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The Law of One Price: Developed and Developing Country Market Integration

Jian Yang, David A. Bessler, and David J. Leatham

ABSTRACT

The Law of One Price (LOP) is important to models of international trade and exchange rate determination. This study investigates a variant of the LOP applied to developed and developing countries. The competing hypotheses are (1) that one price prevails in both developed and developing countries and (2) that one price prevails in developed countries and another single price in developing countries. Using data from an internationally competitive commodity (soybean meal), we found evidence favors the first hypothesis, although two large developing countries under study are active participants in regional trade integration, which may bias them against the first hypothesis.

Key Words: *law of one price, developing countries, error-correction model, directed graphs.*

The law of one price (LOP) states that for a given commodity a representative price adjusted by exchange rates and allowance for transportation costs will prevail across all countries. The LOP plays an important role in models of international trade and exchange rate determination (Protopapadakis and Stoll, 1983, 1986; Michael *et al.*, 1994). The LOP also defines the extent of the market and measures market integration (Stigler and Sherwin, 1985). If a single price exists over several spatially separate markets, it implies that these markets are integrated as a single market. Measurement of market integration can be viewed as basic to understanding how specific markets work (Ravallion, 1986). The extent to which commodity markets are integrated also has important implications for governments' regulation and general economic policy. If a market is internationally integrated, governmental intervention within one nation may be ineffective or very costly.

Recognizing the nonstationarity property of commodity prices, researchers have extensively employed cointegration and error-correction models (ECM) (Engle and Granger, 1987) to test the LOP and market integration on international commodity markets. This is particularly useful because the LOP and market integration are tested as a long-run relationship that is not affected by short-run deviations. Earlier studies (e.g., Protopapadakis and Stoll, 1986, p.336) already found that the LOP almost never holds in the short run. These works include Ardeni (1989), Baffe (1991), Goodwin (1992), Zanias (1993), Michael *et al.* (1994), Diakosavvas (1995), Mohanty *et al.* (1996), Taylor *et al.* (1997), Mohanty *et al.* (1998), and Mohanty *et al.* (1999). Most of these authors found some evidence for the validity of the LOP and international market integration. However, previous studies only considered de-

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veloped countries and little has been done to examine whether or not the LOP holds across both developed and developing countries.

In this study we were concerned with whether or not commodity markets in "southern" developing countries are well integrated with their counterparts in "northern" developed countries. There exist two competing hypotheses over this issue. One is a natural outcome of the original LOP, i.e., one price should prevail across both developed and developing countries (hereafter the hypothesis of *north-south market integration*). The other is the hypothesis of *north-south market segmentation,* which suggests that the LOP may hold separately in each of these two markets, i.e., one price in developed-country markets versus another single price in developing-country markets. The latter suggests considerable variation in the LOP and is supported by some economists. For example, Cristini (1995) argued that when theoretically modeling commodity price linkage between developed and developing countries, the developed countries in the Organization for Economic Cooperation and Development (OECD) should be viewed as a unified bloc which interacts with the developing countries as a whole in primary commodity markets. Cristini's model assumes that there are at least two separate markets for a primary commodity, composed of developed countries and developing countries. Similarly, Monke and Taylor (1985) presented a model where market participants of the world commodity market are classified into two groups depending on whether or not there are quantitative controls on their international trade. In the context of this paper, developed countries as a whole should have relatively little quantitative controls compared to developing countries. Segmentation in international commodity markets was also considered an essential assumption in Hollifield and Uppal's (1997) model of uncovered interest rate parity. Ghosh (1996) also pointed out that though developing countries are more integrated into the global market than before, the price difference for similar products tends to be much larger between the developed and the developing countries than between developed countries. Thus, the inference from these works supports the hypothesis of north-south market segmentation. To our knowledge, no relevant empirical tests based on cointegration analysis have been conducted to address the controversy.

This study contributed to the literature in two ways. First, it addressed the issue of whether developed and developing markets as two different groups are segmented or integrated, which has not been explored. As explained in the next section, the data set of an international competitive commodity such as soybean meal is ideal for exploring this issue. Second, the study modeled price dynamics combining directed graphs (Sprites, Glymour and Scheines, 1993; Pearl, 1995; Bessler and Akleman, 1998) and error-correction modeling. This was an extension of the recent advance in VAR innovation accounting analysis, as done in Bessler and Akleman (1998). The contemporaneous causal flows among prices were explored, which is not only important itself but also crucial to the VAR-type innovation accounting. Applications of directed graph technique in economics are not yet commonplace. The technique is similar to a procedure recently suggested by Swanson and Granger (1997) which sorts out causal flow on innovations from a vector autoregression (VAR). The rest of the paper is organized as follows. Section II describes the data. Section III presents results of hypothesis testing based on cointegration and the error-correction model. Section IV further discusses price dynamics using directed graphs and innovation accounting. Finally, Section V concludes.

Data

Soybean meal prices in the United States (US), United Kingdom (UK), Argentina (AGN), and Brazil (BRZ) were obtained from Datastream International. The data covered January 1, 1991, to March 31, 1998, totaling 1891 daily observations for each price-time series. The prices used included Argentinean export prices (CIF Rotterdam) for soybean meal with 45 percent protein, Brazilian export prices (CIF Rotterdam) for soybean meal with 48 percent protein, U.S. active cash prices for soybean meal with 44 percent protein, and U.K. active cash prices for UK-produced soybean meal with 49 percent protein. Soybean meal prices in the United States, Argentina, and Brazil were originally denominated in terms of U.S. dollars, and soybean meal prices in the United Kingdom were converted into U.S. dollars using the appropriate daily exchange rate of UK pounds against U.S. dollars. The price differences due to quality differences and transportation costs may be captured by a properly defined constant term in the cointegration model, as explained in the next section.

Three features of the data set were unique in empirically investigating the issue of developed and developing country market segmentation or integration. First, compared to previous studies, results of this study are more likely to be free from the influence of governmental price controls. It has been argued that government intervention can fundamentally change cointegration of international commodity prices (Bessler and Peterson, 1996; Yang and Leatham, 1999). For example, the U.S. government historically manages many important agricultural commodities through its farm commodity programs, including soybeans. In contrast, no direct government interventions affect soybean meal; thus, market forces may more fully determine the supply and demand of soybean meal. Thus, soybean meal prices can be significantly more market-driven than many other agricultural commodities under previous studies.

Second, Argentina and Brazil are major producers and exporters of soybean meal, just like the developed countries, i.e., the U.S. and the U.K. (the UK price represents the European Union price, which is usually the fourth largest exporter.) This fact helps prevent a possible compounding effect of sampling smaller open developing economies. The theoretical models of open economies typically suggest that smaller open economies are much more likely to follow the prices determined by the "big players" (usually the large developed countries) in international commodity markets, whether they are already developed or still developing. Thus, previous works based on smaller open developing economies and large developed countries may not have revealed the true price relationship between large developed countries and large developing countries.

Third, Argentina and Brazil have been historically active in participating in regional trade agreements. Currently, they are members of the new Southern Common Market, known as MERCOSUR, which aims to liberalize the trade within the region (including Argentina, Brazil, Uruguay, and Paraguay). Regional economic integration is prevalent among many developing countries and this suggests that the special characteristics of price relationships among developing countries may be well represented in this study. The sample period of this study covers the period when Argentina and Brazil have been members of the MERCOSUR, which was initiated in March 1991.

A precondition of cointegration analysis requires establishing that each individual soybean meal price series is nonstationary and integrated on an order of 1. Two standard procedures were applied to examine the data's time-series properties. The first procedure used was the augmented Dickey-Fuller (ADF) regression model (Dickey and Fuller, 1981). The second test procedure used was one proposed by Phillips and Perron (1988). The null hypothesis of both tests states that the price series has a unit root. Therefore, if the reported test statistics are larger than the critical values, the null hypothesis cannot be rejected. Table 1 reports the unit root test results for price levels and first price differences. The results show that each price series is 1.

Cointegration, Error Correction and South-North Market Integration

The hypothesis testing was based on the framework of cointegration and the error-correction model. The cointegration analysis in this study employs the procedure developed by Johansen and Juselius (1990, 1994) and Johansen (1992). Let X_t denote a vector which includes the market prices (p) for the four countries under consideration

Table 1. Results of Unit Root Tests to Determine Stationarity of Prices

Country	Without Linear Trend		With Linear Trend	
	ADF[a]	PP[b]	ADF[a]	PP[b]
		Level Prices		
Argentina	−1.84	−6.36	−1.97	−9.77
Brazil	−1.59	−5.19	−1.58	−7.88
US	−1.47	−6.25	−0.97	−6.80
UK	−2.06	−8.05	−2.39	−12.89
		1st Difference of Prices		
Argentina	−19.39	−1171.8	−19.45	−1171.8
Brazil	−15.65	−1117.6	−15.69	−1116.6
US	−18.65	−1096.5	−18.69	−1096.7
UK	−17.93	−1036.5	−17.96	−1036.8

[a] Test for the presence of a unit root developed by Dickey and Fuller (1981).
[b] Test for the presence of a unit root developed by Phillips and Perron (1988).
Notes: The optimal lags are selected by applying the principle of AIC+2 (Pantula *et al.*, 1994). The critical value of the ADF unit root tests with constant and without trends is −2.86 at the 5% level. The critical value of the ADF unit root tests with constant and with trend is −3.41 at the 5% level. The critical value of the PP unit root tests with constant and without trends is −14.1 at the 5% level. The critical value of the PP unit root tests with constant and with trends is −21.7 at the 5% level.

$$\left(p = 4 \text{ and } X_t = \begin{pmatrix} X_{1t} \\ X_{2t} \\ X_{3t} \\ X_{4t} \end{pmatrix} \text{ in this study, where } \right.$$

$$\left. \text{1-AGN, 2-BRZ, 3-US, 4-UK} \right)$$

and it can be modeled in an error-correction model (ECM):

$$\text{(1)} \quad \mathrm{H}_0: \Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu + e_t \quad (t = 1, \ldots, T).$$

Including a constant term μ in equation (1) is important when considering transportation costs and price differentials associated with commodity quality differences. The μ in the above ECM may account for relatively constant transportation costs and quality price differentials, or the constant transportation costs and quality price differentials with a time trend.

We first tested the hypotheses of north-south market integration versus segmentation by determining the number of cointegrating vectors, r, as follows:

$$\text{(2)} \quad H_1(r): \Pi = \alpha\beta'.$$

If there are r cointegrating vectors among p markets, this implies the presence of $p - r$ common trends. If we expect all p markets to be integrated as a single market, r should be found equal to $p - 1$. The hypothesis of north-south market segmentation predicts that two common trends exist for these four countries, which may be the sum of one common trend between two developing countries and another common trend between two developed countries. By contrast, the hypothesis of north-south market integration predicts that one common trend may prevail across these developing and developed countries.

A trace test was conducted to determine r. The null hypothesis for the trace test is that there are at most r $(0 \leq r \leq p)$ cointegrating vectors, where p is the dimension of the vector. The trace test results of H_1 are reported in Table 2. Following the sequential testing procedure suggested by Johansen (1992), we found that three cointegrating vectors with a constant are included in the cointegrating space. This clearly rejects the prediction from

Table 2. Johansen Trace Test of H_1 on Four Soybean Meal Markets[a]

	Without Linear Trend		With Linear Trend	
HO[b]:	T[c]	C[d] (5%)	T[c]	C[d] (5%)
r = 0	129.80	53.42	129.70	47.21
r ≤ 1	75.33	34.80	75.24	29.38
r ≤ 2	27.92	19.99	27.84	15.34
r ≤ 3	2.41	9.13	2.40	3.84

[a] The critical values are from Tables B.2 and B.3 in Hansen and Juselius (1995).
[b] r is the number of cointegrating vectors.
[c] T is the trace test statistics.
[d] C is the trace test critical value.

the hypothesis of north-south market segmentation.

However, further evidence for the hypothesis of north-south market integration requires exact identification of cointegrating β vectors. Mathematically, this type of identification can be expressed as:

$$(3) \quad H_2 : R'\beta = 0.$$

In the context of the hypothesis of north-south market integration, we specifically tested such restrictions on β which yield the following restricted β^*:

$$\beta^* = \begin{bmatrix} 0 & 1 & -1 & 0 & * \\ 1 & -1 & 0 & 0 & * \\ 0 & 0 & 1 & -1 & * \end{bmatrix}$$

where * denotes unrestricted constants in the cointegration space. The likelihood ratio test results for H_2 are summarized in Table 3. The χ^2 test statistics suggest no rejection of the projected restrictions. Consequently, this study verified the structure of the LOP as suggested by the hypothesis of north-south market integration, i.e., a single price holds across both the two developed countries and the two developing countries.

We were also interested in investigating which country is the primary information source that drives a single common trend in the long run on the international soybean meal market. This was done by performing a weak exogeneity test of the market price X_i. The weakly exogenous price X_i may be argued to cause other prices in the long run. The hypothesis testing was framed as the following:

$$(4) \quad H_3 : B'\alpha = 0.$$

Results of testing H_3 for weak exogeniety of α, i.e., $\alpha_{ij} = 0$ ($i = 1,2,3,4$; $j = 1,2,3$) are summarized in Table 3 and show that α_{3j} is equal to zero at the 5-percent level, but that α_{1j}, α_{2j}, α_{4j} are not. Considering the identified LOP structure in β matrix, we finally have

$$(5) \quad \alpha\beta' X_{t-1} = \begin{bmatrix} -0.027^* & -0.045^* & -0.018^* \\ -0.036^* & -0.005^* & -0.017^* \\ 0 & 0 & 0 \\ 0.018^* & -0.006 & 0.036^* \end{bmatrix} \times \begin{bmatrix} 0 & 1 & -1 & 0 & -43.5 \\ 1 & -1 & 0 & 0 & 10.9 \\ 0 & 0 & 1 & -1 & 77.0 \end{bmatrix} \times \begin{bmatrix} P_{AGN} \\ P_{BRZ} \\ P_{US} \\ P_{UK} \\ 1 \end{bmatrix}_{t-1}.$$

Responses to perturbations in each of the long-run relations are given in the first matrix on the right-hand-side of equation (5). Perturbations in the long-run equilibrium are given by $\beta' x(t-1)$, the second matrix and vector on the right-hand side of equation (5). The * denotes alpha matrix elements in which the t-statistic is significant at the 5-percent level. Each alpha magnitude can be interpreted based on the particular normalization used on each beta vector.

Changes in the price for Argentina showed a significant negative response to perturbations in all the three cointegrating vectors. When the Brazil price was high relative to its historical, long-run relation to the US price, the Argentina price fell in the subsequent period by $0.027z_1$ $(t-1)$ (where x_2 $(t-1)$ $-$ x_3 $(t-1)$ $-$ $43.5 = z_1$ $(t-1)$). Similar interpretations exist for the response of the Argentina price to perturbations in the long-run relations between the Argentina and Brazil prices (where x_1

Table 3. Test of Hypothesis H_2: $R'\beta = 0$ and H_3: $B'\alpha = 0$

Hypothesis	χ^2 Test Statistics	Degree of Freedom	Results[a]
H_2: Test of market integration hypothesis			
$\beta_{12} + \beta_{13} = 0 \quad \beta_{11} = \beta_{14} = 0$			
$\beta_{21} + \beta_{22} = 0 \quad \beta_{23} = \beta_{24} = 0$	8.61	9	F
$\beta_{33} + \beta_{34} = 0 \quad \beta_{31} = \beta_{32} = 0$			
H_3: Test of weak exogeniety of adjustment coefficients under the restricted β:			
$\alpha_{1j} = 0$ for j = 1, 2, 3	47.68	12	R
$\alpha_{2j} = 0$ for j = 1, 2, 3	46.39	12	R
$\alpha_{3j} = 0$ for j = 1, 2, 3	16.80	12	F
$\alpha_{4j} = 0$ for j = 1, 2, 3	37.80	12	R

[a] R denotes the rejection of the null hypothesis β and F denotes failure to reject the null hypothesis at a 5% significance level.

$(t-1) - x_2\,(t-1) + 10.9 = z_2\,(t-1)$), and the US and UK prices (where $x_3\,(t-1) - x_4\,(t-1) - 77.0 = z_3\,(t-1)$). Not surprisingly, Argentina responded most significantly to perturbations in the Argentina and Brazil long-run equilibrium.

Brazil showed a significant negative response to disturbances in the first and third long-run relations. If the Brazil price was high, relative to the long-run equilibrium with the US (if $z_1\,(t-1)$ is positive), then the Brazil price decreased by $0.036z_1\,(t-1)$ in the next period. Similarly, if the US price was high in period $t-1$, relative to the long-run equilibrium with the UK price, then the Brazil price in period t fell by $0.017\,z_3\,(t-1)$. Interestingly, Brazil did not respond significantly to perturbations in the Argentina-Brazil long-run equilibrium (which may suggest that the Brazil price is exogenous relative to the Argentina price in the long run). Instead, it responded most significantly to disturbances in the Brazil-US equilibrium.

The most interesting findings occurred in the case of the US. The US market appeared not to respond significantly to perturbations in any of the three long-run relations. This is an indication that the US market drives the single common trend across the four country markets in the long run. The larger production and domestic consumption in the U.S. market may explain this finding. Here it is also interesting to note that the export share of Brazil was much larger than that of U.S. in the international market during the sample period, but it did not help Brazil gain the price leadership.

Finally, the UK market responded significantly in a positive manner to shocks in the first and third vectors and insignificantly to shocks in the second vector. Thus, when the Brazil price was high relative to its long-run relation with the US in period $t-1$, the UK price in the subsequent period increased by $0.018\,z_1\,(t-1)$. In addition, when the US price was high relative to its long-run equilibrium with the UK, the UK price responded positively by $0.036\,z_4\,(t-1)$. Similar to Brazil's case, the UK price did not respond significantly to perturbations in the Argentina-Brazil long-run equilibrium, but instead responded most significantly to disturbances in the UK-US equilibrium. In summary, in terms of the adjustment toward the common trend, the results showed that the US is the most exogenous and Argentina is the least exogenous while it was not clear whether the UK or Brazil is more exogenous. However, a more complete picture of exogeneity should also consider short-run dynamics, which will be addressed in the next section using innovation accounting based on the estimated ECM.

Residuals on the ECM estimation are reasonably well behaved. Lagragian multiplier-type tests on first- and fourth-order autocorrelation on residuals (chi-squared tests) reject the null of white noise residuals at p-values of

0.08 and 0.13, respectively. Lagragian multiplier-type tests on five-order ARCH residuals from each equation resulted in the following statistics (which are subject to the chi-squared distribution with five degrees of freedom): 0.81, 0.85, 100.14, and 1.76 for the Argentina, Brazil, US and UK equations, respectively. These statistics suggest a non-constant variance in the innovations from the US equation. Further analysis of the US equation indicates that this ARCH-like behavior in the residuals may have resulted from the weak exogeneity of this market, which was not rejected at any conventional significance level. Following recommendation by Hansen and Juselius (1995, p.12), we conducted the above ECM estimation again, conditioning on the weak exogenous US market prices. In this case, the weakly exogenous variable US prices was still included in levels in the cointegration space and in current and lagged differences in the short-run dynamics. All reported results were confirmed to be qualitatively unchanged.

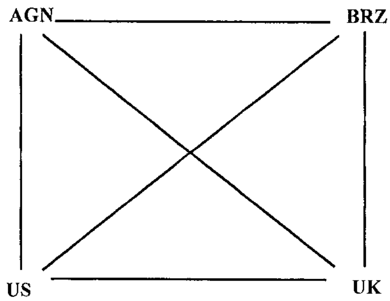
Direct Graphs and Impulse-Response Analysis

To further visualize the dynamic price relationship among the four countries, a directed graph was employed to aid innovation accounting based on the estimated ECM. The estimated cointegrating vectors characterized the stationary long-run equilibrium relationships, and the above ECM was used to summarize the period-by-period influence each market price had on the other market prices of the four variable systems. However, the adjustments that established these relationships in response to various shocks from the international market and the strengths of these dynamic relationships remain unspecified. Because the individual coefficients of the ECM (particularly those of short-run dynamics) are hard to interpret, we inverted the estimated ECM to derive the corresponding level VAR representation. We then conducted impulse-response analysis based on the equivalent level VAR to summarize the dynamic interactions among the four market prices. The manner in which we conducted the innovation accounting addressed the imposition of cointegration constraints in the nonstationary VAR, which was recently proven to be crucial in yielding consistent impulse responses and forecast error decompositions (Phillips, 1998).

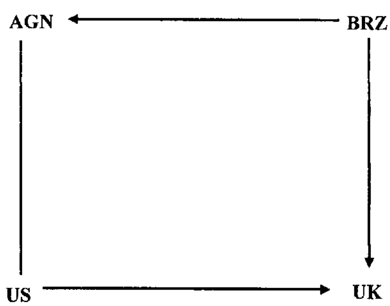
The method for treating contemporaneous innovation correlation is critical to such an impulse-response analysis (or forecast error variance decomposition) (Swanson and Granger, 1997). We followed the factorization commonly referred to as the "Bernanke ordering" which requires writing the innovation vector (u_t) from the estimated error-correction model as $\mathbf{A}u_t = v_t$, where $\mathbf{A}$ is a 4×4 matrix and v_t is a 4×1 vector of orthogonal shocks. It was common in earlier VAR-type (vector autoregression-type) analyses to rely on a Choleski factorization, so that the A matrix is lower triangular, to achieve a just-identified system in contemporaneous time. Similar to Bessler and Akleman (1998) we applied directed graph algorithms such as those given in Spirtes, Glymour and Scheines (1993) to place zeros on the A matrix. A directed graph is an assignment of causal flow (or lack thereof) among a set of variables (vertices) based on observed correlation and partial correlation. Our four-variable error-correction model based on the identifying restrictions resulted in the following innovation correlation matrix (lower triangular entries only are printed in order: Δx_1, Δx_2, Δx_3, and Δx_4):

$$(6) \qquad V = \begin{bmatrix} 1.0 & & & \\ 0.28 & 1.0 & & \\ 0.02 & 0.03 & 1.0 & \\ 0.07 & 0.10 & 0.08 & 1.0 \end{bmatrix}.$$

Directed graph theory explicitly points out that the off-diagonal elements of the scaled inverse of this matrix (V or any correlation matrix) are the negatives of the partial correlation coefficients between the corresponding pair of variables, given the remaining variables in the matrix (Whittaker 1990, page 4). Directed graphs as given in Spirtes, Glymour and Scheines (1993) provided an algorithm (PC algorithm) for removing edges between markets and directing causal flow of information be-



Panel A. Complete Undirected Graph on Innovations from Equation (1).



Panel B. Final Directed Graph on the Model.

Figure 1. Contemporaneous Causal Flow Patterns Using Directed Graphs
Panel A. Complete Undirected Graph on Innovations from Equation (1)
Panel B. Final Directed Graph on the Model

tween markets. The algorithm begins with a complete undirected graph, where innovations from every market are connected with innovations with every other market. Figure 1, Panel A shows this complete undirected graph on innovations from the error-correction model given in equation (1). The algorithm removed edges based on vanishing correlation and partial correlation, the later measure based on the scaled inverse correlation matrix as explained above. Edges between variables were removed sequentially based on either vanishing zero-order correlations (unconditional correlations) or vanishing conditional correlations, where conditioning was done on all possible sets with members 1, 2, ... K-2, where K was the number of variables studied.

The notion of *sepset* is very important to assigning the direction of causal flow between variables which remain connected after all possible conditional correlations have been passed as nonzero. The conditioning variable(s) on removed edges between two variables is called the *sepset* of the variables whose edges have been removed (for vanishing zero-order conditioning information (unconditional correlation) the *sepset* is the empty set). Edges are directed by considering triples X --- Y --- Z, such that X and Y are adjacent as are Y and Z, but X and Z are not adjacent. Direct the edges between triples X --- Y --- Z as X → Y → Z if Y is not in the *sepset* of X and Z. If X → Y, Y and Z are adjacent, X and Z are not adjacent, and there is no arrowhead at Y, then Y --- Z should be positioned as Y → Z. If there is a directed path from X to Y, and an edge between X and Y, then X --- Y should be positioned as X → Y.

In applications, Fisher's z statistic is used to test whether conditional correlations are significantly different from zero. Fisher's z statistic can be applied to test for significance from zero, where $z((i, j|k)n) = 1/2(n - |k| - 3)^{1/2} \times \ln\{(|1 + (i, j|k)|) \times (|1 - (i, j|k)|)^{-1}\}$ and n is the number of observations used to estimate the correlations, (i, j|k) is the population correlation between series i and j conditional on series k (removing the influence of series k on each i and j), and $|k|$ is the number of variables in k (that we condition on). If i, j and k are normally distributed and r(i, j|k) is the sample conditional correlation of i and j given k, then the distribution of $z((i, j|k)n) - z(r(i, j|k)n)$ is standard normal.

We used the software TETRAD II (Scheines *et al.*, 1994) which contains the PC algorithm and its more refined extensions to conduct directed graph analysis. Figure 1 gives both the complete undirected graphs and the final directed graphs on innovations from our four-market error-correction model (Equation 1). Panel A is the starting point from which edges are removed and edges directed according to the plan outlined above (actually according to the TETRAD II programs (Spirtes *et al.*, 1994)). Panel B is the ending point. At the 5-percent level, we found the directed edges as given in panel B. Applying a 5-percent significance level, we saw edges running from Brazil to the UK and from the US to the UK. Because Brazil is a larger producer and exporter than Argentina and there is evidence that Brazil is more exogenous than Argentina in the long-run equilibrium adjustment, we further hypothesized that a causal flow exists from Brazil to Argentina.

We explicitly tested these restrictions using the likelihood ratio test for over-identification given in Doan (1992). Our identification restriction implied three zero restrictions (there were six lower triangular elements or their transpose elements, which can be nonzero in a just-identified model). These restrictions resulted in a chi-squared statistic of 5.91. With three degrees of freedom, we rejected these zero restrictions at a p-value of 0.12, suggesting that the restrictions were consistent with the data.

Under the ordering of innovations as generated by the directed graph at the 5-percent level, 100-day impulse responses associated with the error-correction model are given in Figure 2. All country market prices responses were positive to shocks from other countries (except a few very small negative responses of the US price to shocks from other countries in less than the first 50 days). Obviously, the US is the most exogenous market studied. The US price had little response to price shocks from Brazil and the UK, and some response to price shock from Argentina. In contrast, other countries had much stronger responses to price shocks from the US market. This finding from impulse-response analysis was based on price interactions among four markets in both the short run and the long run, because we incorporated both the short-run dynamics and long-run relationships in generating the impulse-response functions. Another noticeable characteristic of the impulse-response functions was that the effect of a shock from one country to other countries, though with varying strengths, tended to persist in the longer run (100 days). Following Orden and Fisher (1993), we interpreted this as an indication of long-run relationship constraints.

Conclusion

This study evaluated two competing hypotheses on price relationships among developed and developing counties. The hypothesis of north-south market integration was consistent with the original idea of the LOP. As an interesting alternative, some economists (e.g., Monke and Taylor, 1985; Cristini, 1995) have subscribed to the hypothesis of north-south market segmentation, which argues that one price may hold within the developing countries and another single price in the developed countries. There may exist some good reasons to speculate on the north-south market segmentation. For one reason, the economies in the major developed countries were obviously more coordinated with each other than with the economies of the developing countries. The developing countries also focused on strengthening their own economic relationship through regional trade grouping, etc. Particularly, the two developing countries in our study were characterized by similar high inflation experiences and actively participated in the same regional economic integration during the sample period.

The results of this study clearly rejected north-south market segmentation, an interesting variant for the LOP. We found that two developing and two developed countries were fully integrated, and that the LOP holds across these four countries in the long run. This suggests that the market force of international competition may integrate spatially separated markets well. Thus, soybean meal markets in these four countries should be considered as being integrated into one international market in modeling international soybean meal trade. Further, both ECM hypothesis testing and impulse-response analysis indicated that the US is the leading force driving the single common price trends on the international soybean meal market both in the short run and long run.

Finally, further research along this line may consider using commodity future prices in different countries (if available) to test the LOP.

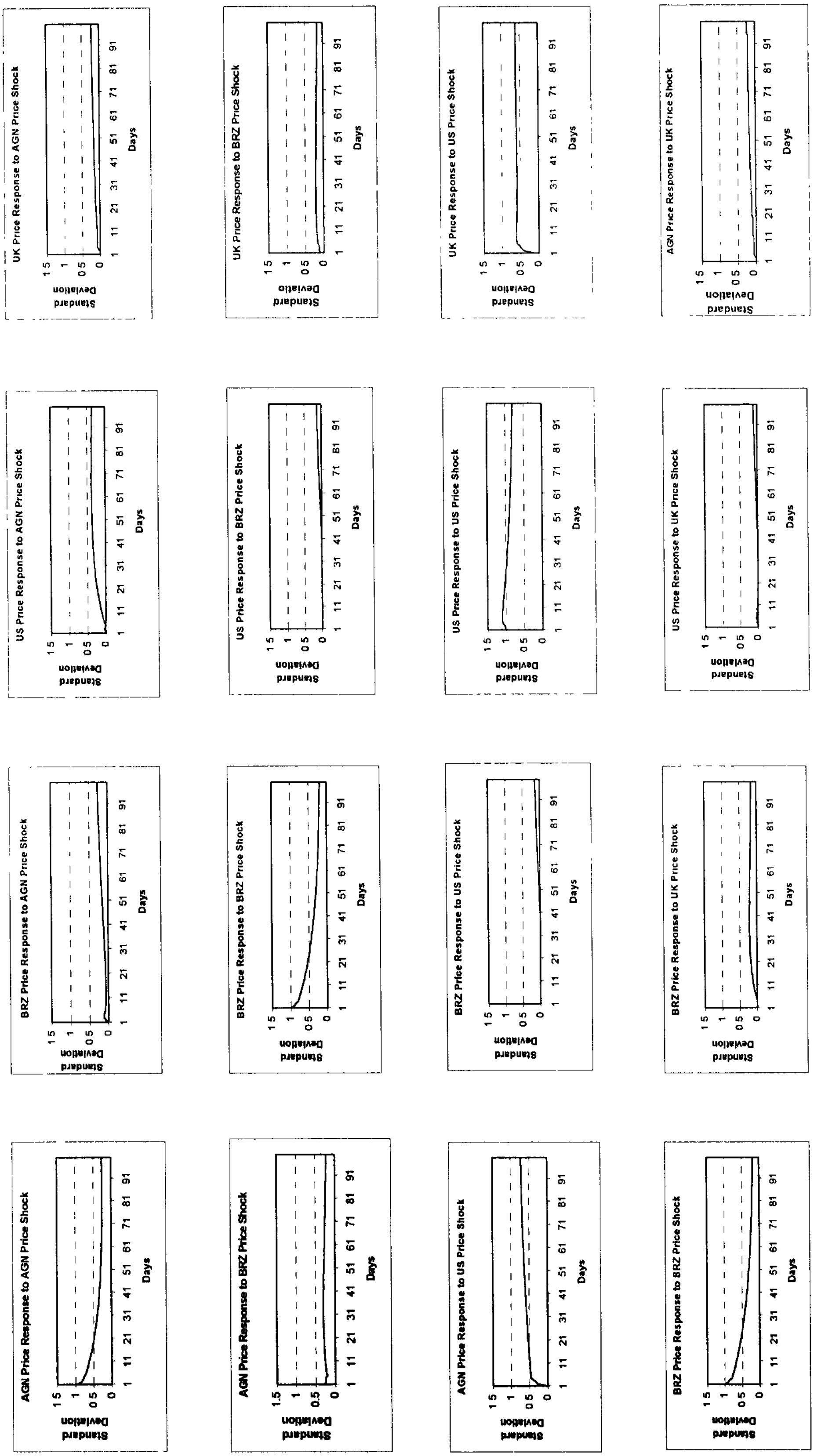


Figure 2. Impulse Response Analysis Results

Protopapadakis and Stoll (1983, p. 1433) argued that the LOP can be investigated in "its purest form" when commodity futures prices are used. Protopapadakis and Stoll (1986) further pointed out that the LOP received strong support when using commodity futures or forward prices, but only modest support when using cash prices. Consistent with these arguments, using cointegration analysis Yang and Leatham (1999) also highlighted the important difference between commodity cash and futures prices in processing and transmitting price information. Similar research should be also conducted on other internationally competitive commodities to further test the robustness of rejecting the hypothesis of north-south market segmentation.

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