Convergence and Divergence in British Housing Space

Abstract

Literature on long run co-dynamics of British house prices is divided about the degree of convergence, which might be a function of the various methods of the empirically driven work and periods considered. Using the rank of the cointegrating matrix and reducibility, two overlapping super-regions emerge, characterised as the augmented North and South. With London found to be deviating from the south east, neither it nor the UK are likely to be indicative of regional series of the south of England. Moreover, with a multi-speed UK, house price convergence of any sort will be incomplete.

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Introduction

Recent literature on long run co-dynamics of British house prices is divided about the degree of convergence, in part reflecting at least three forms of [house price] regional convergence in the literature. First, β-convergence occurs when a region with a relatively low average price experiences a faster growth rate than the average of the group of regions and the gap between it and more expensive regions is reduced; low-priced regions have a higher growth rate than high-priced ones. Second, spread or σ-convergence concerns a reduction in the range of prices, so that the differential between low and high priced declines over time (COOK, 2012). These two consider the convergence hypothesis (ABBOTT and DE VITA, 2013). Third, for there to be convergence between [a pair of] price variables the differential between them should follow a stationary process, regardless of whether the variables are difference or trend stationary (PESARAN, 2007a). This focuses on the cyclical gap hypothesis (ABBOTT and DE VITA, 2013).

A multivariate cointegration approach has revealed a highly convergent system (e.g. ALEXANDER and BARROW, 1994; MACDONALD and TAYLOR, 1993). ABBOTT and DE VITA (2013), using multi-pairwise stationarity, find almost no convergent pairs. Taking a slightly different approach to stationary relations, authors assess convergence to a reference series. HOLMES (2007) and HOLMES and GRIMES (2008) use region differentials with a UK reference. Using a national reference is not without its problems. CARVALHO and HARVEY (2005) argue that a unified market exists in this approach when all regions are converging to the national average. If one is diverging, this will affect the reference national average, reducing the likelihood of finding convergence. As an alternative reference, HOLLY *et al*. (2011) find that London is the driving force behind change in the system and use that as the benchmark. In regional system models, such as cumulative causation or urban systems theory, dominant nodes drive activity in the periphery, so there is a case for this reference.

Somewhere in between complete and no convergence, MONTAGNOLI and NAGAYASU (2015) find some evidence for convergence clubs. Notably, their β-type convergence highlights London as divergent. The post-2008 crisis has led to a prolonged period of relaxed monetary policy throughout much of the developed world. BERAJA, *et al*. (2017) propose that large variations in house price growth are correlated strongly with local economic activity, so that monetary stimulus was likely to exacerbated existing regional consumption inequalities. This is particularly focused on financial centres. FERNANDEZ, *et al.* (2016) see London property as a safe-haven for a wealthy, international elite, undermining locally determined house prices. It would challenge established theory to suggest the diverging region is the principal one, yet this could be possible.

Complementing MONTAGNOLI and NAGAYASU's (2015), the aim of this paper is to consider convergence across British housing space in a cyclical gap sense. The paper addresses the following questions: are there recognisable clusters or super-regions that follow common trends? Does region-nation stationarity provide the same indication of convergence as multivariate cointegration? Can it be shown that London does not cluster with its neighbours?

The paper is structured as follows. First, there is a discussion of the convergence and divergence found in UK house price literature. Next, a bivariate procedure in unit root testing due to PESARAN (2007b) is outlined. The multivariate work entails analysing the number of stable dimensions across the cointegrating space. This is followed by data and results.

The results show that Britain behaves as a collection of linked regions, reflecting MONTAGNOLI and NAGAYASU. Two clusters emerge, envisaged as the augmented North and South. The south is characterised by DE GOEI *et al*. (2010) as a monocentric urban model, so that house price diffusion can be seen in the light of commuting-price adjustments. The group is weakly bonded, with London independent of the rest. Forecast Error Variance Decompositions point to the midlands being both a conduit of shocks between the two and, with Outer South East, drivers of common trends.

It is concluded that regional analysis in clusters of multiple regions reveals more convergence than pairwise. London is found not to co-move with the southern super-region, signifying that neither it, nor the UK price series, would make ideal reference series for other regional house price work.

Theory and Evidence of Convergence and Divergence over the Long Run

ALEXANDER and BARROW (1994) and MACDONALD and TAYLOR (1993) find multiple cointegration amongst UK regional house prices. HOLMES (2007) and HOLMES and GRIMES (2008) consider regional-national house price ratios with panel unit root tests. MEEN (1999) asserts that these ratios should not exhibit much if any of a long-term trend. In other words, the regional-national ratio should be stationary. Using principal components, HOLMES and GRIMES (2008) show all regions are converging. COOK and WATSON (2015) state that the prevalence of the finding of cointegration is to be *expected* as the divergence of regional house prices in the 1980s was succeeded by convergence in the 1990s. HOLMES (2007), using a Seemingly Unrelated Augmented Dickey-Fuller panel unit root test, finds only half of the pairs are stationary.

DIPASQUALE and WHEATON (1996) argue that the structure of local house prices is stable over time. Co-movement emerges from 'arbitrage'; buyers switching search behaviour across commuting space in the face of over or under-priced locales. Urban prices are a function of the city's export and industrial structures. Beyond the commuting shed, price differentials across large expanses of terrain would reflect distinct local economic conditions.

Three of MEEN's (1999) of oft-cited house price 'ripple' explanations rely on [potential] migration to maintain regional house price differentials, intimating migration is both an intra and inter-urban area equilibrating force. It does not explain how that price co-movement will be maintained in the long run. Critically, how can common house price trends be maintained in the face of known long term population drift from the north to south, when there is insufficient construction (ABBOTT and DE VITA, 2013)?

MEEN concludes that the dynamics of spatial differentials have little to do with cross border migration. Rather, it is similarly of the determinants of local house price that explain house price differences. This fails to recognise industrial structure, urban hierarchy or core-periphery, which all affect trend.

HAMNETT (1988) investigates the relationships between British regional housing markets over the period 1969-87. Growth rates were about the same from 1970 to 1981 and, with the exclusion of Northern Ireland, the house price rankings were stable. It is after this period that the most expensive house price regions, the South East, East Anglia and the South West, grew more rapidly than the rest, and the south delinked from the north. JRF (1995) shows an instability in a North-South East ratio from the 1960s onwards.

DRAKE (1995) finds evidence of a 'norm differential' between the house prices of the southern and midlands regions and the South East outside the Capital. He highlights the northern regions, East Anglia and Scotland as regions exhibiting divergent characteristics. Indeed, when assessing the merits of UK against the core regional reference DRAKE (1995) prefers the Rest of the South East to London.

Using PESARAN (2007a), ABBOTT and DE VITA (2013) find that there is little evidence for convergence among regional house prices. MONTAGNOLI and NAGAYASU (2015) find convergence clubs in UK housing of: the South East outer ring plus Northern Ireland; a second is of the midlands; and the third is of four Northern regions. London is distinctive in not converging.

HOLLY *et al*. (2011) apply a spatio-temporal impulse response technique to UK regional house real price diffusion. They find that London leads the UK with temporal and spatial delays (specifically, a ripple) but also, through London, New York prices are found to be weakly exogenous.

Allowing for asymmetry in cointegration, COOK (2003) reveals more dimensions of the long-run relationships between house prices in the different regions of the UK. COOK concludes that regional house price differentials mean-revert; the ratio of regional to national house prices is stationary. Dividing the 1973-2009 period by the cycles in the London house price series, COOK (2012) finds β-convergence across the regions in the downswings and weak evidence of σ-convergence is also reported, suggesting convergence is period dependent. One might conclude that, in a cyclical sense there could be a switch between convergent and divergent periods. This does not preclude that, over an extended period, the differential between London, the south and the rest is ratchet up.

Method

Standard unit root tests (Augmented Dickey-Fuller test (ADF); the tests due to PARK and FULLER (1995); KWIATKOWSKI, *et al.* (1992) (KPSS)) are applied to the regional house price variables and to the differential with the UK in levels. These are well known and will not be discussed.

PESARAN (2007a&b) shows that with panels, where there is cross-sectional dependence, the IM-PESARAN-SHIN (IPS) unit root test suffers from size distortions. To address this he proposes the Cross-sectional Augmented IPS (CIPS), a modification of the ADF. There are four stages to generating the CIPS statistic. The standard unit root ADF test procedure involves the expression:

$$\Delta x_{it} = \alpha_{i0} + (\rho_i - 1)x_{it-1} + \beta_i t + \sum_{j=1}^{p} \alpha_{ij}\, \Delta x_{it-j} + e_{it}\,. \tag{1}$$

To establish cross-sectional interdependence across the N-variables, PESARAN uses a CD statistic, generated from the residuals e_{it} in (1).

$$CD = \sqrt{\frac{2T}{N(N-1)}}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N}\hat{\rho}_{ij}\right) \sim N(0,1) \qquad \hat{\rho}_{ij} = \frac{\sum_{t=1}^{T}\hat{e}_{it}\hat{e}_{jt}}{\sqrt{\left(\sum_{t=1}^{T}\hat{e}_{it}^{2}\right)\left(\sum_{t=1}^{T}\hat{e}_{jt}^{2}\right)}}$$

If the residuals are interdependent, PESARAN recommends the generation of the Cross-sectional ADF (CADF). The CADF procedure involves the expression,

$$\Delta x_{it} = \alpha_{i0} + (\rho_i - 1)x_{it-1} + \beta_i t + \sum_{j=1}^{p} \alpha_{ij}\, \Delta x_{it-j} + (\theta - 1)\, \overline{x}_{t-1} + \sum_{j=1}^{p} \delta_j \Delta \overline{x}_{t-j} + e_{it}, \quad (2)$$

where

p is the order of the lag polynomial

$e_{it} \sim \text{iid}(0, \sigma_{ie}^2)$

t is a time trend.

$\overline{x}_{t-1}$ is the mean of all the values of x at time, $t - 1$.

The fourth stage involves the CIPS statistic, which is defined as $N^{-1}\sum_{i=1}^{N} ADF_i$, where ADF_i is the ADF statistic on the coefficient $(\rho_i - 1)$ in (2). The CIPS, like the IPS test, has a null of unit root.

If difference-stationarity is concluded among the regional price series, they are subject to tests for cointegration using the Johansen estimation procedure. Johansen's representation of a vector error correction model (VECM) of $\boldsymbol{x}_t$ can be written as,

$$\Delta \boldsymbol{x}_t = \boldsymbol{\Phi}_0 + \sum_{j=1}^{p-1} \boldsymbol{\Phi}_j\, \Delta \boldsymbol{x}_{t-j} + \boldsymbol{\Pi} \boldsymbol{x}_{t-p} + \boldsymbol{q}t + \boldsymbol{e}_t\,,$$

where

$\boldsymbol{\Phi}$, $\boldsymbol{\Pi}$ are $n \times n$ matrices of parameters,

$\boldsymbol{\Pi} = \boldsymbol{I} - \sum_{j=1}^{p} \boldsymbol{\Pi}_j$ and $\boldsymbol{I}$ is the $n \times n$ identity matrix.

Assume all the variables in $\boldsymbol{x}_t$ are difference-stationary. If the rank of the long run matrix $\boldsymbol{\Pi} = r$, where $n > r > 0$, some or all the variables are cointegrated with r cointegrating vectors. Define $\boldsymbol{\Pi} = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where both $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ are $n \times r$ matrices, so that the columns of $\boldsymbol{\beta}$ form r distinct cointegrating vectors and $\boldsymbol{\alpha}$, the speeds of adjustment. As such, $\boldsymbol{\beta}'\boldsymbol{x}_t$ is stationary. Thus, each of the r columns in $\boldsymbol{\beta}$ represents a stationary linear combination of non-stationary variables. So long as there is at least one cointegrating vector, the system of variables has an equilibrium relationship and they meander around $n - r$ common trends. The test of the null hypothesis that there are at most r cointegrating relations against n is given by $\lambda_{Trace} = -T\sum_{j=r+1}^{n} ln(1-\hat{\lambda}_j)$, also known as the Trace statistic (JOHANSEN, 1995: p.93). DICKEY *et al.* (1991) discuss multiple cointegration where $n > r > 1$. As each cointegrating vector represents a stationary linear combination, each is a dimension of the equilibrium relation. The greater the number of cointegrating relations, the more stable the n-vector model. This stability is of interest. A higher rank implies a more bonded set of variables in the cointegrating space.

The deterministic component can have both an intercept $\boldsymbol{\Phi}_0$ and trend $\boldsymbol{q}t$, which can be restricted to the cointegrating relations. This restriction allows the possibility of linear trend in regional price levels that cannot be eliminated by the cointegrating relations. A linear trend is allowed in the cointegrating relations, each of which therefore represents a stationary process plus a linear trend, or a trend stationary process (HENDRY and JUSELIUS, 2001). When the n-variables share the same stochastic and deterministic trends, a linear combination cancels out both the trends. The resulting cointegration relation is not trending. The expression $T\sum_{j=r+1}^{n} ln\{(1-\hat{\lambda}_j^*)/(1-\hat{\lambda}_j)\} \overset{a}{\sim} \chi^2(n-r)$, where * denotes the constrained model, can be used to test whether there are n unrestricted intercepts and r trends restricted to the cointegrating space against no trends and n unrestricted intercepts in the cointegrated VAR (JOHANSEN, 1995: p.97).

DAVIDSON (2000) argues that a cointegrating relation with n variables may be *reducible*, implying a sub-set of $n - j$ variables may be cointegrated. A cointegrating relation with n variables is only *irreducible* if omitting any one of them leaves the set not cointegrated. Thus, "it is legitimate to check for cointegrating relationships in sub-sets." (DICKEY and ROSSANA, 1994, p. 342). Following DAVIDSON (2000), a cointegrating space with n variables is *reducible* if a sub-set of $n - j$ variables may be cointegrated where $n > j > r > 1$. A cointegrating matrix of rank $r > 1$ with n variables is *reducible* in rank, or *weakly reducible*, if omitting any variable from n-dimentional cointegrated VAR leaves the remaining set cointegrated with rank r. A cointegrating relation with $n > r > 1$ variables is *weakly irreducible* if dropping any one of them reduces the rank. The set it *irreducible* if omitting any one variable reduces the rank to 0. GONZÁLEZ-RIVERA and HELFAND (2001) follow a similar procedure but require $n - r = 1$.

An irreducible cluster (super-region) is associated with one cointegrating vector. A housing market area is said to be integrated with a cluster if, with the addition of that house price series to the n-dimentional cointegrated VAR, there is an increase in the number of cointegrating vectors by, at least, one. Interpreting increasing rank as indicating market integration follows JONES and LEISHMAN (2006).

BÖSCHEN and MILLS' (1995) interpretation of system comprising three cointegrated variables but the first is not cointegrated with the other two on a pairwise basis is that the non-stationary aspects of the first can only be accounted for by a linear combination of the other two. A pairwise approach would fail to find cointegration if all stationary linear combinations required three variables.

The Johansen approach offers the opportunity to put linear restrictions on the cointegrating vector coefficients. The test of the constraints involves the likelihood ratio statistic, given in JOHANSEN (1995: p.107) as $T\sum_{j=1}^{r} ln\{(1-\hat{\lambda}_j^*)/(1-\hat{\lambda}_j)\} \overset{a}{\sim} \chi^2\left(r(n-s)\right)$, where s is the number columns of the restrictions matrix. If the coefficients for region i in $\boldsymbol{\beta}$ are zero, the region's prices do not form any part of any stationary linear combination within $\boldsymbol{\beta}$. As such, the region is not defined as being part of regional cluster under consideration.

[Generalised] Forecast error variance decomposition (FEVD) assigns the proportional contribution of each innovation in explaining the dependent variable's *m*-step ahead forecast error variance. It measures the relative strength of the influences of innovations due to the shocks to the dependent, and other variables in the system. PESARAN and SHIN (1998) note that because of non-zero covariance between the original innovations at a horizon, the values of the FEVDs do not add up to unity, so they should be interpreted on a relative basis. The values at time zero are viewed as instantaneous overspill, providing a measure of integration. A variable that 'explains' most of its own shock covaries little with other variables and is, therefore, relatively exogenous. Shocks in a VECM can have permanent effects, with stochastic trends, as well as transitory effects. Again, this well-known technique is not outlined.

Data

The regional data are drawn from the Nationwide Building Society's web site for the period 1973Q4 to 2015Q4. This data set is quoted widely in academic papers, such as COOK and WATSON (2015) and ABBOTT and DE VITA (2013).

The extreme values in the UK house prices series of £9,767 occur at the beginning and end of the 42 years of the study. In Table 1, the ratio of the highest to the lowest price is 20.17. In other words, house prices grew in nominal terms around twenty times.

Table 1 House Price Trends

From the growth figures, London is distinct from the south east (South West, Outer South East, East Anglia, Outer Metropolitan, London). The midlands (East and West Midlands) and the north have broadly similar rates with the North West standing out. Northern Ireland features little in the following analysis, partly because it is integrated into the Eire housing market and partly as it is not contiguous with other mainland spaces.

Unit Root Tests

An examination of unit root is undertaken using a variety of methods. In Table 2, the results from three methods on individual expressions with trend are reported for each region. The ADF results indicate that all the series are difference-stationary. This is confirmed by the ADF tests due PARK and FULLER, and the KPSS tests.

Table 2 Unit Root Tests

There is a consideration of a pairwise differential relationship for each of the regions with the nation (without trend in (2)). In the Table 3, the differential is shown to be stationary in the North, North West, East Anglia and Outer Metropolitan. These results can be compared with those of HOLMES and GRIMES (2008, Table 1). In the case of London and the Northern regions, using ADF tests, stationarity in regional house price differentials with the nation is revealed. Table 3 points to some convergence with the UK series

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Table 3 Pairwise Unit Root Tests

Using the lags highlighted in Table 3, a CD statistic of 12.37 [.000] for the residuals indicates that there is cross-sectional dependence amongst the regional markets. As such, pairwise analyses involving the CADF and the CIPS statistics is more appropriate than the IPS panel unit root test. Using the critical value of –3.23 (PESARAN, 2007b, Table Ib), the results in Table 4 suggest that only the North-UK differential is stationary with the North West a marginal, suggesting the other two were misdiagnosed as stationary. Overall, there is not a strong case for arguing convergence. This is confirmed by the CIPS value of –1.81, which is below the critical value of –2.32 (PESARAN, 2007b, Table IIb), indicating the null that all regional differentials have unit root is not rejected. In other words, contrary to HOLMES and GRIMES (2008), convergence among regional house prices to the UK average is not found.

Table 4 Cross ADF Results

Super Region Delineation: The Rank Condition

There is a consideration of clusters of regions. A traditional southern super-region (SSR) could be seen as the South West, East Anglia, Outer South East, Outer Metropolitan and London. As shown in Table 5, the null that the SSR is not cointegrated against forming one cointegrating vector is not rejected (79.60 < 87.17). The Northern super-region revealed by MONTAGNOLI and NAGAYASU (2015) of the North, Scotland, Yorkshire/Humberside, North West is found to be a cluster, with a single cointegrating vector. Their Midlands grouping, by contrast, is not.

Extending the southern group to include the East and West Midlands, the results in Table 5 indicate that this group constitutes a cluster, which is identified as the augmented southern super-region (ASSR). The addition of Wales to the northern group adds an extra vector, intimating what would be viewed as the traditional north of the North-South divide are cointegrated. Taking this one stage further to add West Midlands and East Midlands to the group, a further cointegrating vector is added. In effect, this augmented northern super-region (ANSR) combines MONTAGNOLI and NAGAYASU'S North and Midlands. Constraining the set to having no deterministic trends in the cointegrating space is rejected for both the ASSR ($\chi^2(5)$ 22.7 [.000]) and the ANSR ($\chi^2(4)$ 16.4 [.002]).

Table 5 Rank of Augmented Clusters

In Table 6, the cointegrating vector coefficients for each region are considered. The coefficients associated with London are not different from zero (p = .646) suggesting that the Capital's prices are not constrained to maintain stable relations with those of other regions. By contrast, at least one cointegrating vector coefficient for each region in the ANSR is non-zero, indicating that prices move together in the long run.

Table 6 Augmented Super-Region Coefficients

Taking this one stage further, one could consider whether the ASSR, is weakly reducible; could London be omitted? Reported in Table 5, the null that the ASSR, without London, is spanned by at most one cointegrating vector against two is rejected (96.62 > 87.17). In other words, there is no reduction in rank. From these two tests it is concluded that London does not follow common trends with rest of south east.

Super-Region Delineation: FEVD

The FEVDs are summarised at three stages: the initial period, at horizon zero; after one year; and the final estimated value at 12 years. In the main, they stabilise after 2 years, but some, such as London, often continue to decline over an extended period. Tables 7 and 8 display the FEVD values resulting from a one standard error shock to the equation of the regions of the two super-regions running horizontally. Impacts on other members of its super-region/ cluster are in columns. The largest (Initial) measure of integration favours neighbours in every case apart from London.

Commonly, there is an inverted *U*-shape in the indirect effects (overspill), suggesting an over-reaction followed by an accommodation of a shock. The long term response is smaller than in the one year impact, which is consistent with an overreaction to a shock found in SMITH and TESAREK (1991). A permanent impact would contribute to common trends. The most vigorous responder to any shock in the short and long runs in the southern cluster is the Outer South East, whereas the midlands are most active in the northern group. These could be the conduits of common trends.

There appears to be a distinction between its north facing and south facing dynamics of the midlands. The time profile for each region is similar in both super-regions, but in the north, the values are much higher. This can be explained by higher covariance with northern regions' responses, consistent perhaps with a greater number of cointegrating vectors. Also, the two clusters should have distinctive trends. Given faster growth in the south, a positive shock to the midlands would have a larger impact relatively in the north than in the south. Overall, responses to shocks to the midlands in both the north and south in the long term are relatively large suggesting that those regions are integrated within both super-regions.

Table 7 FEVD - Augmented Southern Super-Region

Table 8 FEVD - Augmented Northern Super-Region

Inferences

The bivariate and the multivariate methods proffer conflicting inferences. The CADF tests indicated that there are few cases of a region mean reverting to a stable differential with the nation, in line with ABBOTT and DE VITA (2013) concerning convergence, and at odds with MEEN (1999), HOLMES (2007) and HOLMES and GRIMES (2008).

The finding of a high order of rank is consistent with ALEXANDER and BARROW (1994) and MACDONALD and TAYLOR (1993). MONTAGNOLI and NAGAYASU's (2015) Midlands and Northern super-regions form the ANSR. The ASSR comprises the southern regions plus East and West Midlands. Even allowing for deterministic trends in the long run vectors, London is distinctive in not converging in the cyclical gap sense.

ALI *et al*. (2011) confirm that larger urban centres have greater 'footprints'. This footprint is very large in the case of London. DE GOEI *et al*. (2010) typify the patterns of interactions within the South East and the East of England as a monocentric urban structure with London as the dominant node. DIPASQUALE and WHEATON (1996) argue that commuting should bond prices across a commuting shed. As such, areas with common commuting time or distance to London could form arcs of cointegrated spaces. Cointegration is less likely to be found on a pairwise basis *because* of the expanse of the London communing shed.

CAMERON and MUELLBAUER (1998) argue that cross-border commuting disperses housing shocks among many regions. They show that the East Midlands has by far the highest rate of cross-border commuting of any region, reflecting its central geographic location and rail links, particularly with London. GRAY (2012) shows the East Midlands is a major conduit of house price diffusion in the run up to 2007. The FEVD concur, indicating that the responses to midlands' shocks have permanent effects.

Consistent with a relatively autonomous region, FEVDs and cointegration London's integration with other cluster members is modest in the longer run. The hardy perennial about the south dislocating from the north seen in, say, HAMNETT (1988) is currently one about London decoupling from the rest of the south and, hence, the rest of the UK (FERNANDEZ, *et al.*, 2016; MONTAGNOLI and NAGAYASU, 2015; TSAI, 2015). CARVALHO and HARVEY's (2005) concern about the emphasis on convergence to the national mean implicit in much of the work above is pertinent. As other regions will not converge to a long run differential with London, using London or the UK as a reference for convergence would present misleading results. Indeed, migration trends propagating greater price inequality is consistent with convergence clubs at best.

ABBOTT and DE VITA (2012) find convergence *within* Greater London is also incomplete. Notably, they find a City of London cluster, which would correspond well with the safe haven thesis, intimating that mobile capital is distorting the Central London housing market. As convergence among London boroughs is not found, and with London growing 50% more quickly than most other regions, convergence for all regions is unlikely.

Conclusion

This paper sets out to consider clustering and common trends among regions of the UK housing space. Using the rank order of cointegrating matrices, it reveals a grouping among the British regions that are classified as a super-region in the north and one in the south that overlap in the midlands of England. London does not converge with its neighbours.

Four contributions are made to price dispersion analysis across the UK. The first is that the UK housing space can be characterised as more tightly bonded northern periphery, which overlaps with a less tightly bonded southern super-region.

Second, regional relationships are found among large groups of regions, raising a question over pairwise work revealing the extent of convergence across UK housing space. Third, overlapping membership of the East and West Midlands provides a conduit of house price diffusion across UK space. Perhaps these could be used as alternative reference regions.

Lastly, the long run price spreads across the UK regions do not point to stable differentials. Treating London as the core region, hegemon or primary node in an urban hierarchy, with a multi-regional commuting shed in a monocentric urban model provides an explanation for cotrending. Concluding that London, the primary region in the UK urban hierarchy, is dislocating from the rest over the study period would undermine the case for national convergence. Regional convergence, or convergence clubs away from London though are not precluded. HOLLY *et al*. (2011) shows that pre-2008 London house prices were linked to other financial centres. It could be that, with mobile capital, this is strengthened, decoupling Central London from the rest of the country, posing both theoretical and methodological questions for regional co-movement analysis. As asset prices respond to Quantitative Easing, the post-2008 era should see stronger links between global financial capitals, and weaker convergence with local regions.

Table 1 House Price Trends

	LON	OMET	OSE	SW	EA	EM	UK
Price High	£456229	£334532	£251296	£219781	£199334	£161398	£197044
Price Low	£12848	£12863	£10871	£9605	£9998	£8637	£9767
Difference	£443381	£321669	£240425	£210176	£189337	£152761	£187276
Growth rate	35.51	26.01	23.12	22.88	19.94	18.69	20.17
	WM	WA	YH	NW	NO	SC	NI
Price High	£167185	£154969	£156429	£159062	£134534	£152479	£227970
Price Low	£9388	£8953	£9517	£8020	£7713	£8972	£7952
Difference	£157797	£146016	£146912	£151042	£126821	£143507	£220018
Growth rate	17.81	17.31	16.44	19.83	17.44	17.00	28.66

Key: EA=East Anglia, EM=East Midlands, LON=Central London, NO=North, NW=North West, OMET=Outer Metropolitan London, OSE= Outer South East, SC=Scotland, SW=South West, WA=Wales, WM=West Midlands, YH=Yorkshire/Humberside, UK=United Kingdom.

Price High=highest price over the period analysed, Price Low=lowest price, Difference=Price High – Price Low, Growth rate=Price High ÷ Price Low.

Table 2 Unit Root Tests

Region	KPSS	ADF-PF	ADF
NO	.32748*	(3) -1.9035	(3) -2.2156
YH	.29509*	(8) -2.2154	(9) -2.4231
NW	.38995*	(8) -2.4360	(8) -2.7138
EM	.35508*	(2) -1.7554	(2) -1.9748
WM	.39037*	(7) -2.1360	(7) -2.2427
EA	.35411*	(8) -2.4430	(8) -2.6974
OSE	.32904*	(9) -2.2972	(9) -2.5723
OMET	.36336*	(10) -2.6065	(10) -2.8976
LON	.27741*	(7) -2.8850	(7) -3.0882

SW	.33410*	(9) -1.7957	(9) -2.2356
WA	.27839*	(2) -1.7061	(2) -2.0969
SC	.35127*	(9) -1.6149	(8) -2.0534
NI	.19199*	(8) -2.4561	(8) -2.4137
UK	.31698*	(9) -2.0252	(9) -2.4608
CV	.15001	-3.2939	-3.4503

For regional abbreviations, see footnote to Table 1

The KPSS tests utilise a kernel of 4, which is approximately $0.75T^{1/3}$. A critical value (CV) for the KPSS tests is presented for with intercept but no trend with a null of stationarity.

The Augmented Dickey Fuller lags are selected using the AIC and are indicated by the values in parentheses

The ADF Test values are OLS *t*-ratios for the region. An indicative critical value for the ADF tests is presented for with intercept but no trend with a null of a unit root. The actual ADF CV depends on the lag length.

* indicates a rejection of a null of stationarity at the 5% level.

Table 3 Pairwise Unit Root Tests

Regions	KPSS	ADF-PF	ADF
NO	.38659	(5) -3.8024#	(5) -3.5933#
YH	2.1671*	(2) -1.6505	(2) -1.6947
NW	.51373*	(3) -3.1612#	(3) -3.0239#
EM	.62104*	(0) -2.2182	(0) -1.9245
WM	.37249	(3) -1.8545	(3) -1.5789
EA	.35146	(3) -2.9856#	(3) -2.8101
OSE	.69509*	(6) -2.5071	(6) -2.3411
OMET	.36857	(3) -2.9636#	(3) -2.8468#
LON	1.7091*	(2) -.96795	(2) -.94593
SW	1.7493*	(5) -1.6057	(5) -1.8843
WA	1.0391*	(3) -2.2434	(5) -2.5453
SC	1.4127*	(6) -3.1321#	(6) -2.9743#
NI	.47932*	(3) -2.8415#	(3) -2.5929
CV	.44178	-2.5454	-2.8217

Pairwise indicates a regional house price – national house price (in natural logs).

The KPSS tests utilise a kernel of 4, which is approximately $0.75T^{1/3}$. A critical value (CV) for the KPSS tests is presented for with intercept but no trend with a null of stationarity.

The Augmented Dickey Fuller lags are selected using the AIC and are indicated by the values in parentheses

The ADF Test values are OLS *t*-ratios for the region-nation pair. An indicative critical value for the ADF tests is presented for with intercept but no trend with a null of a unit root. The actual ADF CV depends on the lag length.

* indicates a rejection of a null of stationarity, # indicates a rejection of a null of a unit root at the 5% level.

Table 4 Cross ADF Results

Region	CADF	Region	CADF
NO	-3.2775#	OSE	-1.2909
YH	-1.4908	OMET	-2.0872
NW	-2.9010	LON	-.85186
EM	-1.3973	SW	-1.0851
WM	-1.9733	WA	-2.3917
EA	-1.2002	SC	-2.3403
NI	-1.2677	CV	-3.23

CADF is the Cross-Sectionally Augmented Dickey Fuller Tests of region-nation pairs.

The values are OLS *t*-ratios for the individual regions.

# indicates a rejection of a null of unit root at the 5% level.

A critical value (CV) for the tests, drawn from PESARAN (2007b), is presented for with intercept but no trend.

For regional abbreviations, see footnote to Table 1.

Table 5 Rank of Augmented Clusters

Super-Region	Members	Hypothesis		Trace	Critical Value
		Null	Alt.		
Midlands	WM, EM, WA	$r = 0$	$r \geq 1$	26.72*	42.34
Northern	NW, NO, SC, YH	$r = 0$	$r \geq 1$	80.39	63.00
Northern SR with Wales	NW, NO, SC, YH, WA	$r = 0$ $r \leq 1$	$r \geq 1$ $r \geq 2$	120.21 63.15	87.17 63.00
Augmented Northern	NW, NO, SC, YH, WA, EM,WM#	$r = 0$ $r \leq 1$ $r \leq 2$	$r \geq 1$ $r \geq 2$ $r \geq 3$	235.92 164.46 106.09	147.27 115.85 87.17
Southern	LON, OMET, OSE, SW, EA	$r = 0$	$r \geq 1$	79.60*	87.17
Augmented Southern	LON, OMET, OSE, SW, EA, WM, EM##	$r = 0$ $r \leq 1$ $r \leq 2$	$r \geq 1$ $r \geq 2$ $r \geq 3$	188.54 126.64 82.24*	147.27 115.85 87.17
Augmented Southern SR without London	OMET, OSE, SW, EA, WM, EM	$r = 0$ $r \leq 1$ $r \leq 2$	$r \geq 1$ $r \geq 2$ $r \geq 3$	143.42 96.62 61.90*	115.85 87.17 63.00

Trace is the trace statistic, r is the number of cointegrating vectors. These values are estimated using a VAR(7) for the ANSR and VAR(9) for the ASSR with unrestricted intercepts and restricted trends in the VARs.
* First time null is not rejected at the 5% level.
# The null that the ANSR space is spanned by two cointegrating vectors against three is rejected (106.09>87.17).
The null that the ASSR space is spanned by two cointegrating vectors against three is not rejected (82.24*<87.17).

Table 6 Augmented Super-Region Coefficients

ASSR	LR Test of Restrictions	ANSR	LR Test of Restrictions
EA	9.0832[.011]*	EM	23.2833[.000]*
EM	34.3092[.000]*	NO	36.5648[.000]*
LON	.87303[.646]	NW	32.7436[.000]*
OMET	10.1464[.006]*	SC	20.2856[.000]*
OSE	10.8019[.005]*	WA	18.5711[.000]*
SW	19.9098[.000]*	WM	31.7342[.000]*
WM	18.0353[.000]*	YH	16.8171[.001]*
Trend	9.1600[.010]*	Trend	14.3706[.002]*

The LR tests of restrictions involve placing a common restriction on a vector coefficient corresponding to a region across the r cointegrating vectors. LR statistic is distributed as $\chi^2(r)$. For ASSR $r = 2$. For ANSR $r = 3$.
p-values in brackets, *sig. at 5% level.
The null that East Anglia has no non-zero cointegrating vector coefficients is rejected at the 5% level; LR stat = 9.0832 [$p = 0.011$]
For regional abbreviations, see footnote to Table 1.

Table 7 FEVD - Augmented Southern Super-Region

Shock	Response	EA	EM	LON	OMET	OSE	SW	WM
EA	Initial	1.000	0.508	0.419	0.498	**0.595**	0.487	0.237
	1 yr	0.824	0.423	0.518	0.718	**0.865**	0.709	0.284
	Final	0.168	0.277	0.124	0.105	**0.284**	0.106	0.083
EM	Initial	**0.508**	1.000	0.383	0.460	0.472	0.402	0.293
	1 yr	0.686	0.509	0.507	0.734	**0.905**	0.794	0.357
	Final	0.251	0.341	0.181	0.198	**0.453**	0.218	0.072
LON	Initial	0.419	0.383	1.000	0.551	**0.636**	0.489	0.206
	1 yr	0.450	0.426	0.902	0.692	**0.828**	0.635	0.210
	Final	0.020	0.113	0.070	0.045	0.078	0.037	**0.171**
OMET	Initial	0.498	0.460	0.551	1.000	**0.706**	0.574	0.362
	1 yr	0.586	0.458	0.672	0.865	**0.883**	0.701	0.219
	Final	0.032	0.140	0.058	0.049	0.110	0.039	**0.159**
	Initial	0.595	0.472	0.636	0.706	1.000	**0.742**	0.338

OSE	1 yr	0.622	0.425	0.632	0.773	0.967	**0.784**	0.287
	Final	0.082	**0.217**	0.104	0.075	0.225	0.078	0.101
	Initial	0.487	0.402	0.489	0.574	**0.742**	1.000	0.383
SW	1 yr	0.638	0.408	0.564	0.757	**0.924**	0.845	0.315
	Final	0.157	0.280	0.148	0.135	**0.342**	0.151	0.075
	Initial	0.237	0.293	0.206	0.362	0.338	**0.383**	1.000
WM	1 yr	0.576	0.374	0.442	0.714	**0.832**	0.742	0.610
	Final	0.288	0.357	0.222	0.282	**0.579**	0.299	0.137

Generalized Forecast Error Variance Decompositions values are estimated using a VAR(9) with unrestricted intercepts and restricted trends in the VAR. The cointegrated vector coefficients are unconstrained. The values for the regions identified vertically resulting from a one standard error shock to the equation of the regions identified horizontally at horizons 0, one year and 12 years (Final). Highlighted in bold are the greatest row values. For regional abbreviations, see footnote to Table 1.

Table 8 FEVD - Augmented Northern Super-Region

Shock	Response	EM	NO	NW	SC	WA	WM	YH
	Initial	1.000	0.218	0.316	0.187	0.341	**0.524**	0.495
EM	1 yr	0.852	0.164	0.332	0.097	0.308	**0.788**	0.335
	Final	0.810	0.245	0.338	0.064	0.167	**0.654**	0.262
	Initial	0.218	1.000	0.250	**0.307**	0.194	0.127	0.322
NO	1 yr	**0.629**	0.642	0.389	0.264	0.501	0.525	0.455
	Final	**0.749**	0.269	0.322	0.043	0.191	0.748	0.302
	Initial	0.316	0.250	1.000	0.304	**0.329**	0.302	0.325
NW	1 yr	0.693	0.342	0.531	0.183	0.427	**0.787**	0.397
	Final	0.649	0.163	0.321	0.033	0.138	**0.852**	0.234
	Initial	0.187	**0.307**	0.304	1.000	0.203	0.063	0.237
SC	1 yr	**0.590**	0.407	0.437	0.623	0.568	0.365	0.403
	Final	**0.766**	0.289	0.348	0.094	0.338	0.772	0.298
	Initial	0.341	0.194	0.329	0.203	1.000	0.252	**0.412**
WA	1 yr	**0.698**	0.303	0.333	0.168	0.623	0.693	0.408
	Final	**0.776**	0.292	0.336	0.069	0.298	0.760	0.309
	Initial	**0.524**	0.127	0.302	0.063	0.252	1.000	0.343
WM	1 yr	**0.639**	0.108	0.326	0.066	0.279	0.933	0.266
	Final	**0.684**	0.155	0.332	0.043	0.134	0.818	0.218
	Initial	**0.495**	0.322	0.325	0.237	0.412	0.343	1.000
YH	1 yr	**0.803**	0.322	0.302	0.123	0.471	0.707	0.543
	Final	**0.800**	0.326	0.320	0.055	0.250	0.681	0.333

Generalized Forecast Error Variance Decompositions values are estimated using a VAR(7) with unrestricted intercepts and restricted trends in the VAR. The cointegrated vector coefficients are unconstrained. The values for the regions identified vertically resulting from a one standard error shock to the equation of the regions identified horizontally at horizons 0, one year and 12 years (Final). Highlighted in bold are the greatest row values. For regional abbreviations, see footnote to Table 1.