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ECONOMETRIC REGIME SHIFTS AND THE US SUBPRIME BUBBLE

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SUMMARY

Using aggregate quarterly data for the period 1975:Q1–2010:Q4, I find that the US housing market changed from a stable regime with prices determined by fundamentals, to a highly unstable regime at the beginning of the previous decade. My results indicate that these imbalances could have been detected with the aid of real-time econometric modeling. With reference to Stiglitz's general conception of a bubble, I use the econometric results to construct two bubble indicators, which clearly demonstrate the transition to an unstable regime in the early 2000s. The indicators are shown to Granger cause a set of coincident indicators and financial (in)stability measures. Finally, it is shown that the increased subprime exposure during the 2000s can explain the econometric breakdown, i.e. the housing bubble may be attributed to the increased borrowing to a more risky segment of the market.

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Supporting information may be found in the online version of this article.

1. INTRODUCTION

Starting in the late 1990s, the US housing market witnessed a tremendous and unprecedented boom. Real four-quarter growth rates were positive for 10 consecutive years between 1997:Q2 and 2007:Q1. Much of this increase was subsequently reversed, and by 2011 real housing prices were back at their 2001 level. The repercussions of the housing collapse have been enormous and it was one of the causes of the recession that still impairs the global economy. There is a great need to understand US housing price formation and dynamics, in order to develop an 'early warning system', to robustify the institutional framework and to prevent such events from repeating in the future.

Furthermore, housing prices play a key role in transmitting shocks to the real economy. Mortgage equity withdrawal (MEW) represents a channel in which gains from soaring housing prices may be capitalized through an increase in private consumption; see Aron *et al.* (2012) for empirical evidence of how it contributed to the US consumption boom of the early 2000s. Leamer (2007) has argued that housing starts and the change in housing starts are the best leading business cycle indicators. The evolution of housing prices may be one important factor that influences the activity in the building and construction sector, i.e. by increasing the profitability of new construction projects through a Tobin-Q effect (Tobin, 1969).

The surge in home prices over the previous decade was paralleled by dramatic changes in banks' lending practices, and securitization of questionable loans increased substantially. Before 2003, most mortgage originations were prime conforming loans, while the share of subprime and Alt-A mortgages increased steadily after this. At the same time, the share of subprime mortgages and Alt-A mortgages

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that were repackaged and sold as private-label asset-backed securities (ABS)[1] rose from 45% of a total value of about \$215 billion in 2001 to 80% of \$2 trillion in 2005/2006 (Hendershott *et al.*, 2010). The enormous increase in lending to more risky borrowers may have caused US housing prices to shoot away from trajectories consistent with underlying fundamentals. Subprime borrowers typically have very high loan-to-value (LTV) ratios and, given the non-recourse option in many US states, the downside risk of taking up a mortgage is quite low. In combination with very low interest rates—so-called teaser rates—the first couple of years, there was not much to stop people from taking on excessive debt.

For some time there has been a discussion in the academic literature about the econometric modeling of US housing prices. Much of this debate has been concerned with the question of whether US housing prices are determined by so-called fundamentals or not, where typical fundamentals are thought to be variables such as housing rents, household income, the cost of financing or owning a property, along with a supply-side measure. In addition to being an interesting and challenging econometric question, the role of fundamentals in determining housing prices is also relevant for the bubble debate.

According to the definition in Stiglitz (1990), a bubble exists 'if the reason why the price is high today is *only* because investors believe that the selling price will be high tomorrow—when "fundamental" factors do not seem to justify such a price' (Stiglitz, 1990, p. 13). In this paper, I combine this definition with the modeling assumption that fundamental factors—if they exist—are non-stationary economic time series. Given this assumption, housing prices are determined by fundamentals *if and only if* there exists a cointegrating relationship between housing prices and these non-stationary economic variables. This approach opens for several insights that are relevant for discussing whether or not—in the Stiglitz (1990) sense—the evolution of US housing prices over the previous decade is best characterized as a bubble. First, if cointegration can be established over the full sample period as well as for different subsamples, the bubble hypothesis is clearly rejected. Conversely, if no evidence for cointegration can be found, we cannot reject a bubble. That said, this may just indicate that our information set does not include the relevant fundamentals. The intermediate case may be even more relevant: if a cointegrating relationship can be established early in the sample but is lost subsequently, we may suspect a structural break. Even more interesting: if cointegration disappears before the onset of a wider financial crisis, the results can be used to test whether the transition from a stable market with equilibrium correction (no bubble) to an unstable market (a bubble) has predictive power for the wider crisis.

Several researchers have estimated equilibrium correction models for US housing prices, but without necessarily drawing the implications for whether or not there is—or has been—a bubble in the housing market. As the literature review in Section 2 will reveal, the results are diverging, which by itself calls for further research in an attempt to consolidate the evidence. Foote *et al.* (2012) argue that the price increase in the 2000s not even in retrospect can be identified as a bubble. My results, based on a system based as well as a single-equation cointegration analysis, suggest otherwise.

In particular, my results demonstrate that a structural break took place in US housing price formation in the early 2000s. While real housing prices are shown to follow fundamentals both in a price-to-rent framework and in an inverted demand equation prior to this, there is no evidence of such a relationship after the break. My econometric results therefore suggest that the conflicting results in the literature may be explained by the transition from a stable to an unstable (bubble) regime in the early 2000s, and thus the diverging results may be ascribed to the different sample periods considered.

The results from the econometric models are used to construct two regime shift indicators that may be interpreted as 'bubble indicators'. Mikhed and Zemcik (2009b) constructed a similar indicator, in

[1] Loans satisfying the conforming loan limits of the Government-sponsored enterprises (GSEs) are eligible for GSE securitization, while subprime and Alt-A mortgages are not. If resold, these loans are repackaged into ABSs by private-label securitizers. For more details on securitization, see the discussion in Hendershott *et al.* (2010).

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which they defined a bubble as a situation where either housing prices are non-stationary and housing rents are stationary, or where both series are non-stationary and the price-to-rent ratio is non-stationary as well. Compared to that approach, the indicators presented in this paper have the advantage of being directly derived from an econometric model linking housing prices to economic fundamentals. I show that these indicators—which could have been calculated in real time—are able to detect the transition to a bubble regime early in the 2000s. Furthermore, these indicators are shown to Granger cause a set of coincident indicators and financial instability measures.

As a final contribution of this paper, I test whether the transition from a stable to an unstable regime—as detected by the bubble indicators—can be explained by the increased exposure to aggressive lending products. The econometric results suggest that the share of subprime loans is an important contributor in that respect. This leads to an interesting interpretation of the recent turmoil in the US housing market: the housing bubble may be attributed to financial innovation and the extension of aggressive lending products, which again lead to increased distress in the financial system.

As already mentioned, the paper starts with a review of the existing literature on the econometric modeling of US housing prices. The literature review is followed by a discussion of how a traditional life-cycle model for housing may be interpreted within an equilibrium correction framework. In Section 4 I turn to a description of the data and their temporal properties. Section 5 documents a structural break in US housing price formation in the early 2000s. The 'bubble indicators' are presented in Section 6. In the same section, I report results from tests for Granger non-causality between the 'bubble indicators' and a set of financial (in)stability measures and coincident indicators. In Section 7 it is shown that the econometric regime shift—interpreted as a bubble—may be ascribed to the increased exposure to subprime lending. Before ending with some concluding remarks, I discuss the role of expectations over the course of the bubble and test whether there are signs of a structural break in previous periods in Section 8.

2. COINTEGRATION OR NOT: AN UNSETTLED DEBATE

There is no consensus in the literature on the question of whether US housing prices and fundamentals are cointegrated. Some papers have found evidence of cointegration, while others have reached the opposite conclusion. In broad terms, the literature can be divided into two groups: those who consider local differences and large panels and those who look at aggregate time series data. Given the level of aggregation, there are two theoretical approaches that are commonly considered when the relationship between housing prices and fundamentals is studied. The first takes as a starting point an inverted demand equation linking housing prices to income, a measure of the cost of housing and a supply measure. The second approach looks at the relationship between housing prices and rents. The present study uses both approaches, but is confined to an aggregate study of the USA, but a brief summary of the findings from both aggregate and regional analyses seems relevant. Table I summarizes the main results of the papers reviewed in this section.

Meen (2002) adopts a single-equation approach to estimate the fundamental determinants of real housing prices at the national level. Based on a sample covering the period 1981:Q3–1998:Q2, he reports evidence of cointegration between real housing prices, real personal disposable income, real net financial wealth, the real interest rate and the housing stock. The author demonstrates that the estimated elasticities are sensitive to the inclusion of the housing stock variable. In fact, the income elasticity turns negative if the housing stock is omitted from the cointegrating relation.

Based on the Johansen (1988) approach, McCarthy and Peach (2004) estimate a stock–flow model for the US housing market. They find the long-run determinants of housing prices to be the stock of dwellings, non-durables and services consumption—which is used as a proxy for permanent income—as well as the user cost of housing. The variables are all measured in real terms. McCarthy and Peach (2004) conclude that there is no evidence of a bubble in the US housing market when the

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Table I. Results from previous studies regarding cointegration between housing prices and fundamentals

Authors	Linear	No evidence	Nonlinear	Regional	National	Sample
Abraham and Hendershott (1996)	x			x		1977–1992
Malpezzi (1999)	x			x		1979–1996
Meen (2002)	x				x	1981:Q3–1998:Q2
Gallin (2006)		x		x	x	1975:Q1–2002:Q2/1978–2002
McCarthy and Peach (2004)	x				x	1981:Q1–2003:Q3
Clark and Coggin (2011)		x		x	x	1975:Q1–2005:Q2
Gallin (2008)		x		x		1970:Q1–2005:Q4
Mikhed and Zemcik (2009b)		x		x		1978:H1–2006:H2
Duca *et al.* (2011a)	x				x	1981:Q1–2007:Q2
Zhou (2010)			x	x	x	1978:Q1–2007:Q4
Mikhed and Zemcik (2009a)[a]		x		x	x	1980:Q2–2008:Q2/1978–2007
Duca *et al.* (2011b)	x				x	1981:Q2–2009:Q3

Note: The table gives a summary of the main conclusions in the literature on whether US housing prices and fundamentals are cointegrated or not. Strictly speaking, Abraham and Hendershott (1996) do not test for cointegration, but the model they derive may be interpreted within an equilibrium correction framework.

[a] For samples ending in 2006:Q4 and 2008:Q2, Mikhed and Zemcik (2009a) find evidence of cointegration between housing prices and construction wages, while housing prices and fundamentals are not found to be cointegrated for samples ending before this.

model is estimated over the sample 1981:Q1–2003:Q3, and argue that housing prices have risen as a result of higher incomes and low interest rates.

An early contribution to the panel data literature is Abraham and Hendershott (1996), who estimate an equilibrium correction type of model for 30 Metropolitan Statistical Areas (MSAs) using annual data for the 1977–1992 period. They find that housing prices depend on construction costs, disposable income and the real interest rate in the long-run, which supports the main conclusions of the aforementioned papers.

Although several authors have found that US housing prices are determined by fundamentals, Gallin (2006) argues that US housing prices cannot be modeled in an equilibrium correction framework. First, he looks at national housing price data over the sample 1975:Q1–2002:Q2 using a two-step Engle and Granger (1987) procedure. Then, the author considers a panel of annual data covering 95 cities over the period 1978–2002. In neither case does he find evidence of cointegration. The findings of Gallin (2006) contradict the results of Malpezzi (1999), who considered a similar panel and found evidence of cointegration on the sample 1979–1996. The same author (see Gallin, 2008) looks at the relationship between housing prices, rents and the direct user cost of housing for a sample covering the period 1970:Qq1–2005:Q4. Estimating a conditional equilibrium correction model, he shows that there is no evidence of cointegration between housing prices and these fundamentals for the full sample.

The main conclusions of Gallin (2006) are supported by Clark and Coggin (2011) and Mikhed and Zemcik (2009a), who both study the long-run determinants of real housing prices at the national and regional levels. Mikhed and Zemcik do, however, find that a cointegrating relationship may be established if the sample ends in 2006 or later, while no such relationship exists when earlier endpoints are considered.

Mikhed and Zemcik (2009b) use semi-annual data on housing prices and rents for 23 MSAs over the period 1978–2006 and find similar results to those of Gallin (2008). Considering the full sample, they do not find evidence of cointegration between housing prices and rents and conclude that there is a bubble. The authors go further and construct a ‘bubble indicator’ based on the relationship between housing prices and rents using 10-year rolling windows. It is assumed that the indicator takes the value one if prices are $I(1)$ and rents are $I(0)$ over a given time interval, while it is equal to zero for stationary housing prices and either stationary or non-stationary rents. If both housing prices and rents are $I(1)$, the value of the indicator is equal to the p-value from the panel unit root test of Pesaran (2007) on the price-to-rent ratio. In other words, they implicitly assume that—if there is cointegration—the

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CI-vector is $(1, -1)$ between prices and rents. For most of the rolling windows considered, this indicator provides no evidence of cointegration and takes a value well above 0.20, which, strictly speaking, should be interpreted as a bubble using their methodology. An alternative approach to constructing such a 'bubble indicator' will be discussed later in this paper.

Contrary to the many recent papers finding no evidence of a cointegrating relationship between housing prices and fundamentals, Duca *et al.* (2011a, 2011b) argue that the reason why most models of US housing prices break down in the 2000s is the exclusion of a measure of exogenous changes in credit availability. In Duca *et al.* (2011b), it is shown that adding a measure of LTV ratio of first-time home buyers in a model linking housing prices to income, the housing stock and the user cost outperform non-LTV models judged by the interpretation of the estimated elasticities as well as the numerical size of the equilibrium adjustment coefficient. Similar conclusions are reached in Duca *et al.* (2011a), where the relationship between the rent-to-price ratio and the user cost is considered.

Finally, Zhou (2010) uses data for the period between 1978:Q1 and 2007:Q4 to test for linear and, if that is not found, nonlinear cointegration between housing prices, income, the mortgage interest rate and construction costs. To determine whether the variables in the information set are linearly cointegrated, both the Engle and Granger (1987) and Johansen (1988) procedures are employed. Only for the case of Cleveland does the author find evidence of linear cointegration, which is also the case when the Johansen procedure is considered. For the country and six cities, he finds evidence of nonlinear cointegration using the two-step procedure of Granger and Hallman (1991) and Granger (1991).

3. A CONCEPTUAL FRAMEWORK FOR EQUILIBRIUM CORRECTING HOUSING PRICES

As mentioned in the literature review, there are generally two different theoretical approaches that are considered when looking at the relationship between housing prices and fundamentals: the inverted demand approach and the price-to-rent approach. To be clear about the origin of these relationships, I will briefly discuss their relation to the life-cycle model of housing (see, for example, Meen, 1990, 2001, 2002, Muellbauer and Murphy, 1997).

Based on the life-cycle model, the following condition must be satisfied in equilibrium:

$$\frac{U_H}{U_C} = \mathrm{PH}\left[(1-\tau^y)(i+\tau^p) - \pi + \delta - \frac{\dot{\mathrm{PH}}}{\mathrm{PH}}\right] \tag{1}$$

The condition in equation (1) follows from the representative consumer's maximization problem, where $\frac{U_H}{U_C}$ is the marginal rate of substitution between housing, H, and a composite consumption good, C. The condition states that the consumer's marginal willingness to pay for housing services in terms of other consumption goods should be equal to the cost in terms of forgone consumption. The term in brackets is usually labeled the real user cost of housing, which can be split into three different components. The first is the sum of the nominal interest rate, i, and the property tax, τ^p, less tax deductions at a rate τ^y, and corrected for an increase in the overall price level, π. The second component is the housing depreciation rate, δ. The final component is the expected real housing price inflation, $\frac{\dot{\mathrm{PH}}}{\mathrm{PH}}$, with PH denoting real housing prices. The sum of the first two components is often referred to as the real direct user cost of housing, which will be my operational measure of the user cost in the econometric analysis.[2]

[2] I have experimented with alternative measures of the user cost, where I also included expected capital gains as a moving average of the housing price growth over previous years or simply as the last period growth (static expectations). What I found was that the results were somewhat sensitive to the number of lags in the moving-average process. For that reason, and because I have no a priori reason to assume a given structure on the moving-average process, I decided to use the real direct user cost instead. Note that this implies that expectations about future price changes are captured by the lags included in the econometric models. This is similar to Abraham and Hendershott (1996), Gallin (2008) and Anundsen and Jansen (2013).

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Market efficiency requires the following no-arbitrage condition to be satisfied:

$$Q = \text{PH}\left[(1-\tau^{y})(i+\tau^{p}) - \pi + \delta - \frac{\dot{\text{PH}}}{\text{PH}}\right] \tag{2}$$

The expression in equation (2) states that the user cost of housing should in equilibrium be equal to the real imputed rent on housing services, Q. That is, the user cost of a given dwelling should be equal to what it would have cost to rent a dwelling of similar quality (the value of living in the property).

Extensions of the life-cycle model of housing include an explicit role of credit constraints (see for example Dougherty and Van Order, 1982; Meen, 1990, 2001; Meen and Andrew, 1998). In that case, the expression in equation (2) would be augmented with an additional term reflecting the shadow price on a mortgage credit constraint. Some house buyers will always be credit constrained, and the composition of which borrowers are credit constrained and which are not—and hence the average value of this variable—may well change over time. In the following, I shall abstract from this hard-to-observe latent variable by assuming it to be constant over time. However, as will become clear in the empirical analysis, my results of an econometric breakdown in the previous decade may be interpreted in the context of a change in credit constraints—and hence the latent variable—caused by the explosion in subprime lending.[3]

Rearranging equation (2) slightly gives the following equilibrium relationship:

$$\frac{\text{PH}}{Q} = \frac{1}{(1-\tau^{y})(i+\tau^{p}) - \pi + \delta - \frac{\dot{\text{PH}}}{\text{PH}}} \tag{3}$$

The real imputed rent is unobservable, and two approximations are custom in the empirical literature. The first approximation is to assume that the real imputed rent can be proxied by the observed rent, i.e. the unobservable Q is replaced by an observable R in equation (3). Since the user cost takes negative values over the sample period considered in this paper, I shall consider equation (3) on a semi-logarithmic form in the empirical analysis. The expression based on the price-to-rent approach therefore reads:

$$\text{ph} = \gamma_r r + \gamma_{\text{UC}}\text{UC} \tag{4}$$

where lower-case letters indicate that the variables are measured on a log artithmic scale and UC denotes the real direct user cost. In contrast to Gallin (2006), Mikhed and Zemcik (2009b) and Duca *et al.* (2011a), I do not impose a unitary coefficient between housing prices and rents from the outset, since the implied unitary elasticity between housing prices and rents is a testable restriction. Finally, it is not clear a priori whether rents can be considered weakly exogenous with respect to the long-run parameters, which is another testable restriction.[4] The equilibrium correction representation of the price-to-rent model can be expressed in the following way:

$$\begin{aligned}\Delta \text{ph}_t &= \mu + \alpha_{\text{ph}}(\text{ph} - \gamma_r r - \gamma_{\text{UC}}\text{UC})_{t-1} \\ &\quad + \sum_{i=1}^{p} \rho_{\text{ph},i}\,\Delta \text{ph}_{t-i} + \sum_{i=0}^{p} \rho_{r,i}\,\Delta r_{t-i} + \sum_{i=0}^{p} \rho_{\text{UC},i}\,\Delta \text{UC}_{t-i} + \varepsilon_t\end{aligned} \tag{5}$$

[3] Another simplification in equation (2) is that constant transaction costs and risk premium are assumed.

[4] Using the price-to-rent ratio instead (imposing $\gamma_r = 1$ in equation (4) from the outset) does not affect the results in this paper.

where—from theory—we would expect that $ph - \gamma_r r - \gamma_{\mathrm{UC}}\mathrm{UC} \sim I(0)$, i.e. that the variables are cointegrated.

The second approach followed in the literature is to assume that the imputed rent is a function of variables such as income, Y and the housing stock, in which case we have

$$Q = g(Y, H) \tag{6}$$

Inserting equation (6) in equation (3), a log-linear approximation becomes

$$\mathrm{ph} = \tilde{\gamma}_y y + \tilde{\gamma}_h h + \tilde{\gamma}_{\mathrm{UC}}\mathrm{UC} \tag{7}$$

where lower-case letters again indicate that the variables are measured in logs. The transformations and approximations imply that equation (7) may not be very different from the demand part of a demand-and-supply model (see Meen, 2002, for more discussion).

Since the housing stock evolves slowly, it is assumed to be fixed in the short-run, i.e. it is assumed that the short-run supply schedule is vertical. In the short-run, it is therefore assumed that prices clear the market, which again implies that short-run price movements reflect changes in demand. The equilibrium correction representation of equation (7) can be formulated in the following way:

$$\begin{aligned}\Delta \mathrm{ph}_t = {} & \tilde{\mu} + \tilde{\alpha}_{\mathrm{ph}} \left(\mathrm{ph} - \tilde{\gamma}_y y - \tilde{\gamma}_h h - \tilde{\gamma}_{\mathrm{UC}}\mathrm{UC}\right)_{t-1} \\ & + \sum_{i=1}^{p} \tilde{\rho}_{\mathrm{ph},i}\,\Delta \mathrm{ph}_{t-i} + \sum_{i=0}^{p} \tilde{\rho}_{y,i}\,\Delta y_{t-i} + \sum_{i=0}^{p} \tilde{\rho}_{\mathrm{UC},i}\,\Delta \mathrm{UC}_{t-i} + \tilde{\varepsilon}_t\end{aligned} \tag{8}$$

where we would expect that $\mathrm{ph} - \tilde{\gamma}_y y - \tilde{\gamma}_h h - \tilde{\gamma}_{\mathrm{UC}}\mathrm{UC} \sim I(0)$.

Whether the underlying theories represented by equations (4) and (7) are sufficient to explain US housing price formation may be judged by the signs and significance of the estimated long-run elasticities and—in particular—the significance and numerical size of the equilibrium correction coefficients, α_{ph} and $\tilde{\alpha}_{\mathrm{ph}}$, in equations (5) and (8), respectively.

From a theoretical point of view, we expect γ_r in equation (5) to be positive. In equation (8), we expect $\tilde{\gamma}_y$ to be positive and $\tilde{\gamma}_h$ to be negative. In both equations (5) and (8), we expect γ_{UC} and $\tilde{\gamma}_{\mathrm{UC}}$ to be negative. Further, we expect the adjustment coefficients to be negative and significantly different from zero if housing prices are determined by fundamentals. In the case of a bubble, one would not expect the adjustment coefficient to be significantly different from zero—or at least that it would change markedly towards zero relative to the value it takes during a period of equilibrium correction (no bubble) dynamics. This is also consistent with Abraham and Hendershott (1996), who distinguish between the bubble builder (the coefficients on lagged housing price appreciation) and the bubble burster (the adjustment coefficient). If the adjustment coefficient is close to (or equal to) zero, deviations from an estimated equilibrium would be restored very slowly—or not at all. Thus, with reference to Stiglitz's definition of a bubble, I will think of a bubble as a situation in which housing prices and fundamentals are not cointegrated.

4. DATA DESCRIPTION AND TEMPORAL PROPERTIES

As the operational measure of housing prices, I use the housing price index of the Federal Housing Finance Agency (FHFA), which is available from 1975:Q1.[5] To measure housing rents, I use the rent component of CPI as reported by the Bureau of Labor Statistics (BLS).

[5] This housing price index is calculated according to the weighted repeat sales method of Case and Shiller (1987) and is the longest quarterly time series available for US housing prices. For further documentation on how the index is calculated, the reader is referred to Calhoun (1996).

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My operationalization of the user cost uses the effective interest rate measured as a weighted average of the effective fixed and flexible mortgage interest rates. These data are based on the Monthly Interest Rate Survey Data as reported by FHFA. The weights are determined by the origination shares of the different mortgages. This detail is important in order to get a precise measure of the financing cost at an aggregate level, since the share of fixed and flexible rate mortgages has changed quite substantially over the time period I consider.

The sum of the property tax rate and the interest rate is corrected for tax deductions using the marginal personal income tax rate (at twice the median family income). Both tax rates are from the database of the FRB-US model. The final component in the direct user cost is the depreciation rate, which is from the National Income and Product Accounts.[6] The real direct user cost is constructed by subtracting the annual inflation rate measured by CPI for all items.[7]

The income series is the disposable personal income series collected from the St Louis Fed's database FRED, and is expressed in per capita terms. To measure the housing stock, I use the replacement cost of residential structures as calculated by the Lincoln Institute for Land Policy.[8] The housing stock series is also expressed in per capita terms.[9]

All data are seasonally *unadjusted*, with the exception of the disposable income series, which was only available seasonally *adjusted*. In the econometric analysis, I used the unadjusted series and included seasonal dummies in the usual way. Housing prices, rents and disposable income are measured in real terms, where the nominal-to-real transformations have been achieved by deflating with the CPI for all items, less shelter. The real per capita housing stock measure was obtained by deflating the series by a price index for residential structures.[10] A detailed data description is given in Table A.1 in Appendix A (supporting information).

To control for the interest rate uncertainty during the inflation period of the late 1970s, I include a dummy, MT, that is equal to one between 1975:Q1 and 1982:Q3. Duca *et al.* (2011a,2011b) used a similar dummy for a sample starting in 1979:Q4 to control for the monetary targeting period between 1979:Q4 and 1982:Q3. Finally, I follow Duca *et al.* (2011a,2011b) and include a dummy for the Tax Reform Act of 1997, which is not properly accounted for by the user cost (see Duca *et al.*, 2011a,2011b, and Cunningham and Engelhardt, 2008, for more discussion). This dummy, CGT, is set equal to one from 1997:Q4. Neither of these adjustments materially affect the other coefficients in the model, or the main conclusions reached in this paper.[11]

It is well known that standard inference theory in general ceases to be valid if the data are non-stationary (see Granger and Newbold, 1974). Because of this, I started by testing for unit roots using both the Augmented Dickey–Fuller (ADF) test (Dickey and Fuller, 1979, 1981) and the Phillips–Perron (PP) test (Phillips, 1987, Phillips and Perron, 1988). In all cases I started with a lag length of 5 and the optimal lag truncation was selected based on the Aikake information criterion (AIC). The results from these tests are summarized in Table B.1 in Appendix B, while Figures C.1 and C.2 in Appendix C display the series in levels and first differences (supporting information).

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[6] I was only able to collect data for the depreciation rate until 2007:Q3. After this, I have assumed that the depreciation rate remains unchanged.

[7] It should be noted that treatment of owner-occupied housing costs in CPI shifted in 1983. In particular, it shifted from a measure including interest rates to being constructed based on imputed rents. This may have distorted the inflation measure.

[8] See http://www.lincolninst.edu/subcenters/land-values/price-and-quantity.asp and Davis and Heathcote (2007) for details.

[9] In an earlier version of this paper, I used the housing stock series from Moody's analytics. Similar conclusions were reached in that version of the paper. I thank an anonymous referee for suggesting that I change my operative measure of the housing stock.

[10] Again the reader is referred to http://www.lincolninst.edu/subcenters/land-values/price-and-quantity.asp and Davis and Heathcote (2007) for details.

[11] Details are available upon request.

Based on the unit root tests, it is clear that all series are non-stationary. With the exception of the housing stock, which according to the tests has an $I(2)$ component, all series are found to be integrated of first order. With this small caveat in mind, I continue the analysis under the assumption that all series are at most integrated of order one.

5. THE RECENT REGIME SHIFT IN US HOUSING PRICE FORMATION

5.1. Methodological Approach

In this section, I present the results obtained when the two theoretical models outlined in Section 3 are confronted with the data. To test for cointegration, I have used the system based approach due to Johansen (1988, 1991, 1995). As a robustness check, I have also considered a single-equation test. The Johansen method relies on a reparametrization of a vector autoregressive (VAR) model. In the case where we consider a pth-order VAR, the vector equilibrium correction model (VECM)—which forms the basis for inference in the cointegrated VAR (CVAR)—takes the following form:

$$\Delta \boldsymbol{y}_t = \boldsymbol{\Pi} \boldsymbol{y}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i \Delta \boldsymbol{y}_{t-i} + \boldsymbol{\Phi} \boldsymbol{D}_t + \boldsymbol{\varepsilon}_t \tag{9}$$

where $\boldsymbol{y}_t$ is a $k \times 1$ vector of endogenous variables, $\boldsymbol{D}_t$ is a vector of deterministic terms (including a constant) and $\boldsymbol{\varepsilon}_t \sim \text{IIN}(\boldsymbol{0}, \boldsymbol{\Omega})$. With reference to a VAR model, we have that $\boldsymbol{\Pi} = \sum_{i=1}^{p} \boldsymbol{\Pi}_i - \boldsymbol{I}$ and $\boldsymbol{\Gamma}_i = -\sum_{j=i+1}^{p} \boldsymbol{\Pi}_j$, with $\boldsymbol{\Pi}_i$ referring to the coefficient matrix attached to lag number i of the vector $\boldsymbol{y}_t$.

A test for cointegration is then to test for the number of independent linear combinations of the variables in $\boldsymbol{y}_t$ that are stationary, which amounts to testing the rank, r, of the matrix $\boldsymbol{\Pi}$. If $\boldsymbol{\Pi}$ has reduced rank, it can be decomposed as $\boldsymbol{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are matrices of dimension $k \times r$ representing the loading factors and the long-run coefficients, respectively.[12] I follow custom and let a deterministic trend enter the space spanned by the matrix $\boldsymbol{\alpha}$.

When considering the price-to-rent-based model, the vector $\boldsymbol{y}_t$ is a 3×1 vector containing real housing prices, real rents and the real direct user cost. The inverted demand equation is tested based on a slightly modified version of equation (9), since I condition on the housing stock in the cointegration space. To illustrate what this implies in terms of the VECM representation, it is convenient to partition $\boldsymbol{y}_t$ into a vector of endogenous variables, $\boldsymbol{x}_t$, and a vector of weakly exogenous variables, z_t. The VECM can then be written in the following way:[13]

$$\Delta \boldsymbol{x}_t = \boldsymbol{\Pi} \boldsymbol{y}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_{\boldsymbol{x},i} \Delta \boldsymbol{x}_{t-i} + \sum_{i=0}^{p-1} \boldsymbol{\Gamma}_{z,i} \Delta z_{t-i} + \boldsymbol{\Phi} \boldsymbol{D}_t + \boldsymbol{\varepsilon}_t \tag{10}$$

where $\boldsymbol{y}_t = \left(\boldsymbol{x}_t', z_t'\right)'$. Thus, when I consider the inverted demand equation, the vector $\boldsymbol{x}_t$ will contain real housing prices, real per capita disposable income and the real direct user cost, while z_t is a scalar containing the housing stock only. Since the housing stock is assumed constant in the short-run, I impose the additional restriction that $\boldsymbol{\Gamma}_{z,i} = \Gamma_{h,i} = 0 \; \forall \; i$.

[12] An additional assumption is needed to rule out the possibility of $I(2)$. More precisely, with reference to the second differenced VAR, we can write $\boldsymbol{\alpha}_\perp' \boldsymbol{\Gamma} \boldsymbol{\beta}_\perp = \boldsymbol{\xi}\boldsymbol{\eta}'$, where $\boldsymbol{\Gamma} = \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_i - \boldsymbol{I}$, while $\boldsymbol{\alpha}_\perp$ and $\boldsymbol{\beta}_\perp$ are the orthogonal complements of $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ (i.e. $\boldsymbol{\alpha}_\perp\boldsymbol{\alpha}' = \boldsymbol{\beta}_\perp\boldsymbol{\beta}' = 0$) with dimension $(k-r)\times s$. In general, if $s < (k-r)$ then there are $k-r-s$ $I(2)$ trends in the data, so under the assumption of no $I(2)$ trends we must have that $s = k-r$, i.e. there are $k-r$ common stochastic I(1) trends.

[13] See Johansen (1994, 1995) and Harbo *et al.* (1998) for details.

5.2. Results from the VAR Analysis

Given the conflicting results in the literature, I started by exploring the stability of the two theoretical relationships for housing price determination described by equations (4) and (7). Relying on the statistical framework described in the previous section, I first estimated the VECM representation ((9) and (10), respectively) of the two models for a sample ending in 1995:Q4. Then, I sequentially added four new observations until both models were estimated over the full sample period, 1975:Q1–2010:Q4.

I started with a VAR of fifth order. Then I tested down the lag length using AIC. In both models and for all endpoints, the appropriate lag length was found to be five.[14] After this, I tested for cointegration using the trace test of Johansen (1988). Finally, I tested the joint restriction of excluding the deterministic trend from the cointegration space and whether weak exogeneity of the other variables in the VAR could be supported. More precisely, when looking at the long-run relationship between housing prices, rents and the user cost (see equations (4) and (9)), I tested whether rents and the user cost could be considered weakly exogenous with respect to the long-run parameters, while the same test was done with respect to per capita disposable income and the user cost when I tested the inverted demand equation (cf. equations (7) and (10)).

In Tables II and III I have summarized the main results from these recursive theory–data confrontations. Columns 1 and 2 report the estimation endpoint and rank of the $\boldsymbol{\Pi}$-matrix. Conditional on a non-zero rank,[15] the next two columns report the p-value from the likelihood ratio test for overidentifying restrictions and the likelihood value for the given endpoint. The final six (eight) columns report the estimated adjustment coefficient (α_{ph}) and the long-run elasticities, along with their corresponding standard errors.

There are several noteworthy results in Tables II and III. Looking first at the results from the price-to-rent approach (Table II), it is seen that there is strong evidence for one cointegrating vector ($\text{Rank}(\boldsymbol{\Pi}) = 1$) until 2002. Furthermore, the overidentifying restrictions cannot be rejected and the estimated coefficients do not change notably as the estimation endpoint is extended gradually from 1995:Q4 to 2001:Q4. However, when 2002:Q4 is included in the sample, that relationship can no longer be supported ($\text{Rank}(\boldsymbol{\Pi}) = 0$). At the end of the sample, there is evidence of a return of equilibrium correction ($\text{Rank}(\boldsymbol{\Pi}) = 1$). It is noteworthy that the adjustment coefficient is much lower and the other coefficient estimates have changed substantially relative to their pre-break values.

An inspection of the results from the inverted demand approach (see Table III) gives a similar impression: the rank of $\boldsymbol{\Pi}$ drops to zero when 2001:Q4 is included in the sample, and there are some signs of restored equilibrium correction at the end of the sample. However, the estimated coefficients change markedly and the overidentifying restrictions are no longer supported.

It is worth mentioning that the signs of the estimated long-run elasticities in the inverted demand model are theoretically consistent and in accordance with the international empirical literature when the estimation endpoint is set to 2000:Q4 or earlier; see Girouard *et al.* (2006) for an overview of results from international studies. An income elasticity of demand $\left(-\frac{\beta_y}{\beta_h}\right)$ of around one—as found for the pre-break period—is in accordance with Meen (2001) and Duca *et al.* (2011b). I also find that the coefficient on housing rents in the price-to-rent model is close to one and that rents are weakly exogenous. This justifies the a priori restrictions made by Gallin (2006), Mikhed and Zemcik (2009b) and Duca *et al.* (2011b). Figure 1 displays the recursively estimated coefficients from both models when the endpoint is set to 2000:Q4.

From Tables B.2 and B.3 in Appendix B (supporting information), it can be seen that the models are well specified over the stable period. The only exception is that there is some evidence of

[14] With four lags used to construct the inflation rate entering the user cost expression and five lags in the econometric model, the full effective sample covers the period 1977:Q2–2010:Q4.

[15] I have used small-sample adjusted test statistics, and—for the inverted demand approach—I have used consistent critical values from Table 13 in Doornik (2003) for the case of one exogenous variable. A 10% significance level was used as a cut-off.

Table II. Results from recursive CVAR analysis using the price-to-rent approach (cf. equations (4) and (9)), 1977:Q2–T

Endpoint (T)	Rank($\mathbf{\Pi}$)	Test for restrictions	Likelihood	α_{ph}	$se_{\alpha_{ph}}$	β_r	se_{β_r}	β_{UC}	$se_{\beta_{UC}}$
1995:Q4	1	0.1720	904.526	−0.232	0.044	0.998	0.157	−1.319	0.383
1996:Q4	1	0.1721	951.498	−0.233	0.042	1.064	0.151	−1.307	0.377
1997:Q4	1	0.3590	1002.863	−0.227	0.041	1.069	0.152	−1.367	0.379
1998:Q4	1	0.2881	1054.442	−0.229	0.040	1.062	0.148	−1.334	0.369
1999:Q4	1	0.1346	1102.853	−0.225	0.039	1.075	0.145	−1.249	0.365
2000:Q4	1	0.2576	1153.330	−0.199	0.037	1.152	0.164	−1.176	0.409
2001:Q4	1	0.8403	1198.694	−0.169	0.033	1.248	0.192	−1.300	0.470
2002:Q4	0	*	*	*	*	*	*	*	*
2003:Q4	0	*	*	*	*	*	*	*	*
2004:Q4	0	*	*	*	*	*	*	*	*
2005:Q4	0	*	*	*	*	*	*	*	*
2006:Q4	0	*	*	*	*	*	*	*	*
2007:Q4	0	*	*	*	*	*	*	*	*
2008:Q4	0	*	*	*	*	*	*	*	*
2009:Q4	0	*	*	*	*	*	*	*	*
2010:Q4	1	0.3175	1544.514	−0.060	0.012	2.184	0.348	0.059	1.270

Note: This table reports a summary of the main results when the system-based approach of Johansen (1988) is implemented by sequentially adding four new observations to the sample. The first endpoint is 1995:Q4, while the last is 2010:Q4. The endogenous variables in the system are real housing prices, ph, real rents, r and the real direct user cost, UC. A deterministic trend is restricted to enter the cointegration space, while a constant, three centered seasonal dummies and the MT and CGT dummies enter unrestrictedly. An asterisk replaces the results for the given endpoint when Rank$(\mathbf{\Pi}) = \mathbf{0}$.

Table III. Results from recursive CVAR analysis based on inverted demand approach (cf. equations (7) and (10)), 1977:Q2–T

Endpoint (T)	Rank($\mathbf{\Pi}$)	Test for restrictions	Likelihood	α_{ph}	$se_{\alpha_{ph}}$	β_y	se_{β_y}	β_{UC}	$se_{\beta_{UC}}$	β_h	se_{β_h}
1995:Q4	1	0.5402	839.190	−0.213	0.042	1.217	0.310	−1.432	0.469	−1.495	0.469
1996:Q4	1	0.5356	883.629	−0.207	0.040	1.370	0.296	−1.454	0.480	−1.739	0.429
1997:Q4	1	0.3425	930.568	−0.208	0.043	1.216	0.284	−1.403	0.480	−1.443	0.398
1998:Q4	1	0.2283	977.966	−0.209	0.042	1.203	0.276	−1.372	0.466	−1.426	0.387
1999:Q4	1	0.1797	1025.115	−0.207	0.040	1.203	0.272	−1.324	0.448	−1.426	0.381
2000:Q4	1	0.1213	1068.996	−0.179	0.038	1.350	0.315	−1.270	0.518	−1.603	0.443
2001:Q4	0	*	*	*	*	*	*	*	*	*	*
2002:Q4	0	*	*	*	*	*	*	*	*	*	*
2003:Q4	0	*	*	*	*	*	*	*	*	*	*
2004:Q4	0	*	*	*	*	*	*	*	*	*	*
2005:Q4	0	*	*	*	*	*	*	*	*	*	*
2006:Q4	0	*	*	*	*	*	*	*	*	*	*
2007:Q4	0	*	*	*	*	*	*	*	*	*	*
2008:Q4	0	*	*	*	*	*	*	*	*	*	*
2009:Q4	1	0.0001	1383.153	−0.027	0.008	0.687	1.792	1.648	3.556	−0.858	2.284
2010:Q4	2	0.0000	1419.476	−0.029	0.009	0.547	1.680	2.859	3.043	−0.349	2.125

Note: This table reports a summary of the main results when the system-based approach of Johansen (1988) is implemented by sequentially adding four new observations to the sample. The first endpoint is 1995:Q4, while the last is 2010:Q4. The endogenous variables in the system are real housing prices, ph, real per capita disposable income, y and the real direct user cost, UC. A deterministic trend and the housing stock, h, are restricted to enter the cointegration space. A constant, three centered seasonal dummies and the MT and CGT dummies enter unrestrictedly. Consistent critical values when including one exogenous variable in the cointegration space are tabulated in Table 13 in Doornik (2003). An asterisk replaces the results for the given endpoint when Rank$(\mathbf{\Pi}) = \mathbf{0}$.

autocorrelation in the inverted demand model. I find that excluding the trend from the model (a restriction that is supported) removes this autocorrelation and the model is well specified over the entire stable period in that case (see Table B.4 in Appendix B, supporting information).

The results from the system-based cointegration analysis strongly suggest a breakdown of both the price-to-rent model and the inverted demand model in the early 2000s. With reference to the discussion

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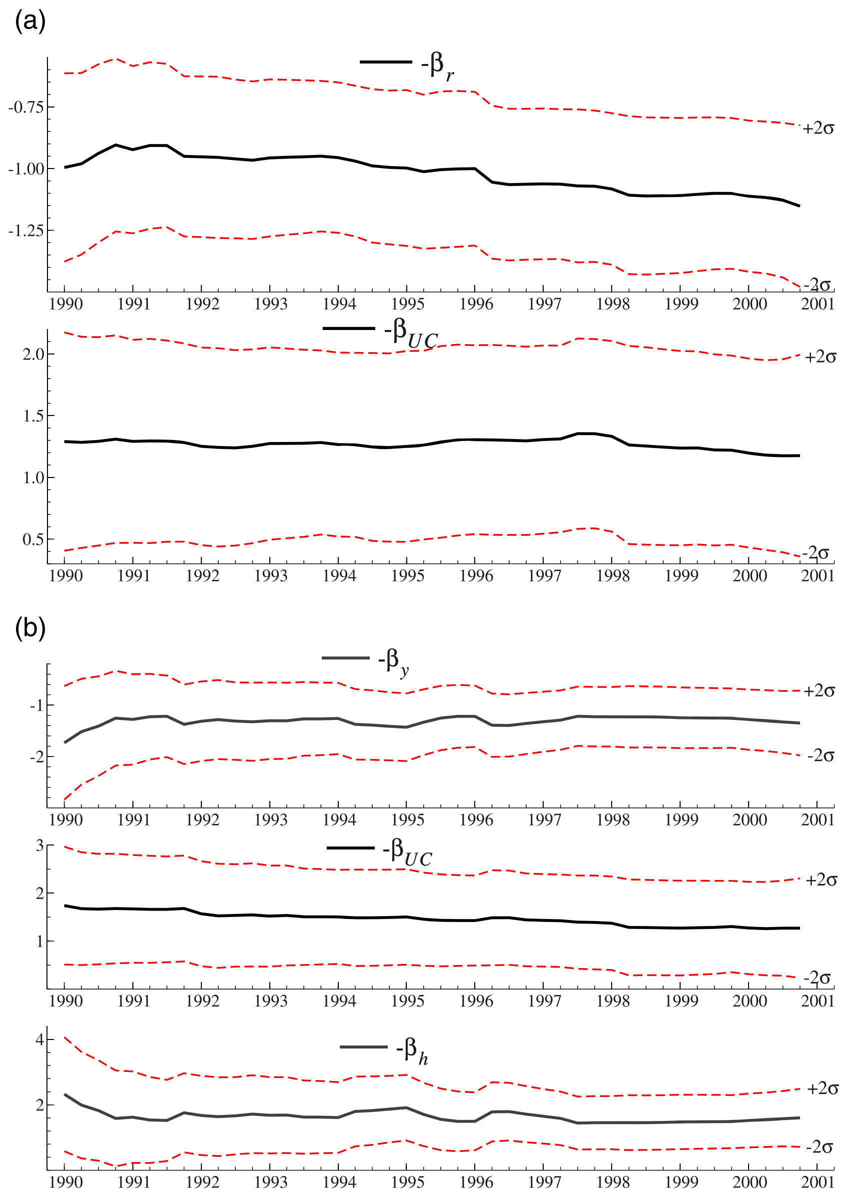

Figure 1. (a) Recursively estimated coefficients for rent and user cost in the price-to-rent model, 1990:Q1–2000:Q4. (b) Recursively estimated coefficients for user cost, per capita disposable income and housing stock from the inverted demand approach, 1990:Q1–2000:Q4

in Section 3, another interpretation of this result is that the latent credit constraint variable remained fairly constant until the early 2000s, but that it changed dramatically after that. In the next section, I will shed some more light on this breakdown, resorting to a single-equation analysis.

5.3. Results from the Conditional Analysis

An alternative approach to testing for cointegration is to estimate equations (5) and (8) directly, and then test the significance of the adjustment coefficient. This follows from the Engle–Granger

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Table IV. Recursive coefficients for price-to-rent model using a single-equation approach (cf. equations (4) and (5)), 1977:Q2–T

Endpoint (T)	β_r	β_{UC}	α_{ph}	p-value	Autocorrelation	Non-normality	Heteroskedasticity	σ	Exp.
1995:Q4	1.164	−0.816	−0.224	0.001	0.3558	0.2184	0.4218	0.0067	0.358
1996:Q4	1.177	−0.796	−0.228	0.000	0.2948	0.2234	0.3259	0.0067	0.337
1997:Q4	1.206	−0.816	−0.219	0.001	0.3749	0.2677	0.2060	0.0067	0.326
1998:Q4	1.200	−0.819	−0.223	0.000	0.3277	0.2470	0.1382	0.0065	0.319
1999:Q4	1.202	−0.819	−0.222	0.000	0.3023	0.1615	0.0766	0.0064	0.318
2000:Q4	1.266	−0.828	−0.203	0.000	0.2861	0.2037	0.0546	0.0064	0.306
2001:Q4	1.409	−1.001	−0.161	0.003	0.4352	0.3051	0.0852	0.0064	0.240
2002:Q4	1.630	−0.909	−0.130	0.005	0.4835	0.1323	0.0639	0.0064	0.238
2003:Q4	1.900	−0.726	−0.105	0.038	0.4757	0.0020	0.0318	0.0069	0.338
2004:Q4	*	*	*	0.589	*	*	*	*	*
2005:Q4	*	*	*	0.848	*	*	*	*	*
2006:Q4	*	*	*	0.660	*	*	*	*	*
2007:Q4	*	*	*	0.129	*	*	*	*	*
2008:Q4	1.919	−1.004	−0.046	0.061	0.7313	0.2221	0.1694	0.0073	0.936
2009:Q4	1.935	−1.470	−0.056	0.013	0.9831	0.3068	0.1135	0.0073	0.970
2010:Q4	2.095	−0.922	−0.061	0.002	0.3098	0.1865	0.3741	0.0075	0.901

Note: This table reports the estimated cointegrating vector along with the loading factor and corresponding p-value when the price-to-rent model is estimated using a single-equation approach. The next four columns report the p-values from tests for autocorrelation, non-normality and heteroskedasticity, as well as the equation standard error. The final column reports the sum of the coefficients on lagged housing price appreciation in the final model. An asterisk replaces the results for the given endpoint when the p-value from the test for cointegration exceeds 0.1.

representation theorem (see Engle and Granger, 1987), which states that equilibrium correction implies cointegration and vice versa.[16]

Since the theoretical models tell us little about the dynamics of housing prices, I have estimated equations (5) and (8) following a *general-to-specific* (Gets) procedure. I used the automatic variable selection algorithm *Autometrics*, which is implemented in PcGive (see Doornik, 2009, Doornik and Hendry, 2009).[17] The lagged levels were restricted to enter the final specification, which ensures theory consistency.[18]

Tables IV and V report the long-run elasticities and the adjustment coefficients along with their finite sample p-values, when I sequentially add four more observations to the sample and use Autometrics to select the relevant variables. The next four columns of the tables report p-values from tests for autocorrelation, non-normality and heteroskedasticity, as well as the equation standard errors. The final column reports the sum of the coefficients on lagged housing price appreciation in the final model, which—given the previous discussion—can be interpreted as measuring an expectation effect.

It is reassuring that these results mimic those I find in the system-based analysis, and the results strongly suggest that the two models for US housing price formation broke down early in the previous decade. The estimated coefficients for the stable period are also close to those I find from the system-based analysis. Furthermore, the same results regarding equilibrium correction are obtained, though this alternative approach seems to support cointegration in the rent-to-price model for a longer period than the system-based approach does. That said, the estimated loading factor changes towards zero already in 2001/2002, which closely resembles the results from the system-based analysis. It is clear that the sum of the coefficients on the retained housing price appreciation terms is positive, which we would expect if these terms are capturing an adaptive expectation channel.

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[16] Ordinary critical values for the t-distribution, however, cannot be used under the null of no cointegration as the distribution of α_{ph} is non-standard and skewed to the left. That said, a program for calculating finite-sample critical values for the conditional equilibrium correction model accompanies the paper by Ericsson and MacKinnon (2002) and is available at http://qed.econ.queensu.ca/pub/faculty/mackinnon/.

[17] This algorithm automatizes the Gets approach and can also handle cases where regressors are not mutually orthogonal. A recent evaluation of the search algorithm is given in Castle *et al.* (2011).

[18] The significance level for the variable selection was set to 5%.

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Table V. Recursive coefficients for inverted demand equation using a single-equation approach (cf. equations (7) and (8)), 1977:Q2–T

Endpoint (T)	β_y	β_h	β_{UC}	α_{ph}	p-value	Autocorrelation	Non-normality	Heteroskedasticity	σ	Exp.
1995:Q4	1.054	−1.183	−0.919	−0.175	0.0240	0.8810	0.5130	0.8467	0.0076	0.628
1996:Q4	1.197	−1.430	−0.905	−0.171	0.0249	0.7493	0.8286	0.9772	0.0076	0.550
1997:Q4	1.104	−1.223	−0.874	−0.175	0.0169	0.6004	0.9420	0.9593	0.0076	0.589
1998:Q4	1.124	−1.251	−0.867	−0.177	0.0119	0.5523	0.9607	0.9098	0.0074	0.577
1999:Q4	1.126	−1.255	−0.864	−0.175	0.0078	0.5424	0.9994	0.8272	0.0072	0.576
2000:Q4	1.208	−1.355	−0.867	−0.161	0.0146	0.4384	0.9886	0.8391	0.0072	0.551
2001:Q4	*	*	*	*	0.2067	*	*	*	*	*
2002:Q4	*	*	*	*	0.3585	*	*	*	*	*
2003:Q4	*	*	*	*	0.9381	*	*	*	*	*
2004:Q4	*	*	*	*	0.9508	*	*	*	*	*
2005:Q4	*	*	*	*	0.9863	*	*	*	*	*
2006:Q4	*	*	*	*	0.9866	*	*	*	*	*
2007:Q4	*	*	*	*	0.7107	*	*	*	*	*
2008:Q4	*	*	*	*	0.7145	*	*	*	*	*
2009:Q4	*	*	*	*	0.8002	*	*	*	*	*
2010:Q4	*	*	*	*	0.5536	*	*	*	*	*

Note: This table reports the estimated cointegrating vector along with the loading factor and the corresponding p-value when the inverted demand model is estimated using a single-equation approach. The next four columns report the p-values from tests for autocorrelation, non-normality and heteroskedasticity, as well as the equation standard error. The final column reports the sum of the coefficients on lagged housing price appreciation in the final model. An asterisk replaces the results for the given endpoint when the p-value from the test for cointegration exceeds 0.1.

5.4. Encompassing the Existing Findings

As discussed in Section 2, the results in the literature show no consensus about the issue of whether an equilibrium correction model can capture the dynamics of US housing prices well or not. There may be several reasons for the divergence of results and my results indicate that the different sample periods used can be one explanation.

In that respect, the results reported in Tables II–V tell an intriguing story:[19] as long as the estimation endpoint is set to 2000:Q4 or earlier, my results suggest that, considering an inverted demand model, housing prices and fundamentals are cointegrated. Interestingly, Meen (2002), Abraham and Hendershott (1996) and Malpezzi (1999), whose samples end prior to this, all reach that conclusion.

However, a researcher estimating the same model for a sample ending in any period between 2001 and 2010 would have been led to the conclusion that an equilibrium correction model cannot possibly explain the fluctuations in US housing prices. That is the case for Gallin (2006), Clark and Coggin (2011) and Zhou (2010), whose sample ends in 2002:Q2, 2005:Q2 and 2007:Q4, respectively. It is interesting to note that while Mikhed and Zemcik (2009a) find evidence of cointegration between housing prices and construction wages for a sample ending in 2006:Q4 but not in 1996:Q4, my results suggest the opposite.

Turning to the price-to-rent approach, neither Gallin (2008) nor Mikhed and Zemcik (2009b) find evidence for cointegration when looking at the relationship between housing prices and rents for samples ending in 2005 and 2006, respectively. This corroborates the findings reported in Tables II and IV.

The above discussion indicates that, to a large extent, the diverging results in the literature can be ascribed to the use of different estimation endpoints. The two studies that stand out from the rest are Duca *et al.* (2011a, 2011b), who document that there is evidence of cointegration in both a price-to-rent model and an inverted demand equation for samples ending in 2007:Q2 and 2009:Q3, respectively. They include a measure of the loan-to-value ratio for first time home buyers in their analysis, which

[19] I compare to both studies that have employed national data and studies that have considered large panels. Although the comparison is not meant to be exact in the sense that start years, operationalizations of the data and test procedures may differ across the studies, it is still interesting to observe that parts of the diverging results in the literature may be attributed to different sample periods.

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may explain why they find cointegration for the period as a whole. Nevertheless, as Figure 1 shows, the cointegrating relations I am able to establish prior to 2001 are very stable when estimated recursively and there is strong evidence of cointegration also prior to this (cf. Tables II–V). With that in mind, another interpretation of the results in Duca *et al.* (2011a, 2011b) is that by conditioning on the LTV ratio they are able to model a structural break.

6. ECONOMETRICALLY BASED REGIME SHIFT INDICATORS

I have constructed two ‘bubble indicators’ (BIs) in the spirit of Mikhed and Zemcik (2009a), but my indicators are based on the relationship between housing prices and fundamentals from recursively estimating and respecifying the models represented by equations (5) and (8) using Autometrics.

I have let my indicators take the values of the finite sample p-values calculated when the variable selection is done recursively quarter-by-quarter all the way back to 1995:Q4.[20] This means that the derived bubble measure is dependent on the extent to which housing prices and fundamentals are cointegrated at different points in time, which can be seen as an operationalization of Stiglitz’s (1990) definition of a bubble. Thus, if we believe that the lack of cointegration corresponds to a bubble (or at least that prices are not responding to deviations from fundamentals in a ‘normal’ way), then any p-value in excess of, say, 10% may indicate a major distortion in the housing market.

Given the data sources and methodology outlined in this paper, my indicators could have been constructed already in 2000 (or earlier) and be used to say something about the temperature in the US housing market in real time, i.e. asserting the role of fundamentals. The two indicators are plotted along with a straight line indicating a 10% (no bubble) significance level in Figure 2.

Although the two indicators are not identical, they both send a quite clear signal already in the early 2000s. In 2004, it is evident that both indicators suggest a bubble in the US housing market. They stay at a high level until 2006, where both start dropping (the price-to-rent-based indicator more so). While the price-to-rent indicator hits the no-bubble line in 2009, that is not the case for that derived from the inverted demand equation. This may either reflect the notion of a negative bubble or simply be the result of the fact that this alternative approach requires more observations to re-establish cointegration.[21]

Again, with reference to the alternative interpretation that the econometric breakdown can be interpreted as a dramatic shift in the previously constant latent mortgage credit constraint variable, the indicators can be thought as measuring the non-constancy of this variable, i.e. that this variable was fairly constant until the early 2000s before changing dramatically.

As a first step to investigate the relevance of these bubble indicators a little further, I have constructed an average indicator that gives equal weight to the two indicators.[22] I then explore whether this composite indicator is leading a set of financial (in)stability measures and coincident indicators. To explore this, I have tested for Granger non-causality (see Granger, 1969). Data definitions for the variables considered are given in Table A.1 in Appendix A (supporting information).

The standard setup to test for Granger non-causality between two variables is to consider a bivariate VAR. The appropriate lag length may be determined by some information criterion. A test for Granger non-causality from one variable to another is to test whether lagged values of this variable

[20] The calculation of finite sample critical values was done using the program accompanying Ericsson and MacKinnon (2002). As they emphasize, the critical values for the conditional equilibrium correction model depends on a number of features such as the sample size, the number of variables in the hypothesized cointegrating vector, what deterministic terms are included, as well as the number of estimated coefficients.

[21] For these indicators to be used as operative measures for the future, they must be ‘brought down’ to no-bubble values before being applied, i.e. cointegration must be re-established. A possible way of doing this would be to put a smaller weight on observations from the bubble of the 2000s when using such measures to monitor the housing market in the future.

[22] Similar results are obtained by considering the two indicators separately, but the composite indicator seems to be a stronger predictor overall.

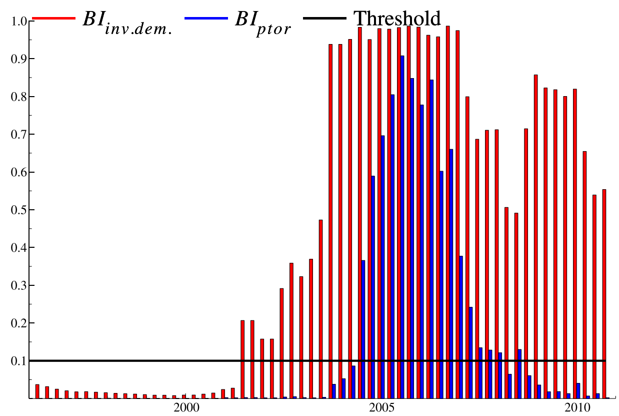

Figure 2. Bubble indicator from price-to-rent approach (blue) and inverted demand approach (red), 1995:Q4–2010:Q4

Table VI. Tests for Granger non-causality

Variable (x)	Lags	Rank($\mathbf{\Pi}$)	$x \to$ BI	BI $\to x$
Unemployment	6	0	0.5010	0.0130
Industrial production	7	0	0.6976	0.2003
Delinquency rates	8	0	0.6293	0.1918
Loan losses	7	0	0.1225	0.0053
Non-performing loans	8	1	0.0005	0.0015
Financial stress index	8	1	0.6706	0.0000
Tightened credit standards	7	1	0.0105	0.0000
Financial conditions index	8	1	0.4127	0.0000
Sample		1997:Q4–2010:Q4		

Note: The table reports the p-values from standard F-tests for Granger non-causality between the composite bubble indicator and a set of financial (in)stability measures and coincident indicators. 'Rank(Π)' signifies the number of cointegrating relationships and 'lags' is the lag truncation for the VAR, which was decided based on AIC. Small-sample corrected critical values have been used for the trace test.

helps predicting the other. That said, several of the variables considered in this paper appear to be non-stationary. For that reason, I start—in the usual way—by determining the optimal lag length by relying on AIC. Then, I test for cointegration between the variables in the VAR.[23] If there is no evidence of cointegration, I consider the variables in first differences. However, if there is evidence of cointegration, I consider the bivariate VAR on VECM form.[24] Cointegration implies Granger causality in at least one direction (Granger, 1986), and in the case of a non-zero rank I move on to test weak exogeneity and the significance of the lagged variables in the VECM jointly.

Initially, I started with a generous lag length of 8. Then I decided the optimal lag truncation based on AIC. Results from these tests for GNC are displayed in Table VI.

The results from the GNC tests suggest that the composite indicator has some predictive power for the different financial (in)stability measures as well as the unemployment rate. There is, however, little evidence of a causal relationship going in the other direction.

This suggests that these indicators can possibly be used—together with other measures—to monitor the risk of financial instability, and in particular the risk of imbalances building up in the housing

[23] Since the sample for the test is relatively short, I used a strict 1% cut-off for the trace test.

[24] It should be noted that, strictly speaking, the trace test indicates a rank of two for the bivariate model with BI and the financial conditions index. This would, however, indicate that both series are stationary, which clearly is at odds with the rest of the results from the GNC tests. For that reason, the rank was set to one also in that case.

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market. The most intriguing finding with regard to the bubble indicators is that they clearly warn of the imbalances in the US housing market at a quite early stage. The relevance of such indicators for monitoring the housing market should, however, be assessed by looking at more countries or possibly by disaggregating to a state or a Metropolitan Statistical Area level for the case of the USA.

7. WAS THE INCREASED SUBPRIME EXPOSURE A CAUSE OF THE BREAKDOWN?

One possible cause of the econometric breakdown documented in Section 5 is that the substantial changes in the subprime market allowed previously constrained and risky borrowers to finance the housing bubble. If that was the case, we should not expect housing prices and fundamentals to be cointegrated. Figure 3 displays the number of subprime loans as a share of total loans serviced by the participants in the mortgage delinquency survey over the period 1998:Q1–2010:Q4.

It is clear from that figure that the explosion in subprime lending comes very close in date to the equilibrium correction breakdown I have documented in the previous sections, with the ratio of subprime loans as a share of total loans going from only 2% in 1998:Q1 to 14% at its peak in 2007.

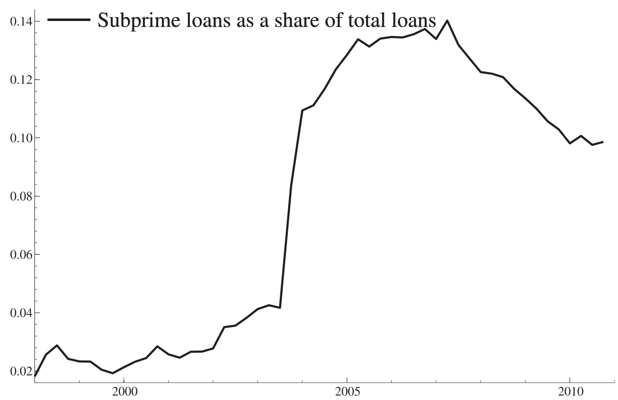

Figure 3. The number of subprime loans as a share of total loans, 1998:Q1–2010:Q4 (Source: Moody's)

Table VII. CVAR analysis for the rent-to-price approach with subprime share in VAR, 1977:Q2–2010:Q4

Eigenvalue	H_0	H_A	λ_{trace}	5% critical value
0.281	$r = 0$	$r \geq 1$	67.81	63.66
0.126	$r \leq 1$	$r \geq 2$	29.94	42.77
0.082	$r \leq 2$	$r \geq 3$	14.41	25.73
0.039	$r \leq 3$	$r = 4$	4.57	12.45

Notes: Results when trend is excluded and weak exogeneity of user cost, rents and subprime share is imposed (standard errors below point estimates):

$\text{ph} + \underset{0.486}{1.201}\text{UC} - \underset{0.167}{1.218}r - \underset{0.179}{1.419}\text{sp}$

$\alpha_{\text{ph}} = \underset{0.023}{-0.143}, \alpha_{\text{UC}} = 0, \alpha_r = 0, \alpha_{\text{sp}} = 0$

Log-likelihood: 2110.57

Likelihood ratio test for overidentifying restrictions:

$\chi^2(4) = 4.7267[0.3165]$

Estimation period: 1977:Q2–2010:Q4

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Table VIII. CVAR analysis for the inverted demand approach with subprime share in VAR, 1977:Q2–2010:Q4

Eigenvalue	H_0	H_A	λ_{trace}	5% critical value
0.284	$r = 0$	$r \geq 1$	89.20	73.13
0.221	$r \leq 1$	$r \geq 2$	50.87	50.08
0.136	$r \leq 2$	$r \geq 3$	22.08	30.91
0.045	$r \leq 3$	$r = 4$	5.34	15.33

Notes: Results when trend is excluded and weak exogeneity of user cost, rents and sp is imposed (standard errors below point estimates):
$\text{ph} + \underset{0.757}{0.559}\text{UC} - \underset{0.463}{1.345}y + \underset{0.668}{1.618}h - \underset{0.335}{2.480}\text{sp}$
$\alpha_{\text{ph}} = \underset{0.022}{-0.116}, \alpha_{\text{UC}} = 0, \alpha_y = 0, \alpha_{\text{sp}} = 0$

Log-likelihood: 1986.11
Likelihood ratio test for overidentifying restrictions: $\chi^2(4) = 17.401[0.0016]$
Estimation period: 1977:Q2–2010:Q4

To investigate the role played by the increased subprime lending a little further, I have included this ratio, sp, as an additional variable in the VECMs of Section 5.[25] I have summarized the results when I redo the cointegration analysis with the subprime measure included in the models in Tables VII and VIII.[26]

It can clearly be seen from the results in Table VII that by including this variable in the price-to-rent model I find evidence for one cointegrating vector over the full sample. In addition, I find that the trend can be excluded and weak exogeneity of all the variables in the VAR (including the new variable) is supported. Most striking is the fact that including this variable, which is positive and highly significant, changes the estimates of the other coefficients for the full sample analysis in such a way that they move very close to their pre-break values (compare to the pre-break results in Table II). Furthermore, the absolute values of the loading factors increase substantially, and now have a more reasonable numerical size. Table VIII tells a similar story (compare the results to the pre-break results in Table III), but there are some signs that the rank might be two and that the weak exogeneity assumption is not supported by the data.[27] To explore the stability of the other coefficients a little further, Figure 4 plots the recursive estimates for the post-break period. It can be observed that the coefficients are quite stable, which suggests that by including the subprime measure I am able to explain the econometric breakdown documented earlier.

Finally, including the subprime measure in a model ending in 2000:Q4 (just before the break), I do not find that this variable enters the cointegrating relationships.[28] This suggests that we can, without loss of generality, exclude this variable from the model in the pre-break period. It further suggests that the breakdown of the stable relationship between housing prices, the user cost and rents as well as the inverted demand equation was caused by the increased exposure to the more risky segment of the market.

A strict interpretation of the combined results from the previous and the current section is that there exists formal statistical evidence implying that the extension of subprime lending caused the breakdown (the bubble) and that this contributed to the instability in the banking sector and the wider financial crisis.

[25] Owing to the lack of data, I have set this series to zero prior to 1998:Q1. That said, since subprime lending is a relatively new phenomenon, this approximation should not be very important for my results.
[26] The sudden jump in this series in 2003 leads to some misspecification in the VARs that was not present earlier, but it is nevertheless interesting to see what happens when this variable is included in the VARs.
[27] The test for overidentifying restrictions fails once weak exogeneity of the subprime variable is imposed. However, if I do not impose that restriction, I get similar results.
[28] Further details and results are available upon request.

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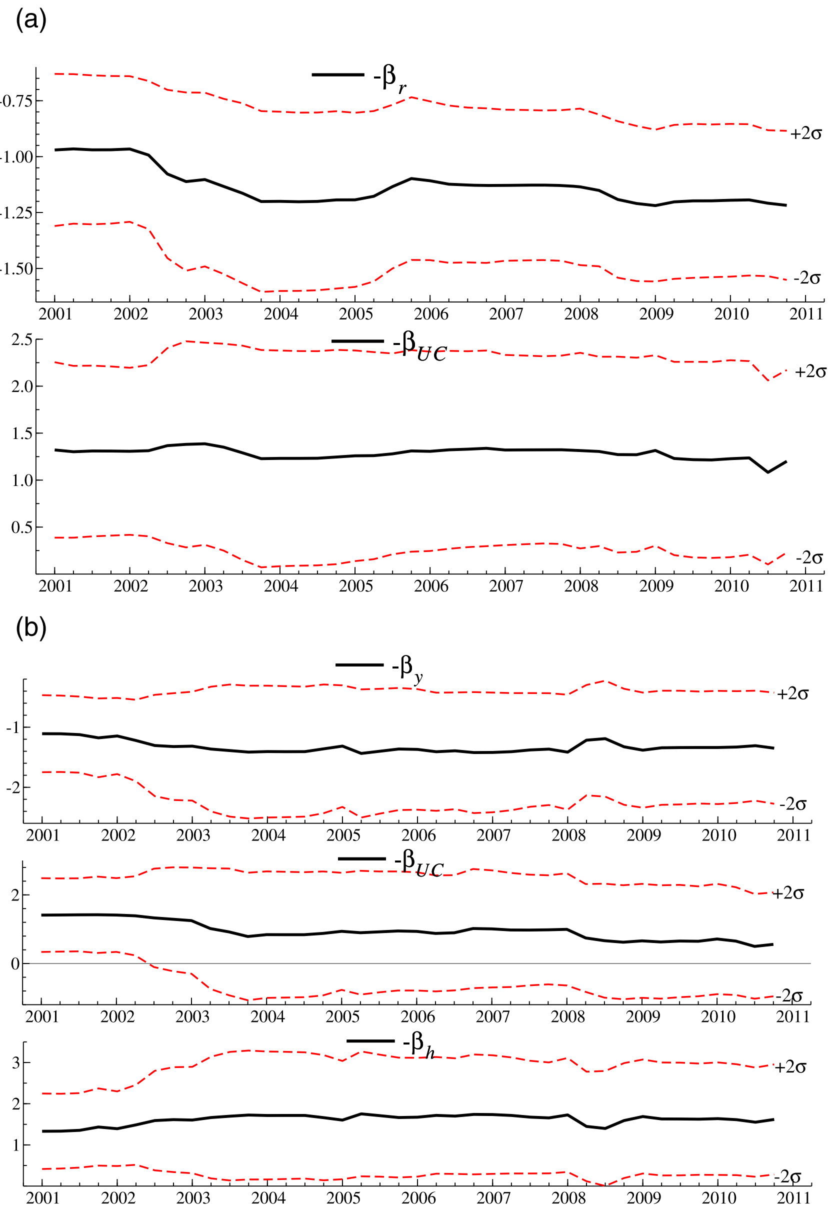

Figure 4. (a) Recursively estimated coefficients for rents and the user cost in the price-to-rent model, 2001:Q1–2010:Q4. (b) Recursively estimated coefficients for per capita disposable income, the user cost and housing stock from the inverted demand approach, 2001:Q1–2010:Q4

8. THE ROLE OF EXPECTATIONS AND TESTING FOR OTHER PERIODS OF STRUCTURAL BREAKS

8.1. The Role of Expectations

As discussed in Section 3, I have followed Abraham and Hendershott (1996), Gallin (2008) and Anundsen and Jansen (2013) and assumed that expectations of future house price increases are captured by lagged housing price appreciation in the dynamic part of the model, instead of including a

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Table IX. Selected lag lengths

Lag	Price-to-rent		Inverted demand	
Δph_{t-1}	*	0.254 (0.061)	0.249 (0.107)	0.438 (0.077)
Δph_{t-2}	*	*	*	*
Δph_{t-3}	0.306 (0.076)	0.412 (0.072)	0.302 (0.095)	0.246 (0.07)
Δph_{t-4}	*	0.294 (0.083)	*	0.504 (0.080)
Sample endpoint	2000:Q4	2010:Q4	2000:Q4	2010:Q4

Notes: This table reports the estimated coefficients on the lagged housing price inflation terms for both the price-to-rent and the inverted demand models. When the sample endpoint is set to 2010:Q4, the subprime variable is included in the models. Autometrics, with a significance level of 5%, is used to select the final models. Standard errors are reported below the point estimates, and an asterisk signifies that the variable was not found significant in the final model.

price expectations term directly in the user cost measure. This modeling choice was made because results were somewhat sensitive to the number of lags included in a moving average process used to represent housing price expectations in the user cost term. In the following, I will discuss three issues in relation to this. First, given that the lags pick up a price expectation effect, one would expect them to enter positively and significantly in the econometric models. Second, the importance of expectation effects may have changed over the course of the bubble. Third, I restricted the number of lags to be at most five, which clearly imposes some a priori structure on the expectation process as well. This was done to avoid overdimensioning the VAR, but it is possible—using Autometrics—to explore how the results from the conditional equilibrium correction models are affected when this assumption is relaxed.

To address the first two issues, Table IX reports the estimated coefficients for the lagged housing price appreciation terms both in the inverted demand model and the price-to-rent model for samples ending in 2000:Q4 (pre-break) and 2010:Q4 (post-break) when I consider the conditional equilibrium correction approach. For the models where the sample endpoint is set to 2010:Q4, I have included the subprime measure in the general unrestricted models (GUMs) to remain consistent with the modeling results presented in previous sections.

An inspection of Table IX reveals that lagged housing prices clearly enter all models both positively and significantly, giving credence to the interpretation that they pick up an expectational effect. Another interesting observation is that—in both models—the (sum of the) lagged housing price coefficients are higher for the models with sample endpoint in 2010:Q4. Also this would be in accordance with the bubble interpretation of the econometric regime shifts documented earlier, since—using the terminology of Abraham and Hendershott (1996)—the *bubble builder* was more important in the latter period.

To explore the third issue regarding the lagged dependent variable, namely the sensitivity of my results to the maximum number of lags allowed for in the GUM, I have redone the entire analysis of Section 5.3 allowing for up to eight lags in lagged housing price appreciation.[29] The results for the price-to-rent and the inverted demand approach are reported in Tables X and XI, respectively.

Comparing the results in those tables to the results reported in Tables IV and V, it is clear that the results are almost unaffected; the estimated long-run elasticities and the adjustment parameters are similar to the previous results, and the finding of a regime shift is still evident in both models. It is evident that the fit is slightly improved when more lags are included, which can be seen by comparing the estimated standard errors (σ) in Tables X and XI to those reported in Table IV and V. Further, we see that the sum of the coefficients on the retained housing price appreciation terms are positive, and that the sum is greater than in the models where only four lags of lagged growth rates were

[29] I am grateful to an anonymous referee for suggesting this extension to me.

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Table X. Recursive coefficients for price-to-rent model allowing for up to eight lags in housing price appreciation, 1977:Q2–T

Endpoint (T)	β_r	β_{UC}	α_{ph}	p-value	Autocorrelation	Non-normality	Heteroskedasticity	σ	Exp.
1995:Q4	0.997	−1.193	−0.244	0.001	0.5144	0.1505	0.9811	0.0062	0.817
1996:Q4	1.005	−1.180	−0.249	0.000	0.4256	0.2117	0.9797	0.0062	0.810
1997:Q4	0.961	−1.107	−0.263	0.001	0.8636	0.0994	0.9932	0.0059	0.966
1998:Q4	1.006	−1.178	−0.249	0.000	0.5423	0.3326	0.9430	0.0060	0.814
1999:Q4	0.954	−1.097	−0.271	0.000	0.7730	0.0583	0.9859	0.0057	0.968
2000:Q4	1.034	−1.106	−0.234	0.000	0.6115	0.1223	0.9587	0.0058	0.971
2001:Q4	1.141	−1.212	−0.191	0.003	0.9884	0.2835	0.9555	0.0060	0.928
2002:Q4	1.319	−1.317	−0.154	0.005	0.7856	0.6995	0.7472	0.0061	0.768
2003:Q4	1.634	−1.229	−0.117	0.038	0.6241	0.0035	0.0785	0.0067	0.660
2004:Q4	*	*	*	0.589	*	*	*	*	*
2005:Q4	*	*	*	0.848	*	*	*	*	*
2006:Q4	*	*	*	0.660	*	*	*	*	*
2007:Q4	*	*	*	0.129	*	*	*	*	*
2008:Q4	1.999	−0.990	−0.041	0.061	0.5809	0.3735	0.3212	0.0073	0.795
2009:Q4	1.894	−1.564	−0.059	0.013	0.3677	0.1878	0.7253	0.0071	1.035
2010:Q4	2.095	−0.922	−0.061	0.002	0.3098	0.1865	0.3741	0.0075	0.901

Note: This table reports the estimated cointegrating vector along with the loading factor and corresponding p-value when the price-to-rent model is estimated using a single-equation approach. The next four columns report the p-values from tests for autocorrelation, non-normality and heteroskedasticity, as well as the equation standard error. The final column reports the sum of the coefficients on lagged housing price appreciation in the final model. An asterisk replaces the results for the given endpoint when the p-value from the test for cointegration exceeds 0.1.

Table XI. Recursive coefficients for inverted demand equation allowing for up to eight lags in housing price appreciation, 1977:Q2–T

Endpoint (T)	β_y	β_h	β_{UC}	α_{ph}	p-value	Autocorrelation	Non-normality	Heteroskedasticity	σ	Exp.
1995:Q4	1.110	−1.230	−1.224	−0.227	0.0017	0.3647	0.5652	0.8706	0.0069	1.218
1996:Q4	1.177	−1.334	−1.221	−0.229	0.0011	0.2027	0.7225	0.9630	0.0069	1.172
1997:Q4	1.076	−1.146	−1.008	−0.252	0.0000	0.5058	0.6682	0.9525	0.0066	1.182
1998:Q4	1.060	−1.104	−0.857	−0.234	0.0001	0.5345	0.9523	0.9036	0.0065	1.273
1999:Q4	1.060	−1.129	−1.003	−0.257	0.0000	0.6084	0.6051	0.8981	0.0063	1.186
2000:Q4	1.092	−1.140	−0.875	−0.222	0.0000	0.5941	0.8366	0.8910	0.0063	1.253
2001:Q4	1.293	−1.484	−1.367	−0.154	0.0206	0.9871	0.2943	0.3799	0.0071	1.163
2002:Q4	1.532	−1.797	−1.335	−0.107	0.0758	0.9559	0.4329	0.6586	0.0072	1.131
2003:Q4	*	*	*	*	0.5516	*	*	*	*	*
2004:Q4	*	*	*	*	0.8477	*	*	*	*	*
2005:Q4	*	*	*	*	0.9863	*	*	*	*	*
2006:Q4	*	*	*	*	0.9866	*	*	*	*	*
2007:Q4	*	*	*	*	0.7107	*	*	*	*	*
2008:Q4	*	*	*	*	0.8451	*	*	*	*	*
2009:Q4	*	*	*	*	0.9223	*	*	*	*	*
2010:Q4	*	*	*	*	0.7578	*	*	*	*	*

Note: This table reports the estimated cointegrating vector along with the loading factor and corresponding p-value when the inverted demand model is estimated using a single-equation approach. The next four columns report the p-values from tests for autocorrelation, non-normality and heteroskedasticity, as well as the equation standard error. The final column reports the sum of the coefficients on lagged housing price appreciation in the final model. An asterisk replaces the results for the given endpoint when the p-value from the test for cointegration exceeds 0.1.

included. This may suggest that up to 2 years of past appreciation is relevant when households form their expectations about future capital gains.

8.2. Testing for Signs of Structural Breaks in Other Periods

As discussed in Section 3, an implicit assumption made for the pre-break period is that the latent mortgage credit constraint variable was fairly constant. As has become evident in the previous sections,

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this cannot be said to be the case for the period that followed the break, where the subprime explosion became a major driver of US housing price dynamics. In order to investigate the plausibility of my implicit assumption that the credit constraint variable was relatively constant prior to the subprime explosion in a bit more detail, I shall utilize the impulse indicator saturation (IIS) algorithm, which is an integrated part of Autometrics.[30]

The IIS algorithm includes an impulse dummy for every single observation in the information set and the model is estimated in blocks to determine which indicators are significant (see Hendry *et al.*, 2008, Johansen and Nielsen, 2009). On average, only αT indicators are retained by chance, where α denotes the chosen significance level and T is the number of observations. This is indeed a low cost to pay for robustifying a model to intermittent structural breaks. Castle *et al.* (2012) show that the IIS algorithm is successful in detecting multiple breaks in the data (even for the case with 20 breaks and only 100 observations).

To investigate whether there are signs of any other major breaks, loosely interpreted as shifts in the latent credit constraint variable, I estimated both the price-to-rent and the inverted demand-based model on the full sample, 1975:Q1–2010:Q4, employing the IIS algorithm with a significance level of 1%. In both cases, I included the subprime variable as part of the information set, and a constant as well as the lagged levels and the MT and CGT dummies were restricted to enter the final models.

Four dummies were picked up in the inverted demand model and 10 in the price-to-rent model. Three of the dummies that were picked up in the inverted demand model, and also four of the dummies in the price-to-rent model, are related to the financial crisis period. This suggests that the assumption of relatively constant credit constraints prior to the subprime explosion can be maintained. It is worth mentioning that the long-run coefficients from this exercise are very close to those I found using the system-based approach to cointegration (without using the IIS routine), and in both cases I find that there are no signs of residual misspecification.

9. CONCLUSION

Based on both system-based and single-equation tests for the absence of cointegration, this paper has documented how two stable equilibrium relationships linking real US housing prices to real rents and the real direct user cost and another one linking real housing prices to the real direct user cost, real per capita disposable income and the housing stock breaks down in the early 2000s. Although there is some evidence of restored equilibrium correction at the end of the sample, the adjustment coefficients and the long-run elasticities are diametrically different in the post-break period.

The breakdown of a cointegrating relationship can often be interpreted as a result of a far-reaching or fundamental change in an interwoven system like the US housing and credit market. It can also be interpreted as a passage from a regime where fundamentals drive housing prices, to a regime dominated by bubble dynamics. In that perspective, I developed two regime shift indicators, which can be interpreted as 'bubble indicators'. According to these indicators, the US housing bubble started in the early 2000s, was pricked in 2007 and by the end of 2010 housing prices were more closely in line with the pre-break fundamentals.

Tests for Granger non-causality showed that the indicators have predictive power for a set of financial (in)stability measures and coincident indicators. This highlights that such indicators possibly can be part of a toolkit when analyzing the stability of the financial system.

Finally, it was shown that including a measure for the number of subprime mortgages as a share of all mortgages, the pre-break relationships were re-established. Furthermore, the long-run coefficients were in this case found to be highly stable when estimated recursively. These findings suggest that

[30] Again, I am grateful to an anonymous referee for suggesting this extension to me.

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it was the expansion of subprime borrowing that caused the econometric breakdown and therefore contributed to the major imbalances in the US housing market in the previous decade.

Given the findings in this paper, a fruitful approach for future research would be to explore the role of subprime lending in explaining regional differences in US housing price dynamics over the recent boom–bust cycle. Another interesting area of research is to explore whether the methodology suggested in this paper applies at a more disaggregate level as well.

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