Measuring Market Integration in the Presence of Threshold Effect: The Case of Bangladesh Rice Markets

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Abstract

Spatial price integration among five major Bangladesh rice markets is examined in the presence of threshold effects to account for the impact of transaction costs in the price adjustment process. Hansen and Seo (2002) threshold cointegration tests and threshold vector error correction models confirm the presence of threshold effects. Results highlight the importance of directing policy goals towards reducing transaction cost to engender greater pricing efficiency in Bangladesh rice markets.

Keywords: market integration, rice markets, transaction cost, Bangladesh

1. Introduction

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 1980's by 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. Moreover, over this same period, transportation infrastructure – roads and communication and mobile networks – have been developed. Hence, in the wake of these reforms 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, different trade barriers and prohibitive transactions 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) in the market integration have been done. The seminal work of Ravallion (1986) showed that there is limited market integration in rice markets in Bangladesh. While Goletti *et al.*, (1995) conclude 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. Dawson and Dey (2002) showed that Bangladesh rice markets were perfectly integrated following the trade liberalization reforms. The authors used a vector auto-regressive error correction model (VECM) to test the Law of One Price (LOP) within the central-regional market, following Ravallion (1986). 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 recent literature such as Enders and Siklos (2001), Enders and Granger (1998), Goodwin and Piggot (2001), Meyer (2004), Sarno *et al.*, (2004) it is argued that the standard cointegration framework is mis-specified if the adjustment process is nonlinear and asymmetric. This is likely the case if TC is significant.

The factors that might contribute to higher TC are inadequate infrastructure, transportation bottlenecks, lack of market information, information asymmetry, market power, menu cost and so on. These kinds of factors are common in developing countries' agricultural markets such as Bangladesh and pose serious challenges to policy makers. So, estimating the threshold in the price adjustment from one market to another or from one level to another in the supply chain should be a rule rather than an exception, especially in the context of developing countries. The present study is an attempt to sequentially test first, whether the domestic rice markets in Bangladesh are integrated using Johansen and Jesulius (1992) method and then testing causality to infer about market dominance. Our study is different from the study of Dawson and Dey (2002) in the sense that we relax the assumption of regional-central market hypothesis. Secondly, our study differs because we are testing threshold cointegration by using Hansen and Seo (2002) methodology in which 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 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. First, it uses 'state of art methodology' to test for spatial market integration by considering the role of transaction costs and secondly, it is the first study of its kind to examine market integration with respect to post-reform era of highly liberalized Bangladesh rice markets.

The remainder of the paper is organized as follows. The next section presents the econometrics methodology of linear cointegration (Johansen-Jesulius, 1992) along with a causality test for market dominance followed by a conceptual basis and econometrics estimation of threshold cointegration and threshold vector error correction model. The data are explained in section 3. Section 4 presents the results and discussions. The last section concludes.

2. Econometrics methodology

2.1 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:

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

Where P_t includes all n variables of the model which are $I(1)$, the Π, Γ_i and Φ are parameter matrices to be estimated, D_t is a vector with deterministic elements (constant, trend) and ω_t is a vector of random error follow Gaussian process. If ΔP_t is $I(0)$ then Π will be a zero matrix except when a linear combination of the variables in P_t is stationary. If rank Π = r = K, the variables in levels are stationary meaning that no integration exists; if rank Π = r =0, meaning that all the elements in the adjustment matrix has value zero, therefore, none of the linear combinations are stationary. According to the Granger representation theorem (1987) that when 0< rank (Π=r) < K, there are r cointegrating vectors. For example if rank (Π = r) = 1, there is single cointegrating vector or one linear combination which is stationary such that the coefficient matrix Π can be decomposed into $\Pi = \alpha\beta'$ where α is the vector of loading factor and β is the cointegrating vector in where $\beta' P_{t-1}$ is $I(0)$. Johansen method is to estimate Π matrix from an unrestricted VAR and to test whether we can reject the restrictions implied by the reduced rank of Π. There are two methods of testing for reduced rank (Π), 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.

2.2 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 Lutkephol, 1996). The underlying principle is that if α matrix in cointegration matrix (Π) has a complete column of zeros, then no casual relationship exist, because there is no cointegrating vector in that particular block. For pair-wise causal relationship, it can be written in the following equation (2)

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

In the equation (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 bi-directional causality and the last two imply uni-directional 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 equation (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 $\Delta P_{2,t}$.

2.3 Threshold cointegration

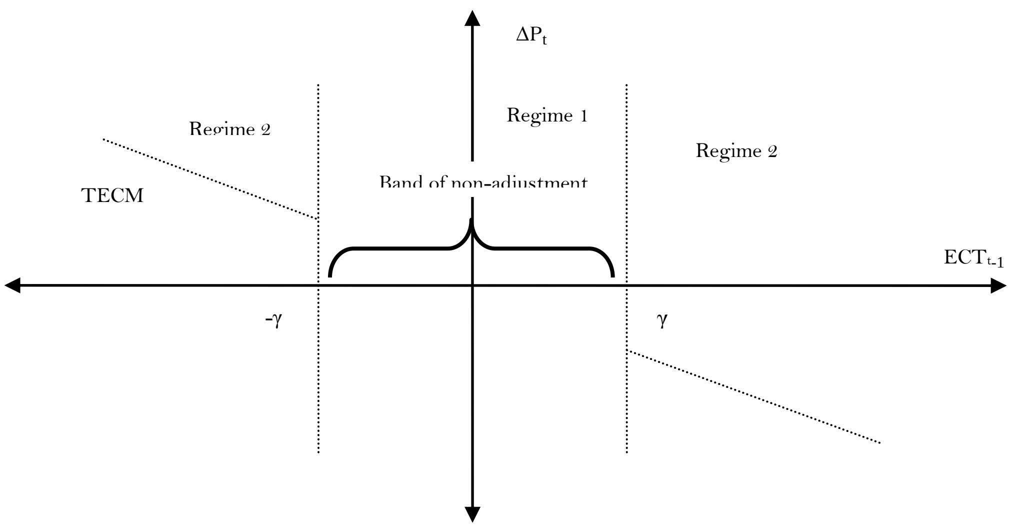
The concept of threshold cointegration was introduced first by Balke and Fomby (1997) as a way of combining cointegration and non-linearity. The authors present the possibility that movements towards the long-run equilibrium might not occur in every time period, due to the presence of TC. After that, the limitation of linear cointegration has been often discussed in recent literature because neglecting of TC may inhibit price integration across spatially separated markets (for example, see Barret and Li, 2002; Fackler and Goodwin, 2001; Goodwin and Piggot, 2001; Abdulai, 2000, 2002; Goodwin and Harper, 2000). Goodwin and Piggott (2001) have used a threshold error correction model to estimate spatial integration in US corn and soybean markets. Ben-Kaabia and Jose (2007) have estimated price transmission between vertical stages of the Spanish lamb market using a threshold model. Sanogo and Maliki (2010) have analysed the rice market integration between Nepal and

India applying a threshold autoregressive model. The conceptual basis of the analysis, along with the econometrics estimation procedures is explained below.

One implicit assumption of the linear model like Johansen and Jesulius (1992) and Engel and Granger (1987) is that adjustment of prices induced by deviations from the long-term equilibrium is a continuous and a linear function of the magnitude of deviations. Thus, every small deviation will always lead to an adjustment. This assumption might mislead the results because it ignores the affect of TC in price adjustment.

Considering the role of TC into account one could use a threshold cointegration model in which the price adjustment could differ based on the magnitude of the deviations from its long-run equilibrium. The speed of adjustment can be different if the deviations are above or below the specific threshold –which would proxy the size of TC.

Figure 1: The effect of transactions costs in the price adjustment



In Figure 1, the price adjustment (ΔPt) is considered to be a function of deviations from the long-run equilibrium (ECT) which can be represented by a two regime threshold vector error correction model (TVECM). We proceed by estimating the two regime TVECM proposed by Hansen and Seo (2002). Here, the regime is defined based on only one threshold (γ) and therefore if the *absolute* price deviation from the long-run equilibrium is bigger than the threshold (γ), the price transmission process is defined by regime 2, while in the case of smaller deviations and thus falling within a '*band of no adjustment*' from the long-run equilibrium, the price transmission process is defined as regime 1 (see Figure 1). Therefore, to estimate a two-regime threshold vector error correction model, the threshold γ must also be estimated. For this, a variant of the Hansen and Seo (2002) model is presented below. 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 2x1 cointegrating vector β 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:

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

Where

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

In equation 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 (MDS) 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 (β, A, Σ) 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:

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

Where, γ represents the threshold parameter. The model in equation (5) may also be written as

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

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

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

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 γ (in other words, whether w_{t-1} is above or below γ) allow all coefficients – except the cointegrating vector β – to switch between these two regimes.

The threshold effect exist if $0 \prec P(w_{t-1} \leq |\gamma|) \prec 1$, otherwise the model belongs to the linear cointegration form. We impose this constraint assuming that $\pi_0 \prec P(w_{t-1(\beta)} \leq |\gamma| \prec (1-\pi_0)$ and by setting $\pi_0 \succ 0$ as a trimming parameter equal to 0.05 (Andrews, 1993)[2] in the empirical estimation. Further it we ensure that the indicator function represented by equations (7) and (8) contain enough sample variation for each choice of γ. The likelihood function of the model in equation (6) under the assumption of *iid* Gaussian error u_t, has the following form:

[2] 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

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

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

The MLE of ($\widehat{A}_1, \widehat{A}_{2,} \widehat{\beta}, \widehat{\Sigma}, \widehat{\gamma}$) are obtained by maximizing the $\ln(A_1, A_2, \beta, \Sigma, \gamma)$. This is achieved by first holding (β, γ) fixed, and computing the constrained MLE for (A_1, A_2, Σ) using the OLS regression and are as follows.

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

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

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

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

Equations (11) and (12) are the OLS regressions of ΔP_t on $P_{t-1}(\beta)$ for two sub-samples where $w_{t-1}(\beta) \le \gamma$ and $w_{t-1}(\beta) \succ \gamma$. In the next step, the estimates $(\widehat{A}_1, \widehat{A}_2, \widehat{\Sigma})$ are utilized to yield the concentrated likelihood

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

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

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

Hansen and Seo (2002) used a grid search algorithm to obtain the MLE estimates of β and γ. The grid searching algorithm is summarized as follows

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

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

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

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

In the empirical application, the grid search procedure is carried out with 130 grid points. Once β and γ 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):

$$SupLM^1 = \underset{\gamma_L \le \gamma \le \gamma_U}{SupLM(\hat{\beta},\gamma)}$$

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

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

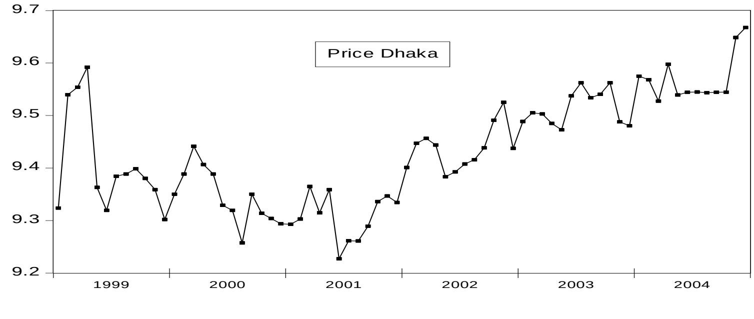
3. The data and their time series properties

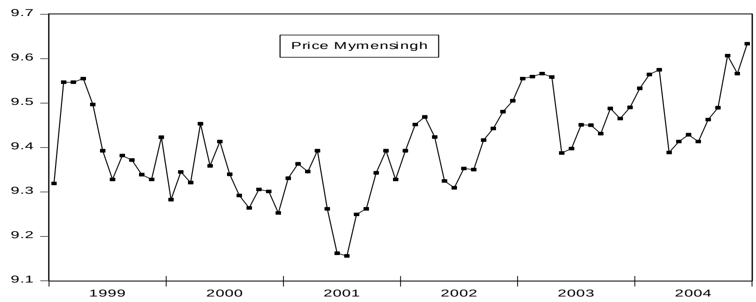
3.1 The data

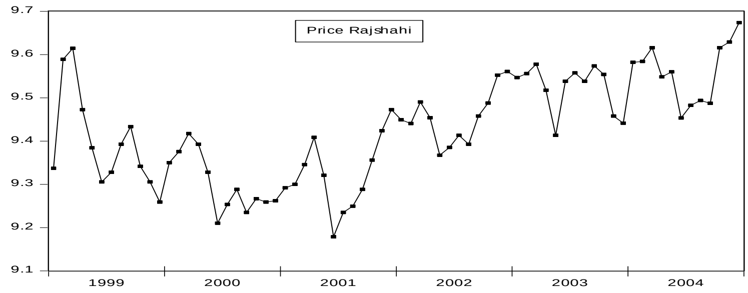
The data cover the period from January 1999 to December 2004 with five main wholesale markets (Dhaka, Chittagong, Rajshahi, Khulna and Mymensingh) of rice in Bangladesh being taken from the Department of Agricultural Marketing (DAM), the people's republic of the government of Bangladesh. Time series econometric price transmission analysis need to allow for the fact that, over time, the domestic price of a commodity at different spatial markets commonly trend together in nominal terms (Dawson and Dey, 2002), that is why the data series considered for the analysis are in nominal terms. The sample period was selected on the basis of data availability. Following Dawson and Dey (2002) only the prices of *Aman* and *Boro* are used to derive the price series for analysis. The three rice varieties in Bangladesh are *Aus*, *Aman* and *Boro*. However, the production share of *Aus* is very small with about 5-10 percent. *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.

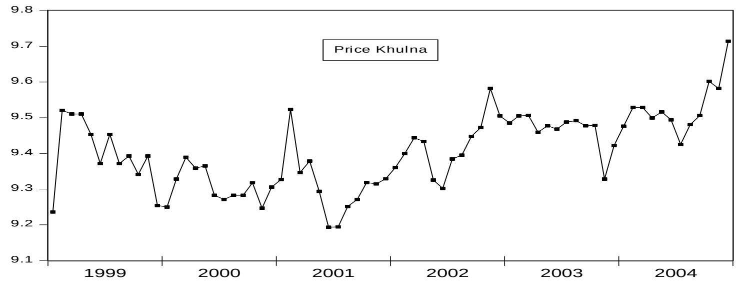
The DAM collects the agricultural food commodity prices by its permanent headquarters in each district of Bangladesh. The collected prices are assumed to be representative of prices in all local markets, and their simple arithmetic average is the weekly wholesale price for the different places of that respective district. However, all price data are transformed into logarithmic forms. Figure 2 presents a plot of wholesale prices for the selected rice markets. The price pattern shows that 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 in Bangladesh.

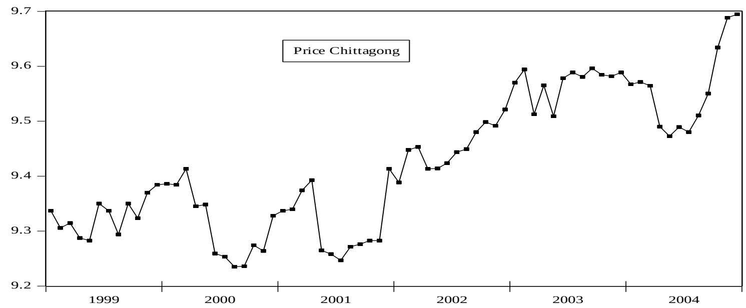
Figure 2: Plots of five markets price series (in log form)











3.2 Time series properties

Looking at the plots of the data, it is clear that none of the series is stationary. Therefore, we test time series data properties to determine the order of integration. We perform this by using ADF and PP tests and the results are reported in Table 1. Our test indicates 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.

Table 1: Unit root tests results

Prices series	Deterministic terms in test equations		First differences (τ_{pw})	Order of integration, I(d)
	τ_c	$\tau_{c,t}$		
Dhaka				
ADF	-1.901	-2.895	-9.929***	I(1)
PP	-1.708	-2.686	-11.539***	I(1)
Mymensingh				
ADF	-2.576	-3.084	-9.804***	I(1)
PP	-2.575	-3.133	-10.205***	I(1)
Rajshahi				
ADF	-2.043	-3.033	-8.415***	I(1)
PP	-1.842	-2.722	-9.071***	I(1)
Khulna				
ADF	-2.565	-3.343	-10.993***	I(1)
PP	-2.565	-3.343	-12.517***	I(1)
Chittagong				
ADF	-0.592	-2.453	-9.272***	I(1)
PP	-0.637	-2.637	-9.219***	I(1)

Notes: *** indicates that unit root in the first differences is rejected at 1% significant level; τ_c, $\tau_{c,t}$ and τ_{pw} indicates *tau*-statistics of random walk with drift (τ_c), random walk with drift and slope ($\tau_{c,t}$) and pure random walk (τ_{pw}) models respectively; Critical values are -3.525 (1%), and -2.903 (5%) with constant only model; -4.093 (1%), and -3.474 (1%) for a model with constant and trend; -2.598 (1%) and -1.945 (5%) for pure random walk model respectively (MacKinnon, 1996).

4. Empirical results and discussions

4.1 Linear cointegration test results

The trace test (λ_{trace}) and the maximum eigenvalue (λ_{max}) test results are presented in Table 2. From the test results, it is seen that all market pairs contains one cointegrating rank (r), meaning that this gives a number of stationary linear combinations of the price pairs. For example, the Dhaka and Mymensingh price shows one cointegrating rank that means that there is a one common factor for which the price of both the markets has a long-run equilibrium relationship. All the market pairs show that the cointegration relationship exists which is consistent with the results of Dawson and Dey (2002). We have tested the models with no linear trend and with linear trend denoted by model 2 and model 3 respectively and have found the same conclusion. When we perform the cointegration in a vector error correction model framework we also conduct the residual analysis (normality test, Ljung Box/Portmanteau test, white heteroscedasticity test) and found in all the case there are no problem of misspecification of estimated models.

Table 2: Johansen cointegration test results

Market pairings	**Model 2 (no linear trend)**		**Model 3 (linear trend)**	
	Test statistics	Critical values	Test statistics	Critical values
Mymensingh-Dhaka				
Trace statistics (λ_{trace})				
$H_0: r=0$ vs $H_1: r\geq 1$	21.449**	20.261	21.332**	15.495
$H_0: r\leq 1$ vs $H_1: r\geq 2$	2.368	9.165	2.278	3.841
Maximum eigenvalue statistics (λ_{max})				
$H_0: r=0$ vs $H_1: r\geq 1$	19.081**	15.892	19.055**	14.265
$H_0: r\leq 1$ vs $H_1: r\geq 2$	2.368	9.165	2.278	3.841
Rajshahi-Dhaka				
Trace statistics (λ_{trace})				
$H_0: r=0$ vs $H_1: r\geq 1$	26.734**	20.162	26.611**	15.495
$H_0: r\leq 1$ vs $H_1: r\geq 2$	2.824	9.165	2.765	3.841
Maximum eigenvalue statistics (λ_{max})				
$H_0: r=0$ vs $H_1: r\geq 1$	23.909**	15.892	23.845**	14.265
$H_0: r\leq 1$ vs $H_1: r\geq 2$	2.824	9.165	2.765	3.841
Khulna-Dhaka				
Trace statistics (λ_{trace})				
$H_0: r=0$ vs $H_1: r\geq 1$	16.841	20.262	16.665**	15.495

$H_0: r \leq 1$ vs $H_1: r \geq 2$	-	-	1.512	3.841
Maximum eigenvalue (λ_{max})				
$H_0: r = 0$ vs $H_1: r \geq 1$	15.206	15.892	15.152**	14.265
$H_0: r \leq 1$ vs $H_1: r \geq 2$	-	-	1.512	3.841
Chittagong-Dhaka				
Trace statistics (λ_{trace})				
$H_0: r = 0$ vs $H_1: r \geq 1$	22.978**	20.262	21.603**	15.495
$H_0: r \leq 1$ vs $H_1: r \geq 2$	1.113	9.165	0.202	3.841
Maximum eigenvalue statistics (λ_{max})				
$H_0: r = 0$ vs $H_1: r \geq 1$	21.864**	15.892	21.402**	14.265
$H_0: r \leq 1$ vs $H_1: r \geq 2$	1.113	9.165	0.202	3.841
Rajshahi-Mymensingh				
Trace statistics (λ_{trace})				
$H_0: r = 0$ vs $H_1: r \geq 1$	25.461**	20.262	25.394**	15.495
$H_0: r \leq 1$ vs $H_1: r \geq 2$	2.956	9.165	2.903	3.841
Maximum eigenvalue statistics (λ_{max})				
$H_0: r = 0$ vs $H_1: r \geq 1$	22.496**	15.892	22.492**	14.265
$H_0: r \leq 1$ vs $H_1: r \geq 2$	2.965	9.165	2.903	3.841
Chittagong-Mymensingh				
Trace statistics (λ_{trace})				
$H_0: r = 0$ vs $H_1: r \geq 1$	17.422	20.262	15.749**	15.494
$H_0: r \leq 1$ vs $H_1: r \geq 2$	-	-	0.297	3.841
Maximum eigenvalue statistics (λ_{max})				
$H_0: r = 0$ vs $H_1: r \geq 1$	15.503	15.892	15.452**	14.265
$H_0: r \leq 1$ vs $H_1: r \geq 2$	-	-	0.297	3.841
Khulna-Mymensingh				
Trace statistics (λ_{trace})				
$H_0: r = 0$ vs $H_1: r \geq 1$	23.444**	20.262	23.219**	15.495
$H_0: r \leq 1$ vs $H_1: r \geq 2$	2.578	9.165	2.423	3.841
Maximum eigenvalue statistics (λ_{max})				

$H_0: r=0$ vs $H_1: r \geq 1$	20.866**	15.892	20.797**	14.265
$H_0: r \leq 1$ vs $H_1: r \geq 2$	2.578	9.165	2.422	3.841
Khulna-Rajshahi				
Trace statistics (λ_{trace})				
$H_0: r=0$ vs $H_1: r \geq 1$	21.870**	20.262	21.668**	15.495
$H_0: r \leq 1$ vs $H_1: r \geq 2$	3.183	9.165	3.087	3.841
Maximum eigenvalue statistics (λ_{max})				
$H_0: r=0$ vs $H_1: r \geq 1$	18.687**	15.892	18.582**	14.265
$H_0: r \leq 1$ vs $H_1: r \geq 2$	3.183	9.165	3.087	3.841
Chittagong-Rajshahi				
Trace statistics (λ_{trace})				
$H_0: r=0$ vs $H_1: r \geq 1$	28.606**	20.262	26.915**	15.495
$H_0: r \leq 1$ vs $H_1: r \geq 2$	1.485	9.165	0.472	3.841
Maximum eigenvalue statistics (λ_{max})				
$H_0: r=0$ vs $H_1: r \geq 1$	27.121**	15.892	26.443**	14.265
$H_0: r \leq 1$ vs $H_1: r \geq 2$	1.485	9.165	0.473	3.841
Chittagong-Khulna				
Trace statistics (λ_{trace})				
$H_0: r=0$ vs $H_1: r \geq 1$	19.648	20.261	18.487**	15.495
$H_0: r \leq 1$ vs $H_1: r \geq 2$	-	-	0.655	3.841
Maximum eigenvalue statistics (λ_{max})				
$H_0: r=0$ vs $H_1: r \geq 1$	18.024**	15.892	17.832**	14.265
$H_0: r \leq 1$ vs $H_1: r \geq 2$	1.624	9.165	0.655	3.841

Note: ** indicates that the null hypotheses are rejected at 5% level of significant

The long-run coefficients can be treated as long-run elasticity estimates (Table 3). The coefficients are close to unity which testifies that the markets are cointegrated almost perfectly. The higher the values of the long-run elasticity in the absolute terms, higher the market price are responsive in the long-run. In the case of the speed of the adjustment, results show that the deviations from the long-run perturbation are corrected within two months or in other words half of the deviations are corrected within a month, the non-zero values of the adjustment coefficient ranged from 0.24 to 0.59. The relatively faster speeds of adjustment minimize the possibility of the spatial scarcity of rice. Our results are consistent with the one of Dawson and Dey (2002) who also found evidence of long-run cointegration in the period after liberalization.

Table 3: Long-run elasticity and the speed of adjustment coefficients

Market pairs (right hand is normalized as explanatory market)	Long-run elasticity (β)	Speed of the adjustments	
		Market I (α_1)	Market II (α_1)
Mymensingh-Dhaka	0.847***	-0.526***	0.096
Rajshahi-Dhaka	1.017***	-0.500**	0.242*
Khulna-Dhaka	0.879***	-0.587**	0.058
Chittagong-Dhaka	1.222***	-0.095	0.371***
Rajshahi-Mymensingh	1.299***	-0.029	0.533***
Chittagong-Mymensingh	1.547***	0.024	0.261***
Khulna-Mymensingh	1.075***	-0.205	0.452**
Khulna-Rajshahi	0.932***	-0.308**	0.324**
Chittagong-Rajshahi	1.163***	-0.063	0.449***
Chittagong-Khulna	1.524***	-0.013	0.325***

Notes: ***, ** & * indicates that the null hypotheses are rejected at 1%, 5% and 10% level of significant; Market I and Market II indicates the first and second market in each market pairs, for example in Mymensingh-Dhaka market pair-Mymensingh is Market I and Dhaka is Market II.

4.2 Causality test results

To determine direction of price causality among our market pairs we used the weak exogeneity Wald test (test specification is specified in methodology section) and the results are presented in Table 4. Of the ten cointegrated bivariate models, results indicate that only two market pairs (Rajshahi-Dhaka and Khulna-Rajshahi) exhibit a bi-directional 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 eight market pairs exhibit a unidirectional price relationship in which one market dominates the other in the price formation process. For example, in the Chittagong-Dhaka pair, Chittagong market Granger causes the price of Dhaka, so any intervention in the Chittagong market will have an impact in Dhaka market. The overall results from the causality tests imply that although all bivariate model shows that markets are cointegrated, there are still some bottlenecks in the interconnectedness between the markets. In that case intervention in any of the markets does not necessarily immediately pass to others markets.

Table 4: Market dominance using Wald test in the VECM

Market pairs	Causality test		Results
	$H_0 : \alpha_1 = 0$ vs $H_1 : \alpha_1 \neq 0$	$H_0 : \alpha_1 = 0$ vs $H_1 : \alpha_1 \neq 0$	
Mymensingh-Dhaka	14.127***	0.589	Uni-directional
Rajshahi- Dhaka	7.329***	2.774*	Bi-directional
Khulna-Dhaka	9.075***	0.122	Uni-directional
Chittagong-Dhaka	1.409	18.356**	Uni-directional
Rajshahi-Mymensingh	0.042	15.242***	Uni-directional
Chittagong-Mymensingh	0.277	14.963***	Uni-directional
Khulna-Mymensingh	1.662	9.073***	Uni-directional
Khulna-Rajshahi	3.546*	3.911**	Bi-directional
Chittagong-Rajshahi	0.617	18.425***	Uni-directional
Chittagong-Khulna	0.049	15.333***	Uni-directional

Note: ***, ** and * indicates the null hypotheses are rejected at 1%, 5% and the 10% level of significant

In terms of market interdependence, two main conclusions emerge. One is that the Chittagong market plays a leadership role (Figure 3). Second, only the Mymensingh market adjusts price from the price changes in all other markets. The geographical locations of these two markets (Chittagong and Mymensingh) could be the main reason. Chittagong is the main and largest sea port in Bangladesh from where the major portion of the total import is taking place (the average rice imports of Bangladesh is 5-10 percent of its total consumption). The possibility to have legal or illegal trade of paddy/milled rice trade with Myanmar, the neighboring country, might be a factor explaining the Chittagong markets' importance in terms of price leadership. In that case any intervention in Chittagong markets would pass to all other markets. This result is very interesting in terms of further investing the cointegration relationship between the Chittagong market prices and the price of rice exporting country (or Myanmar) to Bangladesh. Moreover, investigating the possibility to have an illegal rice importation from Myanmar would shed a light on it.

On the other hand, all markets (Dhaka, Chittagong, Khulna and Rajshahi) Granger cause Mymensingh price. Mymensingh is the nearest market from the capital city Dhaka (which is the biggest demanding market in Bangladesh in terms of total rice consumption) (the distance between Dhaka and Mymensingh is 193 km) and it is the only market that does not represent divisional prices. Therefore, these two factors might be responsible for making this market a follower of other markets.

4.3 Results of threshold cointegration

Table 5 shows the results pertaining to the threshold cointegration. The p-values were computed by 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, a lag of one. The null hypothesis of linear cointegration is rejected in all market pairs except for the market pair of Rajshahi-Dhaka in favour of threshold cointegration at the 10% significance level. But out of 9 markets pairs, five are rejected at the 5% significance level. To check the robustness of our results we also estimated all the market pairs with 2 lags and in this case found threshold cointegration for all market pairs at the 10% significant level.

Table 5: Threshold cointegration test

Market Pairs	Test particulars	$SupLM^1(\hat{\beta} \& \hat{\gamma})$ test	
		k=1	k=2
Mymensingh-Dhaka	SupLM test statistic value	11.927	16.214
	Critical values (0.05 level)	12.238	16.485
	Residual bootstrap *p*-value	0.056*	0.054*
Rajshahi-Dhaka	SupLM test statistic value	11.129	14.189
	Critical values (0.05 level)	13.762	16.799
	Residual bootstrap *p*-value	0.107	0.092*
Khulna-Dhaka	SupLM test statistic value	10.610	13.895
	Critical values (0.05 level)	12.026	16.525
	Residual bootstrap *p*-value	0.078*	0.093*
Chittagong-Dhaka	SupLM test statistic value	16.148	14.258
	Critical values (0.05 level)	10.393	15.694
	Residual bootstrap *p*-value	0.006***	0.066*
Rajshahi -Mymensingh	SupLM test statistic value	15.104	16.973
	Critical values (0.05 level)	12.695	16.635
	Residual bootstrap *p*-value	0.019**	0.045**
Khulna - Mymensingh	SupLM test statistic value	14.424	13.194
	Critical values (0.05 level)	12.804	18.347
	Residual bootstrap *p*-value	0.025**	0.155
Chittagong - Mymensingh	SupLM test statistic value	12.689	14.059
	Critical values (0.05 level)	10.121	15.529
	Residual bootstrap *p*-value	0.020**	0.065*
Khulna-Rajshahi	SupLM test statistic value	11.177	16.102
	Critical values (0.05 level)	13.109	17.270
	Residual bootstrap *p*-value	0.092*	0.069*
Chittagong-Rajshahi	SupLM test statistic value	11.444	14.564
	Critical values (0.05 level)	12.183	17.188
	Residual bootstrap *p*-value	0.063*	0.096*
Chittagong-Khulna	SupLM test statistic value	11.513	18.069
	Critical values (0.05 level)	10.723	15.455
	Residual bootstrap *p*-value	0.040**	0.023**

Note: ***, ** and * indicates the null hypotheses are rejected at 1%, 5% and the 10% level of significance

The estimated long-run elasticity and the threshold parameters are presented in table 6. Based on the estimated threshold parameter, the model is divided into two regimes. Recall that, regime 1 (the band of non-adjustment) is defined when the absolute price deviations from the long-run equilibrium 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 (regime of adjustment), when the absolute price deviation from long-run equilibrium is bigger than the threshold parameter, there will be a cointegrating relationship and prices will realign. For illustrative purposes, consider the Mymensingh-Dhaka market pair. The estimated long-run cointegrating parameter is 0.74 implying that a 10 percent increase in the price in Dhaka brings about 7 percent increase in the price of Mymensingh in the long-run. The value of the SupLM[1] test is 11.927 (k=1) and the p-value is 0.056 for the residual bootstrap supporting the threshold cointegration hypothesis. Here, like the linear VECM, the statistical significance of the speed of the adjustment in the TVECM reveals that Dhaka is the dominant market and the Mymensingh market adjusts from the price changes in the Dhaka market. The estimated threshold is 1.303 Taka[3] which identifies the two regimes in the threshold model. So, when the absolute price deviation from Mymensingh and Dhaka long-run equilibrium exceeds 1.303 Mymensingh prices will adjust to bring the long-run relationship back in line. This adjustment will account for 64 percent or almost 2/3 of the price deviation within one month. However, when the absolute price deviation is less than 1.303, and we are in regime 1, our theoretical model suggests that no price adjustments would occur. In general our results are consistent with our model based a priori expectations and error correction terms are insignificant in regime 1.

[3] Local currency

Table 6: 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 (β)	Threshold value (γ)	Speed of the adjustment in regime 2 (R2)	
			Market I (α_1)	Market II (α_1)
Mymensingh-Dhaka	0.787**	1.303*	-0.64***	0.04
Rajshahi-Dhaka	0.979**	0.137*	-0.47***	0.14
Khulna-Dhaka	0.909**	0.791*	-0.44***	0.10
Chittagong-Dhaka	1.206**	-2.078***	-0.16	0.42***
Rajshahi-Mymensingh	1.209**	-1.962**	-0.26	0.42***
Chittagong-Mymensingh	1.591**	-5.641*	0.04	0.35***
Khulna-Mymensingh	1.082**	-0.702**	0.20	-0.38**
Khulna-Rajshahi	0.842**	1.479*	0.16	0.66
Chittagong-Rajshahi	1.162**	-1.616*	0.10	0.62**
Chittagong-Khulna	1.516**	-4.812**	-0.02	0.54**

Notes: ***, ** and * indicates the null hypotheses are rejected at 1%, 5% and the 10% level of significance; Market I and Market II indicates 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.

5. Conclusions

Market integration studies that have ignored the role of transaction costs have received much criticism in recent literature (see Barret and Li, 2002; Meyer, 2004; Goodwin and Piggot, 2001; Ben-Kaabia and Jose, 2007; Sanogo and Maliki, 2010). Modelling transaction costs is of particular importance when analyzing market integration in developing countries. To address this issue, we employ the two-regime threshold cointegration model of Hansen and Seo (2002) to analyse spatial integration among Bangladesh rice markets over the 1999 to 2004 period. 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 two-three months, or in other words half to two third of the price deviations are corrected within one 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 the markets. 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 to adjust to relatively small price shocks. For example, Mymensingh-Dhaka pair, we showed that only when the absolute price difference is bigger than 1.303 Taka, the price adjustment will occur.

Thus, our results provide important policy implications for Bangladesh rice markets, namely that polices aimed at reducing transaction costs (for example, investing in roads and communications, information delivery center etc.) should be encouraged to further improve market 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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