COTTON, POLYESTER AND OIL PRICES

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Abstract

This article analyzes the relationship between the spot prices of the two major textile fibers -cotton and polyester- and the spot price of crude oil over the period 1991-2008. No stable long-run relationship exists between cotton, polyester and oil prices, or between cotton and polyester prices. Monthly changes in cotton prices are positively related to past changes in cotton prices, but they are not significantly related to past changes in polyester or oil prices. Monthly changes in polyester prices are positively related to past changes in polyester and oil prices, but they are not significantly related to past changes in cotton prices. Monthly changes in oil prices are not significantly related to past changes in polyester or cotton prices. Price shocks are rapidly propagated, and unexpected changes in cotton and polyester prices tend to be self-perpetuating. A shock in oil prices is rapidly propagated to polyester prices, but only a fraction of the shock is transmitted. Short-term price relationships are stable throughout the entire sample.

Introduction

The market for textile fibers underwent substantial transformations over the last five decades: it shifted from a mainly natural fibers market (cotton being the major natural fiber) to a mainly synthetic fibers market (polyester being the main synthetic fiber). The market share of synthetic fibers increased from less than one-third of total world textile fiber use in the 1960s to more than half in the 2000s (ICAC 2009). Scale economies in synthetic fiber production, declining real energy prices, government measures supporting capacity expansion, development of new traits for synthetic fibers, and development of new blending technologies are usually cited among the causes of these structural changes. Not only have the shares of the world textile market shifted through time, but also the size of the total market has more than doubled. Higher income per capita, increased population and lower real textile prices are among the most relevant factors influencing the expansion of the textile fiber market. One important consequence of these phenomena is that while cotton prices used to be the benchmark for all other textile fibers, cotton prices now share that position with polyester prices. Therefore, while textile fiber prices used to be benchmarked to an agricultural commodity whose prices depended on seasonal supply and demand, textile fiber prices are now benchmarked to an agricultural commodity *and* an industrial product, whose production process is continuous and subject to substantial scale economies. Another level of complexity is added to the analysis of textile fiber prices when considering that polyester is obtained from polymers (paraxylene and monoethylene glycol mainly), which are products refined from crude oil, another commodity. So not only are cotton and polyester prices critical inputs to the decision making process of the textile sector, but the price of oil is also sometimes used by companies as a leading indicator of polyester prices.

Previous studies provide mixed conclusions regarding the price relationship between textile fibers and oil. Furthermore, due to structural shifts in textile fiber markets, current price relationships might differ from the previously analyzed relationships. A report by the Food and Agriculture Organization of the United Nations (FAO 2002) analyzes the statistical relationships between oil and cotton spot prices, and between oil and polypropylene spot prices, and concludes that, at most, only weak links exist between oil prices and textile fiber prices. The FAO report applies two alternative methods to analyze these relationships: a cointegration analysis to test for the existence of any stable long-run relationship between oil and fiber prices (the Johansen (1998) method), and a dynamic structural econometric model to test whether oil prices add explanatory power to the proposed model. The analyses are alternatively run on monthly data from January 1980 to December 1999, and on quarterly data from 1977(Q1) to 2000(Q4). Cointegration results indicate that no stable long-run relationship exists between oil and cotton prices, or between oil and polypropylene prices. Analyses of short-run relationships with quarterly prices indicate that changes in oil prices affect cotton and polypropylene prices, but the magnitude of the effects are small and oil price changes are not fully propagated to fiber prices. Short-run analyses with monthly data indicate that only cotton prices adjust to changes in oil prices, but also slowly and only partially. Results from the structural model (fiber prices related to its recent history and levels of demand and stocks) indicate that oil prices do not increase the explanatory power of the model for cotton. The structural model could not be estimated for polypropylene due to a lack of quantity data. The price data used in the FAO article are U.S. cotton prices, Western Europe polypropylene prices and Dubai crude oil prices, deflated by the U.S. consumer price index.

Baffes and Gohou (2005), using monthly data between 1980 and 2002, found a strong co-movement between cotton and polyester spot prices, and a significant effect of oil spot prices in polyester spot prices. The study also found that changes in the price of polyester are more rapidly transmitted to cotton prices than vice-versa: about half of the effects of a shock in the polyester market is transmitted to the cotton market within a 5-month period, while a similar transmission would take about 22 months in the other direction. The Baffes and Gohou article suggests that the difference in the speed of adjustment is a consequence of cotton being a primary commodity subject to both demand and supply shocks and polyester being an industrial product subject mainly to demand shocks; and that while cotton prices are determined in the futures market, polyester prices are determined through contractual agreements. Two caveats apply to Baffes and Gohou: (a) while the price of cotton refers to a world index of prices (the Cotlook A Index), polyester and oil prices refer to indexes of prices in the United States, whose participation in the supply and use of polyester declined significantly over the entire sample; (2) the hypothesis of stability of the long-run relationship between cotton and polyester prices cannot be rejected for the ratio of cotton and polyester prices, but this does not necessarily mean that the long-run relationship between prices in levels is stationary.[1] The authors suggest that future research use the Johansen (1998) method to test for multiple cointegrating vectors among cotton, polyester and oil prices, since the inclusion of oil prices in the analysis reduces the significance of the cotton-polyester relationship, and only increases the overall explanatory power of the model marginally.

Baffes (2007) analyzes the contemporaneous relationship between oil spot prices and the spot prices of other commodities with annual data for 1960-2005. The elasticity of cotton price with respect to oil price is found to be 14%, and the elasticity with respect to inflation 89%. Cotton prices are measured by the Cotlook A Index, oil prices are a world average of spot prices, and inflation is represented by the manufacture unit value index. According to Augmented Dickey-Fuller and Phillips-Perron tests, residuals for the cotton price equation are stationary, indicating that the long-run relationship between annual-average oil and cotton prices is stable.[2]

Fadiga and Misra (2007), using annual world spot prices for cotton, polyester and oil deflated by the World Bank manufacture import unit value for the period 1960-2004 in a multivariate unobserved component model, found that (a) no long-run relationship between cotton and polyester prices exists, (b) no short-term direct relationship between cotton and polyester prices exists, (c) polyester prices respond to past polyester prices and oil prices, and that (d) cotton prices depend on past cotton prices and changes in cotton stocks. However, Fadiga and Misra also found that cotton and polyester prices have synchronous short-term cycles (i.e. the correlation between their cycle disturbances is high), so oil prices indirectly affect cotton prices (through polyester prices) in the short run.

Harri, Nalley and Hudson (2009), using monthly futures prices and cointegration techniques, found that between June 2004 and September 2008 a stable long-run relationship existed between crude oil and cotton future prices.

The objective of the present article is to explore the relationship between oil, polyester and cotton world spot prices and identify regularities that can be used for price risk management in the cotton and textile sectors. The econometric analysis uses world prices (as opposed to U.S. prices used in previous studies), and follows a general-to-specific modeling approach, testing for, rather than imposing, a priori restrictions on the parameters of the model. The general-to-specific econometric approach consists of testing for the existence of stable long-run and/or short-run relationships between polyester, oil and cotton prices, applying Granger causality tests, and estimating the speed of adjustment of each price to changes in other prices.

Methodology

This section describes the steps followed to arrive at a parsimonious representation of the relationships between polyester, oil and cotton prices, starting from a very general unrestricted model. The first step is to analyze the time series properties of each price series. The second step is to test for cointegration among prices. The third step is to test for alternative restrictions on the parameters to arrive at a parsimonious model.

The classical regression model requires that all series be stationary and that the errors have a zero mean and finite variance to avoid the "spurious regression" problem, which consists of high statistical significance of the estimated model but a lack of a causal connection. A series is said to be (weakly or covariance) stationary if the mean and autocovariances of the series do not depend on time. So the first step is to analyze the time series properties of prices with Augmented Dickey-Fuller (ADF) tests before making inferences about their relationships. If prices in levels are non-stationary, the series are differenced n-times until the stationarity hypothesis cannot be rejected. The series, then, are said to be integrated of order n, I(*n*).

If the series are integrated of the same order and n>0, there might exist a long-run linear relationship between prices in levels such that the error term is stationary despite the fact that prices in levels are non-stationary. If such a relationship exists, then the series are said to be cointegrated. When series are cointegrated they cannot move independently from each other. In that case, an error-correction model is used to capture short- and long-term relationships among prices. Two methodologies are followed to test the null hypothesis of no cointegration among prices: the Engle-Granger methodology and the Johansen methodology.

According to the Engle-Granger methodology, the series are cointegrated if the residuals from the estimated equations of the long run relationships between the three variables are stationary. The following three long-run relationships are estimated separately:

$$(1)\quad PC_t = \beta_{01} + \beta_{11}PP_t + \beta_{21}PO_t + e_{PCt}$$

$$(2)\quad PO_t = \beta_{02} + \beta_{12}PC_t + \beta_{22}PP_t + e_{POt}$$

$$(3)\quad PP_t = \beta_{03} + \beta_{13}PO_t + \beta_{23}PC_t + e_{PPt}$$

where PC is the price of cotton, PP is the price of polyester, and PO is the price of oil. The PC, PP and PO series are said to be cointegrated of order (*n*,1), i.e. CI(*n*,1), if the null hypotheses a_{1j} =0, (*j=PP,PC,PO*), are rejected in the ADF tests on the residuals from the equilibrium equations $\{\hat{e}_{jt}\}$:

$$(4)\quad \Delta\hat{e}_{jt} = a_{1j}\hat{e}_{jt-1} + \sum_{i=1}^{k} a_{ji+1}\Delta\hat{e}_{jt-i} + \varepsilon_{jt}$$

where the residuals $\{\hat{\varepsilon}_{jt}\}$ are white noise. The lag length *k* is selected by paring down a model with *k*=24 according to the usual t-tests[3] so that the final model includes only significant lags at the 5% level of significance, and the Ljung-Box Q-statistics cannot reject the sequential hypotheses of no autocorrelation in the residuals $\{\hat{\varepsilon}_{jt}\}$ up to order 1, 2, 3, …, 24.

Sometimes the Engle-Granger tests suggest the existence of cointegration using one set of equations and no-cointegration using another set of equations. To avoid relying on mixed results, when Engle-Granger tests yield mixed results, the Johansen (1998) method is applied, which avoids the two-step approach to testing for cointegration and allows for the estimation and test for the presence of multiple cointegrating vectors. The Johansen (1998) approach requires the following error correction model (ECM) be estimated:

$$(5)\quad \Delta x_t = Bz_t + \pi_0 x_{t-1} + \sum_{i=1}^{p-1}\pi_i \Delta x_{t-i} + \tilde{\varepsilon}_t$$

where x_t is the vector of prices, z_t is a vector of deterministic variables, *B*, π_0 and the π_i's are matrices of coefficients, *p* is the lag length of the vector autoregression (VAR), and $\tilde{\varepsilon}_t$ is the vector of white noise errors. Since results depend on the number of lags considered, the general-to-specific modeling approach delineated in Enders (2004) is followed to determine the appropriate number of lags to consider: unrestricted VAR models in levels ($x_t = \alpha_0 + \alpha_1 x_{t-1} + ... + \alpha_p x_p + \varepsilon_t$) with alternative lag structures are estimated and the appropriate lag structure (*p*) is indicated by the model with the lowest Akaike, Schwartz, and Henderson-Quandt Information Criteria (abbreviated AIC, SIC and HQIC, respectively).

The number of independent cointegrating vectors equals the rank of π_0, r(π_0): if r(π_0)=0 then prices are not cointegrated; if r(π_0)=3, the vector process is stationary, i.e. all prices are jointly stationary; if r(π_0)=1, there is a single cointegrating vector and the expression $\pi_0 x_{t-1}$is the error-correction term; if r(π_0)=2, there are multiple cointegrating vectors. The Trace (λ_{trace}) and Maximum Eigenvalue (λ_{max}) tests are used to test alternative hypotheses on r(π_0).

$$(6)\quad \lambda_{trace}(r) = -T \sum_{i=r+1}^{n} \ln\left(1-\hat{\lambda}_i\right)$$

$$(7)\quad \lambda_{\max}(r, r+1) = -T \ln\left(1-\hat{\lambda}_{r+1}\right)$$

where the $\hat{\lambda}_i$'s are the estimated values of the eigenvalues obtained from the estimated $\hat{\pi}_0$ matrix, T is the number of usable observations, and n=0,1,2,3. λ_{trace} tests the null hypothesis that the number of distinct cointegrating vectors is less than or equal to r against a general alternative (greater than r). λ_{max} tests the null hypothesis that the number of cointegrating vectors is r against the alternative of $r+1$ cointegrating vectors. Critical values of λ_{trace} and λ_{max} are obtained from table E in Enders (2004).

The results from the Johansen (1998) test might differ depending on the assumed form of the cointegrating vector and the existence of trends in the prices in levels, i.e. the form of $Bz_t + \pi_0 x_{t-1}$. If results *do* differ, the likelihood ratio test (LRT) proposed by Johansen (1991) is used to detect whether restrictions on the structure of the data and the cointegrating vector artificially inflate the number of cointegrating relations. The Johansen (1991) method is a 3-step procedure. First, the model is fitted under the following alternative null hypotheses:

- *Null hypothesis 1, unrestricted model*: the level data x_t have no deterministic trends and the cointegrating equations do not have intercepts, i.e. H1: $Bz_t + \pi_0 x_{t-1} = \alpha(\beta' x_{t-1})$, where α and β are two matrices of dimension (n.r) where r is the rank of π_0, β is the matrix of cointegrating parameters and α is the matrix of weights with which each cointegrating vector enters the n estimated equations. α is also the matrix of adjustment coefficients to the long-term equilibrium.

- *Null hypothesis 2*: the level data x_t have no deterministic trends and the cointegrating equations have intercepts, i.e. H2: $Bz_t + \pi_0 x_{t-1} = \alpha(\beta' x_{t-1} + \rho_0)$.

- *Null hypothesis 3*: the level data x_t have linear trends but the cointegrating equations have only intercepts, i.e. H3: $Bz_t + \pi_0 x_{t-1} = \alpha(\beta' x_{t-1} + \rho_0) + \alpha^* \gamma_0$, where $\alpha^* \gamma_0$ is the deterministic trend in x_t and the term in parenthesis is the cointegrating relation.[4]

- *Null hypothesis 4*: the level data x_t and the cointegrating equations have linear trends, i.e. H4: $Bz_t + \pi_0 x_{t-1} = \alpha(\beta' x_{t-1} + \rho_0 + \rho_1 t) + \alpha^* \gamma_0$.

Second, the characteristic roots of the π_0 matrix are calculated for each model and ordered in a decreasing manner: $\hat{\lambda}_{Hj,1} > \hat{\lambda}_{Hj,2} > \ldots > \hat{\lambda}_{Hj,n}$, where Hj indicates the appropriate null hypothesis (j=1,2,3,4).

Third, the null hypothesis that the restrictions from the model estimated under Hm (m=2,3,4) are not binding in the model estimated under Hk (k=1,2,3) is tested with the Johansen (1991) LRT:

$$(8)\quad LRT_{Hm,Hk} = -T \sum_{i=r+1}^{n} \left[\ln\left(1-\hat{\lambda}_{Hm,i}\right) - \ln\left(1-\hat{\lambda}_{Hk,i}\right) \right]$$

where r and m are, respectively, the number of nonzero and the total number of characteristic roots in the less restricted model Hk, and T is the number of observations in the time space. The LRT has a Chi-squared distribution with (n-r) degrees of freedom. The intuition behind the test is that if in the unrestricted model there are only r cointegrating vectors, then that number should not diminish if a non-binding condition is imposed, and the values of $\ln\left(1-\hat{\lambda}_{Hm,i}\right)$ and $\ln\left(1-\hat{\lambda}_{Hk,i}\right)$ should be the same. Therefore, small values of the LRT indicate that it is permissible to include the restrictions in the model, and large values of the LRT imply that the restrictions artificially inflate the number of cointegrating vectors (Enders 2004). Finally, the Trace and Max-Eigenvalue tests on the most restricted model that cannot be rejected by the Johansen (1991) test determine the true number of cointegrating vectors.

If prices in levels are non-stationary and not cointegrated, then a VAR model in the stationary differenced prices is estimated:

$$(9)\quad \Delta x_t = A_0 + \sum_{i=1}^{p-1} \pi_i \Delta x_{t-i} + \varepsilon_t$$

Since differenced prices are stationary, tests of hypothesis can be conducted using classical regression techniques on (9). If prices in levels are cointegrated, then the structural representation of the ECM in (5) selected according to the Johansen (1998) test is used to make inferences about price relationships. Cross-equation restrictions in the final model are tested with the LRT suggested by Sims (1980):

$$(10)\ LR = (T-c)\left(\ln|\Sigma_r| - \ln|\Sigma_u|\right)$$

where $\ln|\Sigma_r|$ is the natural logarithm of the determinant of the variance-covariance matrix of the residuals of the restricted model, $\ln|\Sigma_u|$ is the natural logarithm of the determinant of the variance-covariance matrix of the residuals of the unrestricted model, c is the maximum number of regressors contained in the longest equation, and T is the number of observations in the time space. The LRT follows a Chi-square distribution with degrees of freedom equal to the number of restrictions in the system.

In particular, we are interested in determining whether one or more prices do not receive significant feedback from changes in other prices and therefore do not need a VAR representation, i.e. they can be treated as weakly exogenous and their equation can be eliminated from the system. This is done by testing for block exogeneity. The test for block-exogeneity restricts all lags of one series of prices in the other series of prices to zero. The unrestricted model in (10) consists of the VAR equations of the 2 endogenous prices including p lags of the potentially block-exogenous price. The restricted model excludes all lags of the potentially block-exogenous price. The LRT test has $2p$ degrees of freedom, since p lags are excluded in each of the equations of the model. If the hypothesis of block exogeneity is rejected, then that price is said to Granger-cause the other 2 prices.

The forecasting power of the final model is tested by the Theil Inequality coefficient. The Theil Inequality coefficient (TIC) takes values between 0 and 1, zero indicating a perfect fit of the forecast to the observed series.

The stability of the final model, i.e. the absence of structural breaks, is tested with the Quandt-Andrews (Q-A) and the Chow Forecast tests (Quantitative Micro Software 2007). These tests evaluate whether the parameters of the model are stable across various sub-samples of the data. The Chow's Forecast test estimates two models using the whole sample: the restricted regression uses the original set of regressors, while the unrestricted regression adds a dummy variable for each forecast point. The Chow Forecasts log likelihood ratio statistic compares the maximum of the (Gaussian) log likelihood function of each model and has asymptotic Chi-squared distribution with T_2 degrees of freedom.

The logic behind the Q-A test is that a single Chow Breakpoint test is performed at every observation between two dates, τ_1 and τ_2. The Breakpoint Chow test fits the model separately for each subsample and one (restricted) model for the entire period, and tests whether there are significant differences in the estimated parameters across models. The resulting test statistics are then summarized into one test statistic to test the null hypothesis that there are no breakpoints between τ_1 and τ_2. The test trims a small percentage of observations at the beginning and the end of the full sample period to avoid the degeneration of the non-standard distribution followed by the test. The Maximum Q-A statistic, *MaxF*, is the maximum of the individual Chow F-statistics, calculated as:

$$(11)\ MaxF = \max_{\tau_1 \le \tau \le \tau_2}\left(F(\tau)\right)$$

$$(12)\ F(\tau) = \frac{\left(\bar{u}'\bar{u} - (u_1'u_1 + u_2'u_2)\right)/k}{(u_1'u_1 + u_2'u_2)/(T-2k)}$$

where $\bar{u}'\bar{u}$ is the restricted sum of squares and $u_i'u_i$ is the sum of squared residuals from subsample i. Each F-statistic follows an F-distribution with (k, T-k) degrees of freedom, where k is the number of parameters in the equation, and T is the number of observations in the time space. Therefore, failing to reject the null hypothesis of the Q-A test indicates stability of the model over the trimmed sample.

Results

The data used in the present analysis consists of monthly observations of polyester, cotton and oil prices over the period January 1991 through November 2008. The selection of the sampling period was based on data availability. Polyester prices are the midpoints of a range of prices reported by PCI Fibres[5] as the world average of polyester staple prices, 1.5 denier (North America, Western Europe, and Asia, weighted by annual production volumes). Oil prices are the simple average of crude oil spot prices of Dated Brent, West Texas Intermediate, and Dubai Fateh, reported by the International Monetary Fund (IMF 2009). Cotton prices are the monthly average of the cheapest five quotations from a selection (at present numbering nineteen) of the principal upland cottons traded internationally, with base quality middling 1-3/32", reported by Cotlook Ltd. and known as the A Index. All prices are expressed in U.S. dollars, and deflated by the U.S. consumer price index (1982-84=100). By deflating the series, the effect of inflation is removed from nominal prices, and the resulting real prices are expressed in currency units directly comparable through time (U.S. dollars of 1982-84). Monthly U.S. inflation ranged from -1.9% to 1.2% over the sample period, with an average of 0.2%. However, the accumulated inflation between April 1991 and November 2008 amounted to 57%. The price index used to deflate the series was chosen on grounds of availability of monthly data. Table 1 presents descriptive statistics for the deflated series in levels.[6] All results presented in this article refer to analyses conducted on real prices expressed in natural logarithms.

Table 1. Descriptive Statistics of the Real Prices in Levels (1982-84=100)

Statistic	Polyester(1)	Cotton(2)	Oil(3)
Mean	32.42	39.10	17.95
Maximum	65.79	76.04	60.26
Minimum	19.18	20.93	6.35
Std. Dev.	9.80	11.84	10.38
Observations	215	215	215

Note: (1) Deflated price of polyester, in U.S. cents per pound; (2) Deflated price of cotton, in U.S. cents per pound; (3) Deflated price of oil, in U.S. dollars per barrel.

According to the ADF tests, polyester, cotton, and oil prices are non-stationary in levels but stationary in first differences, i.e. they are I(1) (table 2).[7] Classical regression methods on the variables in levels would provide spurious results, so cointegration tests are performed next: the Engle-Granger tests first, and then the Johansen methodology.

To eliminate autocorrelation of the residuals from regression (4), the equations for cotton, oil, and polyester prices are augmented, respectively, with lags {1, 18}, {1, 10, 13}, and {1, 18}. The Ljung-Box statistics indicate no autocorrelation of the residuals up to lag 24: the lowest p-value in the cotton equation corresponds to the test for lags up to order 14, and equals 0.523; the lowest p-value in the oil equation corresponds to the test for lags up to order 3, and equals 0.605; the lowest p-value in the polyester equation corresponds to the test for lags up to order 1, and equals 0.590.[8]

The critical values at the 5% and 10% significance level for the Engle-Granger cointegration test for 3 variables and 200 observations are, respectively, -3.785 and -3.484 (Enders 2004, p. 441).[9] The estimated values of a_{1j} and the associated t-statistics from the ADF tests on the long-run equilibrium equations are presented in table 3. Since the t-statistics from the equilibrium equations for oil and polyester prices are smaller in absolute value than the critical value, the null hypothesis of no cointegration cannot be rejected with these equations. The null hypothesis is only rejected with the cotton equation. The Engle-Granger tests yield mixed results and provide weak support to the hypothesis of cointegration of polyester, cotton and oil prices. To avoid relying on mixed results, the Johansen method is implemented next.

Table 2. Unit Root Tests

Augmented Dickey Fuller Tests for Cotton Prices

Variable in	Exogenous	T	Lag Length^	t-Statistic	p-value	
Levels	None	213	1	-0.968	0.297	
Levels	Constant	213	1	-2.362	0.154	
Levels	Constant, Linear Trend	213	1	-3.352	0.061	*
First Differences	None	213	0	-8.258	0.001	***
First Differences	Constant	213	0	-8.286	0.001	***
First Differences	Constant, Linear Trend	213	0	-8.264	0.001	***

Augmented Dickey Fuller Tests for Oil Prices

Variable in	Exogenous	T	Lag Length^	t-Statistic	p-value	
Levels	None	213	1	0.178	0.737	
Levels	Constant	213	1	-1.423	0.570	
Levels	Constant, Linear Trend	213	1	-2.444	0.356	
First Differences	None	213	0	-11.490	0.001	***
First Differences	Constant	213	0	-11.475	0.001	***
First Differences	Constant, Linear Trend	213	0	-11.426	0.001	***

Augmented Dickey Fuller Tests for Polyester Prices

Variable in	Exogenous	T	Lag Length^	t-Statistic	p-value	
Levels	None	213	1	-0.990	0.288	
Levels	Constant	213	1	-1.546	0.509	
Levels	Constant, Linear Trend	213	1	-1.915	0.643	
First Differences	None	213	0	-9.069	0.001	***
First Differences	Constant	213	0	-9.100	0.001	***
First Differences	Constant, Linear Trend	213	0	-9.062	0.001	***

Note: *, **, *** indicate rejection of the null hypothesis that the series has a unit root at the 10%, 5%, and 1% level, respectively, using MacKinnon (1996) one-sided p-values; T: number of observations; ^ Lag length based on SIC from lags 1 to 24.

The AIC, SIC and HQIC on unrestricted VAR models in levels with alternative lag structures indicate that a VAR model with 2 lags best represents the data among VAR models with lags 1 through 18. Therefore, the Johansen (1998) method in (5) is applied with p=2 (only 1 lag in differences). Table 4 presents the summary of results from the Johansen (1998) method under null hypotheses 1 through 4. According to the Trace tests, the hypothesis that $r(\pi_0)=0$ cannot be rejected at the 10% significance level for any model. The Max-Eigenvalue test under null hypotheses 2 and 3 indicates one possible cointegrating vector at the 10% and 5% significance level, respectively.

Table 3. ADF Test on the Residuals from the Equilibrium Equations

Long-run Equilibrium Equation	Dependent Variable	Included Lags	a_{1j} Estimate	t-statistic
Cotton	$\Delta\hat{e}_{PCt}$	1 , 18	-0.096524	-3.996**
Oil	$\Delta\hat{e}_{POt}$	1, 10, 13	-0.052423	-2.746
Polyester	$\Delta\hat{e}_{PPt}$	1, 18	-0.072204	-3.364

Note: ** indicates rejection of the null hypothesis that a_{1j} =0 at the 5% level.

Table 5 presents the results of the Johansen (1991) LRT for all possible pairwise comparisons of the models under alternative hypotheses according to the Max-Eigenvalue criteria. The critical values from a Chi-square distribution with 3 degrees of freedom at the 5% and 10% level of significance corrected for the presence of 6 simultaneous tests[10] are, respectively, 11.74% and 10.26%. The critical values from a Chi-square distribution with 2 degrees of freedom at the 5% and 10% level of significance corrected for the presence of 6 simultaneous tests are, respectively, 9.57% and 9.19%. Since the restrictions on the form of the data and the cointegrating vectors from *null hypotheses 2* and *4* are binding at the 5% and 1% level of significance, respectively, and those from *null hypothesis 3* are binding at the 5% level of significance when compared to the restricted model (estimated under *null hypothesis 1*), then it is concluded that models estimated under null hypotheses 2, 3 and 4 artificially inflate the number of cointegrating vectors due to the imposed structure, and the unrestricted model appropriately fits the data.[11] Since the Trace and the Max-Eigenvalue tests for the model fit under *null hypothesis 1* indicate that the number of cointegrating vectors among cotton, oil and polyester prices is zero, it is concluded that these prices are not cointegrated.

Table 4. Johansen (1998) Tests for Cointegration of Cotton, Oil and Polyester Prices under Alternative Null Hypotheses, *p*=2

Null Hypothesis 1: No Deterministic Trends in the Level Data and No Intercepts in Cointegrating Equations

Hypothesized No. of CE(s)	Eigen-value	Trace Test			Max-Eigenvalue Test		
		Statistic	Critical Value	p-value^	Statistic	Critical Value	p-value^
None	0.058389	15.48	24.28	0.419	12.81	17.80	0.240
At most 1	0.009818	2.66	12.32	0.888	2.10	11.22	0.911
At most 2	0.002625	0.56	4.13	0.516	0.56	4.13	0.516

Null Hypothesis 2: No Deterministic Trends in the Level Data and Intercepts in the Cointegrating Equations

Hypothesized No. of CE(s)	Eigen-value	Trace Test			Max-Eigenvalue Test			
		Statistic	Critical Value	p-value^	Statistic	Critical Value	p-value^	
None	0.098012	27.87	35.19	0.247	21.97	22.30	0.055	*
At most 1	0.017978	5.90	20.26	0.954	3.86	15.89	0.964	
At most 2	0.009525	2.04	9.16	0.770	2.04	9.16	0.770	

Null Hypothesis 3: Linear Trends in the Level Data and Intercepts in the Cointegrating Equations

Hypothesized No. of CE(s)	Eigen-value	Trace Test			Max-Eigenvalue Test			
		Statistic	Critical Value	p-value^	Statistic	Critical Value	p-value^	
None	0.096292	25.75	29.80	0.136	21.57	21.13	0.043	**
At most 1	0.017863	4.19	15.49	0.888	3.84	14.26	0.876	
At most 2	0.001634	0.35	3.84	0.555	0.35	3.84	0.555	

Null Hypothesis 4: Linear Trends in the Level Data and Linear Trends in the Cointegrating Equations

Hypothesized No. of CE(s)	Eigen-value	Trace Test			Max-Eigenvalue Test		
		Statistic	Critical Value	p-value^	Statistic	Critical Value	p-value^
None	0.098608	32.96	42.92	0.338	22.11	25.82	0.143
At most 1	0.032612	10.85	25.87	0.884	7.06	19.39	0.896
At most 2	0.017622	3.79	12.52	0.773	3.79	12.52	0.773

Note: ** and * denote rejection of the hypothesis at the 5% and 10% levels, respectively; ^MacKinnon-Haug-Michelis (1999) p-values

In the absence of cointegration, price relationships must be analyzed as a VAR model in first differences according to (9). Ordinary least squares[12] (OLS) estimates of the full VAR model with *p*=2 are reported in table 6. According to this model: (a) cotton prices only depend on past realizations of cotton prices, but not on past realizations of oil or polyester prices; (b) oil prices only depend on past realizations of oil prices, but not on past realizations of cotton or polyester prices; (c) polyester prices depend on past realizations of polyester and oil prices, but not on past realizations of cotton prices.

Table 5. Johansen (1991) LRT to Determine the Appropriate Structure of the Model Using the Max-Eigenvalue Criteria

Hypotheses Tested	LRT Statistic	p-value		$r(\pi_0)$ in Unrestricted model	Df	Unrestricted Model
H1 vs. H2	12.40	0.01	**	0	3	H1
H2 vs. H3	1.72	0.42		1	2	H3
H3 vs. H4	6.66	0.04	**	1	2	H3
H2 vs. H4	4.95	0.08	*	1	2	H2
H1 vs. H4	17.49	0.00	***	0	3	H1
H1 vs. H3	10.28	0.02	**	0	3	H1

Note: T=213. *, **, *** indicate rejection of the null hypothesis that the restrictions in the restricted model are binding at the 10%, 5%, and 1% level, respectively.

Table 6. Vector Autoregression Estimates, *p*=2

	ΔPC	ΔPO	ΔPP
ΔPC(-1)	0.486348	0.170101	0.064627
	(0.06322)	(0.10801)	(0.04967)
	[7.693]***	[1.574]	[1.301]
ΔPO(-1)	-2.55E-06	0.222193	0.145165
	(0.04104)	(0.07011)	(0.03225)
	[-6.2e-05]	[3.169]***	[4.502]***
ΔPP(-1)	0.064605	-0.158349	0.299289
	(0.08625)	(0.14735)	(0.06777)
	[0.749]	[-1.075]	[4.416]***
A_{0i}	-0.002285	0.002424	-0.002625
	(0.00306)	(0.00522)	(0.00240)
	[-0.747]	[0.464]	[-1.093]
R-squared	0.247373	0.059658	0.22607
Adj. R-squared	0.236569	0.04616	0.214961
F-statistic	22.898	4.420	20.350
p-value(F(3,210))	0.000	0.005	0.000

Note: *, **, *** indicate rejection of the null hypothesis that the coefficient equals zero at the 10%, 5%, and 1% level, respectively; Standard errors in ( ) & t-statistics in [ ]; Sample (adjusted): 1991M03 2008M11; T= 213 after adjustments.

The following three tests examine whether each price is block-exogenous. The LR test for the hypothesis that cotton prices do not Granger-cause oil *and* polyester prices is performed using the equations for polyester and oil prices only. The LR test equals 3.30, the p-value with 2 degrees of freedom is 0.192, and the null hypothesis cannot be rejected.

The LR test for the hypothesis that oil prices do not Granger-cause cotton *and* polyester prices equals 20.57, the p-value with 2 degrees of freedom is less than 0.001, and the null hypothesis is rejected. Therefore, oil prices Granger-cause cotton *and* polyester prices.

The LR test for the hypothesis that polyester prices do not Granger-cause cotton *and* oil prices equals 1.95, the p-value with 2 degrees of freedom equals 0.377, and the null hypothesis cannot be rejected. Therefore, polyester prices do not Granger-cause cotton *and* oil prices.

Furthermore, the LR statistic for the null hypothesis that the coefficients of cotton and polyester prices in the oil price equation jointly equal zero is 2.96, the p-value with 2 degrees of freedom equals 0.228, and the hypothesis cannot be rejected at the 10% level of significance. Therefore, the oil price equation can be deleted from the VAR system while oil prices are retained as explanatory variables in the equations for cotton and polyester prices.

Rejection of the cointegration hypothesis for three time series does not rule out the possibility of two of them being cointegrated. Therefore, a cointegration analysis is run for each pair of prices. The results from the ADF on the long-run equilibrium relations are reported in table 7. The critical values at the 5% and 10% significance level for the Engle-Granger cointegration test for two variables and 200 observations are, respectively, -3.368 and -3.067 (Enders 2004, p. 441).[13] Since all t-statistics are lower in absolute value than the critical values, the null hypothesis of no cointegration cannot be rejected for any pair of prices. The Ljung-Box statistics indicate no autocorrelation of the residuals up to lag 24: (a) in Panel a, the lowest p-value in the cotton equation corresponds to the test for lags up to order 12, and equals 0.295, while the lowest p-value in the polyester equation corresponds to the test for lags up to 17, and equals 0.265;[14] (b) in Panel b, the lowest p-value in the polyester equation corresponds to the test for lags up to order 2, and equals 0.136, while the lowest p-value in the oil equation corresponds to the test for lags up to 13, and equals 0.384;[15] (c) in Panel c, the lowest p-value in the oil equation corresponds to the test for lags up to order 13, and equals 0.173, while the lowest p-value in the cotton equation corresponds to the test for lags up to 11, and equals 0.137.[16]

Table 7. ADF Tests on the Residuals from the Equilibrium Equations

Cotton and Polyester (Excluding Oil)

Long-run Equilibrium Equation	Dependent Variable	Included Lags	a_{1j} Estimate	t-statistic
Cotton	$\Delta\hat{e}_{PCt}$	1, 18 ,19	-0.055	-3.030
Polyester	$\Delta\hat{e}_{PPt}$	1, 18 ,19	-0.049	-2.844

Polyester and Oil (Excluding Cotton)

Long-run Equilibrium Equation	Dependent Variable	Included Lags	a_{1j} Estimate	t-statistic
Oil	$\Delta\hat{e}_{POt}$	1	-0.018	-1.540
Polyester	$\Delta\hat{e}_{PPt}$	1	-0.015	-1.645

Cotton and Oil (Excluding Polyester)

Long-run Equilibrium Equation	Dependent Variable	Included Lags	a_{1j} Estimate	t-statistic
Cotton	$\Delta\hat{e}_{PCt}$	1, 5	-0.034	-2.446
Oil	$\Delta\hat{e}_{POt}$	1	-0.031	-2.201

Note: *, ** indicate rejection of the null hypothesis that a_{1j} =0 at the 10% and 5% level, respectively.

Since the prices of cotton and polyester are not cointegrated,[17] an ECM is not attempted and price relationships are modeled as a VAR in first differences. Furthermore, since oil prices were found to Granger-cause cotton *and* polyester prices in the 3-equation VAR in first differences from table 6, the 2-equation VAR in first differences for cotton and polyester is augmented with lagged first differences of oil prices. The number of lagged terms of oil prices in differences to include in the 2-equation model is determined by paring down a 2-equation VAR model with 5 lagged terms of oil prices in differences (corresponding to 6 lags in levels) according to the LRT in (10). The chosen number of lags of oil in first differences to include is 2.

Table 8 presents the OLS estimates of the augmented 2-equation VAR in first differences. Only the lagged change in cotton prices is significant in the cotton equation: current changes in cotton prices are positively correlated with changes in cotton prices during the previous month. Oil and polyester prices are not significant in the cotton equation. In the equation for polyester prices, lagged changes in polyester and oil prices are significant, but not cotton prices: current changes in polyester prices are positively correlated with changes in polyester prices during the previous month, as well as with changes in oil prices during the previous 2 months. The constants at the origin are negative, indicating that cotton and oil prices tend to decline through time. However, the constants are not significant. The R-squares of the cotton and polyester price equations are, respectively, 0.248 and 0.251.

Applying LR tests to the VAR model, two hypotheses are tested and none can be rejected at the 10% level of significance: (a) the LR test for the null hypothesis that oil prices do not influence cotton prices equals 0.202, and the p-value with 2 degrees of freedom equals 0.904; (b) the LR test for the null hypothesis that oil and polyester prices do not influence cotton prices equals 0.788, and the p-value with 3 degrees of freedom equals 0.852. Therefore, it can be inferred that cotton prices do not depend on oil or polyester prices. Note that since the system is block-recursive, each equation can be estimated separately.

Table 8. Vector Autoregression Estimates, *p*=2, Augmented with Lagged Oil Prices in Differences

	ΔPC	ΔPP
ΔPC(-1)	0.486041	0.071672
	(0.06364)	(0.04922)
	[7.638]***	[1.456]
ΔPP(-1)	0.076936	0.246218
	(0.09035)	(0.06987)
	[0.852]	[3.524]***
A_{0i}	-0.002062	-0.003165
	(0.00309)	(0.00239)
	[-0.668]	[-1.325]
ΔPO(-1)	-0.002184	0.139072
	(0.04222)	(0.03265)
	[-0.052]	[4.259]***
ΔPO(-2)	-0.019039	0.089475
	(0.04388)	(0.03394)
	[-0.434]	[2.637]***
R-squared	0.24871	0.250799
Adj. R-squared	0.234192	0.236322
F-statistic	17.131	17.323
p-value(F(4,209))	<0.001	<0.001

Note: *, **, *** indicate rejection of the null hypothesis that the coefficient equals zero at the 10%, 5%, and 1% level, respectively; Standard errors in ( ) & t-statistics in [ ]; Sample (adjusted): 1991M03 2008M11; T= 209 after adjustments.

The estimated parameters are used to calculate the magnitude and speed of propagation of price shocks.[18] A permanent 10% increase (decrease) in oil prices results in a final 3.1% increase (decrease) in polyester prices with respect to the pre-shock level, and the adjustment occurs over a 6-month period. In the first month following the shock, 45% of the adjustment occurs. The adjustment increases to 87% in the second month, and to 97% in the third month following the shock. So the propagation of a shock in oil prices is fast, but only a small proportion of the shock is transmitted. Therefore, if real oil prices increase by 25% in the long term (from $60 to $75) as expected by the World Bank (2008, p.87), polyester prices can be expected to increase by about 8%.

An unexpected 10% increase (decrease) in polyester prices results in a final 14.5% increase (decrease) in polyester prices with respect to the pre-shock level. The adjustment occurs over a 4-month period (excluding the month of the original shock), but only in the first 2 months are changes in polyester prices statistically significant at the 5% level. Seventy-four percent of the adjustment occurs in the following month after the shock, and it accumulates to 93% by the second month. Therefore, unexpected shocks in polyester prices are self-perpetuating and rapidly propagated.

An unexpected 10% increase (decrease) in cotton prices tends to self-perpetuate through time, resulting in a final 19% increase (decrease) in cotton prices with respect to the pre-shock level. Half of the adjustment occurs in the month following the shock, accumulating to 75% in 2 months. Full adjustment occurs over a 7-month period (excluding the month of the original shock), although only changes in the first 5 months are statistically significant at the 5% level. This is relevant in the face of recent shocks in the cotton futures and options markets that possibly originated from non-fundamental factors (Plastina 2008). These shocks are rapidly transmitted to the spot market (Plastina 2009) and might affect spot prices for several months after their occurrence.

Table 9. Quandt-Andrews Stability Tests for the Cotton and Polyester Equations.
Equation: Cotton

Sample [Trim]	MaxF Statistic	p-value	Month	Number of breaks compared
1991m12-2007m09 [5%]	2.33	1	2007m08	190
1992m11-2006m10 [10%]	2.30	1	2004m01	168
1993m09-2005m12 [15%]	2.30	1	2004m01	146

Equation: Polyester

Sample [Trim]	MaxF Statistic	p-value	Month	Number of breaks compared
1991m12-2007m09 [5%]	5.05	0.99	1999m03	190
1992m11-2006m10 [10%]	5.05	0.99	1999m03	168
1993m09-2005m12 [15%]	5.05	0.98	1999m03	146

Note: Null Hypothesis: No breakpoints within the sample; Varying regressors: All equation variables; p-value calculated using Hansen's (1997) method; *, **, *** indicate rejection of the null hypothesis at the 10%, 5%, and 1% level, respectively.

The forecasting power of the model is tested by fitting the model with observations from January 1991 to December 2005, and comparing the forecast values with the actual realizations from January 2006 to November 2008. For the cotton equation the TIC takes a value of 0.62, indicating poor forecasting power. The model correctly forecasts the sign of the price change 57% of the time. From a practitioner's stand point, it can be concluded that monthly changes in cotton prices cannot be consistently anticipated by looking at past changes in oil, polyester or even cotton prices. For the polyester equation the TIC takes a value of 0.403, suggesting poor forecasting power. The model correctly forecasts the sign of the price change 69% of the time. It is concluded that monthly changes in polyester prices cannot be consistently anticipated by looking at past changes in oil, cotton or even polyester prices.

The Quandt-Andrews test on each price equation fails to reject the null hypothesis of no breakpoint of the model over the period 1991-2007 at the 10% confidence level (table 9). Therefore, the models for cotton and polyester prices can be considered stable over those years. But since the cotton market was not an exception to the recent commodity price boom (World Bank 2008) and was apparently affected by non-fundamental factors in 2008 (Plastina 2008), it is useful to test for the existence of structural breaks after September 2007. The Chow Forecast tests fail to reject the null hypothesis of the existence of a structural break in any of the price equations between October 2007 and June 2008 at the 10% comparisonwise level of significance (table 10).[19] Therefore, it can be concluded that the short-run price relationships between cotton, polyester and oil prices are stable throughout the sample. Despite its low forecasting power, the final 2-equation structural model is useful to understand the relationships among prices.

Table 10. Chow Forecast Tests of Stability of the Cotton and Polyester Price Equations

t	Cotton			Polyester		
	F-statistic	DF	p-value	F-statistic	DF	p-value
2007M10	1.433608	14, 193	0.141	0.780418	14, 193	0.690
2007M11	1.551703	13, 194	0.102	0.835019	13, 194	0.623
2007M12	1.687947	12, 195	0.072	0.900632	12, 195	0.547
2008M01	1.850795	11, 196	0.048	0.879289	11, 196	0.562
2008M02	1.883063	10, 197	0.049	0.952399	10, 197	0.487
2008M03	2.102909	9, 198	0.031	1.063287	9, 198	0.392
2008M04	2.198363	8, 199	0.029	1.200613	8, 199	0.300
2008M05	1.85069	7, 200	0.080	1.347637	7, 200	0.230
2008M06	2.154738	6, 201	0.049	1.541413	6, 201	0.166

Note: Null hypothesis: no structural break exists in t; *, **, *** indicate rejection of the null hypothesis at the 10%, 5%, and 1% Bonferroni-adjusted comparisonwise level of significance, respectively; Number of comparisons=9.

Conclusions

In light of recent changes in the structure of the world market for textile fibers, the price relationships between the two major textile fibers -cotton and polyester- and oil are analyzed with robust time series methods and no a priori restrictions on the estimating model.

Real monthly cotton, polyester and oil prices in natural logarithmic form are integrated of order 1. Cointegration tests indicate that there is no *stable* long-run relationship between cotton, polyester and oil prices, as well as no *stable* long-run relationship between cotton and polyester prices. Monthly changes in polyester prices *do* depend on previous changes in polyester and oil prices; but monthly changes in cotton prices *do not* depend on previous changes in oil or polyester prices. Oil prices are found to Granger-cause polyester prices. As expected, monthly changes in oil prices are found *not* to depend on polyester or cotton prices. These results are similar to the findings of Fadiga and Misra (2007), based on real annual prices.

Unexpected changes in polyester and cotton prices are self-perpetuating and rapidly propagated. An unexpected 10% increase (decrease) in the current price of polyester puts upward (downward) pressure on the price of polyester over the following 4 months. Three-quarters of the total adjustment to the shock occurs in the first month after the shock. An unexpected 10% increase (decrease) in the current price of cotton puts upward (downward) pressure on cotton prices over the following 7 months. Three-quarters of the total adjustment to the shock occurs in the first 2 months after the shock.

From a practitioner's stand point, it can be concluded that monthly changes in cotton and polyester prices cannot be consistently anticipated by looking at past changes in oil, polyester or cotton prices.

Finally, it must be noted that (a) the present analysis is conducted in real terms, and any extrapolation to nominal terms must include the effects of inflation; (b) the analysis is based solely on prices, and excludes causality relationships with market fundamentals (demand, supply, stocks, and trade).

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[1] In terms of equation (1) below, the restrictions $\beta_{11} = 1$ and $\beta_{21} = 0$ are imposed in Baffes and Gohou (2005). However, if the first restriction is binding (i.e. $\beta_{11} \neq 1$) then the residuals from the long run relationship are the summation of the true residuals and polyester prices multiplied by the difference between the unrestricted slope coefficient and 1. Therefore, if the true residuals are non-stationary but they are cointegrated with polyester prices, then the test performed in Baffes and Gohou (2005) would incorrectly reject they hypothesis of no stationarity.

[2] The World Bank (2008), extending the analysis in Baffes (2007) to 2007, reports that the adjusted R^2 of a regression between cotton and oil prices (with the manufacture unit value index and a time trend as covariates) is 0.81, indicating a strong co-movement among cotton and crude oil annual average prices.

[3] From Rule 1 of Sims, Stock and Watson (1990) all the coefficients on the expressions $\Delta\hat{e}_{jt-i}$ converge to t-distributions.

[4] The decomposition of $\alpha\ \rho_0+\alpha^*\gamma_0$ (and $\alpha(\rho_0+\rho_1 t)+\ \alpha^*\gamma_0$) is (are) not uniquely identified. EViews identifies the first term by regressing the cointegrating relation $\beta'\ x_{t-1}$ on a constant (and a linear trend), such that the error correction term has a sample mean of zero (Quantitative Micro Software 2007, p. 365-6).

[5] PCI Fibres is a consultancy firm to the fibers and related industries, specializing on the major manufactured fibers and raw materials for acrylic, nylon, polyester and viscose as well as related products.

[6] ANOVA and Welch F-Tests for equality of means indicate that mean oil, cotton and polyester prices are significantly different from each other. Barlette and Levene tests for equality of variances indicate that the variance of cotton and polyester prices are not significantly different, but their variances are significantly different from the variance of oil prices. Therefore, it can be concluded that while polyester prices tended to be lower than cotton prices, their variability around the mean is about the same.

[7] The ADF test for cotton prices in levels including a constant and a linear trend rejects the null hypothesis that the series has a unit root at the 10% level of significance. Based on the ADF test, some researchers might consider the cotton series as trend stationary. However, the Philips Perron test (available from the author upon request) reaches the opposite conclusion.

[8] Normality of the residuals of the augmented equations could not be rejected at the 10% significance level: the p-value of the Jarque-Bera test for the cotton, oil and polyester equations are, respectively 0.46, 0.49, and 0.63.

[9] The extrapolated values for 212 observations are, respectively, -3.784 and -3.482.

[10] The Bonferroni correction provides an approximation to multiple comparison error rates when those comparisons are not independent. The comparisonwise error rate for the test statistic is determined by dividing the maximum desired family error rate by the number of simultaneous test, i.e. 0.05/6= 0.0083 (Kuehl 2000).

[11] The same conclusion is reached if the Bonferroni correction is omitted.

[12] OLS yield consistent and efficient VAR estimates.

[13] The extrapolated values for 212 observations are, respectively, -3.367 and -3.066.

[14] Normality of the series of residuals $\{\hat{\varepsilon}_{jt}\}$ cannot be rejected at the 10% significance level: the p-values of the Jarque-Bera tests for the cotton and polyester equations are, respectively, 0.727 and 0.503.

[15] Normality of the residuals of the oil equation is rejected at the 5% significance level (p-values of the Jarque-Bera test = 0.011), and normality of the residuals of the polyester equation is rejected at the 1% significance level (p-values of the Jarque-Bera test < 0.0001).

[16] Normality of the residuals of the oil equation is rejected at the 5% significance level (p-value of the Jarque-Bera test = 0.025), and normality of the residuals of the cotton equation is rejected at the 10% significance level (p-value of the Jarque-Bera test=0.083).

[17] The Johansen (1998) tests also reject the existence of a cointegrating relationship among cotton and polyester prices.

[18] The (non-significant) constants at the origin are excluded for the following calculations. Including the constants would change the analysis to one of deviations from declining trends in real cotton and polyester prices.

[19] The Bonferroni-adjusted 10% comparisonwise critical value is 0.011=0.1/9.

References

Baffes, J. 2007. “Oil Spills on Other Commodities.” Policy Research Working Paper 4333, World Bank.

Baffes, J., and G. Gohou. 2005. “The Co-movement between Cotton and Polyester Prices.” Policy Research Working Paper 3534, World Bank.

Enders, W. 2004. Applied Econometrics Time Series, 2nd ed. Wiley Series in Probability and Statistics.

Fadiga M. and S. Misra. 2007. “Common Trends, Common Cycles, and Price Relationships in the International Fiber Market.” Journal of Agricultural and Resource Economics 23(1): 154-168.

Food and Agriculture Organization of the United Nations. 2002. “Oil Prices and Agricultural Commodity Prices.” In Commodity Market Review 2001-2002. Rome.

Harri, A., Nalley L. and D. Hudson. 2009. “The Relationship between Oil, Exchange Rates, and Commodity Prices.” Journal of Agricultural and Applied Economics 41(2): 501-510.

International Cotton Advisory Committee. 2009. World Textile Demand. Washington, DC.

International Monetary Fund. 2009. Primary Commodity Prices. Washington DC.

Johansen, S. 1991. “Estimation and Hypothesis Testing of Cointegrating Vectors in Gaussian Vector Autoregressive Models.” Econometrica 59: 1551-80

---. 1988. “Statistical Analysis of Cointegration Vectors.” Journal of Economic Dynamics and Control 12: 231-54.

Kuel, R. O. 2000. Design of Experiments: Statistical Principles of Research Design and Analysis, 2nd ed. California: Duxbury Press.

PCI Fibres. Fibres Report, various issues, available at www.pcifibres.com.

Plastina, A. 2009. “Cotton: Spot Prices Become more Responsive to Futures Prices.” Cotton: Review of the World Situation 63(1): 9-11.

---. 2008. “Speculation and Cotton Prices.” Cotton: Review of the World Situation 61(4): 8-12.

Quantitative Micro Software, LLC. 2007. EViews 6 User’s Guide. California.

Sims, C. 1980. “Macroeconomics and Reality.” Econometrica 48: 1-49.

Sims, C., J. Stock and M.W. Watson. 1990. “Inference in Linear Time Series Models with some Unit Roots.” Econometrica 58: 113-44.

World Bank. 2008. Global Economic Perspectives 2009: Commodities at the Crossroads. Washington DC.