The Journal of
Regional
Analysis
& Policy

JRAP 42(1): 28-35.

The Long-Run Relationship among Regional Housing Prices: An Empirical Analysis of the U.S.

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Abstract. This empirical study tests the ripple effect and the long-run convergence associated with the dynamics of U.S. regional housing prices using the ARDL bounds testing approach and seasonally-adjusted monthly data from 1991:1 to 2010:12 for the nine U.S. census regions. The results support the presence of the ripple effect across regional housing markets and the long-run convergence of regional housing prices. However, the results reveal variation in the degree to which changes in regional housing prices differ across different regions in both the short run and long run. The speed of adjustment toward long-run equilibrium also varies across regions as well.

1. Introduction

The recent collapse of the U.S. housing market as a result of the financial crisis and other factors has rejuvenated interest in the transmission of shocks across regional housing markets. With housing comprising a substantial portion of household wealth, the interest in the dynamics of regional housing markets is quite timely. The transmission of shocks across regional housing markets, known as the ripple effect, occurs when shocks to regional housing prices impact all regions, causing house prices to move together in the long-run (Meen, 1999). The ripple effect in regional housing prices has been attributed to four factors: migration patterns of households, equity conversion, spatial arbitrage, and spatial patterns in the determinants of house prices. Although studies by Tirtiroglu (1992), Clapp and Tirtiroglu (1994), Gatzlaff and Tirtiroglu (1995), and Cho (1996), among others, raise questions as to the efficiency of the housing market to adjust quickly to new information, the ripple effect suggests that regional housing prices should exhibit convergence in the long run, with the speed of adjustment varying across regions.

Although there have been a number of studies examining the ripple effect and the transmission of shocks across regional housing markets, the literature on the ripple effect and long-run convergence of regional housing prices, particularly in the U.S., has yielded mixed results. This empirical study extends the literature on the ripple effect and long-run convergence for U.S. regional housing prices by using the ARDL bounds testing approach. Unlike the standard unit root and co-integration procedures used in previous studies, which require regional housing prices to be integrated of order one, the ARDL approach to co-integration does not require the variables to be integrated of the same order (Pesaran et al., 1997; Pesaran and Shin, 2001). This aspect of the ARDL approach is quite useful given the relatively low power of unit root tests. Moreover, the ARDL procedure allows for the simultaneous modeling of the short-run and long-run dynamics associated with regional housing prices. The empirical study addresses the questions: (1) To what extent are shocks to regional housing markets transmitted to other regions of the country? (2) Is there a tendency for regional housing prices to converge in the long-run? and (3) To what extent does the speed of adjustment toward long-run equilibrium vary across regions?

Section 2 briefly discusses the literature on the ripple effect and convergence. Section 3 presents the data, methodology, and results. Section 4 provides concluding remarks.

2. The ripple effect and the empirical literature

The ripple effect describes the transmission of shocks across regional housing markets resulting in long-run housing price convergence. The interplay between regional housing markets can be explained by household migration patterns, equity conversion, spatial arbitrage, and spatial patterns in the determinants of house prices. In studies by Drake (1995), Meen (1999), Cebula (2007), and Gupta and Miller (forthcoming a, b), the migration patterns of households from one region to another due to regional house price differentials provide an avenue for regional housing markets to be linked. Another source for the relationship among regional housing markets occurs when home owners borrow against the equity in their homes to purchase another home in a different region of the country. In addition to the behaviour of households in response to regional housing price differentials, spatial arbitrage may occur in which financial capital moves from higher priced regional housing markets to purchase houses in lower priced regions in anticipation of higher future prices in the lower priced regions. Another possibility is that the correlation in regional housing prices reflects the correlated movement in the underlying determinants of housing prices across regions.

Much of the literature on the ripple effect begins with an analysis of regional housing markets in the U.K. The focus of these studies is the tendency for house prices in the South East region to increase and respond faster during cyclical upturns, with prices moving outwards to other U.K. regions. Likewise, in cyclical downturns house prices in the South East region decrease sooner and faster than in other regions. Rosenthal (1986) uses a cross-spectral analysis of regional house prices with the results failing to support the ripple effect. Studies by Giussani and Hadjimatheou (1991), MacDonald and Taylor (1993), Alexander and Barrow (1994), Ashworth and Parker (1997), Cook (2005), and Holmes and Grimes (2008) employ a variety of co-integration procedures with varying results. Giussani and Hadjimatheou (1991) find evidence of a ripple effect emanating from the South East region, whereas MacDonald and Taylor (1993) show the ripple effect originating from the Greater London region. Alexander and Barrow (1994) find that regional house prices in the South influence house prices in the North and Midlands regions. Ashworth and Parker (1997) utilize spatial correlation analysis in addition to co-integration analysis to question the existence of ripple effects from the South East region. Cook (2005) reveals that the speed of adjustment is faster when house prices in the South decrease relative to other regions, but slower when house prices in the South increase. Holmes and Grimes (2008) find that regional house prices in the U.K. are driven by a single common stochastic trend, evidence of strong convergence in the long run.

Using a time-varying parameter model to test for convergence between regional housing prices in the U.K., Drake (1995) provides evidence of convergence, with house prices in the North and Scotland revealing greater divergence from prices in the South East region than in the case of the Southern and Midlands regions. In another form of testing convergence, Cook (2003) examines the stationarity of the U.K. regional/national house price ratios via asymmetric unit root tests to show convergence, with asymmetries in the adjustment process as regions in the South East experiencing faster convergence following price downswings while other regions exhibit faster convergence during price upswings. In a related study, Cook and Thomas (2003) employ non-parametric tests and business cycle dating procedures to yield support for the ripple effect in which changes in house prices in the South East region occur earlier and more extensively when compared to other regions.

Larraz-Iribas and Alfaro-Navarro (2008) examine the long-run relationship among house prices in 17 regions within Spain to find that co-integrated relationships exist among regions based on physical proximity and/or similar economic profiles. Chien (2010) tests the stationarity of the regional/national house price ratios for cities within Taiwan to show that only the regional/national house price ratio for Taipei City fails to reject the null hypothesis of a unit root, which suggests that shocks to Taipei City's regional house prices cannot ripple out across other cities.

With respect to the U.S. regional housing market, Pollahowski and Ray (1997) investigate the spatial and temporal dynamics of house prices for the nine U.S. census regions and metropolitan areas to find no difference in price diffusion patterns between neighboring and non-neighboring divisions. Zohrabyan et al. (2008) use co-integration techniques to conclude that regional house prices in the U.S. appear to be led by regions that are influential in financial and economic aspects. Clark and Coggin (2009) examine U.S. regional house price convergence using structural time series models for the nine U.S. census regions with mixed results on the convergence of regional house prices. Gupta and Miller (2012a) employ a battery of co-integration and causality tests to determine the causal relationship between house prices in Los Angeles, Las Vegas, and Phoenix. In a related study, Gupta and Miller (2012b) find that a long-run relationship exists between house prices in eight Southern California MSAs. More recently, Barros et al. (2011) utilize fractional integration and long memory models to examine the ratio of state-level housing prices to national housing prices for the U.S. to find a great deal of variation in the mean-reverting behaviour across states. Apergis and Payne (2012) use club convergence and clustering procedures to find that U.S. states fall within three convergence clubs.

Unlike previous studies which employ standard unit root and co-integration procedures, which require the respective regional house price data to be integrated of order one, this empirical study utilizes the ARDL bounds testing approach which circumvents the pre-testing of unit roots and can be applied irrespective of whether the regional house price data are integrated of order zero or one. The next section presents the data, methodology, and empirical results.

3. Data, Methodology, and Results

Monthly data from 1991:1 to 2010:12 for the regional house purchase price index for each of the nine U.S. census regions is used in the analysis.[1] The data are seasonally adjusted with a base period 1991:1 = 100. The nine census regions include the following states:

East North Central (ENC): Illinois, Indiana, Michigan, Ohio, and Wisconsin;
East South Central (ESC): Alabama, Kentucky, Mississippi, and Tennessee;
Middle Atlantic (MA): New Jersey, New York, and Pennsylvania;
Mountain (M): Arizona, Colorado, Idaho, Montana, Nevada, New Mexico, Utah, and Wyoming;
New England (NE): Connecticut, Maine, Massachusetts, New Hampshire, Rhode Island, and Vermont;
Pacific (P): Alaska, California, Hawaii, Oregon, and Washington;
South Atlantic (SA): Delaware, Florida, Georgia, Maryland, North Carolina, South Carolina, Virginia, and West Virginia;
West North Central (WNC): Iowa, Kansas, Minnesota, Missouri, Nebraska, North Dakota, and South Dakota;
West South Central (WSC): Arkansas, Louisiana, Oklahoma, and Texas.

Figure 1 displays the respective U.S. regional house price indices. As one can see from Figure 1, regional housing prices appear to exhibit co-movement along with the decrease in regional housing prices and heightened volatility in the aftermath of the recent financial crisis.

The natural logarithms of U.S. regional housing prices are denoted in lower case and used in the analysis. This ARDL bounds testing procedure begins by estimating a conditional ARDL-ECM model as follows:

$$\begin{aligned}\Delta enc_t &= \alpha_0 + \textstyle\sum_{i=1}^{p}\gamma_i \Delta enc_{t-i} + \sum_{i=0}^{p}\delta_i \Delta esc_{t-i} \\ &+ \textstyle\sum_{i=0}^{p}\phi_i \Delta ma_{t-i} + \sum_{i=0}^{p}\theta_i \Delta m_{t-i} \\ &+ \textstyle\sum_{i=0}^{p}\lambda_i \Delta ne_{t-i} + \sum_{i=0}^{p}\psi_i \Delta p_{t-i} + \sum_{i=0}^{p}\varpi_i \Delta sa_{t-i} \\ &+ \textstyle\sum_{i=0}^{9}\varphi_i \Delta wnc_{t-i} + \sum_{i=0}^{p}\zeta_i \Delta wsc_{t-i} \\ &+ \pi_1 enc_{t-1} + \pi_2 esc_{t-1} + \pi_3 ma_{t-1} + \pi_4 m_{t-1} \\ &+ \pi_5 ne_{t-1} + \pi_6 p_{t-1} + \pi_7 sa_{t-1} \\ &+ \pi_8 wnc_{t-1} + \pi_9 wsc_{t-1} + \varepsilon_t \qquad (1)\end{aligned}$$

where α_0 is the drift component and ε_t is the random error. The same specification is undertaken for each region by replacing the dependent variable with the specific region of interest and adjusting the lag length associated with the dependent variable on the right-hand side. The null hypothesis of no co-integration (H_0: $\pi_1 = \pi_2 = \cdots = \pi_9 = 0$) is tested against the alternative hypothesis of co-integration (H_A: $\pi_1 \neq 0$; $\pi_2 \neq 0$; .... ; $\pi_9 \neq 0$) via a partial F-test. However, as discussed by Pesaran et al. (2001), the asymptotic distribution of the F-statistic is non-standard, regardless of whether the variables are I(0) or I(1). Pesaran et al. (2001) provide lower and upper bound critical

[1] The data source is the Economagic database, www.economagic.com.

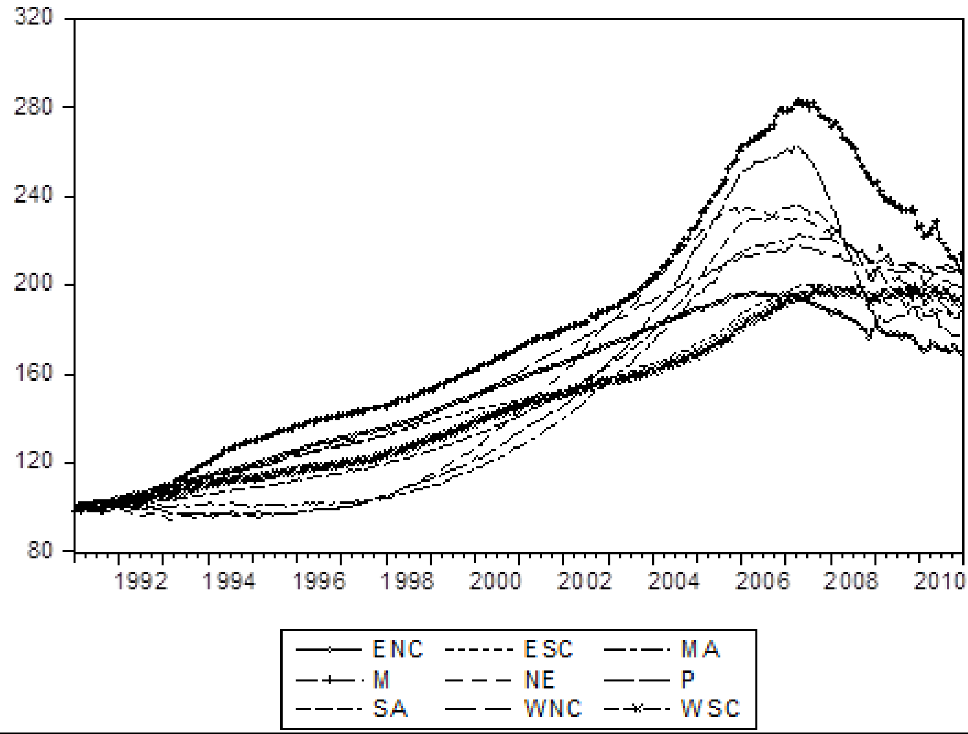

Figure 1. U.S. regional housing price indices, 1991:1 to 2010:12. Base period 1991:1 = 100.

values, where the lower bound critical values assume all variables are I(0) while the upper bound critical values assume all variables are I(1). If the calculated F-statistic is above the upper critical value, the null hypothesis of no co-integration can be rejected irrespective of the orders of integration of the variables. If the calculated F-statistic is below the lower critical value, the null hypothesis of no co-integration cannot be rejected. However, if the calculated F- statistic falls between the lower and upper critical values, the result is inconclusive.

Table 1 displays the partial F-statistics associated with the null hypothesis of a long-run relationship among the regional housing prices along with the asymptotic critical values of the bounds testing procedure. As shown in Table 1, the F-statistic exceeds the upper critical bound at the 1% significance level with respect to housing prices for each region. The results clearly indicate there is a long-run cointegrating relationship among U.S. regional housing prices. Table 2 reports the long-run coefficients with respect to each region. Given that the regional housing prices are in natural logarithms, the coefficients can be interpreted as long-run elasticity estimates.[2] A cursory view of Table 2 reveals quite a bit of heterogeneity in the response of a region's housing prices to the housing prices in other regions. The statistically significant long-run elasticity estimates range from -2.031 to 3.450. In the long-run, the housing prices in the East South Central, New England, and South Atlantic regions impact six regions, while the Pacific region impacts housing prices in only two regions, the East South Central and South Atlantic regions.

The error-correction models are reported in Table 3. The statistically significant short-run elasticity estimates range from -0.175 to 0.510. In the short run, the housing prices in the Mountain, South Atlantic, and West South Central regions impact six regions, while the East North Central, Pacific, and West North Central regions impact housing prices in only three regions. In addition to the variation in the short-run and long-run elasticity estimates of regional housing price changes, there is also variation in the speed of adjustment to restore long-run equilibrium across regions.[3] The Pacific region exhibits the slowest adjustment toward long-run

[2] The long-run relationship is based on the autoregressive distributed lag (ARDL) model using 1 lag based on Schwarz Bayesian criterion (SBC) with estimates of the asymptotic standard errors of the long-run coefficients computed by the regression approach by Bewley (1979).

[3] The speed of adjustment is defined as the reciprocal of the absolute value of the coefficient of the error correction term, ect (-1).

equilibrium, 12.66 months, while the West North Central region yields the fastest adjustment, 1.95 months. The average speed of adjustment across the regions is 5.83 months. Furthermore, the error-correction models show predictive power as evidenced by the statistical significance of the overall F-statistics for the respective models.

Table 1. ARDL bounds testing procedure tests of long-run relationships.

	F-statistic
$F(enc \mid esc, ma, m, ne, p, sa, wnc, wsc)$	6.4703[a]
$F(esc \mid enc, ma, m, ne, p, sa, wnc, wsc)$	4.3962[a]
$F(ma \mid enc, esc, m, ne, p, sa, wnc, wsc)$	7.5154[a]
$F(m \mid enc, esc, ma, ne, p, sa, wnc, wsc)$	10.0072[a]
$F(ne \mid enc, esc, ma, m, p, sa, wnc, wsc)$	6.4365[a]
$F(p \mid enc, esc, ma, m, ne, sa, wnc, wsc)$	5.3274[a]
$F(sa \mid enc, esc, ma, m, ne, p, wnc, wsc)$	6.9965[a]
$F(wnc \mid enc, esc, ma, m, ne, p, sa, wsc)$	6.2082[a]
$F(wsc \mid enc, esc, ma, m, ne, p, sa, wnc)$	3.7973[b]

Notes: Critical value bounds for the partial F-statistics with intercept and no trend for the number of variables k = 9: 99%: I(0) 1.899, I(1) 3.047; 95%: I(0) 2.163, I(1) 3.349; 90%: I(0) 2.716, I(1) 3.989. Significance levels: a(1%), b(5%), and c(10%).

4. Concluding remarks

This empirical study tests the presence of the ripple effect and the long-run convergence of U.S. regional housing prices using the ARDL bounds testing procedure. Specifically, the study answers the following three questions. First, to what extent are shocks to regional housing markets transmitted to other regions of the country? The results indicate that shocks to regional housing markets are not uniform, with some regions having a larger impact than others regardless of geographical proximity in both the short run and long run. Second, is there a tendency for regional housing prices to converge in the long-run? The results from the ARDL bounds testing procedure suggest that regional housing markets are co-integrated, thus lending support for long-run convergence. Third, to what extent does the speed of adjustment toward long-run equilibrium vary across regions? The results show that the speed of adjustment varies across regions, ranging from 1.95 months to 12.66 months, with the average speed of adjustment 5.83 months.

Whereas there has been a great deal of research on the differential impacts of fiscal and monetary policy on regional housing markets, an avenue for future research resides in understanding the determinants of the underlying adjustment process to restore equilibrium across regional housing markets. Understanding such determinants will assist local, state, and national policymakers in the design of the appropriate policy responses to disruptions in the housing market.

Table 2. Long-run estimates.

Independent Variables	Dependent Variable								
	enc	*esc*	*ma*	*m*	*ne*	*p*	*sa*	*wnc*	*wsc*
intercept	1.29	-0.068	0.753	-2.031	-1.486	0.057	0.244	-0.718	1.084
	[0.000][a]	[0.737]	[0.225]	[0.000][a]	[0.001][a]	[0.966]	[0.330]	[0.000][a]	[0.002][a]
enc	-----	0.14	-0.812	1.17	0.449	-1.028	-0.012	0.768	-0.9
		[0.364]	[0.068][c]	[0.034][b]	[0.223]	[0.348]	[0.953]	[0.000][a]	[0.001][a]
esc	0.233	-----	-0.839	-1.062	-1.308	-0.663	0.682	0.103	1.073
	[0.096][c]		[0.091][c]	[0.021][b]	[0.000][a]	[0.445]	[0.000][a]	[0.177]	[0.000][a]
ma	-0.296	-0.005	-----	0.531	0.001	-1.078	0.17	0.238	-0.15
	[0.000][a]	[0.949]		[0.066][c]	[0.993]	[0.020][b]	[0.026][b]	[0.000][a]	[0.288]
m	0.15	-0.111	-0.399	-----	-0.857	-1.386	0.294	0.09	0.216
	[0.046][b]	[0.107]	[0.123]		[0.000][a]	[0.012][b]	[0.000][a]	[0.024][b]	[0.108]
ne	0.154	-0.381	0.078	-1.284	-----	-0.147	0.296	0.112	0.31
	[0.047][b]	[0.000][a]	[0.762]	[0.000][a]		[0.766]	[0.000][a]	[0.007][a]	[0.024][b]
p	0.023	-0.088	-0.197	0.212	0.002	-----	0.186	0.028	0.128
	[0.603]	[0.025][b]	[0.143]	[0.213]	[0.987]		[0.000][a]	[0.252]	[0.126]
sa	-0.011	0.655	1.459	0.905	1.203	3.45	-----	-0.441	-0.464
	[0.946]	[0.000][a]	[0.001][a]	[0.041][b]	[0.000][a]	[0.000][a]		[0.000][a]	[0.106]
wnc	0.935	0.283	1.465	0.444	1.101	2.649	-0.235	-----	0.556
	[0.000][a]	[0.122]	[0.009][a]	[0.448]	[0.010][a]	[0.060][c]	[0.274]		[0.103]
wsc	-0.463	0.522	0.084	0.521	0.721	-0.821	-0.434	0.256	-----
	[0.000][a]	[0.000][a]	[0.827]	[0.124]	[0.003][a]	[0.308]	[0.002][a]	[0.000][a]	

Notes: Probability values are in brackets. The Schwarz Bayesian criterion is used in the selection of one lag in the ARDL specification. Significance levels: a(1%), b(5%), c(10%).

Table 3. ARDL Error Correction Models.

Independent Variables	Dependent Variable Δenc	Δesc	Δma	Δm	Δne	Δp	Δsa	Δwnc	Δwsc
Δenc	-----	0.05	-0.083	0.152	0.092	-0.082	0.51	0.393	0.041
		[0.368]	[0.162]	[0.019][b]	[0.230]	[0.332]	[0.000][a]	[0.000][a]	[0.596]
Δesc	0.064	------	0.086	0.065	0.003	0.219	0.196	0.053	0.192
	[0.100][c]		[0.070][c]	[0.389]	[0.977]	[0.012][b]	[0.000][a]	[0.186]	[0.000][a]
Δma	-0.082	-0.002	-----	0.069	2.9E-3	0.128	0.049	-0.004	0.209
	[0.000][a]	[0.949]		[0.026][b]	[0.993]	[0.174]	[0.063][c]	[0.946]	[0.002][a]
Δm	0.041	-0.04	-0.041	-----	-0.175	0.197	0.306	0.046	0.039
	[0.041][b]	[0.119]	[0.096][c]		[0.000][a]	[0.010][a]	[0.000][a]	[0.029][b]	[0.122]
Δne	0.043	-0.012	0.008	-0.167	-----	-0.012	0.296	0.057	0.056
	[0.049][b]	[0.812]	[0.765]	[0.000][a]		[0.763]	[0.000][a]	[0.011][b]	[0.042][b]
Δp	0.006	0.094	0.077	0.027	3.3E-3	-----	0.053	0.014	0.023
	[0.600]	[0.040][b]	[0.070][c]	[0.153]	[0.987]		[0.000][a]	[0.248]	[0.119]
Δsa	0.333	0.235	0.149	0.444	0.246	0.274	-----	-0.028	-0.083
	[0.000][a]	[0.000][a]	[0.004][a]	[0.000][a]	[0.001][a]	[0.000][a]		[0.588]	[0.117]
Δwnc	0.258	0.101	-0.029	0.058	0.225	0.21	-0.067	-----	0.1
	[0.000][a]	[0.126]	[0.712]	[0.446]	[0.013][b]	[0.026][b]	[0.318]		[0.145]
Δwsc	0.065	0.187	0.205	0.068	0.147	0.178	-0.125	0.131	-----
	[0.199]	[0.000][a]	[0.001][a]	[0.138]	[0.005][a]	[0.051][c]	[0.000][a]	[0.000][a]	
ect(-1)	-0.276	-0.358	-0.102	-0.13	-0.204	-0.079	-0.288	-0.512	-0.179
	[0.000][a]	[0.000][a]	[0.000][a]	[0.000][a]	[0.000][a]	[0.000][a]	[0.000][a]	[0.000][a]	[0.000][a]
adj. R^2	0.569	0.380	0.498	0.544	0.425	0.592	0.663	0.535	0.219
F	36.11	17.4	27.57	32.75	20.69	39.76	53.29	31.62	8.63
	[0.000][a]	[0.000][a]	[0.000][a]	[0.000][a]	[0.000][a]	[0.000][a]	[0.000][a]	[0.000][a]	[0.000][a]

Notes: Probability values are in brackets. The Schwarz Bayesian criterion used in the selection of one lag in the error correction specification. Significance levels: a(1%), b(5%),c(10%)

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