TITLE: Price Linkages in World Cotton Markets and Cointegration Tests, Structural Breaks

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ABSTRACT

This paper examines recent evolution (1994/95-2005/06) of cotton price dispersion among world cotton producing countries in a market integration framework. In order to test market integration and Law of One Price (LOP), among six price series, we use Johansen et al. (2000) cointegration procedure, which permits structural breaks in time series data. The multivariate cointegration tests show that there are three cointegrating vectors in the system, which implies that though the markets are integrated, the LOP does not hold. Vector error correction (VECM) restriction tests also show that structural breaks have impacts on the long-run linkages among the prices.

Key words: Cotton, cointegration, market integration, law of one price, structural breaks.

INTRODUCTION

Cotton which, is extensively traded commodities in the international market, is the most essential staple for textile industry. The world cotton market is characterized by a limited number of major cotton producing countries. The major actors in the world cotton markets are the United States, China, EU, and some developing countries.

An indirect means of analyzing market efficiency is to test for market integration (Hopcraft 1987). Market integration refers to co-movement or a long relationship of prices. It is defined as the smooth transmission of price signals and information across separated markets. Two trading markets are assumed integrated if price changes in one market are manifested to an identical price response in the other market (Goletti et al., 1995; Barett, 1996). There are very few studies of market integration dealing with cotton markets. However, there have been numerous studies testing the law of one price in agriculture sector. In line with our research, other empirical studies investigate long-run relationship using cointegration techniques but they do not employ a cointegration test which allows a structural break. Bessler and Chen pointed out that U.S. and world cotton prices were cointegrated during 1980-84 and 1986-93. They qualified their results for 1977-84, citing a failure to reject weak exogeneity for both series during that period. In other study, MacDonald (2000) used Johansen (1998) cointegrating tests, and found that U.S. and world prices are cointegrated with the law of one price apparently violated during the 1990's. Other researchers, Baffes and Ajwad (1998) used regression analysis to test the degree of cotton price linkages of world cotton market. They found that the main source of improvements in price linkages seems to be a result of short run price transmission and to a very limited extent a result of long run comovement. The results indicate that there is relatively high long-run convergence between Central Asia and West Africa and very limited long-run convergence between the U.S. and other countries. Yang and Leatham (1999) used cointegrating analysis to cotton price relationships. In this research, they used six cotton price series included Turkish cotton prices and they found evidence for market integration among some developing countries (excluding Turkey).

The LOP as a measure of integration is especially important in agricultural markets. The LOP notion of market integration accommodates short-run price differences but requires that price changes across area markets correspond with each other on a one-to-one basis in the long run. The LOP assumes that as long as markets are cointegrated, the current price changes in one market are influenced by the current price changes in another market. Ardeni (1989) found that assumption of the law of one price was counterfactual and argued that much of the empirical support was based on econometric shortcomings such as spurious regressions, nonstationary in data, and inappropriate use of first differences. The nonstationarity test and cointegration tests for the group of commodities showed that LOP, as a long-run relationship, uniformly failed and the deviations from the pattern were permanent. Baffes(1991) tested LOP, using Engle and Granger two-step procedure for several agricultural commodities in four countries. Their empirical results supported the LOP for the specific commodities and time periods examined. The failure of the LOP as a long run relationship was a price specific and time period specific problem rather than a general failure. A possible reason for the failure of LOP was non-inclusion of transportation costs. Their results indicate that the LOP depend on the commodities, time period of data, method of analysis and whether transactions are included.

The main purpose of the paper is to test cointegration that can be used to investigate market integration and test for the LOP. Statistical developments in cointegration testing by Johansen (1988) provide a method for generating test statistics (*i.e.*, likelihood ratios) with exact limiting distributions and allow for direct testing of the LOP hypothesis (Johansen and Juselius 1990). In the other studies, Johansen et al. (2000) suggest a new cointegration procedure which permits structural breaks. We apply a time series approach based on structural change literature which allows for not only structural breaks in data to be identified but also shows how the effects price transmissions evolve over time.

Cotton Production and Trade in World

The majority of world cotton production is located in the developed world where the USA is far the largest producer. Other major national producers are China, India, Pakistan, Uzbekistan and countries in (Francophone) West Africa. Figure 1 shows production, export, consumption and ending stocks in world. World wide cotton production has increased significantly during the last decades, from 18.8 million tons in 1994/95 to 21.5 million in 2002/03 and reached 24.8 million tons in 2005/06 (FAO, 2006). Most of this growth recently was in China, where production tripled. China has long been the world's largest cotton producer and consumer, but in recent years its soaring economy and global textile trade liberalization have driven its cotton imports far beyond any other markets. As recently as marketing year 2000/01, China was a net exporter of cotton, but since then imports have risen to record highs year after year. China's textile industry has been a dominant purchaser in recent years, taking up more than a quarter of global cotton output. Other major users are the EU, India, the US and Turkey, which take half of cotton output.

World cotton production is climbing because of government measures that sustain cotton production in the U.S. and EU, China. Support to cotton producers has been greatest in the US, followed by China and the EU. Cotton subsidies encourage overproduction, which is then sold on the world market. Interventions by U.S. and China, EU have the most impact on the world cotton markets. They have high level assistance. Therefore, recently, these significantly increased share of global cotton markets. This has depressed world cotton prices, damaging those developing countries which rely on exports of cotton for a substantial component of their foreign exchange earnings.

MATERIALS and METHODS

The main data for this study were obtained from Cotton Report published by the U.S. Department of Agriculture and Izmir Mercantile Exchanges, and Cotton Outlook Report published by Cotlook Limited. Prices used for this analysis are monthly quoted CIF prices (cleared, insured and forwarded) for delivery at a Northern Europe port. Cotton price data from analysis starts from January 1994 and ends in December 2006 (156 observations). The specific spot price series include U.S (Memphis Territory) (P_{US}), Turkey (P_{TR}), Central Africa (P_{CA}), Greece (P_{GR}), West Africa (referred to as African "Franc Zone") (P_{WA}), Liverpool A index, which is the measure of the "world" price of cotton (P_{LV}). In this study, China is not included because her cotton prices are not available for the period 1994/95-20001/02. The main criteria for including these countries are their importance to world cotton export markets in terms of market shares and the consistent availability of data for chosen countries. The unit prices which were reported to different markets were converted into U. S. cent per lb equivalents using foreign exchange data obtained from the International Monetary Fund's International Financial Statistics database. As following the common practice in applying prices based tests of market integration and the LOP, the logarithms of the price data were used for the empirical analysis. In addition, all series are adjusted in terms of deterministic seasonality. Stochastic seasonality of price series is not investigated here, since effects of seasonal adjustment for series are still argued in the literature (Hatanaka, 1996).

A graphical illustration of the price data from six markets of cotton in the world is shown in Figure 2. Although there are observable differences in month to month movements, the six price series appear to share a similar pattern over a longer period of time. Notice that for all six markets, prices decline throughout much of the 1990s, and then prices in the first years of the 2000s increase until September 2000 but they return to a low in late 2000. Meanwhile, cotton prices have dropped about 50 percent from 1995 to 2002. Prices from all six markets move thereafter, generally in upward direction in early 2003, at which they decline through the remainder of the 2000s. Cotton prices decline in late 2004 as a result of record output in the major cotton production countries, which together account for more than 70 percent of the world production. World cotton prices recover in the first three months of 2005, mostly due to expectations of lower production in 2005/06 following reduced plantings in response to low prices at sowing time. Figure 2 illustrates the price series with two break times indicated by dashed vertical lines.

When investigating market integration, the first priority is to examine the time series properties of the price series. We test the series for unit roots allowing for a change in level and slope at an unknown time and without such breaks. The break point is chosen so that the absolute value of the t-statistics on the coefficient of the change in slope is maximized according to Perron (1997). As we are dealing with monthly data, a general to specific recursive procedure with initial 12 lags is used to test down the number of lags based on the significance of the t-statistic on the coefficient associated with the last lag in the estimated regression equation. According to series' properties, we employ both innovational outlier model I, which specifies a gradual change in the intercept and innovational outlier model II, which allows for a change in the both the intercept and the slope at time T_b .

Table 1 reports the results of testing for Perron (97) unit root. Results indicate that we can accept the hypothesis of a unit root for all series considered, and conclude that they are all non-stationary. In order to determine the order of integration of these series, we first differentiate the data and retest them. It conclude that they are all integrated of order one, that is, I (1). The implication from these findings is that these series can be modeled in a standard cointegrating framework.

The LOP has been used as a measure of integration in important international agricultural markets where commodities are generally of a homogenous nature across markets. There are two versions of the LOP, the "absolute" and "relative" version. The "absolute" or "short-run" version of the LOP asserts that the foreign price of a commodity, once adjusted for the exchange rate, will be equal to the domestic price of the commodity plus an adjustment for transportation costs and trade taxes. The "relative" or "weak" version of the LOP predicts the proportionate changes in the prices, and can be regarded as the test of the LOP in the long run. Compared with the absolute version, the relative version of the LOP is less restrictive. According to the relative version, common currency prices for a particular product should change in the same way over time in different countries and, therefore, the law is compatible with the existence of a stable price differential across markets. Applied researchers often base their statistical models of market integration upon the relative (or weak) version of the LOP. An econometric representation of the LOP can be explained as follows:

$$\ln P_{1t} = \alpha + \beta \ln P_{2t} + \varepsilon_t \qquad (1)$$

where p_{1t} and p_{2t} are the prices of same or homogenous commodity traded in market one and market two respectively. The weak version of the LOP holds when the restrictions $\alpha \neq 0$ and $\beta = 1$ are satisfied. The law of one price, which is the main requirement for market integration, will hold only when market prices are cointegrated (Engle and Granger 1987). Cointegration holds when the means and variances of two or more variables move together (Enders 2004). This cointegration test have been criticized on the grounds that the cointegration is confined to pairwise comparisons and such tests require that one of the two variables be designated as exogenous. It would not lend itself well to analyzing multivariate systems that characterize for example markets with many sellers and buyers. Hypothesis testing on the estimated cointegrating vector is likewise not possible under this approach. The Johansen method is preferred over the Engle and Granger approach and has proven to be popular in the recent literature on market integration and LOP. However, this procedure does not require that one of the variables be designated as exogenous.

In Johansen multivariate cointegration framework, market integration implies that in distinct locations there must be n-1 cointegrating vectors. In general, in a system with n data series and r cointegration vectors, there will be n-r different stochastic trends (Stock and Watson, 1988). However, since with *n* prices one can find at most $n-1$ cointegrating vectors, all but $n-1$ pairs are redundant. Hence, the weak version of the LOP as implied by pair-wise cointegration (i.e. when all the price series share a common stochastic trend) is a stronger proposition than the general notion of market integration as implied by the presence of at least one cointegrating vector in a multivariate system. This signifies that the number of cointegrating vectors is an important indicator of the extent of co-movement of prices. An increase in the number of cointegrating vectors implies an increase in the strength of market integration (Ghosh, 2003). The Johansen cointegration test is based on a vector auto-regression (VAR) system. Given a price vector P_t it is carried out using the following representation:

$$P_t = \mu + \sum_{i=1}^{k} A_i P_{t-i} + \varepsilon_t \qquad t = 1,2,\ldots,T \qquad (1)$$

where $P_t = \left[P_{US,t} \quad P_{TR,t} \quad P_{CA,t} \quad P_{GR,t} \quad P_{WA,t} \quad P_{LV,t}\right]'$, μ and A_i are matrices of parameters, k is the lag length, and ε_t is an error term; the VAR model does not require the specification of a casual ordering prior to estimation. Let P_t be a 6×1 vector of I(1) process with one cointegrating vector. Suppose we have q sub-sample periods in T observations with T_j - T_{j-1} observation in the j-th period, $j = 1,\ldots,q$ and $T_0 < T_1 < \ldots < T_{q-1} < T_q$. For each sub-sample period, a vector autoregressive model is chosen, so that the parameters of the stochastic components are the same, deterministic trends may change between sub-samples so that the process can have breaks. The model, which is formulated conditionally on the first k observations of each sub-sample period, can be represented in the following vector-error-correction-model (VECM) form:

$$\Delta P_t = \alpha \begin{pmatrix} \beta \\ \gamma \end{pmatrix}' \begin{pmatrix} P_{t-1} \\ tE_t \end{pmatrix} + \mu E_t + \sum_{i=1}^{k-1} \Gamma_i \Delta P_{t-i} + \sum_{i=1}^{k} \sum_{j=2}^{q} \kappa_{j,i} D_{j,t-i} + \sum_{m=1}^{d} \varphi_m W_{m,t} + \varepsilon_t \qquad (3)$$

where Δ is the difference operator; k is the number of lags; $E_t = \begin{bmatrix} E_{1t} & E_{2t} & \cdots & E_{qt} \end{bmatrix}'$ is a vector of q dummy variables with $E_{j,t} = 1$ for $T_{j-1} + k \le t \le T_j$ $(j = 1,\ldots,q)$ and zero otherwise and the first k observation of $E_{j,t}$ are set to zero; $E_{j,t}$ is the effective sample of the jth period. $D_{j,t\text{-}i}$ is an indicator dummy variable for the ith observation in the jth period; that is $D_{j,t\text{-}i} = 1$ if $t = T_{j\text{-}1+i}$ (j=2,...,q, t=...,-1,0,1,…) and zero otherwise. Intervention dummies, $W_{m,t}$ $(m = 1,\ldots,d)$, are included to render the residuals well-behaved following Hendry and Mizon (1993). The β is the cointegrating vector representing the long-run relationship and α is a vector representing the speeds of adjustment toward the long-run equilibrium. $\gamma = \begin{bmatrix} \gamma_1 & \gamma_2 & \ldots & \gamma_q \end{bmatrix}$ is a vector of q long run drift parameters. The short run parameters are μ of order $6\times q$, Γ_i of order 6×6 for $i = 1,\ldots,k-1$, $\kappa_{j,i}$ of order 6×1 for $j = 2,\ldots,q$ and $i = 1,\ldots,k$, and φ_m of order 6×1 for $m = 1,\ldots,d$. The innovations, ε_t, are assumed to be independently and identically distributed with mean zero, and symmetric and positive definite variance-covariance matrix Ω, that is $\varepsilon_t \sim iid(0,\Omega)$.

There are two models suggested in Johansen et al. (2000). The First one is named as Model A in this study, which includes broken level that is restricted in cointegration space. The second one is named Model B, which includes broken linear trend that is restricted in cointegration space. In order to test between the models, the Pantula principle (Harris and Sollis, 2003) is used to test the joint hypothesis of both rank and deterministic components (Johansen, 1992).

Given cointegrating rank, we employ two VECM restrictions on the cointegration space, using log-likelihood ratios. First, we test whether n-price series exist in cointegration space. Second, it is tested if the structural breaks imply changes in joint long-term evolution that is whether the breaks are statistically significant as in Dawson and Sanjuan (2005).

RESULTS

Using Models A and B in equation (3), we employ Johansen et al. (2000) cointegration procedure to test existence of cointegration among the price series (2000). In order to count for the largest interval between two breaks to test long-run relationship among the prices, and since they display the similar results, the break times 2001:1, 2001:3, 2001:7 and 2001:8 are excluded from the analysis. Thus, we included 2000:09 and 2002:11 in the analysis as break times. The break time 2009:9 can be related to over production in Turkey in 2000/01 season. In this season cotton production in Turkey was at the highest level in the last five years, because of increasing irrigated area and cotton production in the Southeastern Anatolian region, and good weather. Therefore, the cotton prices decreased in 2000/01 season in

Turkey. In 2000/01 world production was stable, and world stocks fell, but China accounted for more than the total world decrease. In addition, a slowing world economy cut into cotton consumption during 2000/01. As a result, the world cotton prices also began to decrease at the beginning of 2000/01. The extremely low world prices of 2001/02 resulted from the same pressures that reduced prices in five of the six years after 1994/95, especially the continued disposal of surplus stocks by China. In addition, the low world prices can be associated with very good weather in nearly all of the world's major cotton-producing countries, resulting in a record world yield, and the effects of recession, especially the crisis in U.S. consumer confidence following the terrorist attacks of September 11, 2001 (USTR, 2003). In 2002/03, world production fell sharply, as planted area, responded to lower prices and weather conditions, returned to normal. Then U.S. cotton production decreased about 30 percent. Therefore, U.S. cotton prices increased in this season. This situation took place as a breaking level in U.S. prices in 2002:11. Table 2 presents cointegration rank test results. Three cointegrating vectors $r = 3$ are found for Model B, which includes a change in level and trend in the long-run. The leg-length selected by adopting the minimum value of Schwarz criterion and $k = 2$ here. To produce well-behaved residuals, intervention dummies are included for 1994:12, 1995:8, 1995:9, and 2000:01. Multivariate normality statistics for skewness, kurtosis and joint are 1.762 (*p-value*=0.296); 3.674 (*p-value*=0.072) and 16.136 (*p-value*=0.231)**.** These tests imply that residuals are well-behaved and the model is well-specified.

Since the number of prices in the VAR is six, it is expected that there are five cointegrating vectors in the system for validity of weak-version of the LOP. If all the series contain the same stochastic trend, they are pair-wise cointegrated. According to the cointegration results, however, international cotton markets are integrated, but the LOP does not hold. Trace statistics reveal three cointegrating vectors and three common stochastic trends. Therefore, the prices are not pair-wise cointegrated and the LOP does not hold.

Table 3 reports the log-likelihood test results. According to Table 3, in all cases null hypotheses of individual exclusions of the price series are rejected. Hence, we can say that each price maintains a long-run link in the model, and they are I(1). Table 3 also shows that results of testing for existence of breaks in the long-run equilibrium are $\mu_1 = \mu_2$ and $\mu_2 = \mu_3$. The null hypotheses are rejected and, the breaks in 2000:9 and 2002:11 are statistically significant. Therefore, it can be said that periods in the long-run relationships are different significantly.

CONCLUSIONS

This study is motivated by recent changes in the world cotton markets. The study investigates two aspects of major cotton markets over the world. First, we examine the existence of market integration. Second, we investigate whether the LOP holds among the major cotton markets.

In order to test market integration, we used Johansen et al. (2000) cointegration framework for monthly cotton prices. Using the Perron (1997) test, we examined stationarity properties of price series, and unit root test results indicate that the price series are non-stationary with structural breaks. The cointegration results provide evidence of a long-run relationship among wheat prices. Finding of only three cointegrating vector implies that major cotton markets are integrated but LOP does not hold. Further the VECM restrictions tests also show that, all prices exist in cointegration spaces, and structural breaks included in long-run relationships are significant. Therefore, structural breaks have impacts on long-run links among the cotton prices.

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Table 1. Tests of hypotheses of nonstationary on cotton price from the six markets

Model	*Price*	*Test Statistics*	*Lags*	*Break Point*
Changing Level	P_{US}	**-5.085**	**12**	**2002:11**
	P_{TR}	**-5.283**	**1**	**2002:12**
	P_{CA}	**-3.888**	**13**	**2002:11**
	P_{GR}	**-5.488**	**10**	**2002:10**
	P_{WA}	**-5.210**	**1**	**2002:10**
	P_{LV}	**-4.839**	**10**	**2003:05**
Changing Level and Trend	P_{US}	**-5.282**	**12**	**2001:07**
	P_{TR}	**-4.908**	**1**	**2000:09**
	P_{CA}	**-4.783**	**8**	**2001:01**
	P_{GR}	**-4.761**	**12**	**2001:08**
	P_{WA}	**-4.975**	**12**	**2001:03**
	P_{LV}	**-4.645**	**10**	**2000:10**

Critical values at the 5% significance level: changing level model: -5.05; and changing level and trend model: -5.19 (Perron, 1997).

Table 2 Trace test statistics on number of cointegrating vectors (r)

H_0	*Model A: with intervention dummies* $W_{m,t}$	*Model B: with intervention dummies* $W_{m,t}$
$r = 0$	**213.838 (132.777)***	**277.189 (172.183)**
$r \leq 1$	**134.108 (101.690)**	**188.232 (134.506)**
$r \leq 2$	**87.073 (74.464)**	**111.529 (100.870)**
$r \leq 3$	**53.516 (51.177)**	**67.975 (71.351) §**
$r \leq 4$	**29.880 (31.677)**	**35.520 (45.742)**
$r \leq 5$	**8.538 (15.865)**	**13.201 (23.705)**

*** Critical values in parentheses at 95% confidence level and they can be approximated by Gamma distribution explained in Johansen et al. (2000).**

§ Denotes that the null is not rejected in Models A and B without intervention dummies $(W_{m,t})$ using the Pantula principle.

Table 3. Test of hypotheses on the cointegration space

Null Hypothesis	H_0	***LR-Statistics***
Individual exclusion of:		
$P_{US,t}$	$\beta_{P_{US}} = 0$	**45.464 (0.000)**
$P_{TR,t}$	$\beta_{P_{TR}} = 0$	**62.432 (0.001)**
$P_{CA,t}$	$\beta_{P_{CA}} = 0$	**30.149 (0.000)**
$P_{GR,t}$	$\beta_{P_{GR}} = 0$	**77.252 (0.000)**
$P_{WA,t}$	$\beta_{P_{WA}} = 0$	**44.649 (0.000)**
$P_{LV,t}$	$\beta_{P_{LV}} = 0$	**18.919 (0.004)**
Breaks in long-run equilibrium:		
2000:9	$\mu_1 = \mu_2$	**33.722 (0.000)**
2002:11	$\mu_2 = \mu_3$	**31.146 (0.000)**

Figure 1. World cotton production and trade

Figure 2. Plots of historical cotton prices in the world markets, 1994-2006