Journal of Agricultural and Resource Economics 41(3):444–457 ISSN 1068-5502

Market Integration and Price Transmission in the World Rice Export Markets

Bo Chen and Sayed Saghaian

This paper investigates market integration and asymmetric price transmission in the world rice export markets. Using monthly rice prices from Thailand, Vietnam, and United States, we employ the Johansen test and estimate the threshold vector error correction model (TVECM). Our main findings are that export prices in the three countries are cointegrated, with Thailand and the United States the price leaders, and that the Vietnamese price adjusts faster to long-run equilibrium when it is above its equilibrium level with Thai and U.S. prices. These results suggest market integration and competition rather than collusion are prevalent in world rice markets. Policy implications are also briefly discussed.

Key words: asymmetric adjustment, competition, market integration, price transmission, rice

Introduction

Rice is a main staple food around the world, and consumers in many countries rely on imported rice to meet their daily needs. Rice is also an important source of income for farmers in major producing areas, especially smallholders in Southeast Asia. An adequate rice supply and stable prices in world rice markets contribute to the welfare of rice market participants.

The world rice market is typically characterized as thin, with only 6% of the total production traded across borders (Childs, 2012). This thinness magnifies the impact of large transactions and provides incentives for market manipulation (Anderson et al., 2007). Besides, exports are concentrated in six countries: Thailand, Vietnam, the United States, India, Pakistan, and China; together they have contributed more than 80% of total rice exports since 2000 (U.S. Department of Agriculture, Economic Research Service, 2014). These features of the world rice market render rice prices volatile, and a sudden reduction of supply by one major exporting country could lead to rapid price increases. Several studies have suggested that the rice export restrictions imposed by some major exporters contributed to the steep price hikes in 2008 (Childs and Kiawu, 2009; Headey and Fan, 2008).

Given these market conditions, an important question is whether the international rice market is integrated and efficient. A well-integrated world rice market indicates competitive pricing and price efficiency and implies that the world rice market is a reliable source to meet domestic demands. An integrated world rice market also suggests that major exporters do not have market power to charge higher prices than those prevalent in the market; thus, policies resembling the Thai rice-pledging scheme, the aim of which was to raise global prices by stockpiling, are ineffective or costly (Yang, Bessler, and Leatham, 2000). In this study, we formally test for the market integration of the world rice market, incorporating the three major rice exporting countries: Thailand, Vietnam, and the United States.

Bo Chen is a Ph.D. candidate and Sayed Saghaian is an associate professor in the Department of Agricultural Economics at the University of Kentucky.
This is Kentucky Experiment Station Manuscript Number: 16–04–035. We appreciate the anonymous reviewers' comments. Any remaining errors are the sole responsibility of the authors.

Review coordinated by Hikaru Hanawa Peterson.

Asymmetric price adjustment,[1] frequently observed in agricultural commodity markets, may also be present in world rice markets, providing more insights into the nature of competition in these markets. Specifically, a positive asymmetric price transmission (when price decreases are passed along faster than price increases) suggests competition among the sellers for market share, while a negative asymmetric price transmission (when price increases are passed along faster than price decreases) is likely an indication of market power (Meyer and von Cramon-Taubadel, 2004). We further explore scenarios of asymmetric price transmission in the world rice market with the threshold vector error correction model (TVECM) developed by Hansen and Seo (2002). To the best of our knowledge, the possibility of asymmetric price transmission in world rice markets has rarely been investigated (with only one exception: Ghoshray, 2008). Our main contributions lie in our theory-based empirical model and use of more recent data.

Background and Literature Review

Most global rice exports come from a few countries concentrated in southeast and southern Asia, while rice-importing countries are dispersed across the globe. The international rice trade has substantially benefitted from free trade policies established under the General Agreements on Tariffs and Trade (GATT) and the World Trade Organization (WTO), but nearly all the major rice exporting countries have policies either to stabilize domestic prices in order to ensure domestic food security or to subsidize domestic farmers to increase profits. Hence, government policies likely affect international rice pricing.

The literature in this area mostly focuses on assessing the impact of policy changes in major exporting countries—especially Thailand—on world rice markets. Since 2001, the Thai government has used a rice-pledging program to offer farmers loans with a value of up to the expected harvest-time market value of their rice crop. Depending on the pledging price and market price, farmers can choose to either redeem or forfeit their collateral rice. Though rice trade has been relatively free from direct government intervention, among major rice exporters, only Thailand and the United States did not restrict rice trade in the 2008 food crisis (Childs, 2012). In 2011, the Thai government adopted a policy change to substantially raise the pledging price above the predicted market price so that farmers were incentivized to forfeit, leaving the government with a substantial stock of rice (John, 2013). In doing so, the Thai government expected world rice prices to increase, and covered the costs of the stockpiling (Mahanaseth and Tauer, 2014).

Using price series data from Thai domestic and export markets, John (2013) found that Thai domestic pricing programs do not heavily distort world rice prices. Mahanaseth and Tauer (2014) tested for the existence and extent of market power in a range of major Thai rice export destinations and rice verities using the residual demand elasticity (RDE) approach and found that Thailand's ability to influence its export price is constrained by competition from Vietnam and India, especially for low-quality and generic rice varieties.

In contrast to the free rice trade in Thailand, Vietnamese rice exports are tightly managed by the Vietnam Food Association (VFA), which is responsible for the setup and distribution of export quotas and stipulating minimum export prices (MEPs) (Ngan, 2010). State-owned enterprises, including VINAFOOD-1 and VINAFOOD-2, are unrestricted by export quotas or MEPs and account for nearly half of rice export shares in Vietnam (Fulton and Reynolds, 2015; Slayton, 2009). From a political economy perspective, VFA and VINAFOOD have been structured to benefit from the frequent export restrictions triggered by the volatile rice export markets (Fulton and Reynolds, 2015).

Studies on world rice market integration are scarce. Yavapolkul, Gopinath, and Gulati (2006) studied market integration in the post-Uruguay Round era with export prices from India, Thailand,

[1] We use three terms synonymously in contrast to linear adjustment: threshold, asymmetric, and nonlinear. However, their meanings could differ slightly in the literature.

Vietnam, and the United States and concluded that international rice markets were partially integrated, with U.S. and Thai rice prices leading Vietnamese and Indian prices. However, possible asymmetric price transmission in the international rice market is ignored. To fill this gap, Ghoshray (2008) adopted a methodology analogous to the Engle-Granger two-step estimation: Enders and Siklos's 2001 momentum threshold autoregressive model (M-TAR) is first applied to the deviation from the long-run equilibrium to investigate threshold integration between price series; for those price series with threshold integration, an asymmetric error correction model (AECM) is further estimated to capture asymmetric price transmission. Ghoshray found threshold integration between Thailand and Vietnam prices for high and medium rice. Additionally, Vietnam adjusted its prices faster when the price gap between the two countries was narrowing than when it was widening. These results implied that Vietnam was engaging in price undercutting behavior for world high-quality rice exports in order to compete with Thailand. Though intuitive, Ghoshray's M-TAR-AECM methodology does not have a foundation in statistic theory, and thus the estimator properties are not well understood. We contribute to the literature by studying market integration and possible asymmetric price transmission in Thailand, Vietnam, and the United States with Hansen and Seo's 2002 TVECM model, which is based on well-developed statistical theory. Besides the traditional top two exporters, Thailand and Vietnam, the United States is an interesting case to include in the analysis because it is the only major exporter not in south or southeast Asia.

Market Integration and Asymmetric Price Transmission

In the agricultural economics literature, market integration is generally defined through the Law of One Price (LOP), which states that—given free trade—arbitrage would equalize prices of the same good in different markets up to the transaction cost. This has become the theoretical foundation for a vast body of empirical literature aiming to test market integration based on price data alone, including pioneering works by Ardeni (1989), Goodwin (1992), and Asche, Bremnes, and Wessells (1999). However, the price-based modeling of market integration has drawn some criticism. Barrett, Li, and Bailey (2000) argued that testing market integration with price data alone relies on unrealistic assumptions about the trading behavior and cost of commerce; consequently, rejection of the null hypothesis of market integration cannot be distinguished from the rejection of assumptions of model specification. In addition, Barrett (2001) proposed to distinguish between market integration and market efficiency and redefine market integration with a flow-based indicator of tradability and market efficiency with price-based notion of market equilibrium. Following this distinction, Miljkovic (2009) studied market integration of Canadian and U.S. livestock markets with the above two model specifications, respectively, and further called into question the study of market cointegration with price data alone.

The criticisms of studying market integration without trade quantity data are valid in many market integration analyses but do not apply in this study of international rice markets, which maintains the LOP-based definition of market integration. First, it is interesting to note that the three countries in this study—Thailand, Vietnam, and the United States—are all major exporters of rice to the rest of the world. That is, a direct rice-trade linkage among the three countries is quite small: during the study period, only Thailand exported rice to the United States, which accounted for only 7% of the total Thai rice exports. Therefore, it is not arbitrage among the three countries that brings their prices into the long-run equilibrium. The world rice market includes several major exporting countries and many importing countries, and fully-informed importers switch among exporters to obtain the most competitive pricing. This switching is the mechanism underlying market integration, maintaining the LOP-based definition of market integration.

Second, many countries import rice from multiple exporters, and the export destinations are substantially overlapping, especially for Thailand and Vietnam. This feature incentivizes importers to switch to exporters with lower prices, effectively linking the export markets and resulting in integrated markets with closely linked prices (Goodwin and Vavra, 2009). Thus, the definition of market integration based on LOP can be justified is this research.

Third, a direct trade is not the only mechanism by which to link prices in different markets, and the flow of information can also play a substantial role. Stephens et al. (2012) investigated spatial price transmission of three tomato markets in Zimbabwe during periods with and without trade flows and found larger and more rapid price adjustments when no physical trade linked these markets. In world rice markets, also, the mechanism of importers switching to competitive exporters relies on importers' full price information in the world markets. Rice prices are constantly monitored by importers, exporters, and institutions such as FAO and the USDA due to their strategic importance.

In summary, the prices of the three major exporting countries in the world rice market are expected to establish certain linear relationships in the long run. We test this market integration hypothesis following the widely used Johansen cointegration test (Johansen, 1988; Johansen and Juselius, 1990). Meanwhile, the price transmission process whereby prices adjust to the long-run equilibrium is the norm in the short run. For a perfectly competitive market, price transmission tends to be symmetric, meaning a price increase in one market transmits to other markets with the same speed as a price decrease (Goletti, 1994), which is the assumption underlying the conventional vector error correction model (VECM). However, the world rice market may not be perfectly competitive, which suggests that the price adjustment toward long-run equilibrium may be asymmetric.

Since there are only a few major exporting countries in the world rice market, exporters may be incentivized to collude and rapidly follow one another's price increases under collusion, leading to a negative asymmetric price transmission. Signs of such collusive behavior among the major exporters are not rare. In 2002, the Thai government proposed establishing the Council on Rice Trade Cooperation (CRTC) with Vietnam, India, Pakistan, and China, which was intended to be a rice cartel (Poapongsakorn, 2010). Even though the CRTC never materialized, a similar proposal was again raised by Thailand in the Association of Southeast Asian Nations in 2012 (Mohindru and Phromchanya, 2012).

It is also possible for rice exporters to compete for market share and undercut one another's prices, leading to a positive asymmetric price transmission. Possible strategies include decreasing prices quickly when facing price decreases from other exporters but taking little or no action when other countries raise their prices. Both scenarios indicate asymmetric price transmission but have opposite implications for international rice markets. We first test for the existence of the asymmetric adjustment following Hansen and Seo (2002). If asymmetric adjustment is found, we continue to estimate a TVECM model and examine the adjustment speed for each regime to further infer which scenario leads to asymmetric price adjustment.

Data

We use monthly nominal rice price for Thailand, Vietnam, and the United States from August 2000 to July 2013 and our data source is the USDA/ERS Rice Yearbook. The choice of this dataset is based on both data compatibility and availability. Rice prices depend on numerous factors, including variety, grade, and processing technology. In this analysis, we include the most comparable rice types from these three nations to minimize the impact of quality differences on the price relations. *Oryza sativa* var. *indica* (hereafter "indica") is the most widely traded variety of rice globally, accounting for 80% of global rice trade (Childs and Chambers, 2000), and all three nations export large quantities of indica. Milled rice, as opposed to rough or brown rice, is the most traded rice form. Except for the United States, all of the major exporters do not allow export of rough rice in order to protect their domestic milling industries, and there is no comparable export price for rough rice across major exporters. Thus, for the sake of comparability, we chose milled indica rice for this study. Furthermore, price is heavily influenced by grade, which is mainly measured by broken rate. Not all broken rates are available for all exporters; however, high-grade (less than 5% broken rate) rice is common in international rice trade.

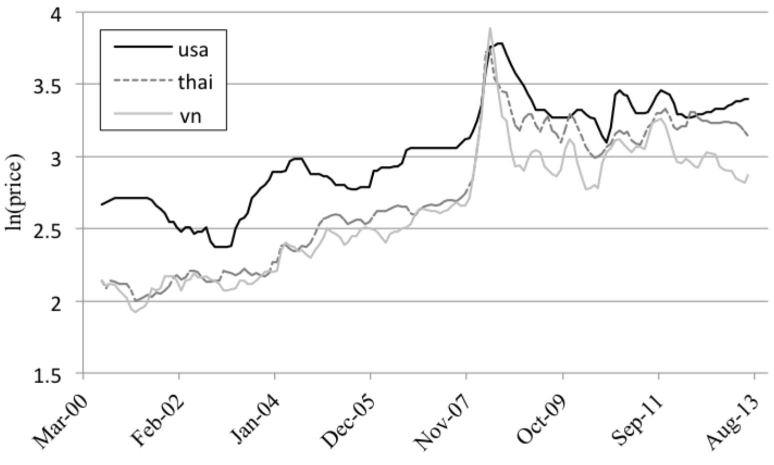

Figure 1. Log Prices of Rice in Thailand, Vietnam, and the United States
Source: Data derived from Rice Yearbook, U.S. Department of Agriculture, Economic Research Service (2014).

The Thai price is for milled rice with 5% broken rate, Free-On-Board (FOB) Bangkok. The Vietnamese price series is for double water-polished (DWP) long-grain rice with 5% broken rate, FOB Ho Chi Minh City. The Vietnam price series has nine missing values; cubic spline interpolation is employed to fill the missing values. The U.S. price is U.S. No. 2 long-grain rice with <4% broken rate, milled in Houston, Texas. All of these prices are converted to dollars per hundredweight ($/cwt) and natural logarithm transformation is applied. Note that using nominal price is common in price transmission literature (Esposti and Listorti, 2013), and the exchange rates used for converting prices to U.S. dollars already reflect the differences in inflation rates between countries.

The main reason for not including all major exporters is either lack of data or low quality of data with many missing observations. Data insufficiency is a major limitation in this area. The three countries included in this study are the important players in the international rice trade with data availability. As a result of sustained production growth, Thailand and Vietnam accounted for half of the world's rice exports in 2011; the pricing of rice in Southeast Asia has considerable influence on world rice markets (Baldwin et al., 2012). The United States is the third largest exporter of rice and half of its production is for exports; U.S. prices are sensitive to price movements in international rice markets (Childs and Chambers, 2000). India, Pakistan, and China are large producers but relatively small exporters during most months in the study period. Figure 1 plots the natural log of the export prices of the three exporters. The price series have generally been stable since the 2000–2001 marketing year, except for the steep hike in the 2007 global food crisis. Also, there exist clear price co-movements, especially between Thai and Vietnamese rice prices. This gives the first hint of market integration in international rice markets.

Empirical Methodology

Our empirical procedure comprises a series of tests and model estimations. We first perform ADF (Dickey and Fuller, 1979) and KPSS (Kwiatkowski et al., 1992) stationarity tests on the three price series. The two tests are complementary to one another due to their opposite null hypotheses. To study possible asymmetric price adjustment, we analyze the three prices in pairs. A Johansen cointegration test (Johansen, 1988; Johansen and Juselius, 1990) is performed on all three country pairs and VECM is estimated if cointegration is found. Furthermore, Hansen and Seo's 2002 *SupLM* test is employed to test for threshold adjustment. For price pairs demonstrating threshold adjustment, TVECM will be estimated.

Johansen Test and VECM

The Johansen cointegration test allows the test of cointegration relationships among times series based on maximum likelihood estimation. Consider a vector auto regressive (VAR) model with i price series under study:

$$\boldsymbol{P}_t = \boldsymbol{\mu} + \sum_{n=1}^{k+1} \boldsymbol{\Pi}_n \boldsymbol{P}_{t-n} + \boldsymbol{\varepsilon}_t, \tag{1}$$

where $\boldsymbol{P}_t$ is a $i \times 1$ price vector, $\boldsymbol{\mu}$ is a $i \times 1$ constant vector, $\boldsymbol{\Pi}$'s are $i \times 1$ parameter matrices, $\boldsymbol{\varepsilon}_t$ is the i.i.d. normal disturbance, and $k+1$ is the number of lags. Equation (1) can be written in error correction form such that

$$\Delta\boldsymbol{P}_t = \boldsymbol{\mu} + \boldsymbol{\Pi}\boldsymbol{P}_{t-1} + \sum_{n=1}^{k} \boldsymbol{\Gamma}_n \Delta\boldsymbol{P}_{t-n} + \boldsymbol{\varepsilon}_t, \tag{2}$$

where $\boldsymbol{\Pi} = \boldsymbol{\Pi}_1 + \boldsymbol{\Pi}_2 + \cdots + \boldsymbol{\Pi}_k - \boldsymbol{I}$ and $\boldsymbol{\Gamma}_k = -\sum_{j=k+1}^{p} \boldsymbol{\Pi}_j$.

The long-run matrix $\boldsymbol{\Pi}$ can be decomposed into $\boldsymbol{\Pi} = \boldsymbol{\alpha\beta}'$, where $\boldsymbol{\alpha}$ is the $i \times r$ adjustment vector, representing the speed of adjustments of the variables toward long-run equilibrium and $\boldsymbol{\beta}$ is the $r \times i$ cointegration vector, reflecting the linear relationships among the variables in the long-run equilibrium. The Johansen test can be used to test for the number of cointegration vectors by testing r, the rank of $\boldsymbol{\Pi}$. If a test of the null $r = 0$ is not rejected, then the price series are not cointegrated and there are no long-run relations among the price series. If $r = 0$ is rejected, the null $r = 1$ is further tested. The price series are cointegrated if $r = 1$ is not rejected. In addition, if cointegration is found among the price series, $\boldsymbol{\Pi}\boldsymbol{P}_{t-1}$ can be rewritten as $\boldsymbol{\alpha\beta}'\boldsymbol{P}_{t-1}$, where $\boldsymbol{\beta}'\boldsymbol{P}_{t-1}$ is also denoted as the error correction term ECT_{t-1}, representing the deviation from the long-run equilibrium at period $t-1$.

In addition, long-run causality can be inferred from the conventional t-tests performed on the coefficients of $\boldsymbol{\alpha}$, which are referred to as weak exogeneity tests. This causality is typically interpreted as price leadership (Motamed, Foster, and Tyner, 2008). Lastly, the $\boldsymbol{\Gamma}$'s represent the effects of price changes in previous periods on current price changes. In contrast to the speed of adjustments, $\boldsymbol{\alpha}$, which is in effect until long-run equilibrium is restored, $\boldsymbol{\Gamma}$ has only a temporary effect on the price-adjustment process.

Testing for and Estimating the Threshold Adjustment

The VECM discussed above assumes that the price pairs would adjust to the long-run equilibrium at a constant speed, $\boldsymbol{\alpha}$. In order to model different adjustment speeds to equilibrium in the international rice export market, a VECM is extended to a TVECM:

$$\Delta\boldsymbol{P}_t = \begin{cases} \boldsymbol{\mu}_1 + \boldsymbol{\alpha}_1 ECT_{t-1} + \sum_{n=1}^{k} \boldsymbol{\Gamma}^1_{t-n} \Delta\boldsymbol{P}_{t-n} + \boldsymbol{\varepsilon}^1_t, ECT_{t-1} \le v \\ \boldsymbol{\mu}_2 + \boldsymbol{\alpha}_2 ECT_{t-1} + \sum_{n=1}^{k} \boldsymbol{\Gamma}^2_{t-n} \Delta\boldsymbol{P}_{t-n} + \boldsymbol{\varepsilon}^2_t, ECT_{t-1} > v \end{cases} \tag{3}$$

where v is one threshold value for the error correction term ECT_{t-1} and all the other notations have similar meanings as in equation (2). Depending on the relative magnitude of the deviation from equilibrium to the threshold value, the long-run adjustment vector $\boldsymbol{\alpha}$ and short-run adjustment coefficients $\boldsymbol{\Gamma}$ may take different values, characterizing distinct adjustment processes toward long-run equilibrium; this is also known as threshold effect.

Table 1. Stationary Tests

	Level		First Difference	
	ADF	**KPSS**	**ADF**	**KPSS**
Thailand	−3.037	0.167**	−7.677***	0.064
Vietnam	−4.050***	0.177**	−6.845***	0.078
U.S.	−2.970	0.133*	−6.702***	0.052

Notes: Single, double, and triple asterisks (*, **, ***) indicate statistical significance at the 10%, 5%, and 1% level. The null hypothesis of the ADF test is that the price series is non-stationary. Critical values for ADF test with intercept and deterministic trend for 1%, 5%, and 10% significance are −4.019, −3.440, and −3.144, respectively. If the ADF test statistic is smaller than the critical value, the null hypothesis is rejected. The null hypothesis of the KPSS test is that the price series is stationary. Critical values for KPSS test with intercept and deterministic trend for 1%, 5%, and 10% significance are 0.216, 0.146, and 0.119, respectively. If the KPSS test statistic is larger than the critical value, the null hypothesis is rejected.

In order to justify the threshold effect in equation (3), Hansen and Seo (2002) developed a supreme Lagrange Multiplier (*SupLM*) test with the null of linear VECM, against which the alternative TVECM is applied:

$$SupLM = \underset{v_U \leq v \leq v_L}{SupLM}(\tilde{\boldsymbol{\beta}}, v), \tag{4}$$

where $\tilde{\boldsymbol{\beta}}$ is the estimated cointegration vector, v_U is the π percentile of ECT_{t-1}, and v_L is the $(1-\pi)$ percentile of ECT_{t-1}.[2] A grid search from v_U and v_L is performed to determine v, which maximizes the *SupLM* test statistic. The *SupLM* does not have a standard distribution, and a fixed regressor bootstrap method is used to obtain the critical values. If the above test rejects the null of linearity adjustment, conditional on the estimates of cointegration vector $\tilde{\boldsymbol{\beta}}$ and the threshold parameter v, the TVECM in equation (3) is estimated with conditional least square (CLS).

Test and Estimation Results

For each country pair, VECM is presented to show the long-run price relations and price transmission between countries in the pair. In addition, TVECM is also estimated and presented for comparison for country pairs demonstrating threshold price adjustment.

Stationarity and Cointegration Tests

The stationarity of both the levels and first differences of the price series are tested and the results are shown in table 1. The Thai and U.S. price are shown to be $I(1)$ in level, and their first differences are $I(0)$ at conventional significance levels (i.e., 1%, 5%, or 10%) in those two tests. The results are mixed for the Vietnamese price series. The ADF test rejects the null that the series is non-stationary at the 1% level, while the KPSS test rejects the stationarity null hypothesis at the 5% level. There is likely a unit root in the price level. Thus, the first difference of the Vietnamese price is tested and the results from ADF and KPSS unambiguously show the stationarity of the first-differenced prices series.

The results of pair-wise Johansen cointegration tests are presented in table 2. For the pairs of Thailand-Vietnam and U.S.-Vietnam, the Johansen trace tests reject that the prices are not integrated but fail to reject that there is at most one cointegration vector at 1%. For the Thailand-U.S. pair, the Johansen test only marginally fails to reject the null of non-cointegration at the 10% significance level, and thus the existence of a long-run price relationship between Thai and U.S. prices is still likely.

[2] $\tilde{\boldsymbol{\beta}}$ is estimated from linear VECM (Stigler, 2010); π is the trimming parameter, which is used to ensure a minimum number of periods within each of the regimes in the model and conventionally takes a value between 0.05 and 0.15 in the literature (Andrews, 1993). We set π equal to 0.15, as a too-small π may introduce an extreme regime during which price relations deviate substantially from long-run equilibrium; this further renders the estimation of adjustment speed based on this extreme period unreliable.

Table 2. Johansen Cointegration Test

	$r = 0$	$r = 1$
Thailand - Vietnam	29.34^{***}	2.46
Thailand - U.S.	16.57	2.88
Vietnam - U.S.	25.18^{***}	3.03

Notes: Single, double, and triple asterisks (*, **, ***) indicate statistical significance at the 10%, 5%, and 1% level. The null hypothesis of r=0 is that price series are not cointegrated; critical values for 1%, 5%, and 10% significance are 24.60, 19.96, and 17.85, respectively. The null hypothesis of r = 1 is that price series are cointegrated with one long-run relation; critical values for 1%, 5%, and 10% significance are 12.97, 9.24, and 7.52, respectively. If the test statistic is larger than the critical value, the null hypothesis is rejected.

Table 3. Linear VECM: Thailand and U.S.

	$\Delta p_{thai,t}$	$\Delta p_{us,t}$
ECT_{t-1}	-0.034	0.051^{***}
	(0.022)	(0.018)
Intercept	0.004	0.002
	(0.004)	(0.003)
$\Delta p_{thai,t-1}$	0.492	0.156^{*}
	(0.190)	(0.081)
$\Delta p_{us,t-1}$	-0.095^{***}	0.452^{***}
	(0.153)	(0.113)
Cointegration Vector	(1, -1.354)	
ADF test on deviation from long-run equilibrium	-3.835^{**}	
SupLM test	14.866	
LB-Q test on disturbances		
Lag 6	9.163	5.456
	[0.165]	[0.487]
Lag 12	24.898^{**}	9.5646
	[0.015]	[0.654]

Notes: Robust standard error in parenthesis for estimated coefficients. Single, double, and triple asterisks (*, **, ***) indicate statistical significance at the 10%, 5%, and 1% level. The number of bootstrap replications for *SupLM* test is 1,000, and the critical values for 1% and 5% significance levels are 21.782 and 18.234, respectively. The null hypothesis of the LB-Q test is that the disturbance series is a white noise process. p-values are in square brackets.

VECM and TVECM Estimations

Before further discussing the model estimation, post-estimation tests reveal that these models perform reasonably well. The ADF tests performed on the deviation from the equilibrium show that all three country pairs are stationary, and thus the estimated cointegration vectors indicate long-run relationships among the prices. Ljung-Box Q (LB-Q) tests are also performed on the disturbances, and the null hypothesis that the disturbance series is a white noise process cannot be rejected except in the Thailand-U.S. pair, where the disturbance series from the Thai price equation demonstrates slight serial correlation after large lags.

The results from the Thailand-U.S. price pair are presented in table 3. Since the *SupLM* test fails to reject the linearity in price adjustment, only the linear VECM is estimated. The coefficient of the speed of adjustment first reveals that 5.1% of the deviation from the long-run equilibrium for each period is corrected in the U.S. price. The Thai price, however, does not significantly adjust toward the long-run equilibrium. This result implies that Thai price leads the U.S. price, which is consistent with the fact that Thailand has been the world's largest rice exporter for decades, while the United States is a smaller player in the market. From 2001 to 2014, Thailand exported more than twice as much rice as the United States on average (U.S. Department of Agriculture, Economic Research Service, 2014). More importantly, the estimated cointegration vector indicates that, given a 1.35% increase in the Thai price, the U.S. price will increase by 1% in the long run; this implies Thai and U.S. markets are not fragmented, despite their geographical separation. It needs to be remembered

Table 4. Linear and Threshold VECMs: Thailand and Vietnam

	Linear VECM		Threshold VECM			
			Regime 1		Regime 2	
Observations (%)	–		81		19	
	$\Delta p_{thai,t}$	$\Delta p_{vn,t}$	$\Delta p_{thai,t}$	$\Delta p_{vn,t}$	$\Delta p_{thai,t}$	$\Delta p_{vn,t}$
ECT_{t-1}	0.085	-0.160^{**}	0.093^{*}	-0.161^{**}	-0.174	-0.500^{**}
	(0.056)	(0.068)	(0.052)	(0.072)	(0.194)	(0.201)
Intercept	0.004	0.004	-0.025^{*}	0.048^{**}	0.090	0.201^{**}
	(0.003)	(0.004)	(0.014)	(0.02)	(0.083)	(0.088)
$\Delta p_{thai,t-1}$	0.354^{*}	0.384^{**}	0.239^{**}	0.069	0.343	0.500^{**}
	(0.206)	(0.151)	(0.096)	(0.126)	(0.402)	(0.224)
$\Delta p_{vn,t-1}$	0.251^{***}	0.512^{***}	0.232^{***}	0.438^{***}	0.308	0.513^{***}
	(0.085)	(0.103)	(0.077)	(0.104)	(0.232)	(0.196)
$\Delta p_{thai,t-2}$	-0.342^{*}	-0.226	-0.211^{*}	-0.145	-0.572^{**}	-0.366
	(0.188)	(0.189)	(0.117)	(0.141)	(0.289)	(0.255)
$\Delta p_{vn,t-2}$	0.036	0.015	-0.008	0.063	0.324^{*}	0.244
	(0.099)	(0.136)	(0.130)	(0.171)	(0.196)	(0.211)
Cointegration Vector			(1, −0.847)			
ADF test on deviation from long-run equilibrium			-4.287^{***}			
SupLM test			32.109^{***}			
Threshold Parameter v			0.38			

LB-Q Test on Disturbances	VECM		TVECM	
	$\Delta p_{thai,t}$	$\Delta p_{vn,t}$	$\Delta p_{thai,t}$	$\Delta p_{vn,t}$
Lag 6	8.499	5.042	8.121	9.975
	[0.204]	[0.538]	[0.229]	[0.126]
Lag 12	16.755	7.784	16.048	13.313
	[0.159]	[0.802]	[0.189]	[0.347]

Notes: Robust standard error in parenthesis for estimated coefficients. Single, double, and triple asterisks (*, **, ***) indicate statistical significance at the 10%, 5%, 1% level. The number of bootstrap replications for *SupLM* test is 1,000, and the critical values for 1% and 5% significance levels are 28.481 and 25.044, respectively. The null hypothesis of the LB-Q test is that the disturbance series is a white noise process. p-values are in square brackets.

that the key driver behind market integration is unlikely to be direct trade, due to the tiny rice trade volume between the two countries, but rather importers' switching between these two exporting countries.

Table 4 presents the results from the Thailand-Vietnam price pair. The *SupLM* test rejects the null of linear VECM, indicating a potential threshold effect. Both linear and threshold VECM are estimated. Both models again suggest Thai price leadership, pointing to Thai's longstanding largest market share in the world rice market, whereas Vietnam only resumed exports in the mid-1990s (Childs and Chambers, 2000). In the long run, a 1% increase in the Thai price will eventually lead to a 0.85% price increase in the Vietnamese price. Export competition between the two countries may also contribute to this market integration. Two factors may intensify the competition. First, rice is the major crop in both countries. Not only it is the major staple food but it also plays an important role in the rural economy and agricultural development. Second, geographic proximity and similarities in exported rice varieties could mean that these countries have a large number of overlapping export destinations, leading to export competition.

Depending on the relative magnitude of the deviation from equilibrium and threshold parameters, the whole period can be divided into two regimes (figure 2): Regime 1, when Vietnamese price is relatively low ($p_{vn} - 0.847 \times p_{thai} \leq 0.38$), and Regime 2, when Vietnamese price is relatively high ($p_{vn} - 0.847 \times p_{thai} > 0.38$) compared to its price level in the long-run equilibrium. In Regime 1,

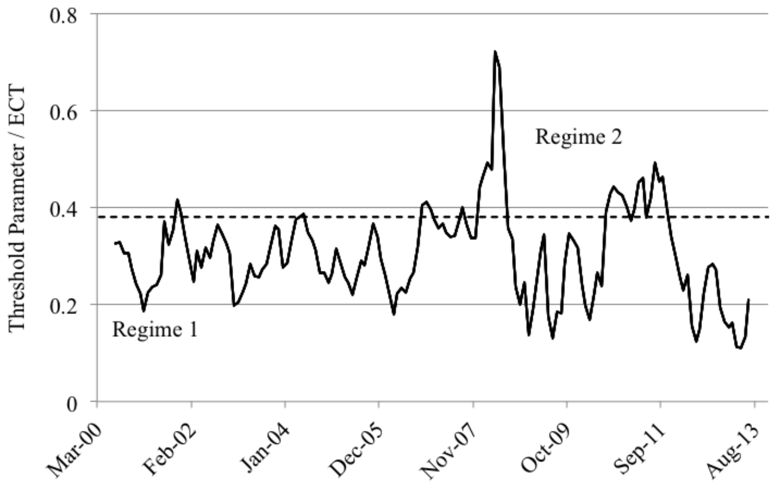

Figure 2. Timing of Threshold Adjustment in the Vietnam-Thailand Price Pair

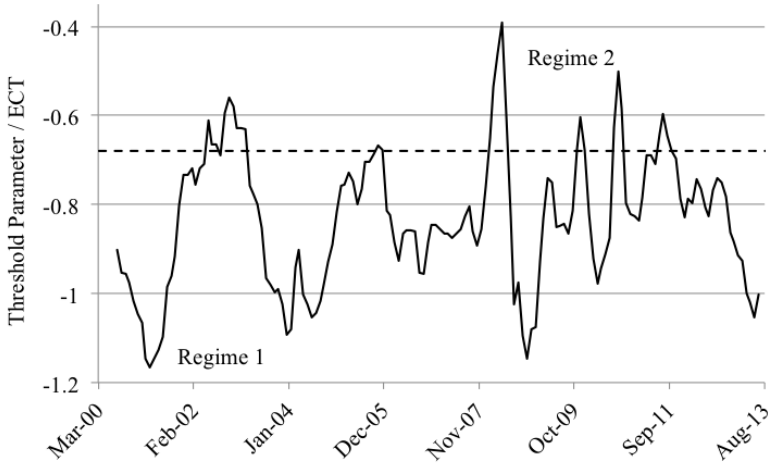

Figure 3. Timing of Threshold Adjustment in the Vietnam-United States Price Pair

the price adjustment process is quite similar to that modeled in VECM (coefficient of adjustment speed -0.16). However, in Regime 2, Vietnamese prices move to the long-run equilibrium at a much faster speed (coefficient of adjustment speed -0.50). The drastic difference in adjustment speeds supports the threshold adjustment toward equilibrium. Importantly, Vietnamese price drops at a faster rate when it is relatively high compared to its equilibrium level. This finding suggests Vietnamese exporters undercut Thai price to compete with Thai exporters for market share. Besides market competition, the quality differences of the rice may contribute to the threshold price adjustment between Thailand and Vietnam. Even though comparable price data have been used in the analysis, the quality differences cannot be entirely eliminated. Thai rice enjoys a quality premium over Vietnamese rice; to maintain market share it might be reasonable for Vietnamese exporters to decrease their price faster in response to a decrease in Thai price.

The results from the Vietnam-U.S. pair are shown in table 5. The *SupLM* test statistics reject linear price adjustment, and there is possible threshold adjustment; both VECM and TVECM are estimated. The estimated cointegration vector shows that a 1% price increase in U.S. rice prices leads to a 1.13% increase in Vietnamese prices, revealing an almost one-to-one price transmission.

Table 5. Linear and Threshold VECMs: Vietnam and U.S.

	Linear VECM		Threshold VECM			
			Regime 1		Regime 1	
Observations (%)	–		83.8		16.2	
	$\Delta p_{vn,t}$	$\Delta p_{us,t}$	$\Delta p_{vn,t}$	$\Delta p_{us,t}$	$\Delta p_{vn,t}$	$\Delta p_{us,t}$
	(0.037)	(0.023)	(0.037)	(0.019)	(0.448)	(0.265)
Intercept	0.003	0.002	-0.059^{*}	-0.003	-0.485^{*}	-0.008
	(0.004)	(0.003)	(0.032)	(0.018)	(0.280)	(0.166)
$\Delta p_{vn,t-1}$	0.549^{***}	0.178^{***}	0.552	0.114^{**}	0.900^{***}	0.057
	(0.122)	(0.063)	(0.103)	(0.051)	(0.154)	(0.120)
$\Delta p_{us,t-1}$	0.061	0.360^{***}	-0.022	0.363^{***}	0.028	0.587^{***}
	(0.612)	(0.099)	(0.110)	(0.087)	(0.246)	(0.225)
Cointegration Vector			$(1, -1.139)$			
ADF test on deviation from long-run equilibrium			-4.801^{***}			
SupLM test			20.949^{**}			
Threshold Parameter v			-0.68			

LB-Q Test on Disturbances	VECM		TVECM	
	$\Delta p_{vn,t}$	$\Delta p_{us,t}$	$\Delta p_{vn,t}$	$\Delta p_{us,t}$
Lag 6	7.005	6.293	8.856	5.879
	[0.320]	[0.391]	[0.182]	[0.437]
Lag 12	10.389	11.976	10.377	11.024
	[0.582]	[0.448]	[0.583]	[0.527]

Notes: Robust standard error in parenthesis for estimated coefficients. Single, double, and triple asterisks (*, **, ***) indicate statistical significance at the 10%, 5%, and 1% level. The number of bootstrap replications for SupLM test is 1,000, and the critical values for 1% and 5% significance levels are 21.642 and 18.470, respectively. The null hypothesis of the LB-Q test is that the disturbance series is a white noise process. p-values are in square brackets.

The linear VECM results in table 5 show both U.S. and Vietnamese prices adjust significantly to the long-run equilibrium over the study period. In the TVECM, however, the study period is divided into two regimes, as indicated above (figure 3). In Regime 1, when Vietnamese price is relatively low ($p_{vn} - 1.139 \times p_{us} \leq -0.68$), the Vietnamese price adjustment speed (-0.072) is comparable to the adjustment speed in the linear VECM (-0.090). However, in Regime 2 ($p_{vn} - 1.139 \times p_{us} > -0.68$), adjustment speed is substantially faster (-0.753). Similar to the Vietnam-Thailand pair, the Vietnamese price decreases at a higher speed when it is relatively high, indicating competition rather than collusion.

Like Thailand, U.S. rice is also of higher quality than Vietnamese rice, which could give U.S. exporters a premium on the price, making Vietnam less competitive in international markets where high-quality rice is demanded. This offers further explanations on the U.S. price leading Vietnamese price in the international markets and the threshold adjustment shown in the TVECM.

Concluding Remarks

We examined market integration and price transmission in the international rice export market using rice price data from the three major rice-exporting countries: Thailand, Vietnam, and the United States. The results indicate that the international rice export market is well integrated, despite market thinness, high exporter concentration, and possible government intervention. The Thai price leads U.S. and Vietnamese prices, while Vietnam follows Thai and U.S. price changes. Further, U.S. prices adjust to the Thai price at a constant speed, while the Vietnamese price adjusts faster to long-run equilibrium when it is above the equilibrium with Thai and U.S. prices.

These results not only corroborate market integration findings but also suggest price competition rather than price manipulation in the world rice market. The results are largely consistent with Yavapolkul, Gopinath, and Gulati (2006) and Ghoshray (2008), though we studied a more recent period. Ghoshray (2008) divided the series into two different adjustment regimes based on whether the changes were positive or negative. In this study, however, the relative magnitude of deviation from the long-run equlibrium and the estimated threshold parameter were the criteria for the different adjustment regimes. This confirms market structure stability in international rice markets.

These results have important implications for rice importers and exporters. For rice importing countries, international rice markets are still a reliable source to meet domestic consumption needs. Due to the existence of strong competition among exporters, the market is integrated and the pricing mechanism is efficient. Moreover, the Thai rice price-pledging program in 2011 serves as a demonstration that implementing policies in an attempt to manipulate international prices may be futile. The Thai government purchased rice from local farmers at a price much higher than the market price and kept large stocks of rice in warehouses. The government increased prices by limiting the quantity sold to international markets and covered the high costs of storage by charging a high price to exporters, driving up export prices. However, not long after implementing this policy, in September 2011, Vietnam started undercutting the Thai price, and India also greatly increased its rice exports, leading rice-importing countries to substitute for the cheaper rice (Kedmey, 2013). Hence the Thai program failed, leaving the government to incur high storage and transportation costs with large stocks of rice. Furthermore, the Vietnamese price dropped rapidly when Thailand and the United States reduced their prices.

Testing for asymmetric price adjustment and market integration are only the first steps for a complete understanding of world rice markets. Due to data limitations, other major exporters like Pakistan, India, and China were not included in this analysis. A more accurate picture of world rice markets could be revealed by collecting and using price as well as quantity data for all major exporters. Also, an array of other fundamental factors could affect global prices. For example, rising fuel and fertilizer prices, a slow yield growth during the past decade, and adverse weather conditions in some growing areas could have contributed to price increases during the global food crisis in 2007 (Childs and Kiawu, 2009). Moreover, the role of various trade instruments, state-owned enterprises, and rice-quality differences on market integration warrants further research.

[Received July 2015; final revision received June 2016.]

References

Anderson, J. D., D. Hudson, A. Harri, and S. C. Turner. "A New Taxonomy of Thin Markets." 2007. Paper presented at the Annual Meeting of the Southern Agricultural Economics Association, February 4-7, Mobile, Alabama.

Andrews, D. W. K. "Tests for Parameter Instability and Structural Change with Unknown Change Point." *Econometrica* 61(1993):821–856. doi: 10.2307/2951764.

Ardeni, P. G. "Does the Law of One Price Really Hold for Commodity Prices?" *American Journal of Agricultural Economics* 71(1989):661–669. doi: 10.2307/1242021.

Asche, F., H. Bremnes, and C. R. Wessells. "Product Aggregation, Market Integration, and Relationships between Prices: An Application to World Salmon Markets." *American Journal of Agricultural Economics* 81(1999):568–581. doi: 10.2307/1244016.

Baldwin, K., N. Childs, J. Dyck, and J. Hansen. "Southeast Asia's Rice Surplus." Outlook Report RCS-121-01, U.S. Department of Agriculture, Economic Research Service, Washington, DC, 2012. Available online at http://www.ers.usda.gov/publications/rcs-rice-outlook/rcs-121-01.aspx.

Barrett, C. B. "Measuring Integration and Efficiency in International Agricultural Markets." *Review of Agricultural Economics* 23(2001):19–32. doi: 10.1111/1058-7195.00043.

Barrett, C. B., J.-R. Li, and D. Bailey. "Factor and Product Market Tradability and Equilibrium In Pacific Rim Pork Industries." *Journal of Agricultural and Resource Economics* 25(2000):68–87.

Childs, N. "Rice Outlook: May 2012." Outlook Report RCS-12e, U.S. Department of Agriculture, Economic Research Service, Washington, DC, 2012. Available online at http://www.ers.usda.gov/publications/rcs-rice-outlook/rcs-12e.aspx.

Childs, N., and B. Chambers. "Rice Situation and Outlook Yearbook." Outlook Report RCS-2000, U.S. Department of Agriculture, Economic Research Service, Washington, DC, 2000. Available online at http://usda.mannlib.cornell.edu/usda/ers/RCS-yearbook/2000s/2000/RCS-yearbook-11-29-2000.pdf.

Childs, N., and J. Kiawu. "Factors behind the Rise in Global Rice Prices in 2008." Outlook Report RCS-09D-01, U.S. Department of Agriculture, Economic Research Service, Washington, DC, 2009. Available online at http://www.ers.usda.gov/publications/rcs-rice-outlook/rcs-09d01.aspx.

Dickey, D. A., and W. A. Fuller. "Distribution of the Estimators for Autoregressive Time Series with a Unit Root." *Journal of the American Statistical Association* 74(1979):427–431. doi: 10.1080/01621459.1979.10482531.

Enders, W., and P. Siklos. "Cointegration and Threshold Adjustment." *Journal of Business & Economic Statistics* 19(2001):166–76.

Esposti, R., and G. Listorti. "Agricultural Price Transmission across Space and Commodities during Price Bubbles." *Agricultural Economics* 44(2013):125–139. doi: 10.1111/j.1574-0862.2012.00636.x.

Fulton, M. E., and T. Reynolds. "The Political Economy of Food Price Volatility: The Case of Vietnam and Rice." *American Journal of Agricultural Economics* 97(2015):1206–1226. doi: 10.1093/ajae/aav019.

Ghoshray, A. "Asymmetric Adjustment of Rice Export Prices: The Case of Thailand and Vietnam." *International Journal of Applied Economics* 5(2008):80–91.

Goletti, F. "Market Liberalization and Integration of Maize Markets in Malawi." *Agricultural Economics* 11(1994):311–324. doi: 10.1016/0169-5150(94)00005-0.

Goodwin, B. K. "Multivariate Cointegration Tests and the Law of One Price in International Wheat Markets." *Review of Agricultural Economics* 14(1992):117. doi: 10.2307/1349612.

Goodwin, B. K., and P. Vavra. "What Can We Learn from Spatial and Vertical Price Transmission Studies? Empirical Examples from U.S. Meat Markets." 2009. Paper presented at the Courant Research Centre "Poverty, Equity and Growth" Inaugural Conference, University of Göttingen, Germany, July. Available online at http://www4.ncsu.edu/∼bkgoodwi/papers/goodwin_gottingen.pdf.

Hansen, B. E., and B. Seo. "Testing for Two-Regime Threshold Cointegration in Vector Error-Correction Models." *Journal of Econometrics* 110(2002):293–318. doi: 10.1016/S0304-4076(02)00097-0.

Headey, D., and S. Fan. "Anatomy of a Crisis: The Causes and Consequences of Surging Food Prices." *Agricultural Economics* 39(2008):375–391. doi: 10.1111/j.1574-0862.2008.00345.x.

Johansen, S. "Statistical Analysis of Cointegration Vectors." *Journal of Economic Dynamics and Control* 12(1988):231–254. doi: 10.1016/0165-1889(88)90041-3.

Johansen, S., and K. Juselius. "Maximum Likelihood Estimation and Inference on Cointegration - with Applications to the Demand for Money: Inference on Cointegration." *Oxford Bulletin of Economics and Statistics* 52(1990):169–210. doi: 10.1111/j.1468-0084.1990.mp52002003.x.

John, A. "Price Relations between Export and Domestic Rice Markets in Thailand." *Food Policy* 42(2013):48–57. doi: 10.1016/j.foodpol.2013.06.001.

Kedmey, D. "How Thailand's Botched Rice Scheme Blew a Big Hole in its Economy." *Time* (2013). Available online at http://world.time.com/2013/07/12/how-thailands-botched-rice-scheme-blew-a-big-hole-in-its-economy/.

Kwiatkowski, D., P. C. Phillips, P. Schmidt, and Y. Shin. "Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root: How Sure Are We that Economic Time Series Have a Unit Root?" *Journal of Econometrics* 54(1992):159–178. doi: 10.1016/0304-4076(92)90104-Y.

Mahanaseth, I., and L. W. Tauer. "Thailand's Market Power in Its Rice Export Markets." *Journal of Agricultural & Food Industrial Organization* 12(2014):109–120. doi: 10.1515/jafio-2014-0003.

Meyer, J., and S. von Cramon-Taubadel. "Asymmetric Price Transmission: A Survey." *Journal of Agricultural Economics* 55(2004):581–611. doi: 10.1111/j.1477-9552.2004.tb00116.x.

Miljkovic, D. "US and Canadian Livestock Prices: Market Integration and Trade Dependence." *Applied Economics* 41(2009):183–193. doi: 10.1080/00036840600994260.

Mohindru, S. C., and P. Phromchanya. "Rice Exporters Eye Cartel to Boost Prices." *Wall Street Journal* (2012). Available online at http://on.wsj.com/PJi2lm.

Motamed, M., K. A. Foster, and W. E. Tyner. "Applying Cointegration and Error Correction to Measure Trade Linkages: Maize Prices in the United States and Mexico." *Agricultural Economics* 39(2008):29–39. doi: 10.1111/j.1574-0862.2008.00312.x.

Ngan, P. H. "The Vietnamese Rice Industry During the Global Food Crisis." In D. Dawe, ed., *The Rice Crisis: Markets, Policies and Food Security,* Rome: FAO, 2010, 219–232. Available online at http://www.fao.org/3/a-an794e.pdf.

Poapongsakorn, N. "The Political Economy of Thai Rice Price and Export Policies in 2007–2008." In D. Dawe, ed., *The Rice Crisis: Markets, Policies and Food Security,* Rome: FAO, 2010, 191–217. Available online at http://www.fao.org/3/a-an794e.pdf.

Slayton, T. "Rice Crisis Forensics: How Asian Governments Carelessly Set the World Rice Market on Fire." Working Paper 163, Center for Global Development, Washington, DC, 2009. Available online at http://www.cgdev.org/publication/rice-crisis-forensics-how-asian-governments-carelessly-set-world-rice-market-fire.

Stephens, E. C., E. Mabaya, S. v. Cramon-Taubadel, and C. B. Barrett. "Spatial Price Adjustment with and without Trade*: Spatial price adjustment and trade." *Oxford Bulletin of Economics and Statistics* 74(2012):453–469. doi: 10.1111/j.1468-0084.2011.00651.x.

Stigler, M. "Threshold Cointegration: Overview and Implementation in R." 2010. Available online at https://cran.r-project.org/web/packages/tsDyn/vignettes/ThCointOverview.pdf.

U.S. Department of Agriculture, Economic Research Service. *Rice Yearbook*. Washington, DC: U.S. Department of Agriculture, Economic Research Service, 2014. Available online at http://www.ers.usda.gov/data-products/rice-yearbook.aspx.

Yang, J., D. Bessler, and D. J. Leatham. "The Law of One Price: Developed and Developing Country Market Integration." *Journal of Agricultural and Applied Economics* 32(2000):429–440.

Yavapolkul, N., M. Gopinath, and A. Gulati. "Post-Uruguay Round Price Linkages between Developed and Developing Countries: The Case of Rice and Wheat Markets." *Agricultural Economics* 34(2006):259–272. doi: 10.1111/j.1574-0864.2006.00123.x.