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# Spatial market integration of rice in Bangladesh in the presence of transaction cost

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## Abstract

Modeling transaction costs is of particular importance when analyzing price integration in developing countries. Spatial price integration among five major Bangladesh rice markets are examined over the post-liberalization period that covered a long time span from January 1999 to December 2021, the longest period covered ever. We account for nonlinear price adjustments induced by transaction costs using a threshold cointegration approach to account for the impact of transaction costs in the price adjustment process. Our results show that large price deviations from long-run equilibrium are corrected within 2–3 months. Our results provide strong supporting evidence of the presence of threshold effects, which means that large transaction costs contribute to slow equilibrium price adjustments. Thus, our results provide important policy implications for Bangladesh rice markets, namely that policies aimed at reducing transaction costs are needed to engender greater pricing efficiency in Bangladesh rice markets.

**Keywords:** Market integration, Rice markets, Transaction cost, Bangladesh

## Introduction

Rice is synonymous with food security in Bangladesh, and economic and political stability is highly correlated with rice price volatility (Kabir et al. 2020; Sayeed and Yunus 2018). Rice prices raise farmers' incomes and livelihoods. However, high prices raise food spending and food insecurity and hurt the impoverished portion of the country. Rice farming has moved away from mostly subsistence to considerably more commercial farming, resulting in significant changes in the domestic trade of rice. The marketed surplus of rough rice produced at the farm level is 54 to 60% of total production depending upon seasons, farm categories and geographical segmentations in Bangladesh (Rahman et al. 2021).

The Bangladesh government has enacted substantial policy reforms over the last 30 years to increase pricing efficiency among its domestic rice markets. These reforms were recommended in the 1980s by the World Bank and the International Monetary Fund, under the structural adjustment program. As a result of the policy reforms,

Bangladesh domestic rice markets were liberalized and all kinds of supports were virtually abolished. However, producers and consumers could not fully benefit from the liberalization reforms, as Bangladesh commodities markets were not well integrated, leading to ineffective price signals across marketing channels. In the late 1980s and early 1990s, Bangladesh's rice markets experienced massive liberalization (Hossain and Verbeke 2010). Since the late 1990s, large-scale automated rice mills have greatly increased milling capacity with large investments in Bangladesh (Reardon et al. 2014). Unhusked rice is processed in semi-automated and automatic rice mills and is supplied to both local and long-distance markets. This form of market expansion has fulfilled the local demand and taste for locally supplied clean rice. This milling revolution also added quality to basic rice through polishing, grading and bagging, allowing millers to better share economic benefits between producers and customers (Murshid 2015).

Until the mid-1990s, Bangladesh's rice market was largely isolated from the world to insulate itself from world price volatility (Sayeed and Yunus 2018). In 1998, the country was self-sufficient in rice production for the first time in history, but in 1999, floods forced the country back to imports (Kabir et al. 2020). In mid-2007, Bangladesh's rice markets were affected by market volatility as world rice and other cereals grain prices surged dramatically. This volatility was exasperated in October 2007 when India ceased exporting rice to Bangladesh because of low public wheat reserves (Dorosh 2009). Bangladesh was able to overcome this shortage by boosting domestic rice production via the use of green revolution-led technology and government support. Since 2009, the country has consistently been close to self-sufficiency in rice production, exporting an average of 5,998 million tons of fragrant rice annually (Ali and Sunny 2020). In Bangladesh, the government's share of the rice market is very low and has a negligible impact on the pricing and distribution system of rice to the end users (Rahman et al. 2020). Private traders, millers and wholesalers mostly procure and sell domestic rice, resulting in a wide price spread between farm and retail levels (Alam et al. 2016). Furthermore, a few well-organized millers and wholesalers wield large market power, leading to market failure (Rahman et al. 2020). Farmers are obligated to sell rough rice at a lower price than the competitive market price, but consumers pay exorbitant costs (Abdullah and Hossain 2013).

Over the last decades, transportation infrastructure—roads and communication and mobile networks—has been greatly improved and developed. In the wake of these infrastructure and communication developments, greater spatial market integration was expected. High levels of spatial market integration are crucial to market performance. Markets that are not integrated may convey inaccurate price information, leading to misguided policy decisions and a misallocation of resources. Sexton et al. (1991) identified three reasons for a lack of market integration: imperfect competition, differential trade barriers and prohibitive transaction costs. With this in mind, we model the impact of transaction costs, which are typically high in developing countries, using a threshold vector error correction model.

Although several studies have examined rice market integration in Bangladesh, to date no comprehensive studies that consider the role of transaction costs (hereafter TC) with respect to market integration have been done. The seminal work of Ravalion (1986) showed that there is limited market integration in Bangladesh rice markets.

In addition, Goletti et al. (1995) concluded that market integration in Bangladesh rice markets is moderate. These conclusions of limited and moderate market integration in the pre-reform era reflected restricted food grain movement, poor infrastructure and inadequate communications. For example, prior to market reforms, Bangladesh government procured rice from surplus regions to maintain a buffer stock and this policy restricted the incentive of private traders to move rice from surplus to deficit regions. In effect, the policy prevented price equalization across regions. In contrast, Dawson and Dey (2002) showed that Bangladesh rice markets were perfectly integrated following the trade liberalization reforms. The authors used a vector autoregressive error correction model (VECM) to test the Law of One Price (LOP) within the central–regional market. However, they did not account for transportation costs. Their standard VECM modeling framework implicitly assumes that the price adjustment process is linear and symmetric. However, in more recent literature such as Enders and Siklos (2001), Enders and Granger (1998), Goodwin and Piggot (2001), Meyer (2004), Sarno et al. (2004), Pede and McKenzie (2008), it is argued that the standard cointegration framework is mis-specified if the true adjustment process is nonlinear and asymmetric. This is likely the case if TC are significant.

The factors that might contribute to higher TC are inadequate infrastructure, transportation bottlenecks, lack of market information, information asymmetry, market power and menu costs. These kinds of factors are common in developing countries' agricultural markets and pose serious challenges to policy makers. Therefore, estimating threshold TC and their effect on price adjustments from one market to another or from one supply chain level to another should be a rule rather than an exception, especially in the context of developing countries. Hossain and Verveke (2010) examined if Bangladesh rice markets have become spatially integrated following the liberalization using wholesale weekly coarse rice prices during January 2004 to November 2006. The authors used Johansen cointegration analysis and found that rice markets in Bangladesh are only moderately integrated. Although this paper examined long-term relationships, it covered only a very short span of time—less than 3 years. This short span of time might not be sufficient to make firm conclusions about long-term adjustments. Siddique et al. (2008) investigated spatial integration between Dhaka, Mymensingh, Sherpur, Kishoregonj, Bogra, Rangpur and Dinajpur districts using monthly wholesale rice prices from January 1994 to December 2004. They concluded that distance between markets was not an important impediment to integration. More recently, Rahman et al. (2021) examined market cointegration among five main rice markets in Bangladesh using monthly wholesale prices from January 2006 to June 2017. They concluded that rice markets in Bangladesh are well integrated. Although Dhaka is the leading consumer market in the country and is home to about 20 million inhabitants, the study concluded that the leading rice market is Chattagram. This is a particularly surprising result given that Chattagram is the lowest rice-producing region.

Akhter (2016) studied market integration between surplus (India) and deficit rice markets (Bangladesh and Nepal) during the global food crisis. They concluded that domestic rice prices of India, Bangladesh and Nepal are integrated in both the short and long run in spite of the imposition of export restriction policies by India. Alam et al. (2012) examined the dynamic relationship between the world and the domestic market price

of rice for Bangladesh over the September 1998–February 2007 period. They found the evidence of a long-run unidirectional equilibrium relationship, meaning that domestic Bangladesh prices adjust to the world prices but not vice versa. However, neither Akhter (2016) nor Alam et al. (2012) analyzed markets integration among different regions within Bangladesh.

Our study first tests whether domestic Bangladesh rice markets are integrated using Johansen cointegration and secondly tests for price causality between these markets. We extend Dawson and Dey (2002), Hossain and Verveke (2010), Siddique et al. (2008) and Rahman et al. (2021), by relaxing the assumptions of linear price adjustments. We explicitly account for nonlinear price adjustments in the presence of transaction costs using the Hansen-Seo (2002) threshold cointegration model. Under this approach, the threshold is estimated by means of a grid search approach. The proposed methodology is appropriate when only price data are available; if trade flow and TC data were also available, the parity bound method of Baulch (1997) would be a more appropriate alternative. Since the sample size is relatively small for threshold cointegration and threshold model estimation, we attempt to estimate the linear model first in order to validate the results from the threshold model.

The paper contributes to the existing literature in two different ways. It is the first paper to consider transaction cost-induced nonlinear price adjustments when testing for spatial market integration in Bangladesh rice markets. In addition, it is the first study of its kind to examine Bangladesh rice market integration with respect to the post-liberalization reform era. Our analysis includes the data from January 1999 to December 2021, which is the longest time series examined to date.

The remainder of the paper is organized as follows. The next section describes the linear cointegration approach, a causality test for market dominance, and the threshold cointegration model. The data are explained in Sect. 3. Section 4 presents the results and discussions, and the last section concludes.

## Econometrics methodology

### Johansen–Juselius (1992) cointegration model

If prices are non-stationary and in same order of integration, then the Johansen–Juselius (1992) likelihood ratio test in the vector autoregressive (VAR) specification is as follows:

$$\Delta P_t = \Phi D_t + \Pi P_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta P_{t-1} + \omega_t \quad (1)$$

where  $P_t$  includes all  $n$  variables of the model which are  $I(1)$ ,  $\Pi$ ,  $\Gamma_i$  and  $\Phi$  are parameter matrices to be estimated,  $D_t$  is a vector with deterministic elements (constant, trend) and  $\omega_t$  is a vector of random error that follows Gaussian process. If  $\Delta P_t \Delta P_t \sim I(0)$  is  $I(0)$ , then  $\Pi$  will be a zero matrix except when a linear combination of the variables in  $P_t$  is stationary. If  $\text{rank } \Pi = r = K$ , the variables in levels are stationary meaning that no integration exists; if  $\text{rank } \Pi = r = 0$ , all the elements in the adjustment matrix have value zero; therefore, none of the linear combinations are stationary. According to the Granger representation theorem (1987) that when  $0 < \text{rank } (\Pi = r) < K$ , there are  $r$  cointegrating vectors. For example if  $\text{rank } (\Pi = r) = 1$ , there is single cointegrating vector or one linear

combination which is stationary such that the coefficient matrix  $\Pi$  can be decomposed into  $\Pi = \alpha\beta'$  where  $\alpha$  is the vector of loading factor and  $\beta$  is the cointegrating vector in where  $\beta'P_{t-1}$  is  $I(0)$ . Johansen method is to estimate  $\Pi$  matrix from an unrestricted VAR and to test whether we can reject the restrictions implied by the reduced rank of  $\Pi$ . There are two methods of testing for reduced rank ( $\Pi$ ), the trace test and maximum eigenvalue tests. The trace statistics tests the null hypothesis that the number of distinct cointegrating vectors ( $r$ ) is less than or equal to  $r$  against a general alternative. Another statistics maximal eigenvalue tests the null that the number of cointegrating vector is  $r$  against the alternative of  $r + 1$ .

**Causality tests from Johansen VECM**

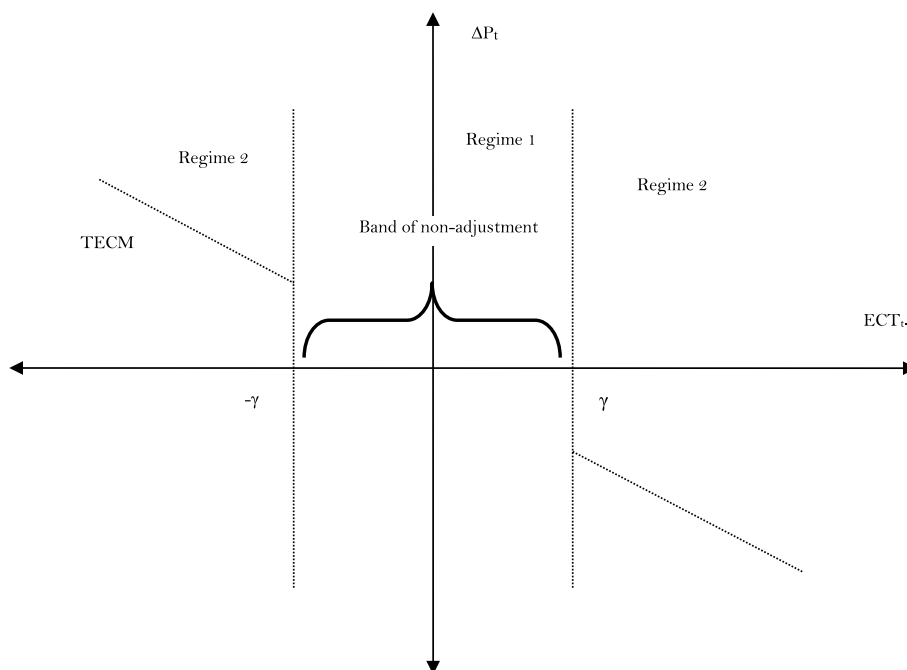
The existence of cointegration in bivariate relationship implies Granger causality which under certain restrictions can be tested within the framework of Johansen VECM by standard Wald test (Masconi and Giannini 1992; Dolado and Lutkepohl 1996). The underlying principle is that if  $\alpha$  matrix in cointegration matrix ( $\Pi$ ) has a complete column of zeros, then no causal relationship exist, because there is no cointegrating vector in that particular block. For pair-wise causal relationship, it can be written in Eq. (2):

$$\begin{bmatrix} \Delta P_{1,t} \\ \Delta P_{2,t} \end{bmatrix} = \begin{bmatrix} \mu_1 \\ \mu_2 \end{bmatrix} + \sum_{i=1}^{k-1} \begin{bmatrix} \Gamma_{i,11} & \Gamma_{i,12} \\ \Gamma_{i,21} & \Gamma_{i,22} \end{bmatrix} \begin{bmatrix} \Delta P_{1,t-i} \\ \Delta P_{2,t-i} \end{bmatrix} + \begin{bmatrix} \alpha_1 \\ \alpha_2 \end{bmatrix} [\beta_1 \ \beta_2] \begin{bmatrix} P_{1,t-k} \\ P_{2,t-k} \end{bmatrix} + \begin{bmatrix} \omega_{1t} \\ \omega_{2t} \end{bmatrix} \tag{2}$$

In Eq. (2), the subscript number refers to the markets. There are three possible cases of causality to be tested, (a)  $\alpha_1 \neq 0, \alpha_2 \neq 0$  (b)  $\alpha_1 = 0, \alpha_2 \neq 0$  and (c)  $\alpha_1 \neq 0, \alpha_2 = 0$ . The first one is bidirectional causality and the last two imply unidirectional causality. To explain how to make implications of the causality decision suppose  $\alpha_1 = 0$  this implies that the error correction term or the third term of the right-hand side of the first equation of Eq. (2) is eliminated and the long-run solution to  $\Delta P_{1,t}$  will not be affected by the deviations from the long-run equilibrium path defined by the cointegrating vector. In the same way, when  $\alpha_2 = 0$  the  $\Delta P_{1,t}$  will not cause  $\Delta P_{2,t}$ .

**Threshold cointegration**

Early research concentrated on linear price cointegration, while subsequent research has moved to the regime-dependent nature of price relationships (Ihle and von Cramon-Taubadel 2008). The concept of threshold cointegration was first introduced by Balke and Fomby (1997) to account of transaction cost-induced nonlinear price dynamics. It is now well understood in the literature that TC may inhibit price integration across spatially separated markets (for example, see Barret and Li 2002; Fackler and Goodwin 2001; Goodwin and Piggott 2001; Abdulai 2000, 2002; Goodwin and Harper 2000). A threshold introduces nonlinearities into the functional relationship between prices of market pairs (Tong 1990). Hansen and Seo (2002) proposed a SupLM test statistic with the null hypothesis of linear cointegration against the alternative hypothesis of threshold cointegration. Hansen and Seo (2002) note that price movements toward a long-run equilibrium might not occur in every time period, due to the presence of TC. Goodwin and Piggott (2001) used a threshold error correction model to estimate spatial integration in US corn and soybean markets. Ben-Kaabia



**Fig. 1** Effect of transaction costs in the price adjustment

and Jose (2007) have estimated price transmission between vertical stages of the Spanish lamb market using a threshold model. Sanogo and Maliki (2010) have analyzed integration between Nepal and Indian rice markets applying a threshold autoregressive model. One implicit assumption of linear cointegration models, like Johansen and Jesulius (1992) and Engel and Granger (1987), is that price adjustments induced by deviations from a long-term equilibrium are a continuous and a linear function of the magnitude of the deviations.

In contrast, a threshold cointegration mode that takes into account TC allows price adjustments to differ based on the magnitude of the deviations from a long-run equilibrium. The speed of price adjustment can also differ depending upon whether deviations are above or below a specific threshold—which proxies the size of TC.

In Fig. 1, the price adjustment ( $\Delta P_t$ ) is considered to be a function of deviations from a long-run equilibrium, which can be represented by a two-regime threshold vector error correction model (TVECM). We proceed by estimating the two-regime TVECM proposed by Meyer (2004), which is an extension of Hansen and Seo (2002). Pede and McKenzie (2005) take this approach to estimate market integration in Benin maize markets.

Following Hansen and Seo (2002), let  $P_t$  be a two-dimensional I (1) price series with one  $2 \times 1$  cointegrating vector  $\beta$  and  $w_t(\beta) = \beta'P_t$  denote the I (0) error correction term. Considering linear relationship, the vector error correction model (VECM) can be written as follows:

$$\Delta p_t = A'P_{t-1}(\beta) + \mu_t \tag{3}$$

where

$$P_{t-1}(\beta) = \begin{pmatrix} 1 \\ w_{t-1}(\beta) \\ \Delta p_{t-1} \\ \Delta p_{t-2} \\ \vdots \\ \Delta p_{t-l} \end{pmatrix} \tag{4}$$

In Eq. 4,  $P_{t-1}(\beta)$  is  $k \times 1$  and the matrix  $A$  is  $k \times 2$  of coefficients. The model assumes that the error term  $u_t$  is a vector of a Martingale difference sequence with finite covariance matrix  $\Sigma = E(u_t u_t')$ . The term  $w_{t-1}$  represents the error correction term obtained from the estimated long-term relationship between two market prices. The two prices are simultaneously explained by deviations from the long-term equilibrium (error correction term), the constant terms and the lagged short-term reactions to previous price changes. The parameters  $(\beta, A, \Sigma)$  are estimated following a maximum likelihood estimate (MLE) approach with the assumption that the errors  $u_t$  are independently and identically Gaussian.

A two-regime threshold cointegration model is given as:

$$\Delta p_t = \begin{cases} A_1' P_{t-1} + u_t & \text{if } w_{t-1}(\beta) \leq |\gamma| \\ A_2' P_{t-1} + u_t & \text{if } w_{t-1}(\beta) > |\gamma| \end{cases} \tag{5}$$

where  $\gamma$  represents the threshold parameter. The model in Eq. (5) may also be written as:

$$\Delta p_t = A_1' P_{t-1}(\beta) d_{1t}(\beta, \gamma) + A_2' P_{t-1}(\beta) d_{2t}(\beta, \gamma) + u_t \tag{6}$$

where

$$d_{1t}(\beta, \gamma) = 1 \text{ if } w_{t-1}(\beta) \leq |\gamma| \tag{7}$$

$$d_{2t}(\beta, \gamma) = 1 \text{ if } w_{t-1}(\beta) > |\gamma| \tag{8}$$

The coefficient matrices  $A_1$  and  $A_2$  govern the dynamics in the regimes. Values of the error correction term, in relation to the level of the threshold parameter  $\gamma$  (in other words, whether  $w_{t-1}$  is above or below  $\gamma$ ), allow all coefficients except the cointegrating vector  $\beta$  to switch between these two regimes.

The threshold effect exists if  $0 < P(w_{t-1} \leq |\gamma|) < 1$ , otherwise the model belongs to the linear cointegration form. We impose this constraint assuming that  $\pi_0 < P(w_{t-1}(\beta) \leq |\gamma|) < (1 - \pi_0)$  and by setting  $\pi_0 > 0$  as a trimming parameter equal to 0.05 (Andrews 1993)<sup>1</sup> in the empirical estimation. Further, we ensure that the indicator function represented by Eqs. (7) and (8) contains enough sample variation for each choice of  $\gamma$ . The likelihood function of the model in Eq. (6) under the assumption of *iid* Gaussian error  $u_t$  has the following form:

<sup>1</sup> For our empirical estimation we fixed the trimming parameter to 0.05 following Hansen and Seo (2002) and Ben-Kaabia and Jose (2007). Therefore, each regime is restricted to contain at least 5% of all observations.



$$\ln(A_1, A_2, \beta, \Sigma, \gamma) = -\frac{n}{2} \text{Log}|\Sigma| + \frac{1}{2} \sum_{t=1}^n u_t(A_1, A_2, \beta, \gamma)' \Sigma^{-1} u_t(A_1, A_2, \beta, \gamma), \tag{9}$$

where

$$u_t(A_1, A_2, \beta, \gamma) = \Delta p_t - A_1' P_{t-1}(\beta) d_{1t}(\beta, \gamma) - A_2' P_{t-1}(\beta) d_{2t}(\beta, \gamma) \tag{10}$$

The MLE of  $(\hat{A}_1, \hat{A}_2, \hat{\beta}, \hat{\Sigma}, \hat{\gamma})$  is obtained by maximizing the  $\ln(A_1, A_2, \beta, \Sigma, \gamma)$ . This is achieved by first holding  $(\beta, \gamma)$  fixed and computing the constrained MLE for  $(A_1, A_2, \Sigma)$  using the OLS regression as follows:

$$\hat{A}_1(\beta, \gamma) = \left( \sum_{t=1}^n P_{t-1}(\beta) P_{t-1}(\beta)' d_{1t}(\beta, \gamma) \right)^{-1} \left( \sum_{t=1}^n P_{t-1}(\beta) P_{t-1}(\beta)' d_{1t}(\beta, \gamma) \right), \tag{11}$$

$$\hat{A}_2(\beta, \gamma) = \left( \sum_{t=1}^n P_{t-1}(\beta) P_{t-1}(\beta)' d_{2t}(\beta, \gamma) \right)^{-1} \left( \sum_{t=1}^n P_{t-1}(\beta) P_{t-1}(\beta)' d_{2t}(\beta, \gamma) \right), \tag{12}$$

$$\hat{u}_t(\beta, \gamma) = u_t(\hat{A}_1(\beta, \gamma), \hat{A}_2(\beta, \gamma), \beta, \gamma) \text{ and}$$

$$\hat{\Sigma}_t(\beta, \gamma) = \frac{1}{2} \sum_{t=1}^n \hat{u}_t(\beta, \gamma) \hat{u}_t(\beta, \gamma)'$$

Equations (11) and (12) are the OLS regressions of  $\Delta P_t$  on  $P_{t-1}(\beta)$  for two subsamples where  $w_{t-1}(\beta) \leq \gamma$  and  $w_{t-1}(\beta) > \gamma$ . In the next step, the estimates  $(\hat{A}_1, \hat{A}_2, \hat{\Sigma})$  are utilized to yield the concentrated likelihood

$$\ln(\beta, \gamma) = L(\hat{A}_1(\beta, \gamma), \hat{A}_2(\beta, \gamma), \hat{\Sigma}(\beta, \gamma)) = -\frac{n}{2} \log \left| \hat{\Sigma}(\beta, \gamma) \right| - \frac{np}{2} \tag{13}$$

The maximum likelihood estimator  $(\hat{\beta}, \hat{\gamma})$  can be obtained by minimizing  $\log \left| \hat{\Sigma}(\beta, \gamma) \right|$  subject to the normalization imposed to the  $\beta$  and the constraints:

$$\pi_0 \leq n^{-1} \sum_{t=1}^n 1(P_t' \beta \leq \gamma) \leq 1 - \pi_0$$

Hansen and Seo (2002) used a grid search algorithm to obtain the MLE estimates of  $\beta$  and  $\gamma$ . The grid searching algorithm is summarized as follows:

*Step 1:* Construct a grid on  $[\gamma_U \gamma_L]$  and  $[\beta_L, \beta_U]$  based on the linear estimate of  $\beta$  and constraint above.

*Step 2:* Calculate  $\hat{A}_1(\beta, \gamma)$ ,  $\hat{A}_2(\beta, \gamma)$ , and  $\hat{\Sigma}(\beta, \gamma)$  for each value of  $(\beta, \gamma)$  on those grids.

*Step 3:* Search  $(\hat{\beta}, \hat{\gamma})$  as the values of  $(\beta, \gamma)$  on those grids which minimize  $\log \left| \hat{\Sigma}(\beta, \gamma) \right|$



Step 4: Estimate  $\widehat{\Sigma} = \widehat{\Sigma}(\widehat{\beta}, \widehat{\gamma})$ ,  $\widehat{A}_1 = \widehat{A}_1(\widehat{\beta}, \widehat{\gamma})$ ,  $\widehat{A}_2 = \widehat{A}_2(\widehat{\beta}, \widehat{\gamma})$ , and,  $\widehat{u}_t = \widehat{u}_t(\widehat{\beta}, \widehat{\gamma})$  as the final estimated parameters.

In the empirical application, the grid search procedure is carried out with 130 grid points. Once  $\beta$  and  $\gamma$  have been estimated, the null of linear cointegration is tested against the alternative of threshold cointegration by means of supremum Lagrange multiplier (SupLM) test following Andrews (1993) and Andrews and Ploberger (1994):

$$\text{Sup } LM^1 = \text{Sup}_{\gamma_L \leq \gamma \leq \gamma_U} LM(\widehat{\beta}, \gamma)$$

Since the asymptotic distribution of the test is not known, it is approximated by means of the residual bootstrap. In the empirical application, the bootstrap is done with 5000 replications. So, the model under null hypothesis is

$$\Delta p_t = A'_1 P_{t-1}(\beta) + u_t$$

with an alternative hypothesis,  $\Delta p_t = A'_1 P_{t-1}(\beta) \cdot d_{1t}(\beta, \gamma) + A'_1 P_{t-1}(\beta) \cdot d_{2t}(\beta, \gamma) + u_t$ .

Empirical results presented in this article are estimated using R algorithm. We have carried out the tests for all market pairs.

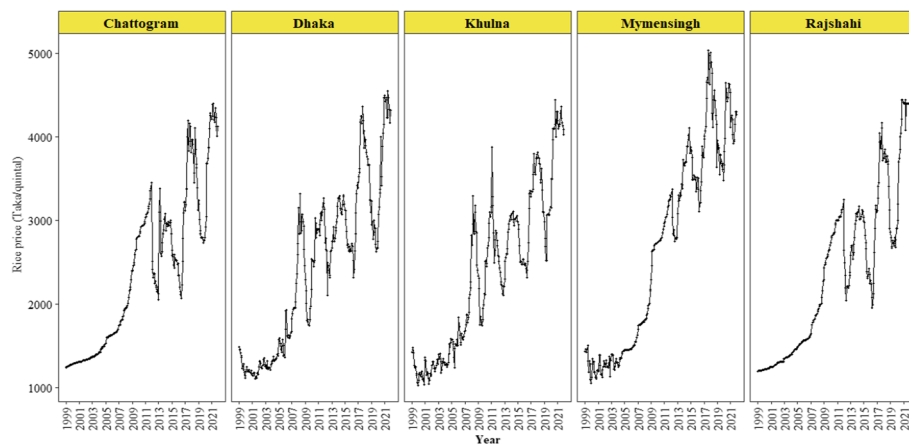
### The data and their time series properties

#### The data

The price data were collected from the Bangladesh Department of Agricultural Marketing (DAM) and cover the period from January 1999 to December 2021, with the data including the five main Bangladesh wholesale rice markets (Dhaka, Chittagong, Rajshahi, Khulna and Mymensingh). Although Bangladesh has three rice varieties, *Aus*, *Aman* and *Boro*, we follow Dawson and Dey (2002) and only use *Aman* and *Boro* prices to derive our time series for analysis. We justify this approach, as the production share of *Aus* is very small accounting for only about 5–10 percent of overall production. *Aman* paddy is harvested in November–December, while *Boro* paddy is harvested in May–June. Accordingly, we select the *Aman* price between November–April when *Boro* is not typically sold and the *Boro* price between May–October when *Aman* is not typically sold.

The DAM collects the agricultural food commodity prices in each district of Bangladesh. The collected prices are assumed to be representative of prices in all local markets, and their simple arithmetic average is the weekly wholesale price for the different places of that respective district. However, all price data are transformed into natural logarithms. Figure 2 presents a plot of monthly wholesale prices for the selected rice markets. The price pattern shows a close relationship or co-movement between the prices of all selected markets. Market selection for our analysis was based on the data availability that covers the whole geographical location as well as represents different divisions within Bangladesh. In addition, the selected markets include a mix of different demand and supply conditions in the country.

We provide descriptive statistics of wholesale price of different divisional markets in Table 1. The wholesale prices are quoted in Taka/quintal. The lowest average wholesale



**Fig. 2** Plots of monthly rice price (Taka/quintal) of five major markets in Bangladesh from 1999 to 2021

**Table 1** Descriptive statistics of major rice markets in Bangladesh

Statistic	Dhaka	Mymensingh	Rajshahi	Khulna	Chattogram
Mean	2439.7	2669.1	2381.8	2377.5	2408.8
Standard deviation	978.2	1192.1	949.5	943.8	921.1
Minimum	1107.0	1050.0	1197.0	1028.5	1240.0
Maximum	4550.0	5040.0	4450.0	4450.0	4400.0
Skewness	0.25	0.16	0.47	0.28	0.42
Kurtosis	-1.01	-1.41	-0.80	-1.04	-0.93
Ljung-Box test	268.11***	273.18**	269.73***	268.51***	267.76***

Ljung-Box test is the null hypothesis of time series containing an autocorrelation

\*\*\*Denotes statistical significance at the 1% critical level

price is 2377.5 Taka/quintal in Khulna, and the highest price is 2669.1 Taka/quintal in Mymensingh.

**Time series properties**

Looking at the plots of the data, it is clear that none of the series is stationary. Therefore, we use augmented Dickey-Fuller (ADF) and Phillip-Perron (PP) tests to determine the order of integration and the results are reported in Table 2. Our results show that wholesale market prices contain unit roots. Our tests indicate that all price series are non-stationary in levels but stationary in first differences. The optimum lag length for the ADF test was decided based on the Schwarz info criteria (SIC), and for PP test it was based on Newey-West (1994). Given that all the price series are integrated of order 1 denoted by I(1), we next proceed to test for cointegration.

**Empirical results and discussions**

**Linear cointegration test results**

The trace test ( $\lambda_{trace}$ ) and the maximum eigenvalue ( $\lambda_{max}$ ) test results are presented in Table 3 for models including a linear trend and without one. From the test results, it is seen that all market pairs contain one cointegrating rank (r). This means there is a one common factor that explains the long-run equilibrium relationship between all

**Table 2** Unit root test results

Prices series		Deterministic terms in test equations		First differences ( $\tau_{pw}$ )	Order of integration, I(d)
		$\tau_c$	$\tau_{c,t}$		
Dhaka	ADF	-0.633	-2.971	-10.687***	I(1)
	PP	-0.866	-3.360	-17.200***	I(1)
Mymensingh	ADF	-0.626	-3.078	-11.744***	I(1)
	PP	-0.646	-3.150	-16.400***	I(1)
Rajshahi	ADF	-0.555	-2.672	-9.078***	I(1)
	PP	-0.633	-2.810	-12.700***	I(1)
Khulna	ADF	-0.825	-3.324	-11.022***	I(1)
	PP	-0.996	-3.380	-16.700***	I(1)
Chattogram	ADF	-1.136	-3.276	-12.365***	I(1)
	PP	-0.942	-3.160	-13.700***	I(1)

\*\*\*Indicates that unit root in the first differences is rejected at 1% significant level;  $\tau_c$ ,  $\tau_{c,t}$ , and  $\tau_{pw}$  indicates tau-statistics of random walk with drift ( $\tau_c$ ), random walk with drift and slope ( $\tau_{c,t}$ ) and pure random walk ( $\tau_{pw}$ ) models, respectively; critical values are -3.44 (1%) and -2.87 (5%) with constant only model; -3.98 (1%) and -3.42 (5%) for a model with constant and trend; -2.58 (1%) and -1.95 (5%) for pure random walk model, respectively (Mackinnon 1996)

of the market pairings. All the market pairings exhibit a cointegrating relationship, which is consistent with the results of Dawson and Dey (2002), Hossain and Verke (2010), Siddique et al. (2008) and Rahman et al. (2021). In addition, residual diagnostics (Jarque–Bera normality test, Ljung–Box/Portmanteau autocorrelation test, and White heteroscedasticity test) indicate our models are well specified.

The long-run coefficients can be treated as long-run elasticity estimates (Table 4). The coefficients are close to unity, which suggests that the markets are almost perfectly integrated in the long run. The higher the values of the long-run elasticity in absolute terms, the more responsive are market prices in the long run. Our speed of the adjustment results shows that deviations from the long-run perturbation are corrected within 2 months, or in other words half of the deviations are corrected within a month. The adjustment coefficients ranged from 0.06 to 0.49. These relatively fast speeds of adjustment would suggest a low possibility of regional rice scarcity being a prolonged issue. Our results are consistent with the Dawson and Dey (2002), Hossain and Verke (2010), Siddique et al. (2008), Rahman et al. (2021), whose studies also found evidence of long-run cointegration in the period after liberalization.

**Causality test results**

To determine direction of price causality among our market pairs we used the weak exogeneity Wald test, as specified in methodology section, and the results are presented in Table 5. Of the ten cointegrated bivariate models, results indicate that only three market pairs (Chattogram-Dhaka, Khulna-Rajshahi and Chattogram-Khulna) exhibit a bidirectional price relationship. This shows interdependence between these two markets, or in other words the price in either market reacts to simultaneous shocks in the other market from its long-run equilibrium path. On the other hand, the remaining seven market pairs exhibit a unidirectional price relationship in which one market dominates the other in the price formation process. For example, in the Rajshahi-Dhaka pair, Rajshahi market Granger causes the price of Dhaka, so any intervention in the Rajshahi market will have

**Table 3** Johansen cointegration test results

Market pairings	Model 2 (no linear trend)		Model 3 (linear trend)	
	Test statistics	Critical values	Test statistics	Critical values
<i>Mymensingh-Dhaka</i>				
Trace statistics ( $\lambda_{\text{trace}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	23.88**	19.96	27.59**	25.32
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.66	9.24	9.09	12.25
Maximum eigenvalue statistics ( $\lambda_{\text{max}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	19.22**	15.67	21.50**	18.96
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.66	9.24	9.09	12.25
<i>Rajshahi-Dhaka</i>				
Trace statistics ( $\lambda_{\text{trace}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	24.59**	19.96	28.86**	25.32
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.37	9.24	7.7	12.25
Maximum eigenvalue statistics ( $\lambda_{\text{max}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	20.23**	15.67	23.15**	18.96
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.37	9.24	7.7	12.25
<i>Khulna-Dhaka</i>				
Trace statistics ( $\lambda_{\text{trace}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	29.06**	19.96	37.45**	25.32
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	1.92	9.24	9.76	12.25
Maximum eigenvalue statistics ( $\lambda_{\text{max}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	27.15**	15.67	27.69**	18.96
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	1.91	9.24	9.76	12.25
<i>Chattogram-Dhaka</i>				
Trace statistics ( $\lambda_{\text{trace}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	21.42**	19.96	28.78**	25.32
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.13	9.24	9.29	12.25
Maximum eigenvalue statistics ( $\lambda_{\text{max}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	19.29**	15.67	19.46**	18.96
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.13	9.24	9.29	12.25
<i>Rajshahi-Mymensingh</i>				
Trace statistics ( $\lambda_{\text{trace}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	24.92**	19.96	27.54**	25.32
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.49	9.24	5.36	12.25
Maximum eigenvalue statistics ( $\lambda_{\text{max}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	16.43**	15.67	22.18**	18.96
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.49	9.24	5.36	12.25
<i>Chattogram-Mymensingh</i>				
Trace statistics ( $\lambda_{\text{trace}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	24.07**	19.96	26.36**	25.32
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.34	9.24	9.74	12.25
Maximum eigenvalue statistics ( $\lambda_{\text{max}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	18.73**	15.67	20.62**	18.96
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.34	9.24	9.74	12.25
<i>Khulna-Mymensingh</i>				
Trace statistics ( $\lambda_{\text{trace}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	23.20**	19.96	30.41**	25.32
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.3	9.24	10.33	12.25
Maximum eigenvalue statistics ( $\lambda_{\text{max}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	20.89**	15.67	22.08**	18.96

**Table 3** (continued)

Market pairings	Model 2 (no linear trend)		Model 3 (linear trend)	
	Test statistics	Critical values	Test statistics	Critical values
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.3	9.24	10.33	12.25
<i>Khulna-Rajshahi</i>				
Trace statistics ( $\lambda_{\text{trace}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	20.82**	19.96	29.31**	25.32
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	1.99	9.24	9.42	12.25
Maximum eigenvalue statistics ( $\lambda_{\text{max}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	18.83**	15.67	19.89**	18.96
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	1.99	9.24	9.42	12.25
<i>Chattogram-Rajshahi</i>				
Trace statistics ( $\lambda_{\text{trace}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	26.90**	19.96	34.81**	25.32
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.33	9.24	7.41	12.25
Maximum eigenvalue statistics ( $\lambda_{\text{max}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	22.57**	15.67	25.40**	18.96
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	2.33	9.24	7.41	12.25
<i>Chattogram-Khulna</i>				
Trace statistics ( $\lambda_{\text{trace}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	25.99**	19.96	35.63**	25.32
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	1.91	9.24	11.88	12.25
Maximum eigenvalue statistics ( $\lambda_{\text{max}}$ )				
$H_0 : r = 0$ vs $H_1 : r \geq 1$	24.08**	15.67	23.75**	18.96
$H_0 : r \leq 1$ vs $H_1 : r \geq 2$	1.91	9.24	11.88	12.25

\*\*Indicates that the null hypotheses are rejected at 5% level of significant

**Table 4** Long-run elasticity and the speed of adjustment coefficients

Market pairs (right hand is normalized as explanatory market)	Long-run elasticity ( $\beta$ )	Speed of the adjustments	
		Market I ( $\alpha_1$ )	Market II ( $\alpha_2$ )
Mymensingh-Dhaka	0.823***	-0.321***	0.107**
Rajshahi-Dhaka	1.358***	0.008	0.079**
Khulna-Dhaka	0.869***	-0.108**	0.143***
Chattogram-Dhaka	0.814***	-0.091**	0.060*
Rajshahi-Mymensingh	2.182***	0.131***	-0.192***
Chattogram-Mymensingh	1.206***	-0.089**	0.019
Khulna-Mymensingh	1.627***	-0.158***	0.120***
Khulna-Rajshahi	0.695***	-0.115***	0.069*
Chattogram-Rajshahi	0.916***	-0.492***	0.109**
Chattogram-Khulna	1.976***	-0.112***	0.040

\*, \*\* and \*\*\* indicate that the null hypotheses are rejected at 1%, 5% and 10% level of significant, respectively; Market I and Market II indicate the first and second market in each market pair, for example in Mymensingh-Dhaka market pair, Mymensingh is Market I and Dhaka is Market II

an impact on the Dhaka market. Overall, the causality test results imply that although all markets are cointegrated, there are still some bottlenecks in regional trade between the markets. Therefore, intervention in any one market does not necessarily have an immediate price impact in other markets.

**Table 5** Market dominance using Wald test in the VECM

Market pairs	Causality test		Results
	$H_0: \alpha_1 = 0$ vs	$H_0: \alpha_2 = 0$ vs	
	$H_1 : \alpha_1 \neq 0$	$H_1 : \alpha_2 \neq 0$	
Mymensingh-Dhaka	1.102	2.767**	Unidirectional
Rajshahi-Dhaka	0.465	6.320***	Unidirectional
Khulna-Dhaka	2.055	8.077***	Unidirectional
Chattogram-Dhaka	4.393***	4.198***	Bidirectional
Rajshahi-Mymensingh	1.112	4.196***	Unidirectional
Chattogram-Mymensingh	6.132***	0.533	Unidirectional
Khulna-Mymensingh	3.529**	1.810	Unidirectional
Khulna-Rajshahi	6.895***	3.643**	Bidirectional
Chattogram-Rajshahi	9.653***	0.737	Unidirectional
Chattogram-Khulna	6.169***	3.596**	Bidirectional

\*\* and \*\*\* indicate the null hypotheses are rejected at 1% and 5% level of significant, respectively

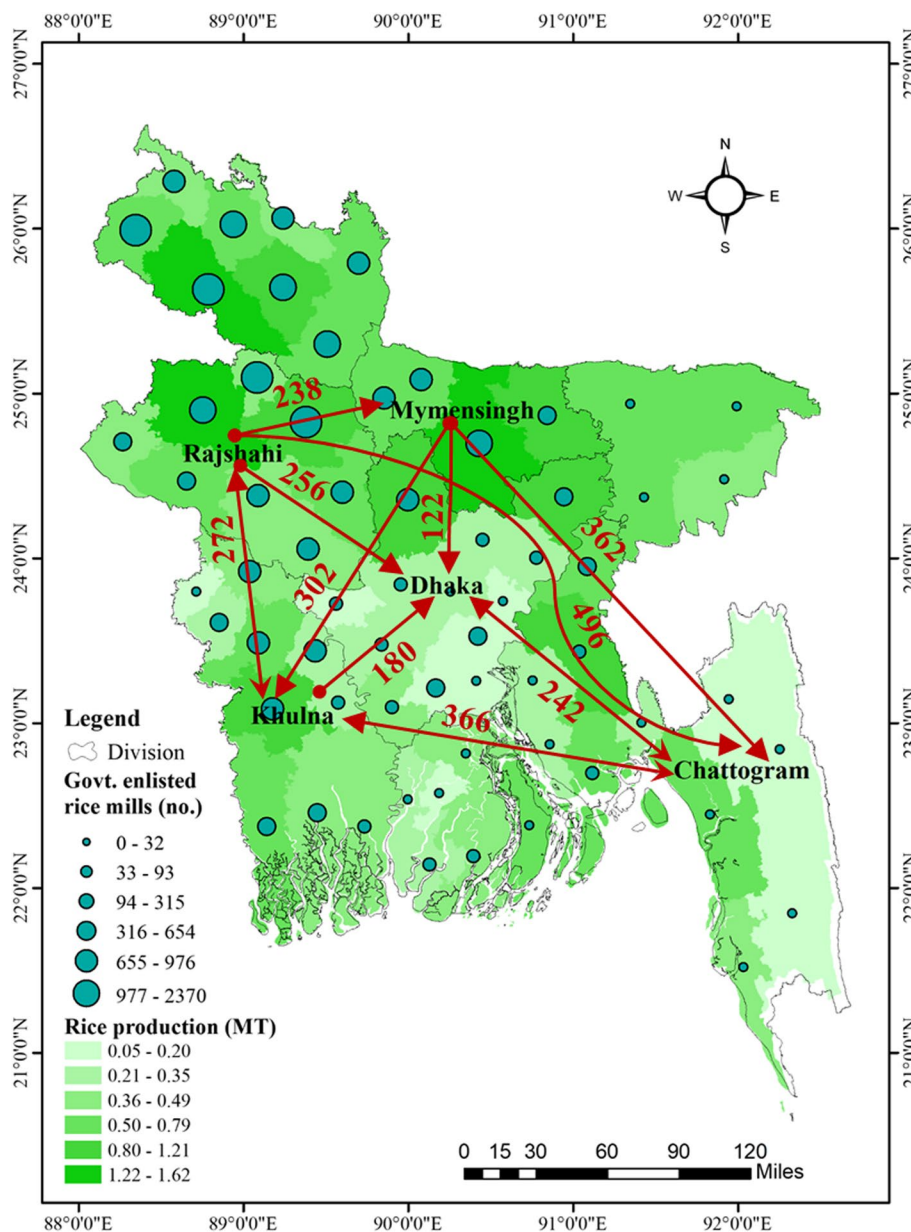
In terms of market interdependence, two main conclusions emerge. One is that the Rajshahi and Mymensingh markets play a leadership role (Fig. 3). Second, only the Dhaka market adjusts to price changes emanating from the other markets (Mymensingh, Rajshahi, Khulna and Chittagong). The geographical locations of these two markets (Rajshahi and Dhaka) could be the main reason for these price relationships. Rajshahi is one of the top rice-producing divisions in Bangladesh and supplies a major portion of overall rice production to other regions in the country. In addition, Indian rice imports—legal and illegal—trade through the Rajshahi region, and this might explain its importance as a price leader. On average, Bangladesh imports around 5 percent of its total rice consumption Thus, any price policy intervention in the Rajshahi market would likely effect prices in other markets. This result is very interesting and helps shed a light on price relationships between Bangladesh and India.

In contrast, Mymensingh, Chittagong, Khulna and Rajshahi markets Granger-cause the Dhaka price. Dhaka is the biggest deficit region with the highest demand in terms of total rice consumption. Given this is a demand-driven market likely explains why it is responsive to price changes in other markets.

**Results of threshold cointegration**

Table 6 shows the results pertaining to the threshold cointegration. The *p* values were computed using a residual bootstrap procedure as in Hansen and Seo (2002) using 5000 simulation replications. To select the lag length of the VAR, we used the Akaike information criteria and the Bayesian information criteria and found in all the cases an optimal lag of one. The null hypothesis of linear cointegration is rejected for all market pairs in favor of threshold cointegration at 10% significance levels. Our results are robust for all the market pairs with 2 lags (*k* = 2).

Giving our findings that market pairings exhibit nonlinear relationships, we estimate two-regime TVECM for each market pairing. The estimated long-run elasticity, speed of adjustment coefficients and threshold parameters are presented in Table 7. For illustrative purposes, consider the Rajshahi-Dhaka market pair. The estimated



**Fig. 3** Direction of price causality among major rice market pairs of Bangladesh. Note: Single and double arrows represent the unidirectional and bidirectional causality, respectively. For example, 'Mymensingh → Dhaka' represents the price of Dhaka market is determined by the price of Mymensingh. Conversely, 'Khulna ↔ Chattogram' represents the price of Khulna market is determined by the price of Chattogram market and vice versa. The numerical value of the graph shows the market distances in kilometer

long-run cointegrating parameter is 0.97 implying that a 10 percent increase in the price in Rajshahi brings about a 9.7 percent increase in the long-run Dhaka price. The value of the SupLM<sup>1</sup> test is 18.950 ( $k = 1$ ) and the  $p$  value is 0.057 for the residual bootstrap, supporting the threshold cointegration hypothesis.

The long-run price transmission elasticity in all the market pairs shows significance at the 5% level. The highest long-run coefficient is found for the Mymensingh-Dhaka market pair (1.098), followed by Chattogram-Rajshahi, Chattogram-Khulna,



**Table 6** Threshold cointegration test

Market pairs	Test particulars	SupLM <sup>1</sup> ( $\hat{\beta}$ & $\hat{\gamma}$ ) test	
		K = 1	K = 2
Mymensingh-Dhaka	SupLM test statistic value	18.950*	29.712***
	Critical values (0.05 level)	19.381	24.789
	Residual bootstrap <i>p</i> -value	0.057	0.002
Rajshahi-Dhaka	SupLM test statistic value	18.401*	23.288*
	Critical values (0.05 level)	19.072	24.212
	Residual bootstrap <i>p</i> -value	0.065	0.073
Khulna-Dhaka	SupLM test statistic value	19.227*	24.354*
	Critical values (0.05 level)	19.735	25.179
	Residual bootstrap <i>p</i> -value	0.059	0.068
Chattogram-Dhaka	SupLM test statistic value	18.914*	22.558*
	Critical values (0.05 level)	19.226	24.322
	Residual bootstrap <i>p</i> -value	0.055	0.089
Rajshahi-Mymensingh	SupLM test statistic value	20.160**	22.873*
	Critical values (0.05 level)	18.571	23.871
	Residual bootstrap <i>p</i> -value	0.025	0.074
Khulna-Mymensingh	SupLM test statistic value	19.062*	23.899*
	Critical values (0.05 level)	19.262	24.625
	Residual bootstrap <i>p</i> -value	0.054	0.066
Chattogram-Mymensingh	SupLM test statistic value	21.426**	25.886**
	Critical values (0.05 level)	18.695	23.800
	Residual bootstrap <i>p</i> -value	0.012	0.018
Khulna-Rajshahi	SupLM test statistic value	23.103***	22.456*
	Critical values (0.05 level)	19.187	24.495
	Residual bootstrap <i>p</i> -value	0.006	0.096
Chattogram-Rajshahi	SupLM test statistic value	17.617*	27.394***
	Critical values (0.05 level)	18.739	24.028
	Residual bootstrap <i>p</i> -value	0.082	0.009
Chattogram-Khulna	SupLM test statistic value	22.793***	23.615*
	Critical values (0.05 level)	19.246	24.494
	Residual bootstrap <i>p</i> -value	0.007	0.069

\*, \*\* and \*\*\* indicate the null hypotheses are rejected at 1%, 5% and the 10% level of significance

Khulna-Rajshahi, Chattogram-Dhaka, Khulna-Dhaka, Rajshahi-Dhaka, Chattogram-Mymensingh, Rajshahi-Mymensingh and Khulna-Mymensingh, pairings, respectively. Analysis shows that the long-run slope parameters of the Mymensingh-Dhaka, Chattogram-Rajshahi and Chattogram-Khulna market pairs all exceeded unity, indicating that these markets are closely integrated. For example, the long-run elasticity of the market pairs Mymensingh-Dhaka is estimated at 1.098, indicating that on average, price changes in Dhaka are transmitted to Mymensingh in equal magnitude. Turning to the other market pairs, we note that the long-run price transmission coefficient is highest between production consumption regions. In some cases, the degree of price transmission is likely determined by the distance and mode of transportation between markets.

The speed of adjustment in regime 2 indicates how fast the market price adjusts to perturbations from long-run equilibrium (Table 7). All markets converge to the long-run equilibrium at different speeds. For Rajshahi and Dhaka, Dhaka prices

**Table 7** Normalized long-run elasticity and speed of adjustment coefficients at threshold vector error correction model

Market pairs (right-hand side market is normalized as explanatory market)	Long-run elasticity ( $\beta$ )	Threshold value ( $\gamma$ )	Speed of the adjustment in regime 2 ( $R^2$ )	
			Market I ( $\alpha_1$ )	Market II ( $\alpha_2$ )
Mymensingh-Dhaka	1.098**	-4.111*	-0.020	0.290**
Rajshahi-Dhaka	0.968**	-3.959*	-0.019	-0.209**
Khulna-Dhaka	0.970**	3.508*	-0.739***	0.304
Chattogram-Dhaka	0.974**	4.291*	0.014	-0.249***
Rajshahi-Mymensingh	0.868**	-3.570**	-0.019	-0.121*
Chattogram-Mymensingh	0.874**	-4.129*	-0.009	0.027
Khulna-Mymensingh	0.863**	-3.649**	-0.044*	-0.006
Khulna-Rajshahi	0.991**	3.151***	-0.118**	-0.041
Chattogram-Rajshahi	1.004**	-2.067*	-0.396***	-0.149**
Chattogram-Khulna	1.001**	-4.132***	-0.128***	0.030

\*, \*\* and \*\*\* indicate the null hypotheses are rejected at 1%, 5% and 10% level of significance; Market I and Market II indicate the first and second market in the market pairs, for example Mymensingh-Dhaka market pair, Mymensingh market is the market I and Dhaka market is the Market II. Eicker-White standard errors are used to get the significance level of the speed of adjustments

adjust toward their long-run equilibrium with almost 20.9% ( $p < 0.05$ ) of total adjustment occurring within a month. Khulna market prices (production region) adjust most quickly to Dhaka price changes with almost 73.9% ( $p < 0.01$ ) of total adjustment occurring within a month. The fact that adjustment primarily takes place in Khulna is consistent with the notion that trade between markets, which are geographically close, should be easier, leading to large and quick price adjustments to bring markets in line with their long-run equilibrium.

Our results show that the weakest adjustment speeds occur between market regions with surplus rice production. This may be attributed to relatively poor transportation systems compared to those linking production and deficit/demand regions. Higher adjustment speeds were found between surplus-deficit market pairs.

The threshold value is a proxy for transaction costs, and transportation cost is an important part of overall transaction costs (Osebeyo and Aye 2014). The average transportation cost for carrying one quintal of rice from market one market to another is 151.5 Taka depending upon market distances, which is below the overall transaction cost. The magnitude of threshold depends on regional demand, mode of transportation, distance, and market-dependent consumers. The transportation modes in the country are Mahindra tractor, pickup, and small truck for local transport, and truck for long-distance trade. Based on the estimated threshold parameter, our model is divided into two regimes. Recall that regime 1 (the non-adjustment band) contains absolute price deviations from the long-run equilibrium that are below the threshold. In this case, we would expect no price adjustments to perturbations in long-run equilibrium. In other words, no cointegrating relationship will exist in that regime. On the other hand, in regime 2 (the adjustment band), contains absolute price deviations from long-run equilibrium that are bigger than the threshold parameter. In this case, there will be a cointegrating relationship and prices will realign. For illustrative purposes, consider the Rajshahi-Dhaka market pair.

Here, like the linear VECM, the statistical significance of the speed of the adjustment in the TVECM reveals that Rajshahi is the dominant market and the Dhaka market adjusts to Rajshahi price changes. The estimated threshold is 3.595 Taka,<sup>2</sup> and when absolute price deviations from the Rajshahi-Dhaka long-run equilibrium exceed 3.959 Dhaka prices will adjust to bring the long-run relationship back into line. Almost 2/3 of the price adjustment will occur within one month. However, when the absolute price deviation is less than 3.595, and we are in regime 1, our theoretical model suggests that no price adjustments would occur. In general, our results are consistent with our a priori economic based expectations.

## Conclusions

Market integration studies that have ignored the role of transaction costs have received much criticism in the recent literature (see Barret and Li 2002; Meyer 2004; Goodwin and Piggot 2001; Ben-Kaabia and Jose 2007; Sanogo and Maliki 2010). Modeling transaction cost is of particular importance when analyzing market integration in developing countries. To address this issue, we employ the two-regime threshold cointegration model of Meyer (2004), which is an extension of Hansen and Seo (2002), to analyze spatial integration among Bangladesh rice markets. Our results provide strong supporting evidence of the presence of threshold effects. Our results show that large price deviations from long-run equilibrium are corrected within 2–3 months, or in other words, half to two-thirds of the price deviations are corrected within 1 month. Thus, although the price adjustment process is relatively slow compared with developed markets, it appears that private sector trade can be relied upon to transfer price signals between Bangladesh rice markets in the long run. These results are consistent with the linear cointegration results presented in Dawson and Dey (2002). However, our results shed additional light on the issue of Bangladesh rice market integration. Importantly, we find evidence of threshold effects for some of our market pairings. In these cases, transaction costs prevent market prices from adjusting to relatively small price shocks. For example, with respect to the Rajshahi-Dhaka market pairing, we showed that only when the absolute price difference is bigger than 3.595 Taka, would adjustment to price shocks take place.

Thus, our results provide important policy implications for Bangladesh rice markets; namely, that policies aimed at reducing transaction costs (for example, investing in roads and communications, information delivery center etc.) should be encouraged to further improve market integration and efficiency. Of course, although increased market efficiency is a desirable outcome, further study would be required to clearly identify and quantify the costs and benefits of reducing transaction costs.

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<sup>2</sup> Local currency.

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### Author contributions

First author developed the concept, prepared data for analysis, contributed to the writing of the manuscript, worked on the estimation procedure and provided critical review. Second and third authors worked in the concept development, contributed to the writing of the manuscript, and worked on the estimation procedure. Fourth and fifth authors provided interpretation and critical review; sixth and seventh authors conducted data analysis and generated tables and figures. The last author provided interpretation and critical review. All authors read and approved the final manuscript.

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## Declarations

### Competing interests

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