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Price Convergence Among Indian Cities: A Cointegration Approach

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Abstract

Price dynamics in Indian cities were examined using cointegration analysis. We identified and calculated a common trend for prices in 25 major cities in India. Impulse response functions were obtained to calculate the rates of convergence to the prices and we found that the half-life of any shock is very small for Indian cities. Although a close to three-month half-life seems too fast, there is some indication in the literature that half-life can be much smaller than the conventional rates of 3 to 5 years. We have calculated half-life using the panel unit root method, and found that estimates of half-life from cointegration analysis provide a faster convergence rate than estimates using the panel unit root method. We also analyzed how shock can be transmitted from one city to another and found no systematic behavior of transmission from one city to another.

JEL Classification Code: F15, E31, C23.

Keywords: Law of One Price, Price Convergence, Half-life, Cointegration.

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I. Introduction

Recent efforts to understand the price behavior of the same good at different locations have produced a body of literature that supports the notion that the Purchasing Power Parity (PPP) holds in the long-run; however, over the short-term, the real exchange rate can deviate from its PPP equilibrium. The consensus among economists is that deviation of the exchange rate from their PPP level damp out at a rate of roughly 15 per cent per year, implying that these deviations have a “half life” of three to five years (Rogoff, 1996; Obstfeld and Rogoff, 2000). The observed deviations may be due to many reasons, including the Balassa-Samuelson effect (Balassa, 1964; Samuelson, 1964) postulating that cross-country productivity differentials between traded and non-traded sectors will lead to changes in real costs and the price of traded goods relative to the non-traded goods, and subsequently affect the real exchange rate, in particular for the medium and long-term. Cheung and Lai (2000), however, found cross-country variations in this half-life of a shock, for example, a smaller half-life in developing countries than in industrial countries.

Virtually all researchers argue that the presence of a nominal exchange rate and trade barriers between locations are important factors in generating these results. This idea has prompted researchers to conduct experiments with data from cities within a country to exploit the benefits of no trade barrier and no nominal exchange rate fluctuations. Parsley and Wei (1996) used a panel of 51 commodity prices from 48 cities in the US to estimate the rate of convergence to PPP and found that the convergence rates were higher for relative prices calculated for cities nearby than for cities farther apart, but that the distance between locations could explain a small portion of the differential rates of convergence. They concluded that the half-life of the price gap for traded goods is roughly four to five quarters, while it is fifteen quarters for services. Using

price data from 24 U.S. cities on individual goods, O'Connell and Wei (2002) found that relative prices across U.S. cities are in general stationary and their reversion to equilibrium is significantly faster than those for relative international prices. However, Cecchetti et al. (2002) investigated the aggregate prices (consumer price indices) for 19 U.S. cities from 1918 to 1995 and found that the half-life of convergence to an average CPI of these 19 U.S. cities was approximately nine years. Sonora (2005), using Mexican city price indices, found that the city relative prices were stationary and that the half-life estimate was around 2.5 years. Imbs et al. (2005), on the other hand, argue that the slow rate of convergence arises because the aggregation in calculating CPI creates a bias that results in the sharp decline in the rate of convergence. Their empirical results with disaggregated data from the European Union (EU) countries indicate that PPP holds even in the short-run. However, Chen and Engel (2005) showed that the aggregation bias suggested by Imbs et al. (2005) might not be responsible for the slow rate of convergence.

Cecchetti et al. (2002) examined the city price movements in the United States with the assumption that the United States can be perceived as a collection of developed economies where monetary policy is conducted by one central bank, the Federal Reserve System. They found that the divergence of city prices was temporary but persistent. Since the EU is a collection of developed countries, and the European Central Bank (ECB) conducts its monetary policy, the study of city price convergence in the USA can improve our understanding about regional variation of prices. Sonora (2005) examined Mexican city prices where Mexican cities can be considered as a collection of middle-income countries. Increasingly, developing countries have been looking into the possibility of economic union. For example, the Korea Monetary and Finance Association (KMFA) held an International Forum on Monetary and Financial Cooperation for Asia in 2004 to discuss the possibility and feasibility of an East Asian monetary

union and integration¹. Similar proposals are being discussed in the South Asian context as well. Oil-rich Gulf Arab states have also stepped up efforts towards their goal of a monetary union. The Gulf Cooperation Council (GCC) custom union was launched in 2003, representing an important step in the process of integration. GCC central bank governors have scheduled the establishment of a monetary union in 2005, a common market in 2007 and a single currency in 2010.

What might be the nature of price dynamics in an economic union among developing countries warrants attention. To the best of our knowledge, there is no study dealing with price dynamics in an economic union of developing countries. Since there is no economic union of developing countries, we believe that by considering Indian cities as a collection of developing countries with no trade barrier and no minimal exchange rate fluctuations, we can examine both short-run and long-run price dynamics. This paper is an attempt to understand price dynamics in more integrated developing countries by empirically examining the price dynamics in Indian cities.

In this paper we investigated the consumer price behavior of 25 major cities in India with monthly data for 156 months beginning with October 1988, and calculated the rate of convergence to PPP using cointegration analysis. Generally, researchers use the same model for panel unit root tests as the model to estimate autoregressive equations and then calculate the half-life of the shocks by employing an approximation technique². However, this calculation of half-life is appropriate only for first order autoregressive processes (Goldberg and Verboven, 2004, footnote 11). For higher order autoregressive processes, this formula would yield biased

¹ A recent Journal of Asian Economics featured a special issue (Vol. 15, Issue5, October 2004) on Asian economic integration.

² The formula to calculate half-life is $\ln(0.5)/\ln(1+b)$, and b is the coefficient of the lagged price in the regression model for the augmented Dickey-Fuller test.

estimates. Consequently, we used impulse response functions to calculate half-lives; it is well known that impulse response analysis is applicable to any autoregressive structure.

We found that there was only one common trend for the chosen 25 cities in India. This unique common trend traced very closely the overall CPI of India. We decomposed the effects of a shock into a stochastic trend effect and a stationary effect. Examining impulse response functions for a shock in the price of the own city, we found that the half-life, defined as the period when the marginal change of the stationary component becomes half of the initial jump, was exceptionally small. In fact, the average half-life was found to be around only three months. This suggests a much faster rate of convergence than that reported in the literature. Moreover, our use of cointegration technique allowed us to examine the effects of a shock in one city in relation to another city. We observed that there is no clear pattern in this context in Indian cities. For example, for a unit shock in Bombay, the half-life of its impact on the price in Nagpur was more than five months, while for a unit shock in Nagpur, the half-life of its impact on the price of Bombay was found to be two months. Consequently, the use of distance between locations in determining the differentials in half-life should be explained with more caution because there may be an asymmetry.

Section II describes the data and its collection in more detail. Section III discusses the appropriate methodology and why it is imperative to adopt the cointegration techniques. After a common trend was calculated, we used impulse response functions to calculate the half-life for shock in the same city and also for shock originated in other cities. We also calculated half-life using panel unit root methods. The results are reported in section IV. A final section summarizes our main conclusions.

II. Data

We collected monthly consumer price indices for industrial workers for 25 large cities in India for the period of 1988-2001 (base year 1982). The Indian Labour Bureau changed its base year to 1982, starting from October 1988. In addition to this, India undertook conscious efforts to open the economy and tried to make the economy more market friendly starting from 1991 (Rodrik and Subramanian, 2004). As a result, we believe that it is imperative to examine more recent price data more carefully. These secondary data were collected from various issues of *Indian Labour Journal*, a monthly publication of the Indian Labour Bureau. These large cities are from 12 states of India and the federal territory of Delhi (shown in Appendix Table 1). The Indian Labour Bureau reports CPI data for about 70 large cities in India. Among them, we have selected 25 cities by the population size.

To generate these consumer price indices, Family Income and Expenditure Surveys were conducted at 76 industrially important centers spread across India during 1981-82. A total of 32,616 families and 226 markets were surveyed. Price data were collected for 260 items. Retail price data were collected regularly regarding items commonly consumed by working class families from fixed shops in the selected markets of each center on a fixed price collection day every week. The weekly prices were pooled to get monthly average prices³. Some descriptive statistics are reported in the following Table 1.

³ For details, see *Indian Labour Journal*, January 1989, Vol. 30, No. 1, pages 57-60.

Table 1
Descriptive Statistics: Annual Inflation Rates (calculated from our sample)

Year	Mean Inflation	Standard Deviation	Coefficient of Variation
1990	5.15	2.91	0.57
1991	16.58	3.59	0.21
1992	12.99	2.97	0.23
1993	5.79	3.25	0.56
1994	8.95	2.45	0.27
1995	9.88	3.09	0.31
1996	8.01	3.18	0.40
1997	11.12	2.72	0.24
1998	9.81	3.27	0.33
1999	10.70	5.99	0.56
2000	2.37	2.02	0.85
2001	3.46	2.66	0.77

The mean annual inflation rates show an inflationary cycle in the 1990s. While the highest yearly inflation was observed in 1991 (16.6%), the lowest inflation was observed in 2000 (2.4%). It is, however, interesting to note that the standard deviations of the rate of inflation in the selected 25 cities are relatively small during both high inflation and low inflation periods (column 2). This indicates that the inflation in one city in India has been more in line with inflation in other cities and thus suggests a possible quick convergence of city prices. Population data are from the 1991 census and were collected from the United Nations Statistics Division from the website <http://unstats.un.org/unsd/citydata/>. We also collected CPI for India from the International Financial Statistics of the International Monetary Fund (IMF). Distance between cities in miles was calculated using latitude and longitude for each city in the website <http://www.indo.com/distance/>.

III. Methodology and Preliminary Data Analysis

Cecchetti et al. (2002) used the average of city price indices as the underlying stochastic trend, which can be viewed as a dynamic factor, and estimated the half-life of convergence of each of the 19 city price indices to this trend. They first employed the panel unit root test to

examine whether the relative prices with respect to this trend were unit root processes, and then estimated the half-lives using the panel autoregressive model. However, if there is more than one stochastic trend, the panel autoregressive model cannot capture long-run dynamics among the cities, as this model captures only one of the trend components. Further, the formula to estimate half-lives is similar to that based on the autoregressive model of order one, AR (1), for short-run dynamics, and thus is not an appropriate formula to use for the type of model considered in their analysis. (This point was addressed in Goldberg and Verboven, 2004.)

Cecchetti et al. (2002) found that the real exchange rates between U.S. cities were stationary, i.e., they do not contain any unit root⁴. However, the deviation from the common trend was highly persistent, and the average half-life estimate was about nine years; the slower rate of convergence was weakly related to the distance between locations⁵.

In order to capture the long-run dynamics, we employed cointegration analysis by considering the log of the price indices of the 25 cities in India as a vector time series. This cointegration analysis enabled us to capture the short-run dynamics among the cities involved. As a by-product of cointegration analysis, we were able to obtain the impulse responses of each of the price indices, attributable not only to the shock to its own index, but also to the shocks to

⁴ This implies that the log of the price index of one city and the log of the price index of the numeraire city, Chicago in this case, are cointegrated with a cointegrating vector $(1, -1)'$ and the cointegrating rank of the log of the price indices of 19 cities including Chicago is 18. Thus, there is one common trend, which is the mean of the log of the price indices.

⁵ Cecchetti et al. (2002) analyzed the time series properties of city real exchange rates in the United States. The city real exchange rates were computed by taking the natural logarithm of the consumer price index for each city and then dividing this by the natural log of the consumer price index in Chicago. Although Papell (1997) showed that the choice of numeraire currency is significant in the context of the international tests of the PPP, Cecchetti et al. (2002) argue that the panel unit root testing with common time trend effect make the choice of numeraire city irrelevant. The common time trend in the panel unit root test takes into account the effects of any change in the numeraire price index.

other cities' price indices. Based on these impulse responses, we determined the half-lives of price index convergences to their own shocks, as well as to the shocks in other cities⁶.

Cointegration analysis is based on the following error correction representation of the vector autoregressive model:

$$\Delta \mathbf{y}_t = \mathbf{d} + \mathbf{a} \mathbf{b}' \mathbf{y}_{t-1} + \sum_{j=1}^{p-1} \Phi_j^* \Delta \mathbf{y}_{t-j} + \mathbf{e}_t, \quad (1)$$

where $\mathbf{y}_t$ is a 25-dimensional vector whose components are the logs of city price indices at time t and $\Delta \mathbf{y}_t = \mathbf{y}_t - \mathbf{y}_{t-1}$, $\mathbf{a}$ and $\mathbf{b}$ are $25 \times r$ matrices. (The cointegrating rank r is to be determined.) The matrix $\mathbf{a}$ is called the speed of adjustment matrix and the columns of $\mathbf{b}$ are linearly independent cointegrating vectors with $\mathbf{b}' \mathbf{y}_{t-1}$ representing the long-run equilibrium errors. As there are structural breaks around July 1998 attributable to regional tensions due to nuclear tests in India and Pakistan and conscious efforts of the central bank of India to offset potential contagion of a financial crisis in East Asia by increasing money supply, the vector series $\mathbf{y}_t$ is adjusted for these structural breaks. Based on this representation, we can obtain the common stochastic trends $\mathbf{b}'_{\perp} \mathbf{y}_t$, where $\mathbf{b}_{\perp}$ is a $25 \times (25 - r)$ matrix such that $\mathbf{b}'_{\perp} \mathbf{b} = 0$ and explain the short-run dynamics through the Φ_j^* . By inverting this model into

$$\mathbf{y}_t = \mathbf{m} + \sum_{j=0}^{\infty} \Psi_j \mathbf{e}_{t-j} \quad ^7,$$

⁶Cecchetti et al. (2002) investigated the half-lives to its own shock only.

we can obtain the impulse response, Ψ_j , through which the half-life is obtained.

Based on the Akaike Information Criterion (AIC), we identified the autoregressive order p as 2. In order to identify the cointegrating rank r , we examined the squared partial canonical correlations (SPCCs) between $\Delta \mathbf{y}_t$ and $\mathbf{y}_{t-1}$ adjusted for $\Delta \mathbf{y}_{t-1}$. The most commonly used test statistics for testing the null hypothesis of cointegrating rank of r are the trace statistic (TR) and the maximum eigenvalue (ME) statistic:

$$TR = -n \sum_{i=r+1}^m \ln(1 - \hat{l}_i) \text{ and } ME = -n \ln(1 - \hat{l}_{r+1}),$$

respectively, where $\hat{l}_i$ is the i -th largest SPCC and n is the sample size. The asymptotic distributions of these statistics are well known and tabulated in, for example, Johansen and Juselius (1990) for the case without structural break, and in Lütkepohl et al. (2003) for the case with structural shifts. Johansen et al. (2000) investigated the asymptotic distributions for the case with structural breaks in the deterministic trend and suggested using approximation based on simulation for the critical values. However, this approach is not directly applicable to our data, as the indices of some cities have different forms of structural breaks and different time points for structural breaks. Investigation of the asymptotic distribution of the above test statistics for the case of our data will be an interesting future econometric study and will not be pursued here. As an exploratory measure, we examined the relative magnitude of the SPCCs that are tabulated in Table 2. The smallest SPCC was only about 14 percent of the second

⁷ This is understood as $\mathbf{y}_t = m + \sum_{j=0}^K \Psi_j \mathbf{e}_{t-j} + \mathbf{z}_{t-K}$ for some large K such that $\mathbf{z}_{t-K}$ embodies the

“initializing” features of $\mathbf{y}_t$. The Ψ_j is obtained based on the recursion $\Psi_j = \sum_{k=1}^p \Phi_k \Psi_{j-k}$ with $\Psi_0 = I$ and $\Psi_j = 0$ for $j < 0$. See p. 103 of Box, Jenkins, and Reinsel (1994).

smallest SPCC, while the others were 70 to 90 percent of the next largest SPCC, except for the third smallest, which was about 47 percent of the fourth smallest SPCC. Therefore, we tentatively identified the cointegrating rank of the logs of the price indices y_t as 24, and thus with one common trend⁸.

Table 2
Squared partial canonical correlations, $\hat{l}_i$ from Model (1)

0.790	0.715	0.694	0.641	0.610
0.547	0.531	0.471	0.430	0.407
0.386	0.325	0.312	0.276	0.267
0.236	0.224	0.192	0.170	0.158
0.117	0.104	0.049	0.035	0.005

Based on the estimates of the model in (1) with the cointegrating rank of 24, we obtained the common trend:

$$\begin{aligned} \hat{b}'_1 y_t = & 0.0374y_{1t} + 0.0396y_{2t} + 0.0382y_{3t} + 0.0405y_{4t} + 0.0427y_{5t} + 0.0473y_{6t} + \\ & 0.0385y_{7t} + 0.0408y_{8t} + 0.0389y_{9t} + 0.0435y_{10t} + 0.0399y_{11t} + 0.0389y_{12t} + \\ & 0.0361y_{13t} + 0.0382y_{14t} + 0.0364y_{15t} + 0.0381y_{16t} + 0.0418y_{17t} + 0.0374y_{18t} + \\ & 0.0383y_{19t} + 0.0362y_{20t} + 0.0415y_{21t} + 0.0393y_{22t} + 0.0413y_{23t} + 0.0446y_{24t} + \\ & 0.0447y_{25t} \end{aligned}$$

where y_{it} is the log of the price index of the i -th city listed in Appendix Table 1. This estimated common trend displayed in Figure 1 is a weighted average of log of price indices of the 25 cities. These weights are closely tied to the sizes of the cities: Cities with a weight more than the average of 0.04 are in general larger cities, such as Ahmedabad (y_{4t}), Bombay (y_{10t}), Madras (y_{17t}), Calcutta (y_{23t}), and Delhi (y_{25t}) with a population of over three million. This common trend closely traces the log of CPI of India, as shown in Figure 2. The estimated generalized

⁸ With a cointegrating rank of 23, one may obtain numerically a second “common trend” that is orthogonal to the common trend obtained based on a cointegrating rank of 24. But this second one does not have a unit root, and thus we conclude that the y_t has one common trend and is of cointegrating rank 24.

least squares regression model of the log of the CPI on the common trend is $LCPI_t = -1.051 + 0.987CT_t$ with R-square of almost one, where $LCPI_t$ is the log of the CPI and CT_t is the common trend. Therefore, the common trend is interpreted as the inflation factor. Cecchetti et al. (2002) and Goldberg and Verboven (2004) estimate a common trend by the simple cross sectional averages, while we have used weighted (cross sectional) averages.

Based on the orthogonal projection, we have the following decomposition:

$$\mathbf{y}_t = \hat{\mathbf{b}}(\hat{\mathbf{b}}'\hat{\mathbf{b}})^{-1}\hat{\mathbf{b}}'\mathbf{y}_t + \hat{\mathbf{b}}_{\perp}(\hat{\mathbf{b}}'_{\perp}\hat{\mathbf{b}}_{\perp})^{-1}\hat{\mathbf{b}}'_{\perp}\mathbf{y}_t,$$

where the first term on the right side is the stationary (transitory) component and the second term represents the nonstationary (permanent) component. The impulse response function of the stationary component can be obtained by $\hat{\mathbf{b}}(\hat{\mathbf{b}}'\hat{\mathbf{b}})^{-1}\hat{\mathbf{b}}'\hat{\Psi}_j$, where $\hat{\Psi}_j$ is the estimated impulse response from the model in equation (1). From this we estimated the half-life of the convergence to the common trend in response to the shock of its own log of the price index of a city by examining the diagonal elements of $\hat{\mathbf{b}}(\hat{\mathbf{b}}'\hat{\mathbf{b}})^{-1}\hat{\mathbf{b}}'\hat{\Psi}_j$ and to the shock of other cities by examining the off-diagonal elements. The latter type of half-life was not examined in the panel autoregressive models because the cross sectional dynamics were not considered in those analyses.

Figure 1. Estimated Common Trend of the Log of the Price Indices of Twenty-five Cities

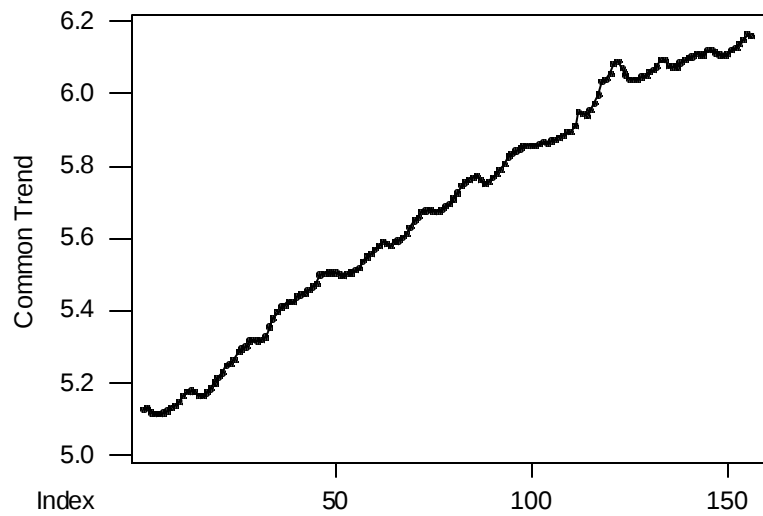
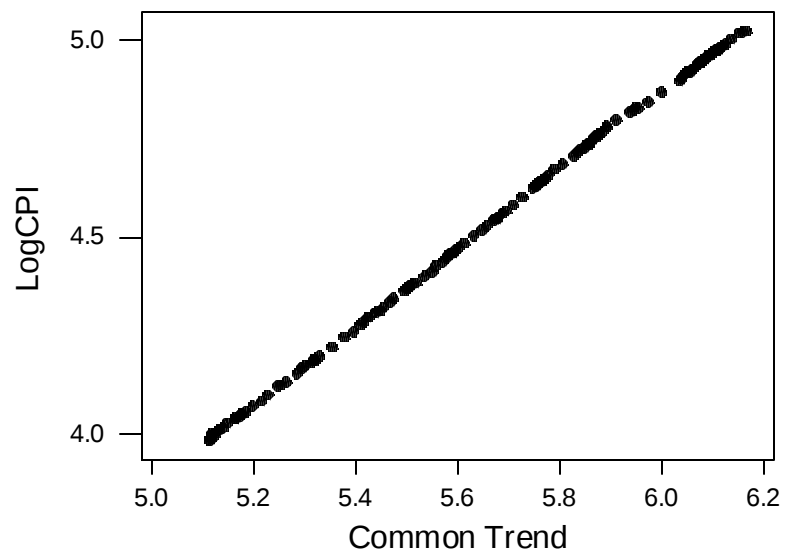


Figure 2. Scatter Plot Between the Log of the CPI and the Estimated Common Trend



IV. Rates of Convergence and Half-life

First, we estimated half-life using the panel autoregressive methods used in Cecchetti et al. (2002) for the Indian cities. The results are shown in Table 3.

Table 3
Half-life for Indian City Prices: Panel Estimates

Two-Four Months	Four-Six Months	Six-Eight Months	Eight-Ten Months	More than Ten Months
Srinagar	Ahmedabad	Hyderabad	Guntur	Bangalore
Bombay	Indore	Bhavnagar	Jamshedpur	Bhopal
Nagpur	Amritsar	Jaipur	Ajmer	Saharanpur
Sholapur	Madras	Coimbatore	Varanasi	Asansol
Madurai	Kanpur		Howrah	
Calcutta	Delhi			

We found that for 12 out of 25 cities, the half-life estimates were less than six months. Only four cities yielded more than 10 months half-life. It is interesting to compare these results with that of U.S. cities (Cecchetti et al., 2002) and Mexican cities (Sonora, 2005). U.S. cities, considered as a collection of developed economies, yielded an estimated half-life of about nine years, while Mexican cities, considered as a collection of middle-income countries, yielded estimates of half-life of about 2.5 years. Indian cities, considered as a collection of less developed countries, yielded a much shorter half-life of less than a year. Interestingly, the same technique was applied in these three studies. This implies that the rate of price convergence declines as economies become more industrialized. Cheung and Lai (2000) observed a similar relationship between real exchange rate behavior in different countries and their level of development.

We, however, believe that the appropriate way to calculate half-life is to examine the impulse response function. Choi et al. (2005) noted that for a non-monotonic impulse response function, one might observe multiple half-lives. Still, we adopted the following rule to calculate half-life. The period in which the marginal change in the stationary component of the impulse

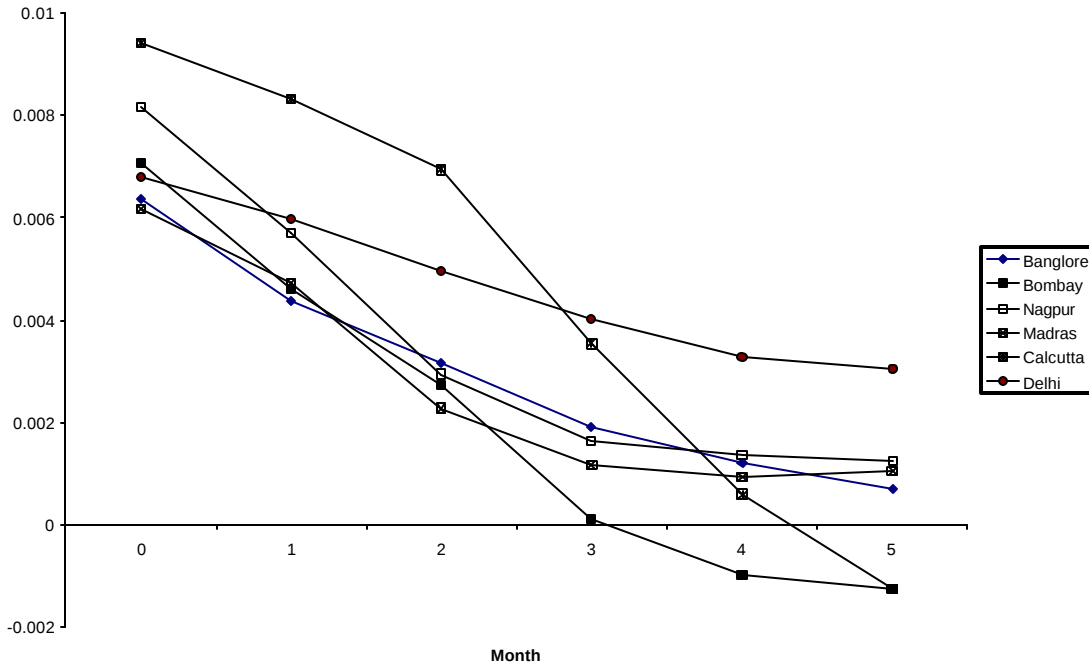
response becomes half of the initial response (in case of non-monotonic impulse response function, the first time the marginal change becomes half of the initial response is selected) is our definition of a half-life. The half-lives for shocks in the same city are reported in Table 4.

Table 4
Half-life Estimates from Impulse Response Functions: Stationary Components (Own Shock)

One Month	Two Months	Three Months	Four Months
Ajmer	Hyderabad, Jamshedpur Ahmedabad, Bangalore Bhopal, Indore Bombay, Nagpur Amritsar, Jaipur Coimbatore, Madras Madurai, Kanpur Saharanpur, Asansol Howrah	Guntur Bhavnagar Srinagar Sholapur Varanasi Calcutta	Delhi

The results shown in Table 4 indicate that the half-life for own price shocks in the Indian cities are much smaller compared to that observed in the literature. Indian capital city Delhi yielded the highest half-life estimate (four months), while Ajmer yielded the lowest half-life estimate (one month). Out of 25 cities, 17 cities yielded a half-life of two months, while six cities yielded a half-life of three months. Thus, our cointegration technique and impulse response functions yielded much faster convergence rate than that observed when panel unit root test was applied. Rangakulnuwat (2005) observed the same phenomenon when the half-lives were estimated by cointegration technique using the same U.S. data that were used in Cecchetti et al. (2002). Figure 3 shows the impulse responses from a shock in the city of six selected cities.

Figure 3: Impulse Response of Selected Six Cities (Own Shock)



We observed that for a number of cities, for example, Nagpur, even though the half-life was estimated to be two months, even after five months the marginal effect of the shock remained around 25% of the initial response, while for Bombay the marginal effect became negative after three months, although the half-life estimate was also two months.

If half-life is defined by how long it would take from the marginal change of both stationary and stochastic trend components together to be lower than the half-way mark of the initial response of both stationary and stochastic trend components together, we can observe a slightly higher average half-life. This is attributable to the slow adjustment of a non-stationary component to a shock. Although we believe that the proper definition for a half-life should be the marginal change related to the stationary component, we report the half-life calculation for total price change in Table 5.

Table 5
Half-life Estimates from Impulse Response Functions: Both Stationary and Trend Components
(Own Shock)

One Month	Two Months	Three Months	Four Months	Five Months and Above
Ajmer	Bhopal	Srinagar	Guntur	Hyderabad
Amritsar	Nagpur	Sholapur	Ahmedabad	Jamshedpur
	Coimbatore	Calcutta	Jaipur	Indore
	Madras		Madurai	Bombay
	Kanpur		Delhi	Varanasi
	Saharanpur			Bangalore
	Asansol			Bhavnagar
	Howrah			

Although we found that the range of half-life was much larger for these estimates, 20 out of 25 cities yielded four month or lower half-life estimates. This implies that Indian cities are substantially integrated. This, in essence, validates the PPP doctrine.

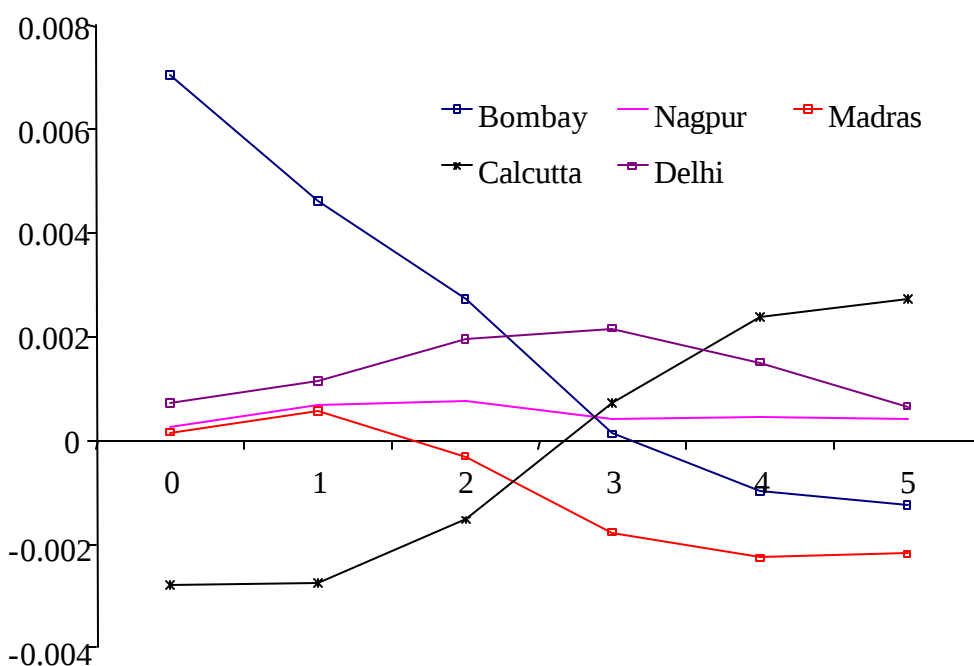
We calculated half-life not only for a price shock originated in the city under consideration, but also for a price shock initially imposed on some other cities. For brevity, the results from only five cities are reported in Table 6. These are the four largest cities in the “four corners” of India; Nagpur was chosen as a city at the geographic center (distance from Nagpur to other cities are reported in Table-6). Diagonal elements are half-life for own shock, while off-diagonal entries are half-life for shock emanating from the cities shown in the first row.

Table 6
Half-life Estimates (Months) from Impulse Response Functions for Five Main Cities
(Own Shock and Shocks in Other Cities)

	Bombay	Calcutta	Delhi	Madras	Nagpur	Distance From Nagpur (Miles)
Bombay	2	4	4	5	2	439
Calcutta	3	3	2	1	3	593
Delhi	>5	1	4	>5	>5	531
Madras	2	2	3	2	>5	561
Nagpur	>5	5	>5	2	2	

We found that effects of a shock in one city have differential impact on other cities. For example, a unit shock in the price of Bombay would have a half-life of two months, while it would have a half-life of more than five months for Nagpur and Delhi, while the half-lives would be two and three months for Madras and Calcutta, respectively. We observed similar variations in cross-city transmission of price shocks. It is also interesting to note that the origin of the shock affects same city pairs. The half-life for a shock originated in Bombay on prices in Madras was two months, while the half-life for a shock originated in Madras on prices in Bombay was found to be five months. Figure 4 shows the impulse responses for a shock in Bombay on itself and four other large cities.

Figure 4: Impulse Responses of Shock on Bombay



We observed no specific pattern of this shock on other cities. For example, Calcutta (the farthest city from Bombay) prices yielded a negative initial impulse from a positive shock in the prices in Bombay, while three other cities registered positive impulse. For Nagpur, which is

much closer to Bombay, the positive impulse was small and it remained similar for all five months shown in the diagram. Although we did not calculate the confidence intervals for half-lives presented here, we may argue from our results that the role of distance in convergence to prices should be analyzed with much care.

V. Conclusions

The main focus of this paper is to examine the nature of price fluctuations in Indian cities. We followed cointegration technique to identify a common trend. Interestingly, the common trend turns out to be closely related to the overall CPI of India. The rate of convergence for a unit shock imposed on a city was calculated by using impulse response functions, which we believe to be a precise way of calculating the half-life. Moreover, our approach is flexible enough that we can calculate the half-life of a shock in one city on the shock imposed on another city. We found that the rate of convergence to the stochastic trend is much faster, with an average half-life of around three months. Although Murray and Papell (2002) and Goldberg and Verboven (2004) reported the half-life of convergence to be in the international context less than one year for a few real exchange rates, and Cheung and Lai (2000) showed that the rate of convergence is faster for developing countries than for developed countries, our results from Indian city data provide a strong support to the PPP theory.

Contributions of this paper to the literature are two-fold. We used cointegration analysis to derive the common trend. Generally, researchers use an average for all locations as the common trend without any effort to ascertain the true common trend. Not only is there a possibility of having more than one common trend for any panel data, but also, even when there is only one common trend, the mean of the price indices may not be the true one. We would, therefore, argue that the common trend should be identified and calculated for an exercise

dealing with the convergence to PPP. In this paper a unique common trend is identified and calculated. We then examine the rates of convergence of the city CPIs to this common trend.

It is also very important to define half-life more rigorously. For example, when one standard deviation shock in prices in a city is added, we decomposed this shock into two components: stochastic trend and stationary component. We calculated the immediate change in the stationary component. How long it takes the marginal change in the stationary component to the half-way mark of the initial response is our half-life estimate. These measures are calculated with better precision by using impulse response functions. This method of calculating a common trend and the use of impulse response function to calculate the half-life of the stationary component is invariant with the order of autoregressive structure. We can get impulse response function not only for the shock to the city's own CPI, but also for shocks emanating from other cities⁹ due to the cointegration technique. This will certainly improve our policy making related to price stability.

In terms of cross-city shock transmissions, we found that the shock in a particular city would have differential effects on different cities. For example, the nature of shock transmission from Bombay to Nagpur seems different from the shock transmission from Nagpur to Bombay. This suggests that the use of city distance to calculate the transport costs in the case of price formation in different cities warrants caution.

⁹ Baskar and Hernandez-Murillo (2003) is an exception; they try to calculate effects of a shock in one city on the prices in a different city.

Appendix

Appendix Table 1
States and Cities in India Covered in the Present Study

State and Federal Territory	Cities
Andhra	Guntur and Hyderabad
Bihar	Jamshedpur
Gujarat	Ahmedabad and Bhavnagar
Jammu and Kashmir	Srinagar
Karnataka	Bangalore
Madhya Pradesh	Bhopal and Indore
Maharashtra	Bombay (Mumbai), Nagpur, and Sholapur
Punjab	Amritsar
Rajasthan	Ajmer and Jaipur
Tamil Nadu	Coimbatore, Madras (Chennai), and Madurai
Uttar Pradesh	Kanpur, Saharanpur, and Varanasi
West Bengal	Asansol, Calcutta (Kolkata), and Hawrah
Delhi	Delhi

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