What can US city price data tell us about purchasing power parity?

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Abstract

We study the dispersion of absolute price levels for US cities since 1918. By absolute price levels, we mean price indices that measure the cost of a given consumption basket at each point in time. We find strong evidence that city price levels converge over time and that the dispersion of price levels is lower for US cities than between OECD countries. We argue that price level convergence for US cities will produce bilateral real exchange rate nonstationarity. In this case, however, nonstationarity is not evidence against Purchasing Power Parity (PPP), rather it is the consequence of improved market integration.

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1. Introduction

In recent years, researchers have explored various aspects of real exchange rate behavior with city price data. The literature in this area begins with Engel and Rogers’s (1996) influential work where they compare relative price variability within
and across the US and Canada using disaggregated city price indices. More recently, Culver and Papell (University of Houston, TX, unpublished) (1999) and Cachetti et al. (2001) tested aggregate versions of Purchasing Power Parity (PPP) with city data. In this paper, we extend the city literature to consider the long run behavior of absolute price levels. By absolute price levels, we mean price indices that measure the cost of a given consumption basket at each point in time. Our first goal is to document the behavior of absolute price levels for US cities since 1918. Our second goal is to explain why researchers find it difficult to reject real exchange rate nonstationarity for city real exchange rates.

We introduce our data in Section 2. In Section 3, we examine the dispersion of US city price levels. We have two findings. First, we show that the dispersion of city price levels has fallen by 40% since the 1920s. In other words, US city price levels are converging over time. Second, we show that price level dispersion is lower for cities within the US than across OECD economies. We argue that these findings are consistent with beliefs that markets are more closely integrated within countries and that market integration for US cities has increased over time.

In Section 4, we discuss recent tests of city bilateral real exchange rate stationarity in the light of city price level convergence. According to previous research, difficulties in rejecting bilateral real exchange rate nonstationarity for US city data are due to slow speeds of adjustment. We argue in contrast that nonstationarity may be the result of the convergence of city price levels. When city price levels are converging, bilateral city real exchange rates cannot return to a fixed mean. Price level convergence thus implies bilateral real exchange rate nonstationarity. Clearly nonstationarity in this case is not evidence against PPP. On the contrary, it supports versions of PPP that allow for reductions in transport costs and improved market integration. We conclude in Section 5.

2. The city price data

The BLS city price indices refer to standard metropolitan areas (SMAs) or Consolidated Metropolitan Areas (CSMAs). The indices are ideal for studying intranational real exchange rate behavior because they cover identical goods and services at each point in time and they are available for long spans.

We have CPI indices for 19 cities from 1918. They are New York, Philadelphia, Boston, Pittsburgh, Chicago, Detroit, St Louis, Cleveland, Minneapolis, Cincinnati, Kansas City, Washington DC, Baltimore, Houston, Atlanta, San Francisco, Los Ang-

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1 Recent work in this area includes Engel and Rogers (2001) and Parsley and Wei (2001). The PPP debates of the 1970s and early 1980s also generated a city price literature; see McCloskey and Zecher (1976) and Gandolfi and Lothian unpublished, (1980). The early work differs from current research in that it focuses on whether PPP holds in the short run.

2 For early accounts of the BLS indices, see Douglas (1930) and BLS bulletins 77 and 156.

3 There are small differences in consumption baskets across cities.
The group includes nine of the ten currently largest US metro areas. The exception is Dallas.

A drawback of the BLS indices is that they do not measure absolute city price levels. Instead, each city CPI has the same base year, 1982–1980 = 100. Hence the indices measure changes in price levels while providing no information on the relative cost of living across cities at a point in time.

Using disaggregated BLS data, Koo et al. (2000) construct absolute city cost of living indices for 1989. We use their estimates as our benchmarks. To obtain absolute price indices for other years, we project the Koo et al. (2000) estimates backwards and forwards in time with inflation rates calculated from individual city price indices. The ratio of the resulting price indices for any two cities gives the relative cost of living for the $i^{th}$ relative to the $j^{th}$ city from 1918 to 2000.

How accurate are our absolute price level estimates? One way to determine their accuracy is to compare them with independent benchmarks for overlapping years. Stecker (1937) gives estimates for 1935 while the BLS provides a benchmark for 1975, (see BLS Bulletin, 1705). As it turns out, our price level estimates are close to the alternative benchmarks. Moreover, our findings still hold if we construct the absolute price measures with the alternative benchmarks.

### 3. City price level dispersion

In this section, we look at price level dispersion for US cities where we measure dispersion by the coefficient of variation of the log absolute price levels. An increase in the coefficient of variation shows that price level dispersion has increased while a decline means that dispersion has fallen. In the literature, a decline in price level dispersion is traditionally seen as evidence of improved market integration.

Fig. 1 looks at overall price level dispersion since 1918. Fig. 1 shows that the dispersion of city price levels is declining over time. By 2000, the coefficient of variation (0.020) is 58% of its average level in the 1920s (0.034). Most of the fall occurs between 1935 and 1955.

The finding of city price level convergence is robust. It is not sensitive to the cities in the sample. Furthermore, dispersion falls within each region. The coefficient of variation declines for the Midwest, the Northeast, the South and the West. The

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4 The BLS provides data for longer spans and higher quality than other sources. Koo et al. (2000) discuss the problems associated with the American Chamber of Commerce (AACRA) survey price data used in some previous work. Researchers have also used Gross State Product deflators, available from the BLS since 1977, as measures of state prices. But these deflators are not true price indices since they use national prices with state industry weights.

5 The Penn Tables use a similar procedure to obtain an annual price series.

6 For a discussion, see Knetter and Slaughter (2001).

7 Of course, not all city prices have converged. The San Francisco and Seattle prices levels, for example, have appreciated relative to the US average since the 1920s.
timing of the price level convergence within the regions is similar to Fig. 1 suggesting that similar forces are at work for all regions.

One explanation for price level convergence is the improvements in transportation and communications, such as the construction of the interstate highway system. If that were the case, we should expect to see price level dispersion to fall by more for tradables than for nontradables.

To proceed we will look at the dispersion of tradables and nontradables. There are two sub-indices of the CPI available for the 1918–2000 period, food, and rent. Food is our proxy for tradables while rent proxies nontradables. We construct the absolute price indices for food and rent using 1975 BLS absolute food and rent benchmarks because Koo et al. (2000) do not provide estimates for sub-components of the CPI.

Fig. 2 gives the dispersion of food price levels since 1918.
There are three features of Fig. 2 worth noting. For all years the coefficient of variation for food prices is lower than for the overall price level. This is expected given that food prices have a larger traded component. Second, food price levels have converged over time. The coefficient of variation for food falls from an average of 0.023 during the 1920s to 0.010 in 2000. Most of the fall is between 1935 and 1948. Lastly, the coefficient of variation falls by more for food than for the overall price level consistent with the claim that traded goods drive overall price convergence.

Fig. 3 looks at the dispersion of rents. Rent dispersion falls before 1950. But thereafter it increases. By 2000, the coefficient of variation for rents is twice that of the overall price level and it is four times that of food. An examination of the rent data shows that rents in the North East and the West have increased relative to the US average after 1950 while rents in the Mid West and South have fallen.

What explains the rent divergence? The empirical evidence suggests that city rent levels are influenced by amenities such as weather (Roback, 1982) along with fiscal differences (Gyourko and Tracy, 1991). There is also evidence that rent differences can reflect agglomeration effects (Rauch, 1993). The divergence in rent levels after 1950 could therefore be due to increased demand for amenities, or divergences in taxes and services and agglomeration effects across US regions.

Finally, how does city price level dispersion compare with price level dispersion internationally? To answer this question, Fig. 4 compares the dispersion of absolute price levels for our 19 US cities with the dispersion of absolute price levels for 18 developed economies. We take the international price data from the OECD. Our data are from 1970 to 2000.

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8 The OECD constructs the estimates by combining its periodic price level comparisons with domestic price indices. Note that the OECD refers to the overall GDP deflator whereas the city data refer to
From 1970 to 2000, international price levels are two to three times more widely dispersed than city prices. The result is not sensitive to the countries or the cities included in the samples.

The finding that price level dispersion is lower across US cities than internationally accords with the Engel and Rogers (1996) argument that markets are more integrated within countries. Note, however, that our results are not strictly comparable to Engel and Rogers (1996). They compare the volatility of bilateral price indices within and across countries whereas we look at the dispersion of absolute price levels. In our data set, bilateral real exchange rates internationally are, on average, seven to ten times more variable than city real rates depending on the country/city base.

4. Price level convergence and real exchange rate nonstationarity

In this section, we re-examine some recent debates on city real exchange rate stationarity. We argue that city price level convergence will produce bilateral real

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consumption. Our sample consists of Australia, Canada, Belgium, Denmark, Finland, France, Germany, Austria, France, Ireland, Italy, Japan, New Zealand, the Netherlands, Norway, Sweden, the UK and the US. The data can be accessed at http://www.oecd.org/std/pppoeecd.xls.

9 There is overwhelming evidence that goods and factor markets are more tightly linked within rather than across economies. McCallum’s (1995) widely cited work, for example, shows that trade linkages are 20 times denser between Canadian provinces as compared to linkages between these provinces and US states of similar size and distance. There is equally strong evidence that factor markets are more integrated within economies, see Helliwell (1998).
exchange rate nonstationarity and that nonstationarity in these circumstances is not evidence against PPP.

To set the stage and to facilitate a comparison with previous work, we will first present ADF tests for city bilateral real exchange rates. But we are not interested here in testing for stationarity per se. Instead our objective is to reinterpret the results of such tests.

Let us define the real exchange rate for the $i^{th}$ US city in logs as $p_i - p_{us}$ where $p_i$ is the log of the price level for the $i^{th}$ city and $p_{us}$ is the log of the US CPI. We use the overall US price level as our base because it is a convenient way to summarize the data\(^{10}\). The null hypothesis is that city real exchange rates follow a random walk with drift while the alternative hypothesis is that real exchange rates are AR(1) processes with a mean. To determine lag length, we rely on the reduction method of Campbell and Perron (1991) starting with eight lags. Following earlier work, we omit time trends.

We give the stationarity results in Table 1.

<table>
<thead>
<tr>
<th>City</th>
<th>ADF t</th>
<th>$\rho$</th>
<th>Half life</th>
</tr>
</thead>
<tbody>
<tr>
<td>Atlanta</td>
<td>6.22(^a)</td>
<td>0.79</td>
<td>2.94</td>
</tr>
<tr>
<td>Baltimore</td>
<td>3.39(^a)</td>
<td>0.94</td>
<td>11.20</td>
</tr>
<tr>
<td>Boston</td>
<td>1.76</td>
<td>0.95</td>
<td>13.51</td>
</tr>
<tr>
<td>Chicago</td>
<td>3.08(^b)</td>
<td>0.85</td>
<td>4.27</td>
</tr>
<tr>
<td>Cincinnati</td>
<td>2.01</td>
<td>0.87</td>
<td>4.98</td>
</tr>
<tr>
<td>Cleveland</td>
<td>2.08</td>
<td>0.92</td>
<td>8.31</td>
</tr>
<tr>
<td>Detroit</td>
<td>3.11(^b)</td>
<td>0.86</td>
<td>4.60</td>
</tr>
<tr>
<td>Houston</td>
<td>2.92(^b)</td>
<td>0.88</td>
<td>5.42</td>
</tr>
<tr>
<td>Kansas</td>
<td>3.49(^a)</td>
<td>0.79</td>
<td>2.94</td>
</tr>
<tr>
<td>Los Angeles</td>
<td>4.01(^a)</td>
<td>0.79</td>
<td>2.94</td>
</tr>
<tr>
<td>Minneapolis</td>
<td>2.04</td>
<td>0.92</td>
<td>8.31</td>
</tr>
<tr>
<td>New York</td>
<td>2.98(^b)</td>
<td>0.86</td>
<td>4.60</td>
</tr>
<tr>
<td>Philadelphia</td>
<td>4.07(^a)</td>
<td>0.74</td>
<td>2.30</td>
</tr>
<tr>
<td>Pittsburgh</td>
<td>3.23(^b)</td>
<td>0.71</td>
<td>2.02</td>
</tr>
<tr>
<td>Portland</td>
<td>2.37</td>
<td>0.92</td>
<td>8.31</td>
</tr>
<tr>
<td>Seattle</td>
<td>1.28</td>
<td>0.96</td>
<td>16.98</td>
</tr>
<tr>
<td>San Francisco</td>
<td>0.65</td>
<td>0.98</td>
<td>34.31</td>
</tr>
<tr>
<td>St Louis</td>
<td>1.16</td>
<td>0.94</td>
<td>11.20</td>
</tr>
<tr>
<td>Washington DC</td>
<td>2.99(^b)</td>
<td>0.86</td>
<td>4.60</td>
</tr>
</tbody>
</table>

\(^a\) Denotes significance at the 1% level.  
\(^b\) Denotes significance at the 5% level.

\(^{10}\) Cachetti et al. (2001) use Chicago while Culver and Papell (1999) experiment with different city bases. None of our results depend on our US base.
We reject nonstationarity at the 10% level for eleven of the 19 cities\textsuperscript{11}. The results are similar to those of Cachetti et al. (2001).

The results of stationarity tests for the food and rent sub-indices are also of interest. For rent, we reject nonstationarity at the 10% level for four cities using the overall US rent price index as our base. For food, we reject nonstationarity at the 10% level for just two cities\textsuperscript{12}. At first glance, the results for food are surprising since we might expect to find more evidence for food stationarity than for the overall price level on the grounds that food is more traded. As we shall see, this argument is flawed.

The third column in Table 1 provides the AR(1) coefficients, $\rho$, while column four gives the associated half-life. We calculate the half-life as $-\ln(2)/\ln \rho$. In accordance with previous work, we find that speeds of adjustment for the overall price level are slow. The mean of the autoregressive coefficients in Table 1 is 0.87 implying a half-life of deviations from PPP of almost five years, slow compared to the three to five years found in international data sets\textsuperscript{13}. Turning to the disaggregated data, the results for food are similar to the overall price level while speeds of adjustment for rents are exceedingly slow with the average around 17 years.

As mentioned, Culver and Papell (1999) and Cachetti et al. (2001) attribute their difficulties rejecting city real exchange rate nonstationarity to slow speeds of adjustment\textsuperscript{14}. An alternative explanation is that city price level convergence has produced nonstationarity. Recall that stationarity requires that the relative price level between any two cities return to a constant level. Clearly this cannot hold when city price levels are converging over time. Given the convergence in the overall price and food price levels revealed by Figs. 1 and 2, we would not expect to find bilateral real exchange rate stationarity for food or for the overall price level.

This line of argument has a striking implication. It is taken for granted in the literature that PPP requires real exchange rate stationarity. It should be apparent that this rule ignores the case where nonstationarity arises because of price level convergence brought about by increased market integration. Put differently, real exchange rate nonstationarity is not evidence against broad versions of PPP if nonstationarity is accompanied by price level convergence. By broad versions we mean versions of PPP that allow for reductions in transport costs and improved market integration.

\textsuperscript{11} It is more difficult to reject nonstationarity using individual city price indices as base. PPP does best with a Chicago base where we reject nonstationarity at the 10% level for 11 of the 18 bilateral real exchange rates. On the other hand, we cannot reject nonstationarity for a single city pair using a San Francisco base. The average for all US city bases is around eight rejections.

\textsuperscript{12} By contrast, Cachetti et al. (2001) reject nonstationarity with pooled city data for tradables and nontradables where they construct their price measures using disaggregated price data from 1972 to 1995.

\textsuperscript{13} Our estimates are higher, on average, than Culver and Papell (1999) who use pooled data covering 12 of the 19 cities in our sample. Cachetti et al. (2001) using long span data for our cities, however, find half-lives equal to nine years with pooled data after they adjust their $\rho$ estimates to control for small sample bias. We find that the average $\rho$ coefficient for the 18 bilateral city real exchange rates using their base of Chicago is 0.87 implying a half-life of five and a quarter years. The average speeds of adjustment for most other city bases lie between five and six years.

To underline this point, let us return to food prices. Relying on the stationarity tests, we would conclude that PPP fails for food or that PPP holds better for the overall price level than for food. But this conclusion makes little sense given that price dispersion is lower for food than for the overall price level and food price dispersion has fallen by 60% since the 1920s. It is more likely that our inability to reject nonstationarity for food is the consequence of improved market integration. In other words, nonstationarity when combined with price level convergence actually supports broad versions of PPP.

Of course, nonstationarity may also be associated with absolute price level divergence as in the case of rents considered earlier. In this case, PPP clearly fails.

Absolute price level convergence has a further implication. When price levels are converging, bilateral real exchange rates will contain a trend or one or more breaks depending on the pattern of convergence. A visual inspection shows that many city real exchange rates appear to have breaks or trends. Using bilateral city data from 1918 to 1987, Sonora (University of Texas unpublished, 2002) finds that speeds of adjustment average two years after allowing for breaks. These speeds of adjustment are faster than those found internationally.\(^{15}\)

5. Summing up

We have studied absolute price level dispersion using long span US city data. The evidence shows that the dispersion of absolute price levels is lower for US cities than internationally, that city price level dispersion is falling over time and that price level dispersion is lower for traded as compared to non-traded goods. These findings are consistent with beliefs that markets are better integrated within countries than internationally and that market integration has increased over time for US cities. Finally, we argue that nonstationarity in city data is not evidence against PPP, rather it may be the result of price level convergence due to improved market integration. Furthermore, once we allow for price level convergence, it appears that speeds of adjustment are faster than those found internationally. In sum, the US city price data support PPP.

Acknowledgements

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References


\(^{15}\) When Cachetti et al. (2001) introduce a time trend, they find that speeds of adjustment are between two and four years.


