Do relative prices of nontraded goods explain relative movements in price indices across U.S. cities?*

August 2006

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Abstract: Inclusion of nontraded goods prices has been shown to be partly responsible for slow convergence in price indices across major cities in the United States (Cecchetti et al 2002). Drawing upon insights provided by well known theories, this paper uses panel econometric methods to seek evidence in U.S. city level Consumer Price Index (CPI) data of a systematic long-run relationship between prices of nontraded goods relative to traded goods, and overall relative price. While the results for convergence of overall relative prices are mixed, the hypothesis of a long-run relationship between overall relative prices and relative prices of nontraded goods is overwhelmingly rejected. However, the evidence of cointegration between relative prices of nontraded goods across cities suggests that relative movements of overall prices could be driven by temporary shocks that move nontraded goods prices away from their long-run paths.

Keywords: Traded goods; Nontraded goods; Panel unit root; Panel cointegration

JEL classifications: E3, R1

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* The authors are grateful to Don Freeman for many useful suggestions and insightful discussions.
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1. Introduction

The existence of a dichotomy between traded and nontraded goods has been shown to be a theoretically plausible and empirically valid explanation for the breakdown of the purchasing power parity (PPP) hypothesis in the cross-country context.[1] For the United States, inclusion of nontraded goods prices has been found to be partly responsible for slower convergence in price indices across major cities (Cecchetti et al 2002). But do prices of nontraded goods relative to traded goods systematically affect the relative movements of overall price indices across cities in the long-run? To address this question, we first investigate the univariate dynamic properties of overall relative prices and relative prices of nontraded goods for 20 major U.S. cities over a period from 1968 to 2005.[2] Furthermore, we conduct cointegration analyses to examine the long-run dynamic relationship between these two relative prices.

The PPP hypothesis is based on *the law of one price* which suggests that prices (expressed in a single currency unit) of identical goods across countries (more generally, across different locations) converge to a single price in the long-run. PPP is the extension of this law to overall prices and it implies that real exchange rate – relative price of a country's goods and services in terms of another country's goods and services - is stationary. Since the empirical literature largely rejects the hypothesis of long-run PPP, the recent studies have focused on the explanation of a non-stationary real exchange rate. Of several propositions, the one that relates inclusion of nontraded goods prices to

[1] For example, DeLoach (1997), Kakkar and Ogaki (1999) and Kakkar (2003) find evidence of unit root in relative price of nontraded goods across countries – which is interpreted as providing support for this line of explanation. The terms 'traded' and 'nontraded goods' are used to include not only tangible goods (or commodities) but also intangible services. We follow this convention throughout the paper.

[2] As will be defined below, 'overall relative price' refers to CPI in one city relative to average CPI in other cities in our sample, and 'relative price of nontraded goods' refers to CPI of nontraded goods relative to CPI of traded goods in a city.

persistent differences in overall prices across countries has generated some interest among researchers. In particular, the evidence of a stochastic trend and unit root nonstationarity of the relative price of nontraded goods has been seen as a source of permanent change in the real exchange rate. This line of research has received theoretical endorsement from the influential works of Balassa (1964) and Samuelson (1964) who emphasize that different trends in technologies in the traded and nontraded goods infuse a permanent component in relative price of nontraded goods. Recently, Kakkar (2003) has shown that the relative price of nontraded goods and sectoral total factor productivity are cointegrated across OECD countries, thereby providing empirical evidence in support of the insights provided by Balassa-Samuelson model.

This explanation of observed long-run behavior of real exchange rate across countries may be useful in understanding persistent differences in relative prices across U.S. cities. In a recent study using Consumer Price Index (CPI) data for 19 major U.S. cities from 1918 to 1995, Cecchetti et al (2002) find slow convergence in overall relative prices across cities. They estimate a half life of nearly 9 years. They ascribe slow convergence to a combination of factors namely, transportation costs, differential speeds of adjustment to small and large shocks, and the inclusion of nontraded goods in the calculation of overall price index. There have been a few other studies that use U.S. city level data to examine convergence in prices. Culver and Papell (1999) and Chen and Devereux (2003) find evidence of slow convergence. In contrast, Parsley and Wei (1996), in a related paper, find that prices converge to purchasing power parity at a much faster rate,using quarterly data on 51 final goods and services prices - in contrast to overall CPI in Cecchetti et al (2002) - across 48 U.S. cities over the period from 1975 through 1992.

None of these other studies, however, explores the effect of the dichotomy between traded and nontraded goods on the relative movements of city level prices.

The objective of this paper is to investigate if there is any evidence of a long-run relationship between overall relative price and relative prices of nontraded goods among major U.S. cities. We use annual CPI data – all items, services and commodities – for 20 cities for a period from 1968 through 2005. We first investigate univariate dynamic properties of overall relative price and of relative prices of nontraded goods by conducting panel unit root tests. We find mixed evidence of unit root stationarity for overall relative prices whereas there is strong evidence of unit root for relative prices of nontraded goods. We use multivariate panel cointegration test procedures to examine the long-run structural relationship suggested by theory. We find little evidence of cointegration. Furthermore, the evidence of a cointegrating relationship between relative prices of nontraded goods across cities suggests that relative movements of overall prices could be driven by temporary shocks that move nontraded goods prices away from their long-run paths.

This paper makes two contributions to the literature. First, drawing upon the PPP hypothesis and the Balassa-Samuelson model it explicitly derives a structural relation between overall relative price and relative prices of nontraded items. This structural relation is then subjected to empirical testing for cointegration. Second, by providing empirical evidence of a cointegrating relationship between relative prices of nontraded goods across cities, this study suggests an alternative interpretation of overall relative prices as representing transitory deviations of nontraded goods prices from their long-run paths.

The rest of the paper is organized as follows. In section 2, we discuss the theoretical background for our empirical investigation. In particular, we derive an empirically testable structural relation. Section 3 discusses empirical methodology. In Section 4, we discuss data and present our empirical results. The next section includes our concluding remarks.

2. Theoretical Background

Real exchange rate is defined as:

$$R_t = \frac{E_t \times P_t}{P_t^*} \tag{1}$$

where R_t is the real exchange rate (price of domestic goods in terms of foreign goods), E_t is the nominal exchange rate (number of foreign currency per unit of domestic currency), P_t is the domestic price level and P_t^* is the foreign price level. t indexes time period. Denoting natural logarithms of the variables by lower cases, we can write this equation as follows:

$$r_t = e_t + p_t - p_t^* \tag{2}$$

The law of one price implies that the price of a good in the foreign country is equal to the price of the good in the domestic country when denominated in the foreign country's currency. If this law holds for *all* goods, then PPP holds across countries. Empirical analysis of the long-run PPP hypothesis typically involves testing for stationarity of the real exchange rate. Any evidence of nonstationarity is interpreted as a verdict against the hypothesis. In the context of a country, r can be interpreted as the overall price level in one city (or region) relative to the overall price level in another city (region).

Furthermore, the nominal exchange rate between cities is unity; that is, $E_t \equiv 1$, and therefore $e_t = 0$ and equation (2) becomes

$$r_t = p_t - p_t^* \tag{3}$$

where p_t and p_t* are now logarithms of price levels in two different cities. For convenience of exposition, we will call them 'home city' and 'foreign city' respectively in next few paragraphs.

The evidence against long-run PPP has given rise to a substantial literature that has suggested several explanations for incomplete adjustments of prices across countries. One such explanation credited to Balassa (1964) and Samuelson (1964), stresses the importance of the dichotomy between traded and nontraded goods for de-linking the law of one price and PPP. If some goods are not tradable, arbitrage condition will not be satisfied and the law of one price will not hold. This explanation has received some empirical support in the study of convergence of price indices across U.S. cities as well (Cecchetti at al 2002). One of the implications of this argument is that long-run PPP will hold only for traded goods. Thus the PPP for traded goods between home and foreign city can be represented as follows:

$$\varepsilon_t = p_{T,t} - p_{T,t}^*, \text{ with } E(\varepsilon_t) = 0 \tag{4}$$

Combining (3) and (4),

$$r_t = (p_t - p_t^*) - (p_{T,t} - p_{T,t}^*) + \varepsilon_t \tag{5}$$

We assume that the overall price level is a weighted average of the traded and non-traded goods prices:

$$p_t = \gamma p_{T,t} + (1-\gamma) p_{N,t} \tag{6}$$

and $$p_t^* = \gamma^* p_{T,t}^* + (1-\gamma^*) p_{N,t}^* \tag{7}$$

where γ and γ^* denote weights assigned to traded goods prices in home and foreign city. Subscript N denotes nontraded goods and T denotes traded goods. Note that since the prices are in logarithms, this assumption implies that the general price level is a geometric mean of traded and nontraded goods prices. Furthermore, both traded and nontraded goods categories can be composed of goods (commodities) as well as services. Thus the prices of traded and nontraded goods can be written as weighted averages of prices of goods and services:

$$p_{T,t} = ap_{G,t} + (1-a)p_{S,t} \tag{6}$$

$$p_{N,t} = bp_{G,t} + (1-b)p_{S,t} \tag{7}$$

where a and b respectively represent weights assigned to prices of goods and services in home city. Subscript G denotes 'goods' and S denotes 'services'. Using analogous expressions for traded and nontraded goods with weights a^* and b^* for goods and services in foreign city, substituting into the expression for r in equation (5), and rearranging terms we obtain

$$r_t = \theta\left(p_{S,t} - p_{G,t}\right) + \phi\left(p^*_{S,t} - p^*_{G,t}\right) + \varepsilon_t \tag{8}$$

where $\theta = (1-\gamma) \times (a-b)$

$$\phi = -\left(1-\gamma^*\right) \times \left(a^* - b^*\right)$$

To extend this formulation to more than two cities we now use slightly different notations. Using i and j to index U.S. cities, we can rewrite equation (8) as:

$$r_t^{i,j} = \theta^i\left(p^i_{S,t} - p^i_{G,t}\right) + \phi^j\left(p^j_{S,t} - p^j_{G,t}\right) + \varepsilon_t^{i,j} \tag{9}$$

where $r_t^{i,j}$ represents the price level in city *i* relative to price level in city *j*.[3] Summing over *j*, and dividing by *n* - 1 where $j \neq i$ and *n* is the number of cities in the U.S., we have the following relationship:

$$r_t^i = \theta^i \left(p_{S,t}^i - p_{G,t}^i\right) + \overline{\phi}\left(\overline{p}_{S,t} - \overline{p}_{G,t}\right) + \varepsilon_t^i \tag{10}$$

where $$\overline{p}_{S,t} = \frac{1}{n-1}\sum_{j\neq 1} p_{S,t}^j \,; \qquad \overline{p}_{G,t} = \frac{1}{n-1}\sum_{j\neq 1} p_{G,t}^j$$

and $$\overline{\phi} = \frac{1}{m-1}\sum_{j\neq i}\phi^j \qquad \text{for } i, j = 1, 2, \ldots, n$$

The first term on the right hand side of equation (10) is the relative price of services in terms of goods in city *i* and the second term is the average relative price of services in all other cities. Note that the dichotomy between traded and nontraded items – which is purported to explain the breakdown of the PPP hypothesis – essentially boils down to a dichotomy between goods and services. The fact that data are readily available for prices of goods and services separately and not for prices of traded and nontraded items, makes equation (10) an empirically testable relationship. For comparability with the existing

[3] When price index such as CPI is used to represent general price level, $r^{i,j}{}_{,t}$ actually represents changes in relative prices in city *i* vis-à-vis city *j* between period *t* and the base year used for calculation of the price index. Thus,

$$R_t^{i,j} = \frac{CPI_t^i}{CPI_t^j}$$

But $$CPI_t^i = \frac{P_t^i}{P^i{}_{base\ \mathbf{year}}} \quad \text{and} \quad CPI_t^j = \frac{P_t^j}{P^j{}_{base\ \mathbf{year}}}$$

where P^i_t is the price or cost of a fixed basket of goods and services in city *i* in period *t*. Substituting for *CPI*s, and rearranging, we obtain

$$R_t^{i,j} = \frac{P_t^i}{P_t^j} \bigg/ \frac{P^i{}_{base\ \mathbf{year}}}{P^j{}_{base\ \mathbf{year}}}$$

This implies

$$r_t^{i,j} = ln\left(\frac{P_t^i}{P_t^j}\right) - ln\left(\frac{P^i{}_{base\ year}}{P^j{}_{base\ year}}\right)$$

literature we continue using the terms 'nontraded items' and 'traded items' to refer to 'services' and 'goods' respectively.

Equation (10) implies that in the long-run overall relative price in a city depends not only on relative price of nontraded goods in that city but also on relative prices of nontraded goods in other cities. Thus a permanent shock to the relative prices of nontraded goods will have persistent effects on the relative movements of the overall price level in that city.

3. EMPIRICAL METHODOLOY

We summarize the implications of our discussion from the previous section in the form of an estimable cointegrating regression equation:

$$r_t^i = \alpha^i + \beta_1^i r_{N,t}^i + \beta_2^i r_{N,t}^{A-} + \varepsilon_t^i \qquad (11)$$

where r_t^i is as before; α^i is an intercept term; $r_{N,t}^i = p_{S,t}^i - p_{G,t}^i$; $r_{N,t}^{A-} = \overline{p}_{S,t} - \overline{p}_{G,t}$ and ε_t^i is a zero-mean stationary random variable that captures short-run deviation of relative price in city *i* from its long run equilibrium value.

Inclusion of the intercept term does not automatically follow from our derivation of equation (10). Note that r_t^i is the 'relative price level in city *i* in year *t*' *relative to* 'relative price level in the base year'.[4] Similarly, $r_{N,t}^i$ ($r_{N,t}^{A-}$) is the 'relative price of nontraded goods in city *i* (all cities other than *i*) in year *t*' *relative to* 'relative price of nontraded goods in the base year'. Thus by construction the variables in equation (10) control for the effects of time invariant base year conditions which could be a source of

[4] See footnote 3 for details.

heterogeneity across cities. Therefore, α^i simply represents fixed effects not associated with base year relative prices. These fixed effects could be city specific demand and supply conditions such as geographic location and climate, or demographic composition. Inclusion of the intercept term gives us more flexibility without being inconsistent with the underlying theory.[5]

Since the sample period for our data set is relatively short we use panel data techniques for the econometric analysis of the theoretical relation derived in the previous section. By pooling data in time and cross-section dimensions we gain in efficiency while estimating the models, and in power while conducting the tests. Furthermore, as Kao (1999) and Phillips and Moon (1999) point out, using panel data we can avoid the problem of spurious regression.

3.1 Panel unit roots tests

Before we conduct the tests for cointegration, we have to determine the order of integration for each series to be used in the estimation of equation (11). We thus perform the panel unit root tests. There are several test procedures suggested in the literature. While each test procedure has its strengths and weaknesses, researchers commonly use the methods proposed by Levin and Lin (1993) and Im, Pesaran and Shin (1995). Following suggestions from Pedroni (1999), we implement three versions of Levin and Lin (*LL*) test procedures. We would call these test statistics *LL-rho*, *LL-trho* and *LL-ADF* respectively. The Im, Pesaran and Shin (*IPS*) procedure is simply a generalization of the

[5] In our empirical analysis, we conduct two separate F-tests: (1) one for the exclusion of intercept term, and (ii) the other for heterogeneity of the intercept term. These tests are based on panel estimation of fixed effects model and panel common intercept model. The test results reject not only the null of no intercept but also the null of homogeneous intercept term.

ADF test applied to pooled data. Both *LL* and *IPS* procedures are based on the following test equation:

$$\Delta y_{i,t} = \alpha_{m,i} d_{m,t} + \delta_i y_{i,t-1} + \sum_{l=1}^{q_i} \beta_{i,l} \Delta y_{i,t-l} + \nu_{i,t} \text{ ; } m = 1, 2, 3 \tag{12}$$

where $y_{i,t}$ is a variable of interest in city i from the set: $\{ r_t^i,\ r_{N,t}^i,\ r_{N,t}^{A-} \}$; $d_{m,t}$ is the vector of deterministic variables and $\alpha_{m,i}$ the corresponding vector of coefficients; $\beta_{i,l}$s are coefficients, and q_i is the appropriate lag length of the augmented terms $\Delta y_{i,t-l}$; $\nu_{i,t}$ is a white noise error term. Although, in general, there could be three different specifications (as indicated by three possible values of m) of vectors $d_{m,t}$ and $\alpha_{m,i}$ (which simply imply test equations with (i) no intercept-no deterministic time trend, (ii) intercept *but* no trend and (iii) both intercept *and* trend), for reasons discussed above, we use the specification with $m = 2$ with $d_{2,t} = \{1\}$. That is, we are including a heterogeneous intercept term in the test equation (12).

The null hypothesis of the test for unit root is: H_0: $\delta_i = 0$. These two test procedures differ in their specifications of the alternative hypotheses. For the *LL* test the alternative hypothesis is H_1: $\delta_i = \delta < 0$ while for the *IPS* test, H_1: $\delta_i < 0$ for some i and $\delta_i = 0$ for rest of i's. Rejection of the null hypothesis implies that the series is unit root stationary while a failure to reject suggests that the series is a unit root process.

3.2 Cointegration tests

The univariate analysis sets the stage for our cointegration analysis. For two or more series to be cointegrated, they have to be of same order of integration. Equation (11) defines the cointegrating relationship. We use panel cointegration test methods suggested by Pedroni (1999). As noted before, pooling of time series and cross-section dimensions help reduce the problem of low power in single-equation tests. These test procedures are based on residuals derived from the hypothesized cointegrating regression represented by equation (11). The tests involve the parameter μ^i in the following hypothesized data generating process:

$$e_t^i = \mu^i e_{t-1}^i + \zeta_t \tag{13}$$

where e^i_t is the estimated ε^i_t of equation (11). Of the seven statistics suggested by Pedroni (1999), four are based on pooling along the within-dimension, and three are based on pooling along the between-dimension.[6] For the within-dimension statistics the test for the null of no cointegration is implemented as a residual-based test of the null hypothesis H_0: $\mu^i = 1$ for all i, *versus* the alternative hypothesis H_1: $\mu^i = \mu < 1$ for all i, so that it presumes a common value for μ^i. On the contrary, for the between-dimension statistics the null of no cointegration is implemented as a residual-based test of the null hypothesis H_0: $\mu^i = 1$ for all *i, versus* the alternative hypothesis H_1: $\mu^i < 1$ for all i, so that it does not presume a common value for μ^i under the alternative hypothesis. Note that the null hypothesis in each case assumes that there is no cointegration. According to Pedroni (1999), under the null hypothesis the test statistics are asymptotically normal. Of the seven tests, the first one is a right tailed test while the others are left-tailed tests.

4. Data And Empirical Results

4.1 Data and variables

The source of city level Consumer Price Index (CPI) data is the U.S. Bureau of Labor Statistics (BLS). We obtain annual data on CPI-all items, CPI-commodities and CPI-

[6] For details on these statistics, see Pedroni (1999).

services for 20 major U.S. cities for a period from 1967 to 2005.[7] The choice of the sample period is guided by the fact that CPI data for commodities and services are separately available since 1967 only.

Using these price indices, we construct data on the following variables. Let $P_{,t}^{i}$ be the consumer price index of all items in city i in year t. Let n be the number of cities in the panel. Then the relative price, r_t^i in city i in year t is defined as

$$r_t^i = 100 \times \left(\ln P_t^i - \frac{1}{n-1} \sum_{j \neq i} \ln P_t^j \right) \tag{14}$$

Relative price of nontraded goods in city i is calculated as follows:

$$r_{N,t}^i = 100 \times \left(\ln P_{S,t}^i - \ln P_{G,t}^i \right) \tag{15}$$

where $P_{S,t}^i$ is the CPI –services and $P_{G,t}^i$ is the CPI-commodities. Finally, average relative price of nontraded goods in other cities is calculated as follows:

$$r_{N,t}^{A-} = 100 \times \frac{1}{n-1} \sum_{j \neq i} \left(\ln P_{S,t}^j - \ln P_{G,t}^j \right) \tag{16}$$

Table 1 presents summary statistics of these constructed series. Note that, by construction, the mean of overall relative price is 0, and the means of relative prices of nontraded items and of their average relative prices are identical. As we can see, there are considerable variations in all three series across the U.S. cities.

Figure 1 illustrates how differences in overall relative prices and relative prices of nontraded goods have evolved over time. As we can see from the figure, the dispersion

[7] The cities in our sample are: Atlanta, Boston, Chicago, Cincinnati, Cleveland, Dallas, Detroit, Honolulu, Houston, Kansas City, Los Angeles, Milwaukee, Minneapolis, New York, Philadelphia, Pittsburg, San Diego, San Francisco, Seattle, St. Louis.

around mean has increased for all three series over time. The year-to-year movements in their dispersions have very similar patterns which may be indicative of some systematic relationship.

4.2 Unit root test results

The panel unit root test results are presented in Table 2. As shown by Pedroni (2004), under the null hypothesis, these test statistics have asymptotic Normal distributions. Furthermore, because of the specification of the alternative hypotheses as discussed in the previous section, they are left-tailed tests. Comparing the estimated test statistics with the critical values, we find that all three versions of *LL* test fail to reject the null of unit root for overall relative price, r_t^i, at conventional levels of significance while the *IPS* test rejects the unit root null for the series at 1 percent level. The result of the *IPS* test is consistent with the results of Cecchetti et al (2002) for comparable sample period. However, our *LL* test results are very different from theirs. For other two series, $r_{N,t}^i$ and $r_{N,t}^{A-}$, both *LL* and *IPS* tests unequivocally indicate that they are unit root processes.

The unit root test results reported above provide mixed evidence of convergence in overall relative prices across cities. Absence of unit root is traditionally interpreted as evidence in support of convergence. Thus the *IPS* test result suggests that overall relative prices converge across cities while the *LL* test results do not find any evidence of convergence.

Not surprisingly, the *LL* model and the *IPS* model also differ in the speed of convergence which is measured by 'half-life'. It is calculated as $-ln(2)/ln(\rho)$ where $\rho_i = 1 + \delta_i$, is the persistence parameter. Since the *LL* model restricts the persistence parameter to be homogeneous across cities, (that is, $\rho_i = \rho$) we use the estimated value $\hat{\rho}$ to calculate the half-life. However, the *IPS* model allows ρ_i to be different across cities, and, therefore, we calculate the average across i. We find that the half-life is approximately 9.8 years using *LL* and 4.4 years using *IPS*. As discussed by Cecchetti et al (2002), because of the small sample size the estimated persistence parameters may be biased downward. If we make the adjustment for this bias as suggested by Nickell (1981), the half-life increased manifold to 147.2 years for the *LL* model and to 7.5 years for the *IPS* model.[8] Although the *IPS* half-life is shorter, yet comparable to those obtained by Cecchetti et al (2002), the *LL* half life is substantially higher.[9]

The evidence of nonstationarity of the relative prices of nontraded items and their average relative prices is a robust result, and very much in accordance with the implication of the Balassa-Samuelson model. The fact that these series have stochastic trend components signifies that the differences in technology may indeed infuse a permanent component in relative movement of prices of nontraded goods vis-à-vis prices of traded goods. In a cross-country context, Kakkar (2003) observed similar behavior for relative prices of nontraded goods.

4.3 Cointegration test results

Although there is some degree of ambivalence regarding the stochastic trending property of overall relative prices let us assume that the *LL* test procedure correctly identifies it so

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[8] We use the following formula suggested by Nickell (1981) to make adjustment for the bias:

$$\underset{n\to\infty}{p\,lim}(\hat{\rho}-\rho)=\left\{\frac{2\rho}{1-\rho^2}-\left(\frac{1+\rho}{T-1}\left(1-\frac{1}{T}\times\frac{(1-\rho^T)}{(1-\rho)}\right)\right)^{-1}\right\}^{-1}$$

[9] Given that n=20, the use of the adjustment factor may not be appropriate.

that we can empirically examine the long-run relationship suggested by theory and represented by equation (11), by conducting cointegration test. As discussed in the previous section, we use seven residual-based test procedures. The results are presented in Table 3. As we can see from the table, none of the tests rejects the null of no cointegration. Thus these tests overwhelmingly demonstrate that there is little evidence of a long-run relationship between overall relative prices and relative prices of nontraded goods across U.S. cities.

This result fails to support one of the implications of the Balassa-Samuelson model which suggests, if the relative prices of traded goods have a stochastic trend component (which, as we discussed in the previous subsection, is the case), any permanent shock to the relative price of nontraded goods will have a persistent effect on overall relative prices through the relationship as represented by equation (11). Thus, the dichotomy between traded and nontraded goods is perhaps not a good explanation for relative movements in price levels across U.S. cities in the long-run. This result *per se* is not in contrast with the findings of earlier studies. For example, Cecchetti et al (2002) find little evidence against convergence and only fragmentary evidence to suggest that slow convergence in overall prices is due to slower adjustment in the prices of nontraded goods.

4.4 Are relative prices of nontraded goods cointegrated across cities?

Since both relative prices of nontraded goods and their average relative prices are $I(1)$ processes, it is natural to ask if they are cointegrated. If they are, what does such a cointegrating relationship suggest? Since overall relative price can be written as a linear combination of these two, as shown in equation (10), and the *IPS* test indicates that it is an $I(0)$ process, this line of investigation is even more pertinent and intriguing.

We conduct panel cointegration test on relative prices of nontraded goods and their average relative prices. The results are reported in Table 4. Three out of seven tests strongly reject the null of no cointegration. If we accept these results as the true characterization of the relationship between relative prices of nontraded goods across U.S. cities, they indicate the presence of a common permanent component in those prices across cities. This accords well with the insights of the Balassa-Samuelson theory: any permanent shock – say, technology shock to one type of goods: traded or nontraded – will permeate through the entire market of that good in the long-run and will have a permanent impact on the relative price of nontraded items in every city.

The existence of a cointegrating relationship also implies that a linear combination of the cointegrating variables is stationary. Contrary to our assumption in Section 4.3 if we now assume that the *IPS* test result reveals the true characterization of the stochastic trending property of overall relative price, the result of this section assumes special significance. Equation (10) shows that overall relative price is a linear combination of relative prices of nontraded goods and their average relative prices. Although there could be more than one linear combination between two variables, the cointegrating combination is unique to any affine transformation. Thus our result suggests that overall relative price can just represent transitory deviations from the long-run relationship between relative prices of nontraded items across cities. Thus relative movements of overall prices are driven by shocks that move relative prices of nontraded items in a city away from their long-run paths. Of course, this is not an explanation for slow convergence of overall relative prices but certainly an interesting finding in itself that warrants further investigation which is beyond the scope of the current study.

5. Concluding Remarks

Inclusion of nontraded goods prices has been shown to be partly responsible for slow convergence in price indices across major cities in the United States (Cecchetti et al 2002). Drawing upon insights provided by well known theories, this paper uses multivariate panel econometric methods to investigate if there is any evidence in city level Consumer Price Index (CPI) data of a systematic long-run relationship between prices of nontraded goods relative to traded goods, and overall relative prices across U.S. cities. While the results for convergence of overall relative prices are mixed, the hypothesis of a long-run relationship between overall relative prices and relative prices of nontraded goods is overwhelmingly rejected. However, the evidence of cointegration between relative prices of nontraded goods across cities suggests that relative movements of overall prices could be driven by temporary shocks that move nontraded goods prices away from their long-run paths.

Table 1: Summary statistics

	Overall relative price (r_t^i)	Relative price of nontraded goods ($r_{N,t}^i$)	Relative price of nontraded goods in other cities ($r_{N,t}^{A-}$)
	(1)	(2)	(3)
Mean	0.00	31.01	31.01
Maximum	24.50	77.13	62.32
Minimum	-10.98	-2.84	1.54
Std. Dev.	4.40	20.21	19.13
Observations	760	760	760
Cross sections	20	20	20

Table 2: Panel unit root test results

Overall relative price (r_t^i)	Relative price of nontraded goods ($r_{S,t}^i$)	Relative price of nontraded goods in other cities ($r_{S,t}^{A-}$)
(1)	(2)	(3)
Levin-Lin rho-stat = -0.66	Levin-Lin rho-stat = -0.97	Levin-Lin rho-stat = -0.97
Levin-Lin t-rho-stat = 1.19	Levin-Lin t-rho-stat = 0.47	Levin-Lin t-rho-stat = 0.47
Levin-Lin ADF-stat = 0.54	Levin-Lin ADF-stat = 0.21	Levin-Lin ADF-stat = 0.22
IPS ADF-stat = -2.55***	IPS ADF-stat = -0.67	IPS ADF-stat = -0.66

Note: ***denotes significance at 1% level; ** denotes significance at 5% level; and * at 10% level.

Table 3: Panel cointegration test results: Relative overall price, relative price of nontraded goods and average relative price of nontraded goods

Hypotheses	Estimated test statistic	
(1)	(2)	
H_0: $\mu_i = 1$ H_1: $\mu_i = \mu < 1$	Panel v-stat Panel rho-stat Panel pp-stat Panel adf-stat	-0.90 0.90 0.18 -0.80
H_0: $\mu_i = 1$ H_1: $\mu_i < 1$	Group rho-stat Group pp-stat Group adf-stat	2.38 1.08 -0.61

Note: ***denotes significance at 1% level; ** denotes significance at 5% level; and * at 10% level.

Table 4: Panel cointegration test results: Relative price of nontraded goods and average relative price of nontraded goods

Hypotheses	Estimated test statistic	
(1)	(2)	
H_0: $\mu_i = 1$ H_1: $\mu_i = \mu < 1$	Panel v-stat Panel rho-stat Panel pp-stat Panel adf-stat	311.21*** -23.76*** 0.70 -0.52
H_0: $\mu_i = 1$ H_1: $\mu_i < 1$	Group rho-stat Group pp-stat Group adf-stat	-20.02*** 2.63 0.91

Note: ***denotes significance at 1% level; ** denotes significance at 5% level; and * at 10% level.

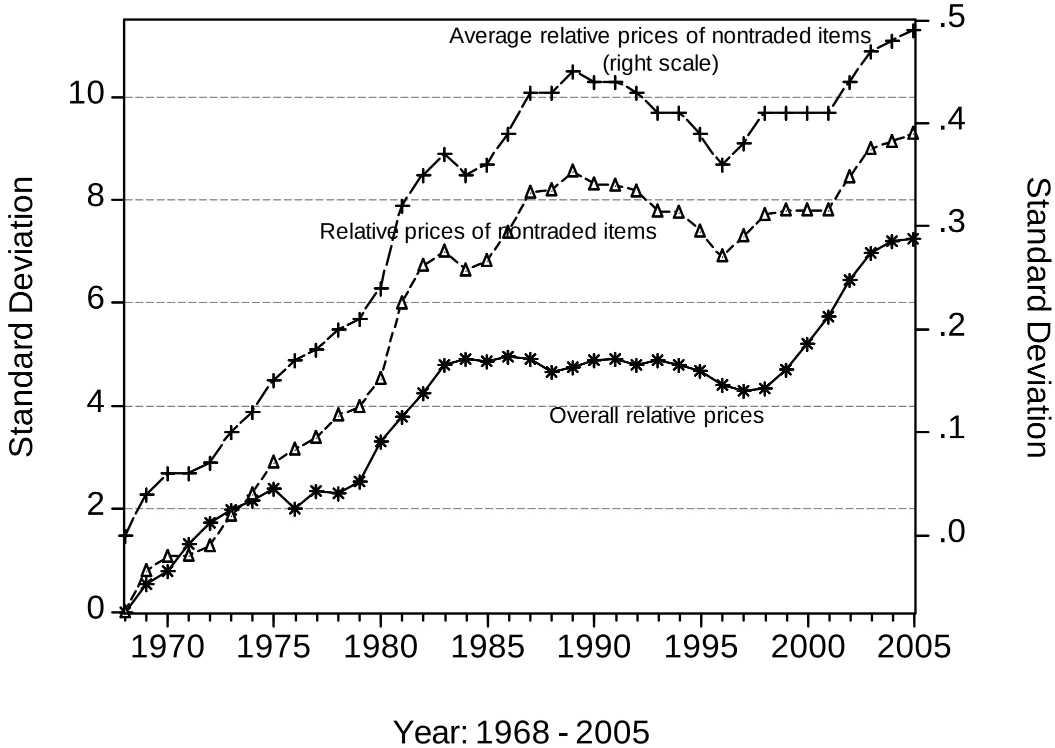

Figure 1: Standard deviation across U.S. cities

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