Home Prices and Fundamentals: Solving the Mystery for the G-7 by Accounting for Nonlinearities

Abstract

Home prices and their fundamental determinants, such as income, should, according to theory and intuition have a stable long run relationship. While bubbles exist, these are deviations from this long run relationship. However, numerous empirical studies have failed to find such long run stationarity between home values and income. This presents a puzzle-could house prices really drift away from income *indefinitely,* with no tendency to return to some equilibrium?

Previous studies have imposed linear adjustment to deviations from equilibrium in their tests. However, it has been clearly established that both home prices and income exhibit nonlinear dynamics. Moreover, tests for stationarity have low power in the presence of nonlinearity. We accordingly address this issue for the G-7 countries. First, we have a longer span of data than previous studies. In addition we employ tests which detect and allow for nonlinear adjustment to shocks. For the US, we are able to reject the null hypothesis of nonstationarity in the price/income ratio with this longer span of data and the use of a linear, but powerful testing procedure heretofore not utilized in this literature. However, we do not find a stable relationship using this procedure for the remaining six G-7 nations. We do find evidence, for these six countries, of nonlinearity, and upon applying a test which posits asymmetric adjustment we obtain results which indicate stationary, long run relationships between values and income in five of the remaining six G-7 countries.

Keywords: House Prices, Asset Prices, Bubbles, G-7
JEL Codes: C22 Times Series Models, G12 Asset Pricing, R31 Housing Supply and Markets

Introduction

The existence of bubbles-periods in which prices exceed levels justified by fundamentals-has plagued housing markets in numerous countries, with negative consequences for capital allocation and, subsequent to the bubbles bursting, investor returns and the macroeconomy. However, while bubbles can likely develop, theory and intuition hold that over time, prices and fundamentals such as income or rents share a long run relationship. It would be contrary to basic asset pricing models to have asset prices become permanently unmoored from their basic determinants. For housing it would seem very odd to have home values become larger and larger relative to income, without any tendency, even over decades, to return to some sustainable level.

However, despite the theory and intuition, several studies on the topic have failed to find evidence that home prices have a long run relationship with income or rents. Of course, finding that prices and fundamentals can deviate from long run relationships, in say a bubble period, presents no puzzle. But failing to find a long run, stationary relationship between the two over a number of decades does present a puzzle.

Previous papers on the topic have either examined the ratio of home prices to income (or in one case rents) and tested it for stationarity or tested the two variables for cointegration (Malpezzi, 1999, Meen, 2002, Gallin, 2006, Mikhed and Zemcik, 2009, Holly Pesaran and Yamagata, 2010). These papers have failed to find a reliable stationary relationship for the variables.

A possible reason for the failure to reject the null hypothesis of no long run relationship between home values and fundamentals may be the lack of power-a low probability of rejecting a false null- in standard unit root and cointegration tests. Gallin (2006) points out that the span of data, which is usually two to three decades in most previous studies, may be too short to detect a long run relationship (the power of these techniques obviously increases when more data-especially a longer span of data which can capture more housing cycles-is employed). Some papers attempt to cope with the lack of power by using panel unit root or cointegration tests, in which home prices and income are tested for a stable relationship in a number of regions simultaneously. However, even the use of more powerful panel methods has either still failed to find stationarity or found that only some, but clearly not all of the regions exhibit a stable long run relationship.

In this paper we take a different approach. All previous studies have imposed linear adjustment on the price/fundamentals relationship. That is, a positive shock to the relationship is restricted to decay at the same rate as a negative shock. It has been established, however that home values exhibit nonlinearities (Miles, 2008, Kim and Battacharya, 2009). This means that periods when prices are higher than is typical exhibit different dynamics compared to when prices are relatively lower. And there is a very large literature on how income exhibits nonlinearity-output behaves differently over the expansion versus recession phase of the business cycle (Hamilton, 1989 is an early and seminal example). Despite these findings on nonlinearity, all existing papers have ignored the possibility that the relationship between housing values and fundamentals itself may exhibit asymmetric adjustment. Indeed, unless home prices and income exhibited identical, perfectly synchronized nonlinearities, the relationship could well also be asymmetric.

We thus apply the following approach. First, we utilize a recently compiled data set on home values and income (Mack and Garcia, 2011) and examine the relationship for the G-7 countries (previous studies on the topic have examined only the US and UK, although Beltratti and Morana (2010) examined a different topic on housing interactions among the G-7). We have a data set of over four decades, which exceeds the sample size of previous studies by about ten to fifteen years, which should increase the power of the tests we use. However, to anticipate our results, when we apply standard unit root tests to the ratio of home prices to income we still cannot reject the null of nonstationarity for any country (with the very partial exception of Italy). We then apply a test-the Ng-Perron method-which de-trends the data and employs a different lag selection procedure to increase power. But this method still, like previous techniques, imposes linear adjustment on the variables. Once again, we fail to find evidence of a long run relationship for home prices and income in all countries save for the US. We then apply tests for nonlinearity and find-for all nations except the US evidence of asymmetry.

We then apply the Enders-Granger unit root test to all countries-this test allows for nonlinearity in adjustment. We can reject the null of a unit root for five of the six remaining G-7 countries. Overall, the results indicate that, contrary to previous findings, there is a long run relationship between house prices and income for six of the seven G-7 nations.

This paper proceeds as follows. The next section describes the previous literature. The third describes our data and methodology. The fourth section presents our results, and the fifth our conclusion.

Previous Literature

The proposition that house prices and income should have a long run relationship, with a tendency to revert to some attractor, comports with intuition. In particular, while bubbles, or periods when values exceed levels justified by fundamentals may exist, it is somewhat hard to conceive intuitively that values could drift ever farther from incomes, over several decades, with no tendency to revert to a more sustainable ratio with income. Theoretically, Leung (2014) uses a DSGE model to show that home prices and income should be cointegrated. In a similar vein, Holly, et al. (2010) specify a theoretical model in which the ratio of home prices and income must be stationary.

Empirically, however, demonstrating that housing and income share a stationary relationship has proven difficult. Malpezzi (1999) uses data on 133 US MSAs over 1979-1996. Applying a panel unit root test, he finds he cannot reject the null of nonstationarity in the price/per capita income ratio. However, Malpezzi then regresses house prices on income, and tests the residuals for a unit root with the Levin Lin Chu panel unit root test, and rejects the null of nonstationarity in these residuals. At first glance then, one might conclude that house prices and income are cointegrated. However Gallin (2006) points out that Malpezzi's method of first running a regression of house prices on income and then testing for a unit root "overstates the likelihood of cointegration because it ignores the first-stage estimation in the residuals-based cointegration test" (Gallin, p. 419).

Gallin applies the Pedroni method which is specifically designed to test for panel cointegration to twenty-three years of US house prices and personal income. However, the author cannot reject the null hypothesis of no cointegration.

Meen (2002) was the only study to examine national-level data on home prices and per capita income for stationarity, which he did for the UK and US. Using data spanning 1969-1996, Meen specifies home values and income in ARDL models, and tests for cointegration by applying the ADF test to the residuals. He states that his test statistics are "close to their critical values" (p. 8). Gallin (p. 419), however, points out correctly that Meen has failed to find cointegration at conventional significance levels, despite Meen's interpretation.

Mikhed and Zemcik (2009) adopt a slightly different approach to the issue of housing values and fundamentals, and examine the relationship between home prices and rents, rather than income, for twenty three US MSAs over 1978-2006. Rental data is not available for many other countries. The authors test for stationarity in the ratio of prices to rents with the CIPS panel unit root test, but find they cannot reject the null hypothesis that the ratio is I(1). Similarly, they test for cointegration between prices and rents with Pedroni's panel cointegration test, but cannot reject the null of no cointegration. Again, it appears home values do not share a long run relationship with fundamental determinants.

Holly, Pesaran and Yamagata (2010) examine home values and per capita disposable income for forty-nine US states, using annual data over 1975-2003. They model home prices as a function of spatial and other factors. The authors test for cointegration by collecting residuals from the specified panel, and applying the CIPS unit root test. The authors are able to reject the null that *all* of the residuals are nonstationary. From this it would seem that, at first glance, the use of panel data, and the CIPS unit root test that takes account of cross-sectional dependence, a long run relationship between home values and income has been found. Unfortunately, such an interpretation is unwarranted. The null hypothesis maintained in the CIPS test is that all of the series-i.e all of the price income ratios in all forty none states, are nonstationary. The alternative is that at least one such series is stationary. Thus, the most we can infer from Holly, et al. is that perhaps one or more states exhibit a long run relationship between home prices and income, but not that home prices and income generally share a stationary relationship.

These results are similar to those from the debate on purchasing power parity in the international finance literature. If purchasing power parity holds, the real, price level adjusted currency exchange rate between two countries should be mean-reverting. Numerous studies were unable to find mean reversion using standard unit root tests between individual countries. Accordingly, some researchers resorted to applying panel unit root tests to a set of real exchange rates between multiple countries, and reported results in which the null hypothesis of nonstationarity in *all* bilateral exchange rates was rejected. Taylor and Taylor (2004) have pointed out that this result only meant that at least one of the real exchange rates was mean-reverting, but that researchers "tended to draw a much stronger inference that all of the real exchange rates were mean-reverting-and this broader inference is not valid" (p. 145). Thus, to reiterate, the results of Holly, et al. do not indicate that a linear combination of home prices and income is generally stationary, only that the combination is mean-reverting for at least one of forty-nine states.

In the light of theory (Leung, 2014, Holly, et al. 2010) as well as intuition, the inability to find a mean-reverting relationship between home prices and income presents a puzzle. Gallin (2006) suggests that the span of data employed in most studies-usually less than thirty years in existing papers-may be "too short to estimate what may be a genuine long run relationship" (p. 419). The addition, then of another decade or more of data may help in uncovering a stationary relationship.

Another possible issue is that all of the studies to date have employed methods which impose linear, or symmetric adjustment on the price-income relationship, regardless of the nature of the deviation from a possible long run attractor. It is quite possible, for instance, that if prices are above their long-term level vis-a-vis income, the adjustment will be different than if prices are lower than "usual" with respect to income.

Upon reflection, this imposition of symmetric adjustment seems odd. Many financial and economic time series are modeled with nonlinear specifications. To take just one example, unemployment is known to rise quickly but fall more slowly (Rothman, 1998). Output variables such as GDP have long been specified as nonlinear processes to capture the different dynamics over the business cycle-see Hamilton, 1989 as an early example. Moreover, home values have also been modeled as nonlinear processes. Miles (2008) finds that nonlinear specifications provide for better forecasts than linear models for state level home price data. Kim and Bhattacharya (2009) also find evidence of nonlinearity for US regional and national home values.

Unless home prices and income follow identical nonlinear processes, it is likely the ratio could exhibit nonlinearity. Unit root tests have low power when the variable in question exhibits asymmetric adjustment. The failure to apply tests which account for nonlinearity is thus an omission in the literature.

Data and Methodology

To examine whether home prices and income have a long run relationship for the G-7 nations, we utilize the data set of Mack and Garcia (2011) from the Dallas Federal Reserve website. A challenge for papers on corss-country home price movements is the inconsistent methodologies employed in calculating housing indices in different countries (Silver, (2012), Hirata, Kose, Otrok and Terrones (2013)). Fortunately, Mack and Garcia have computed home price indices for eighteen different countries. While not obviating all issues of cross-country comparison, they are an improvement on previously available house price data, as they are all created with a method that is used by the US Federal Housing Finance Agency's index.

The data is adjusted for inflation and seasonality, and is quarterly. The series run from 1975:1 through 2017:1 (the indices, originally published in 2011, are updated regularly on the Dallas Fed's website). The authors also compile an index of per-capita disposable income for each of the eighteen countries. We will, as in previous studies, employ the ratio of each country's house price index to per capita income as our metric. Graphs of the ratios are displayed in Figures one through seven. The ratios do display a fair amount of variation over the four decades. In addition, some ratios-all but those for Italy and the US-appear to exhibit trends over the sample period. Of course the house price/income ratio can change over time-people may desire higher quality housing, for example, and land can become more valuable as time proceeds. However, it would be at odds with theory and intuition to have home prices and income move completely apart, even from a pronounced trend.

Given the longer span of data compared to previous studies, we first apply the standard Augmented Dickey Fuller (ADF) test to the ratios, to see if there is stationarity over four, rather than say three decades. To anticipate our results, we find that this method still yields test statistics that largely indicate the ratios are unit root processes.

We therefore next apply the Ng-Perron test to the data. This procedure was specifically developed to have greater power than the ADF method. The Ng-Perron test entails first de-trending the data so that the presence of a constant or trend do not lower the power of the test. The main innovation of the Ng-Perron technique, however, is in the selection of lags for the test specification. While criteria such as the AIC or SIC are often used to select lags for the ADF test (we employ the SIC in our ADF testing as it has been shown to be more likely than the AIC to choose the actual number of lags in the data generating process) Ng and Perron show that such methods often lead to the incorrect number of lags, often too few. They propose the Modified AIC (MAIC) for choosing lags. As will be displayed, this will mean a greater number of lags for two of the G-7 countries (France and Japan).

However, like the ADF, the Ng-Perron test imposes linear adjustment on all deviations from any long-term mean or trend. Thus whether a shock is positive or negative its impact is specified to decay at the same rate. Enders (2014) points out that Balke and Fomby (1997) have demonstrated that unit root tests have low power when the process being tested displays asymmetry in its adjustment. Given the well-documented nonlinearities in home prices as well as income, it may well be the case that such tests falsely lead to the conclusion that the price/income ratio is nonstationary when in reality it is a stationary process with asymmetric adjustment. And again to anticipate our results, the linear Ng-Perron test will indicate that, in six of seven cases, we still cannot reject the null of non-stationarity. We will then test all ratios for evidence of nonlinearity. Next, we will apply the Enders-Ganger unit root test to all seven ratios.

Enders (2014) cited Balke and Fomby (1997) on the lack of power of standard unit root tests in the presence of asymmetric adjustment. Accordingly, Enders and Granger developed their nonlinear unit root tests. The authors point out that the specification of standard unit root tests, such as the ADF, is as follows:

$$\Delta y_t = \rho y_{t-1} + \varepsilon_t \quad (1)$$

For simplicity, the variable y_t is only a first-order autoregressive process and contains no constant or trend, but of course the testing procedure is the same even if these restrictions are relaxed. That procedure would be of course to test the null hypothesis of nonstationarity i.e. that ρ is zero, against the alternative of mean reversion, or that ρ is less than zero. But the specification in (1) imposes linearity on the process-whether y_t is above or below a long-run attractor, the adjustment is the same. And of course if y_t is a stationary but nonlinear process, the power of the test is reduced so one may well conclude that a mean-reverting variable actually contains a unit root.

Enders and Granger propose an alternative specification:

$$\Delta y_t = I_t \rho_1 (y_{t-1} - a_0) + (1 - I_t) \rho_2 (y_{t-1} - a_0) + \varepsilon_t \quad (2)$$

Where $I_t = 1$ if $y_{t-1} \geq a_0$, and $I_t = 0$ if $y_{t-1} < a_0$ (equations 1 and 2 are from Enders and Granger, p. 305). In (2), a_0 is an attractor. The attractor could be zero, or a constant, or a constant and trend. And as with (1), further lags could be employed in modelling the process. The null hypothesis in the Enders-Granger test is that of a unit root in (1), as with standard tests such as the ADF. The alternative hypothesis, specified in (2) is that y_t follows a threshold autoregressive (TAR) process, where the variable follows one type of dynamic when the lag of y is above its threshold, or attractor (a_0) and another when y_{t-1} is below the threshold. The authors have calculated critical values for testing whether y_t has a unit root. In addition, one can test whether the process is asymmetric, i.e. whether $\rho_1 = \rho_2$. As so many variables-GDP, unemployment, home values, interest differentials, etc. have been found to be nonlinear, we will employ the Enders-Granger method to discern whether home values and income have a long-term relationship across the G-7.

Results

Table 1 displays the results from the ADF unit root tests. The lag length was chosen by the SIC criterion. Given figures 1 through 7, linear trends are included for all countries but Italy and the US. Results are not changed based on whether a trend is included, with the partial exception of Italy, for which we cannot reject the null of a unit root at the five percent level with or without a linear trend, but we do reject the null at the ten percent level without a trend but not with a trend.

Consistent with previous papers, for no country's ratio is the null of nonstationarity rejected at the five percent level. Indeed, with the above-noted exception of Italy, for no country is the null rejected at even ten percent. And in Italy's case we can just barely reject the null at ten percent, with a p-value of 0.096.

Thus the use of a longer series, which spans more cycles in housing than the data used in previous papers, has not for the most part changed the finding that house prices and income do not appear to share a long run relationship. Thus given the well-known low power of the ADF method, we next performed the Ng-Perron test. As displayed in Table 2, the modified AIC criteria ends up choosing five lags for France and thirteen for Japan (as opposed to three lags and one for each respective country chosen by the SIC criterion for the ADF). Despite the greater power of this test, we are still unable to reject the null of a unit root for six of the seven countries, even at the ten percent level. The United States is the one exception. We are able to reject the null of nonstationarity in the US ratio at the five percent level. Results for the US are different from those found by Meen (2002) who could not reject the null of no cointegration using the ARDL method. Our use of a longer span of data (42 years versus 27) as well as a more powerful technique have yielded a different finding.

For the remaining countries, we still fail to find a stable long run relationship between prices and income using the Ng-Perron test, which, despite its greater power compared to the ADF, sill imposes linear adjustment on all shocks. We thus test each series for nonlinearity using the Brock, Dechert and Scheinkman (BDS) test. The BDS procedure tests for time-based dependence in the residual of a series. If a series is linear, residuals from the linear model should be iid. This means that the probability that the distance between any two residuals is less than a constant should be the same for all residuals. If this condition does not hold, it is evidence of nonlinearity.

To conduct the test, we obtain residuals from the linear models used to conduct the Ng-Perron test-the difference of the ratio regressed on lags (the number of lags chosen by the MAIC for the test), the lagged level of the ratio and a constant, and except for Italy and the US, a trend. In all cases, we use one standard deviation as the distance between the pairs. Results of the test are shown in Table 3. As displayed, results indicate that the null of linearity is rejected for all ratios at the five percent level, with the exception of the US case. Given that the US is the one country that appeared to have a stable, long run relationship between income and prices by the Ng-Perron test, it is plausible that prices and income share a relationship that is "linear enough" that the Ng-Perron test indicates stationarity. For the other six nations, we will apply the Enders-Granger method (we will also apply it to the United States, just to compare results). Results are displayed in Tables four through ten.

The number of lags for each test is the same as those chosen by the MAIC for the Ng-Perron test. In Tables four through ten the symbols ρ_1 and ρ_2 of course refer to the estimates of ρ_1 and ρ_2 from equation 2. The φ corresponds to the test statistic for the unit root. The equality F-test is the test statistic for the null hypothesis that both regimes are identical, or that the ratio is a linear process. After the F-test we display the coefficients and standard errors for the lags employed in the test. As with the ADF and Ng-Perron tests, all ratios include a trend except those for Italy and the US.

Results in Tables 4 through 10 indicate that the null hypothesis of nonstationarity can be rejected at the five percent level for all countries except France and the US. In the case of France, the test statistic of 5.203 is not close to the five percent critical value of 6.3 (Enders and Granger, p. 306). However, it is not that far from the ten percent significance critical value of 5.27. Also these critical values are for one hundred observations, and here we have one hundred and sixty-nine. The critical value for two hundred and fifty observations is 5.18. We obviously cannot run the simulations of Enders and Granger to determine the exact critical value for our number of observations, but for France we cannot confidently reject nonstationarity.

For the US, our test statistic of 3.521 also does not allow us to reject the null hypothesis at the five percent (critical value 4.64) or even ten percent (critical value 3.79) levels. However, we believe the US price-income ratio is mean reverting, based first on the Ng-Perron results and on the failure to reject the null of nonlinearity with the BDS test earlier. And the US was the only country for which the null of linearity was not rejected with the BDS procedure. In addition, it is true that if the DGP of a series is nonlinear, standard unit root tests will have low power (Balke and Fomby, 1997). However, if a process is linear, or if deviations from symmetry are not large in magnitude, splitting the data into different regimes to estimate an additional parameter will also lower the power of the test. Thus the Ng-Perron results for the US seem credible.

For the five countries besides France and the US, where we can reject the null of a unit root at the five percent level with this Enders Granger method, the F-tests for the null of equality of regimes, which have standard distributions under the alternative (stationary) hypothesis, clearly indicate asymmetric adjustment.

In examining the adjustment, we first note that in the original Enders-Granger paper, the authors apply their method to the interest rate spread between ten year bonds and the fed funds rate. They reject the null of a unit root in the differential, and reject, with the F-test, the null hypothesis of symmetric adjustment. They find that while their estimate of ρ_1 is significant, their estimate of ρ_2 is not. This of course does not mean that adjustment is symmetric, but rather that adjustment is much greater when the differential is above instead of below its attractor.

In our results, for the UK, both ρ_1 and ρ_2 are clearly significant. For Canada and Germany, one regime is clearly significant while the other has a t-stat of close to, but not quite two. Italy and Japan each have one clearly significant regime, while the second regime has a low t-statistic. For the cases of France and the US, where we could not, with this test, reject the null of a unit root, we note that France has one regime that is significant, while the other is nearly so, while the US has one regime that is significant while the other is clearly not.

In five of seven cases, three of which where the null of a unit root is rejected at five percent, ρ_1 measuring the regime above the attractor is larger in absolute value than ρ_2. This means faster adjustment when the ratio is above, compared to below the attractor. In two cases, Germany and Italy, the opposite holds, in that the absolute value of ρ_2 exceeds that of ρ_1. One interpretation would be, that in the nations where ρ_1 exceeds ρ_2 (in absolute value), such as the UK, home prices may rise above income, and income may then rise, perhaps somewhat in response (housing is a leading indicator of the macroeconomy). But when the ratio is below the attractor, adjustment is slow-it may be that house prices are either falling, or growing more slowly in response to a housing downturn, and it takes a long time-i.e. time on market, to adjust to the long run relationship with income, as prices are stickier downward.

In contrast, in nations with more moribund markets such as Germany this impact of housing on the economy may be less pronounced.

In any event, the exact nonlinear dynamics of the house price/income ratios in different countries would be a topic for future research. Our main result is that, for six of the seven countries, there does appear to be a stable, long-run relationship between home values and income.

Conclusion

The exact nature of the nonlinearities found in most of the home price/income ratios for the G-7 are a topic for future research. There are many different types of existing nonlinear models-the TAR investigated here, as well as the bilinear, GAR, ESTAR, LSTAR, Markov Switching, etc. The exact model that best fits the ratio will likely differ across countries.

But our main finding is that the home price/income ratio for six of seven G-7 countries is stationary. This resolves a puzzle at the national level, as while bubbles could be thought of as deviations from a long-run relationship, the notion that there is *no* long run relationship at all runs counter to theory and intuition. This puzzle has largely been solved with the use of a longer data set, more powerful tests and mostly by allowing for nonlinearity in the dynamics of the ratio.

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References

Balke, N. and S. Fomby (1997) "Threshold Cointegration", *International Economic Review*, 38, 627-643.

Beltratti, A. and C. Morana (2010) "International House Prices and Macroeconomic Fluctuations", *Journal of Banking and Finance*, 34, 533-545.

Enders, W. (2014) Applied Econometric Time Series, John Wiley and Sons.

Enders, W. and C. Granger (1998) "Unit Root Tests and Asymmetric Adjustment with an Example Using the Term Structure of Interest Rates", *Journal of Business and Economic Statistics*, 16, 304-311.

Gallin, J. (2006) "The Long-Run Relationship between House Prices and Income: Evidence from Local Housing Markets", *Real Estate Economics*, 34, 417-438.

Hamilton, J. (1989) "A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle", *Econometrica*, 57, 357-84.

Hirata, H., M. Kose, C. Otrok and M. Terrones (2012) "Global House Price Fluctuations: Synchronization and Determinants", *IMF Working Paper WP/13/38.*

Holly, S., M. Pesaran and T. Yamagata (2010) "A Spatio-Temporal Model of House Prices in the USA", *Journal of Econometrics*, 158, 160-173.

Kim, S. and R. Battacharya (2009) "Regional Housing Prices in the USA: An Empirical Investigation of Nonlinearity", *Journal of Real Estate Finance and Economics*, 38, 443-460.

Leung, C. "Error Correction Dynamics of House Prices: An Equilibrium Benchmark", *Federal Reserve Bank of Dallas Globalization and Monetary Policy Institute Working Paper No. 177.*

Mack, A. and E. Garcia (2011) "A Cross-Country Quarterly Database of Real House Prices: A Methodological Note", *Federal Reserve Bank of Dallas Globalization and Monetary Policy Institute Working Paper 99.*

Malpezzi, S. (1999) "A Simple Error Correction Model of House Prices", *Journal of Housing Economics*, 8, 27-62.

Meen, G. (2002) "The Time Series Behavior of House Prices: A Transatlantic Divide?", *Journal of Housing Economics*, 11, 1-23.

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

Miles, W. (2008) "Boom Bust Cycles and the Forecasting Performance of Linear and Non-Linear Models of House Prices", *Journal of Real Estate Finance and Economics*, 36, 249-264.

Ng, S. and P. Perron (2001) "Lag Length Selection and the Construction of Unit Root Tests with Good Size and Power", *Econometrica*, 59, 1519-1554.

Rothman, P. (1998) ´Forecasting Asymmetric Unemployment Rates”, *Review of Economics and Statistics,* 80, 164-68.

Silver. M. (2012) “Why House Price Indices Differ: Measurement and Analysis”, *IMF Working Paper 12/125*.

Taylor, A. and M. Taylor (2004) “The Purchasing Power Parity Debate”, *Journal of Economic Perspectives,* 18, 135-158.

Table 1
ADF Test Results

Country	Lags	Test Statistic	P-Value
Canada	1	-1.114	0.9227
France	3	-2.15	0.5088
Germany	3	-1.485	0.8311
Italy	3	-2.595	0.096
Japan	1	-1.61	0.7838
UK	1	-2.36	0.3986
US	3	-2.21	0.2

In all cases the number of lags was chosen by the SIC criterion. For all countries except Italy and the US a trend (as well as a constant) was included, based on the graphs. However, in all cases the results of failure to reject the null of a unit root were maintained regardless of whether a trend was included, with the partial exception of Italy. As displayed, it is possible to (barely) reject the null of a unit root at the ten, but not five percent level when a trend is not included, while when a trend is included, it is no longer possible to reject even at the ten percent level (the p-value is 0.226, results available upon request).

Table 2
Ng-Perron Test Results

Country	Lags	5% Critical Value	Test Statistic
Canada	1	-17.3	-4.425
France	5	-17.3	-8.717
Germany	3	-17.3	-9.943
Italy	3	-8.1	-1.0625
Japan	13	-17.3	-4.69
UK	1	-17.3	-9.104
US	3	-8.1	-8.617

In all cases the number of lags was chosen by the Modified AIC criterion. For all countries except Italy and the US a trend (as well as a constant) was included, based on the graphs. However, in all cases the results were maintained regardless of whether a trend was included. The null is rejected when the test statistic is less than the critical value.

Table 3
BDS Test Results

Country	BDS Test Stat.	Standard Error	P-Value
Canada	0.017899	0.007362	0.015
France	0.029501	0.006313	0.0000
Germany	0.036035	0.007461	0.0000
Italy	0.036431	0.007037	0.0000
Japan	0.034884	0.009278	0.0002
UK	0.01046	0.005152	0.0423
US	0.008009	0.007444	0.282

These were tests performed on the residuals of the linear specifications for the Ng-Perron test (the results of which are displayed in Table 2).

Table 4
Enders-Granger Test Results for Canada

Attractor	0.888+0.0033t
ρ_1	-1.2993
Std. Error	0.278
ρ_2	-0.0118
Std. Error	0.0062
φ_τ	12.9426
5% Critical Value	6.30
Equality F-test	21.411
Δy_{t-1}	0.4072
Std. Error	0.0718

The symbols ρ_1 and ρ_2 refer to the regimes above and below the attractor. The φ corresponds to the test statistic for the unit root. The equality F-test is the test statistic for the null hypothesis that both regimes are identical, or that the ratio is a linear process.

Table 5
Enders-Granger Test Results for France

Attractor	0.652 + 0.0027t
ρ_1	-0.0848
Std. Error	0.0311
ρ_2	-0.0094
Std. Error	0.0054
φ_τ	5.203
5% Critical Value	6.30
Equality F-test	5.7122
Δy_{t-1}	0.4037
Std. Error	0.0776
Δy_{t-2}	0.0943
Std. Error	0.0824
Δy_{t-3}	0.1953
Std. Error	0.0818
Δy_{t-4}	0.2221
Std. Error	0.0833
Δy_{t-5}	-0.1012
Std. Error	0.0798

The symbols ρ_1 and ρ_2 refer to the regimes above and below the attractor. The φ corresponds to the test statistic for the unit root. The equality F-test is the test statistic for the null hypothesis that both regimes are identical, or that the ratio is a linear process.

Table 6
Enders-Granger Test Results for Germany

Attractor	1.6401-0.0055t
ρ_1	-0.0126
Std. Error	0.0065
ρ_2	-1.6256
Std. Error	0.3161
φ_τ	15.4137
5% Critical Value	6.30
Equality F-test	25.9961
Δy_{t-1}	0.2633
Std. Error	0.0744
Δy_{t-2}	0.3124
Std. Error	0.0677
Δy_{t-3}	0.1763
Std. Error	0.0715

The symbols ρ_1 and ρ_2 refer to the regimes above and below the attractor. The φ corresponds to the test statistic for the unit root. The equality F-test is the test statistic for the null hypothesis that both regimes are identical, or that the ratio is a linear process.

Table 7
Enders-Granger Test Results for Italy

Attractor	0.767
ρ_1	-0.009
Std. Error	0.0065
ρ_2	-0.0853
Std. Error	0.0315
φ_μ	4.6606
5% Critical Value	4.64
Equality F-test	5.6028
Δy_{t-1}	0.7923
Std. Error	0.0639
Δy_{t-2}	-0.5737
Std. Error	0.0781
Δy_{t-3}	0.5157
Std. Error	0.0604

The symbols ρ_1 and ρ_2 refer to the regimes above and below the attractor. The φ corresponds to the test statistic for the unit root. The equality F-test is the test statistic for the null hypothesis that both regimes are identical, or that the ratio is a linear process.

Table 8
Enders-Granger Test Results for Japan

Attractor	2.2248-0.0077t		
ρ_1	-0.4673		
Std. Error	0.1004		
ρ_2	-0.0027		
Std. Error	0.0026		
φ_τ	11.0765		
5% Critical Value	6.30		
Equality F-test	21.493		
Δy_{t-1}	0.4628		
Std. Error	0.0806		
Δy_{t-2}	0.1312		
Std. Error	0.0885		
Δy_{t-3}	0.2628		
Std. Error	0.089		
Δy_{t-4}	0.0186		
Std. Error	0.0907		
Δy_{t-5}	0.1367		
Std. Error	0.0878		
Δy_{t-6}	-0.1031	Δy_{t-10}	0.1031
Std. Error	0.0871	Std. Error	0.089
Δy_{t-7}	-0.1431	Δy_{t-11}	-0.1553
Std. Error	0.0867	Std. Error	0.0881
Δy_{t-8}	-0.1232	Δy_{t-12}	0.0685
Std. Error	0.0864	Std. Error	0.0886
Δy_{t-9}	0.2714	Δy_{t-13}	-0.0509
Std. Error	0.0867	Std. Error	0.0776

The symbols ρ_1 and ρ_2 refer to the regimes above and below the attractor. The φ corresponds to the test statistic for the unit root. The equality F-test is the test statistic for the null hypothesis that both regimes are identical, or that the ratio is a linear process.

Table 9
Enders-Granger Test Results for UK

Attractor	0.553+0.0037t
ρ_1	-0.1405
Std. Error	0.045
ρ_2	-0.0098
Std. Error	0.0049
φ_τ	6.9186
5% Critical Value	6.30
Equality F-test	8.3226
Δy_{t-1}	0.7364
Std. Error	0.0524

The symbols ρ_1 and ρ_2 refer to the regimes above and below the attractor. The φ corresponds to the test statistic for the unit root. The equality F-test is the test statistic for the null hypothesis that both regimes are identical, or that the ratio is a linear process.

Table 10
Enders-Granger Test Results for US

Attractor	0.9834
ρ_1	-0.1272
Std. Error	0.0514
ρ_2	-0.0055
Std. Error	0.0053
φ_μ	3.521
5% Critical Value	4.64
Equality F-test	5.575
Δy_{t-1}	0.4888
Std. Error	0.0722
Δy_{t-2}	-0.1562
Std. Error	0.0806
Δy_{t-3}	0.3921
Std. Error	0.073

The symbols ρ_1 and ρ_2 refer to the regimes above and below the attractor. The φ corresponds to the test statistic for the unit root. The equality F-test is the test statistic for the null hypothesis that both regimes are identical, or that the ratio is a linear process.

Figure 1

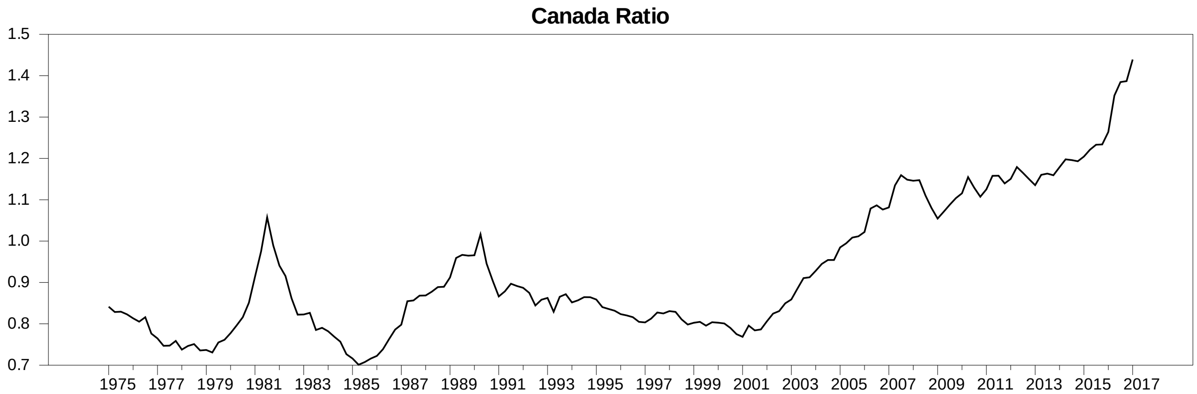


Figure 2

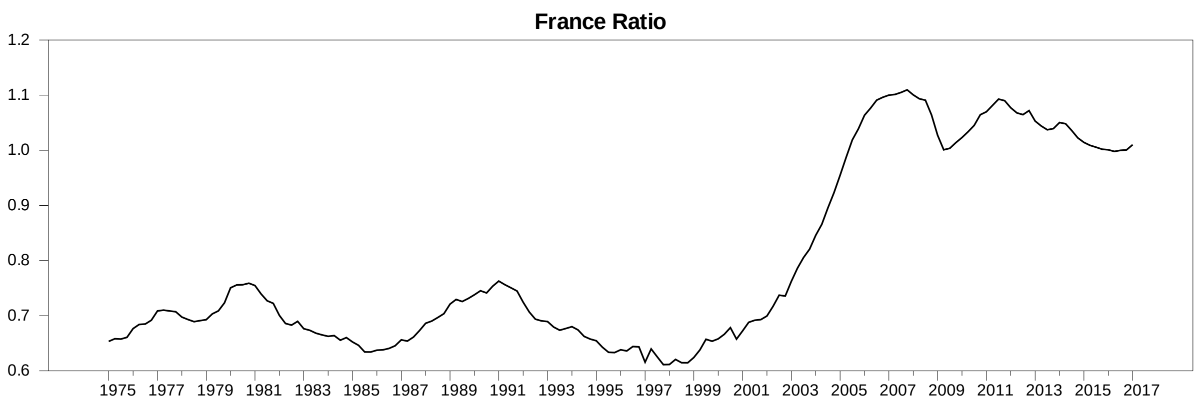


Figure 3

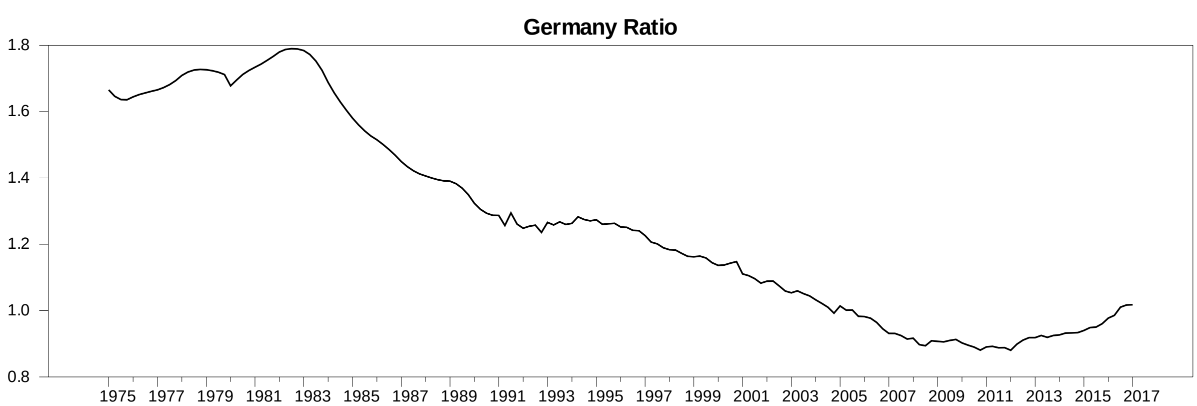


Figure 4

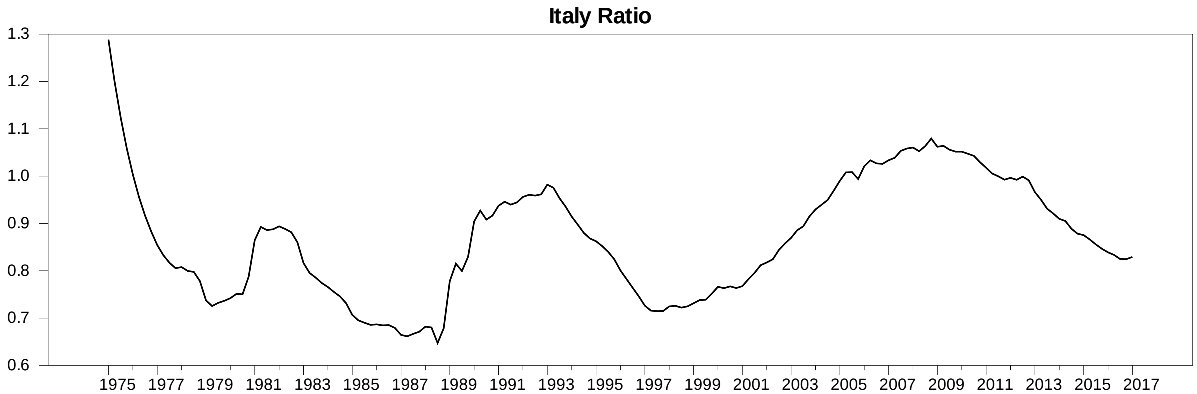


Figure 5

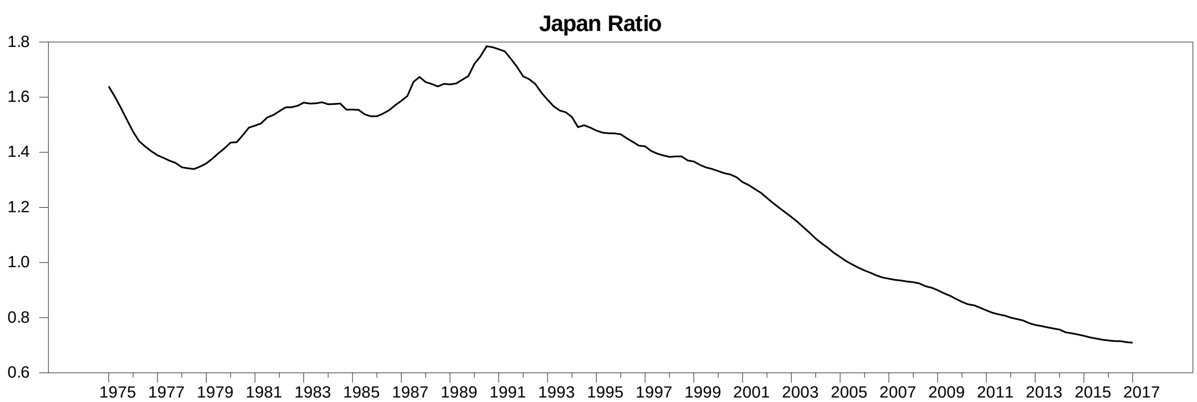


Figure 6

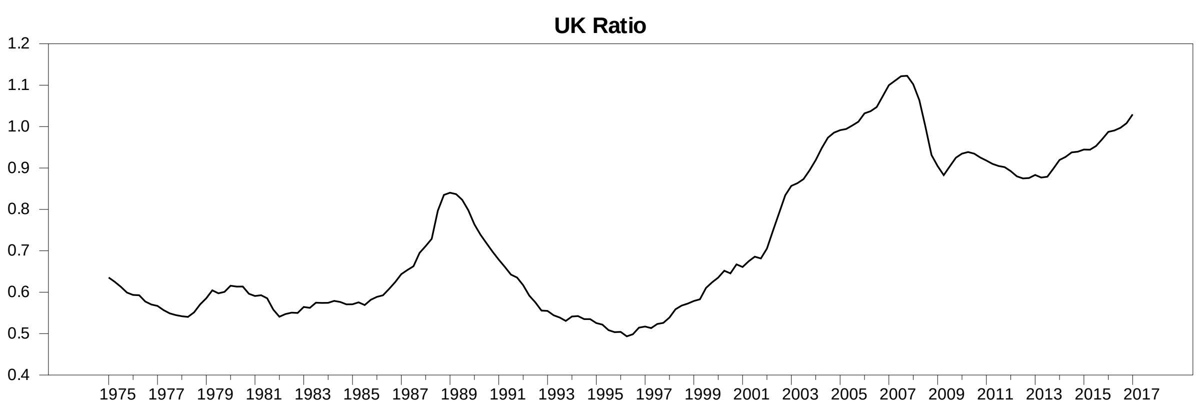
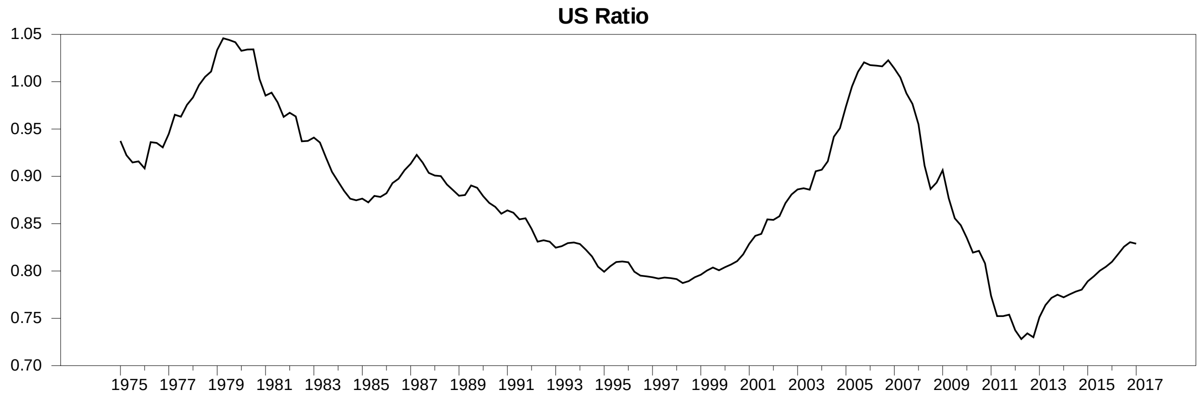


Figure 7



US Ratio
1.05
1.00
0.95
0.90
0.85
0.80
0.75
0.70
1975 1977 1979 1981 1983 1985 1987 1989 1991 1993 1995 1997 1999 2001 2003 2005 2007 2009 2011 2013 2015 2017