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THE INTEGRATION OF THE INTERNATIONAL RICE MARKET

Modeling the international rice market has often appeared to be an exercise in frustration. The rice trade is small—less than 5 percent of world production regularly enters the trade—and is dominated by institutional arrangements that include concessional sales, government-to-government contracts, and other forms of market intervention. The picture is additionally confounded by the variety of grades of rice that are traded regularly. The lack of free competitive behavior and clear linkages in this market might appear to doom the researcher to the exclusive examination of individual market participants and leave the treatment of the international market at the level of an accounting exercise. The purpose of this paper is to suggest that the constraints on the researcher are not so severe as they might at first blush appear.

This essay assesses the role of price in the allocation of the quantity and qualities of rice which enter the world market in any given year. A delineation of this role facilitates the choice of an appropriate price for studies of world market trends (Thai 5 percent brokens have traditionally filled this role), for benefit-cost analysis of domestic agricultural investments, or for the determination of the amount of subsidization or taxation necessary to achieve domestic production and consumption goals.[1]

Markets may be integrated along some dimensions but not others. Geographical links may be stable at one point in time, but break down in others. Markets may react differently to variation in quality or variety ("a rose, is a rose, is a rose" may not hold for rice). This paper examines the issue from three directions. First, geographic links between classes of exporters and importers are considered. Next, an international analysis, with an examination of causality, is performed on time series of Thai and United States prices. Finally, a characteristic, or hedonic, approach is taken with respect to different varieties and qualities of rice.

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[1] Examples of the importance of the world price variable in these areas are provided in R.W. Herdt, A. Te, and R. Barker (1977), E.A. Monke, S.R. Pearson, and N. Akrasanee (1976), and Pearson, C.P. Humphreys, and Monke (forthcoming).

MARKET INTEGRATION AND PRICE ANALYSIS

The analysis of commodity prices across several dimensions has been an important component in the study of market integration. A great deal of the literature has dealt with the issue of competitive links among geographically dispersed markets and has employed a variety of correlation and regression techniques to demonstrate the relationships that exist across space and time. From these results have come conclusions about the degree of competition or integration that exist in a market network. B. Harriss (1979) critiques this literature and observes that most studies fail to distinguish between the concepts of competition and market integration. "[W]e must note a common incidence of tautology in the definition of integrated markets as characterized by high [correlation] coefficients and the conclusion from their occurrence that markets are therefore integrated and competitive" (p. 202). Integration is a spatial concept, and a spatially integrated market does not have to be competitive.

Even casual observation confirms the verity of the above remarks for the international rice market. The prevalence of government monopolies on trade, domestic tax, and subsidy policies, and the widespread use of concessional sales and contractual agreements, are exemplary of the broad range of noncompetitive influences on international rice trade.[2] These factors imply that the rice price observed in the international market bears no necessary resemblance to the price which would prevail in a freely competitive situation, where exports are determined by comparative advantage and all consumers face relative rice prices which differ across countries only by the amount of transport costs.

While correlation and regression techniques cannot indicate the degree of competition in trade, price analysis can still be a useful tool to investigate the degree of interconnectedness among prices in a market at any given time. The use of correlation and regression analysis for this purpose relies on the belief that prices for different locations or qualities cannot move independently in an integrated market system. Substitution effects across qualities in both consumption and production will result in the transmission of price effects across the entire range of qualities in the market. Changes in export availabilities from a particular country will result in cross-country movements in demand by prospective importers and thus transmit the impact of a stochastic disturbance throughout the market. In a market which is not integrated by the price mechanism, deviations in weather patterns or government policies which influence the quantity and f.o.b. price of a particular quality of rice in a particular country should have no impact on the f.o.b. prices of other qualities or other countries. Price movements across countries and qualities would be essentially independent.

Most empirical analyses have presumed that with random price behavior expected of a nonintegrated market, the bivariate correlation coefficients of price movements (for different qualities and countries) will be zero. Conversely, in a

[2] The prominence of government of monopolies in rice trade is demonstrated in Food and Agriculture Organization of the United Nations (FAO), Intergovernmental Group on Rice, Compendium of National Rice Trade Policies, CCP:RI/CONS 74/3, November 1973 and CCP:RI 78/2, December 1977. Further discussion of the role of government controls is contained in W. P. Falcon and Monke (1979-80) and Monke (1980, chs. 2 and 6).

perfectly integrated market the correlation coefficient of price movements is expected to be unity. However, a number of factors suggest that the statistical verification of market integration is a much more complex process than the estimation of bivariate correlation coefficients. Cross-price elasticities of supply and demand are not equal across qualities, and thus cross-quality price movements will not be expected to be equal in an integrated market. Variations in the distances between potential importers and exporters mean that transport cost margins are unlikely to be constant as trade patterns change. P.A. Samuelson (1952, pp. 297-302) has shown that prices at particular locations in spatially-integrated markets can move independently of one another over a certain price range if optimal trading patterns dictate that the countries in question do not physically trade with one another. Finally, Harriss (1979) and C.P. Timmer (1974) point out that markets may be spatially integrated, but demonstrate low price correlations because of changes in the geographical direction of price formation. Within the course of a year, a particular market may serve as a center of supply, nonfinal demand, and final demand, and annual averages of prices may give little indication of the presence of spatial integration. Hence even with perfect spatial integration price variations across countries will not be exactly correlated, and correlations and regression coefficients will be less than one. The finding of significant relationships in spite of additional variation of price movements in an integrated market thus strengthens the case for market integration. Difficulties in interpretation arise only with the use of low significance levels (large standard errors) to reject the hypothesis of market integration.

Since confidence intervals can be determined for both bivariate price correlation and regression, both techniques appear equally attractive as analytical tools. But two additional factors mediate in favor of regression as a preferred technique. As Harriss (1979) points out, correlations in time series data can be spurious, particularly in periods of secular price movements. This problem is essentially one of an underspecified relationship, in which changes in one price are allowed to explain an undue share of changes in the other price. Regression analysis has the capability of overcoming this problem by the addition of new variables, particularly a time trend, to the relationship.

A second factor which favors the reliance on regression techniques rather than correlation analyses involves the requirement of stationary time series. Correlation coefficients can be calculated for first-differences rather than for absolute prices, but this procedure provides no guarantee that problems of nonstationary prices have been resolved. In regression analysis, standard statistical techniques can be used to determine the probability of autocorrelation. If first-order autocorrelation is present, estimation techniques other than ordinary least squares can then be applied to estimate the relationship.

"SPATIAL" INTEGRATION

The "spatial" integration of the international rice market can be examined through analysis of the behavior of annual average c.i.f. and f.o.b. rice prices among major importers and exporters across time.[3]

The form of the equation tested is

$$P_{it} = \beta_0 + \beta_1 P_{jt} + \beta_2 t , \qquad (1)$$

where P_i, P_j = annual average c.i.f. or f.o.b. prices of countries i and j, respectively, $(i \neq j)$, and t is a time trend variable from 1961 to 1977.

The first set of pairwise regressions were performed on the importing countries of Iran, Saudi Arabia, Bangladesh, Indonesia, Sri Lanka, Senegal, Singapore, and South Korea. These are major and consistent importers, accounting for about 30 percent of the import trade throughout the 1960s and 1970s (Falcon and Monke, 1979-80). They also have the distinction of having apparent preferences in the particular type of rice they consume. Iran and Saudi Arabia have imported high quality (*indica*, long-grain) rice since the early 1970s. Bangladesh and Indonesia have been major recipients of concessional aid. Senegal, Sri Lanka, and Singapore form a geographically diverse group of consumers of predominantly broken rice. Since Japan's withdrawal from the market in the late 1960s, South Korea has been the only major importer of round-grain (*japonica*) rice.[4]

The set of countries representing the exporters includes the United States, Egypt, Burma, China, Pakistan, Thailand, Italy, Australia, Japan, and Taiwan. These countries represent 85 percent of world exports. Italy, Taiwan, and Japan are predominantly round-grain producers. The remaining countries produce a mix of indica varieties.

Complete results are provided in the Appendix, and Tables 1 and 2 present only the price coefficients and the corresponding t-statistics.[5] Where first-order autocorrelation was found, Cochrane-Orcutt generalized least squares estimation was used. Examination of the autocorrelation structures gave no indication of higher-order autocorrelation (Monke, 1980, ch. 2).

Of the importers, all the indica countries show significant t-statistics at the 95 percent confidence level with the exception of the Bangladesh-Singapore and Saudi Arabia-Indonesia combinations. The former is significant at the 90 percent level; the poor fit of the latter may be due to the abrupt change in the quality of Saudi imports after 1971. With the exception of Sri Lanka and Senegal, time trends were significant, suggesting that unit values maintained a constant relationship across most countries. Price movements within the indica market appear well-synchronized.

[3] Prices are drawn from the FAO *Trade Yearbook*, with the exception of Taiwanese f.o.b. values. FAO's data are based on official government statistics. Data contained in the trade yearbooks are not entirely consistent across issues. Estimates of the quantity of Indonesian imports for 1968, for example, will differ in the 1969, 1970, and 1971 yearbooks. Where such differences appeared, the most recently published statistic is assumed to represent the best information.

[4] A classification of the imports and exports of the major trade participants is provided in T.F. Moriak, R.P. Strickland, and W.R. Grant (1975).

[5] Only one-half of the possible number of regressions are estimated. Because of the presence of the time trend in the regression, it is likely that the price coefficient estimates and t-statistics will differ between regressions which reverse the choice of the dependent and independent price variable. However, random checks of this possibility produced no contradictions to the conclusions described in the text.

TABLE 1.—IMPORT UNIT VALUES, CROSS-PRICE RELATIONSHIPS, 1961-77[a]

Dependent variable	Independent variable							
	Senegal	Iran	Saudi Arabia	Bangladesh	Singapore	Indonesia	Sri Lanka	South Korea
Senegal		.46	.24	.82[b]	.58[b]	.58	.77	.00
		(6.67)	(3.11)	(3.24)	(8.91)	(2.29)	(5.73)	(-1.44)
Iran			.39	1.82	1.13	1.41	1.59	.00
			(2.40)	(3.95)	(6.14)	(3.34)	(8.01)	(-.79)
Saudi Arabia				1.68	.90[b]	.71	1.19	-.01
				(2.22)	(2.17)	(.94)	(2.12)	(-1.26)
Bangladesh					.22[b]	.61	.52	-.002
					(1.64)	(3.20)	(3.72)	(-.88)
Singapore						.79	1.16	.00
						(2.13)	(6.56)	(-1.09)
Indonesia							.64	.00
							(4.47)	(.03)
Sri Lanka								.00
								(-.72)
South Korea								

[a] All regressions involved 17 observations, unless otherwise indicated.
[b] Cochrane-Orcutt generalized least-squares estimation; 16 observations.

TABLE 2.—EXPORT UNIT VALUES, CROSS-PRICE RELATIONSHIPS, 1961-77[a]

	Independent variables									
Dependent variables	United States	Egypt	Burma	China	Pakistan	Thailand	Italy	Australia	Japan	Taiwan
United States		.45	.93	1.11[b]	.52	.84	.79[b]	1.13[a]	.83[b]	1.24
		(8.11)	(7.32)	(4.39)	(2.53)	(12.6)	(3.29)	(3.76)	(3.61)	(3.71)
Egypt			2.00	2.82[b]	1.46	1.64	1.27	3.18[a]	1.87[b]	1.68
			(11.0)	(6.84)	(4.80)	(9.84)	(3.04)	(8.73)	(2.33)	(2.18)
Burma				.88	.57	.80	.61	1.13	.96	.79
				(4.86)	(3.21)	(14.5)	(3.09)	(5.26)	(2.90)	(1.86)
China					.60	.40[b]	.74	1.19	.70	.81
					(4.24)	(4.44)	(6.31)	(11.2)	(3.41)	(3.15)
Pakistan						.55	.48	1.26[b]	.50	.00
						(2.65)	(1.87)	(5.44)	(.95)	(-.12)
Thailand							.78	1.31	1.09[b]	1.26
							(3.53)	(4.70)	(3.54)	(2.78)
Italy								1.06	.74	1.18
								(3.92)	(4.17)	(4.87)
Australia									.58	.51
									(2.43)	(2.26)
Japan[c]										.82[e]
										(1.45)
Taiwan[d]										

[a] All data are ordinary least-squares estimates, with 17 observations, unless otherwise indicated.
[b] Cochrane-Orcutt generalized least squares; n - 1 observations.
[c] 7 observations.
[d] 14 observations.
[e] 6 observations.

Between the indica and japonica markets, however, the evidence is less clear. All the price coefficients for South Korea, the only japonica importer in the sample, are near zero and statistically insignificant. The time trend, however, is significant at the 99 percent level and the coefficients are positive, suggesting that japonica prices have declined relative to indica prices over the last two decades. This trend in relative prices corresponds to the general decline of import demand for japonica rice, due principally to Japan's withdrawal from the import market.

The group of exporters also demonstrates a close interaction over time. This is true generally even between the block of japonica producers and indica producers. Only three of the 45 estimations yielded insignificant coefficients, with the weakest links occurring between Pakistan, Taiwan, and Japan. Most japonica exporters must sell in indica markets, and the significance level of the price coefficients suggests that, in these cases, japonica prices cannot move independently of indica prices. Significant time trends appeared primarily in the United States, Egyptian, and Japanese estimations. The reasons for the United States result are not understood, but in the latter two cases, the japonica-indica relationship may again account for the trend.

These results suggest that rice markets are integrated geographically, when considering adjustments that occur within the course of a year. The japonica-indica distinction may be significant, but relatively unimportant for the international market due to the small magnitude of the japonica trade. This conclusion is based upon the behavior of annual unit values rather than prices for specific qualities of rice. But given the institutional structure of trade, in which most imports and exports are directly determined by governments rather than consumers, unit values may provide the most relevant variables for an analysis of integration of the rice market.

INTERTEMPORAL LINKAGES

The particular relationships among qualities and the specific time lags involved in price adjustment are not directly addressed by the above analysis. To explore the intertemporal linkages of rice prices, five series of monthly prices are analyzed. Three grades of Thai rice and two grades of United States rice were chosen—5 percent brokens, A-1 super brokens, and parboiled 5 percent brokens, all f.o.b. Bangkok, U.S. No. 2 long grain, f.o.b. Houston, and U.S. No. 1 medium-grain, f.o.b. San Francisco. The series are complete for 1967-78 inclusive, except for March-December 1973 when price quotes were unavailable for Thai rice.[6]

[6] Weekly price data (575 observations) were taken from the *Rice Market News* and aggregated into monthly averages. The United States data were obtained from private traders. Incorporation of export subsidies did not affect the results as the amount of price variation during the subsidy period (1969-72) was an insignificant proportion of the variation during the 1967-78 period. The Thai prices represent quotations from the Rice Price Subcommittee of the Board of Trade's Rice Committee, an organization of private traders established as a liaison between rice traders and the government. Quoted prices do not necessarily agree with actual export prices. If a consistent bias is present in the quoted prices, no problems are created for an estimation in first differences since quoted prices will move in concert with actual prices. If the biases in quoted prices are not consistent, but may be regarded as additional random error, the variance of coefficient estimates will be increased, and the reported t-statistics will understate their true values.

The time series appears to represent two very different periods of price behavior. From 1967–mid-1972 markets were extremely quiet. Nominal prices moved relatively little and probably declined in real terms. During 1970-71, for example, the price of U.S. No. 2's (inclusive of export subsidies) changed only four times, and then by less than 1 percent. In mid-1972, however, the market picture began to change, and during 1973-74 rice prices moved in an unprecedented manner. Fueled by grain-production shortfalls in 1972 and 1973 and the drawdown of stocks in the potential substitutes of wheat and corn, rice prices increased by as much as $200/metric ton (mt) per month. With the return of good weather, prices came down nearly as quickly as they had increased. Price fluctuations continued into 1978, at a somewhat lesser rate than the 1973-74 period, but well above the 1967-72 period (Falcon and Monke, 1979-80).

Because of the shifting character of the market, Equation (1) was estimated using the first differences of the prices. This improves the stationarity of the series making the statistical tests more reliable. An analysis of covariance suggested the *intertemporal* behavior of the prices did not change significantly between the 1967-72 and 1973-78 periods (Monke, 1980, ch. 11, app. 4). Table 3 contains the price coefficients and t-statistics for the Equation (1) regressions performed on the entire sample.

Intra-United States and intra-Thai price movements for different qualities are significant at the 99 percent confidence level. Markets are strongly linked within each country, but the relationship between Thai and United States prices is less clear. Only half the coefficients were significant at the 95 percent level, and all the relations between the Thai rices and U.S. No. 2's (the dominant quality in United States trade) were of negative sign, decidedly contrary to expectation and the behavior suggested by the analysis of the series of annual average unit values. This is an indication that over the period of study, the Thai and United States markets were sufficiently disjoint so that one month periods were not, on average, sufficiently long to allow events in one market to impact fully on the events in the other.

To examine the time lags involved in the transfer of price information, Equation (1) was altered along the lines of an econometric causation model developed by C.W.J. Granger (1969) and employed in macro models by C.A. Sims (1972). These statistical tests of causality look to see if current values of X can be better predicted using past observations on Y than by not including these lagged values. If this is true, Y is said to cause X in the Granger sense. Numerous theoretical associations between X and Y could be suggested, but the Granger definition, examining the direction of the association along the time dimension, offers a powerfultool of analysis in that it is empirically testable. The key requisite for this approach (as applied by Sims) to be valid is that the time series examined are characterized by white noise[7] (or can be filtered to be so). These price series, in their first difference form, meet this requirement (Monke, 1980, ch.2, app. 3).

[7] This is a series that does not exhibit any form of autocorrelation. For a discussion of this and related topics see Box and G.M. Jenkins (1970, p. 46).

The procedure involves regressing current values of the first series on past, current, and future values of the second. Then the regression is reversed. The regression forms are given below:

$$P_{i0} = \beta_0 + \beta_1 P_{j(-t)} + \beta_2 P_{j(-t+1)} + \ldots + \beta_i P_{j0} + \beta_i + 1 P_{jt} + \ldots + \beta_{i+t} P_{jt} \quad (2)$$

$$P_{j0} = \beta'_0 + \beta'_1 P_{i(-t)} + \beta'_2 P_{i(-t+1)} + \ldots + \beta'_i P_{i0} + \beta'_{i+1} P_{it} + \ldots + \beta'_{i+t} P_{it} \quad (3)$$

The choice of t, the lengths of the leads and lags, can be determined statistically by means of an F-test. The regressions that follow are based on leads and lags of two months.

There are several relationships that could exist between the two series. For example, if P_i causes P_j, it should be expected that the lagged values of P_i would help explain current P_j in Equation (3), and that the *future* values of P_j would contribute to explaining P_i in Equation (2). If P_i and P_j demonstrate a consistent relationship in terms of lead-lag behavior, the equations are reestimated with only the significant price variables. To continue the example, Equation (3) is reestimated without future values of P_i, and Equation (2) is reestimated without the lagged values of P_j. This analysis of variance provides an F-statistic to test whether the coefficient of the insignificant price variables can be assumed zero at the 5 percent confidence level.

Results are provided in Table 4. The introduction of the leads and lags removes the puzzle of the negative coefficients found in Table 3, and the statistical results are again consistent with a view of an international market integrated over space and time. United States price movements consistently lag one month behind Thai movements—both the largest coefficients and the largest t-statistics are consistently United States prices (+1) and Thai prices (-1). In only one case out of ten is the F-statistic large enough to suggest that causation might occur in the opposite direction. These results only suggest that the geographical focus for price determination in the international rice market is in Asia. Since Asian countries account for 90 percent of world production and 60 percent of world trade, these findings are consistent with a priori expectations.[8]

The total impact of one price on the other can be thought of as the sum of the regression's coefficients. In the cases of Table 3 where Thai and United States prices were instantaneously negatively correlated, the introduction of the coefficients on leads and lags shows that the total impact over the five-month period was indeed positive, as intuition would suggest. The adjustment between the U.S. long-grain No. 2's and the Thai varieties appeared essentially complete within a two-month period, as the sum of the lagged Thai coefficients ranged between 1.0 and 1.2. Long-grain varieties account for about two-thirds of United States exports. In the remaining cases, the relevant coefficients sum to substantially less than 1, suggesting that the adjustment process is more complicated than the Thai-United States long-grain relationship.

[8] These results are consistent with the observation of trade experts that Thai traders frequently play a leadership role in the process of price formation, in that Thai traders are the quickest to adjust to changes in supply-demand relationships (Efferson, personal communication).

TABLE 3.—NOMINAL PRICE DIFFERENCES, 1967-78[a]

	United States			Thailand	
	Long-grain No. 2, white	Medium-grain No. 1, white	5 percent brokens, white	A-1 super brokens, white	5 percent brokens, parboiled
United States					
Long-grain No. 2, white	1.0	.46 (3.04)	-.36 (-2.47)	-.35 (-1.82)	-.11 (-.73)
Medium-grain No. 1, white		1.0	.12 (1.38)	.21 (1.91)	.02 (.248)
5 percent brokens, white			1.0	.96 (11.8)	.88 (17.1)
Thailand					
A-1 super brokens, white				1.0	.56 (11.3)
5 percent brokens, parboiled					1.0

[a] All estimates utilized generalized least squares; 127 observations

QUALITY LINKAGES

Pairwise comparisons of prices have given clues of the links that exist over space and time, even for varying types and qualities of rice. A more thorough examination of the quality issue can be achieved by moving beyond the two variable world.

The price of rice depends critically on the characteristics it possesses. The percent brokens, the length of grain, and whether it is raw or parboiled may influence a particular sample's price. The analysis of price on the basis of a commodity's characteristics was developed by A.T. Court (1939) and was revived by Z. Griliches (1961) and others. This approach suggests that the price of a good is determined by the value of the attributes it possesses. The major impact of this work has been in the construction of quality adjusted, or hedonic, price indices.[9] This technique has also been usefully employed in demand studies of heterogeneous groups of goods, like automobiles and houses.

The general procedure behind the hedonic approach involves collecting data on both the price and the relevant characteristics that affect the "quality" of the good in question. Then the price is expressed as some function of these characteristics, and the relationship is estimated econometrically. For the study of the quality-price relationship in rice, the Thai-United States data of the previous section are used. Having established a certain linkage among these types, it is possible to place a value on particular characteristics. Unfortunately, the breadth of the examination is limited by the availability of data. For example, the length of grain is either long or medium, as no round-grain prices were available in such a complete series. Even so, the results should be suggestive. Equation (4) is the relationship to be tested.

$$P = f(X_1, X_2, X_3, X_4), \tag{4}$$

where X_1 = country of origin
= 1 if U.S.
= 0 if Thailand

X_2 = length of grain
= 1 if long
= 0 if medium

X_3 = percentage brokens
= 1 if less than 5 percent broken
= 0 if greater than 5 percent broken

X_4 = raw or parboiled
= 1 if raw
= 0 if parboiled.

Unlike previous studies where quantitative variables are used, all the quality variables are dummies. This poses no problem and, in fact, is entirely appropriate given the discrete nature of the rice grades.[10] The one difficulty of this procedure

[9] For an excellent review of this work see Griliches (1971).

[10] Griliches (1961) used the dummy variable approach to handle options on cars like automatic transmissions. The one distinct advantage of dummy variables is the ease of interpretation of the regression coefficients.

TABLE 4.—THAI-UNITED STATES INTERTEMPORAL PRICE REGRESSIONS

Dependent variable	Independent variable		Coefficient	(t-Statistic)	F-Statistic (2,102) from analysis of variance
U.S. long-grain No. 2	Thai 5 percent brokens	(+2)	.232	(1.46)	
		(+1)	.110	(.792)	
		(0)	-.488	(-3.45)	.88
		(-1)	.927	(6.55)	
		(-2)	.076	(.51)	
Thai 5 percent brokens	U.S. long-grain No. 2	(+2)	.060	(.35)	
		(+1)	.278	(6.14)	
		(0)	-.125	(-2.78)	.52
		(-1)	.050	(1.11)	
		(-2)	.025	(.55)	
U.S. long-grain No. 2	Thai A-1 super brokens	(+2)	.000	(-.002)	
		(+1)	.324	(1.77)	
		(0)	-.341	(-1.86)	1.49
		(-1)	1.068	(5.83)	
		(-2)	.140	(.74)	
Thai A-1 super brokens	U.S. long-grain No. 2	(+2)	.041	(1.13)	
		(+1)	.219	(5.95)	
		(0)	-.066	(-1.81)	2.29
		(-1)	.076	(2.05)	
		(-2)	.041	(1.10)	

U.S. long-grain No. 2	Thai 5 percent parboiled	(+2)	.052	(.27)	
		(+1)	.199	(1.15)	
		(0)	-.386	(-2.21)	.67
		(-1)	.847	(4.90)	
		(-2)	.189	(1.11)	
Thai 5 percent parboiled	U.S. long-grain No. 2	(+2)	.105	(2.41)	
		(+1)	.207	(4.70)	
		(0)	-.064	(-1.48)	.31
		(-1)	.037	(.84)	
		(-2)	.011	(.25)	
U.S. medium-grain No. 1	Thai 5 percent brokens	(+2)	.116	(1.12)	
		(+1)	.204	(2.26)	
		(0)	.064	(.69)	3.15
		(-1)	.029	(.31)	
		(-2)	.110	(1.15)	
Thai 5 percent brokens	U.S. medium-grain No. 1	(+2)	-.079	(-.90)	
		(+1)	.295	(6.30)	
		(0)	.074	(2.29)	2.87
		(-1)	.190	(2.29)	
		(-2)	.087	(1.05)	
U.S. medium-grain No. 1	Thai A-1 super brokens	(+2)	.005	(.044)	
		(+1)	.223	(2.02)	
		(0)	.198	(1.78)	1.97
		(-1)	.269	(2.42)	
		(-2)	.075	(.66)	

Dependent variable	Independent variable		Coefficient	(t-Statistic)	F-Statistic (2,102) from analysis of variance
Thai A-1 super brokens	U.S. medium-grain No. 1	(+2)	-.082	(-1.15)	
		(+1)	.230	(5.96)	
		(0)	.105	(1.54)	2.51
		(-1)	.152	(2.26)	
		(-2)	.017	(.25)	
U.S. medium-grain No.1	Thai 5 percent parboiled	(+2)	.015	(.13)	
		(+1)	.245	(2.36)	
		(0)	.056	(.54)	2.74
		(-1)	.004	(.03)	
		(-2)	.128	(1.26)	
Thai 5 percent parboiled	U.S. medium-grain No.1	(+2)	-.063	(-.75)	
		(+1)	.215	(4.68)	
		(0)	.037	(.45)	2.70
		(-1)	.185	(2.29)	
		(-2)	.046	(.58)	

is that no estimable functional form is suggested by Equation (4). G.E.P. Box and Y.R. Cox (1964) suggest ways of testing appropriate transformations of the variables. It was found here that the choice of functional form did not affect significantly the explanatory power of the regression.[11] The results in Table 5 are of regressions of the log of price on the dummy variables for those characteristics.

The first equation gives the results for the entire sample. It was noted in the previous section that there was a radical change in the rice market in 1973. Given this shift, an intercept dummy was placed in the regression equation.

The results show that there is a premium placed on rice that has a low percentage of brokens and that United States-origin rice is more expensive. Interestingly enough raw rice is valued higher than parboiled rice, but this probably reflects the underlying lower quality of rice on the Thai parboiled market. There is no significant difference in the prices paid for long- and medium-grain rices. As expected, the intercept dummy was highly significant, showing a much lower overall rice price in the 1967-72 period as compared to 1973-78.

It should be emphasized that except for the intercept dummy, time does not play a role here, and that the information contained in Equation (5.1) in Table 5 does nothing to test the hypothesis of market linkages. The purpose of the first regression equation is to demonstrate the existence of significant premiums and discounts according to the rice's characteristics.

To shed more light on the linkage issue, Equation (5.1) can be expanded. If the rice market is genuinely linked across all characteristics, the discounts and premiums should be relatively stable (on a percentage basis) across various time periods and market conditions, reflecting constant cross-price elasticities. If the percentage premium for low brokens changed with the market conditions, this would represent an inflexibility for either producers or consumers or both and would suggest that the market for that particular characteristic rice should be examined separately.

In Equation (5.2) in Table 5 slope dummies were used to test for the existence of such shifts. The only shift occurred in the parboiled market. All the other premiums remained stable between the two periods, adding support to the notion that the rice market is well linked on a Thai-United States basis, across grain lengths (long to medium), and across the percentage brokens.

Examination of the price data suggests that consumers demonstrate some inflexibility between parboiled and raw rice. Historically, parboiled Thai rice sold at a discount to high-quality raw rice, while in the United States, parboiled rice sells at a premium to high-quality raw rice. After the mid-1970s, intra-Thai export price relationships changed. While raw 5 percent brokens sold at a premium of $15-20/mt relative to parboiled rice during the February-July period; during the months August-January, parboiled rice was offered at an average premium of $20/mt. This pattern was historically unprecedented and

[11] The Box-Cox transformations alter the algebraic specification of the variables in the linear regression, and hence the implied functional relationship. Straight linear and logarithmic regressions are two special cases of the general Box-Cox form. It should be noted that in this study it is only necessary to transform the price variable, since exponential transformation of the (0.1) dummies would not alter the fundamental "off-on" relationship.

TABLE 5.—HEDONIC PRICE RESULTS FOR RICE[a]
(Independent variables)

Dependent variable log of price	Constant	X_1 = United States/ Thailand	DX_1	X_2 = long/ medium grain	DX_2	X_3 = per- centage broken	DX_3	X_4 = raw/ par- boiled	DX_4	D	$\bar{R}^2$ =
Equation (5.1)	5.26	.285		-.039		.465		.166		-.748	.805
	(113.5)	(9.98)		(-1.40)		(16.0)		(5.69)		(-41.2)	
Equation (5.2)	5.35	.259	.051	-.032	-.015	.440	.044	.085	.146	-.912	
	(81.4)	(6.27)	(.90)	(-.79)	(-.27)	(10.3)	(.77)	(1.98)	(2.53)	(-10.1)	.809

[a]t-statistics in parenthesis; D = 1, 1967-72; D = 0, 1973-78.

reflected large increases in parboiled demand from Nigeria and the Middle Eastern countries. United States exporters were traditional suppliers for these areas, but were unable to meet the recent demand increases due to a lack of parboiled capacity. Three additional parboiling mills are under construction in the United States, and the price relationships between Thai parboiled and raw rice are expected to return to their pre-1975 levels (Efferson, 1980). Thus these results imply that for the short-run, when supplies are relatively fixed by processing capacity, an analysis of the parboiled market should not weigh too heavily the events that effect the dominant indica market.

SUMMARY AND CONCLUSIONS

The purpose of this paper is to investigate linkages in the international rice market. The null hypothesis is that the degree of government intervention and quality variation prohibits the operation of a cohesive market, and tests were employed to search for evidence of integration across several dimensions. First, it was found that on a geographical basis, evidence of firm linkages existed in all but the relatively unimportant japonica import markets. This test captured any cross-country interrelationships that would operate within the space of a year.

The second test examined the structure of Thai-United States price interaction, initially constraining the reaction period to be a month. This test produced evidence of strong links within the Thai and United States markets, but failed to confirm the interaction discovered in the annual country data. With the inclusion of leads and lags of two months, the Asian market was found to be the geographic focus for this twelve-year period of study. Movements in the prices of various qualities and characteristics were closely linked. The hedonic approach to prices allowed further investigation of the linkage across qualities. This analysis demonstrated consistent premiums for certain rice characteristics that bridged two very different market periods. Only parboiled rice prices failed to maintain a consistent relationship with respect to the prices of indica varieties.

The approaches used here identified linkages in space, time, and quality. With the exception of the relatively small round-grain and parboiled markets, the international rice market appears to have been integrated over the period 1961-78. The implications are fairly clear. First, the modeling of the international rice market does not have to be an accounting exercise of adding individual, independent nation's behavior. Second, annual unit values of exports or imports provide a useful indicator of trends in the world market. In the absence of complete data on prices for imports or exports of any given country, the price of a dominantly traded indica variety (such as Thai 5 percent brokens) can be used, not with complete confidence, but with an anticipation that the relationships discovered will be indicative of the true factors under study. Finally, importers and exporters demonstrate sufficient substitution across qualities and trading partners to keep prices in line with one another. This result suggests that policy actions, such as concessional sales, reverberate throughout the market and cannot be considered to have occurred in a vacuum. In many ways this makes life more difficult for the observer of policy in the world rice market, but it emphasizes the critical importance of policy considerations in any examination of the international market for rice.

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TABLE A.1.—EXPORT UNIT VALUES, CROSS-PRICE RELATIONSHIPS

$$P_i = \alpha + \beta P_j + \delta T$$

Dependent variable	United States	Egypt	Burma	China	Pakistan	Thailand	Italy	Australia	Japan	Taiwan
United States		70.9 (4.68)	20.3 1.10)	44.3[a] (.66)	36.4 (1.04)	26.1 (2.33)	56.6 .66	-9.05[a] (-.17)	-803.2[a] (-1.48)	-79.1 (-1.54)
		.45 (8.11)	.93 (7.32)	1.11 (4.39)	.52 (2.53)	.84 (12.60)	.79 3.29	1.13 (3.76)	.82 (3.61)	1.24 (3.71)
		6.18 (3.25)	8.38 (4.40)	-2.14 (-.30)	9.10 (2.29)	6.43 (5.13)	-.24 -.03	3.32 (.59)	75.8 (9.92)	8.92 (2.32)
Egypt			-107.7 (-4.09)	241.9[a] (.89)	-104.9 (-2.01)	-85.4 (-3.04)	-69.4 (-1.08)	-146.5[a] (-1.23)	-1686.2[a] (-4.44)	-174.4 (-1.49)
			2.00 (11.0)	2.82 (6.84)	1.46 (4.80)	1.64 9.84)	1.27 (3.04)	3.18 (8.73	1.87 (2.33)	1.68 (2.18)
			5.47 (2.01)	-46.4 (-2.24)	2.24 (.38)	2.98 (.94)	2.42 (.29)	-21.2 (-2.05)	130.6 (5.67)	8.73 (.99)
Burma				8.62 (.36)	15.5 (.51)	12.2 (1.30)	21.2 (.70)	-24.7 (-.94)	-753.2 (-4.50)	-35.0 (-.54)
				.88 (4.86)	.57 (3.21)	1.80 (14.5)	.61 (3.09)	1.13 (5.26)	.96 (2.90)	.79 (1.86)
				-3.47 (-1.10)	.50 (.14)	-1.06 (-1.01)	-1.10 (-.28)	-4.43 (-1.44)	59.2 (5.66)	3.33 (.68)
China					12.9 (.54)	-37.5[a] (-.51)	10.8 (.60)	-29.9 (-2.29)	-642.4 (-4.96)	-7.17 (-.18)
					.60 (4.24)	.40 (4.44)	.74 (6.31)	1.19 (11.2)	.70 (3.41)	.81 (3.15)
					5.29 (1.94)	14.6 (2.58)	1.98 (.84)	.03 (.02)	57.4 (7.08)	2.76 (.94)
Pakistan						51.7 (1.49)	52.5 (1.32)	-12.0[a] (-.42)	-571.1 (-1.73)	121.4 (2.19)
						.55 (2.65)	.48 (1.87)	1.26 (5.44)	.50 (.95)	-.04 (-.12)
						7.04 (1.81)	5.92 (1.14)	-.73 (-.21)	58.8 (2.86)	8.27 (1.99)
Thailand							8.76 (.26)	-37.9 (-1.12)	-962.9[a] (-5.78)	-97.3 (-1.40)
							.78 (3.53)	1.31 (4.70)	1.09 (3.54)	1.26 (2.78)
							-.54 (-.12)	-3.08 (-.78)	77.9 (7.62)	4.33 (.83)
Italy								-10.7 (-32)	-699.4 (-6.25)	-41.7 (-1.12)
								1.06 (3.92)	.74 (4.17)	1.18 (4.87)
								3.49 (.91)	63.0 (9.00)	2.01 (.72)
Australia									-462.3 (-3.10)	37.2 (1.08)
									.58 (2.43)	.51 (2.26)
									44.6 (4.78)	3.97 (1.54)
Japan										692.0 (3.49)
										.82 (1.45)
										-54.1 (-2.49)
Taiwan										

(Independent variable: United States through Taiwan.)

KEY: $\alpha\ (t_\alpha)$ / $\beta\ (t_\beta)$ / $\delta\ (t_\delta)$

[a] Denotes Cochrane-Orcutt estimation techniques.

TABLE A.2.—IMPORT UNIT VALUES, CROSS-PRICE RELATIONSHIPS

$$P_i = \alpha + \beta P_j + \delta T$$

Dependent variable	Independent variable: Senegal	Iran	Saudi Arabia	Bangladesh	Singapore	Indonesia	Sri Lanka	South Korea
Senegal		39.1 (2.58) .46 (6.67) .66 (.28)	51.0 (2.15) .24 (3.11) 4.30 (1.21)	-4.18 (-0.15) .82 (3.24) 5.19 (1.59)	-53.0[a] (-1.04) .58 (8.91) 10.2 (2.57)	23.7 (.84) .58 (2.29) 2.02 (.38)	6.02 (.33) .77 (5.73) 4.30 (1.95)	-19.9 (-.32) -.004 (-1.44) 18.7 (3.62)
Iran			24.2 (.49) .39 (2.40) 12.9 (1.73)	-121.1[a] (-1.41) 1.82 (3.95) 11.2 (1.24)	-67.0 (-1.99) 1.13 (6.14) 12.1 (3.19)	-39.6 (-.84) 1.41 (3.34) .36 (.04)	-67.4 (-2.48) 1.59 (8.01) 9.03 (2.77)	-160.0 (-1.35) -.004 (-.79) 41.3 (4.21)
Saudi Arabia				-119.8 (-1.41) 1.68 (2.22) 19.2 (1.97)	-1281.5[a] (-1.65) .90 (2.17) 109.6 (2.25)	-42.7 (-.50) .71 (.94) 21.6 (1.35)	-78.0 (-1.01) 1.19 (2.12) 21.6 (2.34)	-241.6 (-1.51) -.01 (-1.26) 54.0 (4.08)
Bangladesh					19.6[a] (0.37) .22 (1.64) 8.95 (1.86)	37.9 (1.99) .61 (3.60) -2.14 (-0.60)	35.3 (1.86) .52 (3.72) 3.53 (1.55)	25.7 (0.63) -.005 (-0.97) 12.5 (3.58)
Singapore						43.3 (1.04) .79 (2.13) -1.95 (-.25)	13.5 (0.56) 1.16 (6.56) .07 (0.02)	33.4 (.33) -.005 (-1.09) 16.9 (2.01)
Indonesia							6.96 (0.4) .64 (4.47) 11.5 (4.93)	-31.4 (-.51) .001 (.03) 24.3 (4.75)
Sri Lanka								2.12 (.03) -.003 (-.72) 15.4 (2.41)
South Korea								

KEY: α (t_α)
β (t_β)
δ (t_δ)

[a] Denotes Cochrane-Orcutt estimation techniques.