Dynamic relationships between US and Thai rice prices [1]

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Abstract

Relationships between the United States Department of Agriculture's (USDA) estimated World Market Price, the Thai milled rice price, US transportation-adjusted cash rough rice price, and the Chicago Rice and Cotton Exchange rough rice futures price are examined for the 1987–1991 marketing years. Specifically, a cointegration analysis is used to address the pricing and informational efficiency of the respective markets. Testing indicates the system is described by two cointegrating vectors. The analysis preformed herein provides insight into the pricing performance of several world rice markets.

1. Introduction

Price discovery is a major issue and concern for agricultural producers. Information is critical for the creation and implementation of production, financial, and marketing strategies. For some commodities information is readily available from trading in competitive auction-type markets (e.g. wheat in the US). For rice, however, this is not the case. A unique marketing structure has evolved in the rice sector. World rice markets have been characterized by noncompetitive pricing (Karp and Perloff, 1989; Yumkella et al., 1994) and a significant degree of price leadership. Cooperatives, bid/acceptance sales offices, and forward sales contracts have evolved and dominate the US rice industry. These have created an information void, leading to a noted lack of price discovery mechanisms for rice producers.

An additional actor in the world rice market is US governmental policy and its pricing institutions. The Food Security Act of 1985 (USDA, 1985) developed and initiated the Rice Marketing Loan (RML) in response to previous government programs that were established in order to reduce rice price volatility and support prices. These programs were instrumental in maintaining the US rice price above the world price of rice, resulting in decreased US exports and increased US government stocks of rough rice. A key provision of the RML is weekly calculation and announcement of the World Market Price (WMP). The function of the WMP is to establish a value of rice based on current confirmed sales prices in the global market. Thus, the WMP is an attempt to provide information on this otherwise uncertain market.

Reintroduction of rough rice futures trading at the Chicago Rice and Cotton Exchange (CRCE) in August 1986 provided an additional price information signal. The CRCE rough rice futures market provides market participants with information regarding the local, national, and international rice markets, as

well as being a primary price discovery mechanism; however, it is not a heavily traded market.

Many within the rice industry have taken the view that the USDA is 'setting' rough rice prices when the WMP is announced (Rosera, 1993). This belief is brought about because, even though the WMP formulation procedure is an average price based on confirmed sales, many of the individual components and the sources are unknown to those outside the USDA. Questions and concerns arise regarding the USDA's role in announcing the WMP. Eugene Rosera, Agricultural Economist, ASCS-USDA, responded, "There isn't any authority in the law for the Secretary to undercut the world price or to be too high. It simply says that he will estimate the world price. There is a lot of judgement in that... if he consciously decided to undercut the world price, he could be faulted just as much for breaking the law as if he... set it too high" (Rice Journal, 1986). Further anxiety and misunderstanding of rice market behavior is compounded, as there tends to be a lack of knowledge among rice producers regarding the rough rice futures market, accompanied by a general distrust of futures trading in general (Pro Farmer, 1990).

The purpose of this article is to examine the statistical relationships that exist between several key rice price series. This study utilizes a time series framework to examine the relationship between the CRCE rough rice futures market price, a local cash rough rice price (the Texas transportation-adjusted cash price), the USDA World Market Price, and the Thai milled rice price. Specifically, the study focuses on the pricing performance and informational efficiency of the respective markets.

As price is an information signal whose purpose is to summarize all relevant information on supply and demand, we expect its time series representation to be close to a random walk (Samuelson, 1965). Accordingly, focus will be on testing for nonstationary behavior and modeling such if found. Recent innovations in time series analysis of such data appear to be useful for such investigations. Cointegration involves the analysis of series which individually appear to be nonstationary, but whose linear combination is, in fact, stationary. It has gained popularity because it allows researchers to address concerns related to short-run and long-run information flows between or among markets.

2. Time series considerations

Given two series, x_t and y_t, each of which is stationary in first differences, i.e. $I(1)$, a linear combination of these two series is generally also stationary in first differences, $I(1)$. If, however, there exists a constant A such that

$$Z_t = x_t - A y_t \tag{1}$$

where Z_t is $I(0)$, then the two series, x_t and y_t, are said to be cointegrated. Engle and Granger (1987) demonstrated that cointegrated data can be modeled as an error correction process. In the following section, we consider a procedure for testing for cointegration and error-correction representation.

2.1. Maximum likelihood estimation

Johansen (1988) and Johansen and Juselius (1990), developed a procedure for testing a multivariate system for cointegration. They derived maximum likelihood (ML) estimation techniques for estimating the parameters of the error correction model (ECM) for the cointegrated series. The ML process considered is defined from a p-dimensional sequence characterized as NID $(0, \Lambda)$. The process X_1 is defined as a vector autoregression (VAR) in p variables with k lags

$$X_t = \boldsymbol{\Pi}_1 X_{t-1} + \cdots + \boldsymbol{\Pi}_k X_{t-k} + \epsilon_t \tag{2}$$

For all $X \sim I(1)$, such that ΔX_t is stationary, Eq. (2) can be represented as

$$\Delta X_t = \sum_{i=1}^{K-1} \boldsymbol{\Gamma}_i \Delta X_{t-i} + \boldsymbol{\Pi} X_{t-1} + \mu + \epsilon_t$$

$$t = 1, \ldots, T \tag{3}$$

where

$$\boldsymbol{\Gamma}_i = -\sum_{j=i+1}^{K} \boldsymbol{\Pi}_j \text{ for } i = 1, \ldots, k-1 \tag{4}$$

and

$$\boldsymbol{\Pi} = -(I - \boldsymbol{\Pi}_1 - \ldots - \boldsymbol{\Pi}_k) \tag{5}$$

Here the $\boldsymbol{\Pi}_i$ are $(p \times p)$ matrices of autoregressive parameters from a VAR in levels of X_t of lag order k. The $\boldsymbol{\Gamma}_i$ are $(p \times p)$ parameter matrices summarizing short-run relationships among the

ΔX_i's, and Π is a $(p \times p)$ long-run parameter matrix (see Engle and Granger, 1991 for a discussion of long-run and short-run economic relationships and their time series representations). The vector μ is a $(p \times 1)$ constant, and ϵ_t is a $(p \times 1)$ white noise innovation term. Eq. (3) is labeled hypothesis H_1. Three cases are permissible:

1. rank(Π) = p, i.e. Π is full rank, indicates the process is stationary, and a VAR in levels (Eq. (2)) is appropriate;
2. rank(Π) = 0, i.e. Π is a null matrix containing no long-run information, indicates the model is a traditional differences times series model;
3. $0 < $ rank(Π) = $r < p$, which implies that there exists $(p \times r)$ matrices α and β such that $\Pi = \alpha \beta'$.

Case (3) is one of cointegration. Here, the Π matrix can be decomposed into the cointegration vector, β, and the adjustment coefficients, α. The matrix β reflects the long-run equilibrium which holds the elements of X together. The adjustment coefficients indicate the speed of adjustment toward long-run equilibrium.

Determination of the rank of Π involves its ordered eigenvalues, $(\lambda_1 > \ldots > \lambda_p)$, which can be derived as solutions to the equation

$$\left| \lambda S_{kk} - S_{k0} S_{00}^{-1} S_{0k} \right| = 0 \tag{6}$$

Here, the product moment matrices of residuals are defined as

$$S_{ij} = T^{-1} \sum_{t=1}^{T} R_{it} R'_{jt} \quad i,j = 0,k \tag{7}$$

The residuals R_{0t} and R_{kt} are determined by regressing ΔZ_t and Z_{t-k} on $\Delta Z_{t-1}, \ldots, \Delta Z_{t-k+1}$. The likelihood ratio test statistic of the null hypothesis (H_2)—that the rank of Π is less than or equal to r—is given as:

$$-2 \ln(Q) = -T \sum_{i=r+1}^{p} \ln\left(1 - \hat{\lambda}_i\right) \tag{8}$$

Testing the rank of Π using Eq. (8) requires one to make explicit the manner in which μ enters into the error correction process—either as a constant in the cointegrating vector or as a time trend in the original levels representation (Eq. (2)). Johansen (1992) considers this problem and suggests a sequential procedure of testing eigenvalues calculated with and without restrictions on the time trend. Testing begins with the rank equal to 0 and the constant restricted. Restrictions on the constant indicate no linear trend exists in the levels data. The order of hypotheses for testing is

$$H_2(0)^*, H_2(0), H_2(1)^*, H_2(1), \ldots, H_2(p)^*,$$
$$H_2(p) \tag{9}$$

where $H_2(k)$ indicates hypothesis H_2 with k cointegrating vectors, $k = 0, \ldots, p$. The asterisk denotes the restricted constant. The hypotheses are tested sequentially until one fails to reject the null hypothesis. Asymptotic critical values for the trace test are reported in Johansen and Juselius (1990, table A3 for the H_2^* hypotheses and table A1 for the H_2 hypotheses). Below we apply these tests to data on US rough rice and CRCE rough rice futures, Thai and the USDA's world rice prices.

2.2. Data

The data consist of the natural logarithms of four separate rice price series—the Texas cash rough rice price, the Chicago Rice and Cotton Exchange (CRCE) rough rice futures price, the USDA weekly-announced WMP, and the Thai milled rice price. The Texas cash rough rice price series (SP) used is a calculated price for US Number 2 rough rice with a milling yield of 55/70, FOB Houston. The price series is based on confirmed bid/acceptance sales of rough rice in Texas for the 1987–1991 marketing years. The observed cash sales prices were adjusted

Table 1
Summary statistics for the rough and milled rice prices, 1987–1991

	Obs.	Mean	Variance	Min.	Max.
Standardized cash price	260	2.10	0.02	1.54	2.55
CRCE futures price	260	2.07	0.02	1.57	2.55
World market price	260	1.81	0.02	1.35	2.20
Thai milled price	260	2.71	0.01	2.42	2.92

for transportation and standardized based on the implicit price schedule exhibited at these market locations for the time period in question and then averaged to obtain weekly observations.

Futures prices (*FP*) for the analysis were obtained from the CRCE rough rice futures market. The CRCE futures prices are for US Number 2 or better long grain rough rice with a par milling yield of 55/70 (Chicago Board of Trade, 1989). Weekly values were acquired by using the Wednesday closing price for the nearby contract month. Currently, the CRCE rough rice futures contract months are September, November, January, March, May, and July. Prior to the 1989 marketing year, the July contract was not traded at the CRCE.

The WMP, as mandated by the 1985 and subsequent farm bills, is announced weekly by the USDA (USDA, 1986; USDA, 1991a). The WMP, as set forth by legislation, is to be announced by the Secretary of Agriculture in order to determine the prevailing world price of rice to be used in calculation of loan repayment. The WMP series employed in this study represents US No. 2 or better long grain rough rice with a milling yield of 55/70. The WMP was obtained from USDA published reports (USDA, 1987–1993).

The fourth price series utilized is the Thai milled rice price (Thai). This price series was obtained from *Rice Market News*, Foreign Agriculture Report, The Hague (USDA, 1992). The price series quoted is for bulk milled Siam SWR 100% Grade B rice. The Siam 100% Grade B milled rice is comparable to US No. 2, 4% broken, long grain milled rice (Schnepf, 1993). The prices are quoted in US dollars per 100 lb.

Weekly observations were obtained for each of the respective price series for the time period extending from August 1987 through July 1992. This time period represents the 1987–1991 marketing years, providing 260 observations on each series. Missing observations were generated by assuming a random walk. Missing observations were present in only two of the four price series utilized, the Texas cash and the Thai milled price. There were 58 missing observations for the Texas cash rough rice price series and 42 for the Thai milled series. Missing observations for the Texas price were the result of a lack of bid/acceptance sales activity for the weeks in question.

tion. This occurred more often at the end of the marketing year when rice supplies were tight. Missing observations for the Thai milled price were the result of a lack of sales or price quotes by the *Rice Market News*. Summary statistics for the four price series are presented in Table 1.

Initial testing for nonstationarity was performed on both levels and first differences. Three procedures, the Dickey–Fuller (DF; Dickey and Fuller, 1979), the augmented Dickey–Fuller (ADF), and the Durbin–Watson (DW) tests were employed in testing the hypothesis of nonstationarity. These results

Table 2
Dickey–Fuller (DF), augmented Dickey–Fuller (ADF), and Durbin–Watson (DW) tests for stationarity on levels and first differences

Levels	DF [a]	ADF [b] (I^*)	DW [c]
Standardized cash price	− 2.46	− 2.46 (0)	0.17
Futures price	− 2.04	− 2.34 (1)	0.08
World market price	− 1.49	− 1.95 (2)	0.03
Thai milled price	− 3.01	− 3.01 (0)	0.12
First differences	DF	ADF (I^*)	DW
Standardized cash price	− 11.96	− 11.96 (0)	2.22
Futures price	− 7.01	− 7.01 (0)	1.66
World market price	− 6.34	− 4.41 (1)	1.38
Thai milled price	− 12.25	− 12.25 (0)	1.91

[a] Dickey–Fuller test of the null hypothesis that the series is generated as a random walk. The test statistics in the table refer to a *t*-test on the coefficient associated with lagged levels in a regression of the differenced data on lagged levels. The hypothesis is rejected for test statistics which are below − 3.4.

[b] The augmented Dickey–Fuller test is of the same form as the DF test. Lags of the dependent variable are on the right-hand side of the test, as well as one lag of levels, in a regression equation. The ADF tests were performed with I^* lags of the dependent variable. I^* was determined by application of an SC search to successive regressions (Hsiao, 1979). I^* represents the minimum SC for each regression. The Schwarz loss function used is defined as $(\log(e'e) + K(\log(T))/T$, where $e'e$ is the residual sum of squares for a model with lag length (I^*), T is the number of observations, and K is the number of regressors. The SC was used instead of the FPE as the SC tends to select univariate and multivariate models of shorter lags (Judge et al., 1988). An approximate 5% critical value is − 2.89. The decision rule is to reject the null hypothesis that the series follows a random walk for values of the test statistic less than the critical value.

[c] The Durbin–Watson test on a regression of the levels (or first differences) on a constant. The null hypothesis is that the series follows a random walk. The approximate 5% critical value is given in Sargan and Bhargava (1983) as 0.259. The decision rule is to reject the null hypothesis for DW statistics greater than this value.

are summarized in Table 2. All tests (DF, ADF, and DW) suggest that the standardized cash price, futures price, and WMP are nonstationary in levels. The test for the Thai milled price was not as conclusive. The DF test indicates that, at the 5% level, one can reject the hypothesis of nonstationarity in levels. The ADF and DW tests indicate the Thai milled price is nonstationary in levels. Whereas the ADF test is considered a more stringent test (Granger and Newbold, 1986), we tentatively conclude that the Thai milled price is nonstationary in levels.

As the price series were all nonstationary in levels, the first differences were then tested. The DF, ADF, and DW tests indicate that the first differences of the respective prices' series are stationary.Estimated autocorrelations and partial autocorrelations on levels and first differences were examined for evidence of seasonality (Box and Jenkins, 1976). None was detected. To save space these estimates are not reported but are available from the authors.

3. Tests for cointegration and error correction

Results of the trace tests for alternative cointegration specifications are presented in Table 3. Hypothesis testing at the 95% level indicates that r, the rank of Π, is 2. Additional testing indicated no linear trend exists.

3.1. Error correction specification

The error correction model (ECM) for the respective price series—the standardized Texas cash rough rice price, the WMP, the CRCE rough rice futures price, and the Thai milled rice price—was estimated using *CATS in RATS* (Juselius, 1991). The model was estimated with a restricted constant (no linear trend), and a lag structure of two periods. Specifically, the estimated ECM is given as

$$
\begin{bmatrix} \Delta SP_t \\ \Delta WMP_t \\ \Delta FP_t \\ \Delta THAI_t \end{bmatrix}
=
\begin{bmatrix}
-0.10 & 0.29 & 0.10 & 0.13 \\
(1.5) & (1.4) & (1.0) & (0.8) \\
0.03 & 0.11 & 0.03 & 0.20 \\
(1.2) & (1.6) & (0.8) & (4.0) \\
0.00 & 0.20 & 0.02 & 0.50 \\
(0.1) & (1.5) & (0.2) & (4.7) \\
0.01 & 0.17 & -0.01 & 0.01 \\
(0.4) & (2.0) & (0.2) & (0.2)
\end{bmatrix}
\begin{bmatrix} \Delta SP_{t-1} \\ \Delta WMP_{t-1} \\ \Delta FP_{t-1} \\ \Delta THAI_{t-1} \end{bmatrix}
+
\begin{bmatrix}
0.04 & -0.05 & 0.05 & 0.08 \\
(0.6) & (0.3) & (0.5) & (0.5) \\
0.01 & 0.16 & 0.01 & 0.18 \\
(0.6) & (2.5) & (0.3) & (3.4) \\
0.06 & 0.01 & 0.02 & 0.16 \\
(1.4) & (0.0) & (0.3) & (1.5) \\
0.01 & 0.04 & 0.01 & 0.04 \\
(0.4) & (0.5) & (0.3) & (0.6)
\end{bmatrix}
$$

$$
\times
\begin{bmatrix} \Delta SP_{t-2} \\ \Delta WMP_{t-2} \\ \Delta FP_{t-2} \\ \Delta THAI_{t-2} \end{bmatrix}
+
\begin{bmatrix}
0.17 & 0.03 & 0.17 & -0.05 & 0.21 \\
(-3.9) & (-4.1) & (3.5) & (-1.6) & (1.8) \\
-0.03 & -0.01 & 0.02 & -0.02 & 0.08 \\
(-1.8) & (-1.7) & (1.3) & (2.3) & (2.1) \\
0.05 & 0.01 & -0.08 & -0.08 & 0.27 \\
(1.5) & (1.0) & (-2.3) & (-3.3) & (3.2) \\
-0.02 & -0.01 & 0.01 & -0.06 & 0.22 \\
(1.2) & (2.0) & (0.3) & (-4.6) & (4.5)
\end{bmatrix}
\begin{bmatrix} SP_{t-1} \\ WMP_{t-1} \\ FP_{t-1} \\ THAI_{t-1} \\ 1.000 \end{bmatrix}
\qquad (10)
$$

The *t*-statistics are listed in parentheses under the appropriate elements of the Γ and Π matrices. Box–Pierce Q statistics are presented in Table 4. All are well below the 5% critical values (for 61 degrees of freedom), indicating autocorrelation among the residuals is not a problem.

3.2. Testing hypotheses on α and β

By cointegration, the Π matrix of Eq. (10) can be factored as $\Pi = \alpha\beta'$. With regard to the matrices, α and β, two hypotheses are of interest. The first hypothesis, denoted as H_3, concerns whether or not

Table 3
Trace test on alternative cointegration specifications, with and without a linear trend

$p - r$ [a]	r	$H_2(r)^*$—No linear trend			$H_2(r)$—Linear trend		
		Trace*	$C^*(5\%)$ [b]	Decision	Trace	$C(5\%)$ [c]	Decision
4	0	66.51	53.347	Reject	66.04	47.181	Reject
3	1	37.30	35.068	Reject	37.20	29.509	Reject
2	2	13.78	20.168	Fail [d]	13.69	15.197	Fail
1	3	6.12	9.094	Fail	6.06	3.962	Fail

[a] p equals the number of series, and r equals the number of cointegrating vectors.
[b] Taken from Johansen and Juselius (1990, table A3).
[c] Taken from Johansen and Juselius (1990, table A1).
[d] First failure to reject the null hypothesis, given the testing order given in Johansen (1992).

$\beta_{ij} = 0$, for $i = 1, \ldots, 4$. Here we are interested in whether all four price series enter the long-run equilibrium. (The notation used here refers to standard row–column labeling of the elements of the cointegrating vector. From our tests reported in Table 3, we have two cointegrating vectors: $\beta_1' = (\beta_{11}, \beta_{21}, \beta_{31}, \beta_{41}, \beta_{51})$ and $\beta_2' = (\beta_{12}, \beta_{22}, \beta_{32}, \beta_{42}, \beta_{52})$, where the rows are in order SP, WMP, FP, and THAI, and β_{51} and β_{52} are constants. Our test, for example that $\beta_{2j} = 0$, is a test that the second element of both vectors is zero; i.e. the second series (WMP) does not enter either cointegrating vector.) We test the hypothesis

$$H_3: \beta = H\varphi \qquad (11)$$

where H is a known design matrix of dimension $(p \times s)$ with rank s, and φ $(s \times r)$ is a matrix of unknown parameters. It is assumed that $r \leq s \leq p$. Johansen (1988) and Johansen and Juselius (1990)

demonstrated that the appropriate likelihood ratio test is

$$-2\ln(Q; H_3 | H_2) = T \sum_{i=1}^{r} \ln\left[(1 - \hat{\lambda}_{3i})/(1 - \hat{\lambda}_i)\right] \qquad (12)$$

where $\hat{\lambda}_{3i}$ is the ith eigenvalue calculated under the restricted hypothesis H_3, and $\hat{\lambda}_i$ is the ith eigenvalue calculated under H_2. The test statistic is asymptotically distributed as χ^2 with $r(p - s)$ degrees of freedom. The appropriate decision rule is to reject the null hypothesis (H_3) if the likelihood ratio statistic exceeds the critical chi-square value.

The tests of hypotheses $H_3: \beta = H\varphi$ are summarized in Table 5. Of the four tests conducted on β, the null hypothesis was rejected in three of the four cases, β_{1j}, β_{3j}, and β_{4j}; the null hypothesis $\beta_{2j} = 0$

Table 4
Box–Pierce Q test for autocorrelation among the cointegration residuals

	B-P Q Statistic [a]	Critical $\chi^2_{0.95,60}$
Standardized cash price	37.33	79.08
World market price	54.25	79.08
CRCE futures price	45.34	79.08
Thai milled price	40.53	79.08

[a] $Q = n \sum_{k=1}^{K} r_k^2(\hat{a})$ is distributed chi-squared with $K = 60$ degrees of freedom under the null hypothesis that the residuals from the error correction model are white noise. Here, $n = 260$, $r_k(\hat{a})$ the autocorrelation of residuals at lag k.

Table 5
Test of hypotheses $H_3: \beta = H\varphi$ and $\alpha = A\psi$

Hypothesis	χ^2	d.f.	Result	Significance
Test of hypothesis $H_3: \beta = H\varphi$				
$H_3: \beta_{1j} = 0$	18.92	2	Reject	0.00
$H_3: \beta_{2j} = 0$	0.89	2	Fail [a]	0.64
$H_3: \beta_{3j} = 0$	16.12	2	Reject	0.00
$H_3: \beta_{4j} = 0$	5.23	2	Reject	0.07
Test of hypothesis $H_3: \alpha = A\psi$				
$H_3: \alpha_{1j} = 0$	11.73	2	Reject	0.00
$H_3: \alpha_{2j} = 0$	5.11	2	Reject	0.08
$H_3: \alpha_{3j} = 0$	9.97	2	Reject	0.01
$H_3: \alpha_{4j} = 0$	14.21	2	Reject	0.00

[a] Fail to reject the null hypothesis, $H_3: \beta_{2j} = 0$, at the 0.10 level of significance.

was not rejected. The latter result indicates the WMP does not enter into the long-run equilibrium, whereas the other price series—standardized cash, CRCE futures, and Thai milled—do enter the long-run equilibrium for the system. The WMP is not conveying any new information to the long-run equilibrium. All pertinent information is contained and conveyed by the Texas cash, futures, and Thai milled prices.

Next, linear restrictions on α of the form

$$H_3: \alpha = A\psi \qquad (13)$$

are tested, where A is a $(p \times m)$ matrix. Interpretation of this test is that some of the rows of α are zero, i.e. $\alpha_i = 0$. If the ith row of α equals 0, this implies that cointegration relations do not enter the ith equation. This is a test of weak exogeneity of X_{it}. Johansen (1988), Johansen and Juselius (1990), Johansen (1991), and Juselius (1991) demonstrated

that the appropriate likelihood test statistic for this hypothesis test is of the form:

$$-2\ln Q = T \sum_{i=1}^{r} \left[\ln(1 - \lambda_i^*) / (1 - \hat{\lambda}_i) \right] \qquad (14)$$

where λ_i^* are the eigenvectors of the restricted model, and $\hat{\lambda}_i$ are the eigenvectors from the unrestricted model. This test statistic is asymptotically distributed as χ^2 with $r(p-m)$ degrees of freedom. Tests of the linear restrictions on α are also summarized in Table 5.

Testing the linear restrictions on α, i.e. $\alpha_{ij} = 0$, indicates that successive rows of the α matrix, α_{1j}, α_{2j}, α_{3j}, and α_{4j}, are not equal to zero (at the 0.08 level and lower). This indicates that the cointegration relationships do enter the equations for each series.

Our representation of the long-run (or equilibrium) information flows present in the four series is summarized by Eqs. (15)–(17)

$$[\alpha] = \begin{bmatrix} -0.12 & -0.05 \\ -0.02 & 0.01 \\ -0.02 & 0.08 \\ -0.05 & 0.03 \end{bmatrix} \qquad (15)$$

$$[\beta] = \begin{bmatrix} 1.00 & 1.00 \\ 0.00 & 0.00 \\ -0.83 & -1.32 \\ 1.00 & -0.66 \\ -3.10 & 2.42 \end{bmatrix} \qquad (16)$$

$$\Pi = \begin{bmatrix} -0.17 & 0.00 & 0.16 & -0.09 & 0.25 \\ (-3.7) & (0.0) & (3.4) & (-2.3) & (2.0) \\ -0.02 & 0.00 & 0.01 & -0.03 & 0.08 \\ (1.4) & (0.0) & (0.9) & (2.0) & (2.0) \\ 0.05 & 0.00 & -0.08 & -0.07 & 0.25 \\ (1.7) & (0.0) & (-2.5) & (-2.8) & (3.0) \\ -0.03 & 0.00 & 0.01 & -0.07 & 0.23 \\ (-1.4) & (0.0) & (0.4) & (-4.5) & (4.5) \end{bmatrix} \qquad (17)$$

As before, the t-statistics are listed in parentheses under the appropriate elements of the Π matrix.

As an administered price calculated by the USDA for US rough rice policy considerations, the informational content of the WMP is limited. All of the

information used in the calculation of the WMP, and depicted by the WMP, has previously been disclosed and disseminated by participants of the cash and futures markets for rice. The WMP conveys no unique information for understanding long-run rela-

tionships among our four price series. In terms of the long-run equilibrium among these series, WMP appears to be redundant.

3.3. Geometric presentation and interpretation of the cointegration relationship

Owing to the fact that the WMP is weakly exogenous in this system, the cointegration relationship evaluated in this paper can be presented in three dimensions—depicted by the standardized Texas rough rice prices (SP), CRCE rough rice futures price (FP), and the Thai milled rice price (THAI). However, the explicit cointegrating vectors are not unique, as only the space spanned by the vectors is unique; any alternative normalization is equally legitimate as the one illustrated in the section. The two cointegrating vectors are

$$Z_{1t} = SP - 3.10 - 0.83FP + 1.00\text{THAI} \qquad (18)$$

$$Z_{2t} = SP + 2.42 - 1.32FP - 0.66\text{THAI} \qquad (19)$$

Here Z_{1t} and Z_{2t} are introduced as perturbations in the respective cointegration vectors. Z_{1t} and Z_{2t} represent stationary deviations from the two long-run relationships. The cointegrating vectors (Eqs. (18) and (19)) are illustrated in Fig. 1.

The first cointegrating vector, Eq. (18), shows a positive relationship between the standardized Texas rough rice price and the CRCE rough rice futures price. In Fig. 1, the representation of this vector extends from near the origin of the X–Y axis out-

ward in X–Y space at an approximate 45° angle. This vector appears to underscore the strong arbitrage potential existing between the Texas cash and the CRCE futures markets for rough rice.

The second cointegrating vector, Eq. (19), is perhaps more interesting. This vector is represented in Fig. 1 as the plane extending downward in X–Z space from left to right. The most notable feature is the inverse relationship between the standardized cash price and the Thai milled price. Plausible explanations for this phenomena stem from the market structure represented in the two respective markets. First, the two markets do not represent a homogeneous product. The Texas price is for rough rice, and a milled series is reported for Thailand. Secondly, due to the oligopsonistic nature of the Thailand marketing system (E.J. Wailes, personal communication, 1993), there is potential for firms in Thailand to exercise market power in the pricing of milled rice. Finally, barriers to trade and agricultural policies of the two respective countries may also contribute to the inverse relationship.

The Royal Thai Government supports its rice industry through a series of government programs consisting of price supports, input subsidies, infrastructure development, and export credit programs (Schnepf, 1993). Similar government programs exist in the US. The basic premise of the farm and food policy of the Royal Thai Government is "to maximize agricultural export earnings while at the same time provide low-priced food for the Thai people" (USDA, 1989). The major instrument employed by
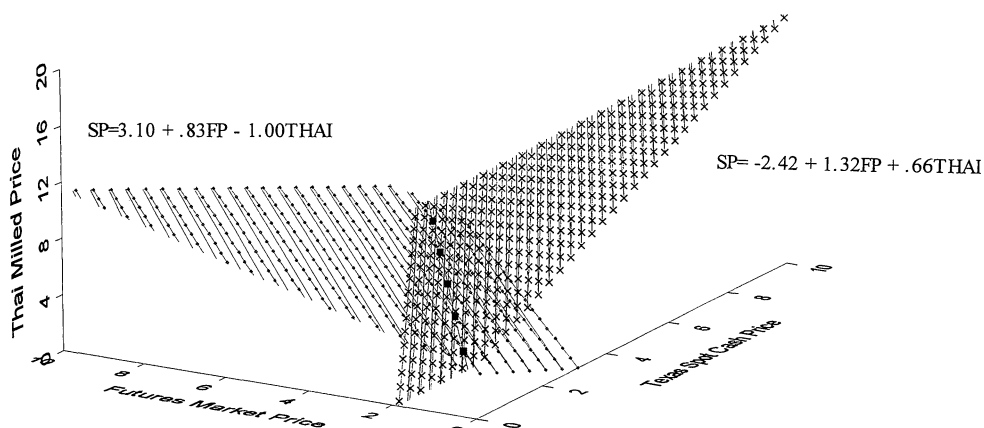

Fig. 1. Graphic illustration of cointegrating vectors (normalized on spot price).

the Thai government has been purchases of rice in the domestic market (Schwartz, 1985; Schnepf, 1993). This strategy of purchasing domestic rice to support internal rice prices has had a substantial impact on the relationship between US and Thai rice prices. Traditionally, the US has been an exporter of high-quality rice. In addition to high-quality rice, the US has been noted for reliability and marketing efficiency. This has allowed US rice to sell at a premium to similar quality Thai rice. Vast purchases of paddy rice by the Royal Thai government, in an effort to support domestic rice prices, have resulted in a diminished premium between Thai SWR 100% Grade B and US No. 2, 4%, long grained milled rice (Schnepf, 1993). Lately, however, there has been some indication the Thai government may be trying to gain market share in the high-quality rice market, as they face increased competition from Vietnam in the intermediate- and low-quality milled rice markets previously dominated by Thailand (Giordano and Raney, 1993; Schnepf, 1993). Direct competition between the US and Thailand should increase as Thai exporters target the higher-quality markets.

The final reason contributing to the breakdown of the traditional arbitrage condition are the barriers to trade existing between the US and Thai markets. Currently, there are no trade barriers prohibiting Thai rice from being imported into the US. Although there are no import barriers, the majority of current Thai imports are of milled aromatic rices for specialty markets (Cramer et al., 1990). Free trade does not exist in Thailand. The Thai government has erected barriers to both the importation and exportation of rice (Schnepf, 1993; Giordano and Raney, 1993; USDA, 1989; USDA, 1991b). Exportation of rough, or paddy, rice is banned by the Royal Thai Government (Giordano and Raney, 1993), as is the importation of rice (Schnepf, 1993; USDA, 1989; USDA, 1991b). Although the US and Thailand are the dominant exporting countries in the world, barriers to trade between the two countries inhibit the arbitrage condition between the US rough rice and Thai milled rice markets.

Given product differences, the oligopsonistic export market in Thailand, and extensive Government involvement, the relationship between Texas cash rough rice prices and Thai milled rice prices becomes convoluted. Furthermore, the arbitrage condi-

tion is obscured as the Royal Thai Government buys and sells rice in an effort to achieve its own strategic goals.

The intersection of the two planes, the bold solid line in Fig. 1, represents the long-run equilibrium. Although the prices can vary in each of the respective planes, they will be drawn to, and tend to converge along, the intersection of these planes. Thus, in equilibrium, the respective price series will be concentrated along this line.

4. Conclusions and implications

The analysis performed herein provides information and insight into the pricing performance of the US and global rice markets. The analysis examined the short- and long-run relationships between Texas (55/70) rough rice price, CRCE rough rice futures price, the USDA's announced WMP, and the Thai milled rice price during the 1987–1991 marketing years. These are modeled as an error correction process which captures cointegration properties of the data.

The first major finding is that the WMP does not enter the long-run equilibrium. This indicates the WMP is not conveying any new information into the rice markets, but merely reflecting information contained in the other price series studied. Secondly, the cointegration vectors from the error correction model provide details on the relationship between the three price series that do enter the long-run equilibrium— the Texas rough rice, CRCE rough rice futures, and the Thai milled prices. The cointegration vectors indicate there exists a positive (arbitrage) relation between the cash and futures markets for rough rice. This finding highlights the price discovery capability of the CRCE rough rice futures market. In addition to price discovery and informational aspects of the market, the findings suggest there exists the potential for risk reduction for producers in utilizing the futures market. The cointegration space indicates an inverse long-run relationship between the Texas and Thai markets. Plausible reasons for this unexpected condition stem from price and quality differentials, the potential for the firms in Thailand to exercise market power in pricing milled rice, and finally, the

existence of barriers to trade and government price supports.

The implications of this study are twofold. First, the key findings of the study indicate the long-run equilibrium in the international rice market is influenced by the Thai, Texas, and futures markets. From the ECM, the long-run price level for the three respective markets can be ascertained by using only one of the three price series. This is a key finding in the study of both price determination and informational efficiency. Secondly, the findings indicate the WMP does not provide any valuable new information to the long-run equilibrium. Current operating procedures within Texas bid/acceptance markets entail buyers submitting bids relative to the lower of the loan or the WMP for each individual lot of rough rice. Considerable attention is being paid to movement of the WMP from week to week. The aforementioned results indicate such movement is due to information and/or market conditions previously disclosed in the cash, futures, or Thai markets. These findings suggest that the WMP is summarizing market information and not making or influencing the fundamental long-run equilibrium in these rice markets.

Acknowledgements

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