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# Price Convergence Patterns across U.S. States

**Summary:** This study examines the convergence patterns of prices across 50 U.S. states over the period 1960-2007, by applying the convergence algorithm developed by Peter C. B. Phillips and Donggyu Sul (2007). The empirical findings suggest the rejection of full convergence across the 50 U.S. states' prices, and the presence of 11 subgroups, or convergence clubs. The main implications of this paper point to the low degree of market integration across the U.S. states, the limitations of using a unique national price deflator to calculate real U.S. state variables, and the different effects that national monetary policy decisions will have on U.S. state prices.

**Key words:** Price convergence, Club convergence, U.S. states.

**JEL:** C33, E31, R10.

According to Purchasing Power Parity (PPP), based on the Law of One Price, in the absence of transport costs and trade restrictions, competitive markets will equalise the price of an identical good in different countries/regions when the prices are expressed in the same currency (Kenneth Rogoff 1996). At an international level, PPP implies that exchange rates would adjust to offset inflation differentials across countries, while within a country, at a regional level, it implies that regional prices would be equal in the long-run; that is, regional price differentials would be temporary and would tend to disappear in the long-run. However, this implication is difficult to reconcile with the empirical data. In fact, Regional Price Parities (RPPs), which measure the differences in the price levels of goods and services across U.S. states, estimated by the Bureau of Economic Analysis (2016)[1] show that the real purchasing power is 34 percent greater in Mississippi than in Hawaii, suggesting large price differences across U.S. states. Actually, in 2013 the states with the highest RPPs were Hawaii (with a price level of

[1] **Bureau of Economic Analysis.** 2016. Regional Price Parities by State (U.S.=100). http://www.bea.gov/newsreleases/regional/rpp/rpp_newsrelease.htm (accessed May 10, 2016).

116.2, compared with a national average of 100), New York (115.3), New Jersey (114.5) and California (112.3), while the states with the lowest RPPs were Mississippi (86.8), Arkansas (87.5), South Dakota (87.6), Alabama (87.7) and West Virginia (88.4). Among the states with RPPs closest to the national average price level were Vermont (100.2), Illinois (101.0), Florida (98.8) and Oregon (98.7).

In this context, the objective of this paper is to examine the convergence patterns of 50 U.S. states' prices over the period 1960-2007, by employing the convergence algorithm developed by Phillips and Sul (2007). The three main contributions of this paper are the following. First, we tested for price convergence across U.S. states. Although the price convergence across U.S. cities has been studied already in several papers (see the above-mentioned literature), price convergence across U.S. states has not been analysed yet, so that, to the best of our knowledge, this is the first paper that analyses price convergence across U.S. states. Second, based on the idea that the degree of convergence has varied over the analysed period, we tested for convergence using the nonlinear varying coefficients factor model developed by Phillips and Sul (2007), which provides the framework for modelling transitional dynamics as well as long-run behaviour. Finally, this methodology allowed us to determine endogenously the existence of different convergence clubs among the different states in the sample.

The remainder of this paper is organised as follows. Section 1 reviews the literature and Section 2 describes the econometric methodology. Section 3 presents the data and discusses the empirical results. Finally, Section 4 summarises and concludes the study.

## 1. Literature Survey

The discrepancy between the theoretical predictions and observed international data explains why PPP is one of the most studied topics in international economics (Jeffrey A. Frankel and Andrew K. Rose 1996; Yangru Wu 1996; David H. Papell 1997, 2006; Ru-Lin Chiu 2002; Chi-Young Choi 2004a). One of the main result of these studies is that, although relative prices revert to a common mean, the rate of mean reversion is quite slow, with a half-life (defined as the time required for any deviation from PPP to dissipate by one half, and commonly used to measure the speed of convergence) estimate of four or five years. Although we would expect to observe a more rapid price convergence across regions or across cities within a country than at an international level, empirical studies for the U.S. case show evidence against regional price convergence, or when convergence is found, they show that the speed of convergence in some cases can be greater than at the international level (Stephen G. Cecchetti, Nelson C. Mark, and Robert J. Sonora 2002; Hiranya K. Nath and Jayanta Sarkar 2009, 2013). There is a large body of literature that has analysed price convergence across a sample of U.S. cities, using different time series techniques (David C. Parsley and Shang-Jin Wei 1996; Cecchetti, Mark, and Sonora 2002; Phillips and Sul 2007; Syed A. Basher and Josep L. Carrion-i-Silvestre 2009, 2011; Viera Chmelarova and Nath 2009, 2010; Nath and Sarkar 2009; Ege Yazgan and Hakan Yilmazkuday 2011; Ho-Chuan Huang, Wei-Han Liu, and Chih-Chuan Yeh 2012; Young Se Kim and Jeffrey J. Rous 2012; Natalie D. Hegwood and Nath 2013). For example, Parsley and Wei (1996) analyse price convergence using a panel of 51 prices (41 tradable and 10 non-tradable goods) from 48 U.S. cities. They find that prices of tradable goods converge quickly, while this is not the case for non-traded goods. Yazgan and Yilmazkuday (2011) extend the analysis by Parsley and Wei (1996) by analysing the bilateral price-level convergence of 48 different goods across 52 U.S. cities for a longer period, 1990:Q1-2007:Q4, and find evidence of convergence at fast rates (their estimated half-lives range from 1.37 for traded goods to 2.75 for non-traded goods).

Cecchetti, Mark, and Sonora (2002) analyse price convergence using annual consumer price indices (CPIs) in 19 major U.S. cities for the period 1918-1995 using panel unit root tests (Andrew T. Levin, Chien-Fu Lin, and Chia-Sang J. Chu 2002; Kyung So Im, M. Hashem Pesaran, and Yongcheol Shin 2003) and find that the divergence across U.S. cities is temporary, although persistent, with a half-life of nearly 9 years. The slow rate of convergence obtained in this study has been revisited in the literature, applying different techniques to the same or very similar datasets. For example, Nath and Sarkar (2009) also use CPI data for a panel of 17 U.S. cities between 1918 and 2006 and, after correcting for different sources of bias (time aggregation bias and small-sample bias), obtain estimates of half-life estimates of 7.5 years, which, although smaller than those obtained by Cecchetti, Mark, and Sonora (2002), also suggest the existence of slow price convergence rates. On the contrary, Basher and Carrion-i-Silvestre (2009, 2011) also revisit the empirical analysis by Cecchetti, Mark, and Sonora (2002), analysing price convergence among 17 U.S. cities for the period 1918-2005, and find that, allowing for structural breaks, U.S. city price level differentials are stationary with an estimated half-life of convergence equal to between 1.5 and 2.65 years, much lower than that obtained in Cecchetti, Mark, and Sonora (2002) and Nath and Sarkar (2009). A similar result is obtained by Hegwood and Nath (2013), who analyse relative price convergence among a sample of 17 U.S. cities during the period 1918-2010 by means of applying panel unit root tests and allowing for the presence of structural breaks. Their main results suggest that when allowing for a structural break in 1985, the speed of convergence they obtain is faster than when structural breaks are not accounted for. In fact, they obtain an estimated half-life of 3.9 years.

While the previous studies investigate relative price convergence by analysing the time series properties of the city price levels relative to the national price average (they do not deal with the problem of the numeraire city), Chmelarova and Nath (2010) show that the choice of the numeraire or base city is a key factor for explaining the price convergence results and the half-life estimates. For example, when using Atlanta, Chicago or Los Angeles as the numeraire cities, the authors find evidence of price convergence and lower half-life estimates than those in Cecchetti, Mark, and Sonora (2002).

The slow rates of convergence obtained in some of the previous studies (Cecchetti, Mark, and Sonora 2002; Nath and Sarkar 2009) are interpreted by Phillips and Sul (2007) as evidence against convergence towards a unique index. They develop a time varying factor model and analyse convergence in the cost of living indices among 19 U.S. cities using annual data from 1918 to 2001, and they obtain evidence against convergence towards a unique index. On the contrary, they find evidence of three different convergence subgroups. The same methodology is used by Kim and Rous (2012), who study house price convergence in 48 U.S. states and 119 metropolitan areas using quarterly data from 1975:Q1 to 2009:Q2, and find that there is little evidence of overall convergence and strong evidence of multiple convergence clubs. The non-trading characteristics of the housing markets is one of the factors explaining the low evidence of convergence found when analysing house price convergence.

## 2. Econometric Methodology

In this section we outline the methodology proposed by Phillips and Sul (2007) (henceforth PS) to test for convergence in a panel of countries/states/regions and to identify convergence clubs. PS propose a new econometric approach for testing the convergence hypothesis and for identifying convergence clubs. Their method uses a nonlinear time-varying factor model and provides the framework for modelling transitional dynamics as well as long-run behaviour. Furthermore, their statistical methodology can test for convergence in economic variables other than output.

Let $y_{it}$ denote a time series index $i$ at time $t$. The new methodology adopts the following simple time-varying common-factor representation for $y_{it}$ of country $i$:

$$y_{it} = \delta_{it}\mu_t, \tag{1}$$

where $\mu_t$ is a single common component and $\delta_{it}$ is a time-varying idiosyncratic element that captures the deviation of state $i$ from the common path defined by $\mu_t$. Within this framework, all $N$ states will converge to the steady state at some point in the future, if $\lim_{k\to\infty} \delta_{it+k} = \delta$ for all $i = 1, 2, ..., N$, irrespective of whether states are near the steady state or in transition. This is important given that the paths to the steady state (or states) can differ significantly across U.S. states.

The goal of PS is to test whether economic variables $y_{it}$, $i = 1, 2, ..., N$, tend to converge to a single steady state as $t\to\infty$. To this end they adopt a factor representation $y_{it} = \delta_{it}\mu_t$ (Equation 1) for each economic variable in the sample. The factor $\mu_t$ is assumed to be common among individuals (economies), while the transition dynamics are captured by the idiosyncratic components $\delta_{it}$, which are allowed to vary across time and cross section. Convergence is a dynamic process. Since $\delta_{it}$ traces the transition paths, convergence can be tested by examining the temporal relative evolution of $\delta_{it}$. PS do not assume any parametric form for $\mu_t$; they simply factor it out and concentrate on $\delta_{it}$.

Since we cannot directly estimate $\delta_{it}$ from Equation (1), due to over-parameterization, i.e. the number of parameters is greater than the number of observations, PS assume a semiparametric form for $\delta_{it}$, which enables them to construct a formal test for convergence. In particular, they eliminate the common component $\mu_t$ through rescaling by the panel average:

$$h_{it} = \frac{y_{it}}{\frac{1}{N}\sum_{i=1}^{N} y_{it}} = \frac{\delta_{it}}{\frac{1}{N}\sum_{i=1}^{N} \delta_{it}}. \tag{2}$$

The relative measure $h_{it}$ captures the transition path with respect to the panel average. Defining a formal econometric test of convergence as well as an algorithm for defining club convergence requires the following assumption for the semiparametric form for the time-varying coefficients $\delta_{it}$:

$$\delta_{it} = \delta_i + \sigma_{it}\xi_{it}, \tag{3}$$

where $\sigma_{it} = \frac{\sigma_i}{L(t)t^{\alpha}}$, $\sigma_i > 0$, $t \geq 0$, and $\xi_{it}$ is weakly dependent over $t$, but is iid(0,1) over $i$. The function $L(t)$ (in this paper we set $L(t)$ = log $t$) varies slowly, increasing and diverging at infinity. Under this particular form for $\delta_{it}$, the null hypothesis of convergence for all $i$ takes the form $H_0 : \delta_i = \delta, \alpha \geq 0$, while the alternative hypothesis of non-convergence for some $i$ takes the form $H_A : \delta_i \neq \delta$ or $\alpha < 0$. PS show that we can test for the null hypothesis of convergence using the framework of the following regression (Appendix B of PS shows the analytic proof under the convergence hypothesis for this regression equation):

$$\log\left(\frac{H_1}{H_t}\right) - 2\log L(t) = \hat{c} + \hat{b}\log t + \hat{u}_t, \tag{4}$$

for $t = [rT], [rT] + 1, ..., T$, and $r > 0$. Following the recommendation of PS, we set $r$ = 0.2. Finally, only a $(1 - r)$ fraction of the sample is used for the regression. In this regression, $H_t = \frac{1}{N}\sum_{i=1}^{N}(h_{it} - 1)^2$ and $\hat{b} = 2\hat{\alpha}$, where $h_{it}$ is defined in Equation (2) and $\hat{\alpha}$ is the least squares estimate of $\alpha$. Under the null hypothesis of convergence, the dependent variable diverges whether $\alpha > 0$ or $\alpha = 0$. In this case, we can test the convergence hypothesis by a $t$-test of the inequality $\alpha \geq 0$. The $t$-test statistic follows the standard normal distribution asymptotically and is constructed using a heteroskedasticity and autocorrelation consistent standard error. PS call the one-sided $t$-test, which is based on $t_{\hat{b}}$, the log $t$ test due to the presence of the log $t$ regressor in Equation (4). The null hypothesis of convergence is rejected at the 5% level if the log $t$ test is less than -1.65.

The empirical convergence literature also deals with the possible existence of multiple equilibriums. In this case, rejection of the null hypothesis, that all states in the sample converge, does not imply the absence of convergence clubs in the panel. In this study we implement the club convergence and clustering procedure proposed by PS, which can be summarised as follows: (1) order the $N$ countries with respect to the last-period value of the time series; (2) form all possible core (club) groups $C_k$ by selecting the first $k$ highest states, with $k$ = 2, 3, ..., $N$, then test for convergence using the log $t_k$ test within each subgroup of size $k$; finally, define the core club $C^*$ of size $k^*$ as the club for which the maximum computed log $t_{k^*}$ statistic occurs, given that the log $t_k$ statistic supports the convergence hypothesis; (3) from the remaining $N$-$k^*$ states, add one state at a time to the core club $C^*$, and test for convergence through the log $t$ test. If the test strongly supports the convergence hypothesis (log $t \geq 0$), then add the state to the group $C^*$. Find all countries that, according to the log $t$ test, converge to the same steady state as the core group $C^*$; these states, together with the states of the core group $C^*$, form the first convergence club in the panel; (4) then, for the remaining states (if any), repeat the procedure described in steps 1-3 to determine the next convergence club, if one exists. Finally, terminate the procedure when the remaining economies fail to converge. However, since the sieve criterion ($\log t \geq 0$) set in step 3 is highly conservative, the club convergence and clustering procedure tends to find more clubs than the true number. In order to avoid such overestimation, PS suggest running log*t* test regressions across the subgroups to provide support for merging clubs into larger clubs.

## 3. Data and Empirical Results

### 3.1 Data

We collected annual observations of a cost of living index (COL index) for the 50 U.S. states for the period 1960-2007 (the median cost of living for 2007 was 100, the average index value for the two "middle states", New Mexico and Wyoming), developed first by William D. Berry, Richard C. Fording, and Russel L. Hanson (2000, 2009) and available through the IQSS data archive (Harvard Dataverse 2016)[2]. As explained in Berry, Fording, and Hanson (2000), the COL index uses the family budgets corresponding to an intermediate standard of living, calculated by the Bureau of Labor Statistics (BLS), which are comparable across space and time. The family budgets tally the costs of food, housing, transportation, clothing, personal items, medical care, and taxes for an urban family of four (see BLS 1976). The authors have assessed the quality and reliability of the state COL index, and conclude that there is strong evidence for the validity of this index.

Table 1 reports the basic statistics of state price levels for each of the 50 U.S. states. The highest average price levels correspond to Connecticut, Hawaii and Massachusetts, while the lowest average prices are those of Montana, West Virginia and Arkansas. This data shows that, on average, real purchasing power in Montana has been 41 percent greater than in Connecticut. Furthermore, the highest price level was observed in Hawaii, and the lowest price level was observed in Montana.

**Table 1** Descriptive Statistics

| | Mean | St. dev. | Min | Max |
|---|---|---|---|---|
| Alabama | 42.42 | 26.01 | 8.00 | 90.38 |
| Alaska | 50.98 | 28.20 | 13.94 | 109.44 |
| Arizona | 45.94 | 27.42 | 10.89 | 101.53 |
| Arkansas | 42.13 | 25.40 | 8.00 | 88.55 |
| California | 50.82 | 32.61 | 10.20 | 119.41 |
| Colorado | 48.29 | 29.86 | 11.30 | 112.15 |
| Connecticut | 58.92 | 34.61 | 14.75 | 124.92 |
| Delaware | 48.61 | 27.50 | 13.79 | 100.41 |
| Florida | 43.88 | 26.43 | 9.28 | 93.08 |
| Georgia | 43.76 | 26.94 | 8.35 | 93.99 |
| Hawaii | 57.83 | 35.52 | 13.15 | 135.11 |

[2] **Harvard Dataverse.** 2016. http://dvn.iq.harvard.edu/dvn/ (accessed May 15, 2016).

| | | | | |
|---|---|---|---|---|
| Idaho | 45.57 | 26.85 | 10.89 | 99.62 |
| Illinois | 49.91 | 29.52 | 11.79 | 103.13 |
| Indiana | 47.80 | 28.10 | 11.24 | 98.06 |
| Iowa | 48.03 | 27.67 | 11.37 | 97.43 |
| Kansas | 48.37 | 27.85 | 11.63 | 97.86 |
| Kentucky | 42.61 | 25.84 | 8.33 | 90.53 |
| Louisiana | 42.88 | 25.76 | 8.66 | 90.77 |
| Maine | 52.75 | 30.80 | 12.92 | 110.96 |
| Maryland | 47.89 | 28.52 | 10.72 | 101.61 |
| Massachusetts | 56.97 | 35.08 | 13.15 | 127.35 |
| Michigan | 48.53 | 28.88 | 11.15 | 100.67 |
| Missouri | 50.11 | 29.47 | 11.55 | 102.99 |
| Montana | 41.79 | 25.32 | 7.27 | 88.10 |
| Nebraska | 47.91 | 28.05 | 11.42 | 97.65 |
| Nevada | 45.66 | 26.53 | 11.34 | 98.87 |
| New Hampshire | 48.84 | 28.03 | 12.20 | 98.88 |
| New Jersey | 49.60 | 27.92 | 14.36 | 107.66 |
| New Mexico | 56.12 | 32.71 | 14.08 | 118.40 |
| New York | 57.06 | 34.35 | 13.42 | 124.28 |
| North Carolina | 45.28 | 27.14 | 10.65 | 99.76 |
| North Dakota | 54.22 | 32.96 | 12.57 | 119.18 |
| Ohio | 43.31 | 26.91 | 7.92 | 93.60 |
| Oklahoma | 49.06 | 27.57 | 11.67 | 97.87 |
| South Carolina | 48.07 | 28.46 | 11.10 | 98.90 |
| South Dakota | 42.65 | 25.06 | 9.36 | 88.57 |
| Tennessee | 47.10 | 28.94 | 11.00 | 108.64 |
| Texas | 50.52 | 30.23 | 11.47 | 109.05 |
| Utah | 56.17 | 32.68 | 14.16 | 118.32 |
| Vermont | 43.06 | 26.36 | 7.93 | 91.79 |
| Virginia | 48.73 | 27.85 | 12.21 | 98.22 |
| Washington | 42.93 | 26.29 | 8.12 | 91.49 |
| West Virginia | 42.03 | 25.58 | 8.22 | 88.96 |
| Wisconsin | 46.81 | 28.52 | 11.16 | 107.24 |
| Wyoming | 54.42 | 31.23 | 13.99 | 114.50 |

**Source:** Authors' calculations.

## 3.2 Empirical Results

It is well known that using consumer price indices (CPI) for testing convergence may lead to misleading evidence of convergence due to the base year problem in CPI time series. To solve this problem, we followed the suggestion of Phillips and Sul (2007): for each one of the time series we took the first observation in our sample as the base year and discarded the first 20% of the observations.

Table 2 reports the results of applying the PS convergence and clustering procedure. The first row of Panel A reports the result for the test of the hypothesis that all countries converge to a single steady state; the convergence hypothesis is rejected at any reasonably levels of significance.

**Table 2** Club Convergence Results

| Subgroup | States | *t*-statistic | *b* coefficient |
|---|---|---|---|
| | **Panel A: club convergence** | | |
| Full sample | Alabama, Alaska, Arizona, Arkansas, California, Colorado, Connecticut, Delaware, Florida, Georgia, Hawaii, Idaho, Illinois, Indiana, Iowa, Kansas, Kentucky, Louisiana, Maine, Maryland, Massachusetts, Michigan, Minnesota, Mississippi, Missouri, Montana, Nebraska, Nevada, New Hampshire, New Jersey, New Mexico, New York, North Carolina, North Dakota, Ohio, Oklahoma, Oregon, Pennsylvania, Rhode Island, South Carolina, South Dakota, Tennessee, Texas, Utah, Vermont, Virginia, Washington, West Virginia, Wisconsin, Wyoming | -14.742*** | -0.914 |
| 1st subgroup | Alabama, Georgia, Mississippi, North Carolina, South Carolina, Tennessee | -0.809 | -0.026 |
| 2nd subgroup | Arkansas, California, Texas | 0.821 | 0.050 |
| 3rd subgroup | Louisiana, Virginia | 1.192 | 0.174 |
| 4th subgroup | Florida, Washington, West Virginia | 0.725 | 0.123 |
| 5th subgroup | Arizona, Colorado, Maryland, New Mexico, Oklahoma, Oregon, Utah | -0.752 | -0.079 |
| 6th subgroup | Idaho, Massachusetts, Michigan, Pennsylvania | 2.828 | 0.147 |
| 7th subgroup | Hawaii, Illinois, Indiana, Iowa, Kansas, Minnesota, Missouri, New York, Ohio, Wisconsin | -0.667 | -0.037 |
| 8th subgroup | Maine, Montana, Nebraska, New Jersey, North Dakota, South Dakota | 1.066 | 0.364 |
| 9th subgroup | New Hampshire, Rhode Island | -0.307 | -0.061 |
| 10th subgroup | Connecticut, Delaware, Vermont, Wyoming | 1.007 | 0.031 |
| 11th subgroup | Alaska, Nevada | 25.367 | 1.098 |
| Non-converging | Kentucky | | |
| | **Panel B: subgroup merging** | | |
| | 1st subgroup + 2nd subgroup | -9.606*** | -0.163 |
| | 2nd subgroup + 3rd subgroup | -2.367*** | -0.085 |
| | 3rd subgroup + 4th subgroup | -3.687*** | -0.247 |
| | 4th subgroup + 5th subgroup | -4.442*** | -0.328 |
| | 5th subgroup + 6th subgroup | -2.340*** | -0.292 |
| | 6th subgroup + 7th subgroup | -2.537*** | -0.122 |
| | 7th subgroup + 8th subgroup | -52.532*** | -0.926 |
| | 8th subgroup + 9th subgroup | -5.953*** | -0.362 |
| | 9th subgroup + 10th subgroup | -8.494*** | -0.330 |
| | 10th subgroup + 11th subgroup | -5.083*** | -0.284 |
| | 11th subgroup + non-converging country | -29.869*** | -1.056 |

**Notes:** *, **, *** denote the rejection of the null hypothesis of convergence at the 10%, 5% and 1% significance levels, respectively.

**Source:** Authors' calculations.

The remaining rows of Panel A show the results obtained from a direct application of the clustering algorithm described above. The algorithm classifies the U.S. states into 11 subgroups (clubs), with Alabama, Georgia, Mississippi, North Carolina, South Carolina and Tennessee in the first; Arkansas, California and Texas in the second; Louisiana and Virginia in the third; Florida, Washington and West Virginia in the fourth; Arizona, Colorado, Maryland, New Mexico, Oklahoma, Oregon and Utah in the fifth; Idaho, Massachusetts, Michigan and Pennsylvania in the sixth; Hawaii, Illinois, Indiana, Iowa, Kansas, Minnesota, Missouri, New York, Ohio and Wisconsin in the seventh; Maine, Montana, Nebraska, New Jersey, North Dakota and South Dakota in the eighth; New Hampshire and Rhode Island in the ninth; Connecticut, Delaware, Vermont and Wyoming in the tenth; and finally, Alaska and Nevada in the eleventh. According to the PS algorithm, Kentucky is the only state that fails to converge to any of the existing steady states.

Panel B reports the results of the test conducted to determine whether any of the original subgroups reported in Panel A can be merged to form larger convergence clubs. These results clearly suggest that no merging can take place. Since no further merging seems possible, we accept the result from the PS clustering procedure reported in Panel A of Table 2.

## 3.3 Convergence among Housing Prices

In order to better understand the lack of price convergence across U.S. states, this subsection analyses housing prices, which are considered to be one of the most significant determinants of geographical differences in the cost of living. For this purpose, we used seasonally adjusted quarterly data for the period 1975:Q1 to 2016:Q3, and applied the same methodology used in the previous section. The data we used, obtained from the Federal Housing Finance Agency (FHFA 2016)[3], is the All Transactions Index, which adds prices from appraisal data to sales transactions from mortgage data.

At this stage it is important to explain why we could expect convergence in house prices across the U.S. states, even when housing is a non-tradable good. Two houses in two different locations are believed to be sold within the same market if house prices in one location impose a competitive constraint on house prices in the other location (Massimo Motta 2004; Dennis W. Carlton and Jeffrey M. Perloff 2005). In the long-run, given this competitive constraint, it is unreasonable to expect that house prices in different states will diverge indefinitely. However, if the prices do tend to diverge, we can conclude that the different housing markets operate as mutually independent local markets and that their prices should diverge. In light of this discussion, two questions arise. First, do the heterogeneity and spatial fixity of houses – implying non-tradability – not undermine the basis of arbitrage among the different geographical regions, leading to separate housing markets? The answer to this is that, even though houses are heterogeneous, they all provide an unobservable and non-tradable commodity called a "housing service"; therefore one can ignore the physical

[3] **Federal Housing Finance Agency (FHFA).** 2016. Datasets. https://www.fhfa.gov/DataTools/Downloads/Pages/House-Price-Index.aspx (accessed December 15, 2016).

heterogeneity (Lawrence B. Smith, Kenneth T. Rosen, and George Fallis 1988). Given that houses are non-tradable, a second question that arises is: why would one expect prices of such non-tradable services to converge across geographical areas? As Philippe Burger and Lizelle Janse Van Rensurg (2008) indicate, the housing market is comprised of institutional and wealthy individual investors who often have investments in more than one property, which are utilised to earn rental income and capital appreciation. Understandably, this is an attempt to reduce risk or to balance risk and return. Given this, if property prices in one area diverge too far from those in another location, an arbitrage opportunity will arise (William Nelson Goetzmann 1993; Joaquim J. Montezuma 2004).

Table 3 reports the results of applying the PS convergence and clustering procedure. As in the previous table, the first row of Panel A reports the result of testing the hypothesis that all prices converge to a single steady state, and suggests that the convergence hypothesis is rejected at any reasonably level of significance. The remaining rows of Panel A suggest the existence of 7 subgroups (clubs), with California, Hawaii, Massachusetts, New Jersey, Oregon and Washington in the first; Colorado, Connecticut, Maine, Maryland, Montana, New Hampshire, New York, Rhode Island, Utah, Vermont and Virginia in the second; Arizona, Illinois, Minnesota, New Mexico, Pennsylvania, Wisconsin and Wyoming in the third; Alaska, Delaware, Idaho, Iowa, Kentucky, Louisiana, Missouri, Nebraska, North Carolina, North Dakota, Oklahoma, South Carolina, South Dakota, Tennessee and Texas in the fourth; Florida, Michigan and Nevada in the sixth; and Alabama, Arkansas, Indiana, Kansas and Ohio in the seventh. According to the PS algorithm, Georgia is the only state that fails to converge to any of the existing steady states.

**Table 3** Club Convergence Results in Housing Prices (Quarterly, Seasonally Adjusted Data)

| Subgroup | States | *t*-statistic | *b* coefficient |
|---|---|---|---|
| | **Panel A: club convergence** | | |
| Full sample | Alabama, Alaska, Arizona, Arkansas, California, Colorado, Connecticut, Delaware, Florida, Georgia, Hawaii, Idaho, Illinois, Indiana, Iowa, Kansas, Kentucky, Louisiana, Maine, Maryland, Massachusetts, Michigan, Minnesota, Mississippi, Missouri, Montana, Nebraska, Nevada, New Hampshire, New Jersey, New Mexico, New York, North Carolina, North Dakota, Ohio, Oklahoma, Oregon, Pennsylvania, Rhode Island, South Carolina, South Dakota, Tennessee, Texas, Utah, Vermont, Virginia, Washington, West Virginia, Wisconsin, Wyoming | -107.967*** | -1.125 |
| 1st subgroup | California, Hawaii, Massachusetts, New Jersey, Oregon, Washington | 10.700 | 0.081 |
| 2nd subgroup | Colorado, Connecticut, Maine, Maryland, Montana, New Hampshire, New York, Rhode Island, Utah, Vermont, Virginia | 3.207 | 0.434 |
| 3rd subgroup | Arizona, Illinois, Minnesota, New Mexico, Pennsylvania, Wisconsin, Wyoming | 3.558 | 0.092 |
| 4th subgroup | Alaska, Delaware, Idaho, Iowa, Kentucky, Louisiana, Missouri, Nebraska, North Carolina, North Dakota, Oklahoma, South Carolina, South Dakota, Tennessee, Texas | 5.654 | 0.193 |
| 5th subgroup | Florida, Michigan, Nevada | 0.691 | 0.029 |
| 6th subgroup | Alabama, Arkansas, Indiana, Kansas, Ohio | 3.729 | 0.225 |
| 7th subgroup | Mississippi, West Virginia | 4.396 | 0.300 |

**Panel B: subgroup merging**

| | | |
|---|---|---|
| 1st subgroup + 2nd subgroup | -7.700*** | -0.432 |
| 2nd subgroup + 3rd subgroup | -8.894*** | -0.312 |
| 3rd subgroup + 4th subgroup | -22.573*** | -0.501 |
| 4th subgroup + 5th subgroup | -9.587*** | -0.185 |
| 5th subgroup + 6th subgroup | -3.210*** | -0.711 |
| 6th subgroup + 7th subgroup | -1.947** | -0.298 |
| 7h subgroup + non-converging country | -1.746** | -0.318 |

**Notes:** *, ** and *** denote the rejection of the null hypothesis of convergence at the 10%, 5% and 1% significance levels, respectively.

**Source:** Authors' calculations.

## 4. Concluding Remarks

The main objective of the paper is to examine the price convergence patterns across 50 U.S. states over the period 1960-2007 by employing the convergence algorithm developed by Phillips and Sul (2007). Although the price convergence across U.S. cities has been extensively studied in the literature, price convergence across U.S. states has not been previously analysed to the best of our knowledge, this is the first paper to study global price convergence across the 50 U.S. states. We believe that the use of state-level data for analysing price convergence complements the literature on price convergence across U.S. cities in different ways. First, it may help us understand how representative the sample of U.S. cities that is used extensively in the literature is for testing price convergence, given that this sample of cities does not represent all 50 U.S. states. Second, even for those states in which city prices are available, these prices are not always representative of state-level prices, especially for those states with a higher proportion of rural areas. In this context, the conclusions of this paper will allow us to analyse the degree of market integration across all of the U.S. states. Furthermore, this study will allow us to understand how prices across U.S. states adjust to monetary policy interventions. The main results are the following.

First, the empirical results suggest the rejection of full convergence of relative prices across the 50 U.S. states. According to the literature, this result could be explained by different factors, such as distance, transportation costs, the inclusion of non-traded goods in regional price indices (i.e. house prices), the differences in prices between urban and rural states, and the differential speeds of adjustment to small and large shocks (Cecchetti, Mark, and Sonora 2002). In this context, it would be interesting to distinguish between convergence in tradable and non-tradable good prices, but the lack of state-level price data for tradable goods makes it difficult to distinguish between convergence processes.

Second, among the 50 U.S. states, 11 subgroups or convergence clubs emerge. Taking a closer look at the states in each of these groups, some results are worth mentioning. For example, and as far as the first convergence group is concerned (Alabama, Georgia, Mississippi, North Carolina, South Carolina and Tennessee), distance seems to be the main factor explaining the convergence in prices among the six states in this group. The same variable, distance, seems to explain the convergence among some of the states in the 5th subgroup (Arizona, Colorado, New Mexico, Oklahoma and Utah) and the 8th subgroup (Montana, Nebraska, North Dakota and South Dakota). The significant relationship between the distance among regions and convergence has already been documented in the literature (Chmelarova and Nath 2009; Mark J. Holmes, Jesus Otero, and Theodor Panagiotidis 2011). Furthermore, a comparison of our results with the BEA classification of regions suggests the following. The above 6 states classified in the first subgroup, the states in the third group (Louisiana and Virginia), and two of the three states classified in the fourth group (Louisiana and West Virginia) are classified in the Southeast Region according to the BEA classification. This suggests the agreement of the results obtained in this paper with the BEA classification criteria. However, they also point to the existence of more than one convergence group within the extensive group of the Southeast region, at least when regional price dynamics are considered. While the distance could be a factor explaining the price convergence in some of the subgroups, the characteristics of the states classified in the eighth group (Maine, Montana, Nebraska, New Jersey, North Dakota and South Dakota) suggest that the factor behind the price convergence phenomenon across these U.S. states would be their larger rural population. This factor has also been identified in the literature as a determinant of price divergence (Cecchetti, Mark, and Sonora 2002).

Third, when we analyse the convergence in housing prices across U.S. states using the same methodology, the results also suggest that the null hypothesis of convergence to a single steady state should be rejected. Furthermore, among the 50 U.S. states, 7 subgroups convergence clubs emerge. As in the previous analysis, the distance among the states and the common characteristics of some of the regions (such as the percentage of rural population) are some of the factors that explain these results.

The main policy implications of the results are the following. First, the lack of evidence of convergence suggests a low degree of market integration across the U.S. states, which could be interpreted as a low competitive pressure in some of the states, although, as mentioned above, an analysis of the convergence process among the prices of tradable goods would help us better understand the forces behind the observed degree of market integration across the U.S. states. Second, the results point to the limitations of using the same national price deflator for calculating real regional variables (income, public expenditure in health, education, etc.) for comparison purposes. This limitation has already been pointed out in the studies on real *per capita* GDP convergence across U.S. states (see, for example, Steven C. Deller, Martin Shields, and David Tomberlin 1996 or Choi 2004b). Finally, national monetary policies aimed at stabilising national prices will affect regional prices differently.

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