Asymmetric Threshold Vertical Price Transmission in Wheat and Flour Markets in Dhaka (Bangladesh): Seemingly Unrelated Regression Analysis

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Abstract

The analysis of price transmission for commodities requiring processing in vertical markets is challenged by fuzzy policy environments in the case of developing countries. However the analyses of threshold and asymmetries in price transmission at different levels of vertical markets provide a good indicator of market efficiency. The paper employs threshold cointegration that takes into account the asymmetric adjustment towards a long-run equilibrium and short-run price transmission. The paper investigates the non-linear price adjustment in short- and long-run in vertical markets of wheat and flour in Bangladesh. Using monthly wholesale and retail prices of wheat and flour for data from FAOStat for the period January 2008 to February 2016 we develop an asymmetry threshold error correction model for three vertical chains namely (i) wholesale and retail markets of flour, (ii) wholesale markets of wheat and flour, (iii) wholesale markets of wheat and retail markets of flour. We find evidence of threshold effects in vertical wheat-flour markets. The speed of adjustment towards the long-run equilibrium is different when the price deviations exceed the threshold value from when price deviations are below the threshold. Additionally, we find evidence of short-run price asymmetries implying that downstream price responds faster when upstream price increases than when the latter falls. This validates the hypothesis of `rocket and feature` principle in the wheat-to-flour markets in Bangladesh. Proximate reasons for these differences are discussed.

Keywords: Price transmission, vertical markets, wheat and flour, cointegration, asymmetry threshold error correction model.

JEL Classification Code: C22, C51, Q13, Q18

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1. Introduction

In Bangladesh, wheat is the second staple next to rice. Although domestic wheat production has been decreasing due to host of factors including economic and climatic factors but due to changes in life styles, fast food consumption habits, considerable expansion of confectionery and bakeries and innovation in processing, future consumption of wheat flour (locally called *atta)* is likely to rise sharply. Flour consumption by the middle and the high income consumers is higher than for low income consumers due to its high price compared to its substitute (rice) in the country including in the capital city, Dhaka, which has a population of more than 18 million.

In marketing and industrial organization literature (Kuiper et al. 2003), a transmission channel behaviour is assumed but not empirically tested. In this literature wholesalers are the vertical price leaders as per the Stackelberg leadership model wherein retailers and wholesalers maximize profits. For example, the retailer maximizes the profit conditional on the wholesale price that has to be paid to the wholesalers, and the wholesalers then determines the wholesale price by maximizing profit while taking the conditional profit maximizing behaviour of retailers into account. Our motivation was to test the price leadership role as per Stackelberg model as the starting point. If two price series are cointegrated in the long-run but diverge from each other in the short-run due to random shocks, an adjustment could restore equilibrium. Most studies assume that the speed of this adjustment behaviour is symmetric. But this may not be true if there is a non-linear adjustment due to threshold effects – in particular, in developing countries where large transaction costs, policy interventions etc. are common. Testing for non-linear behaviour of price adjustment is a primary focus of this paper. Then, given the price leadership role in the vertical chain actors, we test for asymmetries in the short- and long-run based on the identified price leader, e.g., in the wholesale and retail markets of flour. Retail prices respond faster when wholesale price increases than when it falls. This is the so-called `rocket and feather` principle in the literature on price transmission the vertical markets.

Against this backdrop the objective of this paper is to estimate the threshold and test the short- and long-run asymmetries in vertical supply chains of wheat-to-flour markets in Dhaka city, Bangladesh. The vertical markets considered in this paper are (i) wholesale and retail markets for flour, (ii) wholesale markets for wheat and flour, and finally (iii) wholesale markets for wheat and retail markets of flour. The vertical markets are selected based on data availability. The paper is structured as follows. Section 2 presents a brief overview of wheat markets and wheat-to-flour vertical chains. A brief review of literature is presented in section 3. The modelling of threshold and asymmetry in wheat-to-flour markets are presented in section 4. Section 5 presents data followed by results and discussion in section 6. Section 7 concludes.

2. Overview of wheat market and wheat-to-flour supply chains in Bangladesh

Compared to its substitute (rice) the domestic wheat and flour markets in Bangladesh are thin. This is due to declining trend of wheat production in the country attributable mainly to economic and climatic factors. The production of wheat is less profitable than its competitive crop - Boro rice. Climatic factors such as temperature and rainfall are also responsible to this declining trend of wheat production. Since relative price of wheat/flour has increased over time, consumption of wheat flour tends to be low in the rural areas than in urban areas including Dhaka. The food consumption pattern has been changing, urban consumers prefer more processed and fast food than home based traditional food consumption thus increasing the demand for flour and flour made confectionery to rise.

The USDA estimate shows that wheat production in Bangladesh in 2002-03 season was 1.5 million tons whereas it decreased to 1.1 million tons in 2012-13. Despite this drop in production consumption is high and rising steadily. Total wheat consumption in 2002-03 was 3 million tons but increased to 4 million tons in 2012-13. These figures clearly indicate that although wheat production has been decreasing, wheat demand continues to increase. The average per capita wheat consumption has increased 115 percent during last decade (BBS 2012). BIDS (2012) survey shows that average daily per capita wheat consumption was 46 grams (83 grams for urban consumers and 25 grams for rural consumers). This indicates that the changing food consumption pattern from traditional rice based to the processed and fast food based food consumption pushing flour demand to rise. There have been large increases in the number of wheat processing mills and milling capacity with modern roller mills replacing traditional mills. The processing capacity of large scale modern mill is about 500 million tons and that of small and medium-sized mills is about 100 million tons. So, this capacity increases of wheat processing mill can be seen as a direct result from the expansion of fast food restaurants and growth of export oriented food processing sector in the country. Bangladesh is a net wheat importing and has doubled its wheat import in last decade. USDA estimates that Bangladesh imported 1.3 million tons of wheat in 2002-03 which increased to 3.9 million tons in 2011-12. Major exporting countries exporting wheat to Bangladesh are Canada, Australia, Ukraine and United States. The commercial import volume of wheat constitutes 75 percent of total and remaining 25 percent comes as food aid and/or public import under the Public Food Distribution System.

In Bangladesh, vertical supply chains of wheat-to-flour consist of numerous actors including farmers, private importers, public importers, wheat wholesalers, millers, flour wholesalers, flour retailers, consumers. Baulch et al (1998), Farid and Rahman (2002) mapped the marketing channels in where major actors are producers, Beparies (traders), millers, wholesalers, importers, wholesaler of milled wheat, manufacturers, retailers and consumers (Figure 1). Hossain et al. (2004) mapped wheat-to-flour marketing channels in Bangladesh. These are (i) Farmers →Beparies → Aratdar-cum-wholesalers of

Wheat →Millers →Flour wholesalers →Flour retailers → Consumers, (ii) Farmers →Beparies →Millers →Flour wholesalers →Flour retailers →Consumers, (iii) Farias →Beparies →Millers →Retailers →Consumers, (iv) Famers →Millers →Retailers → Consumers. Based on the numerous actors participating in Bangladesh wheat-flour vertical markets, we select three distinct chains as follows: (i) Flour wholesalers →Flour retailers; (ii) Wheat wholesalers →Flour wholesalers, and (iii) Wheat wholesalers →Flour retailers

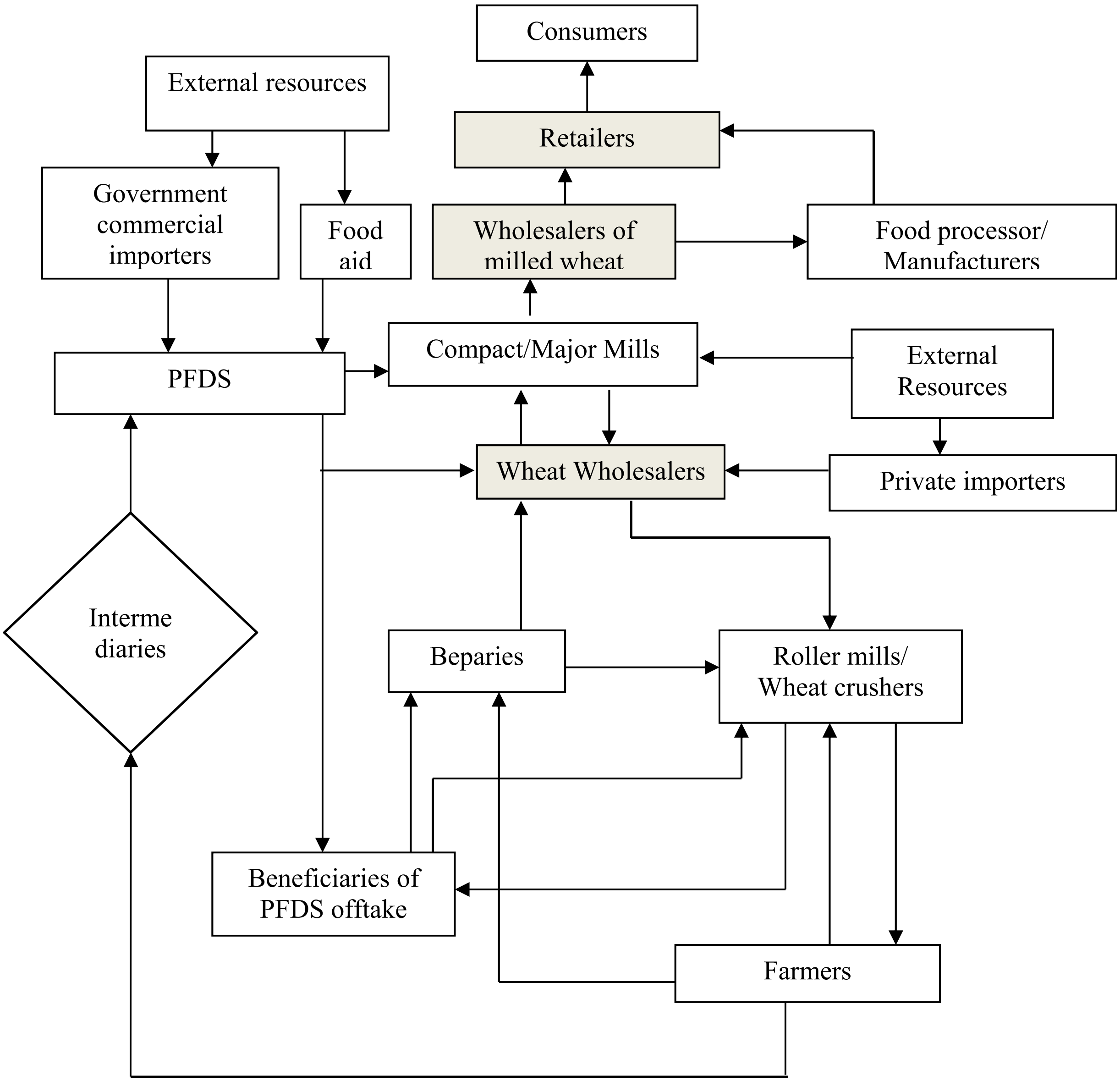

Figure 1: Wheat marketing channels in Bangladesh (Source: Farid and Rahman 2002, Baulch et al. 1998)

3. Review of Literature on the Asymmetry in Vertical Supply Chain

Asymmetry in vertical supply chains was first discussed by Wolffram (1971) and later improved by Houck (1977) who focuses on differences in responses of aggregate supply functions to rises and falls in prices. Many assessments of price transmission asymmetry (PTA) in the food system (Kinnucan and Forker 1987; Boyd and Brorsen 1988; Hansmire and Willett 1992; Zhang et al. 1995) adopted Wolffram (1971) and Houck (1977) to study of price transmission. Von Cramon-Taubadel (1998) argue that these studies may be biased because they ignore the time series properties of the data. Specifically, prices at different levels of the supply chain are often cointegrated which may lead to spurious results.

A number of empirical studies (Mankiw and Romer 1991; Peltzman 2000) identified the presence of PTAs in aggregate price adjustments and led economists to develop theories to explain this phenomenon. The PTAs are viewed as the result of price setting frictions, for example costs associated with price adjustments as well as the staggered timing of price changes and inventory management (Levy et al. 1997). At the aggregate level, PTAs are regarded as the consequence of imperfect competition, including demand externalities and coordination failures (Borenstein et al. 1997; Neumark and Sharpe 1992). These principles have been widely employed to develop testable models of PTAs in vertical and spatial markets of agricultural commodities and food products (Kinnucan and Forker 1987; Bailey and Brorsen 1989; Azzam 1999; Xia 2009).

Recently, attention has turned to empirical procedures based on Engle and Granger (1987) and extended by Granger and Lee (1989) to test for PTAs by different actors in vertical markets. These authors developed a formal model showing that when two price series are cointegrated, there exists an error correction (EC) representation that describes the short- and long-run relationships as well as the inherent price transmission mechanism. Indeed, the second half of the 1990s saw an increasing interest in EC models to study PTAs (Balke et al. 1998; Frost and Bowden 1999; Peltzman 2000). Von Cramon-Taubadel and Loy (1996) pioneered the application of EC models to examine PTAs in markets for agricultural commodities and challenged then existing methods to discuss price asymmetry in international wheat markets. The advantages of EC models to investigate PTAs when price series are cointegrated are formalized by Von Cramon-Taubadel and Loy (1999). Subsequent studies employ EC models to examine PTAs primarily in markets for meats (Ben-Kaabia et al. 2005; Sanjuan and Gil 200l; Miller and Hayenga 2001; Goodwin and Holt 1999; Von Cramon-Taubadel 1998) and dairy products (Lass 2005; Serra and Goodwin 2003; Romain et al. 2002). These studies provide evidence of short-run price asymmetries along the supply chains for agricultural commodities.

PTAs can occur in the short- and long-runs, depending on the stochastic process governing prices. For instance, two price series are assumed to be cointegrated, the differences between positive and negative changes accumulate over time leading to a non stable long-run equilibrium. In contrast, if two time series are integrated and not cointegrated, only short-run asymmetries are possible (Von Cramon-

Taubadel and Loy 1996). Abdulai (2000) developed threshold cointegration that allow for asymmetric adjustment towards a long-run equilibrium relationship to examine price linkage between principal maize markets in Ghana. The paper confirmed that wholesale maize prices in local markets respond more swiftly to increase than to decrease in central markets. Abdulai (2002) employed a similar model to analyse price transmission between producer and retail prices in the swine-pork supply chain in Switzerland and confirmed the PTA between the producer and retail levels is asymmetric. Awokuse and Xiaohong (2009) examined the effect of nonlinear threshold dynamics on asymmetric price transmission for US dairy products using threshold EC models. They found that price transmission of changes between producer and retail stages of the marketing chain is asymmetric for butter and fluid milk. The authors concluded that previous studies that assumed symmetric behaviour and ignored threshold may be misleading. Lee and Miguel (2013) employed the threshold cointegration model to analyze price transmission between international-to-retail prices of coffee and impact of export quota system (EQS) in the coffee supply chain. The authors compared the periods of EQS and post-EQS and confirmed threshold and asymmetries in short- and long-run price transmission from international-to-retail prices of Germany, USA and France. Ghoshray and Ghosh (2011); Listoti (2009) and Ghoshray (2002) also confirmed threshold and PTAs in short- and long-run in wheat-to-flour markets.

Sango and Mohamane (2010) analysed the threshold market integration of rice between Nepal and India. The authors found that coarse rice prices in Nepal respond to shock originating in India and adjustments to negative price deviations from long-run equilibrium are quicker than adjustment to positive ones. The authors emphasized that restrictive food trade policies in India will undermine household food security in Nepal. Ghoshray and Ghosh (2011) investigated the relationship between various wheat prices quoted in different market centers in four Indian states. They used MTAR model to identify the presence of asymmetric adjustment of wheat prices. The authors indicated that the asymmetric price transmission may be a cause from the poor dissemination of knowledge on market conditions and the transaction costs.

For the case of Bangladesh contributions have concentrated on the commodity market, mainly rice market integration at spatial (Alam et al. 2012b; Dawson and Deb 2002; Ravallion 1986; Baulch et al. 1998; Goletti and Farid 1995) and at domestic vs international (Alam et. al. 2012a) levels with mixed evidence. However, none of these studies addressed possible asymmetric behaviour by different market actors in vertical chains. The only exception is Alam et al. (2016) who found that wholesale market plays a leadership role in determining retail prices of rice and confirmed the fear and concerns of consumers about the price asymmetry in the vertical markets of rice. The authors employed asymmetric error correction-EG approach assuming null threshold.

4. Modeling Threshold and Asymmetry in the Price Transmission

In the first stage we consider a linear and symmetric long-run relationship between two price series as in equation (1)

$$\mathrm{RP_t} = \beta_0 + \beta_1 \mathrm{WP_t} + \varepsilon_t \qquad (1)$$

where the $\mathrm{RP_t}$ and $\mathrm{WP_t}$ are retail and wholesale prices assumed to be integrated in order one, β_0 is an intercept that accounts for transportation and quality differences, β_1 is long-run coefficient and ε_t is the error term that can be serially correlated. The error term, $\varepsilon_t = \mathrm{RP_t} - \beta_0 - \beta_1 \mathrm{WP_t}$ indicates deviations from the long-run equilibrium. According to Granger representation theorem if the error term in (1) is stationary, wholesale and retail prices are cointegrated. The second stage advocates testing for unit roots using Augmented Dickey-Fuller on the estimated residual as in equation (2)

$$\Delta\hat{\varepsilon}_t = \rho\hat{\varepsilon}_{t-1} + \sum_{i=1}^{k} \psi_i \Delta\hat{\varepsilon}_{t-i} + v_t \qquad (2)$$

where v_t is a white noise error term, k denotes the number of lags. Rejection of the null ($\rho = 0$) of non-stationarity implies that residuals of equation (1) are stationary and hence, one can conclude that wholesale and retail prices are cointegrated.

In this paper, we extend the threshold cointegration approach developed by Enders and Granger (1998). We incorporate two important properties of price transmission in vertical markets - existence of threshold in the cointegrating vector as well as possible asymmetries in threshold and in short-run price dynamics. In the presence of threshold effects, according to Enders and Granger (1998), the threshold autoregressive (TAR) model can be expressed as

$$\Delta\hat{\varepsilon}_t = \mathrm{I_t}\left[\rho_0{}^{\mathrm{OUT}} + \rho_1{}^{\mathrm{OUT}}\hat{\varepsilon}_{t-1}\right] + (1 - \mathrm{I_t})\left[\rho_0{}^{\mathrm{IN}} + \rho_1{}^{\mathrm{IN}}\hat{\varepsilon}_{t-1}\right] + \sum_{i=1}^{k} \Gamma_i \Delta\hat{\varepsilon}_{t-i} + v_t \qquad (3)$$

where, the autoregressive (AR) term of error term (ε_t) can be separated into two regimes namely the \`IN\` and the \`OUT\` regimes depending on whether the threshold variable $|\varepsilon_{t-d}|$ exceeds a threshold value θ. The \`IN\` regime defines deviations of magnitudes smaller than the threshold θ, i. e., it is inside threshold interval [-θ, θ]. The \`OUT\` regime defines when the deviations are outside the threshold interval [-θ, θ]. The Heaviside indicator function in TAR model is defined as in equation (4)

$$\mathrm{I_t} = \begin{cases} 1 \text{ if } |\varepsilon_{t-d}| \geq \theta \\ 0 \text{ if } |\varepsilon_{t-d}| < \theta \end{cases} \qquad (4)$$

where \` $\theta$ \` represents a threshold by which movement towards the long-run equilibrium are asymmetric, \` d \` is a delay parameter which represents the delay in the change from one regime to the other. This is determined through statistical procedure (Goodwin and Halt 1999; Lee and Miguel 2013). The sufficient conditions for the stationarity of the $\hat{\varepsilon}_t$ are $\rho_0 < 0$, $\rho_1 < 0$ and $(1 + \rho_0)(1 + \rho_1) < 1$ (Petrucelli and Woolford 1984).

The Heaviside indicator function in equation (4) depends on the level of $\hat{\varepsilon}_{t-1}$ but the decay could depend on the previous period change in $\hat{\varepsilon}_{t-1}$. This is especially valuable when the adjustment is asymmetric and process exhibits lopsided `momentum` in one direction (Enders and Granger 1998). In this model if $|\rho_0| < |\rho_1|$, the Threshold Autoregressive (TAR) exhibits little adjustment for positive $\Delta\hat{\varepsilon}_{t-1}$ but substantial decay for negative $\Delta\hat{\varepsilon}_{t-1}$. That means increases tend to persist but decreases tend to revert quickly back to the attractor irrespective of where disequilibrium is relative to attractor. The Heaviside indicator function could be expressed as follows.

$$I_t = \begin{cases} 1 \text{ if } |\,\Delta\varepsilon_{t-d}| \geq \theta \\ 0 \text{ if } |\,\Delta\varepsilon_{t-d}| < \theta \end{cases} \tag{5}$$

The equations (1), (3) and (4) together form TAR model and the equations (1), (3) and (5) together form MTAR model. Taking the `threshold` into account, the threshold error correction model (TECM) in vertical markets can be expressed as follows.

$$\Delta P_{i,t} = I_t\left[\rho_0{}^{OUT} + \rho_1{}^{OUT}\hat{\varepsilon}_{t-1}\right] + (1 - I_t)[\rho_0{}^{IN} + \rho_1{}^{IN}\hat{\varepsilon}_{t-1}] + \sum_{i=1}^{k} \beta_j\, \Delta P_{i,t-i} + v_{i,t} \tag{6}$$

where P_i is a vector of prices in the vertical markets, β_i is the estimates of short-run price dynamics, the error terms v_i follow white noise process with mean zero and constant variance. To identify the existence of threshold effects in the error term ε_t of equation (1), we employ a non-parametric approach following Tsay (1989) wherein we estimate recursive least squares to examine whether the coefficients of autoregressive process of ε_t are constant. We estimate TAR-F (also MTAR-F) statistics to test the null of linear process. Its rejection indicates the existence of threshold (θ). The threshold value can be assumed to be null or can be estimated.

Once the threshold effects in the AR process of ε_t is confirmed the threshold θ can be estimated using Chan`s (1993) grid search approach. The threshold values are estimated through a search over all possible threshold values minimising the residual sum of squares (RSS). The estimated residual series, the threshold variable $|\hat{\varepsilon}_{t-d}|$, is first sorted in ascending order, the largest and smallest 15 percent of residual series are eliminated and remaining 70 percent of the values are considered as possible thresholds. The estimated threshold yielding the lowest RSS is chosen as appropriate threshold. Hansen (1997) argues that null hypothesis of linearity in the AR process of $\hat{\varepsilon}_t$ does not follow a standard distribution, hence, the conventional test is inappropriate. He proposes a Chow-type test for threshold values where he used simulations methods and provided p-values based on bootstrap simulations (Hansen 1997; Goodwin and Holt 1999; Goodwin and Piggott 2001; Lee and Miguel 2013). The method in Hansen (1997) is used to estimate the maximum F-statistics and p-values using bootstrap methods. Subsequently, equation (3) is estimated in both TAR and MTAR specifications.

According to the Granger representation theorem (1987), the existence of threshold cointegration justifies estimating threshold error correction model. Once the presence of threshold effects using Tsay (1989), estimate of threshold using Chan (1993) and significance of threshold using Hansen (1997) are confirmed, a TECM can be estimated. The TECM model allows us to nest together the short- and long-run dynamics in the identified vertical markets. Since it is possible that both prices (i e., wholesale and retail prices of flour in chain I; wholesale prices of flour and wheat in chain II and finally wholesale prices of wheat and retail prices of flour in chain III) are determined simultaneously we employ Zellner`s (1963) seemingly unrelated regression (SUR) estimation in addition to OLS for each price equations in the vertical chains. This is done to select appropriate model and examine the sensitivity to models being employed. A simultaneous representation of the system of equations of TECMs for the chains I, II and III yields the following equations

Vertical chain I: Wholesale and retail prices of flour

$$\Delta RP_{(f)t} = \alpha_0 + \alpha_1{}^{OUT} I_t \hat{\varepsilon}_{t-1}{}^{OUT} + \alpha_1{}^{IN}(1 - I_t)\hat{\varepsilon}_{t-1}{}^{IN} + \sum_{i=1}^{k} \alpha_{2,i}\, \Delta RP_{(f),t-i} + \sum_{i=1}^{k} \alpha_{2,i}\, \Delta WP_{(f),t-i} + v_{1,t} \quad (7a)$$

$$\Delta WP_{(f)t} = \beta_0 + \beta_1{}^{OUT} I_t \hat{\varepsilon}_{t-1}{}^{OUT} + \beta_1{}^{IN}(1 - I_t)\hat{\varepsilon}_{t-1}{}^{IN} + \sum_{i=1}^{k} \beta_{2,i}\, \Delta WP_{(f),t-i} + \sum_{i=1}^{k} \beta_{2,i}\, \Delta RP_{(f),t-i} + v_{2,t} \quad (7b)$$

Vertical chain II: Wholesale prices of wheat and flour

$$\Delta WP_{(f),t} = \alpha_0 + \alpha_1{}^{OUT} I_t \hat{\varepsilon}_{t-1}{}^{OUT} + \alpha_1{}^{IN}(1 - I_t)\hat{\varepsilon}_{t-1}{}^{IN} + \sum_{i=1}^{k} \alpha_{2,i}\, \Delta WP_{(f),t-i} + + \sum_{j=0}^{k} \alpha_{3,j}\, \Delta WP_{(w),t-j} + v_{1,t} \quad (8a)$$

$$\Delta WP_{(w)t} = \beta_0 + \beta_1{}^{OUT} I_t \hat{\varepsilon}_{t-1}{}^{OUT} + \beta_1{}^{IN}(1 - I_t)\hat{\varepsilon}_{t-1}{}^{IN} + \sum_{i=1}^{k} \beta_{2,i}\, \Delta WP_{(w),t-i} + + \sum_{j=0}^{k} \beta_{3,j}\, \Delta WP_{(f),t-j} + v_{2,t} \quad (8b)$$

Vertical chain III: Wholesale prices of wheat and retail prices of flour

$$\Delta RP_{(f),t} = \alpha_0 + \alpha_1{}^{OUT} I_t \hat{\varepsilon}_{t-1}{}^{OUT} + \alpha_1{}^{IN}(1 - I_t)\hat{\varepsilon}_{t-1}{}^{IN} + \sum_{i=1}^{k} \alpha_{2,i}\, \Delta RP_{(f),t-i} + \sum_{j=0}^{k} \alpha_{3,j}\, \Delta WP_{(w),t-j} + v_{1,t} \quad (9a)$$

$$\Delta WP_{(w)t} = \beta_0 + \beta_1{}^{OUT} I_t \hat{\varepsilon}_{t-1}{}^{OUT} + \beta_1{}^{IN}(1 - I_t)\hat{\varepsilon}_{t-1}{}^{IN} + \sum_{i=1}^{k} \beta_{2,i}\, \Delta WP_{(w),t-i} + \sum_{j=0}^{k} \beta_{3,j}\, \Delta RP_{(f),t-j} + v_{2,t} \quad (9b)$$

where, the term $\hat{\varepsilon}_{t-1}{}^{OUT}$ represents the deviations from the long-run equilibrium which is larger than the absolute value of threshold θ and $\hat{\varepsilon}_{t-1}{}^{IN}$ represents the deviations from the long-run equilibrium within the threshold interval [-θ, θ]. The error terms v_1 and and v_2 follow white noise process.

Possible short-run asymmetries in price transmission in vertical markets can be examined by splitting the contemporary and lags of independent variables into possible and negative changes (Von Cramon-Taubadel and Loy 1996; Lee and Miguel 2013). A simultaneous representation of the system of equations of asymmetric threshold error correction models (ATECM) for the chains I, II and III, in where two major properties - threshold and non-linearity in price responses are incorporated, can be re-written as follows.

Vertical chain I: Wholesale and retail prices of flour

$$\Delta RP_{(f)t} = \alpha_0 + \alpha_1{}^{OUT} I_t \hat{\varepsilon}_{t-1}{}^{OUT} + \alpha_1{}^{IN}(1 - I_t)\hat{\varepsilon}_{t-1}{}^{IN} + \sum_{i=1}^{k} \alpha^{+}{}_{2,i}\, \Delta^{+} RP_{(f),t-i} + \sum_{i=1}^{k} \alpha^{-}{}_{2,i}\, \Delta^{-} RP_{(f),t-i}$$
$$+ \sum_{j=0}^{k} \alpha^{+}{}_{3,j}\, \Delta^{+} WP_{(f),t-j} + \sum_{j=0}^{k} \alpha^{-}{}_{3,j}\, \Delta^{-} WP_{(f),t-j} + v_{1,t} \qquad (10a)$$

$$\Delta WP_{(f)t} = \beta_0 + \beta_1{}^{OUT} I_t \hat{\varepsilon}_{t-1}{}^{OUT} + \beta_1{}^{IN}(1 - I_t)\hat{\varepsilon}_{t-1}{}^{IN} + \sum_{i=1}^{k} \beta^{+}{}_{2,i}\, \Delta^{+} WP_{(f),t-i} + \sum_{i=1}^{k} \beta^{-}{}_{2,i}\, \Delta^{-} WP_{(f),t-i}$$
$$+ \sum_{j=0}^{k} \beta^{+}{}_{3,j}\, \Delta^{+} RP_{(f),t-j} + \sum_{j=0}^{k} \beta^{-}{}_{3,j}\, \Delta^{-} RP_{(f),t-j} + v_{2,t} \qquad (10b)$$

Vertical chain II: Wholesale prices of wheat and flour

$$\Delta WP_{(f),t} = \alpha_0 + \alpha_1{}^{OUT} I_t \hat{\varepsilon}_{t-1}{}^{OUT} + \alpha_1{}^{IN}(1 - I_t)\hat{\varepsilon}_{t-1}{}^{IN} + \sum_{i=1}^{k} \alpha^{+}{}_{2,i}\, \Delta^{+} WP_{(f),t-i} + \sum_{i=1}^{k} \alpha^{-}{}_{2,i}\, \Delta^{-} WP_{(f),t-i}$$
$$+ \sum_{j=0}^{k} \alpha^{+}{}_{3,j}\, \Delta^{+} WP_{(w),t-j} + \sum_{j=0}^{k} \alpha^{-}{}_{3,j}\, \Delta^{-} WP_{(w),t-j} + v_{1,t} \qquad (11a)$$

$$\Delta WP_{(w)t} = \beta_0 + \beta_1{}^{OUT} I_t \hat{\varepsilon}_{t-1}{}^{OUT} + \beta_1{}^{IN}(1 - I_t)\hat{\varepsilon}_{t-1}{}^{IN} + \sum_{i=1}^{k} \beta^{+}{}_{2,i}\, \Delta^{+} WP_{(w),t-i} + \sum_{i=1}^{k} \beta^{-}{}_{2,i}\, \Delta^{-} WP_{(w),t-i}$$
$$+ \sum_{j=0}^{k} \beta^{+}{}_{3,j}\, \Delta^{+} WP_{(f),t-j} + \sum_{j=0}^{k} \beta^{-}{}_{3,j}\, \Delta^{-} WP_{(f),t-j} + v_{2,t} \qquad (11b)$$

Vertical chain III: Wholesale prices of wheat and retail prices of flour

$$\Delta RP_{(f),t} = \alpha_0 + \alpha_1{}^{OUT} I_t \hat{\varepsilon}_{t-1}{}^{OUT} + \alpha_1{}^{IN}(1 - I_t)\hat{\varepsilon}_{t-1}{}^{IN} + \sum_{i=1}^{k} \alpha^{+}{}_{2,i}\, \Delta^{+} RP_{(f),t-i} + \sum_{i=1}^{k} \alpha^{-}{}_{2,i}\, \Delta^{-} RP_{(f),t-i}$$
$$+ \sum_{j=0}^{k} \alpha^{+}{}_{3,j}\, \Delta^{+} WP_{(w),t-j} + \sum_{j=0}^{k} \alpha^{-}{}_{3,j}\, \Delta^{-} WP_{(w),t-j} + v_{1,t} \qquad (12a)$$

$$\Delta WP_{(w)t} = \beta_0 + \beta_1{}^{OUT} I_t \hat{\varepsilon}_{t-1}{}^{OUT} + \beta_1{}^{IN}(1 - I_t)\hat{\varepsilon}_{t-1}{}^{IN} + \sum_{i=1}^{k} \beta^{+}{}_{2,i}\, \Delta^{+} WP_{(w),t-i} + \sum_{i=1}^{k} \beta^{-}{}_{2,i}\, \Delta^{-} WP_{(w),t-i}$$
$$+ \sum_{j=0}^{k} \beta^{+}{}_{3,j}\, \Delta^{+} RP_{(f),t-j} + \sum_{j=0}^{k} \beta^{-}{}_{3,j}\, \Delta^{-} RP_{(f),t-j} + v_{2,t} \qquad (12b)$$

where, $RP_{(f)}$, $WP_{(f)}$ and $WP_{(w)}$ are the retail prices of flour, wholesale prices of flour and wholesale prices of wheat, respectively. Δ is the first difference operator. The variables $\Delta^{+}WP_{(f),t-i} = \Delta WP_{(f),t-i}$ if $\Delta WP_{(f),t-i} \geq 0$, zero otherwise; and $\Delta^{-}WP_{(f),t-i} = \Delta WP_{(f),t-i}$ if $\Delta WP_{(f),t-i} < 0$, zero otherwise. Similarly, $\Delta^{+}RP_{(f),t-i} = \Delta RP_{(f),t-i}$ if $\Delta RP_{(f),t-i} \geq 0$, zero otherwise, and $\Delta^{-}RP_{(f),t-i} = \Delta RP_{(f),t-i}$ if

$\Delta RP_{(f),t-i} < 0$, zero otherwise; $\Delta^{+}WP_{(w),t-i} = \Delta WP_{(w),t-i}$ if $\Delta WP_{(w),t-i} \geq 0$, zero otherwise, and $\Delta^{-}WP_{(w),t-i} = \Delta WP_{(w),t-i}$ if $\Delta WP_{(w),t-i} < 0$, zero otherwise. The residuals follow a white noise process.

This paper follows a systematic approach to select the appropriate specification to modelling the threshold and asymmetries in the speed of adjustment and short-run price responses in the vertical markets of wheat-to-flour in Bangladesh. First, the order of integration and cointegration are tested for non-stationarity and cointegration using unit root tests (i. e., ADF, PP, DF-GLS, ERS) and Johansen cointegration (i. e., trace and maximum eigenvalue tests). Second, causality is examined using Wald test. This test is performed to investigate the price discovery role in vertical markets, i.e., to examine if Stackelberg leadership holds true in Bangladesh wheat-to-flour markets. Third, we examine possible threshold effects and delay parameter following Tsay (1989). Once the presence of threshold effects is confirmed, we estimate the threshold θ using Chan`s (1993) grid search approach. Next, we test the significance of threshold θ using Hansen (1997) and use bootstrap simulations for p-values. Both the TAR and MTAR models are estimated for which first, we assume null threshold (θ =0) following Enders and Granger (1998), Enders and Sikolos (2001), Sanago and Mohammed (2010) and then the threshold value is estimated. Fourth, we estimate system of equations i. e., (10a and 10b), (11a and 11b), and (12a and 12b) using Zellner`s (1963) SUR model and OLS for each equation within the specified vertical markets for asymmetric threshold error correction model (ATECM). Finally, the short-run asymmetries are tested using F-statistics under the null of symmetries. We employ AIC to select the model and optimal lag length. Diagnostic checking of the residuals is performed to ascertain if residuals are free from autocorrelation and heteroscedasticity. CUSUM and recursive co-efficient are estimated to ascertain model stability.

5. Data and time series properties of data

We employ monthly data on wheat and flour prices in Dhaka, the capital city of about 18 million consumers. Data are compiled from the FAOStat for the period from July 2008 to March 2016. The price data for wholesale and retail markets of flour and wholesale markets of wheat are collected based on the data availability to analyse the threshold and asymmetries in price transmission along the vertical markets. The wholesale and retail prices of flour and wholesale prices of wheat are expressed as Taka/kilogram. The Figure 2, 3 and 4 shows the wholesale and retail prices of flour and wholesale prices of wheat in Dhaka city. Figures show that wholesale and retail prices of flour/wheat move together indicating that prices follow each other. We provide descriptive statistics of these data in Table 1.

Figure 2: Wholesale and retail prices of flour

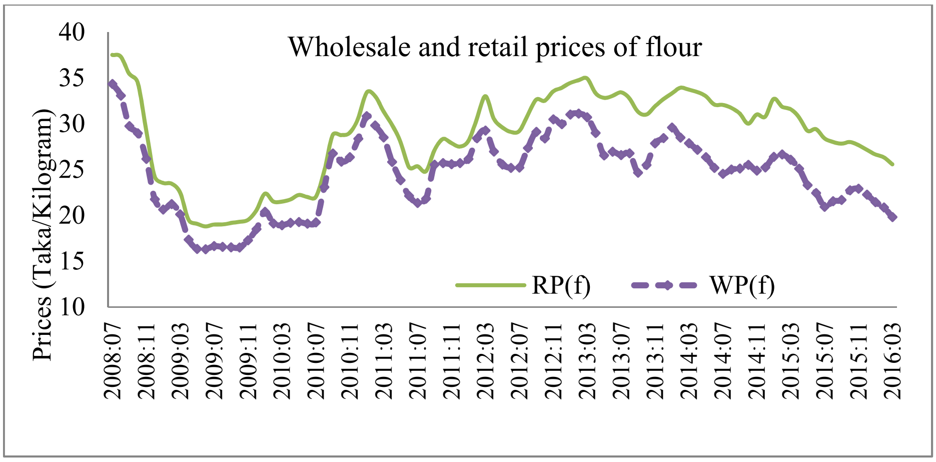

Figure 3: Wholesale prices of flour and wheat

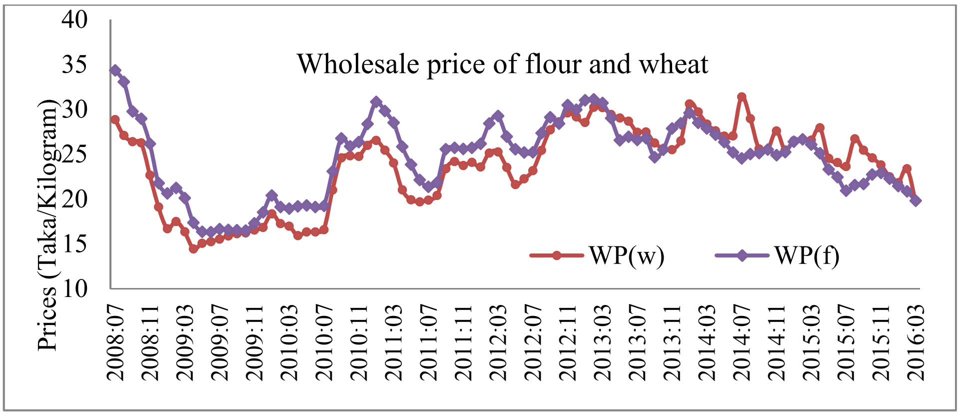

Figure 4: Wholesale prices of wheat and retail prices of flour

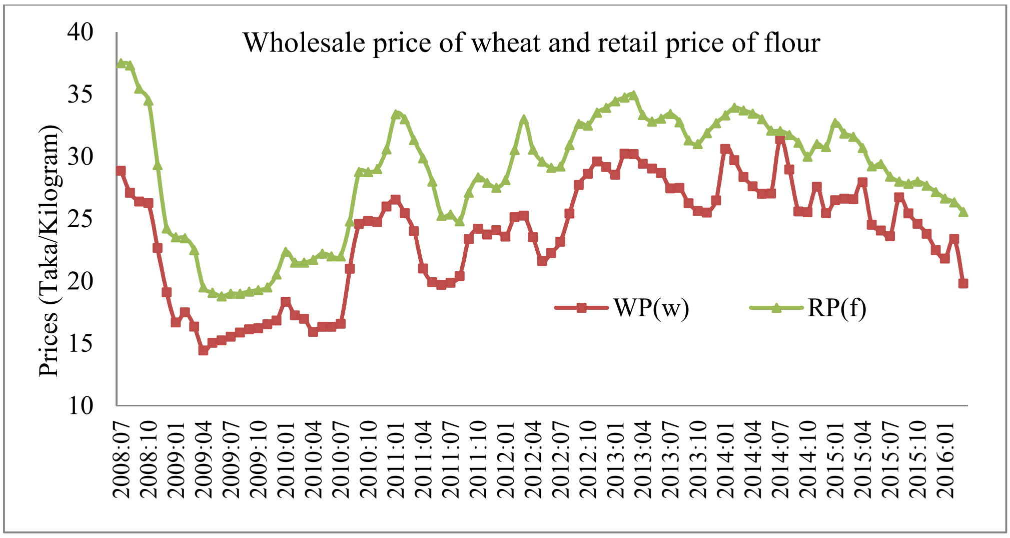

Table 1: Descriptive statistics

Variables	Names	N	Mean	Max	Min	Std. Dev.	CV
Wholesale prices -wheat (Tk/kg)	WP(w)	93	23.62	31.39	14.44	4.55	0.19
Wholesale prices -flour (Tk/kg)	WP(f)	93	24.59	34.34	16.31	4.23	0.17
Retail prices -flour (Tk/kg)	RP(f)	93	28.66	37.50	18.79	4.84	0.17

A set of unit root tests is employed to identify the order of integration of wholesale and retail prices of flour and wholesale prices of wheat. Most tests of integration assume non-stationarity under the null hypothesis. The Augmented Dickey-Fuller (ADF), Phillips-Perron, DF-GLS and ERS tests are examples of this approach. However, simulations have shown that in small samples the ADF and PP tests show lower diagnostic power than the DF-GLS and ERS tests (Elliott et al. 1996; Elliott 1999). Therefore, we conduct ADF, PP, DF-GLS and ERS tests with non-stationarity under the null hypothesis. Test results in Table 2 are robust to the alternative specifications as well as to deterministic processes (i.e. level with only drift; and level with drift and trend). The results suggest that wholesale and retail prices of flour and wholesale prices of wheat in levels contain unit root with drift or with drift and trend. However, the null hypotheses for the price series in first differences are rejected indicating that all price series are I(1).

Table 2: Unit root test results

Tests	Prices	Level with only drift		Level with drift & trend		First differenced	
		H_0	Estimated values	H_0	Estimated values	H_0	Estimated values
ADF	WP(w)	$H_0 \sim (1)$	-1.833 (l=0)	$H_0 \sim (1)$	-2.415 (l=0)	$H_0 \sim (1)$	-7.585*** (l=0)
	WP(f)	$H_0 \sim (1)$	-2.329 (l=1)[3]	$H_0 \sim (1)$	-3.089 (l=1)	$H_0 \sim (1)$	-6.133*** (l=0)
	RP(f)	$H_0 \sim (1)$	-2.621 (l=1)	$H_0 \sim (1)$	-3.314 (l=1)	$H_0 \sim (1)$	-5.883*** (l=0)
PP	WP(w)	$H_0 \sim (1)$	-2.036 (bw=2)	$H_0 \sim (1)$	-2.651 (bw=2)	$H_0 \sim (1)$	-7.511*** (bw=3)
	WP(f)	$H_0 \sim (1)$	-2.329 (bw=0)	$H_0 \sim (1)$	-2.898 (bw=4)	$H_0 \sim (1)$	-6.001*** (bw=4)
	RP(f)	$H_0 \sim (1)$	-2.315 (bw=4)	$H_0 \sim (1)$	-2.790 (bw=3)	$H_0 \sim (1)$	-5.851*** (bw=5)
DF-GLS	WP(w)	$H_0 \sim (1)$	-1.206 (l=0)	$H_0 \sim (1)$	-1.552 (l=0)	$H_0 \sim (1)$	-6.859*** (l=5)
	WP(f)	$H_0 \sim (1)$	-1.143 (l=1)	$H_0 \sim (1)$	-1.951 (l=1)	$H_0 \sim (1)$	-5.297*** (l=0)
	RP(f)	$H_0 \sim (1)$	-1.297 (l=1)	$H_0 \sim (1)$	-1.856 (l=1)	$H_0 \sim (1)$	-5.914*** (l=0)
ERS	WP(w)	$H_0 \sim (1)$	9.101 (l=0)	$H_0 \sim (1)$	23.885 (l=l)	$H_0 \sim (1)$	1.074*** (l=0)
	WP(f)	$H_0 \sim (1)$	9.685 (l=1)	$H_0 \sim (1)$	14.973 (l=1)	$H_0 \sim (1)$	0.874*** (l=0)
	RP(f)	$H_0 \sim (1)$	8.654 (l=1)	$H_0 \sim (1)$	18.634 (l=1)	$H_0 \sim (1)$	0.668*** (l=0)

Notes: MacKinnon (1996) one-sided p-values are used for critical values; Parentheses indicate the number of lags and bandwidths based on the AIC and Bartlett Kernel, respectively; *** indicates significance at 1 percent.

6. Results

Johansen (1992a, 1992b, 1995) proposes tests to determine whether two I(1) time series are cointegrated. One method is the *trace* test (Johansen 1988), which is a likelihood ratio test. The principle is to determine how many eigenvalues equal one and the test is carried out until the null hypothesis cannot be rejected. The second approach, the maximum eigenvalue test, addresses the significance of the estimated eigenvalues. Since the tests of cointegration are sensitive to the structure of data generating process, Johansen and Juselius (1990) consider five possible cases namely (i) no deterministic trend in the data, (ii) no deterministic trend in the data and an intercept in the cointegrating

[3] Modified SIC used in lag selection

equation, (iii) intercept restricted to the cointegration space, (iv) intercept in the short-run model (which corresponds to a model with drift), and (v) linear trend in the cointegration vector. Johansen (1992b) suggests testing the joint hypothesis of both rank order and deterministic components. Consequently, we move from the most restrictive model (iii) to the least restrictive model (v). At each stage the test statistics are compared to their critical values. We keep conducting these tests as long as the null hypothesis is rejected. We conduct trace and maximum eigenvalue tests for downstream prices with respect to the upstream prices for vertical markets being analysed. The tests results are presented in Table 3, where *r* is the number of cointegrating vectors. The optimal lag length and lag selection criteria used in the restricted vector auto-regressive (VAR) model are given in the first column of Table 3.

Table 3: Cointegration rank tests results

Chains	Tests	Hypothesis	Test statistics	Critical values	Decision
WP(f)-RP(f) (l=2; AIC, BIC)	λ_{trace}	$r = 0$ vs $r \geq 1$	36.173***	25.872	Rejected
		$r \leq 1$ vs $r \geq 2$	5.836	12.517	Not rejected
	λ_{max}	$r = 0$ vs $r \geq 1$	30.337***	19.387	Rejected
		$r \leq 1$ vs $r \geq 2$	5.836	12.517	Not rejected
WP(w)-WP(f) (l=1; BIC)	λ_{trace}	$r = 0$ vs $r \geq 1$	29.192**	25.872	Rejected
		$r \leq 1$ vs $r \geq 2$	8.088	12.517	Not rejected
	λ_{max}	$r = 0$ vs $r \geq 1$	21.103**	19.387	Rejected
		$r \leq 1$ vs $r \geq 2$	8.088	12.517	Not rejected
WP(w)-RP(f) (l=1; AIC)	λ_{trace}	$r = 0$ vs $r \geq 1$	39.128***	25.872	Rejected
		$r \leq 1$ vs $r \geq 2$	8.501	12.517	Not rejected
	λ_{max}	$r = 0$ vs $r \geq 1$	30.627***	19.387	Rejected
		$r \leq 1$ vs $r \geq 2$	8.501	12.517	Not rejected

Notes: Model includes both drift and trend in the cointegration space; *** and ** indicates significant at 1 percent and 5 percent, respectively.

We estimate the Johansen (1995) standard linear cointegration model to identify the cointegration relationship. We test the price leadership role with the vertical markets being analysed. The trace and eigenvalue tests indicate one cointegrating vector between the wholesale and retail prices of flour meaning that two price series are cointegrated (first panel in Table 3) indicating that these prices move together in the long-run. The null hypothesis of no cointegration is rejected against the alternative of at least one cointegration vector with p-value of 1 percent (LR_{trace}=36.173 and LR_{max}=30.337) whereas null of one cointegration vector could not be rejected (LR_{trace}=5.836 and LR_{trace}=5.836). But this result does not indicate anything about the direction of causality. We find similar results for chain II (wholesale prices of wheat and flour) and chain III (wholesale prices of wheat and retail prices of flour). We can reach similar conclusions based on the results from chain II and chain III (second and third panels of Table 3). The wholesale prices of wheat and wholesale prices of flour are cointegrated. The null hypothesis of no cointegration is rejected against the alternative of at least one cointegration vector with p-value of 5 percent (LR_{trace}=29.192 and LR_{max}=21.103) whereas null of one cointegrating vector could not be rejected (LR_{trace}=8.088 and LR_{trace}=8.088). In chain III, the wholesale prices of wheat and retail prices of flour are cointegrated. The null hypothesis of no cointegration is rejected against the alternative of at least one cointegrating vector with p-value of 1 percent (LR_{trace}=39.128 and LR_{max}=30.627) whereas null of one cointegration vector could not be rejected (LR_{trace}=8.501 and LR_{trace}=8.501). We check all diagnostic tests - Ljung-Box Q-statistics, auto-regressive conditional heteroscedasticity (ARCH) and Jarque-Bera (JB) for data normality and the CUSUM and recursive coefficient tests for model stability. The results indicate that the estimated models are free from autocorrelation, heteroscedasticity and the residuals are normally distributed. The estimated models are stable. However, since two price series are cointegrated, there is an error correction term (speed of adjustment) that corrects the deviations from its long-run equilibrium.

The estimate of speed of adjustment for retail prices of flour with respect to wholesale prices in chain I is -0.522 with the p-value of 1 percent (Table 4). This indicates that it takes about 2 months to correct the disequilibrium error. The estimate of speed of adjustment in wholesale prices with respect to retail prices is found to be insignificant. This gives additional indication that only the retail price adjusts to the changes in wholesale prices and estimating the retail prices might suffice our purpose. In chain II and chain III, the adjustment estimates are -0.250 and -0.251, respectively. The speed of adjustment estimate in wholesale prices of wheat with respect to flour wholesale prices in chain II is found to be -0.250 with the p-value of 5 percent. The estimate in retail prices of flour equation with respect to wholesale prices of wheat is found to be -0.251 with the p-value of 1 percent. This means that it takes about 4 months to correct the disequilibrium error in chain II. The same applies to chain III. The signs of the estimates are negative implying model convergence.

Table 4: Estimates of speed of adjustment

Vertical markets	Estimates	Standard errors
Flour (retail - wholesale)	-0.522***	0.151
	-0.108	0.720
Flour (wholesale) - Wheat (wholesale)	-0.168	0.094
	-0.250**	0.117
Flour (retail) - Wheat (wholesale)	-0.251***	0.097
	0.237	0.125

Note: *** and ** indicates level of significance at 1 percent and 5 percent, respectively.

The causality (week exogeneity, short-run causality and strong causality) tests are performed based on the estimated VECM models in Johansen framework. The results are presented in Table 5. In chain I, week exogeneity results show that wholesale prices of flour Granger cause retail prices of flour but not vice versa. We clearly fail to reject the null of wholesale price exogeneity. The $\chi 2$-test statistics 9.826 is rejected at 1 percent level. On the contrary, we can reject the null of retail price exogeneity- wholesale prices of flour Granger causes retail prices of flour but not vice versa. In addition, our strong exogeneity supports the same conclusion. This result is expected as upstream prices dominate the prices at downstream meaning that upstream price plays a price discovery role in the vertical markets being analysed. The short-run causality results show that there is no causality relation between the wholesale price of flour and retail price of flour. In chain II, results show that wholesale price of flour Granger causes wholesale price of wheat but not vice-versa. This result is different than what one would expect- downstream prices dominate upstream prices. We clearly fail to reject the null of exogeneity of wholesale prices of flour. The χ2-test statistics 3.001 is rejected at 10 percent level. In addition, our strong exogeneity results support similar conclusion. The χ2-test statistics 9.402 is rejected at 1 percent level. Similar to chain I, we do not find short-run causality relationship between two price series. In chain III, week exogeneity results show bidirectional causality between the wholesale prices of wheat and retail prices of flour. This indicates the possibility that the prices are simultaneously determined and it is justified to estimate SUR model. The χ2-test statistics 4.992 is rejected at 5 percent level and 2.769 is rejected at 10 percent level, respectively. In addition, strong exogeneity supports similar conclusions. The χ2-test statistics 8.608 and 6.865 are rejected at 5 percent level. Similar to chain I and chain II, we do not find short-run causality relationship.

Table 5: Results from the Wald test for causality

Chains	Causations	Hypotheses	χ2-test statistics	Causality
WP(f)-RP(f)	Week exogeneity	$\alpha_1 = 0 \ vs \ \alpha_1 \neq 0$ $\alpha_2 = 0 \ vs \ \alpha_2 \neq 0$	9.826*** (0.000) 0.351 (0.553)	WP(f) → RP(f)
	Short-run causality	$\sum \beta_i = 0 \ vs \sum \beta_i \neq 0$ $\sum \beta_j = 0 \ vs \sum \beta_j \neq 0$	2.202 (0.332) 2.455 (0.292)	WP(f) ≠ RP(f)
	Strong exogeneity	$\sum \beta_i = 0, \alpha_1 = 0 \ vs \sum \beta_i \neq 0, \alpha_1 \neq 0$ $\sum \beta_j = 0, \alpha_2 = 0 \ vs \sum \beta_j \neq 0, \alpha_2 \neq 0$	18.667*** (0.000) 2.669 (0.445)	WP(f)→ RP(f)
WP(w)-WP(f)	Week exogeneity	$\alpha_1 = 0 \ vs \ \alpha_1 \neq 0$ $\alpha_2 = 0 \ vs \ \alpha_2 \neq 0$	2.066 (0.150) 3.001* (0.083)	WP(f) →WP(w)
	Short-run causality	$\sum \beta_i = 0 \ vs \ H_1: \sum \beta_i \neq 0$ $\sum \beta_j = 0 \ vs \ H_1: \sum \beta_j \neq 0$	0.033 (0.845) 1.833 (0.175)	WP(f) ≠ WP(w)
	Strong exogeneity	$\sum \beta_i = 0, \alpha_1 = 0 \ vs \sum \beta_i \neq 0, \alpha_1 \neq 0$ $\sum \beta_j = 0, \alpha_2 = 0 \ vs \sum \beta_j \neq 0, \alpha_2 \neq 0$	3.498 (0.173) 9.402*** (0.009)	WP(f) →WP(w)
WP(w)-RP(f)	Week exogeneity	$\alpha_1 = 0 \ vs \ \alpha_1 \neq 0$ $\alpha_2 = 0 \ vs \ \alpha_2 \neq 0$	4.992** (0.025) 2.769* (0.096)	WP(w) ↔ RP(f)
	Short-run causality	$\sum \beta_i = 0 \ vs \ H_1: \sum \beta_i \neq 0$ $\sum \beta_j = 0 \ vs \ H_1: \sum \beta_j \neq 0$	0.001 (0.968) 2.194 (0.138)	WP(w) ≠ RP(f)
	Strong exogeneity	$\sum \beta_i = 0, \alpha_1 = 0 \ vs \sum \beta_i \neq 0, \alpha_1 \neq 0$ $\sum \beta_j = 0, \alpha_2 = 0 \ vs \sum \beta_j \neq 0, \alpha_2 \neq 0$	8.608** (0.013) 6.865** (0.032)	WP(w) ↔ RP(f)

Notes: →, ↔ and ≠ means unidirectional causality, bidirectional causality and no causality, respectively. ***, ** and * indicates level of significance at 1 percent, 5 percent and 10 percent, respectively.

Next, we estimate for threshold cointegration and test long-run asymmetry in the speed of adjustment. As explained in Section 3, we consider two cases. First, threshold is set to zero (case 1) and second, threshold is estimated (case 2). This is done to find the robust results from the models being analyzed.

Case 1: When θ is known and equal to zero

We use the Heaviside indicator function and estimate the equations (1), (3) and (4) for the TAR model and the equations (1), (3) and (5) for the MTAR model in bivariate framework. We estimate both TAR and MTAR models for six prices in three vertical chains. The results are as follows

Chain I (Wholesale and retail prices of flour): Since our price discovery results provide mixed evidence on the role of price leadership between the chain actors (wholesalers and retailers), we estimate both retail and wholesale price equations separately. First, we estimate the retail prices of flour with respect to wholesale prices of flour. Second, we estimate the wholesale prices of flour with respect to retail prices of flour. The TAR and M-TAR models are estimated and tested for cointegration and long-run symmetry (i.e., asymmetry in the speed of adjustment) where the Heaviside indicator function is identified based on null threshold. Ghoshray and Ghosh (2011), Sanogo and Mohamme (2010), assume the null threshold for estimating TAR and MTAR models. We use AIC and BIC to select the optimal number of lags. The number of lags and deterministic terms (i. e., drift, trends) included in the model are presented in Table 6. The TAR and MTAR models are validated by interpreting the F-statistics of joint null hypothesis, $\rho_1 = \rho_2 = 0$ by Φ_μ. The F-statistics of Φ_μ is compared with the values tabulated by Enders and Siklos (2001). In retail price equation, we reject the null of $\rho_1 = \rho_2 = 0$ only in TAR model implying that the wholesale and retail prices of flour are cointegrated. The F-statistics is found to be 10.962 and significant at 1 percent. This result is similar to Johansen cointegration test results. We find the signs of estimates ρ_1 and ρ_2 are consistent and significant at 1 percent level again only in TAR model. We use t-statistics to test the significance of null hypotheses. The model converges when both estimates ρ_1 and ρ_2 are negative (necessary conditions) (Enders and Siklos, 2001).Estimates of the adjustment speed are $\rho_1 = -0.319$ and $\rho_2 = -0.409$ suggesting model convergence. The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist compared to negative price deviations from the long-run equilibrium relationship. However, we cannot reject the null ($H_o: \rho_1 = \rho_2$) of long-term symmetry. The estimated F-statistics from the TAR model is 0.169 with the probability value of 0.682, suggesting that two speed of adjustments are not statistically different. We fail to reject the null of no cointegration in the case of MTAR model (Table 6). This indicates that TAR model fits the data better than the MTAR model when we estimate retail price equation with respect to wholesale prices. This result supports the Stackelberg model -the upstream price dominates the price at downstream. We find similar results - wholesale and retail prices of flour are cointegrated when we estimate wholesale prices with respect to retail prices, but only in TAR model. The estimates of the speed of adjustment $\rho_1 = -0.322$ and $\rho_1 = -0.339$ are significant at 1 percent. The model converges as the sign of both parameters are negative. We could reject the null of cointegration ($\rho_1 = \rho_2 = 0$) by Φ_μ at 10 percent significant level. The test statistics is found to be 6.017. The critical values are taken from Enders and Siklos (2001). However, we fail to reject the null of symmetry. Similar to the retail price equation, we fail to reject the null of no cointegration meaning that the wholesale and retail prices of flour are cointegrated in the long-run, as expected. Also, the results indicate that the TAR model fits the data better. The results are independent irrespective of normalization of prices (wholesale or retail prices) in the TAR model. We conduct the model diagnostics tests. We test Ljung-box Q-statistics up to 4, 8 and 12 lags and ARCH. The results indicate that the estimated models (TAR and MTAR) for both retail and wholesale prices of flour are free from serial correlation and heteroscedasticity (Table 6). For chain I (wholesale and retail markets of flour), we conclude that wholesale and retail prices of flour are cointegrated in the long-run only in TAR model. Also no evidence of asymmetry. These results point to testing asymmetry with unknown threshold.

Chain II (wholesale markets of wheat and flour): Similar to chain I, both wholesale prices of wheat and wholesale prices of flour are estimated separately. We estimate wholesale prices of flour with respect to wholesale prices of wheat and then wholesale prices of wheat with respect to wholesale prices of flour. The first equation is specified based on the Stackelberg model - upstream prices in vertical markets dominate the prices at downstream. The second equation is estimated to examine the results` robustness. Table 8 presents the number of lags and deterministic terms included in the models. We use AIC and BIC to select the optimal lag length. Recall that we estimate the OLS in first stage and save the residual to estimate the TAR and MTAR models. The models are validated by interpreting the F-statistics of joint null hypothesis, $\rho_1 = \rho_2 = 0$ by Φ_μ. The F-statistics of Φ_μ is compared with the values tabulated by Enders and Siklos (2001). In the equation for wholesale prices of flour, we can reject the null of $\rho_1 = \rho_2 = 0$ in both TAR and MTAR models implying that the wholesale prices of wheat and wholesale prices of flour are cointegrated, as expected. The F-statistics 11.131 and 6.660 in TAR and MTAR model are statistically significant at 1 percent and 5 percent, respectively. This result is similar to Johansen cointegration results. The signs of ρ_1 and ρ_2 are consistent. As previously mentioned, according to Enders and Siklos (2001), the model converges when both estimates ρ_1 and ρ_2 are negative (necessary conditions). We find that the estimates $\rho_1 = -0.345$ and $\rho_2 = -0.454$ suggesting model convergence. The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist compared to negative price deviations from the long-run equilibrium relationship. However, we fail to reject the null ($H_o: \rho_1 = \rho_2$) of long-term symmetry. The estimated F-statistics in TAR model is 0.447 with the p-value of 0.505 suggesting that two speed of adjustments are statistically not different. However, we find evidence of asymmetry in MTAR model implying that two speed of adjustments are statically different. In wheat wholesale price equation, we fail to reject the null of cointegration relationship in TAR model but we can reject the null in MTAR model. Also, we find evidence of asymmetry in speed of adjustment. We check the model diagnostics. The estimated Ljung-box Q-statistics up to 4, 8 and 12 lags and ARCH tests indicate that the TAR and MTAR models for both retail and wholesale prices are free from serial correlation and heteroscedasticity. The test results are presented in Table 8. So, for chain II (wholesale prices of wheat and wholesale prices of flour), we conclude that wholesale prices of wheat and wholesale prices of flour are cointegrated in the case of both TAR and MTAR models. Also, there is evidence of asymmetry in the speed of adjustments (long-term asymmetry) in both wholesale prices of wheat and wholesale prices of flour equations, but only in MTAR model.

Table 6: TAR and MTAR estimates and hypotheses tests

Hypotheses tests/Model parameters/ Model diagnostics	Chain I: WP(f)-RP(f)				Chain II: WP(w)-WP(f)				Chain III: WP(w)-RP(f)			
	RP(f)		WP(f)		WP(f)		WP(w)		RP(f)		WP(w)	
	TAR	MTAR	TAR	MTAR	TAR	MTAR	TAR	MTAR	TAR	MTAR	TAR	MTAR
ρ_1	-0.319*** (-2.713)	-0.169 (-1.113)	-0.322*** (-2.843)	-0.009 (-0.046)	-0.345*** (-2.782)	0.568*** (2.971)	-0.195*** (-2.702)	-501*** (-3.091)	-0.073 (-0.602)	0.416 (1.922)	-0.804*** (-3.587)	-0.440** (-2.567)
ρ_2	-0.409*** (-2.718)	-0.033 (-0.235)	-0.339*** (-2.485)	-0.159 (-0.796)	-0.454*** (-4.016)	-0.507*** (-3.032)	-0.122 (-1.435)	0.629*** (3.199)	-0.642*** (-4.320)	-0.415** (-2.394)	-0.544** (-2.369)	0.381 (1.878)
No. of lags and deterministic terms included in TAR & M-TAR models	l=2; constant	l=0; no constant	l=4; no constant	l=0; constant	l=1; no constant	l=0; constant	l=0; no constant	l=0; constant	l=0; constant	l=0; constant	l=4; constant	l=0; constant
AIC	2.154	2.353	2.091	2.311	3.272	3.357	3.299	3.272	3.149	3.325	3.148	3.251
BIC	2.293	2.408	2.260	2.394	3.355	3.440	3.354	3.355	3.231	3.408	3.345	3.334
Hypothesis tests												
Φ_μ : Cointegration H_o: $\rho_1 = \rho_2 = 0$	10.962***	0.646	6.017*	0.411	11.131***	6.660**	4.682	7.384**	12.987***	3.400	10.284***	3.802
Critical Values[4] (5%)	6.01	5.98	5.20 (10%)	5.98	5.98	5.98	5.04	5.98	5.98	6.51	6.28	5.98
Long-term symmetry (H_o: $\rho_1 = \rho_2$)	0.169 (0.682)	0.421 (0.517)	0.013 (0.909)	0.205 (0.652)	0.447 (0.505)	13.272*** (0.000)	0.424 (0.516	14.741*** (0.002)	6.205** (0.014)	6.420** (0.013)	0.589 (0.445)	7.016*** (0.009)
Model diagnostics												
Q-stat/Ljung Box Statistics (Q4)[5]	0.876	0.591	0.998	0.543	0.645	0.372	0.215	0.396	0.712	0.593	0.983	0.535
Q(8)	0.842	0.256	0.960	0.170	0.659	0.635	0.265	0.642	0.338	0.130	0.289	0.069
Q(12)	0.348	0.126	0.468	0.056	0.432	0.421	0.108	0.397	0.223	0.042	0.209	0.057
ARCH test	0.553 (0.458)	0.041 (0.839)	0.857 (0.357)	0.143 (0.706)	0.001 (0.969)	0.017 (0.896)	0.064 (0.800)	0.026 (0.870)	1.592 (0.210)	0.817 (0.368)	0.557 (0.457)	1.133 (0.289)

Note: Parentheses indicate the number of selected lags; ***, ** and * means significant at 1%, 5% and 10%, respectively.

[4] Critical values are from Enders and Siklos (2001)

[5] Significance level of the Ljung-Box statistics, Q(P) is that the first P of the residuals serial correlations are jointly equal to zero

Chain III (Wheat wholesale prices and flour retail prices): Similar to chain I and chain II, we estimate both wholesale prices of wheat and retail prices of flour separately. First, we estimate the retail prices of flour with respect to wholesale prices of wheat and then we estimate wholesale prices of wheat with respect of retail prices of flour. The number of lags and deterministic terms included in the models are presented in Table 6. We use AIC and BIC to select the optimal lag length. The TAR and MTAR models are validated by interpreting the F-statistics of joint null hypothesis, $\rho_1 = \rho_2 = 0$ by Φ_μ. The F-statistics of Φ_μ is compared with the values tabulated by Enders and Siklos (2001). In the equation for retail prices of flour the null of $\rho_1 = \rho_2 = 0$ in TAR model is rejected implying that the wholesale prices of wheat and retail prices of flour are cointegrated, as expected. The F-statistics of Φ_μ is 12.987 and significant at 1 percent level. The result is similar to Johansen cointegration test results. We find that the signs of estimates ρ_1 and ρ_2 are consistent but only ρ_2 is significant at 1 percent level. The model converges when both ρ_1 and ρ_2 are negative (necessary conditions) (Enders and Siklos, 2001). We find that the estimates $\rho_1 = -0.073$ and $\rho_2 = -0.642$ suggesting model convergence. The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist compared to negative price deviations from the long-run equilibrium. We clearly fail to reject the null ($H_o: \rho_1 = \rho_2$) of long-term symmetry. The estimated F-statistics in TAR model is 6.205 with the probability value of 0.014 suggesting that two speeds of adjustment are statistically different. The results from the MTAR model indicate no cointegration relationship. For wholesale prices of wheat, we fail to reject the null of cointegration relationship in TAR specification but can reject the null in the case of MTAR model. However, we fail to reject the null of symmetry in TAR model whereas we can reject the null of symmetry in MTAR model. We check the model diagnostics - autocorrelation and heteroscedasticity. We estimate the Ljung-box Q-statistics up to 4, 8 and 12 lags and ARCH and find that the estimated models (both the TAR and MTAR) in both retail and wholesale prices are free from serial correlation and heteroscedasticity (Table 6). So, we conclude that wholesale prices of wheat and wholesale prices of flour are cointegrated only in the case of TAR model. Also, there are evidences of asymmetry in the speed of adjustments (long term asymmetry) in both TAR and MTAR models in the retail prices of flour and only in MTAR model in wholesale prices of wheat.

Case 2: When θ is unknown and estimated

We estimate the threshold following Chan`s (1993) grid search approach. The TAR and MTAR models are estimated based on the estimated threshold. In total, we estimate six different prices in three different vertical chains. The results are presented in Table 7.

Chain I (Wholesale and retail prices of flour): Since our price discovery results provide a mix evidences on the role of price leadership between the chain actors, we estimate both retail and wholesale prices separately. First, we estimate retail prices of flour with respect to wholesale prices of flour. Then we change to estimating the wholesale prices of flour with respect to retail prices of flour. We estimate both the consistent-TAR and consistent-MTAR models and test for cointegration and long-run symmetry (symmetry in the speed of adjustments). We identify optimal lag length using SBC. We find that consistent-MTAR model fits better to the data, hence, only the results from the consistent-MTAR model are presented. The model is validated by interpreting the F-statistics of joint null hypothesis, $\rho_1 = \rho_2 = 0$ by Φ_μ. The F-statistics of Φ_μ is compared with the values tabulated by Enders and Siklos (2001). We reject the null of $\rho_1 = \rho_2 = 0$ in the consistent-MTAR model implying that the wholesale and retail prices of flour are cointegrated in the long-run. The result is similar to result found by Johansen cointegration test and model estimated using null threshold. We find that the signs of estimates ρ_1 and ρ_2 are consistent and significant at 1 percent level. The model converges when both ρ_1 and ρ_2 are negative (necessary conditions) (Enders and Siklos, 2001). We find that the estimates of the adjustment speed $\rho_1 = -0.167$ and $\rho_2 = -0.390$ suggesting model convergence. The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist compared to negative price deviations from the long-run equilibrium. We use AIC and BIC to select the optimal lag length. The delay parameter \`d\` is identified based on the Tsay (1989) (i. e choosing \`d\` that maximizes the F-statistics (Goodwin and Holt 1999; Goodwin and Piggot 2001; Lee and Miguel 2013). For retail price of flour, the Tsay (1989) test finds strong evidence of non-linearity in $\hat{\varepsilon}_{t-1}$. The estimated F-statistics is 6.648 and rejected at 5 percent level. This implies that the null of a liner AR process in the cointegrated vector is rejected at 5 percent level. The percent share of observation in the inside regime (i e deviations from the long-run equation in the interval [-θ, θ] is 46 and outside regime is 54 - a well distribution of observations, indicating that identified threshold is useful. Since, nonlinearities are found in the error correction term, we proceed to estimate the threshold value (θ) using Chan\`s (1993) approach. As mentioned in the modelling section, the threshold values are estimated through a search over all possible threshold values minimizing sum of square errors (SSE). The estimated threshold is 0.589 that minimizes the SSE. Hansen (1997) argues that conventional test is not appropriate since null of linearity in the AR process does not follow a standard distribution. Hansen proposes a Chow tests for threshold values using simulations and provides asymptotic p-values based on bootstrapping (Hansen 1997; Goodwin and Halt 1999; Lee and Miguel 2013). Hansen (1997) tests also reject the null hypothesis of no threshold effects at 6 percent level of significance. The max-F statistics value is 5.184. This result provides additional evidence of threshold effects in the cointegrating vector between the retail prices and wholesale prices of flour. The F-statistics to test the null of symmetry (in Table 7) confirms the existence of the long-run asymmetry across regimes supporting the null of presence of nonlinearities in the error correction term. We reject the null ($H_o: \rho_1 = \rho_2$) of long-term symmetry. We can arrive at similar conclusions when we estimate wholesale prices of flour with respect to retail prices (second panel in Table 7). The estimates of speed of adjustment are $\rho_1 = -0.154$ and $\rho_1 = -0.403$. The estimates are statistically significant at 1 percent level. The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist compared to negative price deviations from the long-run equilibrium. The model converges as the signs of both estimates are negative. We can reject the null of no cointegration ($\rho_1 = \rho_2 = 0$) by Φ_μ at 1 percent significant level. The threshold value is found to be 0.588. We find evidence of nonlinearity in the error correction and evidence of long-run asymmetry (asymmetry in the speed of adjustment) and a well distribution of observation in `IN` regime (48 percent) and the OUT regime (52 percent).

Table 7: Consistent-TAR/Consistent-MTAR estimates and hypotheses tests

Chains	Normalized equations & model	Estimates and hypotheses tests	Values
WP(f)-RP(f)	RP(f)=f(WP(f)) (consistent-MTAR)	Optimal lag length	0
		Delay parameter	4
		Tsay test & probaility value (F-stat) (H_0:No linear process)	6.648** (0.011)
		Threshold cointegration Test (bootstrap p-value)	5.184* (0.060)
		Estimated threshold (γ) using Chan`s (1993) grid search	0.589
		Cointegration ($H_0: \rho_1 = \rho_2 = 0$) (F-stat)	8.476*** (0.000)
		Long-run asymmetry across regimes ($H_0: \rho_1 = \rho_2$) (F-stat)	3.276* (0.071)
		ρ_1	-0.167* (0.062)
		ρ_2	-0.390*** (0.000)
		Number and percentage of observation in regime `IN`	42 (45.65%)
		Number and percentage of observation in regime `OUT`	50 (54.35%)
	WP(f)=f(RP(f)) (consistent-MTAR)	Optimal lag length	0
		Delay parameter	4
		Tsay test & probaility value (F-stat) (H_0:No linear process)	3.423* (0.067)
		Threshold cointegration Test (bootstrap p-value)	4.8671*(0.080)
		Estimated threshold (γ) using Chan`s (1993) grid search	0.58849
		Cointegration ($H_0: \rho_1 = \rho_2 = 0$) (F-stat)	8.662*** (0.000)
		Long-run asymmetry across regimes ($H_0: \rho_1 = \rho_2$) (F-stat)	3.076* (0.082)
		ρ_1	-0.154*(0.10)
		ρ_2	-0.403*** (0.000)
		Number and percentage of observation in regime `IN`	44 (47.8%)
		Number and percentage of observation in regime `OUT`	48 (52.2%)
WP(w)-WP(f)	WP(f)=f(WP(w)) (consistent-TAR)	Optimal lag length	0
		Delay parameter	5
		Tsay test & probaility value (F-stat) (H_0:No linear process)	6.787*** (0.010)
		Threshold cointegration test (bootstrap p-value)	2.393 (0.550)
		Estimated threshold (γ) using Chan`s (1993) grid search	1.214
		Cointegration ($H_0: \rho_1 = \rho_2 = 0$) (F-stat)	11.192*** (0.000
		Long-run asymmetry across regimes ($H_0: \rho_1 = \rho_2$) (F-stat)	3.742** (0.050)
		ρ_1	-0.602*** (0.000)
		ρ_2	-0.250*** (0.006)
		Number and percentage of observation in regime `IN`	39 (42.4%)
		Number and percentage of observation in regime `OUT`	53 (57.6%)
	WP(w)=f(WP(f))	Optimal lag length	0

	(consistent-MTAR)	Delay parameter (based on larger MTAR-F stat)	8
		Tsay test & probaility value (F-stat) (H_0:No linear process)	2.939* (0.090)
		Threshold cointegration Test (bootstrap p-value)	5.982**(0.023)
		Estimated threshold (γ) using Chan`s (1993) grid search	0.806
		Cointegration ($H_0: \rho_1 = \rho_2 = 0$) (F-stat)	5.384*** (0.006)
		Long-run asymmetry across regimes ($H_0: \rho_1 = \rho_2$) (F-stat)	1.403 (0.239)
		ρ_1	-0.111 (0.132)
		ρ_2	-0.242*** (0.004)
		Number and percentage of observation in regime `IN`	41 (44.6%)
		Number and percentage of observation in regime `OUT`	51 (53.4%)
WP(w)-RP(f)	RP(f)=*f*(WP(w)) (consistent-TAR)	Optimal lag length	0
		Delay parameter	12
		Tsay test & probaility value (F-stat) (H_0:No linear process)	3.108*(0.08)
		Threshold cointegration Test (bootstrap p-value)	5.412** (0.030)
		Estimated threshold (γ) using Chan`s (1993) grid search	0.418
		Cointegration ($H_0: \rho_1 = \rho_2 = 0$) (F-stat)	15.883*** (0.000
		Long-run asymmetry across regimes ($H_0: \rho_1 = \rho_2$) (F-stat)	10.747*** (0.001
		ρ_1	-0.267*** (0.000)
		ρ_2	-1.299*** (0.000)
		Number and percentage of observation in regime `IN`	78 (84%)
		Number and percentage of observation in regime `OUT`	14 (16%)
	WP(w)=*f*(RP(f)) (consistent-TAR)	Optimal lag length)	0
		Delay parameter	4
		Tsay test & probaility value (F-stat) (H_0:No linear process)	2.866* (0.09)
		Threshold cointegration Test (bootstrap p-value)	2.269 (0.640)
		Estimated threshold (γ) using Chan`s (1993) grid search	1.273
		Cointegration ($H_0: \rho_1 = \rho_2 = 0$) (F-stat)	16.139*** (0.000
		Long-run asymmetry across regimes ($H_0: \rho_1 = \rho_2$) (F-stat)	6.733** (0.011)
		ρ_1	-0.093
		ρ_2	-0.583*** (0.000)
		Number and percentage of observation in regime `IN`	34 (36.95%)
		Number and percentage of observation in regime `OUT`	58 (53.05%)

Notes: Optimal lags are determined by SBC
Delay parameters are chosen by the lags giving the largest TAR-F/MTAR-F statistics from Tsay test.
The null hypothesis of Tsay test is that AR follows a linear process in a recursive least square estimation.
The null hypothesis of Hansen test (1997) is `no threshold effects in autoregressive representation of variable`. The F-test for no threshold effects in autoregressive representation of variable.
***, ** and * indicates level of significance at 1 percent, 5 percent and 10 percent, respectively.
The F-test for no thresholds effects and parenthesis indicates asymptotic p-value of bootstrap simulations with 300 replications.

Chain II (Wholesale prices of wheat and flour): Similar to chain I, since our price leadership result provides a mix evidence on the role of price discovery between the vertical chain actors, we estimate both wholesale prices of wheat and wholesale prices of flour separately. First, we estimate wholesale prices of flour with respect to wholesale prices of wheat. Next, we estimate wholesale prices of wheat with respect to wholesale prices of flour. We estimate both the consistent-TAR and consistent-MTAR models and test the cointegration and long-run symmetry (in other words asymmetry in the speed of adjustment). The optimal number of lags is identified using BIC criteria. We find that consistent-TAR model fits better to the data, hence we present the results from the consistent-TAR model in the case of wholesale prices of flour. The model is validated by interpreting the F-statistics of joint null hypothesis, $\rho_1 = \rho_2 = 0$ by Φ_μ. The F-statistics of Φ_μ is compared with the values tabulated by Enders and Siklos (2001). We reject the null of $\rho_1 = \rho_2 = 0$ in the consistent-TAR model implying that wholesale prices of wheat and wholesale prices of flour are cointegrated. This result is similar to that found by Johansen cointegration test. We find that the signs of estimates ρ_1 and ρ_2 are consistent and significant at 1 percent. The model converges when both ρ_1 and ρ_2 are negative (necessary conditions) (Enders and Siklos, 2001). We find that the estimates of the adjustment speed $\rho_1 = -0.602$ and $\rho_1 = -0.250$ suggesting model convergence. The delay parameter `$d$` is identified based on the Tsay (1989) (i. e choosing `d` that maximizes the F-statistics (Goodwin and Holt 1999; Goodwin and Piggot 2001; Lee and Miguel 2013). For wholesale price of flour, we find that the consistent-MTAR model fits the data better. The Tsay (1989) test finds strong evidence of non-linearity in error correction ε_{t-1}. The estimated F-statistics 6.787 is rejected at 1 percent. This implies that the null of a liner AR process in the cointegrated vector is rejected at 1 percent. The percent share of observation in the inside regime (i e deviations from the long-run equation in the interval [-θ, θ] is 42 percent and outside regime is 58 percent, a well distribution of observations, indicating that identified threshold provides useful information. Since, the nonlinearities in the error correction term are found, we proceed to estimate the threshold (θ) using Chan`s (1993) approach. The estimated threshold is 1.214 that minimizes the RSS. Hansen (1997) test fails to reject the null of no threshold effects. So, we conclude that nonlinearity exists in $\hat{\varepsilon}_{t-1}$. However, the F-statistics to test the null of symmetry confirms the existence of the long-run asymmetries across regimes supporting the null of presence of nonlinearities in the error correction term (Table 7). However, we reject the null ($H_o: \rho_1 = \rho_2$) of long-term symmetry. We can make similar conclusions when we estimate wholesale prices of wheat with respect to wholesale prices of flour (fourth panel in Table 7). The estimates of the speed of adjustment are $\rho_1 = -0.111$ and $\rho_1 = -0.242$. Only the ρ_2 is statistically significant at 1 percent. The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist compared to negative price deviations from the long-run equilibrium. The model is converging as the signs and magnitudes of both parameters are consistent. We can reject the null of no cointegration ($\rho_1 = \rho_2 = 0$) by Φ_μ at 1 percent significance level. We find the threshold value 0.806. But surprisingly we do not find evidence of asymmetry when we estimate wholesale prices of wheat with respect to wholesale prices of flour. This indicates that estimating MTAR model for wholesale prices of flour is the appropriate choice.

Chain III (Wholesale prices of wheat and retail prices of flour): Similar to chain I and chain II, we estimate both wholesale prices of wheat and retail prices of flour separately. First, we estimate retail prices of flour with respect to wholesale prices of wheat. Next, we estimate the wholesale prices of wheat with respect to retail prices of flour. We estimate both the consistent-TAR and consistent-MTAR models and test the cointegration and long-run symmetry (asymmetry in the speed of adjustments). The optimal number of lags is identified using BIC. We find that consistent-TAR model fits the data better for both the wholesale prices of wheat and retail prices of flour equations, hence we present the results only from the consistent-TAR model. The model is validated by interpreting the F-statistics of joint null hypothesis, $\rho_1 = \rho_2 = 0$ by Φ_μ. The F-statistics of Φ_μ is compared with the values tabulated by Enders and Siklos (2001). We reject the null of $\rho_1 = \rho_2 = 0$ in the consistent-TAR model implying that the wholesale prices of wheat and retail prices of flour are cointegrated. This result is similar to Johansen cointegration test result. We find that the signs of estimates ρ_1 and ρ_2 are consistent and significant at 1 percent. We use t-statistics to test the significance of null hypotheses. The model converges when both ρ_1 and ρ_2 are negative (necessary conditions) (Enders and Siklos, 2001). We find for estimates of the adjustment speed $\rho_1 = -0.267$ and $\rho_1 = -1.299$ suggesting model convergence. The delay parameter `d` is identified based on the Tsay (1989) (i. e choosing `d` that maximizes the F-Statistics (Goodwin and Holt, 1999; Goodwin and Piggot, 2001; Lee and Miguel, 2013). For wholesale prices of flour, we find that consistent-TAR model fits better to the data. Tsay (1989) test finds strong evidence of non-linearity in ε_{t-1}. The estimated F-statistics 3.108 is rejected at 8 percent level. This implies that the null of a liner autoregressive (AR) process in the cointegrated vector is rejected at 8 percent level. The percent share of observation in the inside regime (i e deviations from the long-run equation in the interval [-θ, θ] is 84 percent and outside regime is 16 percent. Since, nonlinearities are found in the error correction term, we proceed to estimate the threshold value (θ) using Chan`s (1993) approach. The estimated threshold is 0.418 that minimizes the SSE. Hansen (1997) test rejects the null of no threshold effects. The max-F statistics 5.412 is significant at 3 percent level using the bootstrap p-value. So, we conclude that nonlinearity exists in error correction term. However, the F-statistics (10.747 and significant at 1 percent level) to test the null of symmetry (in Table 7) confirms the existence of long-run asymmetry across regimes supporting the null of presence of nonlinearities in the error correction term. However, we fail to accept the null ($H_o: \rho_1 = \rho_2$) of long-term symmetry. We can make similar conclusion when we estimate wholesale prices of wheat with respect to retail prices of flour (sixth panel in Table 7). The estimates of the speed of adjustment are $\rho_1 = -0.093$ and $\rho_1 = -0.583$. Only the ρ_2 is statistically significant at 1 percent. The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist compared to negative price deviations from the long-run equilibrium. The model converges as sign and magnitude of both parameters are consistent. We reject the null of no cointegration ($\rho_1 = \rho_2 = 0$) by Φ_μ at 1 percent significant level. We find that the threshold value is 1.273. The percent share of observations in the inside regime (i e deviations from the long-run equation in the interval [-θ, θ]) is 37 percent and outside regime is 63 percent.

Table 8: Threshold test results, estimated threshold, bootstrap p-values

Chains	Normalized equations	Max F-statistics	Bootstrap[6] p-values	Estimated threshold
WP(f)-RP(f)	$RP_{(f)} = f(WP_{(f)})$	5.184*	0.060	0.589
	$WP_{(f)} = f(RP_{(f)})$	4.867*	0.080	0.588
WP(w)-WP(f)	$WP_{(f)} = f(WP_{(w)})$	2.393	0.550	1.214
	$WP_{(w)} = f(WP_{(f)})$	5.982**	0.023	0.806
WP(w)-RP(f)	$RP_{(f)} = f(WP_{(w)})$	5.412**	0.030	0.418
	$WP_{(w)} = f(RP_{(f)})$	2.269	0.640	1.273

Notes: The null is no threshold against alternative of threshold under maintained assumption of Homoskedastic errors Hansen (1997); ** and * indicates level of significance at 5 percent and 10 percent, respectively.

The estimated threshold, Max-F statistics and bootstrap p-values are presented in Table 8. We estimated six different prices under three different chains. Out of six estimated thresholds, four estimated thresholds are significant using bootstrap p-values. In chain I, irrespective of normalization of prices in OLS from where we derive the error correction term, the estimated models show existence of threshold and non-linear adjustment, The estimated thresholds are significant at 10 percent. In chain II, we find that estimated threshold is significant at 5 percent only in wholesale prices of wheat equation. In chain III, we find that threshold is significant at 5 percent level only in the case of retail prices of flour equation. All these imply the existence of non-linearity in the speed of adjustment but with mixed evidence on price leadership role. For example, in chain I, we find that non-linear adjustment exists in both prices. In chain II, upstream prices adjust to the price changes in the downstream but not the other way round. In chain III, downstream price adjusts to the price changes in the upstream. This is in line with the marketing and industrial organization literature.

Given the existence of threshold effects in the cointegrating vector in the vertical chains, we examine possible short term asymmetries in contemporary and lagged explanatory variables. We estimate both OLS and the SUR (Zellner 1963) models. We test the short-run symmetries in a way that the coefficients for positive and negative changes in the explanatory variables (contemporary and lagged explanatory variables) are equal ($\text{i.e.}, \alpha^{+} = \alpha^{-}$ and/or $\beta^{+} = \beta^{-}$). If the null is rejected we conclude the existence of short-run asymmetry in price transmission. We present the F-statistics for estimated OLS and χ^2 -statistics for estimated SUR models.

Results from the Asymmetric Threshold Error Correction model (ATECM)

Chain I: Wheat Wholesale-Flour Wholesale

Table 9 and Table 10 show the estimates and tests results from the ATECM for retail prices of flour and wholesale prices of flour equations, respectively. In retail prices of flour, we find that the estimates $\rho^{IN} = -0.433$ and $\rho^{OUT} = -0.411$ are significant at 1 percent (Table 9). The overall goodness of fit

[6] Bootstrap replication is 300

indicates that model is highly significant. The F-statistics 41.717 is rejected at 1% . The signs and magnitudes of the estimates are consistent. We select optimal lag length using AIC. Since one of our main focuses is to test PTAs, we highlight only the asymmetric test results. The null hypotheses of split contemporary variables and lagged variables are tested to find evidences of short-run asymmetries. Rejection of symmetry implies the existence of asymmetry in responding retail prices of flour to the changes in flour wholesale price. Equivalently, the retail prices adjust differently when wholesale price increases than when it falls. Results show that the null of symmetry between positive and negative variables in contemporary price changes is rejected at 5 percent (Table 9). The F-statistics is found to be 6.330. Moreover, the results show that the null of symmetry between positive and negative variables in lagged price changes is rejected at the 5 percent level. The null of symmetry between sum of positive lagged price changes and sum of negative lagged price changes is also rejected at 5 percent level. The Ljung Box Q-statistics (2.205) indicates no serial correlation. ARCH supports no heteroscedasticity. J-B test indicates that the residuals are normally distributed. The CUSUM and recursive coefficients show that the estimated model is stable.

Table 9: OLS estimates of ATECM: Retail price of flour (ΔRPf_t)

Variables	Estimates	S E	t-stat	p-values
Constant	0.069	0.124	0.565	0.574
ε_{t-1}^{IN}	-0.433***	0.090	-4.871	0.000
ε_{t-1}^{OUT}	-0.411***	0.104	-4.009	0.000
ΔRPf_{t-1}^{+}	0.072	0.181	0.398	0.692
ΔRPf_{t-1}^{-}	0.139	0.131	1.061	0.292
ΔRPf_{t-2}^{+}	0.119	0.155	0.764	0.447
ΔRPf_{t-2}^{-}	0.251	0.132	1.900	0.061
ΔWPf^{+}	0.607***	0.075	8.040	0.000
ΔWPf^{-}	0.956	0.096	10.00	0.000
ΔWPf_{t-1}^{+}	0.184	0.162	1.138	0.258
ΔWPf_{t-1}^{-}	-0.359**	0.139	-2.564	0.012
ΔWPf_{t-2}^{+}	-0.344**	0.152	-2.265	0.026
ΔWPf_{t-2}^{-}	-0.109	0.139	-0.789	0.432
F-statistics	41.717***			
Short-run asymmetries (F-statistics)				
$H_o: \Delta WPf^{+}{}_{t} = \Delta WPf^{-}{}_{t}$	6.330**	0.014		
$H_o: \Delta WPf^{+}{}_{t-1} = \Delta WPf^{-}{}_{t-1}$	5.647**	0.020		
$H_o: \Delta WPf^{+}{}_{t-2} = \Delta WPf^{-}{}_{t-2}$	1.107	0.296		
$H_0: \Delta WPf^{+}{}_{t} + \Delta WPf^{+}{}_{t-2} = \Delta WPf_{t} + \Delta WPf^{-}{}_{t-2}$	5.427**	0.023		
Model Diagnostics				
Ljung-Box Q-Stat for autocorrelation	2.206 (0.332)			
Jarque-Bera	1.743 (0.418)			
ARCH	0.019 (0.890)			
CUSUM test	Stable			
Recursive coefficient estimates	Inside the band of ±2 SE			

Note: *** and ** indicates level of significance at 1 percent and 5 percent, respectively.

In wholesale price of flour equation, we find that the estimates of $\rho^{IN}$and ρ^{OUT} regimes are significant at 1 percent level. The signs and magnitudes of the estimates are consistent. We find the estimates $\rho^{IN} = -0.388$ and $\rho^{OUT} = -0.465$ (Table 10). The model converges as the signs of both parameters are negative. The overall goodness of fit indicates that the model is highly significant. The F-statistics of 39.559 is rejected at 1 percent. The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist compared to negative price deviations from the long-run equilibrium relationship. We select optimal lag length using AIC. Results show that the null of symmetry between positive and negative variables in contemporary price changes is rejected at 1 percent (Table 10). The F-statistics is found to be 21.734. Moreover, the results show that the null of symmetry between positive and negative variables in lagged price changes is rejected at the 10 percent level. Also, the null of symmetry between sum of positive lagged price changes and negative price changes is also rejected at 1 percent level. This means that the wholesale prices adjust differently when retail price increases than when it falls. The results are subject to model diagnostics. The Ljung Box Q-statistics (1.659) indicates no serial correlation. ARCH test (0.052) supports no heteroscedasticity. JB test result indicates that the residuals are normally distributed. The CUSUM and recursive coefficients estimates show that the model is stable (Last panel of Table 10).

Table 10: OLS estimates of ATECM: Wholesale price of flour (ΔWPf_t)

Variables	Estimates	S E	t-stat	p-values
Constant	-0.155	0.126	-1.234	0.221
ε_{t-1}^{IN}	-0.388***	0.106	-3.664	0.000
ε_{t-1}^{OUT}	-0.465***	0.110	-4.206	0.000
ΔWPf_{t-1}^{+}	0.006	0.172	0.035	0.972
ΔWPf_{t-1}^{-}	0.224	0.147	1.529	0.131
ΔWPf_{t-2}^{+}	0.401**	0.162	2.482	0.015
ΔWPf_{t-2}^{-}	0.198	0.142	1.404	0.164
ΔRPf_{t}^{+}	1.318***	0.096	13.587	0.000
ΔRPf_{t}^{-}	0.689***	0.082	8.398	0.000
ΔRPf_{t-1}^{+}	-0.302	0.187	-1.609	0.111
ΔRPf_{t-1}^{-}	0.137	0.135	1.017	0.312
ΔRPf_{t-2}^{+}	-0.129	0.163	-0.790	0.431
ΔRPf_{t-2}^{-}	-0.344**	0.132	-2604	0.011
F-Statistics	39.559***			
Short-run asymmetries (F-statistics)				
$H_o: \Delta RPf^{+}{}_{t} = \Delta RPf_{t}$	21.734*** (0.00)			
$H_o: \Delta RPf_{t-1} = \Delta RPf^{-}{}_{t-1}$	3.090* (0.080)			
$H_o: \Delta RPf^{+}{}_{t-2} = \Delta RPf^{-}{}_{t-2}$	0.843 (0.361)			
$H_0: \Delta RPf^{+}{}_{t} + \Delta RPf^{+}{}_{t-2} = \Delta RPf^{-}{}_{t} + \Delta RPf^{-}{}_{t-2}$	8.806*** (0.004)			
Model Diagnostics				
Ljung-Box Q-Stat for autocorrelation	1.659 (0.436)			
Jarque-Bera	1.361 (0.506)			
ARCH	0.052 (0.819)			
CUSUM test	Stable			
Recursive coefficient estimates	Inside the band of ±2 SE			

Note: *** , ** and * indicates level of significance at 1 percent, 5 percent, and 10 percent, respectively.

Next, we estimate SUR model to examine the results robustness. In retail price of flour equation, we find that the estimates of $\rho^{IN}$and ρ^{OUT} regimes are significant at 1 percent level. The signs and magnitudes of the estimates are consistent. We find the estimates $\rho^{IN} = -0.369$ and $\rho^{OUT} = -0.361$ (Table 11). The model converges as the signs of both parameters are negative. We select optimal lag length using AIC. Results show that the null of symmetry between positive and negative variables in contemporary price changes is rejected at 5 percent (Table 11). The F-statistics is 6.195. Similarly, the null of symmetry between positive and negative variables in lagged price changes is rejected at 5 percent. This means that the wholesale prices adjust differently when retail price increases than when it falls. In wholesale prices of flour equation, we find that the estimates of $\rho^{IN}$and ρ^{OUT} regimes are significant at 1 percent. The signs and magnitudes of the estimates are consistent. We find the estimates $\rho^{IN} = -0.408$ and $\rho^{OUT} = -0.400$ (Table 11). The model converges as the signs of both parameters are negative. We select optimal lag length using AIC. Results show that the null of symmetry between positive and negative variables in contemporary price changes is rejected at 1 percent (Table 11). The F-statistics is 12.235. This means that the wholesale prices adjust differently when retail price increases than when it falls.

Table 11: Seemingly Unrelated Regression (SUR) estimates of ATECM

	Retail price of flour (ΔRPf)				Wholesale price of flour (ΔWPf)			
Variables	Coeff	SE	T-stat	p-values	Coeff	SE	T-stat	p-values
Constant	-0.005	0.108	-0.041	0.967	-0.050	0.111	-0.452	0.651
ε_{t-1}^{IN}	-0.369***	0.078	-4.725	0.000	-0.408***	0.090	-4.514	0.000
ε_{t-1}^{OUT}	-0.361***	0.088	-4.097	0.000	-0.400***	0.094	-4.269	0.000
ΔRPf_t^+					1.269***	0.075	17.013	0.000
ΔRPf_t^-					0.933***	0.057	16.395	0.000
ΔRPf_{t-1}^+	0.1238	0.165	0.751	0.452	-0.165	0.172	-0.963	0.335
ΔRPf_{t-1}^-	0.068	0.115	0.595	0.552	-0.016	0.119	-0.135	0.892
ΔRPf_{t-2}^+	0.153	0.143	1.069	0.285	-0.161	0.151	-1.071	0.284
ΔRPf_{t-2}^-	0.249**	0.119	2.096	0.036	-0.240**	0.121	-1.994	0.046
ΔWPf_t^+	0.769***	0.057	13.485	0.000				
ΔWPf_t^-	1.008***	0.066	15.231	0.000				
ΔWPf_{t-1}^+	0.133	0.147	0.907	0.364	-0.126	0.155	-0.809	0.418
ΔWPf_{t-1}^-	-0.332***	0.128	-2.596	0.009	0.288**	0.133	2.162	0.030
ΔWPf_{t-2}^+	-0.341**	0.141	-2.420	0.015	0.409***	0.149	2.752	0.005
ΔWPf_{t-2}^-	-0.099	0.126	-0.787	0.431	0.0841	0.129	0.654	0.513
Short-run asymmetries (χ^2 -statistics)								
$H_o: \Delta WPf_t^+ = \Delta WPf_t^-$			6.195**	0.012	$H_o: \Delta RPf_t^+ = \Delta RPf_t^-$		12.235***	0.000
$H_o: \Delta WPf_{t-1}^+ = \Delta WPf_{t-1}^-$			4.994**	0.025	$H_o: \Delta RPf_{t-1}^+ = \Delta RPf_{t-1}^-$		0.448	0.503
$H_o: \Delta WPf_{t-2}^+ = \Delta WPf_{t-2}^-$			1.391	0.238	$H_o: \Delta RPf_{t-2}^+ = \Delta RPf_{t-2}^-$		0.135	0.713

Note: *** , and ** indicates level of significance at 1 percent, and 5 percent, respectively.

Chain II: Wheat Wholesale-Flour Wholesale

In wholesale prices of flour equation, the estimates of $\rho^{IN}$and ρ^{OUT} regimes are significant at 5 percent. The signs and magnitudes of the estimates are consistent. We find the estimates $\rho^{IN} = -0.357$ and $\rho^{OUT} = -0.1845$ (Table 12). The model converges as the signs of both parameters are negative. The overall goodness of fit indicates that model is highly significant. The F-statistics 8.755 is rejected at 1 percent. We select optimal lag length using AIC. Results show that the null of symmetry between positive and negative variables in contemporary price changes is not rejected (Table 10). The F-statistics is found to be 0.307. Moreover, the results show that the null of symmetry between positive and negative variables in lagged price changes is not rejected. This means that no evidence exists of short-run asymmetry when we estimate the wholesale flour price. The results are subject to model diagnostics. The Ljung Box Q-statistics (1.659) indicates no serial correlation. ARCH test (0.052) supports no heteroscedasticity. JB test result indicate the residuals are normally distributed. The CUSUM and recursive coefficients estimates show that the model is stable (Last panel of Table 12). We conclude that symmetric threshold error correction model presents better fit to the data than asymmetric threshold error correction model.

Table 12: OLS estimates of ATECM: Wholesale flour price (ΔWPf_t)

Variables	Estimates	S E	t-stat	p-values
Constant	0.090	0.239	0.378	0.707
ε_{t-1}^{IN}	-0.357**	0.155	-2.298	0.024
ε_{t-1}^{OUT}	-0.184**	0.096	-1.919	0.059
ΔWPf_{t-1}^{+}	0.084	0.179	0.470	0.639
ΔWPf_{t-1}^{-}	0.495***	0.166	2.986	0.004
ΔWPf_{t-2}^{+}	-0.062	0.182	-0.339	0.736
ΔWPf_{t-2}^{-}	-0.159	0.166	-0.960	0.340
ΔWPw_{t}^{+}	0.537***	0.117	4.598	0.000
ΔWPw_{t}^{-}	0.650***	0.135	4.814	0.000
ΔWPw_{t-1}^{+}	0.087	0.157	0.556	0.580
ΔWPw_{t-1}^{-}	-0.251	0.167	-1.509	0.135
ΔWPw_{t-2}^{+}	0.001	0.152	0.004	0.997
ΔWPw_{t-2}^{-}	0.192	0.162	1.189	0.238
F-statistics	8.755***			
Short-run asymmetries (F-statistics)				
$H_o: \Delta WPw^{+}{}_{t} = \Delta WPw^{-}{}_{t}$	0.307		0.581	
$H_o: \Delta WPw^{+}{}_{t-1} = \Delta WPw^{-}{}_{t-1}$	1.891		0.173	
$H_o: \Delta WPw^{+}{}_{t-2} = \Delta Wpw^{-}{}_{t-2}$	0.623		0.433	
Model Diagnostics				
Ljung-Box Q-Stat for autocorrelation	0.154 (0.926)			
Jarque-Bera	0.069 (0.966)			
ARCH	0.702 (0.404)			
CUSUM test	Stable			
Recursive coefficient estimates	Inside the band of ±2 SE			

Note: *** , and ** indicates level of significance at 1 percent, and 5 percent, respectively.

For wholesale price of wheat equation, we find that only the estimates of ρ^{OUT} regime is significant at 5 percent. We find the estimates $\rho^{IN} = -0.027$ and $\rho^{OUT} = -0.227$ (Table 13). The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations from the long-run equilibrium relationship in previous periods tend to persist compared to negative price deviations. The model converges as the signs of both parameters are negative. The overall goodness of fit indicates that model is highly significant. The F-statistics 7.739 is rejected at 1 percent. We select optimal lag length using AIC. Results show that the null of symmetry between positive and negative variables in lagged price changes is rejected at 10 percent. The F-statistics is found to be 3.193. This means that the wholesale prices of wheat adjust differently when wholesale price of flour increases than it falls. The results are subject to model diagnostics. The Ljung Box Q-statistics (0.071) indicates no serial correlation. ARCH test (0.148) supports no heteroscedasticity. The CUSUM and recursive coefficients estimates show that the model is stable (Last panel of Table 13).

Table 13: OLS estimates of ATECM: Wholesale wheat price (ΔWPw_t)

Variables	Estimates	S.E	t-stat	p-values
Constant	0.096	0.304	0.315	0.754
ε^{IN}_{t-1}	0.027	0.096	0.285	0.776
ε^{OUT}_{t-1}	-0.227**	0.091	-2.487	0.015
ΔWPw^{+}_{t-1}	-0.327*	0.182	-1.799	0.076
ΔWPw^{-}_{t-1}	0.424**	0.190	2.232	0.029
ΔWPw^{+}_{t-2}	0.008	0.189	0.042	0.966
ΔWPw^{-}_{t-2}	-0.392**	0.198	-1.982	0.051
ΔWPf^{+}_{t}	0.761***	0.172	4.433	0.000
ΔWPf^{-}_{t}	0.701***	0.185	3.791	0.000
ΔWPf^{+}_{t-1}	0.313	0.217	1.442	0.153
ΔWPf^{-}_{t-1}	-0.280	0.206	-1.360	0.178
ΔWPf^{+}_{t-2}	-0.076	0.215	-0.354	0.724
ΔWPf^{-}_{t-2}	0.368	0.196	1.879	0.064
Regression F-statistics	7.739***			
Short-run asymmetries (F-statistics)				
$H_o: \Delta WPf^{+}{}_{t} = \Delta WPf^{-}{}_{t}$	0.041			0.841
$H_o: \Delta WPf^{+}{}_{t-1} = \Delta WPf^{-}{}_{t-1}$	3.193*			0.078
$H_o: \Delta WPf^{+}{}_{t-2} = \Delta WPf^{-}{}_{t-2}$	1.806			0.183
Model Diagnostics				
Ljung-Box Q-Stat for autocorrelation	0.071 (0.965)			
ARCH	0.148 (0.701)			
CUSUM test	Stable			
Recursive coefficient estimates	Inside the band of ±2 SE			

Notes: *** , ** and * indicates level of significance at 1 percent, 5 percent, and 10 percent, respectively.

Similar to chain I, we estimate SUR model for chain II to examine the results robustness. In wholesale price of wheat equation, we find that only the estimate of ρ^{OUT} regime is significant at 5 percent level. We find the estimates $\rho^{IN} = -0.008$ and $\rho^{OUT} = -0.141$ (Table 14). The model converges as the signs of both parameters are consistent and negative. The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist compared to negative price deviations from the long-run equilibrium relationship. We select optimal lag length using AIC. Results show that the null of symmetry between positive and negative variables in lagged price changes is rejected at the 5 percent level (Table 14). This means that there is an evidence of short-run asymmetry. In wholesale prices of flour equation, we find that only the estimates of ρ^{OUT} regime is significant at 5 percent level. We find the estimates $\rho^{IN} = -0.203$ and $\rho^{OUT} = -0.124$ (Table 14). The model converges as the signs of both parameters are consistent and negative. The overall goodness of fit indicates that model is highly significant. The F-statistics 7.739 is rejected at 1 percent level. The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist compared to negative price deviations from the long-run equilibrium relationship. We select optimal lag length using AIC. Results show that the null of symmetry between positive and negative variables in lagged price changes is rejected at the 5 percent level. The F-statistics is found to be 4.927. This means that the wholesale prices of flour adjust differently when wholesale prices of wheat increases than when it falls.

Table 14: Seemingly Unrelated Regression (SUR) estimates of ATECM

	Wholesale price of wheat (ΔWPw)				Wholesale price of flour (ΔWPf)			
Variables	Coeff	S.E	T-stat	p-values	Coeff	S.E	T-stat	p-values
Constant	0.013	0.264	0.048	0.962	0.002	0.212	0.009	0.993
ε_{t-1}^{IN}	-0.008	0.072	-0.110	0.912	-0.203*	0.117	-1.735	0.083
ε_{t-1}^{OUT}	-0.141**	0.070	-2.007	0.045	-0.124*	0.075	-1.656	0.098
ΔWPw_t^+				-	0.761***	0.091	8.324	0.000
ΔWPw_t^-				-	0.798***	0.107	7.494	0.000
ΔWPw_{t-1}^+	-0.316**	0.162	-1.949	0.051	0.191	0.139	1.370	0.171
ΔWPw_{t-1}^-	0.379**	0.175	2.159	0.031	-0.306**	0.152	-2.010	0.044
ΔWPw_{t-2}^+	-0.074	0.169	-0.434	0.664	0.081	0.138	0.585	0.559
ΔWPw_{t-2}^-	-0.405**	0.178	-2.266	0.023	0.247*	0.148	1.675	0.094
ΔWPf_t^+	1.063***	0.134	7.906	0.000				
ΔWPf_t^-	0.994***	0.146	6.803	0.000				
ΔWPf_{t-1}^+	0.243	0.194	1.251	0.211	-0.079	0.161	-0.487	0.627
ΔWPf_{t-1}^-	-0.425**	0.188	-2.260	0.024	0.463***	0.153	3.022	0.003
ΔWPf_{t-2}^+	0.057	0.194	0.292	0.770	-0.089	0.164	-0.541	0.589
ΔWPf_{t-2}^-	0.365**	0.181	2.022	0.043	-0.221	0.153	-1.448	0.148

Short-run asymmetries (χ^2 -statistics)

$H_o: \Delta WPf^+{}_t = \Delta WPf^-{}_t$	0.090	0.764	$H_o: \Delta WPw^+{}_t = \Delta WPw^-{}_t$	0.057	0.812
$H_o: \Delta WPf^+{}_{t-1} = \Delta WPf^-{}_{t-1}$	4.972**	0.026	$H_o: \Delta WPw^+{}_{t-1} = \Delta WPw^-{}_{t-1}$	4.927**	0.026
$H_o: \Delta WPf^+{}_{t-1} = \Delta WPf^-{}_{t-1}$	1.050	0.306	$H_o: \Delta WPw^+{}_{t-1} = \Delta WPw^-{}_{t-1}$	0.552	0.458

Note: *** , ** and * indicates level of significance at 1 percent, 5 percent, and 10 percent, respectively.

Chain III: Retail Prices of flour-Wholesale Prices of Wheat

For the retail price of flour equation, we find that estimates of $\rho^{IN}$and ρ^{OUT} regimes are significant at 1 percent level. The signs and magnitudes of the estimates are consistent. We find the estimates $\rho^{IN} = -0.380$ and $\rho^{OUT} = -0.908$ (Table 15). The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist longer compared to negative price deviations from the long-run equilibrium relationship. The model converges as the signs of both parameters are negative. The overall goodness of fit indicates that model is highly significant. The F-statistics 10.704 is rejected at 1 percent. We select optimal lag length using AIC. Results show that the null of symmetry between positive and negative variables in lagged price changes is rejected at 10 percent. The F-statistics is found to be 3.123. Moreover, the null of symmetry between positive and negative variables in lagged (3) price changes is rejected at the 10 percent level. The F-statistics is found to be 3.299. All these mean that the retail prices of flour adjust differently when wholesale price of wheat increases than when price falls. The results are subject to model diagnostics. The Ljung Box Q-statistics (0.542) indicates no serial correlation. ARCH test (0.072) supports no heteroscedasticity. JB test result indicates that the residuals are normally distributed. The CUSUM and recursive coefficients estimates show that the model is stable (Last panel of Table 15).

Table 15: OLS estimates of ATECM: Retail flour price equation (ΔRPf_t)

Variables	Estimates	S.E	t-stat	p-values
Constant	-0.081	0.248	-0.328	0.744
ε^{IN}_{t-1}	-0.380***	0.102	-3.740	0.000
ε^{OUT}_{t-1}	-0.908***	0.263	-3.446	0.001
ΔFLR^{+}_{t-1}	0.356*	0.182	1.962	0.054
ΔRPf^{-}_{t-1}	0.313**	0.148	2.118	0.038
ΔRPf^{+}_{t-2}	0.056	0.174	0.323	0.748
ΔRPf^{-}_{t-2}	-0.035	0.146	-0.242	0.809
ΔRPf^{+}_{t-3}	0.311*	0.162	1.921	0.059
ΔRPf^{-}_{t-3}	-0.208	0.133	-1.569	0.121
ΔWPw^{+}_{t}	0.445***	0.103	4.334	0.000
ΔWPw^{-}_{t}	0.714***	0.124	5.752	0.000
ΔWPw^{+}_{t-1}	0.042	0.138	0.305	0.761
ΔWPw^{-}_{t-1}	-0.328**	0.165	-1.994	0.050
ΔWPw^{+}_{t-2}	-0.231*	0.135	-1.706	0.092
ΔWPw^{-}_{t-2}	0.079	0.155	0.513	0.609
ΔWPw^{+}_{t-3}	-0.168	0.132	-1.280	0.205
ΔWPw^{-}_{t-3}	0.197	0.143	1.379	0.172
Model F-statistics	10.704***			
Short-run asymmetries (F-statistics)				
$H_o: \Delta WPw^{+}{}_{t} = \Delta WPw^{-}{}_{t}$	2.006	0.161		
$H_o: \Delta WPw^{+}{}_{t-1} = \Delta WPw^{-}{}_{t-1}$	3.123*	0.081		
$H_o: \Delta WPw^{+}{}_{t-2} = \Delta WPw^{-}{}_{t-2}$	2.046	0.157		
$H_o: \Delta WPw^{+}{}_{t-3} = \Delta WPw^{-}{}_{t-3}$	3.299*	0.073		
Model Diagnostics				
Ljung-Box Q-Stat for autocorrelation	0.542 (0.762)			
Jarque-Bera	0.644 (0.724)			
ARCH	0.072 (0.789)			
CUSUM test	Stable			
Recursive coefficient estimates	Inside the band of ±2 SE			

Note: *** , ** and * indicates level of significance at 1 percent, 5 percent, and 10 percent, respectively.

In wholesale wheat price equation, we find only the estimate of ρ^{OUT} regime is significant at 1 percent level. The signs and magnitudes of the estimates are consistent. We find the estimates $\rho^{IN} = -0.333$

and $\rho^{OUT} = -0.802$ (Table 16). The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist compared to negative price deviations from the long-run equilibrium relationship. The model converges as the signs of both parameters are negative. The overall goodness of fit indicates that model is highly significant. The F-statistics 8.032 is rejected at 1 percent. We select optimal lag length using AIC. Results show that the null of symmetry between positive and negative variables in lagged price changes is not rejected. The Ljung Box Q-statistics (0.718) indicates no serial correlation. ARCH test (0.372) supports no heteroscedasticity. JB test result indicate the residuals are normally distributed. The CUSUM and recursive coefficient estimates show that the model is stable (Last panel of Table 16). We can conclude that symmetric threshold error correction model presents better to the data than the asymmetric threshold error correction model.

Table 16: OLS estimates of ATECM: Wholesale wheat price (ΔWPw_t)

Variable	Estimates	SE	t-stat	p-values
Constant	0.391	0.295	1.327	0.189
ε_{t-1}^{IN}	-0.333	0.232	-1.431	0.157
ε_{t-1}^{OUT}	-0.802***	0.161	-4.977	0.000
ΔWPw_{t-1}^{+}	0.020	0.173	0.113	0.910
ΔWPw_{t-1}^{-}	0.585**	0.223	2.623	0.011
ΔWPw_{t-2}^{+}	0.125	0.175	0.712	0.479
ΔWPw_{t-2}^{-}	0.288	0.208	1.389	0.169
ΔWPw_{t-3}^{+}	0.333**	0.169	1.975	0.050
ΔWPw_{t-3}^{-}	0.035	0.192	0.181	0.857
ΔRPf_{t}^{+}	0.890***	0.177	5.015	0.000
ΔRPf_{t}^{-}	0.852***	0.150	5.693	0.000
ΔRPf_{t-1}^{+}	-0.226	0.223	-1.012	0.315
ΔRPf_{t-1}^{-}	-0.301	0.200	-1.502	0.137
ΔRPf_{t-2}^{+}	-0.114	0.219	-0.518	0.606
ΔRPf_{t-2}^{-}	-0.067	0.192	-0.350	0.727
ΔRPf_{t-3}^{+}	-0.280	0.200	-1.401	0.166
ΔRPf_{t-3}^{-}	-0.006	0.176	-0.034	0.973
Regression F(12,77)	8.032*** (0.000)			
Short-run asymmetries (F-statistics)				
$H_o: \Delta RPf^{+}{}_{t} = \Delta RPf^{-}{}_{t}$	0.021			0.885
$H_o: \Delta RPf^{+}{}_{t-1} = \Delta RPf^{-}{}_{t-1}$	0.062			0.804
$H_o: \Delta RPf^{+}{}_{t-2} = \Delta RPf^{-}{}_{t-2}$	0.025			0.875
$H_o: \Delta RPf^{+}{}_{t-3} = \Delta RPf^{-}{}_{t-3}$	0.952			0.333
Model Diagnostics				
Ljung-Box Q-Stat for autocorrelation	0.718 (0.869)			
ARCH	0.372 (0.543)			
CUSUM test	Stable			
Recursive coefficient estimates	Inside the band of ±2 SE			

Notes: *** , ** and * indicates level of significance at 1 percent, 5 percent, and 10 percent, respectively.

Similar to chain I and II, we estimate SUR model for chain III to examine the results' robustness. For wholesale price of wheat equation, we find that the estimates $\rho^{IN}$and ρ^{OUT} regimes are significant at 1 percent and 5 percent, respectively. The signs and magnitudes of the estimates are consistent. We find the estimates $\rho^{IN} = -0.381$ and $\rho^{OUT} = -0.584$ (Table 17). The model converges as the signs of both parameters are negative. The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist compared to negative price deviations from the long-run equilibrium relationship. We select optimal lag length using AIC. Results show that the null of symmetry between positive and negative variables for neither the contemporaneous nor the lagged price changes are rejected. For retail prices of flour equation, we find that the estimates $\rho^{IN}$and ρ^{OUT} regimes are significant at 1 percent. The signs and magnitudes of the estimates are consistent. We find the estimates $\rho^{IN} = -0.335$ and $\rho^{OUT} = -0.568$ (Table 17). The model converges as the signs of both parameters are negative. The speed of adjustment to negative price deviations (ρ_2) is higher than the speed of adjustment to positive price deviations (ρ_1) in absolute terms. This implies that positive price deviations in previous periods tend to persist compared to negative price deviations from the long-run equilibrium relationship. We select optimal lag length using AIC. Results show that the null of symmetry between positive and negative variables in lagged price changes is rejected at 5 percent. The F-statistics is 5.622. Also the null of symmetry between positive and negative variables in lagged (3) price changes is rejected at the 5 percent level. The F-statistics is found to be 4.422. All these mean that retail prices of flour adjust differently when wholesale price of wheat increases than when price falls.

Table 17: Seemingly Unrelated Regression (SUR) estimates of ATECM

	Wholesale price of wheat (ΔWPw)				Retail price of flour (ΔRPf)			
Variable	Coeff	SE	T-stat	p-values	Coeff	SE	T-stat	p-values
Constant	0.132	0.255	0.519	0.604	-0.150	0.211	-0.711	0.477
ε_{t-1}^{IN}	-0.381**	0.176	-2.169	0.030	-0.335***	0.082	-4.090	0.000
ε_{t-1}^{OUT}	-0.584***	0.127	-4.595	0.000	-0.568***	0.195	-2.912	0.004
ΔWPw_t^+					0.628***	0.078	8.107	0.000
ΔWPw_t^-					0.782***	0.095	8.194	0.000
ΔWPw_{t-1}^+	-0.101	0.150	-0.673	0.501	0.101	0.119	0.850	0.395
ΔWPw_{t-1}^-	0.524***	0.195	2.688	0.007	-0.343**	0.146	-2.357	0.018
ΔWPw_{t-2}^+	0.102	0.154	0.663	0.507	-0.123	0.118	-1.035	0.301
ΔWPw_{t-2}^-	0.114	0.182	0.626	0.531	0.060	0.136	0.443	0.658
ΔWPw_{t-3}^+	0.285**	0.150	1.906	0.050	-0.178	0.116	-1.534	0.125
ΔWPw_{t-3}^-	-0.144	0.170	-0.846	0.397	0.201	0.128	1.577	0.115
ΔRPf_t^+	1.243***	0.135	9.228	0.000				
ΔRPf_t^-	1.024***	0.119	8.579	0.000				
ΔRPf_{t-1}^+	-0.190	0.197	-0.966	0.334	0.249	0.159	1.572	0.116
ΔRPf_{t-1}^-	-0.362**	0.176	-2.056	0.040	0.287**	0.131	2.196	0.028
ΔRPf_{t-2}^+	0.022	0.194	0.113	0.910	0.017	0.153	0.113	0.910
ΔRPf_{t-2}^-	-0.051	0.172	-0.299	0.765	-0.038	0.131	-0.293	0.769
ΔRPf_{t-3}^+	-0.280	0.179	-1.567	0.117	0.301**	0.144	2.095	0.036
ΔRPf_{t-3}^-	0.091	0.157	0.579	0.563	-0.160	0.119	-1.342	0.180
Short-run asymmetries (χ^2 -statistics)								
$H_o: \Delta RPf^+{}_t = \Delta RPf^-{}_t$			1.257	0.262	$H_o: \Delta WPw^+{}_t = \Delta WPw^-{}_t$		1.206	0.272

$H_o: \Delta RPf^{+}_{t-1} = \Delta RPf^{-}_{t-1}$	0.417	0.519	$H_o: \Delta WPw^{+}_{t-1} = \Delta WPw^{-}_{t-1}$	5.622**	0.018
$H_o: \Delta RPf^{+}_{t-2} = \Delta RPf^{-}_{t-2}$	0.079	0.779	$H_o: \Delta WPw^{+}_{t-2} = \Delta WPw^{-}_{t-2}$	0.911	0.340
$H_o: \Delta RPf^{+}_{t-3} = \Delta RPf^{-}_{t-3}$	2.180	0.140	$H_o: \Delta WPw^{+}_{t-3} = \Delta WPw^{-}_{t-3}$	4.422**	0.035

Note: *** , ** and * indicates level of significance at 1 percent, 5 percent, and 10 percent, respectively.

Model selection and asymmetries in short- and long-run price transmissions

We present the short- and long-run asymmetries in the estimated equations for different vertical chains in Table 18. We find that out of estimated six different models, five models identify the asymmetry in the long-run. Thus the speeds of adjustment to the positive deviations and the negative deviations are statistically different. We also find the mixed evidence on model selection. Out of six models, we find that consistent MTAR model fits better to the data for three price equations namely the wholesale prices of flour and retail prices of flour in chain I, and wholesale price of wheat in chain II. On the contrary, we find that consistent TAR model fits better to the data for remaining three equations namely the wholesale prices of wheat in chain II, and retail prices of flour and wholesale prices of wheat in chain III.

Regarding short run asymmetries in price responses in chain I, we find the evidence of short run asymmetries both in OLS and SUR models. We find evidence of asymmetries between positive and negative variables in contemporary and lagged price changes irrespective of normalization of prices in OLS. Also, the SUR model provides evidence of asymmetry between positive and negative variables in contemporary price changes again irrespective of normalization of prices. The results provide evidence of asymmetry between positive and negative variables in lagged price changes only for the retail price of flour equation. This means that since we find the evidence of asymmetries in different models, we can confirm that the short- and long-run asymmetries exist in wholesale and retail flour markets. In chain II, we do not find any evidence of short run asymmetry when we estimate wholesale price of flour using OLS. On the contrary, we find evidence of short run asymmetries in wholesale prices of flour. We find evidences of asymmetries between positive and negative variables in lagged price changes using OLS. The SUR model provides evidence of asymmetry between positive and negative variables in lagged price changes in both prices- wholesale prices of wheat and wholesale prices of flour. So, for chain II, the SUR model fits the data better than the estimated OLS. For chain III, we find the evidence of asymmetry only in the estimated retail price of flour using OLS. The results provide evidence of asymmetry between positive and negative variables in lagged price changes. We do not find any evidence of asymmetry in the wholesale prices of wheat. We find evidence of asymmetry only in retail price of flour equation using SUR. The results provide evidence of asymmetry between positive and negative variables in lagged price changes (lagged 1 and 3). So, we conclude that SUR model fits the data better than the OLS in the case of chain III. However, we conclude that there are short- and long-run asymmetries in the vertical wheat-to-flour markets.

Table 18: Tests of long-run and short-run asymmetries in wheat-flour chains

Models/chains	Equations	Hypotheses	F-statistics	p-values
Long-run asymmetries				
Chain I: WP(f)-RP(f)				
Consistent-MTAR	RP(f)	$H_0: \rho_1 = \rho_2$	3.276*	0.071
Consistent-MTAR	WP(f)	$H_0: \rho_1 = \rho_2$	3.076*	0.082
Chain II: WP(w)-WP(f)				
Consistent-TAR	WP(f)	$H_0: \rho_1 = \rho_2$	3.742**	0.050
Consistent-MTAR	WP(w)	$H_0: \rho_1 = \rho_2$	1.403	0.239
Chain III: WP(w)-RP(f)				
Consistent-TAR	RP(f)	$H_0: \rho_1 = \rho_2$	10.747***	0.001
Consistent-TAR	WH(w)	$H_0: \rho_1 = \rho_2$	6.733***	0.001
Short-run asymmetries				
Chain I: WP(f)-RP(f)				
ATECM (2): OLS	ΔRPf	$H_o: \Delta WPf^{+}_{t} = \Delta WPf_{t}$	6.330**	0.013
		$H_o: \Delta WPf^{+}_{t-1} = \Delta WPf^{-}_{t-1}$	5.647**	0.020
		$H_o: \Delta WPf^{+}_{t-2} = \Delta WPf^{-}_{t-2}$	1.107	0.295
		$H_o: \Delta WPf^{+}_{t} + \Delta WPf^{+}_{t-2} = \Delta WPf^{-}_{t} + \Delta WPf^{-}_{t-2}$	5.427**	0.022
ATECM (2): OLS	ΔWPf	$H_o: \Delta RPf^{+}_{t} = \Delta RPf^{-}_{t}$	21.734***	0.000
		$H_o: \Delta RPf^{+}_{t-1} = \Delta RPf^{-}_{t-1}$	3.090*	0.080
		$H_o: \Delta RPf^{+}_{t-2} = \Delta RPf^{-}_{t-2}$	0.843	0.361
		$H_o: \Delta RPf_{t} + \Delta RPf^{+}_{t-2} = \Delta RPf^{-}_{t} + \Delta RPf_{t-2}$	8.806***	0.004
ATECM (2): SUR	ΔRPf	$H_o: \Delta WPf^{+}_{t} = \Delta WPf^{-}_{t}$	6.195**	0.012
		$H_o: \Delta WPf^{+}_{t-1} = \Delta WPf^{-}_{t-1}$	4.994**	0.025
		$H_o: \Delta WPf^{+}_{t-2} = \Delta WPf^{-}_{t-2}$	1.391	0.238
	ΔWPf	$H_o: \Delta RPf^{+}_{t} = \Delta RPf^{-}_{t}$	12.235***	0.000
		$H_o: \Delta RPf^{+}_{t-1} = \Delta RPf^{-}_{t-1}$	0.448	0.503
		$H_o: \Delta RPf^{+}_{t-2} = \Delta RPf^{-}_{t-2}$	0.135	0.713
Chain II: WP(w)-WP(f)				
ATECM (2): OLS	ΔWPf	$H_o: \Delta WPw^{+}_{t} = \Delta WPw_{t}$	0.307	0.581
		$H_o: \Delta WPw^{+}_{t-1} = \Delta WPw^{-}_{t-1}$	1.891	0.173
		$H_o: \Delta WPw^{+}_{t-2} = \Delta WPw_{t-2}$	0.623	0.433
ATECM (2): OLS	ΔWPw	$H_o: \Delta WPf^{+}_{t} = \Delta WPf^{-}_{t}$	0.041	0.841
		$H_o: \Delta WPf^{+}_{t-1} = \Delta WPf^{-}_{t-1}$	3.193*	0.078
		$H_o: \Delta WPf^{+}_{t-2} = \Delta WPf^{-}_{t-2}$	1.806	0.183
ATECM (2): SUR	ΔWPw	$H_o: \Delta WPf^{+}_{t} = \Delta WPf^{-}_{t}$	0.090	0.764
		$H_o: \Delta WPf^{+}_{t-1} = \Delta WPf^{-}_{t-1}$	4.972**	0.026
		$H_o: \Delta WPf^{+}_{t-2} = \Delta WPf^{-}_{t-2}$	1.050	0.306
	ΔWPf	$H_o: \Delta WPw^{+}_{t} = \Delta WPw^{-}_{t}$	0.057	0.812
		$H_o: \Delta WPw^{+}_{t-1} = \Delta WPw^{-}_{t-1}$	4.927**	0.026
		$H_o: \Delta WPw^{+}_{t-2} = \Delta WPw^{-}_{t-2}$	0.552	0.458
Chain III: WP(w)-RP(f)				
ATECM (3):OLS	ΔRPf	$H_o: \Delta WPw^{+}_{t} = \Delta WPw^{-}_{t}$	2.006	0.161
		$H_o: \Delta WPw^{+}_{t-1} = \Delta WPw^{-}_{t-1}$	3.123*	0.081
		$H_o: \Delta WPw^{+}_{t-2} = \Delta WPw^{-}_{t-2}$	2.046	0.157
		$H_o: \Delta WPw^{+}_{t-3} = \Delta WPw^{-}_{t-3}$	3.299*	0.073
ATECM (3):OLS	ΔWPw	$H_o: \Delta RPf^{+}_{t} = \Delta RPf^{-}_{t}$	0.021	0.885
		$H_o: \Delta RPf^{+}_{t-1} = \Delta RPf^{-}_{t-1}$	0.062	0.804
		$H_o: \Delta RPf^{+}_{t-2} = \Delta RPf^{-}_{t-2}$	0.025	0.875

		$H_o: \Delta RPf^{+}_{t-3} = \Delta RPf^{-}_{t-3}$	0.952	0.333
ATECM (3): SUR	ΔWPw	$H_o: \Delta RPf^{+}_{t} = \Delta RPf^{-}_{t}$	1.257	0.262
		$H_o: \Delta RPf^{+}_{t-1} = \Delta RPf^{-}_{t-1}$	0.417	0.519
		$H_o: \Delta RPf^{+}_{t-2} = \Delta RPf^{-}_{t-2}$	0.079	0.779
		$H_o: \Delta RPf^{+}_{t-3} = \Delta RPf^{-}_{t-3}$	2.180	0.140
	ΔRPf	$H_o: \Delta WPw^{+}_{t} = \Delta WPw^{-}_{t}$	1.206	0.272
		$H_o: \Delta WPw^{+}_{t-1} = \Delta WPw^{-}_{t-1}$	5.622**	0.018
		$H_o: \Delta WPw^{+}_{t-2} = \Delta WPw^{-}_{t-2}$	0.911	0.340
		$H_o: \Delta WPw^{+}_{t-3} = \Delta WPw^{-}_{t-3}$	4.422**	0.035

Notes: ATECM means asymmetric threshold error correction model; parentheses indicate the number of selected lags using AIC and SBC; SUR means seemingly unrelated regression model; consistent-TAR and consistent-MTAR means the threshold autoregressive and momentum auto-regressive models with estimated threshold (τ); ***, ** and * means significant at 1%, 5% and 10%, respectively.

7. Discussions and Conclusions

Wheat-to-flour markets in Bangladesh are characterized by private and public presence. The domestic and border policies are liberalized allowing the private traders to involve in wheat and flour markets. The government of Bangladesh procures wheat from farmers by paying support price which is set higher than the market price. However, the quantity of procurement is very insignificant. Public importation and food aid constitute a small percent of total supply. The marketing chains are long, numerous actors are in the marketing chains which make the markets to be some extent, atomistic. Due to liberal market policies, the markets are mainly driven by the private traders, from wheat production and importation till end uses such as the consumers and industrial users. In every stages, numerous private traders are involved. But there is also serious concerns and fears that food markets are not that much efficient that one expected due to favorable policies, development of information and communication system and the less presence of the government intervention in the food markets. However, a number of reasons are identified in the literature to explain the asymmetric price transmission in the spatial and vertical food markets. Potential explanations for the asymmetric price transmission are retail level market power (Griffith & Piggot 1994; Bettendorf and Verboven 2000, Moorthy 2005), adjustment costs (Buckle and Carlson 2000; Chavas and Mehta 2002), search costs (Gomez et al., 2013; Richards et al. 2014) inventory holding management costs (Reagan and Witzman 1982; Cui et al. 2008), asymmetric information (Baily and Brorsen 1989; Kumar et al. 2001; Busse el al 2006), market structure (Lee and Miguel 2013), public intervention (Kinnucan and Forker 1987). Many studies argue that asymmetry is related to inventory management and adjustment problem at the downstream, for instance at the retail level. Prices at the downstream may not responds due to menu costs - cost occurring with the re-pricing and pricing strategy. Many of the retailers might be unwilling to re-pricing due to price instability. Also, in many cases, the retailers are uncertain if price increases is permanent or transitory. Peltzman (2000) finds no evidence of relationship between menu cost and price asymmetries but does report evidence of greater asymmetries in more fragmented value chain where one can expect menu costs to be higher. This might be more relevant to the developing countries including Bangladesh where the marketing chains are long and fragmented with many chain actors being are involved. However, Peltzman (2000) also mentions that asymmetry at the downstream is not an exception rather it's a rule and absence of any economic theory explaining the asymmetry in price transmission is a gap that needs to be filled up. Baily and Brorsen (1989) also pointed that the asymmetries at the downstream may be due to asymmetries in the costs of adjustments. The presence of asymmetry in price transmission in vertical markets often considered evidence of market power (Cramon-Taubadel and Meyer, 2004). Although it is widely discussed in the literature that market power is the main cause for imperfect price transmission especially at the retail level but in many cases it may not be true at all especially in developing countries where numerous actors are present in these markets. In Bangladesh, domestic flour markets are mainly driven by wholesalers. The flour wholesaler purchases flour from the millers. In some cases wheat wholesaler uses the milling services before transferring milled wheat to the flour wholesalers. The millers are not many in numbers. Although it is unlikely that millers could manipulate market prices through collusion since in many cases millers provides milling service only and charge milling fees. However, the scope of this paper does not allow us to test if millers are responsible to this price asymmetry. The retailers purchases milled wheat and selling to ultimate consumers. The large scale manufacturer, for example confectioneries, bakeries, cookies industry purchases wheat milled from the wholesaler in large volume. Individual consumers or small scale bakeries purchase milled wheat from the retailers. Since we find that downstream prices are more sticky than the upstream, that means retailer response to positive and negative prices changes are different. We do not claim that this kind of sticky price adjustment is due to retailer`s market power. The retailers are small traders linking wholesalers and consumers. The retailers own premises in the consuming markets and sell flour usually ranging from 1 kilogram to 2 kilograms maximum (Farid and Rahman 2002). In recent years since number of volume of trade by superstores increases in the capital city, the departmental stores pack flour in one/two kilogram packets with different brand name and sell these to consumers. But this is so far a very small proportion of total flour being sold in Dhaka. This form of flour retailing by numerous retailers indicates that retailers are unlikely to exercise market power and also that it is unlikely that non-competitive market structures are responsible for this asymmetric adjustment.

However, price transmission asymmetries can provide valuable information for decision makers about supply chain behavior. We develop threshold and asymmetry error correction models to statistically test threshold and for short- and long-run PTAs in wheat-flour vertical markets in Dhaka, Bangladesh- a city of more than 18 million population. The paper statistically tests the threshold effects in the price adjustment and examines short- and long-run asymmetry in the price adjustment in three distinct vertical chains. The paper focuses on the impact of changes in prices at downstream on the prices in upstream of wheat-flour supply chain. We find evidence of threshold effects in price adjustment and presence of short- and long-run asymmetries in price transmissions. So, we can conclude that there are inefficiencies prevailing in the wheat and flour markets. Since downstream prices, here retail prices are stickier than the wholesale prices, consumers are worse-off particularly - when price decreases at the upstream does not necessarily passes to the consumers via retailers. This means that consumers may not benefit as was expected from the liberalization of agricultural and trade policies if retailers do not pass price decreases to consumers. However, since we do not formally investigate the reasons underlying the asymmetric price adjustment further empirical work needs to be done to explain the reasons for this.

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