Housing over Time and over the Life Cycle: A Structural Estimation*

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ABSTRACT

We estimate a structural model of optimal life-cycle housing and nonhousing consumption with labor income and house price uncertainties as well as housing adjustment cost using data from the Panel Study of Income Dynamics (1984-2005). Our focus is on the substitutability between housing and nonhousing consumption, a key driver of the results in many housing policy studies. The benchmark analysis suggests an intratemporal elasticity of substitution between housing and nonhousing consumption of 0.487, a result largely driven by the moments conditional on state house prices and the moments in the latter half of the sample period. The estimate is robust to different assumptions of housing adjustment cost, but sensitive to the choice of sample period and the degree of aggregation of data moments. We then conduct experiments in which we let house prices and household income take values as those observed between 2006 and 2011. We show that the responses depend importantly on the elasticity of substitution between housing and nonhousing consumption. In particular, compared with the benchmark, the impact of the shocks on homeownership rates is reduced, but the impact on nonhousing consumption is magnified when housing service and nonhousing consumption are highly substitutable or when the house selling cost is sizable.

Keywords: Life cycle, housing adjustment costs, intratemporal substitution

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1. Introduction

The U.S. housing market has experienced dramatic price movements in recent years. These movements, accompanied by substantial changes in household indebtedness, have drawn the attention of policymakers and academicians. Calibrated housing models are now increasingly deployed in studying the effects of housing on consumption and savings, stock market participation and asset allocation, asset pricing, and the transmission channel and effectiveness of monetary policies.[1] Despite this growing interest in housing models in an intertemporal setting, econometric research aimed at identifying the relevant housing preference parameters has been lacking. As a consequence, theoretical models are often calibrated with little empirical guidance regarding the key model input parameters.[2]

One such key parameter is the intratemporal elasticity of substitution between housing and nonhousing consumption, which governs to a large extent the impact of changes in house prices and income and, hence, policies that affect these changes, on household consumption and welfare. Among the few econometric studies of housing preferences, there has been little consensus on the magnitude of this coefficient. Studies based on macro-level aggregate consumption or asset price data frequently suggest a value larger than one — implying that households reduce expenditure share on housing when house prices move up relative to prices of nonhousing consumption (Davis and Martin, 2009, and Piazzesi, Schneider, and Tuzel, 2007). These studies have typically assumed the existence of a representative agent and abstracted from market incompleteness and informational frictions, despite strong evidence of household heterogeneity and housing adjustment cost documented in the literature (Carroll and Dunn, 1997 and Attanasio, 2000).

[1]See Campbell and Cocco (2007), Fernandez-Villaverde and Krueger (2011), Li and Yao (2007), Stokey (2009), Kiyotaki, Michaelides, and Nikolov (2011), and Yang (2009) for consumption and savings; Cocco (2005), Flavin and Yamashita (2002), and Yao and Zhang (2005) for stock market participation and asset allocation; Piazzesi, Schneider, and Tuzel (2007), Siegel (2008), Lustig and Van Nieuwerburgh (2005), and Flavin and Nakagawa (2008) for asset pricing; and Iacoviello (2005) for monetary policy effects.

[2]Many theoretical studies using numerical calibrations adopt the Cobb-Douglas utility function for its simplicity and often abstract from housing adjustment costs. A direct implication of the Cobb-Douglas utility function is that housing expenditure share is constant over time and across regions. The aggregate housing expenditure as a share of total expenditure, however, has fluctuated over time with a sharp rise, leading to the Great Depression followed by a prolonged decline. The share started to recover at the end of World War II and has since fluctuated more mildly than before (Figure A1 in the online appendix). At the micro level, the Consumer Expenditure Survey also indicates that expenditure shares at the metropolitan statistical area (MSA) level have fluctuated over time with many experiencing upward movement until the recent housing crisis (Figure A2). See Stokey (2009) and Kahn (2008) for additional evidence against Cobb-Douglas utility specification.

In contrast, investigations using household-level data recover much lower values for the elasticity parameter, often in the range of 0.15 and 0.60 (For example, Flavin and Nakagawa, 2008; Hanushek and Quigley, 1980; Siegel, 2008, and Stokey, 2009). These studies, however, suffer from selection bias because households endogenously make decisions on both house tenure (renting vs. owning and moving vs. staying) and the quantity of housing service flows. Furthermore, the identification in many of the studies is predicated on households having unlimited access to credit, which contradicts the practice in reality. The lack of robustness to market friction and incompleteness, therefore, complicates the interpretation of the empirical estimates in these studies.

This paper structurally estimates a stochastic life-cycle model of consumption, savings, and housing choices and jointly identifies the intratemporal as well as intertemporal preference parameters by matching average wealth and housing profiles generated by the model with the profiles from micro data. We postulate constant elasticity of substitution (CES) preferences over housing and nonhousing consumption and allow households to make housing decisions along both the extensive margin of homeownership and the intensive margin of housing service flows and house value. The model explicitly admits a housing transaction cost and a collateral borrowing constraint, as well as labor income and house price uncertainties. Our model, therefore, builds on a growing literature examining household house tenure and housing consumption choices in a life-cycle framework (Ortalo-Magne and Rady, 2005; Fernandez-Villaverde and Krueger, 2011; Gervais, 2002, Campbell and Cocco, 2007; Chambers, Garriga, and Schlagenhauf, 2009; Yao and Zhang, 2005; Li and Yao, 2007; and Yang, 2009).

Our estimation of the structural parameters is achieved through the method of simulated moments (MSM). We first construct the average wealth, homeownership rates, moving rates, house value-to-income ratio, and rent-to-income ratio profiles from the Panel Study of Income Dynamics (PSID) data set across three age groups between 1984 and 2005. For homeownership rates, house values, and rent values, we further group households according to the levels of house prices in their state of residence and construct additional moments. We then numerically solve the model for optimal household behavior and simulate the model to generate paths of life-cycle housing and wealth profiles in the same manner as the data moments to eliminate potential bias caused by cohort and time effects as well as selection bias. By minimizing the weighted difference between the simulated model profiles and their empirical counterparts, we identify the parameters of our structural model.

Our simulated wealth and housing profiles match important features in the data over the sample period. Our estimation also reveals that after explicitly accounting for housing adjustment cost, the intratemporal elasticity of substitution between housing services and nondurables is around 0.487, a value that is on the high end of micro empirical estimates but smaller than those assumed in macro models. The estimate is not very sensitive to different parameterizations of housing transaction cost. When we double our exogenously imposed housing transaction cost from 8 percent of the house value to 16 percent, the parameter comes down slightly to 0.485. The low elasticity estimate, however, is largely driven by moments conditional on state house prices and moments in the latter half of the sample period. When we focus on a sample period with little house price movement (1984 to 1993) and drop moments conditional on state house prices, the elasticity estimate more than tripled, increasing to 1.690.

To illustrate the importance of the different estimations of the elasticity parameter, we conduct several policy experiments. Specifically, we investigate how households respond to changes in house prices and income as those observed between 2005 and 2011. We find that large and persistent house price depreciation coupled with declines in income leads to significant decreases in both homeownership rates and nonhousing consumption in the two economies in which the elasticity of substitution between housing and nonhousing consumption is low. When the elasticity of substitution between housing and nonhousing consumption is high, homeownership rates do not move much, but nonhousing consumption declines substantially. Finally, a lower expected house price growth aggravates the negative effect of current house price declines on the homeownership rates, house or rent value, and nonhousing consumption. To conclude, our paper reconciles to some extent the differences in theoretical as well as empirical estimations of the intratemporal elasticity of substitution between housing and nonhousing consumption. It nevertheless remains a challenge as to what is the relevant parameter to use for conducting policy experiments given their different implications.

To the best of our knowledge, our paper represents one of the first structural estimations of housing preference parameters that are consistent with both time series and cross-sectional evidence on households' housing consumption and savings decisions. Estimating a rich life-cycle model allows us to address potential biases directly by replicating them in the simulation. The recent paper by Bajari, Chan, Krueger, and Miller (2013) is the closest in spirit to our paper. There are, however, significant differences. The first difference is methodology. Bajari et al. adopt a two-step approach. In the first step, reduced form decision rules are estimated together with the law of motion for state variables. The structural parameters are then estimated in the second step using simulation based on the reduced form decision rule in the first step. In contrast, we solve the decision rules endogenously instead of imposing reduced forms. Additionally, we explicitly model and estimate households' tenure decisions. The second and more important difference between the two papers is different target moments, which the models are estimated to match. There is much less variation in aggregate house prices during Bajari et al's sample period (1983 to 1993) than during ours (1984 to 2005). Furthermore, we make use of cross-sectional heterogeneity by studying areas with high, medium, and low house prices separately. These additional heterogeneities turn out to be crucial in explaining the differences in our estimates. When we re-estimate our model using similar life-cycle aggregate moments as in Bajari et al. (2013), not surprisingly, we obtain a much larger elasticity of substitution between housing and nonhousing consumption.

The rest of the paper proceeds as follows. In Section 2, we present a life-cycle model of housing choices with an adjustment cost. In Section 3, we lay out our estimation strategy and describe the data sources. Section 4 discusses our main findings and implications. Section 5 presents alternative estimations. We further explore model implications in Section 6. We conclude and point to future extensions in Section 7.

2. The Model Economy

Our modeling strategy extends that of Yao and Zhang (2005) and Li and Yao (2007) by admitting a flexible specification of elasticity of substitution between housing and other consumption.

2.1. Demographics and Preferences

A household lives for a maximum of T periods. Let λ_j denote the probability that the household is alive at time j conditional on being alive at time $j-1$, $j = 0, ..., T$. We set $\lambda_0 = 1$, $\lambda_T = 0$, and $0 < \lambda_j < 1$ for all $0 < j < T$.

The household derives utility from consuming housing services H_t and nonhousing goods C_t, as well as from bequeathing wealth Q_t. The utility function from consumption demonstrates a CES between the two goods, modified to incorporate a demographic effect:

$$
\begin{aligned}
U(C_t, H_t; N_t) &= N_t \frac{[(1-\omega)(\frac{C_t}{N_t})^{1-\frac{1}{\zeta}} + \omega(\frac{H_t}{N_t})^{1-\frac{1}{\zeta}}]^{\frac{1-\gamma}{1-\frac{1}{\zeta}}}}{1-\gamma} \\
&= N_t^{\gamma} \frac{[(1-\omega)C_t^{1-\frac{1}{\zeta}} + \omega H_t^{1-\frac{1}{\zeta}}]^{\frac{1-\gamma}{1-\frac{1}{\zeta}}}}{1-\gamma},
\end{aligned}
\tag{1}
$$

where N_t denotes the exogenously given effective family size to capture the economies of scale in household consumption. The parameter ω controls the expenditure share on housing services, and ζ governs the degree of intratemporal substitutability between housing and nondurable consumption goods. The bequest function $B(Q_t)$ will be discussed later.

The household retires exogenously at $t = J$ $(0 < J < T)$. Before retirement, the household receives labor income Y_t,

$$Y_t = P_t^Y \varepsilon_t, \tag{2}$$

where ε_t is the transitory shock to Y_t and P_t^Y is the permanent labor income at time t that follows

$$P_t^Y = \exp\{f(t)\} P_{t-1}^Y \nu_t^s \nu_t^i. \tag{3}$$

The deterministic component $f(t)$ is a function of age. The term ν_t^s represents the aggregate shock to permanent labor income that is shared by all agents in the economy, while ν_t^i is the idiosyncratic shock to permanent labor income. We assume that $\{\ln \varepsilon_t, \ln \nu_t^s, \ln \nu_t^i\}$ are independently and identically normally distributed with mean $\{-0.5\sigma_\varepsilon^2, -0.5\sigma_{\nu^s}^2 + \mu_\nu^s, -0.5\sigma_{\nu^i}^2\}$ and variance $\{\sigma_\varepsilon^2, \sigma_{\nu^s}^2, \sigma_{\nu^i}^2\}$, respectively.[3] After retirement, a household's permanent labor income does not change any more, and the household receives a constant fraction θ of its permanent labor income before retirement,

$$P_t^Y = P_J^Y, \quad \text{for } t = J, ..., T, \tag{4}$$

$$Y_t = \theta P_t^Y, \quad \text{for } t = J, ..., T. \tag{5}$$

[3] The labor income process largely follows that of Carroll and Samwick (1997), which is also adopted in Cocco, Gomes, and Maenhout (2005) and Gomes and Michaelides (2005), with the exception that we allow part of the permanent income to be driven by an economywide aggregate shock.

2.2. Housing and Financial Markets

A household can either own a house or rent housing services by paying a fraction α ($0 < \alpha < 1$) of the market value of the rental house. A renter's housing tenure is denoted as $D_t^o = 0$, and a homeowner's housing tenure is $D_t^o = 1$.

We assume that house price P_t^H is exogenously given by the following stochastic process:

$$P_t^H = P_{t-1}^H \xi_t^s \xi_t^i, \tag{6}$$

where ξ_t^s represents the aggregate shock to the house price shared by all households and ξ_t^i is the idiosyncratic shock unique to the household. ξ_t^s and ξ_t^i follow an i.i.d. normal process with mean $\{-0.5\sigma_{H^s}^2 + \mu_H^s, -0.5\sigma_{H^i}^2\}$ and variance $\{\sigma_{H^s}^2, \sigma_{H^i}^2\}$. Additionally, we assume that the aggregate house price shock ξ_t^s is correlated with the aggregate income shock ν_t^s with a correlation coefficient of ρ.[4]

A household can finance home purchases with a mortgage. The mortgage balance denoted by M_t needs to satisfy the following collateral constraint at all times:

$$0 \leq M_t \leq (1-\delta) P_t^H H_t, \tag{7}$$

where $0 \leq \delta \leq 1$, and $P_t^H H_t$ denotes the value of the house at time t.[5]

We assume that households can save in liquid assets, which earn a constant risk-free rate r, but they cannot hold noncollateralized debt. We further assume that the mortgage borrowing rate r_m is higher than the risk-free rate r. It follows immediately then that households will always want to repay their mortgages and will not hold mortgages and liquid assets simultaneously. We let S_t denote the net financial wealth. Finally, in order to keep the house

[4]Flavin and Yamashita (2002), Campbell and Cocco (2007), and Yao and Zhang (2005) also make similar assumptions about house price dynamics. The innovation here is that we split the permanent price shock into an aggregate component and an idiosyncratic component and allow the aggregate component to be correlated with the aggregate shock that affects individual permanent income. Campbell and Cocco (2007) assume the correlation to be 1.

[5]By applying collateral constraints to both newly initiated mortgages and ongoing loans, we effectively rule out default. Default on mortgages was, until recently, relatively rare in reality. According to the Mortgage Bankers Association, the annual 90+ days default rate for a prime fixed-rate mortgage loan was around 2 percent prior to 2007. When house prices experience a sequence of adverse shocks (as during the recent crisis) that amounts to a loss of 20 percent or more of house equity, this assumption would imply that households will be forced to sell more frequently in the model than in reality.

quality constant, a homeowner is required to spend a fraction ψ ($0 \leq \psi \leq 1$) of the house value on repair and maintenance.

At the beginning of each period, each household receives a moving shock, D_t^m, that takes a value of 1 if the household has to move for reasons that are exogenous to our model, and 0 otherwise. The moving shock does not affect a renter's decision since he does not incur any moving costs. When a homeowner receives a moving shock ($D_t^m = 1$), he is forced to sell the house.[6] Selling a house incurs a transaction cost that is a fraction ϕ of the market value of the existing house. The selling decision, D_t^s, is 1 if the homeowner sells and 0 otherwise. A homeowner who does not have to move for exogenous reasons can choose to sell the house voluntarily. Following a home sale — for either exogenous or endogenous reasons — a homeowner faces the same decisions as a renter coming into period t and is free to buy or rent for the following period.

2.3. Wealth Accumulation and Budget Constraints

We denote the household's spendable resources upon home sale by Q_t.[7] It follows that

$$Q_t = \max\{S_{t-1}\left[1 + r(S_{t-1})\right] + Y_t + P_{t-1}^H D_{t-1}^o H_{t-1} \xi_t^s \xi_t^i (1-\phi), \eta P_t^Y\}. \tag{8}$$

As discussed earlier, if S_{t-1} is nonnegative, the household has paid off its mortgage and, hence, $r(S_{t-1}) = r$; if S_{t-1} is negative, it represents the mortgage the household holds on the house. As a result, $r(S_{t-1}) = r_m$. The last term ηP_t^Y denotes government transfers. Following Hubbard, Skinner, and Zeldes (1994, 1995) and De Nardi, French, and Jones (2010), we assume that government transfers provide a wealth floor that is proportional to the household's permanent labor income. The intertemporal budget constraint, therefore, can be written as:

[6] We assume that house prices in the old and new locations are the same. This is obviously a simplifying assumption as in reality households can move to states that have very different house price movement. In Figure A3 in the online appendix, using data from the Current Population Survey, we show that the annual rate at which households move across states, however, has been small, averaging between 2 and 3 percent from 1984 to 2005.

[7] Under this definition, conditional on selling the house, a homeowner's problem is identical to that of a renter with same age t, permanent income P_t^Y, housing price $P_t H$, and liquidated wealth Q_t.

$$Q_t = \begin{cases} C_t + S_t + \alpha P_t^H H_t, & \text{rent at } t-1 \text{ and } t\ (D_{t-1}^o = D_t^o = 0); \\ C_t + S_t + \alpha P_t^H H_t, & \text{own at } t-1, \text{ sell and rent at } t\ (D_{t-1}^o = D_t^s = 1, D_t^o = 0); \\ C_t + S_t + (1+\psi) P_t^H H_t, & \text{rent at } t-1, \text{ own at } t\ (D_{t-1}^o = 0, D_t^o = 1); \\ C_t + S_t + (1+\psi) P_t^H H_t, & \text{own at } t-1, \text{ sell and own at } t\ (D_{t-1}^o = D_t^s = D_t^o = 1); \\ C_t + S_t + (1+\psi-\phi) P_t^H H_t, & \text{own the same H at } t-1 \text{ and } t\ (D_{t-1}^o = D_t^o = 1, D_t^s = 0). \end{cases} \tag{9}$$

The third term on the right-hand side of the budget constraint represents housing expenditure when households have various different house tenure status. Note that for the last group of households who stayed in their houses, we need to subtract from their expenditure the housing selling cost, which was subtracted from wealth in hand on the left-hand side defined in equation 8.

2.4. Bequest Function

We assume that upon death, a household distributes its spendable resources Q_t equally among L beneficiaries to finance their nonhousing good and rented housing services consumption for one period. Parameter L thus determines the strength of bequest motives. Under the assumption of CES utility, the beneficiary's expenditure on nonhousing good and housing service consumption is the following function of house price:

$$\frac{C_t}{C_t + \alpha P_t^H H_t} = \frac{(1-\omega)^\zeta}{(1-\omega)^\zeta + \omega^\zeta (\alpha P_t^H)^{1-\zeta}}. \tag{10}$$

Therefore, the bequest function is defined by

$$\begin{aligned} B(Q_t) &= L \frac{\left[(1-\omega)\left(\frac{Q_t}{L}\frac{(1-\omega)^\zeta}{(1-\omega)^\zeta + \omega^\zeta(\alpha P_t^H)^{1-\zeta}}\right)^{1-\frac{1}{\zeta}} + \omega\left(\frac{Q_t}{L}\frac{\omega^\zeta(\alpha P_t^H)^{-\zeta}}{(1-\omega)^\zeta + \omega^\zeta(\alpha P_t^H)^{1-\zeta}}\right)^{1-\frac{1}{\zeta}}\right]^{\frac{1-\gamma}{1-\frac{1}{\zeta}}}}{1-\gamma} \\ &= L^\gamma Q_t^{1-\gamma} \frac{\left[(1-\omega)\left(\frac{(1-\omega)^\zeta}{(1-\omega)^\zeta + \omega^\zeta(\alpha P_t^H)^{1-\zeta}}\right)^{1-\frac{1}{\zeta}} + \omega\left(\frac{\omega^\zeta(\alpha P_t^H)^{-\zeta}}{(1-\omega)^\zeta + \omega^\zeta(\alpha P_t^H)^{1-\zeta}}\right)^{1-\frac{1}{\zeta}}\right]^{\frac{1-\gamma}{1-\frac{1}{\zeta}}}}{1-\gamma}. \end{aligned} \tag{11}$$

2.5. The Optimization Problem

We denote $X_t = \{D^o_{t-1}, Q_t, P^Y_t, P^H_t, H_{t-1}\}$ the vector of state variables. The household at state X_t solves the following optimization problem:

$$V_t(X_t) = \max_{A_t}\{\lambda_t \left[U(C_t, H_t; N_t) + \beta E_t[V_{t+1}(X_{t+1})]\right] + (1-\lambda_t)B(Q_t)\}, \tag{12}$$

subject to the mortgage collateral borrowing constraint (equation 7), wealth processes (equation 8), and the intertemporal budget constraints (equation 9), where β is the time discount factor, and $A_t = \{C_t, H_t, S_t, D^o_t, D^s_t\}$ is the vector of choice variables.

An analytical solution for the optimization problem does not exist. We thus derive numerical solutions through backward induction of value function. The online Appendix A provides details of our numerical method.

2.6. Characterization of Individual Housing and Consumption Behavior

Qualitatively, at a given household age, the effects of wealth-to-income ratio and house value-to-income ratio on the household's optimal decision rules are similar to those reported in Yao and Zhang (2005). A renter's house tenure decision is largely determined by the renter's wealth-to-income ratio as depicted in Figure A4 in the online appendix. In this figure, we plot a renter's house tenure decision as a function of the renter's wealth-to-permanent labor income ratio for a given median house price. Two observations emerge. First, the more wealth a renter has relative to income, the more likely the renter will buy, as more wealth on hand enables the renter to afford the down payment for a house of desired value. The wealth-to-income ratio that triggers homeownership increases late in life, reflecting the short expected duration of an older household.[8] Once a household becomes a homeowner, the household will stay in the house as long as its house value-to-income ratio is not too far from the optimal level it would have chosen as a renter in order to avoid incurring transaction costs.

[8]Note that the wealth-to-income cutoff ratio for homeownership is not tilted upward for the very young as in Yao and Zhang (2005). This is because we target homeownership rates for married stable households between the ages 25 and 34, while Yao and Zhang (2005) target much lower ownership rates for the very young, and a much higher cutoff ratio is needed to generate the low rates.

A renter's consumption and savings functions are very similar to those documented in the precautionary savings literature with liquidity constraints. At low wealth levels, a renter continues to rent and spends all his wealth on nonhousing goods and rent payment. At relatively higher wealth levels, a renter starts saving for intertemporal consumption smoothing and for a housing down payment. For a homeowner who stays in his existing house, the value of the house also affects nonhousing consumption, reflecting the effect of substitution between the two goods.

3. Data and Estimation Procedure

We implement a two-stage MSM to identify parameters in our theoretical model.[9] The mechanics of our MSM approach follow the standard literature. In the first stage, we estimate or calibrate parameters that can be cleanly identified without the explicit use of our model. In the second stage, we take as given the parameters obtained in the first stage and estimate the rest of the model parameters by minimizing the distance between the simulated moments derived from the optimal household decision profiles and those observed in the data. We provide detailed discussions of the first- and second-stage estimations, after describing our data sources.

3.1. Data Sources

The main data we use in this study are from the PSID managed by the University of Michigan. The PSID is a longitudinal survey that followed a nationally representative random sample of families and their extensions since 1968. The survey details economic and demographic information for a sample of households annually from 1968 to 1997 and biannually after 1997. From 1984 through 1999, a wealth supplement to the PSID surveyed the assets and liabilities of each household at five-year intervals. The supplemental survey becomes biennial after 1999, coinciding with the frequency of themain survey.

[9]This methodology was first introduced by Pakes and Pollard (1989) and Duffie and Singleton (1993) to estimate structural economic models without closed-form solutions. Since then, the MSM has been successfully applied to estimations of preference parameters in Gourinchas and Parker (2002), and Cagetti (2003), labor supply decisions by French (2005), and medical expenses and the savings of elderly singles by De Nardi, French, and Jones (2010), among many others.

For households to be included in our data sample, they have to be present in the 1984 survey but not in the 1968 subsample of low-income families. Observations were further deleted for the following reasons:

- The head of the household is female, is a farmer or rancher, or is unmarried in any wave of the survey.
- The age of the household head is younger than 25 or older than 54 in the 1984 survey, or the head of the household does not have a valid age variable.
- Households obtained housing as a gift, or live in housing that is paid for by someone outside of the family unit or is that is owned by relatives; households live in a public housing project owned by a local housing authority or public agency; or households neither own nor rent.
- The real household labor income is less than $10,000 or more than $1 million.
- Information on households' state of residence, net worth, income, or house value for homeowners is missing.

The final sample consists of 17,392 observations for 1,069 households.[10] We use this sample to estimate life-cycle income profiles and to compute sample moments. We supplement the PSID data with information from the American Housing Survey (AHS) and the Office of the Federal Housing Financial Agency (FHFA) for house price information, the PSID for mobility, and the National Center for Health Statistics for life expectancy.

3.2. First-Stage Estimation and Calibration

3.2.1. Life-cycle Income Profiles

The income in our model corresponds to after-tax nonfinancial income empirically. We calibrate the stochastic income process (equations 2–5) as follows. We first compute before-tax income as the total reported wages and salaries, Social Security income, unemployment com-

[10]Table A1 in the online appendix provides summary statistics of the sample. Household heads average 46 years of age with fewer than two children. Almost all of household heads have graduated from high school and over a quarter have college degrees. The majority of them are white and own their houses. Their annual family income, house value, and net worth are comparable to those from the Survey of Consumer Finances for the same married age groups.

pensation, workers compensation, supplemental Social Security, other welfare, child support, and transfers from relatives from both the head of household and his spouse.[11]

We then subtract from the households' pretax income federal and state income tax liabilities as estimated by the National Bureau of Economic Research (NBER)'s TAXSIM program (Feenberg and Coutts, 1993). This program calculates taxes under the U.S. federal and state income tax laws from individual data, including marital status, wage and salary of household head and his or her spouse, and number of dependents. The after-tax income is further deflated using a nonshelter consumer price index (CPI-NS) provided by the Bureau of Labor Statistics, with the year 2004 - 2005 normalized to 100. We refer to this deflated disposable income as household labor income in this paper.

Finally, we apply an approach similar to the one used in Cocco, Gomes, and Maenhout (2005) to estimate coefficients for a sixth-order polynomial function of age and retirement income replacement ratio, as well as standard deviation for permanent and transitory shocks to income. Using individual income and the methodology outlined in the online Appendix B, we estimate the permanent income shock to have a standard deviation of 0.10 and the transitory income shock to have a standard deviation (σ_ε) of 0.22. The income replacement ratio after retirement θ is 0.70. Using aggregate labor income, we estimate that the aggregate income shock has a mean growth rate (μ_ν^s) of 1.2 percent and a standard deviation (σ_ν^s) of 0.022. The standard deviation of the permanent income shock (σ_ν^i) is then backed out to be approximately 0.10 ($\sqrt{0.1^2 - 0.022^2}$).

3.2.2. Mortality, Mobility, and Household Composition

The conditional survival rates ($\{\lambda_j\}_{j=0}^{j=T}$) are taken from the 1998 life tables of the National Center for Health Statistics (Anderson, 2001). The exogenous moving rates are obtained by fitting a fifth-order polynomial of age to the PSID married households moving across the county. The life-cycle profile of family equivalent size for all married couples in the PSID is computed following Scholz et al. (2006). The calibrated life-cycle income, mortality, exogenous mobility, and family size profiles are presented in Figure A5 in the online appendix.

[11] Recall that we use only married households from the PSID in our sample.

3.2.3. The House Price Process

The mean rate of return to aggregate housing shock (μ_H^s) is set to 1 percent and the standard deviation (σ_H^s) is set to 0.015, which are roughly the mean and standard deviation of the annual nationwide real house price growth rates between 1975 and 2005. The overall standard deviation of housing returns across all states is estimated to be 0.10, similar to estimates in Campbell and Cocco (2003) and Flavin and Yamashita (2002). The standard deviation of housing returns in the model (σ_H^i) is then about 0.10 ($\sqrt{0.1^2 - 0.015^2}$). Finally, we set the correlation between the permanent shock to housing returns and the permanent shock to labor income ρ to be 0.30, a number we calculated using state real per capita disposable personal income and real house price growth rates in all 50 states between 1975 and 2005. The online Appendix C provides details on the construction of the state-level house price index over time.

3.2.4. Other Parameters

Other parameters in the first stage are largely chosen according to the literature. The decision frequency is annual. Households enter the economy at age 25 and live to a maximum age of 100 ($T = 75$). The mandatory retirement age is 65 ($J = 40$).

The annual real interest rate is set at 2.7 percent, approximately the average annualized post-World War II real return available on Treasury bills. The mortgage collateral constraint is set at 80 percent.[12] The mortgage interest rate is set at 6.7 percent, implying a spread between the mortgage rate and the risk-free interest rate of 4 percent.[13] The wealth floor η is picked at a low 0.06 of permanent labor income. This number is within the range of those used in the literature (for example, De Nardi, French, and Jones, 2010) and rarely binds in our simulation. Finally, we set housing selling cost at 8 percent. Table 1 summarizes parameters from our first-stage calibrations.

[12] Using the 1995 American Housing Survey, Chambers, Garriga, and Schlagenhauf (2009) calculate that the down payment fraction for first-time home purchases is 0.1979, while the fraction for households who previously owned a home is 0.2462.

[13] The spread between the 30-year fixed mortgage interest rate (plus points and fees) from the Federal Housing Administration and the 10-year constant maturity Treasury bond yield is about 3.2 percent between 1985 and 2005, and between the one-year adjustable mortgage interest rate and one-year Treasury bill yield is about 4.6 percent.

3.3. Second-Stage Estimation

In this subsection, we describe our choices of moment conditions and how they help identify the structure parameters of our model. One major advantage of structural estimation of a rich life-cycle model is that it allows us to explicitly address potential biases by replicating them in the simulation. For example, we account for the endogeneity of homeownership status, market frictions, and incompleteness (borrowing constraints and house adjustment costs) by incorporating these features in our theoretical model. To mitigate potential biases caused by cohort and time effects, we group the households in our simulation by age and calender year, and we subject the households to the same aggregate house price shocks as in the data.

3.3.1. Moment Conditions

We estimate the following seven structural parameters in the second stage estimation: β – subjective time discount factor, γ – curvature measure for the utility function, L – bequest strength measure, ω – weight on housing, ζ – elasticity of intratemporal substitution, ψ – house maintenance costs, and α – rental rate.[14]

To identify our structural parameters, we choose to match the average net worth, mobility rate, homeownership rate, rent-to-income ratio, and house value-to-income ratio profiles for three age cohorts annually between 1985 and 1997 and biennially between 1999 and 2005.[15] The three age cohorts are constructed according to birth year. At the beginning of our sample year 1984, the three cohorts' age ranges are 25 - 34, 35 - 44, and 45 - 54, respectively.

In addition, to exploit the cross-sectional heterogeneity in house prices, for each age cohort–calendar year cell, we also match the average homeownership rate, rents, and house value profile for households residing in the most and least expensive states.[16] We thus have 11

[14]For α, the estimation is performed in terms of rental premium, i.e., $\alpha + \mu_H^s - r - \psi$.

[15]We drop the year 1984 in the moment matching, since we initiate our simulation by randomly drawing households from the 1984 PSID data.

[16]We define the most expensive states as the 18 states with the highest house price level in 1995, the middle year in our sample, and the least expensive states as the 19 states with the lowest house price level in 1995. According to this definition, we have roughly equal numbers of households residing in the most expensive, least expensive, and medium price range states in 1984. The choice of 1995 is inconsequential, since the ranking of house prices did not change much during our sample period. Note that in the model we assume that households never move to states where house prices belong to a different category. In the online appendix Figure A3, we show that the fraction of households who move to different states or regions are much smaller than those that move within the state.

moments at most for each age cohort-year cell, for a potential maximum of $11 \times 17 \times 3 = 561$ moments. We lost 33 moments in 11 years, since wealth variables are only available for 1989 and 1994 before 1997. Further, we lost an additional 18 moments because the rent variable is missing for 1988 and 1989. The number of total matched moments therefore is 510.

3.3.2. Construction of Simulated Moments

In the second-stage estimation, we first choose a vector of structure parameters and solve the optimal decision rules as described in the previous section, taking the first-stage parameters as given. We then simulate households' behavior to construct our simulated moments under the given choice of parameters.

To initialize our simulation, we randomly draw 1,000 households from each age between 25 and 54 in the 1984 PSID data, for an initial simulated sample of 30,000 households. We then assign a series of moving, income, and house price shocks to each simulated path. The aggregate house price and income paths come from the actual realized prices in the household's state of residence. We update the simulated sample path based on the optimal decision rules.

Once all the simulated paths are complete, we compute the average profiles for our target variables in the same way that we compute them from the real data, i.e., by grouping households into different calender year $\times$ age cohort $\times$ house price level cells. Finally we construct a model fitness measure by weighting the differences between the average profile in the simulated model economy and the data with a weight matrix.[17]

The procedure is repeated until the weighted difference between the data and simulated profiles is minimized.[18] The online Appendix D provides more details on our MSM estimation procedure.

[17]The theoretically most efficient weighting matrix is the inverse of a sample variance-covariance matrix. According to Altonji and Segal (1996), however, the optimal weighting matrix, though asymptotically efficient, can be severely biased in small samples. Our weighting matrix, thus, takes the diagonal terms of the optimal weighting matrix for scaling purposes, while setting the off-diagonal term to be zero as in De Nardi, French, and Jones (2010).

[18]The minimization of weighted moment distance is achieved through a combination of a global population-based optimization using a differential evolution method and a local nongradient-based search routine via a simplex algorithm.

4. Benchmark Estimation Results

4.1. Housing and Wealth Profiles over Time and over the Life Cycle

Figures 1 to 5 show the fit of our baseline model to the empirical data profiles for different age cohorts. The solid line with solid dots depicts the empirical data profile, while the dashed line marked with crosses represents the average profile from our model. Additional figures that compare data and model simulation for different house price categories are provided in the online appendix.

Figure 1 charts the evolution of household net worth over time. Households become richer as they age for all cohorts. At the same time, the youngest cohort accumulates wealth at a much slower pace initially than do older cohorts. This is typical of life-cycle models, as young households expect their income to grow in the future and, therefore, do not save. Older cohorts are also richer on average than younger cohorts. Figure 2 shows that the homeownership rate starts at around 75 percent for the youngest cohort and quickly goes up to 90 percent in 10 years.[19] The other two older cohorts also demonstrate slight increases in homeownership rates. By the end of the sample period, most households have achieved homeownership. As shown in Figure 3, the average house value-to-income ratio is flat until 1996 for the two younger cohorts and until 1992 for the oldest cohort. Then, it begins to rise, and the speed of the increase is the fastest for the oldest cohort as the growth of house prices outstrips the growth of income for the oldest cohort.[20] The average rent-to-income ratio is flat for all three cohorts over the sample period (Figure 4). Note that as households age, the number of them that remain renters becomes small. The moving rates are mostly stable over the sample period except for the youngest cohort, which shows a slight decline between 1984 and 1997 (Figure 5).

Across states, as seen in Figure A6 in the online appendix, the homeownership rates of all cohorts in the most expensive states are slightly lower than those in the least expensive states. The average house value-income ratios, however, are much higher for those in the most expensive states than for those in the least expensive states. The ratios also grow much faster over the sample period (Figure A7 in the online appendix). While renters in the most

[19]The overall homeownership rate in our sample is much higher than the country as a whole. This is due to the sample criteria we use in order to maintain household stability. Recall that we only admit married couples with incomes above $10,000 to our sample.

[20]As a matter of fact, retired households no longer receive any income growth in our model.

expensive states on average also spend a larger share of their income on housing services, the time trend is less clear, since we have few renters in the sample, especially for the later years (Figure A8 in the online appendix).

Overall, our model captures some important trends in the data. We miss along some dimensions, though. We do not match mobility rates in 1985. This is because our model equilibrium does not best capture the economy in 1984. As a result, mobility rates are much higher between 1984 and 1985 in the model than in the data not just because of the realization of the shocks but because the house sizes we enter using 1984 data are not viewed as optimal by households in the model. Additionally, the model generates not enough house value-to-income ratio, especially for the youngest cohorts. Finally, as alluded to in our previous discussion, we overpredict homeownership rates for the older cohorts, especially the oldest. The result arises mainly because older households in our economy no longer face income uncertainty and we assume away expense uncertainty. This along with the assumption that house prices have a positive growth trend make ownership especially beneficial for these households.

4.2. Parameter Estimates and Identification

Table 2 shows estimated parameters. The estimate for the time discount factor, at 0.919, is in line with the values identified in the literature.[21] The estimate for the risk aversion parameter, at 7.158, is on the high end of empirical estimates.[22] The bequest strength L is estimated to be 7.562. While the time discount factor and risk aversion are largely determined by households' wealth accumulation earlier in life, the bequest strength is mostly driven by households' wealth profiles later in life.

As for the intratemporal utility function, the intratemporal elasticity of substitution between housing and nonhousing consumption is estimated to be 0.487, while the weight on housing parameter ω is estimated to be 0.011. These two parameters are largely identified through the cross-sectional as well as time series variation of house value-to-income ratios and homeownership rates. Households in expensive states spend more relative to their income

[21]For example, using the PSID, Carroll and Samwick (1997) estimate the rate of time preference to be approximately 11 percent.

[22]There exists a wide range of empirical estimates on risk aversion. Studies that attempt to reconcile the equity risk premium with rational behavior indicate that individuals' constant relative risk aversion may be above 10 (see, for example, Campbell, 1996, and Kocherlakota, 1996). By contrast, Halek and Eisenhauer (2001) examine life insurance purchases and find a median estimate of 0.888 and mean estimate of 3.735.

on housing, both when renting and when owning. The higher house value-to-income ratio requires a larger down payment, which takes longer to accumulate and delays transition to homeownership.[23]

Our point estimate of intratemporal elasticity of substitution is close to many of the micro estimates. For example, Hanushek and Quigley (1980) examine data from the Housing Allowance Demand Experiment, which involved a sample of low-income renters in Pittsburgh and Phoenix. Households in each city were randomly assigned to treatment groups which received rent subsidies that varied from 20 percent to 60 percent and a control group that received no subsidy. The estimated price elasticities were 0.64 for Pittsburgh and 0.45 for Phoenix. Siegel (2008) estimates the elasticity from the PSID over the period 1978-1997. Aggregating across households and using only the time series information, Siegel estimates the elasticity of substitution to be 0.53.[24] Flavin and Nakagawa (2008) use the PSID over the period 1975 to 1985. Instead of using households' self-reported house value, they construct a housing service measure and derive Euler equation conditions on consumption for households that do not move. Their estimate of the elasticity of substitution between housing and nonhousing consumption is a low 0.13.[25]

The house maintenance cost ψ is estimated to be 1.7 percent of the house value, which implies that the user cost of homeownership is $\psi + r_m - \mu_H^s = 7.4$ percent, where r_m is the mortgage interest rate and μ_H^s is the mean house price growth rate. An alternative measure of user cost using risk-free rate is $\psi + r - \mu_H^s = 3.4$. Those two numbers provide the upper and lower bound of user cost and are close to the $[3.3, 7.1]$ range of the user cost for homeownership calculated by Himmelberg, Mayer, and Sinai (2005) for 46 metro areas. While the *cross-section variation* of the house value-to-income ratio helps pin down the intratemporal

[23]To illustrate the implications of our estimated ω and ζ parameters on the cross-sectional house expenditure patterns, we compute the implied renters' housing expenditure shares for all 50 states and the District of Columbia based on the house price in the year 2005 using our parameter estimates and present it in Figure A9 in the online appendix. The share varies from 19 percent for the cheapest state (Illinois) at $44 per square foot, to 51 percent in the most expensive area (Washington, D.C.) at $454 per square foot.

[24]Siegel (2008) limits the sample to only homeowners, uses total household food expenditure as a proxy for nondurable consumption and the self-reported value of the owner-occupied house for housing, and assumes that durable consumption is constant until the household moves.

[25]By focusing on nonmovers, Flavin and Nagagawa's (2008) generalized method of moments methodology is robust to the existence of adjustment cost. However, their empirical estimates, which are based on consumption Euler equations, could be sensitive to assumptions about borrowing constraints and other market incompleteness.

preference parameters, the *average* level of the same ratio identifies the house maintenance parameter.

The rental premium is estimated to be 3.2 percent of the property value each year and is identified through matching homeownership profiles and rent-to-income ratios. The implied α parameter, which is the sum of the cost of capital, maintenance, depreciation, and rental premium, is then 6.6 percent.[26,27]

5. Alternative Estimations

To check the robustness of our estimation results and further understand the importance of admitting house price heterogeneity in the estimation, we now conduct three alternative experiments. In the first experiment, we vary house selling cost and investigate its effect on our parameter estimates. In the rest of the experiments, we explore the sensitivity of our estimates to different target moments. In particular, we test the importance of allowing for heterogeneity of house price movements both across states and over time. The results are reported in Table 3.

In the first experiment, when we double the house selling cost from 8 percent of the house value to 16 percent, the intratemporal elasticity of substitution between housing and nonhousing consumption declines slightly from 0.487 to 0.485. When the house selling cost is high, in order to match the same data moments on house value-to-income ratios and moving rates in the data, housing consumption becomes less substitutable with nonhousing consumption so that households are more likely to adjust house sizes when house prices change. A higher rental premium and a larger weight on housing that tilt preference toward homeownership is needed to encourage homeownership. Although the housing maintenance cost doubles from 1.7 percent of the house value to 3.8 percent, the rental premium increases significantly from

[26]The U.S. Department of Housing and Urban Development and the U.S. Census Bureau report a price-to-rent ratio of 10 or a rent-to-price ratio of 0.1 in the 2001 Residential Finance Survey (chapter 4, Table 4-2). Garner and Verbrugge (2009), using Consumer Expenditure Survey data drawn from five cities over the years 1982-2002, report that the house price-to-rent ratio ranges from 8 to 15.5 with a mean of approximately 12. In terms of rent-to-price ratio, these numbers imply a range of 0.065 to 0.125 with a mean of 0.083 percent. Our estimate of 0.066 therefore is on the low end. The cities included in this analysis are Chicago, Houston, Los Angeles, New York, and Philadelphia.

[27]All parameters have very small standard errors, as is typical with MSM estimations (see De Nardi, French, and Jones, 2010).

3.2 percent to 4.7 percent. The implied α parameter is, therefore, 10.2 percent as opposed to 6.6 percent in the benchmark.

In the second experiment, we match only aggregate moments over the period from 1985 to 2005. Specifically, we match wealth, homeownership rates, house value-to-income ratios, rent-to-income ratios, and moving rates of each cohort, but we no longer use moments that distinguish whether households live in states with high house prices or low house prices. Interestingly, the elasticity parameter estimate is now 0.810. The direct cause of the larger elasticity is that, over time, aggregate moments do not fluctuate nearly as much as their individual components. Put differently, the movement of aggregate moments over time are averages of moments in high house price and low house price states. By using aggregate moments only, we ignore important underlying heterogeneities. With a larger elasticity parameter, households can more easily substitute between housing and nonhousing consumption and, thus, have less desire to consume housing when house prices are high and become homeowners. In order to match the same housing expenditure shares, a higher weight is assigned to housing in the period utility function, 0.10 as opposed to 0.01 in the benchmark. This larger share more than offsets the disincentive for homeownership stemming from the now somewhat higher housing adjustment cost and lower rental premium. Finally, in order to match wealth profiles derived from the data especially for the years after 1996, the discount rate increases, making households more willing to save intertemporally.

In the last experiment, we match aggregate moments as in experiment two but use only the sample data between 1985 and 1993, a period close to what Bajari et al. (2013) used for their estimation. It is worth pointing out that this is a period when income and house prices were both relatively flat, as evident in both the wealth and the house value-to-income ratio profiles depicted in Figures 1 and 3. At 1.690, the elasticity parameter becomes even larger. Although the estimate is still much smaller than the 4.5 estimate in Bajari et al. (2013), it indicates that housing and nonhousing consumption are strong substitutes. Additionally, because wealth does not rise by nearly as much in the new sample period as in the benchmark sample period, the estimate for the curvature parameter declines substantially, apparently more than offsetting the effects from the stronger bequest motive and larger discount factor. Finally, with a higher elasticity of substitution, in order to match the housing expenditure share in the data, a much larger weight for housing expenditure is now required in the preference.

To summarize, our sensitivity analysis indicates that while house selling cost does not affect our model estimates much, admitting additional heterogeneity both over time and across section in the simulation is key for identification. In the next section, we will conduct policy analyses to illustrate the different policy implications associated with these alternative parameter estimates.

6. Implications

Using our estimated models in the benchmark and in the sensitivity analysis, we now conduct experiments that may have policy implications. We first draw the initial population from the 2005 PSID data, and then we simulate it forward using the optimal decision rules obtained in the estimation and the two alternative estimations we conducted in the last section, the one with higher house selling cost and the other using aggregate moments between 1985 and 1993. We simulate the model assuming that shocks to house prices and income follow their realized counterparts at the state and national levels, respectively.[28] It is worth noting that house prices started to decline in 2006 and bottomed out in 2010 for most states. Income, on the other hand, continued to rise until 2008 for the majority of the states, then it declined for one or two years before rising again. For many states, income had recovered to its precrisis 2005 level by 2011. We make two assumptions on households' expectation of mean house price growth rates starting from 2006: 1 percent as in the estimation versus 0 percent. In the second case, we recompute households' optimal decision rules using the estimated parameters in the benchmark estimation. In all experiments, we focus on aggregate homeownership rates, average house value for homeowners, average rents for renters, and average nonhousing consumption for all households.[29]

[28] We supplement these aggregate shocks with idiosyncratic shocks from the computer random number generator governed by their respective stochastic processes.

[29] The reported aggregate statistics are based on a sample constant in age distribution. We achieve this constant by admitting one new young age group into the sample each year while dropping the oldest households from the sample. The aggregate statistics are then computed using population weight for each age group from the 2010 census. Specifically, we only include households between the ages of 31 and 80 for the calculation. In other words, a household that is 30 in 2005 will not appear in the calculation of the aggregate statistics in 2005 but will enter the 2006 calculation as the household turns 31. Similarly, a household that is 80 in 2005 will appear in the 2005 sample but will drop out of the 2006 sample. This choice of age range is based on the initial population in 2005 that satisfies our selection criteria: being married through sample periods, not in government housing, not in the military, etc. See the data section for more details.

Table 4 provides our simulation results for 2006 to 2011. Several observations stand out. First, the biggest adjustment in most variables occurs in 2006. As alluded to previously, house prices began to decline in 2006 for most states. Additionally, our model does not capture 2005 household characteristics perfectly. Households will adjust even if house prices and income do not change.

Second, across the three economies in the benchmark policy analyses, homeownership rates decline the most in the first economy, and then they begin to inch up. Because of the larger housing adjustment cost in the second economy, households have a big inaction region in house sizes. The reduction in homeownership rates in the second economy is slower and more durable than it is in the first economy. By contrast, in the third economy when housing and nonhousing consumption are highly substitutable, low house prices relative to the numeraire nonhousing good actually present buying opportunities for some relatively rich homeowners who then upgrade to bigger houses. Although some homeowners do exit homeownership when they can no longer maintain the required mortgage loan-to-value ratio, more rich renters become homeowners. The end result is that homeownership rates are roughly unchanged through our simulation periods. Consistent with movements in house prices, the average house value for homeowners declines in all three economies but much less so in the third economy. This is due to the selection effect because more homeowners with substantial reduction in house value have exited homeownership in the first two economies than in the third. In addition, higher elasticity in the third economy implies that households that are adjusting would purchase larger houses in response to lower prices, leading to a lower reduction of housing value. Overall, rents increase in all three economies. Some relatively rich homeowners exit homeownership and become renters. These new renters are likely to rent larger houses than do existing renters. This is especially evident in the first two economies. Moreover, existing renters do not own housing and, therefore, they do not suffer any losses. The lower housing price allows them to rent more. The large elasticity of substitution between housing and nonhousing consumption in the third economy gives renters an extra incentive to rent larger houses. The different movements in rents in the first period come from the selection effects and the different calibration in the three economies. For instance, the rental premium is much larger in the second economy when housing adjustment cost is large. As a result, rents increase more in the second economy than in the first economy though not in the third economy because of the large elasticity in the third economy as we just discussed. Nonhousing consumption declines significantly in all three economies, especially in the third economy. Lower house prices lead to reduced wealth for homeowners and, therefore, fewer resources for nonhousing consumption. Lower house prices also make housing cheaper than nonhousing consumption. As a result, households, especially renters, allocate more of their resources to housing consumption than to nonhousing consumption. This is particularly true in the third economy, when housing is highly substitutable with nonhousing good.

Third, when households expect zero average house price growth rate, they are much more likely to exit homeownership in all economies, as housing is no longer an attractive investment asset that yields positive returns on average. The average house value of remaining homeowners, on one hand, declines less than that in the benchmark policy experiment, as those with more negative house price shocks have become renters. On the other hand, the much lower house prices reduce house value much more than in the benchmark policy analyses. The first force dominates in the first two economies while the second force dominates in the third economy. Rents increase more in all three economies. Nonhousing consumption, by contrast, declines more in all three economies.

To summarize, the previous analyses indicate that different parameter estimations resulting from different assumptions in the model or different target model moments have important implications on how the economy responds to economic shocks. In particular, in the presence of negative house price shocks, the homeownership rate decreases much less when the housing adjustment cost is large or when housing and nonhousing consumption are more substitutable. Nonhousing consumption, however, declines much more.[30,31]

7. Conclusions and Future Extensions

In this paper, we estimate a dynamic structural model of household consumption over the life cycle augmented with housing with a focus on the intratemporal elasticity of substitution between housing and nonhousing consumption. We explicitly model housing adjustment along both the extensive margin of owning versus renting and the intensive margin of house sizes. The model also includes a credit constraint in the form of collateral mortgage borrowing. The

[30]We cannot meaningfully compare our model simulations with the data, since the PSID does not cover 2006, 2008, or 2011. The 2005 nonhousing consumption is extrapolated based on food and housing consumption as in Ziliak and Kniesner (2005).

[31]We conduct another analysis in which we let income follow its stochastic process instead of the actual realization. We find that policy results do not change significantly. This suggests our policy results are mostly driven by changes in house prices rather than changes in income.

paper, thus, contributes to the analysis and the understanding of household housing demand and the impact of the housing market on household consumption.

Our estimation indicates that the intratemporal elasticity of substitution between housing and nonhousing consumption is about 0.487. The estimate is robust to different parameterizations of house selling cost but sensitive to the degree of house price heterogeneity that is admitted by the model. When less heterogeneity is admitted, housing and nonhousing consumption are much more substitutable. More importantly, different estimates of the elasticity parameter have important implications on how the economy responds to house price and income shocks. In particular, a higher substitutability between housing and nonhousing consumption and a high housing adjustment cost reduce the impact of house price and income shocks on homeownership but increase the impact on nonhousing consumption.

Despite our reconciliation of the different estimates of the intratemporal elasticity of substitution between housing and nonhousing consumption, it remains a challenge as to what is the relevant parameter estimate to use for conducting policy experiments. There are many natural extensions to our model that potentially help further test the robustness of our results. One is to allow for richer household portfolio decisions by differentiating further between stocks and bonds in a household's liquid asset menu. Another is to model mortgage contracts more explicitly and realistically by imposing mortgage amortization requirements as well as refinancing charges. We leave these to future research.

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Table 1
Calibrated Parameters

Parameter	Symbol	Value
Demographics		
Maximum life-cycle period	T	75
Mandatory retirement period	J	40
Labor income and house price processes		
Mean permanent income growth	μ_ν^s	0.012
Standard deviation of aggregate income shock	σ_ν^s	0.022
Standard deviation of permanent income shock	σ_ν^i	0.10
Standard deviation of temporary income shock	σ_ε	0.22
Income replacement ratio after retirement	θ	0.70
Mean growth rate of house prices	μ_H^s	0.010
Standard deviation of aggregate housing return	σ_H^s	0.015
Standard deviation of individual housing return	σ_H^i	0.10
Correlation between permanent income and housing return	ρ	0.30
Wealth floor	η	0.06
Liquid savings		
Risk-free interest rate	r	0.027
Mortgage interest rate	r_m	0.067
Housing and mortgage		
Down payment requirement	δ	0.20
House selling cost	ϕ	0.08

Table 2
Estimated Structural Parameters – Benchmark

Parameter	Symbol	Value	Std. Err.
Discount rate	β	0.919	0.0005
Curvature parameter	γ	7.158	0.1349
Bequest strength	L	7.562	0.2257
Weight on housing	ω	0.011	0.0001
Intratemporal elasticity of substitution	ζ	0.487	0.0073
Housing maintenance cost	ψ	0.017	0.0002
Rental premium	$\alpha + \mu_H^s - r - \psi$	0.032	0.0002

Table 3
Estimated Structural Parameters – Alternative Estimations

Parameter	Symbol	Value	Std. Err.
Estimation 1: higher housing selling cost			
Discount rate	β	0.901	0.00002
Curvature parameter	γ	7.270	0.1201
Bequest strength	L	7.235	0.2429
Weight on housing	ω	0.017	0.00001
Intratemporal elasticity of substitution	ζ	0.485	0.0042
Housing maintenance cost	ψ	0.038	0.00002
Rental premium	$\alpha + \mu_H^s - r - \psi$	0.047	0.000002
Estimation 2: using aggregate moments only			
Discount rate	β	0.948	0.002
Curvature parameter	γ	6.307	0.148
Bequest strength	L	6.593	0.437
Weight on housing	ω	0.100	0.002
Intratemporal elasticity of substitution	ζ	0.810	0.041
Housing maintenance cost	ψ	0.023	0.0002
Rental premium	$\alpha + \mu_H^s - r - \psi$	0.026	0.0005
Estimation 3: using aggregate moments only (1985 to 1993)			
Discount rate	β	0.981	0.003
Curvature parameter	γ	1.200	0.309
Bequest strength	L	16.485	2.588
Weight on housing	ω	0.407	0.008
Intratemporal elasticity of substitution	ζ	1.690	0.002
Housing maintenance cost	ψ	0.047	0.0003
Rental premium	$\alpha + \mu_H^s - r - \psi$	0.022	0.0004

Table 4
Model Implications

	Base case				Larger house selling cost				Larger elasticity between h and nonh			
Year	Homown (%)	Houseval (%)	Rent (%)	Cons (%)	Homown (%)	Houseval (%)	Rent (%)	Cons (%)	Homown (%)	Houseval (%)	Rent (%)	Cons (%)
Benchmark: 1% expected HPI growth												
2006	-6.76	-17.93	0.74	0.74	-5.39	-16.46	2.17	-3.82	0.11	-39.68	-15.82	-18.45
2007	0.25	-4.36	10.79	-0.07	-1.12	-3.87	11.09	-7.31	0.80	-4.10	18.02	0.86
2008	1.09	-9.57	3.55	-2.21	-0.67	-8.04	1.59	-2.78	0.69	-9.09	12.24	-1.16
2009	1.67	-5.59	4.70	-0.76	-0.19	-6.76	6.03	0.01	0.81	-4.16	9.19	-0.05
2010	0.65	-4.28	5.10	1.48	-0.39	-6.06	2.95	1.66	0.20	-3.39	14.04	2.31
2011	0.54	-5.43	1.12	-0.90	-0.59	-5.80	0.16	-0.45	0.39	-3.99	4.91	-0.29
Alternative economy: 0% expected HPI growth												
2006	-11.73	-16.96	2.56	0.16	-9.11	-15.92	13.74	-10.86	-6.47	-45.36	-1.29	-15.93
2007	-0.85	-3.65	9.35	-0.09	-0.89	-3.94	12.91	0.69	-0.60	-4.16	10.84	0.86
2008	-0.17	-9.11	3.74	-2.12	0.06	-9.35	2.27	-1.90	-0.43	-8.98	11.05	-1.10
2009	0.98	-5.34	3.85	-0.72	1.07	-5.88	3.57	-0.42	0.10	-4.13	6.34	-0.12
2010	-0.08	-4.21	5.33	1.54	0.48	-4.84	5.34	2.48	-0.74	-3.30	10.68	2.28
2011	0.02	-5.50	1.62	-0.80	0.64	-6.06	2.75	-0.43	-0.97	-4.16	4.70	-0.25

Note: The homeownership results are reported as differences in percentage points from the previous year. House value, rent, and nonhousing consumption are reported as percentage changes from the previous year. The average house value is for homeowners only, while the average rent is for renters only. Homeown=homeownership, Houseval=house value, Cons=nonhousing consumption, h=housing consumption, nonh=nonhousing consumption, HPI=house price index.

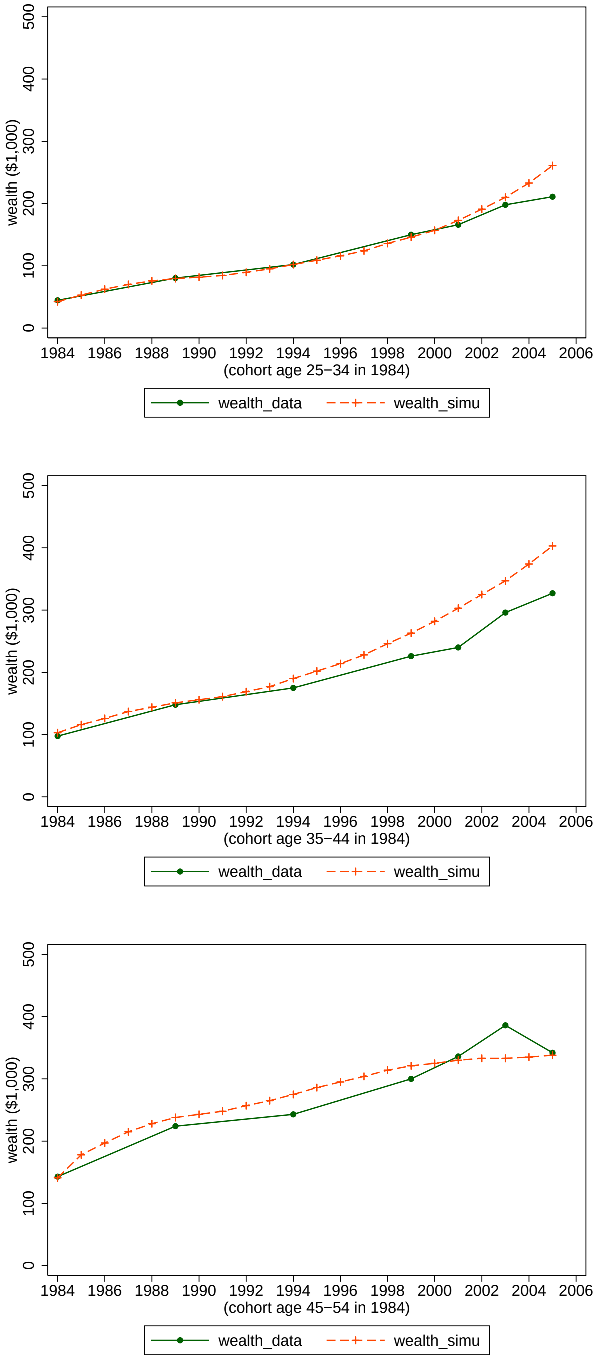

Figure 1. Wealth by cohorts in all states

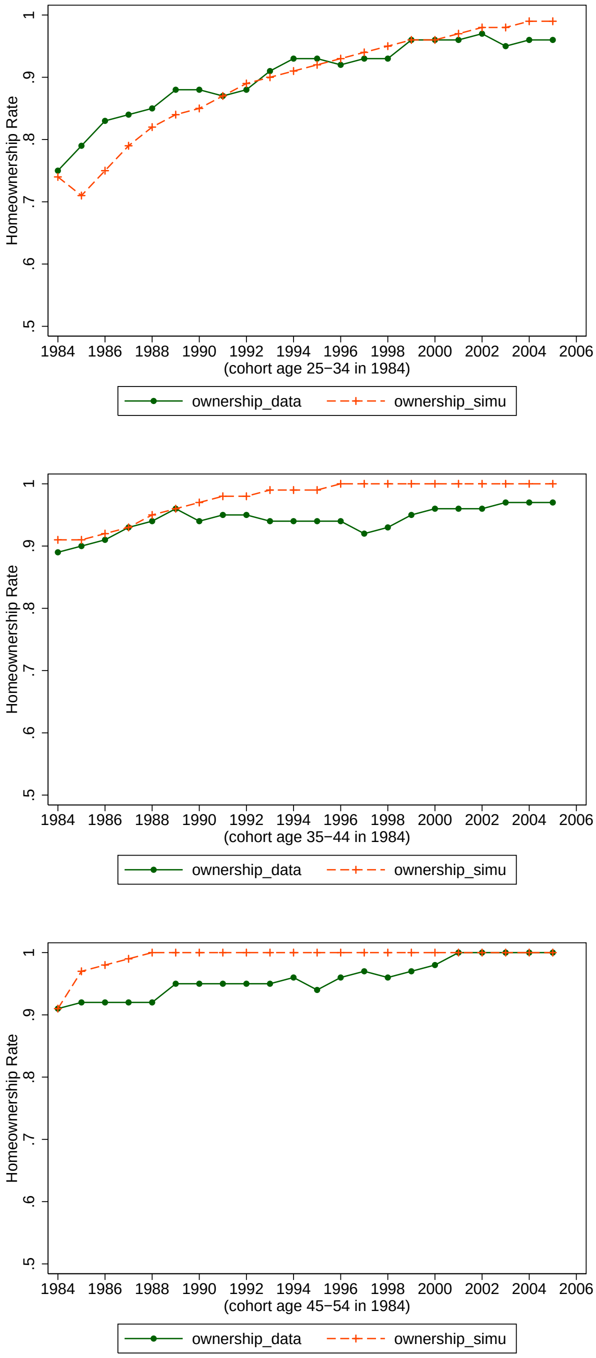

Figure 2. Homeownership by cohorts in all states

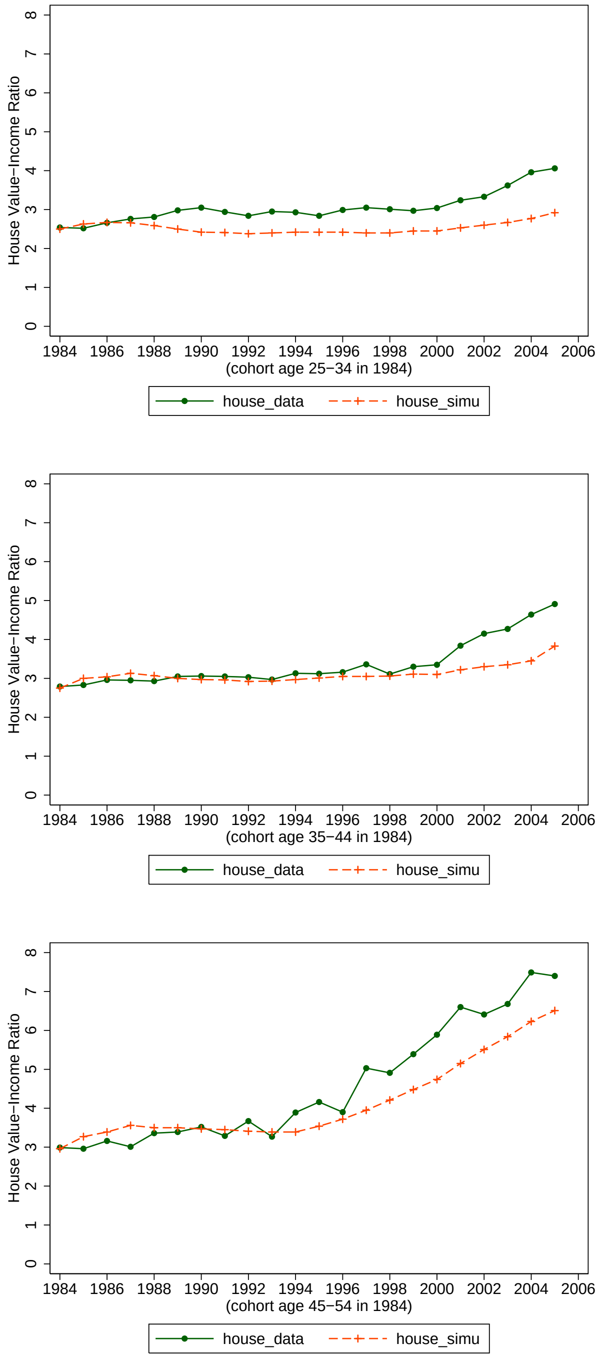

Figure 3. House value-to-income ratio by cohorts in all states

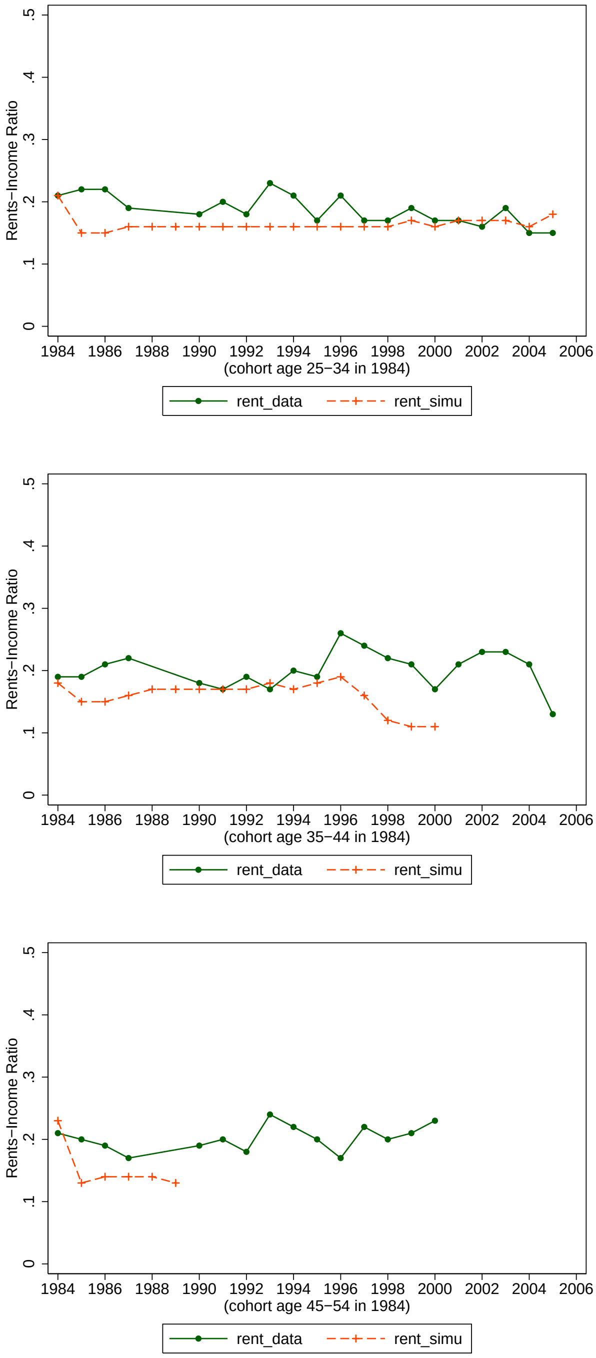

Figure 4. Rent-to-income ratio by cohorts in all states

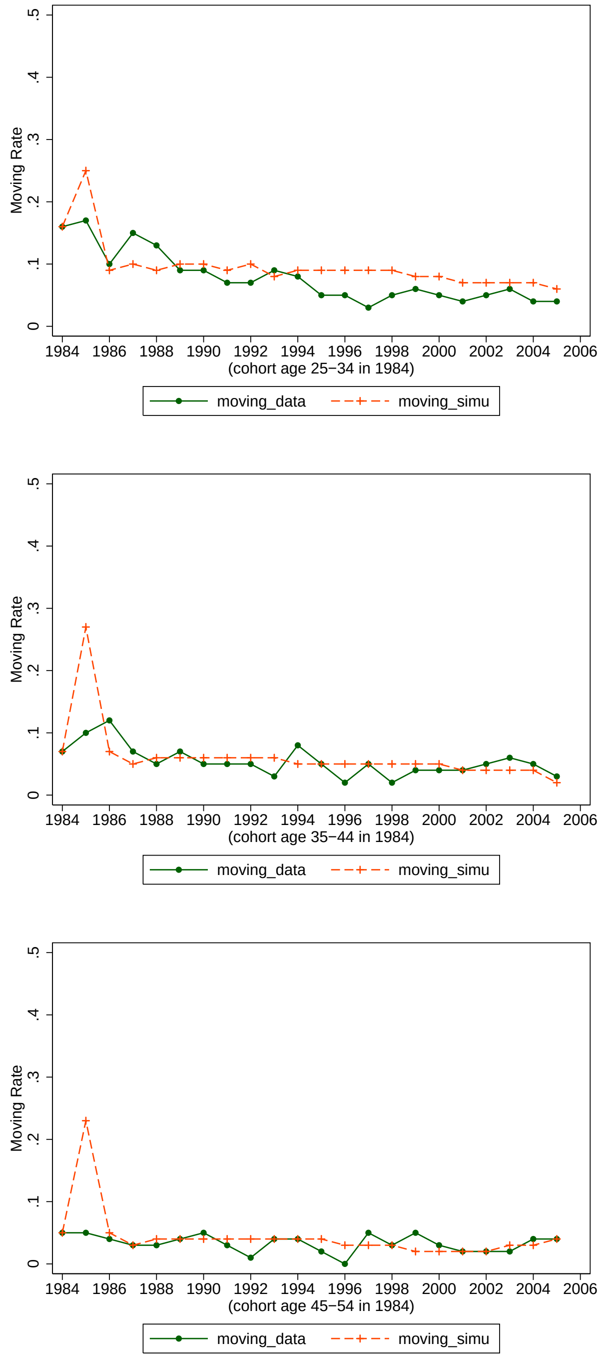

Figure 5. Homeowners' mobility by cohort in all states