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Structural Breaks and Relative Price Convergence among U.S. Cities

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JEL Classifications: C33; E31; R19

Keywords: Relative price convergence; Structural break; Panel unit root test; Half-life; Nickell bias; Time aggregation bias

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1 Introduction

The study of relative price convergence across U.S. cities has been an important topic of research primarily for two reasons. *First*, some of the factors responsible for the breakdown of the international purchasing power parity (PPP) hypothesis, such as tariff and non-tariff barriers, fluctuations in nominal exchange rates, heterogeneity of consumption baskets that are used to construct consumer price indices in different countries, are not present while using price indices from different cities within the same country. Thus, the researchers can conduct a natural experiment with city price data to gain a better understanding of deviations from PPP. *Second*, the differences in the rate of inflation between cities – implied by relative price movements – determine the differences in real wages and real interest rates, which in turn influence the directions of labor and capital flows respectively. Thus, an understanding of relative price movements across cities could be useful in explaining regional growth. The existing literature focuses on the convergence of both aggregate as well as individual prices across different cities in the U.S.[1]

Most previous studies on relative price convergence using city level aggregate CPI data across the U.S. find evidence of convergence. But the estimated speed of convergence reported in these studies is very slow. For example, Cecchetti et al. (2002) estimate a half-life of 9 years using annual data for 19 cities between 1918 and 1995. Of course, this has been an issue that some subsequent studies addressed. For example, Nath and Sarkar (2009) correct for the Nickell and time aggregation bias – usually present in panel estimates of half-life – and are able to find a half-life estimate of about 7.5 years. Chmelarova and Nath (2010) further present unbiased estimates of half-lives that vary depending on the choice of the numeraire city.

All of the above studies use annual CPI data for a long period of time that began in the early 20th century. It is therefore conceivable that the relative price movements during this period were characterized by structural break(s). Any evidence of structural break(s) would be a potential source of upward bias in the estimates of autoregressive coefficients that are used to calculate half-lives.

[1] While Culver and Papell (1999), Cecchetti *et al.* (2002), Chen and Devereux (2003), Nath and Sarkar (2009); Chmelarova and Nath (2010), Sonora (2009), and Basher and Carrion-i-Silvestre (2009, 2011) examine price index convergence across major U.S. cities, Carrion-i-Silvestre *et al.* (2004) investigate it for cities in Spain and Sonora (2005) for Mexico. Chen and Devereux (2003) and Basher and Carrion-i-Silvestre (2009) consider absolute price convergence and others examine relative price convergence. These studies use aggregate price indices. However, they were preceded by two very influential papers by Engel and Rogers (1996) and Parsley and Wei (1996) that examine disaggregate prices of various commodities across the U.S. cities.

The international PPP literature has already incorporated structural breaks in a way of resolving the "PPP puzzle" *a la* Rogoff (1996).

Beginning primarily with Perron (1989), adding shifts in the mean (that represent structural breaks) of a real exchange rate series has been used as a solution to the inability to reject the unit root.[2] This has also been a way to reduce the length of the half-life of a deviation from long run PPP (LRPPP), thereby addressing the "PPP puzzle" of extremely slow mean reversion. Notable examples along this line of research in the literature include Dornbusch and Vogelsang (1991), Perron and Vogelsang (1992), Culver and Papell (1995), and Hegwood and Papell (1998 & 2002). These studies include only one structural break. Lumsdaine and Papell (1997) extend this analysis by adding a second break into the unit root test framework. The next development was to incorporate structural breaks in the panel context. For example, Papell (2002) tests several panels of between 11 and 20 real exchange rates for unit root incorporating multiple structural breaks. Other examples of panel studies of PPP with structural breaks include Im et al. (2005) and Narayan (2008).

To the best of our knowledge, Sonora (2009) and Basher and Carrion-i-Silvestre (2009, 2011) are the only studies that examine price behavior across cities in the U.S. with structural breaks. Using unit root test procedures that incorporate a single, and two structural breaks respectively to each city relative price series for 19 U.S. cities for the period from 1918 to 1997, Sonora (2009) reports average half-life estimates that range between 1.74 years and 2.96 years. In contrast, applying median unbiased estimates of autoregressive coefficients for each price series for 17 cities between 1918 and 2006, Basher and Carrion-i-Silvestre (2011) calculate a half-life for each city with a median half-life ranging between 1.5 and 2.6 years. Further, Basher and Carrion-i-Silvestre (2009) conduct panel stationarity tests with structural breaks to examine price level convergence among U. S. cities.[3] They discuss different concepts and definitions of PPP as they relate to city price convergence with structural breaks. They, however, do not provide any estimate of the half-life.

In this paper, we apply panel unit root test procedures that incorporate structural breaks to CPI data for 17 U.S. cities for a period from 1918 to 2010 to examine convergence behavior of relative prices across cities. In contrast, Basher and Carrion-i-Silvestre (2009) use panel unit root test procedures with structural break (that are different from our method) to investigate absolute price level convergence. Further, they include a time trend in their test equations and allow for both

[2] A rejection of the unit root null has been interpreted as evidence in support of PPP in the literature.

[3] Like Chen and Devereux (2003), they use annual CPI data for 17 U.S. cities between 1918 and 2005 to construct absolute price level series.

intercept and slope breaks. We, however, do not include a time trend and only allow for breaks in the intercept to be consistent with the LRPPP theory. Furthermore, we correct for two types of biases that arise in panel data estimates: small sample bias (the so-called "Nickell Bias") and time aggregation bias. The unit root tests with a single and two breaks respectively find evidence of convergence in relative prices across U.S. cities, which is consistent with the existing literature. The results further indicate that the unbiased estimate of half-life with two structural breaks is less than 3 years, which is substantially lower than most panel half-life estimates for U.S. city prices reported in the literature.

The rest of the paper is organized as follows. Section 2 describes the data. In Section 3, we present the results of panel unit root test procedures. We first report the results with no structural break and we then report the test results with structural breaks. Section 4 presents the unbiased estimates of half-life. Section 5 includes our concluding remarks.

2 Data

We use annual CPI data for 17 U.S. cities from 1918 to 2010, obtained from the Bureau of Labor Statistics (BLS). The cities in our sample are: Atlanta, Boston, Chicago, Cincinnati, Cleveland, Detroit, Houston, Kansas City, Los Angeles, Minneapolis, New York, Philadelphia, Pittsburgh, Portland, San Francisco, Seattle, and St. Louis. The choices with regards to the number of cities, the frequency of data, and the sample period are dictated by the availability of data.[4] We construct the relative price series for each city by using the following equation:

$$r_{i,t} = 100 \times \left(\ln P_{i,t} - \ln P_t\right) \tag{1}$$

where $r_{i,t}$ is the logarithm of relative price and $P_{i,t}$ is the CPI in city i in year t, and P_t is the U.S. city average CPI. Note that this relative price represents the percentage deviation of CPI in a city from

[4] Monthly CPI data are available for Chicago, Los Angeles, and New York only. Bimonthly data are available for Atlanta, Detroit, Houston, Philadelphia, San Francisco, and Seattle for even months (February, April, etc.) and for Boston and Cleveland for odd months (January, March, etc.). Semiannual CPI data are available for Cincinnati, Kansas City, Minneapolis, Pittsburgh, Portland, and St. Louis. However, these high frequency data are not available for the entire sample period that we consider here. For data availability, see http://www.bls.gov/cpi/cpifact8.htm. Cecchetti et al. (2002) and Sonora (2009) use a sample of 19 cities that include Baltimore and Washington, D.C. The BLS has stopped publishing CPI data separately for these two cities and, therefore, we do not include them in our sample.

the national average CPI.[5] The city average CPI captures the common time effect component of relative prices.

3 Panel Unit Root Tests

3.1 With no Structural Break

In order to determine how structural breaks affect relative price convergence, we first run a panel unit root test that does not incorporate structural change. The test is conducted by running the following regression:

$$\Delta r_{i,t} = \mu_i + \rho_i \, r_{i,t-1} + \sum_{j=1}^{k_i} c_{i,j} \Delta r_{i,t-j} + \varepsilon_{i,t} \tag{2}$$

The subscript $i = 1, \ldots, n$ indexes the cities in the panel. We allow for heterogeneous intercepts, μ_i, and lag lengths, k_i. Feasible generalized least squares (FGLS) seemingly unrelated regression (SUR) is used to estimate Eq.(2). This method accounts for contemporaneous and serial correlation, both of which are likely to be present in city relative prices.[6] The number of lagged differences, k_i, for each city i is determined by individual ADF tests using the general-to-specific method suggested by Campbell and Perron (1991) and Ng and Perron (1995). This method involves setting a maximum lag length, k^i_{max}, and paring it down to the number of lags where the lagged difference is significant. We start with a maximum lag of 8 years for each city.

The null hypothesis is that each series contains a unit root, H_0: $\varrho_i = 0$ for all i. The alternative hypothesis is H_1: $\varrho_i = \varrho < 0$, that is, all of the series are stationary. This alternative hypothesis requires a homogenous ϱ, as in Levin et al. (2002).[7] Note that the distribution of the panel unit root test is not standard. Therefore, we use Monte Carlo methods involving 5000 replications to calculate critical values that reflect both the number of cities in the panel and the number of observations, and also account for both serial and contemporaneous correlations.[8]

[5] In the international PPP literature, this relative price would be equivalent to the real exchange rate. Although a numeraire currency is chosen for calculating real exchange rate, we use the average city CPI, an approach previously adopted by Cecchetti et al. (2002), Chen and Devereux (2003), Nath and Sarkar (2009), and Basher and Carrion-i-Silvestre (2011).

[6] See Murray and Papell (2000).

[7] A less restrictive alternative hypothesis that at least one of the series is stationary, which allows ϱ to be heterogeneous, as in Im et al. (2003) would not be any more informative if we do reject the null, which is the case in this paper.

[8] For the details of this method, see Hegwood and Papell (2007)

[Insert Table 1]

As shown in Table 1, the unit root null hypothesis is rejected at the 1% level of significance indicating that each relative price series is stationary around its long run mean. This is what we would expect based on the LRPPP theory, particularly among cities within a single country, and this result is consistent with the results reported in other studies that use panel unit root test procedures (e.g. Cecchetti *et al*. 2002; Nath and Sarkar 2009; Chmelarova and Nath 2010).

3.2 With Structural Breaks

As we discuss in the introduction, although the existing literature finds overwhelming evidence of relative price convergence among U.S. cities, the speed of convergence is extremely slow. One possible reason for the long half-life is that the long run mean relative price may have experienced one or more permanent shifts or structural breaks. An examination of this possibility would first involve conducting panel unit root tests that allow for structural breaks. If we are able to reject the unit root null, we will then use the structural break dates to estimate corresponding half-lives. Thus, we first incorporate one, and then two structural breaks into our panel unit root test. We utilize an Additive Outlier (AO) model framework that allows for instantaneous change(s).[9] This model has been adapted for non-trending data incorporating one or two shifts of the intercept.[10]

The panel unit root test with structural breaks comprises two stages. The first stage in a test with one break involves running the following regression on the panel of relative prices:

$$r_{i,t} = \mu_i + \delta_i DU_t + u_{i,t} \tag{3}$$

The intercept break dummy variable, DU_t, equals 1 for all t greater than the break date, TB, and zero otherwise.[11] In the second stage, the residuals, $u_{i,t}$'s are regressed against their lagged value and lagged differences as follows:

$$\Delta u_{i,t} = \rho u_{i,t-1} + \sum_{j=1}^{k_i} c_{i,j} \Delta u_{i,t-j} + \varepsilon_{i,t} \tag{4}$$

[9] Hegwood and Papell (2007) use a similar framework to incorporate intercept and trend breaks to study the movements in real GDP in three groups of advanced countries.

[10] This is a panel adaptation of the univariate tests in Perron and Vogelsang (1992). They included an additional set of "crash" dummies.

[11] We impose a restriction by forcing the break date(s) to be the same for all cities. However, allowing different breaks for each city will reduce the degrees of freedom and the power of the test.

As in the panel unit root test with no structural break, the number of lagged differences, k_i, is determined by the general-to-specific method. Eq. (3) and (4) are estimated sequentially for each possible break year, $TB = k+2, \ldots T-1$, where T is the number of observations. The year that minimizes the t-statistic on ϱ is chosen to be the break date. The unit root null hypothesis is rejected if the absolute value of the minimum t-statistic on ϱ is greater than the appropriate critical value. As before, the critical values are calculated using Monte Carlo simulations. The result of this panel unit root test with a single structural break as shown in the first row of Table 2 indicates that the unit root null is rejected at the 1% level with the break in 1985.

[Insert Table 2]

It is likely that the behavior of city relative prices is characterized by more than one structural breaks. In order to examine this possibility and the resulting stochastic trending properties, we now extend our analysis to include a second mean shift.[12] We use the same AO model framework but simply include a second dummy variable for a second intercept shift. Appropriate critical values are calculated as before. The results are reported in Table 2. We reject the null hypothesis at the 1 % level with mean shifts in 1943 and 1990.

Overall, these results provide strong evidence of convergence in relative prices across U.S. cities and are consistent with the results reported by other studies in the literature. The break dates: 1985 with a single break, and 1943 and 1990 with two breaks, warrant some explanations as to the significance of those particular years. However, since these breaks are endogenously determined strictly based on statistical criteria, it is often very difficult to speculate on one or more particular reasons for these permanent shifts without further investigation. It is more so because these permanent mean shifts may result from a combination of a number of factors/events. Nevertheless, we plot relative prices for each city with two structural breaks in Figure 1 to have a visual sense of how relative price behavior changes around these breaks.[13]

[Insert Figure 1]

[12] In principle, we can include more structural breaks. But, since we impose the restriction that break dates are the same across cities, more breaks will represent significant mean shifts for increasingly less number of cities. As an illustration of this point, when we include one break, it is visually prominent for 16 out of 17 cities. In contrast, when we include two breaks, both are prominent for only 7 cities (See Fig. 1 and related discussion). Furthermore, two breaks seem to have fitted the data reasonably well for most cities.

[13] We present the graphs with two structural breaks as an illustration. In comparison to one break, two breaks seem to fit the data better. To save space, we do not include the plots with one break. However, they can be obtained from the authors.

We make a few observations. *First*, both structural breaks are visually prominent for seven cities: Boston, Houston, Kansas City, Minneapolis, Pittsburgh, San Francisco, and Seattle. In contrast, none of these two breaks is visually prominent for Chicago and Philadelphia. For the rest of the cities, one break is more prominent while the other is not. *Second*, after the first break in 1943, the mean got closer to 0 for almost all cities, implying that on an average, city CPIs were getting closer to the national average. This shift occurred in the aftermath of the great depression and coincided with the World War II.[14] However, it would require an in depth examination of a number of factors including the fiscal and monetary policies of that time to explain this shift, which is not a focus of this paper.

In contrast to the first break, the mean shifted away from 0 after the second break in 1990 for most cities. While prices, on an average, rose above the national average for some cities, particularly in the east as well as the west coast (e.g. Boston, New York, Los Angeles, San Francisco, and Seattle), prices fell, on an average, below the national average for others, particularly cities in the Midwest and the South (e.g. Atlanta, Cincinnati, Detroit, Houston, Kansas City, St. Louis). Although it needs further investigation, it may be related to the sectoral shifts that took place in the 1990s and afterwards. With the manufacturing sector shrinking and information technology and financial sector growing, there was a spatial redistribution of economic growth and decline across various regions in the U.S.[15]

4 Unbiased Half-Life Estimates with Structural Breaks

Half-life is the time required for any deviation from LRPPP to dissipate by one half and is commonly used as a measure of the speed of convergence. In an AR(1) case, half-life is calculated as follows:

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

[14] Although an attempt to make any connection will be too far-fetched, President Roosevelt coincidently signed an executive order freezing prices, salaries, and wages to prevent inflation.
[15] Interestingly, 1990 is marked by the debut of the worldwide web ushering in the era of internet.

where *h*( ) is the half-life and ϱ is the AR coefficient. In the literature on relative price convergence among U.S. cities, the half-life is estimated to be excessively long - even longer than the commonly accepted range of half-life of 3 to 5 years for international PPP.

As Choi et al. (2006) emphasize, three potential biases may arise in panel data estimation of the half-life. *First*, if there is sufficient heterogeneity in the dynamic behavior of price indices across cities (that is, the autoregressive coefficients are significantly different across cities), then panel estimation of a common autoregressive coefficient will be biased upward and so will be the implied half-life. *Second*, in small samples, inclusion of a constant in the estimation of a dynamic regression introduces a downward bias. In the panel context, it is known as the "Nickell bias", after Nickell (1981) who first discusses this small-sample bias in the panel data. *Finally*, an additional upward bias potentially arises from the fact that the annual CPI data are averages of goods and services prices recorded monthly, rather than point-in-time sampled prices.[16] This time-averaging (also referred to as *time aggregation*) process introduces a moving average structure into the regression error. This is often ignored in the panel estimation of the autoregressive models of city relative prices. Because the magnitude of half-life is very sensitive to the value of autoregressive coefficients, failure to correct for those biases in panel estimation of these coefficients can lead to inaccurate measure of the half-life.

As in Choi et al. (2006) and Nath and Sarkar (2009), we first conduct a test of heterogeneity of estimated autoregressive coefficients of relative prices in different cities in order to determine if there is any possibility of an upward bias in panel estimate of the half-life. We first obtain recursive mean adjusted seemingly unrelated regression (SUR) estimates of the autoregressive coefficients of relative prices and then construct a Wald test statistic with homogeneity restrictions under the null hypothesis.[17] This procedure is shown to have the desirable property of mitigating size distortion due to the small-sample bias. Note that, for the cases with structural breaks, we include dummy variables for the respective breaks in the test equation. The test results are reported in Table 3. Comparing with the 5% critical value, we conclude that there is little evidence of heterogeneity in the autoregressive coefficients of relative prices across cities with and without structural break.

[Insert Table 3]

[16] For a discussion, see Imbs et al. (2005)

[17] For a detailed discussion, see Choi et al. (2004)

Thus, the results presented in Table 3 suggest that, irrespective of whether we include structural break or not, it is appropriate to apply panel estimation techniques to pooled data. Thus, the panel estimates of autoregressive coefficient and half-life for city relative prices involve two potential biases: a downward bias due to small sample size and an upward bias due to the moving average error term introduced by time aggregation of data. We use a fixed effects panel generalized least squares (GLS) estimation technique that incorporates structural breaks and also controls for cross-sectional dependence.[18] To sketch an outline of the procedure, suppose relative price in city *i* follows an AR(1) process:

$$r_{i,t} = \alpha_i + \sum_{j=1}^{m} \delta_j DU_j + \rho_i r_{i,t-1} + u_{i,t} \tag{6}$$

where a_i is a city-specific constant; $i = 1, 2, \ldots, n$; $m = 1$ or 2 depending on whether we include one structural break or two; and $t = 1, 2, \ldots, T$. DU_j is a dummy variable for structural breaks where $DU_j = 1$ if $t > TB_j$ for $j = 1, 2$ and 0 otherwise. In the presence of time aggregation the regression error has a moving average (MA) structure. Suppose $u_{i,t}$ follows an MA(1) process:

$$u_{i,t} = \upsilon_{i,t} + \lambda \upsilon_{i,t-1} \text{ and } \upsilon_{i,t} = \gamma_i \theta_t + \zeta_{i,t} \tag{8}$$

where γ_is are factor loadings, θ_t is the common shock, and $\zeta_{i,t}$s are serially and mutually independent. We estimate the factor loadings and the error covariance matrix by iterative method of moments, and then use the estimated covariance matrix to obtain the feasible GLS estimate of ρ. Note that this estimated covariance matrix includes both the contemporaneous and the long-run covariance. We then adjust the estimated autoregressive coefficient for the Nickell bias, the time aggregation bias, and the combined Nickell and time aggregation bias as discussed in Choi et al. (2006) and use these bias-corrected estimates of autoregressive coefficient in Eq. (5) to obtain various unbiased estimates of the half-life to price index convergence among U.S. cities.[19]

[Insert Table 4]

The results reported in Table 4 indicate that the inclusion of structural breaks lowers the half-life estimate from 10.7 years to 6.5 years with a single break and to 4.5 years with two breaks, which

[18] This technique has been adapted from Phillips and Sul (2004) to include structural breaks.
[19] Time aggregation of the data introduces an interaction between the Nickell bias and the time aggregation bias, which requires additional adjustment in the estimation of the autoregressive coefficient. For a discussion, see Choi *et al.* (2006). The combined Nickell and time aggregation bias correction incorporates this adjustment.

represent a reduction of about 39% and 58% respectively. When the autoregressive coefficient estimates are corrected for Nickell and time aggregation bias, the estimated half-life is further reduced to 3.9 and 2.8 years, a reduction of 64% and 74% respectively.

These results clearly indicate that both structural breaks and bias correction go a long way in resolving the puzzling result of unrealistically slow mean reversion in relative prices across U.S. cities, that has been reported by the previous studies. The panel half-life estimates presented in this paper compare well with those reported in Sonora (2009) and Basher and Carrion-i-Silvestre (2011). Both studies estimate half-life for each city relative price separately. The average half-life estimates in Sonora (2009) with a single break and with two breaks are slightly lower than our estimates. However, if we re-estimate the half-lives using data for 19 cities for a shorter sample period: 1918-1997 (as in Sonora 2009), the estimated half-lives are very similar to those averages. It is intriguing and needs further investigation that inclusion of more recent data yields a slower speed of convergence. In contrast, Basher and Carrion-i-Silvestre (2011) use data for a longer sample period (1918 to 2006) for the same cities as in the current study and report average half-life estimates that range between 2.1 years and 3.9 years. However, they allow for a time trend and multiple structural breaks.

5 Concluding Remarks

Slow mean reversion in real exchange rates has often been referred to as the "PPP puzzle" because it contrasts with the basic premise of the LRPPP theory that market forces should return real exchange rates to their long run equilibrium level after a temporary deviation. Nowhere else is this puzzle more enigmatic than in the literature on price index convergence among U.S. cities, which has been primarily motivated by a desire to enhance our understanding of the reasons for the breakdown of international PPP. Most studies in this literature report excessively long half-life estimates. Although several attempts have been made to resolve this issue of slow convergence, reasonable estimates of half-life have been elusive. This paper re-examines the price index convergence among U. S. cities by applying panel unit root test procedures that allow for structural breaks to annual CPI data between 1918 and 2010 for 17 major cities. With an endogenously determined single break in 1985, and two breaks in 1943 and 1990 respectively, the test results provide strong evidence of convergence of relative prices across cities, which is consistent with the existing literature. Most importantly, this study finds that the speed of convergence with structural break(s) is much faster than that reported by previous studies that use panel methods. Furthermore, correcting for small-sample bias (the so-called "Nickell Bias") and time aggregation bias generates a half-life of 2.8 years, which is 74% shorter than the half-life estimate with no structural break and no bias correction.

Table 1. Panel Unit Root Test Result: No Structural Break

	Estimated test statistic	Critical values		
		1%	5%	10%
	(1)	(2)	(3)	(4)
Panel unit root test with no structural break	-9.25	-7.94	-7.33	-6.97

Note: The critical values are generated from Monte Carlo simulations

Table 2. Panel Unit Root Test Results: Structural Breaks

	Estimated test statistic	Critical values		
		1%	5%	10%
	(1)	(2)	(3)	(4)
One structural break in 1985	-12.17	-8.11	-7.53	-7.27
Two structural breaks in 1943 and 1990	-15.22	-11.77	-11.25	-10.60

Note: The critical values are generated from Monte Carlo simulations

Table 3. Homogeneity Test Results

	No structural break	One structural break	Two structural breaks	5% Critical value $\left(\chi^2_{n-1,0.05}\right)$
	(1)	(2)	(3)	(4)
Wald test statistic	23.04	22.11	9.37	26.30

Note: The null hypothesis is H_0: $\varrho_1 = \varrho_2 = \ldots = \varrho_n$ where ϱ_i is the AR(1) coefficient of relative price in city *i*. Under the null, the estimated test statistic follows a *chi*-squared distribution with *n*-1 degrees of freedom. The Wald test is described in Choi et al. (2004).

Table 4. Panel Feasible GLS Estimation of ϱ and Implied Half-Life

	No bias corrections		Nickell bias corrected		Time aggregation bias corrected		Nickell and time aggregation bias corrected	
	$\hat{\rho}$	Half-life	$\hat{\rho}$	Half-life	$\hat{\rho}$	Half-life	$\hat{\rho}$	Half-life
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
No structural break	0.937	10.65	0.962	17.89	0.905	6.94	0.930	9.57
One structural break in 1985	0.899	6.51	0.898	6.44	0.817	3.43	0.839	3.94
Two structural breaks in 1943 and 1990	0.856	4.46	0.858	4.53	0.761	2.54	0.783	2.83

Figure 1. Relative Prices with Two Structural Breaks in 1943 and 1990

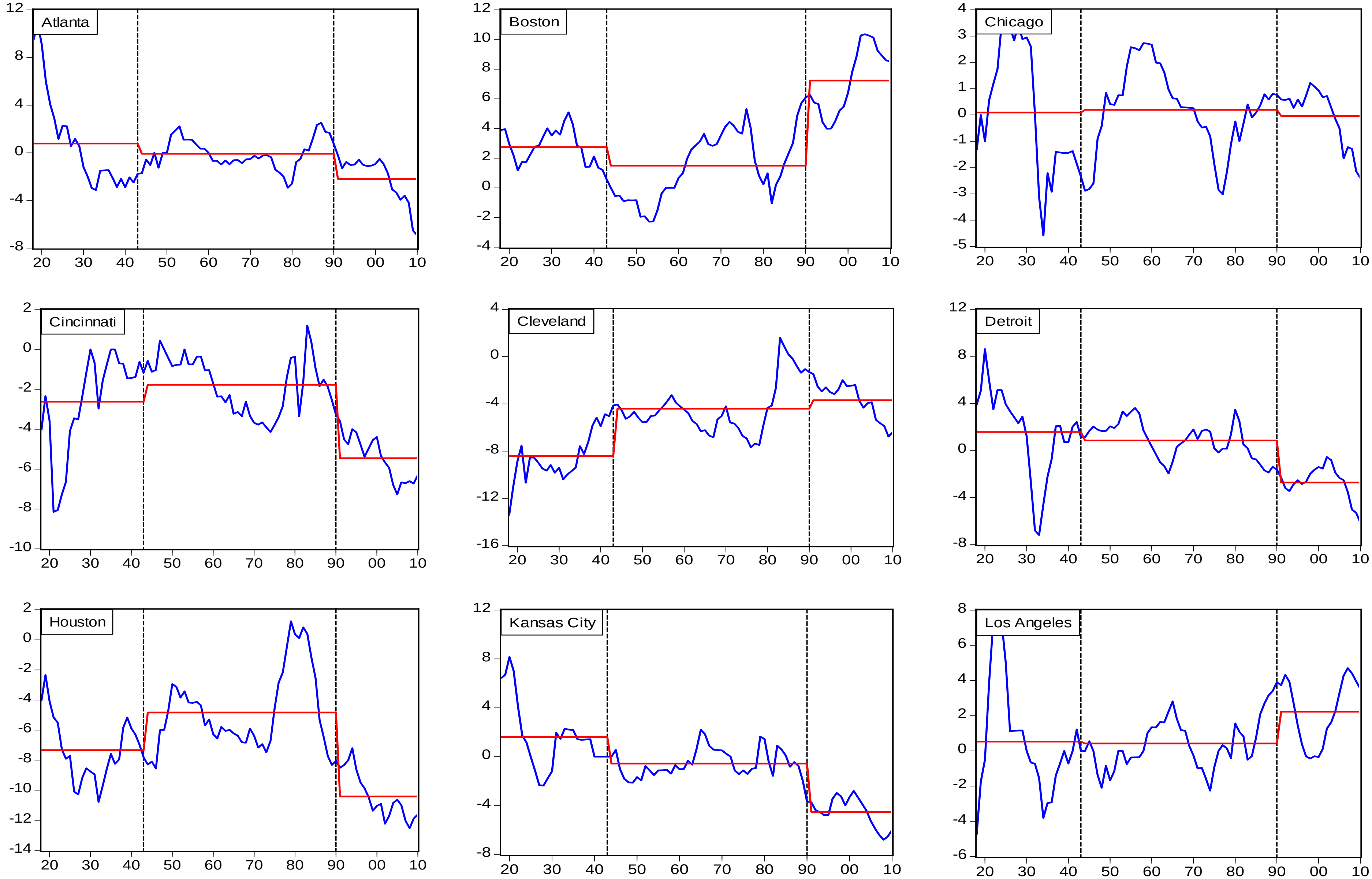
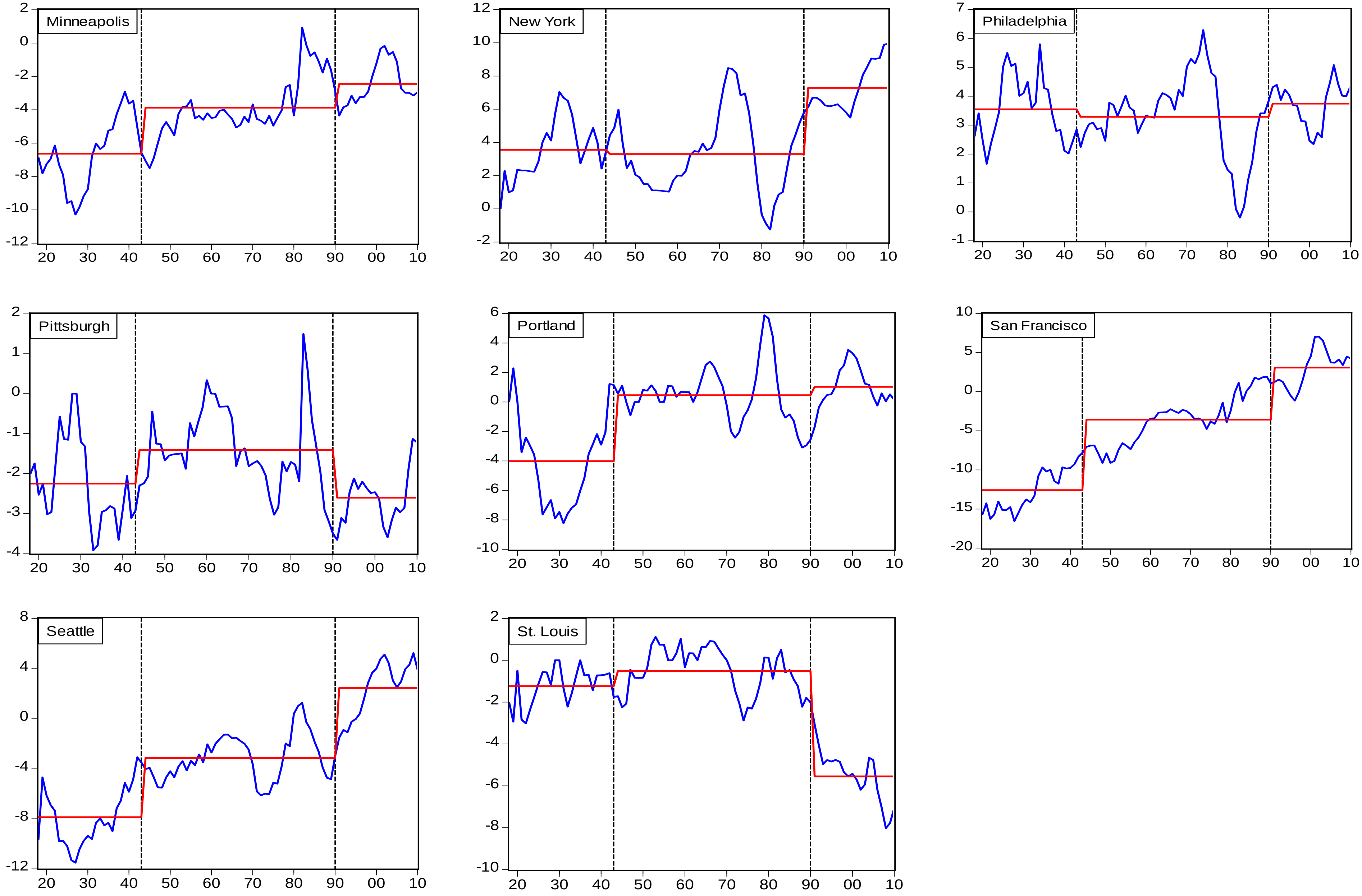


Figure 1. Relative Prices with Two Structural Breaks in 1943 and 1990 (continued)



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