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## Detecting Imbalances in House Prices: What Goes up Must Come Down?

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# WORKING PAPER

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

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## Abstract

With the aid of econometric modeling, I investigate whether rapidly increasing house prices necessarily imply the existence of a bubble that will eventually burst. I consider four alternative econometric methods to construct indicators of housing market imbalances for the US, Finland and Norway. The four approaches are used to study if house prices in these countries in the 2000s can be explained by underlying economic fundamentals, or whether the developments are best characterized by bubble-dynamics. 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. Only one of the measures indicate imbalances in the Finnish housing market, while none of the measures suggest a bubble in Norway.

**Keywords:** *Cointegration; Explosive Roots; Housing Bubbles.*

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

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# 1 Introduction

Is it so 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 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. The motivation for looking at these countries is that 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. House prices increased rapidly in all these 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 such imbalances in real time so that necessary actions can be taken to prevent further imbalances given the strong role that the housing market has in affecting the real economy through consumption wealth effects (Aron et al., 2012; Mian et al., 2013) and through its interactions with the credit market (Hofmann, 2003; Fitzpatrick and McQuinn, 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 form a strong indicator for future recessions in the US economy.

For the case of the US, there has been a long standing discussion in the academic literature on the extent to which the 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 not even in retrospect may be characterized as a bubble. Anundsen (2015) has contested this conclusion by constructing an econometrically based bubble-indicator that clearly signals a bubble in the US housing market in the early 2000s.

The approach in Anundsen (2015) relies on the bubble definition provided by Stiglitz (1990, p.13), that says 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 ‘fun-

damental' factors do not seem to justify such a price". This definition is combined with the modeling assumption that fundamental factors – if they exist – 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 for cointegration can be found, we cannot reject a bubble, but it may also reflect 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).

In this paper, I construct a similar indicator for the US (on a sample covering an additional four years of the housing recovery), Finland and Norway to explore whether the house price behavior in the three countries in the 2000s can be explained by economic fundamentals. In addition to this indicator, I consider three other measures to identify bubble behavior. The first of these measures calculates a fundamental house price relationship using information available in 1999q4, and then asks the question of how actual house prices moved relative to the model-implied fundamental house prices in the period thereafter. The second measure 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 might be interpreted as an overheating of house prices. 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. While both the approach based on deviations of actual prices from fundamental prices and the forecasting exercise may be particularly relevant for an *ex post* evaluation of the extent of overheating in the housing market, the other two measures might be used also for real time monitoring. A caveat with all four measures is that they are conditional on the fundamentals developing along a stable path. If there is a bubble in the fundamentals, the methods may fail to detect a housing bubble, since they would characterize the house price developments as stable relative to a set of non-stable fundamentals.

My results show that all four measures detect a bubble in the US housing market starting in the early 2000s. Using the same measures, I do not find evidence of a bubble in the Norwegian housing market for a sample ending in 2014, 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 development 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 income – is consistent with a fall in house prices. In addition, if the fundamentals are developing along an unstable trajectory, the measures may fail to detect a bubble.

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 3, while results based on each of the indicators are presented in Section 4. The final section concludes the paper.

## 2 Four alternative indicators of house price misalignments

To discuss whether house prices in any given housing market is best characterized as exercising 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.

### 2.1 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) and Muellbauer and Murphy (1997)). 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:

$$\frac{U_H}{U_C} = PH_t \left[ (1 - \theta_t)i_t - \pi_t + \delta - \frac{\dot{P}H_t}{PH_t} \right] \quad (1)$$

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  $\theta_t$ ,  $i_t$  and  $\pi_t$  denote time-varying tax rates, mortgage interest rates and the general CPI inflation rate, respectively. Finally,  $\delta$  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 forgone consumption of other goods. Since the housing market also contains a rental sector, market efficiency requires the following condition to be satisfied in equilibrium:

$$Q_t = PH_t \left[ (1 - \theta_t)i_t - \pi_t + \delta - \frac{\dot{P}H_t}{PH_t} \right]$$

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

$$\frac{PH_t}{Q_t} = \frac{1}{UC_t} \quad (2)$$

where the real user cost,  $UC$ , is defined as:  $UC_t = (1 - \theta_t)i_t - \pi_t + \delta - \frac{PH_t}{PH_t}$ . The real imputed rent is unobservable, but two approximations are common: to proxy the imputed rent by an observable rent  $R_t$ , 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:

$$\frac{PH_t}{R_t} = \frac{1}{UC_t} \quad (2a)$$

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

$$R_t = Y_t^{\beta_y} H_t^{\beta_h}, \quad \beta_y > 0 \text{ and } \beta_h < 0$$

(2) would read:

$$\frac{PH_t}{Y_t^{\beta_y} H_t^{\beta_h}} = \frac{1}{(1 - \theta_t)i_t - \pi_t + \delta - \frac{PH_t}{PH_t}} \quad (2b)$$

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:<sup>1</sup>

$$ph_t = \beta_r r_t + \tilde{\beta}_{UC} UC_t \quad (3a)$$

$$ph_t = \beta_y y_t + \beta_h h_t + \beta_{UC} UC_t \quad (3b)$$

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 (1997); Muellbauer and Murphy (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 extensively used in the US literature, it has been less commonly applied to house price modeling in Europe,

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<sup>1</sup>A semi-logarithmic representation is commonly used, since the user cost may take negative values.



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 consider 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 what is the implied fundamental house price path during the boom period using data until 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 produce (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 equilibrating 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 eventually will revert to the value implied fundamentals anyway), or if there is a tendency that there is no adjustment towards a long-run equilibrium value

### 2.1.1 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 are reverting to a level which is 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  $\mathbf{y}_t = (\mathbf{x}_t', h_t)'$ , where  $\mathbf{x}_t$  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):

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

where the vector  $\mathbf{x}_t$  contains real house prices,  $ph$ , real disposable income,  $y$ , and the real direct user cost,  $UC$ . The vector  $\mathbf{D}$  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 changes only 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,  $\boldsymbol{\varepsilon}_t \sim MVN(\mathbf{0}, \boldsymbol{\Sigma})$ .

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

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

where  $\alpha_{ph}$ ,  $\alpha_y$  and  $\alpha_{UC}$  are the adjustment parameters, i.e., measuring how the three endogenous variables in the VAR responds to last period equilibrium deviations, as measured by  $\boldsymbol{\beta}' \mathbf{y}_{t-1} = (ph - \beta_y y - \beta_h h - \beta_{UC} UC - \beta_{trend} trend)_{t-1}$ . The vector  $\tilde{\mathbf{D}}_t$  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 with this approach is that data from the period being scrutinized are used in developing the measures being 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.<sup>2</sup>

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<sup>2</sup>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:

$$ph_t^* = ph_{t-1}^* + \hat{\beta}_y^{1999q4} \Delta y_t + \hat{\beta}_h^{1999q4} \Delta h_t + \hat{\beta}_{UC}^{1999q4} \Delta UC_t ; t = 2000q1, \dots, 2014q4 \quad (5)$$

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.

### 2.1.2 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).

Irrespective of the econometric model, it can be used to construct forecasts of house price growth, which can be used to evaluate the temperature in 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.

### 2.1.3 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 ec-model as a starting

point:<sup>3</sup>

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

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).<sup>4</sup> Three cases are of particular interest:

1.  $\tilde{\alpha}_{ph} < 0$  for  $t = 1, \dots, s + i$ ;  $\forall i = 0, \dots, T - s$ , where  $T > s$  is the last available time series observation in the data set.
2.  $\tilde{\alpha}_{ph} = 0$  for  $t = 1, \dots, s + i$ ;  $\forall i = 0, \dots, T - s$ , where  $T > s$  is the last available time series observation in the data set.
3.  $\tilde{\alpha}_{ph} < 0$  for a sample  $t = 1, \dots, s$ , but  $\tilde{\alpha} = 0$  for  $t = s + 1, \dots, s + i \forall i = 2, \dots, B - s$  and  $\tilde{\alpha}_{ph} < 0$  for  $t = B + 1, \dots, T$ , where  $\leq T - s$  and the bubble period runs from  $s + 1$  to  $B$ .

In the first case, we would have formal statistical evidence of cointegration independent of what we choose to be the sample end point. This finding is consistent with a stable market, where disequilibrium constellations 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.

Contrary to first case, the second case is consistent with bubble behavior. When  $\tilde{\alpha}_{ph} = 0$ , independent of the sample end point, 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

<sup>3</sup>Note that this approach implicitly imposes weak exogeneity of income and the user cost.

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

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 another model 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.

The third case is interesting, since it gives clear 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 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. Following Anundsen (2015), I will interpret this as a movement from a stable to an unstable market, i.e., a transition to a market exercising 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.<sup>5</sup> 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.

This methodological approach for detecting imbalances in the housing market was suggested by Anundsen (2015), who applied it to US housing market data. Indeed, relying on the two different econometric operationalizations of the theory model outlined in Section 2.1.3 (the price-to-rent approach and the inverted demand approach), the author showed that the US housing market transitioned from a stable market characterized by equilibrium correction behavior to a highly unstable (bubble) market in the early 2000s.

A clear advantage with 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 the extent to which house prices are overvalued. Specifically, successive periods with an indicator value exceeding the threshold value could be taken as an indication that house prices are overvalued.

## 2.2 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 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) 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

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<sup>5</sup>Note that ordinary critical values for the t-distribution can not be used under the null of no cointegration as the distribution of  $\alpha_{ph}$  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/>.

therefore suggests that the price of a house at time  $t$  is given by:

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

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 that is 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:

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

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

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

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

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

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:

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

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

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

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

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

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

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

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

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

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.<sup>6</sup> 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 has an explosive root, the asset pricing theory would suggest that there is a bubble (violation of TVC).

### 2.2.1 Econometric operationalization

I follow Pavlidis et al. (2015) and apply the recursive ADF-based framework suggested by Phillips et al. (2011), 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 standard ADF-regression model for country  $i$ :

$$\Delta X_{i,t} = \mu_i + \rho_i X_{i,t-1} + \sum_{j=1}^p \phi_{i,j} \Delta X_{i,t-j} + \varepsilon_{i,t} \quad (16)$$

When  $\rho_i = 0$ ,  $X_t$  contains one unit root. The standard ADF-test, tests the null of a unit root against the alternative of stationarity ( $\rho_i < 0$ ). With reference to the asset pricing model, the alternative of stationarity seems less relevant, however. The hypothesis we are interested in testing is whether house prices are I(1) v.s. the alternative that they are explosive, i.e.  $\rho_i > 0$ .

The framework suggested by Phillips et al. (2011, 2015a,b) is to consider a recursive version of the ADF-test, so that we can explore whether there are periods when a time

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<sup>6</sup>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.

series exercises I(1) behavior and other periods where it has an explosive root. The general ADF-regression model that this test is based on takes the following form:

$$\Delta X_{i,t} = \mu_{i,r_1,r_2} + \rho_{i,r_1,r_2} X_{i,t-1} + \sum_{j=1}^p \gamma_{i,r_1,r_2} \Delta X_{i,t-j} + \varepsilon_{i,t}, \quad \varepsilon_{i,t} \sim IIN(0, \sigma_{i,r_1,r_2}^2) \quad (17)$$

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. Thus, imposing  $T_1 = 0$  and  $T_2 = T$ , we are back at the standard ADF-regression model in (16). What we are interested in testing is the hypothesis that  $\rho_{i,r_1,r_2} = 0$ , i.e.  $X_{i,t} \sim I(1)$ , against the alternative that  $\rho_{i,r_1,r_2} > 0$ , i.e.  $X_{i,t}$  is explosive. The relevant test statistic is the ordinary ADF statistic, i.e.  $ADF_{r_1}^{r_2} = \frac{\hat{\rho}_{i,r_1,r_2}}{se(\hat{\rho}_{i,r_1,r_2})}$

Phillips et al. (2011) suggested to set  $T_1 = 0$ , while varying  $T_2$  from  $\tilde{T}$  to  $T$  ( $\tilde{T} < T$ ), i.e. an expanding forward recursive window. To test whether there are any periods with evidence of explosive behavior, they suggested to use the sup ADF statistic (SADF), which is given by:

$$SADF(r_1 = 0) = \sup_{r_2 \in [\tilde{r}, 1]} ADF_{r_1=0}^{r_2} \quad (18)$$

with  $\tilde{r} = \frac{\tilde{T}}{T}$ . Like the ordinary ADF statistic, the SADF statistic has a non-standard limiting distribution that is skewed to the left. Moreover, the distribution depends on both  $r_2$  and nuisance parameters. These critical values may, however, be simulated and the null of non-stationarity is rejected in favor of explosivity when the SADF statistic is greater than the corresponding critical value from the right-tail of the relevant Dickey-Fuller distribution.

While this test has been shown to perform well in the case of only one bubble, it has been shown to work less well (low power) when there are multiple bubbles (see Homm and Breitung (2012)). Therefore, Phillips et al. (2015b) and Phillips et al. (2015a) suggested a modified version of the test, where both  $T_1$  and  $T_2$  are allowed to vary, i.e. both the sample starting point and the sample end point vary. The relevant test statistic is called the generalized SADF (GSADF) statistic and is given by:

$$GSADF = \sup_{r_2 \in [\tilde{r}, 1], r_1 \in [0, r_2 - \tilde{r}]} ADF_{r_1}^{r_2} \quad (19)$$

As with the standard ADF statistic and the SADF statistic, the GSADF statistic has a non-standard limiting distribution, and the distribution of GSADF under the null of non-stationarity depends on both  $r_1$ ,  $r_2$ , as well as the inclusion of nuisance parameters.<sup>7</sup> A rejection of the null indicates that there are signs of explosive behavior.

In most cases, it is relevant to ask for what period(s) – if any – the series  $X_{i,t}$  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)):

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



$$BADF(r_2 = \bar{r}_2) = \sup_{r_1 \in [0, \bar{r}_2 - \bar{r}]} ADF_{r_1}^{r_2 = \bar{r}_2} \quad (20)$$

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.:

$$r_{start} = \inf_{r_2 \in [\bar{r}, 1]} r_2 : BADF_{r_2} > CV(\alpha)_{r_2}^{r_1} \quad (21)$$

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, this can be expressed as:

$$r_{end} = \inf_{r_2 \in [r_{start}, 1]} r_2 : BADF_{r_2} < CV(\alpha)_{r_2}^{r_1} \quad (22)$$

### 3 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.<sup>8</sup> 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 both the US and Norway, my sample ends in 2014q4, while 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 1980's, 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.<sup>9</sup> For the US, I use data from 1975q1, which is as far back as the FHFA house price data goes.

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

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<sup>8</sup>For the US, I also include the property tax rate, which is also deductible.

<sup>9</sup>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.

Fuller (1981)) and the Phillips-Perron (PP) test (Phillips (1987) and Phillips and Perron (1988)). For all variables there are evidence of stochastic non-stationarities and I continue my analysis under the assumption that all variables are integrated of order one.

## 4 Econometric results of bubble detection

### 4.1 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 2.1.1. 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 are 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

Variable	<i>US</i>		<i>Finland</i>		<i>Norway</i>	
	Coeff.	SE	Coeff.	SE	Coeff.	SE
User Cost	-1.030	0.618	-4.824	2.568	-13.984	3.810
Disp. income	1.307	0.372	3.045	1.431	3.538	0.952
Housing stock	-1.307	-	-3.045	-	-3.538	-
Asjurement parameter	-0.139	0.033	-0.119	0.036	-0.093	0.023
$\chi^2(4)$	5.6402 [0.2277]		26.632 [0.0000]		6.7227 [0.1513]	

*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 capital, 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 elasticity of house prices with respect to income and the housing stock are of 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 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 differ 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.

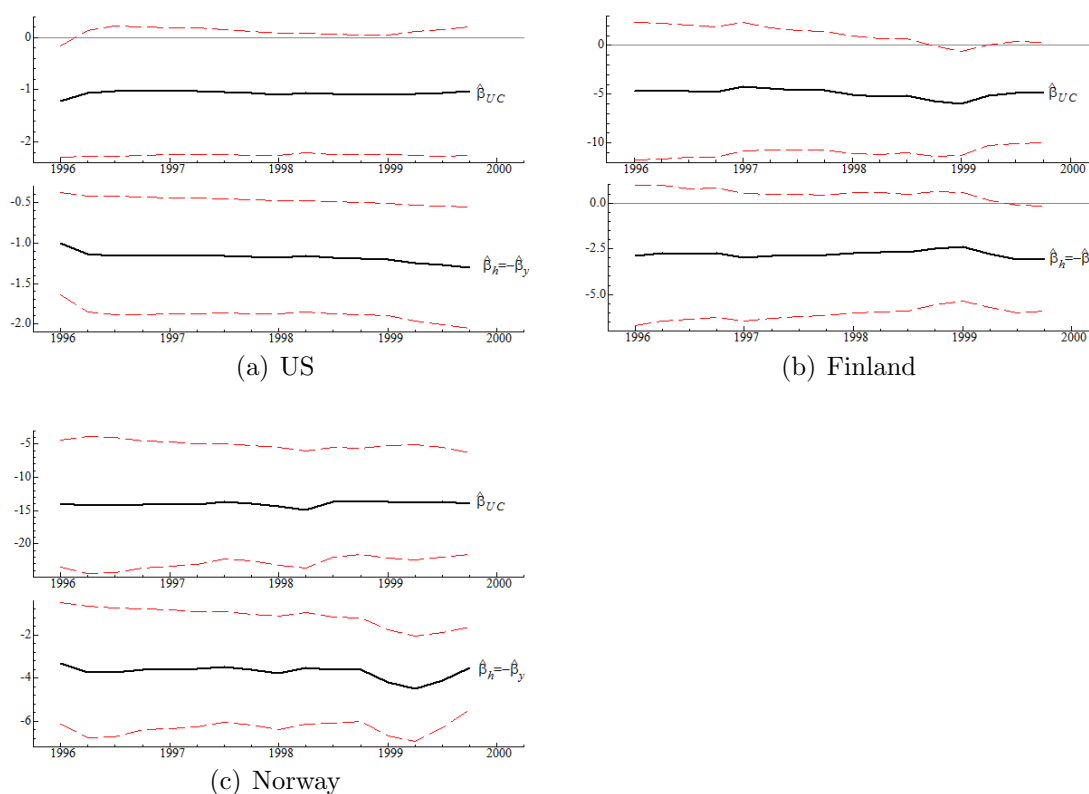


Figure 1: Panel a) Recursively estimated long-run coefficients for the US with 95 percent confidence intervals, 1996q1–1999q4. Panel b) Recursively estimated long-run coefficients for Finland with 95 percent confidence intervals, 1996q1–1999q4. Panel c) Recursively estimated long-run coefficients for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

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 2000q1–2014q4 period. 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.

This first measure clearly suggest that there were sustained equilibrium deviations in

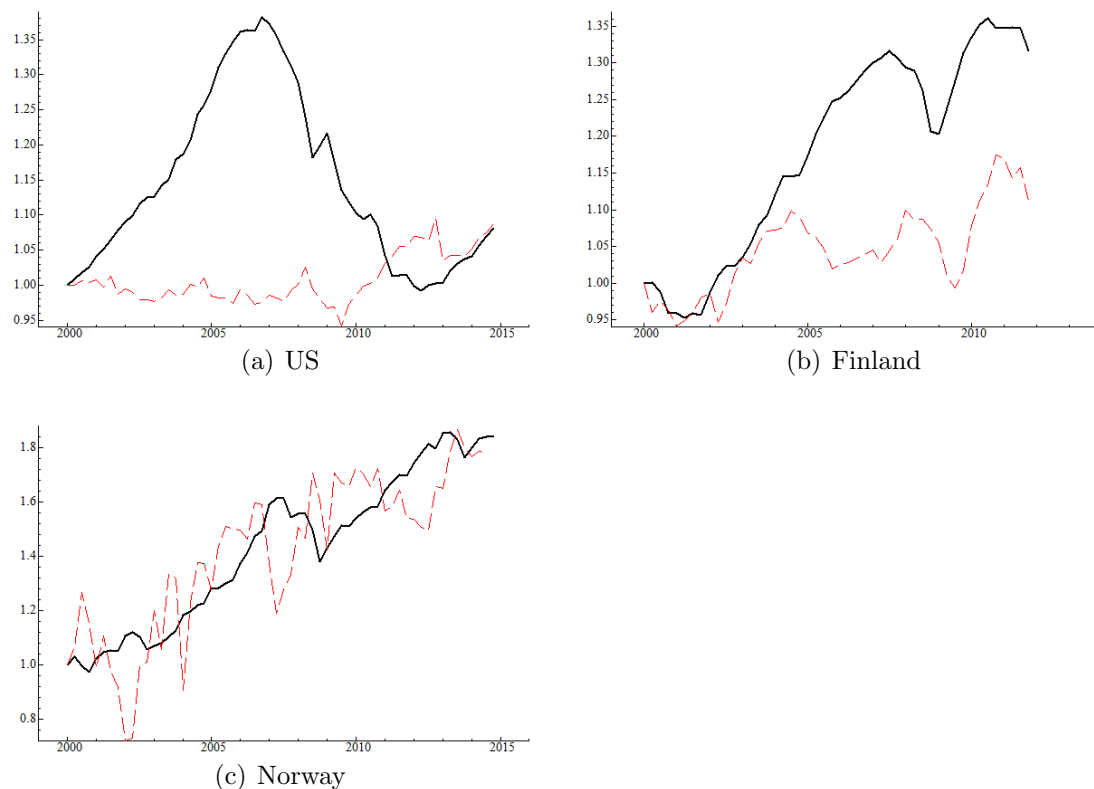


Figure 2: Panel a) Implied equilibrium value (dotted red) against actual house prices (solid black) for the US, 2000q1–2014q4. Panel b) Implied equilibrium value (dotted red) against actual house prices (solid black) for Finland, 2000q1–2011q4. Panel c) Implied equilibrium value (dotted red) against actual house prices (solid black) for Norway, 2000q1–2014q4

the US housing market in the 2000s, see Panel (a). By 2006, the measure suggest that US house prices were overvalued by nearly 50 percent. It is also interesting to see that the same measure indicate 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 suggest 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.

## 4.2 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 the actual house price growth is consistent with bubble behavior.

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

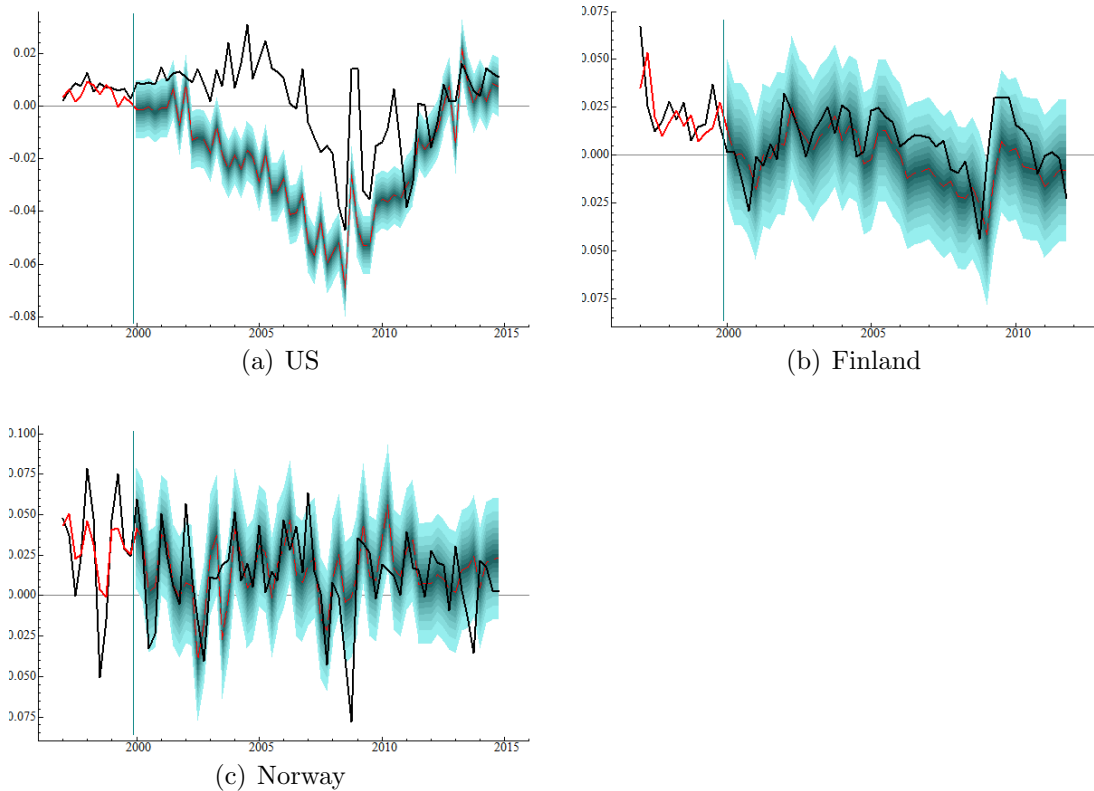


Figure 3: Panel a) Forecasts (dotted red) against actual house price growth (black) for the US. Green fans show 95 percent confidence intervals, 2000q1–2014q4. Panel b) Forecasts (dotted red) against actual house price growth (black) for Finland. Green fans show 95 percent confidence intervals, 2000q1–2011q4. Panel c) Forecasts (dotted red) against actual house price growth (black) for the Norway. Green fans show 95 percent confidence intervals, 2000q1–2014q4.

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 For 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.

### 4.3 Approach # 3: Econometrically based bubble indicators

As explained in Section 2.1.3, one approach to construct 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.<sup>10</sup>

<sup>10</sup>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

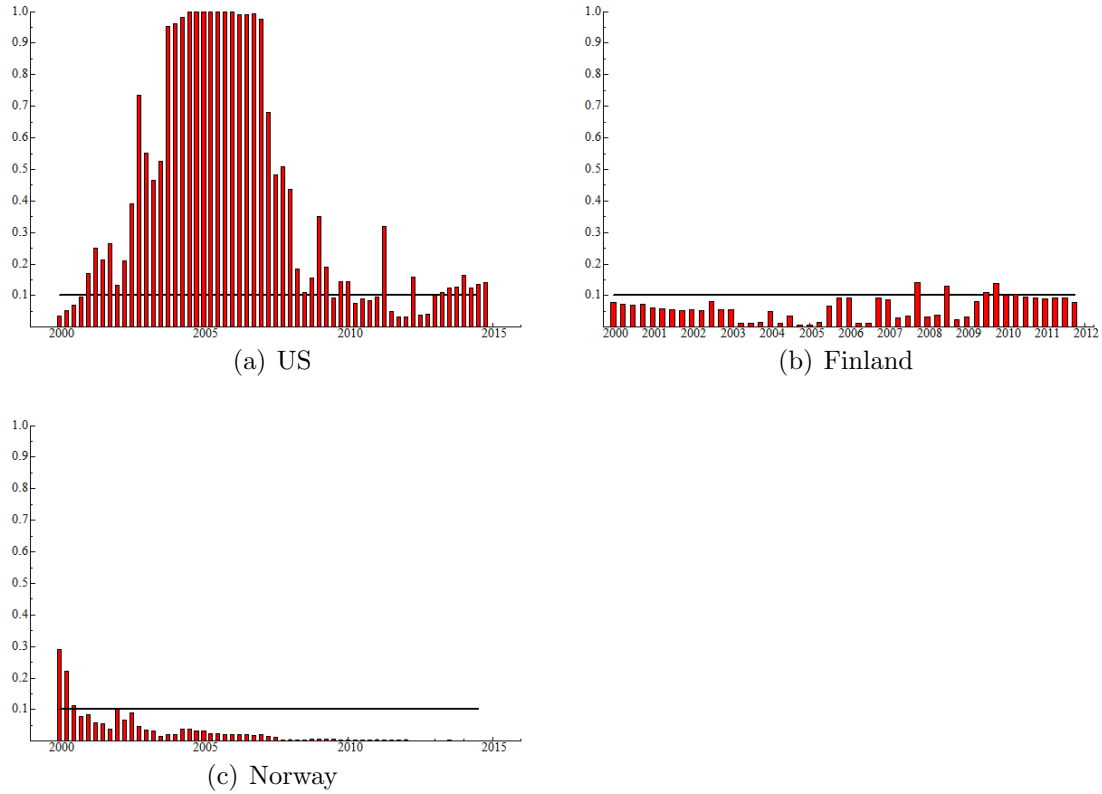


Figure 4: Panel a) Bubble indicator for the US, 2000q1–2014q4. Panel b) Bubble indicator for Finland, 2000q1–2011q4. Panel c) Bubble indicator for Norway, 2000q1–2014q4.

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 2014, and we see that it suggests that current 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.

#### 4.4 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 have been simulated using  $M = 5000$  Monte Carlo replications.

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 multivariate cointegration analysis.

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.

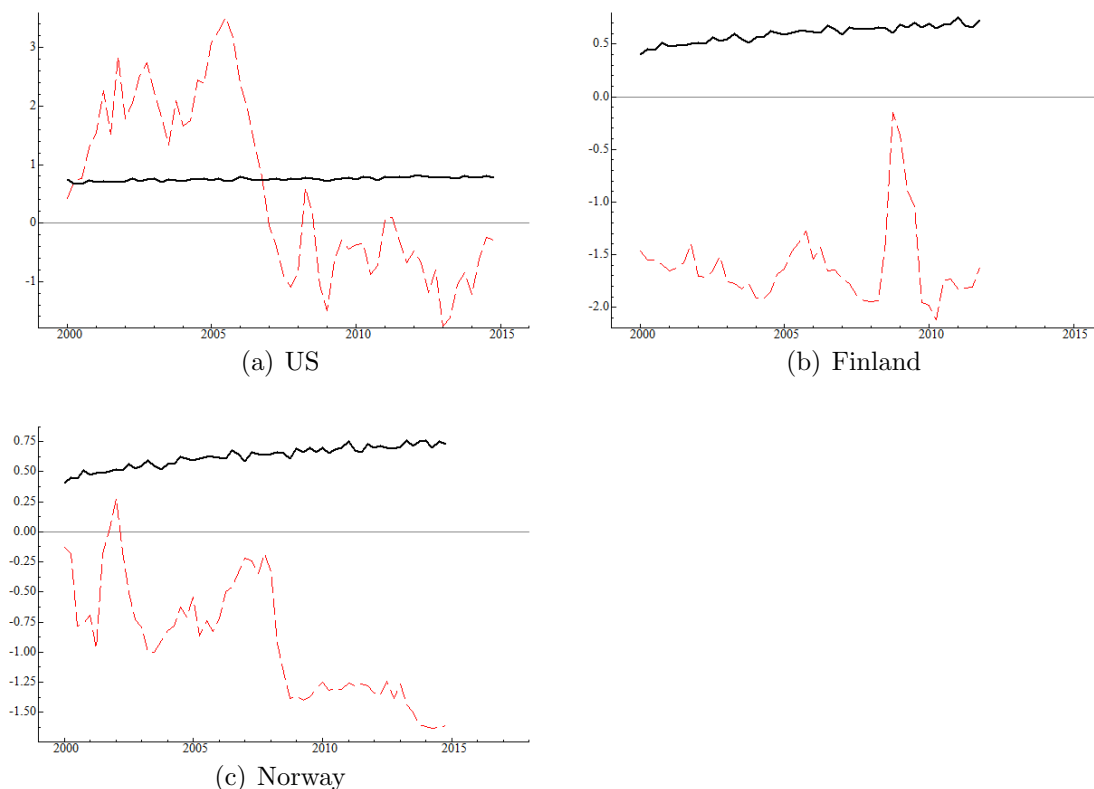


Figure 5: Panel a) Test for transition to explosivity for the US, 2000q1–2014q4. Panel b) Test for transition to explosivity for Finland, 2000q1–2011q4. Panel c) Test for transition to explosivity for Norway, 2000q1–2014q4.

## 5 Conclusions

This paper has considered four alternative indicators of housing market instability for Norway, Finland and the US. While all the indicators have their individual weaknesses, the combined evidence from the indicators may be useful to evaluate the temperature in 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 a less fortunate development 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.

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 to prevent future housing 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

<b>Series</b>	<b>Description</b>	<b>US</b>	<b>Finland</b>	<b>Norway</b>
$PH$	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
$\tau_y$	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

$H_0$	$H_A$	$\lambda_{trace}$			1%-critical value
		US	Finland	Norway	
$r = 0$	$r \geq 1$	55.32	76.97	58.82	56.83
$r \leq 1$	$r \geq 2$	29.39	36.66	33.71	36.44
$r \leq 2$	$r \geq 3$	12.87	13.85	14.65	19.53
<i>Diagnostics</i>					
	Autocorrelation	1.5814 [0.0212]	0.9197 [0.6001]	1.2431 [0.2227]	
	Normality	2.7596 [0.8384]	2.1329 [0.9071]	6.6288 [0.3565]	
	Heteroskedasticity	1.6754 [0.0000]	1.2558 [0.1129]	1.1196 [0.2766]	

*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).