Economica

Economica (2018) **85**, 449–478
doi:10.1111/ecca.12253

House Prices, Wealth Effects and Labour Supply

By Richard Disney† and John Gathergood‡

†*University College London, Institute for Fiscal Studies* ‡*University of Nottingham*

Final version received 12 August 2017.

We examine the impact of house prices on labour supply decisions using UK microdata. We combine household survey data with local-level house price measures and controls for local labour demand. Our microdata also allow us to control for individual level income expectations. We find significant house price effects on labour supply, consistent with leisure being a normal good. Labour supply responses to house prices are concentrated among young married female owners and older owners. This finding suggests that house prices affect the decisions of marginal workers in the economy. Our estimates imply that house prices are economically important in the participation decisions for these workers.

Introduction

Over the course of the past decade, many developed economies experienced sustained house price increases in the run up to the Great Recession, followed by a period of rapid house price decline. These housing market 'booms' and 'busts' were particularly pronounced in the USA and the UK. In these economies, most households are homeowners and housing is the largest single investment for these households. On paper, movements in house prices generate large changes in wealth on household balance sheets.

Do these changes in house prices matter for household behaviour? Recent studies have confirmed that house prices are important for a range of household activity and behaviour. Several studies based on US and UK microdata have shown that changes in housing wealth affect consumption spending and household indebtedness, both by changing lifecycle wealth and through relaxing and tightening borrowing constraints.[1] Other studies have found that house prices have significant effects on educational choices (Lovenheim 2011; Lovenheim and Reynolds 2013) rates of childbirth (Lovenheim and Mumford 2013; Dettling and Kearney 2014), demand for long-term care insurance (Davidoff 2010) and divorce (Farnham *et al*. 2011).

Do house prices also matter for labour supply decisions? In this paper we estimate the size of housing wealth effects on labour supply for a panel of households in the UK. Our results show that labour supply responses to changes in housing wealth are highly heterogeneous across household types. We find small average effects of house prices on labour supply choices, but large effects for subsets of households. The household types that show significant responses to housing wealth changes are those at the margins of labour supply: married women, at the intratemporal margin of household labour supply; and men close to retirement at the intertemporal margin of lifetime labour supply. The prior literature shows that these households are responsive to changes in marginal tax rates. We show that wealth effects are also important for understanding the labour supply decisions of these groups.

The effects that we find are economically significant. For example, we find that a 10% rise in local house prices relative to the national trends is associated with a reduction in the labour market participation rate among young married/cohabiting women of 1.8%, and a reduction in the participation rate among older men of approximately 4.4%. Therefore our results show that house price changes have distributional effects on labour supply (as well as consumption) that correlate with lifecycle characteristics. Hence there is a lifecycle effect as well as an overall effect of house price changes on labour supply.

Why do house prices also matter for labour supply decisions? Basic economic theory tells us that increases in non-labour income or wealth raise consumption of normal goods. Leisure, like consumption, is typically thought of as a normal good, so we might expect housing wealth gains to increase leisure and decrease labour supply for some homeowning households, and vice versa for housing wealth losses. Furthermore, studies based on microdata typically find the aggregate marginal propensity to consume out of housing wealth is small. One reason for this finding might be that for some types of households, housing wealth changes primarily affect labour supply—for example, the decision to retire of older workers—rather than consumption. Hence housing wealth gains might cause some households to reduce income as they take more leisure, instead of increasing consumption for an unchanged income and labour supply.

Existing studies of wealth effects on labour supply exploit exogenous wealth changes such as lottery wins (Imbens *et al*. 2001; Cesarini *et al*. 2013) and inheritances (Joulfaian and Wilhelm 2001; Brown *et al*. 2010). These studies in general confirm the intuition that labour supply falls when wealth increases. Studies on US data have shown that housing wealth changes impact on decisions that have implications for labour supply, although these studies do not estimate labour supply effects directly. Lovenheim (2011) and Lovenheim and Reynolds (2013) show that increases in housing wealth raise college and university enrolments. Lovenheim and Mumford (2013) show that housing wealth gains also raise the likelihood of homeowners choosing to have children. As we explain in the next section, estimating exogenous wealth effects in the context of housing wealth raises some tricky modelling issues; hence the present study has perhaps been the first to consider the effect of house prices on labour supply choices in detail. However, the results presented here have been broadly confirmed by very recent studies for the USA (Milosch 2016) and Australia (Atalay *et al*. 2016).

Prior studies also show that movements in wealth are particularly important at the margin of retirement timing. Blundell *et al*. (2014) for the UK, and French and Benson (2011) and Daly *et al*. (2009) for the USA, all argue that asset price declines may be one reason why labour supply in the post-2008 recession remained higher than in previous recessions due to delayed retirement. However, Coile and Levine (2011) for the USA and Disney *et al*. (2015) for the UK find evidence that local labour market changes dominate asset (wealth) effects in explaining patterns of retirement over the business cycle.

The paper proceeds as follows. Section I describes our econometric modelling strategy. The British Household Panel Survey, which we use to estimate responses to housing wealth for various dimensions of labour supply, is described in Section II. Section III describes our econometric model. Section IV describes our main results concerning participation and hours. We estimate effects of housing wealth first on hours of work and then on participation decisions. Where we find that house price gains (losses) lead to reduced (increased) labour market participation, we then investigate the types of activities individuals undertake when they withdraw from the labour market—including time away from work to care for children, and retirement. In Section V, we discuss our findings and the economic implications of house price movements for labour supply patterns observed during the recent recession. Section VI concludes.

I. Modelling Strategy

Our modelling strategy is predicated on the observation that house price movements are not randomly assigned across localities and households, and are likely to correlate with local economic conditions as well as other factors that influence labour supply decisions. Typically, too, observed house values arise endogenously with lifetime choices of a household. An important issue in this context is that wealth effects on labour supply should be identified only off exogenous shocks. In the canonical lifecycle model, consumption, wealth accumulation (including housing wealth) and labour supply are simultaneously determined. Households may, for example, work more in order to acquire a more expensive house.

Households are 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 prices are broadly procyclical in nature. Figure 1 illustrates this procyclicality of house prices in the UK using detrended data. House prices are strongly procyclical and more variable than GDP. The correlation coefficient between house prices and GDP is 0.6 over the whole period, which includes many business cycle fluctuations.[2] Hence it is reasonable to assume that households understand the trend and cyclicality of house prices. Modelling the 'exogenous' component of house price changes to households is therefore an important practical issue.[3]

It is not possible to randomly assign housing wealth. Therefore our source of identification arises from differential changes in house prices across localities relative to average national house price changes controlling for neighbourhood effects and trends (such as local amenities that may affect house price levels in the area) and household preferences. In our baseline model we utilize changes in local house price indices conditioned on time, household and neighbourhood effects as our measure of exogenous variation in house prices. We adopt an instrumental variables strategy, instrumenting self-reported housing wealth using local-level house prices, as house values may be reported endogenously. We also control for tenure and locality choices.

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 realized local variations in the rate of change of house prices relative to this national trend.[4] 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 widely used free online property search engines that provide valuations of existing properties. A relaxation of our modelling 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, but we nevertheless investigate this possibility. We also, as an additional sensitivity test, see whether our estimated labour supply responses are robust to using house values self-reported by households in the survey.

To examine the effect of asset prices on labour supply, we must also control for variation in local labour demand, given the likely covariance of shocks to asset and labour markets which might codetermine local house prices and labour supply decisions. We control for local labour market conditions by including the local unemployment rate and local median wage rate as controls.

We use an additional identification strategy for our modelling based on estimating the response of renters to house prices and comparing this with the response of owners.

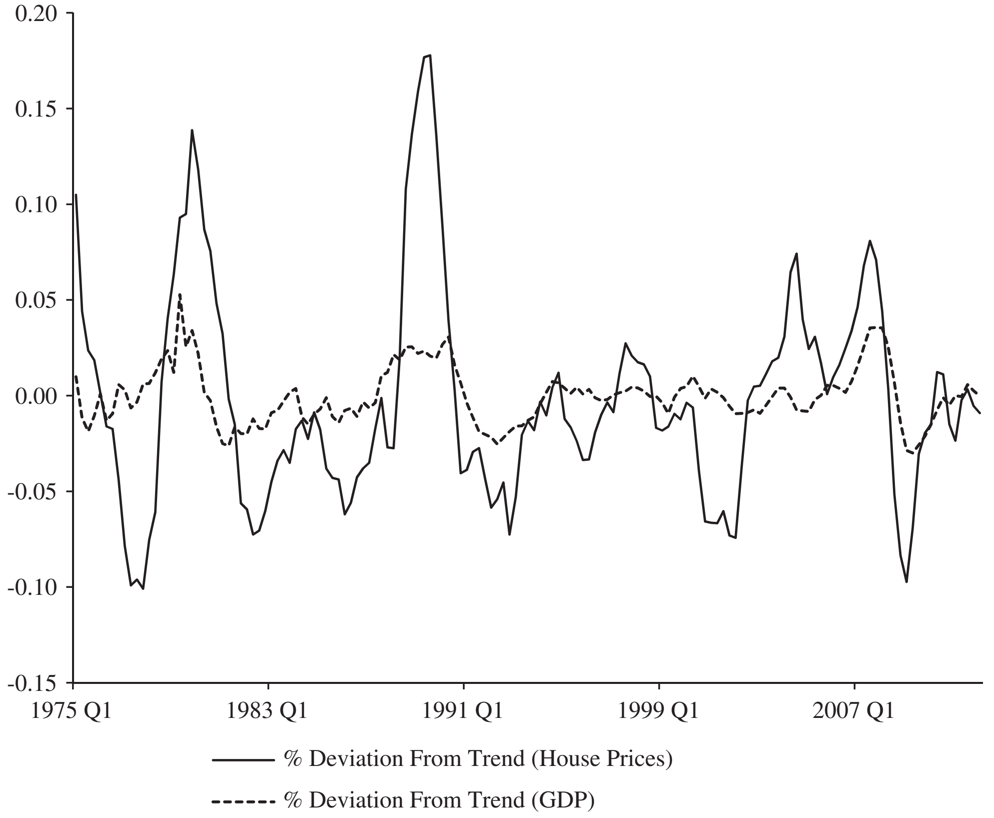

FIGURE 1. Business cycle dynamics of house prices and GDP in the UK, 1975–2012.
Notes: The figure shows percentage deviation from trend for UK real house prices (Halifax quarterly standardized house price index, seasonally adjusted, Q1 1975–Q2 2012) and real gross domestic product (chain weighted measure, ONS coded ABMI, Q1 1975–Q2 2012). Deviations from trend are calculated by applying the Hodrick–Prescott filter.

The rationale for including renters is as follows: owners and renters should respond asymmetrically to house prices—increases which represent gains for owner but losses for renters (through higher rental prices or future house purchase cost).[5] If we observe owners and renters responding symmetrically to house prices, then that would suggest that in the data, house prices are acting as a proxy for uncaptured local economic conditions that affect both owners and renters in the same way. This approach to identification has been used extensively in prior studies of housing and consumption (Campbell and Cocco 2007; Disney *et al*. 2010; Attanasio *et al*. 2011) and also in studies of other non-consumption outcomes (Lovenheim 2011; Lovenheim and Mumford 2013; Dettling and Kearney 2014; Davidoff 2010). It is equivalent to a difference-in-difference estimation across owners and renters.

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 labour supply. In intertemporal models of labour supply, higher expected future income (arising, for example, from higher expected future wages) might induce workers to reduce current labour 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. In contrast, individual-level income expectations data are not available in US household panels covering the working-age population.[6]

Finally, our modelling 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 (Hilber and Vermeulen 2016), 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 that 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.

II. Data Sources

We use UK data combining variation in house prices across geographic localities with household panel data to estimate exogenous housing wealth effects on labour supply—both for total hours and separately at the extensive margin (participation). Our primary dataset 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 labour 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 5000 households, comprising a total of approximately 10,000 individual interviews.

Major topics covered in the survey are household composition and demographics, participation in the labour 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 age 16. The BHPS adopts a following rule such that if a household leaves the survey for a reason other than death, then it is replaced by a similar household. Households that leave the survey due to death are replaced with young households. Hence the sample in each cross-section is representative of the population of the UK. We use 18 waves of data that are available from 1991 to 2009.[7]

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 labour market participation and hours of work. We exclude the self-employed as the relationship between house prices and self-employment has been considered elsewhere (Hurst and Lusardi 2004; Disney and Gathergood 2009). Using the same dataset as this study, Disney and Gathergood (2009) show that house price gains raise the likelihood that an individual becomes self-employed or starts their own business. However, the accumulation of home equity may arise endogenously with the decision to begin a business in the future.

The labour 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 dataset as the sum of hours normally worked per week plus overtime hours for first and second jobs.[8] We define an individual as participating in the labour market if they report their labour market status as 'in paid employment' or 'unemployed'. This is our measure of labour 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 labour 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 changes in the household's financial expectations that might cause changes in labour supply in the current period and is similar to those used in consumer confidence indices.[9] Also, the question is not limited to income but might capture other future financial characteristics related to labour supply and housing choices, such as anticipated childrearing expenses. We take answers to this question and code two 1/0 dummy variables for 'positive financial expectations' and 'negative financial expectations' that we include in our econometric specification, allowing the labour supply responses of individuals to positive and negative expectations to differ in sign and magnitude.

We also match into the BHPS local house price data. This approach, which is similar to that used by Lovenheim and Reynolds (2013), has two purposes: first, it provides an instrument for self-reported house prices reported by owners; 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.

Our house price data come from the recently introduced Land Registry Local Authority level index, which reports average sale values for all new and repeat home sales. Throughout we adjust all financial variables to 2000 prices using the Retail Prices Index. We also match into the BHPS two local authority level variables that capture local labour market conditions: first, registry unemployment data provided by the Office for National Statistics (ONS); second, local authority level average earnings derived from the ONS Annual Survey of Hours and Earnings employer survey.[10]

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/cohabiting 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 homeowners, with the average house value among owners at £133,000. The second and third columns of Table 1 show summary statistics for owners and renters: owners are typically higher-income, more likely to be in work (and have a spouse or partner in work) and have more education.

III. Econometric Model

This section explains in detail our approach to identification and estimation. Our main econometric model is a difference-in-differences specification in which the effect of house prices on hours of work is estimated by homeownership status. The specification is

Table 1
Summary Statistics for BHPS Sample Demographic and Socioeconomic Characteristics

	All	Owners	Renters
Demographics			
N	135,380	100,224	35,156
Age (years)	47.2 (14.62)	42.3 (13.79)	44.3 (16.4)
Male = 1	0.56 (0.50)	0.55 (0.50)	0.059 (0.50)
Racial minority = 1	0.13 (0.11)	0.12 (0.10)	0.15 (0.11)
Married/cohabiting = 1	0.77 (0.42)	0.84 (0.37)	0.57 (0.50)
Divorced = 1	0.08 (0.28)	0.06 (0.24)	0.15 (0.36)
Children aged 0–6 = 1	0.12 (0.33)	0.12 (0.33)	0.13 (0.34)
Children aged 7–16 = 1	0.22 (0.42)	0.21 (0.41)	0.22 (0.41)
Highest educational qualification			
Degree = 1	0.13 (0.33)	0.15 (0.35)	0.07 (0.27)
A levels = 1	0.16 (0.37)	0.17 (0.38)	0.14 (0.34)
O levels = 1	0.29 (0.45)	0.29 (0.46)	0.28 (0.45)
Current employment status			
Employed = 1	0.59 (0.60)	0.65 (0.48)	0.45 (0.50)
Unemployed = 1	0.03 (0.16)	0.02 (0.13)	0.07 (0.26)
Retired = 1	0.26 (0.80)	0.27 (0.40)	0.26 (0.39)
Spouse/partner employed = 1	0.41 (0.41)	0.47 (0.50)	0.24 (0.42)
Household annual income	£33,500 (£18,648)	£42,700 (£32,100)	£29,400 (£18,600)
Housing status and house value			
Owner = 1	0.78 (0.44)	1.00 (0.00)	0.00 (0.00)
Renter = 1	0.22 (0.15)	0.00 (0.00	1.00 (0.00)
House value (£, owners)	£133,000 (£128,000)	—	—
Mortgage value (£, if value > 0)	£53,900 (£45,600)	—	—

Notes
Mean values with standard deviation in parentheses.

$$
\begin{aligned}
h_{ilt} = {} & \alpha + \beta_1 H^*_{lt} O_{ilt} + \beta_2 H^*_{lt} R_{ilt} + \beta_3 O_{ilt} + \beta_4 U_{lt} + \beta_5 E_{lt} + \beta_6 X_{ilt} \\
& + \beta_7 F_{ilt} + \varphi_i + \theta_l + \psi_t + \varepsilon_{ilt},
\end{aligned} \tag{1}
$$

where i denotes an individual, l denotes local authority of residence, and t denotes year. The (log) of annual hours for all employed individuals with non-zero hours is denoted h_{ilt}. H_{ilt} is the local authority house price (the (log) average house price at the local authority level in each year). O_{ilt} is a dummy variable taking the value 1 if the individual is a homeowner and 0 otherwise. R_{ilt} is the reverse dummy variable denoting whether the individual is a renter.

Among the set of control variables, U_{lt} is the local unemployment rate at the local authority level in each year, E_{lt} is (log) average earnings at the local authority level in each year, X_{ilt} is a set of individual level socioeconomic characteristics and control variables, and F_{ilt} is the individual's self-reported financial expectation.[11] Equation (1) also includes local authority fixed effects θ_l, year fixed effects ψ_t, and time-invariant individual characteristics captured by the individual fixed effects φ_i.

The BHPS data also contain self-reported house prices for homeowners (but not for renters). Therefore we also present an instrumental variables (IV) specification in which self-reported prices for homeowners are instrumented using local authority prices. This specification is

$$h_{ilt} = \alpha + \beta_1 \hat{H}^O_{ilt} + \beta_2 H^R_{lt} + \beta_3 O_{ilt} + \beta_4 U_{lt} + \beta_5 E_{lt} + \beta_6 X_{ilt} + \beta_7 F_{ilt} + \varphi_i + \theta_l + \psi_t + \varepsilon_{ilt}, \tag{2}$$

where

$$\hat{H}^O_{ilt} = \beta_1 P_{lt} + \beta_4 U_{lt} + \beta_5 E_{lt} + \beta_6 X_{ilt} + \beta_7 F_{ilt} + \varphi_i + \theta_l + \psi_t + \xi_{ilt}. \tag{3}$$

In this specification $\hat{H}^O_{ilt}$ is the instrumented self-reported house price for owners, where the instrument is the local authority house price P_{lt} (the (log) average house price at the local authority level in each year). H^R_{lt} is the local authority house price for renters. When using self-reported prices, we instrument them using local authority level house prices, to control for potential endogeneity of self-reported prices to labour supply decisions.

Household tenure choice and moving activity may not be exogenous, and we address this in our identification strategy. To interpret the coefficient β_1 as representing the causal impact of housing wealth on labour supply requires that the estimated impact of local house prices on labour supply is not attributable to omitted variable(s) that might drive both house prices and labour supply for which house prices might be a proxy. There may be unobserved differences in local economic conditions not captured by the covariates included in equation (1).

To account for these we incorporate renters into our econometric model, adopting an approach that has been used extensively in the literature on housing and consumption (e.g. Campbell and Cocco 2007; Disney *et al*. 2010; Attanasio *et al*. 2011.) The reason for this approach is as follows. If renters intend to buy in future, then indirect wealth gains and losses among renters arising from local house price changes are in the opposite direction to those experienced by current owners. Indeed, if house price appreciation is ultimately reflected in increased rental prices, then renters face increased cost of renting. Thus, conditioning on controls, if our house price variable is not proxying for unobserved local economic conditions, then renters should respond differently to owners in respect to house price changes.

Hence if the coefficients β_1 and β_2 are both non-zero and equal (i.e. the estimated impacts of local authority house prices on the labour supplies of owners and renters are identical), then we would conclude that local authority house prices proxy for unobserved local conditions. If they are both zero, then 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, then we have identified a negative wealth effect on labour supply arising from (changes in) housing wealth.

This comparison between renters and owners in equation (1) is equivalent to a difference-in-differences model. However, renters and owners have different characteristics, shown in Table 1; for example, our data show that renters typically have lower household income. 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 that 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 labour supply, hence confounding our model.

Accordingly, in our estimates, and in an extension to equation (1), interaction terms between the owner and renter house price variables and *all other* observable covariates are included in the model. To account for the effects of unobservable time-varying differences between owners and renters that might locally correlate with house prices and labour supply, we also incorporate renter-locality time trends and owner-locality time trends into the model shown in equation (2).

In addition, two sources of selection bias might confound estimates of equation (1).

First, local authority level house price changes are not exogenous for individuals who change local authority. Selection bias would occur if individuals moved to higher house price localities and simultaneously changed their labour market participation. To eliminate any bias arising from moving behaviour, we use two strategies.

In the first strategy, we exclude cross-local authority movers (dropping approximately 8.5% of the individual–year observations in our sample). We show that the omission of these households does not change our results. In the second strategy, we keep cross-local-authority movers in the sample but calculate the counterfactual house price change (i.e. that they would have received had they not moved local authority) and use this simulated change in house prices to estimate equation (1) instead of their actual cross-local-authority change. This strategy shows very similar results to our baseline estimates.

Second, selection bias would arise if house price changes caused individuals to change from renting to owning and the if likelihood of changing tenure were related to labour supply. We address this in two ways. 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 to eliminate housing tenure changes that might cause selection bias. Second, we use initial homeownership status as an instrument for contemporaneous housing tenure, assuming that initial homeownership status is exogenous. We show that both strategies yield estimates of β_1 and β_2 that 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 labour market. As we use a fixed effects panel estimator, we are thereby estimating labour market transitions. In similar vein, and corresponding to some of the existing literature, we also estimate transition equations into other non-participation labour-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 local authority level, we calculate standard errors clustered at the local authority 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. Our econometric specification also includes generated regressors in the local-level covariates, which are themselves estimated, hence we also apply a standard bootstrap technique to our econometric estimates.

IV. Results

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 difference-in-differences hours of work equation (equation (1)) for subsamples of individuals defined by marital status (single or married/cohabiting), gender (male/female) and age (less than 40; 40–54; over 54). 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 subsamples defined by gender and the three age categories.

We report coefficients and standard errors on the house price terms for homeowners and renters, the local authority unemployment rate and the financial expectations variable.

TABLE 2
ESTIMATES FOR RELATIONSHIP BETWEEN LOG LOCAL HOUSE PRICES AND LOG HOURS OF WORK FOR WOMEN AND MEN BY MARITAL STATUS AND AGE GROUP: INDIVIDUAL FIXED EFFECTS ESTIMATES

	Women			Men		
	(1) Age < 40	(2) Age 40–54	(3) Age > 54	(4) Age < 40	(5) Age 40–54	(6) Age > 54
Panel A: Individuals in married or cohabiting couples						
(1) log hp–owner	−0.176***	−0.0118	0.0534	0.0456	0.0284	−0.00800
	(0.0282)	(0.0307)	(0.0747)	(0.0323)	(0.0301)	(0.0409)
(2) log hp–renter	0.00510	−0.00385	0.00146	−0.00786	−0.00173	−0.0101
	(0.00492)	(0.00390)	(0.0108)	(0.00542)	(0.00355)	(0.00622)
(3) unemployment (%)	−0.0388	−0.00242	0.247	−0.0484	−0.0959	−0.160
	(0.0679)	(0.0600)	(0.196)	(0.0771)	(0.0569)	(0.0853)
(4) financial expectation	−0.327*	−0.125	0.638	−0.309	0.108	−0.279
	(0.142)	(0.133)	(0.458)	(0.158)	(0.113)	(0.215)
p-value test (1) = (2)	0.0000	0.2223	0.6076	0.4366	0.6230	0.2024
N	12,727	12,597	3636	12,266	11,384	4089
Panel B: Single individuals						
(1) log hp–owner	−0.0267	−0.00280	−1.362	0.0524	−0.0651	−0.512
	(0.0449)	(0.0296)	(1.551)	(0.0846)	(0.0718)	(3.702)
(2) log hp–renter	−0.000655	0.00372	−1.332	−0.0308	0.00143	−0.0628
	(0.00608)	(0.00358)	(1.441)	(0.0192)	(0.00410)	(0.438)
(3) unemployment (%)	−0.227	−0.0667	0.0912	−0.335	−0.00331	0.689
	(0.124)	(0.0583)	(0.301)	(0.230)	(0.0524)	(7.193)
(4) financial expectation	−0.151	0.0467	0.619	−0.138	−0.125	3.615
	(0.245)	−0.00280	(1.382)	(0.481)	(0.129)	(25.01)
p-value test (1) = (2)	0.6589	0.7180	0.8513	0.3165	0.0933	0.9398
N	3747	3064	1098	2620	1778	582

Notes
Dependent variable: natural log of hours of work for sample of individuals with non-zero hours. Sample: head of household plus spouse/partner BHPS 1991–2009. Individual fixed effects estimates. Self-reported house price for owners is instrumented using log local authority mean house price. Additional control variables not shown in table: age, age squared; educational dummies for highest educational achievement (HND, GCSE, A level, degree (or equivalents)); 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, local authority dummies, year dummies, renter–local authority and owner–local authority time trends.
*, **, *** indicate $p < 0.05$, $p < 0.01$, $p < 0.001$, respectively.
Cluster (local authority) standard errors in parentheses.
'*p*-value test' reports values from test for equivalence of coefficients in rows (1) and (2).

Results in Table 2 show that for all groups other than young married/cohabiting women there is no evidence for a statistically significant effect of house prices on 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 hypothesis that the coefficients for the two groups are the same.

However, we do find statistically significant results for young married/cohabiting women. The coefficient on the homeowner 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 less than 0.001, implying that these coefficients are significantly different from one another at a very high level of confidence.

The coefficient on the homeowner house price term takes the value −0.176. This implies that a 10% increase in house prices leads to a reduction in hours for married/cohabiting young female homeowners of 1.8%. Average (non-zero) annual hours for this group in our sample are 1485. Hence a 10% increase in prices reduces annual hours by 27 hours per annum, approximately three-quarters of a working week of hours on average for this group.

For young married/cohabiting 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. Our result that the labour supply of young married women is responsive to housing wealth is unsurprising—this is the labour market groups typically found to operate at the margin of intratemporal household labour supply.

Table 3 shows results from a series of robustness specification tests for our results for the sample of young married/cohabiting women. There are five alternative specifications in the table.[12] The first two specifications relate to home moving activity. If house price changes correlate with labour market changes that induce households to move across local authorities, then our estimates in Table 2 might suffer selection bias. In the first column of Table 3, individuals who move home (approximately 8.8% of the sample) are excluded. In the second column, for individuals who move local authority we construct a counterfactual house price as the price in their former local authority in all waves following their home move (i.e. we allocate to that individual the future house price as if they had not moved local authority).

Results show that when movers are excluded from the sample, the owner local authority house price term remains negative, statistically significant at the 0.1% level, and significantly different from the (not significant) coefficient on the renter local authority house price term. The absolute value of the coefficient is a little larger than in the baseline specification (−1.79 compared with −1.76), confirming our prior that including movers biases the coefficient estimate downwards. When simulated prices are used, the same pattern of statistical significance remains but the absolute value of the coefficient falls a little (to −1.73). 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. If house price changes induce households to change housing tenure and labour supply, then bias will be introduced into our estimates. In column (3), homeownership status of the individual is fixed to be their homeownership status in the first wave in which they are observed in the survey. This is a similar approach to that in column (2) of simulating local authority house prices for movers in that we build a counterfactual status for the individual had they not entered into the activity that might confound our estimates (moving in the previous case, tenure-changing in this case). In column (4), this approach is implemented as an IV 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 that the tenure-changing activity does not confound our main estimates.

The final column of Table 3 shows results from a 'falsification test' where the one-period forward house price is added to the model alongside the contemporaneous house price. A significant coefficient on the one-period forward house price might indicate a spurious relationship—future house prices affecting current labour supply—which might be due to selection, for example, house prices proxying for household wealth. However, in this specification neither of the one-period forward house price terms for owners or renters return statistically significant coefficients, and the coefficients on the contemporaneous house price variable terms are very similar to before.

TABLE 3
ROBUSTNESS ESTIMATES FOR RELATIONSHIP BETWEEN LOG HOUSE PRICES AND LOG HOURS OF WORK FOR YOUNG WOMEN IN MARRIED OR COHABITING COUPLES: INDIVIDUAL FIXED EFFECTS ESTIMATES

	(1) Excluding movers	(2) Simulated prices	(3) Initial owner	(4) IV owner	(5) Forward prices
(1) log hp–owner	−0.179***	−0.173***	−0.175***	−0.173***	−0.174***
	(0.0335)	(0.0327)	(0.0257)	(0.0246)	(0.0346)
(2) log hp–renter	0.00253	0.00246	0.00286	0.00257	0.00578
	(0.00346)	(0.00357)	(0.00435)	(0.00496)	(0.0074)
(3) log hp–owner$_{t+1}$	—	—	—	—	0.0095
					(0.0288)
(4) log hp–renter$_{t+1}$	—	—	—	—	−0.0349
					(0.0287)
p-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.0000
p-value test (3) = (4)	—	—	—	—	0.4916
N	11,206	12,727	12,727	12,727	11,161

Notes
Dependent variable: natural log of hours of work for sample of individuals with non-zero hours. Sample: head of household plus spouse/partner BHPS 1991–2009. Individual fixed effects estimates. Self-reported house price for owners is instrumented using log local authority mean house price. Additional control variables not shown in table: age, age squared; educational dummies for highest educational achievement (HND, GCSE, A level, degree (or equivalents)); 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, local authority dummies, year dummies, renter–local authority and owner–local authority time trends.
*, **, *** indicate $p < 0.05$, $p < 0.01$, $p < 0.001$, respectively.
Cluster (local authority) standard errors in parentheses.
'*p*-value test' reports values from test for equivalence of coefficients in rows (1) and (2); similarly rows (3) and (4).

Using the self-reported house prices of owners in the data, we also estimate IV difference-in-differences models in which (owner) self-reported house prices are instrumented by the local authority house price (following equations (2) and (3)). Results are shown in Tables A1–A4 of the Appendix, which also report first-stage estimates. Coefficient estimates in these models are very similar to those in the non-IV models shown in Tables 2 and 3. Results show labour supply responses only among younger married women (the coefficient value is attenuated slightly in the IV models), with the coefficients again showing statistically significant differences in responses of young female owners compared with young female renters.[13]

This finding for young married/cohabiting women may be driven in part by childcare-related decisions. Labour supply decisions related to childcare needs could potentially be affected by house price increases in either direction. For young families seeking to trade up in the housing market in future, higher house prices imply higher future housing costs, so we might expect labour supply to increase. For those who use paid childcare, higher house prices might correlate with childcare costs (if, for example, house prices affect the cost of local childcare provision), encouraging substitution away from paid childcare and towards providing childcare within the family.

To explore whether the effects that we find for young married/cohabiting women extend beyond those with children, we re-estimate the series of models in Table 4 for a subsample of respondents within the category who either do not have children, or have older children (aged 12 or over). For these groups, childcare needs are reduced or do not exist. Table 4 estimates show that the coefficient on the instrumented house price for owners is negative and statistically significant in these specifications, and slightly larger in absolute magnitude compared with the estimates in Table 2. These results show that the labour supply responses of young married women are not tied to childcare needs, suggesting instead that the wealth effects of house prices increases alter labour supply patterns of second workers, with or without direct opportunity costs of working (in the form of the cost of childcare provision).

Our results from estimates for hours of work show, therefore, that house price gains lead to reduced female labour supply among homeowning married or cohabiting couples, including those without children. This result is consistent with a model in which house price gains operate a wealth effect at the variable margin of adjustment of household labour 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 these wealth effects.

House prices and labour market participation

Next we present results for decision to work on the extensive margin. Table 5 presents estimates from the participation equation, where the labour market participation dummy variable takes value 1 if the respondent is employment or unemployed, and takes value 0 otherwise. We estimate linear probability models with individual fixed effects plus local authority and time effects and renter–local authority plus owner–local authority time trends, following the hours of work specification shown earlier. Results are shown by subgroups using the same convention as in Table 2, with subgroups defined over relationship status, gender and age.

Results show that house price gains decrease the likelihood of participation among young married/cohabiting women and among older men (both married and unmarried). For each of these subsamples, the coefficient on the owner house price term is negative

TABLE 4
ESTIMATES FOR RELATIONSHIP BETWEEN LOG HOUSE PRICES AND LOG HOURS OF WORK FOR WOMEN IN MARRIED OR COHABITING COUPLES WITHOUT CHILDREN, OR WITH CHILDREN AGED 12 OR OVER: INDIVIDUAL FIXED EFFECTS ESTIMATES

	(1) Excluding movers	(2) Simulated prices	(3) Initial owner	(4) IV owner	(5) Forward prices
(1) log hp–owner	−0.148***	−0.147***	−0.141***	−0.150***	−0.152***
	(0.0226)	(0.0208)	(0.0246)	(0.0277)	(0.0264)
(2) log hp–renter	0.00645	0.00335	0.00275	0.00245	0.00424
	(0.00436)	(0.00555)	(0.00834)	(0.00623)	(0.00754)
(3) log hp–owner$_{t+1}$	—	—	—	—	0.0057
					(0.0073)
(4) log hp–renter$_{t+1}$	—	—	—	—	−0.00455
					(0.0374)
p-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.0000
p-value test (3) = (4)	—	—	—	—	0.5026
N	3922	4024	4024	4024	3978

Notes
Dependent variable: natural log of hours of work for sample of individuals with non-zero hours. Sample: head of household plus spouse/partner BHPS 1991–2009. Individual fixed effects estimates. Self-reported house price for owners is instrumented using log local authority mean house price. Additional control variables not shown in table: age, age squared; educational dummies for highest educational achievement (HND, GCSE, A level, degree (or equivalents)); 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, local authority dummies, year dummies, renter–local authority and owner–local authority time trends.
*, **, *** indicate $p < 0.05$, $p < 0.01$, $p < 0.001$, respectively.
Cluster (local authority) standard errors in parentheses.
'p-value test' reports values from test for equivalence of coefficients in rows (1) and (2); similarly rows (3) and (4).

and statistically significant at the 1% level of older single male individuals and at the 0.1% level for older married/cohabiting men and young married/cohabiting 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 shows that female participation among middle-aged and older married/cohabiting women decreases with the unemployment rate, and participation among most groups decreases with a positive financial expectation, though the coefficients on these variables are in each model not statistically significantly different from zero.

The coefficient estimates on the owner house price term for young married/cohabiting women is −0.134, which is statistically significant at the 0.1% level. Hence a 10% increase in house prices causes a 1.3 percentage point reduction in the likelihood of participation for this group. The labour 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/cohabiting 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 that we observe for young women are specific to married/cohabiting young women only. Below we analyse the labour market destinations of this group when they leave the labour force, and consider whether this withdrawal is likely to be temporary or permanent.

We find statistically significant effects for older men and women. For the subgroups of older married/cohabiting and single men, the coefficient estimates on the owner house price variable are −0.153 and −0.127. 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. For older married and single women the coefficients are −0.0436 and −0.0477 (in the latter case statistically significant at only the

TABLE 5
ESTIMATES FOR RELATIONSHIP BETWEEN LOG HOUSE PRICES AND LABOUR MARKET PARTICIPATION FOR WOMEN AND MEN BY MARITAL STATUS AND AGE GROUP: INDIVIDUAL FIXED EFFECTS ESTIMATES

	Women			Men		
	Age < 40	Age 40–54	Age > 54	Age < 40	Age 40–54	Age > 54
Panel A: Individuals in married or cohabiting couples						
(1) log hp–owner	−0.134***	−0.0230	−0.0436	−0.00202	0.0143	−0.153***
	(0.0180)	(0.0238)	(0.0225)	(0.0159)	(0.0187)	(0.0234)
(2) log hp–renter	0.00185	−0.00211	−0.00120	−0.00263	−0.000802	0.00464
	(0.00595)	(0.00380)	(0.00467)	(0.0252)	(0.00293)	(0.00465)
(3) unemployment (%)	0.136	−0.112	−0.280	−0.00461	0.0872	−0.120
	(0.0783)	(0.0579)	(0.656)	(0.0403)	(0.0450)	(0.0659)
(4) financial expectation	−0.112	0.0220	−0.00928	−0.0568	−0.00983	−0.0302
	(0.164)	(0.136)	(0.234)	(0.0806)	(0.100)	(0.207)
p-value test (1) = (2)	0.108	0.074	0.271	0.078	0.098	0.355
Panel B: Single individuals						
(1) log hp–owner	−0.0106	−0.0512	0.0477	0.0550	0.00666	−0.127**
	(0.0144)	(0.0478)	(0.0284)	(0.0463)	(0.0532)	(0.0310)
(2) log hp–renter	0.00639	−0.000236	0.00523	−0.00189	−0.0229**	0.0884
	(0.00410)	(0.00824)	(0.00498)	(0.00306)	(0.00868)	(0.0689)
(3) unemployment (%)	−0.0289	0.0952	0.0620	0.0796	−0.193	−0.0662
	(0.136)	(0.132)	(0.0783)	(0.128)	(0.141)	(0.145)
(4) financial expectation	−0.411	−0.131	0.0264	−0.496	0.295	0.779
	(0.279)	(0.299)	(0.301)	(0.263)	(0.309)	(0.437)
p-value test (1) = (2)	0.218	0.096	0.258	0.376	0.132	0.280

Notes
Dependent variable: 1/0 dummy variable taking value 1 if the individual is in full- or part-time employment and 0 otherwise. Sample: head of household plus spouse/partner BHPS 1991–2009. Individual fixed effects estimates. Self-reported house price for owners is instrumented using log local authority mean house price. Additional control variables not shown in table: age, age squared; educational dummies for highest educational achievement (HND, GCSE, A level, degree (or equivalents)); 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, local authority dummies, year dummies, renter–local authority and owner–local authority time trends.
*, **, *** indicate $p < 0.05$, $p < 0.01$, $p < 0.001$, respectively.
Cluster (local authority) standard errors in parentheses.
'*p*-value test' reports values from test for equivalence of coefficients in rows (1) and (2).

10% level), which translate to less than a 3% decrease in likelihood of participation. Later we discuss these differences in effect sizes across older men and women.

Table 6 presents results from our robustness specifications. As in the hours results, here we show the robustness estimates for subsamples for which the main results returned statistically significant results for the owner house price coefficient (young married/cohabiting women, older married and single men). Results show very similar coefficient estimates 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

TABLE 6
ROBUSTNESS ESTIMATES FOR RELATIONSHIP BETWEEN LOG HOUSE PRICES AND LABOUR MARKET PARTICIPATION FOR SELECTED GROUPS: INDIVIDUAL FIXED EFFECTS ESTIMATES

	Excluding movers	Simulated prices	Initial owner	IV owner	Forward prices
Women in married or cohabiting couples aged < 40					
(1) log hp–owner	−0.142***	−0.137***	−0.132***	−0.135***	−0.137***
	(0.0126)	(0.0213)	(0.0142)	(0.0224)	(0.0224)
(2) log hp–renter	0.00134	0.00125	0.00174	0.00113	0.00153
	(0.00506)	(0.0046)	(0.0042)	(0.00474)	(0.00436)
(3) log hp–owner$_{t+1}$	—	—	—	—	0.0261
					(0.0346)
(4) log hp–renter$_{t+1}$	—	—	—	—	0.00135
					(0.00405)
p-value test (1) = (2)	0.123	0.096	0.094	0.099	0.102
p-value test (3) = (4)	0.0000	0.0000	0.0000	0.0000	0.0000
Men in married or cohabiting couples aged > 54					
(1) log hp–owner	−0.156***	−0.148***	−0.143***	−0.144***	−0.148***
	(0.0235)	(0.0254)	(0.0253)	(0.0245)	(0.0253)
(2) log hp–renter	0.00156	0.00224	0.00223	0.00263	0.00213
	(0.00446)	(0.00406)	(0.00453)	(0.00452)	(0.004t4)
(3) log hp–owner$_{t+1}$	—	—	—	—	0.0164
					(0.0281)
(4) log hp–renter$_{t+1}$	—	—	—	—	0.00231
					(0.00484)
p-value test (1) = (2)	0.357	0.363	0.375	0.353	0.362
p-value test (3) = (4)	0.0000	0.0000	0.0000	0.0000	0.0000
Single men aged > 54					
(1) log hp–owner	−0.128**	−0.128**	−0.125**	−0.127**	−0.122**
	(0.0353)	(0.0364)	(0.0342)	(0.0353)	(0.0324)
(2) log hp–renter	0.00225	0.00226	0.00274	0.00223	0.00203
	(0.00616)	(0.00642)	(0.0064)	(0.00774	(0.0071t
(3) log hp–owner$_{t+1}$	—	—	—	—	0.0219
					(0.0308)
(4) log hp–renter$_{t+1}$	—	—	—	—	0.00227
					(0.00746)
p-value test (1) = (2)	0.330	0.283	0.279	0.283	0.242
p-value test (3) = (4)	0.0000	0.0000	0.0000	0.0000	0.1524

Notes
See Table 4.

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 subsample, the 'forward prices' falsification test yields no evidence of labour 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.

Labour market destinations

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

We might expect that the withdrawal of young married/cohabiting women is temporary due to career breaks for children. Recent studies based on US data have also found that house price increases raise the likelihood of couples having children (Lovenheim and Mumford 2013). They do not examine the labour market consequences of this. Most women undertake some form of 'maternity leave' or other leave following childbirth. In our data we have information on the main activity of non-working individuals so we can estimate whether house price gains induce this form of activity for young women. To do so, we estimate our labour 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 labour market activity described earlier.

Results from these estimates are shown in Table 7. For completeness, we estimate models for four subgroups: young and middle-aged married women plus young and middle-aged 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/cohabiting women return a positive coefficient of 0.0912 for young women and 0.0474 for middle-aged 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 negative and statistically significant at the 5% level, providing some evidence that house price gains decrease the likelihood of leaving the labour force to care for children among young married/cohabiting women renters.

The coefficient estimates imply large proportional effects of house price gains on the likelihood of leaving the labour force to undertake family care activity. The baseline family care rate among young married homeowning women is 18%, hence the impact of a 10% increase in house prices is to raise the likelihood of family care among this group by on average 5%. For middle-aged married women, the baseline rate is 13% and implied effect of a 10% increase in house prices is 3.9%. We present robustness results in Appendix Table A5. These results show that 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 labour supply equation with the dependent variable replaced with a 1/0 indicator for whether the individual is retired. We define retirement as

TABLE 7
ESTIMATES FOR RELATIONSHIP BETWEEN LOG HOUSE PRICES AND NON-WORKING FULL-TIME CHILDCARE FOR WOMEN BY MARITAL STATUS AND AGE GROUP: INDIVIDUAL FIXED EFFECTS ESTIMATES

	Married or cohabiting		Single	
	Age < 40	Age 40–54	Age < 40	Age 40–54
(1) log hp–owner	0.0912***	0.0474*	0.0750	0.0921
	(0.0225)	(0.0224)	(0.0957)	(0.0858)
(2) log hp–renter	−0.0108*	0.000403	−0.00141	0.00358
	(0.00523)	(0.00350)	(0.00346)	(0.00664)
(3) unemployment (%)	−0.0972	0.0605	−0.0594	0.00111
	(0.0687)	(0.0534)	(0.115)	(0.106)
(4) financial expectation	−0.141	−0.124	0.135	0.275
	(0.144)	(0.125)	(0.235)	(0.241)
p-value test (1) = (2)	0.105	0.031	0.139	0.110

Notes
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-labour income, homeowner dummy, local authority dummies, year dummies, renter–local authority and owner–local authority time trends.
*, **, *** indicate $p < 0.05$, $p < 0.01$, $p < 0.001$, respectively.
Cluster (local authority) standard errors in parentheses.

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 labour market.

In Table 8 we report estimates for subsamples 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 0.115 and 0.120 imply that a 10% increase in house prices raises the likelihood of retirement among men by 1.1 percentage points and 1.2 percentage points, respectively, for each group. Baseline retirement rates for these groups are 43% for male married and 34% for male single. Hence a 10% increase in prices causes a 1.9% increase in the likelihood of retirement for male married and a 4.2% increase for male single. Results from robustness specifications shown in Appendix Table A6 confirm very similar coefficient estimates from the alternative specifications.

V. DISCUSSION

Our results show heterogeneous labour 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 homeowners are responsive to house price movements, but strong effects are found for younger married female owners and for older married and single owners. These effects are consistent with labour supply adjustment by marginal workers at the margins of family labour supply (young women) and lifetime labour supply (older men). The economic reasons for these effects may be different, however.

TABLE 8
ESTIMATES FOR RELATIONSHIP BETWEEN LOG HOUSE PRICES AND RETIREMENT FOR OLDER MEN AND WOMEN BY MARITAL STATUS: INDIVIDUAL FIXED EFFECTS ESTIMATES

	Women		Men	
	Married age > 54	Single age > 54	Married age > 54	Single age > 54
(1) log hp–owner	−0.00111	0.0195	0.115***	0.120**
	(0.0287)	(0.0318)	(0.0259)	(0.0488)
(2) log hp–renter	−0.00505	0.00548	−0.00197	−0.00731
	(0.00594)	(0.00677)	(0.00522)	(0.00804)
(3) unemployment (%)	−0.215*	−0.0474	−0.0182	−0.0952
	(0.0836)	(0.106)	(0.0739)	(0.170)
(4) financial expectation	0.406	−0.172	−0.132	0.457
	(0.298)	(0.409)	(0.233)	(0.510)
p-value test (1) = (2)	0.306	0.350	0.281	0.329

Notes
Sample: head of household aged 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-labour income, homeowner dummy, local authority dummies, year dummies, renter–local authority and owner–local authority time trends.
*, **, *** indicate $p < 0.05$, $p < 0.01$, $p < 0.001$, respectively.
Cluster (local authority) standard errors in parentheses.

The response of labour supply of young female owners to housing wealth gains might arise through two routes: either through the effect of having young children (since a wealth gain to the household may allow a young married women to reduce hours, or at least to delay the return to full-time work after the birth of a child, and house prices may correlate with childcare costs), or through alleviating borrowing constraints in the upswing. Table 7 suggests that rising house prices are indeed associated with an increased propensity for the wife or cohabiting partner to be engaged in full-time childcare. Any correlation between childcare costs and house prices might be a potential explanation of this result. However, Table 4 suggests that married women without children are also affected by house price changes. Hence childcare costs do not seem to be the reason for the observed result, and childcare is not the explanation for the result among childless couples. This suggests that other factors are also at work.

Although house price increases for young owners are unlikely to represent significant lifetime net wealth gains as young owners typically trade up to larger houses in future (the price of which also increase with general house price increases), house price gains may loosen borrowing constraints and this may impact on labour supply decisions. This is likely only under certain circumstances: for example, in the upswing rather than during a general reduction in house prices, and where the trade up is not disproportionately large. However, Cooper (2013) shows that among US households, the main route by which house price gains influence consumption is through loosening borrowing constraints; a result reinforcing the disproportionate response of consumption to house price changes among 'collateral-constrained' households shown by Disney and Gathergood (2011). Our present results suggest that this is potentially true also for labour supply among young UK households.

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, labour supply effects are associated with having children, an activity that may have been postponed by households until borrowing constraints relaxed.

The response of older male owners (and, less strongly, among older married women) appears consistent with a pure lifecycle 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 that they intend to downsize after retirement. For these households, house price gains represent pure wealth gains and we can interpret the labour supply response as a pure wealth response similar to the effect of a lottery win or inheritance. Typically, however, older workers tend to reduce labour supply discretely at later ages, either by full retirement or by retiring from a full-time job and switching to part-time work. Hence we expect a stronger effect on the participation margin than on hours of work conditional on retaining the same job. The result may be less strong for married women because many married women in these cohorts have been working part-time for much of their lifetime and hence do not consider 'retirement' as a discrete labour market decision (Table 8).[14]

Our results have implications for the business cycle dynamics of labour supply for the groups of individuals who respond to house price changes. House prices are procyclical, therefore our results suggest that housing wealth gains are a procyclical driver of leisure (for those older men who retire), or family care (for those younger married/cohabiting women who leave the labour force), in contrast to wages, which are a procyclical driver of wealth. However, the specifications that we estimate include time dummies to capture time-specific 'macroeconomic' effects. This means that our estimates for labour 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 necessary for identification, but doing so does not allow us to use our coefficient estimates to calculate the business cycle effects of house price movements on labour 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 −0.138, also statistically significant at the 1% level. For older married men, the coefficient in the model without time dummies is −0.146 (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 labour 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 statistically significant (at the 1% level) in each of these specifications and takes a negative value. Hence time variation in labour 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 on labour market participation during the recent recession. Our estimates imply that housing wealth effects have a strong influence of labour supply over the business cycle compared with local labour market conditions, and can explain a large share of labour supply movements during the recent recession.

Our calculations here can be considered as only 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 labour supply rate among young married women by 1.5 percentage points (pp), among older married/cohabiting men by 1.6pp, and among older single men by 1.3pp. We evaluate these estimated effects against changes in house prices and labour supply during the recent UK recession, the eight-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 labour market participation rate for young women fell from 72.9% to 71.1% (statistics on labour market participation by marital status are not available). Our estimates imply that the 27% fall in price increased labour 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/cohabiting owners, an increase in unemployment of this magnitude leads to a 2.3pp decline in labour market participation. Hence in our estimates, the wealth effect that encourages labour market participation arising from house price changes more than offsets the effect of labour market conditions captured by the local unemployment rate on labour market participation for this group.

Equivalent calculations for older men also show that our estimates imply economically important housing wealth effects during the recent recession. The participation rate of older men (using the same definition of age 55–75 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 that 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/cohabiting and single men implies that a 10% fall in house prices causes a 1.45pp increase in labour market participation rate.

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

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

VI. Conclusion

This paper has presented empirical estimates of the impact of housing wealth on labour supply behaviour among working-age individuals in the UK using individual-level panel data. Results show large responses to housing gains and losses and certain groups that 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/cohabiting women and also among older men close to retirement age.

These results show that housing wealth impacts on household labour supply behaviour 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 labour supply are jointly determined as households evaluating the marginal utility of consumption alongside the marginal utility of leisure. However, our results show that labour supply responses across groups are not solely attributable to pure lifecycle wealth effects whereby older individuals 'win' and younger individuals 'lose' but instead reflect down-payment or liquidity constraint effects that drive labour supply responses of younger individuals. Our results are also of economic significance for understanding the business cycle dynamics of labour supply for those groups that respond to house price movements.

ACKNOWLEDGMENTS

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. We should like to thank the editor and two referees for helpful comments on an earlier draft of this paper. Our thanks also 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.

APPENDIX

TABLE A1
INSTRUMENTAL VARIABLE ESTIMATES FOR RELATIONSHIP BETWEEN LOG HOUSE PRICES AND LOG HOURS OF WORK FOR WOMEN AND MEN BY MARITAL STATUS AND AGE GROUP

	Individuals in married or cohabiting couples					
	Women			Men		
	(1) Age < 40	(2) Age 40–54	(3) Age > 54	(4) Age < 40	(5) Age 40–54	(6) Age > 54
(1) log $\hat{H}$–owner	−0.125***	−0.0113	0.0392	0.0329	0.0201	−0.0153
	(0.0203)	(0.0329)	(0.0703)	(0.0325)	(0.0334)	(0.0463)
(2) log H–renter	0.00514	−0.00129	0.00163	−0.00763	−0.00125	−0.0284
	(0.00435)	(0.00492)	(0.0128)	(0.00734)	(0.00385)	(0.0743)
(3) unemployment (%)	−0.0327	−0.00262	0.262	−0.0425	−0.0463	−0.0407
	(0.0694)	(0.0626)	(0.141)	(0.0701)	(0.0564)	(0.0346)
(4) financial expectation	−0.348**	−0.139	0.699	−0.336	0.362	−0.216
	(0.130)	(0.101)	(0.529)	(0.157)	(0.147)	(0.164)
(5) age	0.0151**	0.0134**	0.0115**	0.0122**	0.0138**	0.0118**
	(0.00592)	(0.00525)	(0.00535)	(0.00402)	(0.00322)	(0.00501)
(6) age squared	−0.00952	−0.00532	−0.00632	−0.00935	−0.00762	−0.00325
	(0.00768)	(0.00348)	(0.00542)	(0.00838)	(0.00984)	(0.00778)
(7) degree = 1	0.263**	0.205**	0.266**	0.253**	0.343**	0.225**
	(0.164)	(0.162)	(0.164)	(0.126)	(0.154)	(0.101)
(8) alevels = 1	0.165**	0.146**	0.165**	0.164**	0.163**	0.152**
	(0.061)	(0.036)	(0.031)	(0.023)	(0.031)	(0.061)
(9) olevels = 1	0.198**	0.174**	0.194**	0.161**	0.131**	0.158**
	(0.0575)	(0.0345)	(0.0325)	(0.0365)	(0.0575)	(0.0575)
(10) local wage rate	2.106	3.295	2.263	2.144	2.823	2.155
	(1.524)	(2.934)	(1.643)	(1.532)	(1.501)	(1.554)
(11) young children = 1	0.265	0.235	0.232	0.258	0.645	0.262
	(0.163)	(0.163)	(0.233)	(0.133)	(0.643)	(0.173)
(12) older children = 1	0.409	0.474	0.464	0.453	0.364	0.526
	(0.223)	(0.223)	(0.223)	(0.225)	(0.423)	(0.203)
(13) spouse employed = 1	0.895**	0.822**	0.801**	0.893**	0.637**	0.800**
	(0.286)	(0.299)	(0.386)	(0.234)	(0.316)	(0.256)
(14) spouse self-employed = 1	0.971**	0.902**	0.923**	0.984**	0.975**	0.931**
	(0.392)	(0.352)	(0.342)	(0.345)	(0.343)	(0.300)
(15) mortgage debt (£)	0.0111	0.0113	0.0164	0.0143	0.0321	0.0412
	(0.0308)	(0.0245)	(0.0547)	(0.0648)	(0.0377)	(0.0508)
(16) health status (1–5)	−0.0877	−0.0834	−0.0467	−0.0747	−0.0857	−0.0787
	(0.145)	(0.164)	(0.165)	(0.25)	(0.146)	(0.133)
(17) smoker = 1	−0.221	−0.275	−0.223	−0.235	−0.283	−0.321
	(0.244)	(0.537)	(0.274)	(0.224)	(0.264)	(0.278)
p-value test (1) = (2)	0.0000	0.3256	0.6219	0.4825	0.6735	0.2734
N	12,727	12,597	3636	12,266	11,384	4089

Notes
Dependent variable: natural log of hours of work for sample of individuals with non-zero hours. Sample: head of household plus spouse/partner BHPS 1991–2009. Individual fixed effects estimates. Self-reported house price for owners is instrumented using log local authority mean house price. Additional control variables not shown in table: age, age squared; educational dummies for highest educational achievement (HND, GCSE, A level, degree (or equivalents)); 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, local authority dummies, year dummies, renter–local authority and owner–local authority time trends.
*, **, *** indicate $p < 0.05$, $p < 0.01$, $p < 0.001$, respectively.
Cluster (local authority) standard errors in parentheses.
'p-value test' reports values from test for equivalence of coefficients in rows (1) and (2).

TABLE A2
FIRST-STAGE INSTRUMENTAL VARIABLE ESTIMATES FOR RELATIONSHIP BETWEEN LOG HOUSE PRICES AND LOG HOURS OF WORK FOR WOMEN AND MEN BY MARITAL STATUS AND AGE GROUP

	Individuals in married or cohabiting couples					
	Women			Men		
	(1) Age < 40	(2) Age 40–54	(3) Age > 54	(4) Age < 40	(5) Age 40–54	(6) Age > 54
(1) local authority house price	0.728**	0.682**	0.812**	0.598**	0.682**	0.842**
	(0.219)	(0.184)	(0.222)	(0.128)	(0.219)	(0.200)
(3) unemployment (%)	−0.0425**	−0.0523**	0.0428**	−0.0453**	−0.0455**	−0.0409**
	(0.0153)	(0.0153)	(0.134)	(0.0102)	(0.0164)	(0.0146)
(4) financial expectation	−0.0243	−0.0432	0.0324	−0.0243	0.0423	−0.0843
	(0.330)	(0.352)	(0.352)	(0.399)	(0.351)	(0.623)
(5) age	0.0134**	0.0124**	0.0163**	0.0146**	0.0155**	0.0134**
	(0.00408)	(0.00566)	(0.00532)	(0.00434)	(0.00632)	(0.00501)
(6) age squared	−0.00634	−0.00645	−0.00633	−0.00435	−0.00637	−0.00654
	(0.00354)	(0.00337)	(0.00654)	(0.00436)	(0.00964)	(0.00678)
(7) degree = 1	0.266**	0.237**	0.275**	0.436**	0.355**	0.301**
	(0.100)	(0.101)	(0.100)	(0.100)	(0.105)	(0.15)
(8) alevels = 1	0.163**	0.246**	0.165**	0.152**	0.153**	0.155**
	(0.061)	(0.066)	(0.031)	(0.063)	(0.041)	(0.031)
(9) olevels = 1	0.0834**	0.0632**	0.0673**	0.628**	0.623**	0.644**
	(0.0565)	(0.0353)	(0.0305)	(0.0328)	(0.0555)	(0.0535)
(10) local wage rate	2.106	3.295	2.263	2.144	2.823	2.155
	(1.524)	(2.934)	(1.643)	(1.532)	(1.501)	(1.554)
(11) young children = 1	0.265	0.235	0.232	0.258	0.645	0.262
	(0.161)	(0.143)	(0.237)	(0.133)	(0.655)	(0.173)
(12) older children = 1	0.419	0.474	0.464	0.456	0.334	0.526
	(0.223)	(0.223)	(0.223)	(0.225)	(0.425)	(0.206)
(13) spouse employed = 1	0.895**	0.824**	0.801**	0.893**	0.637**	0.800**
	(0.286)	(0.699)	(0.386)	(0.234)	(0.346)	(0.256)
(14) spouse self-employed = 1	0.971**	0.902**	0.923**	0.984**	0.975**	0.931**
	(0.392)	(0.352)	(0.342)	(0.345)	(0.343)	(0.300)
(15) mortgage debt (£)	0.0101	0.0163	0.0144	0.0153	0.0321	0.0432
	(0.0358)	(0.0245)	(0.0547)	(0.0648)	(0.0377)	(0.0548)
(16) health status (1–5)	−0.0877	−0.0834	−0.0467	−0.0747	−0.0857	−0.0787
	(0.145)	(0.164)	(0.165)	(0.25)	(0.146)	(0.133)
(17) smoker = 1	−0.221	−0.244	−0.223	−0.235	−0.233	−0.321
	(0.244)	(0.537)	(0.274)	(0.224)	(0.264)	(0.278)
First-stage *F*-statistic	21.04	21.39	20.98	21.77	21.06	21.83
N	12,727	12,597	3636	12,266	11,384	4089

Notes
First-stage estimates for models reported in Table A1.

Table A3
Instrumental Variable Estimates for Relationship Between Log House Prices and Log Hours of Work for Women and Men by Marital Status and Age Group

	Single individuals					
	Women			Men		
	(1) Age < 40	(2) Age 40–54	(3) Age > 54	(4) Age < 40	(5) Age 40–54	(6) Age > 54
(1) log $\hat{H}$–owner	−0.0142	−0.00352	−1.002	0.0509	−0.0635	−0.374
	(0.0153)	(0.0532)	(1.235)	(0.0834)	(0.0719)	(3.253)
(2) log H–renter	−0.00234	0.0033	−1.343	−0.0325	0.00183	−0.0673
	(0.00352)	(0.00323)	(1.435)	(0.0145)	(0.0324)	(0.434)
(3) unemployment (%)	−0.246	−0.0648	0.0932	−0.352	−0.00345	0.664
	(0.124)	(0.0535)	(0.374)	(0.295)	(0.0555)	(7.123)
(4) financial expectation	−0.152	0.0467	0.623	−0.294	−0.123	3.735
	(0.263)	−0.00142	(1.353)	(0.434)	(0.123)	(12.06)
(5) age	0.0141**	0.100*	0.0113**	0.0162**	0.153**	0.0118**
	(0.00563)	(0.00501)	(0.00555)	(0.00702)	(0.00622)	(0.00501)
(6) age squared	−0.00937	−0.00632	−0.00662	−0.00835	−0.00453	−0.00325
	(0.00786)	(0.00364)	(0.00742)	(0.00846)	(0.00934)	(0.00778)
(7) degree = 1	0.243**	0.245**	0.745*	0.237**	0.363**	0.201**
	(0.137)	(0.136)	(0.308)	(0.156)	(0.196)	(0.0975)
(8) alevels = 1	0.167**	0.175**	0.215**	0.185**	0.162**	0.132**
	(0.0541)	(0.0466)	(0.0316)	(0.0553)	(0.0213)	(0.041)
(9) olevels = 1	0.190**	0.144**	0.196**	0.174**	0.151**	0.128**
	(0.0525)	(0.0375)	(0.0345)	(0.036)	(0.0574)	(0.0565)
(10) local wage rate	2.133	3.233	2.633	2.146	2.853	2.255
	(1.542)	(2.963)	(1.643)	(2.632)	(1.522)	(1.563)
(11) young children = 1	0.235	0.245	0.262	0.248	0.665	0.362
	(0.186)	(0.166)	(0.263)	(0.133)	(0.648)	(0.168)
(12) older children = 1	0.423	0.474	0.474	0.434	0.347	0.523
	(0.433)	(0.286)	(0.235)	(0.265)	(0.474)	(0.211)
(13) spouse employed = 1	0.875**	0.833**	0.811**	0.883**	0.677**	0.863**
	(0.215)	(0.289)	(0.308)	(0.2634)	(0.301)	(0.264)
(14) spouse self-employed = 1	0.951**	0.911**	0.944**	0.984**	0.976**	0.831**
	(0.336)	(0.362)	(0.385)	(0.335)	(0.352)	(0.340)
(15) mortgage debt (£)	0.0146	0.0127	0.0134	0.0123	0.0372	0.0422
	(0.0538)	(0.0256)	(0.0427)	(0.0655)	(0.0353)	(0.0548)
(16) health status (1–5)	−0.0647	−0.0874	−0.0667	−0.0701	−0.0822	−0.0763
	(0.135)	(0.146)	(0.187)	(0.251)	(0.156)	(0.134)
(17) smoker = 1	−0.221	−0.295	−0.253	−0.252	−0.289	−0.322
p-value test (1) = (2)	0.6738	0.7946	0.8152	0.3732	0.1846	0.8457
N	3747	3064	1098	2620	1778	582

Notes
Dependent variable: natural log of hours of work for sample of individuals with non-zero hours. Sample: head of household plus spouse/partner BHPS 1991–2009. Individual fixed effects estimates. Self-reported house price for owners is instrumented using log local authority mean house price. Additional control variables not shown in table: age, age squared; educational dummies for highest educational achievement (HND, GCSE, A level, degree (or equivalents)); 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, local authority dummies, year dummies, renter–local authority and owner–local authority time trends.
*, **, *** indicate $p < 0.05$, $p < 0.01$, $p < 0.001$, respectively.
Cluster (local authority) standard errors in parentheses.
'p-value test' reports values from test for equivalence of coefficients in rows (1) and (2).

TABLE A4
FIRST-STAGE INSTRUMENTAL VARIABLE ESTIMATES FOR RELATIONSHIP BETWEEN LOG HOUSE PRICES AND LOG HOURS OF WORK FOR WOMEN AND MEN BY MARITAL STATUS AND AGE GROUP

	Single individuals					
	Women			Men		
	(1) Age < 40	(2) Age 40–54	(3) Age > 54	(4) Age < 40	(5) Age 40–54	(6) Age > 54
(1) local authority house price	0.788**	0.695**	0.846**	0.602**	0.681**	0.863**
	(0.214)	(0.163)	(0.244)	(0.175)	(0.217)	(0.240)
(3) unemployment (%)	−0.0425**	−0.0523**	0.0428**	−0.0453**	−0.0455**	−0.0409**
	(0.0153)	(0.0153)	(0.134)	(0.0102)	(0.0164)	(0.0146)
(4) financial expectation	−0.0243	−0.0432	0.0324	−0.0243	0.0423	−0.0843
	(0.330)	(0.352)	(0.352)	(0.399)	(0.351)	(0.623)
(5) age	0.0134**	0.0124**	0.0163**	0.0146**	0.0155**	0.0134**
	(0.00408)	(0.00566)	(0.00532)	(0.00434)	(0.00632)	(0.00501)
(6) age squared	−0.00634	−0.00645	−0.00633	−0.00435	−0.00637	−0.00654
	(0.00354)	(0.00337)	(0.00654)	(0.00436)	(0.00964)	(0.00678)
(7) degree = 1	0.266**	0.237**	0.275**	0.436**	0.355**	0.301**
	(0.100)	(0.101)	(0.100)	(0.100)	(0.105)	(0.15)
(8) alevels = 1	0.163**	0.246**	0.165**	0.152**	0.153**	0.155**
	(0.061)	(0.066)	(0.031)	(0.063)	(0.041)	(0.031)
(9) olevels = 1	0.0834**	0.0632**	0.0673**	0.628**	0.623**	0.644**
	(0.0565)	(0.0353)	(0.0305)	(0.0328)	(0.0555)	(0.0535)
(10) local wage rate	2.106	3.295	2.263	2.144	2.823	2.155
	(1.524)	(2.934)	(1.643)	(1.532)	(1.501)	(1.554)
(11) young children = 1	0.265	0.235	0.232	0.258	0.645	0.262
	(0.161)	(0.143)	(0.237)	(0.133)	(0.655)	(0.173)
(12) older children = 1	0.419	0.474	0.464	0.456	0.334	0.526
	(0.223)	(0.223)	(0.223)	(0.225)	(0.425)	(0.206)
(13) spouse employed = 1	0.895**	0.824**	0.801**	0.893**	0.637**	0.800**
	(0.286)	(0.699)	(0.386)	(0.234)	(0.346)	(0.256)
(14) spouse self-employed = 1	0.971**	0.902**	0.923**	0.984**	0.975**	0.931**
	(0.392)	(0.352)	(0.342)	(0.345)	(0.343)	(0.300)
(15) mortgage debt (£)	0.0101	0.0163	0.0144	0.0153	0.0321	0.0432
	(0.0358)	(0.0245)	(0.0547)	(0.0648)	(0.0377)	(0.0548)
(16) health status (1–5)	−0.0877	−0.0834	−0.0467	−0.0747	−0.0857	−0.0787
	(0.145)	(0.164)	(0.165)	(0.25)	(0.146)	(0.133)
(17) smoker = 1	−0.221	−0.244	−0.223	−0.235	−0.233	−0.321
	(0.244)	(0.537)	(0.274)	(0.224)	(0.264)	(0.278)
First-stage F-statistic	21.06	21.22	21.99	21.25	21.36	21.75
N	12,727	12,597	3636	12,266	11,384	4089

Notes
First-stage estimates for models reported in Table A1.

Table A5
Robustness Estimates for Relationship Between Non-Working Full-time Childcare for Married Women by Age Group: Individual Fixed Effects Estimates

	Excluding movers	Simulated prices	Initial owner	IV owner	Forward prices
Women in married or cohabiting couples aged < 40					
(1) log hp–owner	0.0708***	0.0693***	0.0801***	0.0899***	0.0712***
	(0.0364)	(0.0214)	(0.0265)	(0.0284)	(0.0207)
(2) log hp–renter	−0.0111*	−0.0131*	−0.0136*	−0.0135*	−0.0199*
	(0.00436)	(0.00474)	(0.00587)	(0.00456)	(0.0065)
(3) log hp–owner$_{t+1}$	—	—	—	—	0.0435
					(0.0643)
(4) log hp–renter$_{t+1}$	—	—	—	—	−0.0167*
					(0.00723)
p-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.0000
p-value test (3) = (4)	—	—	—	—	0.3954
N	18,103	19,026	19,026	19,026	17,283
Women in married or cohabiting couples aged 40–54					
(1) log hp–owner	0.0408*	0.0411*	0.0416*	0.0401*	0.0463*
	(0.0263)	(0.0222)	(0.0315)	(0.0258)	(0.0273)
(2) log H–renter	0.00162	0.00163	0.00188	0.00142	0.00186
	(0.00323)	(0.00634)	(0.00592)	(0.00422)	(0.00342)
(3) log hp–owner$_{t+1}$	—	—	—	—	0.0463
					(0.0339)
(4) log H–renter$_{t+1}$	—	—	—	—	0.00745
					(0.00634)
p-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.0000
p-value test (3) = (4)	—	—	—	—	0.2845
N	17,891	18,775	18,775	18,775	17,238

Notes
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-labour income, homeowner dummy, local authority dummies, year dummies, renter–local authority and owner–local authority time trends.
*, **, *** indicate $p < 0.05$, $p < 0.01$, $p < 0.001$, respectively.
Cluster (local authority) standard errors in parentheses.
'p-value test' reports values from test for equivalence of coefficients in rows (1) and (2); similarly rows (3) and (4).

TABLE A6
ROBUSTNESS ESTIMATES FOR RELATIONSHIP BETWEEN LOG HOUSE PRICES AND RETIREMENT FOR OLDER MEN BY MARITAL STATUS: INDIVIDUAL FIXED EFFECTS ESTIMATES

	Excluding movers	Simulated prices	Initial owner	IV owner	Forward prices
Men in married or cohabiting couples aged > 54					
(1) log $\hat{H}$–owner	0.845***	0.952***	0.839***	0.899***	0.902***
	(0.0242)	(0.0234)	(0.0184)	(0.0105)	(0.0184)
(2) log H–renter	−0.00234	−0.00534	−0.00325	−0.0333	−0.0256
	(0.0173)	(0.00947)	(0.00556)	(0.0822)	(0.0237)
(3) log $\hat{H}$–owner$_{t+1}$	—	—	—	—	0.0645
					(0.0816)
(4) log H–renter$_{t+1}$	—	—	—	—	−0.00534
					(0.00845)
p-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.0000
p-value test (3) = (4)	—	—	—	—	0.2666
N	14,826	15,612	15,612	15,612	14,394
Single men aged > 54					
(1) log $\hat{H}$–owner	0.102**	0.933**	0.845**	0.110**	0.104**
	(0.0463)	(0.0386)	(0.0301)	(0.0274)	(0.0734)
(2) log H–renter	−0.00840	−0.00834	−0.00353	−0.00344	−0.00634
	(0.00863)	(0.00846)	(0.00863)	(0.00733)	(0.00791)
(3) log $\hat{H}$–owner$_{t+1}$	—	—	—	—	0.0237
					(0.0646)
(4) log H–renter$_{t+1}$	—	—	—	—	−0.0341
					(0.0637)
p-value test (1) = (2)	0.0000	0.0000	0.0000	0.0000	0.0000
p-value test (3) = (4)	—	—	—	—	0.3956
N	3015	3462	3462	3462	2839

Notes
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-labour income, homeowner dummy, local authority dummies, year dummies, renter–local authority and owner–local authority time trends.
*, **, *** indicate $p < 0.05$, $p < 0.01$, $p < 0.001$, respectively.
Cluster (local authority) standard errors in parentheses.
'p-value test' reports values from test for equivalence of coefficients in rows (1) and (2); similarly rows (3) and (4).

NOTES

1. Recent studies on the impact of house prices on household consumption and saving include Campbell and Cocco (2007), Disney *et al.* (2010), Attanasio *et al.* (2011), Carroll *et al.* (2011), Browning *et al.* (2013), Mian *et al.* (2013), Cooper (2013); on indebtedness, see Hurst and Stafford (2004), Disney and Gathergood (2011), and Mian and Sufi (2011).
2. 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%.
3. It is also an issue 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.

4. For further discussion of issues concerning the modelling of income and house price expectations, see Browning *et al*. (2013) and Disney *et al*. (2010).
5. This assumes that rents and house prices broadly move in the same direction at the local level, as theory would suggest (Gallin 2008). Using the panel structure of the data, we can calculate the year-on-year growth rates in self-reported house prices and self-reported rents at the local level. Over all waves of our sample period, the simple correlation in these first differences is 0.53. The correlation of first differences over time ranges between a minimum of 0.45 and a maximum of 0.64. Hence we do not observe particular periods of house price changes becoming detached from rents.
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.
7. From 2010 onwards, the BHPS survey sample was incorporated into a new survey, 'Understanding Society'. This resulted in many changes to the survey, including changes to many of the core variables in our analysis. Hence we do not use the Understanding Society sample in this analysis.
8. Individuals who report that they are suffering short-term sickness leave from work or are on vacation from work are classified by their regular labour market status (employed or self-employed).
9. 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?'
10. Local authority 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 are derived from confidential workplace surveys in which employers report wages paid to employees.
11. We model labour supply decisions at the individual level, controlling for the labour market status of an individual's spouse or partner through a set of controls for labour market states. An alternative approach beyond the scope of our analysis would be to jointly model labour supply decisions at the household level.
12. We have estimated models for each of these specifications for each of the subsamples presented in Table 2 (and in the remainder of the paper for the labour 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 show only robustness estimates for subsamples where the main specification returned results of interest. The replication files include robustness estimates (and region-level cluster standard errors estimates that also do not change our main results) for all subsamples.
13. These IV models involve self-reported homeowner home values instrumented by local authority level house prices, entering alongside a local authority level house price interacted for renters only (we do not have self-reported prices for renters in the data). An alternative model is to use the first-stage model estimated using data on homeowners only and predict onto the renter sample (as well as the homeowner sample). When we use this approach we again find very similar results. These results are available from the authors on request.
14. An alternative viewpoint is that older female workers, who typically work part-time, have higher opportunity costs of working. Older female workers are more likely to be a valuable source of childcare for grandchildren, and are less likely to be attached to forms of employment that carry seniority wages due to their lower lifetime attachment to the labour force.

REFERENCES

ATALAY, K., BARRETT, G. and EDWARDS, F. (2016). Housing wealth effects on labour supply: evidence from Australia. Mimeo, University of Sydney.

ATTANASIO, O., LEICESTER, A. and WAKEFIELD, M. (2011). The coincident cycles of house prices and consumption in the UK. *Journal of the European Economic Association*, **9**, 399–435.

BLUNDELL, R., BREWER, M. and FRANCESCONI, M. (2008). Job changes and hours changes: understanding the path of labour supply adjustment. *Journal of Labour Economics*, **26**(3), 421–53.

———, CRAWFORD, C. and JIN, W. (2014). What can wages and employment tell us about the UK's productivity puzzle? *Economic Journal*, **124**, 377–407.

BROWN, J. R., COILE, C. C. and WEISBENNER, S. J. (2010). The effect of inheritance receipt on retirement. *Review of Economics and Statistics*, **92**, 425–34.

BROWNING, M., GØRTZ, M. and LETH-PETERSEN, S. (2013). Housing wealth and consumption: a micro panel study. *Economic Journal*, **123**, 401–28.

CAMPBELL, J. and COCCO, J. F. (2007). How do house prices affect consumption? *Journal of Monetary Economics*, **54**, 591–621.

CARROLL, C., OTSUKA, M. and SLACALEK, J. (2011). How large are housing and financial wealth effects? A new approach. *Journal of Money, Credit and Banking*, **43**, 55–79.

CESARINI, D., LINDQVIST, E., NOTOWIDIGDO, M. J. and OSTLING, R. (2013). The effect of wealth on household labor supply: evidence from Swedish lotteries. NBER Summer Institute Labour Studies Meeting, July 2013.

COILE, C. C. and LEVINE, P. B. (2011). The market crash and mass lay-offs: how the current economic crisis may affect retirement. *B. E. Journal of Economic Analysis and Policy*, **11**, 1–42.

COOPER, D. (2013). House price fluctuations: the role of housing wealth as borrowing collateral. *Review of Economics and Statistics*, **95**, 1183–97.

DALY, M., KWOK, J. and HOBIJN, B. (2009). Labor supply response to changes in wealth and credit. FRBSF Economic Letter 2009-05, San Francisco Fed.

DAVIDOFF, T. (2010). Home equity commitment and long-term care insurance demand. *Journal of Public Economics*, **94**(1), 44–9.

DETTLING, L. J. and KEARNEY, M. S. (2014). House prices and birth rates: the impact of the real estate market on the decision to have a baby. *Journal of Public Economics*, **110**, 82–100.

DISNEY, R. and GATHERGOOD, J. (2009). Housing wealth, liquidity constraints and self-employment. *Labour Economics*, **16**, 79–88.

——— and ——— (2011). House price growth, collateral constraints and the accumulation of homeowner debt in the US. *B. E. Journal of Macroeconomics*, **11**, 1.

———, ——— and HENLEY, A. (2010). House price shocks, negative equity and household consumption in the United Kingdom. *Journal of the European Economic Association*, **8**, 1179–207.

———, RATCLIFFE, A. and SMITH, S. (2015). Booms, busts and retirement timing. *Economica*, **82**, 399–419.

FARNHAM, M., SCHMIDT, L. and SEVAK, P. (2011). House prices and marital stability. *American Economic Review*, **101**(3), 615–19.

FRENCH, E. and BENSON, D. (2011). How do sudden large losses in wealth affect labor force participation? Chicago Fed Letter no. 282, January.

GALLIN, J. (2008). The long run relationship between house prices and rents. *Real Estate Economics*, **36**, 635–58.

HILBER, C. and VERMEULEN, W. (2016). The impact of supply constraints on house prices in England. *Economic Journal*, **126**, 358–405.

HURST, E. and LUSARDI, A. (2004). Liquidity constraints, household wealth and entrepreneurship. *Journal of Political Economy*, **112**, 319–47.

——— and STAFFORD, F. (2004). Home is where the equity is: mortgage refinancing and household consumption. *Journal of Money, Credit and Banking*, **36**, 987–1014.

IMBENS, G. W., RUBIN, D. B. and SACERDOTE, B. I. (2001). Estimating the effect of unearned income on labor earnings, savings and consumption: evidence from a survey of lottery players. *American Economic Review*, **91**(4), 778–94.

JOULFAIAN, D. and WILHELM, M. (2001). Inheritance and labor supply. *Journal of Human Resources*, **29**(4), 1205–34.

LOVENHEIM, F. (2011). The effect of liquid housing wealth on college enrolment. *Journal of Labour Economics*, **29**, 741–71.

——— and MUMFORD, K. (2013). Do family wealth shocks affect fertility choices: evidence from the housing market. *Review of Economics and Statistics*, **95**, 464–75.

——— and REYNOLDS, C. L. (2013). The effect of short-term credit constraints on college choice: evidence from the housing boom. *Journal of Human Resources*, **48**, 1–35.

MIAN, A. and SUFI, A. (2011). House prices, home equity-based borrowing, and the U.S. household leverage crisis. *American Economic Review*, **101**, 2132–56.

———, RAO, K. and SUFI, A. (2013). Household balance sheets, consumption, and the economic slump. *Quarterly Journal of Economics*, **128**(4), 1687–726.

MILOSCH, J. (2016). House price shocks and labour supply changes. Mimeo, California State University, Sacramento.