World and Bangladesh Rice Market Integration: An Application of Threshold Cointegration and Threshold Vector Error Correction Model (TVECM)

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Abstract

This paper analysed the market integration between international and domestic market of rice covering the period of agricultural trade liberalization in Bangladesh. Policy makers are interested to know whether price changes in the world market are transmitted to the domestic market when there are some non-trade factors other than the trade liberalization that might affect the rice markets to be integrated. We used a threshold cointegration and threshold vector error correction model (TVECM) of Hansen and Seo (2002) to account for the affects of transaction cost (which is very substantial in the case of the developing countries) in the market integration. We found from Supremum Lagrange Multiplier (SupLM) test that Bangladesh rice market is partially integrated with the counterpart `world market`. Only one-third of the world price changes are got transmitted to domestic market. We also have found that the presence of the transaction cost affects in the rice market integration. So, it is clear that trade liberalization bring its expected outcomes for markets to be integrated but liberalization alone is not enough to explore the maximum benefit, there are some other factors such as non-tariff barriers and trade facilitations that should be seriously taken into consideration by the policy makers.

Keywords: Market integration, World market, Bangladesh, threshold cointegration

JEL code: Q13, Q17, Q18

1. Introduction

The Bangladesh government has undertaken substantial policy reforms with respect to agricultural trade liberalization, the exchange rate and the import–export procedures. The key objective of these reforms were to increase market integration that will results to the consumer welfare gain and food security for a food deficit country like Bangladesh. Hence, one would expect to observe greater integration between domestic and world agricultural markets because of the policy reforms at domestic and border levels. Highly integrated markets allow for efficient transmission of price signals across markets, thus prevent inefficiencies and misallocation of resources. Markets that are not integrated can convey inaccurate price information, leading to misguided policy decisions and a misallocation of scarce resources. There are three reasons for a lack of market integration such as imperfect competition, different trade barriers and prohibitive transactions costs (Sexton *et al.* 1991).

However, concern arises whether trade liberalization alone is sufficient for markets to be integrated. For example, markets may not be well integrated because of high transaction costs due to poor transportation and communications infrastructure, non-tariff barriers etc. Therefore, investigating long-run price relationship, which measures the degree of market integration of staple commodity \`rice\` between Bangladesh and the world, is crucial for policy makers in formulating optimum policies at domestic and at border levels.

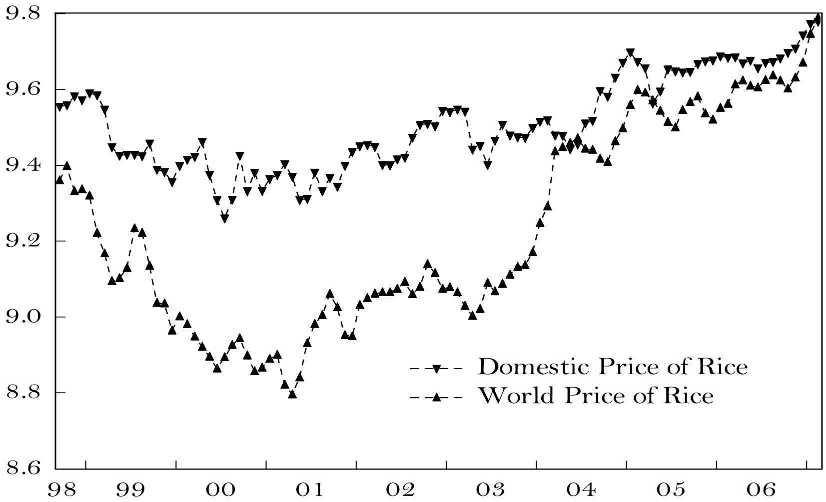
One of the most contentious debates has been whether or not the implementation of the market reforms especially agricultural trade liberalization reform at the border in the developing countries improved price transmission between agricultural commodity markets at the foreign and domestic scenes (Shahidur, 2004). There are two views exist in literature. First, whether the price transmission has been improved from world to the domestic producers\` prices after the reform of agricultural trade liberalization by exporting countries, hence welfare gain of the producers and Second, whether world price has been passed-through to the domestic consumer, hence, gain in consumer welfare and poverty reduction and food security. Peter (2008) found that the cointegration relationship exists between world and domestic Indonesian rice market and found that the elasticity of 0.369 meaning that markets are partially cointegrated. Yavapolkul et al. (2006) observed that the developed and developing countries' rice and wheat markets during the post-Uruguay trade negotiations were only partially cointegrated which means that Uruguay round of the trade negotiation did not improve the world markets to be fully integrated. Baffes and Bruce (2003) presented that only few of the Latin American countries are integrated after the agricultural trade liberalization. Although studies (Ravallion, 1986; Dawson and Dey, 2002) have examined the domestic spatial rice markets integration but to date no studies conducted on the domestic Bangladesh and international rice market except Alam at el. (2012). They used the linear cointegration approach to investigate the market integration between domestic and international rice markets without considering the effects of transaction costs in the market integration study. In recent literature such as Enders and Siklos (2001), Meyer (2004), Sarno *et al.*, (2004), the standard cointegration has been highly criticized. Goodwin and Piggott (2001) have used a threshold cointegration in US corn and soybean markets and found the presence of threshold effects. Sanogo and Maliki (2010) have analysed the market integration between Nepal and India using threshold model and confirmed the presence of threshold effects. However, the evidences from literature are diverse and vary irrespective of methodology (linear cointegration and threshold cointegration) used and importing or exporting country, small or large country case. Apart from the trade liberalization, there are many factors that could influence the market integration outcome. For example- non-trade barriers, the policies of domestic and world markets, the poor communication and infrastructure that leads to higher transaction costs, competition and so on. Little previous research has been investigated the issue especially for the case of Bangladesh domestic and world market by using ex-post methodology of time series econometrics except some *ex-ante* studies (Nabil *et al.*, 2006; Raihan and Razzaque, 2007) using computable general equilibrium (CGE) model. Dorosh (2001) showed that the agricultural trade liberalization reforms in Bangladesh contributed to the country's overall food security and have saved from an unprecedented price hikes that could worsen the poverty situation, especially during the period of domestic supply shocks in 1997. Ninno and Dorosh (2003) showed that how private sector imports contributed to price stabilization and saved from further deterioration in household's purchasing power and calorie consumption following the 1998 flood. The contribution of the private sector trade and agricultural trade liberalization was examined mainly within a static and descriptive framework.

Given this backdrop, the main objective of this paper is to examine whether domestic and international rice markets are integrated and the effect of transaction costs in the market integration process using a sophisticated modelling approach. This is very important to the Bangladesh as well international policy makers and practitioners because the study provides whether markets are integration after the trade liberalization and whether trade liberalization (by tariff reduction) alone is sufficient for the markets to be integrated.

The paper is organized as follows. Following introduction, data are presented in section 2. Section 3 provides the time series properties of data and econometric model i e., threshold cointegration and threshold vector error correction model as a modelling framework. The results and discussions are presented in section 4. Last section concludes and provides policy recommendation.

2. Data

The monthly rice price data are taken from the food outlook of the FAO and the global information and early warning system (GIEWS) of FAO. The exchange rate data are collected from the 'Economic Trends' of Bangladesh Bank. The monthly FOB Thai 100% B prices are used as world price because Bangladesh imports this type of rice. Although, there are some changes of the exporting countries of rice to Bangladesh, the present study used `Thai price` as a world prices for two main reasons. First, Thailand has been the largest rice exporter over the last couple of decades and may be regarded as a price leader in the world rice market. Secondly, In addition, we assume that Thai and Indian rice prices are highly correlated because a recent study by Yavapolkul et al. (2006) found that major importing countries like Thailand and India among others are integrated, therefore, supporting our assumption. So, although Bangladesh imports rice from India, but using Thai price as a proxy for India so far a good choice. The data period covers September 1998 to February 2007. The data periods are chosen because of data availability and also to capture the period of the highest pace of agricultural trade liberalization in Bangladesh. The evolution of the Bangladesh domestic and the world market prices is presented in Figure 1. The co-movement of two price series someway roughly indicates that there might be an existence of long-run equilibrium relationship. The spread between these two prices has been squeezed during the later time period. The graph indicates that prices are more stable in the domestic market than in the world market. However, the number of observations after this period is not considerably enough to model the relationships considering a break point hence we model the dynamic relationships for the whole period in a single model.



3. Time Series Properties of Data and Econometric Model

3.1 Time Series Properties of Data

Since data are time series, the world price and domestic price of rice are tested for their non-stationary. Therefore, we conduct unit root test by the standard augmented Dickey-Fuller (ADF) (1979) and the Philips Perron (PP) (1989) test. The ADF unit root test with an optimal lag length determined by the Akaike information criterion (AIC), Schwarz Bayesian information criterion (SBC) and Lagrangian multiplier (LM) criteria and is used in the following form:

$$\Delta P_{i,t} = c + \rho P_{i,t-1} + \sum_{j=1}^{k-1} \Gamma_i \Delta P_{i,t-j} + \beta T + \varepsilon_{i,t} \qquad (1)$$

Where $P_{i,t}$ is the respective price series, Δ is first difference operator, T is the time trend and ε_t denotes white noise error term. Equation (1) tests the null of a unit root ($\rho = 0$) against a mean-stationary alternative ($\rho \neq 0$). The term $\Delta P_{i,t-j}$ is a lagged first difference to accommodate serial correlation in the errors.

When the time series data are subject to both a deterministic trend (T) and an exogenous shock that causes a structural break, the ADF test tends to under-reject (Perron, 1989).

Therefore, we also test the presence of a unit root using Philips Perron (1989) in the following specification.

$$\Delta P_{i,t} = c + \beta\left\{t - \frac{T}{2}\right\} + \rho P_{i,t-1} + \varepsilon_{i,t} \qquad (2)$$

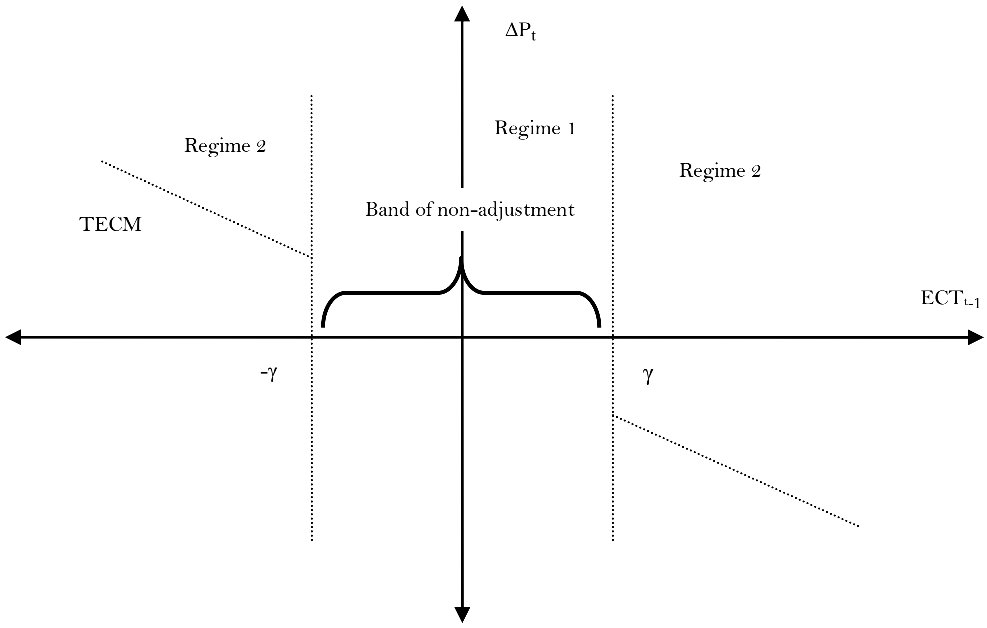
Where $P_{i,t}$ is respective time series, $\left\{t - \frac{T}{2}\right\}$ is the time trend and where T is the sample size, $\varepsilon_{i,t}$ is the error term. This procedure, in fact, uses a non-parametric adjustment to the Dickey–Fuller test statistics and allows for dependence and heterogeneity in the error term.

3.2 Threshold cointegration model

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.

Figure 1: The effect of transactions costs in the 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.

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:

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

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

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

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

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

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 $0 \prec P(w_{t-1} \le |\gamma|) \prec 1$, otherwise the model belongs to the linear cointegration form. We impose this constraint assuming that $\pi_0 \prec P(w_{t-1(\beta)} \le |\gamma| \prec (1-\pi_0)$ and by setting $\pi_0 \succ 0$ as a trimming parameter equal to 0.05 (Andrews, 1993)† 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:

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

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

The MLE of $(\hat{A}_1, \hat{A}_{2,}\hat{\beta}, \hat{\Sigma}, \hat{\gamma})$ 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.

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

† 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

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

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

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

Equations (11) and (12) are the OLS regressions of ΔP_t on $P_{t-1}(\beta)$ for two sub-samples where $w_{t-1}(\beta)\leq\gamma$ 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

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

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:

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

Hansen and Seo (2002) used a grid search algorithm to obtain the MLE estimates of 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 $(\widehat{\beta},\widehat{\gamma})$ as the values of (β,γ) on those grids which minimize $\log\left|\widehat{\Sigma}(\beta,\gamma)\right|$ and **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 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.

4. Results and Discussions

An initial consideration must be to test the logged data for non-stationarity and to determine if the data generating process is difference or trend stationary. It is also important to establish the number of unit roots that a series contains when testing for cointegration. For two non-stationary series to be cointegrated they must be integrated of the same order. Both Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests were employed to determine the stationarity of the data. The optimal number of augmenting lags for the models was determined by using the Akaike information criterion (AIC), the Schwarz Bayesian information criteria (SBC) and Lagrangian multiplier (LM) criteria. Table 1 presents the results of the ADF and PP unit root test for each price series. Based on the critical values reported by MacKinnon (1996), both tests rejected that the price series were stationary in levels (both with and without a trend term). In addition, the ADF and PP tests failed to reject that that the price series were stationary in first differences. Thus, in summary, ADF tests and PP tests indicated that both price series contain a single unit root and therefore may be regarded as difference stationary.

Table 1: Unit root test results

Variables	levels with constant only		levels with constant and trend		Decision
	ADF	PP	ADF	PP	
World price	-0.232	0.127	-3.092	-2.577	Non-stationary
Domestic wholesale price	-0.118	-0.631	-2.735	-2.261	Non-stationary
	First difference with constant only		First difference with constant and trend		
	ADF	PP	ADF	PP	
World price	-6.914	-6.759	-7.409	-7.157	Stationary
Domestic wholesale price	-6.595	-10.64	-7.144	-15.014	Stationary

Notes: 1. Lag length for ADF tests are decided based on Akaike information criterion (AIC). **2.** Maximum Bandwidth for PP tests are decided based on Newey-West (1994) **3.** Critical values are -2.89 (5%), and -3.49 (1%) with constant only model; -3.45 (5%), and -4.05 (1%) for a model with constant and trend (MacKinnon, 1996)

We have found that the prices are in the same order of integration that is I (1), Then we proceed to test the threshold cointegration. Given that the Bangladesh domestic rice price and world rice price are integrated of the same degree I (1), cointegration model can be used to determine if a long-run relationship exists between these prices and whether transaction costs effect the price adjustment process in the long run. We used Hansen and Seo (2002) threshold cointegration and threshold vector error correction model to test the market integration.

Table 2: Threshold cointegration test results

Test particulars	$SupLM^0$ ($\hat{\gamma}$ *estimated*)	
	K=2	Observation in %
SupLM test stat	25.145	
Fixed regressors bootstrap p-value	0.016**	Regime 1 (typical regime)
Residual bootstrap p-value	0.032**	45%
Threshold parameter ($\hat{\gamma}$)	3213	Regime 2 (a-typical regime)
Cointegrating vector (($\hat{\beta}$)	1	55%

Remind that our model is two-regime threshold vector error correction model. Regime 1 is defined as `band of non-adjustment` when the absolute price deviations from the long-run equilibrium are below the threshold. Here no cointegrating relationship exists. Regime 2 is a `regime of adjustment`- when the absolute price deviation from long-run equilibrium is bigger than threshold value. Here cointegrating relationship will exists. Our results show that the adjustment is observed only in the regime 2 but not in regime 1. We have found that Bangladesh and world rice markets are integrated as supported by SupLM test. In our model, only gamma is estimated but beta=1 is given on the basis of the priori known cointegrating vector. From the results we can reject the null of linear cointegration with 1.6% level using the p-value of fixed regressors bootstrap and 3.2% level using the p-value of residual bootstrap. Results of the SupLM tests can be found in Table 2. For getting the probability values we have done 5000 simulations. The regimes are distributed more or less evenly and is 45% and 55% percent of observations. We have found that the value of estimated threshold is 3213. The estimated threshold identifies the two regimes in the threshold model. When the absolute price deviation from world and Bangladesh domestic long-run equilibrium exceeds 3213, the Bangladesh price will adjust to bring the long-run relationship back in line. So in this case, the domestic Bangladesh market and international market is integrated. This adjustment will account for 37 percent or 1/3 of the price deviation within one month. However, when the absolute price deviation is less than 3213, no adjustment will taken place and there will be no market integration. The estimated full model for regime 1 and regime 2 is given below. In sum, our results are consistent with the theoretical model described earlier.

The estimated coefficients of the threshold vector error correction model and the Eicker-White standard errors of the co-efficient are given as follows

<u>Regime 1:</u>

$$\begin{bmatrix} \Delta P_t^D \\ \Delta P_t^W \end{bmatrix} = \begin{bmatrix} 284.292^{**} \\ (113.11) \\ -83.975 \\ (121.939) \end{bmatrix} + \begin{bmatrix} 0.008 & 0.136 \\ (0.131) & (0.126) \\ -0.044 & 0.373^{**} \\ (0.095) & (0.115) \end{bmatrix} \begin{bmatrix} \Delta P^D{}_{t-1} \\ \Delta P^W{}_{t-1} \end{bmatrix} + \begin{bmatrix} 0.163 & 0.303^{**} \\ (0.192) & (0.138) \\ -0.314^{**} & -0.021 \\ (0.124) & (0.116) \end{bmatrix} \begin{bmatrix} \Delta P^D{}_{t-2} \\ \Delta P^W{}_{t-2} \end{bmatrix} + \begin{bmatrix} -0.071 \\ (0.051) \\ 0.080 \\ (0.086) \end{bmatrix} [ECT_{t-1}] +$$

$$\begin{bmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \end{bmatrix}, \quad if\ |ECT_{t-1} = \gamma| \leq 3213$$

<u>Regime 2:</u>

$$\begin{bmatrix} \Delta P_t^D \\ \Delta P_t^W \end{bmatrix} = \begin{bmatrix} 1588.008^{**} \\ (532.616) \\ -460.258 \\ (399.858) \end{bmatrix} + \begin{bmatrix} -0.016 & -0.0915 \\ (0.157) & (0.086) \\ 0.224 & 0.449^{**} \\ (0.181) & (0.134) \end{bmatrix} \begin{bmatrix} \Delta P^D{}_{t-1} \\ \Delta P^W{}_{t-1} \end{bmatrix} + \begin{bmatrix} 0.084 & -0.145^{*} \\ (0.102) & (0.076) \\ 0.323 & -0.034 \\ (0.194) & (0.175) \end{bmatrix} \begin{bmatrix} \Delta P^D{}_{t-2} \\ \Delta P^W{}_{t-2} \end{bmatrix} + \begin{bmatrix} -0.378^{**} \\ (0.126) \\ 0.112 \\ (0.092) \end{bmatrix} [ECT_{t-1}] +$$

$$\begin{bmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \end{bmatrix}, \quad if\ |ECT_{t-1} = \gamma| \geq 3213$$

Notes: Eicker-white s. e. are in the parentheses: *, and **means significant at 10%, and 5% levels

Our results are in the line of our expectation. One could expect greater market integration during the period of market liberalization. As a results of market liberalization policy, we find this strong evidence that is Bangladesh and world rice markets are integrated. This integration is partial however. So, being a net food (here mainly rice) import country, greater market integration is obviously beneficial only when the world market price is stable and less volatile. Also in the case when the world market price is smaller than the border price. In that case, the policy makers in Bangladesh should be concern on the results of partial integration which could results to consumer welfare loss. However, it can be concluded that agricultural trade liberalization policies in Bangladesh that virtually eliminated tariffs so far not enough to fully integrated the domestic market with the world counterpart. This means that the government should design additional policies for greater market integration especially non-tariff barriers for reducing transaction costs, that will promote market efficiency and food security given that the world market prices are stable and less volatile. But, on the other hand, when the international prices are volatile, greater market integration can lead to welfare loss and bring the threat of food security. So government should take into account this unexpected outcome of the market integration.

5. Conclusions and policy recommendations

The studies of market integration 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). Taking into account of the transaction costs is important when analysing the market integration in developing country like Bangladesh. To address this issue, we use two-regime threshold cointegration model of Hansen and Seo (2002) to analyse domestic and international market integration. Our results provide strong supporting evidence of the presence of threshold effects. However, our results shed additional light on the issue of Bangladesh rice market integration. Importantly, we find evidence of threshold effects. In these cases transaction costs prevent market prices to adjust to relatively small price shocks. 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. So, from a policy standpoint, if Bangladeshi government implements only policies related to removing tariff barriers without taking into account the non-tariffs related barriers the effectiveness of such policies for greater market integration would likely to be compromised. Of course although increased market efficiency is a desirable outcome, further study would be required to clearly distinguish the significance of trade and non-trade related factors including the significance of transaction cost so that optimum policy can be undertaken to maximize the gain from agricultural trade liberalization.

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