Is there long-run convergence of regional house prices in the UK?

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Abstract

This paper investigates the long-run convergence of regional house prices in the UK. Using a variety of econometric methods, existing studies have failed to reach a consensus on whether or not regional house prices are cointegrated and exhibit long-run constancy relative to each other. We propose the application of a new test that combines principal components analysis with unit root testing to throw new light on the regional convergence debate. Using mix-adjusted quarterly house price data for 1973-2005, we find that existing unit root and cointegration methodologies indicate the presence of multiple stochastic trends with, at best, very weak evidence of long-run convergence. However, testing for the stationarity of the largest principal component based on regional house price differentials suggests that all UK regional house prices are driven by a single common stochastic trend and can be regarded as exhibiting strong convergence in the long-run. Further analysis suggests there is a high degree of persistence in regional house price differentials.

JEL classification
C5, R0

Keywords
House prices, convergence, unit roots, cointegration, principal components

1 Introduction

The behaviour of regional house prices has constituted a keen area of research in recent years. An important line of investigation has focused on interrelationships between regional house prices and testing the hypothesis that shocks to regional house prices “ripple out” across the economy on account of factors such as migration, equity transfer, spatial arbitrage and spatial patterns in the determinants of house prices (Meen, 1999).[1]

A shock that hits, for instance, on London house prices, may have no immediate impact on house prices in neighbouring regions (e.g. East Anglia) or in more far-flung areas (e.g. Scotland). However those regions may eventually feel the impact of the shock (possibly at different times). If the shocks impact on regional house prices to quite different degrees in the long run, then wealth disparities can widen as those people owning houses in regions with high house price rises increase their wealth relative to house owners in other regions. Under certain conditions, the long run effect of house price shocks may be the same across the entire country. In these circumstances, the housing market may be a source of temporary disparities in wealth across regions but not of long term disparities. In understanding the inter-relationship of the housing market with regional disparities, it is therefore important to test whether house price shocks ripple out fully across all regions or whether the long run effects of shocks are more localised.

If a ripple effect is indeed present, it will be predicated on a degree of long-run relative constancy between regional house prices where the ratio between each regional price and the national house price is stationary. While a large literature now exists supporting the notion of a causal link from house prices in the South East of England to other regions, the literature to date offers only mixed evidence that long-run equilibrium relationships between all regional house prices actually exist. A range of studies employ Engle and Granger (1987) or Johansen (1988) likelihood ratio tests of cointegration in the search of regional–

1 Meen (1999) argues that structural differences are important. Using a model of non-random spatial patterns, a ripple effect is generated irrespective of regional growth patterns. Ashworth and Parker (1997), on the other hand, provide tests of spatial dependence and cast doubt upon the existence of a ripple effect.

national house price convergence (see, *inter alia*, Holmans (1990), MacDonald and Taylor (1993), Alexander and Barrow (1994), Drake (1993), Ashworth and Parker (1997), Meen (1999), Petersen et al (2002)), yet the conclusions drawn from these studies have varied. 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.[2] Holmans (1990) fails to find stationarity over a long span of data starting in the 1930s. In further recent contributions, Cook (2003 and 2005) 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.

Given the lack of consensus on regional house price convergence in the long-run, the key contribution offered by this study is in terms of the econometric methodology that is employed. Our tests for regional house price convergence are on the basis of whether the largest principal component (LPC), based on regional benchmark deviations from the average UK house price, is stationary or not. The use of factor structures to test for unit roots and common trends is reflected in growing literature that includes Snell (1996), Hall et al (1999), Moon and Perron (2002), Phillips and Sul (2002), Bai (2004) and Bai and Ng (2004) and others. On the basis of this literature, one can argue that dynamic factor models are useful in several areas of economic analysis. The first is index modeling and extraction where factors are regarded as unobservable economic indices that capture the co-movement of many variables. Second, factors synthesize information in a way that is capable of aggregating information from many economic indicators. Third, Stock and Watson (1999), Favero and Marcellino (2001) and Artis et al (2001) show how dynamic factor models can be used to improve forecasting accuracy. Fourth, one major source of cross-section correlation in macroeconomic data is common shocks. Dynamic factor models are capable of modeling cross-section correlations allowing for heterogeneous responses to common shocks through heterogeneous factor loadings. Fifth, factor models can be used to study cross-

2 The notion of weak segmentation is further supported by Cook (2005b) who is unable to find in favour of stationary house price differentials across all regions despite the application of more powerful GLS-based unit root tests.

section cointegration in nonstationary panel data. If time series share a common trend, this implies strong cross-section correlation which can be exploited to produce new univariate test statistics.

Unit root testing of the largest principal component based on regional-national house price differentials also offers a number of key advantages over existing tests for convergence. In investigating the number of common shared trends, the advantage of examining the stationarity of the LPC is that, unlike the Johansen (1988) maximum likelihood procedure (and the Stock and Watson (1988) common trend framework), it does not require the estimation of a complete vector autoregression system (VAR). The size and power of the test are not affected by the VAR being constrained to an unreasonably low order on account of data limitations. This method also avoids the need for an entire sequence of tests for the stationarity of a multivariate system. As indicated by Snell (1996), even if each test in the standard sequence of tests for stationarity had a reasonable chance of rejecting the false null, the procedure as a whole is likely to have low power. Snell (1999) presents a small Monte Carlo study where the size and power of the test based on principal components are computed. Experimentation is based on a variety of data generation processes and the test based on principal components is confirmed as having acceptable size and reasonable power compared to the Johansen likelihood ratio cointegration test. In cases of marginal cointegration (i.e. when the cointegrating combination is only borderline stationary) and in small sample sizes, the Johansen (1988) likelihood ratio test has little ability to discriminate between no cointegration and cointegration whereas estimation based on principal components *does* have discriminatory power in such circumstances. Further evidence is provided by Hall et al (1999) who point to favourable size and power qualities of using principle components to ascertain the number of common stochastic trends driving non-stationary series when the number of observations exceeds the number of series.

The paper is organised as follows. The following section discusses the data and econometric methodology. The third section reports and analyses the results. We find that the LPC based on regional house price deviations from the UK average house price is indeed stationary. However, we find that the speed of adjustment back towards long-run equilibrium indicates a very high degree of persistence. The final section concludes.

2 Searching for Regional House Price Convergence

This study employs a two-stage testing procedure for regional (logarithmic) house price convergence. *Stage One* involves computing *n* principal components using the *n* house price differentials. If the LPC is stationary, then all remaining principal components will also be stationary thereby confirming *strong convergence* among all the n regions. If the LPC based on the *n* house price differentials is non-stationary, strong convergence among all regional house prices is ruled out. *Stage Two* involves the cases where the LPC is confirmed as stationary. Using knowledge of the autoregressive parameters estimated from the standard unit root tests on the house price differentials, one may compute the half-life associated with the adjustment back towards long-run convergence and reflect on the degree of persistence in regional house price differentials.

More formally, suppose the benchmark deviations for the n regions are defined as

$$(r_i - r^*)_t = u_{it} \qquad (1)$$

where r_i denotes the (natural logarithm) house price for region i ($i = 1,\ 2, \ldots,\ n$) and r^* denotes the UK or base (natural logarithm) house price. Let X_t be an $(n\text{x}1)$ vector of random variables, namely the u_{it}'s for each of the n regional house price deviations, which may be integrated up to order one. The principal components technique addresses the question of how much interdependence there is in the n variables contained in X_t. We can construct n linearly independent principal components which collectively explain all of the variation in X_t where each component is itself a linear combination of the u_{it}'s.[3] The first principal component explains the greatest part of the variation in X_t, the second principal

[3] See, for example, Child (1970).

component (orthogonal to the first) explains the greatest part of the remaining variation, and so forth. Since I(1) variables have infinite variances, whereas stationary variables have constant variances, it follows that the LPC will be I(1), if I(1) variables are present within X_t, and so corresponds to the notion of a common trend [Stock and Watson (1988)]. However, if the LPC is I(0) then all remaining principal components will also be stationary and there are no common trends, indicating that the u_{it}'s contained in X_t are themselves stationary. The latter finding would confirm strong convergence with respect to the base across the sample of n benchmark deviations. However, if long-run convergence holds for each region with respect to the UK base then it must be the case that convergence holds between all regional pairs.

We may now consider stage one of the LPC methodology in relation to the identification of common trends. Following Stock and Watson (1988) we can argue that each element of X_t may be written as a linear combination of $k \le n$ independent common trends which are I(1), and $(n-k)$ stationary components among the u_{it}'s. The kx1 vector of common trends and $(n-k)$x1 vector of stationary components may respectively be written as

$$\tau_t = \alpha' X_t \tag{2}$$

$$\xi_t = \alpha'_\perp X_t \tag{3}$$

where each of α and $\alpha_\perp$ is an nx$(n-k)$ matrix of full column rank, $\alpha'\alpha = I$ and $\alpha'\alpha_\perp = 0$. If there are k common trends, Snell (1996) demonstrates that the k LPCs of X_t may be written as

$$\tau_t^* = X_t^{*'} \alpha^* \tag{4}$$

where X_t^* is a vector of observations on the u_{it}'s in mean deviation form, α^* represents the k eigenvectors corresponding to the largest eigenvalues of X_t and is defined as αR where R is an arbitrary, orthogonal $(k\text{x}k)$ matrix of full rank. This relationship guarantees that under the null hypothesis of k common trends, each of the k LPCs will be I(1). Similarly, for the $(n-k)$ remaining principal components, it can be shown that

$$\xi_t^* = X_t^{*'} \alpha_\perp^* \tag{5}$$

where $\alpha_\perp^*$ corresponds to the $(n-k)$ eigenvectors that provide the $(n-k)$ smallest principal components and is defined as $\alpha_\perp S$ where S is an arbitrary orthogonal $(n-k)\text{x}(n-k)$ matrix.

The LPC will be I(1) provided there is at least one common trend among the u_{it}'s contained in X_t. We can therefore test the null hypothesis that the LPC is non-stationary against the alternative hypothesis that the LPC is I(0). Rejection of the null means that all principal components are stationary and so there are no common trends among the u_{it}'s contained in X_t. This confirms strong convergence with respect to the UK base across all regions. To test the stationarity of the LPC, this study employs univariate unit root tests advocated by Elliot *et al*. (1996) and Ng and Perron (2001) that offer higher power and less size distortion relative to the more familiar ADF unit root test.

The second stage of the testing procedure involves computing the speed of adjustment associated with deviations from long-run equilibrium. Indeed, speeds of adjustment calculations are common in the general convergence literature. In this study, where stage one confirms strong convergence, the half life of a deviation from long-run equilibrium is computed as $\ln(0.5)/\ln(1+\hat{\rho})$ where $\hat{\rho}$ is the estimated autoregressive parameter from the stage one unit root tests.

Before proceeding to the results discussion, it is important to highlight some caveats associated with this methodology. The advantages over existing methods of testing for long-run regional house price convergence have been discussed above. However, the downside of this methodology concerns a standard criticism of principal component estimation and indeed of common stochastic trends. They are linear combinations of economic variables and so the economic interpretation of a given component can be problematic (although this is not central to our analysis). Also, testing the null of non-stationarity of the LPC leaves one vulnerable to the standard criticisms concerning the low power attached to unit root tests making it difficult to reject the null of non-stationarity. This is a problem with all approaches to testing stationarity. If the null is rejected, this criticism loses its relevance.

3 Data and Results

The data examined are quarterly observations on the natural logarithm of regional house prices for all properties over the period 1973-2005 using two datasets obtained from the Nationwide Bank/Building Society and Halifax Bank.[4] The Nationwide series covers the study period 1973Q4 to 2005Q1 and offers data for the following thirteen regions of the UK: North, Yorkshire and Humberside, North West, East Midlands, West Midlands, East Anglia, Outer South East, Outer Metropolitan, London, South West, Wales, Scotland and Northern Ireland plus the UK. The Halifax series is also a quarterly data set but covers a shorter period of 1983Q1 onwards. All data used in this study are mix-adjusted to allow for variations in housing quality when computing the regional or national house price

4 The Nationwide house price data are downloadable from http://www.nationwide.co.uk/default.htm while the Halifax house price data are downloadable from http://www.halifax.co.uk/home/index.shtml.

series.[5] To begin with, the analysis focuses on the Nationwide series (1973Q4-2005Q1) where we create thirteen regional-national house price ratios through the subtraction of the natural logarithm of the U.K. house price from the natural logarithm of each regional house price. Results based on the Halifax data (1983Q1 onwards) and Nationwide (1983Q1 onwards) are then used as a basis for comparison.

Pre-testing indicated that all regional 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 Ng and Perron (2001) tests are able to identify the stationarity of regional house price differentials involving the North, North West, Scotland and Northern Ireland. This implies that there are only four regions characterized by long-run cointegration with the UK average with a unity coefficient. In the case of the Elliott et al (1996) DF-GLS unit root tests, evidence is weaker still where stationarity is only confirmed for the North and North West regions. Despite the application of unit tests that are relatively more powerful than ADF unit root testing, the initial analysis points to a lack of convergence with the possibility of multiple stochastic trends driving regional house prices.

Alternatively, regional house price convergence may be examined in a multivariate setting using the familiar Johansen (1988) cointegration testing procedure which is more powerful than the regression-based tests. MacDonald and Taylor (1993) use this approach and find evidence of multiple stochastic trends driving the eleven UK regional house price levels they choose to examine over their study period of 1969-87. For the purposes of the current study, evidence of n cointegrating vectors found for the n regional house price differentials would imply that all regional house differentials are stationary. This result would be consistent with all regional house price *levels*, along with the UK house price level, being driven by a single stochastic common trend based on long-run

5 The purpose of mix adjustment is to 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.

cointegrating coefficients of unity. However, the Johansen estimates reported in Table 2 indicate that there are at most six cointegrating vectors (at the 5% significance level) and therefore, no fewer than seven stochastic trends driving the thirteen differentials. This implies the presence of no fewer than eight stochastic trends driving the regional and UK house price levels. It is possible that we may find evidence of fewer stochastic trends if the imposition of long-run homogeneity in the cointegrating vectors is relaxed and we test for a milder version of long-run convergence. Table 2 also reports that if the Johansen procedure is applied to the thirteen regional house price *levels*, rather than differentials, there is evidence of at most seven cointegrating vectors or no fewer than six stochastic trends. In this multivariate setting, relaxing homogeneity facilitates the finding of fewer stochastic trends, yet overall evidence in favor of regional house price convergence is still very weak indeed.

The central theme of this paper is that the univariate unit root and multivariate cointegration tests suffer from low test power making rejection of the null of non-stationary or non-cointegration difficult. To address this issue, we apply the LPC methodology to the regional house price differentials. The LPC explains 59.3% of the variation in regional house price differentials (using the Nationwide data).

Table 3 reports that we are able to reject the null hypothesis of non-stationarity of the LPC. Since the LPC explains the largest variation in the behavior of regional house price differentials, it is this principal component that will be non-stationary if non-stationary differentials are present. Since the LPC is stationary, it follows that all other principal components will also be stationary. This implies that all regional-national price differentials are stationary and since cointegration is a transitive concept, all bivariate regional pairs are characterized by cointegration sharing the same common stochastic trend with a long-run coefficient of unity. This is evidence of strong convergence among regional house prices.

Given the confirmation of convergence, it is of interest to examine the overall speed of adjustment of the regional-national house price ratios towards long-run equilibrium. With regard to the Nationwide data set over the 1973Q4-2005Q1 period, the LPC has a half-life of 32.4 quarters. Since we computed that the LPC explains almost 60% of the variation in regional house price differentials, the majority of variation is characterized by a very high degree of persistence. One possibility why the better known methods of unit root and cointegration testing have been unsuccessful in confirming convergence may be the high degree of persistence, indicating very long memory processes.

The results so far described are based on the use of data provided by the Nationwide Building Society over the full period of 1973Q4-2005Q1. As discussed above, there also exists an alternative mix-adjusted quarterly house price series provided by the Halifax Bank, the largest mortgage lender in the UK mortgage market. The key differences between the two series are (i) the Halifax series only begins at 1983Q1 as opposed to 1973Q4 and (ii) the UK is divided into twelve rather than thirteen regions on account of Greater London and the South East replacing London, Outer Metropolitan and Outer South East regions. To assess the robustness of our findings, we estimate using the Halifax data against the LPC methodology and compare the respective findings. As before, each region is measured as a differential against the UK average.

Initial multivariate estimation using the Johansen cointegration test indicated that there are no more than seven cointegrating vectors among the twelve regional house price differentials. This implies the existence of at least five stochastic trends among the house price differentials or six common stochastic trends driving the regional and UK house price levels. Table 3, however, reports that the LPC based on Halifax data house price differentials is stationary where the non-stationary null is rejected at the 1% significance level. The LPC methodology is able to identify strong convergence irrespective of which data series is used. Table 4 reports that the half life associated with the LPC is computed as 23.3 quarters. This suggests that the speed of adjustment is faster if one excludes data for the 1973-82 period. A similar result is obtained if we estimate for the 1983-2004 period using the Nationwide data. Again, the LPC is stationary and the associated speed of adjustment is measured as 23.7 quarters.

In addressing the faster speed of adjustment during the shorter study period, one might refer to the debate on whether or not the UK mortgage market was subject to rationing and excess demand during the 1970s and early 1980s. It is possible that relative house price adjustment may have suffered some degree of short-run impediment in adjustment towards long-run equilibrium. Moreover, some studies have argued that the use of the monetary policy "corset" during 1973-80 restricted the entry of the clearing banks into the mortgage market. In turn, the building societies cartel, abandoned in 1983, maintained mortgage interest rates at below market clearing levels with subsequent rationing of mortgage advances by non-price means such as variations in the loan-income ratio, the loan-value ratio and the period to maturity.[6]

4 Summary and Conclusion

Our study has approached the debate concerning regional house price convergence from a new perspective based on the application of principal components analysis and unit root testing. In contrast to much of the existing literature that employs more traditional unit root and cointegration testing procedures, we find in favour of regional house price convergence within the UK. This conclusion is based on finding long-run equilibrium relationships, with elasticities of unity, across the regions within a multivariate setting. In turn, this suggests that there is long-run constancy in the house price ratios between all regions. Thus house price shocks that emanate from any region(s) eventually "ripple out" to have the same multiplicative effect on all regional house prices. However, the speed of adjustment towards long-run equilibrium is slow and we calculate that it may be six to eight years before more than half the adjustment is actually achieved.

[6] However, Holmes (1993) and others conclude that rationing did not play a significant role in the provision of regional mortgage finance.

Tables

Table 1. Univariate unit root tests of house price differentials

Region	DF-GLS (no trend)	NP (no trend)
North	-2.629***	-2.994***
Yorkshire and Humberside	-1.636	-1.696*
North West	-2.157**	-2.326**
East Midlands	-1.816*	-1.773*
West Midlands	-1.282	-1.282
East Anglia	-1.657*	-1.627*
Outer South East	-1.874*	-1.857*
Outer Metropolitan	-1.741*	-1.767*
London	-1.443	-1.487
South West	-0.640	-0.732
Wales	-1.336	-1.312
Scotland	-1.891*	-2.120**
Northern Ireland	-1.912*	-2.042**

Notes for Table 1. These are Elliott et al (1996) and Ng and Perron (2001) unit roots tests (respectively denoted by DF-GLS (no trend), and NP (no trend)) on the log regional house prices minus the log UK house price. ***, ** and * denote rejection of the non-stationary null hypothesis at the 1, 5 and 10% significance levels respectively with critical values of -2.58, -1.98 and -1.62 (Ng and Perron) and -2.58, -1.94 and -1.62 (Elliot *et al*.).

Table 2. Johansen Cointegration Tests

Part A. Regional House Price Differentials

Null hypothesis: No. of cointegrating vectors	Eigenvalue	Trace Statistic	Critical value	P-value
None	0.663	610.691***	NA	NA
At most 1	0.531	476.813***	334.984	0
At most 2	0.477	383.701***	285.143	0
At most 3	0.432	303.941***	239.235	0
At most 4	0.397	234.380***	197.371	0.000
At most 5	0.323	172.159***	159.530	0.009
At most 6	0.291	124.168*	125.615	0.061
At most 7	0.188	81.857	95.754	0.305
At most 8	0.145	56.277	69.819	0.367
At most 9	0.118	37.070	47.856	0.344
At most 10	0.090	21.577	29.797	0.323
At most 11	0.069	9.910	15.495	0.288
At most 12	0.009	1.089	3.841	0.297

Part B. Regional House Price Levels

Null hypothesis: No. of cointegrating vectors	Eigenvalue	Trace Statistic	Critical value	P-value
None	0.577	598.376***	NA	NA
At most 1	0.556	492.540***	334.984	0
At most 2	0.491	392.670***	285.143	0
At most 3	0.470	309.628***	239.235	0
At most 4	0.357	231.650***	197.371	0.000
At most 5	0.322	177.394***	159.530	0.004
At most 6	0.280	129.682**	125.615	0.028
At most 7	0.206	89.326	95.754	0.128
At most 8	0.174	60.958	69.819	0.207
At most 9	0.122	37.462	47.856	0.326
At most 10	0.096	21.409	29.797	0.333
At most 11	0.066	8.997	15.495	0.366
At most 12	0.005	0.649	3.841	0.420

Notes for Table 2. These tests are Johansen Trace tests based on Nationwide house price data for the study period 1973Q4-2005Q1. ***, ** and * denote rejection of the null hypothesis at the 1, 5 and 10% significance levels respectively. MacKinnon et al (1999) *p*-values are calculated. Intercept (no trend) in CE and test VAR. Following the application of Schwarz information criteria, a lag length of 2 is employed in the VAR.

Table 3. Analysis of the LPC

LPC	DF-GLS (no trend)	NP (no trend)
Nationwide data (1973-2005)	-2.383**	-2.510**
Halifax data (1983-2004)	-2.640***	-2.933***
Nationwide (1983-2004)	-2.380**	-2.592***

See notes for Table 1.

Table 4. Speeds of Adjustment towards Long-run Convergence

LPC	$\hat{\rho}$	Half-life (quarters)
Nationwide data (1973-2005)	-0.02115	32.428
Halifax data (1983-2004)	-0.02926	23.338
Nationwide data (1983-2004)	-0.02884	23.685

These are cases where the LPC is identified as stationary and the estimates for ρ are obtained through the DF-GLS unit root test reported in Table 3. Half-lives are computed as $\ln(0.5)/\ln(1+\hat{\rho})$.

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