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How Convergent are Regional House Prices in the United Kingdom? Some New Evidence from Panel Data Unit Root Testing

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Abstract. Using a variety of econometric methods, existing studies have failed to reach a consensus on whether or not UK regional house prices are engaged in long-run equilibrium relationships with each other. Using data for the 1973-2005 study period, this study offers a novel approach to this debate through the application of unit root testing within a seemingly unrelated regression framework. It is argued that there exist significant advantages in this approach over and above existing univariate and panel data unit root testing procedures. The results indicate that the majority of UK regions exhibit regional house price convergence. However, there is an east-west split in terms of whether regional house prices have a tendency towards long-run equilibrium relationship with UK prices as whole. There is also evidence of considerable heterogeneity in the regional speeds of adjustment towards long-run equilibrium.

JEL Classification Codes: C5, R0.
Keywords: Panel data unit root, house price, convergence.

1. Introduction

At both national and regional levels, the expansion of owner occupation across the UK over recent decades has enhanced the role of housing wealth in driving consumption expenditure. This is reflected in the attention now paid by the Bank of England to the state of the domestic housing market when commenting on the state of the national economy and setting UK

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interest rates. At the regional level, fluctuations in relative house prices have the potential to influence relative regional economic activity. Fluctuations in relative prices also have the potential to influence labour mobility through the affordability of housing and relocation costs. There is therefore considerable value in understanding how regional house prices behave in relation to each other over time. Starting from the work of Meen (see for example, Meen (1999)), it has been argued that shocks to regional house prices "ripple out" across the economy. While the notion of such a ripple effect may rely on factors such as spatial patterns in the determinants of house prices, migration, equity transfer, and spatial arbitrage, the ripple effect requires some notion of a degree of long-run constancy, or a long-run equilibrium relationship, between regional house prices.

A large literature now exists supporting the notion of a causal link from house prices in the South East of England to other regions. However, the literature to date can only offer mixed evidence that long-run equilibrium relationships between all regional house prices actually exists. Studies that include Holmans (1990), MacDonald and Taylor (1993), Alexander and Barrow (1994), Drake (1995), Ashworth and Parker (1997), Meen (1999), Petersen *et al*. (2002) and Holmes and Grimes (2005) offer varied conclusions in support of a ripple effect. Many of these studies employ Engle and Granger (1987) or Johansen (1988) likelihood ratio tests of cointegration in the search for regional–national house price convergence. For example, MacDonald and Taylor (1993) suggest a ripple effect is present in a limited form where mixed evidence of long-run relationships between regional house prices leads to the notion of weak segmentation of the housing market. The notion of weak segmentation is further supported by Cook (2005b) who is unable to find in favour of stationary house price differential across all regions despite the application of more powerful generalised least squares (GLS)-based unit root tests. In other investigations, Holmans (1990) fails to uncover stationarity over a long span of data starting in the 1930s, while Cook [(2003) and (2005a)] takes a different line of investigation and identifies a consistent pattern of asymmetric adjustment where reversion to equilibrium occurs more rapidly (slowly) when house prices in the South of England decrease (increase) relative to other regions.

This paper investigates the extent to which regional house prices in the UK exhibit long-run convergence. The contribution of this paper to the existing literature is thereby focused on two key factors. First, a key contribution is in terms of the econometric methodology that is employed. A novel approach of this paper is to employ a new test panel data unit root test which estimates augmented Dickey-Fuller (ADF) regressions within a seemingly unrelated regression (SUR) framework in the search for house price convergence between each region and the UK. This methodology is advocated by Breuer *et al.* (2002) and Chung and Crowder (2004) in their studies of international purchasing power parity and real interest parity respectively. This SURADF test offers enhanced test power over the more familiar ADF unit root testing procedure. However, the SURADF test also offers two crucial advantages over existing panel data unit root tests, such as the tests advocated by Levin *et al*. (2002) and Im *et al*. (2003). One advantage is that the SURADF test enables the researcher to identify how many and which series within the panel are responsible for rejecting the null of non-stationarity. A further advantage is that the SURADF procedure is able to address problems associated with the presence of cross-sectional dependency among the series in the panel.

A second contribution to the literature offered by this paper is in terms of examining how regional house prices respond to deviations from their long-run equilibrium relationships. Rather than focus solely on how regional house prices respond to house price changes emanating from the south of England, this study examines regional house prices in relation to UK and focuses on the measured half-lives associated with a given deviation from long-run equilibrium. On this basis, we are able to group regions according to the speed of response back towards equilibrium.

The paper is organized as follows. Section 2 discusses the background literature on panel data unit root testing and highlights the advantages of the SURADF procedure. The Section 3 reports and discusses the results: the application of univariate ADF unit root testing indicates that long-run house price convergence applies to only four out of thirteen regions at most. However, the application of the SURADF procedure is able to identify stationarity in ten regions. The final section contains a summary and the conclusion.

2. Testing for Stationarity in Panel Datasets

This study employs a three-stage testing procedure for regional (logarithmic) house price convergence. *Stage One* involves the application of standard univariate ADF unit root tests on house price differentials. *Stage Two* involves the application of panel data unit root tests which offer enhanced test power over their univariate counterparts. Stage Three involves the application of the SURADF procedure.

Define the regional house price differential for region *i*, d_{it}, as

$$d_{it} = X_{it} - M_{it} \tag{1}$$

where *X* refers to the natural logarithm of the house price index of region *i* and *M* refers to the natural logarithm of the UK house price, $i = 1, 2, \ldots N$ regions and $t = 1, 2, \ldots T$ time periods. Suppose d_{it} is generated by a first order autoregressive process, $d_{it} = \kappa_i + \rho_i d_{it-1} + \omega_{it}$ which can be transformed into the familiar ADF regression

$$\Delta d_{it} = \kappa_i + \tau_i d_{it-1} + \sum_{j=1}^{k_i} \nu_{ij} \Delta d_{it-j} + \varepsilon_{it} \tag{2}$$

where $\tau_i = \rho_i - 1$. Acceptance of the null hypothesis $\tau_i = 0$ ($\rho_i = 1$) means that d_{it} is a non-stationary series whereas rejection of the null means that d_{it} is stationary and therefore the long-run convergence of regional house price indices. There exist a range of panel data unit root tests that offer increased power over methods for univariate unit root testing. Two widely used panel data unit root tests are offered by Levin *et al*. (2002) and Im *et al*. (2003). Both tests rely on subtracting cross-sectional averages to address cross-sectional correlation. Whereas, Levin *et al*. allow for individual specific intercepts and time trends, the Im *et al*. procedure involves the computation of an average ADF statistic for the entire panel.

The early panel data unit root tests advocated by researchers such as Levin *et al*. (2002) offer restrictive joint and null hypotheses where *all* in the panel series are either non-stationary or stationary where all members of the panel have common autoregressive parameters, i.e. ρ ‘s or β ‘s. Both the Im *et al*. and Levin *et al*. tests are of the null of joint non-stationarity across the panel. However, under Levin *et al*. the alternative hypothesis is that all members are stationary. The Im *et al*. test has an advantage in that it allows for heterogeneity under the alternative hypothesis where the autoregressive coefficient can differ across panel members. In addition to this, Im *et al*. (2003) use Monte Carlo results to demonstrate that their test has more favourable finite sample properties than Levin *et al*. However, the Im *et al*. procedure does not provide a beta coefficient that one can meaningfully use to discuss the panel speed of convergence. This is because the panel may comprise both stationary and non-stationary members where the latter are characterised by infinite measures of persistence.

O'Connell (1998) argues that since these panel data unit root tests presume identically and independently distributed disturbances, there may be dramatic implications for statistical size and power to the extent that the null may not be correctly accepted or rejected. O'Connell (1998) evaluates contemporaneous correlation directly by estimating the disturbance covariance matrix and so allows for contemporaneous cross correlation. To allow for correlation across the panel, Im *et al*. (2003) assume that

$$\varepsilon_{it} = \theta_t + u_{it} \tag{3}$$

where θ_t is a time-specific common effect that allows for a degree of dependency across the series and u_{it} is an idiosyncratic random effect that is independently distributed across groups. To remove the effect of the common component θ_t, we can subtract the cross-section mean value for d from both sides of $d_{it} = \kappa_i + \rho_i d_{it-1} + \omega_{it}$ to yield $\tilde{d}_{it} = \tilde{\kappa}_i + \tilde{\rho}_i \tilde{d}_{it-1} + \tilde{\omega}_{it}$ where $\tilde{d}_{it} = d_{it} - N^{-1}\sum_{i=1}^{N} d_{it}$. From this, we can then derive the following demeaned regression

$$\Delta\tilde{d}_{it} = \tilde{\kappa}_i + \tilde{\tau}_i \tilde{d}_{it-1} + \sum_{j=1}^{k_i} \tilde{\nu}_{ij} \Delta\tilde{d}_{it-j} + \tilde{\varepsilon}_{it} \tag{4}$$

where $\tilde{\tau}_i = \left(\tilde{\rho}_i - 1\right)$. However, this demeaning procedure only partially tackles cross-sectional dependence. More recent papers by Bai and Ng (2004), Moon and Perron (2004), and Phillips and Sul (2003) avoid the restrictive nature of the cross-section de-meaning procedure by allowing the common factors to have differential effects on different cross section units. For example, Moon and Perron (2004) propose a pooled panel unit root test based on "de-factored" observations and suggest estimating the factor loadings that enter their proposed statistic by the principal component method. Pesaran (2004) presents the common correlated effects estimator. This procedure filters the individual-specific regressors by means of (weighted) cross-section aggregates such that asymptotically as the cross-section dimension tends to infinity, the differential effects of unobserved common factors are eliminated.

Most of the above-mentioned studies address the issue of contemporaneous correlation of the residuals, but they still provide a single test statistic that does not allow the researcher to identify how many and which of the series in the panel are in fact stationary. To address these concerns, this paper utilizes the alternative test procedure recently advocated by Breuer *et al.* (2002) and Chung and Crowder (2004) that exploits the power of panel data analysis without imposing uniformity across the panel under either the null or alternative hypothesis. This test relies on SUR analysis with no across panel restrictions under either hypothesis. More formally, the SURADF procedure can be represented in the following equations

$$
\begin{aligned}
&\Delta d_{1t} = \kappa_1 + \tau_1 d_{1t-1} + \sum_{j=1}^{i} \nu_{1j} \Delta d_{it-j} + \eta_{1t} \\
&\Delta d_{2t} = \kappa_2 + \tau_2 d_{2t-1} + \sum_{j=1} \nu_{2j} \Delta d_{2t-j} + \eta_{2t} \\
&\ldots\ldots \quad \ldots \quad \ldots. \quad \ldots\ldots\ldots.. \quad .. \\
&\Delta d_{Nt} = \kappa_1 + \tau_N d_{Nt-1} + \sum_{j=1} \nu_{Nj} \Delta d_{Nt-j} + \eta_{Nt}
\end{aligned}
\tag{5}
$$

where η_i denotes the SUR residual for region *i*. The significance of each τ_i is tested against critical values generated through Monte Carlo simulation. This specification allows the significance of the autoregressive parameter to differ across the series by relaxing the restriction $\tau_1 = \tau_2 = ... = \tau_N$ thereby avoiding the joint null hypothesis of non-stationary ($\tau_1 = \tau_2 = ... = \tau_N = 0$) which is used by both the Levin *et al.* and Im *et al.* tests, and the corresponding alternative hypothesis that all series have the same autoregressive coefficient ($\tau_1 = \tau_2 = ... = \tau_N < 0$) which is used by the Levin *et al.* test. The researcher can identify which panel members are stationary or non-stationary because, unlike the previous Levin *et al.* and Im *et al.* tests, the SURADF test is based on individual rather than joint hypotheses.

This SURADF test offers increased power over univariate ADF tests. Moreover, Breuer *et al.* (2001) report some findings from a Monte Carlo analysis of SURADF test power in rejecting a false null hypothesis. They find that the test power associated with the SURADF test procedure is 2-3 times greater than the power associated with the single equation ADF test especially when residual cross-correlations are high and the sample size exceeds 100 observations. In further comparison with the Levin *et al.* and Im *et al*. tests, it should be noted that both the Im *et al*. and Levin *et al*. tests are asymptotically normal under the assumption of cross-sectional independence. Under the likely scenario of cross-sectional dependence, their critical values are not asymptotically normal. The SURADF procedure offers a better way forward through addressing the contemporaneous correlation of the residual terms across each SURADF regression. This gives the SURADF procedure an informational advantage over the univariate ADF, Levin *et al*. and Im *et al*. procedures. In the case of the Levin *et al*. and Im *et al*. tests, for example, simply demeaning the data means that one is only partially addressing cross-sectional dependence with respect to common shocks that affect all panel members together. In addition to this member-specific lag structures are allowed to ensure that each equation is correctly specified with residuals that are white noise.

With regard to comparisons with alternative panel data unit root tests, the relative advantage of the SURADF test procedure lies in its ability to determine which panel members are stationary or non-stationary while also addressing cross-sectional dependence. Existing panel data unit root tests fail to deliver on both these counts. However, at this stage it is informative to consider some disadvantages attached to the SURADF procedure. In particular, the asymptotic distribution of the SURADF test is not free from nuisance parameters, namely the correlation structure of the underlying residuals. Therefore, appropriate critical values are obtained case by case via simulation. A further issue is that the SURADF procedure is limited to cases where the researcher is analyzing a balanced panel where the number of cross-sections is no greater than the number of time-series observations.

3. Results

The data examined are quarterly observations on the natural logarithm of regional house prices and the UK as a whole for all properties over the period 1973Q4-2005Q1 using a dataset provided by the Nationwide Building Society.[1] This series covers the study period 1973Q4 to 2005Q1 and offer a balanced panel data set for the following thirteen regions of the UK: East Anglia, East Midlands, London, North, North West, Outer Metropolitan, Outer South East, Scotland, Northern Ireland, South West, Wales, West

[1] The Nationwide house price data are downloadable from http://www.nationwide.co.uk/default.htm.

Midlands, Yorkshire & Humberside plus the UK. All data used in this study are mixed-adjusted to allow for variations in housing quality when computing the regional or national house price series.[2]

Pre-testing indicated that all house price levels are first difference stationary. Table 1 reports unit root tests on regional house price differentials with respect to the UK. At the 5% significance level, the univariate ADF tests are only able to identify the stationarity of regional house price differentials in the case of London and the Northern regions. Clearly, this is very weak evidence in favour of regional house price convergence and, in fact, is more indicative of long-run divergence with the majority of regions having the tendency to drift further away from the UK as a whole.

The second stage of the empirical investigation is to employ panel data unit root testing. The motivation behind this is to employ more observations and exploit the cross-country variations of the data in estimation thereby yielding higher test power than standard unit root tests based on individual time series. Given that low test power could be responsible for acceptance of the non-stationary null in eleven of the thirteen cases reported in Table 1, the application of panel data unit root tests makes it increasingly likely that stationary can be identified. Table 2 reports the findings from three alternative methods. These are the Levin *et al*. (2002) and Im *et al*. (2003) tests that were described earlier as well as Hadri (2000) who defines a null hypothesis of joint stationarity against the null that all series are non-stationary. Under cross-sectional independence, each of these statistics is distributed as standard normal as both N and T grow large. At the 5% significance level or better, both the Levin *et al*. (2002) and Im *et al*. (2003) tests are able to reject the null hypothesis of joint non-stationarity of the series in the panel.[3] In addition to this, the Hadri (2000) test strongly

[2] The purpose of mix adjustment is to simply isolate pure price changes. One can show how changes in the mixture of properties sold each quarter could give a misleading picture of what is actually happening to house prices. Moreover, the set of properties sold from quarter to quarter will vary by location and design etc. and some adjustment is necessary to make sure these factors do not give a false impression of the actual changes to house prices. A mix-adjusted or ‘standardised’ index is not affected by such changes because the relative weight given to each characteristic of a property in the ‘mix’ (or ‘basket’, to use an analogy with consumer prices) is fixed from one quarter to the next.

[3] It is important to reconcile the results reported in Tables 1 and 2. The ADF statistics (no trend) in Table 1 range from -3.229 to -1.764. However, the Im *et al.* statistic of -1.708, which is an average calculation across the sample, falls outside

rejects the null of joint stationarity. However, these results are based on a single statistic which provides limited information in terms of individual behaviour within the panel. Moreover, the finding with respect to the Im *et al*. (2003) tests is consistent with, say, just a single series from within the panel being responsible for rejecting the null hypothesis of joint non-stationarity. The Hadri (2000) test indicates that one can reject the null that all series are stationary, but the alternative hypothesis is restrictive in that all series are non-stationary. In addition to this, none of these tests adequately account for cross equation correlation. The qualifications that one can attach to the panel data unit root tests provides the case for a technique that enables a reflection of individual series behaviour from within the panel and better handles the possibility of contemporaneous correlation of disturbances.

Table 3 reports the findings from the SURADF test applied to the full sample of thirteen regional house price differentials. This table also reports the 1, 5 and 10% critical values that have been specifically simulated for this panel using knowledge of the variance-covariance matrix of residuals and lag structures across the SUR equations. On this occasion, the non-stationary null is also rejected at 5% significance level or better in seven out of thirteen regional house price differentials. At the 10% significance level, we find that ten out of thirteen regions exhibit convergence. Either way, the SURADF test indicates that a substantially larger proportion of the sample is stationary than was the case of the univariate unit root tests. The three regional house price differentials that are non-stationary include the South West, West Midlands and Wales regions. The implication here is that these regions are segmented from the rest of the UK in the sense that the difference between house prices in these three regions and those elsewhere will tend to grow larger and larger over time, with no tendency for the disparity to settle down at an equilibrium level.

These findings may be seen in the context of MacDonald and Taylor (1993). They look at all possible combinations of regional pairs and find a segmentation of the national housing market with Greater London, the South East, the South West and East Anglia forming one group, and the West and East Midlands, the North, the North-West, Yorkshire-Humberside, Wales

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this range. With this in mind, one should remember that the ADF statistics are based on data for d_i (see equation (2)) whereas the Im *et al*. statistic is based on the use of demeaned data as expressed by $\tilde{d}_i$ (see equation (4)).

and Scotland forming another. This study finds that segmentation is in fact based on a small group of regions for whom there is no long-run relationship with the UK average. Theoretically speaking, it should be noted that cointegration is a transitive relationship. For example, if the UK house prices series is cointegrated with both Scotland *and* Outer Metropolitan house prices, then the Scotland and Outer Metropolitan pair of regional house price series should, in principle, cointegrate. In this study, regional house prices in the South West, West Midlands and Wales are not cointegrated with the UK. Therefore, the transitivity argument suggests that these regions do not exhibit pair-wise cointegration with the remaining ten regions that are cointegrated with the UK series. This analysis offers more support to Cook (2005b) who employs powerful DF-GLS and KPSS unit root tests and detects stationarity in regional-UK house price differentials in the majority of cases. However, the remaining non-stationary cases in Cook's study are the South West and Yorkshire & Humberside. Application of the SURADF methodology in this study indicates that the regions exhibiting divergent tendencies are most likely to be adjacent regions rather than regions that are distant from each other.

Table 4 reports some estimates of the half-lives associated with deviations from long-run equilibrium. These half-life calculations are based on the estimated autoregressive parameter from the estimated SURADF equations and are inversely related to the speed of adjustment towards long-run equilibrium. For those regions where convergence is confirmed, there is considerable variation ranging from 5.7 quarters in the case of Yorkshire & Humberside to 18.4 quarters in the case of Northern Ireland. This is against a background of an average half-life calculated as 9.9 quarters across those regions exhibiting convergence. A number of key characteristics concerning these half-lives can be highlighted. One can point to the presence of regional clustering in the speeds of adjustment towards long-run equilibrium. Moreover, the calculated half-lives enable us to construct geographically-proximate or adjacent groupings. These are (i) Yorkshire & Humberside and the Northern regions (5.7 and 6.1 quarters respectively), (ii) East Anglia and the East Midlands (7.9 and 6.7 quarters respectively) and (iii) London, Outer Metropolitan and Outer South East (11.4, 12.0 and 9.6 quarters respectively). Also, a group comprising the two most distant regions from the South of England comprises Northern Ireland and Scotland (18.4 and 11.4 quarters respectively). On this basis, the usual understanding of the regional house price response may need to be modified. Rather than simply viewing the ripple effect in terms of an economic shock emanating in the South of England affecting local house prices and then rippling out in an orderly fashion across the outer regions. This study finds that the majority of regional house prices have a long-run cointegrating relationship. However, we may also view the UK regions as being in various "convergence clubs" where speeds of adjustment following a deviation from long-run equilibrium with respect to the UK are considerably varied and unrelated to distance from the South of England.

4. Summary and Conclusion

House price linkages have important implications for the mobility of labour within the national economy as well as for regional wealth effects. This study measures the extent to which long-run convergence among regional house prices is present. The novel approach employed in this study is the application of augmented Dickey-Fuller panel data unit root tests within a seemingly unrelated regression framework. Panel data unit root tests offer well-known advantages of increased power over their univariate counterparts and provide some support in favour of convergence. However, the seemingly unrelated regression framework for unit root testing offers clear advantages over existing panel data techniques. These advantages relate to the incorporation of cross equation correlation and the ability to determine which series in panel are responsible to any rejection of the null of joint non-stationarity. Using quarterly data over a thirty two year period, the univariate augmented ADF tests are strongly dismissive of long-run convergence. However, the application of the seemingly unrelated regression approach suggests that the majority of the thirteen regions are linked by long-run homogeneous cointegrating relationships. There is evidence of segmentation because three regions – the South West, Wales and West Midlands – do not exhibit long-run convergence with the UK house price series. Further findings from this study indicate considerable variation in regional speeds of adjustment towards long-run equilibrium where regional clusters based on similar speeds of adjustment exist. This leads us to modify how the concept of a regional ripple effect may be viewed.

Table 1. ADF Unit Root Tests

	ADF (no trend)	ADF (trend)
East Anglia	-2.564	-2.579
East Midlands	-1.798	-2.116
London	-1.841	-3.581**
North West	-2.850*	-2.974
Northern	-3.229**	-3.407*
Northern Ireland	-2.304	-2.353
Outer Metropolitan	-2.507	-2.741
Outer South East	-2.000	-2.151
Scotland	-2.666*	-3.442*
South West	-1.887	-2.708
Wales	-2.111	-3.142
West Midlands	-1.764	-1.993
Yorkshire & Humberside	-1.853	-2.778

Notes for Table 1. In all cases, the lag length is selected according to the AIC. ***, ** and * respectively denote rejection of the non-stationary null at the 1, 5 and 10% significance levels with critical values of -3.484, -2.885 and -2.579 (no trend) and -4.034, -3.447 and -3.148 (trend) respectively.

Table 2. Panel Data Unit Root Tests

Test	Test Statistic
Levin *et al*. (2002)	-2.758***
Im *et al*. (2003)	-1.708**
Hadri (2000)	7.709***

Notes for Table 2. The individual lag lengths for each of these tests are determined by the AIC. All test statistics are distributed as standard normal. In the case of the Im *et el*. and Levin *et al*. tests, *** and ** respectively denote rejection of the joint non-stationary null at the 1 and 5% significance levels where the 1 and 5% critical values are -2.33 and -1.64 respectively. In the case of the Hadri test, *** denotes rejection of the joint stationary null at the 1 % significance level with a 1% critical value of 2.33.

Table 3. The SURADF Test

Region	SURADF	1%	5%	10%
East Anglia	-3.090*	-3.743	-3.173	-2.841
East Midlands	-3.014*	-3.737	-3.140	-2.841
London	-3.128*	-3.825	-3.231	-2.925
North West	-4.639***	-3.883	-3.288	-2.991
Northern	-4.310***	-3.836	-3.226	-2.942
Northern Ireland	-3.153**	-3.618	-3.100	-2.791
Outer Metropolitan	-3.263**	-3.830	-3.204	-2.893
Outer South East	-4.092**	-4.251	-3.541	-3.208
Scotland	-3.423**	-3.849	-3.311	-3.007
South West	-2.676	-4.291	-3.651	-3.290
Wales	-2.715	-3.791	-3.202	-2.903
West Midlands	-1.097	-3.654	-3.146	-2.850
Yorkshire & Humberside	-3.975***	-3.886	-3.276	-2.961

Notes for Table 3. SURADF refers to the ADF statistic obtained through the SUR estimation of ADF regressions for involving the real house price differentials between the thirteen UK regions and the UK house price index. Following Breuer *et al*. (2002), the critical values reported in the three columns on the right have been simulated with 10000 replications where the error series were generated to be normally distributed with the variance-covariance matrix given by the SUR estimation. Each simulated house price differential was then generated from the error series using the SUR estimated coefficients. ***, ** and ** indicate rejection of the null of non-stationarity at the 1, 5 and 10% significance levels respectively. For each equation, further tests were unable to reject the null that the residuals were serially uncorrelated.

Table 4. Speeds of Adjustment

Region	$\tau_i = \rho_i - 1$	Half-life (quarters)
East Anglia	-0.084	7.900
East Midlands	-0.098	6.720
London	-0.059	11.398
North West	-0.066	10.152
Northern	-0.107	6.125
Northern Ireland	-0.037	18.385
Outer Metropolitan	-0.056	12.028
Outer South East	-0.070	9.551
Scotland	-0.059	11.398
South West	-0.085	N/A
Wales	-0.082	N/A
West Midlands	-0.027	N/A
Yorkshire & Humberside	-0.114	5.727
UK Average		9.938

Notes for Table 4. The reported autoregressive coefficients are taken from the SUR results reported in Table 3. For each region, the half-life calculation is $(\ln 0.5)/\ln(1+\tau_i)$.

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