Detecting imbalances in house prices: What goes up must come down?

Running head: House price imbalances

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Abstract

I suggest a toolkit of four bubble-detection methods that can be used to monitor house price developments. These methods are applied to US, Finnish and Norwegian data. For the US, all measures unanimously suggest a bubble in the early to mid 2000s, whereas current US house prices are found to be aligned with economic fundamentals. One of the measures indicates imbalances in Finland, while there are no signs of a bubble in Norway. I find that large parts of the US house price bubble can be explained by the sharp increase in capital inflows and the extension of loans to the subprime mortgage market.

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JEL classification: *C22; C32; C51; C52; C53; G01; R21*

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I Introduction

Is it true that what goes up must come down? Starting in the late 1990s, there was an unprecedented international house price boom accompanying the favorable economic situation in most industrialized countries. The boom was in many cases succeeded by a significant bust, with real house prices falling by more than 30 percent in several countries. The consequences for the real economy following the bust in house prices have been severe, and it was one of the factors contributing to the deepest downturn in the world economy since the Great Depression (see e.g., Mian et al. (2013) and Mian and Sufi (2014)). The collapse culminated with the meltdown of the US housing market and financial system in the autumn of 2008 – the epicenter for the ensuing global financial crisis that is still putting a strain on the global economic recovery. Against this background, I ask whether econometric methods can be used to detect pending imbalances in the housing market.

Using aggregate house price data for the US, Finland and Norway, I consider four alternative econometric approaches to identifying imbalances in the housing market. While one *would want* measures of housing market imbalances to detect a bubble that is building up, one *would not want* the methods to signal a bubble whenever house prices are increasing. This makes the analysis of these three countries particularly relevant, since house prices increased rapidly in all three countries from the beginning of the 1990s and until the global financial crisis. While there was a major and sustained drop in US house prices in the late 2000s, Norwegian and Finnish house prices quickly rebounded after the drop, and reached new historical heights by 2013. This prompts the question of whether house prices in these countries can be explained by underlying economic fundamentals. If not, it is imperative from a policy perspective to detect and quantify such imbalances in real time so that necessary actions can be taken to prevent further overvaluations – especially given the strong role that the housing market has in affecting the real economy, both through consumption wealth effects (Aron et al., 2012; Mian et al., 2013) and through its interactions with the credit market (Hofmann, 2003; Fitzpatrick and Mc-

Quinn, 2007; Gimeno and Martinez-Carrascal, 2010; Anundsen and Jansen, 2013). In addition, Leamer (2007) and Leamer (2015) have shown that large drops in housing investments are a strong indicator of future recessions in the US economy – a result that has gained international support in a recent study by Aastveit et al. (2017).

The first approach I take to evaluate whether house prices were overvalued in these countries in the 2000s is to calculate a fundamental house price using information that would have been available in 1999q4. Then, I investigate how actual house prices developed relative to the model-implied fundamental prices in the period thereafter.

My second measure of housing market imbalances is based on a dynamic forecasting exercise. Since a forecast is a conditional expectation, one would not expect an econometric model including the relevant house price fundamentals to produce forecasts that systematically underpredict house prices if they are close to their equilibrium value. Hence, large and systematic underpredictions of house prices can be interpreted as an overheating of house prices.

As a third approach, I apply the bubble-indicator methodology suggested by Anundsen (2015). This methodology relies on the bubble definition provided by Stiglitz (1990, p.13), which states that 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". This bubble definition is combined with the modeling assumption that fundamental factors are non-stationary economic time series. Given this assumption, house prices are determined by fundamentals *if and only if* there exists a cointegrating relationship between house prices and these non-stationary economic variables. This leads to several possible scenarios. First, if cointegration can be established over the full sample period as well as for different sub-samples, the bubble hypothesis is clearly rejected. Conversely, if no evidence of cointegration can be found, we cannot reject a bubble, but it may also indicate that our information set does not include the relevant fundamentals. If a cointegrating relationship can be established early in the sample but is lost subsequently, we may suspect a structural break. This finding is therefore consistent with the transition from a stable market with equilibrium correction behavior (no bubble) to a market where there are no forces in place to correct disequilibrium constellations (a bubble).

The final measure I consider utilizes recent econometric tools developed by Phillips et al. (2011, 2015a,b) to test for a transition to explosive house price behavior.

All four approaches provide a possible way of identifying imbalances in the housing market. The approach based on deviations of actual prices from fundamental prices and the forecasting exercise are particularly relevant for making quantitative statements about the magnitude of house price misalignments. A drawback with these approaches is that it is far from trivial to translate evidence of overvaluation into bubble behavior. This is because the overvaluation may be temporary and expected to be adjusted relatively quickly. In addition, both these approaches require that the researcher or policymaker takes a position on how large misalignments must be for there to be a bubble. Thus, using these measures in isolation to call bubbles in real time may not be feasible. That said, the bubble indicator and the test for a transition to explosive house prices are particularly useful for real-time monitoring and dating of the onset of a housing bubble, since they directly test for the transition to a bubble regime. Thus, as part of a broader toolkit for monitoring the housing market, these measures may help signal a bubble and – conditional on bubble detection – the deviations of actual house prices from fundamental prices and the forecasting exercise can be applied to quantify the degree of departure from equilibrium.

My results show that all four measures signal a bubble in the US housing market starting in the early 2000s. There are, however, no signs of overheating during the more recent house price boom starting in 2012. Similarly, I do not find evidence of a bubble in the Norwegian housing market for a sample ending in 2016, while only one of the measures suggests that Finnish house prices are overvalued for a sample that ends in 2011. On balance, my results therefore suggest that the strong growth in Norwegian and Finnish house prices may be attributed to the developments in income, interest rates and housing supply. These results do not suggest that house prices may not fall in these countries, since less favorable developments in underlying economic fundamentals – e.g., an increase in the mortgage interest rate or a drop in household income – is consistent with a fall in house prices. In addition, if fundamentals are developing along an unstable trajectory, the measures may fail to detect a bubble.

As a final exercise, I investigate the empirical relevance of two hypotheses of the drivers of US house prices in the 2000s that have received attention in the literature. The first hypothesis states that the price increase in the US housing market in the 2000s was related to an increased inflow of capital from emerging market economies (see Bernanke and Kuttner (2005), Caballero and Krishnamurthy (2009) and Bernanke et al. (2011)). The global savings glut hypothesis states that the increased demand for US assets pushed down long-term interest rates and thereby increased housing demand. Another strand of the literature has suggested that the house price increases in the 2000s were due to an increased access to credit following the subprime explosion, see e.g., Mian and Sufi (2009), Duca et al. (2011a,b), Pavlov and Wachter (2011) and Anundsen (2015). My results suggest that both hypotheses have relevance in explaining the overheating in the US housing market, and my estimates indicate that they together explain nearly 80 percent of the gap between actual and fundamental prices in the 2000s.

There has been a long-standing discussion in the academic literature on the extent to which the US house price increase in the 2000s may be explained by economic fundamentals (see e.g. Meen (2002); Gallin (2006); Clark and Coggin (2011); McCarthy and Peach (2004); Gallin (2008); Mikhed and Zemcik (2009a,b); Zhou (2010)), and Gerardi et al. (2010) point out that few economists predicted the crash in the housing market before it actually occurred. While Himmelberg et al. (2005) found some evidence of overheating in certain US cities in 2004, their main conclusion was that there were few signs of imbalances in most US housing markets at that time. In another study, Foote et al. (2012) have stated that the price run-up in the US in the 2000s may not even in retrospect be characterized as a bubble. Anundsen (2015) contested this conclusion by constructing an econometrically based bubble indicator that clearly signals a bubble in the US housing market in the early 2000s.

This paper contributes to the literature on the US housing bubble in several ways. First, by considering a toolkit of measures of overheating, this paper provides an operative and easily implementable strategy to detect bubbles and to quantify the degree of overheating in the housing market. Both the test for explosive roots and the bubble-indicator methodology have previously been used to show that there was a bubble in the US housing market in the 2000s, see e.g., Pavlidis et al. (2015) and Anundsen (2015). In this paper, I go one step further and suggest two additional indicators that may be used to monitor the housing market. The advantage of considering these additional indicators is that one can go further than a binary categorization of the bubble/no-bubble question in that they may be employed to establish degrees of overheating. As an example of the relevance of this approach, my results corroborate previous studies in that I find evidence of a bubble in the US housing market in the 2000s. Furthermore, my estimates suggest that the gap between actual and fundamental prices reached 50 percent in the midst of the US housing bubble. As a second contribution to the literature on US house price developments, I apply my suggested toolkit to the more recent boom period. The current boom looks remarkably similar to the previous boom in terms of house price growth. That said, my econometric results suggest that current price developments can be attributed to underlying economic fundamentals, suggesting that the current recovery is more healthy – a result that is consistent with the discussion in Glick et al. (2015).

The house price developments in Norway and Finland in the 2000s are reminiscent of those in the US, and while often debated in policy institutions and in the financial press, there is a gap in the academic literature on testing for housing bubbles in other countries. For the case of Finland, the European Commission (Commission (2013)) has estimated a fundamental house price gap based on the relationship between house prices and imputed rents. Their results suggest that Finnish house prices were consistently overvalued over the period 2003-2011 and that the overvaluation reached 15 percent in

2006-2008 and 2010-2011. For the case of Norway, one of the more vocal statements has come from Moody's (Moodys, 2017), claiming that Norway has the largest overvaluation of all countries covered by their analysis.[1] Their approach is to compare the evolution of house prices relative to some rental price series and a measure of the user cost of housing. Their results suggest that house prices in Norway have been consistently overvalued since 2010. Also the IMF has warned about developments in house prices in both Norway and Finland. In IMF (2017), a panel analysis on 20 countries over the period 1990q1-2016q4 is presented. Both Norway, Finland and the US are in the sample. They conclude that house prices in Norway were overvalued by 16 percent in 2016, which – as they emphasize – is among the highest among the countries covered by their analysis. Their results also suggest that house prices in Norway have been consistently overvalued since 2007. Prices in Finland are found to be undervalued since 2008 and – contrary to the analysis of the European Commission – that prices were in line with fundamentals before that. They also argue US prices have been systematically undervalued by about 20 percent since 2008.

Another contribution of this paper is to apply the same toolkit that detected the US housing bubble to the Norwegian and Finnish housing market. Quite contrary to the common belief and the results of the analyses carried out by important policy institutions, I find that there are no signs of housing bubbles in these countries. These results show that it is important to discipline the bubble discussion in a formal econometric framework grounded in theory so that well-informed policy decisions can be made. Both the study by the European Commission and Moody's consider the relationship between house prices and rents, which is theoretically appealing – a similar approach has been used in the US literature (see e.g., Duca et al. (2011a) and Anundsen (2015)). However, this approach is less suitable for European countries, since the rental market is relatively small and illiquid, and since the rental market is heavily regulated in many European countries (Muellbauer, 2012). In this paper, I instead link the evolution of house prices to income,

[1] The countries considered are France, Ireland, Portugal, Spain, Canada, New Zealand, Norway and Sweden.

the housing stock and the user cost of housing, as implied by theory.

The approach taken in the IMF study is related to my first approach of calculating a fundamental house price. However, a weakness of that analysis is that it is carried out in a panel. Thus, there is an implicit assumption that housing demand elasticities are the same across countries. When estimating separate time series models for the US, Finland and Norway, I find considerable heterogeneity in demand elasticities.

Furthermore, as discussed above, finding evidence of misalignments between the actual and an estimated fundamental price does not necessarily imply the existence of a bubble. An advantage with the bubble-detection toolkit suggested in this paper is that it is directly implementable in an operational setting. In addition, looking at several indicators provides a broader information set for central banks and financial supervisory authorities when monitoring the housing market. Two of the approaches used in this paper are particularly useful in detecting the onset and the end of a housing bubble. Whereas this information is important, it is also imperative to know by how much the market is overvalued. Therefore, the other two approaches serve as complementary tools in that they can be directly used to quantify by how much house prices are overvalued, conditional on detecting a bubble in the first place. Thus, one can distinguish between temporary overvaluations that are expected to be corrected by the market and overvaluations that are caused by bubble dynamics. Furthermore, conditional on detecting periods of bubbles, one can use these indicators to investigate the drivers of bubbles, which may be particularly useful in order to prevent such episodes from recurring in the future.

A final contribution of this paper is to shed light on the factors that led to the US housing bubble in the 2000s. This analysis can also help us understand why the Norwegian and Finnish experiences were so different. In contrast to the US, neither Norway nor Finland saw the emergence of subprime lending. At the same time, while the US accumulated a large current account deficit starting in the 1990s, both Norway and Finland had considerable current account surpluses over the same period. My results show that both the current account deficit and the subprime explosion contributed to the overvaluation of the US housing market in the 2000s. Moreover, I quantify that these two factors alone can explain nearly 80 percent of the deviation between actual and fundamental prices over that period.

The rest of the paper is organized as follows. The next section provides a theoretical background to the construction of the four alternative measures of housing market imbalances considered in this paper, and suggests a way of operationalizing each of them econometrically. The data used in the analysis are presented in Section III, while results based on each of the indicators are presented in Section IV. In Section V, I explore the relevance of the global savings glut hypothesis and the relaxation of lending standards in explaining the US housing bubble. The final section concludes the paper.

II Four alternative indicators of house price misalignments

To discuss whether house prices in any given housing market are best characterized as exhibiting bubble behavior, at least two requirements must be satisfied: first, we must have a conceptual understanding of what we define as a house price bubble. Second, given our conceptual understanding of a housing bubble, we need to have a formal (statistical) framework in which the existence of a bubble may be detected conditional on our definition of imbalances.

An inverted demand equation approach

Three of the measures considered in this paper rely on the bubble definition provided by Stiglitz (1990). Operationalizing this definition requires that we also have an understanding of what are the fundamental determinants of house prices. A commonly used theory for the drivers of house prices is the life-cycle model of housing (see e.g. Meen (1990, 2001, 2002)). This theoretical framework takes as a starting point a standard representative agent model, with an agent maximizing her lifetime utility with respect to consumption of housing goods and "other" goods.

The solution to the maximization problem yields the following equilibrium condition:

$$\frac{U_H}{U_C} = PH\left[(1-\theta)i - \pi + \delta - \frac{\dot{PH}}{PH}\right] \tag{1}$$

where U_H measures the marginal utility of housing goods (H), U_C is the marginal utility of other goods (C), PH denotes house prices, while θ, i and π denote the tax rate at which interest expenses are deductible, the mortgage interest rate and the general CPI inflation rate, respectively. Finally, δ is the housing depreciation rate. This condition states that the marginal rate of substitution between housing and the composite consumption good is equal to what it costs to own one more unit of a property, measured in terms of foregone consumption of other goods. Since the housing market also contains a rental sector, market efficiency requires the following condition to be satisfied in equilibrium:

$$Q = PH\left[(1-\theta)i - \pi + \delta - \frac{\dot{PH}}{PH}\right]$$

where Q is the real imputed rent on housing services. Hence, the price-to-rent ratio is proportional to the inverse of the user cost:

$$\frac{PH}{Q} = \frac{1}{UC} \tag{2}$$

where the real user cost, UC, is defined as: $UC = (1-\theta)i - \pi + \delta - \frac{\dot{PH}}{PH}$. In the econometric analysis, I shall assume a constant depreciation rate and I will include the real after-tax interest rate as my operational measure of the (direct) user cost in the models, while I let expectations about future price changes be captured by the lags included in the econometric models, which is similar to the approach in Abraham and Hendershott (1996), Gallin (2008) and Anundsen (2015).

The real imputed rent is unobservable, but two approximations are common: to proxy the imputed rent by an observable rent R, or to assume that it is proportional to income and the stock of housing. Relying on the first approximation, the expression in (2) would read:

$$\frac{PH}{R} = \frac{1}{UC} \tag{2a}$$

while if we instead assume that the imputed rent is determined by the following expression:

$$R = Y^{\beta_y} H^{\beta_h}, \ \beta_y > 0 \ and \ \beta_h < 0$$

where Y measures disposable income and H is the housing stock, (2) would read:

$$\frac{PH}{Y^{\beta_y} H^{\beta_h}} = \frac{1}{(1-\theta)i - \pi + \delta - \frac{\dot{PH}}{PH}} \tag{2b}$$

The expressions represented by (2a) and (2b) are often used as a starting point for building econometric models of house prices, and they will also be central to the econometric modeling used to construct the first three measures of housing market (in)stability considered in this paper.

A natural starting point for an econometric analysis of house price determination is therefore to consider these expressions on a semi-logarithmic form:[2]

$$ph = \beta_r r + \tilde{\beta}_{UC} UC \tag{3a}$$

$$ph = \beta_y y + \beta_h h + \beta_{UC} UC \tag{3b}$$

where we would expect that $\beta_r, \beta_y > 0$ and $\beta_h, \tilde{\beta}_{UC}, \beta_{UC} < 0$. Either or both of these equations form the basis for a series of studies that investigate house price determination, see e.g. Buckley and Ermisch (1983); Hendry (1984); Meen (1990); Holly and Jones

[2]A semi-logarithmic representation is commonly used, since the user cost may take negative values.

(1997); Meen and Andrew (1998); Meen (2001); Duca et al. (2011a,b) and Anundsen (2015) to mention a few of the many empirical studies that are grounded in the life-cycle model of housing.

While the first operationalization (price-to-rent), (3a), has been used in the US literature (see e.g., Duca et al. (2011a)), it has been less commonly applied to house price modeling in Europe, since the rental market is relatively small and illiquid in countries such as e.g., the UK and Norway, and since the rental market is heavily regulated in many European countries (Muellbauer, 2012).

In this paper, I will confine my analysis to considering the inverted demand approach, (3b). This relationship defines equilibrium house prices as a function of the user cost of housing, households' disposable income and the housing stock. To get an operational measure of the degree of overheating in the housing market, I shall consider the following three approaches:

1. **Deviations from an estimated fundamental value:** Based on (3b), I estimate the implied fundamental house price path during the boom period using data up to the beginning of the recent house price boom. Large and systematic deviations of actual house prices from the estimated fundamental value in the boom period will be taken as an indication of unsustainable developments in house prices.

2. **Systematic forecast failures:** A related, but distinct approach is to embed (3b) in a dynamic econometric model and construct forecasts from that model. If house prices are not overvalued, we would not expect such a model to produce large and systematic forecast errors. If, on the other hand, the model systematically produces (significant) underpredictions, we would conclude that the evolution of house prices is not supported by underlying economic fundamentals, which may suggest that house prices are developing along an unsustainable trajectory.

3. **Loss of equilibrium correction:** A third approach is to develop a conditional equilibrium correction model, which by definition has an equilibriating force. This

model may be used to assess the extent to which disequilibrium constellations may be thought of as a short-run phenomenon that we would not want to react to (since prices will eventually revert to the value implied fundamentals anyway), or if there is a tendency for no adjustment towards a long-run equilibrium value.

Approach # 1: Deviations from an estimated fundamental value

All the variables in (3b) exhibit stochastic non-stationarities, i.e., they are integrated time series processes. As a direct consequence of this, it follows that a requirement for (3b) to constitute a stationary equilibrium is that the linear combination $ph - \beta_{UC}UC - \beta_y y - \beta_h h$ is stationary, so that prices revert to a level consistent with the time-varying fundamentals.

Thus, we would expect house prices to cointegrate with the user cost of housing, disposable income and the housing stock.

To establish whether there exists empirical evidence supporting cointegration between house prices and the economic fundamentals suggested by the theory model, I use the system-based test for cointegration in Johansen (1988). In the econometric analysis, I condition on the housing stock in the cointegration space. Consider the following partition $\boldsymbol{y}_t = (\boldsymbol{x}_t', h_t)'$, where $\boldsymbol{x}_t$ is a vector of endogenous variables, while h_t is the housing stock, which is (treated as) weakly exogenous. With this notation, the VECMX representation of an underlying VARX of p^{th} order can be written in the following way (see Johansen (1994, 1995) and Harbo et al. (1998) for details):

$$\Delta \boldsymbol{x}_t = \boldsymbol{\Pi} \boldsymbol{y}_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_{\boldsymbol{x},i} \Delta \boldsymbol{x}_{t-i} + \sum_{i=0}^{p-1} \Gamma_{h,i} \Delta h_t + \boldsymbol{\Phi} \boldsymbol{D}_t + \boldsymbol{\varepsilon}_t \tag{3}$$

where the vector $\boldsymbol{x}_t$ contains real house prices, ph, real disposable income, y, and the real direct user cost, UC. The vector $\boldsymbol{D}$ collects a constant, three centered seasonal dummies and a deterministic trend. In the short-run, I follow custom and assume that the stock of housing is given ($\Gamma_{h,i} = 0 \; \forall \; i$), implying that prices clear the market. This assumption is motivated by the fact that the stock of houses only changes slowly, and it implies that short-run fluctuations in house prices are driven by demand shocks. The disturbances are assumed to follow a multivariate normal distribution, $\boldsymbol{\varepsilon}_t \sim MVN\left(\mathbf{0}, \boldsymbol{\Sigma}\right)$.

Testing for cointegration amounts to testing the rank of the matrix $\boldsymbol{\Pi}$, which corresponds to the number of independent linear combinations of the variables in $\boldsymbol{y}_t$ that are stationary. From the theory discussion, we would expect that $Rank\left(\boldsymbol{\Pi}\right) = 1$, i.e., that there exists one cointegrating relationship among these variables. If the rank of $\boldsymbol{\Pi}$ is found to be one, we also know that $\boldsymbol{\Pi}$ has a reduced rank representation, $\boldsymbol{\Pi} = \boldsymbol{\alpha\beta}'$, where $\boldsymbol{\alpha}$ is a 3×1 matrix, whereas $\boldsymbol{\beta}$ is a 5×1 matrix (since a deterministic trend is also restricted to enter the space spanned by $\boldsymbol{\alpha}$). Thus, (3) may be expressed as:

$$\begin{pmatrix} \Delta ph_t \\ \Delta y_t \\ \Delta UC_t \end{pmatrix} = \begin{pmatrix} \alpha_{ph} \\ \alpha_y \\ \alpha_{UC} \end{pmatrix} \left(ph - \beta_y y - \beta_h h - \beta_{UC} UC - \beta_{trend} trend\right)_{t-1} + \sum_{i=1}^{p-1} \boldsymbol{\Gamma}_{\boldsymbol{x},i} \Delta \boldsymbol{x}_{t-i} + \tilde{\Phi}\tilde{D}_t + \boldsymbol{\varepsilon}_t \tag{4}$$

where α_{ph}, α_y and α_{UC} are the adjustment parameters, i.e., measuring how the three endogenous variables in the VAR respond to last period equilibrium deviations, as measured by $\boldsymbol{\beta}'\boldsymbol{y}_{t-1} = \left(ph - \beta_y y - \beta_h h - \beta_{UC} UC - \beta_{trend} trend\right)_{t-1}$. The vector $\tilde{D}_t$ only contains the constant and three centered seasonal dummies since the trend is restricted to enter the cointegration space. In the representation above, I have normalized with respect to house prices in the cointegrating vector.

Based on this approach, we can construct a model-implied equilibrium (or fundamental) house price path, ph_t^*. The implied equilibrium path constructed based on this approach can then be compared to actual house price developments, ph_t. Large and systematic deviations would be an indication of overheating in the housing market.

A potential limitation of this approach is that data from the period being scrutinized are used in developing the measures used in the assessment. For that reason, I construct my measure based on a model that is estimated using only data from the pre-boom sample, which I will take to be the period up to 1999q4, while I use the period 2000q1–2014q4 as an “out of sample” evaluation period. It should of course be mentioned that this is not a true real-time modeling exercise, as I do not account for potential data revisions that may have been undertaken in subsequent periods. Unless there are systematic (and non-stationary) measurement errors, we would not expect the main conclusions to be materially affected, since that should not affect the cointegrating properties of the data series included in this analysis, i.e., the finding of cointegration should be invariant to adding a stationary measurement error.

Thus, if cointegration can be established on the pre-boom sample, we can construct a measure of disequilibrium behavior for the succeeding period (the suspected bubble period), conditional on the pre-boom equilibrium relationship holding for that period as well. Let $\hat{\beta}^{1999q4}$ denote the estimated cointegrating coefficients from the pre-boom sample. The implied equilibrium path of house prices, ph_t^*, for the period thereafter will then be given by:

$$ph_t^* = ph_{t-1}^* + \hat{\beta}_y^{1999q4}\Delta y_t + \hat{\beta}_h^{1999q4}\Delta h_t + \hat{\beta}_{UC}^{1999q4}\Delta UC_t \ \ ;t > 1999q4 \tag{5}$$

where I shall assume that prices are initially in equilibrium, i.e. $ph_{2000q1}^* = ph_{2000q1}$. Thus, ph_t^* is the implied equilibrium path of house prices over the boom period, conditional on no structural breaks in the cointegrating coefficients. It is therefore important to test the stability of these coefficients on the pre-boom sample to ensure that there are no signs of major instabilities in the coefficients used to construct the equilibrium path in (5).

It should be noted that this approach requires that we are able to establish cointegration on the pre-boom sample, since otherwise the model would never imply that house prices are determined by the fundamentals considered here. It may also be best suited for an *ex post* evaluation of the degree of disequilibrium in the housing market, since it may be hard to determine what large and persistent deviations are in real time, since departures from the price implied by fundamentals may just represent smaller deviations around the underlying equilibrium (steady state) house price value to which prices will eventually converge. That said, this approach may be used to judge by how much actual prices deviate from the equilibrium price, which is particularly useful from a policy perspective.

Approach # 2: Systematic forecast failures

The second approach I shall consider uses forecasts from a dynamic econometric model for house price growth. This model may also embed the cointegrating relationship – if any – estimated using the method described in the previous section. Cointegration is, however, not a requirement for this second approach to work.

If we find evidence of cointegration, then we also know that there exists an equilibrium correction model representation (Engle and Granger, 1987) of the VAR of the form given by (4). If there is no evidence of cointegration ($Rank(\mathbf{\Pi}) = 0$), one may consider a forecasting model in first differences. Starting with a model in first differences is an unnecessary simplification, since this is the same as imposing the *a priori* (and testable) restriction that $Rank(\mathbf{\Pi}) = 0$. Thus, in the case of a non-zero rank, I let my forecasting model be of the equilibrium correction form. If there is no evidence of cointegration, my forecasting model will be a VAR in first differences (which is the same as (4) with the restrictions $\alpha_{ph} = \alpha_y = \alpha_{UC} = 0$ imposed). The forecasting model is estimated on the pre-boom sample and is then used to construct forecasts for the suspected bubble period.

Irrespective of the econometric model, we can then construct forecasts of house price growth to evaluate the temperature of the housing market. In particular, if the model – which includes the relevant theoretical fundamentals as suggested by the theoretical model – produces forecasts that systematically underpredict actual house price growth, I will take this as suggesting that house prices are overvalued.

Approach # 3: Loss of equilibrium correction

The final measure of dis-equilibrium behavior in the housing market that is built on the life-cycle model of housing takes the following single-equation equilibrium correction model as a starting point:[3]

$$\Delta ph_t = \tilde{\mu} + \tilde{\alpha}_{ph}\left(ph - \tilde{\gamma}_y y - \tilde{\gamma}_h h - \tilde{\gamma}_{UC} UC\right)_{t-1} + \sum_{i=1}^{p} \tilde{\rho}_{ph,i} \Delta ph_{t-i} + \sum_{i=0}^{p} \tilde{\rho}_{y,i} \Delta y_{t-i} + \sum_{i=0}^{p} \tilde{\rho}_{UC,i} \Delta UC_{t-i} + \tilde{\varepsilon}_t \quad (6)$$

where $\tilde{\varepsilon}_t \sim IIN(0, \sigma^2)$.

If house prices are determined by fundamentals, we would expect $\tilde{\alpha}_{ph}$ to be negative, i.e., that the variables are cointegrated (Engle and Granger, 1987).[4] Three cases are of particular interest. First, if cointegration can be established independent of the sample end point, disequilibrium constellations in the housing market may occur, but where there also exists a force ($\tilde{\alpha}_{ph}$) ensuring that prices are correcting towards their equilibrium value. This is in line with Abraham and Hendershott (1996), who refer to the equilibrium correction coefficient, $\tilde{\alpha}_{ph}$, as the bubble burster, i.e., the mechanism that ensures that prices will always have a tendency to return to their fundamental equilibrium value, and thus prevent systematic deviations of house prices from the value implied by economic fundamentals. The same authors refer to the coefficients on lagged house price appreciation terms – the $\tilde{\rho}_{ph,i}$ coefficients – as the bubble builder terms, since they capture an extrapolative expectations channel. If there is a one-time increase in income, prices may continue to rise for several periods as long as the sum of the $\tilde{\rho}_{ph,i}$ is greater than zero, but this is counteracted by the bubble burster term (as long as it is negative). Thus, even if there is an extrapolative expectation element in house price formation, a negative $\tilde{\alpha}_{ph}$ ensures that prices converge towards their equilibrium value. With a negatively signed adjustment parameter, I will conclude that the bubble hypothesis is rejected.

[3]Note that this approach implicitly imposes weak exogeneity of income and the user cost.

[4]$\tilde{\alpha}_{ph} < 0$ implies equilibrium correction, and thus – from the Engle-Granger representation theorem – it also implies cointegration between ph, y, h and UC.

In a second scenario, $\tilde{\alpha}_{ph} = 0$ independent of the sample end point. In this case, there is no mechanism in place to ensure that disequilibrium constellations are followed by a correction towards a long-run equilibrium value. For instance, if a positive income shock increases house prices, house prices will increase in succeeding periods as well. That said, the finding that $\tilde{\alpha}_{ph} = 0$ is *consistent with*, but does *not imply*, the existence of a bubble. There are several other features that may be consistent with such a finding. For instance, it may imply that we have omitted some relevant economic variables that are not suggested by the theoretical model we have in mind. It may also suggest that we do not have enough time series observations to empirically detect cointegration, or that a model other than the linear I(1) equilibrium-correction model is a relevant representation of the data. For this reason, I shall not necessarily take this as evidence in favor of the bubble hypothesis. Hence, establishing that $\tilde{\alpha}_{ph} = 0$ for all sample end points is an inconclusive finding.

A third scenario is that $\tilde{\alpha}_{ph} < 0$ for a period before dropping to zero. This provides evidence of a structural break in the econometric model. In particular, it implies that we move from a regime where house prices are characterized by equilibrium correction behavior ($\tilde{\alpha}_{ph} < 0$) to a regime where there is no mechanism in place to ensure that prices will move towards the value consistent with the underlying economic fundamentals ($\tilde{\alpha}_{ph} = 0$). Following Anundsen (2015), I will interpret this as a movement from a stable to an unstable market, i.e., a transition to a market exhibiting bubble behavior.

The above discussion shows that the degree to which the market is stable hinges on the significance of the bubble burster term, $\tilde{\alpha}_{ph}$. To construct an operational measure of the extent to which there are disequilibrium constellations in the housing market, one approach is to calculate the p-value of the $\tilde{\alpha}_{ph}$ coefficient for different sample end points.[5] It is then up to the researcher – or policymaker – to define what is a desirable threshold value. In this paper, I consider a 10 percent significance level.

[5] Note that ordinary critical values for the t-distribution can not 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 on http://qed.econ.queensu.ca/pub/faculty/mackinnon/.

A clear advantage of this approach is that the indicator can be calculated in real time, and as soon as a desirable threshold value has been defined, one can infer whether there are signs of a house price bubble. Specifically, successive periods with an indicator value exceeding the threshold value indicates that there is a bubble.

Approach # 4: An asset pricing approach

While the first three approaches take the inverted demand approach as a starting point, an alternative is to consider housing as an asset like any other asset. In that case, the current value of the asset (the house) should be equal to the expected discounted stream of pay-offs in the next period. This framework is similar to a standard present value model (see e.g., Gordon and Shapiro (1956) and Blanchard and Watson (1982)), and Clayton (1996), who argue that it may equally well be considered for housing.

In a housing context, the relevant alternative cost for an owner-occupier is the imputed rent, i.e., what it would have cost to rent a house of similar quality. Asset pricing theory therefore suggests that the price of a house at time t is given by:

$$PH_t = \mathbb{E}_t\left(\frac{PH_{t+1} + R_{t+1}}{1+r}\right) \tag{7}$$

where $\mathbb{E}_t$ is an expectations operator, PH_t denotes house prices, R_t is the imputed rental price and r is a risk free rate used for discounting. Equation (7) states that the price of a house today is equal to the discounted sum of the price of that house tomorrow and the value of living in the house for one period (as measured by the alternative cost, i.e. the imputed rent).

Equation (7) may easily be solved by forward recursive substitution j times to yield:

$$PH_t = \mathbb{E}_t\left[\sum_{i=1}^{j}\left(\frac{1}{1+r}\right)^i R_{t+i} + \left(\frac{1}{1+r}\right)^j PH_{t+j}\right] \tag{8}$$

The transversality condition (TVC) that rules out explosive behavior is given by:

$$\lim_{j \to \infty} \left(\frac{1}{1+r} \right)^j PH_{t+j} < \infty \tag{9}$$

Imposing the TVC, the unique (no bubble) solution to the difference equation in (8) is given as:

$$PH_t = \mathbb{E}_t \left[\sum_{i=1}^{\infty} \left(\frac{1}{1+r} \right)^i R_{t+i} \right] \tag{10}$$

showing that the value of a house today, PH_t is equal to the expected discounted value of all future rents. The expression in (10) may be thought of as a fundamental house price according to the asset pricing approach. It is important to notice that imposing the TVC rules out explosivity, and thus ensures a unique solution to the difference equation.

If we relax the TVC, it can be shown that the (non-unique) solution to the difference equation in (8) (see Sargent (1987) and LeRoy (2004)) is given by:

$$PH_t = \mathbb{E}_t \left[\sum_{i=1}^{\infty} \left(\frac{1}{1+r} \right)^i R_{t+i} \right] + B_t \tag{11}$$

where B_t is an explosive bubble component. Campbell and Shiller (1987) have shown that (11) may alternatively be expressed as:

$$PH_t - \frac{1}{r} R_t = \frac{1+r}{r} \mathbb{E}_t \left[\sum_{i=1}^{\infty} \left(\frac{1}{1+r} \right)^i \Delta R_{t+i} \right] + B_t \tag{12}$$

If the fundamentals (the imputed rent), R_t, follows a RW process with a drift μ, then:

$$\Delta R_t = \mu + \varepsilon_t, \ \varepsilon_t \sim IIN(0, \sigma^2) \tag{13}$$

It is then easy to see that $\mathbb{E}_t[\Delta R_t] = \mu$. Hence (12) may be written as:

$$PH_t - \frac{1}{r}R_t = \frac{1+r}{r}\left[\sum_{i=1}^{\infty}\left(\frac{1}{1+r}\right)^i \mu\right] + B_t \tag{14}$$

Solving the infinite geometric sequence on the right hand side of the above expression, we find:

$$PH_t - \frac{1}{r}R_t = \frac{1+r}{r^2}\mu + B_t \tag{15}$$

In the absence of explosivity ($B_t = 0$), the asset pricing model implies that house prices should have one unit root, and that house prices and rents are cointegrated.[6] However, conditional on the assumption that $R_t \sim RW$, any explosive behavior in PH_t suggests that $B_t \neq 0$, i.e. that there is an explosive bubble component that drives house prices.

With reference to (14), it is clear that the bubble hypothesis is rejected as long as house prices are integrated of the first order, I(1). However, if house prices have an explosive root, the asset pricing theory would suggest that there is a bubble (violation of TVC).

II.1 Econometric operationalization of Approach # 4

I follow Pavlidis et al. (2015) and apply the recursive ADF-based framework suggested by Phillips et al. (2015a) and Phillips et al. (2015b) to explore whether there are signs that house prices in a given country move from following an I(1) process (TVC satisfied and no bubble) to having an explosive root (violation of TVC and bubble behavior). Consider the following generalized ADF-regression model:

$$\Delta X_t = \mu_{r_1,r_2} + \rho_{r_1,r_2}X_{t-1} + \sum_{j=1}^{p}\gamma_{r_1,r_2}\Delta X_{t-j} + \varepsilon_t,\ \varepsilon_t \sim IIN(0, \sigma^2_{r_1,r_2}) \tag{16}$$

[6] With time-varying risk-free rates, house prices, rents and the risk-free rate should be cointegrated. That said, it seems relatively uncontroversial to assume that the risk-free rate follows an I(0)-process, which implies that it will not help for cointegration.

where $r_1 = \frac{T_1}{T}$ and $r_2 = \frac{T_2}{T}$, with T_1, T_2 and T denoting the sample starting point, end point and the total number of observations, respectively. When $T_1 = 0$ and $T_2 = T$, the model is similar to a standard ADF-regression model. What we are interested in testing is the hypothesis that $\rho_{r_1,r_2} = 0$, i.e. $X_t \sim I(1)$, against the alternative that $\rho_{r_1,r_2} > 0$, i.e. X_t is explosive. The relevant test statistic is the ordinary ADF statistic, i.e. $ADF_{r_1}^{r_2} = \frac{\hat{\rho}_{i,r_1,r_2}}{se(\hat{\rho}_{i,r_1,r_2})}$. The ADF statistic has a non-standard limiting distribution that is skewed to the left under the null. Moreover, the distribution depends on both r_2 and nuisance parameters. These critical values may, however, be simulated using a Monte Carlo simulation.[7]

My interest is to explore for what period(s) – if any – the series X_t is explosive. Consider the case where we keep the sample end point fixed, i.e. $r_2 = \bar{r}_2 < \tilde{r}$, and consider the backward ADF (BADF) statistic (Phillips et al. (2015b)):

$$BADF(r_2 = \bar{r}_2) = \sup_{r_1 \in [0, \bar{r}_2 - \tilde{r}]} ADF_{r_1}^{r_2 = \bar{r}_2} \tag{17}$$

By (forward) recursively changing $\bar{r}_2$, we then obtain a time series for the BADF statistic. Comparing this to the relevant critical values, $CV(\alpha)_{r_1}^{r_2}$, we can determine for what periods there is evidence of explosive behavior.

The starting point of the bubble is defined as the first period at which the BADF statistic exceeds this critical value, i.e.:

$$r_{start} = \inf_{r_2 \in [\tilde{r}, 1]} r_2 : BADF_{r_2} > CV(\alpha)_{r_2}^{r_1} \tag{18}$$

Given the start of the bubble (as a fraction of the number of observations), r_{start}, the end of the bubble (as a fraction of the sample), r_{end}, is defined as the first period after the start of the bubble where the BADF statistic is below the critical value. Mathematically,

[7] In the empirical exercise, I use the Matlab program accompanying Phillips et al. (2015a) to simulate consistent finite sample critical values.

this can be expressed as:

$$r_{end} = \inf_{r_2 \in [r_{start},1]} r_2 : BADF_{r_2} < CV\left(\alpha\right)_{r_2}^{r_1} \tag{19}$$

III Data and temporal properties

All data are collected at the quarterly frequency and are seasonally unadjusted as far as possible. The house price data are national house price indices, and the nominal series are transformed to real measures by deflating with the country-specific CPI deflator. The income data measure households' disposable income, and the housing stock is the value of the existing stock of houses. Both the income variable and the housing stock variable are deflated by the population to obtain the per capita measures. My operational measure of the user cost is the tax-adjusted real interest rate, where the real rate is obtained by subtracting CPI inflation.[8] In the US model, I include a dummy, MT, that is equal to one between 1975q1 and 1982q3. This dummy is included to control for interest rate uncertainty during the inflation period of the late 1970s, and a similar dummy has been used in Duca et al. (2011a,b) and Anundsen (2015).

Data sources for each variable for the three countries are listed in Table A.1 in Appendix A. For the US the sample ends in 2016q1[9], whereas my sample ends in 2016q4 for Norway. I was only able to collect data until 2011q4 for Finland. For Norway and Finland my sample starts in 1986q1, as both countries went through a process of substantial deregulation of the housing and credit markets in the early 1980s, which is likely to have altered the functioning of the housing market, so that a different econometric model would probably be more suitable if we were to consider the period prior to deregulation.[10] For the US, I use data from 1975q1, which is as far back as the FHFA house price data

[8]For the US, I also include the property tax rate, which is also deductible.

[9]The latest available observation for the housing stock series.

[10]See Oikarinen (2009) for background on the Finnish deregulation process and Krogh (2010) for a detailed description of the deregulation of the Norwegian credit market.

go.

I have tested the time series properties of the data series for all of the countries using both the Augmented Dickey-Fuller (ADF) test (Dickey and Fuller (1979) and Dickey and Fuller (1981)) and the Phillips-Perron (PP) test (Phillips (1987) and Phillips and Perron (1988)). For all variables, there is evidence of stochastic non-stationarities and I continue my analysis under the assumption that all variables are integrated of order one.

IV Econometric results of bubble detection

Approach # 1: Deviation of actual prices from fundamental prices

I start by presenting the results obtained when I operationalize the first measure of the degree to which house prices are overvalued using historical housing market data for the US, Finland and Norway, following the approach outlined in Section II. For all three countries, I consider a fifth order VAR, which is also consistent with the value that minimizes AIC. With five lags, normality of the disturbances is satisfied for all countries. There is little evidence of residual autocorrelation, while there is some evidence of heteroskedasticity in the US model. Conditional on the lag length, I tested for cointegration using the trace test of Johansen (1988). I find evidence of one cointegrating vector for all three countries, which we would also expect from theory (see Table B.1 in Appendix B for details on the tests for residual mis-specification and cointegration).

In addition to normalizing the cointegrating vector with respect to real house prices, I impose the (testable) overidentifying restrictions that the series are co-trending, that both disposable income and the real after-tax interest rate are weakly exogenous with respect to the long-run parameters ($\alpha_{UC} = \alpha_y = 0$), and that the demand elasticity of income is equal to one. The latter – an income elasticity of demand – $\left(\frac{\beta_y}{\beta_h}\right)$ around one – is in accordance with what Meen (2001), Duca et al. (2011b) and Anundsen (2015) find on US data and is one of the central estimates put out in Meen (2001). Results when these restrictions are imposed are reported in Table 1.

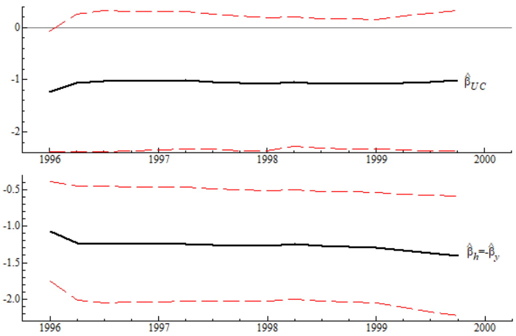
Table 1: Results from recursive CVAR analysis based on inverted demand approach

	US		*Finland*		*Norway*	
Variable	Coeff.	SE	Coeff.	SE	Coeff.	SE
User Cost	-1.015	0.670	-4.824	2.568	-13.813	3.835
Disp. income	1.406	0.407	3.045	1.431	3.536	0.956
Housing stock	-1.406	-	-3.045	-	-3.536	-
Asjustment parameter	-0.130	0.032	-0.119	0.036	-0.093	0.024
$\chi^2(4)$	7.3793	[0.1171]	26.632	[0.0000]	7.1958	[0.1259]

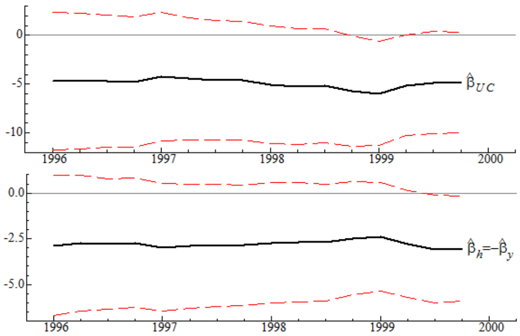
Notes: This table reports a summary of the main results when the system based approach of Johansen (1988) is implemented. The estimation period runs from 1986q1 to 1999q4 for Norway and Finland, while it covers the period 1975q1–1999q4 for the US. The following notation applies: The dependent variable, ph measures real house prices, y is real per capita disposable income, h is the housing stock per capita, while UC measures the real direct user cost.

It is clear that the signs of the estimated coefficients are in accordance with what we would expect from theory in all countries. The responsiveness of house prices with respect to changes in income and the housing stock is of a similar magnitude in Norway and Finland, while it is substantially lower in the US. It is also noticeable that the semi-elasticity with respect to the user cost is much greater in Norway and Finland than in the US. This may reflect the fact that while most loans are fixed-rate mortgages in the US, the majority of loan originations are floating-rate mortgages in Norway and Finland. The finding that the numerical size of the coefficients differs across countries is consistent with the international literature, see Girouard et al. (2006) for an overview of results from a selection of international studies. In Figure 1, I plot the recursively estimated long-run coefficients.

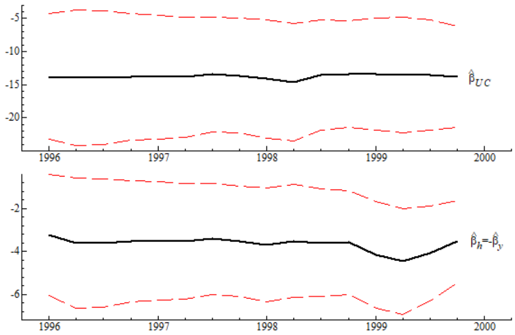
The long-run coefficients are stable when estimated recursively, which is reassuring. It is the estimates reported in Table 1 that I use to construct the first measure of housing market imbalances for the period since 2000q1. In constructing this measure, I assume that house prices were in equilibrium in 2000q1, and I calculate the implied fundamental trajectory of house prices in the ensuing period using (5). This measure is plotted against actual house price developments for the three countries in Figure 2.



(a) US



(b) Finland

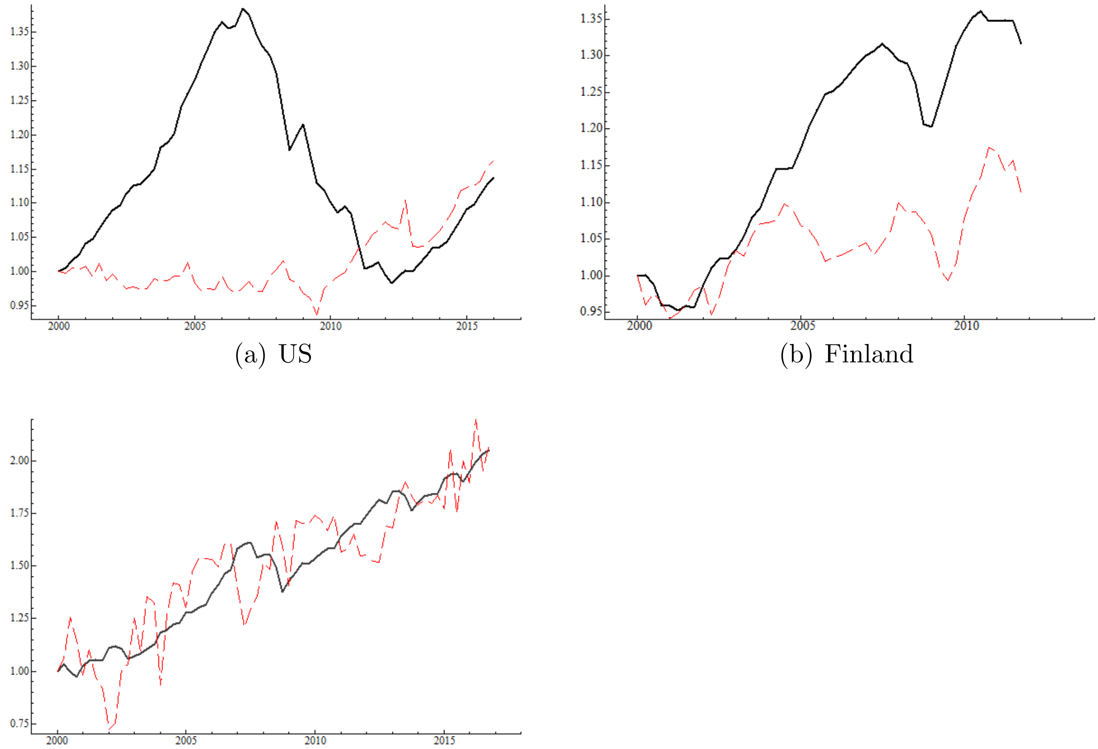


(c) Norway

Figure 1: Panel a) Recursively estimated long-run coefficients for the US with 95 percent confidence intervals, 1996q1–1999q4. Panel b) Recursively estimated long-run coefficients for Finland with 95 percent confidence intervals, 1996q1–1999q4. Panel c) Recursively estimated long-run coefficients for Norway with 95 percent confidence intervals, 1996q1–1999q4

This first measure clearly suggests that there were sustained equilibrium deviations in the US housing market in the 2000s, see Panel (a). By 2006, the measure suggests that US house prices were overvalued by nearly 50 percent. It is also interesting to see that the same measure indicates that current house prices in the US are close to – or even a bit below – the value implied by the underlying fundamentals.

From Panel b), we see that the measure suggests a growing departure from fundamentals in the Finnish housing market in the mid 2000s, and that house prices were overvalued by around 20 percent in 2011. For Norway, see Panel c), we see that there are periods where prices are above what is implied by the fundamentals, but there seems to be no tendency for prices to be systematically overvalued.



(a) US

(b) Finland

(c) Norway

Figure 2: Panel a) Implied equilibrium value (dotted red) against actual house prices (solid black) for the US, 2000q1–2016q1. Panel b) Implied equilibrium value (dotted red) against actual house prices (solid black) for Finland, 2000q1–2011q4. Panel c) Implied equilibrium value (dotted red) against actual house prices (solid black) for Norway, 2000q1–2016q4

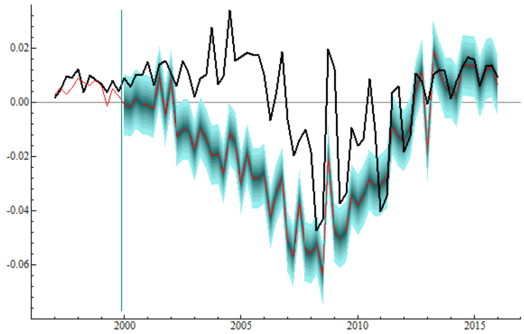
Approach # 2: Systematic underpredictions

In the previous section, I documented that there exists evidence of one cointegrating vector in all countries in the pre-boom period. Further, the results showed that both the interest rate and disposable income are weakly exogenous with respect to the long-run parameters. Thus, we know that we can – without loss of efficiency – abstract from the marginal models of these variables (see e.g., Johansen (1994)). For that reason, the dynamic forecasting model used to construct the second measure of housing market imbalances is a conditional equilibrium correction model of the type described by (6).

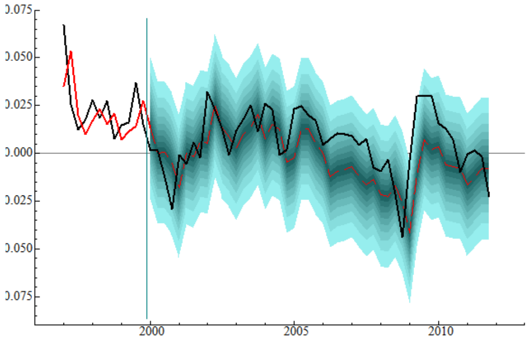
To reduce the dimensionality of this model, I use the automated variable selection algorithm *Autometrics* (see e.g., Doornik (2009) and Doornik and Hendry (2009)) to find a parsimonious model nested in the general unrestricted model (GUM) in (6). This algorithm automatizes the Gets approach and can also handle cases where regressors are not mutually orthogonal and when the number of variables exceed the number of observations. An evaluation of the search algorithm is given in Castle et al. (2011), who consider 59 different Monte Carlo experiments and show that the algorithm is indeed a successful variable selection device. I use a significance level of 5 percent to reduce the dimension of the GUM.

Having obtained a more parsimonious specification, the second measure of housing market imbalances can be constructed. This measure is simply the dynamic forecasts for house prices. Systematic and persistent underpredictions of actual house price growth is consistent with bubble behavior. It is important to note that the forecasting model is specified and estimated using information only from the pre-boom sample. The forecasting model is then used to construct forecasts for the suspected bubble period.

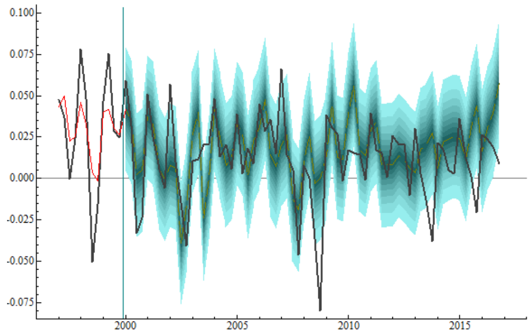
I constructed forecasts based on the models for the three countries, and the conditional forecasts for each of the countries are displayed along with actual developments in house prices in Figure 3.



(a) US



(b) Finland



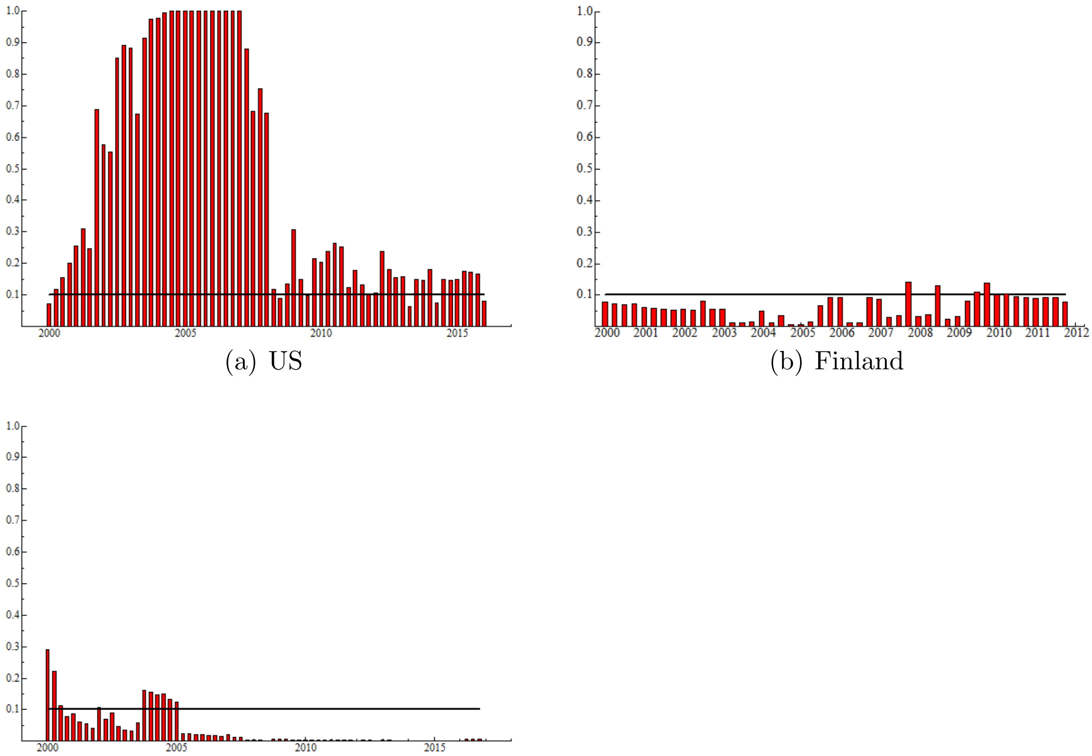
(c) Norway

Figure 3: Panel a) Forecasts (dotted red) against actual house price growth (black) for the US. Green fans show 95 percent confidence intervals, 2000q1–2016q1. Panel b) Forecasts (dotted red) against actual house price growth (black) for Finland. Green fans show 95 percent confidence intervals, 2000q1–2011q4. Panel c) Forecasts (dotted red) against actual house price growth (black) for Norway. Green fans show 95 percent confidence intervals, 2000q1–2016q4.

Exploring this alternative measure for assessing house prices, we reach the same conclusion regarding US house prices as previously; there was an increasing disconnect between predicted and actual house prices starting in the early 2000s. The forecast errors are both significant and very persistent, which is consistent with bubble behavior. For the case of Finland and Norway, there is no evidence that we systematically underpredict house price growth, and the forecasts are indeed within the 95 percent confidence bounds over the entire period. Hence, for Finland and Norway, we conclude that there has not been a bubble in these countries according to this measure.

Approach # 3: Econometrically based bubble indicators

As explained in Section II, one approach to constructing real time measures of the degree of overheating in the housing market is to follow the approach outlined in Anundsen (2015). I have followed this approach to construct bubble indicators for the three countries, and the resulting indicators are displayed in Figure 4.[11]



(a) US

(b) Finland

(c) Norway

Figure 4: Panel a) Bubble indicator for the US, 2000q1–2016q1. Panel b) Bubble indicator for Finland, 2000q1–2011q4. Panel c) Bubble indicator for Norway, 2000q1–2016q4.

Again, we clearly see evidence of bubble behavior in the US housing market, which corroborates the findings of Anundsen (2015), who constructed a similar indicator, and it corroborates the evidence from the other approaches considered in this paper. Compared to Anundsen (2015), I have extended the sample for the calculation of the indicator to also include the years from 2010 through 2015, and we see that it suggests that current

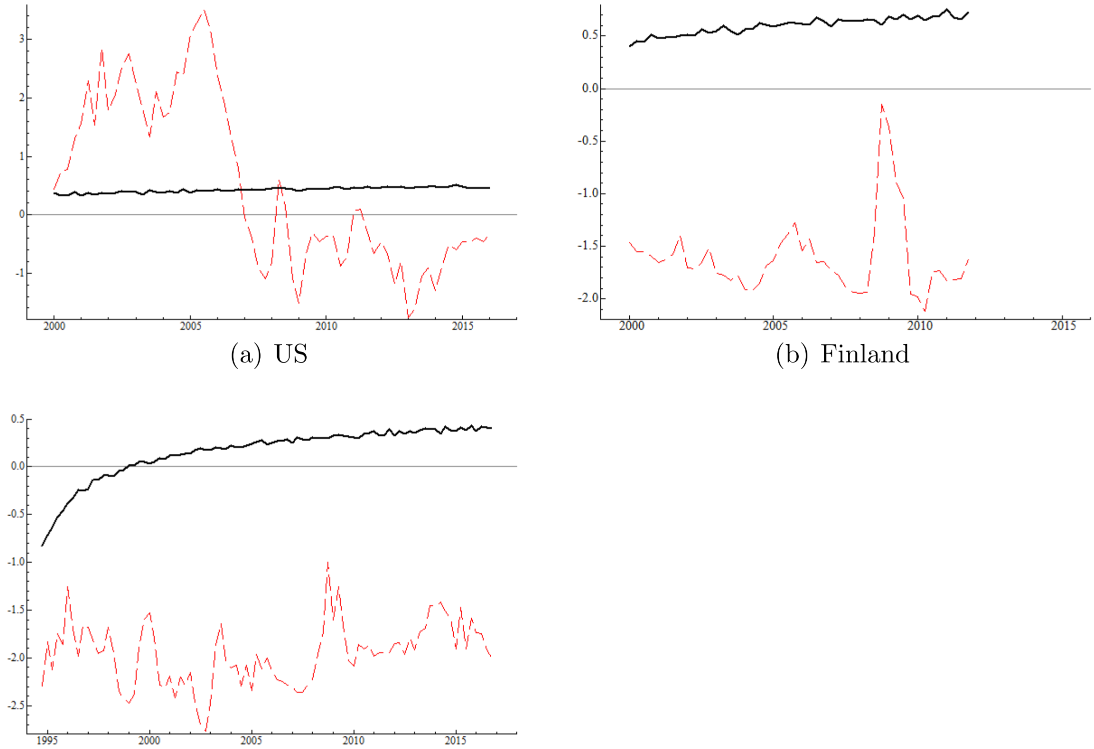
[11] I impose the restriction that the coefficient on income is equal to the negative of the coefficient on the housing stock in the long run. This ensures that the models are consistent with the results from the multivariate cointegration analysis.

house prices in the US are not characterized by bubble behavior. Turning to Finland and Norway, there is no evidence of bubble behavior, which is in line with the results from the other measures.

Approach # 4: Testing for explosiveness

The final measure of housing market imbalances is constructed similar to Pavlidis et al. (2015) and is aimed at testing for explosiveness in the price-to-income ratio. However, as opposed to them, I consider the log of the ratio, which moves the residuals in the ADF regressions closer to satisfying normality. I consider an ADF regression with four lags and a deterministic trend. The sequence of finite sample critical values has been simulated using $M = 5000$ Monte Carlo replications and the minimum window size is set to 36 quarters.

In Figure 5, I have plotted the recursive BADF statistics along with the 5 percent critical values. It is evident that these results corroborate the results from the other approaches, i.e. while there is no evidence of explosive behavior in Norwegian or Finnish house prices over the period considered, there is clear evidence that the US housing market transitioned into a bubble regime in the early 2000s. Moreover, in line with the other measures, the bubble is dated to have started in the first quarter of 2001 and ended in the middle of 2006.



(a) US (b) Finland

(c) Norway

Figure 5: Panel a) Test for transition to explosivity for the US, 2000q1–2016q1. Panel b) Test for transition to explosivity for Finland, 2000q1–2011q4. Panel c) Test for transition to explosivity for Norway, 2000q1–2016q4.

V What explains the US housing bubble?

The econometric evidence strongly supports the view that there was a bubble in the US housing market in the early to mid-2000s. Some authors have attributed the price increase to an increased inflow of capital from emerging market economies (see Bernanke and Kuttner (2005), Caballero and Krishnamurthy (2009) and Bernanke et al. (2011)). The global savings glut hypothesis states that increased demand for US assets pushed down long-term interest rates and thereby increased housing demand. Other strands of the literature have suggested that the house price increases in the 2000s were due to an increased access to credit following the subprime explosion, see e.g., Mian and Sufi (2009), Duca et al. (2011a,b), Pavlov and Wachter (2011) and Anundsen (2015). Unlike the US, neither Norway nor Finland had subprime mortgages, so this phenomonon is particular to the US. Furthermore, whereas the US had been running major current account deficits since the 1990s, both Norway and Finland had been running major surpluses over this period. Thus, both these factors could provide a reasonable explanation as to why the US experience was so different from the Finnish and Norwegian experience.[12]

To investigate the relevance of each of these hypotheses in explaining the US housing bubble, I investigate their importance in explaining the deviation between actual and fundamental house prices, as implied by Approach # 1.

Results are displayed in the second column of Table 2. These results suggest that both hypotheses carry some relevance and that they in total explain nearly 80 percent of the deviation between actual and fundamental prices in the bubble period. I also conducted a similar exercise for the deviation between the house price forecasts and actual realizations (Approach # 2). Results are similar in that case and are reported in the third column of Table 2.

Table 2: Explaining the US bubble

Variable	House price gap	Forecast error
Subprime share	1.333 (0.256)	0.135 (0.043)
Current account deficit	9.203 (0.783)	0.862 (0.132)
Adj. R^2	0.777	0.523

Notes: The second column of table report results from regressing the gap between actual and fundamental prices as calculated in my Approach # 1 on the share of new mortgages given to the subprime segment and the current account deficit as a fraction of GDP. The third column repeats this exercise for the difference between the forecasted house price growth and the actual realization, as implied by my Approach # 2. Standard errors are reported in parenthesis next to the point estimates.

[12]The source of the subprime data is the National Delinquency Survey of the Mortgage Bankers Association and have been collected from Moody's Analytics' Economy.com webpage. Information on the current account deficit as a share of GDP was collected from the FRED data base of the Federal Reserve Bank of St. Louis.

VI Conclusions

This paper has considered four alternative indicators of housing market instability for Norway, Finland and the US. The combined evidence from the indicators may be useful to evaluate the temperature of the housing market. In particular, a relevant evaluation of such indicators is that they do not send a signal as soon as house prices are increasing, since this does not necessarily imply that there is a bubble.

I find that all four indicators strongly suggest that there was a bubble in the US housing market starting in the early 2000s that was pricked in 2006, and that US house prices today are in line with underlying economic fundamentals. The same indicators do not indicate bubble behavior in Norwegian house prices, while one of the measures suggest that Finnish house prices may be overvalued. Though a majority of the measures do not suggest a bubble in Norway and Finland, prices may of course fall in these countries in the case of less fortunate developments in the economic fundamentals, e.g., an increase in the mortgage interest rate or a drop in household income. What the results, however, do suggest is that there are no signs of an expectations-driven bubble in these two countries.

My results suggest that the deviation between actual and fundamental house prices in the US in the 2000s can be attributed to the explosion in subprime lending and the sharp increase in the US current account deficit. In total, I find that these two variables explain about 80 percent of the overvaluation in the US housing market in the 2000s.

The development and assessment of alternative indicators of instabilities in the housing market is important for policy institutions that are constantly monitoring the housing market and it may also be of great importance in preventing future housing market crashes of the type witnessed in many countries in the late 2000s.

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A Data definitions

Table A.1: Variable definitions and data sources

Series	Description	US	Finland	Norway
PH	House price index	FHFA	BoF	NB/SN
P	Price deflator	BLS	SF	NB/SN
H	Housing stock	LILP	SF	NB/SN
Y	Households' disposable income	BEA	BoF	NB/SN
i	Mortgage interest rate	FHFA	BoF	SN
τ_y	Capital gains tax rate	FRB-US	BoF	NB/SN
POP	Population	CB	SF	NB/SN

Notes: This table reports data descriptions and sources for the analyses of this paper. The data period runs from 1986q1 to 2014q4 for Norway, from 1986q1–2011q4 for Finland, while it covers the period 1975q1–2013q3 for the US. The abbreviations are the following: BEA=Bureau of Economic Analysis, BLS=Bureau of Labor Statistics, BoF = Bank of Finland, NB = Norges Bank, CB=Census Bureau, FHFA=Federal Housing Finance Agency, LILP=Lincoln Institute for Land Policy, NIPA=National Income and Product Accounts, SN = Statistics Norway and SF = Statistics Finland.

B Cointegration tests

Table B.1: Trace test for cointegration

H_0	H_A	λ_{trace}			
		US	Finland	Norway	1%-critical value
$r = 0$	$r \geq 1$	55.01	76.97	59.00	56.83
$r \leq 1$	$r \geq 2$	29.55	36.66	33.91	36.44
$r \leq 2$	$r \geq 3$	11.92	13.85	14.67	19.53
Diagnostics					
Autocorrelation			1.5504 [0.0262]	0.9197 [0.6001]	1.1754 [0.2835]
Normality			3.6890 [0.7187]	2.1329 [0.9071]	6.4898 [0.3706]
Heteroskedasticity			1.5543 [0.0002]	1.2558 [0.1129]	1.1901 [0.1779]

Notes: The endogenous variables are real housing prices (ph), real disposable income (ydp) and the real direct user cost (UC). A deterministic trend and the housing stock, h, are restricted to enter the cointegration space. A constant and three centered seasonal dummies enter unrestrictedly. Consistent critical values controlling for the inclusion of one weakly exogenous variable in the cointegration space are tabulated in Table 13 in Doornik (2003).