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Relative price convergence among US cities: Does the choice of numeraire city matter?

Viera Chmelarova [a], Hiranya K. Nath [b,*]

[a] European Central Bank, Frankfurt, Germany
[b] Department of Economics and International Business, Sam Houston State University, Huntsville, Texas, United States

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Using annual consumer price index (CPI) data for 17 US cities between 1918 and 2007, this paper examines the implications of the choice of numeraire city for the behavior of relative prices across cities. A common factor representation of relative price is used to understand the nature of the dynamic behavior of relative prices. This paper finds overwhelming evidence of convergence in relative prices with Atlanta, Chicago, and Los Angeles as the numeraire city. Further, with Boston, Cincinnati, Houston, San Francisco, Seattle, and St. Louis, the lack of convergence in relative prices is found to be driven by the nonstationarity of the common factor. In contrast, with New York, Philadelphia, and Portland, while the common factor is stationary, the idiosyncratic components are unit root processes. This paper further investigates the implications of the choice of numeraire city by estimating the half-life with different numeraire cities, after correcting for the small-sample and the time aggregation bias. These half-life estimates are smaller than those reported in previous studies and they vary depending on the choice of numeraire city. The relative price volatility seems to hold some promise for explaining why the choice of numeraire city matters for the unit root results or the half-life estimates.

1. Introduction

Real exchange rates (relative prices) between cities or regions within a country or a common currency and common economic-political system provide an avenue for conducting a natural experiment to gain a better understanding of deviations from purchasing power parity (PPP) without having to worry about such factors as trade barriers, differences in market baskets, sticky nominal exchange rates, as potentially driving these deviations.[1] This has been the primary motivation behind the relatively recent literature that examines the PPP hypothesis using aggregate price data for the US cities.[2] The current study

* Corresponding author.
E-mail address: ECO_HKN@SHSU.EDU (H.K. Nath).

[1] The 'real exchange rate' between two cities within the same country is defined as the ratio between their respective CPI's and often referred to as the 'relative price'. We will use these two terms interchangeably.

[2] Examples of this literature include Culver and Papell (1999), Cecchetti et al. (2002), Chen and Devereux (2003), and Nath and Sarkar (2009). Applying panel unit root test procedures to monthly data between 1978 and 1997 for 14 US cities, Culver and Papell (1999) find little evidence of PPP in the US city prices. In contrast, with annual data for a much longer sample period, both Cecchetti et al. (2002) and Chen and Devereux (2003) find some evidence of convergence among the US city prices. The former applies panel unit root tests to consumer price index (CPI) data for 19 major cities for a period from 1918 to 1995 while the latter applies univariate unit root test procedures to 'absolute price levels' for 19 cities from 1918 to 2000. They construct the so-called 'absolute price levels' by applying growth rates of CPI backward and forward to actual cost of living data for 1989. These aggregate studies were preceded by two very interesting and influential papers by Engel and Rogers (1996) and Parsley and Wei (1996). Both studies examine relative price behavior of various commodities across the US cities. Carrion-i-Silvestre et al. (2004) and Sonora (2005) investigate price convergence among cities in Spain and Mexico respectively.

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contributes to this literature by examining if the choice of numeraire city has any important implications for the dynamic behavior of relative prices.

Most studies in the literature, such as Cecchetti et al. (2002), Chen and Devereux (2003) and Nath and Sarkar (2009), construct relative price as the deviation of the city price level from the national average price level. Thus, they do not have to deal with the choice of a numeraire or base city. Cecchetti et al. (2002) argue that any movements in a numeraire price index are absorbed into the average price index, which is introduced as a common time effect in their panel unit root test equation. In contrast, Culver and Papell (1999) experiment with different city bases. Except for the panel with Detroit as the numeraire city, they find no evidence of convergence among city prices.[3]

The importance of the choice of numeraire currency in panel analysis of international PPP has been discussed by Papell and Theodoridis (2001). Previously, O'Connell (1998) argues that if there is no serial correlation or if the serial correlation properties are homogeneous across cross-section units then controlling for cross-section dependence by including common time effects in panel tests of PPP makes the choice of numeraire currency irrelevant. Papell and Theodoridis (2001) demonstrate that both these conditions are violated and the choice of numeraire currency matters for the dynamic behavior of real exchange rates. These results also indicate that the influence of a common factor introduced by the numeraire currency cannot be ignored, as discussed recently by Mark and Sul (2008). One of the implications is that differences in dynamic properties of the common factors associated with different numeraire currencies drive some of the differences in PPP behavior.

In this paper, we examine the implications of the choice of numeraire city on the relative price behavior among 17 major US cities, using annual CPI data between 1918 and 2007. We model relative price as being composed of two components: a common factor that is common to all cross-sections in the panel of city prices and an idiosyncratic factor that varies across cross-sections. We examine the stochastic trending properties of the common factor and the idiosyncratic factor separately, using the univariate Augmented Dickey–Fuller (ADF) test and panel Cross-sectionally Augmented Im–Pesaran–Shin (CIPS) test respectively. This paper further investigates the implications of the choice of numeraire city by estimating the half-life with different numeraire cities, after correcting for the biases present in the panels.

The rest of the paper is organized as follows. In Section 2, we discuss the empirical methodology. Section 3 presents the results of our empirical analysis. This section is divided into several subsections. The unit root test results are presented in Sections 3.1 and 3.2 presents the unbiased estimates of half-life. The sensitivity analysis results are presented in Section 3.3. Finally, in Section 3.4, we explore if distance between cities and relative price volatility affect the convergence behavior. Section 4 includes our concluding remarks.

2. Empirical methodology

Using consumer price index (CPI) data for 17 major US cities between 1918 and 2007, we calculate the log relative price in city i relative to city j as $r_{it}^{j} = p_{it} - p_{jt}$ where p_{it} is the logarithm of CPI in city i and p_{jt} is the logarithm of CPI in city j in period t. Here, city j is the numeraire or base city. Thus, there are 17 panels of relative prices, one with each city as the numeraire. In order to capture the dynamics of relative prices, we assume that r_{it}^{j} can be represented by a common factor model as follows:

$$r_{it}^{j} = \lambda_t^{j} F_t^{j} + e_{it}^{j} \tag{1}$$

where F_t^j is the common factor and e_{it}^{j} is the idiosyncratic factor, which is orthogonal to F_t^j. As in Mark and Sul (2008), we assume that both components are AR(p) processes. Thus,

$$F_t^j = \sum_{k=1}^{p} \phi_k^j F_{t-k}^j + u_t^j \tag{2}$$

$$e_{it}^{j} = Y_i^j + \sum_{k=1}^{p} \rho_{ik}^{j} e_{it-k}^{j} + \varepsilon_{it}^{j} \tag{3}$$

where $u_t \sim iid(0, \sigma_u^2)$ and $\varepsilon_{it}^{j} \sim iid\left(0, \sigma^2_{\varepsilon_i^j}\right)$. Bai and Ng (2002, 2004) extensively discuss various issues related to such factor structure of large dimensional panels. Such common factor representation helps determine whether the nonstationarity in a series is pervasive or variable-specific (that is, whether it stems from the common factor or from the idiosyncratic factor).

It is important to recognize that if the common factor, F_t^j, is a unit root process, i.e. $I(1)$, then relative price is also $I(1)$ even if the idiosyncratic component, e_{it}^{j}, is $I(0)$.[4] Therefore, our empirical strategy is to test for stochastic trending properties of F_t^j and e_{it}^{j} separately and to combine the test results to draw conclusions about the dynamic behavior of relative prices. We use the cross-sectional average of relative prices in panel j (j = 1, 2, ..., 17) as the measure of F_t^j for that panel. We then conduct a univariate Augmented Dickey–Fuller (ADF) test on this measure of F_t^j to determine if it is a unit root process.

The use of cross-sectional average as a measure of the common factor also allows us to apply a 'second-generation' panel unit root test – suggested by Pesaran (2007) – to test if the idiosyncratic components are unit root processes. This test pro-

[3] That is, the unit root null is rejected only when Detroit is used as the numeraire city. They also consider nine cities in Canada and 15 member countries of the European Union.

[4] In general, "a series with a factor structure is non-stationary if one or more of the common factors are non-stationary, or the idiosyncratic error is non-stationary, or both" (Bai and Ng, 2004, pp.1128). See Bai and Ng (2004) for a detailed discussion. Also see Bai and Ng (2002) and Mark and Sul (2008).

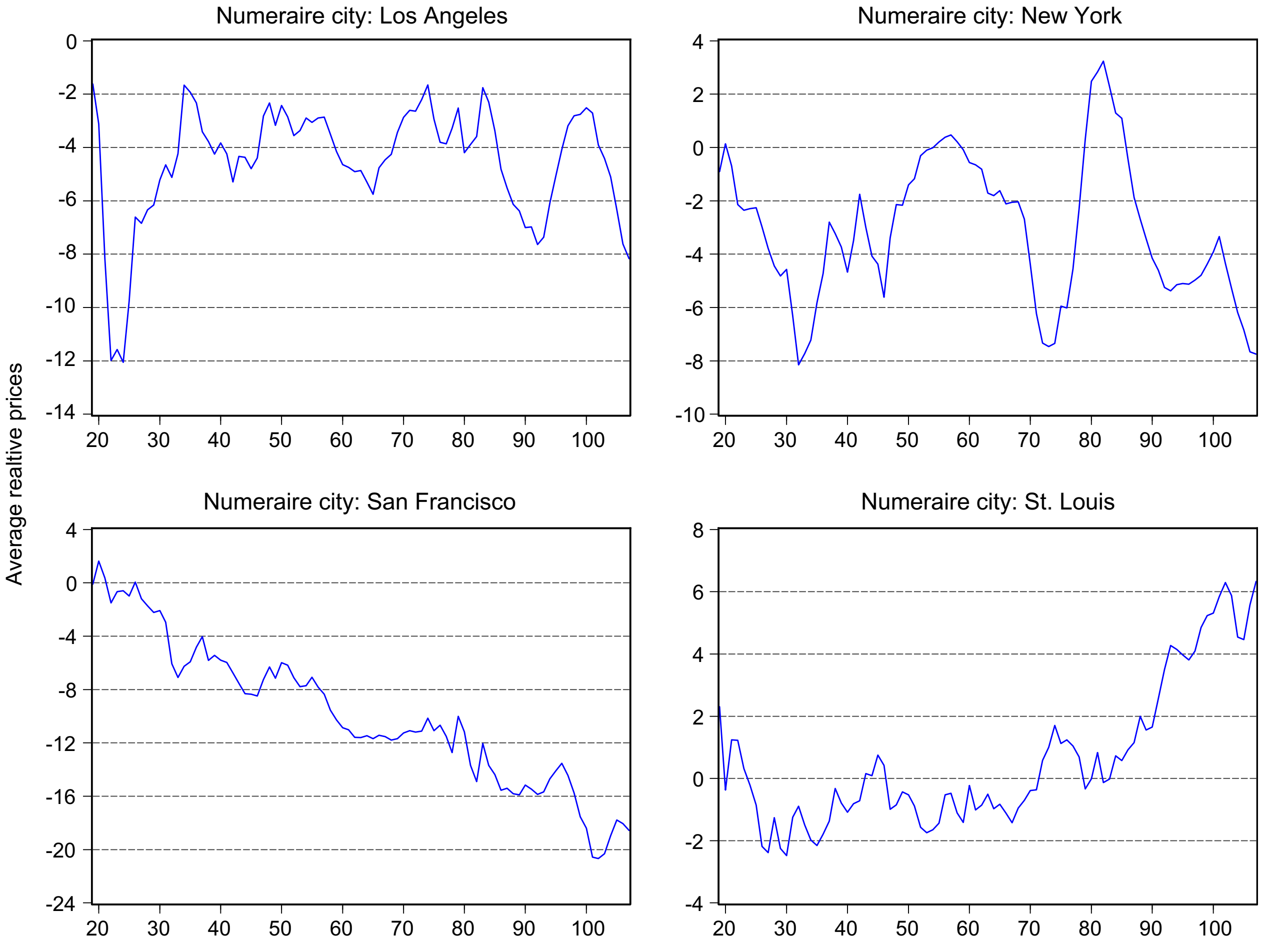

Fig. 1. Average relative prices with selected numeraire cities: 1919–2007.

cedure involves augmenting the standard Dickey–Fuller (DF) or Augmented Dickey–Fuller (ADF) regressions with the cross-section averages of lagged levels and first-differences of the individual series to account for cross-sectional correlation:

$$\Delta r_{it} = \alpha_i + \theta_i r_{i,t-1} + \gamma_i \overline{r_{t-1}} + \sum_{k=1}^{p} \mu_{ik} \Delta \overline{r_{t-k}} + \sum_{k=1}^{p} \phi_{ik} \Delta r_{t-k} + \upsilon_{it} \tag{4}$$

where $\bar{r}_{t-1} = N^{-1}\sum_{i=1}^{N} r_{i,t-1}$ and $\bar{r}_{t-j} = N^{-1}\sum_{i=1}^{N} r_{i,t-j}$. The panel unit root test is based on the simple averages of the individual cross-sectionally augmented ADF statistics (called CADF), given by the t-ratios of the coefficients of $r_{i,t-1}$. By including current and lagged values of the cross-sectional average of relative prices on the right-hand side of the test equation, this procedure has already controlled for the effects of the common factor for the dynamic behavior of relative prices. Thus, this test is essentially a unit root test of the idiosyncratic factor, e^j_{it}, in Eq. (1). Note that this panel unit root test is a modification over the Im–Pesaran–Shin (IPS) panel unit root test (Im et al., 2003) and referred to as cross-sectionally augmented IPS (CIPS) test (see Pesaran (2007)).[5] The null hypothesis of the test is that all time series are unit root non-stationary against the alternative that at least one series is stationary.

3. Empirical results

To give a general idea about how the choice of numeraire city may affect the relative price behavior, we plot the average relative prices (averaged across 16 US cities) for selected numeraire cities and present them in Fig. 1. While average relative prices with Los Angeles and New York as the numeraire city do not exhibit any particular pattern, those with San Francisco and St. Louis clearly display linear trends. Thus, we would expect important differences in dynamic behavior of relative prices based on the choice of numeraire city.

[5] To avoid undue influences of extreme outcomes when T is small (in the range of 10–20), a truncated version of the CIPS test is also used. For details, see Pesaran (2007).

3.1. Unit root test results

The unit root test results for the common factor and the idiosyncratic factor are presented in Table 1. The first six columns present the univariate ADF test results for the cross-sectional averages with different numeraire cities. The test-statistics reported in column 1 are estimated by selecting the lag lengths based on Schwarz Information Criterion (SIC). In next five columns, we report test-statistics estimated with fixed lag lengths ranging between 1 and 5, to be consistent with the CIPS test results. With Chicago, Los Angeles, New York, and Philadelphia as the numeraire city, the common factor is unequivocally $I(0)$. There is overwhelming evidence of unit root stationarity for Atlanta and Portland as well. In contrast, with Boston, Cincinnati, Houston, San Francisco, Seattle, and St. Louis, it is clearly $I(1)$. Furthermore, there is strong evidence of non-stationary common factor in relative prices with Detroit, Kansas City, and Pittsburgh as the numeraire city.

According to the CIPS test results, the null hypothesis of unit root in the idiosyncratic factor of relative prices is rejected for most but not all cases.[6] One problem with this test procedure is that there is no clear guideline as to how to select the lag order for the augmented terms in the test equations. Therefore, as in Pesaran (2007), we experiment with lag length, 1–5.[7] The unit root null is rejected under all specifications with lag order 1–5 at least at the 10% significance level for all but relative prices with Chicago, New York, Philadelphia, and Portland as the numeraire city. However, only for the case with New York, the idiosyncratic factor is unequivocally $I(1)$.

Combining these two sets of results, we may draw the following conclusions. With Los Angeles as the numeraire city, both the common factor and the idiosyncratic factor are $I(0)$. This is suggestive of the convergent behavior of relative prices. There is overwhelming evidence of similar behavior with Atlanta and Chicago as well. With Boston, Cincinnati, Houston, San Francisco, Seattle, and St. Louis, the common factors are $I(1)$ while the idiosyncratic factors are $I(0)$. Thus, the lack of convergence in relative prices in these cases is driven by the nonstationarity of the common factor. In contrast, with New York, Philadelphia, and Portland, while the common factors are $I(0)$, the idiosyncratic factors are overwhelmingly $I(1)$.[8] Therefore, in former cases, the nonstationarity of relative prices is pervasive while, in latter cases, it is city-specific. For the rest of the cities, although the idiosyncratic factors are $I(0)$, the evidence for the common factors is mixed.[9] These results clearly suggest that the dynamic behavior of relative prices depends on the choice of numeraire city.

3.2. Estimates of the half-life

Another way to examine the relative price behavior with respect to the choice of numeraire city is to compare the speed of convergence, which is measured by the half-life. It is the time required for a divergence from PPP to dissipate by one half. In an AR(1) case, half-life is calculated as follows:

$$h(\rho) = \frac{-\ln(2)}{\ln(\rho)} \tag{5}$$

where $h(\rho)$ is the half-life and ρ is the AR coefficient. However, as pointed out by Choi et al. (2006), panel data estimation of the half-life of PPP deviations may involve three potential sources of bias. These biases arise due to inappropriate cross-sectional aggregation of heterogeneous AR coefficients, small-sample estimation of lag coefficients (also known as the Nickell bias), and time aggregation of commodity prices.[10]

To determine if we need to correct for the cross-sectional aggregation bias that may potentially arise due to heterogeneity in the autoregressive (AR) coefficient across cities, we conduct a test of cross-sectional homogeneity by estimating the dynamic panel regression for each numeraire city.[11] The results indicate that cross-sectional heterogeneity is not a significant source of bias. Thus, we need to correct for the small-sample (Nickell) bias and the time aggregation bias in the estimates of half-life. Note that Nickell bias in the estimates of the AR coefficient is a downward bias while the time aggregation bias is an upward bias. Following Choi et al. (2006), we estimate a panel generalized least squares (GLS) model with fixed effects, and use the inverse of a bias function that incorporates both these biases to obtain the unbiased estimate of the AR coefficient.[12]

[6] As discussed before, the CIPS test procedure directly takes account of cross-section dependence among relative prices. We implement the so-called CD (cross-section dependence) test procedure, suggested by Pesaran (2004), to establish that cross-sectional dependence indeed exists in the panels of relative prices. The results of these tests are not reported but can be obtained from the authors.

[7] Pesaran (2007) applies this test procedure to a panel dataset on quarterly real exchange rates in 17 countries between 1974 and 1998 with $p = 1, 2, 3, 4$.

[8] For Portland, the unit root null cannot barely be rejected at the 10% level under the lag order, $p = 4$. The p-value is 0.107.

[9] Applying a factor model to a similar data set for 19 US cities over 1918–2001, Phillips and Sul (2007) show that there is little evidence of overall convergence in cost of living across US cities, Instead, they uncover three convergence clubs among the 19 cities. However, they allow time varying factor loading coefficients for the common factor in their model (unlike a constant λ^j in our model) and they do not examine the effect of the choice of numeraire city on the convergence behavior.

[10] These biases are discussed separately by different authors. For example, Imbs et al. (2005) discusses cross-sectional aggregation bias; Nickell (1981) discusses small-sample estimation bias; and Taylor (2001) discusses time aggregation bias. Choi et al. (2006) discuss all three biases together.

[11] As suggested by Choi et al. (2004), we use recursive mean-adjusted seemingly unrelated regression estimates of ρ so that we can control for the Nickell bias. The test results can be obtained from the authors.

[12] While estimating the CIPS test equation we obtain a different AR coefficient (ρ) for each cross-section within the panel. To obtain a single AR coefficient we need to use some aggregation procedure. Cecchetti et al. (2002) use simple average of the estimated coefficients and then correct for the Nickell bias. We use panel GLS of Choi et al. (2006) to avoid the use of such aggregation.

Table 1
Unit root test results.

Numeraire city	Univariate ADF unit root test results for common factor						CIPS panel unit root test results for idiosyncratic factors				
	Lag length selected by SIC (1)	Lag length = 1 (2)	Lag length = 2 (3)	Lag length = 3 (4)	Lag length = 4 (5)	Lag length = 5 (6)	Lag length = 1 (7)	Lag length = 2 (8)	Lag length = 3 (9)	Lag length = 4 (10)	Lag length = 5 (11)
Atlanta	−5.05***	−4.30***	−3.22**	−2.80*	−2.83*	−2.24	−2.66***	−2.53***	−2.75***	−2.60***	−2.51***
Boston	−1.32	−1.32	−1.81	−1.87	−1.91	−1.95	−2.41**	−2.25*	−2.40**	−2.26**	−2.24*
Chicago	−2.70*	−2.70*	−2.81*	−2.65*	−2.64*	−3.24**	−2.19*	−2.16*	−2.33**	−2.09	−1.92
Cincinnati	−1.33	−1.77	−1.37	−1.41	−1.49	−1.53	−2.68***	−2.74***	−2.96***	−2.80***	−2.66***
Cleveland	−3.04**	−2.34	−1.82	−3.05**	−2.39	−2.41	−2.26**	−2.20*	−2.42**	−2.38**	−2.46***
Detroit	−1.93	−3.55**	−1.93	−2.65*	−2.48	−2.09	−2.52***	−2.58***	−2.99***	−3.21***	−3.03***
Houston	−1.51	−1.51	−1.66	−2.45	−1.75	−1.44	−2.63***	−2.52***	−2.69***	−2.46***	−2.48***
Kansas City	−1.75	−2.86*	−2.51	−1.56	−1.36	−1.11	−2.72***	−2.73***	−3.03***	−2.87***	−2.82***
Los Angeles	−3.87***	−3.87***	−3.59***	−4.50***	−3.81***	−4.24***	−2.36**	−2.30**	−2.41**	−2.28**	−2.22*
Minneapolis	−2.78*	−2.93**	−2.35	−2.24	−2.37	−2.27	−2.67***	−2.69***	−2.95***	−2.80***	−2.89***
New York	−3.09**	−3.09**	−2.90**	−3.02**	−2.86*	−3.52***	−2.08	−1.87	−2.10	−1.90	−1.80
Philadelphia	−3.85***	−2.64*	−3.03**	−3.85***	−3.56***	−2.83*	−2.20*	−1.97	−2.01	−1.85	−1.91
Pittsburg	−2.36	−2.45	−2.39	−2.77*	−2.68*	−2.32	−2.56***	−2.52***	−2.96***	−2.78***	−2.73***
Portland	−2.78*	−2.78*	−2.62*	−2.77*	−2.55	−3.05**	−2.16*	−2.00	−2.20*	−2.06	−1.87
San Francisco	−1.08	−1.54	−1.20	−0.93	−1.12	−0.98	−3.03***	−3.17***	−3.52***	−3.35***	−3.36***
Seattle	−0.86	−0.86	−1.12	−0.98	−1.35	−1.00	−2.88***	−2.81***	−3.19***	−2.97***	−3.08***
St. Louis	−0.40	−0.02	0.26	0.18	0.29	0.46	−2.96***	−3.09***	−3.44***	−3.24***	−3.20***

For the CIPS test, the 1%, 5%, and 10% critical values for $T = 100$ and $N = 15$ generated by stochastic simulation are -2.42, -2.25, and -2.15 respectively. See Table II(b) in Pesaran (2007).
* Significance at the 10%.
** Significance at the 5%.
*** Significance at the 1%.

Table 2
Unbiased estimation of AR coefficients and the half-life to PPP convergence in panel data.

Numeraire city	No bias correction (1)	Estimated half-life (2)	Nickell and time aggregation bias correction (3)	Estimated half-life (4)
Atlanta	0.943	11.811	0.940	11.202
Boston	0.939	11.013	0.926	9.016
Chicago	0.942	11.601	0.930	9.551
Cincinnati	0.935	10.313	0.923	8.651
Cleveland	0.944	12.028	0.943	11.811
Detroit	0.939	11.013	0.923	8.651
Houston	0.937	10.652	0.923	8.651
Kansas City	0.940	11.202	0.926	9.016
Los Angeles	0.943	11.811	0.933	9.995
Minneapolis	0.942	11.601	0.940	11.202
New York	0.936	10.480	0.930	9.551
Philadelphia	0.940	11.202	0.936	10.480
Pittsburg	0.939	11.013	0.930	9.551
Portland	0.940	11.202	0.936	10.480
San Francisco	0.929	9.412	0.919	8.206
Seattle	0.937	10.652	0.923	8.651
St. Louis	0.931	9.695	0.913	7.615
Average		10.982		9.546
Minimum		9.412		7.615
Maximum		12.028		11.811

The results are reported in Table 2. Note that we present the estimates of the AR coefficients with no bias correction and associated half-lives in the first two columns.

The bias-corrected estimated half-life is the shortest with St. Louis as the numeraire city. It is slightly below 8 years. With Cleveland as the numeraire city, the estimated half-life is about 12 years. The average of estimated half-lives with different numeraire cities is 9.54 years. Again, to make a comparison with the results reported in Cecchetti et al. (2002), we also estimate the half-lives for the period 1918 – 1995. Corrected for the Nickell bias, which is directly comparable to Cecchetti et al. (2002), the half-life ranges between 9.02 years (with Seattle as the numeraire city) and 13.51 years (with Cleveland as the numeraire city) with an average of 10.58 years, which is higher than their estimates of 8.54 years (under Levin–Lin procedure) and 9.70 years (under Im–Pesaran–Shin procedure).[13] However, when we correct for the time aggregation bias in addition to the Nickell bias, the half-life estimates for 1918–1995 are much smaller and they range between 5.28 years (Seattle) and 7.27 years (Cleveland) with an average of 6 years. A comparison of the half-life estimates for 1918–2007 and 1918–1995 reveals that with Atlanta, Cleveland, Minneapolis, Philadelphia, and Portland as the numeraire city, the difference in these estimates is more than 4 years. We can make a few inferences from these results. First, depending on the choice of numeraire city, the half-life estimates vary. Second, corrected for both the Nickell bias and time aggregation bias, the half-life estimates are much shorter than those reported by previous studies.[14] Nonetheless, the implied speed of convergence is slow for all numeraire cities. Finally, the events of the past 10–12 years since 1995 seem to have some significant impact on the relative price behavior across the US cities, which need further investigation.

3.3. Sensitivity analysis

In this subsection, we examine if the unit root test results are sensitive to the selection of sample period. We conduct a few experiments. First, we divide the entire time period into two sub-periods: 1918–1955 and 1956–2007. Note that 1955 was the last year when the US experienced a fall in the general price level which has been continuously rising since then.[15] Also, during the second period the transportation costs have gone down significantly and, intuitively, it should have some impact on the relative price behavior across the US cities. Second, we further subdivide the 1956–2007 period into two episodes: 1956–1986 that includes very high and volatile inflation of the 1970s and the early 1980s; and 1987–2007 that has witnessed low and steady inflation.

We summarize the results of the univariate ADF test and the CIPS test for the common factor and the idiosyncratic factors respectively in Table 3. In columns 1 and 2, we list the numeraire cities for which the unit root null is rejected for the common factor and the idiosyncratic factor respectively under all specifications of the lag order, $p = 1–5$. In columns 3 and 4, we

[13] Note that their sample includes two more cities – Baltimore and Washington, DC – and they estimate the AR coefficients differently as discussed in the previous footnote.

[14] In a recent study, Nath and Sarkar (2009) report a half-life estimate of 7.5 years, after correcting for the Nickell bias and time aggregation bias, in annual CPI data for the same 17 US cities as in this paper between 1918 and 2006. Like Cecchetti et al. (2002), they use the deviation of city CPI from the national average (in logarithm) as the measure of relative price.

[15] Cecchetti et al. (2002) also use 1955 as a threshold to split their sample period in one of their experiments.

Table 3
Summary results of panel unit root tests on subsamples.

Periods	Numeraire city(ies) for which unit root null *is* rejected at least at the 10 percent significance level under all values of p (1 through 5)		Numeraire city(ies) for which unit root null *is not* rejected at least at the 10 percent significance level under all values of p (1 through 5)	
	Common Factor (1)	Idiosyncratic Factors (2)	Common Factor (3)	Idiosyncratic Factors (4)
1918–2007	Chicago, Los Angeles, New York, Philadelphia	Atlanta, Boston, Cincinnati, Cleveland, Detroit, Houston, Kansas City, Los Angeles, Minneapolis, Pittsburgh, San Francisco, Seattle, St. Louis	Boston, Cincinnati, Houston, San Francisco, Seattle, St. Louis	New York
1918–1955	–	Atlanta, Boston, Cincinnati, Cleveland, Philadelphia, Portland, San Francisco, Seattle	Boston, Chicago, Houston, Minneapolis, New York, Philadelphia, Portland, San Francisco, Seattle	Detroit
1956–2007	Philadelphia, Portland	Cincinnati, Detroit, Houston, Kansas City, San Francisco, St. Louis	Atlanta, Boston, Chicago, Cincinnati, Cleveland, Detroit, Houston, Kansas City, Minneapolis, Pittsburgh, San Francisco, Seattle, St. Louis	Atlanta, Boston, Chicago, Cleveland, Los Angeles, New York, Philadelphia, Portland
1956–1986	Portland	Boston, Detroit, Philadelphia, San Francisco, St. Louis	Atlanta, Chicago, Los Angeles, Minneapolis, Pittsburgh, San Francisco, St. Louis	Atlanta, Cleveland, Portland
1987–2007	Pittsburgh	Boston, Chicago, Houston, Los Angeles, Seattle	Atlanta, Boston, Chicago, Cincinnati, Cleveland, Detroit, Houston, Kansas City, Minneapolis, New York, Portland, San Francisco, Seattle, St. Louis	Atlanta, Kansas City, Pittsburgh

list the numeraire cities for which these factors are unit root non-stationary or divergent.[16] The cities for which the evidence is mixed (that is, the common or idiosyncratic factors are unit root processes under certain values of p and not under others) do not appear in these columns. For easy reference, we summarize the results from Table 1 in the first row. We can draw at least two conclusions from this summary table. First, the choice of the numeraire city does matter for the dynamic behavior of relative prices across US cities irrespective of the sample period. Second, that the nonstationarity of the common factor drives the relative price behavior with Boston, Houston, San Francisco, Seattle, and St. Louis as the numeraire city is a fairly robust result across almost all sample periods.

The unbiased estimates of half-life range between 4.20 years (Portland) and 6.58 years (St. Louis) for the sub-period 1918–1955 with an average of 5.42 years. The estimated half-life is much longer for the second sub-period (1956–2007) which is counterintuitive. With Seattle as the numeraire city, the unbiased estimate of the AR coefficient turns out to be 1 for this later period, which renders the half-life undefined. As a further check, we also obtain unbiased estimates of half-life for 1956–1986, 1956–1995 (as in Cecchetti et al., 2002), and 1987–2007. These estimates now range between 3.11 years (New York) and 7.62 years (Cleveland) for the period 1956–1986 with an average of 5.30 years and between 5.28 years (Kansas City) and 8.65 years (Cleveland and San Francisco) for 1956–1995. For the period 1987–2007, the estimated AR coefficients turn out to be 1 for panels with six numeraire cities. Because of the short sample period ($T = 21$), however, the estimates are imprecise. Nevertheless, these experiments seem to suggest that there have been important events in most recent years, particularly after 1995, that have significantly affected relative price movements across cities in the United States.

3.4. Distance and volatility as explanatory variables

This subsection is exploratory in the sense that we examine if the implications of the choice of numeraire city for the relative price behavior can be attributed to the 'usual suspects' namely, distance and relative price volatility.[17] Papell and Theodoridis (2001) conclude that distance – which is often used as proxy for transportation costs – explains why there is stronger evidence of the PPP hypothesis when the currency of a country located in geographical proximity to most countries in the sample is used as the numeraire currency. In this subsection we will examine if that is the case when we consider the US cities. We use logarithm of average distance from a numeraire city to all other cities as the explanatory variable, and the absolute values of ADF and CIPS test-statistics, and the unbiased estimate of half-life as the alternate dependent variables.[18] Our hypothesis is that

[16] Unlike the results for the full sample period, the estimated test statistics for the truncated CIPS test differ from the non-truncated version particularly for higher values of p. Also, for shorter sub-periods with large p, there is loss of degrees of freedom in the estimation of the test equation and therefore the estimates become imprecise.

[17] Inclusion of non-traded items in calculation of CPI could be another usual suspect. However, there are two problems: first, CPI data on goods and services which are often used as proxies for traded and non-traded goods, respectively are available only since 1967 and not for the entire sample period; second, with the advent of technology, not all services are non-traded items any more.

[18] We obtain the road distances (in miles) between cities from Rand McNally's Road Atlas and use the logarithmic values as in Cecchetti et al. (2002). Papell and Theodoridis (2001) use the square root of the sum of air distance in statute miles between capital of the numeraire country and all other countries in the sample.

Table 4
Distance and relative price volatility as explanatory variables.

Dependent variables	Log (distance)		Volatility		Adjusted R-squared
	Estimated coefficient (1)	t-Statistics (2)	Estimated coefficient (3)	t-Statistics (4)	(5)
ADF	−0.33	−0.30 (0.77)			−0.060
			0.21	0.34 (0.74)	−0.059
	−0.39	−0.34 (0.74)	0.24	0.37 (0.71)	−0.125
ADF(1)	−0.42	−0.43 (0.68)			−0.054
			−0.21	−0.38 (0.71)	−0.057
	−0.38	−0.37 (0.72)	−0.18	−0.31 (0.76)	−0.121
ADF(2)	0.08	0.10 (0.92)			−0.066
			0.20	0.45 (0.66)	−0.053
	0.03	0.04 (0.97)	0.19	0.42 (0.68)	−0.128
ADF(3)	0.19	0.19 (0.85)			−0.064
			0.43	0.79 (0.44)	−0.024
	0.08	0.08 (0.93)	0.42	0.75 (0.47)	−0.097
ADF(4)	0.18	0.23 (0.82)			−0.063
			0.42	0.94 (0.36)	−0.007
	0.08	0.10 (0.92)	0.41	0.89 (0.39)	−0.078
ADF(5)	0.36	0.40 (0.69)			−0.055
			0.52	1.06 (0.30)	0.008
	0.24	0.26 (0.80)	0.51	0.98 (0.34)	−0.058
CIPS(1)	0.08	0.31 (0.76)			−0.060
			−0.18	−1.33 (0.20)	0.046
	0.13	0.50 (0.62)	−0.19	−1.36 (0.20)	−0.004
CIPS(2)	0.00	−0.02 (0.99)			−0.067
			−0.26	−1.45 (0.17)	0.064
	0.06	0.18 (0.86)	−0.26	−1.41 (0.18)	0.000
CIPS(3)	−0.04	−0.10 (0.92)			−0.066
			−0.35	−1.63 (0.12)	0.095
	0.05	0.12 (0.91)	−0.35	−1.58 (0.14)	0.031
CIPS(4)	−0.12	−0.27 (0.79)			−0.061
			−0.38	−1.73 (0.10)	0.111
	−0.02	−0.06 (0.96)	−0.37	−1.65 (0.12)	0.047
CIPS(5)	−0.04	−0.09 (0.93)			−0.066
			−0.38	−1.67 (0.12)	0.100
	0.06	0.13 (0.90)	−0.38	−1.61 (0.13)	0.037
Half-life	−0.45	−0.43 (0.67)			−0.054
			0.36	0.62 (0.54)	−0.040
	−0.55	−0.51 (0.62)	0.41	0.67 (0.51)	−0.094

Note: Each regression includes an intercept term, the estimate of which is not being reported. The absolute values of the CIPS test-statistics are used as dependent variables. p-values for t-statistics are in brackets.

for the panel of relative prices with a numeraire city located afar from other cities (higher average distance) the likelihood of rejecting the unit root null will be lower and the estimated half-life will be higher.

We further examine if relative price volatility explains why the relative price behavior varies according to the choice of numeraire city. The cities differ from one another by the mix of their economic activities. Therefore, the demand and supply conditions may be tied to the activity(ies) localized in the city. For example, in Detroit, these conditions will depend, to a large extent, on the health of the automobile industry. Intuitively, the cities with more procyclical economic activities are likely to experience higher volatility in demand and supply conditions and, therefore, larger fluctuations in prices. Like Papell and Theodoridis (2001), we use the following measure of average volatility:

$$VOLATILITY_j = \frac{1}{N-1}\left[\sum_{i=1}^{N-1}\left(\sum_{t=1}^{T}\left|\frac{r^j_{i,t+1} - r^j_{it}}{r^j_{it}}\right| \div T\right)\right] \quad (6)$$

where j is the numeraire city, i indexes all other cities and $i \neq j$, and T is the number of periods. This gives us the average volatility of relative prices with respect to each numeraire city j where $j = 1, 2, \ldots, N$.

The regression results are reported in Table 4. In regressions with distance as the only explanatory variable, the estimated coefficients for distance have the wrong sign in some cases, and they are statistically insignificant in all cases. Negative values of the adjusted R-squared also indicate that the estimated models are poor fit. Thus, distance does not seem to be an important determinant of why the relative price behavior varies depending on the choice of numeraire city. This is consistent with the results reported in Cecchetti et al. (2002) who find little evidence of any significant effect of distance on the behavior of the US city prices. In regressions with volatility as the only explanatory variables, the estimated coefficients are mostly positive for ADF test-statistics and all negative for CIPS test-statistics. That is, as relative price volatility increases the probability of rejecting unit root null increases for the common factor and decreases for the idiosyncratic factors. However, the effects are statistically insignificant for common factors and only weakly significant for idiosyncratic factors as we can see from the *p*-values. Furthermore, as volatility increases, the half-life also increases, although the estimated coefficient is not statistically significant. The adjusted *R*-squared measures are larger than those for regressions with distance. Note that the volatility measure is the highest (2.13) when Portland is the numeraire city and the lowest (0.38) when Kansas City is the numeraire city. When we include both distance and volatility together the results do not change very much. One should remember that the number of observations used in these regressions is 17 which represents a very small sample and therefore the results are indicative at best.

These results seem to suggest that relative price volatility, and not distance, hold some promise for explaining why the choice of the numeraire city matters for the behavior of relative prices and future study needs to focus on more city-specific factors that would uncover how these factors change demand and supply conditions in different cities.

4. Conclusions

This paper revisits relative price convergence among the US cities by examining the implications of the choice of numeraire city for the relative price behavior. Relative price is modeled as being composed of a common factor and an idiosyncratic factor, and unit root tests are applied to these components to examine their respective stochastic trending properties. This paper also presents unbiased estimates of half-life, making corrections for two types of biases: the small-sample bias and the time aggregation bias. The results indicate that the choice of numeraire city matters for the dynamic behavior of relative prices both in terms of the unit root tests and the estimates of half-life. With Atlanta, Chicago, and Los Angeles as numeraire cities, the common factors and the idiosyncratic factors are both overwhelmingly $I(0)$, which is indicative of the convergent behavior of relative prices. The results further indicate that the lack of convergence in relative prices with Boston, Cincinnati, Houston, San Francisco, Seattle, and St. Louis is primarily driven by nonstationarity of the common factor. In contrast, with New York, Philadelphia, and Portland, the common factors are unit root stationary while the idiosyncratic factors are not. For other cities, while the idiosyncratic factors are stationary, the unit test results for the common factors are mixed. Corrected for the biases mentioned above, the half-life estimates are smaller than those reported in previous studies and these estimates vary depending on the choice of numeraire city. The relative price volatility seems to hold some promise for explaining why the choice of numeraire city matters for the unit root results or the half-life estimates.

The findings of this study point to several directions for future research. First, future studies need to look at demand and supply conditions specific to each city so that we can learn more about why the choice of a certain city as numeraire has significantly different effects on the relative price behavior than does that of others. That relative price volatility with respect to different numeraire cities has some explanatory power seems to be a pointer in this direction. Second, although the unbiased estimates of half-life are smaller than those reported in previous studies, the implied speed of convergence is still very slow. Thus, the puzzle is not resolved and will need further investigation. Finally, as pointed out in the text, relative price movements across the US cities during the last decade or so seem to have significant effect on overall price convergence. One needs to analyze carefully the data for this period.

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Appendix
Data. and data sources

We obtain annual consumer price index (CPI) data for 17 major cities in the US for the period between 1918 and 2007 from the Bureau of Labor Statistics (BLS). The availability of data dictates the choice of the cities and the sample period. The cities are: Atlanta, Boston, Chicago, Cincinnati, Cleveland, Detroit, Houston, Kansas City, Los Angeles, Minneapolis, New York, Philadelphia, Pittsburgh, Portland, San Francisco, Seattle, and St. Louis. Cecchetti et al. (2002) and Chen and Devereux (2003) include Baltimore and Washington, DC as well. The BLS has discontinued the CPI series separately for these two cities and has been publishing a combined series since 1997. For each city as the numeraire or base, we construct a panel of 16 relative price series. Thus, there are 17 panels with a total of 272 or 136 independent relative price series.

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