TITLE: Cotton Market Efficiency: Turkey, the U.S., and Greece

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ABSTRACT

In this paper, we investigate the validity of the Law of One Price (LOP) to analyze the relative efficiency of the U.S., Greece and Turkish cotton markets via price performance of the products. Johansen et al. (2000) cointegration procedure which allows structural breaks in time series data is utilized to estimate whether the market efficiency mutually between Turkey's cotton market and the U.S.'s and Greece's cotton markets exists or not. Empirical results provide evidences in favor of the LOP and market efficiency.

Keywords: cotton, law of one price, cointegration, structural breaks, vector error correction.

INTRODUCTION

The LOP plays an important role in the theory of international economics. As the world economy becomes more integrated, markets in different regions are also become incorporated through international trade, foreign direct investment, and portfolio investment. In the last two decades, the trend of global market integration has intensified with rapidly growing communication and decreasing transportation costs. Since spatial market integration allows business firms to treat separated areas as one market instead of several disparate markets, increased international trade between developing and industrialized countries has assimilated geographically separate commodity markets. Fama (1976) stated that an asset market is said to be (weak-form) efficient if prices fully and instantly reflect all available information. Since the level and formation of the prices affect the production decisions, market efficiency is significant in commodity trading. In the presence of transactions, for homogenous or very similar commodities which are traded on different markets efficient market hypothesis states that, the LOP should apply to the actual transaction prices among different markets. Since the transaction prices are the prevailing market value, the buyers will not purchase at higher than the market value of commodity and neither will traders give their commodity away at a lower than the market value price. In both case, moving away from the prevailing prices would be a non-rational activity. In a perfectly efficient market the convergence to the one price is assumed to be instantaneous.

In order to test market efficiency, the cointegration approach mostly has been used in the literature. Baillie and Bollerslve (1989), MacDonald and Taylor (1989), Copeland (1991), Lajaunie and Naka (1992) and Lajaunie et al. (1996) researchers test the efficiency for the foreign exchange market. However, their findings are mixed. For the major Asian stock markets Chan et al. (1992a) tests the market efficiency and they conclude that the stock prices in these markets are efficient. With respect to the commodity market, MacDonald and Taylor (1988) find that markets of three primary metals quoted on the London Metal Exchange are efficient. Yang and Leatham (1998) examined the efficiency of the U.S. grain market, and found efficiency using Johansen cointegration procedure.

Little empirical work has been done in the past to test efficiency of agricultural commodity spot markets and no work has been done for cotton market efficiency in Turkey. Here we concentrate on the LOP mutually between Turkey's cotton market and the U.S.'s and Greece's cotton markets as proxy measure of efficiency of cotton markets since Turkey is the main cotton importer of the U.S and Greece. Here our aim is twofold. First, we test that the LOP holds between Turkish and the U.S., and Turkish and Greek cotton markets respectively in the long-run using the cointegration approach. Second, we test that Turkish cotton prices to see whether it tends to converge to the U.S. and Greek prices in an efficient cotton market.

Cotton sector in Turkey

Turkish cotton sector can be analyzed in three periods. Until 1985/86 Turkey was a net exporter country with no cotton imports. From 1986/87 to 1991/92 cotton imports increased while Turkey was still exporting country. However, since 1992/93 marketing year Turkey has become a net importer country. Cotton production increased from 1950 to 1973 due to the investment in textile sector. Following the first oil crisis, turkey experienced some serious economic difficulties. In this period inflation was very high and there was a scarcity in foreign exchange. In 1980 import substitution policy was left, and Turkey witnessed a significant increase in exports, but price support policy was not efficient enough to increase the cotton production, because the procurement prices were not high for producers. In mid-1980s growing body of textile industry caused an increase in cotton needs. Thus, cotton import took place as an alternative. Price support policy was not employed between the periods 1989/90-1990/91, and there were not any restrictions on the exports and imports of cotton. Therefore, producers were open to world competition. In early 1990s Turkish Government reapplied the support price mechanism because of the deterioration of world cotton prices. This mechanism, however, could not achieve an increase in cotton production. Because the level of support prices were inadequate as compared with the other countries' support prices, and so were procurements of agricultural sales cooperatives. In 1993/94 period Turkish Government decided to introduce the direct income support policy in the form of price premium. The upward price developments during that period encouraged the government to leave this system and in the next four seasons no support was given to farmers. As a result exports from Turkey decreased significantly and Turkey became a net importer country. Since high inflation continued during 1990, the production costs were very high. The premium payments from the period 1998/99 to 2001/02 were far from remedying the problem faced by farmers (Gazanfer, 2002). Today the world cotton prices are still below the cost of productions, and some countries continue to subsidy their cotton producers heavily. Thus, Turkey has become the second biggest cotton importer country after Republic of China. As we can see from Figure 1, the cotton imports of Turkey have increased year after year, and in last decade the U.S. and Greece have become significant import partners. Figure 1 illustrates market shares of Turkey's main cotton exporter countries.

MATERIALS and METHODS

The data we use in this study are monthly Turkish $\left(P_{TR,t}\right)$, the U.S. $\left(P_{US,t}\right)$, and Greek $\left(P_{GR,t}\right)$ cotton spot prices (c.i.f.) obtained from Izmir Mercantile Exchange and Cotton Outlook Report published by Cotlook Limited covering the period January 1994 to September 2005. All prices are defined as cent per lb and used in logarithms.

First, we tested deterministic seasonality of price series and we could not find any evidence in favor of deterministic seasonality; stochastic seasonality of price series is not investigated here.[1] Second, we test both non-stationarity and the existence of structural breaks in each price series using Perron's (1997) unit root test procedure. Table 1 shows that the null of a unit root is not rejected for $P_{US,t}$ for the changing level, and for the others for the changing level and trend model.[2] Hence, we can say that all series are non-stationary. We also observe from Table 1 that breaks occur for $P_{TR,t}$ in October 2000; for $P_{GR,t}$ in August 2001; and for $P_{US,t}$ in November 2002. Since irrigated area and cotton production continued to increase in the Southeastern Anatolian region, cotton production in Turkey was at the highest level in the last five years in 2000/01 season. Therefore, cotton prices in Turkey decreased in this season, which can be associated with structural break in 2001:10. According to European Union regulations and the production quota for Greece there were 1,031,000 MT seed of cotton in 2001/02 season. Over this amount the target prices reduced for production. However, amount of reported production was 1,341,000 MT (FAO, 2006). Due to excessive production; Greek farmers were not qualified for full target price. In addition Greece's usual market in Turkey turned to the U.S. sources. Break time 2001:8 for Greece may be explained by these reasons. Since world production dramatically decreased in 2002/03 season, as planted area to responded to lower prices in 2001/02 and weather conditions returned to normal, the U.S. cotton production decreased about 30 percent. This situation took place as a breaking level in the U.S. prices in 2002:11. Figure 2 illustrates the time graph of the series with break times indicated by dashed vertical lines. Although, there are observable differences in month to month movements, the three price series appear to share a similar pattern over a longer period of time.

The estimation of price cointegration is based on the theory of LOP, which states that the prices of a same or homogenous commodity traded in two markets are equal. The LOP in the basic simple bivariate form is tested by estimating:

$$P_{it} = \beta_1 + \beta_2 P_{jt} + \varepsilon_t \tag{1}$$

where P_{it} and P_{jt} are the logarithms of the prices in the market i and j respectively. The strict version of the LOP requires the satisfaction of restrictions $\beta_1 = 0$ and $\beta_2 = 1$. The weak version of the LOP holds when the restrictions $\beta_1 \neq 0$ and $\beta_2 = 1$ are satisfied. Estimating an equation of the above form, however, could lead to spurious regression problems, since regressions of price levels necessitate that the stationarity of the price series. Davidson and Mackinnon (1993) showed some of the consequences of spurious regressions in econometrics induced by the time-series properties of the variables analyzed. They argued that many variables are non-stationary in their levels and appear to be near-random walks. Non-stationarity would make classical asymptotic theory inapplicable and the usual estimation procedures would be invalidated. In this situation, the method of co-integration can provide a better way to test the LOP. The most common approach is Johansen and Juselius (1990)'s ML method of co-integration to test weak version of the LOP. However, since structural breaks occur in economic time series data, this method is not applicable. Therefore, we employ the other cointegration procedure suggested by Johansen et al. (2000) as explained below.

Consider a VAR model describing the relationship between the two variables:

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

Where $P_t = [P_{1t} \quad P_{2t}]'$, 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 2×1 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:

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

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 $T_{j-1} + k \le t \le T_j$ $(j = 1,...,q)$ 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 *j*th period. $D_{j,t-i}$ is an indicator dummy variable for the *i*th observation in the *j*th period; that is $D_{j,t-i} = 1$ if $t = T_{j-1+i}$ (*j*=2,...,*q*, *t*=...,-1,0,1,…) and zero otherwise. Intervention dummies,

$W_{m,t}$ $(m = 1,...,d)$, are included to render the residuals well-behaved following Hendry and Mizon (1993). The β is the co-integrating vector representing the long-run relationship and α is a vector representing the speeds of adjustment toward the long-run equilibrium. $\gamma = \begin{bmatrix} \gamma_1 & \gamma_2 & \dots & \gamma_q \end{bmatrix}$ is a vector of q long run drift parameters. The short run parameters are μ of order $(2 \times q)$, Γ_i of order (2×2) for $i = 1, \dots, k-1$, $\kappa_{j,i}$ of order (2×1) for $j = 2, \dots, q$ and $i = 1, \dots, k$, and φ_m of order (2×1) for $m = 1, \dots, d$. 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 co-integration space. The second one is named as Model B, which includes broken linear trend that is restricted in co-integration 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 the cointegration rank, further restrictions on the cointegration space can be tested using log-likelihood ratios (LR). Harris and Sollis (2003) consider these tests in the standard framework but they are extended here. Assume $r = 1$ and $q = 3$ so that

$$\begin{pmatrix} P_{t-1} \\ tE_t \end{pmatrix} = \begin{pmatrix} P_{1t-1} & P_{2t-1} & tE_{1t} & tE_{2t} & tE_{3t} \end{pmatrix}', \begin{pmatrix} \beta \\ \gamma \end{pmatrix}' = \begin{pmatrix} \beta_1 & \beta_2 & \gamma_1 & \gamma_2 & \gamma_3 \end{pmatrix} \text{ and } \alpha = \begin{pmatrix} \alpha_1 \\ \alpha_2 \end{pmatrix}.$$

First, we test whether each price exists in the co-integration space. The hypothesis of individual exclusion of P_{1t} for example is:

$$H_0 = \begin{pmatrix}\beta \\ \gamma\end{pmatrix}' = (0 \quad * \quad * \quad * \quad *) \text{ or } \qquad \beta_{P_1} = 0 \tag{3}$$

where '*' denote unrestricted parameters, and $LR \sim \chi^2$. Second, we test if the structural breaks imply changes in joint long-term evolution that is whether the breaks are statistically significant. Testing the first break for example, the null is that the intercepts or trends are equal in the first and second periods, that is:

$$H_0 : \begin{pmatrix}\beta \\ \gamma\end{pmatrix}' = (* \quad * \quad 1 \quad 1 \quad *) \text{ or } \qquad \gamma_1 = \gamma_2 \tag{4}$$

and $LR \sim \chi^2$. Third, we test that a 1% change in one variable leads to a change in the other. Here, the null is:

$$H_0 : \begin{pmatrix}\beta \\ \gamma\end{pmatrix}' = (1 \quad -1 \quad * \quad * \quad *) \text{ or } \quad \beta_1 = -\beta_2 \tag{5}$$

and $LR \sim \chi^2$. Fourth, we test for weak exogeneity. To test for the weak exogeneity of P_{1t} for example, the null is:

$$H_0 : \alpha_{P_1} = 0 \tag{6}$$

and $LR \sim \chi^2$. If the null that $\alpha_{P_1} = 0$ is not rejected and $\alpha_{P_2} = 0$ is rejected, we can say that P_{1t} is weak exogenous and P_{2t} is endogenous. In this case, ΔP_{1t} does not react to disequilibrium errors but still may react to lagged ΔP_{2t} (Dawson and Sanjuan 2005).

RESULTS

Using Models A and B in equation (3), we test for cointegration between $P_{TR,t}$ and $P_{US,t}$, and $P_{TR,t}$ and $P_{GR,t}$ following Johansen et al. (2000) which allows for two breaks. Table 3 presents the results of the cointegration rank tests. One cointegration vector is found $(r = 1)$ between Turkey and the U.S., and Turkey and Greece for Model A, which includes change in the level in the long-run. The minimum value of the Schwarz criterion is adopted to select optimum lag-length and in both cases $k = 3$. To produce well behaved residuals, intervention dummies $(W_{m,t})$ are included in Model A and Model B.[3]

We can observe from Table 3 that both of Turkish and the U.S. prices exist in cointegration space so do Turkish and Greek prices, because the null off individual exclusions are rejected. This implies that stationarity comes from a linear combination of $P_{TR,t}$ and $P_{US,t}$, and $P_{TR,t}$ and $P_{GR,t}$ with broken level of a linear trend. Table 3 also shows the results of test for the existence of breaks in the long-run, and we can say that the break in 2000:10 is significant but we can not say same thing for the break in 2002:11 for Turkish and the U.S. prices. For Turkish and Greek prices both breaks are not significant. In both case the null of LOP is not rejected. The null of $\alpha_{TR} = 0$ implying weak exogeneity of $P_{TR,t}$ is rejected for both relationship while $\alpha_{US} = 0$ and $\alpha_{GR} = 0$ are not rejected at 99% confidence level. Thus, we can say the U.S. and Greek cotton markets are the price leader in the long-run and Turkish cotton market is the follower at 99% confidence level. Estimates of adjustment coefficients for Turkish cotton market are $\alpha_{TR} = -0.287$ and $\alpha_{TR} = -0.363$. These imply that 29% any disequilibrium is removed each month by adjustments in $P_{TR,t}$, and 36%. Observing that the $P_{TR,t}$ is endogenous for both relationships, we normalize the cointegrating vectors on $P_{TR,t}$:

$$\begin{aligned} P_{TR,t} &= 0.961 * P_{US,t} + 0.139 * E_{1t} + 0.256 * E_{2t} + 0.222 * E_{3t} \\ P_{TR,t} &= 0.907 * P_{GR,t} + 0.465 * E_{1t} + 0.425 * E_{2t} + 0.481 * E_{3t} \end{aligned} \tag{7}$$

Since cointegrating relationships in equation (7) are identified, β_{US} and β_{GR} are the long-run (price transmissions) elasticities (Johansen, 2002). Thus, they can be interpreted as; 1%

increase in $P_{US,t}$ causes 0.96% increase in $P_{TR,t}$, and 1% increase in $P_{GR,t}$ causes 0.91% increase in $P_{TR,t}$.

CONCLUSIONS

In this paper we address the issue of price behaviors between Turkish and the U.S., and Turkish and Greek cotton markets. The market efficiency is investigated by examining empirical validity of the LOP. Time series analysis is performed on three separate data sets for cotton prices. Using the Perron (1997) unit root tests, stationarity properties of the prices are examined and unit root test results indicate that the price series are non-stationary with structural breaks. The first break time 2000:10 is associated with over production in Turkey in 2000/01. The second break time 2001:8 is related with excessive production in Greece. The last break time 2002:11 is also related with decreasing planted area in the U.S. Johansen et al. (2000) cointegration tests suggest that there is a long-run relationship between Turkish and the U.S. and Turkish and Greek cotton prices. The weak exogeneity tests on the cointegrated prices reveal that the U.S. and Greek cotton markets are short-run efficient for the Turkish cotton markets. According the VECM restrictions tests, all the cotton prices exist in cointegration space and structural break in 2000:10 is significant for Turkey and the U.S. relationship, while the other is not. However, for Turkey and Greece relationship the structural breaks do not have any impact on long-run link.

Cointegration is the necessary but sufficient condition for market efficiency. The long run efficiency in integrated commodity trading also requires the validity of weak-version of the LOP, namely $\beta_2 = 1$ in equation (1). Test of restrictions on the cointegrating vector suggest that the cotton markets are efficient since unit slope restriction cannot be rejected for both cointegration relationships. These findings imply that prices fully and instantly reflect all available information between Turkey’s cotton market and the U.S.’s and Greece’s cotton markets.

NOTES

1. There should be a doubt, if seasonal variations are properly modeled by the (complex) unit roots. Deviations of seasonal fluctuations from the deterministic periodicity are bounded in probability at any time, and the (complex) unit roots are inappropriate to model the deviations (Hatanaka, 1996).

2. Perron (1997) unit root test procedure is applied by using Perron97 source for RATS package which obtained from www.estima.com.

3. Multivariate normality statistics for skewness, kurtosis and joint are 1.161 (*p-value*=0.343); 3.526 (*p-value*=0.061) and 15.181 (*p-value*=0.162) for Turkey and the U.S., and 1.122 (*p-value*=0.227); 3.786 (*p-value*=0.057) and 15.155 (*p-value*=0.175) for Turkey and Greece respectively. These tests imply that both models are well-specified.

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Table 1 Perron (1997) unit root tests

Model	*Price*	*Test Statistics*	*Lags*	*Break Date*
	$P_{TR,t}$	**-5.305**	**1**	**2002:12**
Changing Level	$P_{US,t}$	**-5.039**	**12**	**2002:11**
	$P_{GR,t}$	**-5.146**	**0**	**2002:10**
	$P_{TR,t}$	**-4.343**	**11**	**2000:10**
Changing Level and Trend	$P_{US,t}$	**-5.377**	**12**	**2001:7**
	$P_{GR,t}$	**-5.120**	**12**	**2001:8**

Critical values at the 5% significance level: changing level model: -5.0; and changing level and trend model: -5.19 (Perron, 1997).

Table 2 Trace statistics

$H_0\ (H_1)$	*Model A:*	*Model B:*
$TR-US$		
$r=0\ (r\geq 1)$	**50.917 (31.525)***	**58.742 (15.175)**
$r=1\ (r\geq 2)$	**8.455 (15.760)§**	**11.059 (23.395)**
$TR-GR$		
$r=0\ (r\geq 1)$	**53.856 (30.476)**	**65.884 (42.734)**
$r=1\ (r\geq 2)$	**10.839 (15.126)§**	**14.823 (21.984)**

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

§ Denotes the first that the null is not rejected in Models A and B with intervention dummies $(W_{m,t})$ using the Pantula principle. Intervention dummies are included for 1994:11, 1995:1, 1998:6, 2000:1, 2001:2 and 2001:6 for $TR-US$, and 1994:11, 1995:1, 1998:6, 1998:10, 2000:1, 2001:2 and 2001:6 for $TR-GR$.

Table 3 Likelihood ratio statistics

Null hypothesis	H_0	***LR-statistics***
$TR-US$		
Individual exclusion of:		
$P_{TR,t}$	$\beta_{TR}=0$	**33.138 (0.000)***
$P_{US,t}$	$\beta_{US}=0$	**30.419 (0.000)**
Break in long-run equilibrium:		
2000:10	$\mu_1=\mu_2$	**4.390 (0.036)**
2002:11	$\mu_2=\mu_3$	**0.497 (0.481)**
Law of one price	$\beta_{TR}=-\beta_{US}$	**0.215 (0.642)**
Weak exogeneity of:		
$P_{TR,t}$	$\alpha_{TR}=0$	**20.235 (0.000)**
$P_{US,t}$	$\alpha_{US}=0$	**6.169 (0.046)**
$TR-GR$		
Individual exclusion of:		
$P_{TR,t}$	$\beta_{TR}=0$	**42.418 (0.000)**
$P_{GR,t}$	$\beta_{GR}=0$	**41.206 (0.000)**
Break in long-run equilibrium:		
2000:10	$\mu_1=\mu_2$	**0.577 (0.447)**
2001:8	$\mu_2=\mu_3$	**0.449 (0.503)**
Law of one price	$\beta_{TR}=-\beta_{GR}$	**3.192 (0.074)**
Weak exogeneity of:		
$P_{TR,t}$	$\alpha_{TR}=0$	**21.659 (0.000)**
$P_{GR,t}$	$\alpha_{GR}=0$	**4.704 (0.031)**

*** *P-values* are in parentheses.**

Figure 1 Market shares of Turkey’s main cotton exporter countries

Figure 2 Time graph of cotton prices