WPS 1944

POLICY RESEARCH WORKING PAPER 1944

Detecting Price Links in the World Cotton Market

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Links between international cotton prices have improved in the past decade — in the short run largely because prices are now transmitted more quickly. To a lesser extent, and for different reasons, prices should also converge somewhat more in the long run.

The World Bank
Development Research Group
July 1998

Summary findings

Baffes and Ajwad examine the degree to which international cotton prices are linked and test whether such links have improved over the past decade.

They conclude that the degree of linkage has improved over the past decade, in the short run largely as the result of short-run price transmission — and to a lesser extent because of long-run comovement.

Improvements in information technology have made it much easier for information about demand to be disseminated across markets, so changes in cotton prices attributable to a price shock in one place are soon transmitted to prices in other places.

Moreover, many countries have liberalized their cotton subsectors, and in some countries the government's role has changed substantially.

In East Africa, for example, cotton marketing and trade was handled entirely by government parastatals. Now Tanzania, Uganda, and Zimbabwe have liberalized their marketing and trade regimes, to varying degrees.

In the former Soviet Union (FSU) cotton shipped from Central Asia to other parts of the FSU were considered part of domestic trade. Now cotton exports from Uzbekistan are the most important component of that country's foreign trade.

With such changes, one should expect cotton prices to converge somewhat more in the long run.

Price links between West Africa and Central Asia are much greater than between the United States and other markets — in part because most West African and Central Asian cotton is exported, compared with only 40 percent of U.S. cotton (and 60 percent of Greek cotton). Prices in countries that export most of their cotton are more likely to converge than prices in countries where prices are subject to both domestic and international demand conditions.

To improve price risk management, there should be futures contracts other than those traded on the New York Cotton Exchange, which mostly serves domestic U.S. needs and is not used extensively by non-U.S. hedgers and speculators.

This paper — a product of the Development Research Group — is part of a larger effort in the group to investigate the behavior of world prices. Copies of the paper are available free from the World Bank, 1818 H Street NW, Washington, DC 20433. Please contact John Baffes, room MC3-545, telephone 202-458-1880, fax 202-522-1151, Internet address jbaffes@worldbank.org. July 1998. (36 pages)

DETECTING PRICE LINKS IN THE WORLD COTTON MARKET

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The authors would like to thank Alain d'Hoore and Uma Lele for valuable comments and suggestions on an earlier draft.

I. INTRODUCTION

In the absence of impediments, the comparative advantage argument of trade theory dictates that resources will be allocated in an efficient manner, in turn implying that factor and product prices in different locations will be equalized – subject to transfer costs. Under certain conditions, the existence of strong price linkages, therefore, may be viewed as a necessary requirement for efficient allocation of resources and hence maximum welfare [Samuelson (1952); Takayama and Judge (1964)]. This paper focuses on the degree of price linkages of the world market of cotton.

The issue of price linkages in product markets both at national and international levels has been studied in the literature rather extensively either under the notion of the law of one price [e.g. Protopapadakis and Stoll (1983, 1986), Ardeni (1989), Baffes (1991)] or under the notion of market integration [e.g. Ravallion (1986), Sexton, Kling, and Carman (1991), Gardner and Brooks (1994), Fafchamps and Gavian (1996), Baulch (1997a)]. Moreover, reflecting on the market liberalization and structural adjustment efforts undertaken by a number of developing countries in recent years, the degree to which markets are integrated has been used quite often as a yardstick in assessing the success of policy reforms [e.g. Goletti and Babu (1994), Alexander and Wyeth (1994), Gordon (1994), Dercon (1995)].

As many authors have cautioned, however, price convergence does not necessarily imply efficient allocation of resources unless the setting in which trade takes place is competitive [e.g. Faminow and Benson (1990), Baulch (1997b)]. For example, consider the extreme case of two duopolists who agree to charge the same price in two segmented markets. While convergence in prices (whenever price changes occur) would take place instantaneously, the oligopolistic setting of the market may not necessarily allocate resources in the most efficient manner. The same argument may be advanced for a number of developing countries where parastatals assign panterritorial and panseasonal prices on certain commodities. In such cases the law of one price holds by definition without necessarily implying that resources are allocated efficiently.

The present paper examines the strength of price linkages in the world market of cotton. In pursuing this objective, the paper contributes to the literature of price linkages in two respects. On the theoretical side, it introduces a measure of price linkage and also identifies its source (i.e. short-run price transmission *versus* long-run comovement.) On the empirical side, it applies this measure to the world market of cotton for two different time periods, thereby examining whether improvements in price linkages have taken place over the last decade.

There are at least two reasons as to why one would expect that price linkages may have improved over the last decade. First, improvements in information technology have made it much easier for information on demand conditions to be disseminated across markets; therefore one would expect that cotton price changes from one origin due to a demand shock would be transmitted immediately to the price of the other origins. Second, many countries have undertaken steps to liberalize their cotton subsectors while in other countries the role of the government has been substantially altered. For example, under the Former Soviet Union (FSU) structure, cotton from Central Asia shipped to other parts of the FSU was domestic trade. Currently, cotton exports from Uzbekistan constitute the single most important component of its foreign trade. Changes have also taken place in Africa. For example, until the early 1990s, cotton marketing and trade in East African countries was handled in its entirety by government parastatals. Now, Uganda, Zimbabwe, and Tanzania, operate – to different degrees – within liberalized marketing and trade regimes. One, therefore, would expect faster long-run convergence of cotton prices (if convergence existed) or at least some convergence (if convergence did not exist in the first place).

The remainder of the paper proceeds as follows. In the next section the model along with the explicit measure of the degree of market linkage is outlined. In discussing the model, we undertake a rather extensive literature review on the subject of price linkages. The review indicates that a number of commonly used models are, in fact, restricted versions of the same dynamic specification. The penultimate section discusses the data and presents the empirical findings. The data cover the periods August 1985 through December 1987 and August 1995 through January 1997 and refer to CIF prices in North European ports for US, Greek, West African, and Central Asian types of cotton. The findings indicate that price linkages between W. Africa and C. Asia have been much higher than between the US and the other markets. In fact, in the first period, no comovement between the US and the other markets was detected. The last section concludes by addressing some policy implications and subjects for further research.

II. DETECTING PRICE LINKAGES

Earlier studies examining the relationship between prices either have looked at correlation coefficients [e.g., Lele (1967); Southworth, Jones, and Pearson (1979); Timmer, Falcon, and Pearson (1983); Stigler and Sherwin (1985)] or have used the following type of regression [e.g. Isard (1977), Mundlak and Larson (1992), Gardner and Brooks (1994)]:[1]

$$p_t^1 = \mu + \beta_1 p_t^2 + \varepsilon_t, \tag{1}$$

where p_t^1 and p_t^2 denote prices from two origins of the commodity under consideration, μ and β_1 are parameters to be estimated while ε_t denotes an IID(0, σ^2) term. The hypothesis that the slope coefficient equals unity and (possibly) the intercept term equals zero can be tested; formally, H_0: $\mu + 1 = \beta_1 = 1$. Under H_0 the deterministic part of (1) becomes $p_t^1 = p_t^2$, in turn implying that the price differential, $p_t^1 - p_t^2$, is an IID(0, σ^2) term.

Estimating (1) and testing H_0, while intuitively appealing and computationally implementable, presents two fundamental shortcomings. First, some statistical properties of the series involved in (1), namely nonstationarity, may invalidate standard econometric tests and thus give misleading results regarding the degree to which price signals are being transmitted from one market to another. Second, in primary commodity markets with characteristics such as (small or even perceived) differences in quality, high transfer costs relative to the price, etc., it is rather unlikely that the two prices will only differ by an IID(0, σ^2) term as H_0 of (1) dictates. Therefore, H_0 is expected to be rejected without necessarily ruling out a relatively high degree of price linkage. Consequently, it is deemed necessary to employ a general enough model that imposes no *a priori* requirements on the stationarity properties of the series in question and at the same time allows for some degree of flexibility.

With respect to the nonstationarity problem one can examine the order of integration of the error term in (1) and make inferences regarding the validity of the model (Ardeni, 1989). If prices are indeed nonstationary, the existence of a stationary error term implies comovement between the two prices. However, if the slope coefficient is different from unity, the uniqueness of the cointegration parameter in the bivariate case implies that the corresponding price differential would be growing and such growth would not be accounted for, although prices may move in a seemingly synchronous manner. Hence, stationarity of the error term of (1) (given non-stationary prices) while establishing proportional price movement, should not be considered as a testable form equivalent to that of the H_0 of (1). Note that a number of authors have warned against interpreting non-unity slope coefficient as a sign of market integration (e.g. Barrett (1996)).

To account for the non-unity slope coefficient one can restrict the parameters of (1) according to H_0, in which case the problem is equivalent to testing for a unit root in the following univariate process (Engle and Yoo, 1987):

$$(2) \qquad \left(p_t^1 - p_t^2\right) \sim I(0).$$

If the price differential as defined in (2) is stationary, then one can conclude that price signals are transmitted from one market to another, in the long run. The assumption (or finding) that the cointegration parameter is unity is very crucial, as it ensures that there is no other nonstationary component entering the system. As Meese (1986) and West (1987) observe, the absence of cointegration (with unity slope coefficient in the present setting) can be attributed to omitted nonstationary variables, in turn implying that an additional component would have to be included in (2) in order to fully account for the variability of the price differential.

As a sidelight, it should be emphasized that if the cointegration parameter is unity, it is immaterial for all relevant aspects of the analysis whether (1) or (2) is employed. This is the case because as the sample size increases, regression (1) should yield β_1 equal to unity. However, in finite samples this may not be necessarily the case. For example, Ardeni (1989), using (1) in logarithms for a number of internationally traded primary commodities, found that the corresponding error term was not stationary, thus rejecting the law of one price. Baffes (1991), on the other hand, by using the same data set found that in the majority of cases the price differential was stationary, hence providing supportive evidence for the law of one price as a long run relationship.[2]

From the preceding discussion, it is rather evident that cointegration tests are not very powerful as they only make inferences about the existence of the moments of the distribution of $(p_t^1 - p_t^2)$ and not about certain restrictions that may be required by economic theory [e.g. H_0 of (1)]. Therefore, (2) cannot serve as a substitute for the H_0 of (1); it can only serve as an intermediate step in establishing its validity.

With respect to the restrictive nature of (1), one can circumvent it by introducing a more general autoregressive structure. Appending one lag to (1), gives:[3]

$$(3) \qquad p_t^1 = \mu + \beta_1 p_t^2 + \beta_2 p_{t-1}^2 + \beta_3 p_{t-1}^1 + u_t,$$

where u_t is IID$(0, \sigma^2)$ and $|\beta_3| < 1$. Despite its simplicity, (3) encompasses a wide variety of commonly used dynamic models with different economic interpretations. For example, Hendry, Pagan, and Sargan (1983) discuss a number of testable hypothesis, results of corresponding restrictions on the parameter space of (3). The most important one is the long-run proportionality or homogeneity hypothesis, the validity of which ensures that price movements in one market (p_t^2) will *eventually* be transmitted to the prices of the other market (p_t^1). Such a hypothesis can be tested by restricting all slope parameters of (3) to sum to unity (i.e., $\Sigma_i \beta_i = 1$).

Under long-run proportionality, (3) can be re-parameterized as follows:

$$(4) \qquad \left(p_t^1 - p_{t-1}^1\right) = \mu + (1-\beta_3)\left(p_{t-1}^2 - p_{t-1}^1\right) + \beta_1\left(p_t^2 - p_{t-1}^2\right) + u_t.$$

Relationship (4) belongs to the family of error-correction models (ECM). Because of the equivalence of the existence of cointegration and ECM, stationarity of the price differential (2) implies the existence of (4) (in the sense that $(1 - \beta_3)$ is significantly different from zero) and *vice-versa* [see Appendix A for a formal proof of this argument as well as the steps involved in going from (3) to (4)]. Note that the restriction $|\beta_3| < 1$

implies that $0<1-\beta_3<2$. The sign of β_3 (positive *versus* negative), or alternatively whether $(1 - \beta_3)$ falls between zero and one or between one and two indicates whether the convergence monotonic or oscillatory.

The main feature of (4) is the economic interpretation of its parameters: β_1 indicates how much of a given price change in the price of the commodity in location 2 will be transmitted to location 1 within the first period (referred to as initial adjustment, short-run effect, or contemporaneous effect); $(1 - \beta_3)$ indicates how much of the price difference between the two prices is eliminated in each period thereafter (referred to as error-correction, speed of adjustment, or feedback effect). The coefficient of the short-run effect can, in theory, take any value. The adjustment coefficient, however, is restricted between zero and two. The closer to unity is $(1 - \beta_3)$, the higher the speed at which convergence will take place. Symmetric with respect to unity values of $(1 - \beta_3)$ [e.g. 0.75 and 1.25] indicate that the adjustment speed will be the same but the adjustment path will differ [monotonic in the former and oscillatory in the latter case].

It is worth reemphasizing here that $(1 - \beta_3)$ different from zero is a necessary and sufficient condition for long-run convergence. Conversely, significantly different from zero β_1 is neither a necessary nor a sufficient condition for long-run price convergence; even if $\beta_1 = 1$ (i.e. perfect short-run adjustment) the series may still drift apart in the long run – unless $(1 - \beta_3)$ is significantly different from zero, in which case the series will converge.[4]

Further restrictions on (4) give a number of alternative testable hypotheses. For example, letting $\beta_1 = 0$, (4) becomes:

$$\text{(5)} \qquad \left(p_t^1 - p_{t-1}^1\right) = \mu + (1-\beta_3)\left(p_{t-1}^2 - p_{t-1}^1\right) + u_t .$$

The interpretation of (5) is that while no adjustment is taking place in the current period, prices indeed converge in the long-run with speed $(1 - \beta_3)$.

On the other hand, letting $\beta_1 = 1$, (4) is re-parameterized as follows:

$$\text{(6)} \qquad \left(p_t^2 - p_t^1\right) = -\mu + \beta_3\left(p_{t-1}^2 - p_{t-1}^1\right) + u_t .$$

(6) implies that while price changes in one market are transmitted exactly to prices in

the other market in the short-run, long-run price convergence may be achieved at a lower speed, depending on the size of β_3. Furthermore, if in addition to letting $\beta_1 = 1$ one imposes $\beta_3 = 0$, (6) collapses to the H_0 of (1). A number of studies examining the linkages between futures and cash prices for commodity and asset markets have employed (5) and (6) [e.g., Garbade and Silber (1983); Schroeder and Goodwin (1991); Wang and Yau (1994); Fortenbery and Zapata (1997)].

Alternatively, setting $\beta_3 = 1$ in (4), effectively implying non-convergence, gives:

(7) $$\left(p_t^1 - p_{t-1}^1\right) = \mu + \beta_1\left(p_t^2 - p_{t-1}^2\right) + u_t.$$

Relationship (7), often termed the first difference approach, has also been used in the literature rather extensively [e.g., Richardson (1978), Tomek (1980), Leavit, Hawkins, and Veeman (1983), Hudson, Ethridge, and Brown (1996)].[5] First differences along with detrending have been traditionally the most widely used filters in time series analysis. It should be noted, however, that if $(p_t^2 - p_t^1)$ and $(p_t^2 - p_{t-1}^2)$ are orthogonal, estimating short-run and dynamic adjustment effects through (5) or (7) will yield the same parameter estimates as estimating them jointly through (4).[6]

Finally, by restricting $\beta_2 = \beta_3 = 0$, (3) reduces to (1), which as indicated earlier has been one of the most commonly used models in the literature of price linkages. To these models one should also add Granger-causality tests since a significantly different from zero error-correction term implies Granger-causality with feedback from p_t^2 to p_t^1 [Petzel and Monke (1980) used Granger-causality for the international rice market]. Lastly, if the prices have not been adjusted by transfer costs, one can incorporate them in either (3) or (4) – depending on their stationarity properties – and examine their dynamic effects on prices. At this point it is straightforward to verify that a wide variety of 'law of one price', market integration, or market efficiency models can be derived by (3) (possibly with a higher lag structure) subject to the appropriate restrictions on the parameters.[7]

From the preceding discussion is has become evident that one potential problem with estimating (5), (6), (7), or even (4) for that matter, without ensuring that the appropriate restrictions (or orthogonality conditions where applicable) hold, is that the estimated parameters may be biased and hence give misleading results regarding the underlying economic behavior. For example, $\beta_1 = 1$ in (7) may erroneously lead one to conclude that there exists strong correlation between the two prices where in fact, the prices may not even converge in the long run. On the other hand, (5) imposes the restriction that no adjustment in the short-run takes place, which may not necessarily be the case.

The model outlined above suggests that, given long-run proportionality exists, whether to choose (3) or (4) to recover short- and long-run dynamic price behavior is a matter of stationarity properties. If prices are stationary, (3) would be the preferred structure and long-run proportionality could be tested by restricting the slope parameters to sum to unity. Under non-stationarity, (4) would be the preferred structure and long-run proportionality can be tested by examining the stationarity properties of the price differential (Engle and Yoo, 1987) or equivalently by testing whether $(1-\beta_3)$ is different from zero (Phillips and Loretan, 1991). Then, with appropriate tests one can determine whether any of the underlying restrictions implied by (5), (6), or (7) can be in fact validated by the data.

Having established long-run proportionality and also having recovered the parameter estimates of (4) the next task is to transform the information contained in the parameter space in such a way so that a succinct interpretation of both short-run and feedback effects (and hence price linkage) can be given. Stating the question in a simplified manner: *How long does it take for the price of cotton from origin 1 to adjust to a given price change in origin 2?*

Let n be the period by which k percent of the cumulative adjustment has taken place. In the current period, $n = 0$, k takes the value of β_1 [also equal to $1-(1-\beta_1)$], which is the short-run impact of $(p_t^2 - p_{t-1}^2)$ on $(p_t^1 - p_{t-1}^1)$. In the next period, $n = 1$, k takes the value of $\beta_1+(1-\beta_1)\beta_3$, which is the impact of the previous period, β_1, plus the feedback effect, $(1-\beta_1)\beta_3$ [it can also be written as $1-(1-\beta_1)(1-\beta_3)$]. For $n = 2$, k takes the value of the previous period's adjustment, $\beta_1+(1-\beta_1)\beta_3$ plus $(1 - \beta_3)(1-\beta_1-(1-\beta_1)\beta_3)$ [which can be written as $1-(1-\beta_1)(1-2\beta_3+\beta_3^2)$ or $1-(1-\beta_1)\beta_3^2$]. (Table 1 gives the adjustment for the first five periods, including the current one). Hence, the cumulative adjustment at period n

is given by:

$$k = 1 - (1 - \beta_1)\beta_3^n \qquad (8)$$

Alternatively, solving for n in (8) gives the number of periods required to achieve a certain level of cumulative adjustment, i.e. $n = [log(1\text{-}k) - log(1\text{-}\beta_1)]/log\beta_3$. For values of β_1 and $(1 - \beta_3)$ close to unity, a small n (number of periods) is required for the adjustment to be completed (*i.e.* k close to unity).[8]

Although most models in the literature of price linkages have based the discussion on estimated model parameters and *F*-tests, there have been a number of exceptions which have quantified the linkage in a more explicit manner.[9] Ravallion (1986), for example, using an autoregressive distributed lag model and appropriate parameter restrictions made a distinction among market segmentation, long-run market integration, and short-run market integration and applied it to the rice market in Bangladesh. A number of researchers have applied Ravallion's formulation since then [e.g. Palaskas and Harriss (1993) applied it to the food markets in West Bengal while Gordon (1994) applied it to grain markets in Tanzania].

Timmer (1987) introduced a measure of price linkage based on the unrestricted version of specification (3) and applied it to the corn market of Indonesia. Timmer's Index of Market Connection, converted to the parameters of (4) [i.e. after imposing long-run convergence] is equal to $\beta_3/(1\text{-}\beta_3)$. For values of β_3 close to unity (or alternatively for values of $(1 - \beta_3)$ close to zero), the index takes large values (at the limit approaching infinity), in turn indicating weak price linkage. If $\beta_3 = 0$ (or $1 - \beta_3 = 1$) the index takes the value of zero, which corresponds to error-correction coefficient being equal to unity, consequently indicating strong price linkage – for negative values of β_3 the measure falls within the interval (-0.5, 0). Heytens (1986) and Alderman (1992) used Timmer's measure of market connection to examine the performance of food markets in Nigeria and Ghana, respectively.

Delgado (1986) developed a variance components methodology by making a distinction among different levels of market integration between harvest and post-harvest period and applied it to food markets in Northern Nigeria. Goodwin and Schroeder (1991) in examining spatial price linkages in US regional cattle markets used the magnitudes of cointegration statistics of bivariate regressions as measures of price linkages. More recently, Lutz, van Tilburg, and van der Kamp (1995) introduced a measure of market integration by calculating short- and intermediate-run impact multipliers and then a measure of adjustment (similar to the one proposed in this paper); they applied the model to wholesale and retail food markets in Benin.

III. DATA AND RESULTS

a. Data and Stationarity Tests

Two samples, one covering the period August 15, 1985 to December 24, 1987 (122 observations) and a second covering the period August 3, 1995 to January 9, 1997 (73 observations) were constructed. Thursday price quotations from the following origins were used: US (Memphis Territory), Greece, Central Asia, and African 'Franc Zone' (referred to as W. Africa).[10] In addition to the four quotations, we also included the A Index, the measure of the 'world' price of cotton. However, because the A Index may contain the price which it is paired with, any results related to it should be interpreted with caution. Appendix B contains a detailed discussion of the data. Given that the number of cotton traders in North Europe is sufficiently high to ensure that cotton prices are determined within a competitive environment and therefore any degree of price linkage can be attributed to market efficiency.

To determine the order of integration the augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP) procedures were utilized. The ADF is based on the following regression: $(p_t - p_{t-1}) = \mu + \beta p_{t-1} + lags(p_t - p_{t-1}) + \varepsilon_t$, where p_t denotes the series under consideration (Dickey and Fuller, 1981). A negative and significantly different from zero value of β indicates that p_t is $I(0)$. The PP test is similar to the ADF; their difference lies on the treatment of any nuisance serial correlation aside from that generated by the hypothesized unit root (Phillips and Perron, 1988; Phillips, 1989). To identify the presence of one unit root we test H_0: p_t *is not* $I(0)$ <u>against</u> H_1: p_t *is* $I(0)$. Trend stationarity can be detected by appending a time trend in the relevant regression. Finally, the significance level of the error-correction coefficient itself, $(1 - \beta_3)$, can serve as cointegration test (Phillips and Loretan, 1991).

Stationarity results for both periods are reported in the upper panel of Table 2. The tests indicate that stationarity in levels is rejected in all cases. The middle panel of Table 2 reports results for trend stationarity tests. Here the picture changes considerably since, with the exception C. Asia in the second period, all tests show evidence of trend stationarity.[11]

Because the decision on whether to use (3) or (4) ultimately depends on the stationarity properties of the prices, we supplemented the unit root tests with a variance-ratio test (Cochrane, 1988). This test is based on the statistic defined as $(1/k)Var(p_t-p_{t-k})/Var(p_t-p_{t-1})$, where p_t is the variable of interest and k denotes the lag length; it exploits the fact that the variances of conditional forecasts explode for nonstationary series and converge for stationary (or trend stationary) series as the forecast horizon grows. The idea behind Cochrane's test goes as follows. If p_t is a random walk [i.e., $p_t = \mu + \rho p_{t-1} + \varepsilon_t$, where $\rho = 1$], the variance of its k-differences grows linearly with k, i.e. $Var(p_t-p_{t-k}) = k\sigma_\varepsilon^2$. If, on the other hand p_t is stationary or trend stationary, the variance of its k-differences will eventually approach zero. As a consequence, in the former case $(1/k)Var(p_t-p_{t-k})$ will remain constant at σ_ε^2 as k grows – possibly after an initial jump if ρ is greater than one – while in the latter case it will approach zero – slowly for values of ρ close to but less than one. Dividing by $Var(p_t-p_{t-1})$ (which is independent of k) normalizes the first period to unity.

Figures 1a through 1e provide information regarding the unit root status of the price series under consideration in the form of the variance-ratio statistics. Undoubtedly, the pattern of all variance-ratios is explosive in both periods, therefore pointing to the fact that we are dealing with non-stationary price series.[12] In what follows we proceed under the assumption that prices are non-stationary – an assumption generally consistent with findings in the literature of the subject.

The lower panel of Table 2 reports stationarity statistics of the price differential, a measure of the degree of comovement between pairs of cotton prices. Note that because the cointegration parameter is assumed rather than estimated, the same critical values are used for both levels and price differentials – if the parameter was to be estimated through OLS, more 'demanding' (i.e. higher in absolute levels) critical values would have been used.

Consider first the A index. When compared to the US in period 1 no comovement appears to be in place, while a high degree of comovement is present in the second period, a result reflected in both tests. A very small improvement is detected for A Index-Greece, where the level of significance increases from 5% and 10% in the first period to 1% and 5% in the second period. The A Index-W. Africa price differential, while stationary in period 1, it is non-stationary in period 2. The link between the A Index and the remaining two prices, however, appears to be weakening in period 2. In terms of variance-ratio statistics (depicted in figures 2a and 2j), while for the A Index-US case the pattern is explosive in both periods, it converges at a rather slow rate for the remaining three cases, with no distinguishable pattern between the two periods.

The degree of comovement of prices increased substantially in Greece, W. Africa, and C. Asia, when coupled with the US. In most cases, stationarity statistics more than doubled and in all but one case they exceeded the 5% significance level. However, the variance ratio statistics for US-W. Africa and US-C. Asia, indicate a non-stationary price differential in both periods (more so in the latter than the former case). Comparing Greece with W. Africa and C. Asia, the comovement sharply deteriorates according to stationarity statistics but the differential is stationary in both cases as the variance-ratio statistic indicates. Finally, for W. Africa-C. Asia, while the statistics become lower in absolute value, they are still significant at the 5 and 10% level and in both periods stationary.

To conclude, results from the lower panel of Table 2 indicate that, excluding the A Index, price linkages in the cotton market improved relative to the US but a deterioration was detected among some non-US markets. Although these results are robust with respect to both stationarity tests (PP and ADF), they are in contrast to what was expected, i.e. that improvement should have taken place or at least no deterioration should have been observed. The variance-ratio statistics, however, indicate that while small changes may have taken place between the first and second period, in no case the stationarity properties have been altered for either levels or differentials, as the ADF and PP statistics pointed in a number of cases. Therefore, the error-correction term is expected to yield more insights on the long-run convergence issue and especially regarding the validity of the variance-ratio *versus* ADF and PP stationarity tests.

b. *Goodness of Fit*

Model (4) was estimated for both periods and a Chow test was employed to determine whether the parameters of period 1 were significantly different from those of period 2. The χ^2 testing procedure proposed by Hansen (1982) and White (1980) was utilized to estimate the covariance matrix consistently. Initially we estimated (4) with four lags and subsequently we kept the significant ones.[13] With the exception of six cases, the significance of the higher order lags was very low.

To assess the overall performance of the model, we first examine the goodness of fit (Table 3).[14] Given that (4) can be re-parameterized in terms of current and lagged price differentials as well as one of the two (also current and lagged) price differences (Campbell and Shiller, 1987), one can view the R^2 as a measure of basis risk (i.e. the unpredictable movements in the basis) where basis is defined as the difference between the two prices rather than its traditional definition as the difference between cash and futures price of the same commodity. Then, the lower the R^2 the higher the basis risk and *vice-versa*. The R^2 has been used in the literature extensively as a measure of basis risk [e.g. Lindahl (1989) and Faruqee, Coleman, and Scott (1997)].

With one exception, the R^2 has improved considerably in all cases. On average, about 50% of the price variability from one origin was explained by the variability of another origin's price in period 1. In period 2 the average explanatory power of the model increased to 75%. Excluding the A Index, the relative increase in the explanatory power of the model becomes even greater (from 40% to 71%).[15] Thus, with the evidence at hand, price linkages within cotton markets appear to have improved substantially over the last decade. In what follows we examine whether such result holds if further measures are applied and also identify and quantify the sources of such improvement.

c. *Quantifying Price Linkages*

The upper and middle panels of Table 4 report the adjustment taking place within the first period. A coefficient of one would be interpreted as a perfect transmission of price shocks, while a coefficient of zero represents a short-run invariance to changes in prices elsewhere. Since the short-run effect is in principle unrestricted, β_1 greater than unity for example, would suggest an over-reaction to changes in prices in the current period. The lower panel contains the *p*-values of the hypothesis of equality in the β_1s in the two sub-sample periods, against the two-sided alternative.

At the 5% significance level, six of the nine overall improvements in the short-run effect were significant, while only three of the eight remaining cases represented significant reductions in the amount of adjustment within the first period.[16] Further analysis of the nine significant changes in the short-run effect reveals that the average deviation of the adjustment coefficient from unity fell from 0.32 to 0.25, indicating an overall improvement in the initial adjustment. More specifically, Greece showed the most improvement in the short-run adjustment when coupled to the A Index, US and C. Asia. W. Africa and C. Asia revealed signs of improvement when paired with Greece, while the opposite was true when paired with the US.

The measure of long-run comovement is presented in Table 5, with the upper and middle panels representing period 1 and 2, respectively. In essence, the measure of long-run adjustment captures the correction to a given price change from another origin, subsequent to the current period. In fact, the absolute deviation from the long-run steady-state declines from period to period (i.e., suggesting long-run convergence in prices) when this parameter is statistically different from zero. The lower panel of Table 5 reports the *p*-values for the test of the hypothesis that the dynamic adjustment effect remained the same against the two-sided alternative. Note that for the cases where the error-correction parameter is not significant, specification (7) (i.e. the first difference model) is the valid characterization of the data.

Twelve improvements were observed, while declines in the degree of comovement were present in three cases. The remaining five cases revealed no appreciable change between periods 1 and 2. Significant improvements in the long-run effect were observed when Greece was coupled with A Index and W. Africa at the 2% and 6% levels of significance. All other changes in the measure of long-run comovement between the two sub-samples are not significant at conventional levels (on average the adjustment coefficient increased from 0.08 to 0.11). This result is congruent with the variance-ratio findings (depicted in figures 2a-2j), which do not detect any change in the stationarity properties of the price differentials between the two periods.

Table 6 presents the number of weeks, n, required to achieve 95% of the adjustment to a given price change. Note that n is calculated using equation (8) and, as was mentioned earlier, it is only meaningful when long-run comovement in the Engle-Granger sense is detected. Faster adjustment is observed in fourteen cases while a slower adjustment is observed in only two. Except Greece- US in period 2, with the US as a reference, it is clear that none of the other origins exhibited convergence towards the price levels in the US, a fact which becomes apparent when the insignificant error-correction coefficient is considered.

However, Table 6 reveals that nine of fourteen changes in the number of periods required to be within 5% of complete adjustment, were significant at the 7% level. Hence, price shocks were transmitted at higher speed in period 2 compared with period 1. Additionally, in period 1, nine cases of non-convergence were evident while only three cases appeared in period 2 at the 10% level of significance, indicating that prices from more origins achieved long-run convergence in the second period.

One methodological note is in order. It was argued earlier that a number of studies have used correlation coefficients to examine price linkages. How much does one loose in terms of informational content by using correlation coefficients? Table 6 reports correlation coefficients for levels and first differences. The results indicate that correlation coefficients in levels are in no way capable of detecting the improvement that has taken place. In fact, of the ten cases, eight indicate reduction of linkages in the second period, a result contrary to *a priori* expectations. The picture changes considerably when first differences are considered; with the exception of one case, a substantial increase is detected, similar in direction to the R^2 criterion reported earlier. However, the magnitude of the increase is less than that of the R^2; such result was expected since R^2 also captures improvements in long-run convergence. For example, the correlation coefficients increase by an average of 25% (from 0.69 to 0.86) as opposed to the 50% improvement in R^2. To conclude, therefore, correlation coefficients, properly calculated by considering stationarity properties, capture the short-run effect, but are incapable of capturing long-run comovement.

IV. Summary and Concluding Remarks

This paper examined the degree to which price linkages in cotton markets have improved over the last decade. Weekly data from August 15, 1985 to December 24, 1987 (122 observations) and August 3, 1995 to January 9, 1997 (73 observations) from US, Greece, Central Asia, and West Africa were utilized. Price linkages between Central Asia and West Africa have been the highest in both periods, while the linkages between the US and other markets were non-existent in the first period.

According to the goodness of fit criterion (i.e. the R^2) in almost all cases a substantial improvement in price linkages has taken place. For example, while on average about 50% of the variability of price from one origin was explained by the variability of another origin's price in period 1, the variability explained in period 2 increased to 75%. Moreover, if one excludes the A Index, the relative increase in the explanatory power of the model is even higher (from 40% to 71%).

A number of interesting conclusions emerge from this paper, both policy related and methodological. First, the main source of this improvement in price linkages appears to be a result of short-run price transmission and to a very limited extent a result of long-run comovement. To the degree that short-run price transmission reflects demand conditions while long-run convergence reflects supply conditions, the findings of this paper suggest that over the last decade, information on demand changes have been reflected in price changes much faster now than a decade earlier.

A second conclusion relates to the relatively high long-run convergence between C. Asia and W. Africa observed in both periods (with an estimated adjustment coefficient at 0.17) and the non-existence of convergence between the US and the three other origins (especially in the first period). Cotton produced in W. Africa and C. Asia is exported almost in its entirety, hence making both markets subject to the same world demand conditions (prices respond only to world demand since domestic demand is practically non-existent). On the contrary, only 40% of US cotton is exported while the corresponding figure for Greece is 60% (1996/97 averages), in turn making their respective prices subject to both domestic and world demand conditions. This finding not only reinforces warnings cited in the literature that volume of trade may affect the conclusions regarding the degree of price linkage, but also indicates that exports relative to the size of the domestic market may be an important factor determining the degree of price linkage.

The third finding reflects on a methodological issue. It has been extensively argued in the literature of time series that conventional stationarity tests exhibit low power and may give misleading results regarding the true degree of comovement. This study confirmed this, i.e. stationarity tests by themselves may be incapable of uncovering the comovement. Additional measures, such as Cochrane's variance-ratio tests or Hamilton's advise of looking at the overall sensibility of the results should be used to appropriately assess the presence (or absence) of price linkage.

The results of this paper have also important implications with respect to price risk management. Low comovement between US and non-US cotton prices implies that there is a need for a futures contract other than the one currently traded at the New York Cotton Exchange (NYCE) which is the only contract currently traded (apart from the São Paulo contract at the Brazilian commodity exchange introduced in 1996, albeit with an extremely low liquidity). The NYCE contract serves primarily domestic US needs and is not being used extensively by non-US hedgers and speculators (Lake, 1992). This is not surprising if one considers that in December 31, 1990, the May 1991 contract closed at 76.19 cents, 8.21 cents below the A Index while it expired on May 8, 1991 at 92.22 cents, 8.92 cents above the A Index, a results which is very similar for the individual components of the A Index as this paper indicates.

The need for a futures exchange for non-US hedging needs, has been apparent as noted by *Cotton Outlook* (December 12, 1997, p. 3) which reported: "The lack of an international trading instrument other than the No. 2 [i.e. NYCE] contract – one which consistently reflects broad world cotton market developments but is capable of being used as 'hedge' – continues to be a shortcoming of the current pricing system." An attempt, however, to create a 'world' futures contract by NYCE in 1992 failed.

On the other hand, slow price convergence suggests that a non-US cotton contract is unlikely to attract business from cotton merchants other than the ones that are interested in that particular style of cotton and therefore may be expected to succeed only on national or regional basis.

Finally, a note on further research. One important issue not considered here is endogeneity. For policy related reasons, one would have to first, detect any endogeneity patterns and correct for them through an instrumental variables model.

Endnotes

[1] See Harriss (1980) for a comprehensive (and critical) review of the literature on market integration studies undertaken in the 1960s and 1970s.

[2] Ardeni (1989) and subsequently Baffes (1991) used quarterly averages for the following commodities: wheat, tea, beef, sugar, wool, zinc, and tin. In a more recent study, Zanias (1993) examined the degree of spatial market integration in European Community agricultural product markets by using both unrestricted and restricted versions [i.e. (1) and (2)], without detecting appreciable differences due to specification.

[3] The number of lags to be included in (3) is an empirical question. To make the exposition clear, we only use one lag.

[4] Although the fact that the short-run effect is 1 while there may be no long-run convergence seems counter-intuitive it should not be surprising. Consider the following thought experiment: two series are generated as $p_t^2 = (-1)^{trend}(0.75) + \varepsilon_t$ and $p_t^1 = p_{t-1}^1 + (-1)^{trend}(1.5) + 0.5 + \varepsilon_t$, where *trend* denotes time (1, 2, 3, ...) and ε_t is a white noise. p_t^2 oscillates between ±0.75 (1.5 unit swing) and p_t^1 rises by 2 in one period and falls by 1 in the next period. On average, p_t^1 also demonstrates a swing of 1.5. Estimation of (3) gives a short-run effect of one. However, it is clear that the two series are diverging over time. On the other hand, if $p_t^1 = 1 + \varepsilon_t$, the short-run adjustment is effectively zero, i.e. changes in p_t^2 are completely innocuous to changes in p_t^1 while the error-correction coefficient is one.

[5] Relationship (7) corresponds to case (c) of Table 2.1 of Hendry, Pagan, and Sargan (1984) while (5) and (6) correspond to cases (i) and (g) of their Table 2.2.

[6] On this issue, Kennedy (1992) notes that $(p_t^2 - p_t^1)$ and $(p_t^2 - p_{t-1}^2)$ are " ...closer of being orthogonal than the variables in the original relationship [i.e. (3)]" (p. 264).

[7] Baulch (1997b) discusses in detail the following four specifications: The level/cointegration version of the law of one price [specification (1)]; the first difference version of the law of one price [specification (7)]; the autoregressive distributed lag/error-correction [specifications (3) and (4), possibly with a higher lag structure]; and Granger-causality patterns [inferred from specifications (3) and (4)].

[8] Strictly speaking, (8) should read as $k = (1-\beta_1)|\beta_3|^n$, since as was discussed earlier, the restriction in (3) reads $|\beta_3|<1$, not $0<\beta_3<1$.

[9] As it will be shown later, R^2 and correlation coefficients can also be viewed as measures of price linkages since both fall within the (0, 1) range, with zero corresponding to no linkage and 1 corresponding to perfect linkage.

[10] The four countries/regions considered in the sample account for one third of world exports. If one excludes Argentina and Australia (the two dominant southern hemisphere exporters), they account for more than 85% of world exports. With respect to the sample period, it would have been desirable to have continuous sample throughout the entire decade. However, these were the only large enough subperiods for which the price series were uninterrupted.

[11] The number of observations in the second period span 1.5 years as opposed to the first period which span 2.5 years. It is likely, therefore, that the trend stationarity result of the ADF and PP tests reflects short sample (despite the high frequency of the data).

[12] Hamilton (1994) emphasizes the difficulty in distinguishing truly non-stationary processes from processes that are stationary but persistent. He also suggests comparing estimates obtained under alternative specifications and choose the one that performs better according to several criteria, which is the avenue we pursue in the present case. Following Hamilton and in view of the trend stationarity evidence of the second period, we also estimated the model in levels (i.e. (3) with a trend variable) by testing and subsequently imposing long-run proportionality. In all cases the model exhibited R^2s close to unity and extremely high t-ratios. In no case we found statistically significant difference in the estimates of periods 1 and 2 as the performance of the model was indistinguishable in the two periods. Cochrane (1988) also points to the difficulties of parametric tests in distinguishing between true random walk models and trend stationary models with a small random walk component, which appears to be the case in the period 2.

[13] The lag structure of the six cases (all in the first period) was as follows: Greece-US{2,3}, W. Africa-US{0,3}, C. Asia-US{0,3}, A Index-US{0.3}, W. Africa-Greece{1,0}, and W. Africa- A Index{0,3}. For consistency, we retained identical lag structure in the second period.

[14] The complete set with estimation results is available from the authors upon request.

[15] Because the A Index may contain one of the prices under consideration by definition, it is expected that the regression on the A Index will exhibit superior performance. That explains why when the A Index models are excluded, the average R^2 declines in both periods.

[16] At this point we should clarify the fact that symmetry of the short-run coefficient with respect to unity is interpreted as an equal departure from perfect short-run transmission. For example, the values 1.25 and 0.75 are being viewed as exhibiting the same deviation from unitary short-run elasticity.

TABLE 1: Cumulative Adjustment

Period (n)	**Amount of Cumulative Adjustment (k)**	
0	β_1	$= 1 - (1 - \beta_1)\beta_3^0$
1	$\beta_1 + (1 - \beta_1)(1 - \beta_3)$	$= 1 - (1 - \beta_1)\beta_3^1$
2	$1 - (1 - \beta_1)\beta_3 + (1 - \beta_3)(1 - \beta_1)\beta_3$	$= 1 - (1 - \beta_1)\beta_3^2.$
3	$1 - (1 - \beta_1)\beta_3^2 + (1 - \beta_3)(1 - \beta_1)\beta_3^2$	$= 1 - (1 - \beta_1)\beta_3^3$
4	$1 - (1 - \beta_1)\beta_3^3 + (1 - \beta_3)(1 - \beta_1)\beta_3^3$	$= 1 - (1 - \beta_1)\beta_3^4$

Notes: β_1 and β_3 refer to the parameters of equation (4).

Source: Calculated by the authors.

TABLE 2: Stationarity Tests

	Period 1		*Period 2*	
	ADF	*PP*	*ADF*	*PP*
Levels w/o trend				
A Index	-1.24	-0.70	-1.18	-0.84
US	-1.40	-1.10	-1.38	-1.53
Greece	-1.26	-0.78	-1.18	-0.96
W. Africa	-1.33	-0.72	-1.11	-0.67
C. Asia	-1.30	-0.75	-1.24	-0.86
Levels w/ trend				
A Index	-2.42	-2.03	-3.51***	-3.44**
US	-1.83	-1.55	-4.80***	-4.84***
Greece	-2.55	-2.27	-2.87*	-3.08**
W. Africa	-2.56	-2.07	-3.71***	-3.58***
C. Asia	-2.80*	-2.35	-2.38	-2.59*
Price Differentials				
A Index - US	-1.71	-1.40	-3.11**	-3.26**
A Index - Greece	-2.94**	-2.82*	-3.58***	-3.16**
A Index - W. Africa	-3.70***	-3.77***	-2.48	-2.41
A Index - C. Asia	-3.39**	-3.53***	-2.97**	-2.65*
US - Greece	-1.87	-1.71	-2.81*	-3.03**
US - W. Africa	-1.74	-1.46	-3.42**	-3.55***
US - C. Asia	-1.68	-1.28	-3.31**	-3.36**
Greece - W. Africa	-3.02**	-2.96**	-2.49	-2.35
Greece - C. Asia	-2.85*	-2.80*	-2.26	-2.40
W. Africa - C. Asia	-3.87***	-3.82***	-3.00**	-2.87*

Notes: One (*), two (**) and three (***) asterisks indicate significance at the 10%, 5%, and 1% levels. Critical values are: -2.58 (10%), -2.89 (5%), and -3.51 (1%) (Fuller, 1976).

Source: Estimated by the authors.

TABLE 3: Goodness of Fit

	A Index	*US*	*Greece*	*W. Africa*	*C. Asia*
R^2 in Period 1					
A Index	–	0.62	0.40	0.80	0.88
US	0.50	–	0.18	0.32	0.43
Greece	0.46	0.34	–	0.44	0.44
W. Africa	0.83	0.51	0.41	–	0.72
C. Asia	0.88	0.54	0.35	0.71	–
R^2 in Period 2					
A Index	–	0.73	0.84	0.81	0.87
US	0.74	–	0.62	0.73	0.76
Greece	0.85	0.55	–	0.62	0.80
W. Africa	0.81	0.70	0.61	–	0.75
C. Asia	0.79	0.72	0.79	0.74	–

Notes: Goodness of fit is the R^2 of equation (4).

Source: Estimated by the authors.

TABLE 4: Short-run Effect

	A Index	*US*	*Greece*	*W. Africa*	*C. Asia*
Estimate of β_1 in period 1					
A Index	–	0.48	0.54	0.94	0.83
US	0.98	–	0.49	0.81	0.80
Greece	0.70	0.32	–	0.67	0.56
W. Africa	0.77	0.33	0.43	–	0.72
C. Asia	1.05	0.49	0.58	1.00	–
Estimate of β_1 in period 2					
A Index	–	0.49	0.85	1.05	0.93
US	1.48	–	1.22	1.68	1.50
Greece	1.00	0.48	–	0.99	0.97
W. Africa	0.77	0.40	0.64	–	0.72
C. Asia	0.94	0.59	0.82	1.02	–
Test of equality of β_1 between the two periods: p-values					
A Index	–	0.81	0.00	0.16	0.17
US	0.00	–	0.00	0.00	0.00
Greece	0.00	0.16	–	0.02	0.00
W. Africa	0.99	0.51	0.06	–	0.92
C. Asia	0.13	0.88	0.02	0.80	–

Notes: All reported coefficients are significant at the 1% level. p-value is the significance level of the F-statistic of the hypothesis that β_1 in (4) is the same in the two periods.

Source: Estimated by the authors.

TABLE 5: Dynamic Adjustment

	A Index	*US*	*Greece*	*W. Africa*	*C. Asia*
Estimate of (1 - β_3) in period 1					
A Index	–	0.01	0.01	0.20***	0.13***
US	0.02	–	0.04**	0.03*	0.02
Greece	0.14***	0.02	–	0.16***	0.17***
W. Africa	0.21***	0.02	0.01	–	0.17***
C. Asia	0.14***	0.02	0.02	0.17***	–
Estimate of (1 - β_3) in period 2					
A Index	–	0.01	0.22***	0.16***	0.13***
US	0.10**	–	0.12***	0.09**	0.08**
Greece	0.25***	0.06	–	0.16***	0.11**
W. Africa	0.12**	0.00	0.11**	–	0.17***
C. Asia	0.11**	0.02	0.08*	0.17***	–
Test of equality of (1 - β_3) between the two periods: p-values					
A Index	–	0.92	0.02	0.71	0.96
US	0.12	–	0.15	0.21	0.14
Greece	0.22	0.43	–	0.99	0.51
W. Africa	0.17	0.54	0.06	–	0.91
C. Asia	0.61	0.98	0.44	0.92	–

Notes: One (*), two (**) and three (***) asterisks indicate significance at the 10%, 5%, and 1% levels. *p*-value is the significance level of the *F*-statistic of the hypothesis that (1 - β_3) in (4) is the same in the two periods. Note that for the cases where the adjustment term is significantly different from zero, there exists causality with feedback in Granger's sense (Granger, 1969).

Source: Estimated by the authors.

TABLE 6: Number of Periods Required to Achieve 95% of the Cumulative Adjustment

	A Index	*US*	*Greece*	*W. Africa*	*C. Asia*
Estimate of n in period 1					
A Index	–	n.c.	n.c.	0.8	8.7
US	n.c.	–	54.8	46.4	n.c.
Greece	12.3	n.c.	–	11.2	11.4
W. Africa	4.6	n.c.	n.c.	–	9.6
C. Asia	0.4	n.c.	n.c.	0.0	–
Estimate of n in period 2					
A Index	–	n.c.	4.5	0.0	2.9
US	22.3	–	11.5	26.3	27.0
Greece	0.0	n.c.	–	0.0	0.0
W. Africa	9.6	n.c.	2.6	–	9.0
C. Asia	1.5	n.c.	16.2	0.0	–
Test of equality of n between the two periods: p-values					
A Index	–	n.a.	0.00	0.32	0.37
US	0.00	–	0.00	0.00	0.00
Greece	0.01	n.a.	–	0.06	0.00
W. Africa	0.03	n.a.	0.01	–	0.99
C. Asia	0.30	n.a.	0.07	0.95	–

Notes: p-value is the significance level of the F-statistic of the hypothesis that $(1 - \beta_3)$ and β_1 in (4) (and hence k and n) is the same in the two periods. n is calculated as $[log(0.05) - log(1-\beta_1)]/ log\beta_3$ – a result of setting $k = 0.95$ and solving (8) for n. 'n.c.' indicates that long-run convergence never takes place as the error-correction parameter is not significantly different from zero; 'n.a.' indicates that the test is not reported because the respective prices did not converge (for both cases 10% level of significance was the cut-off point).

Source: Estimated by the authors.

TABLE 7: Correlation Coefficients

	A Index	*US*	*Greece*	*W. Africa*	*C. Asia*
Period 1 **(bold** for levels - NE triangle, *italics* for first differences - SW triangle**)**					
A Index	–	**0.87**	**0.98**	**1.00**	**1.00**
US	*0.72*	–	**0.90**	**0.86**	**0.87**
Greece	*0.64*	*0.43*	–	**0.98**	**0.98**
W. Africa	*0.88*	*0.58*	*0.61*	–	**0.99**
C. Asia	*0.93*	*0.67*	*0.60*	*0.83*	–
Period 2 **(bold** for levels - NE triangle, *italics* for first differences - SW triangle)					
A Index	–	**0.86**	**0.99**	**0.98**	**0.98**
US	*0.86*	–	**0.81**	**0.88**	**0.81**
Greece	*0.91*	*0.78*	–	**0.95**	**0.97**
W. Africa	*0.89*	*0.85*	*0.78*	–	**0.97**
C. Asia	*0.93*	*0.87*	*0.89*	*0.84*	–

Notes: The NE triangular section of each panel reports correlation coefficients for price levels while the SW triangular section reports correlation coefficients for first differences.

Source: Estimated by the authors.

Figure 1a: A Index

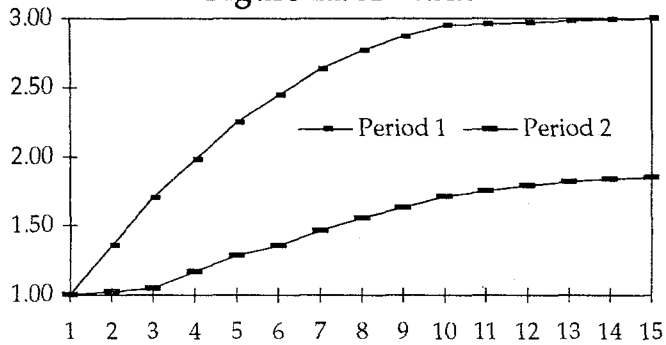


Figure 1b: US

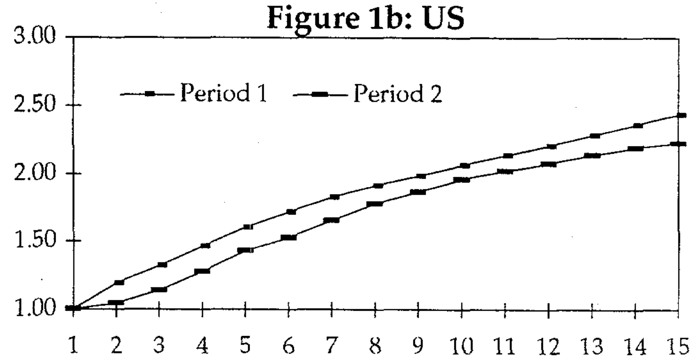


Figure 1c: Greece

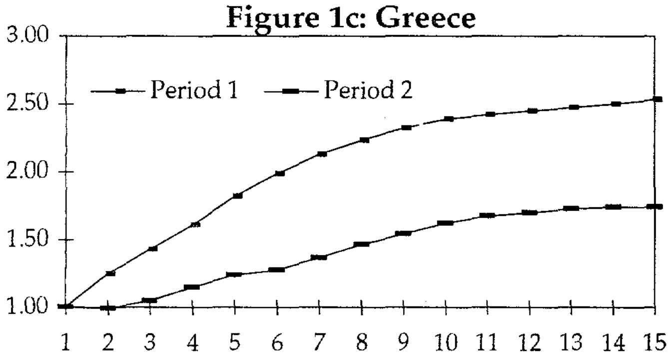


Figure 1d: W. Africa

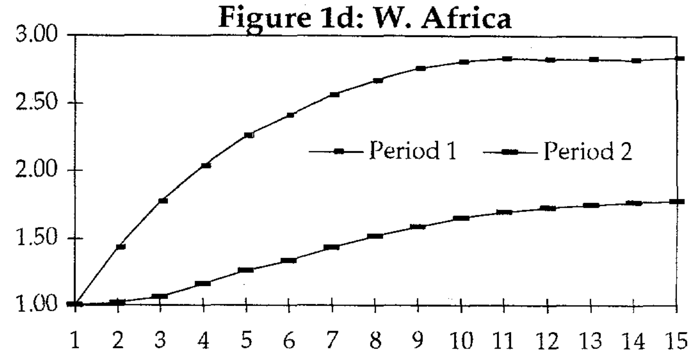


Figure 1e: C. Asia

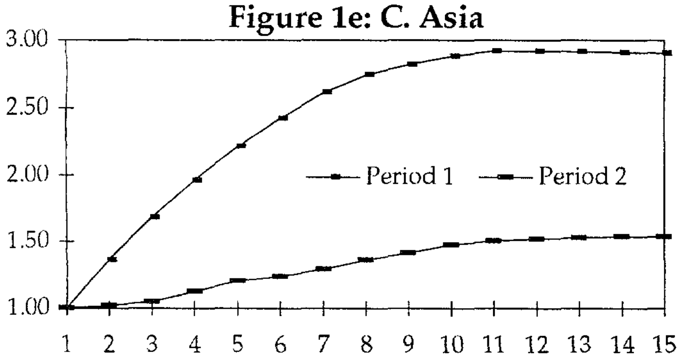


Figure 2a: A Index - US

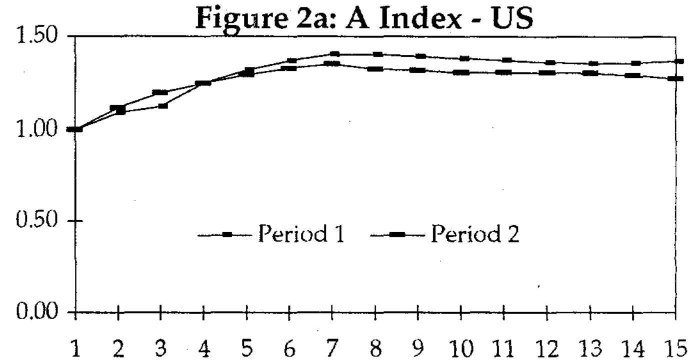


Figure 2b: A Index - Greece

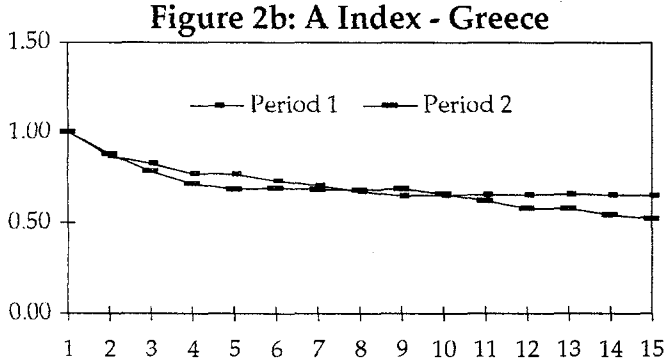


Figure 2c: A Index - W. Africa

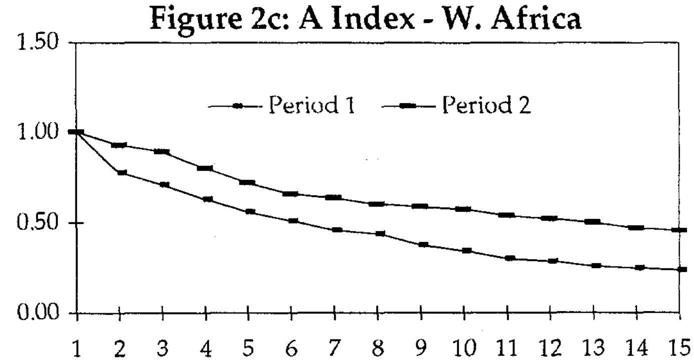


Figure 2d: A Index - C. Asia

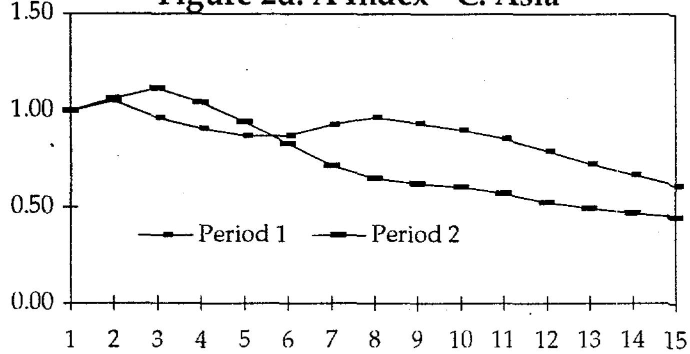


Figure 2e: US - Greece

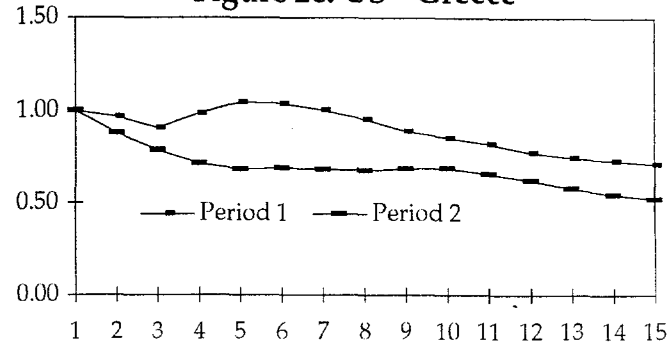


Figure 2f: US - W. Africa

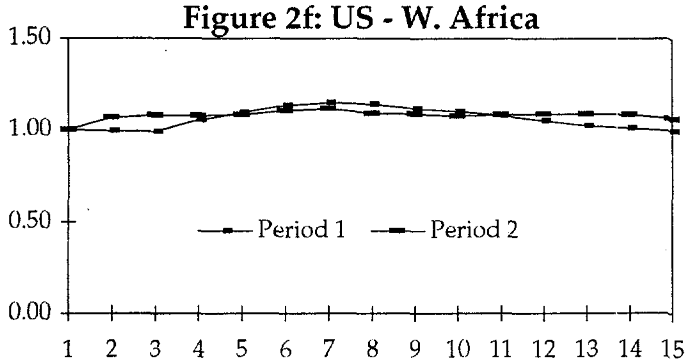


Figure 2g: US - C. Asia

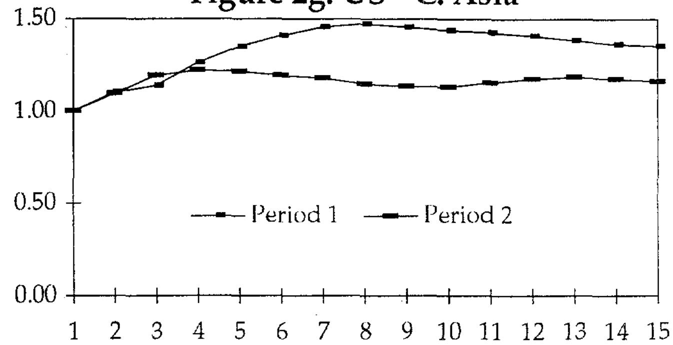


Figure 2h: Greece - W. Africa

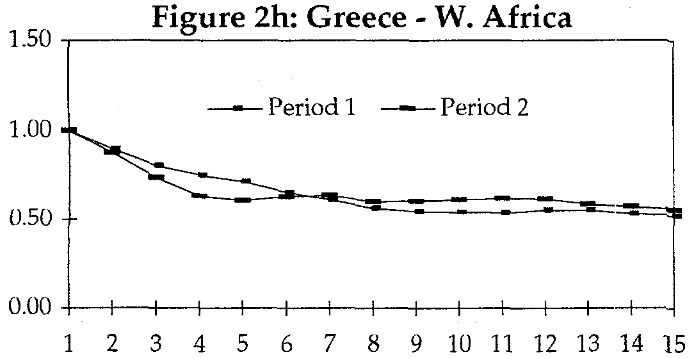


Figure 2i: Greece - C. Asia

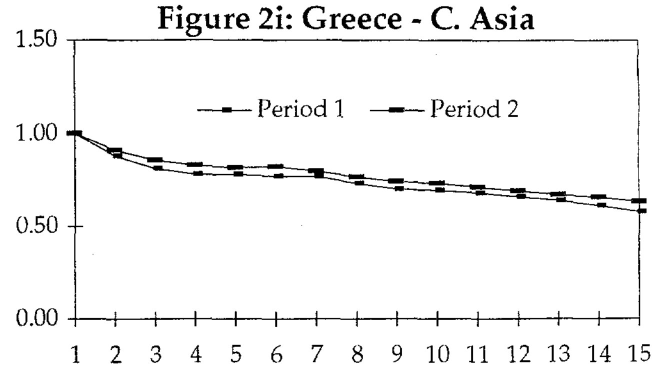
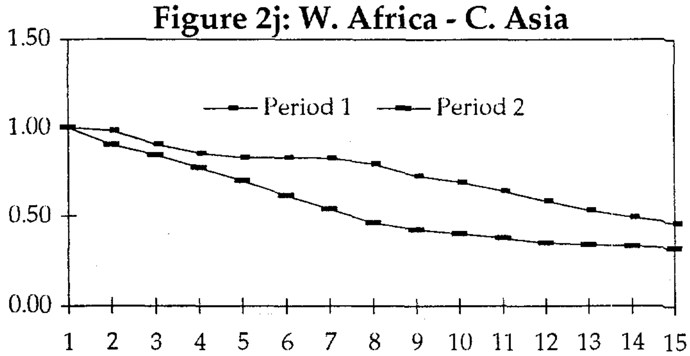


Figure 2j: W. Africa - C. Asia



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Appendix A

PROPOSITION: *Examining long-run proportionality in (3) by testing if $\Sigma_i\beta_i = 1$ when p_t^1 and p_t^2 are I(0), is equivalent to testing for stationarity of the price differential [i.e. $(p_t^1 - p_t^2) \sim I(0)$] when p_t^1 and p_t^2 are I(1).*

Proof: Consider (3) in text again,

$$\text{(A1)} \qquad p_t^1 = \mu + \beta_1 p_t^2 + \beta_2 p_{t-1}^2 + \beta_3 p_{t-1}^1 + u_t,$$

where u_t is IID(0, σ^2) and $|\beta_3|<0$. To solve for the long-run equilibrium set $p_t^1 = p_{t-1}^1$ and $p_t^2 = p_{t-1}^2$ in (A1). Then, the deterministic part of (A1) (excluding the constant term) becomes $(1 - \beta_3)p_t^1 = (\beta_1 + \beta_2)p_t^2$ or $p_t^1 = [(\beta_1 + \beta_2)/(1-\beta_3)]p_t^2$. If p_t^1 and p_t^2 are I(0), the hypothesis of long-run proportionality can be examined by running regression (A1) and then testing H_0: $(\beta_1 + \beta_2)/(1 - \beta_3) = 1$ (or alternatively $\beta_1 + \beta_2 + \beta_3 = 1$).

Setting $\beta_2 = 1 - \beta_1 - \beta_3$ in (A1) gives:

$$\text{(A2)} \qquad p_t^1 = \mu + \beta_1 p_t^2 + p_{t-1}^2 - \beta_1 p_{t-1}^2 - \beta_3 p_{t-1}^2 + \beta_3 p_{t-1}^1 + u_t,$$

Subtracting p_{t-1}^1 from both sides of (A2) and collecting terms results in:

$$\text{(A3)} \qquad \left(p_t^1 - p_{t-1}^1\right) = \mu + (1-\beta_3)\left(p_{t-1}^2 - p_{t-1}^1\right) + \beta_1\left(p_t^2 - p_{t-1}^2\right) + u_t.$$

If p_t^1 and p_t^2 are I(1), then (A3) is a valid error-correction model if and only if β_3 is significantly different from unity (or alternatively $(1-\beta_3)$ is significantly different from zero) (Engle and Granger, 1987), which because of the equivalence between error-correction representation and the existence of cointegration implies that the term in the parenthesis following $(1-\beta_3)$ of (A3) is stationary [i.e. $(p_t^2 - p_t^1) \sim I(0)$]. ■

Appendix B: The 'World' Price of Cotton and its Components

Typically, the world price of a commodity is taken to be the spot price prevailing at a certain location where a substantial part of trade is taking place. While often this location is in a key producing country (e.g. US for maize and Thailand for rice) this may not always be the case (e.g. New York for coffee and London for tea). On the other hand, for several commodities there is more than one dominant trade location (e.g. wheat in US, Canada, and Australia or wool in New Zealand and the UK).

Cotton departs from this tradition in that its 'world' price is not a spot price at which actual transactions take place in one or more locations; instead it is an index, calculated as an average of offer quotations by cotton agents in North Europe. The index is constructed daily by Cotlook Limited, a private information dissemination company based in Liverpool, UK and is published in the weekly magazine *Cotton Outlook.*

The cotton price index (often referred to as the Cotlook A Index or simply the A Index) is an average of the five less expensive out of 14 styles of cotton (Middling 1-3/32'') traded in North Europe originating from: (1) Memphis Territory (US); (2) California/Arizona (US); (3) Mexico; (4) Paraguay; (5) Turkey; (6) Syria; (7) Greece; (8) Central Asia (until June 1997)/ Uzbekistan (since then); (9) Pakistan; (10) India; (11) China; (12) Tanzania; (13) Africa 'Franc Zone'; and (14) Australia (see Table B1 below for a detailed example). These are offer prices, i.e. the price that the agent would quote for the particular type of cotton. To account for the fact that agent's quotation is likely to be above the price at which the actual transaction takes place (since the buyer will probably negotiate a lower price), the index takes the five lowest priced styles.

Some styles are consistently traded with premiums or discounts compared to the A Index. For example, cotton from the US is usually traded above the A Index, cotton from Uzbekistan is traded below the A Index, while that from Africa 'Franc Zone' is traded very close to the A Index. Since not all styles of cotton from the eligible origins are traded in North Europe all year around, not all 14 quotations are available for the A Index at all times. Therefore, the frequency at which a certain cotton style participates in the formation of the A Index depends on whether it is continuously traded in North

Europe and on how inexpensive it is. When less than five eligible quotations are available, the A Index is not reported. This may happen when the Southern Hemisphere runs out of supplies while the Northern Hemisphere is not ready to supply cotton (for example, this was the case in June/July 1995).

TABLE B1: Composition of the Cotlook A Index

	Price of cotton (US ¢/lb.)		
Origin of Quotation	*October 17, 1996*	*January 2, 1997*	*July 31, 1997*
(1) US (Memphis Territory)*	83.00	84.00	85.25*
(2) US (California/Arizona)	83.00	84.00	87.50 N
(3) Mexico	79.75	NQ	NQ
(4) Paraguay	NQ	NQ	NQ
(5) Turkey	NQ	NQ	NQ
(6) Syria	NQ	79.50@	79.50@
(7) Greece*	74.00@	79.00@	NQ
(8) Central Asia*	71.00@	75.50@	79.50@
(9) Pakistan	76.00@	NQ	NQ
(10) India	NQ	NQ	NQ
(11) China	NQ	NQ	NQ
(12) Tanzania	78.00@	NQ	NQ
(13) Africa 'Franc Zone'*	75.00@	79.00@	80.00@
(14) Australia	80.50	84.00@	87.00@
COTLOOK A INDEX	**74.80**	**79.40**	**82.25**

Notes: 'NQ' indicates that cotton from the respective origin was not traded in North Europe. '*' denotes the quotations analyzed in this study. '@' denotes the A Index quotations over which the average is taken. 'N' means that the particular style of cotton is not offered in volume and hence it is not used in the composition of the Index.

Source: Cotton Outlook.

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