Episodes of Exuberance in Housing Markets:

In Search of the Smoking Gun*

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Abstract

In this paper, we examine changes in the time series properties of three widely used housing market indicators (real house prices, price-to-income ratios, and price-to-rent ratios) for a large set of countries to detect episodes of explosive dynamics. Dating such episodes of exuberance in housing markets provides a timeline as well as empirical content to the narrative connecting housing exuberance to the global $2008-09$ recession. For our empirical analysis, we employ two recursive univariate unit root tests recently developed by Phillips et al. (2011) and Phillips et al. (2015). We also propose a novel extension of the test developed by Phillips et al. (2015) to a panel setting in order to exploit the large cross-sectional dimension of our international dataset. Statistically significant periods of exuberance are found in most countries. Moreover, we find strong evidence of the emergence of an unprecedented period of exuberance in the early 2000s that eventually collapsed around $2006-07$, preceding the $2008-09$ global recession. We examine whether macro and financial variables help to predict (in-sample) episodes of exuberance in housing markets. Long-term interest rates, credit growth and global economic conditions are found to be among the best predictors. We conclude that global factors explain (partly) the synchronization of exuberance episodes that we detect in the data in the 2000s.

JEL Classification: C22, G12, R30, R31

KEY WORDS: House Prices; 2008-09 Global Recession; Unit-root Tests; Mildly Explosive Time Series; Sup ADF Test; Generalized Sup ADF Test; Panel GSADF; Probit Model.

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1 Introduction

The latest boom and bust in international housing markets has generated a renewed interest in the dynamics of house prices. A view shared by many academics and policy makers is that the latest boom period in housing was associated with house prices departing from their intrinsic values which led to a misallocation of resources, distorted investment patterns, and eventually precipitated the $2008-09$ global recession. On this basis, in the aftermath of the crisis, international organizations and central bankers have become increasingly more concerned about developments in housing markets across the world (e.g., IMF established the Global Housing Watch and the Federal Reserve Bank of Dallas created the International House Price Database). Central banks have also become more aware of the role of domestic housing markets in financial stability and the real economy (see, e.g., the 2014 U.K. stress testing exercise of the Bank of England).

In this paper we analyze changes in the time series properties of house prices over the last four decades in an attempt to shed light on three research questions: Firstly, when did house price run-ups turn into episodes of exuberance (displaying explosive dynamics) in domestic housing markets, secondly, was there synchronization across countries that led to global exuberance and, finally, which were the fundamental factors that aided in the emergence of housing exuberance. By addressing these research questions we aim at improving our understanding of housing markets in a way that facilitates the development of better monitoring mechanisms and the design of more effective and proactive policies.

An appealing feature of our empirical analysis is the use of panel data from the International House Price Database of the Federal Reserve Bank of Dallas (Mack and Martínez-García (2011)) to detect and date-stamp explosive autoregressive behavior in housing markets. This dataset has a large panel dimension (22 countries), which allows us to have a unique, international perspective on the evolution of housing markets.[1]

With respect to our first research question, we employ two univariate econometric tests (namely, the sup ADF, $SADF$, and the Generalized sup ADF, $GSADF$) recently developed by Phillips et al. (2011) and Phillips et al. (2015) in order to examine whether real house prices and standard long-run anchors of the housing market, such as price-to-income and price-to-rent ratios, exhibit periods of mild explosive behavior.[2] The $SADF$ and $GSADF$ are right-tailed unit root tests that perform recursive ADF regressions to test for explosiveness. Their recursive nature makes the $SADF$ and $GSADF$ tests attractive in our context for two reasons. First, the tests enable us to identify in which (if any) periods a series displayed explosive behavior and, thus, date the emergence of exuberance in domestic housing markets. Second, they exhibit higher power in the presence of boom and bust episodes than standard integration/cointegration tests and tests that allow for a single (permanent) change in persistence.[3]

A timeline of events emerges from the corresponding dating procedure which suggests that the latest boom-bust cycle was unusually widespread across countries and evolved in three phases: A first phase

[1] We complement the Dallas Fed dataset with housing rents from the OECD for 16 of the 22 countries for which there is consistent data over the same sample period (Girouard et al. (2006)).

[2] Mildly explosive behavior is modeled by an autoregressive process with a root that exceeds unity, but remains within the vicinity of one. This represents a small departure from martingale behavior, but is consistent with the submartingale property often used in the rational bubbles literature (see section 2 for further details). Phillips and Magdalinos (2007a), Phillips and Magdalinos (2007b) and Phillips and Magdalinos (2012) provide a large sample asymptotic theory for this class of processes that enables econometric inference in this case, unlike for purely explosive processes.

[3] The $SADF$ and $GSADF$ tests better detect mildly explosive behavior in time series data than standard methods such as unit root/cointegration tests (e.g., Diba and Grossman (1988)), but also variance bound tests (e.g., LeRoy and Porter (1981), Shiller (1981)), specification tests (e.g., West (1987)), and Chow and CUSUM-type tests (e.g., Homm and Breitung (2012)). Few studies have implemented these techniques in the context of housing markets (e.g., Phillips and Yu (2011); Yiu et al. (2013)), but only on domestic markets.

of origination connected with the U.S. experience during the second half of the 1990s; a second phase of propagation that is characterized by widespread and synchronized episodes of exuberance across very different housing markets during the first half of the 2000s; and a final phase where the episode of exuberance burst within a short period of time and was followed by the severe contraction in economic activity of the $2008-09$ global recession. In light of the substantial differences across domestic housing markets and the non-tradability of housing, the pattern of near-simultaneous explosive behavior in the first half of the 2000s is a particularly interesting finding.

However, the $SADF$ and $GSADF$ are univariate testing procedures and, thus, only allow us to draw conclusions at the country level. To address our second research question, we propose a novel extension of the $GSADF$ procedure to a panel setting inspired by the work of Im et al. (2003) which exploits the large cross-sectional dimension of the International House Price Database in order to draw inferences for the global housing market. The extension is easy to compute and, like the univariate methods, provides a date-stamping strategy. The results for our Panel $GSADF$ test provide a very clear picture of global exuberance in housing markets in the period preceding the global $2008-09$ recession between 2000 and 2006.

Having identified periods of explosive behavior at both the domestic and the international levels, we turn to our third research question which deals with the factors that contribute to increase the likelihood of exuberance emerging in housing. A widely used explanation for the presence of boom and bust episodes is the existence of housing bubbles (see, e.g., Case and Shiller (2003); Mayer (2011); Shiller (2015)). In a rational expectations framework, rational bubbles can arise simply due to market participants' expectations of future price increases (Flood and Hodrick (1990)). This type of bubble processes follow explosive paths which makes asset prices diverge from their fundamental values and induces exuberance in housing markets. On the basis of this rationale, several authors have employed integration and cointegration tests to examine the presence of house price bubbles (see, e.g., Hott and Monnin (2008); Mikhed and Zemcik (2009b); André et al. (2014)).

As we show in the following section, non-stationary dynamics in house prices may also arise from causes other than bubbles, such as explosive dynamics in economic fundamentals and time-varying discount rates. For this reason, although we recognize that bubbles can cause explosive behavior in house prices, we refrain from using the term bubbles in the interpretation of our findings. Cappozza et al. (2004), for example, find that the time series properties of house prices in 62 United States metro areas depend on a set of economic variables that proxy for information costs, supply costs, and expectations.[4] Agnello and Schuknecht (2011) and Rousová and van den Noord (2011), among other recent studies, look at fundamental predictors of housing booms and busts as well as turning points in the housing cycle of OECD countries. According to their findings, domestic credit and interest rates are amongst the most important predictors of booms and busts, while macro aggregates (like unemployment) can be helpful in predicting peaks and troughs. In related research, Chen (2009) and Nyberg (2013) examine which macroeconomic and financial variables impact the future likelihood of a bull and bear market. By analyzing data for 18 OECD countries, they find evidence that asset price booms, characterized by rapid price appreciation above estimated trends, generally are preceded by similarly rapid expansions in monetary aggregates and inflation. They also find that financial

[4] The existing empirical evidence points out that house prices may temporarily deviate from fundamentals (e.g., the time series and cross-section evidence in Clayton (1996); Hwang and Quigley (2006); Mikhed and Zemcik (2009a); Cappozza et al. (2004); Adams and Fuss (2010)) and the importance of the bank lending channel in housing (e.g., Mian and Sufi (2009); Pavlov and Wachter (2011); Berkovec et al. (2012)).

variables are useful predictors, especially for bear markets.

We contribute to this literature by assessing whether the probability of a housing market being in a state of exuberance depends on a set of macroeconomic and financial variables that have been reported to be important drivers of house prices. Our results provide evidence in favor of in-sample predictability with the best predictors being long-run interest rates, private credit growth, demand-side factors (such as per capita real disposable income growth) and global economic conditions. These results highlight the important ripple effects of declining long-term interest rates in the synchronization of housing market exuberance and also conform with the anecdotal evidence that during the boom years prior to the $2008-09$ global recession, international investors—reaping the benefits of strong global growth—fuelled house price run-ups across many different countries.

The remainder of the paper proceeds as follows: Section 2 outlines the standard asset-pricing model of housing and illustrates how explosive behavior in house prices may arise. Section 3 describes the $SADF$ and $GSADF$ testing procedures, as well as the extension of the latter to a panel setting. The following section presents and discusses the empirical results for the the univariate and panel unit root tests and the panel probit on the likelihood of an episode of exuberance occurring. Special attention is paid to the U.S., the U.K. and Spain due their economic size and significance, and because these countries exemplify the distinct patterns observed during the three main phases of the timeline of events that we describe in the paper. Finally, Section 5 provides concluding remarks.

2 House Prices, Fundamentals and Rational Bubbles

A conventional framework for the study of explosive behavior in housing markets is provided by the standard asset-pricing model with risk neutral agents (see, e.g., Clayton (1996); Hiebert and Sydow (2011)).[5] In this framework, the price of housing can be derived from the following no-arbitrage condition,

$$\underbrace{\rho}_{\text{constant risk-free rate}} = \underbrace{\mathbb{E}_t\left(R_{t+1}\right)}_{\text{expected return on housing}}, \tag{1}$$

where $\rho > 0$ denotes the (for time being constant) discount rate—the expected net return on an alternative investment opportunity—while $\mathbb{E}_t$ is the expectations operator based on all information available up to time t, and R_{t+1} is the return on housing at time $t+1$ defined as,

$$R_{t+1} \equiv \frac{P_{t+1} + F_{t+1}}{P_t} - 1, \tag{2}$$

where P_t denotes the house price, and F_t is the stream of payoffs (pecuniary or otherwise) derived from housing.

We refer to F_t as the economic fundamentals of the housing market, and work out our analysis with two related specifications. One general specification of F_t includes the payoff stream X_t, which is given by the observed economic rents of housing, including housing services, and another component U_t that is either

[5] The asset-pricing approach to housing builds on the extensive rational bubbles literature (see, e.g., the seminal work of Blanchard (1979) and Blanchard and Watson (1982)).

unobserved or reflects mismeasurement of the housing rents, i.e.,

$$F_t = X_t + U_t. \tag{3}$$

The other specification of F_t relates the payoff stream of housing rents X_t to macroeconomic fundamentals through a demand equation for rental housing but retains an unobserved (or mismeasurement) term U_t. Under a set of constraints on preferences, we can derive a linear expenditure system where the demand for rental housing linearly relates rents X_t to macroeconomic fundamentals such as disposable income Y_t, i.e.,

$$F_t = \theta_F + \delta Y_t + U_t. \tag{4}$$

The Appendix provides details on the derivation of this relationship which captures the affordability determinants of housing.[6]

Using the definition of the return on housing R_t in (2) and re-arranging the no-arbitrage condition in (1), yields the following expression for house prices,[7]

$$P_t = \frac{1}{1+\rho}\mathbb{E}_t\left[P_{t+1} + F_{t+1}\right], \tag{5}$$

which indicates that the price today must be equal to the discounted present-value of the expected fundamentals plus the re-sale price of housing tomorrow. Recursive substitution of this asset pricing equation yields the standard present-value model for the price of housing (see, e.g., Clayton (1996)).

By recursively substituting T periods forward, equation (5) can be re-written as

$$P_t = \mathbb{E}_t\left[\sum\nolimits_{j=1}^{T}\left(\frac{1}{1+\rho}\right)^j F_{t+j}\right] + \mathbb{E}_t\left[\left(\frac{1}{1+\rho}\right)^T P_{t+T}\right]. \tag{6}$$

According to the above expression, the house price at time t is a function of the expected discounted flow of all future payoffs up to time T plus the discounted re-sale value of the house at time T. Letting T go to infinity and imposing the transversality $\lim\limits_{T\to\infty}\mathbb{E}_t\left[\left(\frac{1}{1+\rho}\right)^T P_{t+T}\right] = 0$ to rule out non-fundamental behavior (bubbles), the unique solution to the expectational difference equation in (5) yields

$$P_t^* = \mathbb{E}_t\left[\sum\nolimits_{j=1}^{\infty}\left(\frac{1}{1+\rho}\right)^j F_{t+j}\right], \tag{7}$$

where P_t^* is referred to as the fundamental price of housing due to the fact that it is a function solely of economic fundamentals F_t and the discount rate ρ.[8] In this case, the housing price corresponds to its fundamental-based price—i.e., $P_t = P_t^*$.

The prediction that house prices are driven solely by economic fundamentals depends crucially on the

[6] We implicitly use this demand equation for rental housing to relate house prices to personal disposable income. In doing so, however, the definition of fundamentals has to be augmented with a particular specification of the rental housing demand.

[7] Log-linear approximations are also commonly used but may be less relevant with nonstationary data where sample means do not converge to population constants (see, e.g., Campbell and Shiller (1989) and Chapter 7 in Campbell et al. (1997)). Further discussion on these approximations can be found in Lee and Phillips (2011). In this paper, we work with levels. Using logs does not alter qualitatively the results.

[8] For the standard dividend discount model in which the payoff stream $\{X_t\}_{t=1}^{\infty}$ grows at a constant rate see Gordon and Shapiro (1956). Blanchard and Watson (1982) and Campbell et al. (1997) examine more general processes for $\{X_t\}_{t=1}^{\infty}$.

transversality condition. In the absence of this condition, there exist infinite forward solutions to the difference equation for the price of housing which are of the form (see, e.g., Sargent (1987); Diba and Grossman (1988); LeRoy (2004))

$$P_t = P_t^* + B_t, \tag{8}$$

where P_t^* is the fundamental-based price of housing determined in (7) and B_t is a non-fundamental, bubble term that satisfies the submartingale property

$$\mathbb{E}_t\left(B_{t+1}\right) = (1+\rho)B_t. \tag{9}$$

Since the discount factor is positive ($\rho > 0$), the bubble term B_t is on expectation explosive. Making the discount factor $\rho_t > 0$ either stationary or integrated of order 1 is not going to alter the implications of the submartingale in (9).

A rational bubble occurs when expectations of future price increases rather than fundamentals drive current house prices up (see, e.g., Case and Shiller (2003)). Hence, the emergence of such a bubble creates exuberance in the housing market. Intuitively, in the presence of a bubble, buyers are willing to pay prices increasingly higher than the fundamental-based price P_t^* because they expect to be compensated through future price increases at a rate that equals the discount rate $\rho > 0$.

The fact that the bubble term is explosive has important implications for empirical studies. If economic fundamentals follow either a stationary or an integrated process of order 1, which is a typical assumption in the literature, then the only reason that house prices can display explosive dynamics is because rational bubbles exist. Given that, one can test for rational bubbles by simply applying right-tailed unit root tests to house prices in order to detect and date-stamp periods of (mildly) explosive behavior in the time series. However, the results of such tests must be interpreted with caution since they are only indicative, not conclusive. As we show next, factors other than bubbles can also give rise to explosive dynamics in house prices.

Explosive Fundamentals. Following Campbell and Shiller (1987), we combine equations (7) and (8) to obtain the following expression for house prices,

$$P_t = \frac{1}{\rho}F_t + \left(\frac{1+\rho}{\rho}\right)\mathbb{E}_t\left[\sum\nolimits_{j=1}^{\infty}\left(\frac{1}{1+\rho}\right)^j \Delta F_{t+j}\right] + B_t, \tag{10}$$

where Δ is the difference operator. A plausible assumption is that the economic rents on housing, F_t, follow a general autoregressive process of order 1,

$$F_t = \phi F_{t-1} + \epsilon_t,\ \epsilon_t \sim WN\left(0, \sigma_\epsilon^2\right), \tag{11}$$

where ϵ_t is a white noise process. The stochastic process in (11) is stationary for $|\phi| < 1$, integrated of order one for $\phi = 1$, and explosive for $\phi > 1$. In the absence of bubbles (i.e., if $B_t = 0$ for all t), equation (8) implies that the house price equates its fundamental-based price, so equations (10) − (11) imply that[9]

$$P_t = P_t^* = \left(1 + (1-\phi)\left(\frac{1+\rho}{1+\rho-\phi}\right)\right)\frac{1}{\rho}F_t. \tag{12}$$

It follows from (12) that house prices P_t can display explosive dynamics, even if there is no bubble, due to explosive dynamics in fundamentals—i.e. $\phi > 1$. In this case, exuberance in the housing market is inherited from fundamental factors which might not be directly observable.

It is worth noting that, irrespective of the value of ϕ, the ratio of house prices to fundamentals

$$\frac{P_t}{F_t} = \left(1 + (1-\phi)\left(\frac{1+\rho}{1+\rho-\phi}\right)\right)\frac{1}{\rho} \tag{13}$$

is non-explosive in the absence of bubbles. On the contrary, in the presence of bubbles, house prices increase in expectation faster than fundamentals causing their ratio to explode. This implies that right-tailed unit root tests applied to price-to-fundamental ratios are more informative about rational bubbles than tests applied to house prices alone. For this reason, we also examine the price-to-rent and the price-to-income ratios in our empirical analysis.[10]

It should be noted that working with price-to-fundamental ratios does not make the results of right-tailed unit root tests conclusive for the presence of a bubble. A reason for this is because we do not generally observe all fundamentals, so rather than applying our tests to the unobservable price-to-fundamental ratio $\frac{P_t}{F_t}$ we rely on the observable price-to-rent $\frac{P_t}{X_t}$ and price-to-income $\frac{P_t}{Y_t}$ ratios.[11] Hence, even if there is evidence of explosive behavior in such observable ratios, we cannot truly rule out the possibility that explosiveness is inherited from the unobserved component of fundamentals U_t—a deficiency that plagues virtually all empirical studies.

Time-varying Discount Rates. Another factor that can cause exuberance in the housing market is time variation in the discount rate ρ. Similarly to economic fundamentals, the trajectory of the discount rate can have an important effect on the characteristics of the fundamental-based price of housing. For simplicity, we re-consider the asset pricing model of housing in (5) but with a time-varying discount rate, i.e.,

$$P_t = \frac{1}{1+\rho_t}\mathbb{E}_t\left[P_{t+1} + F_{t+1}\right], \tag{14}$$

and set $\phi = 1$ so that the fundamentals process in (11) follows a random walk process, i.e.,

$$F_t = F_{t-1} + \epsilon_t, \ \epsilon_t \sim WN\left(0, \sigma_\epsilon^2\right). \tag{15}$$

[9] For a discussion of a more general solution with log-linear approximation methods see Engsted et al. (2012).

[10] The price-to-income ratio provides a metric of house prices relative to the ability of households to pay (see, e.g., Himmelberg et al. (2005); Girouard et al. (2006)) and, thus, it incorporates one of the key determinants of the demand for housing.

[11] We note that, apart from income and rent, there are other fundamental drivers of housing prices, such as the cost of foregone interest, the cost of property taxes and maintenance costs (see, e.g., the discussion in Himmelberg et al. (2005)). Lack of consistent and comparable data across countries for fundamental factors like this remains a limitation for applied research in housing.

We are interested in a scenario where there is a gradual and anticipated decline in interest rates over a certain period of time. This change is described by[12]

$$1+\rho_{t+s}=\begin{cases} 1+\rho', \text{ for } 0\leq s\leq k, \\ (1+\rho_{t+s+1})\,g, \text{ for } k+1\leq s<k', \\ 1+\rho, \text{ for } s\geq k', \end{cases} \tag{16}$$

where $0<k<k'<\infty$ defines the time window of decline for the discount rate, and $g\geq 1$ determines the gross rate of decline. This interest rate specification collapses to the constant discount rate case whenever $g=1$, and implies $\rho'>\rho$ whenever $g>1$. The time-variation in (16) captures the idea that declining rates may have been an important factor in the run-up of house prices leading to the $2008-09$ global recession (in line with Bernanke (2005)'s saving glut theory).

Imposing the transversality condition to rule out non-fundamental bubbles, the unique solution to the present value model for house prices is given by

$$P_t=P_t^*=\theta_{t-1}F_t, \tag{17}$$

where θ_{t-1} obeys the following difference equation,

$$(1+\rho_t)\,\theta_{t-1}=(1+\theta_t)\,. \tag{18}$$

Combining the solution in (17) with the specification of fundamentals in (15), we derive the following process for the house price,

$$P_t=\frac{\theta_{t-1}}{\theta_{t-2}}P_{t-1}+\varepsilon_t,\ \varepsilon_t\sim WN\left(0,\theta_{t-1}^2\sigma_\epsilon^2\right). \tag{19}$$

Equation (19) shows the potential impact of time variation in the discount rate on the persistence and volatility of house prices.[13]

In the constant discount rate case where $g=1$, $\theta_t=\frac{1}{\rho}$ and house prices inherit the unit root of fundamentals. Furthermore, house price volatility is that of the fundamentals scaled by a constant related to the discount rate. The solution in the general case where $g>1$ can be characterized by backward induction as follows: For $s\geq k'$, the solution corresponds to the case where the low discount rate remains constant with $\theta_{t+s}=\frac{1}{\rho}$. By taking $\theta_{t+k'}=\frac{1}{\rho}$ as given and using the specification of the discount rate in (16) and the difference equation in (18), we can recover $\theta_{t+k'-1}$. In turn, we can similarly use $\theta_{t+k'-1}$ to back out $\theta_{t+k'-2}$ and so on until we recover the entire trajectory back to time t.[14]

Figure 1 illustrates the impact of time variation in the discount rate ρ_t given in (16) on the time series

[12] The recursive representation of the discount rate is equivalent to the following alternative characterization,

$$1+\rho_{t+s}=\begin{cases} (1+\rho)\,g^{k'-k}, \text{ for } 0\leq s\leq k, \\ (1+\rho)\,g^{k'-s}, \text{ for } k+1\leq s<k', \\ 1+\rho, \text{ for } s\geq k'. \end{cases}$$

[13] For a discussion on the characteristics of the volatility process in house prices with data from the International House Price Database see Mack and Martínez-García (2012). These authors provide empirical evidence of an increase in house price volatility that is consistent with the stylized implications of declining discount rates laid out here.

[14] We can also show that the persistence term $\frac{\theta_{t-1}}{\theta_{t-2}}$ in the house price equation is bounded below by g and above by $g^{k'-k}$ over the period from t up to $t+k'$.

features of the house price series P_t in (19) with a simple numerical example.[15] Our findings suggest that declines in interest rates can rationalize (at least qualitatively) run-ups in house prices, the explosiveness in the time series and the increased volatility without having to appeal to non-fundamental explanations (rational bubbles).

[INSERT FIGURE 1]

In short, the above analysis highlights the importance of taking into consideration all three factors (rational bubbles, time variation in the discount rate, and explosive fundamentals) when explaining changes in the time series properties of house prices.

3 Testing for Episodes of Explosive Behavior

This sections describes the econometric methods that we employ to test for explosive behavior and provides technical details for their estimation. The first part of the section deals with univariate tests—the supremum ADF ($SADF$) of Phillips et al. (2011) and the generalized SADF ($GSADF$) of Phillips et al. (2012) and Phillips et al. (2015)—and the second part with our proposed extension to a multivariate setting.

3.1 The Univariate $SADF$ and $GSADF$ Procedures

Consider the following Augmented Dickey-Fuller (ADF) regression equation

$$\Delta y_t = a_{r_1,r_2} + \beta_{r_1,r_2} y_{t-1} + \sum_{j=1}^{k} \psi^j_{r_1,r_2} \Delta y_{t-j} + \epsilon_t,\ \epsilon_t \overset{iid}{\sim} N(0, \sigma^2_{r_1,r_2}), \tag{20}$$

where y_t denotes a generic time series (using the notation of the previous section, y_t can be either P_t, $\frac{P_t}{X_t}$ or $\frac{P_t}{Y_t}$), Δy_{t-j} for $j = 1, ..., k$ are the differenced lags of the time series, and ϵ_t is the error term. Moreover, r_1 and r_2 denote fractions of the total sample size that specify the starting and ending points of a subsample period, k is the maximum number of lags included in the specification, and a_{r_1,r_2}, β_{r_1,r_2} and $\psi^j_{r_1,r_2}$ with $j = 1, ..., k$ are regression coefficients.

The emergence of explosive behavior in house prices defines a period of exuberance and is indicated by a shift from a random walk—under the assumption that fundamentals are $I(1)$—to mildly explosive behavior. Therefore, we are interested in testing the null hypothesis of a unit root in y_t, $H_0 : \beta_{r_1,r_2} = 0$, against the alternative of mildly explosive behavior, $H_1 : \beta_{r_1,r_2} > 0$. Let

$$ADF^{r_2}_{r_1} = \frac{\widehat{\beta}_{r_1,r_2}}{\text{s.e. } (\widehat{\beta}_{r_1,r_2})} \tag{21}$$

denote the test statistic corresponding to this null hypothesis. Setting $r_1 = 0$ and $r_2 = 1$ yields the standard ADF test statistic, ADF^1_0. The limit distribution of ADF^1_0 is given by

$$\frac{\int_0^1 W dW}{\left(\int_0^1 W^2\right)^{\frac{1}{2}}}, \tag{22}$$

[15] For the numerical example, we set $\rho = 0.02$, $g = 1.0002774397$, $k' - k = 70$, and $\sigma^2_\epsilon = 0.01$.

where W is a Wiener process. The ADF test compares the ADF_0^1 statistic with the right-tailed critical value from its limit distribution. When the test statistic exceeds the corresponding critical value, the unit root hypothesis is rejected in favor of the alternative of explosive behavior.

Although widely employed, the standard ADF test has extremely low power in detecting episodes of explosive behavior when these episodes end with a large drop in prices, i.e., in the presence of boom-bust dynamics. For example, it is well established that nonlinear dynamics, such as those displayed by episodes of explosiveness that are periodically collapsing, frequently lead to finding spurious stationarity even though the process is inherently explosive (see, e.g., Evans (1991)).[16]

In order to deal with the effect of a collapse in house prices on the test's performance, Phillips et al. (2011) proposed a recursive procedure based on the estimation of the ADF regression in (20) on subsamples of the data. Normalizing the end of the original sample to $T = 1$, the authors propose estimating (20) using a forward expanding sample with the end of the sample period r_2 increasing from r_0 (the minimum window size for the fixed initial window) to one (the last available observation). For this procedure, the beginning of the sample is held constant at $r_1 = 0$, and the expanding window size of the regression (over the normalized sample) is denoted by $r_w = r_2 - r_1$. Then, while the starting point of the estimation is kept fixed at $r_1 = 0$, the ADF regression is recursively estimated, while incrementing the window size, $r_2 \in [r_0, 1]$, by adding one additional observation at a time. Each estimation yields an ADF statistic denoted as $ADF_0^{r_2}$.

The Phillips et al. (2011) test statistic, called sup ADF ($SADF$), is defined as the supremum value of the $ADF_0^{r_2}$ sequence expressed as follows:

$$SADF(r_0) \;=\; \sup_{r_2 \in [r_0, 1]} ADF_0^{r_2}. \tag{23}$$

Under the null hypothesis of a random walk, the limit distribution of the $SADF$ statistic is given by

$$\sup_{r_2 \in [r_0, 1]} \frac{\int_0^{r_2} W dW}{\left(\int_0^{r_2} W^2\right)^{\frac{1}{2}}}. \tag{24}$$

Similar to the standard ADF test, when the $SADF$ statistic exceeds the right-tailed critical value from its limit distribution, the unit root hypothesis is rejected in favor of explosive behavior.

The $SADF$ test performs well when there is a single boom-bust episode within the sample. Simulation experiments in Homm and Breitung (2012) show that the $SADF$ test outperforms alternative testing methods—such as the modified versions proposed by Bhargava (1986), Busetti and Taylor (2004), and Kim (2000) and Kim et al. (2002)—in the presence of a single change in the persistence from a random walk to an explosive process.[17]

More recently, Phillips et al. (2015) derive a new unit root test, the Generalized $SADF$ ($GSADF$), that covers a larger number of subsamples than the $SADF$ by allowing both the ending point, r_2, and the starting point, r_1, to change. This extra flexibility on the estimation window results in substantial power gains in

[16] Evans (1991) show using simulation methods that standard unit root and cointegration tests cannot reject the null of no explosive behavior, when such periodically collapsing episodes are present in the data. Price increases during the boom followed by a decline during the correction phase make it look like a mean-reversting (stationary) process. Intuitively, this is the reason why many non-recursive unit root tests wrongly suggest that processes that incorporate periodically collapsing boom-bust episodes are stationary—as indicated by Evans (1991).

[17] These approaches are used to test a permanent change in persistence from a random walk to an explosive process. As a consequence, they perform well only in cases where the series becomes explosive but never bursts in-sample.

comparison to the $SADF$. Furthermore, it makes the $GSADF$ test consistent with multiple boom-bust episodes within a given time series—while the $SADF$ test is consistent only with a single episode.

The $GSADF$ statistic is defined as

$$GSADF(r_0) = \sup_{r_2 \in [r_0,1], r_1 \in [0, r_2 - r_0]} ADF_{r_1}^{r_2}. \tag{25}$$

Under the null hypothesis, the limit distribution of the $GSADF$ statistic is

$$\sup_{r_2 \in [r_0,1], r_1 \in [0, r_2 - r_0]} \left\{ \frac{\frac{1}{2} r_w [W(r_2)^2 - W(r_1)^2 - r_w] - \int_{r_1}^{r_2} W(r) dr [W(r_2) - W(r_1)]}{r_w^{1/2} \left\{ r_w \int_{r_1}^{r_2} W(r)^2 dr - \left[\int_{r_1}^{r_2} W(r) dr \right]^2 \right\}^{1/2}} \right\}, \tag{26}$$

where the window size of each estimation is $r_w = r_2 - r_1$. Again, rejection of the unit root hypothesis in favor of explosive behavior requires that the test statistic exceeds the right-tailed critical value from its limit distribution given by (26).

The Date-Stamping Strategy. If the null of a unit root in y_t is rejected, then the $SADF$ and $GSADF$ procedures can be used to obtain an exact chronology of exuberance in the housing market. The identification of periods where house prices (or price-to-fundamental ratios) displayed mildly explosive behavior is particularly important since it is a necessary condition for shedding light on the factors and developments that led to the $2008 - 09$ global recession. We focus on the date-stamping strategy associated with the $GSADF$ procedure due to its good power properties and it's consistency in the presence of multiple boom-bust episodes.

Phillips et al. (2012) and Phillips et al. (2015) recommend a dating strategy under the $GSADF$ approach based on the backward sup ADF statistic, i.e.,[18]

$$BSADF_{r_2}(r_0) = \sup_{r_1 \in [0, r_2 - r_0]} SADF_{r_1}^{r_2}. \tag{27}$$

The origination date of the period of exuberance is defined as the first observation for which the $BSADF$ statistic exceeds its critical value,

$$\widehat{r}_e = \inf_{r_2 \in [r_0,1]} \left\{ r_2 : BSADF_{r_2}(r_0) > scu_{\lfloor r_2 T \rfloor}^{\alpha} \right\}, \tag{28}$$

and the termination date as the first observation after $\widehat{r}_e$ for which the $BSADF$ falls below its critical value,

$$\widehat{r}_f = \inf_{r_2 \in [\widehat{r}_e,1]} \left\{ r_2 : BSADF_{r_2}(r_0) < scu_{\lfloor r_2 T \rfloor}^{\alpha} \right\}, \tag{29}$$

where $scu_{\lfloor r_2 T \rfloor}^{\alpha}$ is the $100\,(1-\alpha)\,\%$ critical value of the sup ADF based on $\lfloor r_2 T \rfloor$ observations and α is the chosen significance level. The consistency of the above dating strategy in the presence of one or two explosive

[18] The backward sup ADF ($BSADF$) statistic relates to the $GSADF$ statistic as follows,

$$GSADF(r_0) = \sup_{r_2 \in [r_0,1]} \{BSADF_{r_2}(r_0)\}.$$

periods that are periodically collapsing is established in Phillips et al. (2015).

Technical Details. The computation of the $SADF$, $GSADF$ and $BSADF$ test statistics necessitates the selection of the minimum window size r_0 and the autoregressive lag length k. Regarding the minimum window size, this has to be large enough to allow initial estimation but it should not be too large to avoid missing short episodes of exuberance. We follow Phillips et al. (2012) and set the minimum size equal to 36 observations. This minimum window size is also close to the size suggested by the rule of thumb of Phillips et al. (2015), $r_0 = 0.1 + 1.8/\sqrt{T}$.[19]

With respect to the autoregressive lag length k, we evaluate our results primarily for two cases, k equal to 1 and 4.[20] Our findings do not appear very sensitive to the lag length specification and, for this reason, we only report results for $k = 4$. More sophisticated lag length selection procedures based on information criteria (such as the Modified Information Criteria of Ng and Perron (2001)) and sequential hypothesis testing (see, e.g., Ng and Perron (1995)) could, in principle, be applied but with a higher computational cost. Moreover, Phillips et al. (2012) show that such procedures can result in severe size distortions, and a reduction in power of both the $SADF$ and $GSADF$ tests. That is, they frequently lead to rejecting the null hypothesis when in fact the time series follows a unit root process, and not rejecting the null when the time series is explosive.

The implementation of the unit root tests also requires the limit distributions of the $SADF$, $GSADF$ and $BSADF$ test statistics. These distributions are non-standard and depend on the minimum window size. Hence, critical values have to be obtained through Monte Carlo simulations. We obtain finite sample critical values by generating 2000 replications of a driftless random walk process with $N(0,1)$ errors.[21]

Finally, the researcher may choose to neglect very short periods of exuberance by setting a minimum duration period. Phillips et al. (2015) recommend a minimum duration of $\log(T)/T$. Since all our house price time series have 154 quarterly observations, the minimum duration that we adhere to in our empirical evaluation corresponds to 5 quarters.

3.2 The Panel $GSADF$ Procedure

The International House Price Database of the Federal Reserve Bank of Dallas has a large cross-sectional dimension of 22 countries. The $SADF$ and $GSADF$ tests described in the previous sub-section, however, can only be applied country by country and, hence, cannot exploit the panel nature of our dataset. To the best of our knowledge, there is no sequential right-tailed unit root test for panel data in the literature. Inspired by the work of Im et al. (2003), we propose an extension of the $GSADF$ test procedure to heterogeneous panels.[22]

Consider the panel version of the ADF regression equation in (20),

$$\Delta y_{i,t} = a_{i,r_1,r_2} + \beta_{i,r_1,r_2} y_{i,t-1} + \sum\nolimits_{j=1}^{k} \psi^{j}_{i,r_1,r_2} \Delta y_{i,t-j} + \epsilon_{i,t}, \tag{30}$$

[19] Exploring alternative minimum window sizes can be computationally demanding since for each r_0 new critical values must be computed.

[20] The choice of a fixed lag length is appealing because it allows us to employ a recursive least squares approach which substantially reduces the computational cost of estimation.

[21] Using asymptotic critical values doesn't qualitatively change our results. Asymptotic values are provided in Phillips et al. (2015).

[22] We are grateful to an anonymous referee for motivating this extension.

where $i = 1, \ldots, N$, denotes the country index, and the remaining variables are defined as in the previous sub-section. We are interested in testing the null hypothesis of a unit root $H_0 : \beta_{i,r_1,r_2} = 0$ for all N countries against the alternative of explosive behavior in a subset of countries, $H_1 : \beta_{i,r_1,r_2} > 0$ for some i. This alternative allows for β_{i,r_1,r_2} to differ across countries and, in that sense, is more general than approaches based on the homogeneous alternative hypothesis.

We propose a panel unit root test computed by taking the average of the individual $BSADF$ statistics at each time period. To facilitate the analysis, we adjust the notation for univariate test statistics to include the country index i,

$$ADF_{i,r_1}^{r_2} = \frac{\widehat{\beta}_{i,r_1,r_2}}{\text{s.e. } (\widehat{\beta}_{i,r_1,r_2})}, \tag{31}$$

$$SADF_{i,r_1}^{r_2} = \sup_{r_2 \in [r_1, r_2]} ADF_{i,r_1}^{r_2}, \tag{32}$$

and

$$BSADF_{i,r_2}(r_0) = \sup_{r_1 \in [0, r_2 - r_0]} SADF_{i,r_1}^{r_2}. \tag{33}$$

The panel $BSADF$ can now be defined as

$$\text{Panel } BSADF_{r_2}(r_0) = \frac{1}{N} \sum\nolimits_{i=1}^{N} BSADF_{i,r_2}(r_0). \tag{34}$$

The Panel $BSADF$ statistic is particularly appealing because it measures the degree of *overall* exuberance in international housing markets. Having defined the Panel $BSADF$, the definition of the Panel $GSADF$ follows naturally. It is simply the supremum of the Panel $BSADF$, i.e.,

$$\text{Panel } GSADF\,(r_0) = \sup_{r_2 \in [r_0, 1]} \text{Panel } BSADF_{r_2}(r_0). \tag{35}$$

The results of Maddala and Wu (1999) and Chang (2004) show that the distribution of panel unit root tests based on mean unit root statistics is not invariant to cross-sectional dependence of the error terms ϵ_i. In light of the ample evidence of strong financial linkages across countries (see, e.g., Lane and Milesi-Ferretti (2003)), the assumption of uncorrelated shocks seems unrealistic even for international housing markets. In order to draw inferences in this context, we adopt a sieve bootstrap approach that is designed specifically to allow for cross-sectional error dependence. Details of the sieve bootstrap can be found in the Appendix.

Dating episodes of exuberance in the global housing market can be performed by comparing the Panel $BSADF$ with the sequence of bootstrap critical values. The origination date is set equal to the first observation that the Panel $BSADF$ statistic exceeds the $100\,(1-\alpha)\,\%$ critical value obtained from the bootstrap procedure, and the termination date is set equal to the first observation that the Panel $BSADF$ falls below the $100\,(1-\alpha)\,\%$ critical value.

4 Empirical Evidence on International House Prices

The sources and methodology used to construct the 22-country panel of house prices and personal disposable income from the International House Price Database are documented in Mack and Martínez-García (2011).

The data on real house prices and real personal disposable income per capita is reported quarterly, deflated with the PCE deflator, and covers the period from the first quarter of 1975 to the second quarter of 2013. From this data, we construct an affordability index for housing as the ratio (in percent) of real house prices over real personal disposable income per capita for each country. For 16 out of the 22 countries that we consider, a quarterly index of real housing rents is also available from the OECD (see Girouard et al. (2006)). We use this rent index to construct the price-to-rent ratio for the 16 countries with available data.

[INSERT FIGURES 2 & 3]

The sample period covered in our dataset includes several recessions, which makes it ideal for contrasting the timeline of the boom and bust in international housing markets prior to the $2008-09$ global recession against the experience during previous periods of contraction in economic activity. The median, lower and upper quartile of house prices of all 22 countries are displayed in Figure 2. We observe from the figure that for the median country, real house prices troughed in the mid-1990s and peaked around 2006. Furthermore, we can see that the run-up in real house prices during this period is widespread—an observation that has fueled the view that the latest boom-bust episode in housing had a part in the $2008-09$ global recession. This view is further supported by the fact that the time evolution of the ratios of house prices to their long-run anchors (income and rents) display similar patterns (see Figure 3). This motivates our first research question of whether there is formal statistical evidence that substantiates the claim that housing markets were exuberant prior to the global recession.

4.1 Empirical Findings: A Chronology of Exuberance in International Housing Markets

Table 1 reports results for the $SADF$ and $GSADF$ tests on real house prices, the price-to-income ratio and the price-to-rent ratio. A comparison of the results of the two methods reveals large differences. Starting with the $GSADF$ test, there is strong evidence of exuberance in real house prices with the null hypothesis of a unit root being rejected for all countries but three: Finland, Italy and South Korea. The evidence in favor of mildly explosive behavior remains strong when we look at the price-to-income ratio (the null cannot be rejected only for Finland, Italy, South Korea and Norway) and the price-to-rent ratio (the null cannot be rejected only for Germany, France and Italy). Turning to the $SADF$ test, we observe that the number of rejections of the unit root hypothesis is substantially smaller than that for the $GSADF$, which is in line with the higher power of the latter. In particular, the $SADF$ test cannot reject the null for more than half of the countries of our sample (12 out of the 22) when we examine real house prices. When we look at the price-to-income and price-to-rent ratios, the evidence in favor of exuberance becomes even weaker with fewer rejections. Overall, these results indicate that episodes of explosive behavior were widespread across a large number of countries and that powerful testing procedures are required to detect them.[23]

[INSERT TABLE 1]

[23] National house price indices aggregate across dwelling types and diverse locations within a country which may impact on the performance of the econometric tests described in the previous section. In order to examine the effect of aggregation on the properties of the $SADF$ and $GSADF$ tests, we have conducted a large simulation experiment based on the S&P/Case-Shiller 10-City Composite Home Price Index and its constituent series. The results of the simulation experiment (which are available upon request from the authors) illustrate that aggregating lowers the power of both the $SADF$ and $GSADF$ tests. The effect is much larger for the $SADF$ test than for the $GSADF$ test, which gives us another reason to prefer the latter in our econometric strategy.

We now turn to the chronology of exuberance. Figure 4 and Figure 5 show the periods during which the series under examination displayed explosive dynamics (i.e., the periods during which the estimated $BSADF$ statistics exceed the corresponding 95% critical values). Focusing on the results for real house prices, three phases can be identified in connection with the latest boom and bust episode in international housing markets.

In the first phase, between the mid-1990s and the early 2000s, real house prices became explosive in the U.S. and Ireland. There were also concurrent episodes detected for Norway and Switzerland at the time, but those seemed to have evolved and eventually collapsed on their own.

In the second phase of the boom period, i.e., by the early-to-mid 2000s, the number of countries with explosive real house prices shot up to 18 (out of a total of 22). This near-simultaneous exuberance in house prices is particularly interesting given the substantial differences across domestic housing markets and the non-tradability of housing. As evident from figures 4 and 5, this phenomenon has no precedent at least in our sample period. A potential explanation for the first two phases, is that the boom in house prices that originated in the U.S. propagated to international housing markets with perceived greater opportunities or lower risks. The pattern of propagation observed in the data and the high synchronization suggest that a common factor may have contributed to house price exuberance spreading across countries. In this regard, ripple effects from the decline in world interest rates experienced during the 2000s and housing bubbles are prime candidates (see, e.g., Case and Shiller (2003)).

In the third and final phase, the run-up in house prices started to be perceived as not sustainable, and uncertainty about real economic activity in both the U.S. and around the world grew. The episode ended with a collapse of house prices around 2006, making apparent the economic implications of the boom and bust for the U.S. and the world.

Looking at the results for price-to-fundamental ratios, we observe a similar pattern to real house prices. Perhaps not surprisingly, however, the periods of explosive dynamics in the price-to-fundamental ratios are somewhat shorter. For the U.S., in particular, we observe that the house-price-to-income ratio did not become explosive until the early 2000s. The fact that U.S. real house prices were explosive in the late 1990s but the house-price-to-income ratio was not implies that explosiveness may have been driven in part by economic fundamentals. This conclusion is in line with the strong growth (partly due to the new information technologies) in U.S. income in the 1990s.

[INSERT FIGURES 4 & 5]

There is an unusual synchronization in the episodes of explosive behavior across most of the countries in the sample since the early-to-mid-2000s. This period of near simultaneous exuberance was pervasive across very different housing markets whose fundamentals where not necessarily aligned, and it is unprecedented given the sample period and country coverage in our dataset. An overall picture of the behavior of international housing markets is given by the results for the Panel $GSADF$ test. The Panel $GSADF$ statistics are statistically significant for both real house prices and price-to-fundamental ratios which provides strong evidence in favor of global exuberance in our sample.

What is more interesting is the time evolution of the three panel $BSADF$ statistics displayed in Figure 6. Irrespective of the variable under examination (real house prices, the price-to-income ratio and the price-to-rent ratio), the results demonstrate in a clear manner the three phases of the housing market: the $BSADF$ statistics start below their critical values at the beginning of the sample, they increase rapidly after the mid-1990s and eventually exceed their critical values during the second phase starting in the early 2000s, with the period of global exuberance in housing markets continuing until $2006-07$ at which point it quickly collapsed in all three statistics (near-)simultaneously.

[INSERT FIGURE 6]

The Cases of the U.S., the U.K. and Spain. Having established the emergence of global exuberance in housing markets, we now return to the experiences of individual countries and, specifically, the U.S., the U.K. and Spain. This choice is based on the economic size and significance of these countries, and because they exemplify the distinct patterns observed during the three main phases of the timeline of events that we describe in our paper. As can be seen from Figures 7, 8 and 9, the real house price appreciation has been very significant for these three countries since the mid-1990s, with the run-up in the U.K. and Spain being larger over time than that in the U.S. This sets Spain and the U.K. apart, but as our evidence shows, it does not mean that explosive behavior is somehow a weaker phenomenon for the U.S.

[INSERT FIGURES 7 to 9]

Figures 7 to 9 show the estimated $BSADF$ statistics for the three countries together with 95% critical values. Focusing on real house prices, we observe that during the period of global exuberance leading to the $2008-09$ global recession, the U.S. played the leading role with the U.K. and Spain following.

In spite of the differences between the housing markets of Spain and the U.K., our findings show that both countries went through a near simultaneous period of exuberance during the second phase of the timeline—although the U.K. also experienced an echo before the final collapse ahead of the $2008-09$ recession. This pattern of strong synchronization in housing exuberance across countries is in line with the view that exuberance in real house prices in the U.S. migrated and amplified the effect of domestic factors in these two (otherwise very different and distant) housing markets.

4.2 The Predictive Ability of Macro and Financial Variables

Having established a timeline of exuberance in international housing markets, we employ a pooled probit model to assess the in-sample predictive ability of macroeconomic and financial variables. The probit model is described by

$$P\left(EXU_{i,t}=1\right)=\Phi\left(x_{i,t}\beta\right), \tag{36}$$

where $\Phi\left(\cdot\right)$ is the standard normal cumulative distribution function, $x_{i,t}$ is a vector of predictors for all $i=1,...,N$, and $EXU_{i,t}$ is the dependent binary variable for all $i=1,...,N$ that takes the value of 1 if there is evidence of exuberance in country i and 0 otherwise, i.e.,[24]

$$EXU_{i,t}=\begin{cases} 0, \text{ if } BSADF_{i,t}\left(r_0\right)>scu_t^{\alpha}, \\ 1, \text{ if } BSADF_{i,t}\left(r_0\right)<scu_t^{\alpha}. \end{cases} \tag{37}$$

Our choice of predictors is based partly on the theoretical model of house price determination discussed in section 2—particularly in regards to the role of time-variation in interest rates and related financial

[24] The variability of EXU_t within a country is limited because the $GSADF$ methodology does not detect many episodes of exuberance. An advantage of the pooled probit model is that, by incorporating the full variability across countries, it increases the number of episodes leading to more tightly identified results.

variables. We also include a number of macro factors which the existing literature has found to be relevant for predicting booms and busts (or turning points of the housing cycle). A detailed description of the dataset is provided by Grossman et al. (2014). This dataset is complemented with OECD and BIS data, also with national and other sources, whenever necessary.

Among the financial predictors, we consider the spread between the long and short term interest rates. The spread proxies the slope of the yield curve and indicates market expectations of future policy rates. We also include long-term interest rates to proxy for mortgage rates and the effect they may have on the likelihood of a period of exuberance. Long-term rates also provide a financial measure of the opportunity costs of investing in housing.

In addition, using OECD data, we consider quarter-over-quarter changes in stock markets in part because we expect stock prices to be forward-looking and, therefore, to reflect the profitability of alternative asset classes. Furthermore, stock market appreciation is related to changes in households' financial wealth.

We incorporate measures of domestic nominal credit growth to the private sector from the BIS in quarter-over-quarter growth rates as well as the current account to GDP ratio and the quarter-over-quarter changes in the current account. An expansion of private credit can lead to asset price and housing booms and capital inflows from abroad can fuel a rapid expansion of domestic credit. Hence, we aim to determine whether periods of exuberance can be predicted not just from the extent of the preceding credit expansion, but also from whether the funding arises externally.

We also examine quarter-over-quarter changes in real personal disposable income per capita and, based on OECD data, of real housing rents too. Housing rents proxy for time-variation in the expected returns on housing—the relationship between house prices and rents is at the core of our asset-pricing model of housing determination. Regarding real personal disposable income, this variable is meant to capture key demand-side fundamentals that are conventionally viewed as anchoring the housing market over the long run.

As indicators of the state of the business cycle, we use the unemployment rate from national sources, real GDP growth (quarter-over-quarter) and CPI inflation (quarter-over-quarter). These variables are not generally viewed as leading indicators, but they can have predictive power on future consumption and investment on housing. Moreover, we use the updated quarterly average of the global indicator of real economic activity proposed by Kilian (2009) and oil prices on the West Texas Intermediate from the Wall Street Journal to account for global economic developments.

Due to the fact that we do not have data on rent and private credit growth for all countries, our final dataset consists of a subset of 14 countries that have experienced at least one period of exuberance in-sample and for which we have all available data: Australia, Belgium, Canada, Switzerland, Germany, Denmark, Spain, Finland, France, the U.K., Ireland, Japan, Netherlands, and the U.S. To specify the pooled probit model, we start with a general specification including all our explanatory variables. This model is sequentially reduced by deleting after each iteration the insignificant variable with the highest p-value until all remaining variables appear statistically significant.

[INSERT TABLE 2]

The estimation results for the final specification of the pooled probit model with and without fixed effects can be found in Table 2. Overall, there is evidence that financial and macro variables, at least in-sample, have predictive power. Our findings suggest that the factors that matter for predicting an episode of exuberance are: First, mortgage costs and access to mortgage credit, which are proxied by long-term rates and credit growth to the private sector. Second, demand-side factors of the housing market and, in particular, per capita real disposable income growth and the unemployment rate. Third, domestic and global macroeconomic conditions, where domestic conditions are proxied by domestic real GDP growth, while global economic conditions are proxied with the index of global economic activity of Kilian (2009).

These results support the important role of declining long-term interest rates in the synchronization of domestic housing markets, in line with the predictions of our stylized asset-pricing model. Moreover, they highlight that global economic conditions contribute positively to the likelihood of exuberance—a finding that seems to conform with the anecdotal evidence that during the boom years prior to the $2008-09$ global recession, international investors—reaping the benefits of strong global growth—turned their attention increasingly towards housing markets for investment, fueling housing price run-ups across many different countries.

The last row of Table 2 reports McFadden's pseudo R^2 for the models with and without fixed effects. The estimates for the pseudo R^2 are 0.229 and 0.299, respectively, which suggests that although the macro and financial variables examined have strong predictive content, there is ample room for other explanatory factors. According to the theoretical analysis of section 2, unobserved fundamentals and rational bubbles may well be such factors.

5 Concluding Remarks

In this paper, we employed two econometric procedures, namely the $SADF$ and $GSADF$, developed recently by Phillips et al. (2011) and Phillips et al. (2015) in order to examine whether house prices and price-to-fundamental ratios exhibited explosive behavior during the last four decades and, if so, to identify the exact periods of exuberance. An appealing feature of our study is the use of data from the International House Price Database of the Federal Reserve Bank, which has a large cross-sectional dimension and, therefore, allows a unique, international look at the behavior of housing markets.

A consistent timeline of events emerges from our empirical evidence suggesting that the latest boom-bust cycle in international housing markets was unusually widespread and evolved in three phases: One of origination that can be related to the U.S. experience primarily during the second half of the 1990s; a second phase of propagation that is characterized by widespread and synchronized episodes of exuberance across very different housing markets during the first half of the 2000s; and a final phase where this episode of global exuberance burst within a short period of time from the severe contraction in economic activity of the $2008-09$ global recession.

In order to exploit the cross-sectional dimension of our dataset, we also propose an extension of the $GSADF$ test to a panel setting. This extension is appealing because it allows to draw general conclusions about international housing markets. Moreover, it is straightforward to implement and, like the $SADF$ and $GSADF$, allows dating periods of exuberance. The results of the Panel $GSADF$ test provide further support to the view that exuberance in housing markets prior to the $2008-09$ global recession was an international phenomenon.

Finally, we examined whether macroeconomic and financial variables impact the likelihood of exuberance in housing markets. In particular, we employed a panel probit model with a large number of predictors that have been suggested in the literature to be important determinants of housing cycles. Our results suggest that long-run interest rates, credit growth, demand-side factors (such as per capita real disposable income growth), as well as domestic and global macroeconomic conditions impact the likelihood of a housing market being in a state of exuberance. These findings provide (at least partially) an explanation for the widespread exuberance detected in the early- and mid-2000s.

Appendix

A Demand Equation for Rental Housing

Consider the maximization of the Stone-Geary utility function with housing units rented, H_t, and consumption of other goods, C_t, i.e.,[25]

$$U\left(H_t, C_t\right) = \left(H_t - \theta_H\right)^{\alpha} \left(C_t - \theta_C\right)^{1-\alpha}, \ 0 < \alpha < 1,$$

subject to the intratemporal budget constraint,

$$C_t + x_t H_t = Y_t,$$

where the price of the consumption good is normalized to one. $X_t \equiv x_t H_t$ is the housing rents—rental expenditures—paid and x_t the rental rate per unit rented, Y_t refers to disposable income, and $0 < \alpha < 1$, θ_H and θ_C are preference parameters.

From first-order conditions, the Stone-Geary utility function subject to the standard intratemporal budget constraint gives a linear expenditure system where the demand for rental housing takes the following form:

$$H_t = \theta_H + \frac{\alpha}{x_t}\left(Y_t - x_t\theta_H - \theta_C\right), \tag{38}$$

or in expenditure terms,

$$X_t \equiv x_t H_t = \alpha Y_t + \left(1-\alpha\right)\theta_H x_t - \alpha\theta_C. \tag{39}$$

Under the assumption that in equilibrium the units rented are constant (i.e., $H_t = H$) and normalized to one, the demand equation that determines housing rents in (39) reduces to an affine transformation of disposable income Y_t, i.e.,

$$X_t = x_t = \theta_F + \delta Y_{t+1}, \tag{40}$$

where $\delta \equiv \frac{\alpha}{1-(1-\alpha)\theta_H}$ and $\theta_F \equiv -\frac{\alpha}{1-(1-\alpha)\theta_H}\theta_C$.

[25] While the Stone-Geary reduces to the Cobb-Douglas utility function whenever the parameters θ_H and θ_C are both set equal to zero, the specification permits both the rental rate elasticity and the income elasticity to vary with both rental rates and income—unlike the Cobb-Douglas where both elasticities are constant or the constant elasticity of substitution utility function for which the income elasticity is constant.

B The Panel $GSADF$ Test

The bootstrap procedure consists of the following steps:

1. For each country, impose the null hypothesis of a unit root and fit the restricted ADF regression equation,
$$\Delta y_{i,t} = a_{i,r_1,r_2} + \sum\nolimits_{j=1}^{k} \psi^{j}_{i,r_1,r_2} \Delta y_{i,t-j} + \epsilon_{i,t},$$
to obtain coefficient estimates ($\widehat{a}_{i,r_1,r_2}$, and $\widehat{\psi}^{j}_{i,r_1,r_2}$ for $j = 1, \ldots, k$) and residuals ($\widehat{\epsilon}_i$).

2. Create a residual matrix with typical element $\widehat{\epsilon}_{t,i}$.

3. In order to preserve the covariance structure of the error term, generate bootstrap residuals $\epsilon^{b}_{i,t}$ by sampling with replacement raws from the residuals matrix.

4. Use the bootstrap residuals and the estimated coefficients to generate recursively bootstrap samples for first differences,
$$\Delta y^{b}_{i,t} = \widehat{a}_{i,r_1,r_2} + \sum\nolimits_{j=1}^{k} \widehat{\psi}^{j}_{i,r_1,r_2} \Delta y^{b}_{i,t-j} + \epsilon^{b}_{i,t},$$
and for levels,
$$y^{b}_{i,t} = \sum\nolimits_{p=1}^{t} \Delta y^{b}_{i,p}.$$

5. Compute the sequence of Panel $BSADF$ statistics and the Panel $GSADF$ statistic for $y^{b}_{i,t}$.

6. Repeat steps (3) to (5) a large number of times to obtain the empirical distribution of the test statistics under the null of a unit root.

References

Adams, Z. and R. Fuss (2010). Macroeconomic Determinants of International Housing Markets. *Journal of Housing Economics 19*, 38–50.

Agnello, L. and L. Schuknecht (2011). Booms and Busts in Housing Markets: Determinants and Implications. *Journal of Housing Economics 20*, 171–190.

André, C., L. A. Gil-Alana, and R. Gupta (2014). Testing for Persistence in Housing Price-to-Income and Price-to-Rent Ratios in 16 OECD Countries. *Applied Economics 46*(18), 2127–2138.

Berkovec, J., Y. Chang, and D. A. McManus (2012). Alternative Lending Channels and the Crisis in U.S. Housing Markets. *Real Estate Economics 40*(s1), S8ŰS31.

Bernanke, B. S. (2005). The Global Saving Glut and the U.S. Current Account Deficit. At the Sandridge Lecture, Virginia Association of Economists, Richmond, Virginia. http://www.federalreserve.gov/boarddocs/speeches/2005/200503102/.

Bhargava, A. (1986). On the Theory of Testing for Unit Roots in Observed Time Series. *Review of Economic Studies 53*, 369–384.

Blanchard, O. and M. Watson (1982). *Crisis in the Economics Financial Structure*, Chapter Bubbles, Rational Expectations, and Financial Markets, pp. 295–315. Lexington Books, Lexington, Mass.

Blanchard, O. J. (1979). Speculative Bubbles, Crashes and Rational Expectations. *Economics Letters 3*, 387–389.

Busetti, F. and A. Taylor (2004). Tests of Stationarity Against a Change in Persistence. *Journal of Econometrics 123*, 33–66.

Campbell, J., A. Lo, and A. MacKinlay (1997). *The Econometrics of Financial Markets*. Princeton, NJ: Princeton University Press.

Campbell, J. and R. Shiller (1987). Cointegration and Tests of Present Value Models. *Journal of Political Economy 95*, 1062–1088.

Campbell, J. and R. Shiller (1989). The Dividend-Price Ratio and Expectations of Future Dividends and Discount Factors. *Review of Financial Studies 1*, 195–228.

Cappozza, D. R., P. H. Hendershott, and C. Mack (2004). An Anatomy of Price Dynamics in Illiquid Markets: Analysis and Evidence from Local Housing Markets. *Real Estate Economics 32*(1), 1–32.

Case, K. E. and R. J. Shiller (2003). Is There a Bubble in the Housing Market? *Brookings Papers on Economic Activity 34*(2), 299–362.

Chang, Y. (2004). Bootstrap Unit Root Tests in Panels with Cross-sectional Dependency. *Journal of Econometrics 120*(2), 263–293.

Chen, S.-S. (2009). Predicting the Bear Stock Market: Macroeconomic Variables as Leading Indicators. *Journal of Banking and Finance 33*(2).

Clayton, J. (1996). Rational Expectations, Market Fundamentals and Housing Price Volatility. *Real Estate Economics 24*, 441–470.

Diba, B. and H. Grossman (1988). Explosive Rational Bubbles in Stock Prices. *American Economic Review 78*, 520–530.

Engsted, T., T. Pedersen, and C. Tanggaard (2012). The Log-Linear Return Approximation, Bubbles, and Predictability. *Journal of Financial and Quantitative Analysis 47*, 643–665.

Evans, G. W. (1991). Pitfalls in Testing for Explosive Bubbles in Asset Prices. *American Economic Review 81*(4), 922–930.

Flood, R. and R. Hodrick (1990). On Testing for Speculative Bubbles. *Journal of Economic Perspectives 4*, 85–101.

Girouard, N., M. Kennedy, P. van den Noord, and C. André (2006). Recent House Price Developments: The Role of Fundamentals. *OECD Economics Department Working Papers no. 475*(OECD Publishing).

Girouard, N., M. Kennedy, P. van den Noord, and A. C. Christophe (2006). Recent House Price Developments: The Role of Fundamentals. *OECD Economics Department Working Papers No. 475*. OECD Publishing.

Gordon, M. and E. Shapiro (1956). Capital Equipment Analysis: The Required Rate of Profit. *Management Science 3*(1), 102–110.

Grossman, V., A. Mack, and E. Martínez-García (2014). A New Database of Global Economic Indicators. *The Journal of Economic and Social Measurement 39*(3), 163–197.

Hiebert, P. and M. Sydow (2011). What Drives Returns to Euro Area Housing? Evidence from a Dynamic Dividend-Discount Model. *Journal of Urban Economics 70*(2-3), 88–98.

Himmelberg, C., C. Mayer, and T. Sinai (2005). Assessing High House Prices: Bubbles, Fundamentals and Misperceptions. *Journal of Economic Perspectives 19*(4), 67–92.

Homm, U. and J. Breitung (2012, Winter). Testing for Speculative Bubbles in Stock Markets: A Comparison of Alternative Methods. *Journal of Financial Econometrics 10*(1), 198–231.

Hott, C. and P. Monnin (2008). Fundamental Real Estate Prices: An Empirical Estimation with International Data. *Journal of Real Estate Finance and Economics 36*, 427–450.

Hwang, M. and J. M. Quigley (2006). Economic Fundamentals in Local Housing Markets: Evidence from U.S. Metropolitan Regions. *Journal of Regional Science 46*, 425–453.

Im, K. S., M. H. Pesaran, and Y. Shin (2003). Testing for Unit Roots in Heterogeneous Panels. *Journal of Econometrics 115*(1), 53–74.

Kilian, L. (2009). Not All Oil Price Shocks Are Alike: Disentangling Demand and Supply Shocks in the Crude Oil Market. *American Economic Review 99*(3), 1053–1069.

Kim, J.-Y. (2000). Detection of Change in Persistence of a Linear Time Series. *Journal of Econometrics 95*, 97–116.

Kim, J.-Y., J. Belaire-Franch, and R. Amador (2002). Corrigendum to 'Detection of Change in Persistence of a Linear Time Series'. *Journal of Econometrics 109*, 389–392.

Lane, P. R. and G. Milesi-Ferretti (2003). International Financial Integration. Working paper 03/86, IMF.

Lee, J. H. and P. C. Phillips (2011). Asset Pricing with Financial Bubble Risk. *Working Paper, Yale University*.

LeRoy, S. (2004). Rational Exuberance. *Journal of Economic Literature XLII*, 783–804.

LeRoy, S. and R. Porter (1981). The Present-Value Relation: Tests Based on Implied Variance Bounds. *Econometrica 49*, 555–577.

Mack, A. and E. Martínez-García (2011). A Cross-Country Quarterly Database of Real House Prices: A Methodological Note. *Globalization and Monetary Policy Institute Working Paper no. 99*. December, Federal Reserve Bank of Dallas.

Mack, A. and E. Martínez-García (2012). Increased Real House Price Volatility Signaled Break from the Great Moderation. *Federal Reserve Bank of Dallas Economic Letter 7*(1).

Maddala, G. and S. Wu (1999). A Comparative Study of Unit Root Tests with Panel Data and a New Simple Test. *Oxford Bulletin of Economics and Statistics 61*(S1), 631–652.

Mayer, C. (2011). Housing Bubbles: A Survey. *Annual Review of Economics 3*, 559–577.

Mian, A. and A. Sufi (2009). The Consequences of Mortgage Credit Expansion: Evidence from the U.S. Mortgage Default Crisis. *The Quarterly Journal of Economics 124*(4), 1449–1496.

Mikhed, V. and P. Zemcik (2009a). Do House Prices Reflect Fundamentals? Aggregate and Panel Data Evidence. *Journal of Housing Economics 18*, 140–149.

Mikhed, V. and P. Zemcik (2009b). Testing for Bubbles in Housing Markets: A Panel Data Approach. *Journal of Real Estate Finance and Economics 38*, 366–386.

Ng, S. and P. Perron (1995). Unit Root Tests in ARMA Models with Data-Dependent Methods for the Selection of the Truncation Lag. *Journal of the American Statistical Association 90*, 268–281.

Ng, S. and P. Perron (2001). Lag Length Selection and the Construction of Unit Root Tests with Good Size and Power. *Econometrica 69*, 1519–1554.

Nyberg, H. (2013). Predicting Bear and Bull Stock Markets with Dynamic Binary Time Series Models. *Journal of Banking and Finance 37*(9), 3351–3363.

Pavlov, A. and S. Wachter (2011). Subprime Lending and Real Estate Prices. *Real Estate Economics 39*(1), 1–17.

Phillips, P. C. and T. Magdalinos (2007a). Limit Theory for Moderate Deviations From a Unit Root. *Journal of Econometrics 136*, 115–130.

Phillips, P. C. and T. Magdalinos (2007b). *The Refinement of Econometric Estimation and Test Procedures: Finite Sample and Asymptotic Analysis*, Chapter Limit Theory for Moderate Deviations from a Unit Root under Weak Dependence, pp. 123–162. Cambridge University Press, Cambridge.

Phillips, P. C. and T. Magdalinos (2012). Mildly Explosive Autoregression under Weak and Strong Dependence. *Journal of Econometrics 169*(2), 179–187.

Phillips, P. C., S.-P. Shi, and J. Yu (2012). Testing for Multiple Bubbles. *Cowles Foundation Discussion Papers 1843*. Cowles Foundation for Research in Economics, Yale University.

Phillips, P. C., S.-P. Shi, and J. Yu (2015). Testing for Multiple Bubbles: Historical Episodes of Exuberance and Collapse in the S&P 500. *International Economic Review forthcoming*.

Phillips, P. C., Y. Wu, and J. Yu (2011). Explosive Behavior in the 1990's Nasdaq: When Did Exuberance Escalate Asset Values? *International Economic Review 52*(1), 201–226.

Phillips, P. C. and J. Yu (2011). Dating the Timeline of Financial Bubbles During the Subprime Crisis. *Quantitative Economics 2*(3), 455–491.

Rousová, L. and P. van den Noord (2011). Predicting Peaks and Troughs in Real House Prices. *OECD Economics Department Working Papers No. 882*.

Sargent, T. (1987). *Macroeconomic Theory*. Boston: Academic Press.

Shiller, R. J. (1981). Do Stock Prices Move Too Much to be Justified by Subsequent Changes in Dividends? *American Economic Review 71*, 421–436.

Shiller, R. J. (2015). *Irrational Exuberance. Revised and Expanded Third Edition*. Princeton University Press.

West, K. D. (1987). A Specification Test for Speculative Bubbles. *Quarterly Journal of Economics 102*(3), 553–580.

Yiu, M. S., J. Yu, and L. Jin (2013). Detecting Bubbles in Hong Kong Residential Property Market. *Journal of Asian Economics 28*, 115–124.

C Tables and Figures

Table 1: Results for the Univariate SADF and GSADF Tests

Panel A: Test Statistics

	Real House Prices		**Price-to-Income Ratio**		**Price-to-Rent Ratio**	
Country	*SADF*	*GSADF*	*SADF*	*GSADF*	*SADF*	*GSADF*
Australia	2.23***	6.18***	1.08*	2.57***	2.21***	6.40***
Belgium	0.97	2.98***	−0.25	2.92***	0.03	3.41***
Canada	0.32	3.76***	−1.13	2.16**	0.11	3.78***
Switzerland	1.64**	2.3**	1.20*	2.08**	0.47	1.90**
Germany	−0.59	2.10**	0.57	2.55***	0.51	0.85
Denmark	1.31**	2.83***	−0.03	1.76*	0.37	2.98***
Spain	0.39	3.34***	0.01	1.84**	−0.16	2.20**
Finland	0.94	1.45	−0.89	0.96	1.70**	1.70*
France	1.35**	2.21**	−0.03	2.46***	−0.36	1.48
United Kingdom	1.83**	3.34***	1.50**	2.65***	−0.26	3.02***
Ireland	2.59***	3.71***	2.01***	2.19**	2.33***	4.54***
Italy	−1.28	−0.38	−1.52	0.85	−1.93	−0.26
Japan	1.66**	3.76***	0.88	4.63***	1.44**	2.24**
South Korea	−1.11	−0.32	0.49	0.49	–	–
Luxembourg	1.65**	3.89***	−0.27	1.59*	–	–
Netherlands	−0.43	4.13***	−0.17	3.13***	−1.49	3.55***
Norway	0.85	1.75*	0.22	0.31	–	–
New Zealand	1.77**	2.35**	0.43	3.10***	2.82***	3.89***
Sweden	0.18	3.79***	0.23	3.34***	–	–
United States	1.52**	3.81***	−0.78	3.47***	−0.55	3.64***
South Africa	−0.92	3.93***	−1.35	3.44***	–	–
Croatia	0.03	1.64*	0.87	2.23**	–	–
Panel B: Critical Values						
90%	0.98	1.54	0.98	1.54	0.98	1.54
95%	1.25	1.80	1.25	1.80	1.25	1.80
99%	1.89	2.39	1.89	2.39	1.89	2.39

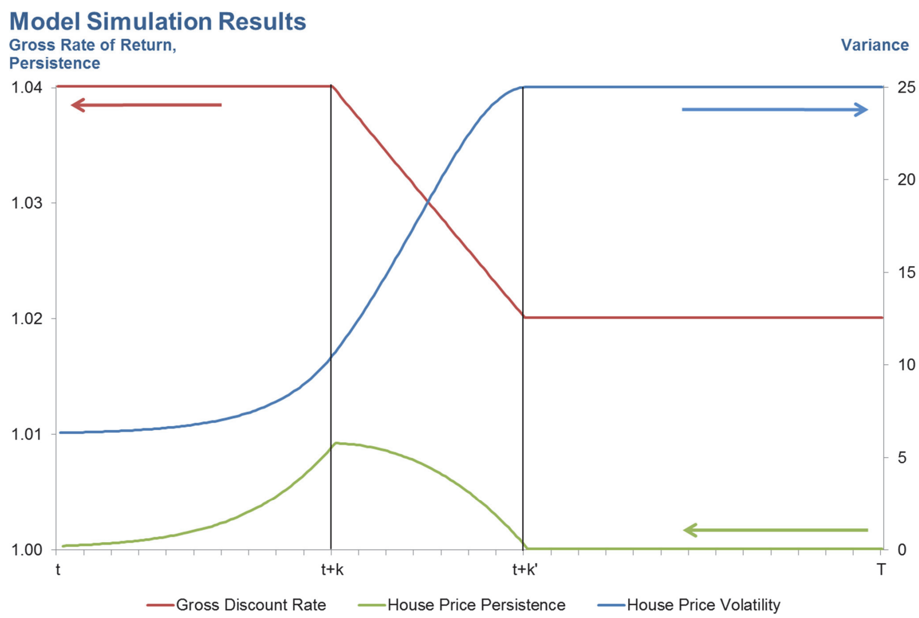
Note: *, ** and *** denote statistical significance at the 10, 5 and 1 percent significance level respectively. All results are for autoregressive lag length k=4.

Table 2: Estimation Results for the Pooled Probit Model

Long-term Interest Rates	-0.166^{***}	-0.228^{***}
Private Credit Growth	0.222^{***}	0.197^{***}
Real Personal Disposable Income Per Capita Growth	0.209^{***}	0.227^{***}
Unemployment Rate	-0.072^{***}	-0.150^{***}
Real GDP Growth	0.170^{***}	0.182^{***}
Global Economic Conditions	0.014^{***}	0.013^{***}
Fixed Effects	No	Yes
McFadden's R^2	0.229	0.299

Note: *, ** and *** denote statistical significance at the 10, 5 and 1 percent significance level, respectively.

Figure 1: Real House Prices: The Dynamic Effects of an Anticipated Decline in the Discount Rate



SOURCE: authors' calculations.

Figure 2: Real House Prices: Cross-Country Characteristics

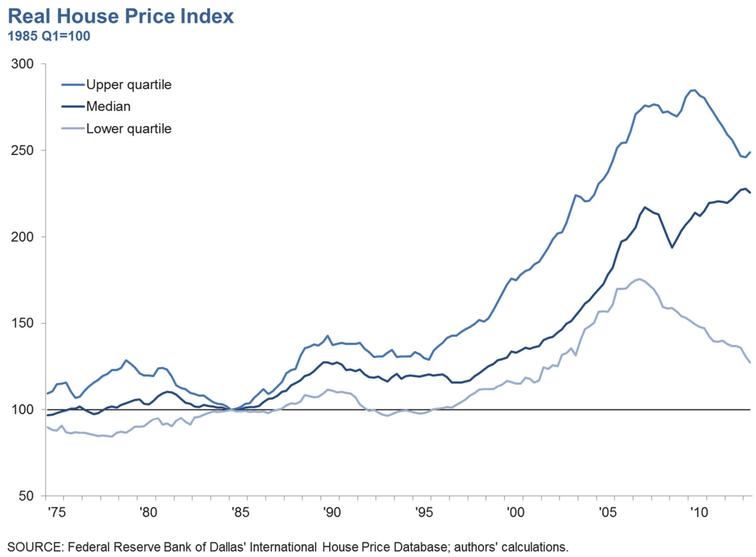


Figure 3: Price-to-Income Ratio and Price-to-Rent Ratio: Cross-Country Characteristics

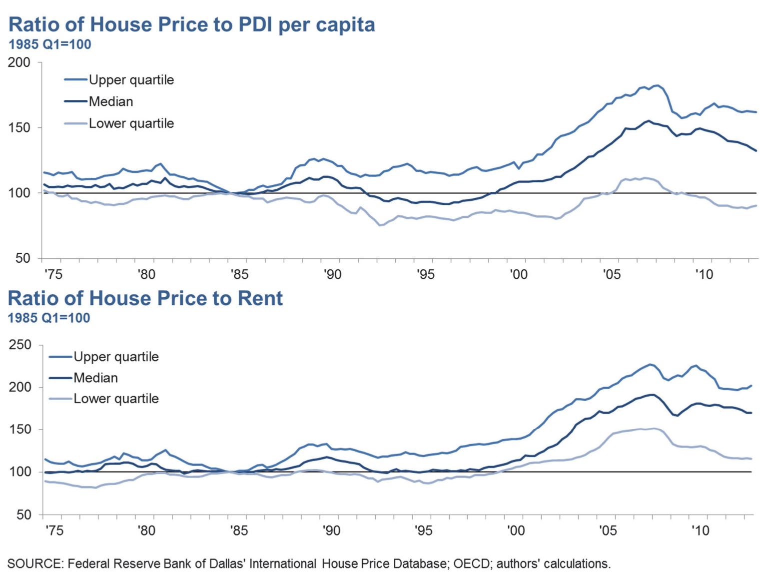
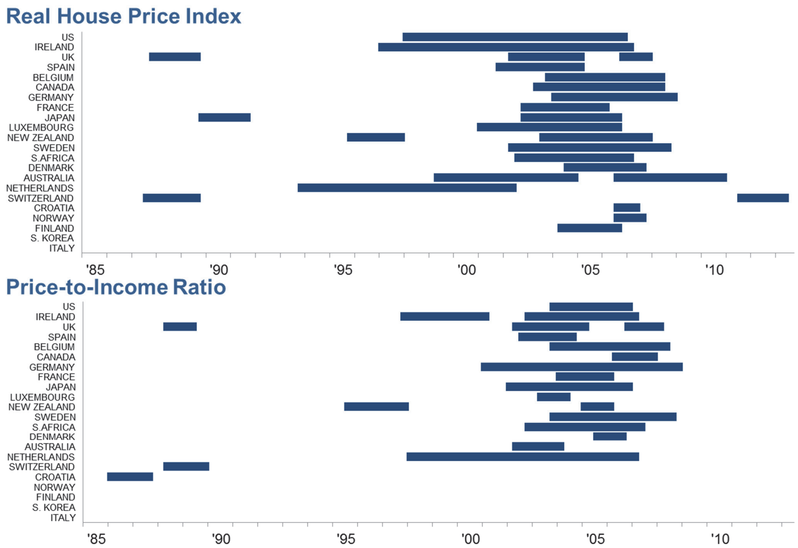


Figure 4: Date-Stamping with Real House Prices and the Price-to-Income Ratios

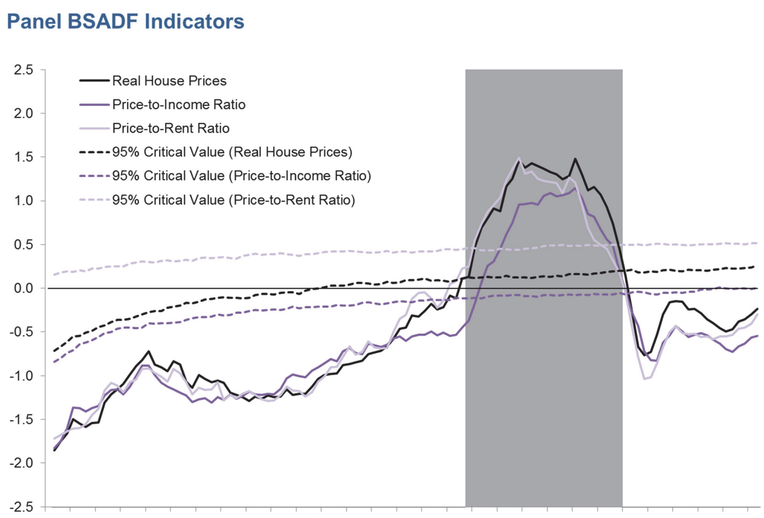


NOTE: Shaded areas indicate periods of exuberance determined by the GSADF test.
SOURCE: Federal Reserve Bank of Dallas international house price database; authors' calculations.

Figure 5: Date-Stamping with Real House Prices and the Price-to-Rent Ratios

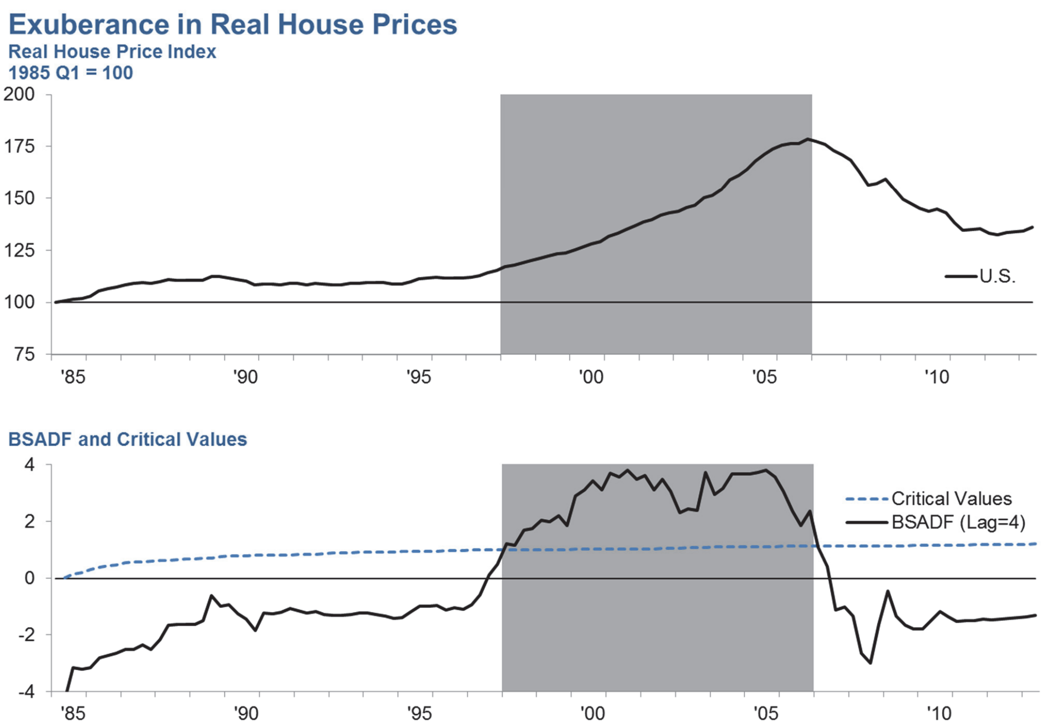
NOTE: Shaded areas indicate periods of exuberance determined by the GSADF test.
SOURCE: Federal Reserve Bank of Dallas international house price database; OECD; authors' calculations.

Figure 6: Date-Stamping Global Periods of Exuberance



NOTE: Shaded area indicates period when real house prices surpass critical value.
SOURCE: Federal Reserve Bank of Dallas' International House Price Database; OECD; authors' calculations.

Figure 7: Date-Stamping with U.S. Real House Prices



NOTE: Shaded area indicates period when real house prices surpass critical value.
SOURCE: Federal Reserve Bank of Dallas' International House Price Database; authors' calculations.

Figure 8: Date-Stamping with U.K. Real House Prices

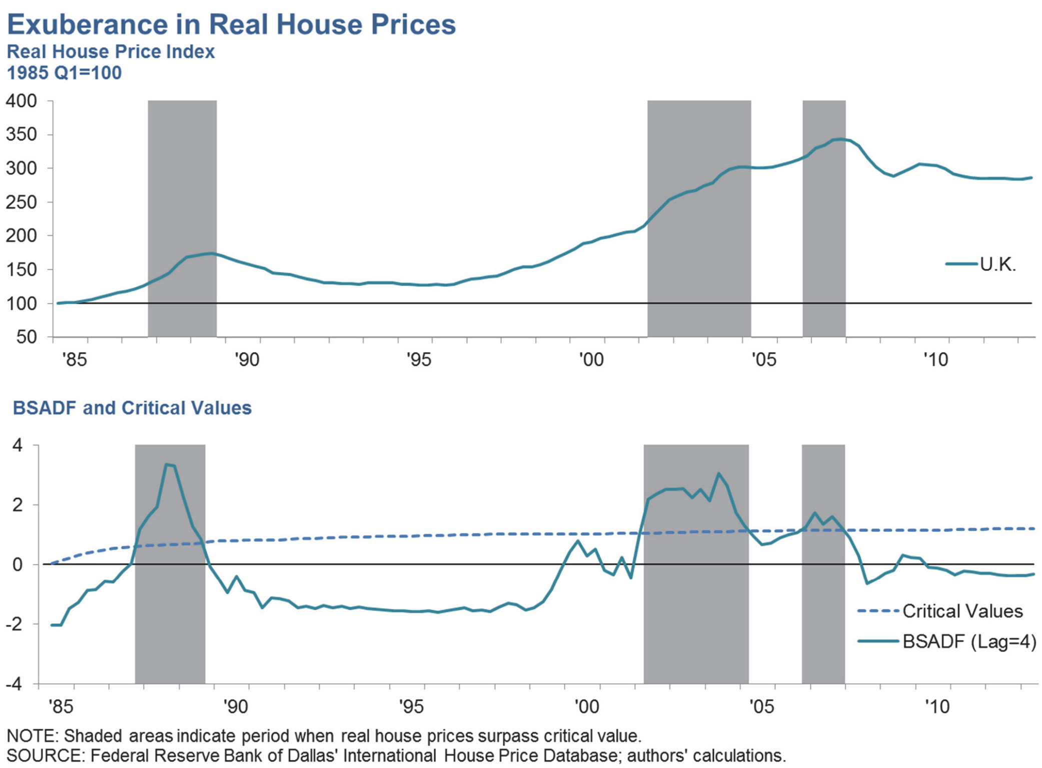


Figure 9: Date-Stamping with Spain Real House Prices