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Asymmetry Price Transmission in the Deregulated Rice Markets in Bangladesh: Asymmetric Error Correction Model

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ABSTRACT

There is a widely held belief among public consumers that rice prices are manipulated in Bangladesh. This manipulation may have led to price asymmetry in the vertical chain of Bangladesh rice markets. This paper is an attempt to investigate the existence of asymmetry between wholesale and retail rice prices in Bangladesh. Maximum likelihood estimation (MLE) based cointegration test was applied to determine long-run equilibrium relationship. We examine whether the wholesale market dominates the retail market—in terms of price discovery and price leadership—or vice versa. Finally, we analyze whether the wholesale-retail price relationship is asymmetric with respect to price increases and price decreases. To test the asymmetric price transmission we used the asymmetric error correction-EG approach. Our results show that wholesale and retail prices are cointegrated, and wholesale market plays a leadership role in determining retail prices, which is in line with industrial organization theory. Our results confirm the fear and concerns of consumers about the existence of price asymmetry. [JEL Classification: Q110; Q113]. © 2016 Wiley Periodicals, Inc.

1. INTRODUCTION

Market liberalization at the domestic level and at the border level has been a dominant feature of market reforms in most developing countries during the last two decades. Like any other developing country, Bangladesh has also undertaken extensive market reforms with respect to its main food grain, rice, by both reducing public intervention for procurement and distribution

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on one hand, and opening-up the private imports since 1992 on the other (Ahmed, 1996). Rice accounts for a high caloric share in the diet of the Bangladeshi population. It is also the most important cereal crop produced and occupies a major share of farmers' agricultural income and employment. A pre-requisite for producers and consumers to benefit from the liberalized market environment is the ability of the market to function efficiently. However, if markets—either spatially or vertically—are constrained by factors such as imperfect market information, lack of credit availability to finance short-run inventories, insufficient transportation, lack of management skills, exercise of market power, hoarding by traders, retailer search behavior, trade promotions, etc., the inferred benefits from reforms can be jeopardized.

The Bangladesh rice market was liberalized in different phases, with the first phase beginning in the mid-1980s following a directive from the World Bank and the IMF under the structural adjustment program (SAP). Under the SAP, domestic public procurement was minimized and transportation restrictions across regions were removed. The second phase of reforms continued in the 1990's when the rationing system, originally designed to support the low-income population in Bangladesh was completely abolished. In the aftermath of these reforms, the Bangladesh government has virtually no role in the procurement and the distribution of rice. All of the reforms significantly changed the structure of the Bangladesh rice market from a publicly controlled market to a free market system dominated by the private traders operating at all levels in rice value chain including, wholesale and retail levels.

Private traders have contributed to the country's overall food security especially after a devastating flood in 1998 (Carl & Dorosh, 2003). The contribution of private sector was mainly to maintain stability in supply when there was a domestic production shortfall due to natural calamities in the country. However, the policy of liberalization has also greatly increased the number of market participants—at the wholesale level—and has created a fragmented marketing system. In this more liberal era, the level of control exercised by private traders in the rice market has led to questions of potential price manipulation. Even though Bangladesh is one of the world's major rice importers, about 85–90% of its total domestic rice consumption comes from domestic production which in essence is procured and then sold to retailers by the private wholesale traders. Under this marketing system a natural question that arises is how fair and efficient is the Bangladesh rice market in delivering product from producers to consumers. This question is of vital importance as Bangladeshi consumers spend a large proportion of their income on rice, while Bangladeshi producers earn the lion share of their income from rice production. Central to answering this question is the market behavior of private traders operating at wholesale and retail levels and their influence on rice prices.

The food grain marketing chains in developing countries tend to be long and complex because of the involvement of many small scale intermediaries. There is a widely held belief that the domestic rice market in Bangladesh can be manipulated by private traders operating at wholesale and retail levels leading to increased and unstable rice prices. Such manipulation would have serious economic consequences for poor Bangladeshi households who are net buyers of rice, which accounts for approximately 40–50% of their total annual expenditures. Given this potential impact, government policy makers should be interested in evidence pointing towards the possible existence and sources of price manipulation. In the case of the Bangladeshi rice market, anecdotal evidence and casual observation supports the idea of price asymmetry. Specifically, it is widely believed that price increases emanating at the wholesale level are quickly passed on in terms of higher prices at the retail level. However, it is also widely believed that wholesale price decreases do not lead to similar price decreases at the retail level. With this in mind, this paper investigates whether the widely held perception of asymmetric price transmission actually exists. So far, no studies have examined potential asymmetries in the Bangladeshi rice market—neither vertical, nor at the spatial level. Studies examining rice price transmission during the period of market liberalization have focused exclusively at the spatial level i.e., testing the law of one price (Dawson & Dey, 2002), and estimating the structural determinants of market integration (Golletti & Farid, 1995), and dynamic relationship between international and domestic markets (Alam, Buysse, McKenzie, Wailes, & Van Huylenbroeck, 2012). So, there is a dearth of information with respect to price transmission through the vertical marketing chain. Examining market efficiency and price transmission across spatial market levels and vertical levels are instrumental in evaluating the economic function and performance of markets. Identifying where price discovery takes place in the marketing chain is also of great importance in guiding policies designed to improve the rice marketing system. The industrial organization literature assumes that retail prices are determined by pricing decisions (e.g., mark-up pricing) at the wholesale level (Tirole, 1988). However, in contrast, much of the development literature suggests that price discovery and transmission occurs downstream at the retail level (Kuiper, Lutz, & Van Tilburg, 2003). But neither of these assumptions has been empirically verified with respect to third-world rice markets.

The paper addresses the following research questions: (1) is there a cointegrating long-run relationship between wholesale and retail rice prices in Bangladesh? Although economic theory would suggest that prices should be related at different levels of the marketing chain it is necessary to formally estimate this relationship taking into account the time series properties of the data (e.g., nonstationarity). This long-run relationship is subsequently used to model pricing asymmetries within a dynamic error-correction model framework (Goodwin & Holt, 1999); (2) does the wholesale market dominate the retail market in terms of price discovery or vice versa?; (3) if a relationship between marketing levels exists, is it linear or nonlinear? In other words, are price increases and decreases symmetric or asymmetric? We are particularly interested in determining if decreases in wholesale prices are passed along to consumers as rapidly or fully as are increases in wholesale prices. Our motivation is to understand if "conventional wisdom" about "rockets and feathers" phenomena in Bangladesh rice market is true. Although popular public opinion in Bangladesh would suggest that the primary source of pricing asymmetries is inventory management, market power, and price manipulation at the wholesale level, the price transmission literature indicates that numerous other potential explanations could equally well apply. Wholgenant (2001) and McCorriston, Morgan, and Rayner, (2001) provide excellent reviews of literature on price transmission. Kinnucan and Zhang (2015) show that if farm price is not expressed on a retail-equivalent basis, competitive market clearing is consistent with an absolute price pass through in excess of unity from farm to retail levels. The extent to which this absolute pass though exceeds unity depends upon the degree of physical transformation that a product goes through between the two marketing levels. Products that undergo little transformation between marketing levels should have lower levels of absolute price pass through in competitive markets (closer to unity). If the absolute price pass through is less than unity, then this would be symptomatic of imperfect between marketing levels. Market power is often assumed to be a potential explanation for asymmetric price behavior (Moorthy, 2005). McLaren (2015) show evidence of pricing asymmetries between international and domestic prices in the presence of strong monopsony power of agricultural intermediaries. Specifically, he finds that as international price falls, local prices fall proportionally more than when international prices increase. However, Weldegebriel (2004) shows that under the assumption of variable input proportions oligopoly and oligopsony powers do not necessarily lead to imperfect price transmission across market levels. Other typical explanations for asymmetric price behavior include: search costs (Gomez, Richards, & Lee, 2013; Richards, Gomez, & Lee, 2014); inventory management costs (Cui, Raju, & Zhang, 2008); information asymmetries (Busse Silva-Russo & Zettelmeyer, 2006; Kumar, Rajiv, & Jeuland, 2005); market structure (Lee & Miguel, 2013); and cost conditions (McCorriston et al., 2001). We do not claim we have an answer as to why this pricing behavior might occur in Bangladeshi rice markets. However, as Goodwin (2006) notes documenting asymmetric price adjustment in itself, is valuable information, if only to demonstrate that positive price shocks elicit a response that differs from negative shocks, irrespective of why such price asymmetries might occur.

The remainder of the paper is organized as follows: Section 2 provides a brief overview of the rice market participants in Bangladesh. The data are presented in Section 3, results and discussions are outlined in Section 4. The last section concludes.

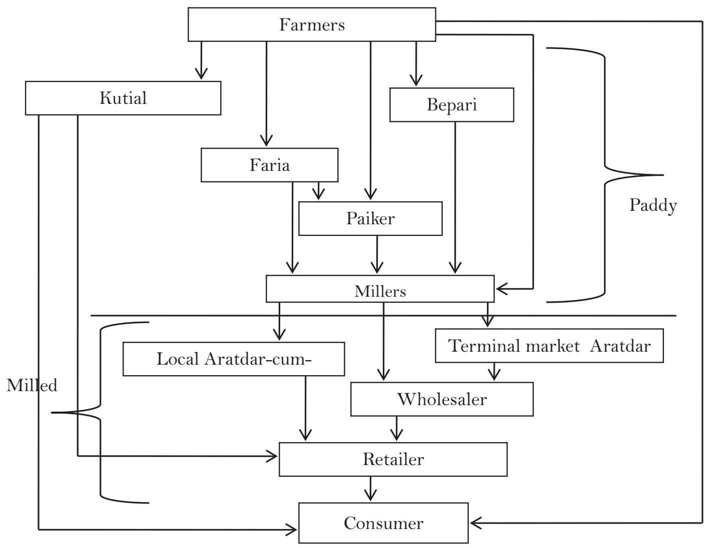

Figure 1 Marketing Channels of Paddy/Rice in Bangladesh.

2. A BRIEF OVERVIEW OF RICE MARKET PARTICIPANTS IN BANGLADESH

There are numerous intermediaries in Bangladeshi rice markets. Marketing channels of paddy/rice in Bangladesh are presented in Figure 1. *Kutials* are generally landless small marginal farmers and laborers who purchase paddy from farmers at farmyards or at the small village markets. After milling (manually) they sell rice to the retailers and consumers at the local markets. *Faria*, who operate in local villages, procure paddy rice from local farmers and subsequently sell it to *Paiker or Millers. Bepari* are long distance traders, who usually perform some marketing functions, such as sorting, grading, packaging, etc. *Bepari* typically supply paddy rice to *Millers*, who are private traders located in bigger markets some distance away from production areas. Millers dry, unhusk and polish the paddy before selling it to *wholesaler*, the *Aratdar-cum-wholesaler* and wholesaler at terminal market (Dhaka). The *Aratdar-cum-wholesaler* typically have large capital reserves with which to run their businesses; they stay mainly at the divisional level and perform different marketing functions; buying milled rice from *Millers*, sorting, grading, and packing, etc. There are some terminal markets where the *Aratdars* perform similar activities but are located in the capital city, Dhaka. Bangladesh imports 10–15% of its rice supplies from India, Myanmar, and Thailand. This imported rice is bought by *Aratdars* located mainly in the capital city of Dhaka. The *Aratdars* sell milled rice to *wholesaler* and then *wholesaler* sell milled rice to retailer, who in turn sell rice to consumers. There are many participants in the paddy/rice marketing channels, and there is no single channel transferring paddy and milled rice to consumers. Retailers can buy milled rice from millers and/or Aratdar-cum-wholesalers. The nature of the marketing system gives the private traders operating at wholesale and retail levels much power with which to influence prices. Given the key role played by these private traders and their powerful entrenched influence on the marketing system, it is of great interest to determine their potential impact on the price transmission and price discovery process.

3. DATA

We use monthly average wholesale and retail prices of coarse rice. The wholesale price data are taken from Global Information on Early Warning System, FAO. The monthly retail prices are

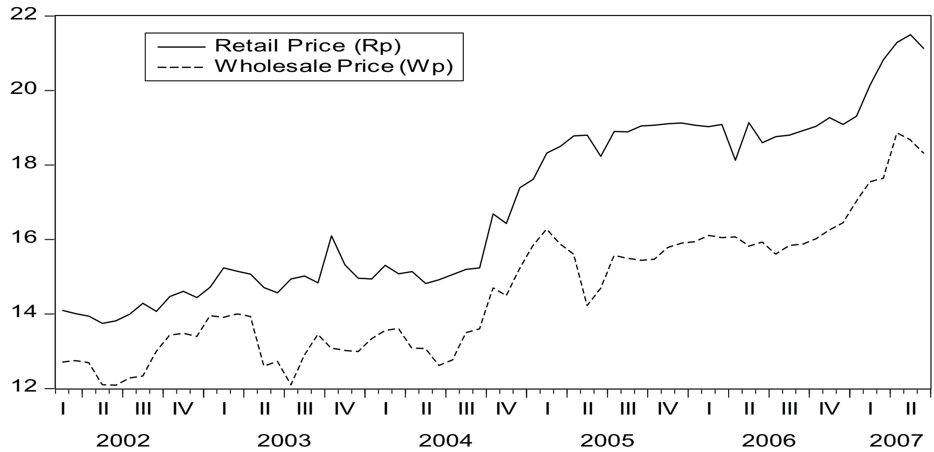

Figure 2 Evolution of Wholesale and Retail Prices in Taka Per Kilogram of Rice in Bangladesh.

taken from different published issues of statistical yearbooks and the economic trends published by Bangladesh Bureau of Statistics and Bangladesh Bank. These data sources are well regarded since there are many screenings done prior to publishing data from all over the world. The data period covers from February 2002 to June 2007. This data period yields the longest continuous available time series data of wholesale and retail price of rice in the published data bases. The prices are measured in Taka[1] per kilogram. The wholesale and retail prices are presented in Figure 2. It can be observed from the figure that the two prices are highly correlated and increase over time.

4. ECONOMETRIC MODELS AND RESULTS

We first used augmented Dickey–Fuller (1979) and Phillips–Perron (1988) tests to determine if the price series contain unit roots. The tests indicated that wholesale and retail prices contain one unit root but are stationary in first differences (Table A1). Given that prices are $I(1)$, we proceed to estimate a long-run equilibrium relationship using standard cointegration procedures.

4.1 Johansen Cointegration Rank Tests

Using Johansen (1995) and Johansen and Juselius (1990) procedure, tests for cointegration were performed. Johansen's procedure is a multivariate approach based on maximum likelihood estimates of the cointegrating regression. A vector autoregressive (VAR) model is specified, in which a vector of prices ($P \times 1$), observed at time t, are determined by a vector of past prices. According to Engle and Granger (1987) representation theorem, the vector P_t has a vector autoregressive error correction representation in the following specification:

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

where $\Pi = \sum_{i=1}^{p} A_i - I$ and $\Gamma_i = -\sum_{j=i+1}^{p} A_j$, P_t is a ($P \times 1$) dimension vector (P_1, P_2) corresponding to the number of price series (P_1 = retail price and P_2 = wholesale price) in which both the prices are $\sim$ I(1), the Π, Γ_i, and Φ are parameter matrices with ($P \times P$) dimension to be

[1]Local currency (79 Taka = 1$).

TABLE 1. Johansen Cointegration Rank Test Results

	Model 2		Model 3		Model 4	
Cointegration Rank (r)	Test statistics (λ_{trace}), (λ_{max})	Critical values ($\lambda_{0.95}$)	Test statistics (λ_{trace}), (λ_{max})	Critical values ($\lambda_{0.95}$)	Test statistics (λ_{trace}), (λ_{max})	Critical values ($\lambda_{0.95}$)
Trace statistics (λ_{trace})						
$H_0: r = 0$ vs $H_1: r \geq 1$	22.665**	20.262	17.649**	15.495	25.759	25.872
$H_0: r \leq 1$ vs $H_1: r \geq 2$	4.143	9.165	0.035	3.841	6.167	12.517
Maximum eigenvalue (λ_{max})						
$H_0: r = 0$ vs $H_1: r \geq 1$	18.522**	15.892	17.615**	14.265	19.592**	19.387
$H_0: r \leq 1$ vs $H_1: r = 2$	4.143	9.165	0.035	3.841	6.167	12.518

Note. **Indicates the hypothesis is rejected at 5% significant level.

estimated, D_t is a vector with deterministic elements (constant, trend) and ω is a $(P \times 1)$ vector of white noise residuals. Cointegration is determined by the rank of matrix (Π) which contains long-run information (also the loading factors) about the variables. We can three different cases from Equation (1). First case, if rank $(\Pi) = P$, then Π is invertible and all the variables in levels are stationary meaning that no cointegration exists. Second, if rank $(\Pi) = 0$, that is, Π is a null matrix means that all the elements in the adjustment matrix has value zero, therefore, none of the linear combinations are stationary, and can be estimated unrestricted VAR to identify the short-run dynamics only. Third, according to the Granger representation theorem, when $0 <$ rank $(\Pi = r) < P$, there are r cointegrating vectors or r stationary linear combinations of the variables. For example if rank (r) of matrix Π is equal to one, there is single cointegrating vector or one linear combination of variables which is stationary. In this case the coefficient matrix Π can be decomposed into $\Pi = \alpha\beta$, where α is the speed of adjustment vector (factor loading vector), which measures the speed at which the variables adjust towards the long-run equilibrium following a price shock. The term β is the long-run equilibrium (cointegrating) vector. In this case, P_1 and P_2 are but the linear combination βP_{t-1} is $I(0)$. So, the Johansen method is to estimate the Π matrix from an unrestricted VAR and to test whether we can reject the restriction implied by reduced rank of Π. Standard trace and maximum eigenvalue tests are used to determine rank of Π.

The lag length used to specify Equation (1) was determined by Schwarz information criterion (SIC). Since inclusion of the deterministic terms (constant, trend) in the cointegration space is sensitive to the identification of cointegration rank (Harris & Sollis, 2003), we estimated three potential models, which are denoted by M2, M3, and M4. M2 restricts all the deterministic components to a constant in the cointegration space; M3 allows linear trend in the level of the variables; and M4 allows linear trend in the cointegration space. Results presented in Table 1 indicate that a single long-run cointegrating relationship exists between wholesale and retail prices. In addition, each of our VAR models satisfies the stability condition.[2] It is also important to mention that the above analysis assumes that the price adjustment process is symmetric. However, before proceeding to test for asymmetry, we test the direction of price causality within the Johansen VECM framework.

4.2 Testing Causality in the Johansen Vector Error Correction Model (VECM)

Jayanta, Sajjad, and Baulch, (1997) is the only previous study to measure the relationship and causality between wholesale and retail prices of coarse rice across the two major Bangladesh cities (Dhaka and Chittagong). Jayanta et al. (1997) used conventional Granger causality F-tests in a simple regression framework. However, if prices are cointegrated, the conventional Granger causality test results, which ignore the long-run equilibrium relationship, are

[2]Because of brevity we do not provide here all specification test results.

mis-specified as they omit an error-correction term variable. Thus, we extend the extant literature by employing causality (Wald) tests within the Johansen VECM framework (Mosconi & Giannini, 1992; Dolado & Lutkepohl, 1996). We perform three different causality tests. To identify the causal relationship (short-run, long-run, and strong causality), Equation (1) can be rewritten as Equations (2) and (3)

$$\Delta R_{t=}\mu_1 + \sum_{i=1}^{K} \beta_{i(w)} \Delta W_{t-i} + \sum_{j=1}^{L} \beta_{j(r)} \Delta R_{t-j} + \alpha_1 Z_{t-1} + \varepsilon_{t,1} \tag{2}$$

$$\Delta W_{t=}\mu_2 + \sum_{i=1}^{K} \beta_{i(w)} \Delta W_{t-i} + \sum_{j=1}^{L} \beta_{j(r)} \Delta R_{t-j} + \alpha_2 Z_{t-1} + \varepsilon_{t,2} \tag{3}$$

where Z_{t-1} is the lag of error correction term (ECT), R is retail price, and W is wholesale price. In Equations (2) and (3), there are three possible cases for testing *first*, long-run Granger causality given that the variables are cointegrated:

(a) $\alpha_1 \neq 0, \alpha_2 \neq 0$, which implies bidirectional causality, mean that there exists a feedback long-run relationship between the variables and no individual price play a leadership role.
(b) $\alpha_1 = 0$ but $\alpha_2 \neq 0$ implies a unidirectional causality and the retail price Granger causes the wholesale price, the retail price is weakly exogenous.
(c) $\alpha_2 = 0$ but $\alpha_1 \neq 0$ implies unidirectional causality and the wholesale price Granger causes the retail price, the wholesale price is weakly exogenous.

Second, to test short-run Granger causality, that is, ΔW does not Granger cause ΔR in the short (or, ΔR does not Granger cause ΔW), we can examine the statistical significance of lagged dynamic terms by testing the null: all $\beta_{i(w)} = 0$ (or, all $\beta_{j(r)} = 0$) using a Wald test. Nonrejection of the null implies that ΔW does not Granger cause ΔR (or ΔR does not Granger cause ΔW).

Finally, we can test for strong exogeneity (overall causality) by imposing joint restrictions on the lagged dynamic terms and the ECT. In this case, ΔW does not Granger cause ΔR if the null, all $\beta_{i(w)} = 0$ and $\alpha_1 = 0$ (or ΔR does not Granger cause ΔW, if the null, all $\beta_{j(r)} = 0$, and $\alpha_2 = 0$).

Results with respect to long-run causality are presented in Table 2. We clearly reject case (a) and case (b), but fail to reject case (c), implying that wholesale price is weakly exogenous; or in other words wholesale price Granger causes retail price. Test results for short-run causality and strong exogeneity are consistent with our long-run causality result, wholesale price Granger causes retail price but not vice versa. These results are in line with the concept of mark-up pricing. We conclude that wholesale market plays a leadership role in price discovery and influences prices at retail level.

4.3 The Asymmetric Error Correction Model

Seminal work of Houck (1977) first developed a test for detecting potential asymmetry in the price transmission process. The test splits price movements across variables into increases and decreases. Other studies, notably Boyd and Brorsen (1988); Kinnucan and Forker (1987); Baily and Brorsen (1989); Zhang, Fletcher, and Carley (1995); Mohanty, Peterson, and Kruse (1995); Willett et al. (1997); Peltzman (2000); Bart and Stevan (2000); Aguiar and Santana (2002), followed this approach. All of these studies focused on asymmetric price transmission in developing countries' agricultural and food markets. However, none of these studies addressed the inherent time series properties of data, that is, nonstationarity and long-run cointegrating equilibrium relationships. Von Cramon-Taubadel and Loy (1996) and Von Cramon-Taubadel

TABLE 2. Restrictions on the Johansen VECM for Testing Causality between Retail and Wholesale Prices

Sources of Causations	Causality Test (Likelihood Ratio Test) Dependent Variable: ΔRP	Dependent Variable: ΔWP	Causality Decision
Long-run causality	H_0: $\alpha_1 = 0$ vs H_1: $\alpha_1 \neq 0$	H_0: $\alpha_2 = 0$ vs H_1: $\alpha_2 \neq 0$	
	9.847** [0.001]	0.654 [0.419]	Unidirectional (wholesale → retail)
Short-run causality	H_0: $\sum^{\beta_{i(w)}} = 0$ vs H_1: $\sum^{\beta_{i(w)}} \neq 0$	H_0: $\sum^{\beta_{j(r)}} = 0$ vs H_1: $\sum^{\beta_{j(r)}} \neq 0$	
	4.013** [0.045]	0.395 [0.529]	Unidirectional (wholesale → retail)
Strong exogeneity	H_0: $\sum^{\beta_{i(w)}} = 0,\ \alpha_1 = 0$ vs H_1: $\sum^{\beta_{i(w)}} \neq 0,\ \alpha_1 \neq 0$	H_0: $\sum^{\beta_{j(r)}} = 0,\ \alpha_2 = 0$ vs H_1: $\sum^{\beta_{j(r)}} \neq 0,\ \alpha_2 \neq 0$	
	13.973** [0.000]	3.037 [0.219]	Unidirectional (wholesale → retail)

Note. **Indicates the significant level at 5%; probability levels are in the parentheses.

(1998) were the first papers to take account of unit root properties of price series by developing the asymmetric error correction models (ECM-EG) approach. Two recent articles by Meyer and Von Cramon-Taubadel (2004) and Giliola (2007) provided a comprehensive discussion on the possible causes and types of asymmetry. Although most of the price asymmetry literature has focused attention on agricultural markets in developed countries, Abdulai (2000), Michele and Kirsten (2006), and Van Campenhout (2007) have identified the existence of asymmetric price behavior in developing markets.

Consistent with this more recent literature, we use the asymmetric ECM-EG modeling approach to test for price asymmetry between Bangladesh wholesale and retail rice markets. Following on our cointegration and causality results we focus on testing for asymmetry with respect to short-run price dynamics. Following Von Cramon-Taubadel (1998), Von Cramon-Taubadel and Loy (1999), Lajos and Imre (2005), Oral and Pablo (2007), and Ioanna and Yannis (2008), we account for potential price asymmetry using the Engle–Granger's (1987) two-step approach. Based upon our causality results and similar to Gomez et al. (2013) and Kuiper et al. (2003) we assume the wholesale price may be considered an exogenous variable and hence Equation (4) can be estimated to explain the long-run equilibrium relationship between wholesale and retail prices:

$$R_t = \beta_0 + \beta_1 W_t + \psi T + \varepsilon_t \tag{4}$$

where R is retail price, W is wholesale price, T is a time trend, and ε is Gaussian white noise error term. The short-run dynamic price adjustments modified by an ECT are specified in terms of a prototypical error correction model (ECM):

$$\Delta R_t = \mu_1 + \sum_{i=1}^{k} \beta_{i(r)}\ \Delta R_{t-i} + \sum_{j=0}^{L} \beta_{j(w)} \Delta W_{t-j} + \alpha\hat{\varepsilon}_{t-1} + \epsilon_t \tag{5}$$

From Equation (5), β_i and β_j measures the short-run impact of past movements in retail and wholesale prices on current retail price changes; α measures the speed of the adjustment to perturbations in long-run equilibrium; ε_{t-1}, the ECM term, measures the size of last periods departure (price perturbation) from long-run equilibrium ($\hat{\varepsilon}_{t-1} = R_{t-1} - \beta_0 - \beta_1 W_{t-1} - \psi T$).

To take account of potential price asymmetries, Equation (5) is respecified as Equation 6 where wholesale price increases and price decreases are measured separately. Equation 6 also reparameterizes the ECT into positive and negative values based on a Heaviside indicator function show in Equation (7). Thus, our ECM-EG asymmetric model has the following form:

$$\Delta R_t = \mu_1 + \sum_{i=1}^{K} \beta_{i(r)} \Delta R_{t-i} + \sum_{j=0}^{L} \left(\beta^{+}_{j(w)} \Delta W^{+}_{t-j}\right) + \alpha_1^{+} \hat{\varepsilon}^{+}_{t-1} + \sum_{j=0}^{L} \left(\beta^{-}_{j(w)} \Delta W^{-}_{t-j}\right) + \alpha_1^{-} \hat{\varepsilon}^{-}_{t-1} + \omega_t \tag{6}$$

where, superscript "+" on the coefficients and the variables is relevant when price increases and superscript "–" is relevant when price decreases. The ECT term ($\hat{\varepsilon}_{t-1}$) $Z_{1,t-1}$) is decomposed into positive and a negative values, $\mathrm{ECT}_{t-1} = \mathrm{ECT}^{+}_{t-1} + \mathrm{ECT}^{-}_{t-1}$. They are defined as $\hat{\varepsilon}^{+}_{t-1} = I_t(R_{t-1} - \beta_0 - \beta_1 W_{t-1} - \psi T)$ and $\hat{\varepsilon}^{-}_{t-1} = (1 - I_t)(R_{t-1} - \beta_0 - \beta_1 W_{t-1} - \psi T)$. I_t is a Heaviside indicator function where,

$$I_t = \begin{cases} 1 & \text{if } \hat{\varepsilon}_{t-1} \geq 0 \\ 0 & \text{if } \hat{\varepsilon}_{t-1} < 0 \end{cases} \tag{7}$$

Our ECM-EG asymmetric model specification allows us to test for short-run and the long-run price asymmetry. Specifically, a Wald χ^2 test is used with the null hypothesis (H_0: $\alpha_1^{+} = \alpha_1^{-}$). This in effect determines if the absolute size of the speed of adjustment parameters differs with respect to price increases and decreases. Rejection of the null would provide evidence of price asymmetries. In a similar vein potential asymmetries in the short-run price dynamics were tested using Wald χ^2 test with the null hypothesis (H_0: $\beta^{+}_{j,(w)} = \beta^{-}_{j,(w)}$). In this case rejection of the null infers retail price responds differently to past wholesale price increases compared to past wholesale price decreases. Lag length of short-run price dynamics are determined using the Akaike information criterion (AIC). Two lags are found to be optimal. In addition, we test asymmetric retail price behavior based upon cumulative changes in lagged wholesale prices. A joint F test is used with the null hypothesis (H_0: $\sum_{j=0}^{L} \beta^{+}_{j(w)} = \sum_{j=0}^{L} \beta^{-}_{j(w)}$)

Table 3 presents the results of the asymmetric error correction model. In the model, we find that the coefficients of ECTs are significantly different from (H_0: $\alpha_1^{+} = \alpha_1^{-} = 0$) zero which is a necessary condition for cointegration and the existence of a long-run equilibrium. The sign and the magnitude of the estimated ECTs are consistent with our *a priori* expectations. The ECT^{-}_{t-1} (α_1^{-}) term induces a significantly greater change in the retail price than does ECT^{+}_{t-1} (α_1^{+}) term . The coefficient of ECT^{-}_{t-1} is significantly greater in absolute terms than ECT^{+}_{t-1} (χ^2 (1) = 3.220 × (0.07)). This indicates that retail price reacts faster to disequilibria induced by wholesale price shock increases compared with disequilibria brought about by wholesale price shock decreases. Thus, when the retail profit margin is squeezed (wholesale prices are relatively high in comparison to retail prices) it appears that retail prices adjust quickly to restore margins. In contrast retail prices show little or no response when retail profit margins are relatively large.

Turning to our short-run price dynamic results, our findings reject the null hypothesis (H_0: $\beta^{+}_{w} = \beta^{-}_{w}$). This implies that retail prices adjust differently in the short-run to positive contemporaneous wholesale prices changes compared to negative contemporaneous wholesale

TABLE 3. Results of the Asymmetric Error Correction Model

Regressors	Coefficients	t-Statistics
ΔR_{t-1}	−0.370**	2.776
ΔR_{t-2}	−0.103	−0.903
ΔW_t^+	0.498**	3.266
ΔW_t^-	−0.102	−0.623
ΔW_{t-1}^+	0.389**	2.298
ΔW_{t-1}^-	0.167	0.979
ΔW_{t-2}^+	0.493**	2.997
ΔW_{t-2}^-	−0.086	−0.484
α_1^+	−0.010	−0.072
α_1^-	−0.532**	−2.602
Ramsey RESET		3.416 [$\sim F(1, 50)$]
Lagrange multiplier autocorrelation test		0.178 [$\sim F(2, 49)$]
Normality test (Jarque–Bera)		1.40[c] [0.496]
ARCH test		0.175 [$\sim F(1, 59)$]

Cointegration equation in the first stage of ECM-EG

$$R_t = \underset{(4.455)}{4.759} + \underset{(7.859)}{0.704} W_t + \underset{(6.812)}{0.056} t + \varepsilon_t$$

Test for cointegration (from asymmetric model)	H_0: $\alpha_1^+ = \alpha_1^- = 0$ (No cointegration)	$\chi^2(2) = 8.155^{**}$ (0.017)
Wald test for symmetry		
Hypothesis 1	H_0: $\alpha_1^+ = \alpha_1^-$ vs H_1: $\alpha_1^+ \neq \alpha_1^-$	$\chi^2(1) = 3.220^{*}$ (0.07)
Hypothesis 2	H_0: $\beta_w^+ = \beta_w^-$ vs H_1: $\beta_w^+ \neq \beta_w^-$	$\chi^2(1) = 5.169^{**}$ (0.02)
Hypothesis 3	H_0: $\beta_{w(t-1)}^+ = \beta_{w(t-1)}^-$ vs H_1: $\beta_{w(t-1)}^+ \neq \beta_{w(t-1)}^-$	$\chi^2(1) = 0.703$ (0.401)
Hypothesis 4	H_0: $\beta_{w(t-2)}^+ = \beta_{w(t-2)}^-$ vs H_1: $\beta_{w(t-2)}^+ \neq \beta_{w(t-2)}^-$	$\chi^2(1) = 4.963^{**}$ (0.02)
Hypothesis 5	H_0: $\sum_{j=0}^{L} \beta_{j(w)}^+ = \sum_{j=0}^{L} \beta_{j(w)}^-$ vs H_1: $\sum_{j=0}^{L} \beta_{j(w)}{}^+ \neq \sum_{j=0}^{L} \beta_{j(w)}^-$	F-stat: 9.625*** (0.003)

Note. ***, **, and * indicate that the hypotheses are rejected at 1%, 5%, and 10% level of significance, respectively, the probability level are in the parentheses, c mean residuals are normal, ARCH means autoregressive conditional heteroscedasticity, RESET means regression equation specification error test.

price changes. Our results show that the null of the symmetric short-run adjustment is not rejected with lag 1. Finally, our joint F-test rejects the null hypothesis that all positive price lags (lags one and two) coefficients are equal in absolute terms to all negative price lags (lags one and two) coefficients. Again, the result implies that retail prices adjust differently to past positive and negative wholesale price changes. We observe statistically significant price responses at retail level to positive price shocks emanating at wholesale level over the previous three months. This cumulative price response calculated as an elasticity of price transmission is 1.22. In other words, a 1% price increase at the wholesale level eventually leads to a 1.22% increase in retail prices. Given that rice is consumed in its raw commodity form and requires little processing or product transformation between wholesale and retail levels (rice has already been milled prior to trade at wholesale level) this level of pass through appears excessive and is consistent with market power at retail level. Similarly, insignificant negative price shocks from wholesale to retail levels, presented in Table 3, are consistent with market power at retail level. Our results provide strong supporting evidence for the widely held belief—by consumers and government—that retail prices adjust asymmetrically to long-run disequilibria induced by wholesale price shocks. The diagnostic test results reported in Table 3 shows that there is no problem of autocorrelation, no heteroscedasticity and residuals do not have any problems of nonnormality.

5. CONCLUSIONS

Our paper uses time series estimation methods to test asymmetric price transmission in vertical chain of Bangladesh rice markets. The objective of this paper is to estimate the long-run relationship between two different levels of rice market prices (wholesale and retail), and to investigate the commonly held belief among consumers and government that retail prices adjust asymmetrically to changes in wholesale prices, that is "rockets and feathers" phenomena. The vertical long-run price relationship was analyzed using the Johansen maximum likelihood approach. Results show that the wholesale and retail prices are integrated in the long run. The exogeneity tests established a unidirectional causality from wholesale price to retail price, which is consistent with the mark-up pricing model. Prices are discovered at the wholesale levels and influence the retail price adjustments. Our short- and long-run asymmetry tests strongly support the notion of asymmetric price transmission.

Our results suggest that retailers respond more quickly when their margins are squeezed than when they are expanded. In this paper we do not formally test for the underlying cause of price asymmetry in Bangladesh rice markets. Irrespective of the underlying cause of the price asymmetry, our empirical results provide strong evidence of systemic pricing inefficiencies within Bangladeshi rice markets. A potentially fruitful topic for future research would be to try and uncover the main causes driving these price asymmetries and inefficiencies. Although testing the reasons of asymmetry in Bangladesh rice markets is beyond scope of this study, our research represents an important first step in this direction, and provides policy makers with additional insights about the nature of pricing irregularities in Bangladesh rice markets.

APPENDIX

The augmented Dickey–Fuller (ADF) tests use the following equation:

$$\Delta Y_{t=} c + \rho Y_{t-1} + \beta t + \sum_{j=1}^{k} d_j \Delta Y_{t-j} + \varepsilon_t \tag{A1}$$

TABLE A1. Unit Root Test Results

	Level Data		First Differences	Order of Integration
Prices Series/Tests	$\tau-\text{stat}_c$	$\tau-\text{stat}_{c,\,t}$	$(\tau-\text{stat}_w)$	I(d)
Wholesale price (W)				
ADF	–0.253	–2.457	–7.314***	I(1)
PP	–0.341	–2.577	–7.314***	I(1)
Retail price (R)				
ADF	–0.084	–2.447	–9.232***	I(1)
PP	0.114	–2.468	–9.144***	I(1)

Critical Values	τ_c		$\tau_{c,t}$		τ_w	
Level of Significance	5 %	10 %	5 %	10 %	5 %	10 %
ADF	–2.907	–2.591	–3.482	–3.169	–1.946	–1.613
PP	–2.907	–2.591	–3.482	–3.169	–1.946	–1.613

Note. Lag length for ADF test is decided based on Schwarz info criteria; maximum bandwidth for PP test is decided based on Newey-West (1994); ***indicates that unit root in the first differences are rejected at 1% level; τ-stat$_c$, τ-stat$_{c,\,t}$, and τ-stat$_w$ indicate τ-statistics of random walk with drift, random walk with slope, and pure random walk models, respectively; critical values are from MacKinnon (1996).

where Y_t is the respective price series and Δ is first difference ($Y_t - Y_{t-1}$) operator and ε_t denotes white noise error term. Equation A1 tests the null of a unit root ($\rho = 0$) against a stationary alternative ($\rho < 0$).

We also test the presence of a unit root using Phillips–Perron (PP) in the following specification:

$$X_t = c + \beta \left\{ t - \frac{T}{2} \right\} + \rho X_{t-1} + v_t \tag{A2}$$

where X_t is respective time series, $\{t - \frac{T}{2})$ is the time trend, and T is the sample size, v_t is the usual white noise error term. This procedure, in fact, uses a nonparametric adjustment to the DF test statistics and allows for dependence and heterogeneity in the error tem. Equation A2 also tests the null of a unit root ($\rho < 0$) against a stationary alternative ($\rho < 0$).

The results of ADF and PP unit root test on each of the variables are reported in the Table A1. The results indicate that all price series are nonstationary at their level but stationary at their first differences. Note that the ADF and PP tests were done both only with drift and slope models. In time series econometrics, it is said that prices are integrated of order one denoted by presenting $P_t \sim I(1)$ and prices of integrated of order zero denoted by $\Delta P_t \sim I(0)$. Here, the order of the integration is one.

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