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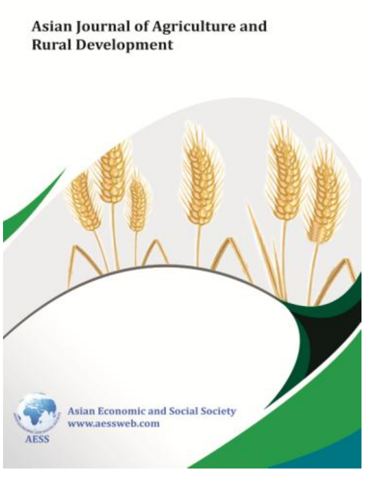

Vertical Price Transmission in Local Rice Markets in Côte d'Ivoire: Are Consumers Really Right?

Yaya KEHO (National School of Statistics and Applied Economics (ENSEA), 08 PO BOX 03 Abidjan 08, Côte d'Ivoire)

Aïssata SOBIA CAMARA (Ivorian Centre for Economic and Social Research (CIRES), University of Abidjan, 08 PO BOX 1295 Abidjan 08, Côte d'Ivoire)

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Author(s)

Yaya KEHO
National School of Statistics and Applied Economics (ENSEA), 08 PO BOX 03 Abidjan 08, Côte d'Ivoire

Aïssata SOBIA CAMARA
Ivorian Centre for Economic and Social Research (CIRES), University of Abidjan, 08 PO BOX 1295 Abidjan 08, Côte d'Ivoire

Abstract

This paper analyses vertical relationships between wholesale and retail prices in three local rice markets in Côte d'Ivoire. The aim of the paper is to ascertain whether the popular complaint of consumers about the asymmetric price transmission holds true. Our empirical analysis makes use of threshold cointegration and error correction models and monthly data for the period 1990-1999. We found that wholesale and retail prices are cointegrated and increases in wholesale prices are passed on to retail prices more quickly than decreases.

Keywords: Asymmetric price transmission; Threshold cointegration; Rice; Côte d'Ivoire

Introduction

Prices play an important role in explaining economic behaviours and well-being. They are central in reducing poverty. With different food products, marketing channels can geographically be long or may consist of many vertically integrated levels involving several agents such as producers, wholesalers and retailers. Many studies conducted by agricultural economists have analysed price transmission at various stage in the production and distribution chain. One strand of these studies investigated spatial price linkages focusing on two issues. Are markets integrated? Are prices dominated by a central market? Another strand of studies considered the vertical price transmission at different stages in the distribution chain and analysed the farm-wholesale and wholesale-retail price relationships. The relationship between wholesale and retail food prices has important economic and political implications. It is commonly felt that price transmission in the marketing chain is not always symmetric. Asymmetric price transmission refers to the phenomenon that downstream prices respond to upstream price increases and decreases at different rates. Usually, consumers complain that retailers quickly increase food prices as wholesale prices rise, but do not reduce retail prices as quickly when prices drop. The extent of increase in retail prices is sometime deemed excessive. This view is strongly supported by the extensive study of Peltzman (2000). A good example of asymmetry is where retail prices remain sticky although farm or wholesale prices have fallen due to increase in primary production. Asymmetry is often perceived by some economists and politicians as the outcome of non-competitive food markets or market power of traders requiring government intervention. In an asymmetric price transmission context, consumers do not benefit from reduction in input and wholesale prices as a consequence of agricultural policy reform.

Recently, a growing body of studies has emerged aiming at testing the asymmetry in upstream and downstream prices relationship for several products and assessing the market power exercised by wholesaler or retailers. Results from this empirical literature are inconclusive. Studies generally differ in terms of the goods analysed, countries and model specification. Consequently, it is difficult to draw conclusions on which to base policy decisions. In this study, we examine the wholesale-retail local rice price relationships for Côte d'Ivoire in order to determine how wholesale price changes are passed on to retail prices. More precisely, we look at whether we can find some empirical evidence supporting the consumers' complaint that retail prices rise faster than they fall. The empirical investigation makes use of the threshold cointegration approach proposed by Enders and Granger (1998) and Enders and Siklos (2001). The appealing aspect of this approach is that it allows one to test for cointegration taking into account the asymmetric nature of the adjustment mechanism, which varies according to the size and sign of the equilibrium error. Even though this econometric approach is widely used in the empirical literature, its application for analyzing vertical price transmission has been very limited. The present study is the first attempt of its kind for Cote d'Ivoire. We consider three regional markets located at the South, Centre and North. To ensure comparability of the findings, we use monthly data and similar econometric model specifications for all markets.

The rest of the article is organized as follows. Section 2 reviews the related economic literature. In Section 3, the main characteristics of the Ivorian rice market are described along with some preliminary analysis of non-linearities. Section 4 describes the econometric methodology used in the study. Section 5 reports and discusses the empirical results. Finally, Section 6 provides summary of the main findings and concluding remarks.

Literature Review

In a competitive market, price changes between wholesale and retail levels are expected to be symmetric. Nonetheless, many studies report that the wholesale-retail price transmission process might be asymmetric. The economic literature identifies several plausible reasons for asymmetric price transmission[1]. Some of the more prominent explanations include imperfect competition (Boyd and Brorsen, 1988; Bettendorf and Verboven 2000; Meyer and von Cramon-Taubadel 2004), adjustment costs at the retail level (Ball and Mankiw 1994), input substitution at the processing level (Bettendorf and Verboven 2000), stocks at both the production and retail level (Reagan and Weitzman 1982), search costs (Bénabou and Gertner 1993), and public intervention in terms of price regulation or subventions (Kinnucan and Forker 1987). In many African countries, inadequate transport and communication infrastructures can hinder the transmission of price signals and prevent arbitrages (Rapsomanikis et al. 2003). Wholesale prices may fall but retail prices remain unchanged because transportation costs are high. The role of market power associated with the organization of wholesalers has been put forward to explain the asymmetric relationship between wholesale and retail prices (Griffith and Piggott 1994; Meyer and von Cramon-Taubadel 2004; Gohin and Guyomard 2000; Vindel 2005). This is particularly true for areas where asymmetric information prevails and search costs associated with that asymmetry are high. In general, consumers have not full information about the prices of products at the relevant markets. Consequently, sellers take advantage of this asymmetry and retain their old retail price when the wholesale price decreases. Gohin and Guyomard (2000) showed that over 20% and 17% of the margins between wholesale and retail prices of dairy products and meat in France can be attributed to oligopolistic behaviour of wholesalers. In a study on price transmission in Benin, Kuiper et al. (2003) highlighted the influence of wholesalers in the formation of maize prices in local markets. Fiamohé and De Frahan (2008) found that the main cause of asymmetric price transmission in the food markets in Benin is the presence of associations of traders who behave like oligopolies or cartels. The key point to note from this theoretical literature is that there are several potential causes of asymmetric price transmission.

Economists have conducted hypothesis tests for asymmetric price transmission in several agricultural product markets using different methods and data, and they have generally reached different conclusions depending on the goods analysed, countries, time frequencies and model specification (Frey and Manera 2007). Peltzman (2000) conducted a wide-ranging vertical price transmission analysis utilizing a

[1] See Meyer and Von Cramon-Taubadel (2004) and Frey and Manera (2007) for a comprehensive review of the literature dealing with potential causes of asymmetries in vertical price transmission.

large data set consisting of prices from over two hundred product markets. His empirical evidence indicates that asymmetric adjustment is prevalent with retail prices rising faster, as compared to falling, whilst this asymmetry is not related to inventory costs, menu costs and imperfect competition. Such findings not only suggest that asymmetry is the rule rather than the exception in market price adjustment but raise a number of questions related to the suitability of empirical price-based tests and the conventional theory of prices. Kinnucan and Forker (1987) examined farm–retail price transmission in the dairy sector and found significant evidence of asymmetric pricing. Ward (1982) showed that wholesale fresh vegetable price decreases are reflected at the retail level more quickly than wholesale price increases, and wholesale price decreases are more fully transmitted to the shipping point farm level relative to wholesale price increases. More recently, Ben-Kaabia and Gil (2007) investigated the non-linear adjustment between farm and retail prices in the lamb sector in Spain, using a three-regime threshold auto-regressive model. The results indicate that, in the long run, price transmission is perfect and any supply or demand shocks are fully transmitted along the marketing chain. In the short run, price adjustments between the farm and the retail levels are asymmetric and reveal a demand-pull transmission mechanism.

Researchers have examined other product markets for asymmetric price transmission. Given the considerable importance of beef and pork in the US food marketing system, much research effort has been devoted to under-standing price transmission across levels of those markets. For example, Boyd and Brorsen (1988) find that wholesale pork prices respond similarly to farm pork price decreases and increases, and pork retailers respond to wholesale pork price increases and decreases in ways that are not significantly different. Schroeder (1988) investigated price trans-mission for specific pork cuts between wholesale and retail markets. Although there is some evidence that retailers respond more quickly to increasing wholesale prices than decreasing wholesale prices, the cumulative effects are very similar. Evidence of asymmetric price transmission have been found by Von Cramon-Taubadel (1998), Goodwin and Holt (1999), Goodwin and Harper (2000), Abdulai (2002), and Miller and Hayenga (2001). Other studies providing evidence for asymmetric price transmission in non–agricultural product markets include Borenstein et al. (1997), Balke et al. (2002) and Al-Gudhea et al. (2007).

Local rice market in Cote d'Ivoire and potential sources of asymmetries

Local rice is produced mainly for domestic consumption and is one of the staple food crops grown throughout the country. Rice production has increased from 634 228 tons in 2001 to 703 931 tons in 2005. Its consumption is around 36 kg per capita per year. It constituted 49% of cereal production and 48% of cereal demand in the country (MINAGR, 2007). Major producing areas are located in the South, the West and the Centre. Although rice was traditionally marketed by a private commercial system, the government intervened in the distribution of staple food crops to help reduce price variability among regions. In practice, however, government failed to achieve this objective and prices varied across regions and time. Since the 1990s government has embarked on economic reforms that include dismantling of price and wage controls and reducing its role in domestic and external food marketing with the goal to create a competitive environment for private agents including farmers and traders. Such market liberalization reforms are consistent with economic theory, which postulates that the proper functioning of markets and marketing channels is essential for the optimal allocation of resources. Today the rice market has been completely liberalized. However, government still intervenes through the Aid Office to the Marketing of Food Products (OCPV) to facilitate and promote food marketing throughout the country.

Food trade is largely organized through networks of traders linked by personal and ethnic ties. Two actors dominated the local rice marketing system in Côte d'Ivoire: wholesalers and retailers. Wholesalers obtain rice either directly from farmers with whom they have long-standing relationships or through resident assemblers. The local wholesalers then sell to long-distance traders serving the urban markets throughout the country. The Abidjan markets are supplied from internal areas because Abidjan is not a production area like Bouaké and Korhogo. Asymmetries in these three markets may arise due to the retailers' use of market power to transmit price increases faster than price decreases. Retailers try to maintain their 'normal' profit margin when prices rise, but they try to capture the larger margins that arise, at least temporarily, when downstream prices fall. The situation lasts as long as consumer search costs are present.

Figure 1 shows the evolution of monthly wholesale and retail prices over the period 1990–1999 for each of the three markets under study. An overall look at this figure shows that wholesale and retail prices move in parallel. This gives an indication that the two series are cointegrated in the long run. However, this visual inspection does not give the nature of this relationship.

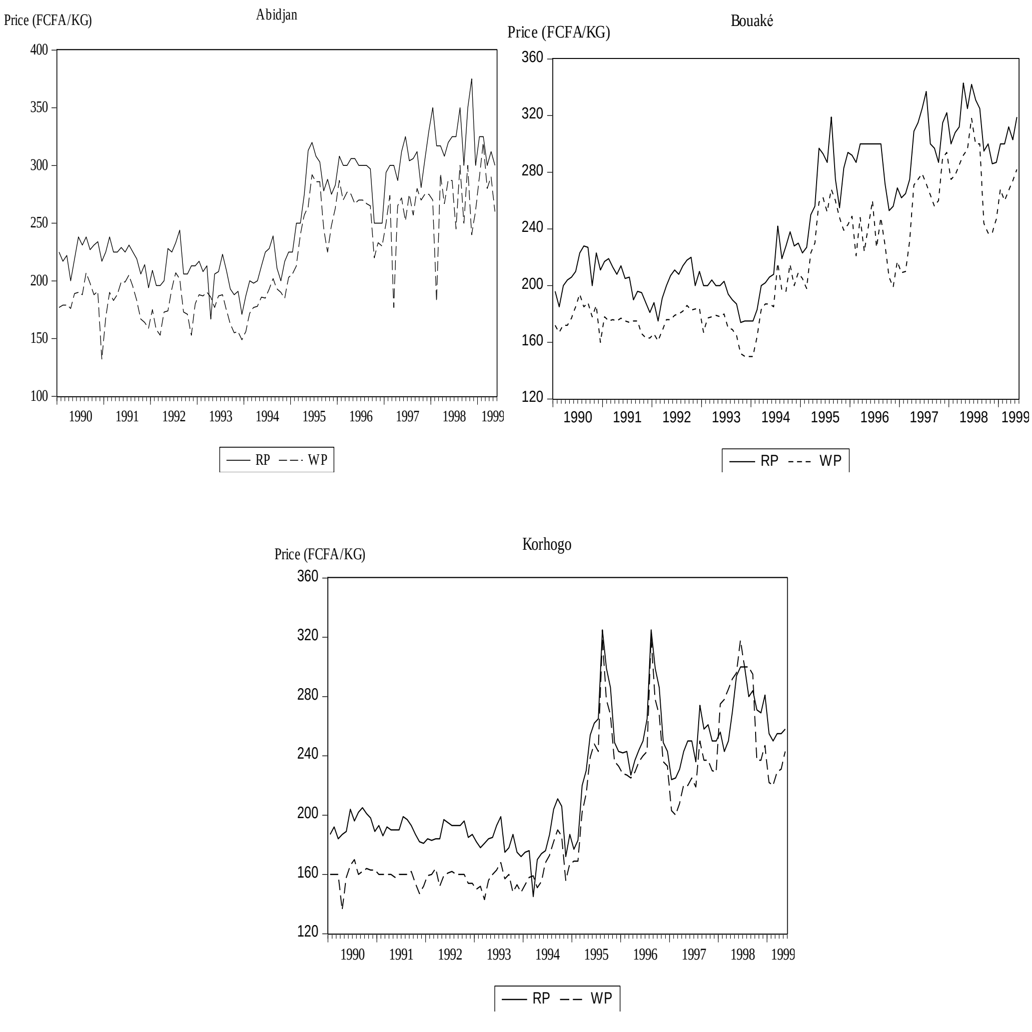

Figure 1: Evolution of wholesale and retail prices over the period 1990-1999

Therefore, we divided the sample period into two sub-periods depending on whether the wholesale price was increasing or decreasing. Then, we calculated the monthly average percentage change in wholesale and retail prices in each sub-sample to obtain, although very roughly, some indication about potential asymmetric relationship. As can be seen from Figure 2, on average, positive changes in wholesale and retail prices seem to be larger than negative changes. In the Abidjan market, wholesale price has increased by 8.84% while retail price has increased by 3.51%. If the wholesale-retail price relationship was symmetric the 6.73% decrease in wholesale price would lead to a 2.67% decrease in retail price instead of a decrease of 2.72%. Thus decrease in retail price is greater than that is required. Similarly, increases in wholesale price are transmitted moderately to retail price. In the other two markets, retail prices rise faster than they fall. For example, if symmetric price transmission holds in the Bouaké market retail price should decrease by 2.65% instead of 2.3% and increase by 2.75% instead of 3.18%. The price transmission mechanism in Abidjan market seems different from that of the other markets in the sense that changes in retail prices are moderate in the Abidjan market. However, the retail-marketing margin in Abidjan market is greater (37 FCFA/KG) than that of the Bouaké (33 FCFA/KG) and Korhogo (20 FCFA/KG) markets.

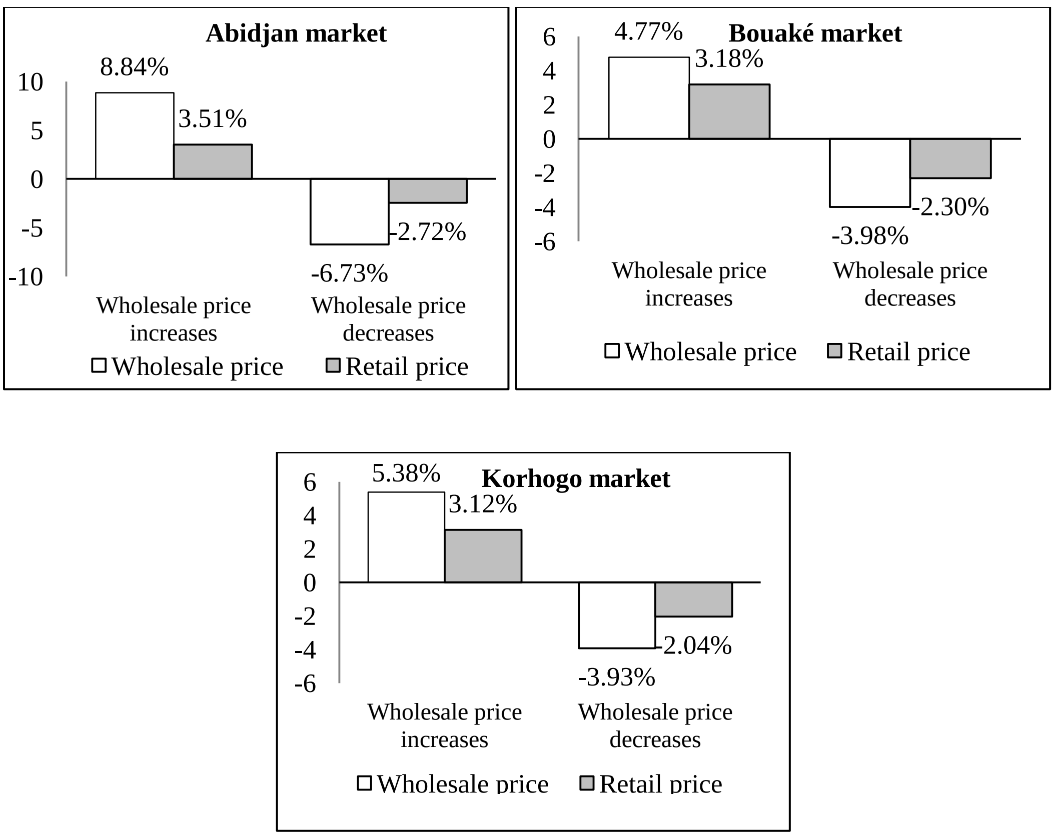

Figure 2: Average monthly changes in wholesale and retail prices when wholesale price is increasing or decreasing

Overall, these figures provide some kind of asymmetric price transmission in the local rice market. However, the issue has to be further explored with appropriate methodological tools, which will be the aim of the following section.

Econometric Methodology

This paper employs the Enders and Siklos (2001) test for threshold cointegration to examine the temporal relationship between wholesale price (W_t) and retail price (R_t). This approach extends the Engle and Granger's (1987) cointegration procedure to encompass possible asymmetric adjustment to disequilibrium. The cointegration approach is based on the fact that deviations from equilibrium conditions for two non-stationary variables should be stationary. A significant implication is that, while individual price series may wander extensively, they should not diverge from one another in the long run. Assuming both variables are I(1) process, the Engle-Granger two-step procedure considers first the retail-wholesale price relationship among the variables in levels:

$$R_t = \beta_0 + \beta_1 W_t + \mu_t \qquad (1)$$

where β_i are parameters to be estimated, and μ_t is a disturbance term which may be serially correlated. The existence of long-run relationship involves stationary μ_t. The augmented Dickey-Fuller (ADF) statistic is used to ascertain whether the residuals μ_t are stationary. To accept stationarity of μ_t, we have to obtain $-2 < \rho < 0$ in the regression given by:

$$\Delta\mu_t = \rho\mu_{t-1} + \sum_{i=1}^{p} \phi_i \Delta\mu_{t-i} + \varepsilon_t \qquad (2)$$

where ε_t is a white-noise process.

However, a crucial limitation of Eq. (2) is that it does not capture the asymmetric nature of the responses of retail prices to wholesale prices. It assumes the short-run adjustment towards long-run equilibrium to be always present and time invariant. Nevertheless, according to Balke and Fomby (1997), the movements towards equilibrium value do not always appear or at least do not have the same intensity. Eq. (2) is particularly inappropriate if prices are sticky in the downward direction, but not in the upward direction. In addition, Pippenger and Goering (1993), Balke and Fomby (1997) and Enders and Granger (1998) showed that standard tests for unit root and cointegration have low power in the presence of non-linear adjustment towards the long-run relationship. The key point of our study is that if we presume the presence of asymmetric behavior in wholesale and retail prices, then these tests must be modified to account for such asymmetry.

As in Enders and Granger (1998) and Enders and Siklos (2001), we introduce asymmetric adjustment to the model by letting the deviation from the long-run equilibrium μ_t behave as a two regimes Momentum Threshold Autoregressive (MTAR) process. Thus, we replace Eq. (2) with:

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \sum_{i=1}^{p} \phi_i \Delta\mu_{t-i} + \varepsilon_t \qquad (3)$$

where I_t is the indicator function such that:

$$I_t = \begin{cases} 1 & if \quad \Delta\mu_{t-1} \geq \tau \\ 0 & if \quad \Delta\mu_{t-1} < \tau \end{cases} \qquad (4)$$

and τ the value of the threshold. It is also possible to allow the adjustment to depend on the level of μ_{t-1} rather than on the change in μ_{t-1}. In this case, the indicator of Equation (4) becomes:

$$I_t = \begin{cases} 1 & if \quad \mu_{t-1} \geq \tau \\ 0 & if \quad \mu_{t-1} < \tau \end{cases} \qquad (5)$$

Asymmetric adjustment is implied by different values of ρ_1 and ρ_2. The sufficient and necessary conditions for μ_t to be stationarity are $\rho_i < 0$, for i = 1, 2, and $(1+\rho_1)(1+\rho_2) < 1$ (Petrucelli and Woolford 1984). If the threshold parameter enters the model unrestrictedly, the problem of how to consistently estimate it along with the values of ρ_1 and ρ_2 emerges here[2]. To obtain a consistent estimate of the threshold, we follow the grid search procedure suggested by Chan (1993). Typically, the residuals μ_{t-1} and $\Delta\mu_{t-1}$ are sorted into an increasing order and the central 70% of observations are then considered as potential thresholds. For each possible threshold, the underlying model is estimated by OLS and the consistent threshold value $\hat{\tau}$ is found by selecting the value that minimizes the sum of squared residuals.

Once $\hat{\tau}$ is obtained the next step is to perform testing for threshold cointegration. The null hypothesis of no cointegration is $H_0 : \rho_1 = \rho_2 = 0$. This restriction can be tested through the F-statistic called the Φ-statistic. Notice, however, that the distribution of Φ is not standard and appropriate critical values have been tabulated by Enders and Siklos (2001). If the null hypothesis of no cointegration is rejected, it is worthwhile to further test for symmetric adjustment ($\rho_1 = \rho_2$). This can be done through a standard F-test because the least squares estimates of ρ_1 and ρ_2 converge to multi-variate normal distributions and the system is stationary (Tong, 1983; Enders and Siklos, 2001). Rejecting both the null hypotheses of $\rho_1 = \rho_2 = 0$ and $\rho_1 = \rho_2$ implies the existence of threshold cointegration and asymmetric adjustment.

The short-run dynamics of the price series can be examined with asymmetric error correction models to obtain additional confirmation of the threshold cointegration findings. These models can be represented as follows:

$$\Delta R_t = \varphi_{10} + \sum_{i=0}^{n} \alpha_{1i}^{+} \Delta W_{t-i}^{+} + \sum_{i=0}^{n} \alpha_{2i}^{-} W_{t-i}^{-} + \sum_{i=1}^{n} \beta_{i1} \Delta R_{t-i} + \gamma_1^{+} \hat{\mu}_{t-1} + \gamma_1^{-}(1 - I_t)\hat{\mu}_{t-1} + e_{1t} \tag{6}$$

$$\Delta W_t = \varphi_{20} + \sum_{i=0}^{n} \alpha_{2i}^{+} \Delta R_{t-i}^{+} + \sum_{i=0}^{n} \alpha_{2i}^{-} R_{t-i}^{-} + \sum_{i=1}^{n} \beta_{2i} \Delta W_{t-i} + \gamma_2^{+} \hat{\mu}_{t-1} + \gamma_2^{-}(1 - I_t)\hat{\mu}_{t-1} + e_{2t} \tag{7}$$

where $\hat{\mu}_{t-1}$ is the one-period lagged error term for the cointegrating Eq. (1), ΔW_t is decomposed as $\Delta W_t^{+} = \max(0, \Delta W_t)$ and $\Delta W_t^{-} = \min(0, \Delta W_t)$; ΔR_t^{+} and ΔR_t^{-} are similarly defined. The estimates of γ^{+} and γ^{-} determine the rate at which positive and negative deviations from the long-run equilibrium are eliminated. If $\gamma_1^{+} \neq \gamma_1^{-}$, R_t exhibits asymmetry in long-run adjustment. We can also test for short-run asymmetric price transmission based on the coefficients on ΔW_t^{+} and ΔW_t^{-}. An asymmetric impact for ΔW upon ΔR is possible if coefficients are not the same. This implies that retail prices are differently affected by positive and negative changes in wholesale prices.

Empirical Results

Data and unit root tests

The study uses monthly data from January 1990 to May 1999, giving a total of 113 observations. The end date of the sample is dictated by the political crisis of December 1999 that made difficult data collection across regions. The main variables under study are nominal retail and wholesale prices. Three local rice markets are concerned with the empirical analysis: Abidjan, Bouaké and Korhogo. These cities are located in the South, Centre and North, respectively. All data are compiled from the Aid Office to the Marketing of Food Products (OCPV) and are expressed in F CFA per kg.

[2] Some studies use a threshold value of zero. However there is no *a priori* reason to expect the threshold to be zero.

For the empirical analysis, variables are expressed in natural logarithms. From an economic point of view, this transformation also allows the estimated coefficients to be interpreted as constant elasticities.

The empirical estimation strategy involves the following steps. Before performing cointegration analysis, we first test for unit root in individual price variables. Second, we examine whether a long-run relation exists between the two price series. Third, if cointegration is found, we determine whether the relationship exhibits a threshold-type non-linearity. Four, we estimate the short-run dynamic behavior of the system.

Table 1: Unit root test results

	Level			First Difference		
Market/Series	**ADF**	**PP**	**DF-GLS**	**ADF**	**PP**	**DF-GLS**
Abidjan						
Retail price	-2.779 (-3.450)	-3.313 (-3.450)	-2.441 (-3.019)	-13.509* (-2.887)	-15.367* (-2.887)	-1.081 (-1.944)
Wholesale price	-3.500* (-3.450)	-5.179* (-3.450)	-3.330* (-3.019)	-3.767* (-2.890)	-26.888* (-2.887)	-3.789* (-1.944)
Bouaké						
Retail price	-2.753 (-3.450)	-2.714 (-3.450)	-2.637 (-3.018)	-12.568* (-2.887)	-12.553* (-2.887)	-0.822 (-1.944)
Wholesale price	-3.172 (-3.450)	-3.172 (-3.450)	-2.962 (-3.018)	-12.439* (-2.887)	-12.456* (-2.887)	-4.492* (-1.943)
Korhogo						
Retail price	-2.999 (-3.450)	-2.999 (-3.450)	-2.892 (-3.018)	-5.588* (-2.889)	-11.236* (-2.887)	-10.608* (-1.943)
Wholesale price	-3.018 (-3.450)	-3.124 (-3.450)	-2.788 (-3.018)	-11.886* (-2.887)	-11.848* (-2.887)	-11.940* (-1.943)

Notes: The numbers in parenthesis are 5% critical values. * indicates the rejection of the null hypothesis at the 5% level of significance.

To ascertain the orders of integration of the two price series, we apply the unit root tests of Dickey-Fuller (1979), Phillips-Perron (1988) and Elliott *et al.* (1996). These tests are denoted as ADF, PP and DF-GLS respectively. The test results reported in Table 1 show that both wholesale as well as retail prices have unit root. However the first differenced price series are stationary. This integration property readily lends itself to cointegration analysis.

Threshold cointegration test results

Table 2 presents the cointegration test results assuming threshold adjustment. The table reports value of the adjustment coefficients ρ_1 and ρ_2, their *t*-values and the Φ and Φ^*-statistic for the null hypothesis of a unit root in μ_t against the alternative of cointegration with asymmetric adjustment. The lag length is selected such that the Akaike Information Criterion (AIC) is minimized. The F-statistic for symmetric adjustment $\rho_1 = \rho_2$, the underlying long-run equations and the consistent estimate of the threshold are also reported in the table. As can be seen from the statistics, the estimates of ρ_1 and ρ_2 satisfy the stationary conditions. More importantly, the null of no cointegration ($\rho_1 = \rho_2 = 0$) can be rejected for both the TAR and the MTAR models for all markets, indicating that retail price and wholesale price are cointegrated. Given this finding, the null hypothesis of symmetric adjustments is tested using the standard F-statistic. The F-statistic rejects the null hypothesis of symmetric adjustment ($\rho_1 = \rho_2$) for Abidjan and Bouaké under the MTAR model at the 10% significance level. The evidence for Korhogo favours symmetric adjustment. As measured by the AIC and the SC, both TAR and MTAR models exhibit similar goodness of fit and, therefore, the one modelling framework is not significantly superior over the other. In what follows our discussion focus on MTAR models.

Table 2: Results of threshold cointegration tests

	Abidjan		Bouaké		Korhogo	
	TAR	**M-TAR**	**TAR**	**M-TAR**	**TAR**	**M-TAR**
ρ_1 [a]	-0.861 (-7.608)	-0.970 (-8.708)	-0.255 (-2.121)	-0.378 (-3.611)	-0.486 (-3.816)	-0.609 (-3.695)
ρ_2 [a]	-0.503 (-3.373)	-0.343 (-2.431)	-0.368 (-2.835)	-0.058 (-0.335)	-0.394 (-3.988)	-0.376 (-4.255)
$\rho_1 = \rho_2 = 0$ (Φ [b] or Φ^*)	34.630*	40.870*	5.251*	6.556*	15.238*	15.885*
$\rho_1 = \rho_2$ (F-test[c]) (Sig.)	3.644** (0.058)	12.118* (0.000)	0.516 (0.473)	2.902** (0.091)	0.321 (0.572)	1.545 (0.217)
Estimated threshold $\hat{\tau}$	0.072	0.078	0.046	-0.034	-0.036	0.026
Number of lags	0	0	2	2	0	0
Cointegrating relationship[d]						
Constant	1.180 (5.948)		0.308 (2.520)		1.519 (13.703)	
Wholesale price	0.809 (21.899)		0.969 (42.307)		0.732 (34.850)	
AIC	-2.214	-2.278	-3.550	-3.573	-3.440	-3.444
SC	-2.166	-2.229	-3.452	-3.474	-3.391	-3.396

Notes: [a] Coefficients and t-statistics for ρ_1 and ρ_2.
[b] Φ denotes the F-test for the null hypothesis $\rho_1=\rho_2=0$. Critical values are from Enders and Siklos (2001).
[c] *F* shows the sample F-statistic for the null hypothesis tests for symmetry $\rho_1=\rho_2$. *p*-values are in parenthesis below.
[d] The numbers in parenthesis are t-statistics.
* and ** denote significance at the 5% and 10% levels, respectively.

Given that $|\rho_1| > |\rho_2|$ in the MTAR model, the speed of adjustment is faster for positive deviations (above threshold) from the long-run relationship than for negative deviations (below threshold). Hence, positive deviations from the long-term equilibrium resulting from increases in retail price or decreases in wholesale price are eliminated much faster than deviations resulting from decreases in retail price or increases in wholesale price. For example, the point estimate of ρ_1 for Abidjan market indicates that approximately 97% of a positive deviation from the long-run equilibrium relationship is eliminated within a month. On the other hand, the point estimate of ρ_2 indicates that 34% of a negative deviation from the long-run relationship is eliminated within a month. Similarly, the estimates for the Bouaké market indicate that the adjustment toward the long-run relationship tends to persist more when wholesale prices are increasing and reverts more quickly when they are decreasing. The evidence provided here indicates that the transmission of prices in the Abidjan and Bouaké markets displays some asymmetry.

Short-run dynamics

Since cointegration exists among the variables, their short-run dynamics can be modelled within error correction models (Engle and Granger 1987). However, the positive finding of threshold cointegration implies that it is incorrect to estimate a symmetric error correction model. Such a model would not reveal differential adjustment under the two regimes. Hence, an asymmetric version of the error correction model is estimated for Abidjan and Bouaké markets to allow for asymmetric adjustment in response to positive and negative price changes. Using the long-run relations reported in Table 2, the results of the fitted error-correction models are presented in Table 3 along with test statistics regarding long and short-run symmetry. An important finding is the statistical significance of the error correction terms in the asymmetric error correction model. While wholesale prices adjust in the wrong direction, retail prices adjust in the right direction by acting to eliminate deviations from long run equilibrium. The *t*-statistics show that the coefficients on the positive and negative error correction terms (i.e. $\hat{\mu}_t^+$ and $\hat{\mu}_t^-$) are significant and noticeably different in Abidjan market. This indicates that retail prices respond differently to positive and negative discrepancies in the long-run price relationship between retail and wholesale prices. Retail price growth is faster when there is a positive deviation from long-run equilibrium than when there is a negative deviation.

The point estimates show that retail prices in Abidjan adjust so as to eliminate more than 84% of a unit positive change, but about 53% of a negative change in the deviation from the equilibrium relationship created by changes in wholesale prices. Similarly, retail prices in Bouaké adjust so as to eliminate approximately 45% of a unit positive deviation and 29% of a unit negative deviation from the long run relationship created changes in wholesale prices. For Korhogo, retail prices adjust so as to eliminate 43% of disequilibrium from long-run relationship.

Table 3: Estimates of the error correction models and symmetry testing results

	Abidjan		**Bouaké**		**Korhogo**	
	ΔR_t	ΔW_t	ΔR_t	ΔW_t	ΔR_t	ΔW_t
Constant	-0.011* (-2.039)	-0.004 (-0.405)	-0.006 (-1.376)	-0.008 (-1.488)	-0.005 (-1.085)	-0.006 (-1.031)
ΔW_t^+	0.750* (5.990)	-	0.876* (8.786)	-	0.725* (8.650)	-
ΔW_t^-	0.142** (1.809)	-	0.455* (4.551)	-	0.457* (4.819)	-
ΔR_t^+	-	0.401 (1.723)**		1.008* (7.931)	-	1.048* (8.194)
ΔR_t^-	-	0.919* (4.964)		0.607* (4.697)	-	0.681* (5.169)
$\hat{\mu}_{t-1}^+$	-0.844* (-6.122)	1.188* (9.002)	-0.452* (-5.521)	0.526* (5.834)	-0.429* (-5.651)	0.417* (4.276)
$\hat{\mu}_{t-1}^-$	-0.524* (-4.363)	-0.012 (-0.066)	-0.288** (-1.895)	0.179 (1.034)	-	-
	Symmetry Testing					
Long-run[a] ($\gamma^- = \gamma^+$)	3.350** [0.067]	26.300* [0.000]	0.887 [0.346]	3.255** [0.071]	-	-
Short-run [b] ($\alpha^- = \alpha^+$)	13.132* [0.000]	2.298 [0.129]	6.667* [0.009]	3.557** [0.059]	3.577** [0.058]	3.137** [0.076]
R^2	0.413	0.497	0.587	0.582	0.588	0.539
AIC	-2.827	-2.012	-3.766	-3.542	-3.482	-3.116
SC	-2.704	-1.890	-3.644	-3.420	-3.385	-3.019

Note: t-statistics are in parentheses. Asymmetric error correction models are based on the MTAR. [a] Wald test for the null hypothesis that the adjustment coefficients are equal. Significance levels are in brackets. [b] Wald test for the null hypothesis that positive and negative changes in wholesale prices are the same effect on retail prices. Significance levels are in brackets below.

Considering the individual coefficients on price change variables, the coefficients of positive and negative price change are all significantly positive. This implies that wholesale price increases (decreases) lead to retail price increases (decreases). The size of the short-run coefficients indicates a relatively high response of retail price to a positive change in wholesale price in all markets. Thus, increases in wholesale prices are transmitted faster to retail prices than decreases. Overall, these results indicate that price transmission asymmetries do exist in Ivorian local rice markets. This implies greater costs for consumers than would occur with symmetric adjustment.

Conclusion

This paper investigated the vertical price transmission in three important local rice markets in Côte d'Ivoire in order to ascertain whether the complaint of consumers about the asymmetric price transmission holds true. To this end, the concept of price transmission was decomposed into co-movement, dynamics and speed of adjustment within threshold cointegration and asymmetric error correction model framework. Using monthly data over the 1990–1999, we found evidence supporting the asymmetric price responses hypothesis in two markets: increases in wholesale prices are passed on to retail prices more rapidly than decreases. The results also show that speed and magnitude of retail prices response to changes in wholesale prices are much higher for Abidjan market than Bouaké market. This is consistent with the fact that the Abidjan market is mainly a market for consumption and exhibits a higher trading intensity than the Bouaké market, which is both a consumption and production area. However, these findings do not imply either market power or supernormal profits among traders in the Abidjan and Bouaké markets. Some of the hypotheses discussed in the literature do suggest that temporary market power could explain the asymmetry. The explanation for the underlying price transmission asymmetries is an issue worthy of further research. We are preparing a survey to investigate this issue. Meantime, our findings suggest government intervention in Abidjan and Bouaké markets in order to control price variations and preserve consumer purchasing power in a context of widespread poverty.

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