

PERSISTENCE AND CYCLICAL DYNAMICS OF US AND UK HOUSE PRICES: EVIDENCE FROM OVER 150 YEARS OF DATA

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Abstract

This paper provides a new and unique look at the dynamics and persistence of historical house prices in the USA and the UK using fractional integration techniques not previously applied to housing markets. Unlike previous research, we consider two components of persistence of house prices: the component associated with the long-run trend and the component associated with the cycle. We find evidence of cyclical and long-run persistence in the UK housing markets. In contrast, we fail to find evidence of cyclical persistence for the USA. For the sub-samples, which account for a structural break in each series, an important difference is the asynchronous pattern of the breaks, an indication of heterogeneity in the house price dynamics of the two countries and a sign that national rather than global events have played an important role. Although the house price movements of the last decade are dramatic, the greatest structural changes in the overall nominal and real price dynamics of the UK and the USA appear to have taken place much earlier, in the late 1970s and early 1980s in the UK and in the mid-1950s and early 1970s in the USA. An important result, common to the whole and sub-samples, is that long-run persistence plays a greater role than cyclical persistence in explaining the dynamics of house prices in both countries. These findings have substantial implications for policy decisions.

Keywords

cyclical behaviour, fractional integration, house prices, persistence

Introduction

In recent years, considerable interest has focused on the housing markets, and a sizeable literature recognises that housing markets play a critical role in the economy, the business cycle and the financial system. Evidence from recent economic history includes Case et al. (2005), Carroll et al. (2011), Attanasio et al. (2011), Chen et al. (2018), Davis and Heathcote (2005), Leamer (2007), Funke and Paetz (2013), and many others.¹ The important role of housing markets in the business cycle became painfully clear during the collapse of the sub-prime mortgage market in late 2006 and the resulting severe recession and financial crisis of 2007–2009, the worst since the Second World War (Mian and Sufi, 2010). Shiller (2007) claims that the housing bubble that began in the mid-1990s is the major, if not the only, cause of the sub-prime mortgage crisis and the worldwide economic and financial crisis of 2007–2009. Leamer (2007) offers a more provocative assertion, arguing

that for the USA, ‘housing is the business cycle’ or, more precisely, that house prices drive the US business cycle.²

Arguably, the recent financial crisis, more than any other macroeconomic event, underscores the importance of understanding the dynamics of house prices and, in particular, the role of persistence and the effect of shocks on house price dynamics. Numerous empirical studies have analysed these issues using alternative time-series methods, including univariate and panel unit-root tests, and fractional integration. This literature is not only relevant to our understanding of the dynamics of house prices, but also sheds some light on and at times questions the appropriateness of theoretical urban and housing models. For instance, Capozza and Helsley (1989, 1990) suggest that an equilibrium relationship exists between real house prices and real income. If real income has a unit root and house prices are stationary, however, then the equilibrium relationship

between real house price and real income does not seem plausible, given the time-series characteristic of the two data series.

Meen (1999) and Peterson et al. (2002) find that the UK house prices follow a unit-root process. Meen (2002) fails to reject the unit root in house prices in the UK and the USA. Muñoz (2004) and Clark and Coggin (2011) fail to reject the unit-root hypothesis of house prices in the USA. Arestis and González (2014) confirm the presence of a unit root in house prices of 18 OECD countries. In contrast, Cook and Vougas (2009) support the stationarity of UK housing prices but with structural change. More recently, Zhang et al. (2016) present evidence from a 120-year national data set that US house prices are trend stationary.

Unit-root tests do not completely measure the degree of persistence of a series. Unit-root tests discriminate between stationary and non-stationary processes, but do not allow for fractional alternatives, where the non-stationarity property of the data may overlap their mean-reversion property, and where stationarity may not exclude persistence. The standard practice to achieve stationarity differences the data. It is possible that to achieve stationarity, however, only fractional differencing is required (Granger and Joyeux, 1980). In this case, the process is fractionally integrated, or $I(d)$. The fractional integration approach is more general than the standard method that only considers $I(0)$ and $I(1)$ processes, since it allows d to be any real number, including a fractional value. If $d = 0$, the process exhibits ‘short memory’ and the values of the autocorrelations show a fast exponential decay. In contrast, if $d > 0$, the process displays ‘long memory’ and the values of the autocorrelations show a slow hyperbolic decay. If $0 < d < 0.5$, the process is stationary, while $d \geq 0.5$ implies non-stationarity. Moreover, if $d < 1$, the process exhibits mean reversion, which implies that if $0.5 \leq d < 1$, the

process is non-stationary, but mean reverting, while if $d \geq 1$ the process is non-stationary, but not mean reverting. Examples of papers that model house prices as fractional integration processes include Barros et al. (2012, 2015), Gil-Alana et al. (2013, 2014) and Gupta et al. (2014). Two observations, however, are warranted regarding this empirical literature. First, none of these studies tests for the presence of structural breaks in the series. Second, all these studies test only for the presence of long-run persistence in house prices. That is, the failure to include all relevant stochastic characteristics may lead to a biased estimate of the long-run persistence.

Our paper provides a new and unique look at the dynamics and persistence of historical house prices in the USA and the UK, using methods not previously applied to housing markets. We use yearly data on real and nominal house prices over a period from 1830 to 2016 for the USA, and from 1845 to 2016 for the UK, which provides a much longer perspective on the behaviour of house prices than commonly appears in the literature, where most empirical work uses data starting from the 1980s or later. We also differ, however, from previous fractional integration research as we extend the fractional integration methodology by taking into account two components of house price persistence (i.e. the component affecting the long-run trend, and the component affecting the cyclical structure).³ In spectral analysis, persistence related to the long-run trend is persistence at frequency zero, while persistence related to the cyclical pattern of the data is persistence at a frequency away from zero.

We hypothesise that persistence of house prices may play different roles in the long run and in the cycle and that modelling jointly these two closely related components of the house price provides a much broader and more comprehensive view of the

housing market dynamics and persistence. Typically, house prices exhibit a peak in the periodogram at zero frequency, which indicates long-run persistence, but also at a frequency away from zero, indicating cyclical dynamics. Testing for persistence while ignoring the cyclical component of persistence tends to overestimate long-run persistence. The available evidence suggests that the periodicity of economic and financial data ranges from five to ten years and, in most cases, researchers estimate a periodicity of about six years (e.g. Baxter and King, 1999; Canova, 1998; King and Rebelo, 1999).

We consider three different fractional integration models – a standard model, defined by a process with a pole in the spectrum at the zero frequency, a process with a pole at the non-zero frequency, and a composite model by incorporating poles at zero and non-zero frequencies in a single framework. Thus, the third model estimates jointly the two components of persistence in house prices. We estimate each of the three models using the parametric procedure of Robinson (1994). This approach has two distinctive features compared with other methods. First, it does not require normality, which is an assumption rarely satisfied by economic data, and second and most importantly, the tests exhibit standard null distributions.

Finally, we examine the possibility of a structural break in the data. This is a relevant issue, not only because of the historical breadth of the data, but also because fractional integration and structural breaks can easily be confused. We account for this issue by re-estimating the fractional models using two sub-samples, with the dates identified by the Bai and Perron (2003) methodology.

The outline of the paper is as follows. The following section describes the models and outlines the main aspects of the fractional integration methodology. The third section presents the data. The fourth section reports

the full sample results, while the fifth deals with the analysis of breaks. Policy implications appear in the final section.

The models

Let d_L and d_C be, respectively, the long-run and cyclical orders of integration. We consider three fractional integration models. The first $I(d_L)$ model is the standard model of the form advocated, for example, in Gil-Alana and Robinson (1997). The model incorporates two equations. The first accommodates the deterministic terms, while the second expresses the conventional fractional integration model.

$$y_t = \beta_0 + \beta_1 t + x_t, (1 - L)^{d_L} x_t = u_t, t = 1, 2, \dots \quad (1)$$

where y_t is the observed time series, β_0 and β_1 are the coefficients corresponding, respectively, to the intercept and linear time trend, L is the lag operator ($Lx_t = x_{t-1}$), and x_t is $I(d_L)$, where d_L refers to the zero (long-run) frequency order of integration.

Note that the specification in equation (1) includes the standard $I(1)$ case, which is employed in the literature for unit-root testing, when $d_L = 1$. In such cases, shocks are permanent. The fact that x_t is $I(d_L)$ implies that we can express its spectral density function as follows:

$$f_x(\lambda) = \frac{\sigma^2}{2\pi} |1 - e^{i\lambda}|^{-2d_L}, -\pi \leq \lambda < \pi. \quad (2)$$

Thus,

$$f_x(\lambda) \rightarrow \infty \text{ as } \lambda \rightarrow 0^+ \quad (3)$$

We observe this feature in many aggregated data. The spectrum, however, may display a pole or singularity at a non-zero frequency. In this case, the process may still display long memory, but the autocorrelations exhibit a cyclical structure that decays slowly.

This is a property of the Gegenbauer processes (Gil-Alana, 2001), defined as

$$(1 - 2 \cos w_r L + L^2)^d x_t = u_t, t = 1, 2, \dots, \quad (4)$$

where $w_r = 2\pi r/T$ with $r = T/j$, j indicates the number of periods per cycle and r the frequency with a singularity or pole in the spectrum. Note that if $r = 0$, the fractional polynomial in equation (4) becomes $(1 - L)^{2d}$, which is the polynomial associated with the $I(d_L)$ model. Gray et al. (1989) show that x_t in equation (4) is stationary if $|\mu| < 1$ and $d < 0.50$ or if $|\mu| = 1$ and $d < 0.25$, where $\mu = \cos w_r$. These authors also show that we can express the polynomial in equation (4) in terms of the orthogonal Gegenbauer polynomials $C_{j,d}(\mu)$ such that for all $d \neq 0$,

$$(1 - 2\mu L + L^2)^{-d} = \sum_{j=0}^{\infty} C_{j,d}(\mu) L^j \quad (5)$$

Thus, the process in equation (4) becomes:

$$x_t = \sum_{j=0}^{t-1} C_{j,d}(\mu) u_{t-j}, t = 1, 2, \dots,$$

and when $d = 1$, reduces to

$$x_t = 2\mu x_{t-1} - x_{t-2} + u_t, t = 1, 2, \dots \quad (6)$$

which is a cyclical $I(1)$ process of the form proposed earlier by Ahtola and Tiao (1987), Bierens (2001), and others to test for unit-root cycles in AR(2) models. Note that in this model, the spectral density of x_t is given by:

$$f_x(\lambda) = \frac{\sigma^2}{2\pi} |1 - 2\mu e^{i\lambda} + e^{2i\lambda}|^{-2d_L}, -\pi \leq \lambda < \pi, \quad (7)$$

Thus, the second model is the cyclical d_C model (Gil-Alana, 2001), which can be specified as follows:

$$(1 - 2\mu L + L^2)^{d_C} x_t = u_t, t = 1, 2, \dots, \quad (8)$$

where d_C refers to the cyclical order of integration.⁴ As in the $I(d_L)$ model, the fractional order of integration can be any real number and u_t is assumed $I(0)$.

Finally, in the third model, $I(d_L, d_C)$ incorporates the two structures dealing with the degree of persistence in a single framework. That is, we include a structure producing a singularity at the zero frequency (long-run trend) along with another one corresponding to the cyclical frequency. The model is given by:

$$(1 - L)^{d_L} (1 - 2\mu + L^2)^{d_C} x_t = u_t, t = 1, 2, \dots \quad (9)$$

Caporale and Gil-Alana (2014a, b, 2016, 2017) provide detailed technical explanations about estimation and testing procedures suggested by Robinson (1994).

Data

We compile a data set of annual time series for the USA and the UK spanning 1830–2016 and 1845–2016, respectively, which includes nominal and real house prices, with real values obtained by deflating the nominal house prices with the consumer price index. Thus, the US sample contains 187 observations while the UK sample contains 172 observations.

The nominal house price index (i.e. Winans International Real Estate Index, WIREI) for the USA comes from the Global Financial Database (<https://www.globalfinancialdata.com/>). We deflate this index by the Consumer Price Index (CPI) to derive the US real house price index. The CPI data come from the website of Robert Sahr (<http://oregonstate.edu/cia/polisci/sahr/sahr>). The nominal house price and the Consumer Price Index data for the UK come from the database *A Millennium*

of *Macroeconomic Data* maintained by the Bank of England at: <https://www.bankofengland.co.uk/statistics/research-datasets> as part of the Three Centuries of Macroeconomic Data project. For a summary overview of the methodology and construction of this database, see Thomas and Dimsdale (2017).

As in the US case, we obtain the UK real house price index by deflating the nominal index by the CPI. An advantage of these historical samples is the ability to examine how the housing markets of these two countries evolve over time, covering almost their entire modern economic history. These series are the longest available annual data on house prices in the USA and the UK. From the perspective of fractional integration, however, they are relatively small samples. The US sample contains 187 observations while the UK sample contains 172 observations.

Figure 1 plots the US and UK real and nominal price series in their log-transformed form as well as the first differences of the log-transformed data. Several observations come from the descriptive analysis of the data. First, real and nominal house prices increased in both the UK and the USA over the sample periods. Between 1845 and 2016, UK house prices rose at an average annual rate of growth of 3.8% in nominal terms and 1.1% in real terms. By comparison, between 1830 and 2016, US house prices rose at an average annual rate of growth of 3.5% in nominal terms and 1.7% in real terms. Second, the growth of nominal and real US and UK house prices has experienced different rates over time. UK house prices in real and nominal terms remained relatively stable from 1845 to 1898. Between 1899 and 1941, however, UK house prices fell on average by 1.2% per year in real terms, although they increased by 1.1% per year in nominal terms. After the Second World War, UK house prices began a positive trend, with particularly high growth rates in the 1990s until the Great Recession.

During the Great Recession (2007–2009), UK house prices declined on average by 6.3% per year in real terms and 4.5% per year in nominal terms and did not recover at the end of the Great Recession, reaching new lows in 2012.

By comparison, US house prices in nominal terms remained relatively stable until the 1950s. US house prices in real terms increased by 1.6% per year until the First World War, contracted during the war, and recovered during the interwar period. During the Great Depression (1929–1939), US house prices fell by 1.6% per year in real terms and by 3.5% in nominal terms. Following the Second World War, US house prices first surged then remained remarkably stable until the early 1990s. During the past two decades, US house prices increased substantially before falling steeply during the Great Recession and beginning to recover only five years after the end of the Great Recession.

Since 2012, the increase in house prices in the USA was more dramatic than that in the UK. The real estate bubble, where house prices peaked in early 2006, started to decline in 2006 and 2007 and reached new lows in 2012, appears pronounced in both countries.⁵

Empirical results for the whole sample

Results from the long-run $I(d_L)$ model

Table 1 reports the whole sample estimates of the degree of fractional integration $d = d_L$ in the first model, $I(d_L)$, which considers only the long-run component of persistence of the series. We assume that the disturbances are uncorrelated (white noise) (top panel of Table 1) and autocorrelated (bottom panel of Table 1). In the latter case, we use a non-parametric approach proposed by Bloomfield (1973) that approximates highly

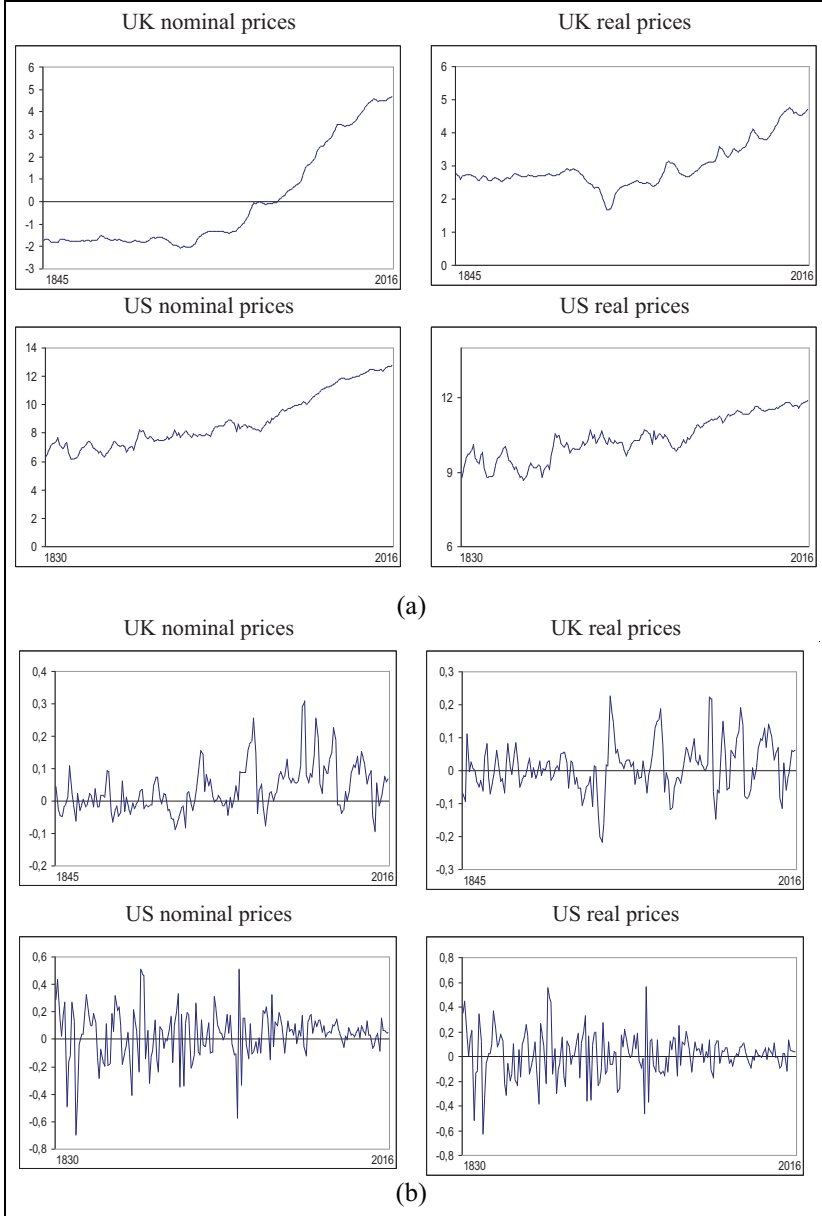


Figure 1. Data. (a) Log-transformed data. (b) First differences on the log-transformed data.

parameterised ARMA processes with a small number of parameters and that accommodates extremely well in the context of fractional integration (Gil-Alana, 2004; Velasco and Robinson, 2000). For each series, we

consider the three standard cases examined in the literature: (1) no deterministic terms (i.e. $\beta_0 = \beta_1 = 0$), (2) an intercept and no trend (β_0 unknown, and $\beta_1 = 0$), and (3) a constant with a linear time trend (β_0 and β_1

Table 1. Estimates of d_L for the whole sample using a parametric approach.

Series	No terms	An intercept	A linear time trend
(i) <i>White noise</i>			
UK nominal prices	1.13 (1.05, 1.22)	1.60 (1.46, 1.82)	1.61 (1.46, 1.82)
UK real prices	1.02 (0.93, 1.15)	1.61 (1.41, 1.87)	1.61 (1.41, 1.88)
US nominal prices	1.03 (0.93, 1.15)	1.03 (0.92, 1.18)	1.04 (0.91, 1.19)
US real prices	1.02 (0.93, 1.15)	0.98 (0.84, 1.15)	0.98 (0.84, 1.15)
(ii) <i>Autocorrelation (Bloomfield)</i>			
UK nominal prices	1.17 (1.06, 1.34)	1.14 (1.10, 1.36)	1.21 (1.11, 1.37)
UK real prices	0.96 (0.80, 1.18)	0.93 (0.82, 1.15)	0.92 (0.78, 1.17)
US nominal	1.00 (0.83, 1.22)	0.89 (0.78, 1.10)	0.88 (0.72, 1.11)
US real	0.98 (0.82, 1.21)	0.70 (0.58, 1.02)	0.67 (0.44, 1.02)

Notes: In bold, the selected models according to the deterministic terms using the t -values of the corresponding estimated coefficients. In parenthesis, the 95% band of non-rejection values of d . For the confidence bands, we use Robinson (1994).

unknown). We obtain estimates of d_L by using the Whittle function in the frequency domain (Dahlhaus, 1989). Together with the estimates, we also report the 95% confidence bands of the non-rejection values of d_L , using the parametric procedures outlined in Robinson (1994). See also, Gil-Alana and Robinson (1997). We mark in bold in Table 1 the selected cases according to the significance of the alternative deterministic terms. Note that Robinson's (1994) parametric approach does not require preliminary differencing. Thus, it allows us to test any real value d_L encompassing both stationary and non-stationary hypotheses.

The results of the estimation of the $I(d_L)$ model use the log-transformed house prices. In Table 1, under the assumption of no autocorrelation, the empirical results suggest that the house-price dynamics in the USA and UK differ substantially. We observe that the time trend does not achieve statistical significance for the UK nominal and real house prices nor for the US real house price. For the nominal house price in the USA, however, the time trend achieves significance.

We also observe that the estimates of d_L are much higher for the two UK house prices than for the US prices. For the UK, the estimated values of d_L equal 1.60 and

1.61, respectively, for the nominal and real prices, implying that we can decisively reject the unit-root null hypothesis in favour of $d_L > 1$ as the confidence bands in these cases all exceed one. We cannot reject, however, the unit-root null hypothesis for the US house prices, where the estimated values of d_L are 1.04 and 0.98, respectively, for the nominal and real prices.

When we permit autocorrelated disturbances the differences are somewhat reduced. The time trend becomes statistically significant in all four cases. The estimates of d_L are substantially reduced: 1.21 and 0.92 for the nominal and real UK series, and 0.88 and 0.67 for the corresponding US series. We cannot reject the unit-root hypotheses for the two US house prices and for the real UK price. For the UK nominal price, however, we still reject the unit-root hypothesis in favour of $d_L > 1$.

Given that the disparities in the results in Table 1 depend on whether we permit autocorrelation or not, we further estimate d_L using a semi-parametric approach, where we make no assumption on the structure of the error term. Table 2 displays the estimates of d_L based on the 'local' Whittle semi-parametric method (Robinson, 1995a, 1995b). The estimation requires the selection

Table 2. Estimates of d_L for the whole sample using a semi-parametric approach.

Series	11	12	13	14	15	16
UK nominal prices	1.418	1.339	1.292	1.331	1.352	1.397
UK real prices	0.925	0.937	0.890	0.892	0.907	0.926
US nominal prices	0.755	0.668	0.632	0.659	0.679	0.708
US real prices	0.500	0.500	0.500	0.522	0.577	0.502
Lower 5% $I(1)$	0.752	0.762	0.771	0.780	0.794	0.800
Upper 5% $I(1)$	1.247	1.237	1.228	1.219	1.212	1.205

Notes: In bold, evidence of unit roots at the 95% level.

of the bandwidth parameter. The table presents results for selected bandwidth values ($m = 11, 12, \dots, 16$), reported at the top.⁶ Bold type identifies evidence of unit roots. The 95% confidence bands for the unit-root hypothesis are reported at the bottom of the table. The semi-parametric estimates of d_L are generally robust across the bandwidth parameters, but lower than the corresponding parametric estimates. For the UK house prices, we find no evidence of mean reversion. For any reported value of the bandwidth parameter, we reject the unit-root hypothesis for the UK nominal house price in favour of the alternative of $d_L > 1$, but cannot reject the unit-root hypothesis for the UK real house prices for any reported value of the bandwidth parameter. By contrast, we reject the unit-root null hypothesis for the US real house price for any reported value of the bandwidth parameter and, except for the first value of the bandwidth parameter, also for the nominal US house price. The estimates of d_L for the US house prices are below one and we find mean reversion for almost any reported value of the bandwidth parameter. In general, we detect much less consistency between the parametric and semi-parametric estimates. This may indicate that the model is incorrectly specified. In particular, the UK estimates, more than the US estimates, may include an upward bias, since the model does not include the cyclical component.

Thus, the evidence based exclusively on the $I(d_L)$ model indicates some degree of heterogeneity between the US and the UK house price dynamics, although the results vary substantially depending on the methodology employed. The UK house prices are either unit-root processes or display orders of integration significantly above one. In contrast, the US house prices are unit-root process in one case and mean-reverting non-stationary processes in all other cases.

Results from the cyclical $I(d_C)$ model

Table 3 reports the whole sample estimates of the second model, the $I(d_C)$ model, which considers only the cyclical component of persistence. The high values of the d_L estimates in the $I(d_L)$ model lead us to estimate the $I(d_C)$ model using first differences of the logarithm of house prices (Panels A and C) and the mean-subtracted first differences (Panels B and D). As in the $I(d_L)$ model, we assume that the error term is $I(0)$, and consider, once more, the two cases of no autocorrelation (Panels A and B) and autocorrelation (Bloomfield-type) (Panels C and D). Little variation in the results exists across the two alternative assumptions on the error structure. Substantial differences in the cyclical component of persistence between the UK and the US house prices do exist. We observe that the cyclical component is much lower than the long-run

component in both the UK and the USA from Tables 1 and 2. For the UK, the estimates of d_C are positive and < 0.5 , indicating that the cyclical component of persistence in UK house prices is stationary but has ‘long-memory’ behaviour. In contrast, the US estimates of d_C are positive but not significantly different from zero, indicating that the cyclical component of persistence is stationary and displays ‘short-memory’ behaviour. Moreover, in the case of the UK, the cyclical component of persistence is much higher for nominal prices than for real prices. In the US case, instead, we see no significant differences.

We also observe in Table 3 that the housing cycle presents more variability in the USA than in the UK. The estimated periodicity (the value of j) ranges between five and six years for the UK and between five and eight years for the USA, which is consistent with the empirical literature on business cycles. These results are robust to changes in the assumptions of the error term and the treatment of the data.

Results from the $I(d_L, d_C)$ model

Finally, we examine the model given by equation (9), which is more general than the previous two specifications in the sense that it includes two fractional integration parameters, one at the zero (long-run) frequency and the other at the cyclical frequency. Table 4 focuses on white-noise errors (Panels A and B), as well as the autocorrelated (Bloomfield) case (Panels C and D).

The estimated periodicity (the value of j) ranges between five and six years for the UK and the USA, the US estimate is slightly smaller than the US estimate obtained by the $I(d_C)$ model, but is still consistent with the empirical literature on business cycles. We find striking differences between the house price dynamics of the USA and the UK as well as substantial similarities. For both the US and UK house prices, the estimates of d_L substantially exceed the estimate

Table 3. Estimated coefficients in (4) assuming white noise errors (Panels A and B) and autocorrelated (Bloomfield) errors (Panels C and D).

Panel A		
Series	J	d_C
UK nominal prices	6	0.42*
UK real prices	5	0.14*
US nominal prices	6	0.05
US real prices	7	0.01
Panel B		
Series	J	d_C
UK nominal prices	6	0.43*
UK real prices	5	0.14*
US nominal prices	8	0.04
US real prices	7	0.01
Panel C		
Series	J	d_C
UK nominal	6	0.41*
UK real	5	0.14*
US nominal	6	0.05
US real	6	0.05
Panel D		
Series	j	d_C
UK nominal	6	0.43*
UK real	5	0.14*
US nominal	5	0.02
US real	7	0.01

Notes: *Significance at the 95% level.

of d_C in all cases, implying that the long-run component plays a more important role than the cyclical component in explaining house price dynamics in the two countries. In both countries, the long-run component is < 1 and > 0.5 , suggesting that long-run house prices are non-stationary, but mean reverting. The long-run component of the UK is much higher than that of the USA, especially when we include the assumption of autocorrelation in the residuals, implying that UK house prices take longer to revert to the initial equilibrium. Significant

Table 4. Estimated coefficients in (9) assuming white noise errors (Panels A and B) and autocorrelated (Bloomfield) errors (Panels C and D).

Panel A			
Series	d_L	j	d_C
UK nominal	0.79	6	0.14*
UK real	0.86	4	0.03
US nominal	0.79	5	0.07
US real	0.80	5	0.09
Panel B			
Series	d_L	j	d_C
UK nominal	0.60	5	0.10*
UK real	0.65	4	0.01
US nominal	0.60	4	0.02
US real	0.60	4	0.03
Panel C			
Series	d_L	j	d_C
UK nominal	0.68	4	0.40*
UK real	0.86	4	0.03
US nominal	0.51	6	0.10
US real	0.52	5	0.09
Panel D			
Series	d_L	j	d_C
UK nominal	0.68	4	0.38*
UK real	0.80	5	0.04
US nominal	0.51	5	0.09
US real	0.50	5	0.08

Notes: *Significance at the 95% level.

differences between the UK and the USA also exist in the estimates of the cyclical component of persistence of house prices. For the UK, the estimates for the nominal series are positive and < 0.5 , indicating that the cyclical component of persistence in the UK nominal house prices is stationary, but has ‘long memory’. In contrast, for the UK real home price and for the US nominal and real home prices, the estimates are positive but not significantly different from zero, indicating that the cyclical component of

persistence is stationary, but has ‘short memory’. For the UK, the cyclical component is much higher for the nominal price than for the real price. For the USA, instead, no significant difference exists. Thus, the cyclical component is only relevant for the UK nominal house price; for the US nominal and real house prices and for the UK real house price, the $I(d_L)$ model sufficiently describes the persistence in the data.

An obvious but important caveat to these results is, however, in order. The analysis of historical data sets, such as the ones used in this work, is particularly vulnerable to the problem of structural change, which may limit the relevance of our conclusions. Housing markets in the USA and the UK have experienced remarkable political and economic reforms, such as financial deregulation and liberalisation, and technological advances, such as mortgage securitisation. Our estimates in the fourth section have ignored this problem and, consequently, may include bias due to the presence of structural breaks in the data. We attempt to deal with this issue in the next section.

Structural breaks and sub-sample results

This section addresses the issue of structural breaks in the data. As earlier argued, this is a relevant issue not only because of the historical breadth of the data, but also because fractional integration and structural breaks are intimately related to and easily confused with each other (Diebold and Inoue, 2001; Gil-Alana, 2008; Gouriéroux and Jasiak, 2001; Granger and Hyung, 2004; Sibbertsen, 2004; Smith, 2005, among others). Housing markets often experience large shocks that could create a structural break. The empirical literature provides evidence that structural changes can affect house price dynamics. Cook and Vaugas (2009) find

structural change in UK house prices and show that contrary to standard unit-root tests, smooth-transition threshold autoregressive tests reject the presence of a unit root in UK house prices. Canarella et al. (2012), in turn, find structural breaks in house prices in the USA.

Thus, to complete our analysis, we adopt the approach developed by Bai and Perron (2003) that estimates endogenously a number of potential breaks in the data along with their respective break dates. After identifying break dates, we re-estimate the fractional parameters in each sub-sample defined by the break dates.⁷ Another caveat, however, is in order. The estimation of multiple sub-samples corresponding to more than one break is constrained by the sample-size problem. Allowing for more break dates would produce sub-samples with a small number of observations, invalidating the analysis based on fractional integration. Thus, we present the results that define a dominant (i.e. main break), but do not exclude the possibility of other non-dominant breaks. Still, even allowing only one break, we cannot eliminate the sample-size problem. In particular, the sample size in the second sub-samples may lead to unreliable estimates and other estimation problems.

An interesting finding of this analysis is that while the house price swings of the last decade are dramatic, the greatest structural changes in the overall nominal and real price dynamics of the UK and the USA appear to occur much earlier and seem to match both macroeconomic shocks (for the UK) and specific political legislative outcomes (for the USA). Most importantly, the breaks are asynchronous, lending further credence to the view that the housing markets in the UK and the USA are not homogeneous in the sense that they do not share the same dynamics. We posit structural changes in the fractional integration coefficients of the models. Recent interest, however, focuses on

a decline in the post-war business cycles' volatilities of overall economic activity, that is, the 'Great Moderation' (Fang et al., 2008). Does this moderation reflect a sharp break or part of a longer trend to increasingly moderate cycles (Stock, 2003)? We follow the conventional interpretation of the structural break tests, which is standard in the literature (Bai and Perron, 2003). We cannot reject, however, the possibility (raised by one referee) that, at least for the UK, the break associates with changing cyclical turning points.

For the UK, the break dates occur at 1976 and 1983 for the real and nominal house prices, respectively. These break dates are consistent with some of the extant research. For example, Miles (2015) finds that, while large price swings exist in the 2000s, the 1980s exhibit sharp episodes of boom and bust. Zhang et al. (2017) using the Bai and Perron (2003) methodology identify statistically significant structural breaks at 1973, 1987 and 1997.

Moreover, these dates roughly associate with important national macroeconomic events, such as oil price shocks (1973–1974, 1978–1979), the banking crisis (1976), the deregulation of the housing market in the early 1980s, the deep recession of the early 1980s, the large escalation in interest rate and inflation in the late 1970s and early 1980s, and, most importantly, the Labour Party's incomes policies. Harold Wilson's government instituted the incomes policy in 1965, attempting to solve the inflation problem using wage/price controls. The policy ended with the election of Margaret Thatcher in 1979. The incomes policy regime exerted profound influences and is associated with significant macroeconomic structural breaks. Meenagh et al. (2009), for example, show that the incomes policy produced a structural change in inflation persistence in 1981. Additionally, major regime changes in the UK credit market may have

Table 5. Estimates of d_L for each sub-sample using a parametric method.

Series	First sub-sample			Second sub-sample		
	No terms	Intercept	T. trend	No terms	Intercept	T. trend
<i>(i) White noise</i>						
UK nom. prices	1.08 (0.99, 1.22)	1.60 (1.45, 1.85)	1.61 (1.46, 1.87)	0.93 (0.66, 1.26)	1.73 (1.41, 2.13)	1.65 (1.33, 2.06)
UK real prices	0.98 (0.89, 1.10)	1.66 (1.41, 1.98)	1.64 (1.41, 1.94)	0.90 (0.67, 1.20)	1.60 (1.22, 2.14)	1.64 (1.41, 2.23)
US nom. prices	1.01 (0.91, 1.15)	1.00 (0.86, 1.19)	1.00 (0.86, 1.19)	0.94 (0.75, 1.21)	1.32 (1.09, 1.56)	1.22 (1.07, 1.42)
US real prices	1.02 (0.91, 1.16)	0.98 (0.78, 1.19)	0.98 (0.81, 1.19)	0.95 (0.79, 1.18)	0.72 (0.56, 1.27)	0.86 (0.58, 1.25)
<i>(ii) Autocorrelation (Bloomfield)</i>						
UK nom. prices	1.03 (0.88, 1.25)	1.18 (1.07, 1.35)	1.20 (1.06, 1.38)	0.56 (0.19, 1.36)	0.66 (0.27, 1.93)	0.68 (-0.26, 1.72)
UK real prices	1.02 (0.82, 1.25)	0.89 (0.68, 1.20)	0.89 (0.68, 1.20)	0.67 (0.13, 1.34)	0.61 (0.40, 1.07)	0.19 (-0.42, 1.10)
US nom. prices	0.99 (0.80, 1.25)	0.74 (0.46, 1.07)	0.72 (0.48, 1.07)	0.80 (0.29, 1.26)	1.42 (0.34, 1.96)	1.26 (0.91, 1.84)
US real prices	0.98 (0.77, 1.29)	0.61 (0.43, 1.14)	0.63 (0.34, 1.10)	0.85 (0.55, 1.25)	0.49 (0.37, 0.63)	0.22 (-0.01, 0.75)

Notes: In bold the selected models according to the deterministic terms using the t -values of the corresponding estimated coefficients. For the confidence bands, we use Robinson (1994).

contributed to the structural change. Muellbauer and Murphy (1997) argue that deregulation of the housing market in the 1980s made mortgages more easily available than in the 1960s and 1970s, when credit was often subject to rationing.

For the USA, the break dates occur at 1955 and 1972 for the real and nominal house prices, respectively. These dates are roughly associated with the major post-Second World War developments in US housing policy, which include the National Housing Act of 1949,⁸ which expanded the federal role in mortgage insurance, the 1955 Amendment to the National Housing Act of 1949, the Housing and Urban Development Act of 1965, which produced major revision of the US federal housing policy and instituted several major expansions in federal housing programmes as part of Johnson's 'Great Society' programmes, and the Housing and Urban Development Act of 1968 which created the Government

National Mortgage Association (commonly referred to as GinnieMae). Interestingly, we do not find a dominant break associated with the Great Moderation, which literature documents as a substantial reduction of volatility in major US macroeconomic time series since the 1980s.

In this section, we present the sub-sample results using the same three fractional integration models that we considered for the whole sample. Each sub-sample is uniquely defined by the corresponding break date, and this results in a different sample size for each sub-sample. Our analysis of the sub-samples, however, is incomplete, since the limited length of the data does not permit estimation of the second sub-sample in the third model.

Sub-sample results from the long-run $I(d_L)$ model

Table 5 reports the results of the estimation of the long-run $I(d_L)$ model for the two sub-

Table 6. Estimated coefficients in (4) for each subsample.

Series	First sub-sample		Second sub-sample	
	j	d_C	j	d_C
<i>(i) Original data</i>				
UK nominal	6	0.48*	5	-0.09
UK real	5	-0.18	5	-0.22
US nominal	5	-0.19	4	0.31
US real	4	0.32*	4	0.41*
<i>(ii) Mean subtracted data</i>				
UK nominal	6	0.54*	6	0.02
UK real	6	0.47*	5	0.30
US nominal	6	0.53*	6	-0.08
US real	6	0.48*	6	0.31

Notes: *Significance at the 95% level.

samples and the three specifications of the deterministic component, under the two cases of uncorrelated (top panel) and autocorrelated errors (bottom panel).

Under the assumption of white-noise disturbances, we observe that the trend is not required for the first sub-sample and is only required for the US data in the second sub-sample. The estimates, however, are similar independently of the inclusion or exclusion of the trend. The orders of fractional integration are significantly higher than one for the UK data in both the first and the second sub-samples. The unit-root hypothesis is accepted for the US data in the first sub-sample; in the second sub-sample, however, the orders of fractional integration exceed one for the nominal series and do not differ from one for the real series. Thus, the results generally mirror the whole sample estimates, and we find no evidence of mean reversion in any case.

In contrast, under the assumption of autocorrelated (Bloomfield) disturbances, the trend is significant in several cases, especially in the second sub-samples. We observe a reduction of the estimate of d_L when we move from the first sub-sample to the second, with the exception of the US nominal

series. We cannot reject the unit-root hypothesis in any case, except for US real prices in the second sub-sample. Nevertheless, we do not observe significant differences in the orders of integration across the sub-samples, regardless of the assumptions about the disturbances.

Sub-sample results from the cyclical $I(d_C)$ model

Table 6 reports the sub-sample results for the cyclical $I(d_C)$ model. The length of the cycles lies between four and six years in all cases, which is consistent with the whole sample estimates. Evidence of substantial differences in the estimates of d_C exists, however, between the whole sample and the two sub-samples. In Panel A, the estimates lie between zero and 0.5 in the first sub-sample for the UK nominal series and the US real series, suggesting cyclical mean reversion. In the second sub-sample, only the estimate for the US real series is significant and < 0.5 . In Panel B, all estimates are significant in the first sub-sample, suggesting high cyclical persistence and mean reversion. In the second sub-sample, the estimates of both the real series are > 0 but are not significant. This lack of significance, however, may reflect the smaller size of the second sample, which likely produces large confidence intervals. Overall, however, the results of the estimates of the cyclical component in the two sub-samples are not consistent with the corresponding results of the whole sample.

Sub-sample results from the $I(d_L, d_C)$ model

Finally, Table 7 displays the results for the $I(d_L, d_C)$ model, which includes both orders of integration, zero and the cyclical one, once more, for the two cases of uncorrelated errors (Panels A and B) and autocorrelated (Bloomfield) errors (Panels C and D). In this

Table 7. Estimated coefficients in (9) assuming white noise errors (Panels A and B) and autocorrelated (Bloomfield) errors (Panels C and D) with the first sub-samples.

Panel A: Original data			
	d_L	j	d_C
UK nominal	1.29	7	0.22*
UK real	1.16	4	-0.03
US nominal	0.59	5	-0.07
US real	1.13	5	-0.04
Panel B: Mean subtracted data			
	d_L	j	d_C
UK nominal	0.78	5	0.24*
UK real	0.82	4	0.03
US nominal	0.50	4	0.00
US real	0.48	4	-0.03
Panel C: Original data			
	d_L	j	d_C
UK nominal	1.28	4	0.33*
UK real	1.04	4	-0.03
US nominal	0.66	6	0.06
US real	0.71	4	-0.04
Panel D: Mean subtracted data			
	d_L	j	d_C
UK nominal	0.58	4	0.19*
UK real	0.66	5	-0.01
US nominal	0.47	5	0.04
US real	0.61	4	-0.08

Notes: *Significance at the 95% level.

case, however, we only report the estimates for the first sub-samples, since the number of observations in the second sub-samples was not sufficient to guarantee significant results. Panels A and B of Table 7 assume white-noise disturbances. The number of periods per cycle varies from four to seven years in the UK data, and from four to five in the US data. In Panel A, the estimates of d_L exceed one in all series except the US nominal data, and the estimate of d_C is significantly positive and < 0.5 only for the UK nominal

series, and insignificant for the remaining series. In Panel B, in contrast, all the estimates of d_L exhibit mean reversion, but only the US real data exhibit non-stationarity.

Panels C and D of Table 7 assume autocorrelated (Bloomfield) disturbances. The error structure does not appear to affect the periodicity of the series, as the number of years per cycle varies from four to six in the UK and from four to five in the USA. In Panel C, the estimates of d_L exceed one in the UK data and are < 1 in the US data, implying non-stationarity and non-mean reversion for the UK data and non-stationarity and mean reversion for the US data. The estimate of d_C is significantly positive and < 0.5 only for the UK nominal series. In Panel D, all the estimates of d_L fall below one, suggesting mean reversion, but only the estimate for the US nominal series falls below 0.5. As in Panel A, the estimate of d_C is significantly positive and < 0.5 only for the UK nominal data.

In conclusion, the results for the first sub-sample generally mirror those of the whole of the samples. We cannot conclude, however, that we observe significant differences when we account for structural breaks because we lack evidence from the second sub-samples.

Conclusions

In the past decade, the US and UK housing markets have experienced significant housing price booms, followed by sharp declines. The temptation exists, because of the apparent similarities, namely the existence of sub-prime lending and the use of mortgage backed securities, to conclude that the US and UK markets mirror each other and share the same experience. From a historical viewpoint, the differences between the two markets are just as important as their similarities. Most literature on housing markets generally accepts the idea that house prices

are non-stationary. In this literature, however, house prices are specified in a stochastic model that assumes only the presence of a pole at the zero frequency. Such models only describe the long-run persistence of house prices. In this paper, we suggest that such models may be mis-specified, since they fail to account for the cyclical component of persistence in house prices. In this paper, we provide a new and unique look at the dynamic and persistence structure of historical house prices in the USA and the UK, using fractional integration techniques not previously applied to housing markets. We suggest that the US and the UK historical house prices may conform to a stochastic process that includes two poles in the spectrum: one at the zero frequency, corresponding to the long-run dependency of the series, and another away from the zero frequency, corresponding to the cyclical dependency of the series.

We use annual data from 1830 to 2016 for the USA and 1845 to 2016 for the UK, which provides a much longer perspective on the behaviour of house prices than commonly implemented in the literature, where most empirical work uses data starting from the 1980s or later. We consider three fractional integration models: (1) a standard $I(d_L)$ model with a pole at the zero frequency, which captures only the long-run component of persistence; (2) a cyclical $I(d_C)$ model that incorporates a pole at a non-zero frequency and captures only the cyclical component of persistence; and (3) the composite $I(d_L, d_C)$ model that incorporates both poles and captures simultaneously the component associated with the long-run trend and the component associated with the cycle.

We find that each country exhibits rich house-price dynamics, at the level of the whole sample and sub-samples, with the break dates estimated using the Bai and Perron (2003) methodology. The sub-sample

analysis is necessary not only because of the historical breadth of the data, but also because fractional integration and structural breaks are intertwined issues. Interestingly, although the house-price swings of the last decade are dramatic, the greatest structural changes in the overall nominal and real price dynamics of the UK and the USA appear to occur much earlier, in the late 1970s and early 1980s in the UK, and in the mid-1950s and early 1970s in the USA. This asynchronous pattern of the breaks indicates heterogeneity in house-price dynamics of the two countries and a sign that national rather than global events played an important role. Sub-sample estimation, however, presents some unique challenges in a fractionally integrated setting, resulting from the small sample size problem. In particular, the sub-sample analysis is only partial in the third model as the sample size after the break is not large enough to produce meaningful estimates. We find, however, that structural breaks affect the estimates of the long-run and cyclical components.

For the whole sample, we find convincing evidence that in the UK housing markets, nominal house prices incorporate two distinct poles in house-price dynamics, at the zero (long-run trend) and non-zero (cyclical) frequencies. In contrast, we fail to find evidence of cyclical persistence for the USA and the real house price in the UK. In contrast, the cyclical model provides evidence that significant cyclical persistence exists in the first sub-sample for both the UK and the USA.

An important result, common to the whole sample and the sub-samples, is that the long-run component of persistence plays a greater role than the cyclical component in explaining the dynamics of house prices in both countries. In no instance, however, are shocks permanent. These findings have substantial implications for policy decisions. Shocks affecting the long-run component will persist for a long

time, while those affecting the cyclical component will not. Thus, policymakers should adopt stronger policies with respect to long-run house-price movements to create an environment whereby housing markets can readily revert to their original trends.

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Notes

1. Evidence of the strong link between housing markets and the economy is not only provided by recent economic history, but also by the entire post-war era (Holly and Jones, 1997) and even pre-dates the Industrial Revolution (Eichholtz et al., 2012).
2. Álvarez et al. (2010) for the Euro area, Ferrara and Vigna (2010) for France, and Álvarez and Cabrero (2010) for Spain provide empirical evidence on the leading nature of housing markets and house prices with respect to the business cycles.
3. Several studies document the relevance of the cyclical structure of many economic data. See Gray et al. (1989).
4. The parameter μ is defined as $\cos w$, where $w = 2\pi/r$, r indicating the number of time periods per cycle.
5. The periodograms of the log-transformed data show the highest values in the close vicinity of the zero frequency, while the periodograms of the first differences on the log-transformed data display the highest values at a non-zero frequency, providing evidence of cyclical patterns, with the exception of the UK log-transformed nominal price.
6. The choice of the bandwidth (m) shows the trade-off between bias and variance: the asymptotic variance and the bias decrease and increase, respectively, with m .

7. We also apply the methodology developed by Gil-Alana (2008). Interestingly, the results are exactly the same.
8. The National Housing Act was the US government's response to the severe shortage of housing in the post-Second World War America, when almost 11 million men and women left the armed services. The Act is best remembered for its declaration that every American deserves a 'decent home and a suitable living environment', which helped millions of Americans realise the 'dream' of homeownership. The Levitt brothers' approach to building homes put the American dream within grasp of the middle class family. By the end of the 1950s, no less than 15 million homes were under construction nationwide.

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