[10] Choice of spatial aggregation for a house price index involves a trade-off between locality of the house price index (i.e. an index which provides very localised house price data) and volume of observations (which are larger at the broader geographic level). The level of disaggregation may be one reason why studies differ in the magnitude of their house price effects. Regional data for the 9 English regions offers more observations per geographic unit but is too aggregated in a UK context. Local Authority (district) data is more localised with 326 individual authorities. However, house sale sample sizes are very low in some authorities. The Nationwide Building Society provides a local authority house price dataset but omits price data for 16% of authority-year cells due to small sample size. Therefore we choose the county level with 60 English county units as an appropriate geographic aggregation which balances locality with sample size.

[11] County level average earnings from the Annual Survey of Hours and Earnings (named the New Earnings Survey pre-1997) is calculated as average full-time monthly pay for all individuals participating in the survey which covers a 1% sample of employee jobs in the UK on an annual basis. Earnings data is derived from confidential workplace surveys in which employers report wages paid to employees.

4. Econometric Model

This section explains in detail our approach to identification and estimation. We incorporate local house prices into our estimation strategy because changes in self-reported housing wealth may be endogenous to individual labor supply decisions if individual work decisions cause changes in housing wealth, such as if a worker increases hours of work to purchase a larger house or, possibly, reduces hours of work to undertake home improvement. Changes in local level house prices are exogenous to individual preferences for leisure, housing and non-housing consumption, though moving activity may not be and we address this in our identification strategy.

The baseline econometric specification that we use to model the relationship between housing wealth, proxied by local house prices, and labor supply is:

$$h_{ict} = \alpha + \beta_1 H_{ct}{*}O_{ict} + \beta_2 H_{ct}{*}R_{ict} + \beta_3 O_{ict} + \beta_4 U_{ct} + \beta_5 E_{ct} + \beta_6 X_{ict}$$

$$+ \beta_7 F_{ict} + \beta_8 \hat{W}_{ict} + \varphi_i + \theta_c + \psi_t + \varepsilon_{ict}$$

$$\hat{W}_{ict} = \beta_1 Z_{ict} + \varepsilon_{ict} \qquad (1)$$

Where i denotes an individual, c denotes county of residence and t denotes year. The (log) of annual hours for all employed individuals with non-zero hours is denoted h_{ict}. O_{ict} is a 1/0 dummy variable indicating that the respondent is a home owner and R_{ict} is a 1/0 dummy variable indicating the respondent is a renter. The variable H_{ct} is the (log) average house price at the county level in each year, U_{ct} is the local unemployment rate at the county level in each year, E_{ct} is (log) average earnings at the county level in each year, X_{ict} is a set of individual level socio-economic characteristics and control variables and F_{ict} is the individual's self-reported financial expectation. Since self-reported hourly wages, $W_{ict,}$ may be endogenous to labor supply if individuals face downward sloping labor demand curves (i.e. reducing hours of work increases the hourly wage), we instrument hourly wages using a human capital regression as in MaCurdy (1981) and Altonji (1986). Hence Z_{ict} is a vector of first-stage instruments in the wage equation. Following MacCurdy (1981) we use age and human capital measures as instruments.

To interpret the coefficient β_1 as representing the causal impact of housing wealth on labor supply requires that the estimated impact of local house prices on labor supply is not attributable to omitted variable(s) which might drive both house prices and labor supply for which house prices might be a proxy. The identifying assumption in Equation 1 is that, conditional on county fixed effects θ_c and year fixed effects $\psi_{t,}$ plus the vector of time-varying control variables X_{ict}, the local unemployment rate $U_{ct,}$ local average earnings $E_{ct,}$ the individual's financial expectation $F_{ict,}$ predicted wages $\hat{W}_{ict}$ and time-invariant individual characteristics captured by the individual fixed effects φ_i, house price variation across counties over time is exogenous to individual labor supply.

A further robustness check incorporates renters into the estimation. Renters experience the same local economic conditions as home owners but do not experience direct wealth gains and losses from house prices. Thus, conditioning on controls, renters should respond differently to owners in respect to house price changes. If renters intend to buy in future then indirect wealth gains and losses arising from local house price changes are in the opposite direction to those experienced by current owners. Hence, if the coefficients β_1 and β_2 are both non-zero and equal (i.e. the estimated impact of county house prices on the labor supply of owners and renters is identical) then we would conclude that county house prices proxy for unobserved local conditions. If they are both zero, we would conclude that house prices have no impact on work decisions. If β_1 is negative and β_2 is either zero or positive, we would argue that we have identified a negative wealth effect on labor supply arising from (changes in) housing wealth.

Incorporating renters into our estimation as a comparison group, however, requires that the coefficients on the interaction terms β_1 and β_2 reflect the differential responses of owners and renters to house price gains and losses due to their homeownership status and not due to other characteristics which differ between owners and renters (such as age and income). Where owners and renters differ in these other characteristics, the coefficients on β_1 and β_2 might reflect the impact of these other characteristics in the relationship between house price and labor supply, hence confounding our model. Accordingly, in our estimates, and in an extension to equation (1), interaction terms between the house price variable and *all other* covariates are included in the model.

In addition, two sources of selection bias might confound estimates of Equation (1). First, county-level house price changes are not exogenous for individuals who move county. Selection bias would occur if individuals moved to higher house price counties and simultaneously changed their labor market participation. To eliminate any bias arising from moving behavior we use two strategies.

In the first strategy, we exclude cross-county movers (dropping approximately 5% of the individual-year observations in our sample). We show the omission of these households does not change our results. In the second strategy, we keep cross-county movers in the sample but calculate the counterfactual house price change (they would have received had they not moved county) and use this simulated change in house prices to estimate Equation (1) instead of their actual cross-county change. This strategy shows very similar results to out baseline estimates.

Second, selection bias would arise if house price changes caused individuals to change from renting to owning and the likelihood of changing tenure were related to labor supply. We address this in two ways following the approach of Lovenheim (2011). First, we use initial homeownership status of the household (i.e. homeownership status in the first wave in which the individual is observed) rather than contemporaneous housing tenure in our specifications in order to eliminate housing tenure changes that might cause selection bias. Second, we use initial home ownership status as an instrument for contemporaneous housing tenure, assuming initial home ownership status is exogenous. We show both strategies yield estimates of β_1 and β_2 which are very similar to those using contemporaneous housing status.

We also run equations at the extensive margin where we estimate the linear probability of an individual participating in the labor market. As we use a fixed effects panel estimator we are thereby estimating labor market transitions. In similar vein, and corresponding to some of the existing literature, we also estimate transition equations into other non-participation labor-market inactive states, specifically the categories of 'retirement' and 'family care'.

We estimate all the models using (within) fixed effects estimation and use a linear estimator throughout. As the house price variable and unemployment variable are both defined at the county level we calculate standard errors clustered at the county level. We have also calculated estimates with standard errors clustered at the region level to allow for wider geographic house price correlation and find very similar results. We also apply a standard bootstrap technique to our econometric estimates.

5. Results

5.1 House Prices and Hours of Work

We first show results for the impact of house prices on hours of work. Table 2 shows estimates for the hours equation for sub-samples of individuals defined by single or married / co-habiting, gender and age. Only individuals with non-zero hours of work are included in the estimation sample. Each column of Panels A and B shows results from a separate model where Panel A includes individuals who are married / cohabiting and Panel B includes single individuals. Within each panel, results are shown for sub-samples defined by gender and three age categories. We report coefficients on the house price interaction terms for home owners and renters, the county unemployment rate and the financial expectations variable. Variables not shown in the table of results are listed below the table.

Results show that for all groups other than young married / co-habiting women house prices have no impact upon hours of work. None of the estimated coefficients on either the owner or renter house price interaction terms are statistically significant at the 5% level and the p-values from t-tests for equivalence of means between the renter and owner coefficients fail to reject the null that coefficients for the two groups are the same.

However, we do find statistically significant results for young married/ co-habiting women. The coefficient on the home owner house price term is negative and statistically significant at the 0.1% level. The coefficient on the renter house price term is positive and statistically not significantly different from zero. The p-value from the test for equivalence of coefficients is below 0.0000 implying these coefficients are significantly different from one another at a very high level of confidence.

The coefficient on the home owner house price term takes a value of -0.182. Hence a 10% increase in house prices leads to a reduction in hours for married / co-habiting young female home owners of 1.8%. Average (non-zero) annual hours for this group in our sample is 1,485. Hence a 10% increase in prices reduces annual hours by 27 hours per annum, a little below one working week of hours on average for this group.

For young married / co-habiting men and women the coefficient on the financial expectations variable is negative and statistically significant at the 5% level. This provides some evidence for intertemporal substitution of hours of work: individuals with positive expectations about their future finances work fewer hours in the current period. The coefficients on the financial expectations variable are also negative for young single men and women but in both cases are not statistically significant.

Table 3 shows results from a series of robustness specifications test our results for young married / co-habiting women. There are five alternative specifications in the table[12]. The first two specifications relate to home moving activity. In the first column individuals who move home (approximately 10% of the sample) are excluded. In the second column for individuals who move county we use the counterfactual house price of their former county in all waves in which they are present in the panel (i.e. allocate to that individual a house price as if they had not moved county).

Results show that when movers are excluded from the sample the owner county house price term remains negative, statistically significant at the 0.1% level and significantly different from the (not significant) coefficient on the renter county house price term. The absolute value of the coefficient is a little larger than in the baseline specification (-1.86 compared with -1.82) confirming our priors that including movers biases the coefficient estimate downwards. When simulated prices are used, the same pattern of statistical significance remains by the absolute value of the coefficient falls a little (to -1.78). Overall, therefore, we find no evidence for moving activity confounding the main estimates presented in Table 2.

The next two columns of Table 3 show results from the robustness specifications relating to housing tenure and tenure-switching activity. In the first column homeownership status of the individual is fixed to be their home ownership status in the first wave in which they are observed in the survey. This is a similar approach to that of simulating county house

[12] We have estimated models for each of these specifications for each of the sub-samples presented in Table 2 (and in the remainder of the paper for the labor market participation models). Due to space constraints we do not show all estimates in the tables accompanying the paper (the full set of robustness estimates for Table 1 alone sums to 60 extra models) but instead only show robustness estimates for sub-samples where the main specification returned results of interest. The replication files include robustness estimates (and region level cluster standard errors estimates which also do not change our main results) for all sub-samples.

prices for movers in that we build a counterfactual status for the individual had they not entered into the activity which might confound our estimates (moving in the previous case, tenure changing in this case). In the second column 'IV Owner' this approach is implemented as an Instrumental Variables regression where current housing tenure is instrumented using initial housing tenure. Coefficient estimates in both columns are quantitatively very similar to the main specification results and show the tenure changing activity does not confound our main estimates.

The final column of Table 3 shows results from the 'falsification test' where the contemporaneous house price is replaced with the one-period forward house price. This is to test whether future house price affect current labor supply, which might indicate a spurious relationship due for example house prices proxying for household wealth. However, in this specification neither of the house price terms for owners or renters return statistically significant coefficients.

Results from estimates for hours of work show, therefore, house price gains cause reduced female labor supply among home-owning married or co-habiting couples. This result is consistent with a model in which house price gains operate a wealth effect at the variable margin of adjust of household labor supply, which is typically hours of work for the female worker. Later we return to the issue of what form of activity (or leisure) females might substitute towards as a result of this wealth effects.

5.2 House Prices and Labor Market Participation

Next we present results for decision to work on the extensive margin. Table 4 presents estimates from the participation equation, where the labour market participation dummy variable takes a value of 1 if the respondent is employment or unemployed, and takes a value of 0 otherwise. We estimate Linear Probability Models with individual fixed effects plus county and time effects following the hours of work specification shown earlier. Results are shown by sub-groups using the same convention as in Table 2 with sub-groups defined by relationship status, gender and age.

Results show house price gains decrease the likelihood of participation among young married / co-habiting women and older men both married and unmarried. For each of these sub-samples the coefficient on the owner house price term is negative and statistically significant at the 1% level of older single male individuals and at the 0.1% level for older married / co-habiting men and young married / co-habiting women. In each case these estimated coefficients are statistically significantly different from the renter house price coefficients at the 0.01% level of significance. The pattern in coefficient estimates also show female participation among middle-age and older married / co-habiting women decreases with the unemployment rate and participation among most groups decreases with a positive financial expectation, though the coefficients on the financial expectation variable are in each model not statistically significantly different from zero.

The coefficient estimates on the owner house price term for young married / co-habiting women is -0.132, statistically significant at the 0.1% level, hence at 10% increase in house prices causes a 1.3 percentage point reduction in the likelihood of participation for this group. The labor market participation rate among this group is 76%, so the 1.3 percentage point fall equates to a 1.7% fall in the likelihood of participation against the baseline participation rate. The renter house price term is positive but not statistically significant, so we see no evidence of a symmetric response among married / co-habiting renters who lose out when house prices increase. Results for young single women show no statistically significant effects of house prices on the participation decisions of either owners or renters, so the effects we observe for young women are specific to married / co-habiting young women only. Below we analyze the labor market destinations of this group when they leave the labor force and consider whether this withdrawal is likely to be temporary or permanent.

For the sub-groups of older married / co-habiting and single men the coefficient estimates on the owner house price variable are -0.149 and -0.134. These imply 1.5 percentage point and 1.3 percentage point reductions in the likelihood of participation into response to a 10% increase in house prices. Evaluated against the baseline participation rates for these groups (which are 36% and 25% respectively) these magnitudes imply that a 10% increase in house prices causes a 4.2% and 5.2% decrease in likelihood of participation. The coefficient estimates are statistically significantly different from the renter house price coefficients at the 0.01% level in both cases.

Table 5 presents results from the alternative robustness specifications. As in the hours results, here we show the robustness estimates for sub-samples of for which the main results returned statistically significant results for the owner house price coefficient (young married / co-habiting women, older married and single men). Results show very similar coefficient estimate on the house price variables for the first four columns which examine sensitivity to home moving and home tenure. As with the hours estimates, excluding movers causes the absolute value of the coefficient to increase confirming that moving activity biases the main result downwards. The specifications for tenure changes return very similar estimates to the main results. For each sub-sample the ‘forward prices’ falsification test yields no evidence of labor market participation responding to forward house price movements. On this basis we are confident that our main estimates are robust to moving activity and home tenure.

5.3 Labor Market Destinations

The results for labor market participation show labor supply elasticities with respect to house prices are significant and large for young married women and older men. These effects are consistent with labor supply adjustment by marginal workers located at the margins of family labor supply (young married / co-habiting women) and lifetime labor supply (older married / cohabiting men and older single men). In this section we explore these transitions further through analysis of the labor market destinations of these groups when leaving the labor force in response to house price gains.

We might expect that the withdrawal of young married / co-habiting women is temporary due to career breaks for children. Recent studies based on U.S. data have also found that house price increases raise the likelihood of couples having children (Lovenheim and Mumford, 2013). They do not examine the labor market consequences of this. Most women undertake some form of ‘maternity leave’ or other leave following childbirth. In our data we can estimate whether house price gains induce this form of activity for young women. To do so, we estimate our labor supply equation in which the dependent variable is a 1/0 dummy for whether a woman undertakes ‘family care’ activity (instead of working). We construct this measure from the survey question on labor market activity described earlier.

Results from these estimates are shown in Table 6. For completeness we estimate models for four subgroups: young and middle-age married women plus young and middle-age single women. Estimates for single women yield no statistically significant coefficients for either the owner house price or renter house price terms. Estimates for married / co-habiting women return a positive coefficient of 0.0926 for young women and 0.0474 for middle-age women, though the latter is statistically significant only at the 5% level. In both cases the owner house price coefficients are statistically significantly different from the renter house price coefficients. The renter house price coefficient for young married women is positive and statistically significant at the 5% level, providing some evidence that house price gains decrease the likelihood of undertaking family care activity for young married women renters.

The coefficient estimates imply large proportional effects of house price gains upon the likelihood of leaving the labor force to undertake family care activity. The baseline family care rate among young married home owning women is 18%, hence the impact of a 10%

increase in house prices is to raise he likelihood of family care among this group by on average 5%. For middle-age married women the baseline rate is 13% and implied effect of a 10% increase in house prices is 3.8%. We present robustness results in appendix table A1. These results show alternative specifications for moving activity and tenure yield very similar results to the main specification.

Finally, we present estimates of the effect of house prices on retirement decisions for older men. We again modify the labor supply equation with the dependent variable changed to a 1/0 indicator for whether the individual is retired. We define retirement as permanent exit from working and check our data to exclude observations for individuals who report themselves as retired in (at least) one wave but subsequently re-enter the labor market.

In Table 7 we report estimates for a sub-samples of older men and women, married and single. Results for women indicate no statistically significant coefficients on either the owner or renter house price terms. Results for men show statistically significant coefficients on the owner house price terms for both married and single men. The coefficient values of 0.156 and 0.143 imply a 10% increase in house prices raises the likelihood of retirement among men by 1.6 percentage points and 1.4 percentage points respectively for each group. Baseline retirement rates for these groups are 43% for male married and 34% for male singles. Hence a 10% increase in prices causes a 3.6% increase in the likelihood of retirement for male married and a 4.2% increase for male singles. Results from robustness specifications shown in appendix Table A2 confirm very similar coefficient estimates from the alternative specifications.

5.4 Discussion

Our results shows heterogeneous labor supply responses to house prices by housing tenure, gender, age and marital status. There is little evidence that participation or hours of work among middle-aged home owners are responsive to house price movements, but strong effects for younger married female owners and for older married and single owners. These effects are consistent with labor supply adjustment by marginal workers at the margins of family labor supply (young women) and lifetime labor supply (older men). The economic reasons for these effects may be different, however.

The response of young female owners suggests housing wealth gains influence labor supply decisions through an impact on borrowing constraints. Housing price increases for young owners are unlikely to represent significant wealth gains as young owners typically trade-up to larger houses in future (the price of which also increase with general house price increases). However, house price gains loosen borrowing constraints and this may impact on labor supply decisions. Cooper (2013) shows that among US households, the main route by which house price gains influence consumption is through loosening borrowing constraints. Our results suggest this is also true for labor supply.

House price gains allow owners who were previously borrowing constrained to extract home equity (e.g. through a larger mortgage) or to reduce mortgage financing costs by refinancing to a mortgage with a lower interest rate previously unavailable due to leverage constraints. Among young households labor supply effects are associated with having children; an activity which may have been postponed by households until borrowing constraints relaxed.

The response of older male owners appears consistent with a pure life-cycle wealth effect. Older male owners towards the end of their mortgage amortization are unlikely to be borrowing constrained. Instead, they are more likely to be holding above lifetime-average housing which they intend to downsize after retirement. For these households, house price gains represent pure wealth gains and we can interpret the labor supply response as a pure wealth response similar to the effect of a lottery win or inheritance.

Our results have implications for the business-cycle dynamics of labor supply for the groups of individuals who respond to house price changes. House prices are pro-cyclical and therefore our results suggest housing wealth gains are a pro-cyclical driver of leisure (for those older men who retire), or family care (for those younger married / co-habiting women who leave the labor force), in contrast to wages which are a pro-cyclical driver of wealth. However, the specifications we estimate include time dummies to capture time specific 'macroeconomic' effects. This means our estimates for labor supply effects of house price movements are net of national movements in prices (and identified off local variation against the national trend).

The inclusion of time dummies is our preferred approach to identification, but doing so does not allow us to use our coefficient estimates to calculate the business cycle effects of house price movements upon labor supply. Therefore we re-estimate the models shown in the previous section and exclude time dummies so that a business cycle interpretation can be applied to the estimated coefficients. We do this for the extensive margin estimates for young married women and older men. For young married women the coefficient value in the specification including time dummies (Table 4) was -0.132. Removing the time dummies results in a coefficient value of -0.141, also statistically significant at the 1% level. For older married men the coefficient in the model without time dummies is -0.154 (compared with -0.149 in the model without dummies) and -0.139 (compared with -0.134).

Why do these coefficient estimates move very little when the time dummies are removed? We should expect that labor supply dynamics have a strong aggregate level component. However, analysis of the coefficient on the unemployment variable provides an answer. With the removal of the time dummies the coefficient on the unemployment variable becomes highly statistically significant (at the 1% level) in each of these specifications and takes a negative value. Hence time variation in labor supply patterns is mostly captured by local unemployment rates, which can be seen as a measure of local macroeconomic conditions. We now use these estimates to calculate the implied aggregate effects of house prices and local unemployment conditions upon labor market participation during the recent recession. Our estimates imply housing wealth effects have a strong influence of labor supply over the business cycle compared with local labor market conditions and can explain a large share of labor supply movements during the recent recession.

Our calculations here can only be considered as illustrations of the importance of housing wealth effects. The coefficient estimates from models without time dummies imply that a 10% increase in house prices lower the labor supply rate among young married women by 1.4pp, among older married / co-habiting men by 1.5pp and among older single men by 1.4pp. We evaluate these estimated effects against changes in house prices and labor supply during the recent UK recession, the 8-quarter period of persistent decline in GDP beginning in the first quarter of 2008 and ending in the first quarter of 2010.

During this period the sale price of homes purchased by first time buyers fell in real terms value by on average 27% (figure derived from the first-time purchaser sales prices in the Halifax house price index used in our analysis). The labor market participation rate for young women fell from 72.9% to 71.1% (statistics on labor market participation by marital status are not available). Our estimates imply the 27% fall in price increased labor supply among young married women by 3.8pp. Hence had house prices seen no change, all other things being equal, the participation rate among young married women would have fallen to 67.3%, nearly three times the observed fall in participation.

Over the same period the unemployment rate rose by 2.5pp. Our coefficient estimates show that for young married / co-habiting owners an increase in unemployment of this magnitude leads to a 2.3pp decline in labor market participation. Hence in our estimates the wealth effect which encourages labor market participation arising from house price changes more than offsets the effect of labor market conditions captured by the local unemployment rate upon labor market participation for this group.

Equivalent calculations for older men also show our estimates imply economically important housing wealth effects during the recent recession. The participation rate of older men (using the same definition of age 55 to 75 as we use in our microdata analysis) fell from 40.7% in the first quarter of 2008 to 38.7% by the first quarter of 2010. We assume house prices facing this group fell in line with the all-sale Halifax index as we do not have a detailed house price index for older households. The index shows a 21% fall over the period. The mid-range of our coefficient estimates on the owner house price variable for older married / co-habiting and single men implies a 10% fall in house prices causes a 1.45pp increase in labor market participation rate.

A 21% fall in house prices therefore implies a 3pp increase in the labor market participation rate. Hence without the decrease in house prices, ceteris paribus, the labor market participation rate among older men would have fallen to 35.7%. For older men the average increase in county unemployment rate over the period of 2.5pp implies a 2.1pp decline in labor market participation. Therefore, as with young married / co-habiting female owners, the effect of house price falls increase labor market participation is larger than the decrease in participation arising due to labor market conditions.

These estimates for the business cycle effects of house price movements upon the labor market participation rate of younger married women and older men show that house price gains and losses may be economically important for understanding the labor supply dynamics of these groups. In particular, ‘wealth effects’ substantially (though not wholly) compensate for the effects of labor demand fluctuations, as proxied by the unemployment rate, over the business cycle.

6. Conclusion

This paper has presented empirical estimates of the impact of housing wealth on labor supply behavior among working-age individuals in the United Kingdom using individual level panel data. Results show large responses to housing gains and losses and certain groups which are unequally distributed among individuals by housing tenure and age. Changes in housing wealth have no significant impact on participation or hours decisions among middle-aged homeowners or renters, but decrease the likelihood of working among young married / co-habiting women and also among older men close to retirement age.

These results show that that housing wealth impacts on household labor supply behavior as well as consumer spending. Consumers partially spend housing wealth gains on both leisure and consumption. These results are consistent with standard models in which consumption and labor supply are jointly determined as households evaluate the marginal utility of consumption alongside the marginal utility of leisure. However, our results show labor supply responses across groups are not solely attributable to pure life-cycle wealth effects whereby older individuals 'win' and younger individuals 'lose' but instead reflect down-payment or liquidity constraint effects which drive labor supply responses of younger individuals. Our results are also of economic significance for understanding the business cycle dynamics of labor supply for those groups that respond to house price movements.

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Figure 1: Business Cycle Dynamics of House Prices and GDP in the UK, 1975-2012

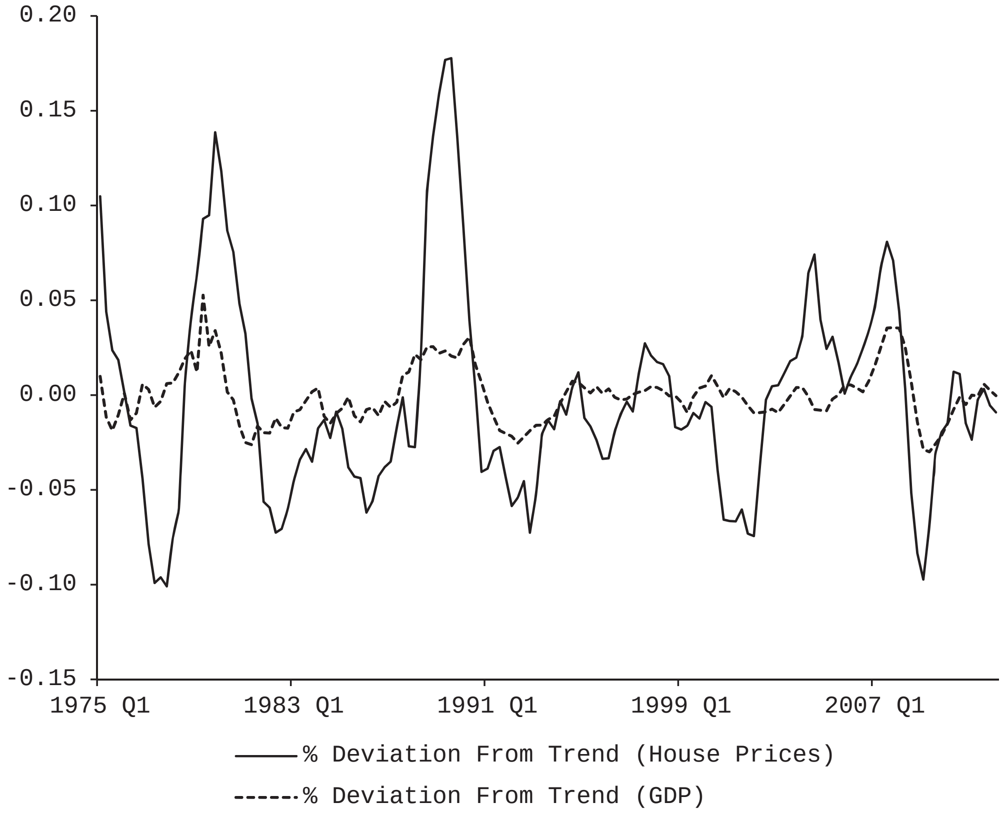

Figure shows percentage deviation from trend for UK real house prices (Halifax quarterly standardised house price index, seasonally adjusted, Q1 1975 – Q2 2012) and Real Gross Domestic Product (chain weighted measure, ONS coded ABMI, Q1 1975 – Q2 2012). Deviations from trend are calculated by applying the Hodrick-Prescott filter.

Table 1
Summary Statistics for BHPS Sample Demographic and Socio-Economic Characteristics

Demographics	
N	135,380
Age (years)	47.2
Male=1	0.56
Racial Minority=1	0.13
Married / Co-Habiting=1	0.77
Divorced=1	0.08
Children age 0-6=1	0.12
Children age 7-16=1	0.22
Highest Educational Qualification	
Degree=1	0.13
A-levels=1	0.16
O-levels=1	0.29
HND (technical college)=1	0.07
Current Employment Status	
Employed=1	0.59
Unemployed=1	0.03
Retired=1	0.26
Spouse / Partner Employed=1	0.41
Household Annual Ancome	£33,500
Housing Status and House Value	
Owner=1	0.78
Renter=1	0.22
House Value (£, owners,)	£133,000
Mortgage Value (£, if value > 0)	£53,900

Table 2 Estimates for Relationship Between Log House Prices and Log Hours of Work for Women and Men by Marital Status and Age Group. IV Fixed Effects Estimates.						
	Panel A: Individuals in Married or Co-Habiting Couples					
	Women			Men		
	Age <40	Age 40-54	Age >54	Age <40	Age 40-54	Age >54
(1) log hp - owner	-0.182***	-0.0143	0.0670	0.0397	0.0368	-0.00437
	(0.0201)	(0.0298)	(0.132)	(0.0313)	(0.0319)	(0.0322)
(2) log hp - renter	0.00264	0.00361	0.00138	0.00672	0.00134	0.0125
	(0.00491)	(0.00383)	(0.0181)	(0.00530)	(0.00364)	(0.0109)
(3) county unem.	0.0312	0.0337	0.333	-0.00880	-0.0809	-0.112
	(0.0596)	(0.0503)	(0.327)	(0.0674)	(0.0497)	(0.0783)
(4) financial expectation	-0.360*	-0.145	0.447	-0.332*	0.101	-0.226
	(0.142)	(0.131)	(0.805)	(0.156)	(0.115)	(0.219)
P-value test (1) = (2)	0.0000	0.2223	0.6076	0.4366	0.6230	0.2024
N	12727	12597	3636	12266	11384	4089
	Panel B: Single Individuals					
	Women			Men		
	Age <40	Age 40-54	Age >54	Age <40	Age 40-54	Age >54
(1) log hp - owner	-0.0171	0.0140	-1.115	0.0429	-0.0282	-1.118
	(0.0392)	(0.0310)	(4.320)	(0.0596)	(0.0164)	(14.80)
(2) log hp - renter	0.00455	0.00317	0.00335	0.0193	0.000220	0.0959
	(0.00533)	(0.00366)	(0.0106)	(0.0107)	(0.00323)	(1.270)
(3) county unem.	-0.194**	-0.0745*	-0.153	-0.180	0.0139	-1.595
	(0.0749)	(0.0371)	(0.788)	(0.121)	(0.0735)	(18.78)
(4) financial expectation	-0.141	0.0541	-1.422	-0.169	-0.108	9.613
	(0.117)	(0.134)	(8.829)	(0.138)	(0.122)	(136.5)
P-value test (1) = (2)	0.6589	0.7180	0.8513	0.3165	0.0933	0.9398
N	3747	3064	1098	2620	1778	582

Sample: Head of household plus spouse/partner BHPS 1991-2009. Individual fixed effects estimates. Instrumental Variable specification in which age, age squared and educational dummies for highest educational achievement (HND, GCSE, A-level, degree (or equivalents)) enter as instruments. Additional control variables: marital status dummies (married, divorced, widowed), number of children, health status (self-reported on 1-5 scale), spouse employment dummies (employed, unemployed, retired), natural log of annual non-labour income, homeowner dummy, county dummies, year dummies. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Cluster (county) standard errors in parentheses. P-value row reports values from test for equivalence of coefficients in rows (1) and (2).

Table 3 Robustness Estimates for Relationship Between Log House Prices and Log Hours of Work for Women in Married or Co-Habiting Couples. IV Fixed Effects Estimates.

	Women in Married or Co-Habiting Couples Age <40				
	Excluding Movers	Simulated Prices	Initial Owner	IV Owner	Forward Prices
(1) log hp - owner	-0.186***	-0.178***	-0.188***	-0.180***	-0.0140
	(0.0306)	(0.0316)	(0.0234)	(0.0275)	(0.0184)
(2) log hp - renter	0.00240	0.00218	0.00206	0.00201	0.00301
	(0.00316)	(0.00306)	(0.00416)	(0.00409)	(0.00308)
P-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.2105
N	11206	12727	12727	12727	12727

Notes: as Table 2.

Table 4 Estimates for Relationship Between Log House Prices and Labor Market Participation for Women and Men by Marital Status and Age Group. Fixed Effects Estimates.

Panel A: Individuals in Married or Co-Habiting Couples

	Women			Men		
	Age <40	Age 40-54	Age >54	Age <40	Age 40-54	Age >54
(1) log hp - owner	-0.132***	-0.0249	-0.0295	-0.00660	0.0173	-0.149***
	(0.0193)	(0.0235)	(0.0220)	(0.0157)	(0.0183)	(0.0223)
(2) log hp - renter	0.00151	0.00216	0.00113	0.00680	0.000250	0.00355
	(0.00596)	(0.00381)	(0.00468)	(0.0252)	(0.00294)	(0.00464)
(3) county unem.	0.0628	-0.115*	-0.152**	0.0642	0.0498	-0.0716
	(0.0694)	(0.0520)	(0.0568)	(0.0362)	(0.0403)	(0.0578)
(4) financial expectation	-0.153	-0.0108	0.0667	-0.0765	0.0118	-0.0477
	(0.164)	(0.136)	(0.236)	(0.0811)	(0.101)	(0.209)
R-squared	0.098	0.059	0.260	0.066	0.082	0.343
P-value test (1) = (2)	0.0000	0.3324	0.1947	0.9990	0.3360	0.0000
N	19026	18775	15820	15051	15499	15612

Panel B: Single Individuals

	Women			Men		
	Age <40	Age 40-54	Age >54	Age <40	Age 40-54	Age >54
(1) log hp - owner	-0.0106	-0.0444	-0.00942	0.0459	0.00940	-0.134**
	(0.0144)	(0.0447)	(0.0220)	(0.0453)	(0.0491)	(0.0390)
(2) log hp - renter	0.00689	0.00199	0.00574	0.00342	0.0201*	0.00235
	(0.00412)	(0.00825)	(0.0502)	(0.00308)	(0.00851)	(0.00683)
(3) county unem.	0.0551	0.124	-0.0848	0.0446	-0.225	0.0367
	(0.103)	(0.115)	(0.0678)	(0.100)	(0.121)	(0.123)
(4) financial expectation	-0.331	-0.0479	-0.00357	-0.381	0.249	0.731
	(0.278)	(0.300)	(0.302)	(0.265)	(0.303)	(0.437)
R-squared	0.218	0.096	0.258	0.376	0.132	0.280
P-value test (1) = (2)	0.0153	0.3364	0.7568	0.2762	0.5442	0.0000
N	6532	4879	8287	4055	2906	3462

Sample: Head of household plus spouse/partner BHPS 1991-2009. Individual fixed effects estimates. Additional control variables: age (in years), age squared (in years), marital status dummies (married, divorced, widowed), highest educational achievement dummies (HND, GCSE, A-level, degree (or equivalents)), ethnic minority group dummy variable, number of children, health status (self-reported on 1-5 scale), spouse employment dummies (employed, unemployed, retired), natural log of annual non-labor income, homeowner dummy, county dummies, year dummies. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Cluster (county) standard errors in parentheses.

Table 5 Robustness Estimates for Relationship Between Log House Prices and Labor Market Participation for Selected Groups. Fixed Effects Estimates.					
	Women in Married or Co-Habiting Couples Age <40				
	Excluding Movers	Simulated Prices	Initial Owner	IV Owner	Forward Prices
(1) log hp - owner	-0.138***	-0.128***	-0.131***	-0.135***	0.0261
	(0.0149)	(0.0246)	(0.0182)	(0.0206)	(0.0346)
(2) log hp - renter	0.00146	0.00118	0.00168	0.00124	0.00135
	(0.00506)	(0.00476)	(0.00431)	(0.00465)	(0.00405)
R-squared	0.110	0.092	0.091	0.097	0.082
P-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.0819
N	18102	19026	19026	19026	19026
	Men in Married or Co-Habiting Couples Age >54				
	Excluding Movers	Simulated Prices	Initial Owner	IV Owner	Forward Prices
(1) log hp - owner	-0.152***	-0.146***	-0.141***	-0.145***	0.0164
	(0.0201)	(0.0254)	(0.0283)	(0.0246)	(0.0281)
(2) log hp - renter	0.00164	0.00274	0.00209	0.00264	0.00231
	(0.00416)	(0.00401)	(0.00478)	(0.00462)	(0.00484)
R-squared	0.351	0.341	0.349	0.342	0.302
P-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.1324
N	14726	15612	15612	15612	15612
	Single Men Age >54				
	Excluding Movers	Simulated Prices	Initial Owner	IV Owner	Forward Prices
(1) log hp - owner	-0.139**	-0.133**	-0.133**	-0.135**	0.0219
	(0.0319)	(0.0346)	(0.0306)	(0.0321)	(0.0308)
(2) log hp - renter	0.00216	0.00224	0.00209	0.00203	0.00227
	(0.00616)	(0.00674)	(0.00669)	(0.00716	(0.00746)
R-squared	0.310	0.276	0.274	0.279	0.234
P-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.1524
N	3046	3462	3462	3462	3462

Notes: as Table 4

Table 6 Estimates for Relationship Between Log House Prices and Non-Working Full-Time Childcare for Women by Marital Status and Age Group. Fixed Effects Estimates.				
	Married or Co-Habiting		Single	
	Age<40	Age 40-54	Age <40	Age 40-54
(1) log hp - owner	0.0926***	0.0474*	0.0750	0.0921
	(0.0258)	(0.0216)	(0.0957)	(0.0858)
(2) log hp - renter	-0.0106*	0.000482	-0.00115	0.00530
	(0.00524)	(0.00350)	(0.00346)	(0.00661)
(3) county unem.	-0.0216	0.0618	-0.120	-0.0349
	(0.0610)	(0.0478)	(0.0867)	(0.0924)
(4) financial expectation	0.186	0.108	0.0862	0.164
	(0.145)	(0.125)	(0.234)	(0.241)
R-squared	0.092	0.022	0.113	0.081
P-value test (1) = (2)	0.0000	0.0299	0.0950	0.0141
N	19026	18775	6532	4879

Sample: female head of household plus spouse/partner BHPS 1991-2009. Individual fixed effects estimates. Additional control variables: age (in years), age squared (in years), marital status dummies (married, divorced, widowed), highest educational achievement dummies (HND, GCSE, A-level, degree (or equivalents)), ethnic minority group dummy variable, number of children, health status (self-reported on 1-5 scale), spouse employment dummies (employed, unemployed, retired), natural log of annual non-labor income, homeowner dummy, county dummies, year dummies. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Cluster (county) standard errors in parentheses.

Table 7 Estimates for Relationship Between Log House Prices and Retirement for Older Men and Women by Marital Status. Fixed Effects Estimates.

	Women		Men	
	Marr Age >54	Single Age >54	Marr Age > 54	Single Age > 54
(1) log hp - owner	0.00178	-0.00640	0.156***	0.143**
	(0.0279)	(0.0295)	(0.0250)	(0.0455)
(2) log hp – renter	-0.00525	0.00689	-0.00128	-0.00712
	(0.00594)	(0.00673)	(0.00521)	(0.00795)
(3) county unem.	-0.0836	-0.0748	0.0788	0.0175
	(0.0721)	(0.0910)	(0.0649)	(0.143)
(4) financial expectation	0.395	-0.0952	-0.124	0.277
	(0.300)	(0.405)	(0.234)	(0.509)
R-squared	0.299	0.265	0.337	0.293
P-value test (1) = (2)	0.8000	0.6461	0.0000	0.0031
N	15820	8287	15612	3462

Sample: Head of household age over 54 plus spouse/partner BHPS 1991-2009. Individual fixed effects estimates. Additional control variables: age (in years), age squared (in years), marital status dummies (married, divorced, widowed), highest educational achievement dummies (HND, GCSE, A-level, degree (or equivalents)), ethnic minority group dummy variable, number of children, health status (self-reported on 1-5 scale), spouse employment dummies (employed, unemployed, retired), natural log of annual non-labor income, homeowner dummy, county dummies, year dummies. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Cluster (county) standard errors in parentheses.

Table A1 Robustness Estimates for Relationship Between Non-Working Full-Time Childcare for Married Women by Age Group. Fixed Effects Estimates.

	Women in Married or Co-Habiting Couples Age <40				
	Excluding Movers	Simulated Prices	Initial Owner	IV Owner	Forward Prices
(1) log hp - owner	0.0946***	0.0920***	0.0918***	0.0916***	0.0135
	(0.0256)	(0.0284)	(0.0279)	(0.0261)	(0.0193)
(2) log hp - renter	-0.0101*	-0.0126*	-0.0134*	-0.0167*	-0.0172*
	(0.00516)	(0.00564)	(0.00543)	(0.00556)	(0.00559)
R-squared	0.095	0.089	0.091	0.088	0.082
P-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.2926
N	18103	19026	19026	19026	19026
	Women in Married or Co-Habiting Couples Age 40-54				
	Excluding Movers	Simulated Prices	Initial Owner	IV Owner	Forward Prices
(1) log hp - owner	0.0479*	0.0464*	0.0468*	0.0471*	0.0216
	(0.0220)	(0.0219)	(0.0223)	(0.0242)	(0.0349)
(2) log hp - renter	0.00106	0.00108	0.00143	0.00109	0.00106
	(0.00318)	(0.00364)	(0.00309)	(0.00351)	(0.00326)
R-squared	0.026	0.021	0.020	0.020	0.017
P-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.1942
N	17891	18775	18775	18775	18775

Sample: female head of household plus spouse/partner BHPS 1991-2009. Individual fixed effects estimates. Additional control variables: age (in years), age squared (in years), marital status dummies (married, divorced, widowed), highest educational achievement dummies (HND, GCSE, A-level, degree (or equivalents)), ethnic minority group dummy variable, number of children, health status (self-reported on 1-5 scale), spouse employment dummies (employed, unemployed, retired), natural log of annual non-labor income, homeowner dummy, county dummies, year dummies. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Cluster (county) standard errors in parentheses.

Table A2 Robustness Estimates for Relationship Between Log House Prices and Retirement for Older Men by Marital Status. Fixed Effects Estimates.

	Men in Married or Co-Habiting Couples Age <40				
	Excluding Movers	Simulated Prices	Initial Owner	IV Owner	Forward Prices
(1) log hp – owner	0.161***	0.154***	0.153***	0.156***	0.0746
	(0.0216)	(0.0267)	(0.0284)	(0.0264)	(0.0816)
(2) log hp – renter	-0.00103	-0.00197	-0.00146	-0.00182	-0.00191
	(0.00503)	(0.00519)	(0.00586)	(0.00549)	(0.00506)
R-squared	0.342	0.328	0.329	0.331	0.294
P-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.2864
N	14826	15612	15612	15612	15612
	Women in Married or Co-Habiting Couples Age 40-54				
	Excluding Movers	Simulated Prices	Initial Owner	IV Owner	Forward Prices
(1) log hp – owner	0.147**	0.140**	0.141**	0.138**	0.0497
	(0.0416)	(0.0408)	(0.0417)	(0.0421)	(0.0516)
(2) log hp – renter	-0.00780	-0.00816	-0.00809	-0.00703	-0.00701
	(0.00816)	(0.00846)	(0.00879)	(0.00761)	(0.00708)
R-squared	0.312	0.292	0.291	0.289	0.271
P-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.4165
N	3015	3462	3462	3462	3462

Sample: head of household age over 54 plus spouse/partner BHPS 1991-2009. Individual fixed effects estimates. Additional control variables: age (in years), age squared (in years), marital status dummies (married, divorced, widowed), highest educational achievement dummies (HND, GCSE, A-level, degree (or equivalents)), ethnic minority group dummy variable, number of children, health status (self-reported on 1-5 scale), spouse employment dummies (employed, unemployed, retired), natural log of annual non-labor income, homeowner dummy, county dummies, year dummies. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. Cluster (county) standard errors in parentheses.