Price Convergence and Market Integration in China*

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Abstract

Whether China has successfully transformed from a formally centrally-controlled economy to a market economy is an issue of immense interest and considerable debate. The current study complements the received literature by testing price movements in China using panel data of monthly price indexes across 36 major Chinese cities over a thirteen-year period. By so doing, it extends the early work by Fan and Wei (2006) and thereby adds to the literature in four main aspects. First, price index data provide a wider coverage of all goods traded in the market than a group of individual commodity prices. Thus, it has more relevant implications for monetary and other macroeconomic policies. Second, as Taylor (2001) notes, there are crucial differences of methodologies in the study of price convergence between individual commodity data and index data. Thus, the present study not only further examines market integration in China, but also provides a comparative study of price index movements in a major transitional economy with for example the study for the US by Cecchetti, Mark, and Sonora (2002). Third, a recently developed new nonlinear unit root test is applied in this paper, which is more relevant to our study than other traditional linear unit root tests. Fourth, we developed a new non-linear half-life estimation procedure that is consistent with the concept of the linear counterpart. We find that not only price indexes do converge to the law of one price in China but also the half life of price index convergence is relatively quick (about 5 to 16 months in the linear cases, and 7 to 24 months in the nonlinear situation).

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1. Introduction

Amongst transitional economies, China has adopted a fairly unique gradualism approach for its economic reform. Whether such an approach has successfully transformed the Chinese economy into a market economy, which has significant policy implications, is still under heated debate. This debate appears to start with an influential and provocative contribution by Young (2000), who argues that the incremental reform adopted by China resulted in the fragmentation of Chinese domestic markets. Young (2000) supports his argument with three major pieces of evidence on provincial industrial structures, price dispersions, and agricultural production in response to weather changes, respectively. However, as pointed out by Holz (2006), it can be a complicated empirical task of testing market integration/fragmentations. For example, with the increasing open-up of China to the rest of the world, a Chinese firm may find that it can utilize its comparative advantage better by exporting to the USA and other developed countries than by selling its products in other provinces in China, Thus, different provinces in China may be less specialized with the increase in international trade, although there is little barrier of inter-provincial trade in China.[1] Also, while weather changes in a province affect its agricultural productivity, it does not provide enough information about

[1] In fact, this point is also noted by Young (2000). Young (2000, pp. 1115-1116) states: "One might argue that the factor endowments and relative productivities of China's diverse provinces appear quite dissimilar when viewed in isolation, but surprisingly homogeneous when contrasted with the rest of the world. If so, then when China was a closed economy, its provinces would have specialized in different sectors, but once China opened to the international market each province would naturally finds its comparative advantage in a common industrial structure. Alternatively, it is possible that interregional patterns of comparative advantage converged during the reform period and that this, without any appeal to internal trade barriers or the opening to the international market, explains the convergence in the composition of output." Young (2000) provides a counter-argument by examining a simple variance decomposition of primary versus secondary industry in China. However, as Holz (2006) points out, this simple method cannot derive convincing conclusions.

its comparative advantage since the changes of industrial productivity of the province (e.g, due to the changes of FDI) are not included in the study of Young (2000), and in fact, may be difficult to measure with accuracy. Thus, Holz (2006) concludes that neither argument nor evidence presented in Young (2000) is compelling.[2]

The current paper tests whether there is price convergence across markets in China using price index data. In terms of the examination of market integration, there are two main reasons/advantages for the empirical approach adopted by this study. First, since an effective price system is a fundamental component for a market economy (Roland, 2000), whether and to what extent the law of one price (LOP) holds in China are important indicators of the extent of its marketization.[3] Second, in contrast to most other empirical methods described in Young (2000) and Holz (2006), the study of LOP based on unit root tests and non-linear mean reversion is conceptually rigorous and is econometrically straightforward to implement.

This paper is related to Fan and Wei (2006), who presented evidence in support of LOP in China using individual commodity prices.[4] However, the number of commodities included in Fan and Wei (2006) (and in fact, in any given study using specific product prices) is only a small fraction of the total number of commodities of the whole economy. Thus, the current study complements Fan and Wei (2006) by using price indexes, which provide aggregate price information in China. As stated by Cecchetti, Mark, and Sonora (2002; CMS hereafter), the wider literature on PPP, exchange rates,

[2] Other related empirical studies in relation to Young (2000) include Bai et al (2005), Poncet (2005), and Fan and Wei (2006).
[3] In fact, testing the convergence to the law of one price is also the standard method in examining market integration in OECD countries (e.g. Engel and Rogers, 2004).
[4] It is also related to other studies of price adjustments and movements in transitional economics in general, which includes Berkowitz, DeJong, and Husted (1998), Berkowitz and DeJong (1999, 2001), and Conway (1999).

and other issues of monetary economics and international macroeconomics is mainly based on price index data.[5] The examination of price index convergence within a country therefore links with the literature on international macroeconomics and monetary policy more closely than studies that are based on the prices of individual commodities.

Our empirical analysis investigates price index movements in China with a monthly dataset that consists of the aggregate retail price index and six sub-indexes of retail prices in 36 major cities over 158 months. As in the study of CMS, we employ two recently developed econometric methods of panel unit root tests – those of Levin, Lin and Chu (2002) and Im, Pesaran, and Shin (2003) – to the balanced panel data of monthly price indexes for 36 major cities in China. Furthermore, this study extends the study of CMS in that we also apply a new nonlinear unit root test constructed by Kapetanios, Shin, and Snell (2003) to test the possibility that price convergence follows a non-linear rather than linear pattern.[6] Our econometric analysis shows that, based on the criteria of price convergence that are commonly used in the literature, there is strong evidence of convergence toward the law of one price in the Chinese domestic markets, which complements the literature in support of the view that the economic reform implemented in China has been generally successful in turning the economy towards a market orientation.

Moreover, this paper adds to the general literature on LOP by providing an improved method of estimating the half-life of price convergence when the underlying time series exhibit both linear and nonlinear convergence. This simple (yet new)

[5] For example, a better understanding of price movements in different regions of a currency area is crucial for the prediction of the impact of monetary policies on regional inflation differentials. Thus, economic analyses at the macroeconomic level need to depend on aggregate information, such as price indexes.

[6] Taylor and Taylor (2004) show that neglecting the possibility of nonlinear price convergence can seriously bias estimates of price convergence.

framework starts with testing if a price index series follows linear convergence. If it does then we calculate the half-life of price convergence using the conventional method, such as that in Im, Pesaran, and Shin (2003). If instead we find that the price index series exhibits nonlinear convergence, we then estimate the half-life of convergence using Monte Carlo simulation method along the lines of Granger and Teräsvirta (1993) and Taylor, Peel, and Sarno (2001). We find that the half life of price index convergence in China is about 6 to 19 months in the linear cases, and 6.6 to 24 months in the nonlinear situations. Interestingly, these estimators are significantly shorter than the findings obtained in the developed countries. For example, CMS report that the half life of price index convergence in the USA is 9 years (i.e., 108 months).Such a finding may be interpreted as being largely due to our use of higher frequency data than most received literature, which may result in more accurate estimations. As demonstrated by Taylor (2001), the use of low frequency data, such as annual or quarterly time-series data, are more likely to suffer from the "aggregation bias".

The rest of this study is organized as follows. The next section describes the background information and the data. Section 3 discusses the empirical methodologies, and in particular develops a new non-linear half-life estimation procedure that is consistent with the concept of the linear counterpart. Section 4 reports and analyzes the main empirical findings. We give our conclusions in Section 5.

2. Data Description

Our data are collected from the China Price Information Centre, a division of the State Planning Committee of the People's Republic of China, which began publishing monthly retail price index data in January 1998. The data that are used in this study cover 36 major Chinese cities[7] for the period of January 1994 to February 2007. In addition to the aggregate retail price index, there are six sub-indexes that are available in this dataset: food, beverages, clothing, medicines, electrical appliances, and fuel. It is argued by Imbs, Mumtaz, Ravn, and Rey (2005) that when the components of an aggregated time series (in our case, overall price index) have heterogeneous dynamics, the estimated persistence of the aggregated time series may be biased upward. Hence, the use of these sub-indexes not only allows us to check the robustness of the findings with the overall retail price index, but also enables us to detect and reduce this potential bias.

The quality of Chinese data is often criticized because data reporting in China is likely to be affected by political factors (e.g. Rawski, 2001; Holz, 2004). In particular, some types of data (e.g. GDP, unemployment) are often manipulated by local government officials as those data may be used as a measure of local officials' performance by the Chinese central government. However, price data are rarely used as such a measure and hence local officials seldom have incentives to manipulate them. Besides, as described in Fan and Wei (2006), there has been a rigorous mechanism in collecting price data in China. Thus, the data set employed in the current paper should be more reliable than most aggregate data obtained from China.

[7] These cities are Beijing, Changchun, Changsha, Chengdu, Dalian, Fuzhou, Guangzhou, Guizhou, Harbin, Haikou, Hangzhou, Hefei, Huhehaote, Jinan, Kunming, Lasha, Lanzhou, Nanchang, Nanjing, Nanning, Ningbo, Qingdao, Shenzhen, Shenyang, Shijiazhuang, Taiyuan, Tianjin, Wulumuqi, Wuhan, Xi'an, Xining, Xiamen, Yinchuan, Zhengzhou, Chongqing, and Shanghai. This includes four municipalities and all of the capital cities of the 28 provinces and autonomous regions in mainland China. The map in Appendix 1

3. Methodology

A common approach to the examination of price convergence is to apply the unit root test to examine whether price differential series are stationary. The rejection of the unit root hypothesis implies that the time series of relative prices is stationary, and that relative prices will converge in the long run. If these tests fail to reject this hypothesis, then the relative prices are deemed to follow a random path, in which case any deviation from the "one price" becomes permanent.

However, it has been noticed that the unit root test for a single time series, such as the Augmented Dickey-Fuller (ADF) test, has low power in the sense that it is too quick to reject the hypothesis of the stationary nature of a time series. However, Levin, Lin, and Chu (2002) show that the use of a unit root test of pooled time-series and cross-sectional (panel) data can significantly increase the power of the unit root test. They developed their method from a multivariate generalization of the ADF test, and provide statistical foundations for panel unit root tests. However, a limitation of the Levin-Lin-Chu test is that it imposes a cross-equation restriction on the first-order autocorrelation coefficients. Recognizing this problem, Im, Pesaran, and Shin (2003) put forward a panel unit root test that allows the autocorrelation coefficients to differ across panel members. Non-linear models of price movements have also been developed to minimize the complication of the dynamics of price convergence that may arise due to significant costs of transportation and transaction in inter-regional trade.

This study applies the econometric methods of panel unit root tests and a non-linear mean-reversion test. As in the study of CMS, we apply the panel unit root tests of

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shows the locations of these cities.

Levin, Lin and Chu (2002) and Im, Pesaran, and Shin (2003) first, and then apply the nonlinear unit root test of Kapetanios, Shin, and Snell (2003) to test the possibility that price convergence follows a non-linear, rather than a linear, pattern.

3.1. Levin-Lin-Chu Test

Similar to CMS, we define price variation as $p_{ij,t} = \ln(g_{ij,t} / \overline{g}_{j,t})$, where i, j and t stand for city, product and time, respectively; $g_{ij,t}$ denotes the raw price of product j in city i at time t; $\overline{g}_{j,t}$ denotes the mean of $g_{ij,t}$ over cities at time t. The Levin-Lin-Chu test is carried out by estimating the following equation:

$$\Delta p_{ij,t} = c_{ij} + \alpha_j p_{ij,t-1} + \sum_{h=1}^{K(j)} \beta_{hj} \Delta p_{ij,t-h} + \varepsilon_{ij,t} \tag{1}$$

where Δ is a first difference operator; and ε is an identically independently distributed (i.i.d.) error term. The number of lags, *K(j)*, to be included in (1) for each product and city series, is determined individually on a product by product basis as in a univariate ADF test using the Modified Akaiki Information Criterion (MAIC) and/or the Modified Schwarz Information Criterion (MSIC) developed by Ng and Perron (2001).[8]

Central to the test of convergence is the estimated value of α_j. If $\alpha_j \geq 0$, the price differential $p_{ij,t}$ is non-stationary, implying persistent or explosive price divergence for product j. A negative and statistically significant value of α_j implies price convergence, and its magnitude determines the speed of convergence. (1) can be estimated by generalized least square method with panel data. Levin-Lin-Chu (2002) has developed a

[8] Ng and Perron (2001) develop a new methodology to determine the number of lags of the autoregressive

panel version of ADF test (a t-test) and produced the corresponding critical values for their test. If the test-statistic is significant that means α_j is indeed negative and statistically significant, the half-life of a shock to the price differential (for product *j*) is computed as -ln(2)/ln(1+α_j).

3.2. Im-Pesaran-Shin Test

The main extension by Im, Pesaran, and Shin (2003) of the Levin-Lin-Chu test is to allow for heterogeneity in the value of α under the alternative hypothesis. The Im-Pesaran-Shin test starts with carrying out the Augmented Dickey-Fuller (ADF) test for every product and over each city. The regression takes the form:

$$\Delta p_{ij,t} = c_{ij} + \alpha_{ij} p_{ij,t-1} + \sum_{h=1}^{K} \beta_{ijh} \Delta p_{ij,t-h} + \varepsilon_{ij} \quad (2)$$

where the number of lags, *K*, to be included in (2) for each product and city series, is determined individually by using the MAIC or the MSIC on a city by city and product by product basis. As in the Levin-Lin-Chu test, all our ADF specifications include a constant term to capture city-specific fixed effects. Such effects may cover, for example, city-specific transportation cost and wage cost.

Let t_i denote the "*t*-statistic" for α_{ij}. Then, Im, Pesaran, and Shin (2003) design the following test statistic:

$$Z = \sqrt{N}(t - E(t)) / \sqrt{Var(t)}$$

variable, which improves the methods of AIC and SIC particularly when there are negative moving average errors.

where $t = (1/N)\sum_{i=1}^{N} t_i$; $E(t)$ and $Var(t)$ are the mean and variance of *t*, respectively. Following CMS, the half life or price convergence (for product *j*) is calculated as: –ln(2)/ln(1+$\bar{\alpha}_j$), where $\bar{\alpha}_j \equiv \frac{1}{N}\Sigma_{i=1}^{N}\alpha_{ij}$.

The panel unit root tests in this paper serve as our first step to check price convergence in China. These methods allow us to compare our results directly with those found for the US by CMS that use the same models of linear convergence.

3.3. Non-Linear Mean-Reversion Test

Dumas (1992) and Taylor (2001) have shown that the existence of significant transaction costs in inter-regional trade often results in a "band of inaction" within which inter-regional price differentials are too small to cover the transaction costs. In other words, if inter-regional price differentials are inside the "band of inaction," the activities of arbitrage are few and hence price differentials are likely to persist. In this case, price movements may experience non-linear dynamics.

The non-linear price convergence approach was initiated by Granger and Teräsvirta (1993) and Teräsvirta (1994)[9] and has gained popularity in a number of recent empirical studies, such as Michael, Nobay, and Peel (1997), Baum, Barkoulas, and Caglayan (2001), Taylor, Peel, and Sarno (2001) and Fan and Wei (2006). Our empirical study of non-linear price convergence here is based on a recently developed econometric test by Kapetanios, Shin, and Snell (2003). Following Kapetanios, Shin, and Snell (2003),

[9] There are also other nonlinear models testing price convergence. For example, see Canjels, Prakash-Canjels, and Taylor (2004).

the deviations from the law of one price (LOP) can be described in a standard ESTAR (exponential smooth threshold autoregressive) model, which takes the following form:

$$\Delta p_{ij,t} = \lambda_{ij} p_{ij,t-1} + \lambda_{ij}^{*} p_{ij,t-1} F(p_{ij,t-d}) + \sum_{h=1}^{m} \phi_{ijh} \Delta p_{ij,t-h} + \varepsilon_{ij,t} \qquad (3)$$

where

$$F(p_{ij,t-d}) = 1 - \exp(-\theta p_{ij,t-d}{}^{2})$$

θ is a positive coefficient, and $p_{ij,t}$ is a zero-mean series.[10] The ESTAR model was chosen as it allows for smooth and symmetric adjustments in response to both negative and positive deviations from the LOP. When $p_{ij,t-d} = 0$, $F(\cdot) = 0$ and (3) reverts to a standard linear model of (2) with $\alpha_{ij} = \lambda_{ij}$. When $p_{ij,t-d}$ departs far away from its zero-mean, $F(\cdot) \approx 1$ and (3) becomes a new linear *AR(m)* model with $\alpha_{ij} = \lambda_{ij} + \lambda_{ij}^{*}$. Thus, the critical parameters are λ_{ij} and λ_{ij}^{*}. When deviations from the LOP are small, λ_{ij} mainly determines the movements of relative prices. When deviations from the LOP become greater and greater, λ_{ij}^{*} becomes more and more important in governing the adjustment process.

By imposing λ_{ij}=0 and *d*=1, Kapetanios, Shin, and Snell (2003) applied the Taylor series approximation to (3) and derived the following specific ESTAR model:

$$\Delta p_{ij,t} = \delta_{ij} p_{ij,t-1}^{3} + \sum_{h=1}^{m} \phi_{ijh} \Delta p_{ij,t-h} + \varepsilon_{ij,t} \qquad (4)$$

A limiting nonstandard *t*-statistic is then developed to test the null hypothesis that $p_{ij,t}$ is

[10] Kapetanios, Shin, and Snell (2003) also studied series with non-zero mean and with time-trend, respectively.

non-stationary with a unit root (i.e., no price convergence):

$$H_0: \delta_{ij} = 0$$

against the alternative of a globally stationary ESTAR process (i.e., nonlinear price convergence):

$$H_1: \delta_{ij} < 0.$$

The *t*-statistic is given as:

$$t_{NL} = \hat{\delta}_{ij} / \text{s.e.}(\hat{\delta}_{ij}),$$

where $\hat{\delta}_{ij}$ is the OLS estimate of δ_{ij} and s.e.($\hat{\delta}_{ij}$) is the standard error of $\hat{\delta}_{ij}$. Asymptotic critical values of the t_{NL} statistics have been tabulated via stochastic simulations by Kapetanios, Shin, and Snell (2003, p.364, Table 1). A significant test-statistic of t_{NL} would indicate that the price is nonlinear stationary.

3.4. Estimation of non-linear half-life

The non-linear half-life of price convergence can be calculated by the Monte Carlo method along the lines of Granger and Teräsvirta (1993) and Taylor, Peel, and Sarno (2001). Unlike a linear model, the half-life of a nonlinear ESTAR model is dependent upon the initial values (path-dependence) and the size of the shock (shock-dependence). The Monte Carlo procedure employed in this paper is described as follows. First, a dynamic reference series of 200 periods is constructed by simulating the ESTAR model (3) for each price variation $p_{ij,t}$, with initial value set to its mean $\overline{p}_{ij,t}$, and with random draws of a Gaussian innovation that has an estimated variance of the residual from model (3). Second, this procedure is then repeated with the same sequence of random draws but with a new initial value $\{|p_{ij,1994m01} - \overline{p}_{ij,t}| + \overline{p}_{ij,t}\}$, where 1994m01 is the first observation of our sample. This new initial value in fact is equal to the mean of $p_{ij,t}$ plus a (positive) deviation from the mean. As Taylor, Peel, and Sarno (2001) argued, this transformation is necessary since our ESTAR model is symmetric. Hence we only need to consider positive shock to the initial value. Third, the difference of the second simulated series and the reference series at each period is stored. We repeat this procedure for 200 times. An average over all of the 200 simulated sequences of differences is then estimated. The half-life of the first observation is found when the averaged difference path hits the value of $0.5|p_{ij,1994m01} - \overline{p}_{ij,t}|$. Then we move to the second observation of our sample, i.e., 1994m02, to estimate its half-life, and repeat this procedure until the last observation at 2007m02. Our sample size is 134 months over the period of 1994m01 to 2007m02. Once all the simulations have been completed, we have 134 estimated half-lives. An average over all of the 134 estimated half-lives is the simulated half-life of the ESTAR model.

The value-added of our calculation of a nonlinear half-life is two-folded. First, it completes the convergence picture. For linear convergence cases we can estimate the speed of convergence. If there were no estimation of nonlinear half-life, we would not know their speed of convergence. Furthermore, without nonlinear half-life estimation, the ESTAR PPP literature only addresses half of the PPP non-convergence puzzle. That is, they can only show many prices are convergent nonlinearly, rather than divergence as linear model suggested. However, they did not provide how fast the nonlinear convergence is. Hence they did not address the slow convergence puzzle of the linear model.

Second, our new method of nonlinear half-life estimation is the first one in literature that is comparable to the concept of linear half-life. One may estimate the half-life of different shocks like the literature of the nonlinear impulse response of arbitrary shocks. This method did provide an average half-life of a nonlinear series, since the range of the shocks is ad hoc and is not the actual shock of the data. Our new method effectively estimates how long the ESTAR model takes from an initial-value of each observed price $p_{ij,t}$ converging towards $0.5|p_{ij,t} - \overline{p}_{ij,t}|$. This is exactly the concept in the linear half-life.

4. Results

This section is divided into two parts. In Subsection 4.1, we report the results that are based on the panel unit tests of Levin, Lin and Chu (2002) and Im, Pesaran, and Shin (2003). As this subsection employs the same empirical methodology as that of CMS, it is easy to compare our findings for China with their findings for the United States. For those city price indexes that do not show linear convergence, the non-linear unit root test are applied to find out if they have non-linear price convergence based on Kapetanios, Shin, and Snell (2003).

4.1. Panel Unit Root Tests

As in the study of CMS, this study aims to examine the movements of the retail price index in different regions of China. It also examines the movements of the price indexes for six categories of goods – food, beverages, clothing, medicines, electrical appliances, and fuel – as a robustness check of the results that are obtained for the aggregate retail price index. The following table displays the results from the price convergence tests that are based on the Levin-Lin-Chu test.

Table 1: Results based on Levin-Lin-Chu (LLC) Test

Price Index	LLC Statistics (MAIC)	P-value	LLC Statistics (MSIC)	P-value
Aggregate RPI	5.89069	1.0000	4.80419	1.0000
Food	3.90502	0.9997	1.16039	0.8771
Beverages	3.67213	0.9998	1.94584	0.9742
Clothing	8.10211	1.0000	6.41455	1.0000
Medicines	4.78465	0.9986	0.72484	0.7657
Electrical Appliance	10.4195	1.0000	8.42794	1.0000
Fuel	6.73924	0.9999	-0.10635	0.4577

Note: P-value>0.05 indicates insignificance at the 5% level.

From Table 1, we can see that neither the aggregate retail price index nor any of the sub-indexes converges based on the Levin-Lin-Chu test when MAIC and MSIC are used to select lags. However, as discussed in the previous section, the Levin-Lin-Chu test is commonly viewed as having a low power to reject the unit root null hypothesis. We next turn to the results that are based on the Im-Pesaran-Shin test, which allows for heterogeneity in the rate of convergence among the time series of different city pairs. The calculated Im-Pesaran-Shin test statistics are presented in the following table.

Table 2: Results based on Im-Pesaran-Shin (IPS) Test

Price Index	**IPS Statistics (MAIC)**	**IPS Statistics (MSIC)**
Aggregate RPI	-1.98081**	-2.48819**
Food	-3.12483**	-4.38756**
Beverages	-2.13148**	-2.33868**
Clothing	0.00182	0.70554
Medicines	-1.79751**	-3.53065**
Electrical Appliance	-4.14809**	-1.8259**
Fuel	-2.41244**	-4.31767**

Note: ** means significance at the 5% level.

Comparing Table 2 with Table 1, we can see that the Im-Pesaran-Shin test significantly increases the power to reject the unit root null hypothesis relative to the Levin-Lin-Chu test. The Im-Pesaran-Shin test indicates that the aggregate price index in China experienced clear convergence during the period under study. In addition, most of the sub-indexes of goods (food, beverages, medicines, electrical appliances, and fuel) are found to converge to the law of one price at the 5% significance level based on both the MAIC and the MSIC lag selection criteria. The only exception is clothing, the price indexes of which did not seem to converge across regions. These findings here are largely consistent with those of CMS for the US.

4.2. Non-Linear Mean-Reversion Tests

For the purposes of comparison, we first look at the single time-series (linear) unit root (ADF) test results, which are presented in Table 4.

Table 4: Proportion of Cities that Exhibit Linear Convergence in ADF Test

Price Index	Percentage
Aggregate RPI	19.4%
Food	25.0%
Beverages	19.4%
Medicines	16.7%
Electrical Appliances	22.2%
Fuel	27.8%

Note: Based on Appendix Table B.

Table 4 reports for each product the proportion of cities in which the univariate ADF test rejects the unit root hypothesis at the 5% significance level and therefore exhibit the *linear* convergence. For the aggregate retail price index this proportion is 19.4%, and for the individual categories it ranges from 16.7% (for medicines) to 27.8% (for fuel).

We now turn to the analysis of non-linear price movements. The empirical findings are summarized in Table 5.

Table 5: Non-Linear Tests

Price Index	% of cities follow non-linear convergence	Simulated nonlinear half-lives (unit: month)	% of convergent cities (linearly or non-linearly)
(1)	(2)	(3)	(4)
Aggregate RPI	72.4	13.8	77.8
Food	96.3	6.6	97.2
Beverages	82.8	24.0	86.1
Clothing	75	15.8	75
Medicines	70	14.4	75
Electrical Appliance	65.5	19.6	75
Fuel	76.9	15.7	83.3

Note: Based on Appendix Table C and D. For column (2), the percentages are calculated for all cities that fail to exhibit the *linear* convergence in the ADF unit root tests.

We first examine the patterns of price adjustments of the time series that failed to exhibit the *linear* convergence in the ADF unit root tests. We apply the non-linear unit root test t_{NL} of Kapetanios, Shin, and Snell (2003) to them, the results of which are reported in column (2) of Table 5. We make the following observations. First, the proportion of cities that fail the linear ADF test but follow a non-linear convergence to the law of one price is high: 72.4% for the aggregate retail price index. Moreover, individual price indexes tend to be more likely to exhibit non-linear convergence. For example, the sub-indexes for food 96.3% cities that fail to exhibit linear convergence in the ADF unit root test follow the non-linear convergence. In addition, the proportion of nonlinear convergence among the remaining five categories of goods is also fairly large. Specifically, it is quite high for beverages (82.8%), fuels (76.9%), clothing (75%), medicines (70%) and electrical appliances (65.5%).

Finally, we report in the last column of Table 5 the proportion of cities with indexes that display overall convergence, either linearly or non-linearly. For the aggregate retail price index, overall convergence is exhibited for 77.8% of the cities. This is a large increase from the case in which linear price movement is imposed based on the univariate ADF test (19.4%). For the six categories of goods, the difference in the proportion of cities in which prices converge under the linear ADF test and under both the linear and the non-linear analysis is even greater: 25% to 97.2% for food, 19.4% to 86.1% for beverages, 0% to 75% for clothing, 16.7% to 75% for medicines, 22.2% to 75% for electrical appliances, and 27.8% to 83.3% for fuels, respectively. The results therefore show that the incorporation of non-linear dynamics into the analysis significantly increases the power to dictate the LOP.

In summary, our econometric analysis shows strong evidence of price convergence – and hence market integration – in China. This finding confirms Fan and Wei (2006) and supports the view that China's transition to a market economy has been largely successful. It is also consistent with the evidence reported by Dougherty and Herd (2005), who find from the surveys of businesses in China that price and quantity controls are of little importance in restricting inter-provincial trade in recent years. Hence, the current study complementing the literature in support of the view that the gradualism reform approach implemented by the Chinese government has been quite successful in transforming its economy into a market-oriented one.

5.3. The Speed of Convergence

This subsection looks at the speed of retail price index convergence in China, which is measured in half lives. As the Levin-Lin-Chu test generally does not reject the unit root null hypothesis for the price indexes, it is not meaningful to estimate the half life of price convergence based on this test. Therefore, in the linear case we only focus on the half-life estimates that are based on the Im-Pesaran-Shin test. As discussed in the introduction, the results will be modified if the underlying dynamics of the time series is non-linear. But we might as well present the results here for a direct comparison with CMS. The results for the aggregate retail price index and for the five categories of goods that are shown to exhibit convergence based on the Im-Pesaran-Shin test are listed in the following table.[11]

Table 5: Estimated Half-lives (unit: month)

(1)	(2)	(3)	(4)	(5)	(6)	(7)
Price Index	**Linear half-life (ADF-MAIC)**	**Linear half-life (ADF-MSIC)**	**No. of cities follow linear convergence**	**Simulated nonlinear half-life (KSS)**	**No. of cities follow non-linear convergence**	**Average half-life (linear and non-linear)**
Aggregate RPI	15.6	14.9	7	13.8	21	14.1
Food	4.7	4.9	9	6.6	26	6.2
Beverages	10.5	9.8	7	24.0	24	20.8
Clothing	-	-	0	15.8	27	15.8
Medicines	12.9	11.2	6	14.4	21	13.7
Electrical Appliance	9.4	10.3	7	19.6	19	16.8
Fuel	6.3	6	10	15.7	20	12.5

Note: Column (2) and (3) show the average half-lives across those cities that exhibit linear price convergence according to the ADF unit root test based on MAIC and MSIC lag selection criteria respectively. Column (5) contains the average half-lives across those cities that follow non-linear convergence (based on the KSS

[11] The half life for the clothing index is not reported, as it did not converge in the first place.

test) instead of linear convergence (based on the ADF unit root tests). Column (7) shows the average linear and non-linear half-lives in col. (3) and (5). Price index of clothing does not exhibit any linear convergence by the IPS test and therefore no linear half-life is calculated. As a result, its average half-life in col. (7) is based on the non-linear half-life only. The details of the half-lives of each price index for each city are reported in Appendix Table A, B, C and D.

We first look at the estimated half-lives for different price indexes based on the linear IPS test. We note that the calculated half-lives in Column 2 and 3 of Table 5 are similar regardless of whether they are based on the MAIC or the MSIC. In particular, the half life of the aggregate retail price index is estimated to be about 15 months, and the half-lives of the five categories of goods with convergent price indexes are found to range between 5 months and 13 months. Thus, we have a fairly robust finding that the half life of price index convergence in China is within 0.5 to 1.5 years.

This finding is striking compared to the evidence in the literature. In cross-country studies, the half life of price index convergence is normally found to be about 2 to 5 years (Froot and Rogoff, 1995). This slow speed of convergence is described by Rogoff (1996) as the "PPP puzzle." In the study of CMS, the half life of price index convergence in the United States is estimated to be 6.18 years based on the Levin-Lin-Chu test and 5.57 years based on the Im-Pesaran-Shin test.[12] As price convergence should take place more rapidly across regions within a single country than across countries, their results seem counter-intuitive. Our empirical results appear to be more in line with the common expectation.

[12] Cecchetti, Mark, and Sonora (2002) calculate the *adjusted* half life of price index convergence based on Nickell (1981). Their estimate of the *adjusted* half life is 8.54 years based on the Levin-Lin-Chu test and 9.70 years based on the Im-Pesaran-Shin test. Their finding of a half life that is even longer than that found in most cross-country studies of PPP is also contrary to their expectations. They expected to observe a more rapid price convergence across regions within a single country than across countries, as within-country markets for products, labor, and capital are presumably better integrated.

Next, we relate our findings to those of Imbs, Mumtaz, Ravn, and Rey (2005), which show that under certain technical conditions the speed of convergence becomes much faster if the heterogeneity of the dynamics of the components of an aggregated time series is taken into account. Specifically, they find that when the heterogeneity of different components is taken into account in constructing the price index, the half life of real exchange rates in several OECD countries decreases dramatically, and may fall to as low as 11 months. It is interesting to note that our estimated half life of price index convergence in China is comparable with their estimated half lives of cross-country price convergence. Our empirical findings also show that the half life of the aggregate retail price index is somewhat longer than the half lives of the sub-indexes. This result provides some evidence to support the argument of Imbs, Mumtaz, Ravn, and Rey (2005).

We now turn to the estimated speed of convergence if the underlying time series is non-linear. The results are presented in Column 5 of Table 5. The average simulated half-lives of the overall retail price index is 13.8 months, which is very similar to the estimated half-life when the underlying time series are assumed to follow linear convergence. The simulated half-lives for the sub-indexes range from 6 to 24 months, which are also not far away from those figures in Column 2 and 3. Finally, we report the hybrid half-lives in Column 7 of Table 5, which is the average of the linear and non-linear half-lives. These half-lives range from 6 months to 20 months, which is not much different from the half-lives based on the linear convergence alone, and are still significantly shorted than most of the half-lives found in the PPP literature.

In summary, it is interesting to note that the speed of price indexes convergence in China is considerably faster than those in OECD countries, and in particular, than that in the USA as reported by CMS. We have the following two main explanations for this finding. First, it is demonstrated by Taylor (2001) that the estimated speed of price convergence can be seriously biased upward if low-frequency data are used. Unlike CMS, we use monthly price index data instead of annual price index data. We believe that the monthly frequency of our price data matches well with the actual time that is needed for price arbitrage in China, and hence is more likely to avoid or reduce the estimation bias pointed out by Taylor. Thus, in light of Taylor's study, it is not surprising to find that the half life of convergence is longer in the study of CMS, despite the fact that they examine price convergence in the United States, where market efficiency should be higher than in China.

Second, although a rigorous empirical study to explain the underlying reasons is not provided in our current paper and clearly deserves a separate paper in future research, we conjecture two main explanations for the obtained result that the speed of price convergence in China is faster than that in some OECD countries. First, relative to the OECD countries, the market economy in China is still at the early stage of development, in which there is usually more competition among firms and less monopoly. For example, supermarket is only a very recent phenomenon in China for most Chinese cities. In contrast, most OECD countries have national supermarket chains, which can to some extent control prices and engage in "price discriminations" in different regions. Second, if the activity of arbitrage is labor intensive, then there will be more activities of arbitrage in poor countries, which is labor abundant. Consequently, we may observe faster price convergence in less developed countries.

Moreover, the above explanations are consistent with the findings of the related empirical studies of other developing/transitional countries that also utilized monthly price data. For example, the study by Das and Bhattacharya (2004) on price index convergence across regions in India finds that the half-lives for price index convergence vary from 8 to 22 months, depending on whether the price shock is due to the common factor or idiosyncratic factors. The study by Gluschenko (2004) for Russia finds that the overall mean half-life of price convergence is 6.5 month there.

5. Concluding Remarks

This study applies the panel unit root tests of Levin, Lin and Chu (2002) and Im, Pesaran, and Shin (2003) and the non-linear mean-reversion test that was developed by Kapetanios, Shin, and Snell (2003) to study price convergence to the law of one price in China. Using a panel dataset that consists of an aggregate retail price index and six sub-indexes in 36 major cities over 158 months, our econometric analysis shows that the overall price index and most of the sub-indexes exhibit the tendency to converge to the law of one price in the Chinese domestic markets. This finding supports and also complements our earlier finding of price convergence in China using individual commodity prices, and contradicts the claim of Young (2000).

The study of the dynamics of price indexes in China has interesting policy implications. First, the examination of price movements is the strategy that is most commonly used in the literature to test market integration (for example, Engel and Rogers, 2004). By showing that the Chinese domestic markets perform fairly efficiently, our study provides empirical support for the view that the economic reform that has been implemented in China has been generally successful in turning the economy toward a market orientation (Ma, 2007). Second, the Chinese government has frequently urged its trading partners to recognize China as a market economy, and the failure of these partners to do so is damaging China's trade due to the harsh anti-dumping policies that are apply to non-market economies under the rules of the World Trade Organization. This study provides direct evidence to support China's claim as a market economy. Third, national economies have become increasingly globalized and world markets increasingly integrated.[13] In this era of globalization, the Chinese economy is playing an increasingly important role in the world economy. In particular, Chinese monetary policies appear to be having a growing impact on the global financial markets. Thus, the study of the dynamics of price indexes in China is not only academically interesting, but also has profound policy implications.

We also find that the average half life of price convergence in China is only about 5 to 16 months in the linear cases, and 7 to 24 months in the nonlinear situation. Our findings of linear and non-linear convergence speed contrast sharply with most of the existing cross-country studies of PPP that find the half life of price convergence to range from 2 to 5 years, and also with the within-country study of Cecchetti, Mark, and Sonora (2002) for the United States that estimates a half life of 5 years. In line with Taylor (2001), we interpret this difference as being largely due to the employment of high-frequency data in our study, which matches well with the time that is needed for domestic price arbitrage. However, in our cases of non-linear convergence, it exhibits much slower convergence speed in terms of longer half lives than the linear counterparts. It remains for future research to compare our non-linear results with that from other countries.

[13] For example, Chen, Imbs, and Scott (2004) demonstrate that increased openness, as measured by import shares, led to lower prices in the European Union over the period 1988-2000.

Appendix

Table A: Estimated Half-lives based on MAIC (unit: month)

	City	Aggregate RPI	Food	Beverages	Clothing	Medicines	Electric Appliances	Fuels
1	Beijing	.	.	.	.	.	.	.
2	Tianjin	.	4.48	5	.	.	2.13	.
3	Shijiazhuang	29.5	.	.	.	.	.	3.09
4	Taiyuan	5.5	.	.	.	.	.	.
5	Hohhot	36.5	2.44	.	.	2.53	.	.
6	Shenyang	.	1.02	9.54	.	.	.	4.32
7	Dalian	1.86	.	.	.	.	.	.
8	Changchun	.	.	.	.	3.18	.	.
9	Harbin	.	.	.	.	.	.	.
10	Shanghai	.	2.09	.	.	.	.	.
11	Nanjing	.	5.01	22.5	.	.	.	.
12	Hangzhou	.	.	.	.	.	.	.
13	Ningbo	.	.	.	.	.	.	.
14	Hefei	5.72	.	.	.	.	.	.
15	Fuzhou	.	.	.	.	.	.	.
16	Xiamen	.	.	.	.	6.65	.	.
17	Nanchang	.	.	.	.	.	.	.
18	Jinan	.	.	.	.	.	.	.
19	Qingdao	.	.	.	.	.	.	.
20	Zhengzhou	.	.	.	.	.	.	2.1
21	Wuhan	.	.	.	.	.	.	3.85
22	Changsha	.	.	.	.	.	.	.
23	Guangzhou	.	7.44	.	.	.	35.3	11.1
24	Shenzhen	.	.	7.87	.	.	12.4	.
25	Nanning	.	.	5.94	.	.	.	5.06
26	Haikou	.	6.21	.	.	.	.	10.5
27	Chongqing	.	.	.	.	.	2.44	5.11
28	Chengdu	11.5	4.34	.	.	40.8	12	.
29	Guiyang	.	.	18.2	.	19.9	.	.
30	Kunming	.	.	.	.	.	7.65	.
31	Lhasa	.	.	.	.	.	0	.
32	Xi'an	.	.	.	.	.	.	.
33	Lanzhou	.	.	4.32	.	.	.	.
34	Xining	.	.	.	.	.	3.33	.
35	Yinchuan	18.7	9.04	.	.	4.62	.	6.97
36	Urumqi	.	.	.	.	.	.	10.9
	Average half-life	15.6	4.7	10.5		12.9	9.4	6.3
	% of cities with linear convergence	19.4%	25%	19.4%	0%	16.7%	19.4%	27.8%

Table B: Estimated Half-lives based on MSIC (unit: month)

	City	Aggregate RPI	Food	Beverages	Clothing	Medicines	Electric Appliances	Fuels
1	Beijing	.	.	.	.	.	.	.
2	Tianjin	.	4.48	5	.	.	3.1	.
3	Shijiazhuang	25.8	.	.	.	.	.	3.82
4	Taiyuan	5.01	.	.	.	.	.	.
5	Hohhot	36.5	2.44	.	.	2.53	.	.
6	Shenyang	.	1.02	6.94	.	.	.	2.97
7	Dalian	1.86	.	.	.	.	.	.
8	Changchun	.	.	.	.	3.68	.	.
9	Harbin	.	.	.	.	.	.	.
10	Shanghai	.	2.09	.	.	.	.	.
11	Nanjing	.	5.01	22.5	.	.	.	.
12	Hangzhou	.	.	.	.	.	.	.
13	Ningbo	.	.	.	.	.	.	.
14	Hefei	5.04	.	.	.	.	.	.
15	Fuzhou	.	.	.	.	.	.	.
16	Xiamen	.	.	.	.	3.37	.	.
17	Nanchang	.	.	.	.	.	.	.
18	Jinan	.	.	.	.	.	.	.
19	Qingdao	.	.	.	.	.	.	.
20	Zhengzhou	.	.	.	.	.	.	1.84
21	Wuhan	.	.	.	.	.	.	3.85
22	Changsha	.	.	.	.	.	.	.
23	Guangzhou	.	7.44	.	.	.	35.3	11.1
24	Shenzhen	.	.	5.35	.	.	13.1	.
25	Nanning	.	.	5.94	.	.	.	5.06
26	Haikou	.	7.33	.	.	.	.	12.7
27	Chongqing	.	.	.	.	.	5.5	5.11
28	Chengdu	11.5	5.33	.	.	40.8	12	.
29	Guiyang	.	.	18.8	.	12	.	.
30	Kunming	.	.	.	.	.	8.22	.
31	Lhasa	.	.	.	.	.	0	.
32	Xi'an	.	.	.	.	.	.	.
33	Lanzhou	.	.	3.95	.	.	.	.
34	Xining	.	.	.	.	.	4.91	.
35	Yinchuan	18.7	9.04	.	.	4.62	.	3.63
36	Urumqi	.	.	.	.	.	.	10.1
	Average half-life	14.9	4.9	9.8		11.2	10.3	6.0
	% of cities with linear convergence	19.4%	25%	19.4%	0%	16.7%	19.4%	27.8%

Table C. Simulated nonlinear half-lives (unit: month)

	City	Aggregate RPI	Food	Beverages	Clothing	Medicines	Electric Appliances	Fuels
1	Beijing	40.4	4.59	47	20.1	.	23.9	.
2	Tianjin	13.5	.	.	19.2	.	.	.
3	Shijiazhuang	.	3.31	.	28.5	13.8	17	.
4	Taiyuan	.	4.26	35.3	2.45	22.2	25.7	22
5	Hohhot	.	.	25.3	.	.	.	22.5
6	Shenyang	6.01	.	.	8.08	25.1	14	.
7	Dalian	.	3.44	25.1	7.56	13.5	10.2	20.2
8	Changchun	.	8.09	.	.	.	.	9.67
9	Harbin	5.54	3.1	8.36	6.69	.	.	8.97
10	Shanghai	14.4	.	.	21	6.43	54.7	21
11	Nanjing	.	.	.	3.1	14	.	.
12	Hangzhou	.	7.92	48	23	19.7	16.4	27.8
13	Ningbo	.	9.44	49	33.5	33.3	22.6	13.7
14	Hefei	.	5.39	22.9	.	9.67	47.3	16.8
15	Fuzhou	5.69	3.96	18.7	14.6	5.42	.	32.1
16	Xiamen	.	19.3	26.3	.	.	18.6	13.7
17	Nanchang	4.59	3.46	9.01	4.09	.	.	14.9
18	Jinan	3.68	3.62	45.3	9.72	.	5.56	6.87
19	Qingdao	5.25	3.25	.	10	7.84	23.7	8.34
20	Zhengzhou	.	7.53	20.5	16.8	.	25.9	.
21	Wuhan	19.2	3.91	10.4	24.4	.	23.9	.
22	Changsha	11.7	3.4	50	8.1	.	.	3.64
23	Guangzhou	23.9	.	.	.	13.7	.	.
24	Shenzhen	8.72	12.8	.	.	.	.	20.2
25	Nanning	38.6	.	.	5.3	10.7	12.2	.
26	Haikou	15.7	.	42	47.5	5.42	5.16	.
27	Chongqing	12	9.22	4.86	.	18.2	.	.
28	Chengdu	.	.	10.2	18.5	.	.	.
29	Guiyang	9.87	8.7	.	11.2	.	5.82	11.8
30	Kunming	11.8	8.87	11.5	30.2	4.18	.	17.9
31	Lhasa	17.6	7.97	9.21	14.2	8.19	.	.
32	Xi'an	6.31	1.74	13	20	9.84	9.29	14.6
33	Lanzhou	15	4.31	.	14.7	40.8	.	.
34	Xining	.	14.8	13.5	.	11.1	.	6.92
35	Yinchuan	.	.	23.2	.	.	11	.
36	Urumqi	.	6.22	8.27	3.6	9.45	.	.
	Average half-life	13.8	6.6	24.0	15.8	14.4	19.6	15.7
	% of cities with nonlinear convergence	58.3	72.2	66.7	75	58.3	52.8	55.6
	% of cities that regarded as I(1) by ADF test but with nonlinear convergence	72.4	96.3	82.8	75	70	65.5	76.9

Table D. Summary table of convergence and non-convergence

		Aggregate RPI	Food	Beverages	Clothing	Medicines	Electric Appliances	Fuels
1	Beijing	NlinC	NlinC	NlinC	NlinC	x	NlinC	x
2	Tianjin	NlinC	LC	LC	NlinC	x	LC	x
3	Shijiazhuang	LC	NlinC	x	NlinC	NlinC	NlinC	LC
4	Taiyuan	LC	NlinC	NlinC	NlinC	NlinC	NlinC	NlinC
5	Hohhot	LC	LC	NlinC	x	LC	x	NlinC
6	Shenyang	NlinC	LC	LC	NlinC	NlinC	NlinC	LC
7	Dalian	LC	NlinC	NlinC	NlinC	NlinC	NlinC	NlinC
8	Changchun	x	NlinC	x	x	LC	x	NlinC
9	Harbin	NlinC	NlinC	NlinC	NlinC	x	x	NlinC
10	Shanghai	NlinC	LC	x	NlinC	NlinC	NlinC	NlinC
11	Nanjing	x	LC	LC	NlinC	NlinC	x	x
12	Hangzhou	x	NlinC	NlinC	NlinC	NlinC	NlinC	NlinC
13	Ningbo	x	NlinC	NlinC	NlinC	NlinC	NlinC	NlinC
14	Hefei	LC	NlinC	NlinC	x	NlinC	NlinC	NlinC
15	Fuzhou	NlinC	NlinC	NlinC	NlinC	NlinC	x	NlinC
16	Xiamen	x	NlinC	NlinC	x	LC	NlinC	NlinC
17	Nanchang	NlinC	NlinC	NlinC	NlinC	x	x	NlinC
18	Jinan	NlinC	NlinC	NlinC	NlinC	x	NlinC	NlinC
19	Qingdao	NlinC	NlinC	x	NlinC	NlinC	NlinC	NlinC
20	Zhengzhou	x	NlinC	NlinC	NlinC	x	NlinC	LC
21	Wuhan	NlinC	NlinC	NlinC	NlinC	x	NlinC	LC
22	Changsha	NlinC	NlinC	NlinC	NlinC	x	x	NlinC
23	Guangzhou	NlinC	LC	x	x	NlinC	LC	LC
24	Shenzhen	NlinC	NlinC	LC	x	x	LC	NlinC
25	Nanning	NlinC	x	LC	NlinC	NlinC	NlinC	LC
26	Haikou	NlinC	LC	NlinC	NlinC	NlinC	NlinC	LC
27	Chongqing	NlinC	NlinC	NlinC	x	NlinC	LC	LC
28	Chengdu	LC	LC	NlinC	NlinC	LC	LC	x
29	Guiyang	NlinC	NlinC	LC	NlinC	LC	NlinC	NlinC
30	Kunming	NlinC	NlinC	NlinC	NlinC	NlinC	LC	NlinC
31	Lhasa	NlinC	NlinC	NlinC	NlinC	NlinC	LC	x
32	Xi'an	NlinC	NlinC	NlinC	NlinC	NlinC	NlinC	NlinC
33	Lanzhou	NlinC	NlinC	LC	NlinC	NlinC	x	x
34	Xining	x	NlinC	NlinC	x	NlinC	LC	NlinC
35	Yinchuan	LC	LC	NlinC	x	LC	NlinC	LC
36	Urumqi	x	NlinC	NlinC	NlinC	NlinC	x	LC
	% of cities with linear convergence	19.4	25	19.4	0	16.7	22.2	27.8
	% of cities with nonlinear convergence	58.3	72.2	66.7	75	58.3	52.8	55.6
	% of cities with linear or nonlinear convergence	77.8	97.2	86.1	75	75	75	83.3

	% of cities without any convergence	22.2	2.78	13.9	25	25	25	16.7

Note: 'LC' and 'NlinC' indicate linear and non-linear convergence respectively. 'x' indicates neither liner nor non-linear convergence.

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Appendix 1: The 36 Cities in the Map of China