# TRADE AND INTEGRATION OF THE US AND CHINA´S COTTON MARKETS

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## Abstract

The cotton market in China is highly interactive with international markets, especially, the US market. The prices in these two markets can reveal important market relations. Investigating the data of futures prices from the New York Board of Trade (NYBOT) and the Zhengzhou Commodity Exchange (CZCE) using several time series methods, we find a long-run cointegration relationship between these I(1) series. Furthermore, a bi-directional Granger Causality between these two futures markets is detected with Generalized Autoregressive Conditional Heteroskedasticity (GARCH) error specifications. We also find the relationship is impacted by the Chinese exchange rate policy change in the 2005.

**Keywords:** cotton futures prices, cointegration, granger causality test, AR-GARCH.

## Introduction

Cotton, throughout history, stands as a prime commodity in the world economy. As a cotton pricing arena, the cotton futures exchange plays an important role in price discovery and risk shifting. Among all the countries which are involved in cotton production, consumption, and trade, China and the US represent the two most important powers. China is the world's largest cotton producer and consumer, while the US is the world's largest cotton exporter and supplier to China. As seen in Figure 1, China's cotton productions accounted for around 28% averagely of the world total from 2004 to 2006. The percentages are even higher in terms of import. Figure 2 shows the major sources of cotton exported to China in recent years, among which are 1.7 million tons from the US in 2006, accounting for 46.9% of China's total import. Besides the cotton trade, the textile industry, which uses cotton as the major raw material, has also benefited from the surging trade between China and the US in recent years. [1]

Because of the dominance and interdependence of the US and China in the global cotton and textile industries, it is of great value to analyze the long-run relationship of the cotton futures prices between these two countries, since the futures market prices reveal the spot prices and will enormously influence the world cotton market. Intuitively, it is expected that there exists a long-run equilibrium between these markets because of this strong interdependence, but the short history of Chinese cotton futures market and the trade and financial policy interventions might have prevented the formation of such equilibrium. This study will provide useful conclusions not only to the cotton and textile producers, consumers, and traders, but also to market speculators and the exchanges and regulatory agencies in these two and other countries.

A brief introduction to the US and China's exchanges in which cotton futures contracts are traded is helpful as a precursor to the explanation of procedure and methodology in this paper.

Cotton futures have been traded since 1870 in New York, the original futures exchange in the US. The New York Board of Trade (NYBOT) now is the one of the most active cotton futures exchanges in the world. The contract size for cotton is 50,000 lbs and the prime trading months for cotton futures at the NYBOT include March, May, July, October, and December. In 2005, around 5.5 million cotton futures contracts were traded in the NYBOT. [2]

China started its cotton futures trading in June of 2004. The Zhengzhou Commodity Exchange (CZCE) is the only exchange for cotton futures trading in China today. The contract size for cotton is 5 metric tons. Contract months for cotton futures in the CZCE include all months but February. In 2005, the total trading volume reached 21,741,400 contracts, while the trading value reached 1568.5 billion Yuan. [3]

With the development of technology and the improvement of market mechanics, various agricultural commodity markets have exhibited a strong integration trend. Many studies have shown a long-run linkage between pairwise prices of several agricultural commodities (Malliaris & Urrutia, 1996; Booth & Ciner, 2001; Liu, 2005). However, all of the above only analyzed the different commodities markets within one country, especially in the US. In recent years, the integration of the global agricultural market has become a popular research topic, and studies on the relationships of agricultural commodity markets across countries have attracted a lot of attention. Several studies have been extended to explore the bilateral relationship of the agricultural commodities markets for some areas in the world, specifically between the US and Canada, between the US and EU, and among EU countries (Booth et al., 1998; Yang et al., 2003; Viju et al., 2006). These studies focused primarily on wheat and other major grain commodities, and they find strong long run relationships between US and Canada and among some EU countries.

With the prominent growth of China's agricultural commodity markets and a tight linkage between China and the global economies, a number of articles have examined the degree of market integration between the emergent Chinese commodity futures markets and other futures markets in the world. The results are mixed. Quite a few studies have found that China's major domestic futures markets for commodities such as rice, corn, and soybeans exhibit cointegration relationships with other foreign markets (see Wu, 2001; Zhao, 2002). Others found no such relationship for wheat futures markets between China and US (see Fung et al., 2003; Du & Wang, 2004).

To our knowledge, no research has been found examining the linkage of different cotton futures markets, despite the fact that cotton is a key agricultural product in the world. This paper explores the cotton futures price linkage between China and the US, and thus makes a significant contribution to the literature and practice. Furthermore, since July 21, 2005, China has adopted a new market-oriented exchange rate system, and this reform has shown deep influences on every aspect of China's economy. First, this paper examines the structural change effect of China's new exchange rate mechanism on cotton futures prices.

**Theory and Methodology**

First, cointegration and error-correction techniques are adopted in this paper because they are the primary methods used for market integration analysis, as applied in most of the studies mentioned above. The cointegration analysis is used in the beginning to check if prices between these two cotton futures markets are related by a long-run equilibrium. Second, two bivariate Error Correction models (ECMs) with Generalized Autoregressive Conditional

Heteroskedasticity (GARCH) structures are estimated to examine the effect of exchange rate policy and pattern of information flows across the US and China's cotton futures markets. Finally, a Granger Causality test is employed to analyze the direction and pattern of this relationship.

***Cointegration Test***

Before the cointegration test, each individual price series needs to be tested for a unit root, that is, for an I(1) process. The Augmented Dickey Fuller test (ADF) and the Phillips-Perron Test (PP)[4] are commonly used for a unit root test. These tests are employed to assure the I(1) property for cointegration analysis thereafter.

If both of the time series data are I(1) processes, a cointegration test is employed to determine the long run relationship between two price series. Cointegration is used instead of regular regression methods because of its capacity in dealing with non-stationary variables. The most popular cointegration test method, developed by Johansen (1988) and Johansen and Juselius (1990), can be applied. This test is based on maximum likelihood estimates of the cointegrating regression system as in (1).

$$\Delta Y_t = \mu + \Phi X_t + \Pi Y_{t-1} + \Gamma_i \sum_{i=1}^{k-1} \Delta Y_{t-i} + \varepsilon_t \tag{1}$$

where $Y_t$ is the vector to be tested for cointegration; $X_t$ are some stationary variables chosen by the model; $\Phi$, $\Pi$ and $\Gamma$ are coefficient matrices; *k* is the order of the lagged differenced dependent variables. The optimal number of lags, *k,* is determined by the Akaike Information Criteria (AIC). Equation (1) is in the Error Correction form.

The Johansen likelihood ratio test is based on the rank of the coefficient matrix, Π. For a Vector Autoregressive (VAR) system with *p* dependent variables, denote the rank of the (p x p) matrix $\Pi$ by $r$, i.e., rank($\Pi$) = $r$. If $r$=0, there is no cointegration for these variables. In other words, no stationary linear combination can be identified. If $r$=$p$, $\Pi$ is of full rank and all $Y_t$ must be stationary. If $r$ is between zero and $p$, there exist $r$ stationary linear combinations. This is the case of the presence of cointegration, and $\Pi$ can be decomposed as $\Pi = \alpha\beta'$ where $\alpha$ and $\beta$ are $p \times r$ full rank matrices. The columns of $\beta$ are $r$ cointegrating vectors such that $\beta' Y_t$ is stationary. The matrix $\alpha$ can be interpreted as the speed of adjustment towards equilibrium and it shows how fast a series will revert to its equilibrium from a deviation.

Two test statistics: the trace statistic and the maximum eigenvalue statistic are commonly used for testing cointegrating vectors:

$$\lambda_{trace} = -T \sum_{i=r+1}^{p} \ln(1-\hat{\lambda}_i)$$

$$\lambda_{\max} = -T \ln(1-\hat{\lambda}_{r+1})$$

where $T$ is the number of observations and $\hat{\lambda}_i$ is the $i$-th largest squared canonical correlation between the two residuals generated by regressing $\Delta Y_t$ and $Y_{t-1}$ on $\Delta Y_{t-1}$, $\Delta Y_{t-2}, \ldots, \Delta Y_{t-(k-1)}$, respectively.

The null hypothesis is that there are at most $r$ distinct cointegrating vectors. For each of $r$ starting from zero, both of the statistics are calculated. For a bivariate time series model, that is, for $p$ = 2, there are only two cases for the value of $r$: $r$ = 0 implies no cointegration relationship, while $r$ = 1 implies one long-run relationship between the processes, i.e., they maintain the equilibrium over time. As Seo (1998) mentioned in his paper, the distribution of the cointegration rank test with stationary covariates is a mixture of the chi-squared distribution and the nonstandard distribution found by Johansen (1988, 1991). In light of this, the asymptotic critical values from Seo (1998) are chosen for the cointegration rank test, which incorporates several stationary exogenous variables.

***Error Correction Model with a Structural Change and ARCH/GARCH Effects***

On July 21, 2005, China initiated exchange rate reform and moved toward a more flexible, market-based exchange rate mechanism from the US dollar pegging system. This structural change has shown a great impact on the international trade between China and US. Such a change is also expected on cotton market prices. To capture the effect of this reform, we set up two models which include two different structure change variables respectively. Model 1 incorporates a dummy variable representing exchange rate policy, under the assumption that the exchange rate reform happened on that day in 2005, separating time into a pre-reform era and a post-reform era. Model 2 incorporates a slow structure change variable of exchange rate policy, under the assumption that the exchange rate reform started on that day in 2005 and is in progress slowly. This model is supported by the fact that the exchange rate between Chinese Yuan and US dollar was kept at the same rate prior to that date and has kept decreasing up to today.

The extended ECM from equation (1) can be expressed as:

$$\begin{pmatrix} \Delta \boldsymbol{NY_t} \\ \Delta \boldsymbol{CZ_t} \end{pmatrix} = \begin{pmatrix} \boldsymbol{\mu}_1 \\ \boldsymbol{\mu}_2 \end{pmatrix} + \mathrm{A}\begin{pmatrix} \boldsymbol{TR} \\ \boldsymbol{S*TR} \\ \boldsymbol{S} \end{pmatrix} + \mathrm{B}\begin{pmatrix} \boldsymbol{z_{t-1}} \\ \boldsymbol{S*z_{t-1}} \end{pmatrix} + \boldsymbol{C}\begin{pmatrix} \Delta \boldsymbol{NY_{t-1}} \\ \dots \\ \Delta \boldsymbol{NY_{t-(k-1)}} \end{pmatrix} + \\ \boldsymbol{D}\begin{pmatrix} \Delta \boldsymbol{NY_{t-1}*S} \\ \dots \\ \Delta \boldsymbol{NY_{t-(k-1)}*S} \end{pmatrix} + \boldsymbol{E}\begin{pmatrix} \Delta \boldsymbol{CZ_{t-1}} \\ \dots \\ \Delta \boldsymbol{CZ_{t-(k-1)}} \end{pmatrix} + \Psi\begin{pmatrix} \Delta \boldsymbol{CZ_{t-1}*S} \\ \dots \\ \Delta \boldsymbol{CZ_{t-(k-1)}*S} \end{pmatrix} + \begin{pmatrix} \boldsymbol{\varepsilon}_1 \\ \boldsymbol{\varepsilon}_2 \end{pmatrix} \tag{2}$$

In this representation, *NY* and *CZ* represent the price at NYBOT and CZCE respectively; *TR* is the weekly cotton trade volume between US and China; and $z_{t-1} = \beta' \begin{pmatrix} NY_t \\ CZ_t \end{pmatrix}$ is the error correction term with $\beta$ defined for equation (1). $\boldsymbol{A}, \boldsymbol{B}, \boldsymbol{C}, \boldsymbol{D}, \boldsymbol{E}, \boldsymbol{\Psi}$ are coefficient matrices. In model 1, $S$ is defined as a dummy variable which has the value:

$$S_t = \begin{cases} 0 & if\ t < t_1 \\ 1 & if\ t \geq t_1 \end{cases}$$

For model 2, $S$ is defined as a slow structure change variable which has the values as follows:

$$S_t = \begin{cases} 0 & if \quad t \leq t_1 \\ (t - t_1)/(t_2 - t_1) & if\ t_1 < t < t_2 \\ 1 & if \quad t = t_2 \end{cases}$$

where $t_1$ is the date when the change starts (July 21, 2005 in this case) and $t_2$ is the date when the change completes. Because the exchange rate between the Chinese Yuan and the dollar is still under the administrative control of the Chinese government, it seems the change has not been completed. As a result, $t_2$ is the date of the last observation of price data.

The mean equation (2) can capture the price transmission and structure change effect in markets and across the markets. The error terms, however, may not be independently and identically distributed. As Park and Bera (1987) pointed out, most economic and financial time series data encounter the presence of heteroskedasticity in the second moment. The Multivariate Generalized Autoregressive Conditional Heteroskedasticity (GARCH) model was developed by Bollerslev et. al (1988) to account for the Autoregressive Conditional Heteroskedasticity (ARCH) effects in the residuals of the ECM. In this paper, bivariate GARCH models are used to explore the volatility behavior of cotton futures prices.

The conditional variance and covariance equations for GARCH effect are specified as:

$$H_{NN,t} = \pi_0^N + \sum_{q=1}^{Q} \gamma_q^N \varepsilon_{N,t-q}^2 + \sum_{p=1}^{P} \delta_p^N H_{NN,t-p} \tag{3}$$

$$H_{CC,t} = \pi_0^C + \sum_{q=1}^{Q} \gamma_q^C \varepsilon_{C,t-q}^2 + \sum_{p=1}^{P} \delta_p^C H_{CC,t-p} \tag{4}$$

$$H_{NC,t} = \lambda_0 + \lambda_1 (\varepsilon_{N,t-1} \varepsilon_{C,t-1}) + \lambda_2 H_{NC,t-1} \tag{5}$$

where $\begin{pmatrix} \varepsilon_{N,T} \\ \varepsilon_{C,T} \end{pmatrix} = \varepsilon_T$ is from the equation (2), $\Omega_{t-1}$ is the information set obtained at time $t$-1 and $\varepsilon_T \mid \Omega_{t-1}$~N $(0, H_t)$, and $H_t$ is conditional variance-covariance matrix.

Estimates of the parameters in the VAR system with structural change and GARCH error terms for the cotton futures prices are obtained by maximizing the log-likelihood function. The GARCH specification is known as the diagonal VECH model, and equations (3)-(5) are defined to capture the appropriate information flows in the markets.

### *Granger Causality Test*

When a cointegration relationship is present for two variables, a Granger Causality Test (Granger, 1969) can be used to analyze the direction of this co-movement relationship. Theoretically, Variable $Y_1$ is said to Granger-cause another variable $Y_2$, if the current value of $Y_2$ is conditional on the past value of $Y_1$ ($Y_{1,t-1}$, $Y_{1,t-2}$, … $Y_{1,0}$) and thus the history of $Y_1$ is likely to help predict $Y_2$.

The following two hypotheses are tested to determine the Granger Causality relationship between two price series in this paper:

$C_{2,1} = C_{2,2} = ... = C_{2,k-1} = D_{2,1} = D_{2,2} = ... = D_{2,k-1} = B_{21} = B_{22} = 0$ (no causality from US to China).

$E_{1,1} = E_{1,2} = ... = E_{1,k-1} = \Psi_{1,1} = \Psi_{1,2} = ... = \Psi_{1,k-1} = B_{11} = B_{12} = 0$ (no causality from China to US).

where all the parameters are defined in equation (2) before.

## Data Description

The data used in this article include the nearby futures close prices obtained from the Zhengzhou Commodity Exchange (www.czce.com.cn) in China and the New York Board of Trade (www.nybot.com) in the US, covering the period from December 1, 2004 to November 30, 2007, a total of 148 weekly observations. The prices reported on every Thursday are represented as weekly prices [5] and, similar to many other studies (see Liu, 2005), the nearby futures contract prices are used here because of their high liquidity and activity. Cotton futures prices are quoted in Chinese Yuan in the CZCE and the US dollar in the NYBOT. To be consistent, the price in the CZCE is converted into dollars according to the exchange rate between these two currencies.

As mentioned earlier, since the Chinese government conducted the reform of exchange rate policy on July 21, 2005, the rate between Dollar and Yuan has been gradually decreased from the previously fixed rate. Therefore, the fixed rate of ¥8.28/$ is used for dates before July 21, 2005 and the flexible rate thereafter. The exchange rate had fallen from ¥8.28/$ to ¥7.40/$ as of the end of the observation period. After the adjustment, both prices are listed as cents per pound. In this study, the scaled cotton trade volume between the U.S. and China is also included as an important explanatory variable.

The time series plot of the cotton futures prices in the CZCE and NYBOT are given in Figure 3. Both series indicate strong nonstationarity. The sample autocorrelation functions and sample partial autocorrelation functions also suggest the presence of a unit root.

## Empirical Results

### *Cointegration Tests*

Two standard unit-root test procedures are applied to examine whether Chinese and US cotton futures prices are nonstationary. The null hypothesis for both tests is that a unit root is present, which can be rejected if the value of test statistics is smaller than the corresponding critical value. The optimal lag length of each test is determined by the SIC. Here lag $k = 2$ and $k=3$ for the original NYBOT and CZCE data respectively so 1 and 2 lagged difference terms are added for the ADF test. For the PP test, Newey-West bandwidths of 1 and 4 are chosen using the Bartlett Kernel method for each set of data.

From table 1, both of the ADF and PP tests, with an intercept and deterministic time trend included in each test equation, indicate that these two futures price series are I(1) processes, confirming the conjecture obtained from casual examination of the time series plot. Furthermore, from the time series plot, there seems to be a long run co-movement pattern for the two series. Therefore, the cointegration analysis between the two series is conducted and a bivariate ECM is explored for the cotton futures prices in the NYBOT and CZCE. If the null hypothesis of $r=0$ cannot be rejected, there is no cointegration. If $r=0$ is rejected and $r=1$ cannot be rejected, then there exists a cointegration relationship.

Table 2 gives the cointegration test results for the case where the level data have deterministic trends and the cointegration equations have intercepts indicated by the time series plot of the data. Based on the AIC, 2 lags are chosen for this test, but other choices of lag-length do not change the results qualitatively. The null hypothesis of $r=0$ is rejected and $r=1$ cannot be rejected under slow structural change case. For the 0-1 dummy variable case, $r=1$ is barely rejected at the 5% significance level, but $r=1$ cannot be rejected at the 2.5% significance level.

Using the 2.5% level, it suggests that the US and Chinese cotton futures prices are cointegrated with a cointegrating rank of one, implying a common stochastic trend that derives these two series. In other words, although each of the series is nonstationary, they move together in the same direction for the long-run and their linear relationship is stationary.

The estimated cointegrating vector is $\hat{\beta} = (1 \quad -0.67)'$, so the long-run relationship between $P_{NYBOT}$ and $P_{CZCE}$ can be represented as $P_{NYBOT}$ =0.67 $P_{CZCE}$. The reason why the long-run cotton futures price in the US is much lower than that in China can be explained by transportation costs, the customs tariff and the quota imposed in the trade. Since the cotton trade between these two countries is still not totally free, this price gap is expected to last for a long time. The estimated speed of adjustment coefficient $\hat{\alpha} = (-0.22 \quad 0.034)'$. This indicates that a one unit increase in the US price will results in a 0.22 unit decrease in its own price and a 0.034 unit increase in the Chinese price for the next period; On the other hand, a one unit increase in the Chinese price will result in a 0.15 (=(-0.22)x(-0.67)) unit increase in the US price and a 0.023 (=0.034x(-0.67)) unit decrease in its own price. In the long run, both prices move together in that if one price increases, then the other price eventually increases. In the short run, the reaction to the other country's price is in the same direction, but is opposite to its own.

***Error Correction Model with GARCH Effect***

In estimating the ECM, the appropriate lag length should be determined. Using the AIC, the two-period lagged model is chosen.

Table 3 presents the estimate results from different model specifications. For each model, the structure change terms are removed if they are statistically insignificant. Since the dummy variable method is normally used for a structure change, this model (model 1) is estimated first.

In the volatility equations, only the ARCH effect is significant in each market when assuming no cross-market effect. This result indicates inertia in adjustment to past shocks in unanticipated volatility.

For model 1, the trade volume does not have a significant effect before the exchange rate reform. But it does show some effect on China's market afterwards. This indicates the cotton trade between China and the US is not completely free of barriers, but the exchange rate reform has facilitated free trade to some extent. The current tariff and quota involved in the trade restrict the price reaction in the two markets in response to the supply changes resulting from the physical flow. However, the cointegration relationship still holds because of the vibrant market participation of traders in these two markets and the financial flow they bring into the markets. Although Chinese futures markets are not directly open to foreign investors, prominent Chinese cotton or financial companies can trade in both Chinese and US markets.

Nevertheless, model 1 is based on the assumption that the exchange rate reform takes place at only one time period. Thus it does not capture the structure change effect on cotton futures markets very well, so a slow structure change model (model 2) is introduced as an improvement. From table 3, the cointegration error term has a significant effect on each market before the policy change but it only has a different effect to the US market after the policy change. This means that the structure change of this policy has altered the speed of adjustment in the US market, but in the Chinese market, the speed of adjustment to the equilibrium remains the same and is not affected by the exchange rate policy change.

For the own-market effect, both of the two lagged terms are significant for the US market before the transition but only the second lagged term is significant during the transition. For China's market, no significant lagged term is observed during this transition. For the cross market effect, the US market is only affected by the first lagged term before the structure change, but is affected by both lagged terms afterwards.  On the other hand, the cross effect is weaker in the Chinese market and only the first lagged cross term is significant before the structure change. All of the above implies that the RMB exchange rate policy reform has much greater influence in the US market than the Chinese market.  The significance of cross effect coefficients in each market also indicates that there is a mutual price transmission effect between US and Chinese cotton markets.

Model 2 also shows the presence of volatility clustering of price change for each market. However, the volatility in the US market is significantly affected by the past shocks, but the impact of previous volatility is not seen.  For China's market, both the past shocks and past volatility have significant impacts on the current process, although the influence of the unanticipated shocks on the innovation is relatively small.

***Granger Causality Test***

The existence of a cointegration relationship between these two price series means that at least one of them Granger-causes the other (Narayan & Smyth, 2004 p.31).  It is consequently relevant to study the direction of the causality of this price relationship.

The null hypothesis for this Granger Causality Test is no causality.  From table 4, each of the hypotheses that the price in China does not Grange-Cause the price in US and the price in US does not Grange-Cause the price in China is rejected, using 10% significance level.  Therefore, there is a bi-directional causality in these two futures markets.

This result is consistent with reality. The NYBOT is the most mature cotton futures market, and its cotton price is always a critical reference for governments making trade policies and producers making production decisions around the world. On the other hand, although the CZCE only has a three-year history in trading the cotton futures contracts, and though the trading volume in this market is relatively small, China depends highly on the cotton import from the world, especially the US due to its enormous domestic consumption. With this increasing trade between these two countries and the improved futures markets, the price in the CZCE has played a more and more important role in the world's cotton pricing.

**Conclusions**

The noticeably increased cotton trade between China and the US has had a profound impact on the national economies of these countries. As the Chinese government is committed to reduce protective quotas, tariffs and subsides in the cotton international trade, China's futures markets are getting more integrated with the world market. This quantitative investigation of cross-market interactions between China and the US cotton futures markets has confirmed such a conjecture.

The conclusions of this study can be summarized as follows: the US and China's cotton futures prices are cointegrated, thus there exists a long-run equilibrium between the CZCE and NYBOT cotton futures prices. Additional result reveals that the NYBOT and CZCE share price transmission, but so far, the cotton trade volume does not contribute explaining this price relationship. The result from the estimation of AR-GARCH model indicates China's recent exchange rate policy change is affecting each of these two markets, with more influence on the US market. In terms of volatility, both markets show the presence of ARCH/GARCH effect. These results provide useful information for the cotton producers, traders, commodity exchanges and the relevant regulatory agents both in China and the US.

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## Endnotes

[1] Data source: the annual report of China Customs.
http://www.customs.gov.cn/YWStaticPage/default.htm

[2] Data source: NYBOT(www.nybot.com), 2006

[3] Data source: CZCE(www.czce.com.cn), 2006

[4] Augmented Dickey-Fuller (Dickey&Fuller 1981) test is a commonly used method with the following formulation:

$$\Delta y_t = \mu + \gamma^* y_{t-1} + \sum_{i=1}^{k-1} \phi_i \Delta y_{t-i} + \varepsilon_t$$

where $\gamma^* = \phi - 1$ from the model:

$$y_t = \mu + \phi y_{t-1} + \sum_{j=1}^{p-1} \phi_j \Delta y_{t-j} + \varepsilon_t$$

The unit root test is carried out by testing the null hypothesis $\gamma^*$ =0 vesus $\gamma^*$ <0. If we fail to reject the null hypothesis, a unit root is present.

The Phillips-Perron (PP) test has been suggested as an alternative to the ADF test to improve on the finite sample properties. Phillips and Perron (1988) proposed the unit root test of the OLS regression model: $y_t = \mu + \beta t + \alpha y_{t-1} + \varepsilon_t$. Let $s^2 = \sum_{t=1}^{T} \hat{\varepsilon}_t^2 /(T-k)$ and let $\sigma^2$ be the variance estimate of the OLS estimator $\hat{\alpha}$, where $\hat{\varepsilon}_t$ is the OLS residual. The asymptotic variance of $\sum_{t=1}^{T} \hat{\varepsilon}_t^2 / T$ using the truncation lag $i$ is: $\hat{\lambda} = \sum_{j=0}^{i} k_j * [1 - j/(l+1)]\hat{\gamma}$ where $k_0$=1, $k_j$=2 for j>0 and $\hat{\gamma}_j = \sum_{t=j+1}^{T} \hat{\varepsilon}_t \hat{\varepsilon}_{t-j}$. Then the Phillips-Perron test is $Z(\hat{\alpha}) = T(\hat{\alpha} - 1) - T^2 \hat{\sigma}^2 (\hat{\lambda} - \hat{\gamma}_0) / 2s^2$.

[5] For some specific weeks in which Thursday data is unavailable, we use Wednesday data for that week.

**Table 1. ADF and PP Unit Root Tests for the Data.**

| Test | ADF | PP | Critical value (1%) |
|---|---|---|---|
| NYBOT | -3.91 | -3.50 | -4.02 |
| CZCE | -2.84 | -2.52 | -4.02 |
| $\Delta NYBOT$ | -10.52 | -10.49 | -4.02 |
| $\Delta CZCE$ | -9.31 | -9.91 | -4.02 |

**Table 2. Bivariate Cointegration Test with Different Exogenous Variables.**

**Test under 0-1 dummy variable case**

| $H_0$: rank=r | $H_1$: rank>r | ρ * | Trace | 5% CV** | 2.5% CV** |
|---|---|---|---|---|---|
| 0 | 0 | 0.36, 0.2 | 26.65 | 11.78 | 13.69 |
| 1 | 1 | 0.35 | 6.07 | 6.04 | 7.64 |

**Test under slow structural change case**

| $H_0$: rank=r | $H_1$: rank>r | ρ * | Trace | 5% CV** | 2.5% CV** |
|---|---|---|---|---|---|
| 0 | 0 | 0.3, 0.2 | 17.84 | 11.08 | 12.92 |
| 1 | 1 | 0.29 | 4.63 | 5.25 | 6.73 |

Based on r=1, the beta vector is (normalized on $P_{NYBOT}$ ):

| $P_{NYBOT}$ | $P_{CZCE}$ |
|---|---|
| 1.00 | -0.67 |

Note. * **ρ** is the canonical correlation.

**The critical values are derived from Seo (1998). See table 3 and table 4 in Seo's paper.

**Table 3. AR-GARCH Models for Cotton Futures Prices in NYBOT and CZCE.**

| US-China market | Model 1 | | Model 2 | |
|---|---|---|---|---|
| | US | China | US | China |
| **Error Correction Model** | | | | |
| $\mu$ | 0.076 | 0.44** | -0.10 | 0.26 |
| $A_1$ (TR) | -0.023 | 0.015 | 0.015 | -0.013 |
| $A_2$ (TR*S) | 0.044 | -0.04** | 0.033 | 0.0002 |
| $B_1$ ($z_{t-1}$) | 0.38** | -0.18* | 0.52** | -0.11* |
| $B_2$ ($S*z_{t-1}$) | 0.04 | 0.15 | -0.46** | 0.14 |
| $C_1$ ($\Delta NY_{t-1}$) | -0.20* | 0.25** | -0.27* | 0.22** |
| $C_2$ ($\Delta NY_{t-2}$) | -0.19** | 0.11** | -0.36** | 0.074 |
| $D_2$ ($\Delta NY_{t-2}*S$) | N/A | N/A | 0.43* | N/A |
| $E_1$ ($\Delta CZ_{t-1}$) | 0.10 | -0.066 | 0.53** | 0.01 |
| $E_2$ ($\Delta CZ_{t-2}$) | 0.57** | 0.024 | 0.29 | 0.27* |
| $\Psi_1$ ($\Delta CZ_{t-1}*S$) | 0.22 | 0.082 | -0.71* | 0.16 |
| $\Psi_2$ ($\Delta CZ_{t-2}*S$) | -0.56** | -0.28 | -0.73* | 0.14 |
| **ARCH-GARCH Structure** | | | | |
| $\pi_0$ | 1.24** | 0.51** | 1.55 | 0.31** |
| $\gamma_1$ (ARCH) | 0.64** | 0.57** | 0.29** | 0.22** |
| $\delta_1$ (GARCH) | 0.027 | 0.028 | 0.13 | 0.52** |

Note. **indicates significance at the 0.05 level.

*indicates significance at the 0.1 level.

**Table 4. Granger Causality Wald Test.**

| Test | F-Statistic | Prob. |
|---|---|---|
| 1 | 3.16 | 0.07 |
| 2 | 4.72 | 0.03 |

| Test 1 | Test 2 |
|---|---|
| Group 1 Variables: $P_{NYBOT}$ | Group 1 Variables: $P_{CZCE}$ |
| Group 2 Variables: $P_{CZCE}$ | Group 2 Variables: $P_{NYBOT}$ |

**Figure 1. The major cotton production countries in the world in 2004-2006 (Data source: USDA, 2007).**

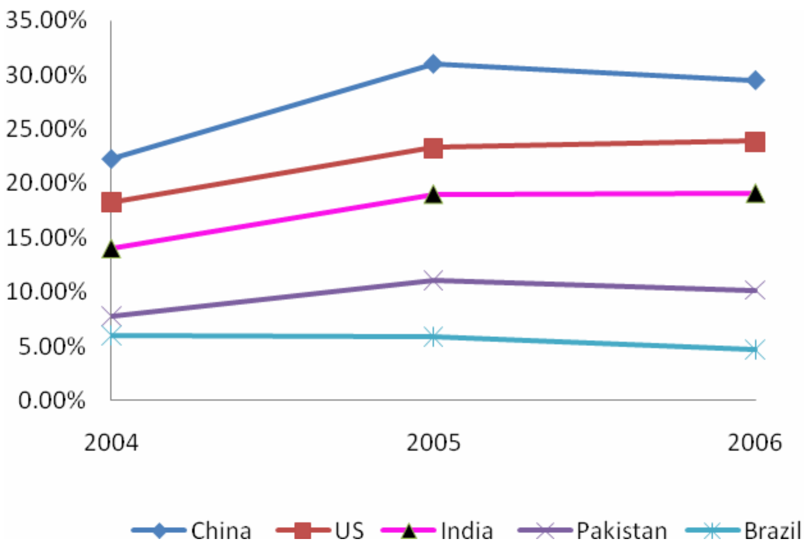

**Figure 2. The major sources of cotton export to China in 2004-2006 (Data source: the annual report of China Customs, 2007).**

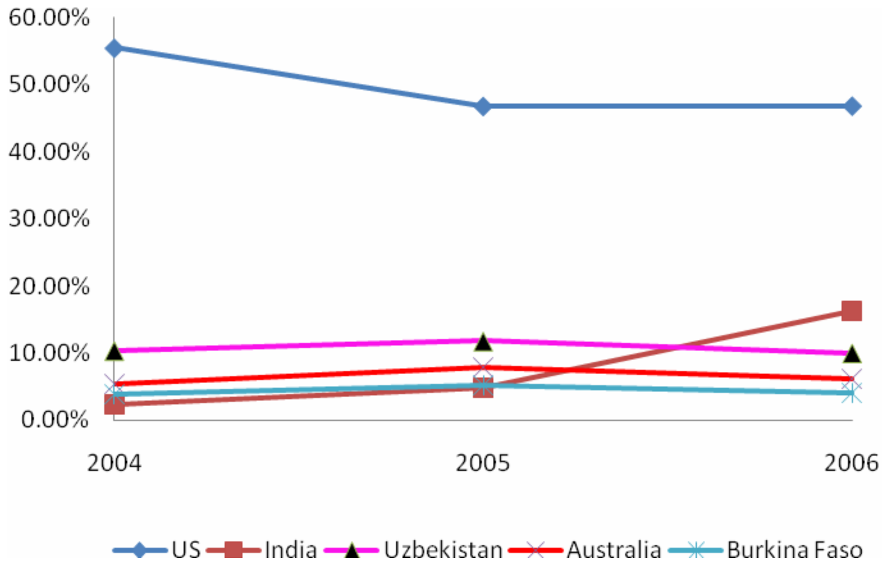

**Figure 3. Time series plot of the price series.**