Nominal Loss Aversion, Housing Equity Constraints, and Household Mobility: Evidence from the United States

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Abstract

This paper exploits the recent variation in U.S. house prices to examine the effect of equity constraints and nominal loss aversion on household mobility. Detailed data from the 1985-1996 National Longitudinal Survey of Youth (NLSY79) were matched with house price data from 149 metropolitan areas to estimate instrumental variables linear probability and semi-parametric proportional hazard models of intra-metropolitan mobility. Household mobility is significantly influenced by nominal loss aversion. There is little evidence that low equity because of fallen house prices constrains mobility.

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Loss Aversion, Equity Constraints, and Mobility

I. INTRODUCTION

Housing markets often exhibit behavior that cannot be explained by standard asset-market models (Poterba [43]). For example, they display rapid swings in prices, strong positive correlation of prices and trading volume over the housing cycle, and the observed reluctance of prospective sellers to reduce asking prices in down markets (Stein [53], Genesove and Mayer [13,14]). An important part of recent research has been to propose and empirically test new theories that explain these puzzles. Two seemingly related, but competing, theories have emerged. The first is housing equity (or collateral) constraints, analyzed by Stein [53], Genesove and Mayer [13], Henley [24], Lamont and Stein [31], Ortalo-Magne and Rady [40], and Chan [7], among others. The second, and quite provocative, is nominal loss aversion, analyzed first by Genesove and Mayer [14].

Although both theories rely on the same propagation---a decline in nominal house prices---they have distinct implications for housing market behavior and government policy. Equity constraints occur because of down payment requirements in mortgage lending. Specifically, because most home purchases are mortgage-financed, housing is a highly leveraged asset. A nominal price decline can result in equity- (or down payment-) constrained households who cannot move, which decreases market demand, and results in further price declines that further constrain household mobility. This suggests some sort of market failure that, in principle, could be addressed through government policy to ease binding equity constraints. In contrast, nominal loss aversion, whereby households are averse to realizing nominal housing market losses, and, hence treat gains and losses asymmetrically, is a characteristic of preferences, which typically are not thought of as affected by government policy instruments.

This paper empirically examines the effect of equity constraints and nominal loss aversion on household mobility. The data used are on young homeowners from the National Longitudinal

Survey of Youth (NLSY79). Young homeowners are an ideal group to study. They have high job and geographic mobility and are highly leveraged. They are the most susceptible to equity constraints. In addition, the analysis exploits the significant metropolitan variation in housing market performance in the United States for 1985-1996. This period encompasses the well-known recessions in the energy states, the Northeast, and California, as well as the rising economic tides in the South, Midwest, and Pacific Northwest. These data display rich variation in nominal losses and gains.

Moreover, the data are unique. Detailed data on demographics, employment, and wealth come from the public-use version of the NLSY79. These were matched to administrative address data to construct mobility histories that are more detailed than those available from the restricted-access NLSY79 Geocode data. In addition, housing value and mortgage data from the NLSY79 master file were used. These data are top-coded on the public-use version of the data. The NLSY79 data were matched with weighted-repeat-sales house price indices for 149 metropolitan areas provided by the Federal National Mortgage Association (Fannie Mae) and the Federal Home Loan Mortgage Corporation (Freddie Mac). Analysis with data drawn from a broad sample of metropolitan areas is an important contribution of this paper, because most of the existing empirical analyses of equity constraints and nominal loss aversion used data from the New York metropolitan area (Chan [7]) and the downtown Boston condominium market (Genesove and Mayer [13,14]), respectively. The final data set was used to estimate instrumental variables linear probability and semi-parametric proportional hazard models of intra-metropolitan mobility.

There are two primary findings. First, household mobility is significantly influenced by nominal loss aversion. Second, there is little evidence that low equity *because* of fallen house prices constrains mobility. Overall, these findings confirm those in Genesove and Mayer [14] and suggest that loss aversion is an important housing market phenomenon across a broad spectrum of metropolitan areas.

The paper is organized as follows. Section II describes the economic relationship between housing equity, nominal loss aversion, and mobility, and discusses related empirical work from the literature. Section III discusses the data and preliminary analysis. Section IV outlines the econometric framework. Section V discusses the estimation results. There is a brief conclusion.

II. PREVIOUS LITERATURE

The basic theory of equity constraints was laid out in Stein [53] and is illustrated most clearly with an example. Consider a household that purchases a $100,000 home with a 10 percent down payment.[1] If house prices rise by 10 percent, the home is worth $110,000 and the household has $20,000 in equity (or 18.2 percent). For the same down payment requirement, and as long as the household had sufficient income, the household could trade up significantly even though housing is more expensive, because it now could use that equity to purchase a $200,000 home with 10 percent down. However, with a 10 percent decline in prices, the home is worth $90,000, and the household has no equity. Although housing is now cheaper, the household could not make the down payment on the same home without other wealth.[2] With no other wealth, the household cannot move and remain a homeowner.[3] Thus, nominal house price changes can have asymmetric

[1] The down payment requirement for conventional mortgages ranges from 10 to 20 percent. Changes in secondary mortgage market underwriting guidelines have made 5 percent down mortgages more prevalent.

[2] Another possible source of down payment funds is transfers from family or friends. Engelhardt and Mayer [12] found that about 25 percent of first-time buyers receive such transfers, but among repeat buyers they were rare (4 percent).

[3] This example abstracts from other costs that may further deter households. These include private mortgage insurance (discussed later in the text), closing costs, broker costs, and moving costs. Moving costs can be at least $750-$4500, depending on the area, distance, and volume. For a $155,000 mortgage in the greater New York metropolitan area, Caplin, Freeman, and Tracy [4] estimated the closing costs to be between $5,100 and $8,400, or 3.3 and 5.4 percent of the loan, respectively. In general, brokers are paid 6 percent of the sale price of the home.

effects on mobility: households can lever capital gains to purchase larger homes, but become constrained by capital losses.[4]

Stein [53] formalized this intuition into a liquidity-based model of the housing market. He showed that this asymmetric effect can result in within-equilibrium housing market multipliers and multiple equilibria. These results hold even if constrained households have the option of moving and renting. The strength of the multipliers depended on the fraction of owners that were constrained movers---only in markets with a sufficiently large fraction of constrained owners could there be significant feedback effects to house prices.[5] Ortalo-Magne and Rady [40] have generalized Stein's findings in an overlapping-generations framework.[6]

Recent research on Boston and New York has provided some evidence in support of equity constraints.[7] One implication of Stein's model is that constrained owners may “go fishing,” i.e., offer their houses for sale at an above market price that would allow a move in the (low probability) event that a buyer arrives to pay that price. Genesove and Mayer [13] examined the effect of

[4] It is important to note that this leverage effect is independent of a wealth effect from the capital gain (or loss). To illustrate this, consider a model with no mortgages---all homes must be purchased in cash. A household owning a $100,000 home that experienced 10 percent appreciation could afford to purchase just a $110,000 home. The fact that households can borrow to finance the purchase of a home means that a dollar of capital gain can buy more than a dollar of housing.

[5] Stein [53] referred to this as “packing.” Mayer [36] has provided striking evidence on this. He examined the level of housing leverage in Massachusetts (predominantly the greater Boston metropolitan area) during the boom and bust of the 1980s and 1990s. He found that of the 580,000 households that purchased single family homes between 1982 and 1992, more than 150,000 had less than 5 percent equity in 1992, and the majority of these had no equity. Even lower equity levels were found in the condominium market. These findings suggest the great potential for collateral constraints. Lamont and Stein [31] observed substantial variation in packing across metropolitan areas using data from the American Housing Surveys.

[6] In this light, housing equity plays a role similar to collateral for firms. There, temporary economic shocks that depress the value of assets used for productive purposes and collateral can reduce the net worth of firms, reduce the asset demand for constrained firms, and result in lower asset prices. This further reduces net worth and feeds back into prices. This link between asset prices and collateral has been examined recently by Kiyotaki and Moore [30] and Shleifer and Vishny [51], among others.

[7] Kiyotaki and Moore [30] provide a summary of the empirical evidence consistent with collateral constraints for firms. Henley [24] examined the effect of negative housing equity on the mobility of British households in the early 1990's. He used a sample of 3,530 households from the 1991-94 waves of the British Household Panel Survey and estimated a semi-parametric duration model of mobility with competing risks. The three risks modeled were a move to an owner-occupied home, a move to a public-sector rental, and a move to a private-sector rental. He found significant lock-in

equity on the time-to-sale and listing behavior of potential sellers in the downtown Boston condominium market. They found that constrained sellers (i.e., with less than 20 percent equity) were more likely than unconstrained sellers to ask above market prices, and that this resulted in an inverse correlation between prices and time-to-sale.

Chan [7] has provided the only direct evidence on equity and mobility in the United States. She examined the experience of homeowners in the New York metropolitan area with a unique data set of Chemical Bank adjustable-rate mortgages (ARM). She found that constrained households (i.e., less than 20 percent equity) experienced a 24 percent reduction in mobility relative to unconstrained owners in the four years after the decline in prices. Her study is noteworthy in that the mortgage data were of unusually high quality, with arguably no measurement error in the mortgage spell length and explanatory variables. A "move" was defined as a mortgage termination. Chan argued that ARM refinancing was rare because of a low-cost option of conversion to a fixed-rate loan at the market rate, and, as such, mortgage duration was a good proxy for residence duration. In addition, she provided evidence from public records that 95 percent of ARM terminations in her sample were from moves.[8]

Nominal loss aversion is a central feature of the prospect theory of Kahneman and Tversky [26]. Barberis, Huang, and Santos [3] have integrated nominal loss aversion into an asset pricing model that explains the large equity premium, volatility in expected returns, and level of average return in the stock market better than traditional consumption-based models. Genesove and Mayer [14] were the first to examine nominal loss aversion in the housing market. The basic hypothesis is that homeowners treat gains and losses differently, and are reluctant to realize nominal losses;

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from fallen prices. His estimates suggested that the mobility of negative equity households would have been 50 percent higher if they had had a positive equity position.

[8] Caplin, Freeman, and Tracy [4], Peristiani, Bennett, Monsen, Peach, and Raiff [41], and Archer, Ling, and McGill [2] found evidence that low equity constrains mortgage refinancing in the United States.

hence, they will set higher list prices and have longer time on the market in the hope that they will find a buyer with an offer high enough to attenuate the nominal loss. Genesove and Mayer [14] used similar (but updated) data from the downtown Boston condominium market as in Genesove and Mayer [13]. However, in the updated analysis, they found that most of seller behavior seemed to be driven by nominal loss aversion. Only about one-quarter of the effect of declining nominal house prices on listing, pricing, and time on the market in Boston operated through equity constraints. The estimates in Chan [7] also suggested some nominal loss aversion in the New York area.

III. DATA AND PRELIMINARY ANALYSIS

The data used in the current paper are unique and are described in detail in the appendix. The primary data cover the 1985-96 period and are from the public-use version of the National Longitudinal Survey of Youth (NLSY79). These have been supplemented by data from three additional sources. First, administrative data on addresses were used to construct mobility histories for each household. This could not be done with data on the public-use and restricted-access Geo-code files. Second, housing value and mortgage data from the NLSY79 master file were used. These data are top-coded in the public use version. Finally, Fannie Mae/Freddie Mac weighted-repeat-sales metropolitan house price indices were matched to each household-year observation. The final sample consists of 6,461 household-year observations that comprised 3,112 residence spells for 2,348 households.[9] Sample descriptive statistics are given in Table A-1 in the appendix.

[9] This sample differs from the "potential" sample of *all* NLSY79 households for the following reasons. First, the analysis in this paper focuses on the mobility of homeowners, and, for the young age ranges studied here, the majority of NLSY79 households do not own homes. Second, the analysis focuses only households in the 149 metropolitan areas for which there are Freddie Mac/Fannie Mae price indices. Third, this analysis relies on matched administrative address histories to track intra-metropolitan moves within counties as described in detail the appendix. A small fraction of households were not matched to the address data and do not appear in the sample. In an early version of this paper, I examined intra-metropolitan own-to-rent and inter-metropolitan moves, as well. Analyses of those moves required only the NLSY79 Geocode data and not the matched address data. The Geocode data are available for all NLSY79 households. To determine whether there was any sample selection bias from using the slightly smaller sample with

This sample has a number of advantages. First, it focuses on young households, many of which own their first home. They are the most mobile and the most leveraged, hence the most likely to be equity-constrained when house prices decline. This sets up a strong empirical test for equity constraints. If there is little evidence in favor of constraints in this sample, then equity is likely unimportant for mobility, because one would not expect it to affect the mobility of older, wealthier households. Alternatively, if there is evidence in favor of constraints, then equity is important for mobility, at least for young households.

Second, the sample spans a period of substantial variation in metropolitan housing market performance that can be used to identify any equity and loss aversion effects. This is particularly important for two reasons. First, the model and empirical work of Stein [53] and Lamont and Stein [31] on equity constraints have indicated the possibility of multiple equilibria, such that equity constraints may be at play in some metropolitan areas, but not others. Second, while both the empirical tests by Genesove and Mayer [13,14] and Chan [7] for equity constraints and loss aversion used incredibly rich data and have provided very persuasive evidence in favor of these phenomena, in the end, they apply only to the experiences of the condominium market in downtown Boston and the greater New York metropolitan area housing market. A question of first-order importance is whether these results generalize to the nation as a whole.

Third, there are standard omitted variable issues. Previous studies have not been able to track changes in the demographic and economic circumstances of the households under study that almost surely affect mobility behavior strongly and may happen to be correlated with local housing

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matched address data, I estimated mobility models for intra-metropolitan own-to-rent and inter-metropolitan moves using the full dataset based on the Geocode data and the slightly smaller dataset based on the matched address data, and the results did not differ, which suggested little role for sample selection bias. In the current version of this paper, intra-metropolitan own-to-rent and inter-metropolitan moves are treated as censored in period in which the move occurs.

and labor market conditions.[10] These include divorce, unemployment, and family decisions. Thus, another key question is whether once one controls for these other factors evidence of equity constraints and loss aversion remains.

From the discussion in the previous section, there are many requirements for equity constraints to bind. First, most wealth must be in housing. If the household in the example above had $10,000 in other wealth and house prices fell 10 percent, the household is not constrained because it could have used its other wealth to make a down payment. Empirically, highly leveraged households are predominantly young, first-time owners with little other wealth. Table 1 shows the fraction of wealth in housing at the time of first home purchase for a sample of NLSY79 households. Overall, 80.5 and 90.6 percent of liquid assets go into housing at purchase, at the mean and median, respectively. Column 4 indicates that 25 percent of first-time homebuyers essentially have no other wealth after first purchase. These figures are higher when wealth is measured as liquid assets less debts.

Second, periods of declining nominal house prices are required for binding equity constraints, and, obviously, for nominal loss aversion.[11] Table 2 summarizes recent episodes of falling prices for selected U.S. metropolitan areas (grouped geographically). These episodes were measured using the Freddie Mac/Fannie Mae price indices described above. Price declines have been quite large. In Houston, there was a 27 percent decline from the market peak to the trough

[10] For example, the LINK data used by Genesove and Mayer [13] had very detailed information on the listing behavior of sellers in the downtown Boston condominium market that were necessary for the very clean empirical tests for equity constraints on seller behavior. But these data do not provide information on motivations for sale (Glower, Haurin, and Hendershott [15]). Chan [7] used Chemical Bank mortgage records for the New York metropolitan area, but these data only provided detailed financial and demographic data at the time of underwriting. In both studies, nothing is known about what precipitated the move or the type of move (intra- vs. inter-metropolitan or own-to-own vs. own-to-rent).

[11] It is important to emphasize that equity effects require declining nominal prices. Real prices can fall even when nominal prices are flat or rising, as long as inflation exceeds nominal appreciation. In addition, loss aversion is with respect to nominal, not real, losses. Genesove and Mayer [14] provide evidence that it is nominal losses that matter.

(column 3). Peak-to-trough declines of 15 percent or more occurred in Texas, New England, and Southern California. These declines were large enough to have constrained most owners with less than 20 percent equity at the market peak.

Third, house prices must decline sufficiently rapidly for equity constraints to bind. If not, forward-looking owners could increase saving to maintain the option of purchasing another home.[12] The market peaks and troughs indicated in columns 1 and 2 of Table 2 show that the duration of price declines has varied greatly. For example, the declines in Texas and New York lasted 2-3 years. In contrast, prices declined for 5 and 7 years in California and Connecticut, respectively. To control for declines of different duration, column 4 presents the annual average decline from peak-to-trough. By this metric, house prices declined rapidly: an annual average rate of 4.5-10 percent in the energy states, for example. Because many owners have 10 percent equity at purchase, these declines may have constrained a large number of households within one year.[13]

Table 3 examines the effect of falling nominal house prices on the distribution of housing equity in the NLSY79 sample. Panel A, column 1 describes equity in the purchase year for all observations. 55.2 percent of owners had 20 percent or more in equity; 20.5 percent had between 10 and 19 percent equity; and, 24.3 percent had less than 10 percent equity. Viewing the spell data as an (unbalanced) panel, column 2 shows contemporaneous equity. As expected, equity grows over time: 63.1 percent of owners had 20 percent or more; 23.5 percent had between 10 and 19 percent; and just 13.4 percent had less than 10 percent.

Seller behavior is only weakly affected by real losses. In the analysis below, there was no evidence that real losses mattered.

[12] This issue is not addressed in the static model of Stein [53], but is in the dynamic model of Ortalo-Magne and Rady [40]. Engelhardt [10] examined the effect of housing gains and losses on homeowner saving behavior.

[13] Furthermore, surveys of homeowners in boom and bust housing markets by Case and Shiller [6] have provided provocative evidence that homeowners are not forward-looking. Rather, they seem to base their expectations of future price movements on past price behavior, not fundamentals, so that it may take a number of years of declining prices until households expect prices to decline and adjust their behavior. Based on the declines in column 4, even two years

Equity growth can come from a decline in mortgage debt through normal repayment, full or partial prepayment, and house price appreciation. Columns 3 and 4 repeat the tabulations in columns 1 and 2 for the metropolitan areas with stable or rising housing markets, respectively. Columns 5 and 6 are defined similarly for the metropolitan areas with weak housing markets. A comparison of columns 3 and 4 to 5 and 6, respectively, clearly illustrates the effect of appreciation. In falling markets, the distribution of equity shifts toward low equity between the initial and current periods. In stable and rising markets, the distribution shifts toward greater equity. Rising prices confer positive shocks; falling house prices confer negative shocks.

Because housing is the largest component of non-pension household wealth, house price fluctuations can have a large impact on the overall household balance sheet. Column 1 in panel *B* describes net worth relative to home value in the purchase year for all observations. 81.4 percent had net worth of 20 percent or more; 10.7 percent had between 10 and 19 percent; and, 7.9 percent had less than 10 percent. Column 2 shows the contemporaneous ratio. As expected, net worth grows over time: 84.8 percent had 20 percent or more; 8.0 percent had between 10 and 19 percent; and just 7.2 percent had less than 10 percent. A comparison of columns 3 and 4 with 5 and 6, respectively, clearly illustrates the effect of appreciation. In weak markets, the distribution of net worth remains very close between the initial and current periods. In stable and rising markets, the distribution shifts toward more net worth. Again, weak markets confer negative shocks to the whole balance sheet.

IV. ECONOMETRIC FRAMEWORK

The 6,461 household-year observations in the sample described above represent an unbalanced panel on 2,348 households that comprise 3,112 residence spells. When viewed as

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of prices falling at an annual rate of 2.5 percent (roughly the lower bound in column 4) would have been enough to have constrained highly leveraged owners.

duration data, the sample has no left-censored spells. The longest (right-censored) spell is 12 years. Of the 3,112 spells, 596 were completed by an intra-metropolitan own-to-own move. Figure 1 shows the Kaplan-Meier empirical hazard for these moves. The hazard rises and falls across spell periods, with peaks at almost 15 percent in year 5 and 4 percent in year 8.[14]

To test for the presence of equity constraints and nominal loss aversion more formally, the econometric analysis employs both linear probability and proportional hazard models of mobility. Mobility is defined as a transition out of the current residence spell. Let i index households and t time periods, then the linear probability model is

$$D_{it}^{MOVE} = Z_{it}^{\prime}\beta + P_{it}^{\prime}\kappa + u_{it}, \qquad (1)$$

where D^{MOVE} is a dummy variable that is one if the household moves within a metropolitan area and remains a homeowner and zero otherwise, Z is a vector of explanatory variables, and P is a vector of spell period dummies. In the duration model, let index t spell periods, then the hazard, λ, which measures the probability of moving in period t conditional on having not yet moved, is

$$\lambda_{it} = \lambda_{0t}\exp(Z_{it}^{\prime}\beta), \qquad (2)$$

where λ_0 is the baseline hazard.[15]

In specifying the functional form of Z, there is a fundamental identification issue. As noted in the introduction, both equity constraints and loss aversion have the same propagation mechanism---a decline in house prices. Households with low *contemporaneous* equity are composed of two types: those who began with low equity and those who began with moderate equity but saw that eroded by the price decline. As long as contemporaneous equity is used to

[14] The shape of this hazard is similar to those in Sinai [52], who used the Panel Study of Income Dynamics (PSID), and Chan [7], who used Chemical Bank mortgage records. This is not inconsistent with young homeowners making lifecycle adjustments to housing consumption by trading up, perhaps to accommodate an increase in family size (Henderson and Ioannides [23]).

define the equity-constrained and unconstrained groups, it will not be possible to determine the effect of loss aversion distinct from equity constraints by examining the mobility behavior of the constrained (low equity) group alone. However, the effect of loss aversion can be identified by examining the mobility behavior *within* the unconstrained (high equity) group, because even when prices decline, these households still have enough equity to avoid being constrained.

There is an additional reason why contemporaneous equity should not be used in the specification. In essence, contemporaneous equity is a function of initial equity at purchase and price appreciation. Households with low contemporaneous equity began with low equity, and had little or no appreciation, or had low-to-moderate equity, but saw that eroded by price declines. However, because initial low equity may be correlated with expected duration in the home, the initial equity position itself may have an effect on mobility independent of the contemporaneous equity position. This means that controls for initial equity position must appear in the specification. This implies that only those households with low initial equity *and* who experienced fallen house prices are truly at risk of being equity constrained.

With these issues in mind, I defined the household's home equity stake by the loan-to-value (*LTV*) ratio at purchase. A high *LTV* means low equity. Mortgage underwriting guidelines suggest that households with an *LTV* greater than 0.80 (i.e., less than 20 percent equity) might be constrained. Mortgages with less than 20 percent down require the purchase of private mortgage insurance (PMI), which is expensive: 0.75 percentage points applied to the entire mortgage, not just the increment of the mortgage that would bring the down payment up to the 20 percent level.[16] Therefore, in the empirical analysis, all households with an *LTV* at the time of purchase of greater than 0.80, or less than 20 percent in equity, are considered potentially at risk of being constrained

[15] Here, *Z* includes a full set of calendar-year dummies.

should nominal prices fall. This is the same definition used by Genesove and Mayer [13,14], Lamont and Stein [31], Caplin, Freeman, and Tracy [4], and Chan [7]. However, because one would expect that the higher the *LTV*, the more likely the household is to become constrained when prices fall, the specification allows for *differential* effects of *LTV* on mobility for *LTV*s between 0.80-0.90, 0.90-0.95, and greater than 0.95. This is done with a linear spline in *LTV* for *LTV*s of greater than 0.80 with knot points at 0.90 and 0.95. This results in a three-segment spline, with segments indexed by $j = 1,2,3$.

The following flexible functional form was specified

$$\begin{aligned} Z_{it}'\beta = X_{it}'\alpha + \sum_{j=1}^{3} \theta_j LTV_{it}^j + \gamma(D_{it}^{LOSS} \times LOSS_{it}) + \sum_{j=1}^{3} \phi_j (LTV_{it}^j \times D_{it}^{LOSS} \times LOSS_{it}) \\ + \delta(D_{it}^{GAIN} \times GAIN_{it}) + \sum_{j=1}^{3} \xi_j (LTV_{it}^j \times D_{it}^{GAIN} \times GAIN_{it}). \end{aligned} \tag{3}$$

The first term on the right-hand side of (3) includes *X*, a vector of explanatory variables. The second term includes the three-segment spline for high *LTV* households. The third term has two parts: D_{it}^{LOSS} is a dummy variable that is one if the household has experienced a nominal loss in house price in current period t relative to the initial purchase period 0, and $LOSS$ measures this nominal loss in percentage terms and expresses it as a *positive* number (e.g., for a five percent nominal loss, $LOSS$ takes on a value of 0.05). The fourth term is an interaction of the loss variable with the *LTV* spline.

The parameters on these variables can be interpreted as follows. Because desired leverage may be correlated with expected duration, initial *LTV* may have an effect on mobility *independent* of equity constraints. Hence, the parameters θ_j represent the effect of high initial *LTV* (low initial

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[16] In fact, for an individual purchasing a home with 15 percent down, the shadow rate of return on the additional down payment of 5 percent would be equal to the mortgage interest rate plus an additional 12 percentage points for PMI.

equity) on mobility. The ϕ_j parameters represent the differential effect of having had high *LTV* *and* experienced a nominal loss in housing value, and, hence, measure the impact of equity constraints on mobility.[17] Under the null hypothesis of no equity constraint, $\phi_1 = \phi_2 = \phi_3 = 0$. In addition, the equity constraint should become more binding as *LTV* rises, so that $\phi_3 < \phi_2 < \phi_1 < 0$ if the effect of the constraint is monotonic.[18] Furthermore, because the *LTV* spline only covers *LTV*s over 0.80, the parameter γ measures the effect of nominal losses for households with *LTV*s under 0.80, who *a priori* would not be expected to be equity constrained. Therefore, γ measures the pure effect of loss aversion on mobility. Under the null hypothesis of no loss aversion, $\gamma = 0$, versus the alternative of loss aversion, $\gamma < 0$.

Overall, the theoretical effect of gains on mobility is ambiguous. Under the equity constraints hypothesis, gains increase mobility due to leverage, as described above. However, when the per unit cost of housing rises, there may be substitution toward non-housing goods in consumption. Offsetting this is the wealth effect that may increase the demand for housing services (if, plausibly, housing is a normal good). An increase in the consumption of housing services can be achieved through a move or modification of the existing structure, through an expansion or renovation, for example (Potepan [42], Montgomery [39]). In particular, because of transaction costs of moving, it may be optimal for households not to move for relatively small gains.

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[17] Moreover, because the parameter ϕ varies by *LTV* category (i.e., by *j*), this effect is allowed to vary by the *size* of the *LTV* within the high *LTV* group.

[18] In principle, the effect of the nominal loss on low equity households may be non-linear if the loss is so large that the mortgage default option is in the money. This is not addressed in this paper because the cell sizes became small as loan-to-value rose above one, which resulted in unreliable estimation and inference in this sample. Moreover, there is an extensive literature on the effect of house price fluctuations on mortgage default, most of which uses data that are superior for default models than these data. Mattey and Wallace [34,35] review much of this literature and discuss the role of house price declines on mortgage defaults.

Consequently, the fifth term in (3) has two parts: D_{it}^{GAIN} is a dummy that is one if the household has experienced a nominal gain in house price in excess of moving costs in current period t relative to the initial purchase period 0, $GAIN$ measures this nominal gain in percentage terms (e.g., for a five percent nominal gain, $GAIN$ takes on a value of 0.05). I assume that transaction costs are 10 percent of the value of the home.[19] Therefore, $GAIN$ measures the nominal gain in excess of 10 percent. The final term is an interaction of the gain variable with the *LTV* spline. Therefore, the ξ_j parameters represent the differential effect of having had high *LTV* *and* experienced a nominal gain of more than 10 percent in housing value. Furthermore, because the *LTV* spline only covers *LTV*s over 0.80, the parameter δ measures the pure effect of nominal gains on mobility for households not expected to be equity constrained.

The vector X contains other characteristics of households that influence mobility. These variables include real income, income-squared, net worth, age, age-squared, education, and dummy variables for whether the household is married, became divorced (in the last year), became unemployed, and white, respectively. In addition, there are dummy variables for whether there are children of various ages: age 5 or under, 6 to 10, and 11 years and older, respectively. These dummies are particularly important. Households with school-age children are thought to be less mobile. Also, the equity-constraints hypothesis focuses on the ability of the household to make a down payment on a desired home. But households contemplating a move must also be able to meet the flow cost of housing services out of their income. Although the hypothesis does not address this directly, it is important that this be controlled for in the estimation. Therefore, a dummy

[19] The assumption of 10 percent is typical of transactions costs assumed and estimated in previous studies, including Linneman [32], Venti and Wise [54], Weinberg, Friedman, and Mayo [55], Goodman [16], Cunningham and Hendershott [8], Rosenthal [48], Malatesta and Hess [33], and Haurin and Gill [19], among others. The findings presented in Table 4 below were robust to the level of transaction costs chosen. In fact, if no transaction costs were assumed, so that the gain variable reflected the actual nominal gain, the parameter estimates for the gain variable and its interaction with the *LTV* spline were not statistically different from the ones presented.

variable that is one if the household has housing expenditure greater than one-third of income and zero otherwise appears. This variable is meant to capture the flow cost of housing services relative to income. This is measured as the user cost of owner-occupied housing multiplied by house value, then divided by income.[20] The construction of this measure is described in detail in the appendix. Finally, it is important that the gain and loss variables not just reflect local labor-market conditions (Chan [7]). For example, metropolitan areas with good local labor markets also have stable or rising nominal house prices. Thus, the specification includes the unemployment rate for the county of residence as an explanatory variable and a full set of calendar year and area fixed effects.[21]

V. ESTIMATION RESULTS

Column 1 of Table 4 shows the OLS estimates for the linear probability model described in equations (1) and (3) above. In this model, the loss and gain variables are measured using self-reported home values in the NLSY79. Heteroscedasticity-robust standard errors that account for the unbalanced nature of the panel are shown in parentheses. The estimate of the pure effect of losses on mobility is $\hat{\gamma}_{OLS} = 0.112$ and significantly different from zero. This suggests that losses *increase* rather than reduce mobility. When evaluated at the sample mean nominal loss of 5 percent, this estimate implies an increase in the likelihood of a move of 6.6 *percent*, *ceteris paribus*. This is shown at the bottom of the table. The estimate of the pure effect of gains on mobility is $\hat{\delta}_{OLS} = 0.097$ and significantly different from zero. This suggests that gains increase mobility. When evaluated at the sample mean nominal gain of 13.5 percent, this estimate implies an increase in the likelihood of a move of 14.2 *percent*, *ceteris paribus* (also shown at the bottom of the table).

[20] Different cutoffs defining constrained households were used and did not change the empirical findings. Overall, the results were quite robust to changes in the definition of this variable.

[21] The calendar year effects in the linear probability and hazard models control for general aggregate effects on mobility as well as factors that time-vary. One such factor is the mortgage interest rate which may affect mobility if households are locked into a low rate. See Quigley [46] and Green and Shoven [18], for example.

As expected, based on the $\hat{\theta}_{jOLS}$ estimates in column 1, higher initial *LTV* is associated with monotonically lower mobility. However, there is little evidence of binding equity constraints. Specifically, $\hat{\phi}_1$ and $\hat{\phi}_3$ are negative, as predicted, but $\hat{\phi}_2$ is positive, violating the prediction of monotonicity.

Unfortunately, there may be measurement error in the self-reported home values in the NLSY79 used to construct the gain and loss measures.[22] In addition, a careful examination of the NLSY79 data shows that reported house values were electronically miscoded from the interview information in some cases. The most typical coding error stems from the omission of the last digit of a reported value. For example, the same home worth \$100,000 in one year is coded as being worth only \$10,000 in the next year, followed by \$100,000 in the subsequent year. Naturally, this results in a large portion of the variation in self-reported home value across time to be due to measurement error. To the extent measurement error in home values is classical in nature, then OLS estimates of the parameters on the gain and loss variables (and their interactions with the *LTV* spline) in column 1 will be subject to the standard attenuation bias. Specifically, if loss aversion is present, and $\gamma < 0$, then the OLS estimate, $\hat{\gamma}$, will be biased upward, toward zero.

To address this, I use the average metropolitan area appreciation measured by the Freddie/Fannie indices from the year of purchase to the current period to construct gain and loss variables to be used as instruments. Descriptive statistics for these variables are shown in Table A-1 in the appendix. Under the assumption that each household's mobility decision does not affect average price performance in the housing market, variation in average metropolitan area appreciation is exogenous to a given household's mobility decision. Therefore, this instrument is

[22] Kish and Lansing [29], Kain and Quigley [27], Robins and West [47], Ihlanfeldt and Martinez-Vazquez [25], Goodman and Ittner [17], DiPasquale and Somerville [9], and Kiel and Zabel [28], among others, have studied measurement error in self-reported home values.

correlated with the self-reported home value but is plausibly uncorrelated with the household-level measurement error.[23]

Column 2 presents the IV estimates.[24] As expected, losses reduce mobility, as $\hat{\gamma}_{IV} = -0.759$ and is significantly different from zero. This suggests that measurement error biased the OLS estimates in column 1 upward significantly. When evaluated at the sample mean nominal loss of 5 percent, this estimate implies a reduction in the likelihood of a move of 44 *percent*, *ceteris paribus*. This is shown at the bottom of the table. Because the sample mean probability of moving (within a metropolitan area) is 0.09, this estimate implies that a 5 percent loss would reduce the probability of moving from 0.09 to 0.05 [i.e., 0.09 $\times$ (1-0.44)=0.04], or four *percentage points*. This is an economically large effect.[25] Gains reduce mobility, $\hat{\delta}_{IV} = -0.359$, but this effect is not statistically different from zero.

The parameter estimates on the interaction between the *LTV* spline and the loss variable, $\hat{\phi}_1, \hat{\phi}_2$ and $\hat{\phi}_3$, represent the effect of equity constraints on mobility. There is no evidence that households that purchased with low equity and then experienced nominal price declines had

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[23] The instrument varies both across metropolitan areas and across calendar time. The specification also controls independently for calendar year effects and area effects, so that the identifying variation in the instrument comes from time-series variation in housing market performance within metropolitan area. Importantly, because survey respondent reporting error in total mortgage debt and house value in the purchase year is unlikely because the household went through the mortgage application and closing process prior to the NLSY79 interview for that year, I assume that the initial *LTV* is measured without error. However, Kiel and Zabel [28] found that homeowners tend to increase their self-assessment of the value of the property just after home purchase thinking that they got a "good deal." This would tend to push some truly high *LTV* households into lower *LTV* categories according to my measure of initial *LTV*.

[24] The standard errors are clustered to account for the fact that the instrument only varies across, and not within, metropolitan area for a given calendar year.

[25] Alternate measures of wealth were used to check the robustness of the estimates for this specification. These measures included non-housing wealth, financial wealth, financial assets, and highly liquid assets. These alternative specifications produced economically very similar estimates to those shown in Table 4. In addition, specifications were estimated that used the measure of "extended *LTV*" from Chan [7], defined as loan balance less other assets, divided by house value. The results were similar in economic magnitude and statistical significance to those shown here. Furthermore, a number of alternative measures of the housing expenditure variable were specified. These are described in detail in the appendix. Again, all of these specifications produced estimates quantitatively similar to those presented here. These specifications are available upon request. Overall, these findings were very robust to alternative specifications. Finally, quadratic terms in the loss and gain variables were added to the model, but the null hypothesis of linearity in each could not be rejected at conventional significance levels.

reduced mobility. In fact, $\hat{\phi}_2$ and $\hat{\phi}_3$ are positive but not statistically significant. However, the cell sizes are relatively small: only 8.8 percent of the sample experienced a loss and had initial *LTV* greater than 0.80.[26] So, these estimates may be somewhat imprecise. In addition, as expected, households with pre-school-age children have higher mobility. Given the importance of the quality of public schools in location decisions in a metropolitan area, these households may be moving to get into a suitable school district. Recently divorced and older households have higher and lower probabilities of moving to another owner-occupied house in the same metropolitan area, respectively. Income raises mobility, but the effect is hump-shaped.

Column 3 shows the estimates from the reduced-form model in which the instruments are substituted directly for the self-reported gain and loss variables and the model is estimated by least squares. Again, there is evidence in favor of loss aversion. When evaluated at the mean loss, the parameter estimate, $\hat{\gamma}_{RF} = -0.53$, implies a reduction in mobility of 31 percent. Gains reduce mobility, $\hat{\delta}_{RF} = -0.075$, but this effect is only marginally statistically different from zero. When evaluated at the mean gain, this parameter estimate implies a reduction in mobility of 11 percent. Again, based on the $\hat{\phi}_{jRF}$ estimates, there is no evidence in favor of equity constraints; in fact, $\hat{\phi}_2$ and $\hat{\phi}_3$ are positive but not statistically different than zero.

One advantage of these linear probability models is that it is easy to circumvent the measurement error in gain and loss variables using instrumental variables. However, a disadvantage is that they assume the disturbance term is normally distributed. To examine the robustness of the findings to the estimator and parametric assumptions, I present estimates of the

[26] This 8.8 percent of the sample can be decomposed as follows: 3.9 percent experienced a loss and had initial *LTV* of 0.80-0.90, 1.7 percent experienced a loss and had initial *LTV* of 0.90-0.95, and 3.2 percent experienced a loss and had initial *LTV* of greater than 0.95.

proportional hazard model in column 4.[27] Specifically, the semi-parametric estimator of Prentice and Gloeckler [45] and Meyer [37,38] is used to estimate the parameters in equations (2) and (3). With this estimator, the baseline hazard is modeled flexibly as a vector of dummy variables, one for each spell period. Because instrumental variable procedures for hazard models are not well developed in the econometrics literature, estimates for the reduced-form model are displayed in column 4, where, again, the instruments are substituted directly for the self-reported gain and loss variables.[28] Hence, the estimates in column 4 are most comparable with those in column 3.

For the proportional hazard model, $\hat{\gamma}_{PH} = -6.506$ and is significantly different than zero. Because of the proportional hazard specification, this parameter estimate implies that nominal losses shift the baseline hazard downward by 29.6 percent when evaluated at the sample mean loss of 5 percent. This is shown at the bottom of the table. Specifically, the estimated percentage shift in the baseline hazard was calculated as

$$[\exp(\hat{\gamma}_{PH} \times \overline{LOSS}^{L})] - 1, \tag{4}$$

where $\overline{LOSS}^{L}$ is the sample mean loss for the sub-sample with losses, i.e., $D^{LOSS} = 1$. This estimate is very similar to the OLS reduced form estimate of –31 percent in column 3, which is suggestive that the OLS reduced-form and IV estimates are robust. As with the other models, there is no evidence in favor of binding equity constraints: $\hat{\phi}_2$ and $\hat{\phi}_3$ are actually positive but not statistically significant. Finally, $\hat{\delta}_{PH} = -0.095$ but is not significantly different than zero. This parameter estimate implies that nominal gains shift the baseline hazard upward by 14 percent when evaluated at the sample mean gain of 13.5 percent (also shown at the bottom).

[27] Also, note that all specifications in column 1-3 in Table 4 include a full set of spell period dummies, so that the linear probability estimates are interpreted as the probability of moving conditional on having lived in the house for that given number of periods. Hence, the estimates are comparable to those in column 4 for the hazard model.

[28] That is, simple two-stage methods, akin to two-stage least squares, do not exist for hazard models.

To sum up, across the specifications in columns 2-4 in Table 4 there is no evidence of binding equity constraints. In addition, there are mixed results for the effect of gains on mobility. While some specifications indicate that gains raise mobility, others indicate the opposite, and in none of the specifications is the estimate of the effect of nominal gains, $\hat{\delta}$, statistically different from zero at standard significance levels. There are, however, economically large and statistically significant effects of nominal losses on mobility. The reduced-form and IV estimates suggest that a 5 percent nominal loss is associated with a 30-44 percent reduction in the probability of a move.

VI. CONCLUSION

The results in this paper suggest that loss aversion is an important phenomenon in metropolitan housing markets in the United States. There are some clear directions for housing research. First, like Barberis, Huang, and Santos [3] have done for the stock market, a complete theoretical model of the housing market with loss averse agents is needed in order to better understand price and volume dynamics and equilibrium. Second, the presence of loss aversion in marketing (Genesove and Mayer [14]) and mobility decisions (this paper) should be reconciled with the mortgage prepayment and default behavior of households with losses. Finally, loss aversion should have some fundamental implications for the demand for and pricing of home equity insurance products, options, and futures (Case and Shiller [5], Shiller and Weiss [49,50]).

APPENDIX

This appendix describes the construction of the data set. The primary data cover the 1985-96 period and are from the 1985-1998 waves of the National Longitudinal Survey of Youth (NLSY79). Engelhardt [11] and Zagorsky [56] discussed the quality of these data. Haurin, Hendershott, and Kim [20] and Haurin, Hendershott, and Wachter [21,22] have used these data to analyze housing decisions of young households. The NLSY79 started as a national, stratified, random sample of 14-21 year olds in 1979. The survey was conducted every year from 1979-1994; after 1994, it was conducted every two years. It asked detailed questions about education, employment, income, home ownership, family background, etc. Since 1985, questions about assets and debts were asked, including mortgage debt and home value. Because home ownership status has been asked each year since 1985, it is possible to completely track residence transitions from early adulthood. As a result, the sample has no left-censored spells. This means that the earliest a spell could have begun and been included in the sample is 1985. In turn, the longest spell observable with these data is 12 years. A potential criticism of this sample is that the distribution of completed residence spells may be poorly estimated if the average spell length of homeowners of this age is greater than 12 years, the maximum spell length in the sample. Sinai [52] studied housing mobility for homeowners of all ages in the 1970-91 waves of the Panel Study of Income Dynamics (PSID). His analysis showed that the hazards for homeowners declined steadily for spells of 10 years or less. For spells greater than 10 years, the hazards were roughly flat for all types of transitions. In addition, he estimated the average duration for a homeowner as 6.8 years. Under the assumption that young homeowners (age 20-41 in my sample) have shorter completed spells than the average-aged homeowner in Sinai's analysis---which seems plausible---then the NLSY79 sample may estimate the underlying spell distribution well. In addition, Sinai presented sensitivity analyses that showed that truncation of the spell length at 8 years had little effect on the hazard estimates.

Mobility Histories – Generally speaking, there are three possible transitions: intra-metropolitan moves from own-to-own, intra-metropolitan moves from own-to-rent, and inter-metropolitan moves. This study focuses on intra-metropolitan own-to-own mobility. Unlike other panel studies, such as the *Panel Study of Income Dynamics* (*PSID*), the NLSY79 did not ask a question each year about whether the respondent had moved since the previous interview. The public-use version of the NLSY79 has information on home ownership in each year. Hence all own-to-rent moves can be tracked in this data set. The restricted-access Geo-code data set gives information on state, county, and the metropolitan area of residence. When combined, the public-use and Geo-code data can track own-to-rent and inter-county moves. Unfortunately, intra-county own-to-own moves cannot be tracked. Because most moves are local, and most metropolitan areas are comprised of just a few (and, in some cases, one) counties, the combined public-use and Geo-code data significantly understate the number of actual transitions. To overcome this problem, I obtained permission from the United States Department of Labor, Bureau of Labor Statistics, to use administrative address data on the NLSY79 respondents in each survey year to construct a mobility history for each respondent. These mobility data were provided graciously by Patricia Reagan, who assembled them at the Center for Human Resources Research at the Ohio State University. When comparing the county of residence from the address records to that in the Geo-code file, a number of errors were found in the Geo-code data. All state and county codes used in this study were based on the administrative address records.

Metropolitan House Price Indices - Because the empirical analysis focuses on the effect of house prices on mobility, only respondents in metropolitan areas with available house price information were included in the sample. Metropolitan house prices were measured by the Fannie Mae/Freddie Mac Weighted Repeat Sales Price index. This index is discussed in detail in Abraham and Hendershott [1]. This is available in the 1985-96 period for 149 metropolitan areas (the list of which is available upon request).

Income – The household income measure used is real total net family income (in 1993 dollars). Interviews typically were conducted in the spring of the calendar year. The survey asked about income earned in the previous calendar year. For example, the 1993 wave contains information on 1992 income. This means that the 1986-1994 surveys provided information on income in calendar years 1985-1993. After 1994, the survey went to an every-other-year format, but questions on income still referred to the previous calendar year. This means that the 1996 survey year gives information on calendar year 1995 income, and the 1998 survey year gives information on calendar year 1997 income. Incomes from calendar years 1994 and 1996 were not asked. Therefore, for this study, income for calendar years 1995 and 1997 proxy for those in 1994 and 1996, respectively. All income values in the paper are in real 1993 dollars, deflated by the All-Items CPI.

House Value and Mortgage Data – The public use NLSY79 top-coded housing value and mortgage debt at $150,000 in nominal terms for the 1985-94 waves (Engelhardt [11]). Initial loan-to-value cannot be calculated for observations with top-coded values, and these observations must be excluded from the sample. In 1985, less than 2 percent of observations had top-coded values for house value and mortgage debt. But because the top-code threshold was fixed in *nominal* terms, over time with inflation, a growing fraction of observations had top-coded values: 17.38 percent for housing value and 6.35 percent for mortgage debt in 1994, respectively. However, because of the substantial regional variation in house price levels, the truncated cases came disproportionately from high-cost markets, such as Boston, San Francisco, Los Angeles, New York, etc. Furthermore, these markets were the ones that experienced steep declines in nominal house prices in the late 1980s and early 1990s. Therefore, the exclusion of observations with top-coded values results in differential sample selection and potentially biased estimated equity and loss aversion effects. To overcome this problem, I obtained permission from the United States Department of Labor, Bureau of Labor Statistics, to use housing value and mortgage data from the NLSY79 master file at the Center for Human Resources Research (CHRR) at the Ohio State University. Specifically, CHRR released to me new house value and mortgage data in which the top-code thresholds were adjusted upward so that only approximately 2 percent of the observations in each of the 1985-1998 survey years had top-coded house values and mortgage amounts. These new data were used in the empirical analysis. All asset and debt values in the paper are in real 1993 dollars, deflated by the All-Items CPI.

Assets and Debts – For budgetary reasons, questions on assets and debts were not asked in the 1991 wave of the NLSY79 (Engelhardt [11]). However, income from assets in 1991 was asked in the 1992 wave, and asset and debt questions were asked in the 1990 and 1992 waves. So, for 1991, the asset income was capitalized at the prevailing annual return. This along with information from 1990 and 1992 was used to impute assets and debts for each household in 1991. The empirical results were not sensitive to the exclusion of all 1991 observations.

Dummy if Housing Expenditure Greater than One-third – Let i index households and t index calendar years, then, following Poterba [44], whether or not the household claims itemized deductions for mortgage interest and property taxes paid depends on the tax saving from itemizing, ξ,

$$\xi_{it} = \theta_{it}[\tau^s_{it} + (\tau^p + i_0 LTV_{it})V_{it}] - S_{it}, \tag{A1}$$

where τ^p is the property tax rate, i_0 is the nominal mortgage interest rate in the year of purchase, LTV is the loan-to-value rate, and S is the standard deduction amount. V is house value. θ is the household's federal marginal tax rate on the first dollar of itemized deduction. τ^s is the household's state marginal tax rate on the first dollar of itemized deduction. If $\xi_{it} \geq 0$, then the household will itemize and the marginal user cost of owner-occupied housing (as a fraction of the house price) is

$$u^m_{it} = (1-\theta_{it})(i_t + \tau^s_{it} + \tau^p) + d + a + m - \pi^e_t, \tag{A2}$$

where d is the physical rate of decay, m is maintenance expenditure, a is a risk factor, and π^e is expected appreciation. If $\xi_{it} < 0$, then the household will not itemize and the marginal user cost is

$$u^m_{it} = ((1-LTV_{it})(1-\theta_{it}) + LTV_{it})i_t + \tau^p + d + a + m - \pi^e_t. \tag{A3}$$

Following Poterba [43,44], the user cost is calibrated for each household under the following assumptions: $\tau^p = 0.02$; $d = 0.014$; $a = 0.05$; and, i is the rate on a 30-year fixed rate mortgage. The federal and state tax first-dollar marginal tax rates were calculated using the NBER TAXSIM calculator. The dummy if housing expenditure greater than one-third was constructed to take on a value of 1 if the flow cost of housing relative to income,

$$\frac{u^m_{it} V^*_{it}}{y_{it}}, \tag{A4}$$

(where y is household income) was greater than 0.33 and zero otherwise. A key assumption in calculating (A4) is what *LTV* to use in (A1) and (A3). Variants of this variable were constructed using an *LTV* of 0.80, 0.90, 0.95, as well as contemporaneous *LTV*; the estimation results were remarkably robust across these alternative specifications. The results in the Tables 4-7 in the paper used an *LTV* of 0.95, which is akin to assuming that the household would take out a mortgage on the next home with just 5 percent down. Such mortgages were prevalent by the end of the sample period. This assumption helps insure that the variable really picks up expenditure-constrained households, for these would be households that could not buy back their current residence and spend less than one-third of their income on housing. Variants of this variable were also constructed using flow-cost-of-housing-to-income ratios of 0.25, 0.28, 0.30, and 0.40 to define the dummy. Again, the estimation results were quite robust to these alternatives.

Table A-1
Sample Means (Standard Deviations)
for the Explanatory and
Instrumental Variables

A. Explanatory Variables	
Dummy if Nominal Loss ×Nominal Loss	0.123 (0.165)
Dummy if Nominal Loss ×Nominal Loss × Initial LTV>.95	0.988 (0.017)
Dummy if Nominal Loss ×Nominal Loss × Initial LTV .90-.95	0.923 (0.015)
Dummy if Nominal Loss ×Nominal Loss × Initial LTV .80-.90	0.845 (0.030)
Initial LTV>.95	0.988 (0.017)
Initial LTV .90-.95	0.926 (0.148)
Initial LTV .80-.90	0.851 (0.031)
Dummy if Nominal Gain ×Nominal Gain	0.270 (0.262)
Dummy if Nominal Gain ×Nominal Gain × Initial LTV>.95	0.990 (0.015)
Dummy if Nominal Gain ×Nominal Gain × Initial LTV .90-.95	0.925 (0.015)
Dummy if Nominal Gain ×Nominal Gain × Initial LTV .80-.90	0.851 (0.033)
Dummy if Married	0.799

Table A-1 (Continued)

Dummy if White	0.818
Age	30.6 (3.48)
Age-Squared	948.5 (211.6)
Education	13.9 (2.2)
Dummy if Children Age 5 and Under	0.432
Dummy if Children Age 6 to 10	0.330
Dummy if Children Age 11 to 18	0.185
Dummy if Children Over Age 18	0.021
Dummy if Became Unemployed	0.172
Dummy if Became Divorced	0.067
Dummy if Housing Expenditure Greater than One-Third	0.195
Real Income	62.173 (98.336)
Real Net Worth	67.736 (74.364)
County Unemployment Rate	6.23 (2.46)

Table A-1 (Continued)

B. Instrumental Variables	
Dummy if Nominal Loss ×Nominal Loss	0.054 (0.046)
Dummy if Nominal Loss ×Nominal Loss × Initial LTV>.95	0.991 (0.015)
Dummy if Nominal Loss ×Nominal Loss × Initial LTV .90-.95	0.928 (0.013)
Dummy if Nominal Loss ×Nominal Loss × Initial LTV .80-.90	0.852 (0.032)
Dummy if Nominal Gain ×Nominal Gain	0.135 (0.131)
Dummy if Nominal Gain ×Nominal Gain × Initial LTV>.95	0.988 (0.017)
Dummy if Nominal Gain ×Nominal Gain × Initial LTV .90-.95	0.925 (0.015)
Dummy if Nominal Gain ×Nominal Gain × Initial LTV .80-.90	0.851 (0.031)
Number of Observations	6,461

Note: Sample means of the explanatory and instrumental variables used in Table 4, with standard deviations for all continuous variables shown in parentheses. Income and net worth are in thousands of 1993 dollars. The county unemployment rate is measured in percentage points. The sample means for the gain and loss variables and the LTV variables are calculated for the sub-sample of observations with gains and losses, respectively, by LTV category.

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Table 1
Percent of Wealth in Housing at First Home Purchase

Wealth Measure	(1) Mean	(2) 25th Percentile	(3) Median	(4) 75th Percentile
Liquid Assets	80.5	68.9	90.6	99.0
Liquid Assets Less Debts	83.6	72.7	94.4	100.0

Note: Author's calculations from the sample of all first-time homebuyers in the 1985-90 waves of the NLSY79. Total liquid assets is the amount of the down payment plus the value of financial assets in the form of savings accounts, money market deposit accounts, certificates of deposit, interest-earning checking accounts, U.S. saving bonds, individual retirement accounts (IRAs), 401(k)-type pension arrangements, Keogh plans, mortgages held by the household, non-interest earning checking accounts, money market funds, U.S. government securities, municipal and corporate bonds, stocks and mutual fund shares, money owed by others to the household, and other interest-earning assets. Debts are the sum of farm, business, and non-owner occupied real estate debt and non-vehicle related debt.

Table 2
Episodes of Falling Nominal House Prices for Selected Metropolitan Areas

Metropolitan Area	(1) Market Peak (Yr:Qtr)	(2) Market Trough (Yr:Qtr)	(3) Peak-to-Trough Total Decline (%)	(4) Peak-to-Trough Annual Average Decline (%)
Houston, TX	83:2	87:4	27.2	5.5
Dallas, TX	86:2	89:1	14.4	5.0
Austin, TX	86:2	88:4	26.9	10.0
San Antonio, TX	86:1	90:2	20.0	4.4
Oklahoma City, OK	86:2	88:3	22.7	9.5
Tulsa, OK	83:3	89:1	15.3	2.6
New Orleans, LA	86:2	88:4	11.9	4.6
Baton Rouge, LA	86:1	89:1	14.1	4.5
Denver, CO	86:2	89:1	6.8	2.4
Boston, MA	88:4	92:2	9.7	2.7
Portsmouth, NH	89:1	92:3	15.3	4.2
Providence, RI	89:4	94:4	11.8	2.3
Hartford, CT	88:3	95:1	19.7	2.8
New Haven, CT	88:2	95:1	20.8	2.8
New York, NY	89:1	91:3	6.4	2.5
Nassau-Suffolk, NY	88:3	91:2	9.0	3.2
Middlesex-Somerset, NJ	88:2	91:3	10.5	3.1
Bergen-Passaic, NJ	88:2	91:3	10.0	3.0
Los Angeles, CA	90:1	95:1	21.5	4.0
Orange County, CA	90:1	95:1	17.9	3.4
Riverside, CA	91:1	95:1	19.1	4.5
Santa Barbara, CA	90:3	95:1	12.4	2.6
San Diego, CA	90:3	95:1	10.2	2.2
San Francisco, CA	90:1	94:4	11.0	2.2
San Jose, CA	89:4	94:3	12.2	2.4

Note: Author's calculations using Freddie Mac/Fannie Mae weighted repeat sales quarterly house price indices for each of the metropolitan areas shown.

Table 3
The Effect of Falling Nominal House Prices on the
Distribution of Housing Equity and Net Worth

	(1)	(2)	(3)	(4)	(5)	(6)
	Percent in Cell					
	All Observations		Observations in Stable or Rising Markets		Observations in Falling Markets	
Category	Purchase Year	Current Year	Purchase Year	Current Year	Purchase Year	Current Year
A. Housing Equity as a Percent of Home Value						
20% or more	55.2	63.1	54.8	64.3	59.0	52.1
10-19%	20.5	23.5	21.2	17.7	14.0	15.5
Less than 10%	24.3	13.4	24.0	18.0	27.0	32.4
B. Net Worth as a Percent of Home Value						
20% or more	81.4	84.8	81.8	85.6	77.1	77.8
10-19%	10.7	8.0	10.7	7.7	10.9	10.9
Less than 10%	7.9	7.2	7.5	6.7	12.0	11.3

Note: Author’s calculations from the sample of 6,461 household-year observations described in the text.

Table 4. Parameter Estimates of Intra-Metropolitan Own-to-Own Mobility

	(1)	(2)	(3)	(4)
	Estimator			
Explanatory Variable (Parameter)	OLS	IV	OLS Reduced Form	Proportional Hazard Reduced Form
Dummy if Nominal Loss × Nominal Loss (γ)	0.112 (0.069)	-0.759 (0.244)	-0.530 (0.132)	-6.506 (3.178)
Dummy if Nominal Loss × Nominal Loss × Initial LTV >.95 (ϕ_3)	-10.213 (4.351)	17.219 (20.548)	10.886 (8.776)	101.115 (208.032)
Dummy if Nominal Loss × Nominal Loss × Initial LTV .90-.95 (ϕ_2)	15.102 (4.278)	6.500 (17.304)	5.810 (12.003)	100.436 (291.091)
Dummy if Nominal Loss × Nominal Loss × Initial LTV .80-.90 (ϕ_1)	-3.824 (1.247)	-1.627 (5.332)	-0.426 (4.156)	-10.627 (105.228)
Initial LTV >.95 (θ_1)	-0.254 (0.389)	0.185 (0.627)	-0.216 (0.368)	-1.494 (5.524)
Initial LTV .90-.95 (θ_2)	-0.169 (0.409)	-0.535 (0.627)	-0.110 (0.368)	-2.664 (5.741)
Initial LTV .80-.90 (θ_3)	-0.0002 (0.132)	-0.065 (0.231)	-0.078 (0.144)	-0.447 (1.844)
Dummy if Nominal Gain × Nominal Gain (δ)	0.097 (0.050)	-0.359 (0.340)	-0.075 (0.051)	0.095 (0.154)
Dummy if Nominal Gain × Nominal Gain × Initial LTV >.95 (ξ_3)	5.774 (4.697)	-36.554 (26.269)	-1.049 (0.876)	-19.410 (13.181)
Dummy if Nominal Gain × Nominal Gain × Initial LTV .90-.95 (ξ_2)	-1.978 (3.965)	21.331 (27.551)	0.841 (1.102)	12.600 (10.758)

Table 4 (Continued)

Explanatory Variable	(1)	(2)	(3)	(4)
Dummy if Nominal Gain × Nominal Gain × Initial LTV .80-.90 (ξ_1)	-1.262 (0.989)	-4.519 (6.138)	-0.173 (0.266)	-3.108 (3.364)
Dummy if Married	0.003 (0.011)	-0.005 (0.013)	0.002 (0.011)	0.044 (0.144)
Dummy if White	0.001 (0.009)	-0.007 (0.012)	0.0007 (0.010)	-0.018 (0.118)
Age	-0.037 (0.015)	-0.047 (0.177)	-0.035 (0.014)	-0.628 (0.145)
Age Squared	0.0006 (0.0002)	0.0007 (0.0003)	0.0005 (0.0002)	0.010 (0.002)
Education	0.0015 (0.0018)	-0.0004 (0.011)	0.0012 (0.0015)	0.015 (0.021)
Dummy if Children Age 5 and Under	0.016 (0.008)	0.016 (0.010)	0.017 (0.009)	0.204 (0.090)
Dummy if Children Age 6 to 10	-0.005 (0.008)	-0.010 (0.009)	-0.005 (0.008)	-0.068 (0.097)
Dummy if Children Age 11 and Over	0.0006 (0.010)	-0.0003 (0.002)	0.0009 (0.010)	-0.014 (0.122)
Dummy if Became Unemployed	0.004 (0.009)	0.008 (0.011)	0.005 (0.010)	0.043 (0.110)
Dummy if Became Divorced	0.042 (0.018)	0.032 (0.018)	0.041 (0.017)	0.491 (0.193)
Dummy if Housing Expenditure Greater than One-third	-0.007 (0.010)	-0.009 (0.011)	-0.007 (0.009)	-0.132 (0.122)
Real Income	0.00031 (0.00016)	0.00020 (0.00017)	0.00027 (0.00014)	0.0022 (0.0016)
Real Income Squared	-0.00000013 (0.00000015)	-0.00000023 (0.00000016)	-0.00000028 (0.00000014)	-0.000000002 (0.000000001)

Table 4 (Continued)

Explanatory Variable	(1)	(2)	(3)	(4)
Real Net Worth	-0.00002	0.00006	-0.000000014	0.000058
	(0.00006)	(0.0001)	(0.0000053)	(0.00065)
County Unemployment Rate	0.0002	-0.0003	0.0001	0.0094
	(0.0017)	(0.0021)	(0.0018)	(0.0223)
R^2	0.106	---	0.104	---
Log Likelihood	---	---	---	-1646.9
Estimated Percentage Change in the Probability of a Move due to Nominal Loss	6.6	-44.0	-31.0	-29.6
Estimated Percentage Change in the Probability of a Move due to Nominal Gain	14.2	-52.5	-11.0	14.0

Note: The dependent variable in columns (1)-(3) is a dummy for whether the household moves. Columns (1) and (3) present OLS estimates. Column (2) presents instrumental variable estimates where the Fannie Mae/Freddie Mac indices are used to instrument for self-reported house price changes as described in the text. The standard errors, shown in parentheses, are heteroscedasticity-robust and account for the fact that the instrumental variables only vary across metropolitan area. Column (4) presents Prentice-Gloeckler-Meyer semi-parametric proportional hazard model estimates. All estimates were calculated on 6,461 household-year observations that comprised 3,112 residence spells, 2,348 households, and had 596 moves. All specifications were estimated with a full set of calendar year, spell period, and area fixed effects.