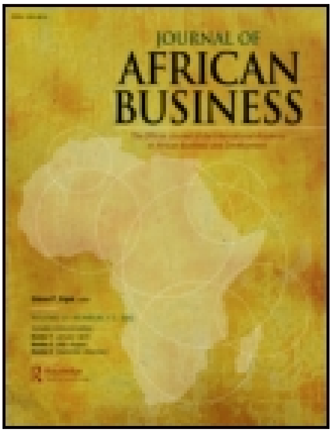

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ABSTRACT

The integration of West African rice market to the world market is assessed in order to derive the implication for food security. To this end, the transmission of rice price changes on the world market to selected markets in West Africa was examined to test for the presence of transaction costs. Using the two-regime threshold cointegration procedure on monthly price data, evidence in support of the hypothesis of asymmetric price transmission was found between Thailand and some West African markets. Price increases on the world market were more quickly transmitted to domestic price than were price decreases in Benin and Mali, suggesting short-run dynamic inefficiencies and the presence of transaction costs. In Senegal, the adjustment was linear, suggesting greater integration with the world rice market. The results suggest that West African governments should design and implement adequate policies to develop the domestic rice sector, improve market infrastructures in order to reduce their country dependency to international markets and ensure food security.

KEYWORDS
rice price transmission; threshold cointegration; transaction costs; West Africa

Introduction

Between 2007 and 2008, the international prices of rice and other major food commodities reached levels never seen before (Luckmann, Ihle, Kleinwechter, & Grethe, 2015). To a large extent, these abrupt increases in world prices have been transmitted to the domestic prices in most countries, particularly those that depend heavily on imports of these commodities. The exposure of Sub-Saharan Africa to rice market shocks became manifest when soaring international prices pushed the domestic price so high as to trigger violent riots across many countries. In West Africa, these riots affected Burkina Faso, Côte d'Ivoire, Guinea, Mauritania, and Senegal (Seck, Tollens, Wopereis, Diagne, & Bamba, 2010). The consequences of the crisis were further aggravated by the inherent instability of the international rice market and the distortive interventions by governments in both importing and exporting countries to insulate their domestic markets from the unstable international market and stabilize domestic price (Calpe, 2003;

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Wailes, 2004; Mendez del Villar, 2006; Fiamohe, Bamba, Seck, & Diagne, 2012). Such public interventions and panicked hoarding decisions by governments, traders, and millions of households and farmers have significant ripple effects on the international market price of rice and consequently on poverty and food insecurity (Timmer, 2008).

The crisis also highlighted the importance of rice for West African countries, which depend largely on imports to fill the growing gap between domestic demand and domestic production, and revealed that affordable access is critical for both social and political stability (Seck et al., 2010; AfricaRice, 2012). The frequent swings and surges in the international price of rice and their consequences in West African countries prompt the question of countries' integration into the world market and most importantly the functioning of their domestic rice markets. This study aims to understand the extent and speed of the transmission of international rice prices to the local markets of selected countries (Benin, Mali, and Senegal) in West Africa. The selection of countries was motivated by data availability and the need for representativeness. In all three countries included in the analysis, as in most West African countries, rice is an important staple food, the consumption of which is growing rapidly. Rice consumption per capita and per year over the period (2000–2011) exceeded 83kg in Benin, 86kg in Mali, and 104kg in Senegal. Unlike most other staple foods widely consumed in the region (maize, sorghum, millet, cassava, and bean), rice has the particularity of being imported in huge quantities from Asia (mostly from Thailand and to some extent from Vietnam, Pakistan, and India) and the USA.

There is a large literature examining the relation between international rice price and domestic price in developing countries. The theoretical basis of price transmission analysis is the Law of One Price (LOP) theory, which states that a homogeneous good in geographically separated markets should sell for identical prices after adjusting for transportation cost and exchange rate (Burdett & Judd, 1983; Ardeni, 1989; Lamont & Thaler, 2003). If this holds, the markets are said to be perfectly integrated and changes in the price on one market should perfectly transmit to the price on the other. Several studies have tested the LOP for different products and markets with mixed results. However, only a few studies have focused on developing countries and particularly on rice. Our research seeks to contribute to this literature by testing to what extent West African countries' rice markets are integrated with the world rice market and assess to what extent and how fast changes in the international price of rice are transmitted to the price of imported rice on these markets. This study also helps to shed further light on the functioning of West African and world rice markets by empirically examining whether there is asymmetric adjustment in the transmission of price shocks from world market to domestic market. The rest of the paper is organized as follows. We first review the literature related to our research and highlights the main contributions to the literature. We then present the empirical methodology used in the analysis. We describe the selected markets and analyze price series data. We present and discuss the main findings. Finally, we present some concluding remarks and policy implications.

Background: Price transmission and rice markets

The literature on spatial or horizontal price transmission is large and still growing, particularly in the aftermath of the 2008 food-price crisis (Fackler & Goodwin, 2001;

Greb, Jamora, Mengel, von Cramon-Taubadel, & Würriehausen, 2012). The theoretical basis of this whole literature is the Law of One Price (LOP). Many studies have tested the LOP and assessed it for various products and markets. There has been little evidence in support of the LOP and the growing consensus is that many factors can affect the transmission of price signal from one market to another even under perfect competition (Pippenger & Phillips, 2008). Some of the most commonly cited factors include the presence of agents with market power, distortive border and domestic policies, and product homogeneity and differentiation.

Analyses focusing on price transmission have referred to non-competitive market structures as an explanation for the failure of the LOP (Meyer & von Cramon-Taubadel, 2004; Abdulai, 2000; Pede & McKenzie, 2008; Fiamohe & Henry de Frahan, 2012). In all major rice exporting and importing countries, state trading companies and government-to-government contracts play active roles. These state actors have substantial market power on the supply side and their interventions introduce some distortions on international prices and consequently disturb the perfect transmission of price signal from the international market to the domestic market. For instance, in Thailand (the major rice exporter and supplier to West African countries), the Thai rice exporters' association works closely with the Thai Government to maintain the country's lead in rice exports. As one of the top rice-exporting countries, Thailand and its exporting firms possess a high degree of market power (Warr, 2001). On the demand side, in most West African countries only a small number of relatively large firms tend to dominate import and wholesale of rice (Baris, Zaslavsky, & Perrin, 2005; PAM, 2008; USAID, 2009; Fiamohe et al., 2012). Unlike their counterparts in exporting countries, firms in West Africa are not well organized and go to the world market individually with little bargaining power. Also, the countries in the region do not coordinate their imports and there is no regional instrument in the Economic Community of West African States (ECOWAS) for regulating rice imports (in both quantity and quality). While governments play little to no direct role in rice import, the sector is not competitive. Acquisition of rice import licenses is not always fair due to corruption and poor governance. Given the level of government intervention in global rice markets and in domestic markets in West Africa, it could be hypothesized that the responses of domestic market prices in importing countries in West Africa to changes in the prices in exporting countries are far from perfect.

Product homogeneity and differentiation are also an important sources of failure of the LOP. In the case of rice in West African countries, there is the possibility of mixing of type of rice imported before sale to domestic consumers. For instance, there are three major types of Thai rice generally imported (Thai 25%, Thai 100, and Thai A1 Super). These types of rice differ depending of their broken rice content and quality. However, at the retail level, traders sometimes mix the rice types and set a unique price. Even when rice types are not mixed, their prices are willfully interlinked. This creates a non-homogeneity such that the product sold is not necessary the same as the product on the international market or from Thailand.

Failure of the LOP due to any of these factors causes asymmetries in the price-transmission mechanism and a sluggish adjustment between the price on the international market and the retail price on domestic markets (Vavra & Goodwin, 2005). Existence of asymmetric price transmission can be a manifestation of market failure, which induces an imperfect pass-through of price between exporting price on the international market and retail price on the domestic market (Vavra & Goodwin, 2005). The domestic rice price responds asymmetrically and nonlinearly to price change on the exporting market. It could also be expected that response to price increases may be different from response to price decreases, or the speed of the transmission faster for the former than for the later. Generally, it is assumed that asymmetry in price transmission between markets occurs when transaction costs are high (Balke & Fomby, 1997; Balcombe, Bailey, & Brooks, 2007).

Since the seminal work of Ardeni (1989), cointegration techniques have become the main tool to analyze price co-movement and market integration (Baffes, 1991; Pippenger & Phillips, 2008). In the presence of anomalies causing asymmetric price transmission, traditional cointegration techniques using Engel and Grander or Johansen methods are no longer appropriate. Goodwin & Piggott (2001) point out that ignoring the presence and the extent of transaction costs between two markets may affect the results of price-transmission analysis and inference made concerning market integration. In such a case, prices between markets could deviate and show no tendency to revert back to a long-run equilibrium unless price differences exceed some transaction cost level (Pede & McKenzie, 2008). From a welfare perspective, Meyer & von Cramon-Taubadel (2004) argue that, in the presence of asymmetric price transmission, consumers in importing countries might not benefit from a price reduction on the international market that would have taken place under conditions of symmetric price transmission (Balke & Fomby, 1997).

Several econometric methods have been developed to account for the possibility of asymmetry in price transmission between two markets. Under the assumption of asymmetric price transmission, the data-generation process of the price series can be represented by a nonlinear specification. Balke & Fomby (1997) introduced the idea of threshold cointegration as a feasible means to combine nonlinearity and cointegration, and account for asymmetry in the relation between two series and particularly the presence of transaction costs between market pairings. A fundamental assumption in their approach is that the cointegrating vector and also the threshold is exogenously given and known. Hansen & Seo (2002) extend Balke & Fomby's approach to account for the case of an unknown cointegrating vector and threshold. We adopt this last approach to examine the ability of selected West African imported-rice markets to respond to pricing signals in the presence of transaction costs in rice markets. The threshold autoregressive presumes that a significant response to a shock is triggered only after a deviation from the long-run equilibrium induced by the shock exceeding a certain threshold. This is precisely the main contribution of our research, as it helps to shed further light on the functioning and efficiency of selected West African markets by empirically examining whether there is asymmetric adjustment in the transmission of price shocks from the world market to domestic markets.

Methodology

The empirical method used to analyze the transmission of rice price from Thailand, the world's top rice exporter, to the selected West African countries and test for the presence of transaction costs follows the two-regime threshold cointegration procedure developed by Hansen & Seo (2002). The first step in this approach consists of examining the time-series properties of the different series by testing for the presence of unit root in the different variables. To this end, we use the augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP) tests. Second, the existence of a linear cointegrating relationship between Thailand price and domestic price of imported rice in the selected West African countries is tested using both Engle-Granger (Engle & Granger, 1987) and Johansen (1991) methods. The Engle and Granger's test consists in estimating using ordinary least square the relation between the two prices and test the stationarity of the residuals. If the residuals are stationary, the two series are cointegrated. The error correction model is estimated in a second step using the variables in difference and the lag residuals as an additional explanatory variable. The Johansen's test generalizes this approach by allowing a one-step estimation of cointegrating relation and the error correction model. Two statistics are commonly use in the Johansen's test: the trace statistics and the eigen value statistics. There is cointegration if the estimated value of these statistics exceed the critical theoretical value for a given level of confidence. If evidence of linear cointegration is found, it implies that the markets are integrated with the world market, but we need to test whether the transmission of price signal from Thailand to West Africa is symmetric or not.

We follow the exposition of Hansen & Seo (2002) to present the general setting of the two-regime threshold cointegration. Let us consider a vector x_t of P time series all integrated of order I (1) and assume that the series are linearly cointegrated with one unknown cointegrating vector β. Thus, the stationarity I (0) error correction term is $w_t(\beta) = \beta' x_t$ and the linear vector error correction model of order $l + 1$ can be written:

$$\Delta x_t = A' X_{t-1}(\beta) + u_t \tag{1}$$

$$\text{where } X_{t-1}(\beta) = \begin{pmatrix} 1 \\ w_{t-1}(\beta) \\ \Delta x_{t-1} \\ \Delta x_{t-2} \\ \vdots \\ \Delta x_{t-l} \end{pmatrix}, \tag{2}$$

$X_{t-1}(\beta)$ are the regressors and u_t the error assumed to be a vector martingale difference sequence with finite covariance matrix $\sum = E(u_t u_t^{'})$.

The two-regime extension of equation (1) takes the form:

$$\Delta x_t = \begin{cases} A_1^{'}\ X_{t-1}(\beta) + u_t \ if \ w_{t-1}(\beta) \le \gamma \\ A_2^{'}\ X_{t-1}(\beta) + u_t \ if \ w_{t-1}(\beta) > \gamma \end{cases} \tag{3}$$

Or alternatively,

$$\Delta x_t = A_1^{'} X_{t-1}(\beta) d_{1t}(\beta,\gamma) + A_2^{'} X_{t-1}(\beta) d_{2t}(\beta,\gamma) + u_t \tag{4}$$

where $d_{1t}(\beta,\gamma) = \mathbb{1}(w_{t-1}(\beta) \le \gamma)$ and $d_{2t}(\beta,\gamma) = \mathbb{1}(w_{t-1}(\beta) > \gamma)$ and $\mathbb{1}$ (.) is the indicator function with the value 1 if the condition in parenthesis is satisfied and the value 0 otherwise; γ is the unknown threshold that defines the two regimes, and A_1 and A_2 determine the dynamics in the first and second regimes, respectively. For the threshold model to be valid it is necessary that none of the two regimes could be empty. That is, $0 < P(w_{t-1}(\beta) \leq \gamma) < 1$. In practice, it is useful to impose a minimum observation per regime by imposing $\pi_0 < P(w_{t-1}(\beta) \leq \gamma) < 1 - \pi_0$, where π_0 is the trimming parameter. Following Hansen and Seo (2002), we set $\pi_0 = 0.05$.

Assuming that the error is Gaussian *iid*, Hansen & Seo (2002) propose the estimation of the equation by the maximum likelihood method. The Gaussian likelihood function is given as:

$$\mathcal{L}_n(A_1, A_2, \Sigma, \beta, \gamma) = -\frac{n}{2}\log\left|\sum\right| - \frac{1}{2}\sum\nolimits_{t=1}^{n}\mu_t(A_1, A_2, \Sigma, \beta, \gamma)'\Sigma^{-1}\mu_t(A_1, A_2, \Sigma, \beta, \gamma) \tag{5}$$

where

$$\mu_t(A_1, A_2, \beta, \gamma) = \Delta x_t - A_1^{'}X_{t-1}(\beta)d_{1t}(\beta, \gamma) - A_2^{'}X_{t-1}(\beta)d_{2t}(\beta, \gamma) \tag{6}$$

The maximum likelihood estimates (MLE) $\left(\widehat{A_1}, \widehat{A_2}, \hat{\beta}, \hat{\gamma}\right)$ are obtained by maximizing$(A_1,\ A_2, \ ,\ \beta,\ \gamma)$. Hansen & Seo (2002) show that concentrating out (A_1, A_2, Σ) by holding (β, γ) and maximizing the constrained MLE yield the ordinary least squares (OLS) estimators for (A_1, A_2, Σ):

$$\hat{A}_1(\beta, \gamma) = \left(\sum\nolimits_{t=1}^{n}(X_{t-1}(\beta)X_{t-1}(\beta)'d_{1t}(\beta, \gamma)\right)'\left(\sum\nolimits_{t=1}^{n}(X_{t-1}(\beta)\Delta x_t' d_{1t}(\beta, \gamma)\right) \tag{7}$$

$$\hat{A}_2(\beta, \gamma) = \left(\sum\nolimits_{t=1}^{n}(X_{t-1}(\beta)X_{t-1}(\beta)'d_{2t}(\beta, \gamma)\right)'\left(\sum\nolimits_{t=1}^{n}(X_{t-1}(\beta)\Delta x_t' d_{2t}(\beta, \gamma)\right) \tag{8}$$

$$\hat{\mu}_t(\beta, \gamma) = \mu_t\left(\hat{A}_1(\beta, \gamma), \hat{A}_2(\beta, \gamma), \beta, \gamma\right) \tag{9}$$

and

$$\hat{\Sigma}(\beta, \gamma) = \frac{1}{n}\sum\nolimits_{t=1}^{n}\hat{\mu}_t(\beta, \gamma)\hat{\mu}_t(\beta, \gamma)' \tag{10}$$

Next, the MLE of β and γ are obtained by minimizing the concentrated MLE:

$$\mathcal{L}_n(\beta, \gamma) = \mathcal{L}_n\left(\widehat{A_1}(\beta, \gamma)\right), \widehat{A_2}(\beta, \gamma), \widehat{\Sigma}(\beta, \gamma), \beta, \gamma) = -\frac{n}{2}log|\widehat{\Sigma}(\beta, \gamma)| - \frac{np}{2} \tag{11}$$

For this estimation, Hansen and Seo (2002) argue that the objective function is not smooth, thus not differentiable, and it is not possible to find an explicit expression for the optimization problem. They suggest a grid-search method, over all pairs (β, γ) to overcome this non-smoothness using a sequential approach which is described as follows. In the first step, we estimate the linear error-correction model and construct a large interval $[\beta_L, \beta_U]$ within which the optimal value of β is searched. Next, the support of the long-run residuals $(\widehat{w}_{t-1}\left(\hat{\beta}\right))$ is trimmed to form an interval $[\gamma_L, \gamma_U]$ within which the optimal value for γ is searched. For each value of (β, γ), we estimate $\widehat{A_1}(\beta, \gamma)$, $\widehat{A_2}(\beta, \gamma)$, $\widehat{\Sigma}(\beta, \gamma)$, and $log|\widehat{\Sigma}(\beta, \gamma)|$. Finally, $\left(\hat{\beta}, \hat{\gamma}\right)$ are chosen to minimize

$log|\widehat{\Sigma}(\beta,\gamma)|$ and the corresponding estimates of all other parameters are derived as $\hat{\Sigma}=\widehat{\Sigma}(\beta,\gamma)$, $\hat{A}_1=\widehat{A_1}(\beta,\gamma)$, and $\widehat{A}_2=\widehat{A_2}(\beta,\gamma)$.

The test for the threshold effect consists of the null hypothesis linear cointegration against the alternative of threshold cointegration and formulated as (H_0) is: $A_1 = A_2$. Hansen and Seo (2002) propose a supremum test statistic. Due to the difficulty of determining the exact asymptotic properties of supremum test statistics, it is common to use a bootstrap procedure to obtain the p-value of the test. Under the hypothesis of Gaussian errors, the estimates $\hat{A}_1$ and $\hat{A}_2$ have normal asymptotic distribution and conventional standard errors are still valid (Hansen & Seo, 2002).

In our empirical application, the vector x_t is formed by a price of a given type of rice on the international market and the domestic price of imported rice in one of the selected West African markets. More specially, we seek to estimate the transmission from the international price PI_t to the price of imported rice PL_t on the domestic markets of the selected countries. With three international prices and three countries, we end up examining nine price pairings. We use the R package tsDyn implemented by Di Narzo, Aznare, and Stigler (2013) for the econometric analysis.

Data and descriptive analysis

The selection of West African countries to include in our analysis was mainly motivated by the availability of long price series. The choice of Thailand graded rice prices to approximate the international price is motivated by the dominant position held by Thailand on the rice export market and as the main supplier of West African countries, at least for the time period considered (see below), which is before Thailand lost its world-leading rice-export position in the wake of its paddy-pledging program policy of 2011.

In the application of the threshold cointegration methodology, the domestic retail prices per kilogram of imported rice in the commercial capital of each of the three selected countries are paired with the international price of different types of Thai rice. The cities considered are Cotonou for Benin, Bamako for Mali, and Dakar for Senegal. We focus on these cities because they constitute the main entry points of imported rice and the largest consumer markets. So, it could be expected that imported rice price in the coastal cities is more directly linked to the international market than are other domestic markets. The price on other domestic markets may reflect many internal factors other than the transmission from the international market. Data for Benin were provided by the Office national d'appui à la sécurité alimentaire (ONASA). Prices for Mali were obtained from the Observatoire des marchés agricoles (OMA), and the Commissariat à la sécurité alimentaire (CSA) provided the data for Senegal. All the West African prices were converted to U.S. dollars using nominal exchange rate from USDA (2013).

Three marked grades of Thai rice are considered: Thai 25% broken rice, Thai 100 (with B grade ≥60% whole kernels), and Thai A1 Super (100% broken rice). These are the most commonly imported Thai rice types in West African countries. However, it is important to stress that the types of imported rice differ greatly among countries (Fiamohe et al., 2012). In Senegal, Thai A1 Super (100% broken rice) is the most imported rice because consumers have strong preference for broken rice (USAID, 2009). In Benin and Mali, consumers have strong preference for whole rice, so these countries import more Thai 25% and Thai 100. The Thai rice price in US dollars was obtained from OSIRIZ and Centre de coopération internationale en recherche agronomique pour le développement (CIRAD).

The data used were monthly and covered the period January 2000 to December 2011. All data on rice price and exchange rates were expressed in their nominal terms. As suggested by Hanawa-Peterson and Tomek (2000), deflated series were not used to avoid altering the time-series properties of the original series. Furthermore, following Alderman (1993), Abdulai (2000), and Fiamohe, Seck, Alia, and Diagne (2013), we used the moving average method to systematically filter out the seasonal-adjusted component of all price series in order to capture the intrinsic relation between the prices of the different markets.

Figure 1 presents the evolution of the price series over time. It shows evidence of co-movement between rice prices in the three selected West African markets (Benin, Mali, and Senegal) and those in Thailand for the selected market grades of rice (Thai 25% broken rice, Thai 100, and Thai A1 Super). The plots also highlight large fluctuations in the prices on the selected markets. Historically, rice prices have an upward overall

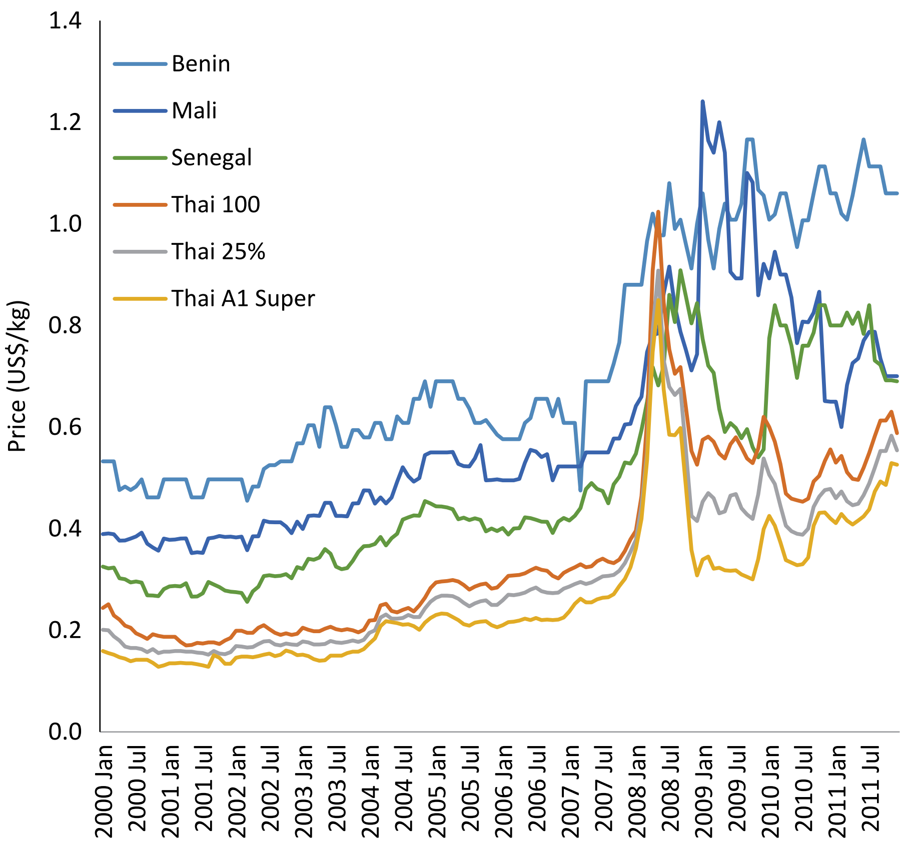

Figure 1. Rice price trends in Benin, Mali, Senegal, and Thailand markets (US$ kg^{-1}).

trend. Since 2000, Thai graded rice international prices have increased and consequently domestic prices in West African countries have kept rising. The graphs highlight the price surges of 2008, when price more than doubled within a few months. It can be seen from the graph that domestic price also responded to change of international price, but not simultaneously. Also, the magnitude of domestic price response seems to depend on the country, the type of Thai rice, and when international prices have increased or decreased.

RESULTS

Stationarity tests and test of linear cointegration

First, the standard tests for stationarity were conducted on all series in level and in difference. The model with intercept and no trend appeared to best fit the data. Table 1 presents the results of the Augmented Dickey-Fuller and Phillips-Perron tests. Both tests detected the presence of a unit root in all series in level, but no evidence of unit root in first difference. Thus, the series of the prices of imported rice in West African markets and the price of Thai rice are integrated of the order 1.

Next, the existence of linear cointegration between each of the three domestic prices and the different Thailand market prices of rice was tested using Engle-Granger and Johansen approaches. A total of nine price pairings were tested: the results are presented in Table 2.

Using the Engle–Granger cointegration test, we found that all pairs of prices were cointegrated ($p < 0.05$). The results were slightly different when using the Johansen test. The trace statistic (λ-trace) indicated the existence of a cointegrating vector between the retail price of rice in Senegal and the prices of Thai 100, Thai 25%, and Thai A1 Super rice in Thailand ($p < 0.10$). There is also evidence that the retail prices of rice in Benin and Mali were cointegrated with the export price of Thai A1 Super and Thai 100 rice, respectively ($p < 0.10$). The results are identical when using the eigen value statistics.

The lack of cointegration between the prices of a good on two separated markets may be explained by several factors. For instance, Barrett (1996) explains that failure to find cointegration between two prices may be a result that is completely consistent with market integration if transaction costs are non-stationary in general. Other factors include low volume and frequency of trade flows and the existence of trade barriers. In the absence of data on transaction costs, it is difficult to interpret a lack of cointegration between commodity prices in two markets as a non-integration of these markets. The lack of linear cointegration between the retail market prices of rice in

Table 1. Stationarity test on log of seasonally adjusted prices.

Market	ADF level	ADF difference	PP level	PP difference
Benin	−1.685	−11.458**	−2.225	−16.714**
Mali	−2.485	−12.085**	−2.508	−12.274**
Senegal	−3.131	−10.384**	−2.807	−10.484**
Thai 100	−2.844	−5.906**	−2.822	−7.929**
Thai 25%	−2.933	−5.523**	−2.666	−8.227**
Thai A1 Super	−2.862	−6.916**	−2.871	−8.154**

MacKinnon critical values: 5%: −3.45. ** $p < 0.05$.
Note: ADF = Augmented Dickey-Fuller test statistic, PP = Phillips-Perron test statistic.

Table 2. Engle-Granger and Johansen tests of linear cointegration.

	Engle–Granger test		Johansen test		
Market pairings	Test stat	Critical value	H_0	Trace statistic	Critical value
Benin–Thai 100	−3.396**	0.000	r= 0	24.837	25.872
Lag=3 months			r= 1	6.222	12.517
Benin–Thai 25%	−3.243**	0.001	r= 0	25.498	25.872
Lag=2 months			r= 1	6.101	12.518
Benin–Thai A1 Super	−3.031**	0.002	r= 0	26.286*	25.872
Lag=3 months			r= 1	5.539	12.518
Mali–Thai 100	−3.561**	0.000	r= 0	21.543*	15.495
Lag=2 months			r= 1	1.215	3.8,415
Mali–Thai 25%	−3.179**	0.001	r= 0	12.983	15.495
Lag=2 months			r= 1	0.309	3.841
Mali–Thai A1 Super	−2.608**	0.009	r= 0	13.993	15.495
Lag=3 months			r= 1	1.457	3.841
Senegal–Thai 100	−4.061**	0.000	r= 0	21.000*	15.495
Lag=3 months			r= 1	0.664	3.841
Senegal–Thai 25%	−4.041**	0.000	r= 0	23.715*	15.495
Lag=2 months			r= 1	0.435	3.841
Senegal–Thai A1 Super	−3.831**	0.000	r= 0	32.499*	15.495
Lag=4 months			r= 1	0.836	3.841

* $p < 0.1$; ** $p < 0.05$.
Note: H_0 = null hypothesis, r = number of cointegrating vector.

Benin and the prices of Thai 100 and Thai 25% rice in Thailand seems intuitive given that the high-quality rice imported by Benin is directly re-exported to Nigeria due to the lower tariff imposed on imported rice in Benin (Hashim & Meager, 1999; Cadoni & Angelucci, 2013). From a methodological perspective, in the absence of linear cointegration, the Hansen and Seo (2002) threshold cointegration procedure cannot be applied. In fact, an important input to the method is the long-run residual estimated in the presence of linear cointegration.

The Johansen test for cointegration proved to be more adequate than the Engel-Granger test, as it also tests for the existence of cointegration and jointly estimates the number of cointegrating vectors. In selecting the price pairings for which we would test for threshold cointegration, we used the results provided by the Johansen test and considered only the pairings for which it detected a linear cointegration. Thus, the remainder of the analysis focused on the pairings Benin–Thai A1 Super, Mali–Thai 100, Senegal–Thai 100, Senegal–Thai 25%, and Senegal–Thai A1 Super.

Threshold cointegration analysis

Table 3 presents the results of the test of two-regime threshold cointegration against the linear cointegration. Following the methodology developed by Hansen and Seo (2002) and their application to interest rates in the USA, we considered two versions of the bivariate model: one with the error correction coefficient β freely estimated and the second with the constraint $\beta = 1$. The grid search was conducted on 300 × 300 points. All p-values were computed using 5,000 bootstrap replications.

The results from the cointegration tests showed evidence of threshold cointegration between the retail prices of imported rice in Benin and the price of Thai A1 Super when β was freely estimated ($p < 0.05$) and when β was constrained to 1 ($p < 0.10$). There was also evidence of threshold cointegration between prices in Mali and Thai 100 for the

Table 3. Test of linear versus threshold cointegration of Hansen and Seo (2002).

Price (bivariate)		Test of Hansen & Seo (2002) β estimate Test statistic	p-value	β=1 Test statistic	p-value	Type of cointegration
Benin	Thai A1 Super	24.034	0.048**	23.669	0.058*	Threshold
Mali	Thai 100	23.436	0.051*	28.851	0.003***	Threshold
Senegal	Thai 100	18.767	0.448	31.611	0.005***	Linear
Senegal	Thai 25%	15.212	0.296	17.469	0.161	Linear
Senegal	Thai A1 Super	14.806	0.333	13.206	0.532	Linear

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$.
Note: β = cointegrating parameter.

Table 4. Estimated cointegrating vectors and thresholds.

Country	Rice	Estimated cointegrating vector	Estimated threshold
Benin	A1 Super	1. 224	−0.1782239
Mali	Thai 100	1. 123	−0.09578717
Senegal	Thai 100	0.794	No threshold
Senegal	Thai 25%	0.808	No threshold
Senegal	Thai A1 Super	0.823	No threshold

specification with β unconstrained ($p < 0.05$) and for the one with β constrained ($p < 0.01$). In the case of Senegal, there was no evidence of threshold cointegration when β was estimated for all Thai prices. When β was constrained to 1, only one evidence of threshold cointegration was detected – with Thai 100 ($p < 0.01$). Hansen and Seo (2002) argue that having β freely estimated is more realistic and therefore preferred. In summary, out of the initial nine pairings of prices, five were linearly cointegrated and among these five cases there was evidence of threshold cointegration for only two parings of prices: Benin–Thai A1 Super and Mali–Thai 100.

Table 4 shows the cointegrating vector for the price pairings with evidence of threshold cointegration along with the estimated thresholds. The cointegrating vectors were higher than unity for both Benin–A1 Super and Mali–Thai 100. This is evidence of greater and faster response of the domestic price to change in the price of imported rice from Thailand. In the long run, a 1% increase in the Thai rice price on the world market is transmitted to the domestic market prices of Benin and Mali at a factor higher than 1%. The negative sign and magnitude of the estimated thresholds provide some indication of the direction and level of transaction costs occurring in these markets. Thus, it can be inferred that importing rice from Thailand induced transaction costs of a magnitude of \$0.18 for Benin (Thai A1 Super import) and \$0.10 for Mali (Thai 100 import). In the case of Senegal, since there was no evidence against the linear cointegration model, there was no threshold to estimate. The linear cointegrating vectors were slightly less than unity, suggesting a weaker and slower response of domestic price to change in Thailand price.

Short-run dynamics of price response

In addition to knowing that there is a long-run stable linear relationship between selected grades of rice in Thailand and the retail market price of imported rice in

Table 5. Linear price responses for Senegal to change in Thailand prices.

Market pairings	Senegal–Thai 25%	Senegal–Thai 100	Senegal–Thai A1 Super
Variable	Coef.	Coef.	Coef.
ΔPL_{t-1}	0.088	0.119	0.053
	(0.079)	(0.081)	(0.078)
ΔPL_{t-2}	0.171***	0.189***	0.132*
	(0.078)	(0.081)	(0.077)
ΔPI_{t-1}	−0.018	−0.014	−0.003
	(0.064)	(0.065)	(0.058)
ΔPI_{t-2}	0.056	0.016	0.008
	(0.067)	(0.068)	(0.061)
Constant	0.001	0.001	0.001
	(0.002)	(0.002)	(0.002)
w_{t-1}	−0.144***	−0.129***	−0.157***
	(0.036)	(0.035)	(0.036)
No. observations	144	144	144

* $p < 0.1$; *** $p < 0.01$.
Standards errors are given in parenthesis ().
Note: PL = the imported rice price on the domestic market, PI = the international rice price paired to the domestic price, ΔPL and ΔPI = dependent variables, w = error-correction term.

West Africa, it is crucial to have some indication of the short-run dynamic adjustments of domestic prices to world market prices. To address this issue, we present the representation of the univariate error correction models for all cointegrated markets. For Senegal, the linear short-run price dynamics are shown in Table 5, while for Benin and Mali the results of the two-regime price adjustments are presented in Table 6. The linear error-correction effect was strongly significant for all pairings of prices in Senegal, but of relatively small magnitude. The negative sign of the error-correction effect suggests a downward adjustment. If the relation between imported rice price in Senegal and the price prevailing on the international market for the different Thai graded prices deviates upward from its long-run equilibrium, in the short run the domestic price reacts but adjusts slowly back toward the long-run equilibrium. This behavior is observed for all Thai prices with almost the same amplitude. The symmetry

Table 6. Price responses for Benin–Thailand and Mali–Thailand market pairings (2000–2011).

Market pairings	Benin–Thai A1 Super		Mali–Thai 100	
Regime	$w_{t-1} \leq -0.178$	$w_{t-1} > -0.178$	$w_{t-1} \leq -0.096$	$w_{t-1} > -0.096$
Variable	Coef.	Coef.	Coef.	Coef.
ΔPL_{t-1}	−1.104***	−0.099	−0.115	0.077
	(0.000)	(0.237)	(0.454)	(0.446)
ΔPL_{t-2}	−0.445	−0.158*	0.064	−0.147
	(0.361)	(0.035)	(0.610)	(0.191)
ΔPI_{t-1}	0.009	−0.115	−0.087	0.094
	(0.912)	(0.056)	(0.484)	(0.496)
ΔPI_{t-2}	−0.1,128	−0.053	0.004	0.293*
	(0.272)	(0.375)	(0.979)	(0.032)
Constant	0.013	0.001	0.031*	−0.003
	(0.479)	(0.699)	(0.047)	(0.288)
w_{t-1}	−0.0301	−0.026	0.064	0.015
	(0.637)	(0.152)	(0.409)	(0.771)
No. obs.	8.5%	91.5%	20.6	79.4

* $p < 0.1$; *** $p < 0.01$.
p-values are given in parentheses ().
Note. PL = the imported rice price on the domestics market, PI = the international rice price paired to the domestic price, ΔPL and ΔPI = dependent variables, w_{t-1} = error-correction term.

of the adjustment implies that Thailand price increases and decreases are transmitted at the same speed and magnitude to the domestic price in Senegal.

Table 6 presents the results of the short-run asymmetric price adjustment between the retail price in Benin and Thai A1 Super (first two columns) and between the retail price in Mali and Thai 100 (last two columns). The presence of a threshold distinguishes two separate adjustment regimes. For the pair Benin–Thai A1 Super, the first regime occurs when $w_{t-1} \leq -0.178$. This regime contains 8.5% of the data points. The second regime occurs for $w_{t-1} > -0.178$ and contains 91.5% of the observations. Following Hansen and Seo's (2002) terms, the first regime with few observations could be labeled as "extreme" and the second regime with a predominant number of observations labeled as "typical." In the case of Mali–Thai 100, the "extreme" regime occurs for $w_{t-1} \leq -0.096$, but is slightly more important (than that in Benin) with 20.6% of the observations, and the "typical" regime had 79.4%.

Despite the existence of a long-run relationship, most coefficients in the short-run equations appeared not to be significant at conventional levels. It is important to remember that the estimation method is based on a quasi-MLE algorithm. Since there is no formal distribution theory for the parameters estimated, Hansen and Seo (2002) warn that the standard errors should be interpreted cautiously. Following this warning, we put less weight on the interpretation of the significance of the parameters estimated.

In the case of Benin, the error-correction terms were negative for both regimes with an adjustment slightly higher in the "extreme" regime. Thus, above the threshold level, the price adjustment seems to be weak, confirming the effect of the transaction costs. In the case of Mali, the error-correction coefficients were (surprisingly) positive suggesting a short-run correction upward to the long-run relation between the two prices.

Our results are consistent with a large strand of the price transmission literature which seldom finds supporting evidence for the LOP. In an extensive survey of the literature on price transmission, Peltzman (2000) found that asymmetric price transmission was largely more prevalent than symmetric price transmission. Focusing on African countries, Greb et al. (2012) found that, unlike other cereal such as maize and wheat, international rice prices were more cointegrated with domestic rice price.

Conclusion

The primary purpose of this research was to examine the transmission of rice prices from the world market dominated by Thailand to domestic markets in selected West African countries and to test for the presence of transaction costs. The main contribution was the use of threshold cointegration techniques developed by Hansen and Seo (2002) to account for the possible nonlinearity and asymmetry in the transmission of price signal from Thailand to West African countries. The empirical evidence indicates that, in the long run, rice prices in Benin, Mali, and Senegal share a common linear trend with prices in Thailand for select types of rice. It appears that the changes in the world market price of rice – as represented by the export market prices of Thai Super

A1, Thai 100, and Thai 25% broken rice – are differently transmitted to the African countries. While there is evidence of co-movements between the three grades of Thai rice and price in Senegal, the evidence of non-linear co-movement was significant only for Thai A1 Super for Benin and Thai 100 for Mali.

The evidence of the existence of threshold cointegration, which could be induced by the presence of high transaction costs, was found only between the retail prices of imported rice in Benin with the price of Thai A1 Super, and between price in Mali and Thai 100. In both countries, we identified one typical regime with minimal error correction and price adjustment, and an unusual regime with fast adjustment. This would imply that price increase on the world market was more quickly transmitted to domestic price than price decrease in Benin and Mali, suggesting short-run dynamic inefficiencies and presence of transaction costs. In Senegal, the adjustment was linear, suggesting deeper integration with the world market.

The imperfection of the global rice market may be reflected in the limited integration of price between Thailand and selected West African countries. Moreover, the transmission of price shocks from the world market to the domestic markets was not instantaneous but happened with some lag. These results show that Benin and Mali were more vulnerable to international price surges, and that even decrease in international prices was not fully and quickly transmitted to domestic prices. Our results suggest that West African governments should design and implement adequate policies to develop the domestic rice sector, improve market infrastructure in order to reduce their countries' dependency on international markets, and limit "imported" food insecurity. Many studies have already shown the potential for rice development and rice self-sufficiency in the region and suggest some measures to boost the sector (AfricaRice, 2012). Among these measure are closing the yield gap between actual yield attained by smallholders under poor enabling conditions and the high potential yield of high-yield varieties developed and diffused by research centers. This requires substantial public and private investment in irrigation schemes, access to fertilizer and equipment, access to credit and adequate training, and the adoption of good pest and disease management practices and postharvest technologies. On the demand side, efforts to improve the quality of local rice compared to imported rice, and the development of the market through investments and effective marketing strategies should be pursued to restore the reputation of the locally produced African rice (Fiamohe, Nakelse, Diagne, & Seck, 2014).

The detection of high transaction costs in the functioning of the rice market in Benin and Mali calls for further studies to disentangle the source of these transaction costs and identify the most effective policies to address these inefficiencies. Another potential area for further study is the analysis of the role of external factors such as the exchange rate between the US dollar and the CFA franc in the transmission of price shock between rice exporting and importing countries.

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