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Price transmission analysis using threshold models: an application to local rice markets in Benin and Mali

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Abstract In most African countries, spatial dispersion of production and consumption often results in high transaction costs that prevent farmers from accessing markets and causes asymmetry in price transmission. The objective of this study was to provide the baseline information on local rice price transmission between paired producer and consumer markets in Benin and Mali. To achieve this, we used Enders and Siklos's threshold models on monthly price series from 2000 to 2010 to examine the nature of price transmission between selected markets in the surplus zones and the nearest important consumption markets. The results for Benin indicated that price transmission between markets in the surplus zone and the consumption markets was asymmetric, probably because of the prevalence of high transaction costs. These results showed that increases in price in the surplus-zone market were more quickly transmitted to the consumer market than decreases in price. Conversely, the results for Mali indicated symmetric price transmission between the market in the surplus zone and the consumer market, suggesting the prevalence of lower transaction costs. These results highlight the need for policies aiming to lower transaction costs observed in selected local rice markets in Benin. Specific policies, such as investment in public infrastructure, e.g. roads, could promote the vertical integration of local rice production with marketing. This would be crucial to achieving rice farmers' food security and hence their wellbeing.

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Introduction

Spatial dispersion of production and consumption leads to a complex set of trade interactions among deficit and surplus regions (Fackler and Tastan 2008). This dispersion often causes high transaction costs that prevent farmers from accessing certain markets. It also creates a large number of intermediaries (Cutts and Kirsten 2006), which slows the flow of information between markets and, consequently, increases transaction costs. In most African countries, high transaction costs exist in many market exchanges. These include high transport costs and are affected by the scarcity of marketing information (McNew and Fackler 1997; Barrett 2001; Fackler and Goodwin 2001; Barrett and Li 2002; Rapsomanikis et al. 2004). Kuiper et al. (2003) pointed out that high transaction costs hamper market integration and as a consequence not all spot markets perform equally well. This implies that with the high transaction costs, farmers in surplus regions cannot be informed on time about price changes in deficit regions.

Transaction costs pose barriers for poor smallholders to participate in some market activities undertaken by better-off small operators (Delgado 1999). This also raises the issue of farmers' food security, and hence their well-being. The importance of market participation for the economic well-being of smallholder farmers in developing countries has been widely discussed. Barrett (2008) has pointed out that the primary theme in the literature on smallholder market participation is the importance of transactions costs. Fafchamps's (1992) analytical model shows that when food markets are present but not well integrated over space, price volatility and high covariance of prices with household production limit the extent to which smallholders will adopt income and food security strategies that are more commercially oriented. Holloway et al. (2004) examined these facts empirically in Ethiopia and concluded that growth in dairying by small-holder farmers in peri-urban areas has been limited by transactions costs inherent in production and marketing of dairy products. Key et al. (2000) argued that household crop supply and welfare response to exogenous market price changes are heavily affected by transaction costs, which create important discontinuities in supply response and non-convexities commonly associated with poverty traps. Barrett and Dorosh (1996) also found that more variable rice prices induced by economic reforms in Madagascar likely imposed additional instantaneous welfare losses by threatening household food security and destabilizing incomes. Yet, better access to remunerative markets is necessary for promoting growth of smallholder agriculture (Holloway et al. 2000), hence their well-being. Moreover, the perfect integration of markets and increased market participation also holds the promise of improving aggregate food availability and efficiency of distribution in urban areas served by those smallholders. Thus, it appears clear that food security is linked to farmers' better market participation in addition to productivity increase.

It is generally assumed that asymmetry in price transmission between markets occurs where there are high transaction costs (Balke and Fomby 1997; Balcombe et al. 2007). Differential transaction costs among households stem from asymmetries in access to assets, information, services and remunerative markets (Delgado 1999). Meyer and von Cramon-Taubadel (2004) observe that a possible implication of asymmetric price transmission is that consumers are not benefiting from price reduction at the producers' level, or producers might not benefit from price increase at the retail level. Consequently, according to Baulch (1997) and Boughton et al. (2007), better market integration leads to faster transmission of price signals and encourages producers to specialize according to comparative advantage and thereby enjoy welfare gains from trade. Barrett (2008) mentioned that maintaining efficient market integration is most important in order to ensure producers benefit in the long-run from technological change. Vavra and Goodwin (2005) identified several types of asymmetry. The first type refers to the difference in the speed of transmission of positive and negative price changes from one market to another. The second type refers to the different thresholds of relative magnitude of price change in the positive or negative directions that must be reached before triggering a response. Thus, the effects of transmission of price shocks might be different from one level to another level of the marketing chain.

Improved market infrastructures should be introduced in order to reduce the number of market intermediaries and the dispersion of marketing information that result from high transaction costs. Improved market infrastructures would enable farmers and other stakeholders along the value chains to benefit from better information flow and linkages among one other. They would improve access to markets and institutions as well as access to information, increase the income of smallholder farmers, and favor food availability for urban consumers. For the minority of farmers who already participate in food grains markets, it is necessary to study patterns of market integration and price transmission to establish where markets do and do not function effectively in transmitting excess demand and supply across space (Barrett 2008). Yet, it is also important to provide baseline information on the nature of local rice price transmission between selected local markets. This analysis will be useful in identifying the marketing areas where high transaction costs are present, and where policy interventions may be needed.

As a contribution to providing such baseline information, we use the threshold models of Enders and Siklos (2001) on monthly series of local rice prices from 2000 to 2010 to examine the nature of rice price transmission between selected markets in the surplus zones and the nearest important consumption markets in Benin and Mali. The basic assumption made here is that local rice markets in surplus zones are asymmetrically integrated (involving asymmetric price transmission) with those in deficit zones in the same country because of high transaction costs. The existence of high transaction costs shows that the marketing margin is high, but when transaction costs are low, the marketing margin is low: if the marketing margin is low, the market operates freely, but when the marketing margin is high, there is no feedback due to excessive rises in the transaction costs which reinforce traders' marketing margins.

In addition to providing baseline information, we also analyze the relationship between the prices of locally produced rice among local markets in West African countries. This relationship has received little attention, while, since the 2008 food crisis, there has been growing interest in the dynamic linkage between rice prices of Asia and Africa (Daviron et al. 2008; Aker et al. 2009; Diallo et al. 2009; Bamba et al. 2010). The main reason explaining the lack of these kinds of studies is the reported unavailability of data on domestic rice prices in rural, suburban and urban zones in the same country. However, a better understanding of markets for locally produced rice in the same country or between African countries can promote domestic rice markets, and encourage farmers to produce more and to participate in the markets in order to improve their income levels. Another reason for this lack of interest in the price of locally produced rice is that many Sub-Saharan African farmers produce rice mostly for their own family consumption because there are no market outlets for their surpluses (WARDA 2001).

A review of the available literature indicates that the threshold models for price transmission analysis can identify the existence of transaction costs which involve nonlinearity in series. The traditional co-integration models of Engle and Granger (1987) and Johansen (1996) were criticized because they did not consider the issues of nonlinear or asymmetric characteristics caused by transaction costs. Threshold models generally allow a variable to display differing amounts of autoregressive decay depending on whether it is increasing or decreasing. This is in contrast to the Engle and Granger (1987) and Johansen (1996) tests, which implicitly assume a linear adjustment mechanism. Pippenger and Goering (1993), Balke and Fomby (1997), Enders and Granger (1998) as well as Enders and Siklos (2001) showed that all the tests for unit-roots and for co-integration have low power in the presence of asymmetric adjustments. In dealing with this problem, a number of studies have used nonlinear techniques to capture asymmetry effects when the variable is adjusting towards its long-run equilibrium. Barrett (1996) observed that failure to find co-integration between two markets' price series may be a result that is completely consistent with market integration if transaction costs are non-stationary.

Most of the analyses focusing on asymmetric price transmission refer to non-competitive market structures as an explanation for asymmetry (Abdulai 2000, 2002; Pede and McKenzie 2008; Fiamohe and Henry de Frahan 2012). According to Abdulai (2002), price is the primary mechanism by which various levels of the market are linked. Evidence of asymmetric price transmission has been considered as a manifestation of market failure (Meyer and von Cramon-Taubadel 2004) and thus provides justification for regulatory public interventions. The literature indicates that asymmetric price transmission can be caused by several factors, such as transport and transaction costs, market power, adjustment costs and domestic policies (Meyer and von Cramon-Taubadel 2004; Campa and Goldberg 2005; Ghosh and Rajan 2006). Many of these factors are at play in the West African food market. Abdulai (2000, 2002), Kuiper et al. (2003), Fiamohe and Henry de Frahan (2012) found that farmers suffer when there is imperfect competition among intermediaries, which allows the latter to exercise market power.

Threshold models are one approach for accommodating such transaction costs which exist in many agricultural markets in Sub-Saharan Africa. Balke and Fomby (1997) point out that the concept of threshold co-integration captures the essence of nonlinear adjustment processes. These are estimated to hold true for many economic phenomena including the behavior of inventories, money balances, consumer durables, prices and employment. These models can account for the effects of the marketing margin in the price transmission analysis even when transaction-cost data are unavailable. Enders and Siklos (2001) indicate that the threshold model, in particular the momentum threshold (M-TAR) adjustment, can be especially useful when policy makers are viewed as attempting to smooth out any large changes in a price series. Alderman (1993) rightly argues that there is a direct relationship between the ease with which stabilization policies can be implemented and the extent to which internal markets are integrated.

The remainder of the paper is organized as follows. "Modeling Asymmetric Price Transmission" describes the threshold co-integration models developed by Enders and Siklos (2001). In "Market Description and Data Analysis", we describe the selected markets and analyze price series data. "Testing for Asymmetric Price Transmission in Local Rice Markets in West Africa" applies the Enders and Siklos threshold models to analyzing the nature of price transmission between markets in Benin and Mali. Section 5 presents some "Concluding Remarks and Policy Implications".

Modeling asymmetric price transmission

We used the threshold model of Enders and Siklos (2001) to identify the prevalence of transaction costs by identifying asymmetry in price transmission. The prices of local rice in separate markets in the same country, at time *t*, is defined as X_{it} Such prices are assumed to have a long-run relationship which, according to Engle and Granger (1987), takes the form:

$$X_{it} = \beta_0 + \sum_{i=2}^{n} \beta_i X_{it} + \mu_t \tag{1}$$

where β_0 and β_i are respectively the constant and co-integrating parameters to be estimated, μ_t is a random error term with constant variance that can be contemporaneously correlated, and *t* is the time index. Long-run market integration test within this framework verifies whether any stable long-run relationship exists between the time-series variables. This implies that μ_t should be constant.

Engle and Granger (1987) consider the simple linear relationship used as the basis for the Dickey–Fuller test as follows:

$$\Delta\mu_t = \rho\mu_{t-1} + \varepsilon_t \tag{2}$$

where ε_t is a white-noise disturbance and ρ is a coefficient to be estimated for the unit tests (Dickey–Fuller test). The standard procedure is to estimate ρ and verify whether $-2 < \rho < 0$ using the appropriate critical values. The Granger representation theorem guarantees that if $\rho \neq 0$, Eqs. (1) and (2) jointly imply the existence of an error-correction representation. The standard models of co-integration and error-correction assume linearity and symmetric adjustment. Considering the above criticisms of the traditional tests, Enders and Siklos (2001) consider an alternative specification modifying the error term μ_t to allow for two types of asymmetric error corrections based on a co-integrating relationship as depicted in Eq. (1). The residuals μ_t obtained from Eq. (1) are used in:

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \sum_{i=1}^{p-1}\gamma_i\Delta\mu_{t-1} + \varepsilon_t \tag{3}$$

where I_t=[T_t, M_t] so that:

$$T_t = \begin{cases} 1\, if\;\; \mu_{t-1} \geq \tau \\ 0\, if\;\; \mu_{t-1} < \tau \end{cases} \tag{4a}$$

$$M_t = \begin{cases} 1\, if\;\; \Delta\mu_{t-1} \geq c \\ 0\, if\;\; \Delta\mu_{t-1} < c \end{cases} \tag{4b}$$

where $I_t[T_t, M_t]$ is the Heaviside indicator function, ε_t is the residual of the white-noise disturbance and τ and c denote the unknown threshold values in Eqs. (4a) and (4b), respectively. Equations (3) and (4a) represent the threshold autoregressive model (TAR), where the indicator function T_t depends on the previous period's μ_{t-1}. Equations (3) and (4b) represent the momentum threshold autoregressive model (M-TAR), where the indicator function M_t depends on the difference with the value of the previous period, $\Delta\mu_{t-1}$. The TAR or M-TAR model is called 'consistent model' when the threshold value τ or c is not zero. M-TAR implies that the adjustment mechanism of μ_{t-1} is dynamic, since the momentum of the series is greater in one direction than the other. Thus, for any large and smooth changes, the M-TAR model can explain the series more efficiently than the TAR model (Enders and Siklos 2001).

In Eq. (3), model-selection criteria, such as the Akaike information criterion (AIC) or the Bayesian information criterion (BIC), can be used to determine the appropriate lag length. Petrucelli and Woolford (1984) showed that the necessary and sufficient conditions for the stationarity of μ_t in Eq. (3) is ρ_1<0, ρ_2<0 and $(1+\rho_1)(1+\rho_2) < 1$ for any value of τ or c. If these conditions are met, μ_t=0 can be considered as the long-run equilibrium value of the system in Eq. (1). Since adjustment is symmetric if ρ_1=ρ_2, an important feature of Eq. (3) is that it retains its equivalence to the Engle–Granger specification. Moreover, Tong (1983, 1990) showed that the least squares estimates of ρ_1 and ρ_2 have an asymptotic multivariate normal distribution. In such circumstances, the adjustment is $\rho_1\mu_{t-1}$, if $\mu_{t-1}/\Delta\mu_{t-1}$ is above the threshold value and $\rho_2\mu_{t-1}$ if $\mu_{t-1}/\Delta\mu_{t-1}$ is below the threshold value.

In addition, because of the use of regression residuals rather than the actual error values, ρ_1 and ρ_2 have inference problems. Hansen (1997) and Enders and Falk (1999) considered issues related to this inference in TAR models. Enders and Falk (1999) state that when the value of the threshold is known, bootstrap t intervals and classic t intervals work well enough to be recommended in practice. They indicate that, for unknown values of thresholds, inference concerning the individual values of ρ_1 and ρ_2, and the restriction ρ_1=ρ_2 is problematic. The property of asymptotic multivariate normality has not been established for such a case. In discussing the difficulty of establishing the distribution of the parameter estimates, Chan and Tong (1989) conjectured that using a consistent estimate should establish the asymptotic normality of the coefficients. Moreover, Enders and Falk (1999) found that the inversion of the bootstrap distribution for the likelihood ratio statistic provides reasonably good coverage in small samples.

In general, the value of the threshold is unknown and needs to be estimated along with the values of ρ_1 and ρ_2. However, in a number of economic applications, it seems natural to assign a zero value to the threshold so that the co-integrating vector coincides with the attractor. For instance, Eq. (1) is an attractor such that its pull is strictly proportional to the absolute value of μ_t (Enders and Siklos 2001). This approach, that takes into account the TAR model with the attractor was used by Enders and Granger (1998), Abdulai (2000, 2002) and Fiamohe and Henry de Frahan (2012). According to Enders and Siklos (2001), in many applications, there is *no a priori reason* to expect the threshold to coincide with the attractor. In such circumstances, it is necessary to estimate the value of the threshold along with the values of ρ_1 and ρ_2. Chan (1993) showed that searching over the potential threshold values, so as to minimize the sum of squares errors from the fitted model, yields a super-consistent estimate of the threshold. Chan's (1993) grid search method sorts the values of the residuals μ_{t-1} and $\Delta\mu_{t-1}$ obtained from Eq. (1) in ascending order and excludes the smallest and largest 15 %. In this method, the consistent estimate of the threshold is the parameter that yields the lowest residual sum of squares over the remaining 70 % of the residuals.

Based on the threshold co-integration validation, the transmissions can be tested using the threshold error-correction model (TECM). Following Enders and Granger (1998) and Enders and Siklos (2001), the TECM can be written as:

$$\Delta X_{it} = I_t\emptyset_1\varepsilon_{t-1} + (1 - I_t)\emptyset_2\varepsilon_{t-1} + \sum_{i=2}^{n}\sum_{j=1}^{k}\beta_{ij}\Delta X_{i,t-J} + \vartheta_{1t} \tag{5}$$

where $\emptyset_1$ and $\emptyset_2$ are the adjustment coefficients for positive and negative discrepancies, respectively.

Market description and data analysis

Description of local rice markets in Benin and Mali

In Benin and Mali, there are many local markets engaged in local rice trade with other markets both within the same country and in neighboring countries. Among these local markets, we selected a pair in each country according to their size and the importance of local rice trade, and also on the basis of the availability of consistent data. The pair of markets comprised one market located in a surplus zone and the other located in an urban area. Our concern here is principally with the extent to which urban market prices are transmitted to the surplus zone market.

In Benin, Malanville and Parakou were chosen. Malanville market is in the irrigated rice production zone, which contributed about 23 % of the total rice production in Benin in 2011. Parakou market is an urban consumption market and accounts for 10 % of Benin's population. Both markets are in the North of Benin – Malanville in the Borgou region and Parakou in the Atacora region. The distance between Parakou and Malanville is about 319 km on the main north–south asphalt road, which was in a very bad condition until January 2012.

Segou Centre and Bamako were chosen for Mali. Segou Centre market is in the surplus zone of the Office du Niger. This irrigated rice production zone contributes about 25 % of Mali's total rice production. Bamako market is an urban consumption market and accounts for 12 % of Mali's population. The distance between Segou Centre and Bamako is about 238 km. The road linking the two markets is in good condition.

Data and unit root tests

The data used in this analysis are based on 127 monthly observations of retail prices from January 2000 to December 2010 for the selected local rice markets. Data were obtained from the Office National d'Appui à la sécurité alimentaire (ONASA) for Benin and Observatoire des Marchés Agricoles (OMA) for Mali. Both Benin and Mali belong to the Union economique et monétaire ouest africaine (UEMOA) economic region. Consequently, all prices used are in CFA francs per kilogram. Missing data points in the series of domestic rice prices were replaced by the average value of the preceding three observations. All price series were transformed into their natural logarithm.

The time plots in Figs. 1 and 2 show the coherence between prices in the local markets in the surplus zones (Segou and Malanville) and those in the deficit zones (Bamako and Parakou). These plots demonstrate large price fluctuations in the local markets.

Because, the price series are monthly, it likely that they might have seasonal components due to the non-linearity of the rice production season, increase of price due to high transportation costs during the rainy season, and other factors. In price transmission analysis, it is usual to remove such seasonal effects in order to capture the intrinsic relation between the prices of the different markets (Alderman 1993; Abdulai 2000; Fiamohe and Henry de Frahan 2012). We used the moving average method to systematically filter out the seasonal adjusted component of all prices series.

The hypothesis that the seasonally adjusted price series are non-stationary time series over whole periods was tested using the Augmented Dickey–Fuller (ADF) *t*-test (Dickey and Fuller 1979) and Phillips and Perron's (1988) (PP) method. The no-constant model was used to perform the unit root test, because it appeared to fit the series better than the other models developed for the Dickey–Fuller test. The lag length was selected on the basis of the Schwartz Bayesian Criteria (SBC) to ensure that the residuals are white noise. This was found to be two lags in all cases. All price series failed to reject the hypothesis of unit roots in levels, while stationarity could not be rejected for all the first differenced series at conventional significance levels according to the ADF and PP tests (Table 1). Consequently, these two tests showed that the series are integrated in order one [I (1)].

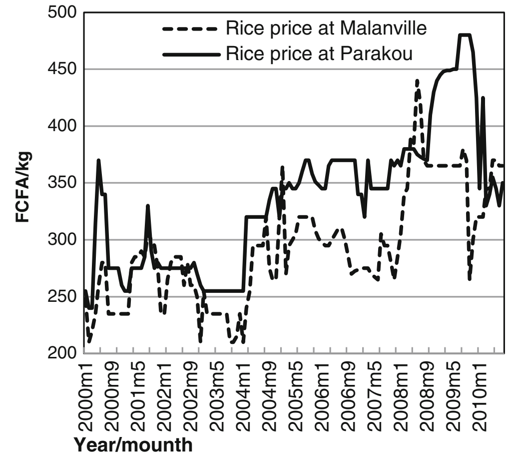

Fig. 1 Local rice price in selected markets in Benin. *Note:* m = month: 1–12 = January–December

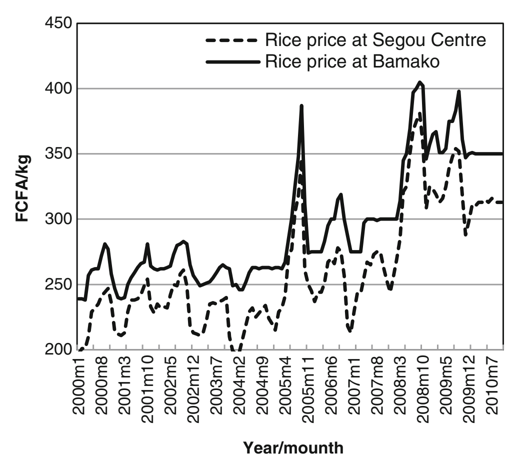

Fig. 2 Local rice price in selected markets in Mali. *Note:* m = month: 1–12 = January–December

Table 1 Stationary test on log of seasonally adjusted prices

Country	Market	ADF level	ADF difference	PP level	PP difference
Benin	Parakou	−0.10	−2.31**	0.61	−3.28***
	Malanville	0.89	−2.87***	1.54	−3.84***
Mali	Bamako	0.98	−3.45***	2.11	−3.47***
	Segou Centre	0.73	−3.02***	1.96	−3.31***

MacKinnon critical values:1 %: −2.597; 5 %: −1.950; and 10 %: −1.612

Statistical significance at 10 %, 5 % and 1 % denoted by *, ** and ***

ADF Augmented Dickey–Fuller test statistic, *PP* Phillips–Perron test statistic

Testing for asymmetric price transmission in local rice markets in West Africa

Asymmetry co-integration estimations

Given that the adjusted prices are I (1), the null hypothesis of no co-integration and symmetry of the co-integration relationship between local rice markets was tested using the residual from ordinary least squares (OLS) estimation of Eq. (1) for $n=2$. The estimated coefficients (Table 2) are highly significant for each pair of selected markets for Benin and Mali, showing that long-term changes in prices in one local market would result in price changes in the other market.

Following the Engle-Granger procedure, we used the classical model of Engle-Granger in the form of Eq. (2), the TAR model according to Eqs. (3) and (4a) and the M-TAR model according to Eqs. (3) and (4b). For each pair of models, we selected the models with two-lagged changes using the Schwartz Bayesian Criteria (SBC). The parameters estimated for all models have the correct signs for the convergence. Moreover, the speed of adjustment is more rapid for negative than for positive discrepancies according to the TAR and M-TAR models.

According to the Engle-Granger co-integration model (Table 3), the estimated values of parameter ρ are −0.03 (Benin) and −0.04 (Mali). The *t*-statistics for the null hypothesis that $\rho=0$ are −2.47 (Benin) and −2.52 (Mali). The critical values for the Engle-Granger test are −2.59, −1.95, and −1.61 at the 1 %, 5 % and 10 % significant levels, respectively. Hence, at the 5 % significance levels, the Engle–Granger test indicates that local rice markets in Benin and Mali are co-integrated.

Given that the Engle-Granger test can be incorrectly specified if the adjustment is asymmetric, we applied asymmetric co-integrating models. The TAR model according to Eqs. (3)–(4a) and M-TAR model according to Eqs. (3)–(4b) are estimated when the threshold $\tau=0$. The sample values of Φ statistic are 3.98 (Benin) and 3.09 (Mali) for the TAR model (Table 3). These sample values are smaller than the 10 % critical value (4.99), indicating that local markets in Benin and Mali are not co-integrated according to the TAR model when the threshold $\tau=0$.

The sample values estimated for the Φ statistic using the M-TAR model are 4.44 for Benin and 4.42 for Mali (Table 3). As in the previous case, these sample values are smaller than the 10 % critical value (5.47), indicating that local markets in Benin and Mali are not co-integrated under the M-TAR model when the threshold $\tau=0$.

Table 2 Estimate coefficients for cointegration relationships

Country	Pared markets	Estimated coefficient	Constant	Number of observation	R^2
Benin	Parakou-Malanvile	1.02***	0.01	128	0.781
		(21.19)	(0.03)		
	Malanville-Parakou	0.76***	1.24***	128	0.781
		(21.19)	(5.90)		
Mali	Bamako- Segou Centre	0.92***	0.58***	130	0.985
		(92.99)	(10.61)		
	Segou Centre- Bamako	1.07***	−0.54***	130	0.985
		(92.99)	(–8.29)		

The values without parentheses are the estimated coefficients and those in parentheses are *t*-statistics

The notations *, ** and *** correspond to 10 %, 5 % and 1 % significance levels

Table 3 Estimates of co-integration relationships for local rice markets in Benin and Mali

	Benin					Mali				
	EG	T	CT	MT	CMT	EG	T	CT	MT	CMT
ρ_1	−0.03	−0.05	−0.05	−0.04	−0.03	−0.04	−0.05	−0.06	−0.03	−0.04
	(−2.63)	(−2.54)	(−3.04)	(−2.78)	(−1.96)	(−2.46)	(−1.81)	(−2.30)	(−1.07)	(−2.03)
ρ_2	NA	−0.02	−0.01	−0.02	−0.11	NA	−0.04	−0.02	−0.06	−0.14
		(−1.13)	(−0.49)	(−1.06)	(−4.73)		(−1.75)	(−1.04)	(−2.78)	(−1.84)
γ_1	0.70	0.69	0.69	0.70	0.71	0.65	0.65	0.65	0.66	0.63
	(6.12)	(6.13)	(6.17)	(6.32)	(6.37)	(6.79)	(6.72)	(6.72)	(6.80)	(6.32)
γ_2	0.14	0.13	0.13	0.13	0.15	−0.07	−0.07	−0.06	−0.08	−0.06
	(1.07)	(1.06)	(1.08)	(1.04)	(1.23)	(−0.76)	(−0.74)	(−0.65)	(−0.90)	(−0.70)
Φ/Φ^*		3.987	4.890	4.442	13.92		3.095	3.165	4.425	3.677
$\mathrm{F-stat}(\rho_1=\rho_2)$		0.95	2.13	0.47	9.16		0.20	1.12	1.21	1.73
		(0.33)	(0.15)	(0.49)	(0.003)		(0.65)	(0.29)	(0.27)	(0.19)
τ/c		0	−0.093	0	0.029		0	−0.032	0	0.008

EG, T, CT, MT, and MCT = Engel-Granger, Threshold, Consistent Threshold, Momentum Threshold and Consistent Momentum Threshold, respectively

The terms ρ_1 and ρ_2 are the estimated values of positive and negative adjustment parameters with the corresponding t-statistics in parentheses for the null hypotheses ρ_1=0 and ρ_2=0, respectively. For the linear adjustment, ρ_1 is considered as a parameter in the Ender and Granger approach

The terms γ_1 and γ_2 are the estimated values of lagged variables parameters with the corresponding t-statistics in parentheses.

The terms Φ is the sample values statistic for the TAR and M-TAR models while the term Φ^* the sample values statistic for the consistent TAR and M-TAR models denoting the test for null hypothesis $\rho_1=\rho_2=0$

For the Engle–Granger model, the critical values of MacKinnon (1991) are −2.59, −1.95 and −1.61 at 1 %, 5 % and 10 % significant levels, respectively.

The critical values of the statistic Φ are 4.99, 6.01, 8.30 for the TAR model; and 5.47, 6.51, 8.85 for M-TAR model at 10 %, 5 % and 1 %, respectively. These critical values are from Enders and Siklos (2001)

The critical values of statistic Φ^* are 6.02, 7.08, 9.51 for the consistent TAR model; and 5.76, 6.86, 9.29 for the consistent M-TAR model at 10 %, 5 % and 1 %, respectively. These critical values are from Enders and Siklos (2001)

F-stat is the sample value of F-statistic for the null hypothesis that the adjustment coefficients are equal ($\rho_1=\rho_2$). Significance levels are in parentheses below

The terms τ and c are the estimated values of thresholds respectively for the TAR and M-TAR models

The consistent TAR model according to Eqs. (3)–(4a) and the consistent M-TAR model according to Eqs. (3)–(4b) are estimated with the threshold $\tau\neq0$, which is determined using Chan's (1993) method. Chan (1993) demonstrated that searching over all values of possible attractors to minimize the sum of squares errors from the fitted model yielded a super-consistent estimate of the threshold.

In searching over the possible thresholds lying in the middle 70 % of the sorted values of $\Delta\mu_t$, we found for the consistent TAR model, the threshold values of −0.09 for Benin and −0.03 for Mali in the smallest residual sum of squares for selected local rice price pairs. Using these estimated consistent threshold values, the corresponding sample values estimated for Φ^* statistic are 4.89 for Benin and 3.16 for Mali. These values are also smaller than the 10 % critical value (7.08), indicating that the local markets in Benin and Mali are not co-integrated under the consistent TAR model.

With the consistent M-TAR model, the threshold values estimated for selected local rice price series are 0.028 for Benin and 0.008 for Mali. Using these estimated consistent threshold values (Table 4), the corresponding sample values estimated for Φ^* statistic are 13.92 for Benin and 3.68 for Mali. The value of Φ^*=3.68 for Mali is smaller than the 10 % critical value (5.76), indicating that we cannot reject the null hypothesis ($\rho_1=\rho_2=0$) of no co-integration. However, the value of Φ^*=13.92 for Benin is greater than the 1 % critical value (9.29), rejecting the null hypothesis of no co-integration under the consistent M-TAR model.

Since differential local rice prices are co-integrated for markets in Benin, the null hypothesis of symmetric adjustment ($\rho_1=\rho_2$) can be tested using a standard F-distribution (Enders and Granger 1998; Enders and Siklos 2001). The sample value of the F-statistic=9.16 with the corresponding p value of 0.003 indicates a rejection of the null hypothesis of symmetric adjustment at the 5 % level of significance. This suggests that adjustment is asymmetric in the sense that positive shocks to the marketing margin (or negative shocks to the local rice prices) tend to persist, but negative shocks revert quickly towards the attractor.

Asymmetric and symmetric error correction results

The positive finding of asymmetric co-integration with the consistent M-TAR adjustment for Benin local rice markets means that we need to examine the short-run dynamics of local rice prices in Benin with an asymmetric error-correction model (AECM). For the sake of comparison, estimated results are reported for both asymmetric and symmetric ECMs (Table 4).

For the consistent M-TAR error-correction model estimated for the Benin local rice market pair, the *t*-statistics for positive and negative adjustments indicate that when the marketing margin is lower than the threshold value, the market system operates freely and there is feedback between the Malanville surplus market and the Parakou consumption market. However, consumer prices in Parakou market adjust significantly more quickly so as to eliminate approximately 12 % of a unit negative change in the deviation from the equilibrium relationship created by changes in prices observed in the Malanville surplus market (Table 4). By contrast, prices in the Malanville market adjust significantly only by 5 % of a negative change in deviation from the equilibrium created by positive changes in prices observed in the Parakou consumption market. However, when the marketing margin is higher than the threshold value, there is no longer any feedback between this pair of markets.

The *t*-statistics from the standard ECM indicate that the error-correction terms are only significant at conventional levels for the Malanville surplus market. The estimated adjustment parameter is −0.04. This implies that there is convergence towards long-run equilibrium when price changes are observed in the Parakou consumer market. However, according to the asymmetric results obtained with a consistent M-TAR ECM, we cannot validate the result of convergence using the standard model for Benin.

Considering the causality test, the F-statistic for the null hypothesis shows that for each price pair, changes in the

Table 4 Asymmetry price transmission results for Benin markets

TECM estimated for Parakou Prices (PP)			TECM for Malanville Prices (MP)		
Coefficient	Engel-Granger	Consistent momentum	Coefficient	Engel-Granger	Consistent momentum
$\varnothing_1$	−0.01	−0.00	$\varnothing_1$	−0.04***	−0.03
	(−1.30)	(−0.34)		(−2.76)	(−1.24)
$\varnothing_2$	–	−0.12***	$\varnothing_2$	–	−0.05*
		(−3.91)			(−1.89)
$\beta_{11}[\Delta PP\ (t-1)]$	0.74***	0.78***	$\beta_{11}[\Delta MP\ (t-1)]$	0.75***	0.75***
	(4.47)	(5.05)		(8.59)	(7.84)
$\beta_{12}[\Delta PP\ (t-2)]$	0.12	0.13	$\beta_{12}[\Delta MP\ (t-2)]$	0.09	0.09
	(0.64)	(0.72)		(1.00)	(0.95)
$\beta_2[\Delta MP\ (t)]$	0.10	0.08	$\beta_2[\Delta PP\ (t)]$	0.11	0.12
	(1.39)	(1.20)		(1.11)	(1.26)
$\beta_{21}[\Delta MP\ (t-1)]$	−0.11	−0.11	$\beta_{21}[\Delta PP\ (t-1)]$	0.16	0.15
	(−1.28)	(−1.29)		(1.16)	(1.23)
$\beta_{22}[\Delta MP\ (t-2)]$	0.05	0.02	$\beta_{22}[\Delta PP\ (t-2)]$	−0.25**	−0.24**
	(0.77)	(0.31)		(−2.27)	(−2.34)
Constant	−0.00	−0.00	Constant	0.00	0.00
	(−0.24)	(−0.72)		(0.85)	(0.07)
Q-stat	0.62	0.19	Q-stat	3.02	2.978
	(0.43)	(0.66)		(0.08)	(0.09)
F-stat	0.816	0.962	F-stat	2.634	2.759
	(0.445)	(0.385)		(0.076)	(0.067)

PP and MP represent the Parakou consumer Prices and Malanville Prices, respectively. β is a coefficient estimated for ΔPP or ΔMP

$\varnothing_1$ and $\varnothing_2$ are the coefficients of error correction terms showing adjustments to increasing and decreasing deviations from the long-run, respectively. The *t*-statistics are in parentheses

Q-stat denotes the Ljung-Box statistic that the first two of the residual autocorrelations are jointly equal to zero. The significance levels are in parentheses below

F-stat denotes the Granger causality test with the significance levels in parentheses below

Malanville market is caused by changes in the Parakou consumption market, while the Malanville market has no influence on the Parakou consumption market. These results suggest there is uni-directional causation from the Parakou consumer market to the Malanville surplus market. Hence, the Parakou consumption market reacts more quickly to negative marketing margin shocks than does the Malanville market. This is clearly consistent with the Parakou consumer market being the dominant market. These results are consistent with those of many empirical analyses of price transmission for agricultural products, which show that, in some periods, price increases (decrease in marketing margin) are more rapidly and fully transmitted than price decreases (Abdulai 2000, 2002; Goodwin and Harper 2000; Aguiar and Santana 2002; Fiamohe and Henry de Frahan 2012). The finding of asymmetry in price transmission when the consistent marketing margin is lower than the threshold (0.029) means that a higher threshold of transaction costs causes an excessive rise in selected prices of local rice in Benin, which allows traders to adjust their marketing margin upward.

For Mali, the positive finding of market co-integration with linear adjustment means that we need to use standard ECM to examine the local rice markets. The results of this examination are reported in Table 5.

For the standard ECM estimated, the *t*-statistics for the long-run adjustment parameter indicate that the pairs of local rice prices in the Bamako and Segou Centre markets mutually and significantly respond to any shock in prices observed in either market. However, the adjustment parameter estimated for the Segou Centre market (−0.05) is slightly greater than the one estimated for the Bamako market (−0.04). These results imply that, in the short run, traders of locally produced rice in both the Bamako consumer market and the surplus market of Segou Centre can immediately respond to any changes in prices by adjusting their prices.

Moreover, the F-statistic of causality test for the null hypothesis that changes in prices in the Segou Centre market do not affect consumer prices in the Bamako market is 5.96 with a corresponding *p* value of 0.003, while the F-statistic for the null hypothesis that changes in consumer prices do not affect prices in Segou Centre market is 32.48 with a corresponding *p* value of 0.00. These results suggest bi-directional causation from the Bamako consumer market to the Segou Centre surplus market and *vice versa*, suggesting perfect price transmission between the two markets, likely due to lower transaction costs.

Concluding remarks and policy implications

The objective of this paper was to provide baseline information on the nature of local rice price transmission between paired local markets in Benin and Mali over the period 2000–2010, and to identify where markets do and do not function well. Although the transmission of local rice prices for Mali paired markets was symmetric, the evidence presented for Benin using Enders and Siklos threshold

Table 5 Standard price transmission results for local rice markets in Mali

Coefficient	ECM estimated for Bamako Prices (BP)	Coefficient	ECM estimated for Segou Prices (SP)
$\varnothing_1$	−0.04**	$\varnothing_1$	−0.05***
	(−2.38)		(−2.96)
$\beta_{11}[\Delta BP\ (t-1)]$	0.68***	$\beta_{11}[\Delta SP\ (t-1)]$	0.66***
	(7.19)		(7.16)
$\beta_{12}[\Delta BP\ (t-2)]$	−0.13	$\beta_{12}[\Delta SP\ (t-2)]$	0.01
	(−1.60)		(0.14)
$\beta_2[\Delta SP\ (t)]$	0.65***	$\beta_2[\Delta BP\ (t)]$	1.04***
	(10.02)		(15.84)
$\beta_{21}[\Delta SP\ (t-1)]$	−0.27***	$\beta_{21}[\Delta BP\ (t-1)]$	−0.84***
	(−2.88)		(−7.49)
$\beta_{22}[\Delta SP\ (t-2)]$	−0.04	$\beta_{22}[\Delta BP\ (t-2)]$	0.16
	(−0.55)		(1.57)
Constant	0.00	Constant	−0.00
	(0.56)		(−0.00)
Q-stat	3.01	Q-stat	3.727
	(0.08)		(0.05)
F-stat	5.96	F-stat	32.48
	(0.003)		(0.00)

BP and SP are the Bamako consumer Prices and Segou Centre Prices, respectively. β is a coefficient estimated for ΔBP or ΔSP

$\varnothing_1$ is the coefficient of error correction terms showing adjustments from the long run

The *t*-statistics are in parentheses

Q-stat denotes the Ljung-Box statistic that the first two of the residual autocorrelations are jointly equal to zero. The significance levels are in parentheses below

F-stat denotes the Granger causality test with the significance levels in parentheses below models strongly supports the hypothesis of asymmetric price responses.

The results for Benin show that increases in local rice prices in the Parakou consumption market are transmitted more rapidly to the Malanville surplus market than the opposite. Furthermore, Granger causality tests indicate a unidirectional relationship from the Parakou consumption market to the Malanville surplus market with no evidence of the reverse causality feedback. These results mean that rice farmers in rice growing areas at Malanville were not informed on time or not at all of the increases in price in the Parakou consumption market. These results highlight important imperfections that increase transaction costs and hamper farmers' participation in local rice markets. They suggest that policies targeting smallholder producers' welfare need to address the reduction of transaction costs to enable them to benefit from better market price in consumption areas. The asymmetric price transmission observed between the paired local markets in Benin may indicate the high level of transaction costs probably due to the long distance between the two markets and the bad state of the road linking them. Moreover, many recent studies of price transmission in food marketing chains have found that trader associations use market power to maintain their marketing margins and this strongly influences price transmission. Therefore, the structure of markets, e.g. the existence and functioning of trader associations, can also influence Benin market behavior. The results of this study are consistent with those from previous studies, indicating that agricultural products markets in West African countries are well integrated but exhibit asymmetry in price transmission (Alderman 1993; Abdulai 2000, 2002; Pede and McKenzie 2008; Fiamohe and Henry de Frahan 2012).

Conversely, the positive symmetric co-integration and price transmission results for Mali suggest that its paired local rice markets function well, with easy access and good connection between them. Although these selected local rice markets are well integrated in Mali, further research is needed to know if this integration contributes to the effective improvement of rice farmers' incomes and favors permanent access to local rice by Bamako consumers. Research on these aspects will be carried out with baseline surveys which are presently ongoing in Mali.

At this stage of the study, we cannot confirm with certainty that the long distance and bad state of the road linking the surplus market to the consumption market are the only components of high transaction costs that influence local rice markets in Benin. The baseline surveys, which are ongoing in many countries, including Benin, should help to identify which components of transaction costs cause the asymmetric price transmission found. Nevertheless, given also the consistency of sources of asymmetry reported by many recent studies, the results of this study for Mali, as concerns the effects of good roads on price transmission, could argue in favor of policy interventions covering investments on the rehabilitation of the main and rural roads. Such investments would facilitate the connection of local rice traders from separate markets. Specifically, the development of infrastructure in rural areas at Malanville where rice is grown would strongly reduce the transaction costs, and then asymmetry in price transmission for farmers' better participation in the local rice markets. These investments could intensify local rice trade flow between these selected markets. By investing in public infrastructure, the vertical integration of farm production with marketing could specifically be promoted by reducing these transaction costs, and then improving rice farmers' incomes. This would be crucial to achieving food security for rice farmers and consequently assure their wellbeing.

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