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PRICE CONVERGENCE ON WORLD COMMODITY MARKETS: FACT OR FICTION?

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This article examines the degree to which commodity prices have converged on world commodity markets over recent decades. Ideally, increases in communications, central bank activities, and globalization would suggest that commodity prices in spatially dispersed markets should become similar over time. To measure convergence, correlation, regression, cointegration, and vector autoregressive methods are employed. Comparable geographic data were assembled for six commodities: coffee, cotton, wheat, lead, copper, and tin, covering the period 1930 through 1998. Overall, the empirical results do not support the convergence hypothesis but rather a pattern of fluctuating divergences.

Keywords: *commodity markets; price convergence; spatial price behavior*

The issue of price convergence on commodity markets both at national and international levels has been suggested by Baffes and Ajwad (2001) to fall under either the notion of the law of one price (e.g., Protopapadakis and Stoll 1986; Ardeni 1989; Goodwin 1989; Miljkovic 1999) 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 1997). Other attempts to examine price integration include Goletti and Babu (1994), Alexander and Wyeth (1994), Gordon (1994), and Dercon (1995). Studies that have emphasized the importance of market competition for convergence can be found in Faminow and Benson (1990) and Baulch (1997).

There are several reasons as to why one would expect that price linkages might have improved over the past half-century. First, gains in information technology have made it much easier for information on demand and supply conditions to be disseminated across markets; therefore, one might expect that commodity price changes from one origin due to a quantity shock would be transmitted immediately

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to prices in other origins. Second, the centralization of commodity markets particularly for metal commodities has increased. In recent years, the London Metal Exchange (LME) has become the major world market for many metals, suggesting that the localization of transactions might cause metal prices to converge over time. Third, central bank activities in developing countries have increased, and globally more coordination between central bank activities can be seen. Fourth, economic globalization such as through the organization of common markets has caused greater congruency between national and international business cycles.

Finally, some countries have undertaken steps to liberalize their production subsectors, while in other countries, the role of the government has been substantially altered. For example, under Former Soviet Union (FSU) rules, cotton shipped from Central Asia to other parts of the FSU was considered domestic trade. Currently, cotton exports from Uzbekistan, for instance, 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 would expect, therefore, a faster long-run convergence of cotton prices (if convergence existed) or at least some convergence (if convergence did not exist in the first place).

In this article, we focus on the degree of spatial price convergence in world commodity markets for coffee, cotton, wheat, copper, tin, and lead using time series and cross-sectional analysis. In pursuing this objective, the article contributes to research on 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 tests the convergence hypothesis using commodity price series covering the period 1930 through 1998. This long period has been selected to permit the variety of possible converging forces to come into play to determine whether improvements in price linkages have taken place.

The remainder of the article is organized as follows. The first section presents the background to spatial price linkage in theory and practice. The second section identifies the commodity price series and interprets their behavior. The third section explains the convergence methodologies we employ. Finally, the empirical findings are given in the fourth section, and conclusions then follow.

Background to Price Linkages

Spatial Price Relations

In his paper on spatial competition, Hotelling (1929) opened the way to the development of imperfect competition and spatial price discrimination theories.

Later on, Hoover (1948) and Isard (1956) advanced spatial price analysis in their classic theoretical works on location and space economy. Subsequently, the process of spatial price arbitrage was formalized as equilibrium among more than two markets by Enke (1951) and Samuelson (1952).

Spatial price relations also are the subject of international trade theory. In developing the Heckscher-Ohlin principle, Samuelson (1948, 1949) formulated the factor price equalization theorem, which states that unless initial factor endowments are too unequal, commodity mobility will always be a perfect substitute for factor mobility. In other words, an increased level of market integration for basic commodities is an essential condition for achieving free mobility of factors of production. Since then, there have been a number of attempts to model the impact of variables such as distance, location, and transport cost on markets and flows (Roehner 1996; Baulch 1997; Fafchamps and Gavian 1996).

Another relevant aspect of trade theory involves the law of one price. Initially intended to examine the closeness of prices of traded goods, it has been applied more recently to integration among primary commodity markets and prices. Such studies have been reviewed in Miljkovic (1999), and examples include Ardeni (1989), Baffes (1991), and Goodwin, Grennes, and Wohlgenant (1990).

The Concept of Market Integration

In recent years, the analysis of market and regional integration has led regional economists and economic historians to examine price convergence; examples include the European Historical Workshop on Market Integration (1993), the Eleventh International Economic History Congress (1994) session on market integration, and various issues of the *Journal of Common Markets*. Different definitions exist of market integration, but Roehner (1995a, 1995b) has reduced these to two alternative conceptions: (1) In the first one, a region (or a market) is said to be integrated if "enough" arbitragers are present in the markets and if they are acting "efficiently" in a sense that supposes a number of conditions such as, for instance, the requirement of perfect information. In this conception, a market either is integrated or it is not; there is no room for a measure to reflect a certain degree of integration. (2) In the second conception, the degree of market integration is identified with the level of intermarket price differentials (or some equivalent variable). If these differentials are large (in relative terms), then the market is poorly integrated; if on the contrary they are small, the market would be well integrated.

The first conception has its origin in financial markets, and one must examine whether this concept of an efficient market can be transposed to commodity markets. If so, one should in particular be able to give an operational meaning to the notion of perfect information. A clear operational criterion of an efficient market would consist of observing that price differentials do not exceed transactions and transport costs. While information on transactions costs is easily available in financial markets, this is not necessarily true for transportation costs in commodity markets. Information on freight rates very often is not made public, though United Nations Conference on Trade and Development (UNCTAD) (2001) has attempted to make limited data available. In contrast to financial markets where only a few parameters are required, a commodity contract involves many parameters such as quality of the product (within a given grade) and specifications regarding storing, transportation, loading and discharging, and so on. Almost none of these parameters are usually made public. To summarize the case for international commodity markets separated over space, it is very difficult to define market efficiency with a clear operational criterion. Statistical investigation, therefore, has to rely on the second conception.

Evidence of Integration

At the international level, empirical evidence on price convergence has been sought primarily in the context of the law of one price (Goodwin 1989; Goodwin and Schroeder 1991; Drame et al. 1991; Roehner 1995b; Protopapadakis and Stoll 1986). This law maintains that the foreign price of a commodity once adjusted for exchange rates and transportation costs will be equal to the domestic price of the commodity. Once prices are converted to a common currency, the same commodity will sell for the same price in different countries. While empirical investigations of the law of one price typically have been pursued at a general level, as tests of purchasing power parity using aggregate data and price indexes, Goodwin (1989) and Goodwin and Schroeder (1991), among others, have considered tests using individual agricultural commodity prices and quantities. Their contribution was to show that price linkages can be temporal as well as spatial, emphasizing the dynamics of expected prices. Based on U.S. commodity prices taken relative to selected foreign market prices, they found some support for price linkages, but this depended on the use of a variable rate of discounting and variable transportation costs.

Among other empirical approaches for evaluating integration, several of these apply simple correlation measures to long-run historical price series. For example, Persson (1994) discovered some evidence of integration in European grain (wheat, rye, barley) markets between the fifteenth and nineteenth centuries. So did Froot (1995) dealing with some seven centuries of data. Also, Drame et al. (1991) and Roehner (1995b) examined spatial price differentials for grain markets in France and across some European countries in the nineteenth century. It is worth mentioning that some of the work on global price integration has been conducted among different commodities rather than among different markets. That is, tests have been performed on the tendency of different primary commodity prices to move together. Earlier research by Labys and Perrin (1976) rejected this possibility, but Pindyck and Rotemberg (1990) and Labys, Achouch, and Terraza (1999) have found some evidence based on tests of the comovement hypothesis.

Finally, most recent empirical research on market integration has emphasized the application of cointegration analysis between spatially separated prices. While several examples exist of domestic market integration, fewer exist at the international level. Examples of such applications are explored in the next section.

In this study, market and price integration are evaluated to the extent that primary commodity prices are compared across markets or exchanges that exist in different countries. Because these markets reflect the competitive paradigm, no attempt is made to explain the influence of geographic differences among them.

THE PRICE DATA

Typically, the world price of a commodity is taken to be the spot price prevailing at a certain market or location where a substantial part of trade is taking place. While often this location is in a key producing country (e.g., the United States for maize or Thailand for rice), this may not always be the case (e.g., New York for coffee or London for copper). On the other hand, there may be more than one major market and major trading country for each commodity (e.g., wheat in the United States, Canada, and Australia; or wool in New Zealand and the United Kingdom). Much as this is true for most commodities, others such as cotton depart from this tradition. The cotton "world" price in particular 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.

Markets for commodities such as grains and cotton, where the proportion of international trade is lower relative to aggregate world production, show a greater divergence in relative prices due to the greater influence of local market demand. A certain amount of integration, however, has been found for metal prices. For example, Labys, Achouch, and Terraza (1999) showed the conformity of certain metal price cycles to business cycles using dynamic factor analysis. In this study, our interest is in comparing prices emerging from different competitive markets for the same commodity. In some cases, such market activity is "hidden" and trade prices (CIF/FOB) are used, though they may be contaminated by shipping costs.

Furthermore, despite the valuation of many commodities in U.S. dollars, only a small proportion of world trade in these products involves bilateral exchange with the United States. This phenomenon stems from commodities that are denominated in other currencies. The prices observed in local markets show significant divergence from nominal market prices during periods of exchange rate volatility, and this results in a spread between nominal market price indices and trade-weighted prices. This spread is not uniform among all commodity markets; hence, possible price asymmetry effects may cause prices to diverge. Convergence between trade-weighted prices and nominal market prices is more significant where there is a greater proportion of international trade in a given commodity relative to total global production.

In past studies, the issue of whether exchange rates impact directly on the behavior rather than just the levels of commodity prices has not been effectively decided. On one hand, changes in exchange rates such as the U.S. dollar are likely to affect prices in different geographic markets much to the same extent (Chambers and Just 1992). On the other hand, researchers have argued the reverse, that is, that changes in commodity prices can be strong enough to affect not only exchange rates but also domestic prices and capital flows. And where exchange rates might affect commodity prices, Adams and Vial (1988) suggested that it is a long-run effect more important on the supply side than on the demand side, thus affecting only future prices through structural adjustments such as in production capacity. In the present case, considerable effort was made to untangle the effects of exchange rates from the price variables, so that our results would reflect only price variations. We examined both exchange rate-converted and unconverted price series to identify any spurious exchange rates effects by performing a preliminary computation of the convergence indexes, with and without relative price exchange rate adjustments. No distortion of the convergence test results was found.

The price data that we could possibly employ have been limited by the availability of relatively few long-run series whose composition has been fairly regular over time. As shown in Figures 1 through 6, we begin with metal prices (copper, lead, and tin) covering the period 1930 through 1998 for Germany, the United Kingdom, and the United States. Also employed are agricultural prices (cotton, wheat, and coffee) covering the period 1950 through 1998 for Argentina, Australia, Brazil, Canada, Colombia, Egypt, Sudan, Uganda, the United Kingdom, and the United States. The quotations for these prices appear in Table 1 and represent trades on important major exchanges, that is, New York Mercantile Exchange; LME; New York Mercantile Exchange; the New York Coffee, Sugar & Cocoa Exchange; and so on.

DEFINING CONVERGENCE AND INTEGRATED TIME SERIES

Past attempts to measure commodity price convergence have been made mainly by regional economists (Persson 1994; Baffes and Ajwad 1998; Jeong 1995; Lefebvre and Poloz 1996). Such studies have been employed primarily to demonstrate various aspects of regional integration. The methodologies developed have revealed that it is usually relevant to distinguish between stationary and nonstationary series. If a series is nonstationary, it has the potential for very large variation over time, so that for convergence to exist between two nonstationary series, cointegration must be a necessary but not a sufficient condition. If the series under consideration are I(1), it may be reasonable to define convergence in terms of the differences between them being of a lower order of integration than other series under consideration (Camarero, Esteve, and Tamarit 2000).

For instance, Bernard and Durlauf (1995) defined long-run convergence between countries i and j if the long-term forecasts of the considered variable for both countries are equal at a fixed time t:

TABLE 1. Characteristics of the Price Series

Commodity	*Country*	*Quotations and Exchanges*	*Data Sources*
Tin, 1947-98	UK	LME, high grade, cash (Bp/ton)	Schmitz
	USA	NY, Straits price to 1975, then Metals Week composite ($/pound)	USGS
	Germany	Hamburg, Min 99.9 percent cash (DM per 100kgs)	Metallst
Copper, 1930-98	UK	LME, wire bars, cash (Bp/ton)	Schmitz
	USA	NYME, producer, wire bars to 1977, then cathode (U.S. cents/pound)	USGS
	Germany	Electrolytic, wire bars (DM per 100kgs)	Metallst
Lead, 1930-98	UK	LME, 99.87 percent pure, settlement (Bp/ton)	Schmitz
	USA	NYME, North America, pure, producers (U.S. cents/ pound)	USGS
	Germany	Soft Pig Lead, min 99.7 percent cash (DM per 100kgs)	Metallst
Coffee, 1950-98	Brazil	NYCSE, Aribacas, Santos No.4 (U.S. cents/pound)	UNCTAD
	Colombia	NYCSE, Mild Aribacas, Colombian Mams (U.S. cents/pound)	IMF
	Uganda	NYCSE, Robusta, Standard (U.S. cents/pound)	
Cotton, 1950-98	USA	NYCE, Memphis, medium staple, middling (U.S. cents/pound)	UNCTAD
	Egypt	Giza 45, extra long staple (U.S. cents/pound)	

Source: Schmitz (1979); U.S. Geological Survey (2000); *Platt's Metals Week* (1996); United Nations Conference on Trade and Development (UNCTAD; 1950-98); World Bank (1950-98); International Monetary Fund (1950-98).
Note: LME = London Metal Exchange; NYME = New York Metal Exchange; NYCE = New York Cotton Exchange; NYCX = New York Coffee, Cocoa & Sugar Exchange.

$$\lim_{k\to\infty} E(p_{i,1+k} - p_{j,t+k}\xi_t) = 0, \tag{1}$$

where ξ_t stands for the information available at time t. This definition is satisfied if $p_{i,1+k} - p_{j,t+k}$ is a mean zero stationary process. This implies that variables for countries i and j to converge, the two series must be cointegrated with cointegrating vector [1, –1]. In addition, if the variables are trend-stationary, then the definitions imply that the trends for each country must be the same. All these conditions have been applied extensively to study the existence of nominal convergence with the main problem being that convergence is a gradual and ongoing process (Camarero, Esteve, and Tamarit 2000). In the present international context, we employ three tests of convergence: (1) correlation analysis, (2) stationarity and cointegration analysis, and (3) vector autoregression (VAR) analysis.

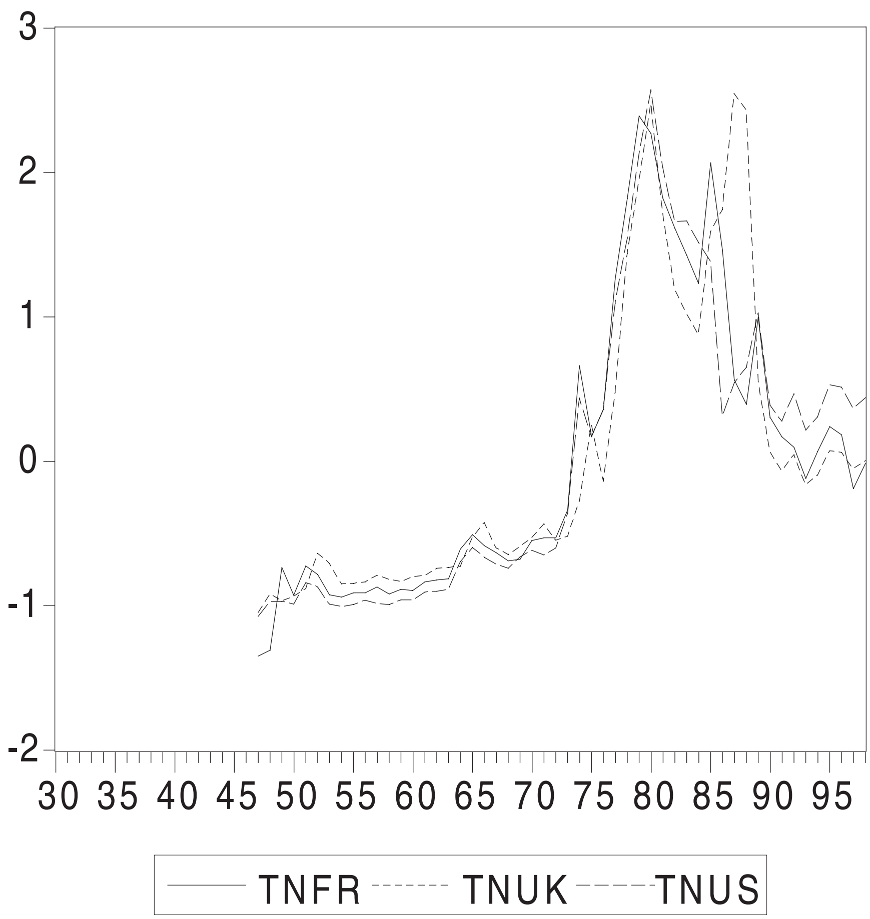

FIGURE 1. Tin Prices

Note: TNFR = tin France; TNUK = tin United Kingdom; TNUS = tin United States.

CORRELATION ANALYSIS

The computation of simple correlation coefficients within different subperiods of a total sample period can be employed to test the concept of converging correlation over time between variables separated by space (e.g., Lele 1967; Southworth, Jones, and Pearson 1979; Timmer, Falcon, and Pearson 1983; Stigler and Sherwin 1985). However, since correlation analysis is static rather than dynamic, it is also important to examine cross-correlations with a lag structure between the

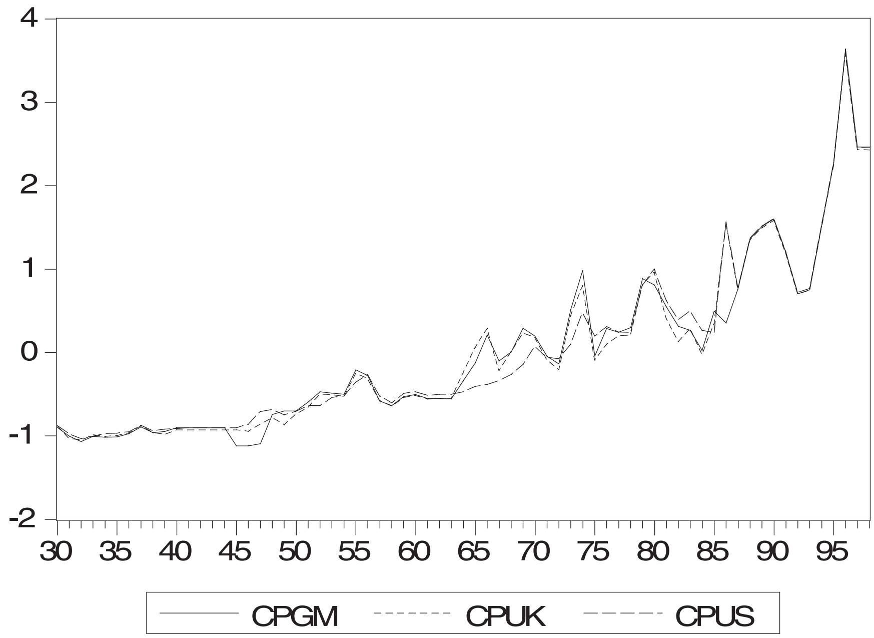

FIGURE 2. Copper Prices

Note: CPUK = copper United Kingdom; CPGM = copper Germany; CPUS = copper United States.

variables of interest. To accomplish this, simple correlation coefficients (r_i^2) within different subperiods of the total sample are calculated. The estimated r_i^2 coefficients are then used to estimate the convergence indexes, C_{ij} and C_{iT} as follows:

$$C_{ij} = \frac{r_{12}^2 + r_{13}^2 + r_{23}^2}{n_{c_2}} \tag{2}$$

$$C_{iT} = \frac{\dfrac{r_{12}^2 + r_{13}^2 + r_{23}^2}{n_{c_2}}}{\dfrac{r_{12}^2 + r_{13}^2 + r_{14}^2 + \cdots}{n_{c_2}}} = \frac{C_{ij}}{C_{11}}, \tag{3}$$

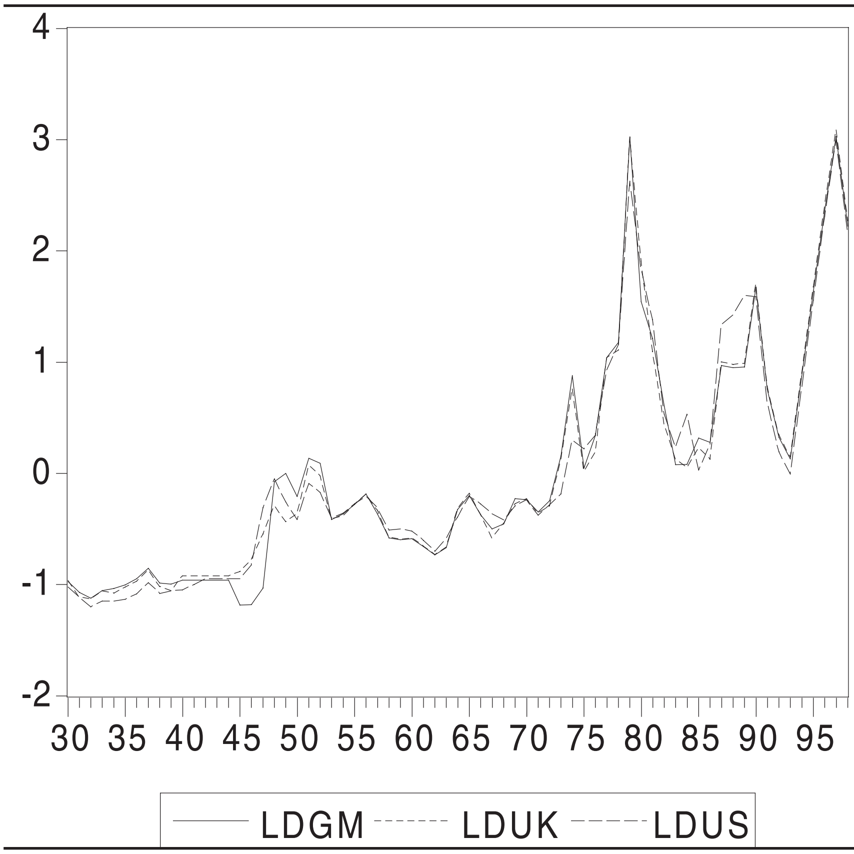

FIGURE 3. Lead Prices

Note: LDGM = lead Germany; LDUK = lead United Kingdom; LDUS = lead United States.

where $i = 1, 2, 3$; $j = 1, 2, 3, \ldots, 6$; $T = 1, 2, \ldots, 14$; and C_{11} is the C_{ij} for the first subperiod. Here, i represent a market, j represents a commodity, and T represents subperiods in each case. In the above equations, a coefficient of C equal to one would be interpreted as a perfect transmission of price shock, while a coefficient of zero would represent a short-run invariance to changes in price elsewhere. Since the short-run effect is in principle unrestricted, a value of C_{iT} greater than unity, for example, would suggest an overreaction to changes in price in the current period.

Stationarity and Cointegration Analysis

One of the conditions of the stochastic definition of convergence is stationarity. Price disparities between economies or markets should follow a stationary process.

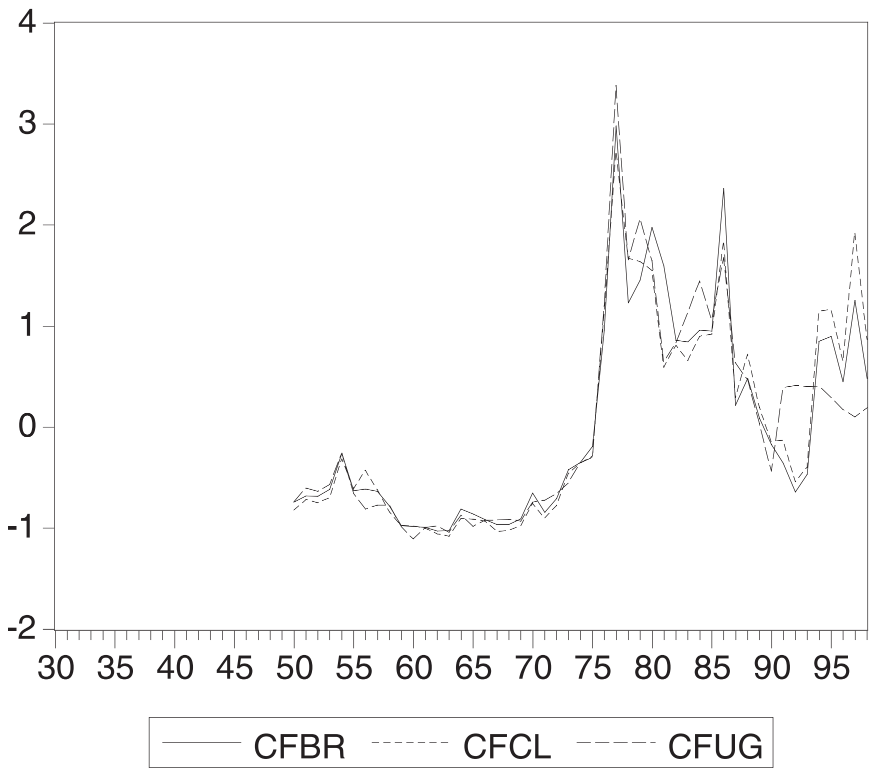

FIGURE 4. Coffee Prices

Note: CFBR = coffee Brazil; CFCL = coffee Colombia; CFUG = coffee Uganda.

Without stationarity, relative price shocks could lead to permanent deviations in any tendency toward convergence (Carlino and Mills 1993; Bernard and Durlauf 1995). Thus, before conducting cointegration analysis, one must confirm that all the price series are nonstationary and integrated of the same order. This is done with Dickey-Fuller and Augmented Dickey-Fuller tests employing the regression

$$\Delta y_t = a + gy_{t-1} + \sum_{i=1}^{k} b_i \Delta y_{t-i} + e_t. \tag{4}$$

The lag length k is chosen to generate a white noise error term e_t. To determine whether y_t is nonstationary, the null hypothesis of nonstationarity is evaluated by testing whether $g = 0$ against the alternative of stationarity $g < 0$.

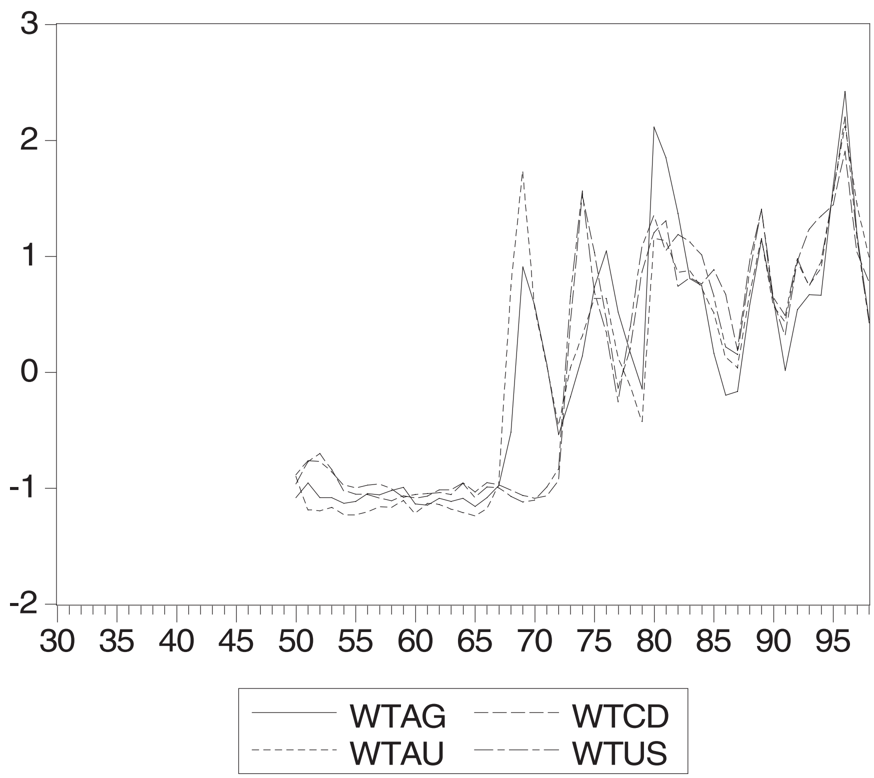

FIGURE 5. Wheat Prices

Note: WTAG = wheat Argentina; WTAU = wheat Australia; WTCD = wheat Canada; WTUS = wheat United States.

Turning to cointegration analysis, the correlation between two time series variables can be better evaluated by employing linear or nonlinear regression analysis (see Isard 1977; Mundlak and Larson 1992; Gardner and Brooks 1994). Previous applications of cointegration analysis to spatial commodity prices include Alexander and Wyeth (1994); Ardeni (1989); Asche, Bremnes, and Wessells (1999); Baffes (1991); Goodwin (1992); Goodwin, Grennes, and Wohlgenant (1990); and Zanias (1993). The cointegrating equation employed here is

$$P_t^1 = \mu + \beta_1 P_t^2 + \varepsilon_t, \tag{5}$$

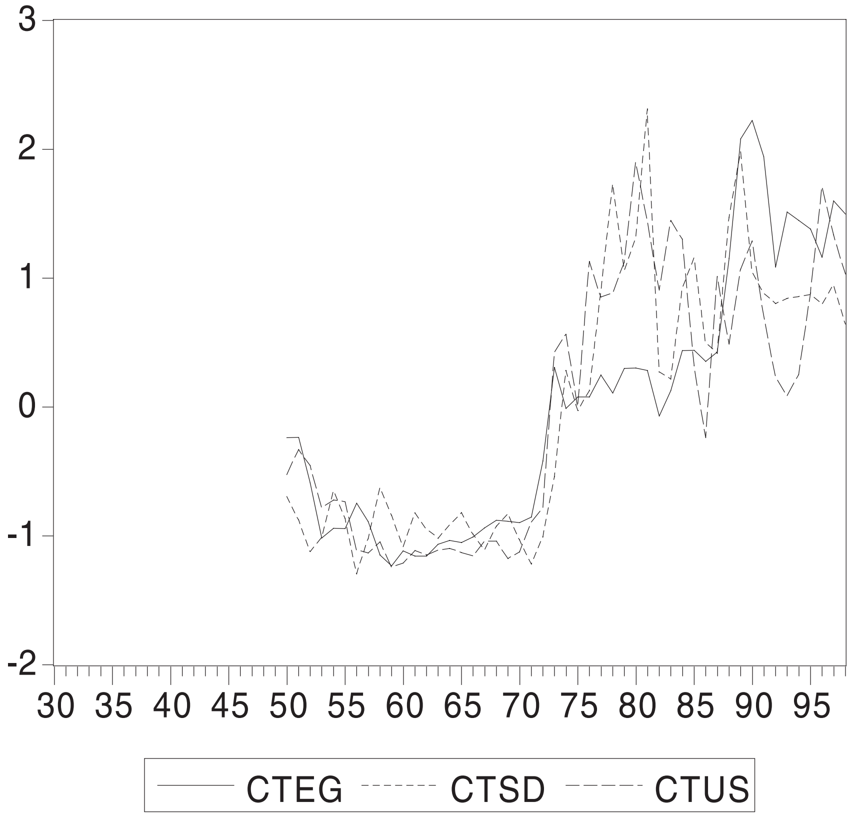

FIGURE 6. Cotton Prices

Note: CTEG = cotton Egypt; CTSD = cotton Sudan; CTUS = cotton United States.

where P_t^1 and P_t^2 denote prices from two different geographical locations of the commodity under consideration; μ and β_1 are parameters to be estimated, while ε denotes a *iid*(0, σ^2) term. Using such a model, the hypothesis that the slope coefficient equals unity and (possibly) the intercept term equals zero is tested; formally, H_0: $\mu + 1 = \beta_1 = 1$. Under H_0, the deterministic part of equation (5) becomes $P_t^1 = P_t^2$, in turn implying that the price differential, $P_t^1 - P_t^2$, is a *iid*(0, σ^2) term. With respect to the nonstationarity problem, one can examine the order of integration of the error term in equation (5) and make inferences regarding the validity of the model. If prices are indeed nonstationary, the existence of a stationary error term implies cointegration between the two prices.

VAR Analysis

Another useful means of summarizing the broad correlation in the variables of a system is VAR. Because most of the convergence results previously obtained from cointegration analysis have not been conclusive, Miljkovic (1999) considered that approach to be only a pretest for convergence and suggested more elaborate methods such as VAR. VAR is commonly used for forecasting systems of interrelated time series and for analyzing the dynamic impact of random disturbances on the system of variables (Cromwell et al. 1994). The VAR approach sidesteps the need for structural modeling by modeling every endogenous variable in the system as a function of the lagged values of all of the endogenous variables in the system. The estimated VARs are used to calculate the percentages of each endogenous variable that can be explained by innovations in each of the explanatory variables and provides information about the relative importance of each random innovation to the variable in the VAR. The mathematical form of a VAR is

$$Y_t = A_1 Y_{t-1} + \ldots + A_p Y_{t-p} + BX_t + \varepsilon_t, \tag{6}$$

where Y_t is a k vector of endogenous variables, X_t is a d vector of exogenous variables, $A_1, \ldots, A_p$ and β are matrices of coefficients to be estimated, and ε_t is a vector of innovations that may vary contemporaneously.

In this article, the VAR model is used to highlight the impact of changes in prices among the dispersed markets in two ways: decomposition of the variance in forecast errors and the analysis of impulse shocks. Variance decomposition involves decomposing the variance of the forecasts error into components that can be attributed to each of the endogenous variables. Impulse shocks involve tracing the response of each variable to a shock, or innovation, in one variable in the system.

Empirical Results

Correlations

The estimated correlation convergence indexes (C_{ij} and C_{iT}) are reported in Table 2, and the graphical representations of the indexes are depicted in Figures 7 through 12. As noted earlier, a coefficient of one in Table 2 represents 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, C_{iT} greater than unity suggests an overreaction to changes in prices in the current period. The results in Table 2 show a higher occurrence of an overreaction to changes in prices in the current period for agricultural commodities than for the prices of metal commodities. This observation is somewhat embedded by the nature of the agricultural commodities. In general, the estimated correlation indexes[1] do not confirm the convergence hypothesis.

TABLE 2. Correlation Market Convergence Index

Years		*Lead*	*Copper*	*Tin*	*Coffee*	*Cotton*	*Wheat*
1930-35	C_{ij}	.87	.95				
	C_{iT}	1.00	1.00				
1936-40	C_{ij}	.82	.87				
	C_{iT}	.95	.92				
1941-45	C_{ij}	–.33	.50				
	C_{iT}	–.38	.52				
1946-50	C_{ij}	.82	.72	.59			
	C_{iT}	.95	.76	1.00			
1951-55	C_{ij}	.98	.93	.38	.96	.29	.43
	C_{iT}	1.14	.98	.65	1.00	1.00	1.00
1956-60	C_{ij}	1.00	1.00	.07	.84	–.05	.09
	C_{iT}	1.15	1.05	.12	.88	–.18	.20
1961-65	C_{ij}	.99	.99	.89	.82	.12	.33
	C_{iT}	1.14	1.04	1.52	.86	.40	.77
1966-70	C_{ij}	.83	.56	.70	.95	.05	–.31
	C_{iT}	.96	.59	1.20	.99	.17	–.72
1971-75	C_{ij}	.87	.88	.73	.98	.82	.73
	C_{iT}	1.01	.93	1.25	1.02	2.83	1.69
1976-80	C_{ij}	.98	.98	.97	.93	.28	.59
	C_{iT}	1.14	1.03	1.65	.97	.96	1.36
1981-85	C_{ij}	.90	.63	.44	–.20	.15	.58
	C_{iT}	1.04	.66	.75	–.21	.50	1.35
1986-90	C_{ij}	.90	.54	–.05	.95	.60	.94
	C_{iT}	1.04	.57	–.09	.99	2.05	2.18
1991-95	C_{ij}	1.00	1.00	.83	–.07	.62	.88
	C_{iT}	1.15	1.06	1.41	–.07	2.13	2.04
1996-98	C_{ij}	1.00	1.00	1.00	–.31	–.03	.99

Stationarity

The findings of the stationary tests suggest that the null hypothesis of a random walk in the levels series cannot be rejected in two of the metal commodities, copper and tin, and in two of the agricultural commodities, coffee and cotton (Table 3). Critical values at the 5 percent level of significance require *t*-statistics in excess of 3.48 in absolute value for rejection of the null hypothesis (Fuller 1976, 373); here the estimated *t*-statistics for these commodities are below 3.48 in absolute values. However, the null hypothesis of a random walk in the first differences is rejected for all commodities. That is, the Augmented Dickey-Fuller *t*-statistics on the first difference series with a trend for all commodities are all in excess of 5.0 in absolute value. These findings suggest that the first differences of all series are stationary.

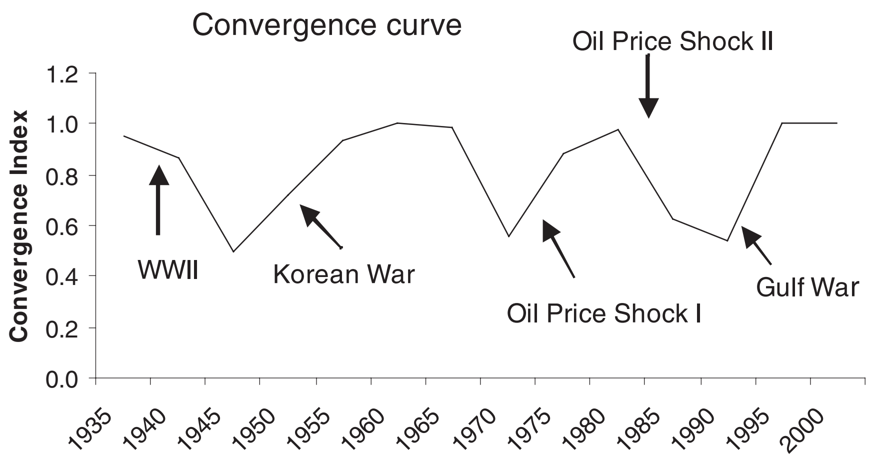

FIGURE 7. Estimated Correlation Convergence Index (C_{iT}) for Copper

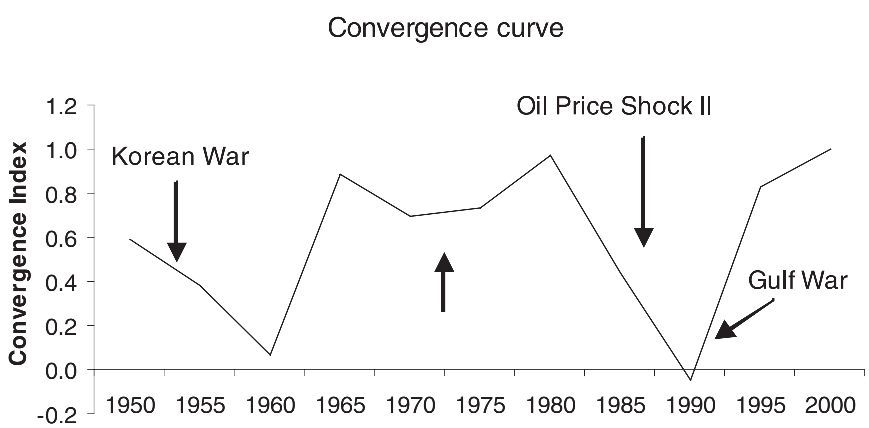

FIGURE 8. Estimated Correlation Convergence Index (C_{iT}) for Tin

Turning to the no time trend specification, the results for the level series are consistent with the earlier findings for the time trend specification (Table 3). Under the no time trend specification, an approximate 5 percent critical value of –2.89 is used, and the null hypothesis of a random walk in the levels series is not rejected since the test statistics are not greater than the critical values for all commodities. On the contrary, however, the null hypothesis of a random walk in the first difference series is rejected for all commodities. Similar to the trend specification, the first differences of each series under the no time trend specification are stationary for all commodi-

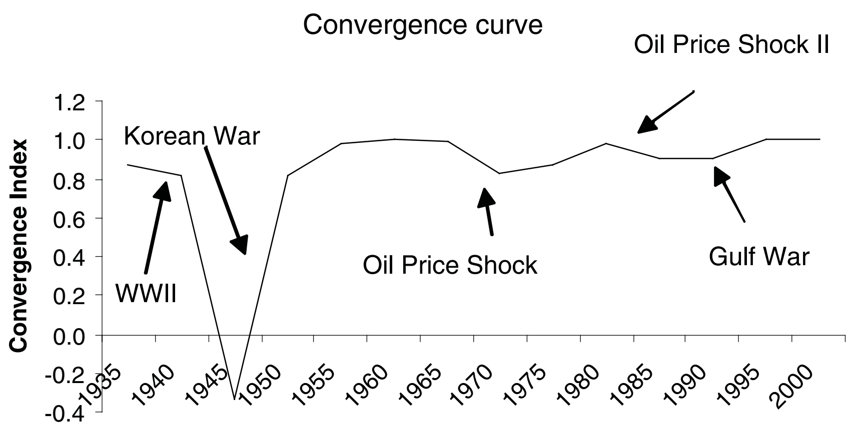

FIGURE 9. Estimated Correlation Convergence Index (C_{iT}) for Lead

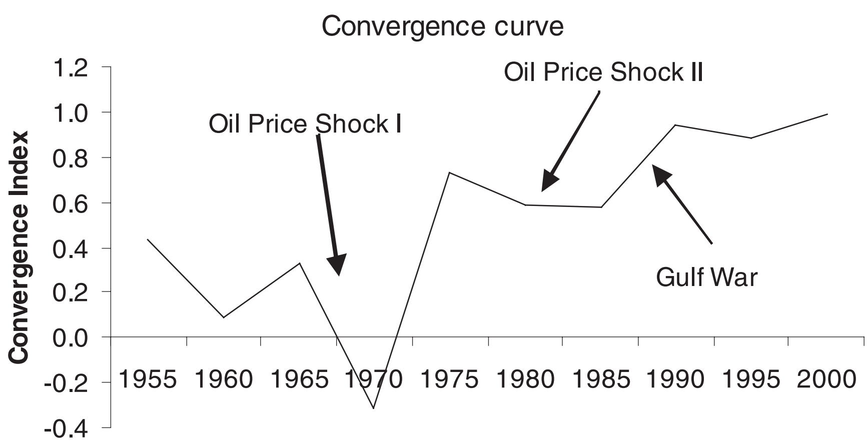

FIGURE 10. Estimated Correlation Convergence Index (C_{iT}) for Wheat

ties. Stationarity in the first difference series would thus suggest a tendency toward convergence.

COINTEGRATION

The Johansen Cointegration test statistics are reported in Tables 4 and 5. First, the test statistics with a time trend in the regression are reported and then the test with no time trend. Based on the critical values for the trace statistics reported by

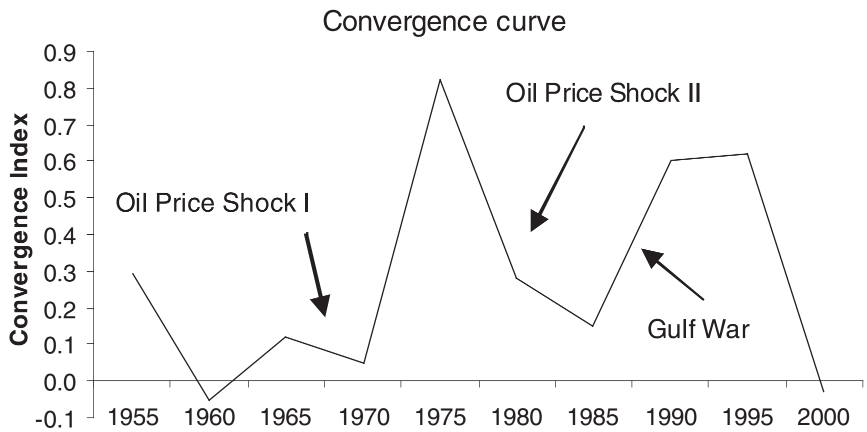

FIGURE 11. Estimated Correlation Convergence Index (C_{iT}) for Cotton

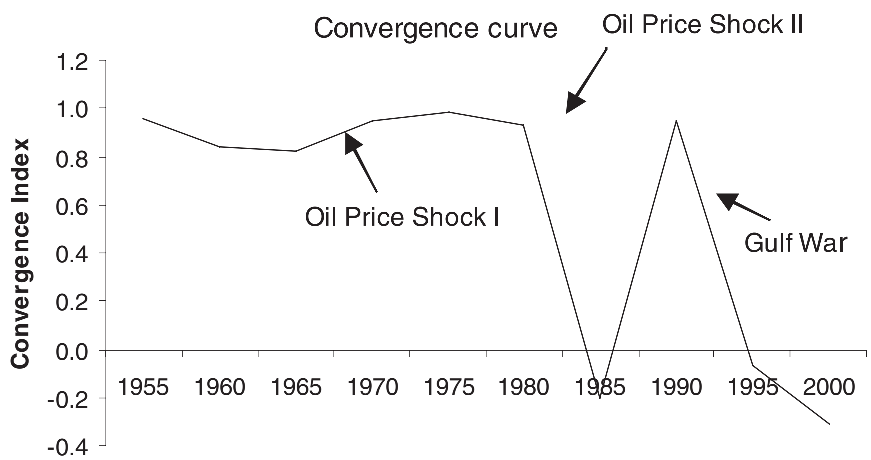

FIGURE 12. Estimated Correlation Convergence Index (C_{iT}) for Coffee

Osterwald-Lenum (1992), the results suggest that the null hypothesis of no cointegration cannot be rejected for copper, cotton, and coffee (with and without time trends) at the 5 percent level of significance. Since the cointegrating vector is not identified unless some arbitrary normalization is imposed, the first r series[2] in the vector is normalized to an identity matrix. The normalized cointegration relation assuming one cointegration relation $r = 1$ and the log likelihood values are reported

TABLE 3. Augmented Dickey-Fuller (ADF) Tests (with and without Trend) on Levels and First Differences of Metal and Agricultural Commodity Prices

	ADF (Trend)		*ADF (No Trend)*	
Series	*Levels*	*First Differences*	*Levels*	*First Differences*
CPUK	–3.40 (1)	–9.06** (0)	–0.61 (1)	–9.04** (0)
CPGM	–3.30 (1)	–8.37** (0)	–0.57 (1)	–8.35** (0)
CPUS	–2.74 (1)	–9.60** (0)	–0.20 (1)	–9.53** (0)
TNFR	–1.42 (1)	–6.77** (0)	–1.68 (1)	–6.73** (0)
TNUK	–2.42 (1)	–5.22** (0)	–2.08 (1)	–5.25** (0)
TNUS	–1.59 (1)	–6.04** (0)	–1.37 (1)	–6.08** (0)
LDGM	–3.99** (1)	–8.38** (0)	–1.37 (1)	–8.42** (0)
LDUK	–3.74**(1)	–7.72** (0)	–1.55 (1)	–7.76** (0)
LDUS	–4.37** (1)	–6.36** (0)	–1.87 (1)	–6.42** (0)
WTAG	–4.39** (1)	–5.96** (0)	–2.42 (1)	–6.03** (0)
WTAU	–4.81** (1)	–5.73** (0)	–2.24 (1)	–5.81** (0)
WTCD	–4.18** (1)	–5.36** (0)	–1.84 (1)	–5.42** (0)
WTUS	–3.99** (1)	–5.17** (0)	–2.10 (1)	–5.23** (0)
CFBR	–2.22 (1)	–8.75** (0)	–1.78 (1)	–8.85** (0)
CFCL	–2.31 (1)	–8.07** (0)	–1.61 (1)	–8.17** (0)
CFUG	–2.07 (1)	–7.53** (0)	–1.80 (1)	–7.60** (0)
CTEG	–3.66 (1)	–5.81** (0)	–0.76 (1)	–5.76** (0)
CTSD	–2.98 (1)	–7.47** (0)	–1.65 (1)	–7.56** (0)
CTUS	–2.65 (1)	–7.50** (0)	–1.27 (1)	–7.56** (0)

Note: CPUK= copper United Kingdom; CPGM = copper Germany; CPUS = copper United States; TNFR = tin France; TNUK = tin United Kingdom; TNUS = tin United States; LDGM = lead Germany; LDUK = lead United Kingdom; LDUS = lead United States; WTAG = wheat Argentina; WTAU = wheat Australia; WTCD = wheat Canada; WTUS = wheat United States; CFBR = coffee Brazil; CFCL = coffee Colombia; CFUG = coffee Uganda; CTEG = cotton Egypt; CTSD = cotton Sudan; CTUS = cotton United States. The numbers in parentheses indicate the number of lags on the dependent variable as determined by a final prediction error (FPE) search in the augmented Dickey-Fuller test. Critical values = –3.48 for ADF (Trend) and –2.89 (ADF No Trend).
**Denotes significance at the 1 percent level.

in Table 5. The results of the cointegration test show only one cointegration equation for lead and wheat prices. This thus implies a tendency toward convergence only in the lead and wheat price series.

VAR Measures

The estimated VARs are used to calculate the percentage of the total variation in each endogenous variable that can be explained by innovations in each of the variables.

This measure, accordingly, can provide a clear picture of the economic importance of a given market for the behavior of the price variable in question. Such an exercise requires that a causal ordering of the system variables be chosen; however,

TABLE 4. Johansen Cointegration Test

Series	*Eigenvalue*	*Likelihood Ratio*	*5 Percent Critical Value*	*Included Observations*
Intercept (no trend)				
Copper	0.27	29.48	29.68	66
Tin	0.47*	37.59	29.68	49
Lead	0.25*	37.26	29.68	66
Cotton	0.19	13.23	29.68	46
Coffee	0.40	28.65	29.68	46
Wheat	0.65*	76.38	47.21	46
Intercept (with trend)				
Copper	0.27	41.71	42.44	66
Tin	0.54*	51.29	42.44	49
Lead	0.30*	56.54	42.44	66
Cotton	0.26	22.85	42.44	46
Coffee	0.45	37.85	42.44	46
Wheat	0.67*	92.91	62.99	46

Note: Lag interval is 1 to 2.
*Denotes rejection of the hypothesis at the 5 percent significance level.

some experimentation revealed that the ordering assumption was not critical to the results that emerged. The estimated variance decompositions are reported in Tables 6 and 7.

A common finding in these tables is that most of the variations in a given variable are explained by lags of the variable itself, and this is a feature of the regional VARs. Looking at the agricultural price series, we note a number of instances where the proportions of variances explained are not trivial. For instance, we find that the wheat price movements in the U.S. market explain a fairly large proportion of the world wheat price movements, between 18 and 76 percent, depending on the market. Similarly, the cotton and coffee price movements in Sudan and Colombia, respectively, explain a fairly large proportion of the world price movements for these commodities. Looking at the metal commodities, the results show that copper, tin, and lead price movements in the U.S. and U.K. markets explain a fairly large proportion of the world market prices for these commodities, respectively. Overall, our results show a varying percentage of the variance due to each innovation within commodities. This suggests that shocks to economic conditions do contribute to the forecast variances of the price series.

So far, the signs of the various linkages in the VAR have not been evaluated. One way of accomplishing this is through impulse response functions that simulate the impacts of a shock to commodity prices (leaving all variables endogenous) and that compute the predicted dynamic responses of each commodity price (market). By treating the residuals of each equation as unexplained innovations, the impacts of innovations are traced through the system by shocking the error terms. Tables 8 and

TABLE 5. Normalized Cointegrating Coefficients, One Cointegrating Equation(s)

Series	*Intercept (No Trend)*	*Intercept (With Trend)*	*Log-Likelihood*
Copper			
Constant	–260.89	–93.41	–1,323.62
Trend	—	–26.23 (51.52)	–1,323.51
CPUS	1.00	1.00	
CPUK	–9.03 (10.40)	–10.54 (15.36)	
CPGM	8.19 (10.61)	10.31 (16.53)	
Tin			
Constant	–112.97	–812.23	–1,229.90
Trend	—	25.93 (7.28)	–1,226.41
TNFR	1.00	1.00	
TNUK	–0.63 (0.10)	–0.51 (0.054)	
TNUS	–0.31 (0.077)	–0.46 (0.042)	
Lead			
Constant	–9.53	–58.33	–1,079.64
Trend	—	–12.98** (14.18)	–1,077.27
LDGM	1.00	1.00	
LDUK	–1.22** (0.16)	–2.29** (1.58)	
LDUS	0.21** (0.16)	2.39** (2.70)	
Cotton			
Constant	–57.89	118.96	–539.07
Trend	—	–5.97 (1.88)	–536.78
CTEG	1.00	1.00	
CTSD	–19.25 (22.32)	–2.64 (1.60)	
CDUS	21.90 (28.84)	3.64 (2.13)	
Coffee			
Constant	7.44	–0.64 (0.14)	–587.12
Trend	—	0.34	–585.21
CFBR	1.00	1.00	
CFCL	–0.45 (0.16)	–0.62 (0.12)	
CFUG	–0.77 (0.18)	–0.66 (0.14)	
Wheat			
Constant	40.02	56.72	–727.17
Trend	—	–0.74* (0.49)	–726.31
WTAG	1.00	1.00	
WTAU	–0.68 (0.059)	–0.58* (0.089)	
WTCD	2.96 (0.47)	3.64* (0.73)	
WTUS	–3.99 (0.58)	–4.71* (0.85)	

Note: Critical values = 25.32 and 30.25 for lead and 42.44 and 48.45 for wheat. Standard errors, in parentheses, are reported only for variables that are identified under one cointegration equation(s). CPUK= copper United Kingdom; CPGM = copper Germany; CPUS = copper United States; TNFR = tin France; TNUK = tin United Kingdom; TNUS = tin United States; LDGM = lead Germany; LDUK = lead United Kingdom; LDUS = lead United States; WTAG = wheat Argentina; WTAU = wheat Australia; WTCD = wheat Canada; WTUS = wheat United States; CTEG = cotton Egypt; CTUS = cotton United States; CTSD = cotton Sudan; CFBR = coffee Brazil; CFCL = coffee Colombia; CFUG = coffee Uganda.
*Denotes rejection of the hypothesis at the 5 percent significance level. **Denotes rejection of the hypothesis at the 1 percent significance level.

TABLE 6. Variance Decomposition for Agricultural Commodities (Wheat, Cotton, Coffee)

	Standard Error	*WTAG (%)*	*WTAU (%)*	*WTCD (%)*	*WTUS (%)*
Variance decomposition WTAG					
Period 1	21.3	100	0	0	0
Period 4	38.9	55.5	14.4	29.7	0.4
Variance decomposition WTAU					
Period 1	22.1	63.8	36.2	0	0
Period 4	41.1	32.3	47.4	19.8	0.5
Variance decomposition WTCD					
Period 1	19.8	17.3	1.0	81.7	0
Period 4	41.9	10.4	2.3	85.2	2.1
Variance decomposition WTUS					
Period 1	17	23.9	0.2	65.2	10.7
Period 4	34.5	17.8	0.7	75.9	5.6

	Standard Error	*CTEG (%)*	*CTUS (%)*	*CTSD (%)*
Variance decomposition CTEG				
Period 1	21.9	100	0	0
Period 4	48.2	94.5	2.0	3.6
Variance decomposition CTUS				
Period 1	8.6	7.4	92.6	0
Period 4	14.5	24.5	64.6	10.9
Variance decomposition CTSD				
Period 1	11.7	12.6	4.9	82.5
Period 4	20.0	31.3	31.7	37

	Standard Error	*CFBR (%)*	*CFCL (%)*	*CFUG (%)*
Variance decomposition CFBR				
Period 1	32	100	0	0
Period 4	48	78	5.7	15.9
Variance decomposition CFCL				
Period 1	28	84	16	0
Period 4	44	71	19	10
Variance decomposition CFUG				
Period 1	22	63	0.4	37
Period 4	37	59	4	36

Note: WTAG = wheat Argentina; WTAU = wheat Australia; WTCD = wheat Canada; WTUS = wheat United States; CTEG = cotton Egypt; CTUS = cotton United States; CTSD = cotton Sudan; CFBR = coffee Brazil; CFCL = coffee Colombia; CFUG = coffee Uganda.

TABLE 7. Variance Decomposition for Metal Commodities (Copper, Tin, Lead)

	Standard Error	*CPUK (%)*	*CPGM (%)*	*CPUS (%)*
Variance decomposition CPUK				
Period 1	373	10	6	87
Period 5	706	4	7	89
Variance decomposition CPUS				
Period 1	335	0	0	100
Period 5	678	0.2	3.3	97
Variance decomposition CPGM				
Period 1	350	0	34.5	0
Period 5	691	1.9	13.4	1.8
	Standard Error	*TNFR (%)*	*TNUK (%)*	*TNUS (%)*
Variance decomposition TNFR				
Period 1	1,448	100	0	0
Period 4	2,820	82	1.2	16.5
Variance decomposition TNUK				
Period 1	1,137	14.8	85.2	0
Period 4	2,971	81.9	16	2
Variance decomposition TNUS				
Period 1	1,589	57.9	0.6	41.5
Period 4	3,348	62.7	0.4	36.8
	Standard Error	*LDGM (%)*	*LDUK (%)*	*LDUS (%)*
Variance decomposition LDGM				
Period 1	127	100	0	0
Period 5	263	89	0.35	10.6
Variance decomposition LDUK				
Period 1	122	93.8	6.2	0
Period 5	260	87.6	2.3	10
Variance decomposition LDUS				
Period 1	107	80.7	4.3	15
Period 5	252	80.8	2.7	16.5

Note: CPUK = copper United Kingdom; CPGM = copper Germany; CPUS = copper United States; TNFR = tin France; TNUK = tin United Kingdom; TNUS = tin United States; LDGM = lead Germany; LDUK = lead United Kingdom; LDUS = lead United States.

9 present the estimated impulse responses for the metal and agricultural prices, respectively. Summarizing these results, it is observed, for instance, that after five lags, the impulse responses do not approach zero, implying that the underlying data are nonstationary. Thus, the results in this study concur with the conclusions by

TABLE 8. Impulse Response for Metal Commodities (Levels Series): Impulse Response to One Standard Deviation Innovations (Copper, Tin, Lead)

	CPUK	*CPGM*	*CPUS*
Impulse response of CPUK			
Period 1	164	339	0
Period 5	115	258	103
Impulse response of CPGM			
Period 1	0	354	0
Period 5	117	261	93
Impulse response of CPUS			
Period 1	159	271	121
Period 5	118	265	107
	TNFR	*TNUK*	*TNUS*
Impulse response of TNFR			
Period 1	1,448	0	0
Period 4	1,135	–103	692
Impulse response of TNUK			
Period 1	437	1,050	0
Period 4	1,628	124	111
Impulse response of TNUS			
Period 1	1,208	123	1,024
Period 4	1,495	–65	718
	LDGM	*LDUK*	*LDUS*
Impulse response of LDGM			
Period 1	127	0	0
Period 5	93	–1.41	30
Impulse response of LDUK			
Period 1	118	30	0
Period 5	88	–4.48	29
Impulse response of LDUS			
Period 1	96	22	42
Period 5	89	–4.48	29

Note: CPUK= copper United Kingdom; CPGM = copper Germany; CPUS = copper United States; TNFR = tin France; TNUK = tin United Kingdom; TNUS = tin United States; LDGM = lead Germany; LDUK = lead United Kingdom; LDUS = lead United States.

Carlino and Mills (1993) and by Bernard and Durlauf (1995) that without stationarity, relative price shocks could lead to permanent deviations in any tendency toward convergence.

TABLE 9. Impulse Response for Agricultural Commodities (Levels Series): Impulse Response to One Standard Deviation Innovations (Wheat, Cotton, Coffee)

	WTAG	*WTAU*	*WTCD*	*WTUS*
Impulse response of WTAG				
Period 1	21	0	0	0
Period 4	3	6	13	–1.9
Impulse response of WTAU				
Period 1	18	13	0	0
Period 4	4	8	12	–1.8
Impulse response of WTCD				
Period 1	8	2	18	0
Period 4	7	3	15	–1.4
Impulse response of WTUS				
Period 1	8	0.67	14	6
Period 4	7	1.7	11	–2.9
	CTEG	*CTUS*	*CTSD*	
Impulse response of CTEG				
Period 1	22	0	0	
Period 4	23	4	7	
Impulse response of CTUS				
Period 1	2	8	0	
Period 4	4	4	3.8	
Impulse response of CTSD				
Period 1	4	3	11	
Period 4	6	5	3	
	CFBR	*CFCL*	*CFUG*	
Impulse response of CFBR				
Period 1	32	0	0	
Period 4	15	3.95	12	
Impulse response of CFCL				
Period 1	26	11	0	
Period 4	14	4.9	8	
Impulse response of CFUG				
Period 1	18	1.4	14	
Period 4	12	–0.57	8	

Note: WTAG = wheat Argentina; WTAU = wheat Australia; WTCD = wheat Canada; WTUS = wheat United States; CTEG = cotton Egypt; CTUS = cotton United States; CTSD = cotton Sudan; CFBR = coffee Brazil; CFCL = coffee Colombia; CFUG = coffee Uganda.

CONCLUSIONS

This study has examined the degree to which commodity prices have converged on world commodity markets over recent decades. Comparable geographic data were obtained for six commodities: coffee, cotton, wheat, copper, lead, and tin. Our results indicate that correlation convergence indexes themselves are not capable of detecting the convergence that might have taken place. Cointegration analysis was employed, beginning with unit root tests for the levels and first difference series. In the levels series under a time trend specification, lead and wheat prices were found to be stationary, while the first differences were found to be stationary for all price series. The cointegration test results showed one cointegrating equation for lead and wheat prices, suggesting a tendency toward convergence in the lead and wheat markets, while the VAR impulse response results confirmed nonstationarity in all cases.

Altogether the presence of nonstationarity, the lack of common trends (no cointegration), and the evidence for increasing variance suggest the lack of convergence. Our empirical results, therefore, do not support the convergence hypothesis but rather a pattern of fluctuating or random divergences in all markets with the exception of lead and wheat. Unfortunately it has not been possible to determine the causes of these divergences. Among possible confounding factors, the following would be worth investigating: commodity market conditions (such as political unrest, wars, and climate), international business cycle conditions (such as inflations/recessions, monetary crises, exchange rate risks), exogenous shocks (such as the oil price shocks of 1974, 1978, 1981, and 1991), and the diversity of transport costs and other factors reflecting geographical separation.

NOTES

1. The convergence properties of these indexes were tested using prices generated by an autoregressive process. The results indicated no tendency toward convergence or other aberration of interpretation.

2. Because of space limitation, only the first r estimates—where r is determined by the likelihood ratio (LR) tests—for the estimated $k - 1$ cointegrating relations are reported.

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