House Prices and Credit Constraints:

Making Sense of the U.S. Experience

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Abstract

Inverting the effective demand for housing implies home prices depend on credit constraints (Meen (2001), Muellbauer and Murphy (1997) and Cameron, Muellbauer and Murphy (2006)), a theoretical result also demonstrated in Kim's (2007) home price-to-rent framework. Previous U.S. home price models lack data on credit constraints facing first-time home-buyers (and regional housing stocks), likely accounting for the poor performance of home price models based on interest rates and income (Gallin, 2006). We incorporate such omitted data into home price models which yield stable long-run relationships, more precisely estimated income and interest rate coefficients, reasonable speeds of adjustment, and improved model fits.

JEL Codes: R31, G21, E51, C51, C52.
Key Words: home prices, credit standards, subprime mortgages

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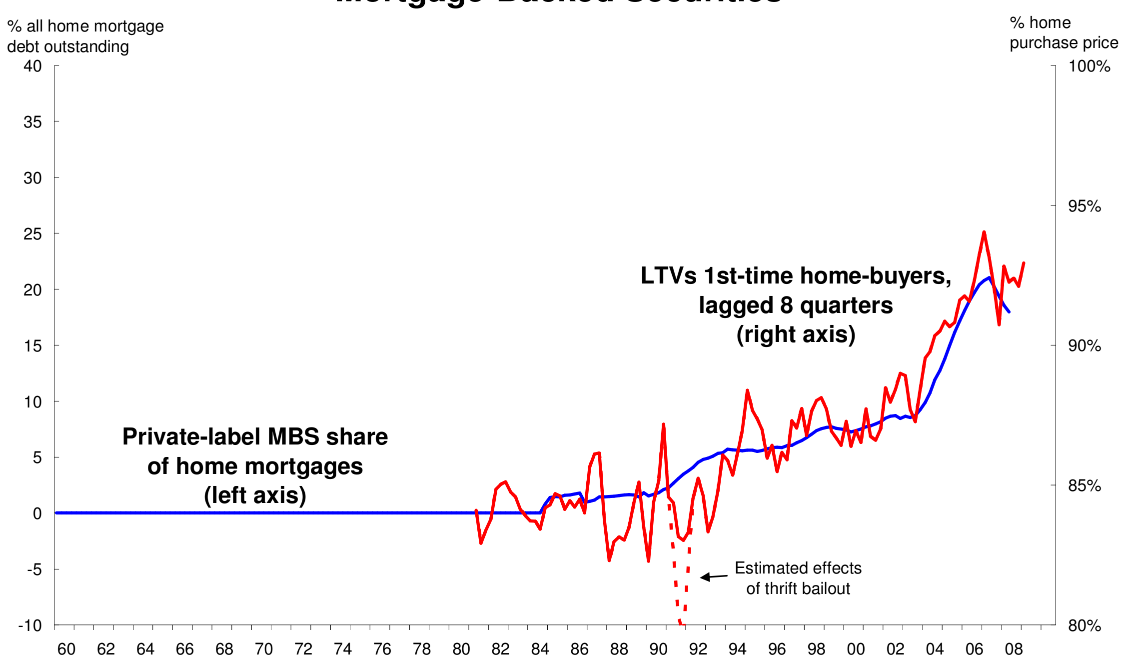
The recent boom and bust in U.S. housing markets has sparked interest in modeling the links between home prices and credit standards. As Meen (2001), Muellbauer and Murphy (1997) and Cameron, Muellbauer and Murphy (2006) stress, inverting the effective demand for housing services implies that home prices are a function of credit constraints, as well as income, the housing stock, and real user capital costs. Kim (2007) shows theoretically that down-payment or loan-to-value (LTV) constraints also help determine home prices using the price-to-rent approach to modeling home prices. Models of U.S. home prices have been hindered by a lack of consistent time series measures of the credit constraints facing the marginal, first-time home-buyer and—at least at the regional level—of the housing stock. This shortcoming suggests that estimates of U.S. home price and consumption models suffer from omitted variable bias.

This issue is addressed using Duca, Johnson, and Muellbauer's (2009) data on average LTV ratios over 1979-2007 for first-time home-buyers, the marginal buyers most likely affected by down-payment constraints. Derived from the American Housing Survey, this series implies that down-payment constraints were eased early this decade, in line with Doms and Krainer's (2007) finding that homeownership rates rose among the young. Consistent with the subprime boom reflecting a weakening of credit standards, LTVs for the young are positively correlated with the share of mortgages outstanding that were securitized into private-label mortgage-backed securities (MBSs, Figure 1). MBSs issued by Fannie Mae and Freddie Mac were generally conforming loans, which met credit standards generally applied to most mortgages in earlier years. In contrast, nonprime mortgages were mainly funded by being packaged into private label" MBSs because they did not conform to Fannie Mae or Freddie Mac standards, or were too risky to be held by regulated depositories (Credit Suisse, 2007). Because our LTV series reflects loan originations, it leads the private MBS share of home mortgages by roughly two years.

The rise in LTV ratios from 2000 to 2005 likely reflects two financial innovations: the adoption of credit scoring technology that enabled the sorting and pricing of nonprime mortgages and funding of such loans in the form of collateralized debt obligations (CDOs) or with protection from credit default swaps (CDSs). The later failure of CDOs to protect investors from unanticipated default losses and the soaring cost of using CDSs induced a reversal of the earlier easing of credit standards. Abstracting from the 8-quarter lag, LTVs peak in 2005q2, and are modestly lower through 2007:q2, before subprime difficulties kicked off the financial crisis that started in August 2007. Reports imply that mortgage standards have tightened further, which will likely be confirmed after we update the series after the 2009 AHS data are released.

These events occur after the end of our LTV series. Abstracting from the 8-quarter lag, the peak in the LTV series is in 2005q2, with a modestly lower level through 2007:q2, just before subprime loan losses induced the financial crisis that began in August 2007. Industry

Figure 1: Share of Mortgages Packaged into NonConforming Mortgage-Backed Securities



reports that mortgage credit standards have tightened further imply that our LTV series will likely decline much, once we can update the series with the release of the 2009 AHS data base.

Including LTV data on first time home-buyers notably improves home price models by yielding stable long-run relationships, sensible and more precisely estimated income and user cost coefficients, reasonable speeds of adjustment, and better model fits. This is true even before the post-2001 subprime explosion raised LTVs and appears to reflect an earlier, more modest rise in LTV ratios enabling us to identify such an effect in pre-2002 samples. Before including LTV data in our models, we regressed them on variables to remove the estimated effects of cyclical and other variables, such as changes in the overall unemployment. In a related study, we find that adding data on LTVs and regional housing stocks qualitatively improves regional home price models, paralleling the UK regional results of Cameron, Muellbauer and Murphy (2006).

This study is organized as follows. Section 2 presents the models and the data. These, in turn, are estimated using cointegration methods in section three and more general, one-stage models in section four where models are less constrained by practical aspects of using cointegration techniques. The fifth section forecasts to what extent U.S. home prices are over-valued. The conclusion discusses the links between credit and asset market bubbles.

II. House Price Models and Data

(a) House Price Models Using the Inverted Demand Approach

Perhaps the simplest theory of what determines house prices is to treat supply—the stock of houses—as given in the short run, with prices driven by the inverted demand for housing services (h) that are proportional to the housing stock (hs).[1] Let log housing demand be given by

$$\log h = -\alpha \log hp + \beta \log y + z \tag{1}$$

[1] Inverse demand functions have a long history, particularly in the analysis of markets for natural resources. Theil (1976) refers to a 1909 Danish study as the first empirical study of inverse demand functions.

where hp = real house price, y = real income and z = other demand shifters. The own price elasticity of demand is -α and the income elasticity is β. Solving yields

$$\log hp = (\beta \log y - \log h + z) / \alpha \tag{2}$$

Reasonable priors for the values of the key long run elasticities are the "central estimates" set out in Meen (2001) and Meen and Andrews (1998), *inter alia*. For example, many estimates of the income elasticity of demand suggest that β is in the region of 1, in which case the income and housing stock terms in above equation simplify to log income per house, i.e., log y -log h.

The demand shifters included in z cover a range of other drivers. Since housing is a durable good, inter-temporal considerations imply that expected or 'permanent' income and 'user cost' are important drivers. The user cost takes into account that durable goods deteriorate, but may appreciate in price and incur an interest cost of financing as well as tax. The usual approximation is that the real user cost is $uc = hp\left(r + \delta + t - h\dot{p}^e / hp\right)$, where r is the real after-tax interest rate of borrowing, possibly adjusted for risk, δ is the depreciation rate, t is the property tax rate, and $h\dot{p}^e / hp$ is the expected real rate of capital appreciation.

Ex-post user costs can be negative if appreciation rates in house price booms exceed nominal user costs. An important issue is how to track expectations of house price appreciation. Many studies find that lagged rates of appreciation are good proxy, suggesting an extrapolative element in household expectations. Our real user cost measure (*RUSER*) uses the lagged annual rate of appreciation in the US house price index over the prior 4 years. Given assumptions on transactions costs, *RUSER* is always positive making log *RUSER* defined over the sample. The log transformation implies that at low values, variations in *RUSER* have a more powerful effect than at high values, reflecting the idea that when appreciation is high relative to tax and interest costs, the market gets into a 'frenzied' state. Hendry (1984) and Muellbauer and Murphy (1997) capture similar effects using a cubic in the recent or fitted rate of appreciation. In results not shown, we found that models using log (*RUSER*) and models linear in *RUSER* but which include a cubic in lagged appreciation, yield similar long run solutions and adjustment speeds.

Other factors could be relevant, given that many mortgage borrowers face limits on their borrowing and may be risk averse. These could include nominal as well as real interest rates, credit supply conditions, demography, and proxies for risk, particularly of mortgage default.

In the dynamics, lagged price adjustment is plausible, given the inefficiency of house prices.[2] The rate of change in the housing stock relative to the population, as well as the per capita stock, are also likely relevant in helping explain house price movements. One interpretation is through expectations: households observing much new construction might lower expectations of future appreciation. Another interpretation is in terms of prices adjusting both to stock and flow disequilibria, for which error or equilibrium correction models are well suited.

(b) Models Using the House Price-to-Rent Ratio Approach

Home prices have been modeled using the price-to-rent approach, especially in the U.S., where regional housing stock measures are not readily available and rents are market-determined, in contrast to the UK. This approach is more grounded in finance and assumes that, absent substantial frictions and credit restrictions, arbitrage between owner-occupied and rental housing markets implies the house rent-to-price ratio depends on the real user cost of capital, defined as the nominal user cost of mortgage finance (*NOMUSER*) minus expected appreciation:

$$RENT/HP = NOMUSER - \text{(expected home price appreciation)} = RUSER, \qquad (3)$$

where *NOMUSER* is tax-adjusted and can reflect tax effects on rents relative to home prices. As

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[2] Hamilton and Schwab (1985), Case and Shiller (1989, 1990), Poterba (1991) and Meese and Wallace (1994) find that house price changes are positively correlated and past information on housing fundamentals can forecast future excess returns. Hamilton and Schwab (1985), Capozza and Seguin (1996), and Clayton (1997), find significant evidence against the hypothesis of rational home price expectations.

shown by Kim (2007), this result also obtains when agency costs make renting housing services more expensive than owning a home. Inverting and taking logs implies:

$$\log(HPRENT) = -\log(RUSER), \qquad (4)$$

where the elasticity of the price-to-rent ratio equals -1 and the price-to-rent ratio is invariant to the housing stock and deviations of income from trend.

However, Kim (2007) has recently theoretically demonstrated in an equilibrium model that when rental agency costs are accompanied by binding, maximum LTV ratios on marginal home buyers, the equilibrium log price-to-rent ratio is more complicated:

$$\log(HPRENT) = f(\log(RUSER), \text{max LTV}, \text{income deviations}), \qquad (5)$$

where income deviations equal actual minus permanent income, and the size of the negative real user cost elasticity can be smaller than 1 in line with empirical results (e.g., Gallin, 2006). We also test the price-to-rent approach using error or equilibrium correction models initially using log specifications (variables in logs are denoted with an "L" before their level names).

(c) Comparison of the Two Approaches

The price-to-rent approach is more grounded in finance and arbitrage, whereas the inverted demand approach is more grounded in consumer demand theory. Empirically, the relative advantages of the price-rent approach are that it is applicable where rental markets are flexible (e.g., U.S.), does not require housing stock data, and uses rents that track factors special to housing that are not controlled for by variables in the inverted-demand approach. Conversely, the inverted demand approach is better in: being practical for countries where rental markets are regulated; not ignoring that income shocks drive rents and home prices; and in tracking home prices in markets where both rents and home prices might be over- or under-valued. It is a priori unclear which approach is better for the U.S. and for robustness, we try both..

(d) Data

The variables used fall into the following categories: home prices and rents, real user cost, household income, housing stock, mortgage credit standards, capital gains and depreciation taxes, monetary/regulatory, and household expectation variables. So far, shifts in demographics variables were not found to be statistically or economically significant in regressions not shown, perhaps reflecting a number of breaks in the population data stemming from diennial censuses. We plan to further investigate adding demographic effects in subsequent versions of this paper.

Home Prices and Rents

We use Freddie Mac data on nominal home prices from repeat sales of homes and omit prices from mortgage refinancings, which are distorted by appraisers' incentives to inflate prices. We seasonally adjusted these data and then deflated them using the personal consumption expenditures (PCE) deflator to measure real home prices (*HP*). To construct the house price-to-rent ratio (*HPRENT*), we deflated nominal home prices with the PCE index for renting fixed dwellings, which closely parallels the owner-equivalent rent series from 1983-present.

Household Income

As in the FRB-US model, per capita income (*Y*) equals the tax adjusted sum of labor and transfer income, deflated by the overall personal consumption expenditures (PCE) deflator. Non-property income is used because it accords with standard consumer theory and avoids simultaneity bias by omitting property income, which includes rents that reflect property values.

Housing Stock

For the inverted demand approach, we tracked the real, per capita housing stock (*HSTOCK*) using Federal Reserve estimates of the replacement cost of residential housing structures owned by households deflated by the price index for housing construction.

Real User Cost of Mortgage Capital

The real user cost of capital (*RUSER*) is the after-tax sum of the effective conventional mortgage interest rate (*NOMRMORT*) and the property tax rate from the Federal Reserve Board (FRB) model plus the FRB depreciation rate for housing minus the annualized home price appreciation over the four prior years adjusted for an assumed 8 percent cost of selling a home. This resulting real rate exceeds zero in the sample, allowing real user costs to enter in logs, an appealing aspect stressed by Meen (2001). Some models split the user cost term into nominal user cost (*NOMUSER*) and appreciation (*APP*) terms to assess issues related to speculation.

Mortgage Credit Standards

Mortgage credit standards are tracked by the average LTV for homes bought by first-time home buyers (Duca, Johnson, and Muellbauer, 2009), based on American Housing Survey data since 1979. This series consistently measures LTV ratios on conventional mortgages, which corresponds to the Freddie Mac home price series that is based on homes bought with conforming conventional mortgages. This LTV series shifted up slightly, from a range around 85% in the late 1970s and the 1980s to a range around 87 percent in the 1990s (Figure 1), before jumping after 2002. Although the discontinued series of the Chicago Trust and Title company is from only one month (November) per year, both series move similarly before 2000 (Figure 2).[3]

We adjust raw quarterly AHS data for two reasons. First, we adjust the raw quarterly data for shifts in average age, seasonality, some unusually small quarterly samples, and regional composition that introduce noise and debt demand factors from which we wish to abstract. Second, we examined whether LTVs were endogenous to several macroeconomic variables over 1979-2007, finding no significant link with income and interest rates, but a statistical link with changes in the overall unemployment rate (*U*). To estimate these effects, Duca, Johnson, and

Muellbauer also regressed the raw, simple mean average LTV ratio on the above variables in the presence of the Hodrik-Prescott filtered LTV (*LTVHP*) to control for LTV trends and controlling for two unusual episodes which would otherwise distort estimates of other coefficients. The latter were the quarter following the September 11, 2001 terrorist attacks (*SEPT11*), which apparently induced a temporary plunge, and a dummy equal to one during the quarter of and following the passage of thrift bailout legislation in 1989:q3, which temporarily disrupted lending as savings and loan institutions were initially seized before being later closed (*FIRREA*):

$$\underset{}{LTV(raw)} = \underset{(1.37)}{0.082622} - \underset{(-3.02)}{.017889*\Delta U} + \underset{(-2.44)}{0.002296*AGE} + \underset{(2.67)}{0.074932*WEST} - \underset{(-6.20)}{0.061997*SEPT11}$$

$$+ \underset{(15.66)}{0.977258*LTVHP} - \underset{(-3.70)}{0.047266*FIRREA} + \underset{(1.42)}{0.082512*\Delta LTV_{t-1}}$$

where t-statistics are in parentheses, R^2 = 0.866, standard error = 01145, LM(2) = 1.64, and the estimation was done in the presence of quarterly seasonal dummies and dummy variables for quarters with less than 20 observations. *WEST* is the western share of first-time buyers in a quarter, which was the only regional share variable that was close to being statistically significant. The positive coefficient on *WEST* plausibly reflects the impact of higher home prices in that region on preferences with respect to LTV ratios and the tendency for faster home price appreciation in that region, which may make lenders feel comfortable with smaller down-payment cushions. The positive coefficient on age plausibly reflects that older households have somewhat more wealth, and would either be able to or would prefer to borrow at a lower LTV.

The adjusted series equals the raw series minus all of the above effects except that of the lagged dependent variable, *FIRREA*, and the H-P filtered LTV. To keep the adjusted LTV near its equilibrium, (1-coefficient on *LTVHP*)**LTVHP* was also deducted from the raw series. We then took a three-quarter, weighted average moving average of the resulting series using quarters

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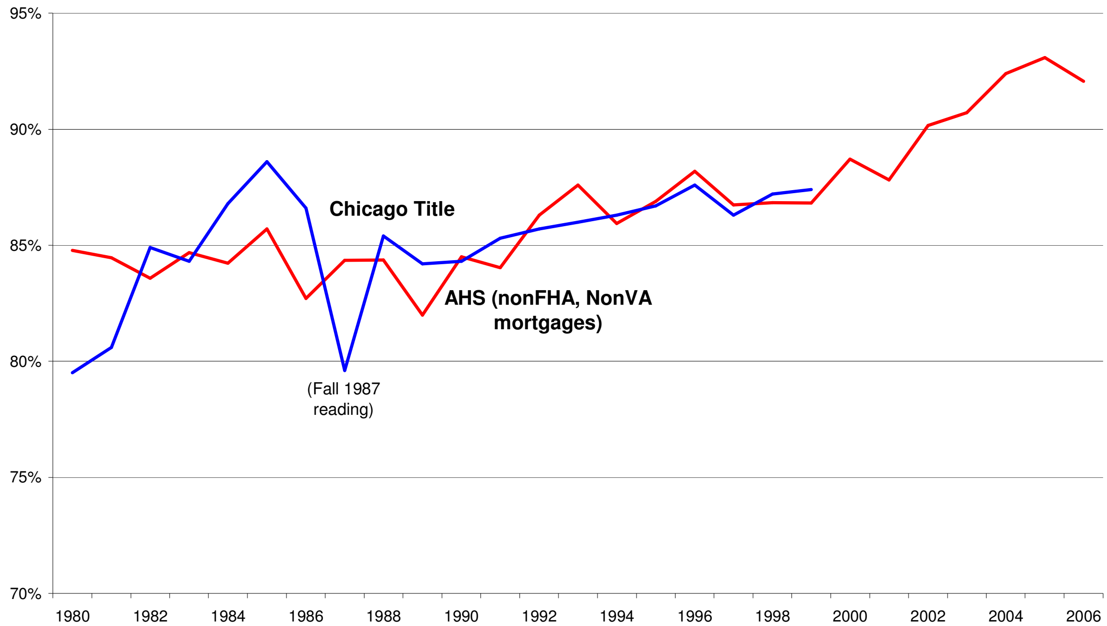
[3] This helps explain the low 1987 reading which followed the October 1987 stock crash seems unrepresentative.

Figure 2: Average First Time Homebuyer LTV Ratios, Chicago Title vs. AHS



t through t-2, where weights are the relative share of observations in each of the three quarters. This smoothes the series, with the observation weights treating individual borrowers equally.

The 2000-05 rise in LTV ratios likely reflects two financial innovations that fostered the securitized financing of riskier mortgages. The adoption of credit scoring technology enabled lenders to sort nonprime borrowers and attempt to price the risk of nonprime mortgages. Since these loans were too risky for banks to hold, they were funded by securities markets, where investor demand for the mortgage-backed securities funding nonprime loans was temporarily boosted by two unusual developments. First, the combination of low interest rates and expanded credit availability in the early 2000s fueled a rise in house prices that plausibly led investors and analysts to under-estimate the default risk on nonprime mortgages. As DiMartino and Duca (2007) argue, the short history of subprime mortgages over 1998-2006 may have tempted analysts to forecast problem loans on labor market conditions while not having enough sample to disentangle the impacts of house prices and interest rates. In addition, the tendency for vintages of Alt A mortgages to have progressively higher proportions of no- or low-documentation of income (Credit Suisse 2007) and to post progressively worsening loan quality (Mayer, Pence, and Sherlund, 2009) suggests that an errors-in-variables problem from overstatements of borrower income contributed to the underestimation of nonprime mortgage defaults.

Second, financial innovations that sorted borrowers or funded their loans were accompanied by changes in regulations and public policies leading to noncyclical increases in the demand for the securities funding these mortgages. This included a 2004 SEC decision that doubled the 1935 limits on investment bank leverage from 15:1 to 33:1 and the rise of hedge funds and SIVs that used short-duration debt to fund risky long positions in nonprime mortgages. Also important were large purchases of nonprime MBS by Fannie Mae and Freddie Mac to meet their public policy goals of expanding home ownership (Frame, 2008) even though they did not issue much nonprime MBS. Technological and policy innovations fostered originations of nonprime mortgages which were sold to the GSEs or private investors in the form of collateralized debt obligations (CDOs) or with protection from credit default swaps (CDSs). The later failure of CDOs to protect investors from unanticipated defaults and the soaring cost of using CDSs led to a collapse in the funding and availability of nonprime mortgages.

Capital Gains

Although income tax rates are in the user cost of capital variable, capital gains tax changes have notably affected home prices. Before 1998, net capital gains on home sales were taxable for households under age 55 if the seller did not purchase a home of equal or greater value. The Tax Reform Act of 1997 largely eliminated this tax by exempting the first $500,000 ($250,000) of gains for married (single) filers, raising the after-tax value of homes and encouraging turnover (Cunningham and Englehardt, 2007). To control for this, we included a dummy (*CAPGAINTAX*) equal to 1 since 1998:q1 and 0 before then.[4]

Household Expectation Variables

Labor expectations could have marginal information about home prices not captured by other variables. From the Conference Board survey of consumer confidence, we include the log of an index equaling 100 plus the percent of households who expect more minus fewer jobs over the next six months, (*LCONFLABOR*), which may reflect information not in current income.

Monetary/Regulatory Variables

One variable (*MONEYTARGET*) equals 1 over the money targeting regime of 1979:q4-1982:q3, which may have reduced the supply of or demand for mortgages by raising interest rate uncertainty. Another is the two-quarter change in Duca's (1996) measure of how much Reg Q ceilings on deposit interest rates were binding (*Δ2RegQ*), which controls for negative short-run disintermediation effects that are not consistently tracked by the user cost of capital (Duca and Wu, 2009). Another variable is the change in the highest upfront insurance premium for Federal Housing Administration (FHA) loans. A rise in the upfront premium (from 0 to 3.8% of the mortgage principal) was announced to take effect in late 1983:q3, inducing many renters to leave rental housing and purchase "starter" homes. This hike was later partially offset by smaller, but sizable cuts. We include the change in the FHA upfront premium (*ΔFHAFEE*) for such short-run effects in price-to-rent models, for which the combined short run effects on prices and rents have a statistically significant effect on the price-to-rent ratio but not the real price of houses.

III. Long-Run and Short-Run Results from Cointegration Models

We first present findings using cointegration methods given the nonstationarity of long-

[4] In other runs, we found that another tax variable to have non-robust effects. This was the time over which rental properties can be depreciated for taxes which may raise the after-tax cost of renting relative to home prices.

run variables, and then relax some of the cointegration restrictions in simpler one-stage OLS models containing lagged levels and first differences of housing variables. Both types of models control for tax effects beyond including simple income and property tax rates in calculating the user cost of mortgage capital, as well as control for the money targeting regime of 1979-1982 that imparted more interest rate risk to house prices beyond that reflected in simple user cost of capital variables. By addressing these important influences, we try to avoid omitted variable bias that can obscure long-term, qualitative relationships and lead to poorly estimated coefficients.

(a) Results Based on the Inverted Demand Approach

Long-Run Results

The long-run solutions implied by this approach in eq. (2) involve the income, user cost, housing stock, and credit variables in the unique, estimated cointegrating vectors of Table 1. Results are from models with and without LTV ratios. The first two vectors are for models 1 and 4 which only include the monetary targeting regime variable as an additional exogenous variable and are estimated over a full common sample 1981:q1-2007:q3. This sample reflects the number of lags of first difference variables needed to obtain a unique and significant cointegrating vector and which minimized the AIC statistic under time series assumptions allowing for possible time trends in the endogenous variables but not a time trend in the cointegrating vector. The second set of two vectors from models 3 and 6 also include the three extra tax, Regulation Q, and job expectations variables, and are estimated over the same full common sample. The third set of vectors includes the same exogenous variables, but has a common end date of 2001:q4, which ends just before LTV ratios jumped during the subprime boom that started in 2002. Lagging the LTVs by one more quarter (t-2) improved the models, perhaps reflecting that it takes longer for homebuyers to detect a change in credit standards than in more visible mortgage interest rates.

Although unique cointegrating vectors are obtained with the expected signs for some non-LTV and all LTV models, the latter are better in several ways. First, for the full sample, LTV vectors have a higher degree of statistical significance (99% confidence level) and unlike the LTV model, a unique cointegrating vector could not be found using the maximal eigen statistic for the nonLTV model in the presence of the full set of exogenous variables. Second, in the simplest vectors that include only the monetary targeting variable as an extra term, the real user cost is statistically insignificant in the non-LTV model. Third, including LTV ratios yields more plausible coefficients especially in the full sample: the coefficient on log housing stock, interpretable as the inverse of the long run price elasticity of demand, is between -1.4 and -0.8 for the LTV models, but implausibly ranges from -2.5 to -0.6 in the non-LTV models. Furthermore, the ratio of coefficients on log income and log housing stock estimates the long run income elasticity of demand, which reasonably lies between 1.5 and 1.8 in LTV models, versus an implausible span of 1.5 to 2.5 in non-LTV models. Finally, for models including a full set of controls, LTV model coefficients are very similar in the pre-subprime boom and full samples.

Short-Run Results

Table 2 reports results from ECM models of the first difference in real home prices, where the error-correction terms equal the gap between actual and equilibrium house prices, where the latter correspond to numbered vectors in Table 1. In several ways the LTV models outperform corresponding non-LTV models. First, LTV models yield corrected R squares that are .04 to .07 higher than from corresponding non-LTV models. Second, the adjustment speeds for LTV models are higher (12 to 18 percent per quarter), and the error-correction term in the non-LTV model 1 is insignificant. Third, in the full sample model having all short-run variables, the speed of error-correction is a higher 16 percent in model 6 versus 6 percent in model 3.

In Table 2, the coefficients on lagged changes in the log per capita housing stocks tend to be negative, as expected and in line with UK results (Cameron et al., 2006). The income dynamics are consistent with a moving average of income. The dynamics in log real user cost also suggest a moving average. Short run dynamics in lagged house price changes suggest a positive short term momentum effect, aside from that embedded in the real user cost term.

We checked the robustness of these results in one-step estimation with alternative lag lengths, confirming the long-run solutions and estimated adjustment speeds. Furthermore, results are generally unaffected by breaking the log real user cost into its two main elements, the tax adjusted interest rate and the annualized rate of home price appreciation over the prior four years. One exception is when we added a new non-linear term, the cubic of home price appreciation over the last four quarters. Without this extra term, the estimated speed of adjustment drops and the residuals suffer from serial autocorrelation. These results are available on request.

Exogeneity

A natural question is whether the LTV series is driven by house prices, which would greatly complicate the interpretation of the above findings. In a vector-error correction system using the lag length of model 6 in Tables 1 and 2, the error correction term is highly insignificant (t statistic of 1.40) in modeling the LTV ratio, indicating that the LTV ratio is weakly exogenous to the other variables, as is the case for the real user cost (t-statistic of 1.75) and income (t-statistic of -0.35). In contrast, house prices are not weakly exogenous (t-statistic of -7.04 on the EC term), as is the case for the stock of housing (t-statistic of -3.41). These results point to an asymmetry to how the vector components adjust to disequilibria, with house prices and the housing stock making the significant adjustments. Thus, consistent with theory, equilibrium house prices and the housing stock are driven by income, user costs, and credit availability.

(b) Results Using the Home Price-to-Rent Ratio Approach

For robustness, we assess the importance of mortgage availability using a home price-to-rent approach. In this approach, exogenous increases in mortgage availability, that are unrelated to income and interest rate movements, alter the relative demand for owner-occupied versus rental housing by increasing the effective demand for owner-occupied housing of the credit constrained and lowering their effective demand for rental housing.[5] Such a relative demand shift can alter the equilibrium price-to-rent ratio by affecting the land intensity of housing since the supply of land is not as price elastic as is the cost of building structures (Davis and Heathcote, 2005). Home price-to-rent models generally estimate a long-run relationship between mortgage interest rates and a price-to-rent ratio, and often find that U.S. home prices are over-valued. Exceptions to the latter are regional or city models that either (1) use unusually low user cost of capital rates arising from assumptions that unusually high rates of past local price appreciation will persist (e.g., Himmelberg, Mayer, and Sinai, 2006) or (2) argue that rental rates are higher in high cost locales than implied by official rent data (Smith and Smith, 2006). Following convention, we use standard measures of rents and use national price appreciation rates to construct real user cost of capital measures. We depart from published models by including our cyclically adjusted measure of LTV ratios for first-time home buyers. We add this variable to cointegrating vectors containing the home price-to-rent ratio (HPRENT) and user cost of mortgage and compare long-run and short-run results to models that omitting LTV ratios.

Long-Run Results

Table 3 reports cointegrating vectors of national price-to-rent ratios estimated over the full sample (data from 1979-2007) under assumptions allowing for deterministic trends in the long-run variables, but not in the cointegrating vector. The lag lengths were long enough to yield statistically significant unique vectors, minimize the AIC statistic, and yield reasonably clean residuals. For the non-LTV models, this was also done when possible, and when not, the vector minimizing the SIC statistic was selected. The first two vectors (numbered 1 and 4) include the monetary targeting dummy, and respectively omit and include the LTV ratio. The third and fourth vectors (numbers 3 and 6) also include terms for the other tax effects, job expectations, and the Regulation Q variable. The fifth and sixth vectors also use these variables, but in a shorter sample ending in 2001:4 prior to the subprime boom starting in 2002.

As implied by trace and maximal eigenvalues, unique cointegrating vectors were found in each of the LTV vectors, with either mixed or less strong evidence of cointegration in models omitting the LTV ratio. In addition, in vector error correction models, the error-correction term was significant in the price-to-rent equation and was insignificant in the income and user cost models. As in the inverted demand models, these two findings are consistent with the view that LTV ratios are largely exogenous drivers of home prices. Consistent with priors, the estimated long-run coefficients in all the vectors indicated that home prices are negatively and significantly related to real user cost of capital, and are positively and significantly related to the LTV ratio.

The LTV models also show less parameter instability when the sample is extended from 2001.q4 to 2007:q3, where the former quarter precedes the 2002-05 jump in LTV ratios. In non-LTV models, the long-run coefficient on real user costs rises notably as the end of sample is extended from 2001 to 2007. The equilibrium price to rent ratios implied by the full non-LTV model (vector 3) differ notably using model estimates through 2001, with a tendency for the earlier period coefficients to under-predict home prices in the mid-2000s (Figure 3).[6] In contrast

[5] Higher LTV ratios in the early 2000's coincided with a jump in subprime mortgages and the homeownership rate.
[6] Note that equilibrium is defined using the current value of real log user cost, reflecting the annual rate of lagged house price appreciation. A plausible alternative is to use historical average appreciation, as discussed in Section IV.

to the parameter instability of non-LTV models, LTV models exhibit parameter stability given the large swings in LTV and price-to-rent ratios of recent years. The equilibrium price-to-rent ratios implied by the full LTV model (vector 6) do not differ much when estimated through 2001 rather than through 2007:q3 (Figure 4). A plausible interpretation of this result, as well as the tendency for the real mortgage rate coefficients to rise in the non-LTV models, is that non-LTV models omit important information and that the rise in LTV ratios is an important driver, along with low real interest rates, of the large increases in U.S. home prices over the period 2002-2005.

Short-Run Results

An easing of mortgage credit standards also has large short-run effects on home prices, as shown in Table 4 which reports the error-correction model results for the change in the home price-to-rent ratio which use the long-run equilibrium relationships corresponding to vectors in Table 3. In each LTV model, the error-correction term is very significant with plausible speeds of error-correction of 6-17 percent per quarter. By contrast, the error-correction speeds of the noncredit models range between 1.2 and 8.2 percent, reflecting the lower ability of non-LTV models to track long-run relationships. This is particularly the case in full sample models that include all the tax, monetary policy, regulatory and labor expectations variables, where the speed of error- correction is 17 percent in the LTV model versus only 7 percent in the non-LTV model. Comparing conventional with credit models that contain the same short-run exogenous variables indicates that including the LTV ratio and its lagged first-difference improves the R-squares of corresponding models by 4-7 percentage points and lowers standard errors by 6-18 percent. Estimates suggest some short term persistence in the dynamics, in that last quarter's change in the log house price to rent ratio has a significantly positive effect on the current log price to rent ratio, though lagged house price appreciation over a longer period is incorporated in log RUSER.

Figure 3: Implied Equilibrium Log Home Price-To-Rent Ratios Differ for Models Omitting LTV Ratios Estimated with and Without 2002-2007

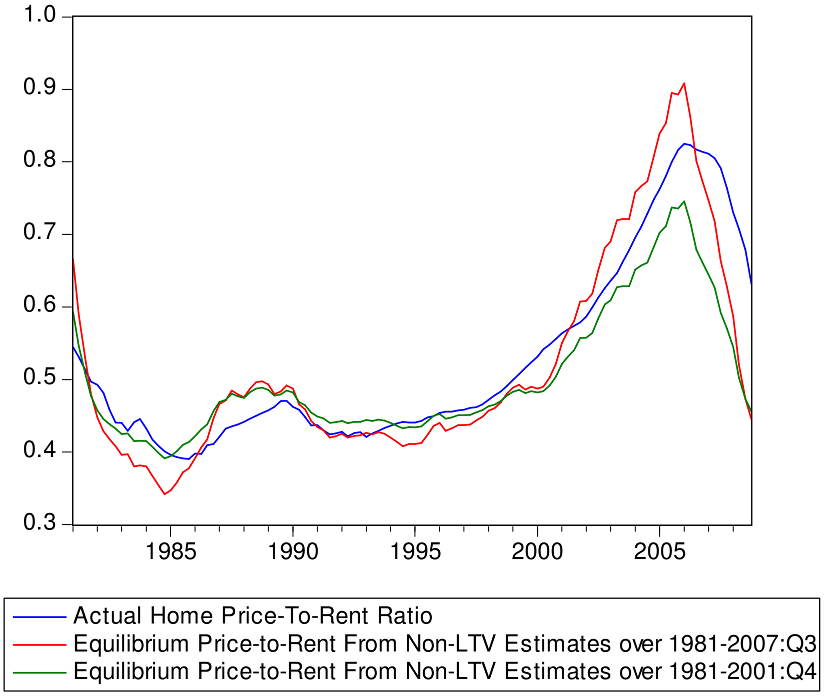
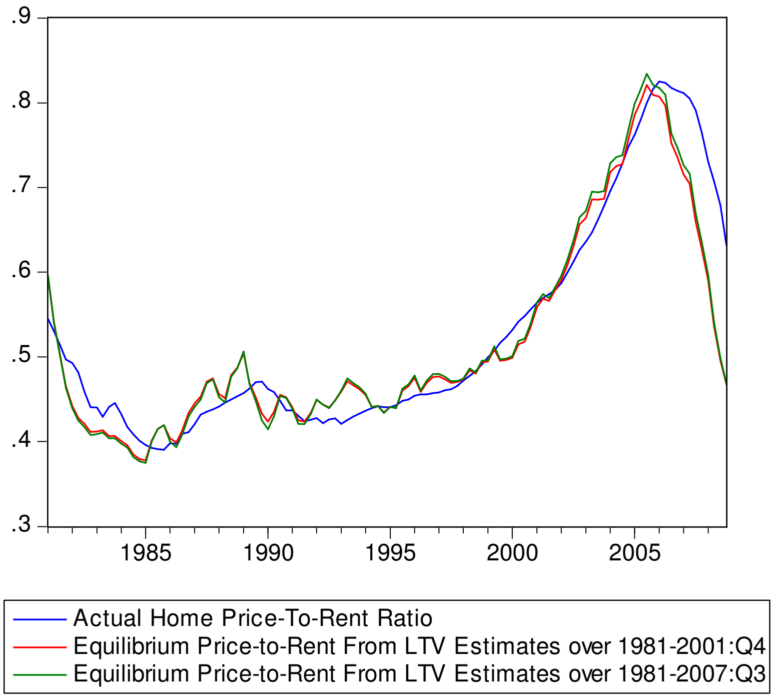


Figure 4: Implied Equilibrium Log Home Price-To-Rent Ratios Similar for LTV Models Estimated with and Without 2002-2007



The LTV models outperform non-LTV counterparts in yielding stronger and more stable long-term home price-to-rent relationships. Also, with respect to explaining short-run changes in the price-to-rent ratio, LTV models yielded better model fits, faster speeds of adjustment, better behaved residuals, and more sensible estimates of changes in the tax treatment of capital gains. In line with the view that easier mortgage credit standards significantly fueled the home price boom of the mid-2000s, the non-LTV models exhibit the symptoms of omitted variable bias. Consistent with this interpretation, in a vector-error correction system using the lag length of model 6 in Tables 4 and 5, the error correction term is highly insignificant (t statistic of 1.19) in modeling the LTV ratio and the real user cost (t-statistic of -0.06), implying that the LTV and real user cost variable are weakly exogenous to other variables. In contrast, the house price-to-rent ratio is not weakly exogenous (t-statistic of -5.09 on the EC term).

(c) Results Using One-Step Models

Using the inverted demand approach, we performed robustness checks with similarly satisfying results: one-step estimates and the alternative specification of non-linear dynamics give similar long-run solutions and estimated speeds of adjustment. In particular, one-step inverted demand models estimated through 2007 show that, relative to corresponding non-LTV models, LTV models have better fits, have faster speeds of adjustment over the full sample, imply long-run price elasticities of the housing stock that are near unity, and have more plausible income elasticities. (Compare models 1 vs. 4, and models 3 vs. 6 in Table 5).[7]

We also examine whether the qualitative results are similar in one-step models and in specifications allowing for nonlinear dynamics. Results show that, relative to corresponding

[7] Because the LTV is a 3 quarter moving average, we include the 3 quarter change in LTV ratios (lagged 2 quarters) to control for short-run dynamics rather than include one-quarter changes that contain similar information.

non-LTV models, LTV models have better fits (.04 to .05 higher corrected R^2's) and faster speeds of adjustment reflecting the statistical significance of the LTV ratio.

d) Alternative One-Stage Specifications

There are two alternative specifications that can help us further assess the robustness of the findings. The first concerns a short-coming of the repeat sales index: its use of repeat sales as a means of controlling for shifts in quality, when home improvements that are linked to rising incomes can cause such indexes to overstate house prices and effectively lead to over-estimates of the income elasticity of house prices. To address this we construct and then analyze a house price index which adjusts for home improvements. We first use a cumulative sum of home improvements from a Census quarterly series on home structure improvements (which unfortunately ends in 2007) and then adjust that sum for depreciation. The depreciation adjusted stock of home improvements is then scaled by Federal Reserve Board Flow of Funds estimates of the replacement cost value of residential home structures, which gives a time series of the relative importance of home improvements as a source of housing stock accumulation. We then multiply this series by the Freddie Mac repeat home sales price index to adjust repeat home sales prices for home improvements (*LRHPadj*).

The second alternative model adds in changes in the share of the age 16 and over population between the prime house buying years of 25 and 44. The share in levels is I(2), so we difference the data. In addition, because age composition data are really available annually, quarterly interpolations are used. To avoid overinterpreting the data, the year-over-year change in quarterly age composition is added to select regressions (*Δpop2544*). This also helps smooth out some breaks in annual data readings as the incorporation of diennial Census information introduces some detectable breaks in age compostion data, particularly in 1972 and 1992-94.

One-stage regressions using the inverted housing demand approach are given in Table 7. Similar two-stage estimation was performed using cointegration techniques, which yielded qualitatively similar results to these one-stage results and to earlier results with respect to the improved model performance that arises when LTVs are incorporated into house price models. Models 1-2 are full sample nonLTV models having a full set of exogenous controls in which model 1 repeats model 3 from Table 5, model 2 replaces the price index terms with home-improvement analogues and model 3 adds the year/year change in the adult population share of age 25-44 persons. Models 4-6 are then just comparable models that add the t-2 log level of LTVs and the t-2/t-5 change in log LTVs to capture first difference effects not contaminated by the three-quarter nature of the moving average smoothing used in constructing the LTV series.

Several patterns emerge in the results. First, the same qualitative findings as before are essentially obtained: LTV models yield better fitting models with faster speeds of adjustment and more sensible parameters. Second, switching to the use of home-improvement-adjusted prices lowers the implied income elasticity of housing demand from about 1.5 to a more believable 1.25. This is consistent with the concern that repeat sales price indexes likely overstate the rise over time in real house prices. Third, the change in age composition is significant in the LTV but not the nonLTV model (models 3 vs. 6). Fourth, including the age composition variable results in faster speeds of equilibrium adjustment (especially in the LTV model (6)), but unwinds the improvement in income elasticities, leading to higher implied income elasticities. The inclusion of population and home-improvement adjusted house prices does not undermine the gains from augmenting house price models by including LTVs for first-time homebuyers. Nevertheless, issues regarding breaks in the population data may partly obscure the true role of demographic changes and future work will address such breaks.

IV. How Much Are U.S. Home Prices Overvalued?

To throw light on how much U.S. home prices may be overvalued, we now examine the implications of the two home price models incorporating LTV terms for the deviations of prices from their 'equilibrium' or 'long-run' values. As noted in section III there is more than one concept of equilibrium. The narrow concept is conditional on the observed log real user cost as used in our econometric models. Consider model 6 (Tables 3 and 4) for the home price-to-rent approach as an example. The long run solution is LHPRENT = 0.9 -0.17*LRUSER + 0.81*LLTV + fitted effects of persistent terms or step dummies for tax variables and the 1979-82 monetary policy regime. Then, conditional on LRUSER, the deviation from equilibrium is:

LHPRENT - 0.9 + 0.17 LRUSER - 0.88 log LTV - fitted step function dummies,

which reflects I(0) variables (e.g., lagged changes in LHPRENT, residuals, and other variables.

By this metric, U.S. home prices were over-valued by 9% using LTV model 6 versus 5% using non-LTV model 3. Similar deviations can be computed from inverted demand models, such as models 3 and 6 from Tables 1 and 2, which suggest 15 and 13 percent overvaluations in 2007:q2, respectively. However, these calculations have the shortcoming that RUSER contains the annual house price appreciation over the prior 16 quarters, which cannot be regarded as permanent and is part of the 'bubble builder' in the model's dynamics, as Abraham and Hendershott (1996) discuss. We address this issue and alternative assumptions about post-2007:q2 values of LTV. We then calculate how much U.S. home prices are over-valued by simulating a two equation model which endogenizes rents and house prices in several scenarios.

A model for rents

In the long run, one would expect residential rents to depend on the general price level, house prices, the cost of capital and property taxes. These propositions, as well a long-run price homogeneity, are supported by an econometric model estimated for 1972 to 2008, and for comparison, 1981-2008, shown in Table 7. In the long-run solution, the consumer expenditure deflator has around three quarters of the weight, while the level of house prices has around one quarter. The tax adjusted nominal mortgage rate also appears in the long run solution, but is not balanced by offsetting inflation terms which potentially might have defined a real interest rate. The model therefore implies some nominal inertia. In the dynamics, the lagged change in log rents has a positive effect, while other important variables are the annual change in the log of real energy prices, the annual growth rate of personal income, and the change in the tax adjusted mortgage rate. Dummies for the ending of the Nixon price controls, the Carter credit controls of1980Q1 and the change in the period in years over which apartments can be depreciated for tax purposes, also appear. The speed of adjustment is slow, less than 0.05 per quarter.

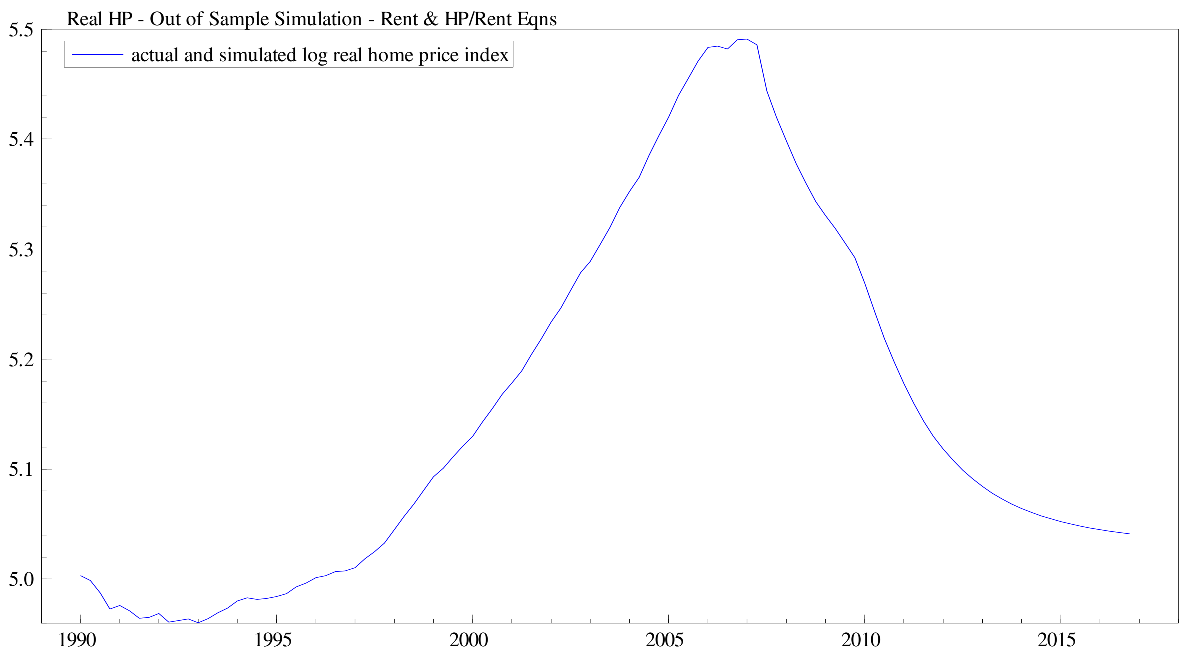
The implications of the two equation model, conditional on paths for the loan-to-value ratio, interest rates, income growth, real energy prices, and inflation are interesting. A return of loan-to-value ratios for first time buyers to late 1990s levels, together with the downward momentum of falling home prices since 2006, leads to further falls. As house prices fall, rents are eventually and gradually dragged down also. Since rents appear in the long-run solution for house prices, this prolongs the fall in house prices. Lower interest rates are only marginally helpful, according to the two equation model, since while lower rates raise house prices given rents, they also lower rents. To the extent that lower mortgage rates raise income growth and prevent deflation, however, they could positively affect on rents and thereby house prices.

We review two scenarios. The first is quite negative: the LTV ratio returns to its 1998 level early in 2009; inflation and nominal per capita income growth are zero from 2009 onwards; interest rates fall one percentage point in 2008Q3 relative to 2008Q2 and stays there through

2016. In this scenario, real house prices reach their floor level in 2016 at roughly the same level they were at in 1998. In an alternative scenario under which real per capita income grows at 2 percent per annum from 2010, the 2016 level of real house prices is close to the 1999 level. Figure 5 shows the former scenario for the log of real home prices, which fall sharply until about 2012, after which they begin to stabilise. This implies that the house price correction is not even half over by the Spring 2009 and appears to be consistent with an extended subdued period for the U.S. housing market, its banking system and consumer spending in the U.S.

The robustness of these findings to alternative formulations of house price dynamics and the implications of the inverted demand approach remain to be explored. On the face of it, the inverted demand approach offers more hope for the scope of low interest rates to affect house prices and suggests that a less pessimistic picture may result. In this approach, endogenizing housing supply would seem a minimal requirement, and this work remains to be done. The provisional conclusions drawn above should therefore be treated with considerable caution.

Figure 5: Simulating Real U.S. House Prices



V. Conclusion

Our findings provide a theoretically appealing and empirically consistent account of the behavior of U.S. home prices. In the two main theoretical approaches to modeling house prices, the inverted demand and home price-to-rent frameworks, credit standards for first-time home-buyers are important determinants of home prices. Our results indicate that a substantial easing of U.S. mortgage standards, as reflected in the LTV ratios for first-time home-buyers, raised the effective demand for housing in the first half of the decade. Between early 2005 and mid-2007, there was a partial reversal of that easing, which had likely intensified since mortgage-related financial market turmoil started in August 2007 and later intensified during the Fall of 2008.

Most empirical models of US home prices lack a measure of mortgage credit standards and thus suffer from a meaningful omitted variable bias, rendering them less capable of tracking the earlier surge of home prices during the mortgage boom and the unwinding of much of that appreciation during the early phases of the subprime bust. In contrast, models including a cyclically adjusted LTV measure for first time home-buyers yield sensible and statistically significant long-run relationships, and in models of short-run movements in home prices, more precise estimates of key coefficients, reasonable speeds of adjustment, and better model fits. Furthermore, our credit-augmented models imply that much of the boom-bust cycle in U.S. home prices stemmed from an easing and subsequent tightening in U.S. mortgage standards affecting potential marginal home-buyers. From a broader perspective, our results are consistent with the view that many asset bubbles are linked to an unsustainable easing of credit standards or adoption of risky financial practices that eventually unwind during a subsequent bust.

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Table 1: Cointegration Results for Home Prices, Inverted Demand Approach, 1980-2007:q3

Model &Vec.	# Sign. Vectors	Cointegrating Equilbrium Relationship (assumption: no trend in vector or in variables)	Log Likelihood (AIC)	Eigenvalue (Trace Statistic, Max-Eigen Statistic) 0 vectors	1 vector
		Sample: 1981q1-2007q3 (only monetary policy regime variable added)			
1	1	$LHP_t = 3.314824+2.693261LY_t^{*}-.045382LRUSER_t-1.771903LHS_t^{*}$	1615.119	0.236924	0.152405
		(2.55) (-1.13) (-2.15)	(-29.06764)	(54.61402*)	(25.68148)
				(28.93254*)	(17.69274)
4	1	$LHP_t = 4.260417+1.746419LY_t^{**}-.149536RUSER_t^{**}-1.134352LHS_t^{**}+1.147345LLTV_{t-1}^{**}$	2018.878	0.329854	0.214482
		(6.99) (-15.42) (-5.51) (4.99)	(-35.25700)	(87.22345**)	(43.99539)
				(43.99539**)	(26.07244)
		Sample: 1981q1-2007q3 (tax, monetary policy, Reg Q, and expectations variables)			
3	1	$LHP_t = 3.721727+1.590178LY_t^{**}-.129213LRUSER_t^{**}-0.892345LHS_t^{*}$	1634.209	0.223388	0.167210
		(3.04) (-7.32) (-2.33)	(-29.42448)	(51.87766*)	(24.82657)
				(27.05109)	(19.57825)
6	1	$LHP_t=4.230976+1.881746LY_t^{**}-.148335LRUSER_t^{**}-1.234785LHS_t^{**}+1.332139\ LLTV_{t-1}^{**}$	2042.370	0.475761	0.203049
		(9.06) (-19.74) (-7.50) (7.55)	(-36.02562)	(108.0653**)	(38.96381)
				(69.10145**)	(24.28488)
		Sample: 1981q1-2001q4 (tax, monetary policy, Reg Q, and expectations variables)			
2	1	$LHP_t = 3.756295+0.999328LY^{**}-.125247LRUSER^{**}-.397243LHS^{**}$	1351.695	0.350621	0.131978
		(5.28) (-10.30) (-2.97)	(-30.75464)	(55.94109**)	(19.67501)
				(36.26608**)	(11.88923)
5	1	$LHP_t =4.111441+1.294386LY_t^{**}-.143123LRUSER_t^{**}-0.719778LHS_t^{**}+0.716542\ LLTV_{t-1}^{**}$	1645.672	0.514093	0.145417
		(6.95) (-12.15) (-5.14) (4.85)	(-37.03981)	(88.32033**)	(27.69436)
				(60.62597**)	(13.19986)

	Level (AIC lag in parentheses)	*5% Critical level for lag*	*1% Critical level for lag*	*First Diff. SIC lag in parentheses)*	*5% Critical level for lag*	*1% Critical level for lag*	*Assumptions*
ln(HP)	-1.868602 (4)	-3.449716	-4.040532	-3.741875* (7)	-3.449716	-4.040532	constant/trend
ln(RUSER)	-2.389194 (2)	-3.449716	-4.040532	-4.124070** (1)	-3.449716	-4.040532	constant/trend
ln(LTV)	-2.060300 (3)	-3.451959	-4.045236	-10.78284** (3)	-3.451959	-4.045236	constant/trend
ln(Y)	-2.292525 (1)	-3.449716	-4.040532	-13.50677** (0)	-3.449716	-4.040532	constant/trend

Notes. *(**) denotes significant at the 95% (99%) level. t-statistics in parentheses except when AIC statistic is reported. For vectors numbered 1-6, lag lengths of (3, 3, 3, 4, 3, and 4), respectively, minimized the AIC and yielded significant and unique vectors allowing time trends in the variables. Lag lengths in the ADF unit root tests based on the Schwartz Information Criterion. Data used span 1979-2007:2.

Table 2: 2nd-Stage ECMs of the Percent Change in Real U.S. Home Prices

	No LTV Terms			*LTV Terms*		
	81:1-07:3	81:1-01:4	81:1-07:3	81:1-07:3	81:1-01:4	81:1-07:3
Variable	**Model 1**	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
Constant	0.0021	-0.0662*	-0.0515+	0.0079**	-0.0770**	-0.1031**
	(1.52)	(-2.12)	(-1.68)	(4.27)	(-2.73)	(-3.86)
ECM_{t-1}	-0.0095	-0.1664**	-0.0554**	-0.1238**	-0.1824**	-0.1589**
	(-1.38)	(-3.74)	(-3.55)	(-4.74)	(-6.48)	(-7.04)
$MONEYTARGET_{t-1}$	-0.0028	-0.0105**	-0.0054*	-0.0147**	-0.0155**	-0.0141**
	(-1.17)	(-3.11)	(-2.04)	(-4.14)	(-4.69)	(-4.82)
$CAPGAINTAX_t$		0.0072**	0.0062**		0.0064**	0.0054**
		(4.31)	(4.04)		(4.59)	(4.30)
$\Delta 2RegQ_{t-1}$		-0.0043**	-0.0032*		-0.0047**	-0.0038**
		(-3.01)	(-2.37)		(-3.55)	(-3.30)
$LCONFLABOR_t$		0.0157*	0.0117+		0.0186**	0.0244**
		(2.28)	(1.75)		(2.96)	(4.14)
ΔLHP_{t-1}	0.9021**	0.8032**	0.7339**	0.8751**	0.7271**	0.6141**
	(8.26)	(7.11)	(6.62)	(8.40)	(7.02)	(6.33)
ΔLHP_{t-2}	-0.0879	0.0338	-0.0721	0.0986	0.0781	0.0482
	(-0.60)	(0.23)	(-0.53)	(0.72)	(0.58)	(0.41)
ΔLHP_{t-3}	0.0736	0.1039	0.0143	0.2183	0.1158	0.1353
	(0.58)	(0.78)	(0.11)	(1.62)	(0.98)	(1.17)
ΔLHP_{t-4}				0.0228**		-0.1040
				(3.27)		(-0.96)
$\Delta LRUSER_{t-1}$	0.0135	0.0315*	0.0164+	0.0337**	0.0360**	0.0259**
	(1.30)	(2.21)	(1.64)	(3.27)	(2.78)	(2.98)
$\Delta LRUSER_{t-2}$	-0.0071	0.0001	-0.0042	0.0149+	0.0051	-0.0116
	(-0.82)	(0.01)	(-0.47)	(1.67)	(0.51)	(-1.49)
$\Delta LRUSER_{t-3}$	-0.0139+	-0.0131	-0.0123	0.0027	-0.0087+	-0.0017
	(-1.80)	(-1.36)	(-1.58)	(0.33)	(-1.00)	(-0.24)
$\Delta LRUSER_{t-4}$				-0.0010		-0.0062
				(-0.13)		(-0.98)
ΔLY_{t-1}	-0.0660	-0.1839**	-0.1659*	-0.1933**	-0.2271**	-0.3016**
	(-1.04)	(-2.71)	(-2.62)	(-2.75)	(-3.61)	(-4.81)
ΔLY_{t-2}	-0.0141	-0.1627*	0.1265+	-0.0911	-0.1768**	-0.2191**
	(-0.21)	(-2.28)	(1.91)	(-1.40)	(-2.74)	(-3.67)

ΔLY_{t-3}	0.0319	-0.0047	-0.0297	-0.0167	-0.0275	-0.0877^{+}
	(0.54)	(-0.08)	(-0.53)	(-0.28)	(-0.50)	(-1.64)
ΔLY_{t-4}				-0.0315		-0.0541
				(-0.59)		(-1.18)
$\Delta LHSTOCK_{t-1}$	-0.4154	-1.0627*	-0.4164	-1.0535*	-1.3728**	-1.1154**
	(-0.87)	(-2.11)	(-1.53)	(-2.27)	(-2.96)	(-2.77)
$\Delta LHSTOCK_{t-2}$	0.3854	0.3811	0.4425	0.0681	0.3334	0.1842
	(0.66)	(0.64)	(0.82)	(0.12)	(0.62)	(0.40)
$\Delta LHSTOCK_{t-3}$	-0.2587	-0.2992	-0.2173	-0.2807	-0.1650	-0.1760
	(-0.62)	(-0.75)	(-0.57)	(-0.52)	(-0.46)	(-0.39)
$\Delta LHSTOCK_{t-4}$				0.0197		0.0876
				(0.05)		(0.27)
$\Delta LLTV_{t-2}$				-0.0190	-0.0393	-0.0964*
				(-0.40)	(-1.11)	(-2.19)
$\Delta LLTV_{t-3}$				-0.0421	-0.0335	-0.0903*
				(-1.07)	(-0.95)	(-2.55)
$\Delta LLTV_{t-4}$				-0.0805*	-0.0828*	-0.1299**
				(-2.01)	(-2.27)	(-3.60)
$\Delta LLTV_{t-5}$				-0.0422		-0.0966**
				(-0.95)		(-2.47)
R^2	.743	.782	.786	.795	.822	.854
S.E.	0.004273	0.003522	0.003901	0.003815	0.003177	0.003221
VECLM(2)	16.45	21.40	15.29	23.18	34.36	20.35
VECLM(8)	10.54	15.93	19.86	19.47	26.85	21.89

(**,$^{+}$) significant at 95% (99%, 90%) level. t-statistics in parentheses. EC terms from VECMs estimating the long and short-run relationships. EC terms in the numbered models are derived from the corresponding numbered vector in Table 1. Δ lags later than t-4 omitted to conserve space.

Table 3: Cointegration Results for the U.S. Home Price-to-Rent Ratio, 1981 -2007:q3

Model &Vec.	# Sign. Vectors	Cointegrating Equilibrium Relationship (assumption: no trend in vector or in variables)	Log Likelihood (AIC)	**Eigenvalue (Trace Statistic, Max-Eigen Statistic)** 0 vectors	1 vector
		Sample: 1981q3-2007q3 (only monetary policy regime variable added)			
1	1	ln(HPRENT) = 1.202309 -.387976*ln(RUSER)**	585.8222	0.252518	0.022486
		(-12.10)	(-10.96804)	(32.94761**)	(2.387928)
				(30.55969**)	(2.387928)
4	1	ln(HPRENT) = 1.046840 -.202192*ln(RUSER)** + 1.213479*ln(LTV_{t-1})**	944.8603	0.224725	0.086631
		(-12.85) (4.52)	(-17.31163)	(36.24323**)	(9.516735)
				(26.72649**)	(9.514627)
		Sample: 1981q3-2007q3 (all tax and monetary policy regime variables present)			
3	2	ln(HPRENT) = 0.914841-.222595*ln(RUSER)**	597.9308	0.222391	0.135384
		(-15.77)	(-11.19868)	(41.68523**)	(15.27438**)
				(26.41085**)	(15.27438**)
6	1	ln(HPRENT) = 0.941008 -.161188*ln(RUSER)** + 0.982973*ln(LTV_{t-1})**	987.6804	0.264799	0.084278
		(-18.48) (6.29)	(-17.61296)	(42.35350**)	(10.05425)
				(32.29925**)	(9.244480)
		Sample: 1981q3-2001q4 (all tax and monetary policy regime variables present)			
2	1	ln(HPRENT) = 0.773861 -.155038*ln(RUSER)**	514.5196	0.146331	0.024864
		(-4.31)	(-12.11023)	(15.03797^{+})	(2.064622)
				(12.97335^{+})	(2.064622)
5	1	ln(HPRENT) = 0.9191747 -.160483*ln(RUSER)** + 0.850918*ln(LTV_{t-1})**	816.3698	0.289978	0.062066
		(-6.00) (3.77)	(-18.36980)	(34.33252*)	(6.250848)
				(28.08167**)	(5.065014)

	Level (AIC lag in parentheses)	*5% Critical level for lag*	*1% Critical level for lag*	*First Diff. AIC lag in parentheses)*	*5% Critical level for lag*	*1% Critical level for lag*	*Assumptions*
ln(HPRENT)	-2.105505 (1)	-3.450436	-4.042042	-5.198989** (0)	-3.451959	-4.045236	constant/trend
ln(RUSER)	-2.389194 (2)	-3.449716	-4.040532	-4.124070** (1)	-3.449716	-4.040532	constant/trend
ln(LTV)	-2.060300 (3)	-3.451959	-4.045236	-10.78284** (3)	-3.451959	-4.045236	constant/trend

Notes. $^{+}$ (*,**) denotes significant at the 90% (95%, 99%) level. t-statistics in parentheses except when AIC statistic is reported. For vectors numbered 1-6, lag lengths of (1, 3, 1, 3, 6, and 6), respectively, minimized the AIC, yielded clean residuals, and yielded significant and unique vectors allowing time trends in the variables. Lag lengths in the ADF unit root tests based on the Schwartz Information Criterion. Data used span 1979-2007:2.

Table 4: 2nd-Stage EC Models of the Percent Change in the U.S. Home Price-To-Rent Ratio, 1981-2007q3

	No LTV Terms			*LTV Terms*		
	81:3-07:3	81:3-01:4	81:3-07:3	81:3-07:3	81:3-01:4	81:3-07:3
Variable	**Model 1**	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
Constant	0.0014^{+}	0.0127	-0.0217	0.0017^{**}	-0.0554	0.0858^{**}
	(2.40)	(0.34)	(-0.63)	(2.74)	(-1.50)	(2.72)
EC_{t-1}	-0.0122^{*}	-0.0822^{**}	-0.0665^{**}	-0.0630^{**}	-0.1529^{**}	-0.1652^{**}
	(-2.10)	(-2.65)	(-3.88)	(-3.02)	(-4.07)	(-4.88)
$MONEYTARGET_t$	-0.0044	-0.0057	-0.0080^{**}	-0.0032	-0.0061	-0.0062^{+}
	(-1.50)	(-1.39)	(-2.81)	(-0.96)	(-1.55)	(-1.89)
$CAPGAINTAX_t$		0.0068^{**}	0.0040^{**}		0.0066^{**}	0.0061^{**}
		(3.26)	(3.37)		(3.85)	(4.33)
$\Delta 2RegQ_{t-1}$		-0.0050^{**}	-0.0048^{**}		-0.0054^{*}	-0.0047^{*}
		(-2.83)	(-2.88)		(-2.46)	(-2.61)
$LCONFLABOR_{t-1}$		-0.0029	0.0048		0.0120	0.0186^{**}
		(-0.36)	(0.63)		(1.49)	(2.71)
$\Delta FHAFEE_t$		0.0046^{**}	0.0035^{**}		0.0039^{**}	0.0038^{**}
		(3.61)	(2.81)		(3.38)	(-3.50)
$\Delta\ln(HPRENT)_{t-1}$	0.5025^{**}	0.3626^{**}	0.4324^{**}	0.4257^{**}	0.2983^{**}	0.3680^{**}
	(5.46)	(3.52)	(4.91)	(4.25)	(2.97)	(4.11)
$\Delta\ln(HPRENT)_{t-2}$		-0.0799		-0.1229	-0.1469	-0.2127^{*}
		(-0.67)		(-1.08)	(-1.29)	(-2.08)
$\Delta\ln(HPRENT)_{t-3}$		0.1948^{+}		0.0813	0.1040	0.0895
		(1.73)		(0.76)	(0.93)	(0.89)
$\Delta\ln(RUSER)_{t-1}$	-0.0259^{**}	0.0018	-0.0134	-0.0224^{*}	-0.0095	-0.0095
	(-2.71)	(0.10)	(-1.52)	(-2.19)	(-0.56)	(-1.00)
$\Delta\ln(RUSER)_{t-2}$		-0.0152		-0.0105	-0.0169	-0.0103
		(-0.82)		(-0.97)	(-0.92)	(-1.07)
$\Delta\ln(RUSER)_{t-3}$		-0.0005		0.0042	0.0150	0.0105
		(-0.03)		(0.39)	(0.83)	(1.07)
R^2	.644	.639	.718	.688	.727	.811
S.E.	.005215	.004480	.004636	.004881	.003894	.003798
VECLM(2)	3.40	9.21^{+}	2.69	8.84	12.50	9.43
VECLM(8)	3.07	1.32	3.07	6.59	7.26	8.80

* (**,$^{+}$) significant at 95% (99%, 90%) level. t-statistics in parentheses. EC terms from VECMs estimating the long and short-run relationships, corresponding to vector numbers from Table 3. Coefficients on lagged changes on lags longer than t-3 and on LTV ratios are omitted to conserve space.

Table 5: One-Stage Models of the Percent Change in Real U.S. Home Prices, 1980-2001 & 1980-2007

	No LTV Terms			*LTV Terms*		
	80:1-07:3	80:1-01:4	80:1-07:3	80:1-07:3	80:1-01:4	80:1-07:3
Variable	**Model 1**	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
Constant	0.2402**	0.4540**	0.2328**	0.4180**	0.5732**	0.5188**
	(2.97)	(3.99)	(3.14)	(4.78)	(4.41)	(5.92)
LHP_{t-1}	-0.0695**	-0.1467**	-0.0806**	-0.0991**	-0.1588**	-0.1377**
	(-3.38)	(-3.99)	(-4.17)	(-4.78)	(-4.43)	(-6.80)
LY_{t-1}	0.1184**	0.1963**	0.1331**	0.2025**	0.2649**	0.2486**
	(3.05)	(3.41)	(3.07)	(4.76)	(4.77)	(5.65)
$LHSTOCK_{t-1}$	-0.0605*	-0.0919**	-0.0696*	-0.1332**	-0.1615**	-0.1601**
	(-2.20)	(-2.75)	(-2.41)	(-4.24)	(-4.62)	(-5.24)
$LRUSER_{t-1}$	-0.0111**	-0.0144**	-0.0118**	-0.0139**	-0.0210**	-0.0197**
	(-2.79)	(-3.20)	(-3.29)	(-3.68)	(-4.22)	(-5.40)
$LLTV_{t-2}$				0.1541**	0.1585*	0.1777**
				(4.23)	(4.18)	(5.42)
$MONEYTARGET_{t-1}$	-0.0081*	-0.0078*	-0.0098**	-0.0100**	-0.0138**	-0.0136**
	(-2.51)	(-2.29)	(-2.91)	(-3.32)	(-3.81)	(-4.45)
$CAPGAINTAX_t$		0.0064**	0.0042+		0.0053**	0.0049**
		(3.12)	(2.05)		(2.81)	(2.74)
$\Delta 2RegQ_{t-1}$		-0.0019+	-0.0014		-0.0036**	-0.0035**
		(-1.80)	(-1.31)		(-3.30)	(-3.29)
$LCONFLABOR_t$		0.0139+	0.0115+		0.0168*	0.0143*
		(1.98)	(1.72)		(2.34)	(2.24)
$LLTV_{t-2}$-$LLTV_{t-5}$				-0.0500+	-0.0633*	-0.0599*
				(-1.77)	(-2.30)	(-2.37)
ΔLHP_{t-1}	0.9330**	0.7598**	0.8098**	0.8151*	0.6846*	0.6833**
	(8.64)	(6.70)	(7.60)	(7.96)	(6.44)	(7.20)
ΔLHP_{t-2}	-0.0575	0.0353	-0.0153	-0.0454	0.0182	-0.0208
	(-0.43)	(0.25)	(-0.12)	(-0.37)	(0.14)	(-0.19)
ΔLHP_{t-3}	0.1242	0.1119	0.0646	0.1407	0.1391	0.1074
	(1.15)	(1.05)	(0.62)	(1.36)	(1.37)	(1.14)
$\Delta LRUSER_{t-1}$	0.0218*	0.0259**	0.0223**	0.0236**	0.0369**	0.0308*
	(2.40)	(3.03)	(2.86)	(2.82)	(3.63)	(3.88)

$\Delta LHSTOCK_{t-1}$	-0.6429^{+}	-0.9007^{*}	-0.6162^{+}	-0.9051^{**}	-1.1321^{**}	-0.9445^{**}
	(-1.93)	(-2.54)	(-1.89)	(-2.91)	(-3.43)	(-3.24)
ΔLY_{t-1}	-0.1375^{*}	-0.2081^{**}	-0.1772^{**}	-0.1522^{**}	-0.2701^{**}	-0.2532^{**}
	(-2.17)	(-3.17)	(-2.94)	(-2.63)	(-4.29)	(-4.53)
ΔLY_{t-2}	-0.0386	-0.1517^{*}	-0.1013^{+}	-0.0635	-0.1729^{**}	-0.1407^{*}
	(-0.66)	(-2.35)	(-1.76)	(-1.17)	(-2.91)	(-2.75)
R^2	.767	.788	.792	.806	.831	.848
S.E.	0.004111	0.003557	0.003928	0.003750	0.003146	0.003328
LM(2)	0.62	4.58	5.18^{+}	0.56	1.05	2.71
Q(24)	35.88^{+}	35.96^{+}	35.92^{+}	30.23	35.36^{+}	24.84

* (**,$^{+}$) significant at 95% (99%, 90%) level. t-statistics in parentheses.

Table 6: One-Stage Models of the Percent Change in U.S. Price-Rent Ratio, 1980-2001 & 1980-2007

	No LTV Terms			*LTV Terms*		
	80:2-07:3	80:2-01:4	80:2-07:3	80:4-07:3	80:4-01:4	80:4-07:3
Variable	**Model 1**	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
Constant	0.0163	0.0620**	0.0567**	0.0776**	0.1255**	0.1260**
	(1.01)	(2.92)	(3.22)	(3.24)	(5.01)	(5.65)
$LHPRENT_{t-1}$	-0.0171	-0.0979**	-0.0731**	-0.0739**	-0.1520**	-0.1383**
	(-0.97)	(-3.20)	(-3.52)	(-3.09)	(-4.98)	(-4.65)
$LRUSER_{t-1}$	-0.0036	-0.0090*	-0.0117**	-0.0102*	-0.0174**	-0.0193**
	(-0.91)	(-2.00)	(-2.80)	(-2.35)	(-3.75)	(-4.65)
$LLTV_{t-2}$				0.1390**	0.1359**	0.1516**
				(3.45)	(3.91)	(4.52)
$MONEYTARGET_{t-1}$	-0.0070*	-0.0068+	-0.0091**	-0.0071**	-0.0086*	-0.0096**
	(-2.37)	(-1.90)	(-3.27)	(-2.54)	(-2.58)	(-3.81)
$CAPGAINTAX_t$		0.0090**	0.0072**		0.0087**	0.0074**
		(3.79)	(3.87)		(4.07)	(4.45)
$\Delta 2RegQ_{t-1}$		-0.0033*	-0.0033*		-0.0042**	-0.0041**
		(-2.31)	(-2.36)		(-3.23)	(-3.18)
$\Delta FHAFEE_t$		0.0045**	0.0046**		0.0049**	0.0049**
		(3.69)	(3.64)		(4.41)	(4.37)
$\Delta LHRENT_{t-1}$	0.5399**	0.3771**	0.5065**	0.4715**	0.3185**	0.4221**
	(5.32)	(3.93)	(5.48)	(4.80)	(3.64)	(4.99)
$\Delta LHRENT_{t-2}$	0.0126	-0.0129	-0.0446	-0.0446	-0.0677	-0.1066
	(0.04)	(-0.13)	(-0.44)	(-0.41)	(-0.73)	(-1.16)
$\Delta LHRENT_{t-3}$	0.1304	0.2567*	0.1587	0.0883	0.2116*	0.1258
	(1.24)	(2.45)	(1.60)	(0.86)	(2.15)	(1.37)
$\Delta LRUSER_{t-1}$	-0.0096	0.0142	0.0054	-0.0079	0.0181+	0.0088
	(-0.99)	(1.27)	(0.59)	(-0.86)	(1.76)	(1.07)
$LLTV_{t-2}$-$LLTV_{t-5}$				-0.0430	-0.0330	-0.0335
				(-1.31)	(-1.14)	(-1.22)
R^2	.657	.671	.735	.693	.734	.788
S.E.	.005338	.004538	.004688	.005051	.00471	.004199
LM(2)	2.15	4.46	3.72	5.82+	3.95	2.78
Q(24)	25.49	18.07	24.20	26.31	24.64	28.11

* (**,+) significant at 95% (99%, 90%) level. t-statistics in parentheses.

Table 7: One-Stage Models of the Percent Change in Real U.S. Home Prices, Home Improvement Adjusted

	No LTV Terms			*LTV Terms*		
	Regular	Imp.Adj.	Imp.Pop.	Regular	Imp.Adj.	Imp.Pop.
Variable	**Model 1**	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
Constant	0.2328**	0.2543**	0.2643**	0.5188**	0.5648**	0.5954**
	(3.14)	(3.75)	(3.38)	(5.92)	(6.00)	(6.45)
LHP_{t-1}	-0.0806**	-0.0801**	-0.0996**	-0.1377**	-0.1375**	-0.1890**
	(-4.17)	(-4.29)	(-3.40)	(-6.80)	(-6.83)	(-6.66)
LY_{t-1}	0.1331**	0.1436*	0.1515**	0.2486**	0.2724**	0.3071**
	(3.07)	(3.07)	(3.17)	(5.65)	(5.61)	(6.24)
$LHSTOCK_{t-1}$	-0.0696*	-0.0872**	-0.0674+	-0.1601**	-0.1953**	-0.1570**
	(-2.41)	(-2.73)	(-1.71)	(-5.24)	(-5.55)	(-4.19)
$LRUSER_{t-1}$	-0.0118**	-0.0129**	-0.0247**	-0.0197**	-0.0339**	-0.0251**
	(-3.29)	(-3.39)	(-3.00)	(-5.40)	(-3.99)	(-6.15)
$LLTV_{t-2}$				0.1777**	0.1833**	0.2073**
				(5.42)	(5.24)	(5.87)
$MONEYTARGET_{t-1}$	-0.0098**	-0.0106**	-0.0094*	-0.0136**	-0.0150**	-0.0118**
	(-2.91)	(-2.98)	(-2.47)	(-3.66)	(-4.55)	(-3.43)
$CAPGAINTAX_t$	0.0042*	0.0042+	0.0053*	0.0049**	0.0049*	0.0076**
	(2.05)	(1.91)	(2.09)	(2.74)	(2.56)	(3.53)
$\Delta 2RegQ_{t-1}$	-0.0014	-0.0013	-0.0015	-0.0035**	-0.0035**	-0.0039**
	(-1.31)	(-1.16)	(-1.29)	(-3.29)	(-3.10)	(-3.48)
$LCONFLABOR_t$	0.0115+	0.0114	0.0121+	0.0143*	0.0133+	0.0166*
	(1.72)	(1.63)	(1.72)	(2.24)	(1.96)	(2.46)
$LLTV_{t-2}$-$LLTV_{t-5}$				-0.0599*	-0.0621*	-0.0698**
				(-2.36)	(-2.29)	(-2.64)
ΔLHP_{t-1}	0.8098**	0.8156**	0.8103**	0.6833**	0.6936**	0.6558**
	(7.60)	(7.62)	(7.55)	(7.20)	(7.20)	(6.92)
ΔLHP_{t-2}	-0.0153	-0.0133	-0.0113	-0.0208	0.0151	-0.0128
	(-0.12)	(-0.10)	(-0.09)	(-0.19)	(0.13)	(-0.12)
ΔLHP_{t-3}	0.0646	0.0926	0.0905	0.1074	0.1136	0.1066
	(0.62)	(0.89)	(0.87)	(1.14)	(1.19)	(1.15)
$\Delta LRUSER_{t-1}$	0.0223**	0.0243**	0.0247**	0.0308**	0.0339**	0.0338*
	(2.86)	(2.96)	(3.00)	(3.88)	(3.99)	(4.09)

$\Delta LHSTOCK_{t-1}$	-0.6162^{+}	-0.7206*	-0.6713^{+}	-0.9445**	-1.0548**	-0.9660**
	(-1.89)	(-2.13)	(-1.95)	(-3.24)	(-3.42)	(-3.20)
ΔLY_{t-1}	-0.1772**	-0.1981**	-0.2059**	-0.2532**	-0.2812**	-0.3079**
	(-2.94)	(-3.10)	(-3.18)	(-4.53)	(-4.64)	(-5.15)
ΔLY_{t-2}	-0.1013^{+}	-0.0965	-0.1003	-0.1407**	-0.1405*	-0.1566**
	(-1.76)	(-1.57)	(-1.62)	(-2.75)	(-2.54)	(-2.89)
Δ4POP2544			0.5072			1.3111*
			(0.86)			(2.51)
R^2	.792	.794	.793	.848	.846	.855
S.E.	0.003928	0.004129	0.004135	0.003328	0.003537	0.003437
LM(2)	5.18^{+}	3.67	4.48	2.71	4.01	6.12*
Q(24)	35.92^{+}	34.78^{+}	36.54*	24.84	27.00	29.29

* (**,$^{+}$) significant at 95% (99%, 90%) level. t-statistics in parentheses.

Table 7: One-Stage Models of the Percent Change in Real U.S. Home Prices, Home Improvement Adjusted

	No LTV Terms		*LTV Terms*				
	Regular	Imp.Adj.	Regular	Imp.Adj.	Imp.ΔPop.	Imp.Pop.	Imp.ΔPop.Inflation
Variable	**Model 1**	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**	**Model 5**
Constant	0.2328** (3.14)	0.2543** (3.75)	0.5188** (5.92)	0.5648** (6.00)	0.5954** (6.45)		
LHP_{t-1}	-0.0806** (-4.17)	-0.0801** (-4.29)	-0.1377** (-6.80)	-0.1375** (-6.83)	-0.1890** (-6.66)		
LY_{t-1}	0.1331** (3.07)	0.1436* (3.07)	0.2486** (5.65)	0.2724** (5.61)	0.3071** (6.24)		
$LHSTOCK_{t-1}$	-0.0696* (-2.41)	-0.0872** (-2.73)	-0.1601** (-5.24)	-0.1953** (-5.55)	-0.1570** (-4.19)		
$LRUSER_{t-1}$	-0.0118** (-3.29)	-0.0129** (-3.39)	-0.0197** (-5.40)	-0.0339** (-3.99)	-0.0251** (-6.15)		
$LLTV_{t-2}$			0.1777** (5.42)	0.1833** (5.24)	0.2073** (5.87)		
MONEY-$TARGET_{t-1}$	-0.0098** (-2.91)	-0.0106** (-2.98)	-0.0136** (-3.66)	-0.0150** (-4.55)	-0.0118** (-3.43)		
CAPGAIN TAX_t	0.0042* (2.05)	0.0042+ (1.91)	0.0049** (2.74)	0.0049* (2.56)	0.0076** (3.53)		
$\Delta 2RegQ_{t-1}$	-0.0014 (-1.31)	-0.0013 (-1.16)	-0.0035** (-3.29)	-0.0035** (-3.10)	-0.0039** (-3.48)		
LCONF $LABOR_t$	0.0115+ (1.72)	0.0114 (1.63)	0.0143* (2.24)	0.0133+ (1.96)	0.0166* (2.46)		
$LLTV_{t-2}$ $-LLTV_{t-5}$			-0.0599* (-2.36)	-0.0621* (-2.29)	-0.0698** (-2.64)		
ΔLHP_{t-1}	0.8098** (7.60)	0.8156** (7.62)	0.6833** (7.20)	0.6936** (7.20)	0.6558** (6.92)		
ΔLHP_{t-2}	-0.0153 (-0.12)	-0.0133 (-0.10)	-0.0208 (-0.19)	0.0151 (0.13)	-0.0128 (-0.12)		
ΔLHP_{t-3}	0.0646 (0.62)	0.0926 (0.89)	0.1074 (1.14)	0.1136 (1.19)	0.1066 (1.15)		
$\Delta LRUSER_{t-1}$	0.0223** (2.86)	0.0243** (2.96)	0.0308** (3.88)	0.0339** (3.99)	0.0338* (4.09)		

$\Delta LHSTOCK_{t-1}$	-0.6162^{+}	-0.7206^{*}	-0.9445^{**}	-1.0548^{**}	-0.9660^{**}
	(-1.89)	(-2.13)	(-3.24)	(-3.42)	(-3.20)
ΔLY_{t-1}	-0.1772^{**}	-0.1981^{**}	-0.2532^{**}	-0.2812^{**}	-0.3079^{**}
	(-2.94)	(-3.10)	(-4.53)	(-4.64)	(-5.15)
ΔLY_{t-2}	-0.1013^{+}	-0.0965	-0.1407^{**}	-0.1405^{*}	-0.1566^{**}
	(-1.76)	(-1.57)	(-2.75)	(-2.54)	(-2.89)
Δ4POP2544					1.3111^{*}
					(2.51)
R^2	.792	.794	.848	.846	.855
S.E.	0.003928	0.004129	0.003328	0.003537	0.003437
LM(2)	5.18^{+}	3.67	2.71	4.01	6.12^{*}
Q(24)	35.92^{+}	34.78^{+}	24.84	27.00	29.29

* (**,$^{+}$) significant at 95% (99%, 90%) level. t-statistics in parentheses.