Exploiting the Cointegration Properties of China's Monthly Cotton Import Market and World Apparel Market Conditions: A Preliminary Analysis

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Textile and apparel products have long been among China's main exports. The abolishment of the Multi-Fiber Agreement (MFA) in January 2005 offered additional export growth opportunities for these products originating in China. To meet the derived demand for cotton, it has been predicted that cotton imports would increase too. However, in its WTO accession agreement China actually scheduled import restrictions on cotton in the form of a tariff-rate quota (TRQ). This TRQ serves the purposes of ensuring a certain amount of market access for cotton exporters with a nominal in-quota tariff. At the same time it provides a shield for domestic cotton producers through the high over-quota tariff (40 percent as opposed to the one-percent in-quota rate). When the quota is filled, the TRQ will likely restrict China's access to the imported cotton needed by its textile and clothing industry, thereby undermining the industry's ability to realize the export opportunities provided through the cessation of the MFA. In responding to the end of the MFA, it would be reasonable for the textile and apparel industry to lobby for the reduction of import restrictions on cotton. Indeed, at least for 2005, the Chinese government decided to treat the 2.57 million tons of cotton import as in-quota import, which was scheduled to be only 0.894 million tons (WTO 2006).

Against this background, a set of interesting research questions emerges: to what extent are cotton imports into China linked to the development of world textile and apparel markets? Has China become more responsive to world market conditions following the end of the MFA? In addition, surges of Chinese textile and apparel exports have already caused uneasiness among its trading partners and resulted in temporary trade sanctions. How do these temporary setbacks influence China's import demand for cotton? Answers to these questions will improve the understanding of important policy trends in China and the input-demand behavior of Chinese textile and apparel industry.

To answer the above, an econometric study with copious time-series data on the domestic demand, supply, and stocks of cotton as well as on exports of its downstream products would be desirable, coupled with data on detailed policy changes in China and the multilateral trading system. However, monthly or quarterly time-series data on Chinese agricultural markets with adequate sample points for econometric analyses are rare, and when they do exist are logistically difficult to collect. We procured data on monthly Chinese imports of cotton and the values of such imports from January 2001 through January 2007 (hereafter, 2001:01–2007:01), although extraction of such data were difficult.[1] Other monthly data on Chinese usage and beginning and ending stocks are not available. Nonetheless, the quantities and values of cotton exports do provide the imported quantity and its price proxied by its unit value. Our analysis may be the literature's first monthly econometric analysis on China's cotton import market.

Given our limited sample of Chinese cotton imports and prices, which the evidence below show to be non-stationary in levels, we propose a rather simple cointegrated model of the following monthly series:

- Chinese cotton imports in tons, hereafter denoted as QCHINA
- Price of such Chinese cotton imports proxied by unit values and denoted as PCHINA, since unit import prices were not published.

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[1] These data were obtained from the Ministry of Agriculture of China (http://www.agri.gov.cn). This website only offers the possibility of extracting the data for a single month at a time.

- World apparel price, proxied by the U.S. wholesale price of apparel,[2] denoted as PAPPAREL.

The model provides the following evidence-based results or estimates: market-propelling parameter estimates, the empirical nature of the dynamic interactions among the cotton import market and world apparel market conditions captured by PAPPAREL's behavior, and empirical estimates of market event-specific effects. Our results are perhaps the first monthly econometric estimates of Chinese cotton-market parameters.

The Underlying Statistical Model: Unrestricted Levels VAR and VEC Equivalent[3]

Hendry (1986) notes that potential adverse econometric consequences of failing to utilize information inherent in the modeled endogenous data's non-stationarity elements include compromised inference and spurious regressions. Johansen and Juselius (1990) provide a procedure for exploiting the cointegration properties existing among three or more individually non-stationary variables. This procedure has been widely used and further refined by Juselius (2006) and Juselius and Toro (2005). This procedure's first stage is to fit an unrestricted levels VAR and its unrestricted vector error correction or VEC equivalent to the data, and then utilize the information inherent in many of the non-stationarity properties to avoid the above-discussed adverse econometric consequences.

A VAR model posits each endogenous variable as a function of k lags of itself and of each of the remaining endogenous variables in the system. Tiao and Box's (1978) likelihood-ratio lag-search method, corrected for small samples, was applied and results suggested a two-lag structure. Deterministic and trend components are also added to each equation as the analysis unfolds. Johansen and Juselius (1990) and Juselius (2006, pp. 59–66) demonstrated that the VAR described above may be rewritten compactly and equivalently as the unrestricted VEC

$$(1)\quad \Delta x(t) = \Gamma(1) * \Delta x(t-1) + \Pi * x(t-1) + \Phi * D(t) + \varepsilon(t)\ .$$

The $\varepsilon(t)$ are residuals distributed as white noise. The $x(t)$ and $x(t-k)$, $k = 1, 2$, are $p \times 1$ vectors of the above three variables in current and lagged levels. $\Gamma(1)$ is a $p \times p$ matrix of short-run regression coefficients on the lagged differences, and Π is a $p \times p$ error-correction term to account for endogenous variable levels. The $\Phi * D(t)$ is the set of deterministic variables mentioned above (a trend, seasonals, other permanent shift, and outlier binary variables) added to address the data issues identified as the analysis unfolds. The rank-unrestricted Π or error-correction term is decomposed as

$$(2)\quad \Pi = \alpha * \beta'\ ,$$

where α is a $p \times r$ matrix of adjustment-speed coefficients and β is a $p \times r$ vector of error-correction coefficients.

The $\Pi = \alpha * \beta'$ term is interchangeably denoted as the levels-based long-run component, error-correction term, or cointegration space of the model. The $[\Gamma(1) * \Delta x(t-1), \Phi * D(t)]$ comprises the short-run/deterministic-model component.

Data analysis suggested possible inclusion of a linear trend (TREND) and various permanent shift dummies (presented below) in Equation 2's cointegration space. These same variables in differenced form and a set of 11 centered seasonals were considered for Equation 1's short-run/deterministic component. Analysis will also lead to consideration, where relevant, of outlier dummy variables in the short-run/deterministic component.

Data analysis and consultation with experts on Chinese cotton, textile, and apparel markets led to the following three permanent shift binary variables for possible inclusion:

- ENDMFA, the end of the Multi-Fiber Agreement: it is valued as unity for 2005:01–2007:01 and zero otherwise, and was an extremely important event where quotas on textiles and clothing exports were abolished, having resulted in huge commercial opportunities for Chinese textile and clothing exporters. In responding to these

[2] Specifically, we chose the following to serve as a proxy for apparel price: U.S. producer-price index (PPI), apparel, within the textile products and apparel products group. Series ID WPU0381, from the U.S. Department of Labor, Bureau of Labor Statistics (2007).

[3] This section draws heavily on Juselius (1998; 2006, chapters 1-8) and Juselius and Toro (2005).

export opportunities, Chinese textile and clothing producers increased demand for imported cotton, resulting in over-quota cotton imports at the national level.[4] Perhaps for the purposes of keeping Chinese textile and clothing exports competitive, the Chinese government actually decided not to charge the high over-quota tariff on the over-quota cotton imports for 2005 (WTO 2006).

- TEMPQTAS, the EU's imposition of temporary quotas on certain clothing imports originating in China to avoid serious disruptions to the EU markets. These import quotas are meant to be temporary and are expected to be eliminated in 2008. It is valued as unity for 2005:06–2007:01 and zero otherwise.
- STRANDED, the period after September 2006, when EU and China agreed to deal with "stranded" Chinese products at EU ports and on how these consignments would factor into the temporary quotas noted above. STRANDED is valued at unity during 2005:09–2007:01 and zero otherwise.

The initial starting point for the unrestricted VEC was Equation 1 with no deterministic binary, seasonal, or trend variables. A well-specified unrestricted VEC was ultimately achieved in a series of sequential estimations, judged on evidence from a battery of diagnostic tests. Juselius (2006, chapters 4 and 7) and Juselius and Toro (2005) suggest the following array of diagnostics: (a) trace correlation as an overall goodness of fit, (b) Doornik-Hansen tests for normality of equation residual estimates, and (c) indicators of skewness and kurtosis. A statistically adequate model was achieved in two sets of successive estimations. First we added the trend, centered seasonals, and the three above-specified permanent shift binaries, re-estimated the model, and retained these added variables since the array of diagnostic values suggested improved specification. The second set of sequential estimations further improved specification by having included outlier binaries to account for potentially extraordinary impacts of transitory or observation-specific "outlier" events. An outlier was deemed potentially and extraordinarily influential based on a large absolute standardized residual.[5] The appropriately specified variable was then included in Equation 1's short-run/deterministic component, and ultimately retained if the diagnostic values suggested improved specification. An adequately specified levels VAR and unrestricted VEC equivalent emerged from these sequential estimations with a trend, centered seasonals, three permanent shift binaries, and a single transitory outlier binary.[6]

Space limitations precluded reporting the diagnostics and estimation results of the sequential estimations. However, Table 1 provides diagnostic results of the unrestricted VEC estimated before and after efforts at achieving statistical adequacy, and the results demonstrate the statistical value of such efforts. Table 1 suggests a marked improvement in the model's statistical adequacy from efforts at specification improvement. Efforts to improve specification increased the model's ability to explain data variation by more than 150 percent with the trace correlation, a system-wide goodness-of-fit indicator, having risen from 0.22 to 0.57. The Doornik-Hansen values suggest that all three equations'

[4] For more discussion of these TRQs and China's WTO accession agreement, see Yu and Frandsen (2005).

[5] We followed a procedure for analysis of potentially extraordinary effects of transitory or observation-specific events using "outlier" binary variables (see Juselius 2006, chapter 6). A transitory or observation-specific event was judged as potentially "extraordinary" one if its standardized residual was 3.4 or more. Such a rule for outliers was designed based on the sample size using the Bonferoni criterion: INVNORMAL$(1.0-1.025)^T$ where INVNORMAL is the function of the inverse of the normal distribution function that returns the variable for the c-density function of a standard normal distribution (Estima 2004, p. 503). Note that sample size T equals 73. Here, the Bonferoni variate equals 3.4, and observations with (absolute) standardized residuals of 3.4 or more were considered outliers and transitory or observation-specific binary variables were specified.

[6] The outlier binary was defined as DT2001_0405 and generated a standardized residual of −3.7 as PCHINA markedly declined by about 75 percent in a single month from March 2001 levels, remained low for April–May of 2001, and then recovered in value by 71 percent in a single month to June 2001 levels. In non-differenced form, such a levels-based binary would be valued as unity for 2001:02–2001:05 and zero otherwise. Given this clearly transitory behavior, the differenced form should be included in the short-run/deterministic component of Equation 1. Thus DT2001_0405 in differenced form was defined as unity for 2001:04, −1.0 for 2001:06, and zero otherwise, and placed in Equation 1's short run/deterministic component. This outlier effect likely arose from expectational influences of the then-imminent Chinese admission to the World Trade Organization in November 2001.

Table 1 . Mis-specification Tests for the Unrestricted VEC: Before and After Specification Efforts.

Test and/or equation	Null hypothesis and/or test explanation	Prior efforts at specification adequacy	After efforts at specification adequacy
Trace correlation	system-wide goodness of fit: large proportion desirable.	0.22	0.57
Doornik-Hansen test for normal residuals (univariate)	Ho: equation residuals are normal. Reject for values above 9.2 critical value.		
ΔPAPPAREL		8.5 (p = 0.014)	3.6 (p = 0.13)
ΔQCHINA		0.58 (p = 0.75)	3.5 (p = 0.18)
ΔPCHINA		2.87 (p = 0.24)	1.7 (p = 0.44)
Skewness (kurtosis) univariate values	skewness: ideal is zero; "small" absolute value acceptable kurtosis: ideal is 3.0; acceptable is 3–5.		
ΔPAPPAREL		0.15 (4.3)	0.043 (3.7)
ΔQCHINA		−0.19 (2.9)	0.19 (3.7)
ΔPCHINA		−0.42 (3.4)	−0.06 (3.3)

residual estimates ultimately behave normally, with the initially non-normally behaving apparel-price equation residuals having improved particularly from specification efforts. And finally, skewness and kurtosis indicators achieved acceptable levels for all three equations. Table 1's diagnostics suggest a statistically adequate VAR and its equivalent, an unrestricted VEC model.

Cointegration: Determining and Imposing Reduced Rank on the Error-Correction Space

The three endogenous variables are shown below to be I(1), with first differences being I(0). With cointegration there are from one to two possible linear combinations which are stationary. Cointegrated variables are driven by common trends, and stationary linear combinations (or cointegration) arise when the nonstationarity of one variable corresponds to the nonstationarity in another (Juselius 2006, p. 80).

Equation 1's Π-matrix is a 3 × 3 matrix equal to the product of a p × r matrix, β of long-run error-correction coefficients and a p × r matrix, α, of adjustment-speed coefficients (Johansen and Juselius 1990, 1992). Under cointegration, the rank of Π is reduced ($r < p$), and $\beta'*x(t)$ is I(0) or stationary, even though the three series in x(t) are individually nonstationary. The stationary linear combinations of the individually nonstationary series are the cointegrating vectors/relations (CVs) that render the system as stationary.

Johansen and Juselius' (1990) nested trace tests suggest that reduced rank, or r, is one, with one single CV in the error-correction space.[7] Evidence therefore suggests a reduced rank of $r = 1$ and consideration of a single CV (with t-values in parentheses):

[7] The trace tests values were provided by Dennis' (2006) CATS2 software. The first null hypothesis that r is zero is rejected at the five-precent significance level as the 51.0 trace value exceeds the 47.2 fractile. Evidence at the five-percent level was insufficient to reject the second null that r is less than or equal to 1, since the 17.5 trace value fell below the 31.1 fractile value. Given the nested nature of these tests, evidence suggests that r = 1.

$$(3)\ \ \underset{(25.8)}{QCHINA = 1.002*PCHINA} + \underset{(3.7)}{22.4*PAPPAREL} + \underset{(3.3)}{0.38*ENDMFA} + \underset{(0.20)}{0.03*TEMPQTAS} - \underset{(-1.88)}{0.24*STRANDED} - \underset{(-4.4)}{0.02*TREND}\ .$$

Equation 1 appears to be a global-cotton-supply function to the Chinese market. We acknowledge that a richer data set would be more desirable, with the potential of demand and supply relationships emerging from a richer cointegration space. But our PCHINA and QCHINA variables were the only monthly series located for the Chinese cotton market with enough data points for an econometric analysis. Such richer analyses with more time series must be relegated to future research. We proceed with Equation 3.

Hypothesis Tests and Inference on the Economic Content of the Cointegrating Relation

We begin with Equation 3, the unrestricted CV, and conduct a series of hypothesis tests on the $\Pi = \alpha'*\beta$ or error-correction matrix, then impose the statistically supported restrictions. Johansen and Juselius (1990, pp. 194-206) and Juselius (2006, chapter 10) detail these procedures.

Hypothesis tests on the beta coefficients take the form

$$(4)\ \ \beta = H*\varphi.$$

The β is a p1 × r vector of β-coefficients on variables in the cointegration space; H is a p1 × s design matrix, with s being the number of unrestricted or free beta coefficients; and φ is an s × r = 1 matrix of the unrestricted beta coefficients. Johansen and Juselius' (1990; 1992) well-known hypothesis test value is provided in Equation 5:

$$(5)\ \ -2\ln(Q) = T*\sum[(1-\lambda^*)/(1-\lambda)] \text{ for } i = 1,2\ (= r).$$

The asterisked (non-asterisked) eigenvalues (λ_i, i = 1) are generated by the model estimated with (without) the tested restriction(s) imposed.

We first test for the stationarity of each endogenous variable. We then implement a set of seven exclusion tests on all CV-included variables, followed by a set of hypothesis tests on individual ß-estimates.

We follow Juselius (2006, p. 297) and Juselius and Toro (2005) and use a multivariate likelihood-ratio test within a system setting that depends on rank. This test uses Equation 4 rewritten as

$$(6)\ \ ß^c = [b,\varphi]\ .$$

For each of the three tests of stationarity for each endogenous variable, the β^c is the p1 × r (here, 7 by 1) beta matrix with one of the endogenous variables being tested for stationarity restricted to a unity value (Juselius 2006, p. 183). The b vector is a p1 × 1 (here, 7 × 1) vector: there is a unity value corresponding to the variable being tested for stationarity, zeros for the other two non-tested endogenous variables, and unity for the four deterministic components restricted to the error-correction space (ENDMFA, TEMPQTAS, STRANDED, and TREND). Evidence was sufficient to reject the four null hypotheses that each of the endogenous variables is stationary. [8]

Equations 4 and 5 are used to test if each of the seven variables should be excluded. Evidence rejected the null hypothesis of zero-valued betas for PAPPAREL, QCHINA, PCHINA, ENDMFA, and TREND, and failed to reject the null for TEMPQTAS and STRANDED.[9] The latter two binaries were thus excluded, suggesting a lack of evidence that the temporary quotas and resolution of issues concerning stranded consignments did not have

[8] With four deterministic components retained and the imposed rank of r = 1, then Equation 6's test value is distributed under the null hypothesis of stationarity as a chi-squared variable with two degrees of freedom. Test values with parenthetical p-values are as follows, with the null of stationarity rejected for p-values less than 0.05: 22.8 (p = 0.000) for PAPPAREL, 20.9 (p = 0.000) for QCHINA, and 19.7 (p = 0.000) for PCHINA.

[9] The following are the test values to test the null hypothesis of zero-valued beta estimates (that is, that the variable should be excluded) that are rejected for p-values less than 0.01: 9.13 (p=0.003) for PAPPAREL; 13.6 (p = 0.000) for QCHINA; 15.9 (p = 0.000) for PCHINA; 8.2 (p = 0.004) for ENDFMA; 0.04 (p=0.85) for TEMPQTAS; 0.3.1 (p=0.08) for STRANDED. Clearly, evidence is insufficient to reject exclusion of the latter two variables, TEMPQTAS and STRANDED.

sustained long-run effects**.** This can be explained by the temporary nature of the quotas, as they are meant to be phased out within a short period, according to the legal text specified in the Chinese WTO accession agreement.[10] Chinese apparel producers as well as cotton exporters are likely aware of this and therefore these two binaries play no role in influencing the supply and demand of cotton in the long run.

Re-estimation with exclusion restrictions for STRANDED and TEMPQUOTAS imposed generated a chi-square test value of 3.44 and a p-value of 0.18. With a p-value far above 0.05, evidence at the five-percent significance level is insufficient to reject the cointegrating relation provided below as Equation 7 that appears to be a global cotton-supply function to China.

Discussion

Equation 7 emerged after full restriction with the statistically supported restrictions imposed (Dennis 2006), and appears to be a global cotton-supply function to the Chinese market:

$$(7)\quad \underset{}{QCHINA} = \underset{(28.9)}{1.02*PCHINA} + \underset{(3.7)}{30.0*PAPPAREL}$$

$$+ \underset{(4.0)}{0.0.51*ENDMFA} - \underset{(-3.3)}{0.015*TREND}\ .$$

The estimated adjustment-speed coefficients are as follows, with parenthetical t-values suggesting that they are statistically significant: 0.005 ($t = 3.5$) for PAPPAREL; 0.92 ($t = 3.2$) for QCHINA; and 1.29 ($t = 3.7$) for PCHINA. The parenthetical t-values suggest that the ß- and α-estimates achieved strong statistical strength.

A number of findings emerge from Equation 7 and related results. First, a price-elasticity of cotton export supply to the Chinese market of about unity emerges. Second, conditions in world apparel markets seem influential in China's cotton market, with supply to China positively affected by rising world fashion prices. This result suggests that despite the import barrier in place, the Chinese government is at least flexible enough to relax the restriction when needed to allow the textile industry to respond to world market conditions. Of course, this is likely a reduced-form result without clear demand or supply interpretation possible until, as noted earlier, a richer set of monthly variables can be found and used to define demand and supply forces in the error correction space. Third, the beta estimate on ENDMFA suggests that, on average, the cessation of the MFA (and other concurrent events) resulted in China-bound cotton exports that were about 67 percent higher than before the dismantling of the agreement.[11] This is further evidence suggesting China's ability to adequately respond to changes in the multilateral trading system. Lastly, statistical tests show that the EU's temporary import quotas and the settlement on stranded products failed to appreciably impact cotton supplied to the Chinese market.

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[10] China's WTO Accession Protocol to the WTO allows the EU to apply the textiles-specific safeguard clause if serious market disruption and irreparable damage to EU industry arise. Under the terms of safeguard clause, import restrictions can only be applied until the end of 2008. The measures are designed to give EU industry time to adapt to the changed levels of competition and are not meant to be permanent.

[11] We used Halvorsen and Palmquist's (1980) well-known convention for interpreting regression estimates on binaries in models using logged data. One takes *e*, the base of the natural logarithm, and raises it to the power of the binary coefficient's value, subtracts 1.0, and multiplies by 100. What results is an average percentage-change effect on the dependent variable of the event (and collectively of other concurrent events) for which the binary was defined.

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