Shifting Credit Standards and the Boom and Bust in U.S. House Prices: Time Series Evidence from the Past Three Decades*

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Abstract: The U.S. house price boom has been linked to an unsustainable easing of mortgage lending standards. However, time series models of U.S. house prices ignore changes in mortgage lending standards and perform poorly in the 2000s. We incorporate data on mortgage standards for first time buyers into a model of US house prices based on the (inverted) demand for housing services. Our first time buyer loan-to-value series is weakly exogenous. It captures shifts in the supply of mortgage credit and not expectations of future house price appreciation. Using this series, we estimate a U.S. house price equation that yields a stable long-run cointegrating relationship, plausible income and price elasticities and an improved fit. Our findings suggest that swings in credit standards played a major, if not the major, role in driving the recent boom and bust in U.S. house prices.

JEL Codes: R31, G21, E51, C51, C52.

Key Words: house prices, credit standards, subprime mortgages

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1. Introduction

Standard house price models perform poorly during the recent boom and bust in U.S. housing markets (see Duca, Muellbauer and Murphy (2011), Gallin (2006), and Geanakoplos (2010), *inter alia*). This paper incorporates new data on mortgage lending standards into a time series model of U.S. house prices, based on the demand for housing services. Consistent with our pre-housing bust results from price-to-rent models (Duca, *et al.* 2011) we find that U.S. house prices were substantially affected by shifts in credit standards, in line with theory and the empirical results for other countries where mortgage credit standards were relaxed.[1] In contrast to other time series models of U.S. house prices, we find a stable long-run cointegrating vector, reasonable speeds of adjustment, plausible income and price elasticities, better model fits, and sensible simulations of when house prices may stabilize.

This paper contributes to the existing literature on the determinant of house prices and the role of credit by using time series data on mortgage credit standards to model U.S. house prices over the past three decades, including the recent subprime boom and bust period. The longer time period is helpful since the recent literature – which mainly focuses on the subprime boom and bust period from 2001 to 2010 and exploits differences across areas - does not agree about the importance of changing mortgage credit standards.

We find that downpayment ratios for first-time home-buyers fell sharply during the subprime boom, consistent with the view that lending standards were weakened and larger mortgage loans were made to riskier borrowers, many of whom would previously have been rejected (Mian and Sufi, 2009, 2010). The house-price rises, set in train at the time by very low interest rates and these credit-supply changes, fooled many people into thinking that such increases would continue. Fundamentals began changing in 2003 as interest rates began to return

[1] For example, Cameron *et al.* (2006), Meen (2001) and Muellbauer and Murphy (1997) show that changes in mortgage credit conditions were important in the UK.

to more 'normal' levels, and high rates of building expanded the housing stock, while house prices became increasingly overvalued. As the extent of bad loans became clear, the fundamentals changed again as the supply of credit for all types of mortgages contracted, inducing an unwinding of earlier rises in house prices (Duca, *et al*., 2010).[2]

In our house price models, we proxy mortgage credit standards using American Housing Survey (AHS) data on average loan-to-value (LTV) ratios for *first time* home buyers, the marginal group most affected by credit constraints. We adjust the LTV series for cyclical and other factors and find that the resulting series captures exogenous shifts in credit standards, albeit with some feedback from foreclosures. The crucial point is the adjusted LTV series we use is not simply picking up borrower or lender expectations of recent or future house price appreciation, at least at the national level. Moreover, since we model house prices over three decades, we can show that the LTV series is not just proxying the sub-prime house price boom and bust. Our models estimate stable and significant mortgage credit effects in samples that omit the subprime mortgage boom, since an earlier modest rise in LTV ratios enables us to identify the effect of credit standards.

Other researchers have used changes in LTV ratios averaged over *all* buyers as a proxy for shifting mortgage credit standards. For example, Glaeser *et al*. (2010) found no convincing evidence that movements in the average-buyer LTV ratio on GSE-backed mortgages, which changed only modestly between 2001 and 2005, explained the recent house price boom. Unfortunately, the average-buyer LTV ratio is highly endogenous. It also masks different trends in the LTV ratios of former owner occupiers and first time-buyers, the latter most likely to be credit constrained.[3] Former owners benefitted from the house price boom so their average LTV ratio fell as they rolled over their capital gains into a new property. By contrast, according to our

[2] Geanakoplos (2010) also finds that leverage was an important causal driver of the boom and bust in house prices.
[3] In addition, the average-buyer LTV ratio, calculated using FHFA data, also misses some of rise in first time buyer LTV ratios in the early 2000s since subprime and Alt A mortgages are not included in the dataset.

estimates, the average first-time buyer LTV ratio rose sharply from about 88% in the mid to late 1990's to a peak of 94% in 2005. This is the reason why we measure shifts in mortgage credit standards using the LTV ratio for first-time home buyers, rather than all buyers.

The approach used in this paper, based on the inverted demand for housing services, is more structural than the price-to-rent approach used in an earlier study (Duca, *et al*., 2011). For example, we obtain estimates of the income and price elasticities of the demand for housing that match the central estimates in the literature. The house price-to-rent approach depends on arbitrage between highly substitutable rental and owner-occupied properties. Good rent data are also required (Glaeser and Gyourko, 2007). This means that, unlike the housing demand results in this paper, our earlier house price-to-rent model results cannot be readily compared with results for countries with a small private rental market and/or large government intervention.

This study is organized as follows. Section 2 briefly reviews the recent literature on the role of changing mortgage credit standards in the recent house price boom and bust. Our model of house prices, based on the demand for housing services, is set out in Section 3. We describe and justify our LTV based measure of mortgage credit standards in Section 4. The other data series we use are discussed in Section 5. Vector error correction estimates of the long run properties of the demand for housing are discussed in Section 6. Various robustness checks, including models using more volatile CoreLogic house price data, are considered in Section 7. The possible path of future house prices is examined in Section 8. Finally, Section 9 summarizes our findings and draws some conclusions.

2. Review of Recent Literature

Previous studies, such as Coleman *et al.* (2008), Wheaton and Nechayev (2008) and Glaeser *et al*. (2010), have found little role for market fundamentals in explaining the run up in house prices in the early to mid-2000s. However, there is disagreement about the contribution of easier mortgage credit standards to higher house prices. On the one hand, Coleman *et al.* (2008) did not find any association of house prices with the subprime share of mortgages and the mortgage approval rate.[4] On the other hand, Wheaton and Nechayev (2008) show that the estimated deviations of house prices from 'fundamentals' were larger in MSAs with substantial subprime lending, whilst Pavlov and Wachter (2011) show that house prices rose faster in areas with a higher incidence of subprime and other non-traditional lending. Dell'Ariccia *et al.* (2008) find evidence of a decrease in lending standards associated with a substantial increase in the number of loan applications, especially in areas with many competing lenders. Dagher and Fu (2011) conclude that lightly regulated independent lenders, in particular, lowered their lending standards, and contributed disproportionately to the recent boom and bust in house prices.

The causation between easier credit and high house prices can run in both directions as argued by Brueckner *et al.* (2011), among others. Coleman *et al.* (2008), noting the weakening connection between market fundamentals and house prices after 2003, suggest that price momentum generated a 'bubble' psychology where market participants grew to expect continuing price increases, weakening the perceived risks of subprime lending. Dell'Ariccia *et al.* (2008) and Goetzmann *et al.* (2011) find that past house price appreciation, viewed as a proxy for price expectations, is positively associated with the supply of subprime mortgage credit, as measured by loan approval rates and LTV ratios. In Brueckner *et al.* (2011), past house price appreciation is positively associated with lower credit scores, especially for repeat buyers and refinancers as opposed to first-time buyers. However, none of these papers test whether past appreciation is an appropriate measure of lender expectations of future house appreciation.

[4] However, their focus on subprime mortgages omits Alt-A mortgages—a type of nonprime mortgage that was just as prevalent in the mid-2000s—which were disproportionately used in many areas now suffering from high rates of mortgage default and large price declines. The interpretation of mortgage approval rates can also be complicated by changes in credit standards and by lenders dissuading the likely to be rejected from applying (see Jappelli, 1990).

Mian and Sufi (2009) indirectly test the expectations effect. They consider areas with elastic housing supply, and find no significant difference in the rates of growth in house prices between areas ripe for subprime lending and other 'non-subprime' areas. They conclude that the faster expansion of mortgage credit in subprime areas could not have been driven solely by house price expectations. Instead, they conclude that supply factors, including the relaxation of underwriting standards and the growth of mortgage securitization, played an important role.

3. A Model of the Demand for Housing

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*). Let log housing demand be given by

$$(1) \qquad \ln hs = -\alpha \ln hp + \beta \ln y + z$$

where *hp* = real house price, *y* = real income and *z* = other demand shifters including the real user cost of housing, *uc*, and our LTV based measure of mortgage lending. The own price elasticity of demand is $-\alpha$ and the income elasticity is β. Solving and substituting in for *z* yields:

$$(2) \qquad \ln hp = \frac{1}{\alpha}\left(\beta \ln y - \ln hs - \gamma \ln uc - \delta \ln cc\right)$$

Such inverted demand equations can be explicitly derived from an inter-temporal optimization problem.

Our income measure is an estimate of real per capita permanent income, since households tend to abstract from temporary income fluctuations and buy housing and other durable goods based on expected future income. We base the proxy on non-property disposable income—the sum of labor and transfer income adjusted for temporary tax effects (Blinder and Deaton, 1985). Non-property income is used because it accords with theory and avoids simultaneity bias by omitting property income, which partly reflects current house prices. Our permanent income measure y^p equals the discounted sum of current and expected future non-property income.

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 $uc = hp\left(r + t + d - h\dot{p}^e / hp\right)$, where r is the real after-tax interest rate of borrowing, possibly adjusted for risk, t is the property tax rate, d is the depreciation rate, and $h\dot{p}^e / hp$ is the expected real rate of capital appreciation. Many studies argue that lagged rates of appreciation are good proxies, suggesting a role for extrapolation in the formation of household expectations. We measure real user costs using the lagged annual rate of appreciation over the prior four years, adjusted for the annualized cost of selling a home.

Other factors may also be relevant, given that many borrowers face limits on their mortgages and may be risk averse. These could include nominal as well as real interest rates, demography and proxies for risk, particularly of mortgage default. In the dynamics, lagged price adjustment is plausible, given the inefficiency of housing markets.[5] The level and growth rate of the per capita housing stock also likely helps explain house prices. One reason is that households observing much construction might lower expectations of future appreciation, while another is that house prices adjust to stock and flow disequilibria, for which error or equilibrium correction models are well suited.

[5] 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. We find that user costs based on the four-year lagged appreciation rate outperformed those based on other lag lengths in both our LTV and non-LTV models.

4. Measuring Changes in Mortgage Credit Standards since the Late 1970's

We measure shifts in mortgage lending standards using the average LTV ratio on conventional mortgages for first time buyers. The raw first time buyer LTV data, which is available from 1979, was extracted from the American Housing Survey (AHS) public use data files (Duca, Johnson, Muellbauer, and Murphy, 2012).[6] We then adjusted the raw data for seasonality and 'noise' due to unusually small quarterly samples. Since we want to track exogenous changes in the supply of mortgage credit, we also purged the first time buyer series of significant cyclical effects.

Other good measures of the credit standards facing marginal homebuyers are unavailable. Some indicators have too short a history for time series analysis (e.g., credit scores and mortgage lending standard answers from the Fed's Senior Loan Officer Opinion Survey) or suffer from measurement error issues (e.g., the 'gaming' of FICO credit scores during the subprime boom). In addition, loan-to-income or debt payments-to-income ratios are too endogenous owing to their dependence on cyclical variables like income or interest rates.

The resulting, cyclically adjusted average LTV ratio for first time buyers is shown in Figure 1. Our estimates suggest the LTV ratio shifted up from about 85% in the late 1970s and the 1980s to 94% in 2006 before falling back to about 90% in 2009. The steep rise of the LTV ratio in the early to mid-2000s represents a very sharp easing of mortgage credit standards.[7] Figure 1 also includes the average LTV ratio for all buyers with conforming mortgages which, as noted before, fails to track the relaxation in mortgage lending standards in the subprime boom years.

[6] We only use conventional mortgages when calculating our first time buyer LTV ratio. This is consistent with the conforming, conventional mortgages used in the Freddie Mac repeat sales house price index.

[7] Geanakoplos (2010) examined the LTV ratios for private label subprime and alt-prime mortgages. He shows that, for mortgages with LTV ratios above the median, the LTV ratio rose strongly from 2000, peaked in 2006q2, before plunging below the 2000 level by 2009.

--- Figure 1 About Here ---

We adjust the raw LTV data using a local level, state space model with fixed seasonal factors and a range of explanatory variables, including the change in the civilian unemployment rate, and dummies for the quarter following the September 11, 2001 terrorist attack and last two quarters of 1989, corresponding to the Savings and Loan / Thrift bailout. Income, lagged rates of growth of house prices and interest rates were not significant.

--- Table 1 About Here ---

The estimated model is shown in the first column of Table 1. The results in columns 2 and 3 show that the LTV measure is not proxying expectations of future house price gains or losses. In column 2, forecasts of future house prices changes, generated using a fairly simple but well-fitting reduced form model, are insignificant. Forecasts using different horizons are also insignificant. In column 3, we show that the University of Michigan / Reuters consumer survey responses to the question of whether or not it is a good time to buy a house fail to explain changes in the LTV ratio. Our adjusted first time buyer LTV measure consists of the smoothed level from the model in column (1) of Table 1, netting out the estimated cyclical effects of unemployment and exogenous shock effects of the Savings and Loan bailout and the September 2001 terrorist attacks. We leave in the estimated impact of foreclosures because we interpret them as ultimately arising as a consequence of the earlier easing of credit standards.

The timing of movements in our first time buyer LTV series in Figure 1 match up nicely with well documented financial innovations and changes in housing policy. The rise in the LTV ratio in the early to mid-1990s occurs shortly after Congress directed Fannie Mae and Freddie

Mac to bolster homeownership by providing low down-payment mortgages, either by easing underwriting credit standards or by buying private label MBSs (Gabriel and Rosenthal, 2010).

The major rise in LTV ratios between 2000 and 2005 may be attributed to a combination of factors – financial innovations, regulatory changes and policy changes – which fostered the securitized financing of riskier mortgages, especially high LTV loans. Innovations in structured finance enabled subprime and other non-prime mortgages to be packaged into private-label mortgage backed securities (MBSs), which were sold to a range of investors who believed the risks to be low.[8] Issuance of private-label, residential MBSs soared in the early to mid-2000s, along with issuance of commercial MBSs, both of which plunged in the summer of 2007, when problems with pricing subprime instruments spurred some hedge funds to suspend redemptions (Figure 2). Investors in such securities were ostensibly protected by either purchasing higher-rated tranches of CDOs or by credit default swaps (CDSs), both recent innovations. When subprime loan quality proved worse than expected, investors in CDOs discovered they were less protected than anticipated by the tranche structure. Investors also doubted the viability of major CDS issuers (e.g., AIG) and CDS premiums soared and CDS volumes started falling (Duca, et al., 2010).

--- Figures 2 and 3 About Here ---

As Roe (2011) and Stout (2008) stress, the initial rapid growth in derivatives trading during the early and mid-2000s was spurred by major changes in securities laws in the 2002 Commodity Futures Modernization Act (CFMA). The CFMA made derivative contracts enforceable, and gave derivatives contracts claims on collateral enforceable before a court

[8] Cho, Kim, and Wachter (2011) find that MBS issuance, as well as standard explanatory variables such as income and interest rates, helps to explain metro house prices in the U.S. and South Korea.

decided which claims to honor in a business bankruptcy. As illustrated in Figure 3, interest rate and currency swaps, which existed before CFMA, grew rapidly and maintained much of their early growth through the financial crisis, reflecting the underlying need of many firms to hedge such risks. In contrast, CDSs were nonexistent before CFMA, surged after its passage, but then largely crashed when CDS market participants upwardly reassessed the risks of insuring many new products, such as subprime mortgages and CMBS. Viewed together, Figures 2 and 3 illustrate the linkages between the rises and falls in private label MBS issuance and outstanding CDS's.

Other policy and regulatory changes reinforced the impact of financial innovations. The Basel II capital requirements, announced in 2004, induced banks to buy more investment-grade, private-label MBSs (Blundell-Wignall and Atkinson, 2008). Holdings of private label residential MBSs were also boosted by a 2004 SEC decision to double the 1935 limits on the leverage of brokerage units at investment banks, and by the rise of hedge funds and SIVs funding long positions in nonprime MBS with short-term debt. In addition, Congress set higher homeownership goals for the GSEs, who then bought more private label residential MBSs.[9]

Consistent with a weakening of credit standards in the subprime boom, our LTV series is positively correlated with the share of outstanding mortgages securitized into private-label MBS, which funded nonprime mortgages that had higher average LTV ratios than conforming mortgages (Figure 1). Because the LTV series reflects credit standards on new mortgages, it leads the private MBS share of the stock of home mortgages by about two years. The rise of the LTV ratio through the mid-2000's also coincided with a rise in homeownership, especially among younger households (Bardhan *et. al*, 2009). Since many young households have limited savings, the rise of LTVs for first-time homebuyers in the early 2000's eased credit constraints for the marginal home-buyer, and bolstered the demand for housing.

[9] Frame (2008) estimates imply that the GSEs funded one-quarter of nonprime residential mortgages.

5. Other Data

Apart from our measure of mortgage credit standards, the models use data on house prices and the housing stock, permanent income, and the real user cost of housing.[10] We primarily track nominal house prices using the FHFA purchase-only house price index. This index, which is available since 1991, covers homes bought with Freddie Mac or Fannie Mae mortgages, and omits upwardly distorted price appraisals from mortgage refinancings. We splice the seasonally adjusted FHFA series onto a seasonally adjusted Freddie Mac series that starts in 1970.[11]

For robustness, we also estimated models using the two CoreLogic repeat house price indices, i.e., excluding and including distressed sales. These indices start in 1976, and include Ginnie Mae and private-label mortgages in addition to Fannie Mae and Freddie Mac mortgages. Since the Core-Logic house price indices include both the low-end (FHA, VA, and many subprime mortgages) and the upper-end (jumbo and Alt A mortgages) of the housing market, it displays larger swings than the FHFA index.[12]

We used the CoreLogic house price indices rather than the Case-Shiller repeat sales index because they extend further back in time and cover a more representative sample of the U.S. than the 10-city Case-Shiller index. Real house prices, *hp*, are obtained by deflating the FHFA and CoreLogic house price indices using the personal consumption expenditures (PCE) deflator.

The housing stock is measured using the Flow of Funds' estimates of the replacement cost of household residential housing structures. The real, per capita housing stock (*hs*) is

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[10] We did not find any significant demographic effects in our models, an issue we will examine in the future. Non-housing wealth-to-income ratios were also insignificant, possibly for the reasons discussed by Tracy *et al*. (1999).

[11] Similar in-sample results were obtained using the FHFA and Freddie Mac series. We adopted the spliced series because the Freddie Mac series ends in 2010q4, and we wanted to simulate house prices to 2015.

[12] The larger swings in the CoreLogic indices occasionally generate negative user costs of housing. As a result, we use the level, rather than the log, of real user costs when modeling CoreLogic house prices.

obtained by deflating this series by the price index for housing construction and by the population.

We generally use an estimate of permanent income as our income measure, since households tend to abstract from temporary income fluctuations and buy houses and other durable goods based on expected future income. We base the proxy on per capita real disposable labor and transfer income, adjusted for temporary tax effects.[13] We use labor and transfer income, because we wish to avoid possible simultaneity bias resulting from the correlation between property income and house prices. Our permanent income measure y^p equals the discounted path of expected non-property income.

We estimate equations for the log of ratio of 'permanent' income to current income: $y_t^p / y_t = \left(1 + \sum_{s=1}^{\infty} \delta^s E_t\left(y_{t+s} / y_t\right)\right) \Big/ \left(1 + \sum_{s=1}^{\infty} \delta^s\right)$, a weighted moving average of forward-looking income growth rates. In practice, we used a 10% quarterly discount rate ($\delta = 0.9$), consistent with the view that households view the future with a high degree of uncertainty, and replaced expectations of future income growth rates up to 40 quarters ahead with the filtered estimates from a common 'local level' unobservable components model with two observed economic drivers. These are the four-quarter change in the 3-month Treasury bill yield representing the impact of monetary policy, and the University of Michigan / Reuters survey measure of consumer expectations, which has the advantage of being based on a survey of actual consumers. The unobservable component model means that we do not have to specify the dates of changes in productivity a priori, and that the forecasts of future income growth are only based on information available at the time the forecast was made.

[13] These include the tax surcharges during the Vietnam War, temporary tax cuts in 1975, 2001, 2005 and 2008; but not Blinder and Deaton's (1985) estimates for the phased-in tax cuts of the early 1980s. The details are available upon request.

The user cost of housing (uc) is the sum of the after tax, effective mortgage interest rate, property tax rate and depreciation rate, all from the Federal Reserve Board FRB/US model database, minus the lagged annual rate of house price appreciation over the prior four years. The rate of appreciation was adjusted for cost of selling a home (8% on average, implying 2% in annualized terms).[14] We also adjusted the user cost for the impact of the first-time homebuyer tax credit in 2009.[15] We did this by reducing the user cost of housing by 4.1 percentage points from 2008q4 to 2010q2. This basically treats the tax credit as having an effect proportional to its impact on real user costs facing the marginal (i.e., first-time) home buyer.[16]

Our samples all start in 1983, which eliminates the need to add controls for the monetary targeting regime of 1979-82 or for the non-screening type of credit rationing (Jaffee and Rosen, 1979, and Stiglitz and Weiss, 1981) induced by binding ceilings on bank deposit interest rates under Regulation Q (Duca, 1996, and Duca and Wu, 2009).[17] Our models are estimated over two samples – a 'long' sample ending in 2009:q2 when our LTV series ends and a 'short' sample ending in 2001q3 before the September 11 terrorist attacks affected housing demand and before LTVs jumped in the subprime boom. If the results of models using LTVs are similar across the two samples, then the LTVs are not simply acting as a dummy shift variable for the subprime mortgage boom of the 2000s.

[14] The capital appreciation component of the user cost is $\sqrt[4]{(1-0.08)(hp_{t-1}/hp_{t-17}-1)}$.

[15] The tax credit was the minimum of 10% of the house price or $8,000. This is about 3.3% of the average price of existing single-family homes sold in the four quarters (2007q4 to 2008q3) before most home buyers expected that the credit would become law. First-time buyers, on average, bought homes that were 20% less expensive (AHS, 2005). Applying this 20% adjustment implies that the credit was about 4.1% of the average price of homes bought by first-time home-buyers. The tax credit was passed in 2009q1 and was extended for sales through June 2010.

[16] The start date of 2008q4 compensates for the t-1 dating on the user costs terms in our model.

[17] Similar results were obtained when the sample was extended back to 1981q1 and a measure of the impact of Regulation Q was included in the model.

6. Econometric Results

This Section presents our vector error correction (VEC) estimates of the demand for housing. We estimated VEC models since the data have unit roots and are cointegrated. We also estimated simpler, single equation autoregressive distributed lag (ARDL) models which yielded similar results. To preserve space, we only report the VEC results here - the ARDL results are available on request. The VEC models are estimated in two stages. In the first stage, the long-run cointegrating regressions are estimated using the Johansen (1991, 1995) method. The second stage is akin to second step of the Engle-Granger (1987) procedure. The general (inverted) housing demand equation is:

$$\begin{aligned} \Delta \ln hp_t = {} & \alpha_1 \left(\ln hp_{t-1} - \beta_0 - \beta_1 \ln y^p_{t-1} - \beta_2 \ln hs_{t-1} - \beta_3 \ln uc_{t-1} - \beta_4 \ln LTV_{t-2} \right) \\ & + \sum\nolimits_s \gamma_s \Delta \ln hp_{t-s} + \sum\nolimits_j \gamma_{1s} \Delta \ln y_{t-s} + \sum\nolimits_s \gamma_{2s} \Delta \ln hs_{t-s} + \sum\nolimits_s \gamma_{3s} \Delta \ln uc_{t-s} \\ & + \sum\nolimits_s \gamma_{4s} \Delta \ln LTV_{t-s-1} + \gamma_0 + \sum\nolimits_k \delta_k z_{k,t} + u_t \end{aligned} \qquad (3)$$

This specification includes lagged levels of and first differences terms in the logs of house prices, current or permanent income, the housing stock, the user cost of housing, the lagged loan-to-value ratio for first time buyers - our proxy for mortgage lending standards - as well as other exogenous variables (the z_k's).[18] The equation also includes a random error term *u*.

--- Table 2 About Here ---

The estimates in Table 2 are of the long-run, unique cointegrating relationship, corresponding to equation (2), between house prices, permanent income, the real user cost of housing, the housing stock and our mortgage standards proxy, the adjusted LTV ratio for first time buyers. Columns 1 and 2 exclude the latter variable using short and long sample periods,

[18] In the long run cointegrating regressions, the log LTV ratio is lagged one period, which reflects the time lag between mortgage applications/approvals and home purchases.

respectively, whereas columns 3 and 4 include it. To control for some large short-run outliers in real house price changes emanating from energy price volatility, we included the current and lagged change in log real energy prices as short-run exogenous, stationary variables. Most of the VEC models were estimated with a common lag length of 5, which yielded unique cointegrating vectors in the LTV models and clean residuals.

The LTV models are superior to their non-LTV counterparts in several dimensions. First, while the results for non-LTV model 1 reveal that a unique cointegrating vector can be found in the short, pre-subprime boom sample, the long-run relationship breaks down when the sample is extended to encompass the subprime boom and bust (model 2). In contrast, unique and significant cointegrating vectors can be identified for both samples using the same LTV specification, as models 3 and 4 show. Second, the LTV variable is correctly signed and highly significant, consistent with our view that the adjusted, first time buyer LTV ratio is proxying credit constraints and consistent with the house price-to-rent model results in Duca *et al*. (2011).

Third, including LTV ratios yields more plausible income and price elasticities of demand, especially in the full sample. The inverse of the coefficient on log housing stock, interpretable as the long run price elasticity of demand, ranges between -0.51 and -0.64 in models 3 and 4, respectively. These are closer to the -0.5 average time series estimate reported in Meen (2001, p. 129), and more stable than the -0.31 to -1.01 estimates from the non-LTV models. Furthermore, the implied income elasticity of housing demand – the ratio of the estimated, long run coefficients on log income and log housing stock – ranges tightly between 1.23 and 1.27 in LTV models 3 and 4, compared with the higher, and less stable, 1.53 and 1.28 estimates in non-LTV models 1 and 2. The former estimates are very close to the average time series estimate of 1.3 in Meen (2001).

Another advantage of the LTV over the non-LTV models is that the implied long run / equilibrium levels of real house prices are closer to actual prices in recent years (conditional on the observed first time buyer LTV ratio). This can be seen from equilibrium values constructed from coefficients estimated over the shorter sample period ending in 2001q3 - before the subprime boom - and extended forward in time (Figure 4). Here the 4-quarter lagged, estimated equilibrium real house price from the LTV model lines up more closely with actual real house prices.

--- Figure 4 About Here ---

The LTV models also fit a little better than the non-LTV models - their adjusted R^2's are .01 to .03 higher and standard errors are 8 to 10 percent smaller (lower panel of Table 2). The (unreported) income dynamics are consistent with a moving average of income, whilst the short run house price dynamics are consistent with a positive momentum effect.

Moreover, the speed of adjustment is stable in the LTV models (models 3 and 4) at about 14% per quarter, but not for the non-LTV models (models 1 and 2). In the non-LTV models, the speed of adjustment plunges from 13% in the short sample to 3% in the full sample, which is highly suggestive of a breakdown in the long-run relationships due to the omission of a measure of changing mortgage lending standards The LTV results clearly indicate that changing mortgage lending standards were an important driver of U.S. house prices, especially during the recent house price boom and bust.

As a robustness check, the full sample non-LTV and LTV models were re-estimated using current non-property income (adjusted for temporary tax changes) in place of permanent income in models 5 and 6. The qualitative results are similar. The first time buyer LTV ratio is a highly significant determinant of long-run house prices in model 6, which has a much higher estimated speed of adjustment (13%) than the non-LTV model (8%), as well as having significant evidence of cointegration. Nevertheless, reflecting the greater theoretical and empirical appeal of using permanent income, the LTV model using permanent income (model 4) fits better than its current income counterpart (model 6), and yields much more sensible implied income and price elasticities.

A natural question to ask is whether or not our LTV based measure of mortgage lending standards is really exogenous, in the broad sense. If the LTV series were driven by expectations of future or current house price gains or losses, this would greatly complicate and alter the interpretation of our findings. We address this issue in a number of ways. First, note that our house price models include measures of real user costs based on past appreciation, as well as permanent income that reflects expectations of future income. Second, as noted earlier, the two major shifts in the LTV series coincided with notable changes in public policy and financial practices, consistent with the view that exogenous policy and financial innovations were the main drivers of swings in credit standards. Third, we tested whether survey or constructed measures of expected house price appreciation were significant in the model we used to adjust the LTV ratio.[19] In all cases, these measures were statistically insignificant.

Fourth, the econometric evidence suggests that our LTV series is weakly exogenous for house prices. For example, in the complete VEC system underlying the Column (4) results, the house price error correction term is insignificant (t stat. = -1.17) in the LTV equation, indicating that the LTV ratio is weakly exogenous to the other variables (Urbain, 1992), as is the case for the real user cost (t stat. = 0.34) and permanent income (t-stat. = 0.82). Only the housing stock is not weakly exogenous (t-stat. = -2.94), reflecting the fact that housing supply tends to respond with a lag to movements in housing demand. Thus, consistent with theory, equilibrium house prices are indeed driven by shifting credit standards as well as by changes in income, user costs and the housing stock. Finally, as another check, we added model-based measures of expected

[19] Expectations of future house price appreciation over several horizons were generated from regressions using lags of real income, real house prices, nominal mortgage rates, real house price appreciation, and real income growth.

house price appreciation from the earlier section over different horizons to the models in Table 2. None of the expected house price terms were significant and their inclusion did not alter the qualitative or quantitative results, especially with respect to the LTV variable.

7. Additional Robustness Checks

As an additional robustness check, we also estimated inverted demand models using two CoreLogic house price indices series, one including distressed sales and the other excluding them. The CoreLogic series are based on repeat sales of homes financed with any mortgage securitized by Fannie Mae/Freddie Mac (essentially the FHFA series), GNMA (including FHA and VA mortgages), as well as private-label RMBS issuers (which covered the bulk of subprime and Alt A mortgages, and many 'jumbo' mortgages). Distressed sales include those of homes in foreclosure or under the threat of foreclosure (e.g., 'short-sales'), and may affect house price expectations. Up to 2007, the two CoreLogic series moved closely together. Since then the series including distressed sales has been more volatile.[20] In addition, both CoreLogic series exhibit much larger price swings than the FHFA series. As a result, lagged house price appreciation rates calculated using the CoreLogic series are much higher in the mid-2000s, rendering the real user cost of housing negative in a number of quarters. Since we cannot take the log of a negative real user cost, we estimated inverted housing demand models for the CoreLogic series using the same setup as in Table 2, apart from substituting the *level* of the real user cost for the *log* of the real user cost.

In Table 3, models 1 and 2 of CoreLogic house prices excluding distressed sales, and models 3 and 4 of the series including distressed sales, correspond to the model 2 and 4 full-sample FHFA house price models in Table 2. The FHFA house price results in models 5 and 6,

[20] House prices from non-distressed sales tend to be higher than from distressed sales, since homeowners consume housing services until they sell their homes and, as result, tend to maintain them better than unoccupied repossessed homes.

with the real user cost entering in levels rather than logs, may be directly compared with the model 1 to 4 CoreLogic results.

--- Table 3 About Here ---

It is reassuring that, using both the distressed and non-distressed CoreLogic house price series, our mortgage lending standards proxy – the first time buyer LTV ratio - is highly significant and exogenous in models 2 and 4, and results in a much faster speed of adjustment than in its non-LTV counterpart (e.g., 8.6% in model 4 versus 4.8% in model 3). The implied long-run income and price elasticities in the CoreLogic LTV models are also close to those in Table 2. The standard errors of the CoreLogic LTV models are higher than their counterparts using the FHFA series, reflecting the greater volatility of the CoreLogic house price series. Note also that the log housing stock is insignificant in the cointegrating vector for the non-LTV model.

The FHFA model 5 and 6 results price series, using the level rather than the log of the real user cost, may be directly compared with the CoreLogic results. There is stronger evidence of cointegration in the LTV model, with only mixed evidence that a significant unique cointegrating vector could be found in the non-LTV model. In addition, all of long-run determinants of house prices have statistically significant coefficients in the LTV model, while the real user cost of housing is insignificant in the non-LTV model. Overall, it is very encouraging that the qualitative results for both distressed and non-distressed CoreLogic house prices in Table 3 match those in Table 2.

8. Where Are House Prices Heading?

In our housing demand models, we identified and estimated a unique, long run cointegrating vector or equilibrium relationship, implying that house prices should, in principle, head back over time towards this equilibrium. Likewise, the extent of overvaluation (or undervaluation) of house prices may be measured by the deviation of current house prices from their long run, equilibrium level. In practice, care must be taken when using the estimated cointegrating vector for this purpose, since the user cost of housing - one component of the long run equilibrium - is a function of the lagged rate of change in house prices. Thus, the forecast level of house prices in 2015q1 depends on forecast rate of change in house prices between 2011q4 and 2014q4.[21]

In sample, we can condition on the observed user cost of housing and obtain a partial answer to the over-valuation question. Consider the long run equilibrium of our preferred model, Table 2, model (4): $\ln hp_t = 5.02 - 0.24 \ln uc_t + 2.47 \ln y_t^p - 1.95 \ln hs_t + 0.66 \ln LTV_{t-1}$. Conditional on $\ln uc_t$, $\ln hp_t - (5.02 - 0.24 \ln uc_t + 2.47 \ln y_t^p - 1.95 \ln hs_t + 0.66 \ln LTV_{t-1})$ is the deviation from equilibrium. Using this measure, real house prices were over-valued in 2009q2 by about 6%.[22]

Out of sample, we can simulate the likely future path of house prices treating the user cost of housing as endogenous. To do this, we have to specify reasonable values for future income, the housing stock, interest rates, loan to value ratios, etc. We assume that the first time buyer LTV ratio remains at its 2009q2 level, which is similar to its level in 2002 and close to the average prevailing over the late 1990s. For the other variables, we use actual data through 2011q4. From 2012q1 onwards, non-property income is assumed to grow in line with the average personal income growth forecasts from the May 2012 Blue Chip Economic Indicators survey of forecasters. Our assumed path for the real housing stock is based on private sector projections that total housing starts would return to a long-run equilibrium pace of 1.4 million units along a linear path by 2014q4, up from the 0.6 million unit pace of late 2011.[23] We

[21] Feedback effects from house prices to consumption and investment may be quantitatively significant at times.
[22] The estimated overvaluation is 9% in the non-LTV model.
[23] We assume that annualized housing starts were 0.6 million units through 2010q4, and rise to 1.4 million units in 2014q4. The shortfall of starts from their long-run pace (1.4 million units) averaged 0.8 million per quarter over

assumed that the tax credit for home-buyers expired for good in 2010q2, and that depreciation and other tax variables stayed at their 2012q1 levels. The future mortgage interest rate path was based on average of forecasts in the June 2012 Blue Chip Financial Forecasts.

--- Figure 5 About Here ---

The simulated path of nominal house prices is shown in Figure 5. Under the assumptions set out above, nominal FHFA house prices may fall 5 percent further from their 2011q4 levels before hitting bottom in 2012q3 (Figure 5).[24] In the simulations, the nominal level of house prices only reverts to its 2007q2 subprime boom peak in late 2014 / early 2015. The time path in Figure 5 is not surprising given the 'bubble builder' and 'bubble burster' (equilibrium correction) dynamics in our house price model (Abraham and Hendershott, 1996).

We also ran other simulations in which we varied the paths of mortgage interest rates and income. These had only minor effects on the contours of the simulations. Only sizable shifts in the assumed time path of the LTV ratio made any notable differences to the simulated path of house prices. For these reasons, we believe that the simulation results are reasonably robust even though we treat housing supply as exogenous.

The simulations, of course, are based on projections of house price determinants which are hard to predict and should be treated with caution. One source of such uncertainty stems from changes in public policy. *Inter alia,* changes in foreclosure policies and/or federal mortgage programs could affect the speed at which house prices adjust.[25]

2008q3 to 2010q2. Lagging one quarter for time to build lags, the real per capita housing stock fell by about 0.238/0.8 percent per 0.1 million shortfall in housing starts. Using this ratio and housing start projections, the real per capita housing stock falls until 2013q1. Thereafter, we assume it rises with real per capita income.

[24] This result is not far from the simulated 5 percent decline in nominal house prices that we discussed in an early draft of our related house price-to-rent paper (Duca *et al*., 2011).

[25] Before 2008, FHA mortgage size limits were lower than those on conventional mortgages securitized by Freddie Mac and Fannie Mae. In the spring of 2008, the FHA increased the maximum size of the loans it guarantees to the

9. Summary and Conclusions

Modelling the demand for housing in the U.S. over the past three decades, we show that mortgage credit standards for first-time home-buyers are important determinants of house prices, along with income, real user costs and the housing stock. Our first-time buyer loan-to-value series is weakly exogenous and captures shifts in the supply of mortgage credit associated with policy changes and financial innovations, and not expectations of future house prices.

The findings indicate that swings in credit standards played a major, if not the major, role in driving the boom and bust of real U.S. house prices in the first decade of this century. Long-term movements in mortgage credit standards for the marginal home-buyer have not been tracked, or incorporated into standard time series models of U.S. house prices, before now. This omission means that these models are mis-specified and explains why they perform poorly during the recent house price boom and bust. In contrast, models including a cyclically adjusted LTV measure for first time home-buyers – our measure of mortgage standards - have much better short- and long-run properties. These range from faster speeds of adjustment to better model fits and more sensible and plausible long-run price and income elasticities of housing demand.

Overall, these findings are consistent with the view that many asset bubbles are fueled by unsustainable increases in the availability of credit or the use of non-robust financial practices. Underlying the unsustainable easing of mortgage credit standards during the subprime boom were regulatory changes and financial innovations that temporarily induced the funding of many nonprime mortgages until the underlying risks became clearer. The resulting swings in credit

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limits used by Freddie Mac and Fannie Mae. This effect may not be captured by our LTV measure which omits FHA mortgages, since the number of first time buyers with FHA mortgages in the AHS was very small. In addition, the FHA's share of mortgage originations rose from 15 percent or so before 2008, to about 50 percent by late 2008.

standards were not tracked by standard gauges of credit standards on conforming mortgages, which gave policymakers and investors a misleading picture of stable mortgage credit quality and availability. In related work (Aron, *et al*., 2012, and Duca, Muellbauer, and Murphy, 2012), we show that the recent housing boom and bust had large effects on housing wealth, which together with shifts in the ability to borrow against housing equity, had large effects on consumer spending in the U.S. and contributed to the depth of the Great Recession.

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Table 1: State Space Estimates of a Local Level Model of the First Time Buyer Loan-to-Value (LTV) Ratio

Dependent Variable: *LTV*.
Sample: 1979 Q1 to 2009 Q2.

Regressors	(1)	(2)	(3)
Δ_2Unemployment Rate	-0.0107** (0.005)	-0.0107** (0.005)	-0.0097* (0.005)
Lagged Foreclosure Rate, 4 Quarter Ave.	-0.0771** (0.032)	-0.0845** (0.034)	-0.0845** (0.034)
S&L Bailout Dummy	-0.0495*** (0.014)	-0.0486*** (0.014)	-0.0498*** (0.015)
9/11 Dummy	-0.0649*** (0.019)	-0.0626*** (0.020)	-0.0666*** (0.015)
Forecast Change in Log Real House Prices in Next Four Years	-	0.0410 (0.51)	-
Good versus Bad Time to Buy House (U. Michigan / Reuters Consumer Survey)	-	-	0.0001 (0.0014)
Dummies for Small Sample Quarters	Yes	Yes	Yes
Std Error	0.019	0.019	0.019
R^2	0.66	0.67	0.66
Normality Statistic	0.71	0.69	0.45
DW	2.00	2.01	1.94
Box-Ljung Q(11) Statistic	8.10	8.43	6.57
Ratio of Estm'd Level to Irregular Variances	0.094	0.112	0.096

Notes: The unobservable components LTV model consists of a local level, fixed seasonal and regression effects, and an irregular error term. The estimated seasonal and small sample quarter (1981 q4, 1982 q2 to q4, 1983 q4, 1993 q4, 2005 q4, 2007 q4 and 2008 q4) dummy coefficients are not reported to save space. Statistically significant effects at the 1%, 5% and 10% levels are denoted by ***, ** and * respectively. Root mean squared errors are shown in parentheses. The forecast 16 quarter change in real house prices is the fitted value from a reduced form model with the following explanatory variables – the change in real house prices (lags 1 to 4); the lagged 4, 8, 12 and 16 quarter change in real house prices; the change in real income (per capita disposable personal income, lags 1 to 4); nominal mortgage rates (lags 1 to 4) and the first four principal components of the aggregate responses to the University of Michigan / Reuters Survey of Consumers question on the reasons why it is good or bad time to buy a house.

Table 2: Vector Error Correction Models of Log Real U.S. FHFA House Prices

Maximum likelihood estimates of the long-run cointegrating relationship and speed of adjustment in a VEC house price model assuming, at most, one cointegrating vector.

	Permanent Income No LTV		Permanent Income LTV		Current Income No LTV	Current Income LTV
Sample	83q1-01q3	83q1-09q3	83q1-01q3	83q1-09q3	83q1-09q3	83q1-09q3
	(1)	(2)	(3)	(4)	(5)	(6)
Long Run Vector						
β_0, constant	4.371	4.522	5.385	5.019	3.907	4.414
β_1, $\ln y_t^p$	1.512** (4.20)	4.149** (4.86)	1.939** (4.53)	2.468** (9.36)	-	-
β_1, $\ln y_t$ (current)	-	-	-	-	0.987** (3.43)	1.386** (9.20)
β_2, $\ln hs_t$	-0.990** (3.06)	-3.250** (4.08)	-1.572** (4.01)	-1.945** (7.77)	-0.406 (1.63)	-0.873** (5.76)
β_3, $\ln uc_t$	-0.213** (13.91)	-0.223** (7.64)	-0.273** (11.69)	-0.239** (27.50)	-0.204** (16.47)	-0.177** (26.92)
β_4, $\ln LTV_{t-1}$	-	-	0.989** (3.62)	0.656** (4.50)	-	0.745** (5.76)
α_1 'speed of adjustment'	-0.133* (2.58)	-0.034* (2.21)	-0.142** (3.72)	-0.138** (4.02)	-0.081** (2.87)	-0.126** (3.32)
Income Elasticity	1.53	1.28	1.23	1.27	2.43	1.59
Price Elasticity	-1.01	-0.31	-0.64	-0.51	-2.46	-1.15
Cointegration Tests						
Trace, 1 Vector	50.30*	40.91	88.95**	82.28**	44.96^{+}	87.48**
Trace, 2 Vectors	22.73	20.79	44.24	40.83	24.05	39.48
λ Max, 1 Vector	27.56*	20.12	44.71**	41.45**	20.90	48.00**
λ Max, 2 Vectors	18.24	15.21	24.40	18.10	16.94	20.17
Adjusted R^2	0.750	0.888	0.784	0.898	0.882	0.888
SE × 100	0.341	0.369	0.317	0.353	0.379	0.370
LM(1)	16.87	15.07	19.65	27.30	31.21*	26.43
LM(2)	21.16	15.81	32.15	27.96	28.15*	25.62
LM(8)	11.04	13.60	21.33	30.80	17.95	26.37

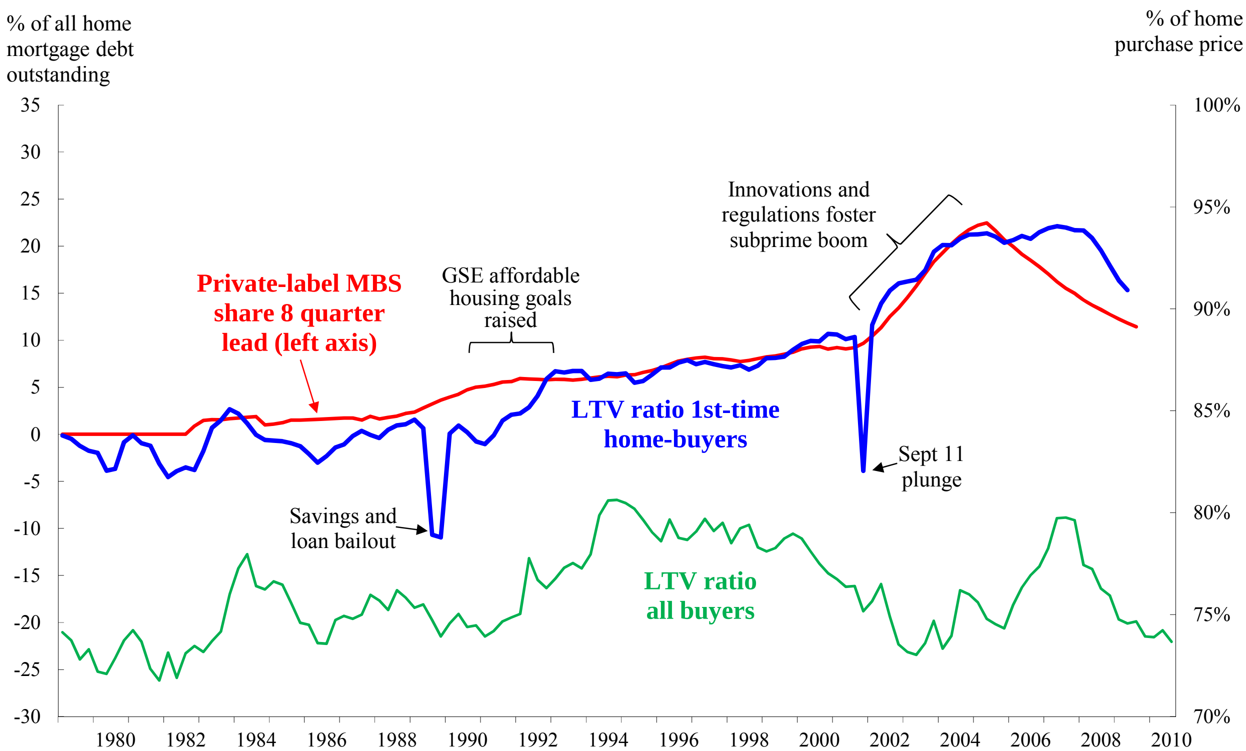
Notes: (i) Within the VEC model, the house price equation is equation (3). The dependent variable is $\Delta \ln hp$, and the long run cointegrating relationship is $\ln hp_t = \beta_0 + \beta_1 \ln y_t^p + \beta_2 \ln hs_t + \beta_3 \ln uc_t + \beta_4 \ln LTV_{t-1} + u_t$. (ii) Statistically significant effects at the 5% and 10% levels are denoted by ** and *, respectively. Absolute t-statistics are shown in parentheses. (iii) The LM statistics are system Lagrange Multiplier tests statistics for 1st, 2nd and 8th order AR/MA autocorrelation. (iv) Apart from model 5, the VEC models were estimated using 5 lags of the first differenced terms, while 4 lags were needed to identify a cointegrating vector in model 5 with sensible signs on long-run income and housing stock coefficients. (v) In order to save space, the estimated coefficients on the lags of the first difference terms, as well as the *first differences of log real energy prices,* are not reported. A complete set of estimations results is available upon request.

Table 3: Vector Error Correction Models of Log Real CoreLogic and FHFA House Prices (Note: The real user cost of housing is not logged.)

Maximum likelihood estimates of the long-run cointegrating relationship and speed of adjustment in the VEC house price model assuming, at most, one cointegrating vector.

	CoreLogic Prices Excld. Distressed		**CoreLogic Prices Incld. Distressed**		**FHFA Prices Semi-Log User Cost**	
	No LTV	LTV	No LTV	LTV	No LTV	LTV
Sample	83q1-09q3	83q1-09q3	83q1-09q3	83q1-09q3	83q1-09q3	83q1-09q3
	(1)	(2)	(3)	(4)	(5)	(6)
Long Run Vector						
β_0, constant	1.959	5.145	1.777	4.790	6.934	6.915
β_1, $\ln y_t^p$	3.269*	2.554**	2.487^{+}	2.453**	26.569**	1.450**
	(2.28)	(3.34)	(1.80)	(2.74)	(4.19)	(2.78)
β_2, $\ln hs_t$	-1.946	-2.135**	-1.234	-1.959*	-22.928**	-1.589**
	(1.51)	(2.89)	(1.00)	(2.29)	(3.94)	(3.18)
β_3, uc_t	-0.016*	-0.035**	-0.018**	-0.035**	-0.020	-0.049**
	(2.33)	(9.47)	(2.86)	(8.43)	(0.46)	(14.19)
β_4, $\ln LTV_{t-1}$	-	2.625**	-	2.418**	-	2.815**
		(4.74)		(3.81)		(9.27)
α_1 'adjustment speed'	-0.042**	-0.087**	-0.048**	-0.086**	-0.006**	-0.045*
	(3.77)	(4.75)	(3.38)	(4.34)	(3.16)	(2.38)
Income Elasticity	1.68	1.23	2.02	1.25	1.28	0.91
Price Elasticity	-0.51	-0.53	-0.81	-0.51	-1.35	-0.63
Cointegration Tests						
Trace, 1 Vector	56.11*	77.15**	55.40**	74.37*	49.04*	83.74**
Trace, 2 Vectors	29.38	44.18	31.69*	45.26	24.14	42.52
λ Max, 1 Vector	26.72	33.97*	23.71	29.10	24.90	41.22**
λ Max, 2 Vectors	16.52	20.75	18.16	20.37	14.87	23.70
Adjusted R^2	0.874	0.895	0.885	0.901	0.883	0.868
SE × 100	0.541	0.494	0.622	0.577	0.378	0.402
LM(1)	9.74	26.18	10.31	26.95	17.07	25.19
LM(2)	2.35	16.98	4.36	19.42	16.01	19.96
LM(8)	10.38	20.72	12.95	22.09	9.44	29.82

Notes: See notes to Table 2. (i) The user costs of housing is not logged, so the long run cointegrating relationship is $\ln hp_t = \beta_0 + \beta_1 \ln y_t^p + \beta_2 \ln hs_t + \beta_3 uc_t + \beta_4 \ln LTV_{t-1} + u_t$. (ii) The CoreLogic price models in columns (1) to (4) were estimated using 6 lags of the first differenced terms. The FHFA models in columns (5) and (6) were estimated using 6 and 5 lags respectively.



Sources: LTV 1st-time home buyers from American Housing Survey, authors' calculations, and Duca, *et al*., (2012). LTV ratio for all buyers from FHFA, based on conforming mortgages insured by the GSEs. Private-label MBS share from Flow of Funds and authors' calculations.

Figure 1: LTV ratios for first time and all homebuyers.

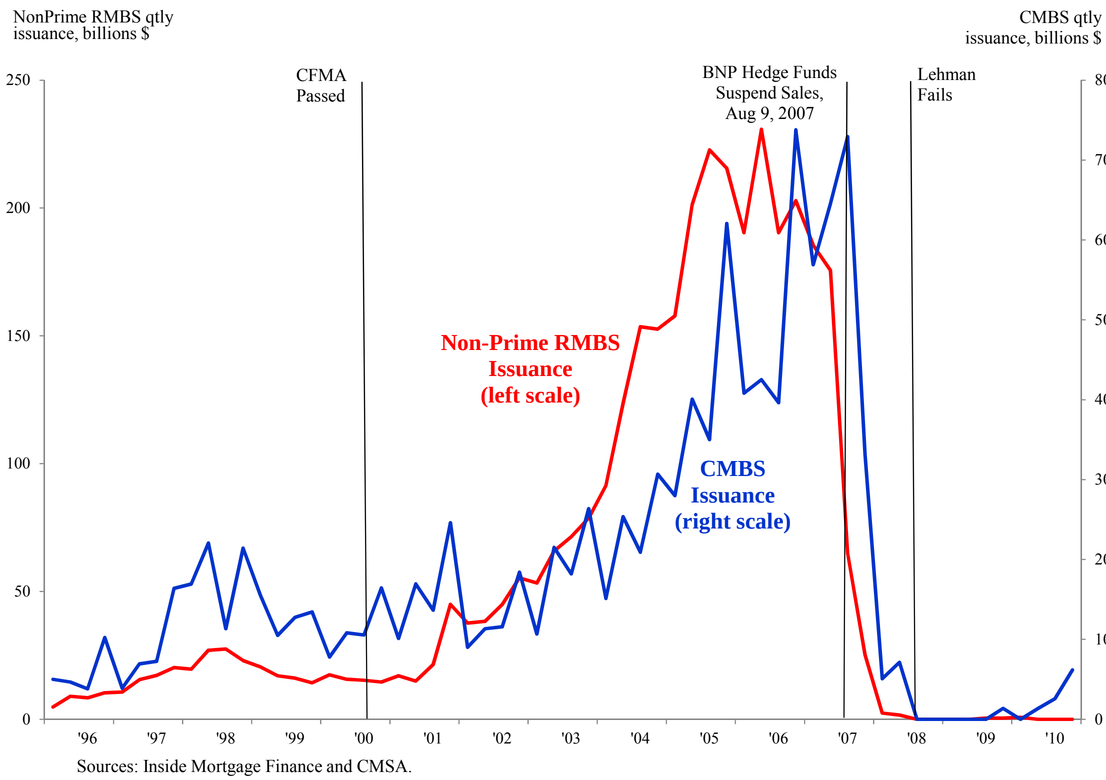


Figure 2: Real CMBS and non-prime RMBS issuances surge in the mid-2000s and plunge in 2007-08.

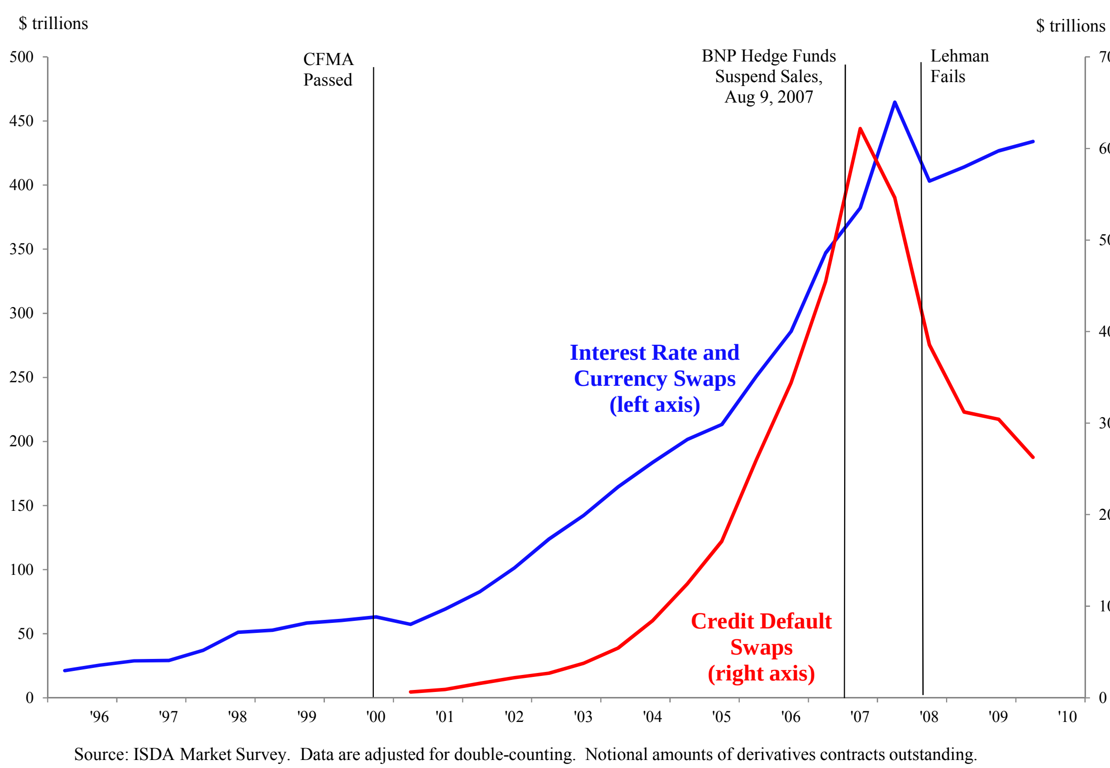


Figure 3: Outstanding derivatives surge after the passage of the 2000 Commodity Futures Modernization Act and CDS's plunge since the Fall of 2007.

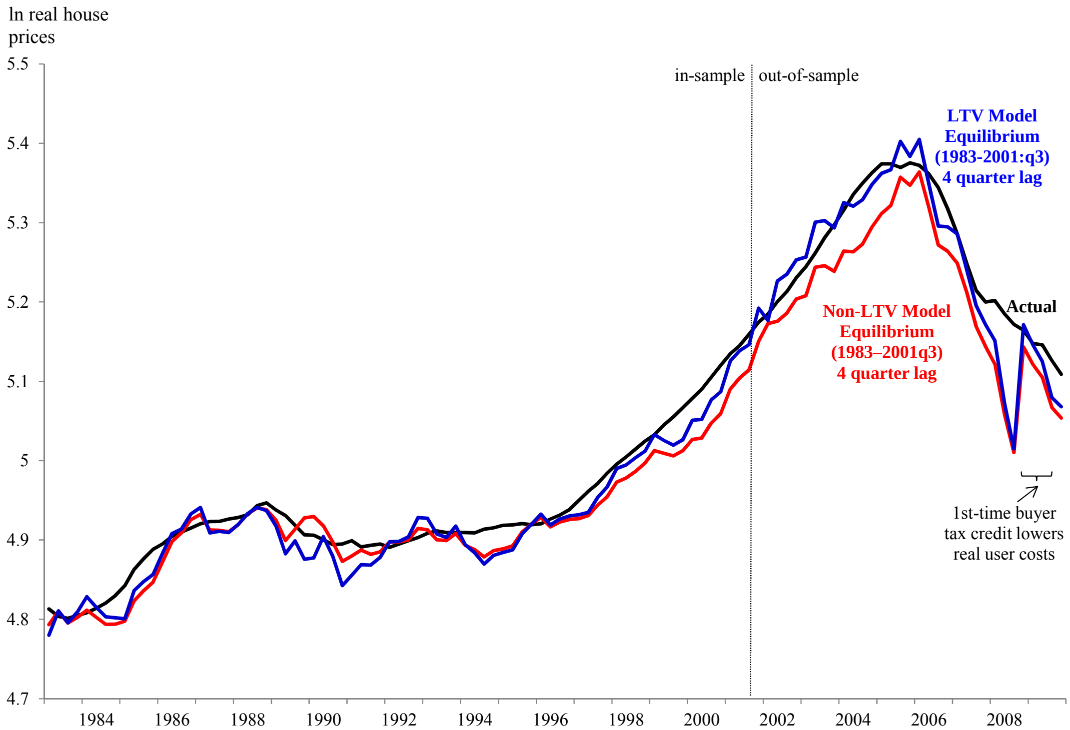


Figure 4: The long run cointegrating relationship from the pre-subprime sample LTV model tracks the house price boom and bust since 2001.

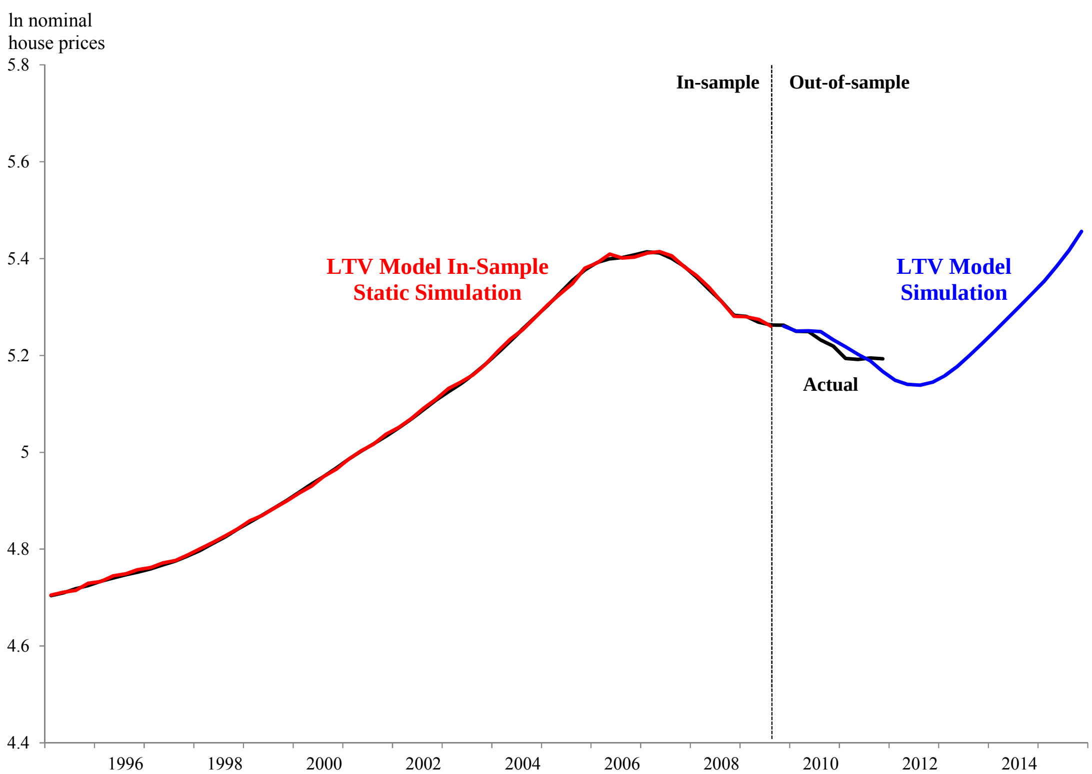


Figure 5: Simulation of nominal FHFA house prices from 2011 to 2015 suggesting that house prices are likely to bottom in 2012 and then recover slowly.