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# Aggregate and Regional House Price to Earnings Ratio Dynamics in the UK

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**Abstract**

This paper examines the time series properties of house price to earnings ratio (HPER) in the UK using aggregate and regional data. Specifically, we utilize a series of unit root tests to examine the null hypothesis of nonstationary HPERs. These include linear tests as well as a nonlinear test and also a test which accounts for abrupt structural change. The results are overall only weakly supportive of stationarity in HPERs. This implies that house prices may permanently diverge from earnings.

*JEL classification*: C22; R31.
*Keywords*: House price to earnings ratio; Unit root tests; Structural breaks; ESTAR models.

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"*How much further will house prices fall? The best guide is the ratio between average earnings and average house prices. This is a measure of affordability.*" (David Blanchflower, 2009)

## 1. Introduction

Recent developments in the property market have reminded investors and policymakers that house prices can not only be very volatile but also significantly impact upon banks' financial health and households' finances. In particular, house prices have declined significantly after reaching a historical high at 2007. The lower property valuations have had important effects across the financial spectrum with the value of bonds and derivative products that were ultimately backed by property falling in tandem. Moreover, since a great amount of an average household's wealth is held in property[1], households are not expected to significantly increase their level of consumption until a strong recovery in house prices materialises.[2] Property market dynamics also play a key role within the balance-sheet monetary policy transmission channel (Bernanke and Gertler, 1989) since property is typically the collateral that firms and households use in order to secure borrowing.

The experience of a significant crisis in the property market followed by a recessionary episode is not unknown to the UK. The late 1980s-early 1990s featured such a combination of events. It is not surprising then that house prices in the UK have been extensively investigated with a large number of previous studies using aggregate and disaggregate data to examine which fundamentals underlie property valuation and

[1] As Goodhart and Hofmann (2000) point out, housing assets account for a greater share of household wealth than equity in most G7 countries.

[2] Note, however, that the relationship between house prices and consumption may not be stable over time. Farlow (2005) argues that consumption has been much less responsive to house price changes during the most recent house price boom than in the past. He argues that credit conditions have become significantly less constrained and that the role of the collateral channel has weakened. Therefore, given the lower shadow price of credit, "an increase in housing equity would not be expected to have as large an impact on consumption at the margin as it would have had in the past" (Farlow, 2005; p. 10). Benito *et al.* (2006) also argue that the role of house prices in loosening spending constraints has weakened, thereby reducing the strength of the collateral channel.

to test for bubbles (see e.g. Cameron *et al.*, 2006). This paper contributes to the literature on the UK property market by investigating the time series properties of the house price to earnings ratio (HPER) using both aggregate and regional data over the time period 1983-2009. We should stress that 'earnings' in our HPER variable correspond to the income of employees, rather than 'rents', so that the HPER is essentially a housing affordability measure.[3] The central question under investigation is whether the HPER is mean-reverting, so that prices do not permanently diverge from earnings. If the HPER follows a stationary mean-reverting process, house prices will fall if current prices are high as compared to earnings, and vice versa.[4]

As Muellbauer and Murphy (2008, p.5) explain, "The deviation of prices from long-run fundamentals is then the 'bubble-burster'". Specifically, house prices may rise due to a series of positive shocks to fundamentals such as earnings. The expectation of further appreciation leads to overvaluation, but in due course the realisation that the improvement in fundamentals has been outpaced by house prices increases, leads to a slowdown in the rate of appreciation. On the other hand, if the HPER is nonstationary then a shock, due e.g. to the global financial crisis, would have permanent effects and it would be therefore unlikely for the series to return to its initial level. Hence, investigating the stationarity property of the HPER will shed some important light on the long-run outlook for the property market.

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[3] The house price to income ratio is one the earliest and most widely used measures of housing affordability (Andre, 2010).

[4] In the asset pricing literature a large number of studies use the stock market price to earnings ratio in an effort to predict future movements in stock prices. For example, Campbell and Shiller (1998) argued that the high stock price to earnings ratios observed in the late 1990s implied that the stock market was overvalued and that stock prices would decline in the future in order to bring the prices closer to the underlying companies' earnings. The measure of earnings used here is different since it corresponds to the labor income of individuals, as opposed to the income received from renting the property. We do not compute price to rent ratios since, to the best of our knowledge, data on rents is not available over the full sample period that we examine (1983-2009), at matching frequency to the one we use (quarterly), and across all the UK regions. This issue is also highlighted in Muellbauer and Murphy (2008) who argue that, due to its small size, the private rented sector in the UK is not a very reliable proxy of the private housing sector as a whole, and the publicly available rent data is often of poor quality. Nevertheless, the idea of mean-reversion once a high level has been reached for the price to earnings ratio is similar in both stock market and property market applications.

We test for mean-reversion in the HPER of the UK as whole as well as regional ratios using unit root tests. Previous UK literature has utilised unit root tests in testing for the 'ripple effect', whereby shocks to UK house prices first hit London and the South East of England before spreading to other regions (see e.g. Cook, 2005), but no study, to the best of our knowledge, has applied the unit root testing framework to HPERs.[5] We start our empirical investigation by utilising standard linear unit root tests. We then consider a unit root test which allows for shifts in the HPER in an effort to account for the potential loss of power of standard no-structural breaks tests in the presence of such breaks.

Visual investigation of the HPER series indicates boom-bust cycles in the context of which reversals in the ratio do not appear to take place abruptly (see also Black *et al.*, 2006; Andre, 2010). This is consistent with the view of Muellbauer and Murphy (2008) that systematic mispricing can persist and therefore the movement from the property market's peak to trough (and vice versa) make take some considerable time to materialise. [6] Hence, this paper further contributes to the literature by presenting empirical evidence which explicitly allows for the possibility that HPERs can be characterized by a smooth nonlinear mean reverting process, captured by the nonlinear unit root test of Kapetanios *et al.* (2003). This process may exhibit near unit root behaviour in a specific range, so that HPERs may appear nonstationary from the perspective of test procedures, which specify a linear nonstationary process as the null hypothesis.

The rest of the paper is structured as follows. Section 2 provides a selective review of previous related research. Section 3 describes and discusses the data. Sections

[5] As Cook (2005) explains, the diffusion of changes in house prices that the 'ripple effect' implies is consistent with a constant long-run ratio of regional to aggregate house prices. He finds that the aforementioned ratio is stationary for a number of regions thereby supporting the notion of the 'ripple effect'.

[6] See also Cook (2006) for an analysis of regional UK house prices with models that allow for asymmetric behaviour. His main finding is that cyclical peaks in house prices are greater in magnitude than troughs.

4, 5 and 6 present respectively the linear, two-break and nonlinear unit root tests and results. Section 7 concludes.

## 2. A selective review of previous research

In the existing literature on UK house prices a large number of studies have utilised aggregate and/or regional data to examine the relationship between house and fundamentals and test for the presence of bubbles. The set of fundamentals typically includes variables such as income, housing stock, demography, credit availability and interest rates, see e.g. Muellbauer and Murphy (1997, 2008) and Cameron *et al*. (2006).[7] In terms of the relationship between house prices and income, Muellbauer and Murphy (2008) find that the long-run elasticity of house prices with respect to non-property income relative to the housing stock is positive and its value exceeds the value of one. They argue that most of the increases in real house prices since 1997 can be attributed to rise in the average real income per household. Cameron *et al*. (2006) utilise a dynamic panel data model of UK regional house prices over the period 1972-2003 and find that income dynamics are important determinants of house prices, especially in London and the South East. They also highlight that their evidence cannot rule out bubble behaviour in UK house prices in the late 1980s.

Black *et al*. (1996) analyse the relationship between house prices and income using a present value approach, motivated from the stock market literature as in Campbell and Shiller (1998), which allows them to compute the fundamental value of housing based on the present value of real disposable income (see also Case and Shiller, 2003).[8] They test for cointegration between UK house prices and income and also examine the stationarity properties of their ratio. Using standard linear unit root tests

[7] See Cameron *et al*. (2006) for a review of the UK regional house price literature.

[8] Note that, as we explain in footnote 4, the theoretically appropriate variable for the present value approach should have been rents from housing rather than disposable income.

they provide very weak evidence for stationarity of the ratio[9]; on the other hand, stronger evidence is obtained for cointegration between house prices and income, which suggests that the UK housing market is not characterised by explosive rational bubbles. Black *et al*. (1996) point out that the results from the unit root test that they apply on the house price to income ratio may be influenced by long swings in the series (see also Andre, 2010). This type of time series behaviour is in line with the view of Muellbauer and Murphy (2008) that systematic mispricing in the property market can be long-lasting and therefore swings in the ratio are rather gradual, and will be accounted for in our empirical estimations.

## 3. Data

Data were collected from Halifax for the UK as a whole and twelve regions: North, Yorkshire and the Humberside, North West, East Midlands, West Midlands, East Anglia, South West, South East, Greater London, Wales, Scotland and Northern Ireland. The HPER variable provided by Halifax is measured as the ratio of the Halifax standardised average house price (all houses, all buyers) to average earnings for full-time male employees.[10] The variable is revised to reflect new data in the Annual Survey of Hours and Earnings. [11] The sample period is 1983Q2 – 2009Q1 providing us with 104 quarterly observations.

[INSERT TABLE 1 HERE]

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9 See also Girouard *et al*. (2006) for international evidence supporting HPER nonstationary.

10 As explained in the appendix that contains the Halifax house price index technical details (www.lloydsbankinggroup.com/media/word/HPI/13.08.09TechDetails.doc) the house price index provided by Halifax is standardised to account for differences in various quantitative and qualitative characteristics that are related to the physical attributes of the houses themselves or their locations, e.g. purchase price, age of the property, number of rooms, garden, etc. The methodology employed to obtain the typical house price over time on a like-for-like basis is based on the “hedonic” approach to price measurement.

11 Note that national (regional) earnings are used for the calculation of the national (regional) HPER. It should be pointed out that the use of male-only earnings in the calculation of the HPER may result to an overestimation of the degree of non-affordability since it is often the case that both partners in a couple work to meet mortgage repayments. This is a constraint imposed by the Halifax dataset, which does not account for female earnings in the HPER measure. Nevertheless, we believe the benefits arising from the use of the Halifax dataset in terms of early starting point and comprehensiveness and harmonization across regions, outweigh those costs. We would like to thank an anonymous referee for raising this point.

Table 1 provides some descriptive statistics. The average HPER for the UK as a whole is 4.02, while in the regions it ranges from 3.34 in Yorkshire and the Humberside to 5.18 in the South East. The data indicates a north-south divide with housing generally becoming less affordable as we move from the northern regions to the southern ones. This heterogeneity also manifests itself in affordability differences across the UK's regions. In particular, the average HPER in the nine English regions is 4.11, as compared to 3.76 in Wales, 3.52 in Scotland and 3.35 in Northern Ireland. The Northern Irish series is the most volatile in the sample, followed by the South West. The least volatile HPER is observed in Scotland.

Figure 1 plots the HPERs. Overall, both the aggregate and the regional UK ratios appear to be characterised by cyclical behaviour. Two major boom-bust episodes are apparent.[12] The horizon of the cycles is relatively long with a correction, involving a large change in the underlying slope, occurring at the peak and trough of the cycle. For the UK as a whole, the first HPER boom-bust cycle commenced at the mid-1980s, reaching a peak (housing affordability low-point) at 1989Q2 and a trough in 1995Q4. Housing was quite affordable throughout the mid-to-late-1990s, setting the scene for the second period of expansion which commenced early in the new millennium, with the ratio peaking in 2007Q3.

[INSERT FIGURE 1 HERE]

As the results in Table 2 indicate, the regional evidence is consistent with the aggregate UK HPER dynamics with the first housing affordability low-point reached around 1988-1990 and the second in 2007.[13] In six regions (East Midlands, East Anglia, South West, South East, Greater London and Wales) housing was most affordable

[12] The exception is Northern Ireland, where the HPER remained bellow its long-run average value throughout the 1980s and the 1990s, and starting exhibiting large increases only since around 2003, reaching a maximum of 8.59 at 2007Q2. Another dimension of regional heterogeneity that emerges from Figure 1 is that in East Anglia and the South East the second upswing in the HPER peaks at a lower level than does the peak of the previous cycle, while in all other regions the opposite pattern prevails.
[13] It is interesting to notice that in most cases the HPER was higher during the peak of 2007 as compared to the late-1980s peak. Only in East Anglia, South East and Scotland, the opposite holds.

around the mid-1990s, in line with the UK as a whole. In five regions (North, Yorkshire and the Humberside, North West, West Midlands and Scotland), though, the housing affordability highpoint was reached later, around 1999-2001.

[INSERT TABLE 2 HERE]

## 4. Linear unit root tests

### 4.1 ADF unit root test

The standard linear ADF test (Dickey and Fuller, 1979; Said and Dickey, 1984) uses the following regression model to test the stationarity of the HPER:

$$\Delta he_t = \gamma_0 + \gamma he_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta he_{t-i} + \varepsilon_t \qquad (1)$$

where $he_t$ denotes the log of the house price to earnings ratio at time period $t$, the $\gamma's$ are constants and $\varepsilon_t$ is a random disturbance term: $\{\varepsilon_t\} \sim iid(0, \sigma_\varepsilon^2)$. The terms in $\Delta he_{t-i}$ are included to remove any serial correlation in $\varepsilon_t$.[14] The null hypothesis of unit root is consistent with the notion of an infinitely persistent HPER. That is, following a shock, due e.g. to the global financial crisis, the HPER will be unlikely to return to its initial level. In other words, a temporary shock could have permanent effects. Rejecting the null hypothesis requires the estimates of $\gamma$ to be negative and significantly different from zero.

The ADF results can be seen in Table 3, columns 2-4. We observe that the unit root null hypothesis is not rejected in all cases. This finding is robust to the manner in which the deterministic component of the ADF regression model is specified. Thus, linear ADF tests provide evidence for unit root behavior in both the aggregate and the regional UK HPERs. Our ADF-based finding that the national HPER is nonstationary at

[14] Note that we also consider a specification where a trend is also added to the set of deterministic components.

the 5% level of significance is in line with previous evidence by Black *et al*. (2006).

In the following set of results we will examine whether the non-rejection of the null hypothesis of a unit root in the HPER is related to the testing procedure employed. Particularly, we will use more recently developed unit root tests which overcome some of the deficiencies associated with the ADF test, such as low power and not accounting for structural change.

[INSERT TABLE 3 HERE]

### 4.2 Ng Perron unit root test

The Ng and Perron (2001) $MZ_\alpha$ test modifies the Phillips and Perron (1988) $Z_\alpha$ test in a number of ways in order to increase the test's size and power. This testing procedure ensures that non-rejections of the null-unit root are not due to a low probability of rejecting a false null hypothesis, while rejections are not related to size distortions. The test statistic is defined as[15]:

$$MZ_a = \left[T^{-1}he_T{}^2 - s_{AR}^2\right]\left[2T^{-2}\sum_{t=1}^{T} he_{t-1}{}^2\right]^{-1} \quad (2)$$

where $t$ = 1…T , $s_{AR}^2 = \hat{\sigma}_k^2 / [1-\hat{\gamma}(1)]^2$ is an autoregressive estimate of the spectral density at frequency zero of $\upsilon_t = \theta(L)\varepsilon_t = \sum_{j=0}^{\infty}\theta_j \varepsilon_{t-j}$ with $\sum_{j=0}^{\infty} j\left|\theta_j\right| < \infty$; $\hat{\gamma}(1) = \sum_{i=1}^{k}\hat{\gamma}_i$ and $\hat{\sigma}_k^2 = (T-k)^{-1}\sum_{t=k+1}^{T}\hat{\varepsilon}_t^2$ are calculated using the OLS estimates from Eq. (1). Ng and Perron (2001) employ the local-to-unity GLS detrending procedure in order to benefit from increased power. They also suggest that the autoregressive truncation lag, $k$, should be chosen using the Modified Akaike Information Criterion to avoid size distortions while maintaining power.

The Ng Perron linear unit root test results are presented in columns five and six of Table 3. In contrast to the ADF findings, the HPER for the UK as a whole is now

[15] The test statistic corresponds to the case where the variable into consideration $(he_t)$ contains no deterministic term. If we allow for a constant, or constant and trend, then $he_{t-1}$ and $he_T$ in Eq. (2) should be replaced by their detrended counterparts.

stationary. With constant only, the unit root null hypothesis is still not rejected though in seven out of twelve regions. When, in addition to the constant, a linear trend is incorporated the unit root evidence in the regional data becomes quite prevalent, since East Midlands is the only region where the HPER appears stationary. Thus, while the Ng Perron unit root test provides some evidence for regional stationarity, overall, both linear unit root tests suggest that the HPERs follow a unit root process in the majority of the regions under investigation.

This finding may be related to the fact that the HPERs exhibit structural shifts which are not accounted for in the Ng Perron and ADF unit root tests. In particular, visual inspection of the HPER series in Figure 1 indicates the presence of boom-bust cycles whereby upward and downward trends are broken. Thus, in the next section we will use a unit root testing framework which allows for structural change in order to examine the robustness of the findings from the aforementioned tests.

## 5. Two-break unit root test

The standard no-breaks unit root tests are subject to the drawbacks of low-power and biases in the presence of structural breaks, with the need to determine the breaks endogenously being emphasised in the literature (see e.g. Perron, 1997). The endogenous two-break minimum LM unit-root test of Lee and Strazicich (2003) counterbalances the potential loss of power of tests that ignore more than one break. The Lee and Strazicich test includes breaks under both the null and the alternative hypotheses, with rejections of the null unambiguously implying trend stationarity.[16] Allowing for breaks in the form of two shifts in the level of HPER, the null and alternative hypotheses are:

$$he_t = \mu_0 + d_1 B_{1t} + d_2 B_{2t} + he_{t-1} + \upsilon_{1t} \qquad \textit{Null} \qquad (3)$$

[16] Structural breaks under the unit root null can be interpreted as large permanent shocks or outliers.

$$he_t = \mu_1 + \gamma t + d_1 D_{1t} + d_2 D_{2t} + \upsilon_{2t} \qquad \textit{Alternative} \qquad (4)$$

where the error terms ($\upsilon_{1t}, \upsilon_{2t}$) are stationary processes; $B_{jt} = 1$ for $t = T_{bj} + 1$ ($j = 1,2$) and 0 otherwise; $D_{jt} = 1$ for $t \geq T_{bj} + 1$ ($j$=1,2) and 0 otherwise. An LM score principle is used to compute the Lee and Strazicich (2003) unit root test statistic based on the following regression model:

$$\Delta he_t = \delta' \Delta Z_t + \varphi \tilde{S}_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta \tilde{S}_{t-i} + u_t \qquad (5)$$

where $Z_t = [1, t, D_{1t}, D_{2t}]'$, $\tilde{S}_t = he_t - \tilde{\psi}_x - Z_t \tilde{\delta}$; $t$ = 2,…,$T$; $\tilde{\delta}$ are coefficients in the regression of $\Delta he_t$ on $\Delta Z_t$; $\tilde{\psi}_x = he_1 - Z_1 \tilde{\delta}$, where $he_1$ and $Z_1$ denote the first observations of $he_t$ and $Z_t$, respectively, and $\Delta \tilde{S}_{t-i}$ terms ($i$ = 1,…,$k$) are included to account for serial correlation. We can consequently test the unit root null hypothesis by examining the t-statistic ($\tilde{\tau}$) associated with $\phi = 0$.

[INSERT TABLE 4 HERE]

Table 4 contains the results from Lee and Strazicich's two-break unit root test. The estimated break dates suggest that in most cases at least of one of the shifts took place around 1989-1991, when the first housing affordability low-point (HPER first peak) was reached. The second estimated break date in the model with shifts in the constant and the trend of the HPER captures the beginning of the new millennium boom. On the basis of the time-series behaviour of the HPERs in Figure 1, it can be argued that the broken trends model may be more appropriate. Focusing on this set of our results, it seems that the second boom in the HPER ratio commenced around 2000-2001 in London and the South of England. This boom propagated to the northern regions and Wales with around a year delay, while Scotland and Northern Ireland lagged by more two than years. The timing of events that emerges from our structural break analysis is in line with the 'ripple effect'.

The results in Table 4 suggest that the unit root null hypothesis can only be rejected in Northern Ireland when the model with breaks in the intercept and the trend is used. For the rest of the regional data, as well as the aggregate UK ratio, the two-break unit root test provides evidence in support of nonstationary behaviour. Thus, our analysis indicates that accounting for structural change within the framework of Lee and Strazicich cannot overturn the unit root evidence obtained by the ADF unit root test.

The dummy variable approach upon which the Lee and Strazicich test is based assumes the transition from one regime to another takes place abruptly. In the next section we will use a model whereby structural change takes place in a non-abrupt manner.

## 6. Nonlinear unit root test

Failure to reject nonstationarity using the ADF and the Ng-Perron tests may be the result of lack of power of linear unit root tests if the true data generating process is nonlinear. Furthermore, the Lee and Strazicich test requires that the breaks, captured through the use of time dummies, are sharp. However, as the evidence in Figure 1 indicates, reversals in the HPER appear to not to be very abrupt. For instance, it took the aggregate UK ratio six years to move from the 1989-peak to the 1995-though. In order to take to take this property of the data into account, a unit root test will be utilised which allows for structural change to take place in a smooth, rather than abrupt, manner. This is the main novelty of the nonlinear approach that we utilise in this section. In particular, Kapetanios *et al*. (2003) developed a test where the null hypothesis of a unit root is tested against an alternative of nonlinear Exponential Smooth Transition Autoregressive (ESTAR) process, which is globally mean reverting. The ESTAR model assumes that the adjustment of the HPER towards its average value is characterized by a symmetric nonlinear process:

$$he_t = \beta he_{t-1} + \delta he_{t-1}\left(1 - \exp[-\alpha he_{t-1}{}^2]\right) + u_t \quad (6)$$

Under the null-non stationarity, $\beta = 1$ and $a = 0$, the HPER follows a random walk. Computing a first-order Taylor series approximation to (6) under the null and allowing for serial correlation in $u_t$, the following auxiliary regression model can be obtained (see Kapetanios *et al.*, 2003):

$$\Delta he_t = \gamma he_{t-1}{}^3 + \sum_{i=1}^{k} \gamma_i \Delta he_{t-i} + v_t \quad (7)$$

where $v_t$ is the error term and the other variables are defined as previously. As Kapetanios *et al.* (2003) explain, if the raw data exhibits an intercept or trend, then the original series must be replaced by its demeaned or detrended counterpart. The unit root null hypothesis in equation (7) is that $\gamma = 0$.

The nonlinear unit root test results are presented in Table 5. Overall, the results are not strongly in favour of nonlinear mean reversion. Specifically, the unit root null hypothesis can only be rejected at the 10% level of significance for the UK as a whole when demeaned and/or detrended data are used. Furthermore, only a small number of regions appear to follow a stationary smooth transition process with the null being rejected in around 30% of the total cases. Thus, the findings from the nonlinear unit root test are not dissimilar to those from the linear no-break and the two-break unit root tests in that the evidence does not overall support the presence of stationary HPERs. Nevertheless, we should point out that the difference between the nonlinear unit root test results shown in this section and those from the Lee and Strazicich test is informative. Using the latter, the evidence that we obtain is overwhelmingly supportive for unit root behaviour in HPERs, while the former provides more nuanced evidence. Thus, our findings indicate that the distinction between abrupt vs. smooth transition of the HPER from one regime to the other is rather important.

[INSERT TABLE 5 HERE]

## 7. Conclusions

This paper investigates the time series properties of the ratio of house prices to earnings using aggregate and regional data from the UK. This ratio is a proxy for the affordability of housing. The results from a series of unit root tests are only weakly supportive of stationarity in HPERs implying that house prices may permanently diverge from earnings. This finding is robust to allowing for shifts or cyclical behaviour in the HPER in an effort to account for the potential loss of power of standard linear no-breaks unit root tests. Evidence for stationarity is stronger when a nonlinear unit root test, as opposed to a test that allows for abrupt structural breaks is used, thereby indicating that the econometric assumption about the type of transition across regimes matters.

The slowness or lack thereof of mean reversion in HPERs implies that there would be considerable costs for the wider economy if monetary policy was used to improve housing affordability by mitigating upward swings in the housing market. More specifically, monetary conditions would need to be tightened for a prolonged period with the resulting high interest rates crowding out other potentially productive investments (see also Andre, 2010). Regional differences, such as the ones identified in our analysis regarding the timing of breaks and nature of the time-series processes involved, further complicate the task of attempting to stabilise the HPER using a common interest rate tool.

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**Table 1: Descriptive statistics of the house price to earnings ratio**

| | Mean | Standard Deviation |
|---|---|---|
| UK | 4.02 | 0.83 |
| North | 3.49 | 0.96 |
| York & Humbers | 3.34 | 0.76 |
| North West | 3.42 | 0.75 |
| East Midlands | 3.85 | 0.85 |
| West Midlands | 4.20 | 0.96 |
| East Anglia | 4.17 | 0.84 |
| South West | 4.78 | 1.15 |
| South East | 5.18 | 1.01 |
| Greater London | 4.52 | 0.98 |
| Wales | 3.76 | 0.97 |
| Scotland | 3.52 | 0.56 |
| Northern Ireland | 3.35 | 1.55 |

Note: The descriptive statistics were calculated over the time period 1983Q2-2009Q1.

**Table 2: Peaks and trough of the house price to earnings ratio**

| | Peak | | Trough | | Peak | |
|---|---|---|---|---|---|---|
| UK | 4.99 | 1989Q2 | 3.09 | 1995Q4 | 5.86 | 2007Q3 |
| North | 3.81 | 1990Q3 | 2.56 | 2000Q4 | 5.92 | 2007Q2 |
| York & Humbers | 4.08 | 1989Q3 | 2.45 | 2001Q1 | 5.13 | 2007Q2 |
| North West | 4.10 | 1990Q1 | 2.59 | 2001Q1 | 5.06 | 2007Q2 |
| East Midlands | 5.22 | 1989Q2 | 2.94 | 1996Q1 | 5.41 | 2007Q3 |
| West Midlands | 5.48 | 1989Q1 | 3.18 | 1999Q1 | 6.16 | 2007Q4 |
| East Anglia | 6.87 | 1988Q4 | 3.14 | 1996Q3 | 5.39 | 2007Q3 |
| South West | 6.72 | 1988Q4 | 3.31 | 1995Q4 | 6.95 | 2007Q1 |
| South East | 7.54 | 1988Q4 | 3.73 | 1995Q4 | 6.70 | 2007Q3 |
| Greater London | 6.12 | 1988Q4 | 2.93 | 1995Q4 | 6.43 | 2007Q3 |
| Wales | 4.55 | 1989Q2 | 2.84 | 1995Q4 | 6.10 | 2007Q1 |
| Scotland | 4.83 | 1989Q2 | 2.60 | 2001Q1 | 4.69 | 2007Q4 |
| Northern Ireland | - | - | 1.77 | 1990Q3 | 8.59 | 2007Q2 |

Note: This Table shows the value of the house price to earnings ratio at the peak or the though of the boom-bust cycle and the date upon which the peak or through was reached.

## Table 3: Linear unit root test results

| | Linear ADF t-test statistic | | | Ng Perron $MZ_\alpha$ test statistic | |
|---|---|---|---|---|---|
| | Constant | Constant and Trend | None | Constant | Constant and Trend |
| UK | -2.49 [1] | -2.41 [1] | -0.40 [5] | -14.41 [1] ** | -40.93 [5] ** |
| North | -1.77 [2] | -2.29 [2] | 0.02 [2] | -7.42 [2] | -11.42 [2] |
| York & Humbers | -1.87 [4] | -2.02 [4] | -0.10 [5] | -6.94 [4] | -11.02 [4] |
| North West | -2.15 [4] | -2.13 [4] | -0.33 [9] | -13.92 [4] ** | -14.98 [4] |
| East Midlands | -1.87 [8] | -1.95 [8] | 0.02 [8] | -11.75 [8] * | -17.37 [8] * |
| West Midlands | -1.63 [4] | -1.75 [4] | -0.04 [5] | -4.46 [4] | -7.85 [4] |
| East Anglia | -1.97 [4] | -1.93 [4] | -0.15 [4] | -8.24 [4] * | -8.92 [4] |
| South West | -1.89 [1] | -2.04 [1] | 0.01 [1] | -6.49 [1] | -8.83 [1] |
| South East | -2.05 [3] | -1.94 [3] | -0.37 [3] | -10.65 [3] * | -12.02 [3] |
| Greater London | -1.87 [5] | -1.48 [4] | -0.25 [5] | -4.49 [4] | -5.61 [4] |
| Wales | -2.06 [2] | -2.12 [2] | -0.28 [4] | -8.34 [2] * | -12.48 [2] |
| Scotland | -1.81 [1] | -1.81 [1] | -0.19 [2] | -6.46 [1] | -6.81 [1] |
| Northern Ireland | -0.83 [4] | -1.76 [7] | 0.09 [3] | -4.13 [3] | -8.99 [3] |

Note: The linear unit root tests were undertaken over the time period 1983Q2-2009Q1. The number in the square bracket shows the number of lagged difference terms in the ADF and Ng-Perron unit root test, chosen by the Modified Akaike Criterion. The reported ADF *t*-statistic, Ng-Perron $MZ_\alpha$ statistic test the null hypothesis that the log HPER contains a unit root. **, * indicate rejection of the null hypothesis at the 1%, 5% level of significance, respectively.

## Table 4: Two-break unit root test results

| | Lee and Strazicich test statistic | | | | | |
|---|---|---|---|---|---|---|
| | Breaks in constant | | | Breaks in constant and trend | | |
| | $\tilde{\tau}$ -stat | Break dates | | $\tilde{\tau}$ -stat | Break dates | |
| UK | -3.48 [10] | 1991Q4 | 1992Q4 | -4.84 [10] | 1994Q2 | 2002Q1 |
| North | -2.80 [11] | 2002Q1 | 2003Q1 | -5.49 [11] | 1990Q2 | 2002Q2 |
| York & Humbers | -3.30 [2] | 2001Q4 | 2002Q4 | -5.04 [10] | 1990Q2 | 2002Q1 |
| North West | -3.84 [11] | 1992Q3 | 2002Q2 | -5.39 [11] | 1990Q2 | 2002Q2 |
| East Midlands | -3.07 [12] | 1989Q2 | 1990Q2 | -5.00 [12] | 1994Q2 | 2001Q4 |
| West Midlands | -3.11 [10] | 1991Q4 | 1992Q4 | -4.70 [11] | 1992Q2 | 2001Q4 |
| East Anglia | -3.79 [2] | 1989Q2 | 1990Q1 | -4.32 [2] | 1990Q1 | 1999Q2 |
| South West | -2.82 [2] | 1989Q2 | 1990Q1 | -4.41 [9] | 1992Q2 | 2001Q4 |
| South East | -3.33 [11] | 1989Q3 | 1997Q1 | -4.72 [2] | 1990Q4 | 2000Q4 |
| Greater London | -3.50 [10] | 1996Q2 | 1997Q2 | -4.31 [10] | 1992Q3 | 2001Q4 |
| Wales | -3.15 [11] | 2002Q4 | 2004Q3 | -4.72 [10] | 1993Q3 | 2002Q2 |
| Scotland | -2.57 [3] | 1989Q2 | 1990Q2 | -4.59 [3] | 1991Q1 | 2003Q4 |
| Northern Ireland | -2.60 [3] | 1993Q3 | 2003Q1 | -6.59 [3] ** | 1990Q1 | 2005Q4 |

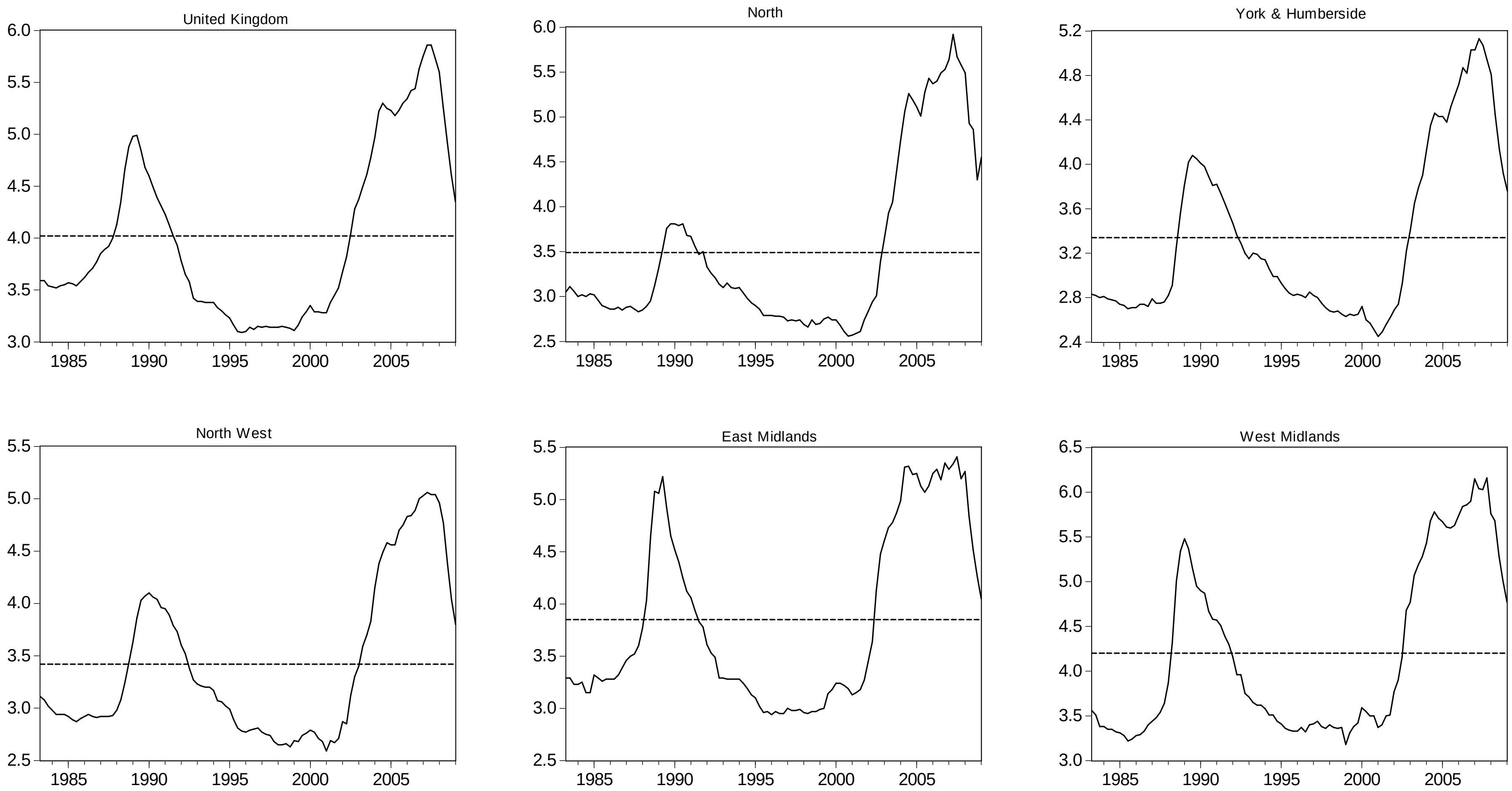
Note: The two-break unit root test was undertaken over the time period 1983Q2-2009Q1. The number in the bracket shows the number of lagged difference terms in the two-break unit root test, chosen by the 't-sig' approach. Particularly, we set an upper bound of twelve for the lag length and test down until a significant (at the 5% level) lag is found. The reported Lee and Strazicich statistic tests the null hypothesis that the log price to earnings ratio contains a unit root. **, * indicate rejection of the null hypothesis at the 1%, 5% level of significance, respectively.

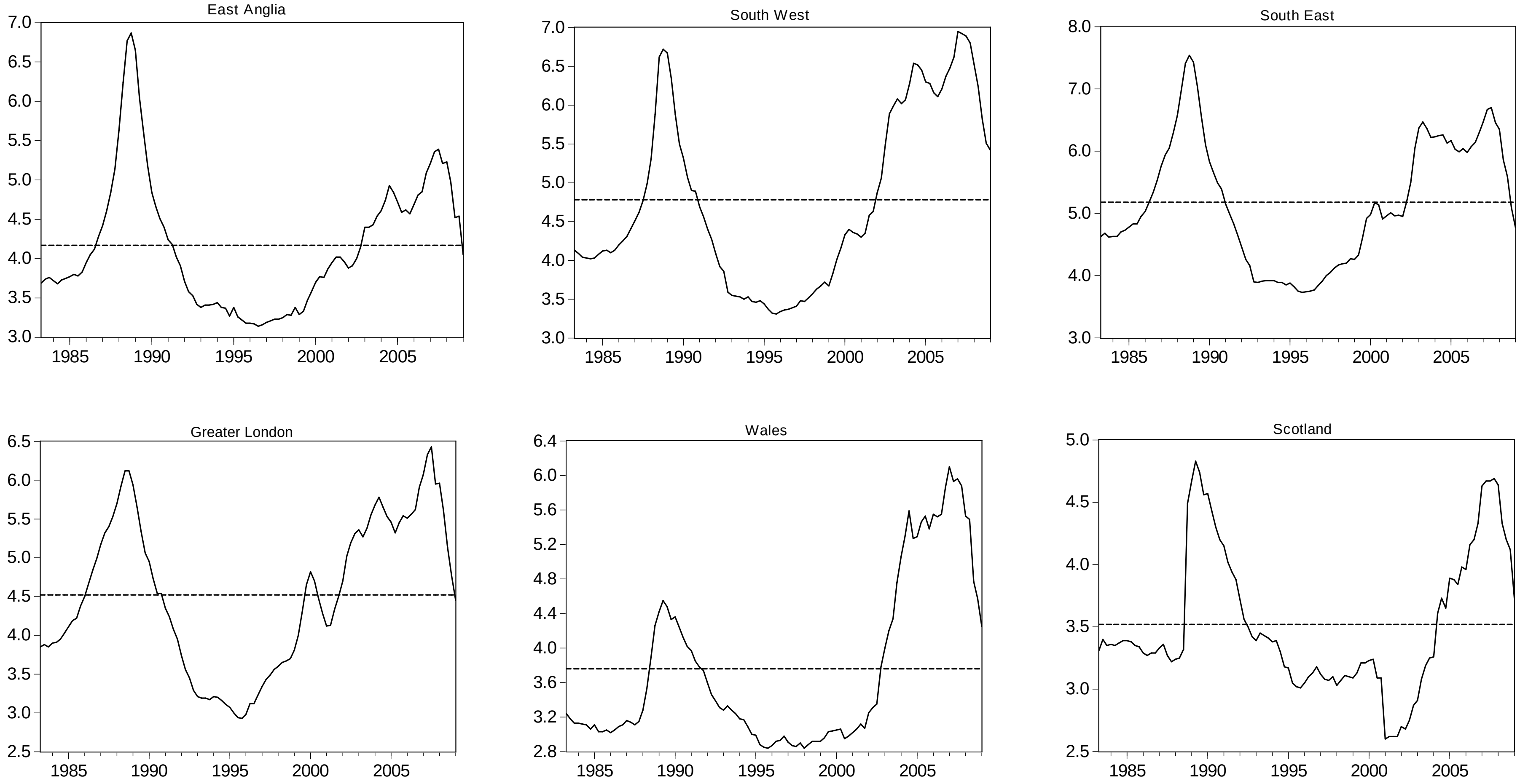
**Table 5: Nonlinear unit root test results**

| | Nonlinear ADF t-test statistic | | |
|---|---|---|---|
| | Constant | Constant and Trend | None |
| UK | -2.66 [5] * | -3.15 [5] * | -1.08 [5] |
| North | -2.05 [2] | -2.52 [2] | -0.74 [2] |
| York & Humbers | -1.77 [4] | -2.36 [4] | -0.51 [4] |
| North West | -2.51 [9] | -3.00 [9] | -1.16 [9] |
| East Midlands | -2.10 [8] | -3.03 [8] | -0.56 [8] |
| West Midlands | -1.94 [5] | -3.60 [5] ** | -0.64 [5] |
| East Anglia | -4.22[5] *** | -3.71 [4] ** | -1.02 [5] |
| South West | -2.84 [3] * | -3.41 [1] * | -0.62 [1] |
| South East | -2.80 [1] * | -2.55 [3] | -0.84 [3] |
| Greater London | -2.13 [5] | -1.79 [5] | -0.75 [5] |
| Wales | -2.19 [2] | -2.93 [2] | -0.94 [2] |
| Scotland | -2.40 [2] | -2.46 [2] | -0.71 [2] |
| Northern Ireland | -2.48 [3] | -3.25 [3]* | -0.83 [3] |

Note: The nonlinear unit root tests were undertaken over the time period 1983Q2-2009Q1. The number in the square bracket shows the number of lagged difference terms in the nonlinear unit root test. The reported nonlinear ADF *t*-statistic tests the null hypothesis that the log HPER contains a unit root. Asymptotic critical values are obtained from Table 1 in Kapetanios *et al*. (2003). **, * indicate rejection of the null hypothesis at the 1%, 5% level of significance, respectively.

.**Figure 1: House price to earnings ratio**




East Anglia
7.0
6.5
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1985
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2005
South West
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Greater London
6.5
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Wales
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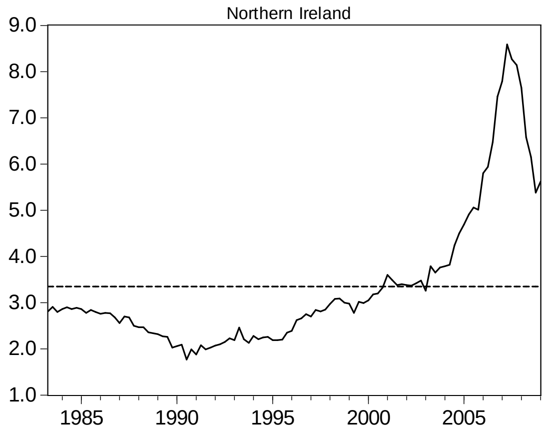


Note: The dotted line represents the average value of the series over the time period 1983Q2-2009Q1.