Club Convergence in Condominium Prices:
Evidence from Major U.S. Metropolitan Areas

Nicholas Apergis, Ph.D.
Professor of Economics
University of Piraeus
80 Karaoli & Dimitriou,
18534 Piraeus
Greece
napergis@unipi.gr

Research Professor
University of Derby
Derby Campus, Kedleston Road
Derby, DE22 1GB
n.apergis@derby.ac.uk

and

James E. Payne, Ph.D.
Provost and Executive Vice President for Academic Affairs
Office of Academic Affairs
Benedictine University
5700 College Road
Lisle, IL 60532
jpayne@ben.edu

January 2019

Club Convergence in Condominium Prices:
Evidence from Major U.S. Metropolitan Areas

Structured Abstract

Purpose:

The purpose of the study is to examine the long-run convergence properties of condominium prices based on the ripple effect for five major U.S. metropolitan areas (Boston, Chicago, Los Angeles, New York, and San Francisco) to identify both overall convergence in condominium prices and the possibility of distinct convergence clubs to ascertain the interdependence of geographically dispersed metropolitan condominium markets.

Design/Methodology/Approach:

Unlike methodological approaches based on the pre-testing of unit roots and cointegration in which the stationarity of the respective condominium prices comes into question, we employ the Phillips-Sul (2007; 2009) time-varying nonlinear approach.

Findings:

The results reveal the absence of overall convergence, but the emergence of two distinct convergence clubs with clear geographical segmentation: on the east coast with Boston and New York and the west coast with Los Angeles and San Francisco while Chicago exhibits a non-converging path.

Research Limitations/Implications:

The results highlight the distinct geographical segmentation of metropolitan condominium markets, which provides useful information to local policymakers, financial institutions, real estate developers, and real estate portfolio managers. The limitations of the research is the identification of the underlying sources for the convergence clubs identified due to monthly data availability for a number of potential variables.

Originality/Value:

This is the first study to examine the convergence of condominium prices relative to the literature which has focused on single-family housing.

Keywords: condominium prices; unit roots; club convergence

JEL Codes: C22, C23, R21, R31

Club Convergence in Condominium Prices: Evidence from Major U.S. Metropolitan Areas

I. Introduction

Understanding the pricing dynamics of various segments of the housing market across space and time can provide useful insights for policymakers and those intimately involved in the real estate market, ranging from financial institutions to real estate developers. The price diffusion of housing, better known as the ripple effect, across regions postulates that shocks in house prices in one region may be transmitted across other housing markets. Given that housing as an asset is heterogeneous and non-transportable, the transmission of house price shocks rests with the mobility of the population and financial capital in response to the relative house prices across regions. As outlined by Meen (1999) and Apergis and Payne (2019), factors underlying population and financial capital mobility include the migration of households, the conversion of home equity, spatial arbitrage, spatial correlation of housing determinants, and relative construction costs. Indeed, if these factors are driving the linkage of housing markets, there should be long-run convergence in regional house prices.

As noted by Apergis and Payne (2019), the convergence pattern in regional house prices reflect, in part, the differences in the sensitivities to regional demand and supply considerations. Relative house prices have an impact on household wealth, which in turn influences household consumption spending, and ultimately economic activity in a region. Also, relative house prices impact the mobility of labor through housing affordability and relocation costs. In light of the relevance to examining the convergence of house prices, and the extensive research undertaken on the subject thus far, the literature has not addressed an important segment of the housing market for metropolitan areas, namely, the condominium market. Given the premium on land usage, especially toward the city-center within large metropolitan areas, condominiums serve a critical role in meeting a metropolitan area's housing needs.[1]

Thus, our study will extend the literature on the long-run convergence in regional house prices on several fronts. First, as the first study of the condominium market, we focus our attention on five major geographically dispersed U.S. metropolitan areas (Boston, Chicago, Los Angeles, New York, and San Francisco). Second, while previous studies have primarily implemented unit root and cointegration tests, which rely on assumption of variable stationarity, we employ the time-varying nonlinear approach of Phillips-Sul (2007; 2009). The Phillips-Sul approach to identifying convergence does not depend on the specific assumptions regarding the stationarity of variables and allows one to test both overall convergence and the identification of convergence clubs across metropolitan areas. The presence of convergence clubs reflects the similar fundamental characteristics among members of the club, in our case metropolitan areas, in generating a steady state equilibrium in condominium prices.

Section 2 surveys the literature on house price convergence in the U.S. Section 3 presents the data, methodology, and results while Section 4 provides concluding remarks.

2. House Price Convergence Literature

The early literature on the ripple effect and the long-run convergence of regional house prices focused primarily on the U.K. More specifically, studies by Rosenthal (1986), Giussani and Hadjimatheou (1991), MacDonald and Taylor (1993), Alexander and Barrow (1994), Drake (1995), Ashworth and Parker (1997), Meen (1999), Cook (2003, 2005a, 2005b), Cook and

[1] As multiple unit residential structures, condominiums have several differences relative to single-family homes. First, the owners of the individual units within a condominium collectively own the shared area of the property through the condominium association with the maintenance of such areas financed through condominium fees and assessments which condominium owners are unlikely to recoup in the selling price. On the other hand, owners of single-family homes are more likely to recoup the expenses associated with home improvements when selling the property (Gil-Alana and Payne, 2019).

Thomas (2003), Holmes (2007), Holmes and Grimes (2008), Abbott and Devita (2012), and Cook and Watson (2016) examine the role of the Greater London and South East regions in the transmission of house price shocks across other U.K. regions. This line of research has also been expanded to other countries: Stevenson (2004, Ireland), Oikarinen (2006, Finland), Lou et al. (2007, Australia), Burger and Van Rensbury (2008, South Africa), Larraz-Iribas and Alfaro-Navarro (2008, Spain), Liu et al. (2008, Australia), Liu et al. (2009, Australia), Shi et al. (2009, New Zealand), Chien (2010, Taiwan), Hui (2010, Malaysia), Ma and Liu (2010; 2013a,b, Australia), Montero and Larraz (2010, Spain), Vandeenkiste and Hiebert (2011, Euro Area Countries), Lean and Smyth (2013, Malaysia), Gupta et al. (2015, European countries), and Holmes et al.(2017, Francis), and Blake and Gharleghi (2018, Australia).

With respect to the U.S. there have been a number of studies exploring the price diffusion (ie. ripple effect) and long-run convergence in house prices across census regions, states, and select metropolitan areas. Pollakowski and Ray (1997) apply both ARIMA and VAR modeling of the conventional mortgage home price indices for the nine U.S. census regions along with the five largest primary metropolitan statistical areas in the New York-Northern New Jersey-Long Island consolidated metropolitan statistical area to determine the degree of price diffusion. Their results show that shocks in house prices in one area cause subsequent shocks in house prices in other areas. The analysis of the Greater New York area reveals price diffusion between contiguous areas but not necessarily in the case of census regions. Gallet (2004) utilizes unit roots to investigate the convergence of county-level median single-family home prices for five counties (Orange, Los Angeles, Riverside, San Bernardino, and Ventura) that comprise the Los Angeles region. The results demonstrate convergence of house prices for the coastal and inland counties, but not overall convergence. Clark and Coggin (2009) use an unobserved components structural time series model to isolate the trend and cyclical components of house prices in the nine U.S. regions and through principal component analysis identify two super-regions. However, further analysis using bivariate unit root tests for absolute and relative convergence yielded mixed results.

Holmes et al. (2011) implement a pairwise testing approach to examine convergence with respect to U.S. house prices for 48 states and 81 metropolitan areas to show long-run price convergence with spatial effects a consideration in the adjustment towards long-run equilibrium. Moreover, Holmes et al. (2011) suggest that house price relationships are more likely stronger in contiguous than non-contiguous regions. Barros et al. (2012) apply fractional integration and cointegration techniques to determine the relationship between house prices in each of the 50 U.S. states relative to the overall U.S. house price index to find the absence of long-run convergence in house prices. On the other hand, Apergis and Payne (2012) employ the Phillips and Sul (2007; 2009) club convergence approach to the 50 U.S. states to find the emergence of three convergence clubs. Canarella et al. (2012) applies several unit root tests with structural breaks to the ratio of 10 metropolitan house prices to the S&P/Case Shiller Composite10 home price index to yield mixed results with respect to the stationarity of the respective house price ratios. Kim and Rous (2012) deploy the Phillips-Sul (2007) club convergence tests for 48 U.S. states and several metropolitan area panels of house prices to demonstrate that house prices do not exhibit homogenous dynamics, but distinct convergence clubs. Upon further examination Kim and Rous (2012) indicate that housing supply regulation and climate are important factors underlying the presence of convergence clubs.

Gupta and Miller (2012a) investigate the time series dynamics between house prices in Los Angeles, Las Vegas, and Phoenix. Within a vector error correction model, Gupta and Miller (2012a) reveal that Los Angeles house prices Granger-cause Las Vegas and Phoenix house prices while Las Vegas house prices Granger-cause Phoenix house prices. Los Angeles house prices are exogenous to Las Vegas and Phoenix house prices whereas Phoenix house prices do not have a statistically significant impact on either Los Angeles or Las Vegas house prices. Gupta and Miller (2012b) investigate the time series dynamics of house prices for eight Southern California metropolitan areas to reveal a long-run relationship between the metropolitan areas exhibiting causal impacts of varying degrees. Payne (2012) estimates an ARDL model for the house prices of the nine U.S. Census regions to render support for the ripple effect and long-run convergence in regional house prices, though the speed of adjustment toward long-run equilibrium varies across regions.

Barros et al. (2014) use fractional integration techniques to test the stationarity of the ratio of regional house prices relative to overall national house prices is stationary across all 50 U.S. states. Their results vary quite a bit with respect to the stationarity of the house price ratios: stationary and mean reverting in two states; nonstationary, but mean reverting in 14 states; a unit root in 13 states; and 21 states with a fractional differencing parameter larger than one. Applying measures of synchronicity and similarity to the house prices of the nine U.S. regions, Miles (2015) shows the integration of housing markets decreasing until 2005 with greater integration at the tail-end of the recent U.S. housing bubble. Using a VAR model and forecast error variance decompositions along with copula analysis, Chiang and Tsai (2016) examine the ripple effect in house prices across eight U.S. metropolitan areas. The results reveal that shocks to house prices in Los Angeles have the greatest on the house prices in Boston, Chicago, and New York with feedback effects to Los Angeles via San Francisco, which are driven by the metropolitan house prices in Boston, Chicago, and New York. Apergis and Payne (2019) test the convergence of 22 metropolitan house prices in the state of Florida to find four distinct convergence clubs that illustrates the geographical segmentation of the housing market in the state.

3.Data

We utilize the S&P/Case-Shiller Condo Price Indices for five U.S. metropolitan areas: Boston, Chicago, Los Angeles, New York, and San Francisco obtained from the St. Louis Federal Reserve Bank database, FRED II. The condominium price data is monthly and seasonally adjusted with a base year 2000 = 100 for the period 1995:1 to 2018:5. For our empirical analysis, we converted condominium prices to natural logarithms. Figure 1 displays the respective time series for condominium price indices by metropolitan area. As one can see from Figure 1, condominium prices peaked leading up to the global financial crisis in 2007-2008, declining through 2012 after which condominium prices began to increase.

[Insert Figure 1 here]

Panel A of Table 1 reports the descriptive statistics associated with the U.S. metropolitan area condominium prices. Los Angeles has the highest average condominium price and Chicago the lowest. In terms of relative variation in prices, the coefficient of variation is quite similar between Los Angeles, New York, and San Francisco followed by Boston with Chicago exhibiting the least amount of relative variation. With the exception of San Francisco, condominium prices are negatively skewed with kurtosis measures less than 3 in magnitude. Panel B of Table 1 displays the correlation matrix for the five metropolitan areas to show condominium prices in general are highly correlated.

[Insert Table 1 here]

Note that prior to the implementation of the club convergence approach, the trend component of the respective time series is extracted via the Hodrick and Prescott (1997) filter.

In an attempt to assess the hypothesis of convergence, the literature has made great efforts to develop the appropriate econometric tools. The concept of convergence pertains to whether countries/regions/sectors that are below their mean (in absolute terms or in relation to their own steady-state position) tend to go faster upwards than those that are above their mean. This concept is called β-convergence and is distinguished from another form (σ-convergence) that relates to a possible tendency for the cross-sectional dispersion of a variable to decline over time. In that strand of the literature, these two methods are applied in the case where the converging variables, i.e. condo prices, are identical across all aspects, including initial structural characteristics in relevance to this part of the real estate market. Given, however, the presence of differentiated idiosyncratic characteristics of those prices across the regions under study, the expected results imply the presence of certain clubs/groups that share these idiosyncratic characteristics and cannot be uniformly pertained to all regional condo prices. In addition, β-convergence ensures that the variance will converge to the steady-state level, but it does not say whether the variance diminishes or increases over time, while although σ-convergence typically implies that the variance of per capita output decreases over time, however, the initial variance converges to its steady-state level, which may be higher or lower than the initial level, depending on whether the dispersion of initial condo prices are lower or higher than their steady state, which is, in turn, a function of the shocks experienced by this part of the housing market.

Therefore, in this paper we attempt to consider and avoid all the above-mentioned deficiencies in order to examine the presence of different regional groups/clubs converging to a steady state, with this convergence being associated with the emergence of 'convergence clubs' as it was put forward by various researchers in the literature (Durlauf and Johnson, 1995; Galor, 1996). To this end, we employ the Phillips and Sul (2007; 2009) club convergence procedure. Unlike other convergence tests which rely on unit root and cointegration tests, the Phillips and Sul (2007; 2009) approach offers several advantages, namely, no specific assumptions regarding the stationarity of the variables and/or the presence of common factors are required. Given the heterogeneity of the time series in the panel, the Phillips and Sul (2007; 2009) approach identifies clubs (i.e., groups of metropolitan areas), each possibly converging towards a common club trend.[2]

4. Methodology

The full panel is formed by N = 5 (metropolitan areas) and T = 281. The method uses a time-varying common factor cp_{it} defined as:

$$cp_{it} = \delta_{it}\mu_t \tag{1}$$

where cp_{it} is the log of condominium prices of metropolitan area i at time t and μ_t is a common condominium price factor (i.e., common trend component in the panel). δ_{it} is a time-varying idiosyncratic component that captures both time and individual specific effects, meaning the distance between cp_{it} and the common factor, μ_t, is the common stochastic trend in the panel. In semi-parametric form, δ_{it} is given as follows:

$$\delta_{it} = \delta_i + \sigma_i\xi_{it}L(t)^{-1}t^{-\alpha} \tag{2}$$

where δ_i is fixed and $\xi_{it} \sim iid(0,1)$ across condominium prices of metropolitan areas i = 1, 2, …, N and weakly dependent over time t; σ_i is an idiosyncratic scale parameter; and $L(t)$ is a slowly varying function of time whereby $L(t) \to \infty$ and $t \to \infty$.[3] The null hypothesis of convergence can be written as follows: $H_O\colon \delta_i = \delta$ and $\alpha_i \geq 0$ against the alternative hypothesis of: $H_A\colon \{\delta_i = \delta \text{ for all i with } \alpha_i < 0\}$ or $\{\delta_i \neq \delta \text{ for some i with } \alpha_i \geq 0, \text{or } \alpha_i < 0\}$. Furthermore, cp_{it} and μ_t do not need to be trend stationary since (1) does not require either variable to be

[2] The methodological presentation parallels Kim and Rous (2012) and Apergis and Payne (2012; 2019).

[3] The parameter α is the rate at which the cross-section variation over the transitions decays over time to zero.

specified as stationary or nonstationary. Phillips and Sul (2007) employ the quadratic distance measure, H_t, as follows:

$$H_t = N^{-1}\sum_{i=1}^{N}(h_{it} - 1)^2 \tag{3}$$

where h_{it} represents the relative transition coefficient,

$$h_{it} = \frac{cp_{it}}{N^{-1}\sum_{i=1}^{N} cp_{it}} = \frac{\delta_{it}}{N^{-1}\sum_{i=1}^{N}\delta_{it}} \tag{4}$$

which captures the transition path with respect to the panel average. When there is a common or limiting transition behavior across individual metropolitan areas, $h_{it} = h_t$ across i, condominium price convergence occurs when $h_{it} \to 1$ for all i as $t \to \infty$; however, in the case that convergence does not hold, the distance remains positive as $t \to \infty$. Following Phillips and Sul (2007), we set L(t) = logt in the decay model (2) so the empirical log t regression can be used to directly test for convergence and convergence clubs as follows:

$$\log\left(\frac{H_1}{H_t}\right) - 2\log(\log t) = a + \gamma \log t + \varepsilon_t \tag{5}$$

for $t = rT, rT+1, \ldots, T$ where $r > 0$ set on the interval $[0.2, 0.3]$. For $\gamma = 2a$, the null hypothesis is conducted as a one-sided test of $\hat{\gamma} \geq 0$ against $\hat{\gamma} < 0$. To avoid estimates in (5) that may be weakly time-dependent, the least squares estimates of γ are based on heteroskedasticity and autocorrelation consistent standard errors.[4] The Phillips and Sul (2007) procedure is based on a clustering algorithm in the identification of clubs within the panel defined as follows:

(1) order the N condominium prices of the metropolitan areas based on their final values of their times series observations;

(2) starting from the highest-order metropolitan area, add adjacent metropolitan areas from the ordered list and estimate regression (5) each time followed by selecting a

[4] $-2\log(\log t)$ acts as a penalty function. The omission of this term under the alternative hypothesis would have an upward bias on the least square estimator of γ.

core group using the cut-off point criterion: $k^* = \text{ArgMax}_k\{t_{\widehat{\gamma_k}}\}$ subject to $\text{Min}_k\{t_{\widehat{\gamma_k}}\} > -1.65$, for k = 2, 3, …N;

(3) add one metropolitan area at a time of the remaining metropolitan areas to the core group, and re-estimate (5) using the sign criterion ($\hat{\gamma} \geq 0$) to determine whether a metropolitan area should join the core group; and

(4) in the case of the remaining metropolitan areas, repeat the above steps iteratively until clubs can no longer be formed.

Based on this iterative procedure each club is associated with its own convergence path. If the last group of metropolitan areas does not exhibit a convergence pattern, those metropolitan areas form a non-converging club.[5]

5. Empirical Results

Panel A of Table 2 reports the panel convergence results for r = 0.30.[6] The first row displays the results testing for overall convergence (i.e., convergence across all 5 metropolitan areas). The coefficient estimate is $\gamma =$ -0.187 with a t-statistic, $t_{\hat{\gamma}}$ = -31.884, and statistically significant at the 1% level, indicating the rejection of the null hypothesis of overall panel convergence. Based on this finding, we proceed with the algorithm to determine whether we can identify club clusters. The results from the club clustering algorithm illustrate the presence of two distinct clubs. The first club (Club 1) consists of the two east coast metropolitan areas of Boston and New York City, with $\gamma = -1.516$ and $t_{\hat{\gamma}} = -0.582$, which fails to reject the null hypothesis of convergence. The second club (Club 2) encompasses the two west coast metropolitan areas of

[5] Some clubs may be weakly divergent, i.e. $-1.65 < t_{\hat{\gamma}} < 0$.

[6] Phillips and Sul (2007) suggest that r = 0.3 is a satisfactory selection in terms of both size and power.

Los Angeles and San Francisco, with $\gamma = -1.839$ and $t_{\hat{\gamma}} = -0.506$. Finally, the non-converging club rests with the Chicago metropolitan area. Note that Phillips and Sul (2007; 2009) suggest that using the sign criterion in step (2) may lead to over-estimation of the true number of clubs. To address this potential issue, we also perform club-merging tests via regression (3) to determine whether merging adjacent numbered clubs into larger clubs is relevant. Panel B of Table 2 shows that tests of merging clubs does not support the merger of the respective convergence clubs**.**

[Insert Table 2 here]

Given the emergence of two distinct convergence clubs, we take the results from Table 2 to provide a visual representation of the relative transition paths associated with metropolitan area condominium prices for the respective convergence clubs as shown in Figures 2 and 3. Transition curves illustrate the tendency of the cluster participants, in our case condominium prices of metropolitan areas, to converge or diverge from above or below one, which is the convergence reference point as noted previously by (4). Figure 1 displays the relative transition paths of the two metropolitan areas within convergence club 1 (Boston and New York). It appears each of the metropolitan areas are converging to one from above, which confirms the presence of these two regions within the same club. The relative transition paths for the two metropolitan areas associated with convergence club 2 (Los Angeles and San Francisco) in Figure 2 comprise the West Coast. The transition paths are also quite distinct and above one in magnitude, while converging towards their mean, also confirming the convergence results on having these two regions within the same club. Finally, Figure 4 shows the relative transition paths for all five metropolitan areas, clearly indicating the convergence path of the condominium prices in the four cities under study (i.e., Boston, New York, Los Angeles and San Francisco), as well as the divergence path of Chicago condominium prices.

[Insert Figures 2-4 here]

The literature has also used typical unit root test (i.e., Augmented Dickey-Fuller) to determine the presence of convergence under the null hypothesis; however, such typical unit root testing suffers from the fact that it does not have a standard distribution (Chumacero, 2002). Consequently, making inference based on these regressions using the traditional statistics and associated critical values can lead to erroneous results (based on the fact that in the presence of a structural break, the standard/typical unit root tests are biased towards the non-rejection of the null hypothesis, Perron, 1989), unless the analysis makes use of unit root testing that explicitly considers the presence of break(s). Therefore, this part of the empirical analysis attempts to provide robust evidence on the convergence of the US metropolitan city condominium housing prices by making use of the endogenous two-break LM unit root test as derived by Lee and Strazicich (2003) in order to investigate whether the ripple effect exists in the case of these condominium housing price ratios. This particular test not only endogenously determines structural breaks, but also avoids any problems associated with bias and spurious rejections issues. Furthermore, this testing procedure considers breaks associated with changes in both the level and the trend of the variable. In addition, the two-break Lee and Strazicich (2003) procedure not only allows for the breaks to be determined endogenously from the data, but unlikely other similar tests, breaks are allowed under both the null and the alternative hypothesis (Perron, 2005). It is possible that structural breaks have occurred which might affect the result of the convergence outcome. Size properties of the Lee and Strazicich (2003)'s test are unaffected by breaks under the null. As Lee and Strazicich (2003) indicate, the test has higher power with two structural breaks, while it is unaffected by spurious rejections of the null when the series has a unit root with breaks. Hence, results applying this minimum LM test have a significant advantage, whereby the rejection of the null unambiguously shows convergence. The model of the two-break minimum LM unit root (Lee and Strazicich, 2003) can be expressed as follows. According to the LM principle, a unit root test statistic is obtained from the following equation:

$$\Delta y_t = \delta' \Delta Z_t + \varphi^d y_{t-1} + \sum_{i=1}^{k} \gamma_i \, \Delta y^d_{t-i} + \mu_t$$

where the detrended series y^d_t is defined as follows:

$y^d_t = y_t - \tilde{\upsilon}_\chi - Z_t\,\tilde{\omega}$, $t = 2, \ldots, T$; $\tilde{\omega}$ = equals the coefficients in the regression of Δy_t onto ΔZ_t; $\tilde{\upsilon}_\chi$ equals $y_1 - Z_1\,\tilde{\omega}$ correspond to the first observations of y_t and Z_t, respectively. The lagged terms ΔS_{t-i} are included to correct for serial correlation. Considering structural breaks in both the intercept and the slope of the trend function, $\Delta Z_t = [1, t, DU_{1t}, DU_{2t}, DT_{1t}, DT_{2t}]$, where DU_{1t} and DU_{2t} are indicator variables for the intercept changes in the trend function occurring at times TB_1 and TB_2, respectively, and DT_{1t} and DT_{2t} are indicator variables for the slope changes in the trend function occurring at times TB_1 and TB_2, respectively. We test the null hypothesis of a unit root in the above equation that $\varphi = 0$ with a t-ratio. Lee and Strazicich (2003) provide the critical values, which depend on the location of the breaks. The minimum LM unit-root test of Lee and Strazicich (2003) incorporates structural breaks under the null hypothesis, and the rejection of the minimum LM test null hypothesis provides genuine evidence of stationarity. Table 3 reports the results of the Lee-Strazicich unit-root test, which allows for two breaks in the constant and the trend. The findings suggest that condominium housing price ratios are nonstationary for a subset of the US metropolitan cities under consideration. In particular, the findings reject the unit-root hypothesis at the 1-percent level for the cases of Boston and New York; by contrast, the findings accept the unit root hypothesis at the 1-percent level again for the remaining three major cities, i.e. Los Angeles, Chicago and San Francisco. Moreover, the two date breaks that minimize the LM statistics provide suggestive information. The significance of the breaks is determined using a conventional t-statistic. The bust of housing prices in 2006 roughly coincides with the first break in the Lee-Strazicich procedure across all five series under consideration. The second break is clustered in around the Global Financial Crisis event characterized by the collapse of the Lehman Brothers; the results clearly indicate that this break point corresponds to the dramatic fall in house prices nationally. Therefore, the empirical results of Lee and Strazicich (2003) test clearly indicate that ripple effects exist only for the cases of Boston and New York; in other words, there is no overall convergence for the condominium house price ratios across the five major US cities under study, a finding that corroborates the results reached previously.

[Insert Table 3 here]

6. Conclusion

Though the condominium market is subject to many of the same factors that influence the supply and demand for single-family housing (i.e., financing costs, construction costs, population and job growth dynamics, migration patterns, cultural amenities, among others), there are some subtle differences as condominiums have additional costs beyond mortgage and utility expenses such as condominium fees and special assessments. However, unlike owners of single-family homes, which are likely to recoup the costs of home improvements in the sales prices, condominium owners are not likely to recoup expenditures on condominium fees and assessments in the sales price. As multiple unit structures each separately owned and, in many cases, occupied by the owner, condominiums can also be used as investment property.

While there have been a number of studies on price diffusion (i.e. ripple effect) and the long-run convergence in house prices, this study is the first to do so with respect to metropolitan condominium markets. Specifically, we apply the Phillips-Sul (2007; 2009) club convergence procedure to the condominium prices of five major metropolitan areas (Boston, Chicago, Los Angeles, New York, and San Francisco) to answer the following questions: (1) Do condominium prices across all five metropolitan markets converge to a steady state equilibrium? and (2) If overall convergence is not identified, are there distinct convergence clubs among the five metropolitan areas? The answers to these questions should be informative to local policymakers, financial institutions, real estate developers, and real estate portfolio managers. Our results illustrate that condominium prices across the five metropolitan areas do not exhibit overall convergence to a steady state equilibrium, but the presence of two distinct convergence clubs with clear geographical segmentation. The first convergence club includes Boston and New York representing the east coast while the second convergence club includes Los Angeles and San Francisco comprising the west coast. However, the condominium prices in Chicago appear to be following a non-converging path. To some extent, the geographical segmentation of the condominium markets is not surprising as one could argue that regional economic bases in the respective convergence clubs have similarities.

References

Abbott, A. and G. DeVita (2012), "Pairwise Convergence of District-Level House Prices in London", *Urban Studies*, 49, 721-740.

Alexander, C. and M. Barrow (1994), "Seasonality and Cointegration of Regional House Prices in the UK", *Urban Studies*, 31, 1667-1689.

Apergis, N. and J.E. Payne (2012), "Convergence in U.S. House Prices by State: Evidence from the Club Convergence and Clustering Procedure", *Letters in Spatial Resource Science*, 5, 103-111.

Apergis, N. and J.E. Payne (2019), "Florida Metropolitan Housing Markets: Examining Club Convergence and Geographical Market Segmentation", *Journal of Housing Research*, forthcoming.

Ashworth, J. and S.C. Parker (1997), "Modelling Regional House Prices in the U.K.", *Scottish Journal of Political Economy*, 44, 225-246.

Barros, C.P., L.A. Gil-Alana, and J.E. Payne (2012), "Comovements among U.S. State Housing Prices: Evidence from Fractional Cointegration", *Economic Modelling*, 29, 936-942.

Barros, C.P., L.A. Gil-Alana, and J.E. Payne (2014), "Tests of Convergence and Long Memory Behavior in U.S. Housing Prices by State", *Journal of Housing Research*, 23, 73-87.

Blake, J.P. and B. Gharleghi (2018), "The Ripple Effect at an Inter-Suburban Level in the Sydney Metropolitan Area", *International Journal of Housing Markets and Analysis*, 11, 2-33.

Burger, P. and L.K. Van Rensburg (2008), "Metropolitan House Prices in South Africa: Do They Converge?", *South African Journal of Economics*, 76, 291-297.

Canarella, G., S.M. Miller, and S. Pollard (2012), "Unit Roots and Structural Change: An Application to U.S. House Price Indices", *Urban Studies*, 49, 757-776.

Chumacero, R. (2002), "Is There Enough Evidence against Absolute Convergence?", Working Paper, No. 176, Santiago: Central Bank of Chile.

Chiang, M.-C. and I.-C. Tsai (2016), "Ripple Effect and Contagious Effect in the U.S. Regional Housing Markets", *Annals of Regional Science*, 56, 55-82.

Chien, M.-S. (2010), "Structural Breaks and the Convergence of Regional House Prices", *Journal of Real Estate Finance and Economics*, 40, 77-88.

Clark, S.P. and T.D. Coggin (2009), "Trends, Cycles, and Convergence in U.S. Regional House Prices", *Journal of Real Estate Finance and Economics*, 39, 264-283.

Cook, S. (2003), "The Convergence of Regional House Prices in the UK", Urban Studies, 40, 2285-2294.

Cook, S. (2005a), "Regional House Price Behaviour in the U.K.: Application of a Joint Testing Procedure", *Physica A*, 345, 611-621.

Cook, S. (2005b), "Detecting Long-Run Relationships in Regional House Prices in the U.K.", *International Review of Applied Economics*, 19, 107-118.

Cook, S. and C. Thomas (2003), "An Alternative Approach to Examining the Ripple Effect in U.K. House Prices", *Applied Economics Letters*, 10, 849-851.

Cook, S. and D. Watson (2016), "A New Perspective on the Ripple Effect in the U.K. Housing Market: Comovement, Cyclical Subsamples, and Alternative Indices", *Urban Studies*, 53, 3048-3062.

Drake, L. (1995), "Testing for Convergence between UK Regional House Prices", *Regional Studies*, 29, 357-366.

Durlauf, S. N. and P. A. Johnson (1995), "Multiple Regimes and Cross-Country Growth Behaviour", *Journal of Applied Econometrics,* 10, 365-384.

Gallet, C.A. (2004), "Housing Market Segmentation: An Application of Convergence Tests to Los Angeles Region Housing", *Annals of Regional Science*, 38, 551-561.

Galor, O. (1996), "Convergence? Inferences from Theoretical Models", *The Economic Journal*, 1056-1069.

Gil-Alana, L.A. and J.E. Payne (2019), "Persistence and Long Memory Behavior in Condominium Prices: Evidence from Major U.S. Metropolitan Areas", *Journal of Housing Research*, forthcoming.

Giussani, B. and G. Hadjimatheou (1991), "Modeling Regional House Prices in the United Kingdom", *Papers in Regional Science*, 70, 201-219.

Gupta, R. and S.M. Miller (2012a), "'Ripple Effects' and Forecasting Home Prices in Los Angeles, Las Vegas, and Phoenix", *Annals of Regional Science*, 48, 763-782.

Gupta, R. and S.M. Miller (2012b), "The Time Series Properties of House Prices: A Case Study of the Southern California Market", *Journal of Real Estate Finance and Economics*, 44, 339-361.

Gupta, R., C. Andre, and L.A. Gil-Alana (2015), "Comovement in Euro Area Housing Prices: A Factional Cointegration Approach", *Urban Studies*, 52, 3123-3143.

Hodrick, R.J. and E.C. Prescott (1997) "Postwar U.S. Business Cycles: An Empirical Investigation", *Journal of Money, Credit and Banking*, 29, 1-6.

Holmes, M.J. (2007),"How Convergent are Regional House Prices in the United Kingdom? Some New Evidence from Panel Data Unit Root Testing", *Journal of Economics and Social Research*, 9, 1-17.

Holmes, M.J. and A. Grimes (2008), Is There Long-Run Convergence among Regional House Prices in the U.K.?", *Urban Studies*, 45, 1531-1544.

Holmes, M.J., J. Otero, and T. Panagiotidis (2011), "Investigating Regional House Price Convergence in the United States: Evidence from a Pair-Wise Approach", *Economic Modelling*, 28, 2369-2376.

Holmes, M.J., J. Otero, and T. Panagiotidis (2017), "A Pair-Wise Analysis of Intra-City Price Convergence within the Paris Housing Market", *Journal of Real Estate Finance and Economics*, 54, 1-16.

Hui, H.C. (2010), "House Price Diffusion across Three Urban Areas in Malaysia", *International Journal of Housing Markets and Analysis*, 3, 369-379.

Kim, Y.S. and J.J. Rous (2012), "House Price Convergence: Evidence from U.S. State and Metropolitan Area Panels", *Journal of Housing Economics*, 21, 169-186.

Larraz-Iribas, B. and J.-L. Alfaro-Navarro (2008), "Asymmetric Behavior of Spanish Regional House Prices", *International Advances in Economic Research*, 14, 407-421.

Lean, H.H. and R. Smyth (2013), "Regional House Prices and the Ripple Effect in Malaysia", *Urban Studies*, 50, 895-922.

Lee, J., and M. Strazicich (2003), "Minimum LM Unit Root Test with Two Structural Breaks", *Review of Economics and Statistics*, 85, 1082-1089.

Liu, C., Z. Lio, L. Ma, and D. Picken (2008), "Identifying House Price Diffusion Patterns among Australian State Capital Cities", *International Journal of Strategic Property Management*, 12, 237-250.

Liu, C., L. Ma, Z.Q. Luo, and D. Picken (2009), "An Interdependence of Australian House Prices Using Variance Decomposition", *International Journal of Housing Markets and Analysis*, 2, 218-232.

Luo, Z.Q., C. Liu, and D. Picken (2007), "Housing Price Diffusion Pattern of Australia's State Capital Cities", *International Journal of Strategic Property Management*, 11, 227-242.

Ma, L. and C. Liu (2010), "The Decomposition of Housing Market Variations: A Panel Data Approach", *International Journal of Housing Markets and Analysis*, 3, 6-16.

Ma, L. and C. Liu (2013a), "Ripple Effects of House Prices: Considering Spatial Correlations in Geography and Demography", *International Journal of Housing Markets and Analysis*, 6, 284-299.

Ma, L. and C. Liu (2013b), "A Panel Error Correction Approach to Explore Spatial Correlations Patterns of the Dominant Housing Market in Australian Capital Cities", *International Journal of Housing Markets and Analysis*, 6, 405-421.

MacDonald, R. and M.P. Taylor (1993), "Regional House Prices in Britain: Long-Run Relationships and Short-Run Dynamics", *Scottish Journal of Political Economy*, 40, 43-55.

Meen, G. (1999), "Regional House Prices and the Ripple Effect: A New Interpretation", *Housing Studies*, 14, 733-753.

Miles, W. (2015), "Regional House Price Segmentation and Convergence in the U.S.: A New Approach", *Journal of Real Estate Finance and Economics*, 50, 113-128.

Monetro, J.M. and B. Larraz (2010), "Estimating Housing Prices: A Proposal with Spatially Correlated Data", *International Advances in Economic Research*, 16, 39-52.

Oikarinen, E. (2006), "The Diffusion of House Price Movements from Center to Surrounding Areas", *Journal of Housing Research*, 15, 3-28.

Payne, J.E. (2012), "The Long-Run Relationship among Regional Housing Prices: An Empirical Analysis of the U.S.", *Journal of Regional Analysis and Policy*, 42, 28-35.

Perron, P. (2005), "Dealing with Structural Breaks", Mimeo In Vol. 1, *Handbook of Econometrics: Econometric Theory*, Palgrave.

Perron, P. (1989), "The Great Crash, the Oil Price Shock, and the Unit Root Hypothesis", *Econometrica*, 57, 1361-1401.

Phillips, P.C.B. and D. Sul (2007), "Transition Modeling and Econometric Convergence Tests", *Econometrica*, 75, 1771-1855.

Phillips, P.C.B. and D. Sul (2009), "Economic Transition and Growth", *Journal of Applied Econometrics*, 24, 1153-1185.

Pollakowski, H.O. and T.S. Ray (1997), "Housing Price Diffusion Patterns at Different Aggregation Levels: An Examination of Housing Market Efficiency", *Journal of Housing Research*, 8, 107-124.

Rosenthal, L. (1986), "Regional House Price Interactions in the UK, 1975-81: A Cross-Spectral Analysis", *Applied Economics*, 18, 1011-1023.

Shi, S., N. Young, and B. Hargreaves (2009), "The Ripple Effect of Local House Price Movements in New Zealand", *Journal of Property Research*, 26, 1-24.

Stevenson, S. (2004), "House Price Diffusion and Inter-Regional and Cross-Border House Price Dynamics", *Journal of Property Research*, 21, 301-320.

Vandeenkiste, I. and P. Hiebert (2011), "Do House Price Developments Spillover across Euro Area Countries? Evidence from a Global VAR", *Journal of Housing Economics*, 20, 299-314.

Figure 1
Metropolitan Condominium Prices
(Base Year 2000=100)

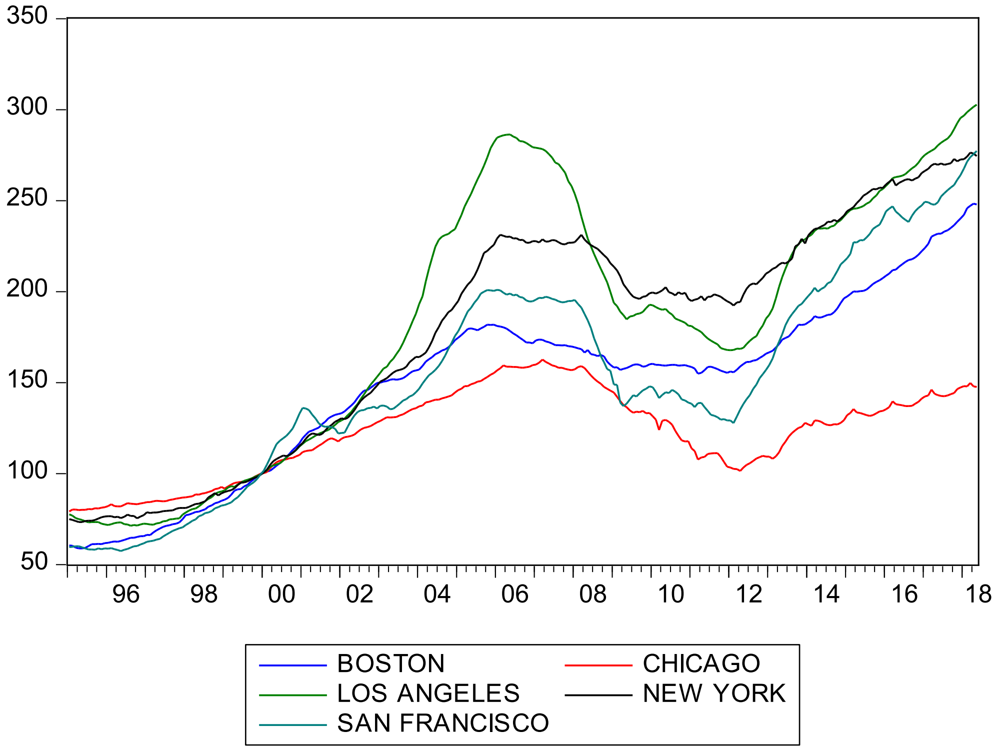

Table 1
Summary Statistics and Correlation Matrix
Condominium Price Indices

Panel A: Descriptive Statistics

Metropolitan Areas	**Mean**	**SD**	**CV**	**Min**	**Max**	**S**	**K**
Boston	149.72	48.63	32.48	58.80	248.30	-0.351	2.40
Chicago	122.60	24.12	19.67	79.11	162.65	-0.249	1.95
Los Angeles	183.25	72.84	39.75	71.27	302.64	-0.148	1.70
New York	177.56	64.38	36.26	73.26	276.38	-0.337	1.74
San Francisco	152.90	58.06	37.97	57.35	277.20	0.071	2.20

Panel B: Correlation Matrix

Metropolitan Areas	Boston	Chicago	Los Angeles	New York	San Francisco
Boston	1.000				
Chicago	0.827	1.000			
Los Angeles	0.940	0.910	1.000		
New York	0.969	0.822	0.957	1.000	
San Francisco	0.970	0.831	0.952	0.949	1.000

Notes: SD denotes the standard deviation, CV coefficient of variation, S skewness, and K kurtosis.

Table 2
Tests of Club Convergence

Panel A: Club Convergence

Full Sample (5 metropolitan areas): Boston, Chicago, Los Angeles, San Francisco, New York

$\hat{\gamma}$ = -0.187[a]
$t_{\hat{\gamma}}$ = -31.884

Club 1 (2 metropolitan areas): Boston, New York

$\hat{\gamma}$ = -1.516
$t_{\hat{\gamma}}$ = -0.582

Club 2 (2 metropolitan areas): Los Angeles, San Francisco

$\hat{\gamma}$ = -1.839
$t_{\hat{\gamma}}$ = -0.506

Non-converging club (1 metropolitan area): Chicago

Notes: Testing for the one-sided null hypothesis $\hat{\gamma} \geq 0$ against $\hat{\gamma} < 0$, the analysis makes use of the critical value $t_{0.05} = -1.65156$ across all cases. p ≤ 0.01 denoted by a.

Panel B: Merging Clubs

Clubs	$\hat{\boldsymbol{\gamma}}$	$\mathbf{t}_{\hat{\boldsymbol{\gamma}}}$
Club 1 + Club 2	-0.065	-7.16[a]

Notes: Testing for the one-sided null hypothesis $\hat{\gamma} \geq 0$ against $\hat{\gamma} < 0$, the analysis makes use of the critical value $t_{0.05} = -1.65156$. p ≤ 0.01 denoted by a.

Table 3

Lee-Strazicich Minimum LM Two-Break Unit-Root Test for the Ratios of US Metropolitan City Condominium Prices

City	Condominium prices (t-statistic)	Break dates	k
Boston	-6.79***	2006:6, 2008:10	7
New York	-7.04***	2006:6, 2008:9	6
Chicago	-5.12	2006:8, 2008:10	5
Los Angeles	-5.24	2006:7, 2008:10	8
San Francisco	-5.19	2006:7, 2008:11	8

Notes: The critical values for the unit-root test are tabulated in Lee and Strazicich (2003, Table 2). They depend upon the location of the breaks. In our case they are: -6.32 (1-percent level), -5.71 (5-percent level), and -5.33 (10-percent level). k denotes the optimal number of lagged first-differenced terms included in the unit root test to correct for serial correlation.

Figure 2
Relative Transition Curves
Convergence Club 1

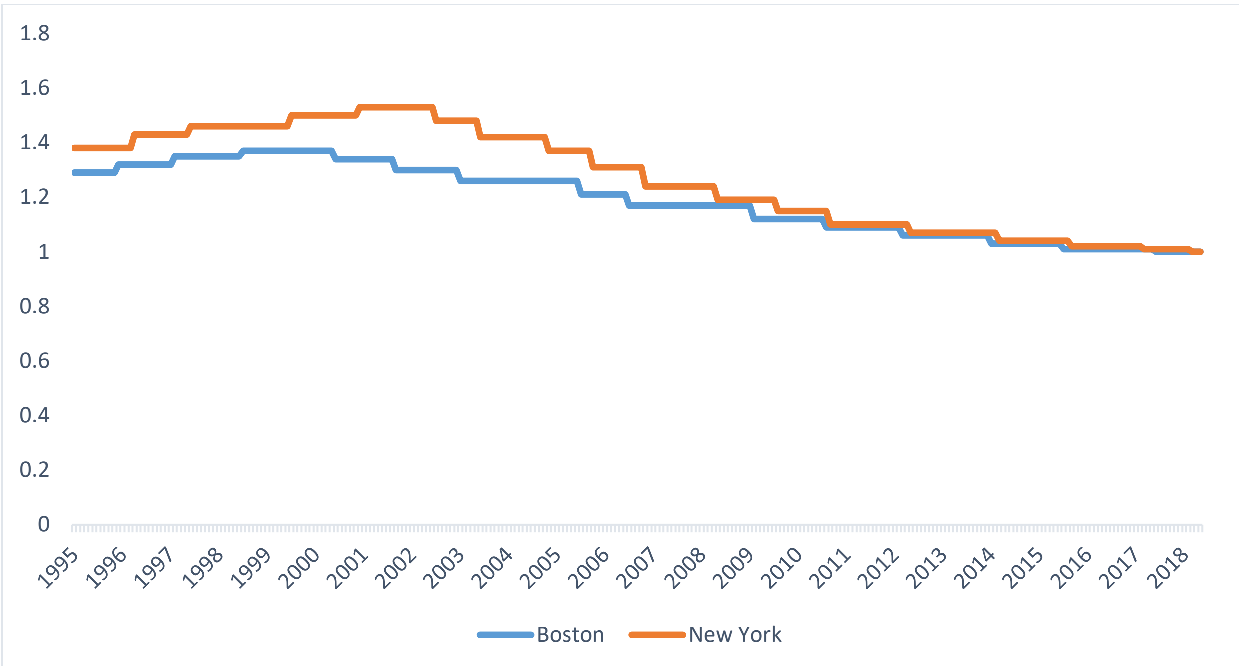

Figure 3
Relative Transition Curves
Convergence Club 2

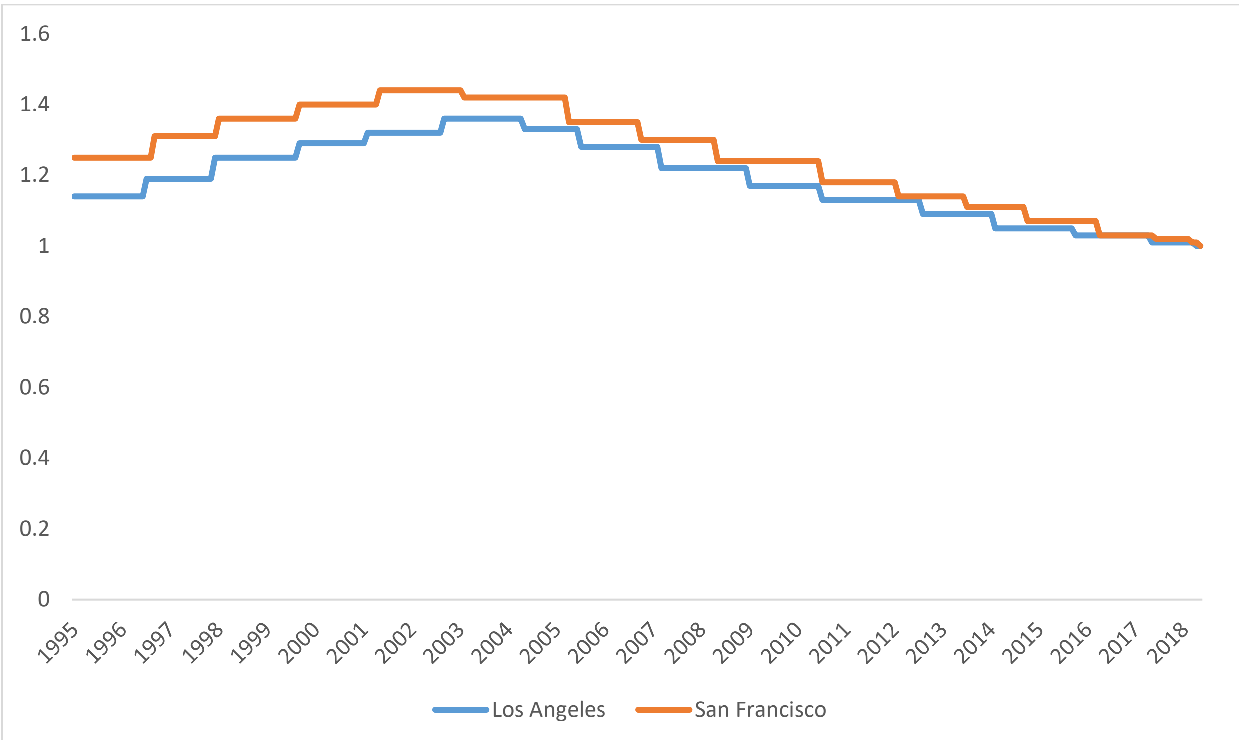

Figure 4
Relative Transition Curves
Full Sample