MTID DISCUSSION PAPER NO. 76

POST-URUGUAY ROUND PRICE LINKAGES BETWEEN DEVELOPED AND DEVELOPING COUNTRIES: THE CASE OF RICE AND WHEAT MARKETS

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ABSTRACT

The Uruguay Round Agreement on Agriculture attempted to lower distortions in the global agricultural markets. However, the significant fall in commodity prices in late 1990s may have reduced the incentives for both developed and developing countries to better integrate into the world markets. This study analyzes price linkages and adjustment between developed and developing countries during the post-Uruguay Round period. Prices of two key commodity markets, long-grain rice and medium-hard wheat, are assembled for major exporters and producers. Results from the multivariate cointegration analysis suggest partial market integration between developed and developing countries in the post-Uruguay Round period. Developed countries are found to be price leaders in these two markets, and in most cases, the changes in their prices have relatively large impacts on those of the developing countries. The new entrants into world markets (Vietnam and Argentina) have faced considerable price adjustment due to changes in the developed countries' prices.

Key Words: Price Linkage, Developing Countries, Rice and Wheat Markets

JEL Classification Codes: Q17, O13, Q11

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POST-URUGUAY ROUND PRICE LINKAGES BETWEEN DEVELOPED AND DEVELOPING COUNTRIES: THE CASE OF RICE AND WHEAT MARKETS

Navin Yavapolkul[1], Munisamy Gopinath,[2] and Ashok Gulati[3]

1. INTRODUCTION

The Uruguay Round Agreement on Agriculture (URAA) in 1994 initiated multilateral reform of the agricultural sector to better integrate global markets. The current Doha Developmental Round of the World Trade Organization (WTO) expects to deepen the reforms of the URAA despite the setbacks at Cancun, Mexico. However, significant differences between the views of developed and developing countries persist on market access, export subsidies and domestic support (IFPRI Forum, 2003; The Economist, 2003; The World Bank, 2003; The WTO, 2003; New York Times, 2003). For instance, the reform of the domestic support policies of developed countries, especially the increasing use of income support and its effects on global commodity prices, has been a key component of developing countries' proposals to WTO.

Simultaneously, commodity prices have fallen significantly after a peak during 1995-1997. In the case of long-grain rice, for example, fob prices at US gulf ports fell from a high of \$300/ton in early 1997 to about \$120/ton by 2001 (Economic Research Service, US Department of Agriculture). Similar, but less stark trends have been observed for wheat, soybeans, cotton and corn. Many countries maintain domestic or

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trade policies, which protect or safeguard their farmers from declining world market prices. The United States maintains price floors or loan rates for major commodities, which are supported by (loan) deficiency payments. Several developing countries have bound tariffs on agricultural commodities, which have allowed them to raise applied tariffs to counter the decline in world market prices.

The key question we address in this paper is whether price linkages exist between developed and developing countries in two key commodity markets – long-grain rice and hard wheat – in the post-URAA era. In addition, if such linkages exist, how have the commodity prices of developed countries impacted those of their developing counterparts and vice versa? Have developed or developing countries adjusted the most to world market changes during the second half of the 1990s? Answers to these questions have important implications for the current round of trade negotiations. The partial or full integration of commodity markets between developed and developing countries in an era of significant price decline can confirm channels for the transmission of a country's policies into other countries. If the adjustment to declining world prices were to fall disproportionately on one set of countries, it would affect price stability, income and welfare of their agricultural households. If agriculture is a key foreign exchange earner, the adjustment would affect also the overall economy.

Prior to distinguishing our effort from other studies, note that we define a spatially integrated market in the spirit of Stigler (1969) as the "area within which the price of a good tends to uniformity, allowances being made for transportation cost." This condition is generally consistent with the weak version of law of one price (LOP), which states that a linear and proportional relationship exists between prices of two markets. While the definition of market integration is a common feature between our study and prior investigations (Mohanty, Peterson and Smith, 1996; Taylor et al., 1996; Goodwin , 1992a and 1992b),[4] we differ from the latter by focusing on developed and developing countries in the post-URAA era of declining commodity prices (Yang et al., 2000). We also focus on Granger-type causality among prices and the adjustment by individual country prices to stay in these relationships.

The long-grain rice and hard wheat markets are selected because they have significant participants from developed and developing countries. The countries chosen for the rice market analysis are the United States, Thailand, Vietnam and India, while those for the wheat market include the United States, Canada, Australia, Argentina and India. The choice on countries and commodities depended also on data availability and comparability. Monthly price data from 1996/1997 to 2002 have been assembled for this purpose. A vector error correction process, which allows for the identification of cointegration or long-run price equilibrium/relationship, is estimated for each of the two markets. We tested the possibility of perfect integration/arbitrage, and the magnitude of adjustment of individual country prices to stay on the long-run relationships. Finally, impulse response analyses based on directed graphs are carried out to track the short- and long-term responses of prices to economic shocks.

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[4]Spatial competitive equilibrium and market integration are not interchangeable concepts. This study's interpretation of market integration is similar to that of Ravallion (1987), who noted, "…one can be interested in testing empirically for spatial market integration, without wishing to rest the case for or against Pareto optimality on the outcome. Measurement of market integration can be viewed as basic data for an understanding of how specific markets work." See also Dawson and Key (2002).

2. METHODOLOGY

This section briefly describes a vector error correction model (ECM) to identify the cointegrating vectors in the rice and wheat markets. The procedures to specify and test the ECM, and to identify the cointegrating vectors and the adjustment coefficients are also outlined in the following section. More details on the procedures and tests can be found in Johansen (1995).

2.1 NON-STATIONARY PROCESS AND COINTEGRATION

Time-series data are tested for stationarity, which means each observation of the data independently takes on a single random event. Stationary data have stable and observable mean and variance, but most economic time series including commodity prices fail to satisfy this criterion. Non-stationary data have a strong trend and their mean and variance shift overtime reflecting changes in the data generating process (Granger and Newbold, 1974). A non-stationary process is said to be integrated of order d, referred to as I(d), if first differencing for d times produces a stationary process. Stationary processes are simply denoted as $I(0)$.

Given two $I(1)$ series, a linear combination of these two series is generally $I(1)$. If, however, there exists a linear combination of the two series that is $I(0)$ then the two series are cointegrated. Alternatively, non-stationary series of a common order are said to be cointegrated, if there exists one or more linear combinations, which transform the series to a stationary system.

2.2 MULTIVARIATE COINTEGRATION ANALYSIS

To establish long-run equilibrium relationships among prices in a multivariate context, we draw on Johansen and Juselius' (1988 and 1990) multivariate cointegration technique. It begins with the description of a vector autoregressive model (VAR) model, where a vector of prices ($p \times 1$) at time t are related to vectors of past prices. The VAR model at time t can be written as:

$$P_t = \Pi_1 P_{t-1} + \Pi_2 P_{t-2} + \ldots + \Pi_k P_{t-k} + \mu + \Phi D_t + \varepsilon_t \qquad (2.1)$$

where $t = 1, \ldots, T$, denotes the number of observations. P_t is a vector of dimension $(p \times 1)$ corresponding to the number of price series. The matrix Π_q $(q = 1,\ldots,k)$ has a $(p \times p)$ dimension corresponding to an autoregressive relationship with its own past values and other lagged variables in the system. Parameter μ is a constant term and variable D_t denotes centered, seasonal dummies which sum to zero over the sample period.

If all prices are $I(1)$, equation (2.1) can be rewritten in first differences as:

$$\Delta P_t = \Gamma_1 \Delta P_{t-1} + \Gamma_2 \Delta P_{t-2} + \ldots + \Gamma_{k-1} \Delta P_{t-k+1} + \Pi P_{t-k} + \mu + \Phi D_t + \varepsilon_t, \qquad (2.2)$$

where $\Gamma_j = -(I - \Pi_1 - \Pi_2 - \ldots - \Pi_j)$ for $j = 1,\ldots, k-1$, and $\Pi = -(I - \Pi_1 - \Pi_2 - \ldots - \Pi_k)$ are p x p matrices of parameters. This model of first differences is often referred to as a vector error correction process or model (ECM). The Γ_j are parameter matrices summarizing short-run relationships among the prices. The Π matrix contains the long-run parameters and 3 cases of it are feasible: (i) Π has full rank, which indicates that the process P_t is stationary and thus the VAR representation in levels is appropriate; (ii) Π has rank zero, a null matrix containing no long-run information (no cointegration); and (iii) the rank of Π is r, where $0<r<p$, and the system is said to have cointegrating rank r. In this case, there exists two $(p \times r)$ matrices α and β, such that $\Pi = \alpha\beta'$. The Π matrix can be decomposed into the cointegration vector, β, and the adjustment coefficients, α.

To determine the rank of Π we require its ordered eigenvalues $\hat{\lambda}_1 > ... > \hat{\lambda}_p$, which can be derived as solutions to the equation:

$$\left|\lambda S_{kk} - S_{k0} S_{00}^{-1} S_{0k}\right| = 0 \tag{2.3}$$

where

- $S_{ij} = T^{-1} \sum_{i=1}^{n} R_{it} R'_{jt}$ for $(i,j = 0,k)$ and S_{ij} is a $p \times p$ matrix of residual product.
- R_{0t} and R_{kt} are p-dimensional vector of residuals from the regression of ΔP_t on Z_t, and P_{t-k} on Z_t, respectively, corresponding to t$^{\text{th}}$ observation and k^{th} lag:

$$R_{0t} = \Delta P_t - \left(\sum_{t=1}^{T} \Delta P_t Z'_t\right)\left(\sum_{t=1}^{T} Z_t Z'_t\right)^{-1} Z_t$$

$$R_{kt} = P_{t-k} - \left(\sum_{t=1}^{T} P_{t-k} Z'_t\right)\left(\sum_{t=1}^{T} Z_t Z'_t\right)^{-1} Z_t \tag{2.4}$$

This procedure allows testing the cointegration rank (r) sequentially from zero to (p-1). The likelihood ratio (LR) or trace test statistic for the null hypothesis (H_2) – that the rank of Π is less than or equal to r – is given as:

$$\lambda_{trace} = -2\ln(Q; H_2 \mid H_1) \equiv -T \sum_{i=r_0+1}^{p} \ln(1-\hat{\lambda}_i^2), \qquad (2.5)$$

where $\hat{\lambda}_i^2$ are the solutions to the eigenvalue problem in (2.3) under H_2. The alternative hypothesis, H_1, denotes no cointegrating vectors, i.e., $r = p$, meaning Π has full rank. The trace test statistic is distributed as a modified chi-square with (p-r) degree of freedom, whose critical values are provided by Johansen and Juselius (1990).

Before applying the multivariate cointegration technique two specification issues are considered: the deterministic components of the ECM (intercepts and dummies) and the lag length. The LR tests for the deterministic components are outlined in Johansen and Juselius (1990, 1992, and 1994), and Johansen (1988). The determination of lag length based on a residual analysis and a LR ratio test is presented in Johansen (1995).

3. DATA DESCRIPTION

The production and export of rice and wheat are concentrated in a few developed and developing countries. In this section, a description of prices used in the cointegration analysis is presented.

3.1 RICE MARKET

The international rice market is often described as thin, volatile, risky, and concentrated with a few major exporters (Child, 2001). The six major rice exporters are Thailand, Vietnam, US, India, China, and Pakistan, respectively accounting for 27.5, 16.1, 11.8, 10.9, 9.0, and 8.3 percent of world exports (1996-2002 average, Foreign

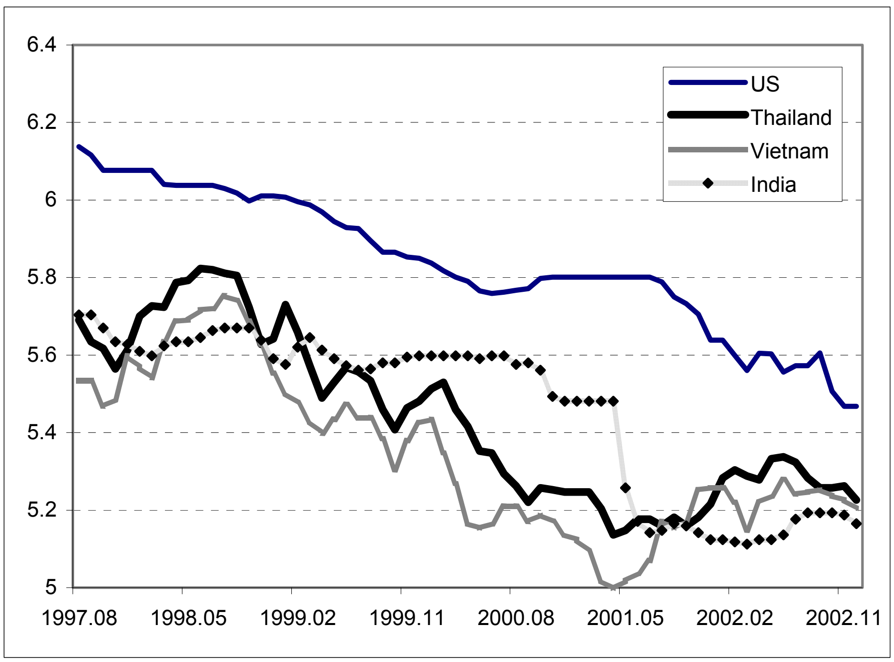
Agricultural Service, USDA). The major world rice producers are China and India respectively accounting for 33.1 and 22.1 percent of share of world production (1996-2002 average, Foreign Agricultural Service, USDA), while the combined share of Thailand, Vietnam and the US is only 10.7 percent. Thus, international rice prices are likely determined by world supply and stocks.

International trade in rice is subject to a high degree of product differentiation. Two important attributes of rice are often emphasized– percent-broken and type of grain. The percent broken indicates the quality of the milling process and, therefore, the lower percent-broken carries a premium. Four types of rice are traded worldwide: indica, japonica, aromatic, and glutinous. We chose indica rice for our study since it accounts for about 75 percent of global rice trade annually. An ideal study of rice market integration would include all six rice exporters. However, two countries, Pakistan and China, could not be included in our study. Given the focus on post-URAA era, i.e., post-1995, sufficient observations are required to establish the asymptotic properties of ECM's parameter estimates. The use of monthly price data rules out Pakistan, whose price data contain several "no quote" (missing) observations due to a significant drop in exports (USDA Rice Yearbook, 2002). Although China imports indica rice, the sparsely available monthly export price data are of poor quality.

The data sample contains 65 observations during the period of August, 1997 to December, 2002. The US price series is long grain No. 2 broken (not to exceed 4 percent), FOB Gulf milled at Houston, Texas. The data are based on monthly average of the midpoint for reported weekly low and high quotes. Thailand rice price is 100%

Grade B, Thai white milled rice, calculated by simple average of weekly price quote, including cost of bags. There is evidence that long-grain indica rice from these regions are close substitutes. Brorsen et al., (1984) studied the dynamic relationships of rice import prices in Europe and concluded that US and Thai rice of this type are close substitutes. The 5% broken milled rice for both Vietnamese and Indian price series are often chosen to compare with the two prior series.[5] Thus, the selected Vietnamese and Indian price are 5% broken fob vessel at Ho Chi Minh City and Bombay port, respectively, and simple averages of weekly price quotes. Figure 1 presents a time series of the rice market's prices.

Figure 1—Time series of international rice prices



Note: all prices are expressed in terms of natural logarithms.
Source: Complied by author.

[5]Due to non-availability of the Vietnamese data for 1996, the 1997-2002 period is chosen. Sources: Thai rice: weekly price reports, US Embassy, Bangkok; US rice: Rice Market News, Agricultural Marketing Service, USDA; Vietnamese rice: Vietnam Industry source, Saigon; India: the Creed Rice Market Report, Creed Rice Co., Inc., Houston, Texas.

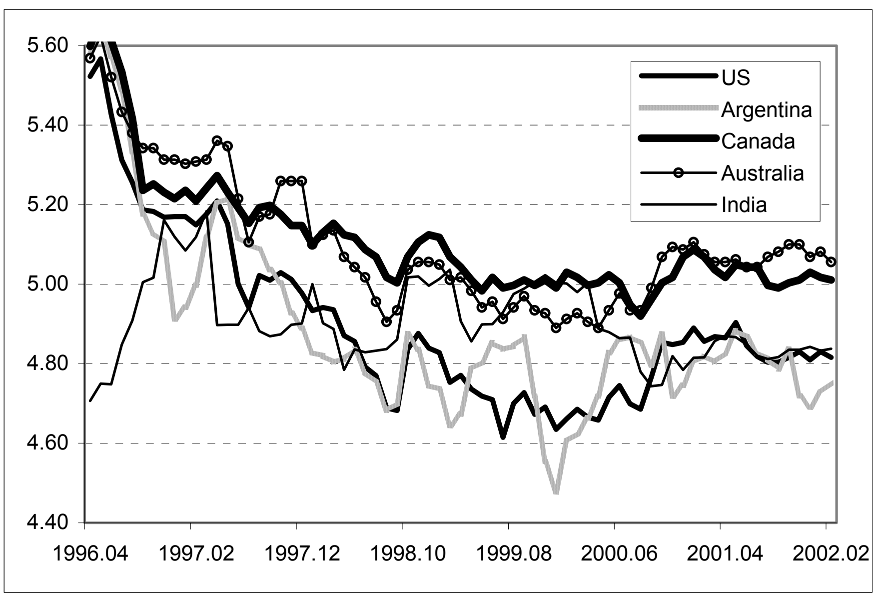
3.2 WHEAT MARKET

The world export of wheat is approximately 105 million tons each year, much larger than the 22 million tons of rice exports. Eighteen percent of the world production of wheat is annually traded, ranked second among agricultural commodities (USDA Wheat Yearbook, 2002). There are a large number of wheat suppliers, but almost 80 percent of exports come from five major exporting countries – the US (26.1%), Canada (15.7%), Australia (14.8%), EU (13.9%) and Argentina (10.1%). Similar to the case of rice, some of the major wheat producers mostly serve domestic demand. China, EU, and India, account for 18.1, 16.9, and 11.9 of the world production, respectively (Foreign Agricultural Service, USDA).

The main criterion used for classifying wheat is "hardness," which is a milling characteristic determined by the protein content. The hardest varieties produce elastic dough appropriate for blending with lower-protein wheat to produce bread flour. Soft wheat with lowest protein content is milled into flour used for cakes, pastries, cookies and crackers. Highly specialized durum wheat is used to produce coarse flour for pasta (Ghoshray, Lloyd and Rayner, 2000; USDA Wheat Yearbook 2002). Veeman (1987), Wilson (1989) and Alston et al., (1994) pointed out that product characteristics and end uses vary by wheat type and so, any analysis treating all wheat as a homogeneous commodity would provide misleading results (Larue, 1991).

In this study, hard/medium hard wheat prices from US, Argentina, Canada, Australia, and India are analyzed. Thus, only wheat varieties for making bread are considered, which eliminates EU, a soft-wheat exporter. China's wheat price series are not available on a monthly basis. Monthly data used in this study cover the period April, 1996 to February, 2002, with 71 observations. The price series are for of US No.2 Hard Red winter, 12.5 % protein (FOB from Gulf port); high protein No.1 Canadian Western Red Spring wheat, 13.5% protein (FOB from St. Lawrence port); Argentinean Trigo Pan (FOB from Buenos Aires); Australian Standard White wheat, which contains lower protein content. The price data for India are Common Spring bread wheat at wholesale level (near ports) since FOB data are not readily available. Any remaining quality differences can be captured using the intercept in the ECM. Figure 2 presents the wheat market's price series.

Figure 2—Time series of international wheat prices



Note: all prices are expressed in terms of natural logarithms.
Source: Complied by author.

4. RESULTS

This section describes the empirical results from the cointegration analysis. The results of the non-stationarity and specification tests are presented in appendix 1 and 2. The Dickey-Fuller and Augmented Dickey-Fuller test results in Appendix 1 show that every price included in our study is an *I*(1) process. For both markets, the lag length of ECM is 6 (Appendix 2).

4.1 COINTEGRATION RANK

The trace test results (λ_{trace}) from equation (2.5) are presented in table 1 for rice and wheat markets. For each market, two specifications of the ECM with and without an intercept, case I and II respectively, are estimated. In most studies, the first failure to reject the null hypothesis by the trace test, given the testing order 0 to (*p-1*), is sufficient to identify the cointegrating rank (Goodwin, 1992a, 1992b; Brester and Goodwin, 1993; Mohanty et al., 1996; Taylor et al., 1996; Ismet, et al., 1998; Bierlen et al., 1998; Dawson and Key, 2001). Results in table 1 indicate that the rice and wheat markets contain one and two cointegrating vectors, respectively. Additional testing indicated evidence of an intercept in the ECM (linear trends in level data), case I, which can account also for transfer costs/quality differences in the prices (Dawson, 2002). The cointegrating rank (*r*) gives the number of stationary linear combinations of the price data, i.e., cointegrating vectors. Market integration is a linear and proportional relationship between prices, and thus, is consistent with the identification of at least one linear combination of prices that exhibits stability over time.

Table 1—Trace tests and Maximum Eigenvalue Tests for Rice and Wheat

	Case 1 (linear trend in level) H_2/1			Case II (no linear trend in level) H^*_2/2		
	λ trace test	λ trace (0.95)	Decision	λ^* trace test	λ^* trace (0.95)	Decision
The rice data (T-k=59 observation)						
$r = 0$	**58.66**	**47.18**	**Reject**	**76.34**	**53.35**	**Reject**
$r \leq 1$	14.66	29.51	Fail to Reject	28.08	35.07	Fail to Reject
$r \leq 2$	4.17	15.20	Fail to Reject	9.38	20.17	Fail to Reject
$r \leq 3$	0.29	3.96	Fail to Reject	1.00	9.09	Fail to Reject
The wheat data (T-k=65 observation)						
$r = 0$	**106.05**	**68.91**	**Reject**	**124.35**	**75.33**	**Reject**
$r \leq 1$	**53.85**	**47.18**	**Reject**	**69.46**	**53.35**	**Reject**
$r \leq 2$	17.80	29.51	Fail to Reject	32.99	35.07	Fail to Reject
$r \leq 3$	7.19	15.20	Fail to Reject	10.03	20.17	Fail to Reject
$r \leq 4$	2.34	3.96	Fail to Reject	2.41	9.09	Fail to Reject

[1] The critical values of λ_{trace} are taken from Johansen (1995, table 15.3).

[2] The critical values of λ^*_{trace} are taken from Johansen (1995, table 15.2).

The estimated p-dimensional matrix Π under H_2 (equation 2.2) contains two types of linear combinations – stationary and non-stationary. The cointegrating rank indicates the number of stationary combinations (r). Since non-stationary combinations are trending, they are referred to as common trends $(p-r)$. Studies of market integration commonly agree that a perfect integrated (strong LOP) system of prices should contain only one common trend and $(p-1)$ long-run relationships for all possible combinations of prices. The reason is that $(p-1)$ cointegrating vectors provide stronger support for the concept of a single international price. A single cointegrating vector implies that any single price can be solved for in terms of the other (p-1) prices. Thus, if one price in the system is fully representative of the set of p prices, then one would expect to find

$(p-1)$ cointegrating vectors (Goodwin, 1992b). However, if more than one common trend is found or if the number of cointegrating vectors is less than $(p-1)$, the market is said to be partially integrated. For instance, Goodwin (1992a) suggests that the international wheat market may be partially integrated since he found only one cointegrating vector among prices of the US, Canada, Australia, EU, and Japan.

4.2 COINTEGRATING VECTORS AND ADJUSTMENT PARAMETERS

Recall that the solution to the problem in equation (2.3) yields the estimated Π matrix as a product of two matrices, $\hat{\Pi} = \hat{\alpha}^2 \hat{\beta}^{2\prime}$ and the ordered eigenvalues $\hat{\lambda}_1^2 > ... > \hat{\lambda}_p^2$, where the superscript 2 refers to the solution under H_2. The corresponding eigenvectors are given by $\hat{V}^2 = (\hat{v}_1^2, \hat{v}_2^2, ..., \hat{v}_p^2)$, where $\hat{v}_i^2$ is a normalized p-dimensional vector such that $\hat{V}^{2\prime} S_{kk} \hat{V}^2 = I$. Once the cointegrating rank (r) is identified from the trace test, the first r columns of $\hat{V}^2$ become the cointegrating vectors, i.e., $\hat{\beta}^2 = \left(\hat{v}_1^2, ..., \hat{v}_r^2\right)$, while the matrix $\hat{\alpha}^2 = S_{0k}\hat{\beta}^2$ represents the adjustment coefficients. For the rice market:

$$\hat{\beta}^{2\prime} = \begin{pmatrix} -31.92 & -49.56 & 60.35 & 23.16 \end{pmatrix}; \qquad \hat{\alpha}^{2\prime} = \begin{pmatrix} 0.005 & -0.003 & -0.017 & 0.005 \end{pmatrix},$$

and the estimated $\hat{\Pi}_{MLE} = \hat{\alpha}^2 \hat{\beta}^{2\prime}$ is given by:

$$\hat{\Pi}_{MLE} = \begin{bmatrix} 0.005 \\ -0.003 \\ -0.017 \\ 0.005 \end{bmatrix} \begin{bmatrix} -31.92 & -49.56 & 60.35 & 23.16 \end{bmatrix} = \begin{bmatrix} \underset{(-1.47)}{-0.15} & \underset{(-1.47)}{-0.24} & \underset{(1.47)}{0.29} & \underset{(1.47)}{0.11} \\ \underset{(1.01)}{0.09} & \underset{(1.01)}{0.14} & \underset{(-1.01)}{-0.17} & \underset{(-1.01)}{-0.06} \\ \underset{(4.91)}{0.54^*} & \underset{(4.91)}{0.84^*} & \underset{(-4.91)}{-1.02^*} & \underset{(-4.91)}{-0.39^*} \\ \underset{(-2.26)}{-0.15^*} & \underset{(-2.26)}{-0.23^*} & \underset{(2.26)}{0.28^*} & \underset{(2.26)}{0.11^*} \end{bmatrix}, \tag{4.1}$$

where the *t*-statistics are listed in parenthesis and the significance of the parameters at the 5% level in $\hat{\Pi}_{MLE}$ is indicated by an asterisk (Johansen, 1995b).

According to the trace test, there are two cointegrating vectors ($\hat{r} = 2$) in the wheat data:

$$\hat{\beta}_1^{2\prime} = (-55.05 \;\; 20.48 \;\; 36.54 \;\; 26.32 \;\; -15.08); \qquad \hat{\alpha}_1^{2\prime} = (-0.011 \;\; -0.024 \;\; -0.010 \;\; -0.009 \;\; 0.008)$$
$$\hat{\beta}_2^{2\prime} = (-100.18 \;\; -29.71 \;\; 52.20 \;\; 130.70 \;\; 13.43); \qquad \hat{\alpha}_2^{2\prime} = (-0.010 \;\; -0.010 \;\; -0.002 \;\; -0.008 \;\; -0.016)$$

Thus, $\hat{\Pi}_{MLE} = (\hat{\alpha}_1^2 \quad \hat{\alpha}_1^2)\begin{pmatrix} \hat{\beta}_1^{2\prime} \\ \hat{\beta}_2^{2\prime} \end{pmatrix}$ is given by:

$$\hat{\Pi}_{MLE} = \begin{bmatrix} -0.011 & -0.010 \\ -0.024 & -0.010 \\ -0.010 & -0.002 \\ -0.009 & -0.008 \\ 0.008 & -0.016 \end{bmatrix} \begin{bmatrix} -55.05 & 20.48 & 36.54 & 26.32 & -15.08 \\ -100.18 & -29.71 & 52.20 & 130.70 & 13.43 \end{bmatrix}$$

$$
= \begin{bmatrix} 1.58* & 0.08 & -0.91* & -1.58* & 0.03 \\ (5.64) & (0.90) & (-5.80) & (-4.84) & (0.52) \\ 2.35* & -0.18 & -1.41* & -1.99* & 0.22* \\ (5.23) & (-1.26) & (-5.64) & (-3.78) & (2.77) \\ 0.79* & -0.15* & -0.49* & -0.55* & 0.13* \\ (4.63) & (-2.80) & (-5.19) & (-2.79) & (4.28) \\ 1.30* & 0.06 & -0.75* & -1.28* & 0.03 \\ (4.23) & (0.57) & (-4.36) & (-3.59) & (0.50) \\ 1.20* & 0.65* & -0.57* & -1.93* & -0.34* \\ (3.03) & (5.16) & (-2.55) & (-4.16) & (-4.83) \end{bmatrix}, \tag{4.2}
$$

where the *t*-statistics are listed in parenthesis and asterisk indicates statistical significance (5%).

4.3 STRUCTURAL HYPOTHESIS TESTS

The $\hat{\Pi}_{MLE}$ matrices for both rice and wheat model [equation (4.1) and (4.2)] show that some of the coefficients of $\hat{\beta}$ and $\hat{\alpha}$ matrices might not be statistically different from zero. If some coefficients of $\hat{\beta}$ and $\hat{\alpha}$ matrices are zero, it would affect both that stationary and non-stationary linear combinations of price data. For example, if the first of element of $\hat{\beta}^{2\prime}$ in the rice model is zero, it suggests that country 1's price does not enter the cointegrating vector and therefore, is not a part of the spatial price linkage. Moreover, within the $\hat{\beta}$ matrix some coefficients might be proportional to each other. For instance, if $\hat{\beta}_1^2$ and $\hat{\beta}_3^2$ are equal to 1 and –1 in the rice market, respectively, prices of countries 1 and 3 have a one-to-one relationship, i.e., perfect arbitrage. Geographical proximity (e.g., Thailand and Vietnam) may suggest the existence of such one-to-one relationship due to lower transport costs. Therefore, three types of hypotheses on β and α matrices are of interest.

4.3.1 Exclusion/Zero Restrictions

The first set of hypothesis on β, denoted as H_3, take the form $\beta_p^2 = 0$ for *p=1,…4* for rice and $\beta_{pr}^2 = 0$ for *p=1,…,5* and *r = 1, 2* for wheat (*r* is the number of cointegrating vectors). Here we are interested in identifying whether or not all prices series enter the long-run equilibrium. Non-excludability of the *p*th price series indicates its inclusion/relevance in the spatial price relationship. Since the wheat market contains 2 cointegrating vectors, the *p*th exclusion restriction is a joint test on the *p*th coefficient in both vectors. That is, $H_3 : \beta_{.5}^2 = 0$ hypothesizes that the fifth element of both cointegration vectors, $\hat{\beta}_1^{2'}$ and $\hat{\beta}_2^{2'}$ is zero.

The exclusion restriction on β correspond to the null hypothesis of $H_3 : \beta^2 = H\varphi$ where *H* is a known matrix of dimension $(p \times s)$ with rank *s* and φ is a $(s \times r)$ matrix of unknown parameters. Given $r \leq s \leq p$, the LR test is given by:

$$-2\ln(Q; H_3 \mid H_2) = T\sum_{i=1}^{\hat{\gamma}_0} \ln\left\{\frac{(1-\hat{\lambda}_i^3)}{(1-\lambda_i^2)}\right\} \tag{4.3}$$

where $\hat{\lambda}_i^3$ and $\hat{\lambda}_i^2$ are respectively, the *i*th eigenvalue calculated under H_3 and H_2 (Johansen, 1995) . The LR test statistic is distributed as $\chi^2_{r(p-s)}$, which equal $\chi^2_{1(4-3)}$ for rice and $\chi^2_{2(5-4)}$ for wheat. Table 2 shows that every null hypothesis of zero coefficient in the cointegrating vector(s) is rejected, indicating that each of the rice and wheat prices is relevant for the respective long-run relationships in the two markets (weak LOP).

Table 2—Exclusion Restrictions on ∃ in the Rice and Wheat Model

	Rice Model				Wheat Model		
	Null Hypothesis	LR/1			Null Hypothesis/2	LR/3	
US	$H_{31:}\ \exists_{11} = 0$	28.48	Reject	US	$H_{31:}\ \exists_{.1} = 0$	21.45	Reject
Thailand	$H_{31:}\ \exists_{21} = 0$	9.95	Reject	Argentina	$H_{31:}\ \exists_{.2} = 0$	17.01	Reject
Vietnam	$H_{31:}\ \exists_{31} = 0$	16.18	Reject	Canada	$H_{31:}\ \exists_{.3} = 0$	18.90	Reject
India	$H_{31:}\ \exists_{41} = 0$	30.64	Reject	Australia	$H_{31:}\ \exists_{.4} = 0$	18.08	Reject
				India	$H_{31:}\ \exists_{.5} = 0$	33.04	Reject

[1] $\chi^2_{(1)}$ at 0.05 level of significance is 3.84.

[2] Denotes column vectors corresponding to the respective element of the two co-integrating vectors.

[3] $\chi^2_{(2)}$ at 0.5 level of confidence is 5.99.

4.3.2 Proportionality Restrictions

The cointegration results in section IV.1 indicated that rice and wheat markets appear to be partially integrated in the sense that there exists less than $(p-1)$ linear and stable relationships among prices. This result, however, doesn't rule out the possibility that any two prices within a market may exhibit a one-to-one relationship, i.e., unit proportionality, which implies perfect arbitrage between the two countries. It is possible for instance, that geographical proximity between Thailand and Vietnam may lead to unit price proportionality between the two countries in the rice market. Some prior studies of wheat market integration have tested unit price proportionality between US and Canadian prices (Goodwin, 1992a, 1992b; Mohanty et al., 1996). Therefore, pair-wise proportionality is hypothesized for all possible price pairings and the results of the corresponding LR tests are presented in table 3.

The calculated LR test statistics for the rice and wheat model, respectively, reject the null hypothesis of proportionality at the 0.05 significance level. Although perfect arbitrage condition seems to be likely between Thai and Vietnamese rice price, it is not valid statistically. In an earlier study there was evidence of Thailand-US unit proportional movement prior to URAA, but it is not detected in our study (Taylor et al., 1996; Bierlen, Wailes, and Cramer 1998).

Unlike the rice market, the international price linkage in wheat market is described by two cointegrating vectors, which span the cointegration space. The exclusion and proportionality restrictions above are joint tests of the respective elements of both cointegrating vectors. Thus, in the wheat model, these tests assume that the two vectors contain the same information. It might be the case that each of these two cointegrating vectors represents a distinct relation in the cointegration space. For the wheat model, the exclusion and proportionality can be tested in each of the cointegrating vectors. Repeating all of the hypotheses in sections IV.3.1-2 for each of the cointegrating vectors, we arrived at:

$$H_5^1 = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ -1 & 0 & 0 \\ 0 & 0 & 0 \\ 0 & 0 & 1 \end{bmatrix}, \text{ and } H_5^2 = \begin{bmatrix} 1 & 0 \\ 0 & 0 \\ 0 & 1 \\ -1 & 0 \\ 0 & 0 \end{bmatrix},$$

where H_5^1 and H_5^2 are restriction matrices for the first and second wheat cointegrating vector, respectively. The matrix H_5^1 imposes (i) US-Canadian price proportionality and

(ii) zero restriction on Australian price. The matrix H_5^2 imposes (iii) zero restriction on Argentinean and Indian prices, and (iv) US-Australian price proportionality, while Canadian price is unrestricted. These linear restrictions are expressed in terms of $H_5 : \beta^5 = (H_5^1\varphi_5^1, H_5^2\varphi_5^2)$ where φ_5^1 and φ_5^2 are parameters to be estimated, corresponding to each cointegrating vector. The LR test statistic, distributed as χ^2 with 3 degree of freedom, is given by:

$$-2\ln\left(Q; H_5 \mid H_2\right) = 86.24 - 81.62 = 4.62\,,$$

which suggests failure to reject the null hypothesis (H_5^1 and H_5^2) at the 0.05 significance level.

Note that pairwise proportionality involving one price each from developed and developing countries is not found in rice and wheat markets. These results further confirm partial integration between developed and developing countries in rice and wheat markets and that strong LOP does not hold. Other studies have often cited informational inefficiencies in international commodity markets as a reason for partial integration. In our case, however, the significant changes in the policy regimes of developed and developing countries in the post-URAA era may be a reason for our results. Developed economies have switched from price support to income support and price floors, while developing countries have continued to use tariffs for protection. A fall in world price, triggered by any exporter, may then cause increased support payments in developed economies possibly adding to further price declines. However, developing countries may have raised tariff barriers within the bounds imposed by the URAA suggesting partial and weakening integration between its and developed countries' markets.

4.3.3 Weak Exogeneity Restrictions

The third hypothesis relates to linear restrictions on α, which take the form $H_4 : \alpha^2 = A\psi$, where A is a $(p \times m)$ restriction matrix, $r \le m \le p$, and ψ is a matrix of unknown parameters. The hypothesis here is that some of the coefficients in $\hat{\alpha}^{2\prime}$ in section IV.2 are equal to zero. In the context of wheat market, as noted in section IV.3.1, $H_4 : \alpha^2_{.1} = 0$ hypothesizes that the first element of both adjustment vectors, $\hat{\alpha}^{2\prime}_1$ and $\hat{\alpha}^{2\prime}_2$ is zero, which implies that the cointegrating relations do not enter the *i*th equation. For example, if $\hat{\alpha}^2_{.1}$ is statistically not significant from zero, it would indicate that country 1's price doesn't adjust to any shocks to other prices in the system. Essentially, this is a test of weak exogeneity of country 1's price, P_{1t}. For H_4, the LR test statistic takes the form:

$$-2\ln(Q; H_4 \mid H_2) = T\sum_{i=1}^{\hat{\gamma}_0} \ln\left\{\frac{(1-\hat{\lambda}_i^4)}{(1-\lambda_i^2)}\right\} \tag{4.4}$$

where $\hat{\lambda}_i^4$ and $\hat{\lambda}_i^2$ are respectively, the *i*th eigenvalue calculated under H_4 and H_2. The LR test statistic is distributed as $\chi^2_{1(4-3)}$ and $\chi^2_{2(5-4)}$ for the rice and wheat model, respectively.

Table 3—Proportionality Tests for all Possible Pairs of Rice and Wheat Prices

	Rice Model				Wheat Model		
	Null Hypothesis	LR/1			Null Hypothesis /2	LR/3	
US-Thailand	$H_{32:}\ \exists_{11}+\exists_{21} = 0$	21.23	Reject	US-Argentina	$H_{32:}\ \exists_{1}+\exists_{2} = 0$	17.5	Reject
US-Vietnam	$H_{32:}\ \exists_{11}+\exists_{31} = 0$	4.00	Reject	US-Canada	$H_{32:}\ \exists_{1}+\exists_{3} = 0$	13.13	Reject
US-India	$H_{32:}\ \exists_{11}+\exists_{41} = 0$	4.37	Reject	US-Australia	$H_{32:}\ \exists_{1}+\exists_{4} = 0$	25.37	Reject
Thailand-Vietnam	$H_{32:}\ \exists_{21}+\exists_{31} = 0$	10.79	Reject	US-India	$H_{32:}\ \exists_{1}+\exists_{5} = 0$	22.9	Reject
Thailand-India	$H_{32:}\ \exists_{21}+\exists_{41} = 0$	4.12	Reject	Argentina-Canada	$H_{32:}\ \exists_{2}+\exists_{3} = 0$	34.9	Reject
India-Vietnam	$H_{32:}\ \exists_{31}+\exists_{41} = 0$	20.18	Reject	Argentina-Australia	$H_{32:}\ \exists_{2}+\exists_{4} = 0$	21.2	Reject
				Argentina-India	$H_{32:}\ \exists_{2}+\exists_{5} = 0$	11.4	Reject
				Canada-Australia	$H_{32:}\ \exists_{3}+\exists_{4} = 0$	18.8	Reject
				Canada-India	$H_{32:}\ \exists_{3}+\exists_{5} = 0$	18.3	Reject
				Australia-India	$H_{32:}\ \exists_{4}+\exists_{5} = 0$	18.8	Reject

[1] $\chi^2_{(1)}$ at 0.05 level of significance is 3.84; [2] Denotes column vectors corresponding to the respective element of the two co-integrating vectors; [3] $\chi^2_{(2)}$ at 0.5 level of confidence is 5.99.

Table 4 shows that the null hypothesis of weak exogeneity cannot be rejected at the 0.05 significance level individually for Thailand and Indian prices in the rice model. However, p-value of the test statistic for US price is 0.03. As type I error is reduced to 0.01, the weak exogeneity hypothesis of the US price is not rejected. These results are also consistent with the *t*-statistics of the maximum likelihood estimator of $\hat{\Pi}_{MLE}$ in equation (4.1), where all coefficients of level variables in the US and Thailand equation are not significant. For the wheat model, the null hypotheses of weak exogeneity of all prices in both cointegration vectors are rejected.

Table 4—Weak Exogeneity Tests for the Rice and Wheat Model

	Rice Model				Wheat Model		
	Null Hypothesis	LR/1			Null Hypothesis/2	LR/3	
US	$H_{4}: \alpha_1 = 0$	4.65	Fail to Reject	US	$H_{4}: \alpha_{.1} = 0$	27.45	Reject
Thailand	$H_{4}: \alpha_2 = 0$	0.99	Fail to Reject	Argentina	$H_{4}: \alpha_{.2} = 0$	27.48	Reject
Vietnam	$H_{4}: \alpha_3 = 0$	18.66	Reject	Canada	$H_{4}: \alpha_{.3} = 0$	30.93	Reject
India	$H_{4}: \alpha_4 = 0$	1.77	Fail to Reject	Australia	$H_{4}: \alpha_{.4} = 0$	16.37	Reject
				India	$H_{4}: \alpha_{.5} = 0$	20.49	Reject

[1] $\chi^2_{(1)}$ at 0.05 level of significance is 3.84; [2] Denotes column vectors corresponding to the respective element of the two co-integrating vectors; [3] $\chi^2_{(2)}$ at 0.5 level of confidence is 5.99.

4.4 FINAL LONG-RUN EQUILIBRIUM/RELATIONSHIP

The structural hypotheses results are used to derive rice and wheat market's final long-run price relationships. The final estimated $\hat{\Pi}_{MLE}$ for the rice model becomes:

$$\hat{\Pi}_{MLE} = \hat{\alpha}\hat{\beta}' = \begin{bmatrix} 0 \\ 0 \\ 0.53^{*} \\ (4.90) \\ -0.21^{*} \\ (-2.05) \end{bmatrix} \begin{array}{cccc} P_{US} & P_{TH} & P_{VT} & P_{IN} \\ \multicolumn{4}{c}{\begin{bmatrix} 1.00^{*} & 1.61^{*} & -1.96^{*} & -0.71^{*} \end{bmatrix}} \end{array} = \begin{bmatrix} 0 & 0 & 0 & 0 \\ (n.a.) & (n.a.) & (n.a.) & (n.a.) \\ 0 & 0 & 0 & 0 \\ (n.a.) & (n.a.) & (n.a.) & (n.a.) \\ 0.53^{*} & 0.85^{*} & -1.04^{*} & -0.38^{*} \\ (6.60) & (6.60) & (-6.60) & (-6.60) \\ -0.21^{*} & -0.34^{*} & 0.41^{*} & 0.15^{*} \\ (-2.13) & (-2.13) & (2.13) & (2.13) \end{bmatrix}, \quad (4.5)$$

where $\hat{\beta}$ is normalized by the US price coefficient. A comparison with original estimated coefficients in *(4.1)* shows that all insignificant parameters have dropped out of $\hat{\Pi}_{MLE}$.

The cointegrating vector $\hat{\beta}$ from equation (4.5) defines the equilibrium error in the rice market for time *t* in terms of past values of individual prices as follows:

$$P_{us.t-1} + 1.61P_{th.t-1} - 1.96P_{vt.t-1} - 0.71P_{in.t-1} = \hat{\upsilon}_t \quad (4.6)$$

This linear combination yields an equilibrium error, $\hat{\upsilon}_t$, whose expected value is zero. The system is temporarily in disequilibrium whenever $|\hat{\upsilon}_t| > 0$, meaning arbitrage profits exist. In the long run, adjustment of prices in the system will force $\hat{\upsilon}_t$ to be zero, exhausting arbitrage profits and restoring the equilibrium. The magnitude of adjustment of individual country prices to restore equilibrium is given by the $\hat{\alpha}$ vector in equation (4.5). For example, the Indian adjustment coefficient of –0.21 means that Indian price will decrease 21 percent of the disequilibrium value, in response to any exogenous changes in prices of all four countries. Vietnam has the largest price adjustment coefficient ($\hat{\alpha}_{31} = 0.53$). Note that US and Thai prices do not respond to equilibrium error, which is consistent with price leadership established also in some earlier studies (Hellwinckel and Ugarte, 2003; Alaouze et. al, 1987; Bredahl and Green, 1983; Petzel and Monke, 1979; Barker and Herdt, 1985; Warr and Wollmer, 1997).

The final long-run equilibrium/relationship for wheat prices, $\hat{\Pi}_{MLE}$, can be written as:

$$\hat{\Pi}_{MLE} = \hat{\alpha}\hat{\beta}' \begin{bmatrix} 0.05 & 1.34^{*} \\ (0.58) & (4.76) \\ 0.33^{*} & 1.99^{*} \\ (2.24) & (4.51) \\ 0.25^{*} & 0.43^{*} \\ (4.50) & (2.56) \\ 0.03 & 1.14^{*} \\ (0.31) & (3.73) \\ -0.60^{*} & 1.50^{*} \\ (-4.50) & (3.71) \end{bmatrix} \begin{array}{c} \begin{array}{ccccc} P_{US} & P_{AR} & P_{CA} & P_{AU} & P_{IN} \end{array} \\ \begin{bmatrix} 1.00 & -0.63 & -1.00 & 0 & 0.56 \\ 1.00 & 0 & -0.50 & -1.00 & 0 \end{bmatrix} \end{array}$$

$$= \begin{bmatrix} 1.39^{*} & -0.03 & -0.72^{*} & -1.34^{*} & 0.03 \\ (5.59) & (-0.58) & (-5.79) & (-4.76) & (0.58) \\ 2.32^{*} & -0.21^{*} & -1.32^{*} & -1.99^{*} & 0.19^{*} \\ (5.93) & (-2.24) & (-6.75) & (-4.51) & (2.24) \\ 0.69^{*} & -0.16^{*} & -0.47^{*} & -0.43^{*} & 0.14^{*} \\ (4.57) & (-4.50) & (-6.24) & (-2.56) & (4.50) \\ 1.17^{*} & -0.02 & -0.60^{*} & -1.14^{*} & 0.02 \\ (4.32) & (-0.31) & (-4.43) & (-3.73) & (0.31) \\ 0.90^{*} & 0.38^{*} & -0.15^{*} & -1.50^{*} & -0.34^{*} \\ (2.50) & (4.50) & (-1.81) & (-3.71) & (-4.50) \end{bmatrix} \qquad (4.7)$$

Consider the cointegrating vectors, $\hat{\beta}_1$ and $\hat{\beta}_2$, for the wheat model. Two equilibrium errors are defined as:

$$\begin{array}{lllllll} P_{us.t-1} & -0.63P_{ar.t-1} & -P_{ca.t-1} & & +0.56P_{in.t-1} & = & \hat{\upsilon}_{1t} \\ P_{us.t-1} & & -0.50P_{ca.t-1} & -P_{au.t-1} & & = & \hat{\upsilon}_{2t} \end{array} \qquad (4.8)$$

In the case of more than one cointegrating vector, a representation of the co-movements can be very complicated, since equilibrium errors are not uniquely defined. However, the first equality in equation (4.8) can be thought of as describing the comovement of prices between developed and developing countries, while the second defines price comovement within developed countries. The first adjustment vector in equation (4.7) indicates that India has the highest adjustment parameter (-0.60), while it is Argentina for the second adjustment vector (1.99). This implies that any external shock that leads to an equilibrium error can significantly change Argentinean and Indian wheat prices.

5. IMPULSE RESPONSE ANALYSIS

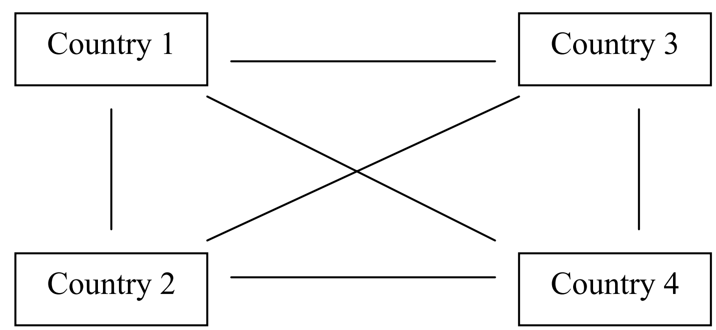
To further visualize the dynamic price relationships among the four (five) countries in the rice (wheat) market, we utilize an impulse response function based on the ECM in (2.1) but with the final, estimated long-run relationships in (4.5) and (4.7) (Bessler and Akelman, 1998; Yang, Bessler and Leatham, 2000). Here, the response of current and future values of one variable in the system (e.g., Vietnam rice price) to a one-standard deviation shock to one of the innovations (e.g., US rice price) are traced. The presence of contemporaneous correlation among the corresponding innovations can distort the calculation of impulse response functions because of the effects of innovations in another variable in the system at the same time.

Directed graphs, as given in Spirtes, Glymour and Scheines (1993), provide an algorithm for directing causal flow of information between countries or within a market. The algorithm begins with a complete undirected graph, where innovations from each country are connected with innovations in every other country (figure 3, panel A). The software TETRAD III (Spirtes et al., 2003), which contains the algorithm, removes edges when partial correlations are not statistically significant from zero and assigns causal flow directions for the remaining edges (figure 3, panel B).

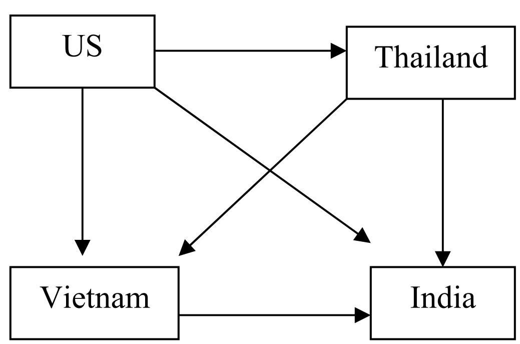
The directed graph results indicate that the causal structures of the four prices in the rice market are different from those of the five prices in the wheat market (figure 3, panel B and C). The results from the rice market suggest that the changes in the US rice price directly affect that of Thailand, Vietnam and Indian prices. Similarly, changes in Thailand (Vietnam) price directly affects prices of Vietnam and India (India). Changes in Indian rice price, however, do not directly affect other prices in the system. In the wheat market, again the changes in the US price have direct effects on all other prices followed by that of Argentina. Canadian and Australian prices affect Indian prices, whose changes, as before, do not affect other prices in the system.

Figure 3—Undirected and Directed Graph on Innovations

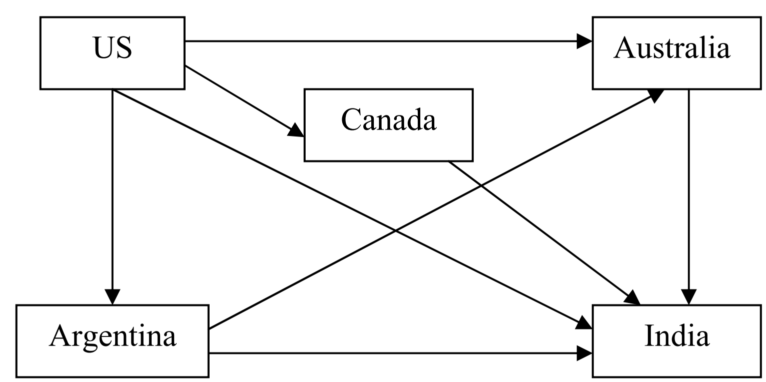
Panel A: An Example of Undirected Graph



Panel B: Directed Graph for Rice Market



Panel C: Directed Graph for Wheat Market



Figures 4 and 5 respectively present the impulse responses for the rice and wheat markets based on the ordering of innovations from directed graphs. They incorporated both short-run dynamics ($\hat{\Gamma}$ matrices in equation (2.1)) and long-run relationships (equations (4.5) and (4.7)). In the case of the rice market, changes in the US price bring about relatively large responses in Thailand, Vietnam and Indian prices. While Thailand and Vietnam prices increase immediately with the US price, the Indian price has a delayed response of up to 16 months. Changes in Thailand's price brings about an almost equal change in Vietnam and Indian price, but the response of US price is smaller consistent with weak exogeneity and directed graphs' causal structure. The Vietnam and Indian price changes have little effects on US and Thailand's prices. Although the effects of a shock from one country to other countries persisted in our 36-month responses, they eventually tended to zero.

In the case of the wheat market, the large effects of changes in developed countries' price on that of developing are clearly evident. The US price changes bring about larger responses in Argentinean and Indian price relative to those in Canadian and Australian prices. The Argentinean price change has relatively little effects on Canadian and Australian prices, while its effect on Indian price is slightly larger than that on the US price. Australian and Canadian price changes bring about larger responses in the US, Argentinean and Indian prices. The response to Indian price changes by the other four countries are less than their responses to developed countries price changes (e.g., Canada, Australia), but more than the Argentinean price effects. It appears that the Canadian price is the most exogenous in the sense it responds less to changes in other countries prices. In general, a developed country's price change has larger effects on other developed and developing countries. The changes in developing countries' prices bring about relatively lower responses in developed countries' prices (e.g., Argentinean price change and Canadian response and vice versa).

Figure 4—Impulse Responses for Rice Market

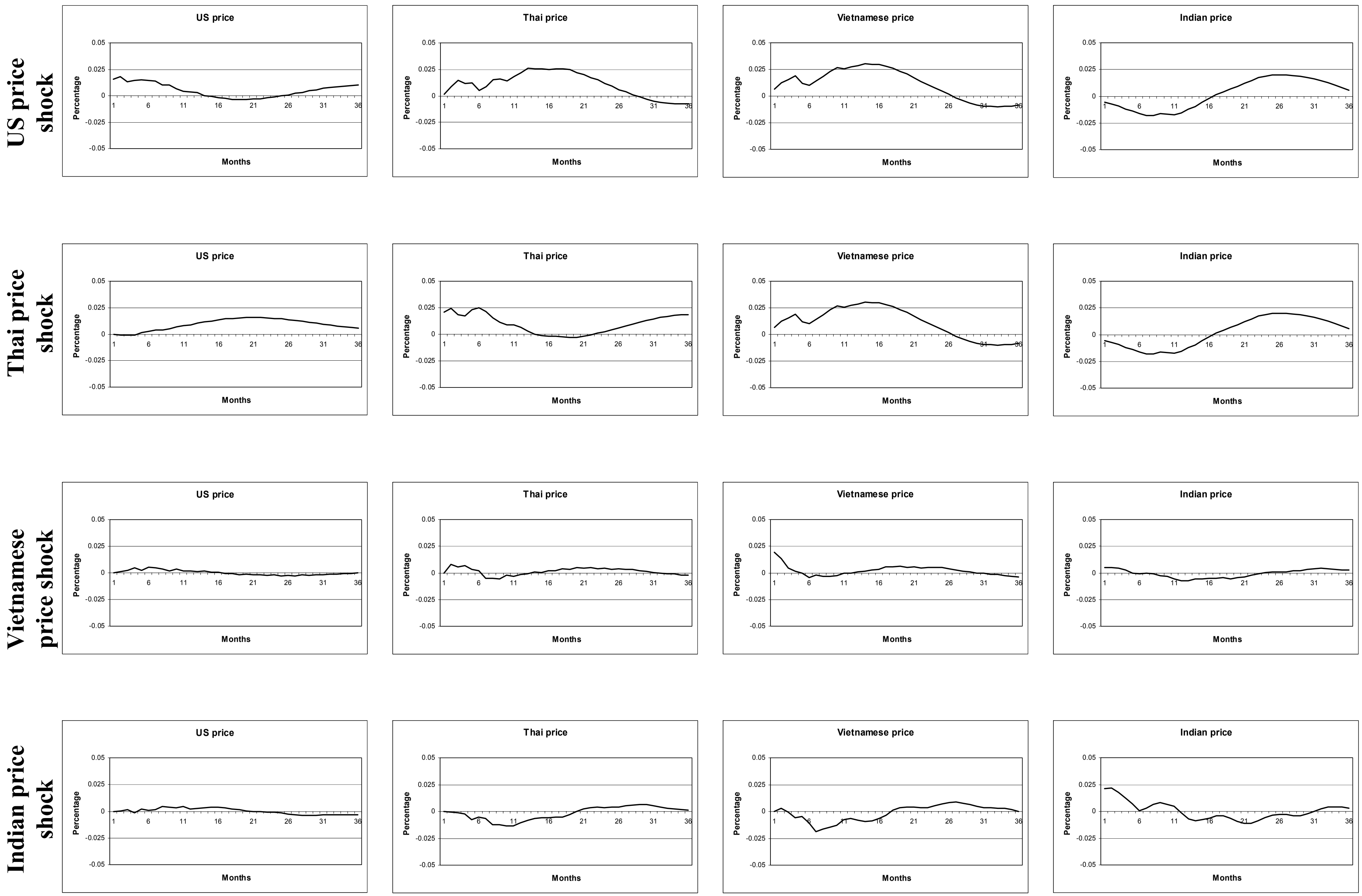
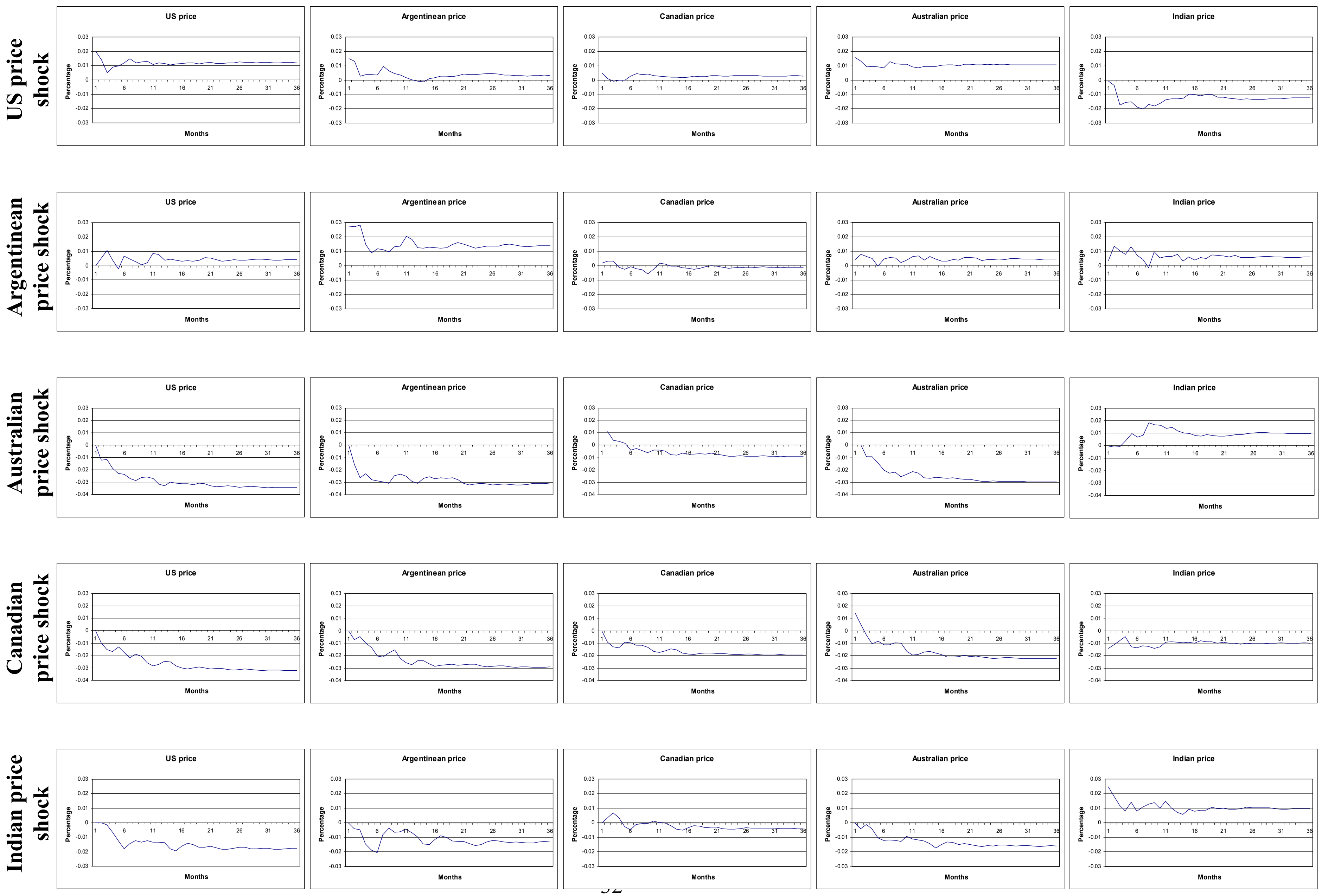


Figure 5—Impulse Responses for Wheat Market



6. SUMMARY AND CONCLUSIONS

This study analyzed spatial price linkages between developed and developing countries in two key commodity markets –long grain rice and hard wheat – during the post-Uruguay Round era using multivariate cointegration techniques. For the long grain rice market, prices of the United States, Thailand, Vietnam and India are chosen. The hard wheat market in this study included the United States, Canada, Australia Argentina and India.

Results suggest that the prices of developed and developing countries are cointegrated, i.e., they exhibit a stable, long-run relationship during the post-Uruguay Round period. We found two cointegrating vectors in the wheat market and both included all 5 prices. In the rice market, only one cointegrating vector is identified, which included prices from all 4 countries. Within the long-run relationship in rice prices, we did not find pairwise unit proportionality among the four countries. In the case of wheat, unit proportionality is found between developed countries prices, but not found for any pair involving one each of developed and developing countries. These results confirm partial integration between developed and developing countries in rice and wheat markets. The dramatic decline in commodity prices and the accompanied increase in protection and/or domestic support in several countries during late 1990s might be a reason for our results.

Additional tests suggest that the US and Thai rice prices are weakly exogenous in the sense that they do not adjust their respective prices to stay in the long-run relationship, i.e., price leadership. Instead, Vietnam and Indian prices adjust to restore the long-run relationship. An impulse response analysis aided by directed graphs on the causal flow of information among the four prices confirmed US and Thailand's price leadership. Changes in the prices of leaders, especially that of the US, have relatively large effects on Vietnam and Indian rice prices. For the wheat market, the adjustment parameters from the two cointegrating vectors suggest that all countries' prices adjust to any shock to the long-run relationship. However, in the first cointegrating vector, Argentinean price adjusts most to perturbations in the long-run price linkage, while the Indian price adjusts the most to any changes in price relationship illustrated by the second cointegrating vector. In the case of the wheat market, the relatively large effects of changes in developed countries' price on that of developing are clearly evident from the impulse response analysis.

To conclude, developed and developing country prices in rice and wheat markets are linked imperfectly, but in most instances the changes in the prices of the former lead to significant adjustments in those of the latter group. Developing countries, especially new entrants into world rice and wheat markets (Vietnam and Argentina, respectively) have faced significant price adjustment especially during the post-Uruguay Round period. Rapid and large adjustment of developing countries' prices to changes emanating from other countries would certainly affect price stability, income and welfare of their agricultural households. Risk-averse producers in developing countries would respond to the higher and fast adjustment (variability) of prices by reducing supply, which may erode their international export competitiveness. In some cases, the falling prices may force sustenance farmers to leave agriculture and aid in the consolidation of farm assets. The livelihood of the relatively low-skilled labor out of agriculture would then depend crucially on the capacity of non-farm sectors to productively employ these resources. If these countries excessively depend on rice or wheat for foreign exchange, external changes passed through the long-run price linkage with other markets could potentially impact their exchange rate and the non-farm sectors. A better understanding of the sources of price changes in developed countries, i.e., demand- or supply- (productivity-) or policy-based, is necessary since the Peace clause of the URAA expired earlier this year. There is the possibility that the changes in the developed countries' prices can be interpreted as policy impacts, especially by developing countries. Such interpretations would probably lead to more protection/support moving away from a multilateral commitment to reform agriculture and alleviate poverty.

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APPENDIX 1: DICKEY-FULLER AND AUGMENTED DICKEY-FULLER UNIT ROOT TEST RESULTS

The unit root tests suggest that the US, Thai, Vietnamese, and India rice prices are non-stationary in level (table A1, top panel). For wheat, the tests indicate that the US, Argentinean, Canadian, and Australian prices have unit roots in level. The result of the tests for Indian wheat price is inconclusive. That is, the ADF_{γ} indicates rejecting the hypothesis of non-stationarity in level at $\alpha = 0.05$, while the DF_t, DF_{γ}, and ADF_t statistics suggest the opposite. Since the ADF_{γ} test is considered a more restrictive test (Granger and Newbold, 1986), the conclusion here is that the Indian price is non-stationary in level as well.

First differences of prices are then tested for unit roots (table A1, bottom panel). The four test statistics indicate the first differenced data of rice and wheat prices are stationary; i.e., $I(0)$. Therefore, we conclude that every price series in our sample is I(1), which is a prerequisite for the multivariate cointegration analysis.*

Table A1—Dickey-Fuller Unit roots test in levels and first differences

Test Statistics	DF_t	$DF\gamma$	ADF_t	$ADF\gamma$	DF_t	$DF\gamma$	ADF_t	$ADF\gamma$
	Level				First difference			
Rice prices:/1								
U.S.	-1.60 (F)	-6.62 (F)	-1.47 (F)	-6.67 (F)	-7.82 (R)	-61.91 (R)	-3.61 (R)	-54.69 (R)
Thailand	-1.54 (F)	-5.09 (F)	-1.82 (F)	-6.17 (F)	-5.67 (R)	-44.69 (R)	-5.25 (R)	-61.32 (R)
Vietnam	-1.40 (F)	-4.31 (F)	-1.87 (F)	-6.07 (F)	-5.92 (R)	-46.52 (R)	-3.89 (R)	-45.64 (R)
India	-1.47 (F)	-4.10 (F)	-1.92 (F)	-5.46 (F)	-5.37 (R)	-41.21 (R)	-4.42 (R)	-48.75 (R)
Wheat price:/2								
U.S.	-1.69 (F)	-5.64 (F)	-1.76 (F)	-6.09 (F)	-8.97 (R)	-104.33 (R)	-5.29 (R)	-138.79 (R)
Argentina	-1.96 (F)	-7.71 (F)	-1.97 (F)	-8.54 (F)	-9.02 (R)	-105.26 (R)	-5.91 (R)	-161.30 (R)
Canada	-1.92 (F)	-7.37 (F)	-1.86 (F)	-7.71 (F)	-10.73 (R)	-129.37 (R)	-5.52 (R)	-161.59 (R)
Australia	-1.67 (F)	-5.50 (F)	-1.82 (F)	-6.21 (F)	-8.92 (R)	-103.32 (R)	-4.76 (R)	-125.66 (R)
India/3	-3.56 (F)/3	-17.58 (F)	-3.59 (F)/3	-30.74 (R)	-8.73 (R)	-73.83 (R)	-5.04 (R)	-104.99 (R)

[1]For rice data at sample size, T =50, critical values for DF_t and ADF_t are -3.50 and for DF_γ and ADF_γ are -19.7 at α=0.5.
[2]For wheat data at sample size, T =100, critical values for DF_t and ADF_t are -3.45 and for DF_γ and ADF_γ are -20.6 at α=0.5. (Fuller, 1976, p.373; 1996, Table 10.A.2)
[3]Conclusions based on critical value of DF_t and ADF_t for T=50 at α=0.025, which is -3.69.
[*]In fact, not all price series should be $I(1)$ as is often incorrectly assumed. Cointegration tests can be carried out so long as two of the variables in the system are $I(1)$. However, the trace test will not identify cointegration between series of different order.

APPENDIX 2: DETERMINISTIC COMPONENTS AND LAG LENGTH OF THE ECM

I. Lag Length of the ECM

The ECM is specified with six lags for the rice and wheat models. The choice on lag length is based on (i) Residual Analysis, (ii) Likelihood ratio test, and (iii) eigenvalues of companion matrices.

(i) The residual analysis is carried out using the Shenton-Bowman test for normality. Residuals from rice and wheat ECMs with six lags generated white noise residuals (table A2.1). The Lagrange Multiplier (LM) test suggests that the wheat model with seven lags yields serially correlated residuals.

Table A2.1—Rice and Wheat Multivariate Test Statistics for Residual Analysis

	Rice Model						Wheat Model					
Lag	**Test for Autocorrelation**				**Test for Normality**		**Test for Autocorrelation**				**Test for Normality**	
Length	LM(1)		LM(4)		Shenton-Bow man Stat		LM(1)		LM(4)		Shenton-Bow man Stat	
k=2	12.02	(0.74)	14.8	(0.54)	64.73	(0.00)	41.73	(0.02)	18.50	(0.82)	35.59	(0.00)
k=3	11.11	(0.80)	18.32	(0.31)	66.36	(0.00)	17.94	(0.84)	18.41	(0.82)	50.14	(0.00)
k=4	12.29	(0.72)	20.07	(0.22)	57.57	(0.00)	21.58	(0.66)	23.39	(0.55)	39.90	(0.00)
k=5	16.28	(0.43)	14.36	(0.57)	62.80	(0.00)	34.89	(0.09)	20.71	(0.71)	19.31	(0.04)
k=6	12.17	(0.73)	18.82	(0.28)	**21.31**	**(0.008)**	20.73	(0.71)	28.73	(0.28)	**11.66**	**(0.31)**
k=7	-	-	-	-	-	-	33.33	(0.09)	32.89	(0.09)	10.30	(0.41)

p-values are in parentheses

(ii) Whether or not the omission of a lag induces a significant loss of fit in the ECM is verified by a LR test. Table A2.2 shows that omitting the sixth lag, indeed, induces a significant loss of fit.

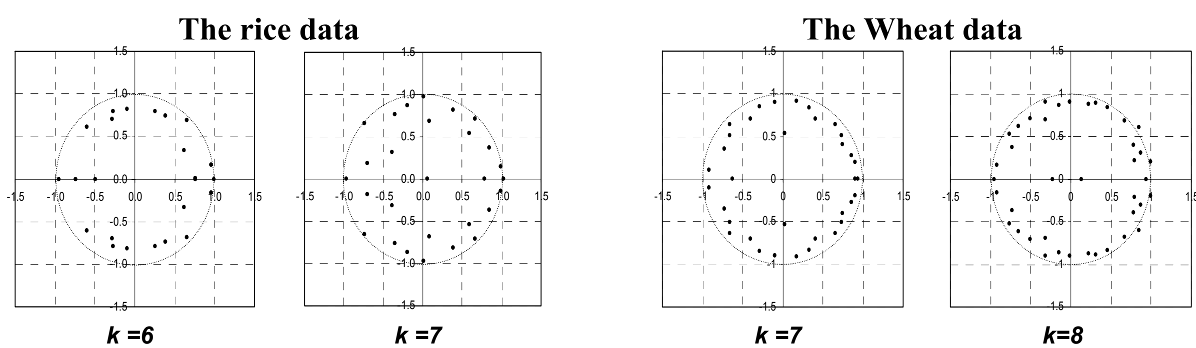
Table A2.2—Log-likelihood ratio for misspecification test

	rice models		wheat model	
Lag Length	*lnL(k)*	Ho:π6 =0	*lnL(k)*	Ho:π6 =0
k=6	596.96	ln[L(5)/L(6)] = 28.62 (R)	841.60	ln[L(5)/L(6)] = 76.15
k=5	582.65		803.53	

For rice models, χ^2_{16} at $\alpha = 0.05$ is 26.3 and for wheat models, χ^2_{25} at $\alpha = 0.05$ is 37.65

(iii) Whenever the eigenvalues of companion matrices are bounded by the unit circle, the corresponding ECM is said to be stable (Johansen, 1995). The model with six and seven lags respectively for the rice and wheat models satisfies this criterion (figure A2.1). However, the model with six lags is chosen for the wheat market to avoid serially correlated residuals (table A2.1)

Figure A2. 1—Eigenvalues of Companion Matrices



II. Deterministic components of the ECM

A presence of a linear trend in the VAR, which is equivalent to an intercept term in the ECM is confirmed using the LR test. Let the absence of a trend (intercept) in the VAR (ECM) be denoted by the hypothesis H_2^*. Testing $H_2^*(r \le 1)$ against $H_2(r \le 1)$ for the rice model yields:

$$-2\ln\left[Q;H_2^*(r=1)\,|\,H_2(r=1)\right]\equiv -T\sum_{i=2}^{4}\ln\left\{\frac{(1-\hat{\lambda}_i^*)}{(1-\hat{\lambda}_i)}\right\}=28.08-14.66=13.42\,.$$

For the wheat data, testing $H_2^*(r\leq 2)$ against $H_2(r\leq 2)$ yields:

$$-2\ln\left[Q;H_2^*(r=2)\,|\,H_2(r=2)\right]\equiv -T\sum_{i=3}^{5}\ln\left\{\frac{(1-\hat{\lambda}_i^*)}{(1-\hat{\lambda}_i)}\right\}=33-17.8=15.19\,.$$

Both tests reject the null hypothesis at $\alpha = 0.05$ ($\chi^2_{(3,0.95)} = 7.81$) suggesting the inclusion of an intercept term in the respective ECMs (case I, table 1).

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