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Abstract

This paper considers vertical price relationships between wholesalers and retailers on five local maize markets in Benin. We show that if the common factor and the long-run disequilibrium error are not explicitly taken into account in testing the channel model, one can easily be wrong about how restrictions on the error-correction structure must be interpreted in terms of economic power in the channel. The empirical results show interesting differences between markets and reveal that retailers play a more prominent role in the price formation process than generally assumed in the literature. Retailers in the two major towns do not allow wholesalers to behave as vertical price leaders, but in the two larger rural centers, wholesalers involved in arbitrage among urban markets are able to influence price formation.

JEL classification: C32; D40; L10; O18; Q13

Keywords: Vertical price leadership; Marketing channels; Cointegration; Common Factor; Benin

1. Introduction

In the literature on industrial organization retail prices are often assumed to be determined by wholesale market conditions (see, for example, Tirole, 1988, Chapter 4; and Martin, 1993, Chapter 12). Likewise, in the marketing literature on the functioning of food markets in tropical countries, the vertical price leadership of wholesalers is often assumed. This is often based on popular complaints from retailers and market organizations. Looking at the Benin maize market, Lutz (1994) found that retail and wholesale price series in the same market place cohere, which implies that retail margins are stationary. This result suggests that retailers are indeed passive decision makers, following wholesale prices without taking local supply and demand conditions into account. On the contrary, a survey among traders showed that a large number of wholesalers from different surplus regions supply the urban market and that the buying strategy of most retailers is to buy from the wholesaler who proposes the lowest price (Lutz, 1994). Moreover, in rural areas retailers can choose to buy either from wholesalers or at the farm gate. Buying directly from farmers may provide retailers some freedom to set prices. Consequently, it is not a settled matter whether wholesalers or retailers or both have some market power and are able to influence price formation.

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In an earlier study on price arbitrage in the wholesale segment of the maize market, we concluded that all wholesale markets played a role in the price formation process (Kuiper et al., 1999). None of the price series of any of the wholesale markets were found to be dominant: all price series were interdependent. The arbitrage process corresponded to a network with a number of interdependent wholesale markets; there were no autarkic markets and transportation costs did not show a stochastic trend. The study, however, did not incorporate the price series observed on the retail segments. In the present paper we focus on this omission, questioning the relationship between prices on wholesale and retail market segments in various markets for the same sample period as in Kuiper et al. (1999). The questions we set out to answer are: is there a difference in wholesale-retail price relationships in towns and rural centers, and is there any evidence for wholesale market dominance vis-à-vis the retailers?

Most studies on vertical price relationships published to date in marketing and industrial organization (see, for example, Gerstner and Hess, 1991; Lee and Staelin, 1997 and the references they cite) have used comparative statics to study channel behavior; the long-run relationships derived have not been empirically tested. Our study differs in that its main focus is empirical analysis. In order to verify whether the price formation process is driven by retailers, wholesalers, or both, we can distinguish two segments in the market: the retail segment and the wholesale segment. We assume that actors in both segments try to maximize profits. To examine whether or not wholesalers are price leader vis-à-vis the retailers in the sense of Stackelberg leadership, one can consider the long-run equilibrium (i.e., cointegrating) relationship between the wholesale and retail prices and test whether or not wholesale prices and retail prices respond to deviations from the equilibrium price.

The basic assumption we make is that the common stochastic factor observed in the cointegrated wholesale-retail price series is generated by local supply and demand conditions (seasonal price trend). Three models then become interesting for the study of price adjustment: *Model* 1 in which both retailers and wholesalers have some freedom to respond to deviations from the equilibrium price, *Model* 2 in which only retailers have the power to respond to deviations from the equilibrium price; and *Model* 3 in which only wholesalers have the freedom to respond to deviations from the equilibrium price.

In *Model* 1 wholesalers have sufficient power vis-à-vis the retailers to behave as vertical price leaders, although retailers can still maximize their profits conditional on the wholesale price being set by the wholesalers. This model applies if both retail and wholesale traders exercise some market power; for example, if alternative market opportunities exist for both actors. In contrast, in *Model* 2 the retailers do not allow wholesalers to influence short-run retail price deviations and leave them with only the option of setting wholesale prices on the basis of the wholesalers' unit costs (i.e., farm gate price plus a margin to enable the wholesaler to survive), which represents the common factor that drives the two prices, the retail price and the wholesale price, in the long run. Market power for retailers may be the result of a temporarily abundant supply in the wholesale segment and a lack of alternative market opportunities for wholesalers. Lastly, in *Model* 3, only wholesalers are able to set their prices in the sense of Stackelberg leadership and to respond to price deviations from the equilibrium. The situation applying in this model is one in which large numbers of retailers buy from farmers and wholesalers to serve local consumers, whereas the local wholesalers are also involved in regional market arbitrage and ship to urban markets. Consequently, the retail price is getting stuck to the common factor from which the wholesale price can deviate in the short run by price arbitrage among the spatially dispersed wholesale markets.

In deriving the testable implications of the hypotheses about economic power in marketing channels, we explicitly take the common factor and the deviations from the long-run vertical price equilibrium into account This is the major contribution of our paper to the debate on vertical price leadership and we will show that if the common factor and disequilibrium error are not explicitly assigned to certain variables in the channel model, one can easily be wrong about how restrictions on the error-correction model must be interpreted in terms of vertical price leadership. Furthermore, since we wish to test between the three theoretical models outlined above, another important advantage of our empirical method is that it nests these tests in one procedure.

The article analyzes the process of price formation for maize in five market places in the south of Benin: two towns (Bohicon and Cotonou) and three rural centers (Azové, Kétou and Dassa). Section 2 will discuss the relevance of the three above-mentioned models. In Section 3 the method of analysis is presented. We formulate the long-run model and derive its testable implications on the short-run price system. Section 4 presents the empirical results and Section 5 the conclusions.

2. Relevance of the distinguished market models

The market for maize, the staple food crop in the south of Benin, consists of a number of market places, scattered throughout the region. Most transactions take place in spot markets; buyers and sellers meet in the market place where the maize for sale is displayed on the market day. Maize is transferred from producers to final consumers through conventional marketing channels, where more or less homogenous products are traded between actors who are not involved in recurrent trade relationships. In each market place a retail and a wholesale segment can be distinguished. In the large towns, local retailers generally buy on the wholesale segment of the market, while retailers in rural centers can choose to obtain their stocks either from the wholesale segment or directly at the farm gate.

Cotonou (Dantokpa) and Bohicon are two important urban markets in the country. Both market places are a centre for retail trade. As urban price levels are relatively high, supply on the local wholesale segment is directed to serve only local retail demand. Sometimes wholesalers organize themselves in order to address specific problems. However, because of the large number of wholesalers and brokers active in the market, there are no enforcement mechanisms to control entry or prices. Moreover, the wholesalers originate from all regions in the country which means that they have no real common interest. Consequently, entry into the wholesale market-segment is free (Lutz, 1994). On the other hand, for retailers entry is constrained by a lack of space on the market: most retailers have a permanent place. Retailers try to tie clients by selling on credit and by negotiating the amounts per unit of measurement. Apparently, retailers have some freedom to deviate from equilibrium prices. Based on the literature we were inclined to expect *Model* 3 to simulate wholesaler-retailer relationships. However, based on our observation that a large number of wholesalers supply the wholesale market segment in the towns and that there is some room for monopolistic competition among the retailers, we argue that also *Model* 2 can hold for both Cotonou and Bohicon.

Important surpluses of maize are traded from Azové and Dassa to the two towns. Consequently, the wholesalers in these rural markets are involved in regional market arbitrage, and hence anticipate supply and demand conditions in the different, but spatially price-integrated, wholesale markets. On both markets a large number of retailers are found. They buy directly from farmers that supply early in the morning a part of their surplus on the wholesale segment (they need money to finance that days' purchases on the market), or buy from local wholesalers. From these observations our empirical results with respect to Azové and Dassa are expected to be in line with the assumptions made in the literature and to comply with *Model* 3.

Kétou is considered to be a rural market place, trading large maize surpluses. On the wholesale market segment a relatively large number of local wholesalers sell to non-resident urban wholesalers, in particular from Cotonou. Wholesalers in Kétou do not have a local alternative for the demand from Cotonou, because there are very few retailers in Kétou (approximately five per market day), who mainly buy at the farm gate. The local retail market is thin as most residents buy directly from farmers or are farmers themselves. However, some local consumers depend on the market in Kétou. This implies that the small number of retailers may exercise some monopolistic behavior. Therefore, *Model* 2 can be expected to be the most appropriate for describing the situation.

3. Method

3.1. Theoretical framework

Let us consider a two-stage channel with M ($M \geq 1$) wholesalers upstream and N ($N \geq 1$) retailers downstream ($M \leq N$). We model the long-run supply decision behavior of these channel members. During the period covered by one time series observation t (e.g., a day in case of daily observations), each wholesaler j ($j = 1, \ldots, M$) exclusively supplies M_j retailers ($M_j \geq 1 \wedge N = \sum_{j=1}^{M} M_j$). The retailer buys an amount of q_i ($i = 1, \ldots, N$) of an intermediate good from the wholesaler at a wholesale price p_{wi}. The wholesaler acquired the intermediate good at a constant unit cost $p_{f\,i}$ (the weighted average of the farm gate price faced by the wholesaler with respect to q_i) and distributed it at a constant unit cost c_{wi}, Retailer i faces constant unit retailing cost, c_{ri}, and resells the product to the consumers at a price p_{ri} on the retail market. It is assumed that the wholesalers and retailers do not throw away any of the intermediate good. Consequently, the quantity bought by the wholesalers is equal to the quantity finally consumed.

Let the consumer behavior faced by retailer i be given by the following flexible demand function (see, e.g., Lilien and Kotler, 1983, p. 74):

$$p_{ri} = s_i q_i^{\delta} + x_i, \tag{1}$$

where q_i is the quantity sold by retailer i and s_i and x_i capture exogenous shifts in the demand curve and may also contain a constant term.

We first consider the Stackelberg model in which the wholesalers are the vertical price leaders, i.e., each retailer i maximizes profit conditional on the wholesale price that has to be paid to the wholesaler, and the wholesaler then determines q_i or, similarly, p_{wi}, by maximizing profit while taking the conditional profit-maximizing behavior of retailer i into account. The conditional profit-maximization problem of retailer i can be written as:

$$\max_{q_i} \, (p_{ri} - c_{ri} - p_{wi})q_i \,, \tag{2a}$$

or equivalently,

$$\max_{p_{ri}} (p_{ri} - c_{ri} - p_{wi})q_i \tag{2b}$$

subject to (1). The first-order condition for this problem is:

$$p_{ri} + (dp_{ri}/dq_i)q_i - c_{ri} - p_{wi} = 0 \ , \tag{3a}$$

or equivalently,

$$q_i + (dq_i/dp_{ri})(p_{ri} - c_{ri} - p_{wi}) = 0 \ . \tag{3b}$$

From each of both (3a) and (3b) it follows that:

$$p_{wi} = (1 + \delta)p_{ri} - c_{ri} - \alpha x_i \tag{4}$$

Wholesaler j maximizes individual profit while taking the conditional profit-maximizing behavior of the retailers into account, so that

$$\sum_{k=1}^{M_j} \max_{q_k} (p_{wk} - c_{wk} - p_{fk})q_k, \tag{5a}$$

or equivalently,

$$\sum_{k=1}^{M_j} \max_{p_{wk}} (p_{wk} - c_{wk} - p_{fk})q_k \tag{5b}$$

is subjected to (4) and has the following M_j first-order conditions:

$$p_{wk} + (dp_{wk}/dq_k)q_k - c_{wk} - p_{fk} = 0 \ , \tag{6a}$$

or equivalently,

$$q_k + (dq_k/dp_{wk})(p_{wk} - c_{wk} - p_{fk}) = 0 \tag{6b}$$

with $k = 1, \ldots , M_j$. Using (4), from each of both (6a) and (6b) we can derive M_j linear combinations of the wholesale and retail prices without q_k included:

$$p_{wk} + \delta(1 + \delta)p_{rk} = c_{wk} + p_{fk} + \delta(1 + \delta)x_k \tag{7}$$

Recall that $N = \sum_{j=1}^{M} M_j$. Consequently, the total number of relations given by (7) equals N.

If we express the price relationships (4) and (7) in weighted average market prices, we obtain:

$$p_w = (1 + \delta)p_r - c_r - \delta x \ , \tag{8a}$$

$$p_w + \delta(1 + \delta)p_r = c_w + p_f + \delta(1 + \delta)x \ , \tag{8b}$$

where

$$p_r = \sum_{i=1}^{N} p_{ri} q_i / \sum_{i=1}^{N} q_i \; ; \;\; p_w = \sum_{i=1}^{N} p_{wi} q_i / \sum_{i=1}^{N} q_i \; ; \;\; c_r = \sum_{i=1}^{N} c_{ri} q_i / \sum_{i=1}^{N} q_i \; ;$$

$$c_w = \sum_{i=1}^{N} c_{wi} q_i / \sum_{i=1}^{N} q_i \; ; \;\; p_f = \sum_{i=1}^{N} p_{fi} q_i / \sum_{i=1}^{N} q_i \; ; \;\; x = \sum_{i=1}^{N} x_i q_i / \sum_{i=1}^{N} q_i \; ;$$

which shows that it is important to collect the data for each retail account instead of taking the wholesaler as an account, because if we define

$$p_{wj} \equiv \sum_{k=1}^{M_j} p_{wk} q_k / \sum_{k=1}^{M_j} q_k \text{ and } q_{wj} \equiv \sum_{k=1}^{M_j} q_k \text{ , then } \sum_{j=1}^{M} p_{wj} q_{wj} / \sum_{j=1}^{M} q_{wj} \neq p_w \, .$$

Solving (8a) and (8b) for p_r and p_w gives:

$$p_r = (1 + \delta)^{-2}\{c_r + c_w + p_f + [(1 + \delta)^2 - 1]x\} \tag{9a}$$

$$p_w = (1 + \delta)^{-1}\{c_w + p_f - \delta c_r + \delta x\}. \tag{9b}$$

In this study it is interesting to observe that if prices are set according to (9a) and (9b), then the wholesalers have enough power vis-à-vis the retailers to behave as vertical price leaders in choosing p_w. However, if the retailers dominate, then we may have a situation in which each retailer maximizes profit and forces the wholesaler to set prices on the basis of total unit costs alone, leading to the following expression in weighted averages:

$$p_w = c_w + p_f. \tag{10}$$

So far, two models have been considered: the model made up by (9a) and (9b), *Model* 1, according to which the wholesaler is able to manipulate the retail price by dp_{wk}/dq_k in (6a) (or, similarly, by dq_k/dp_w in (6b)), and the model formed by the weighted average of (4):

$$p_w = (1 + \delta)p_r - c_r - \alpha x \tag{11}$$

and (10), *Model* 2, which says that the retailers dominate. In addition, a third model, *Model* 3, is obtained if we assume that the retailer is able to buy directly from the farmer and set p_r on the basis on p_f as follows:

$$p_r = c_w + p_f \tag{12}$$

where c_w is added to cover the costs that would otherwise be made by the wholesaler. Competition among retailers tends to be pure and as a consequence, a retailer is not able to charge a price that is different from the one determined by (12). Nevertheless, in spite of the retailer's ability to buy directly from the farmer, the retailer can also buy from the wholesaler while having p_r still determined by (12). Although the wholesaler is not able to influence the local retail market because p_r is fixed by (12), the wholesaler can still determine p_w by (9b) if involved in market arbitrage, so that p_w can be based on the reaction function of retailers in other local markets that are unable to buy directly from the farmer. The testable implications of the three models being considered will be discussed in the next subsection.

3.2. *Econometric considerations*

Many economic time series, like p_{rt} and p_{wt} ($t = 0, 1, \ldots, T$), do not fluctuate around a constant in a seemingly random way, but their first differences, $\Delta p_{rt} = p_{rt} - p_{r,t-1}$ and $\Delta p_{wt} = p_{wt} - p_{w,t-1}$, do (Nelson and Plosser, 1982; Granger and Newbold, 1986). Consequently, the variables in levels, p_{rt} and p_{wt}, are assumed to be nonstationary by containing a unit root, while in first differences they will be stationary. In time series analysis this is expressed by saying that p_{rt} and p_{wt} are integrated of order one, denoted $p_{rt} \sim I(1)$ and $p_{wt} \sim I(1)$, and Δp_{rt} and Δp_{wt} are integrated of order zero, denoted $\Delta p_{rt} \sim I(0)$ and $\Delta p_{wt} \sim I(0)$.

The nonstationarity in case of a unit root is caused by a so-called 'stochastic trend' (Banerjee et al., 1993, p. 153), which can be interpreted as the driving force of the variable. If two variables are driven by the same stochastic trend, then a linear combination of the two will be stationary, which is expressed by saying that the two variables are 'cointegrated' (Engle and Granger, 1987) or, equivalently, have a 'common stochastic trend' (Stock and Watson, 1988).

At first sight, there appear to be three variables by which a stochastic trend could enter the price system derived from (1): x_t, c_{rt}, and $c_{wt} + p_{ft}$. For now we simply assume that x_t and c_{rt} do not contain a stochastic trend of importance when compared with the stochastic trend generated by the prices of the raw product as represented by the farm gate price p_{ft}. Consequently, we assume that $c_{wt} + p_{ft}$ introduces the stochastic trend in the price system, expressing local supply and demand conditions and seasonal factors. In the empirical analysis the stationarity assumption of x_t and c_{rt} is tested by the concept of cointegration.

To illustrate the relationship between the concept of a stochastic trend and the concept of cointegration, let us consider the retail price p_{rt} and the wholesale price p_{wt} in a vector autoregression of order k, denoted VAR(k), as follows:

$$\Delta \mathbf{X}_t = \mathbf{\Pi X}_{t-1} + \sum_{j=1}^{k-1} \mathbf{\Gamma}_j \Delta \mathbf{X}_{t-j} + \mathbf{\Phi D}_t + \boldsymbol{\varepsilon}_t \tag{13}$$

where $\mathbf{X}_t = [p_{rt}, p_{wt}]' \sim I(1)$, $\Delta \mathbf{X}_t = \mathbf{X}_t - \mathbf{X}_{t-1}$, the $\mathbf{\Pi}$ and $\mathbf{\Gamma}_j$ ($j = 1, \ldots, k-1$) are (2×2) parameter matrices, $\mathbf{\Phi}$ is a (2×m) parameter matrix, $\mathbf{D}_t$ is an (m×1) vector with deterministic elements, $\boldsymbol{\varepsilon}_t = [\varepsilon_{rt}, \varepsilon_{wt}]'$ are disturbances that follow a two-dimensional Gaussian white noise process, and the values of $\boldsymbol{X}_{-k+1}, \ldots, \boldsymbol{X}_0$ are fixed. Notice that there can never be a relationship between a variable with a stochastic trend and a variable without a stochastic trend. So, if $\Delta \mathbf{X}_t \sim I(0)$ since $\mathbf{X}_t \sim I(1)$ (and hence, $\mathbf{X}_{t-1} \sim I(1)$), then $\mathbf{\Pi}$ will be a zero matrix except when a linear combination of the variables in $\mathbf{X}_t$ is stationary, i.e., when p_{rt} and p_{wt} are cointegrated (or when one of the prices is stationary so that we should also test for the absence of each individual price in the cointegrating relation to justify our assumption that both prices are $I(1)$). Because this linear combination is unique, the rank of $\mathbf{\Pi}$ will be equal to one, i.e., rank($\mathbf{\Pi}$) = 1. Hence, rank($\mathbf{\Pi}$) = 0 if there is no cointegration and rank($\mathbf{\Pi}$) = 2 if $\mathbf{X}_t \sim I(0)$. The Johansen procedure (for example, Johansen and Juselius, 1990; and Johansen, 1995) estimates the parameters in (13); to test for cointegration, trace statistics are used to determine the rank of $\mathbf{\Pi}$, and asymptotic t statistics are used to test for the absence of each individual price in the long-run equilibrium, in order to check whether both price series are $I(1)$.

Clearly, the result of interest will be rank($\mathbf{\Pi}$) = 1. In this case $\mathbf{\Pi}$ can be decomposed into $\mathbf{\Pi} = \boldsymbol{\alpha\beta}'$, where $\boldsymbol{\alpha} = [\alpha_r, \alpha_w]'$ is the adjustment vector and $\boldsymbol{\beta} = [\beta_r, \beta_w]'$ is the cointegrating vector, so that (13) becomes a vector error-correction model (VECM):

$$\Delta\mathbf{X}_t = \alpha\beta'\mathbf{X}_{t-1} + \sum_{j=1}^{k-1} \Gamma_j\Delta\mathbf{X}_{t-j} + \Phi\mathbf{D}_t + \varepsilon_t \tag{14}$$

where $\beta'\mathbf{X}_t \sim I(0)$ and represents the deviations from the long-run equilibrium, that is, cointegrating, relationship between p_{rt} and p_{wt}, and the changes in at least one of the prices, $\Delta\mathbf{X}_t$, respond to these deviations from the previous period, $\beta'\mathbf{X}_{t-1}$, through the adjustment parameters α in such a way that the disequilibrium errors $\beta'\mathbf{X}_t$, $\beta'\mathbf{X}_{t+1}$, ... converge to zero.

Premultiplying (14) by β' and rearranging, gives:

$$\beta'\mathbf{X}_t = (1 + \beta'\alpha)\beta'\mathbf{X}_{t-1} + \sum_{j=1}^{k-1} \beta'\Gamma_j\Delta\mathbf{X}_{t-j} + \beta'\Phi\mathbf{D}_t + \beta'\varepsilon_t\,. \tag{15}$$

Because $\Delta\mathbf{X}_{t-j}$ ($j = 1,...,k-1$) and ε_t are stationary, the condition $|1 + \beta'\alpha| < 1$, or equivalently, $-2 < \beta'\alpha < 0$, allows $\beta'\mathbf{X}_t$ to be stationary as well. If we return to our theoretical framework in the previous subsection, then given the assumption that x_t and c_{rt} are stationary, (11) is the linear combination of p_{rt} and p_{wt} in *Model* 1 that represents $\beta'\mathbf{X}_t$: $c_{rt} + \delta x_t = (1 + \delta)p_{rt} - p_{wt} = \beta'\mathbf{X}_t$. In *Model* 2 the cointegrating relationship is the same one as in *Model* 1, but now p_{wt} is given by (10) instead of (9b). Lastly, in *Model* 3, substituting (12) for $c_{wt} + p_{ft}$ in (9b) yields the following long-run relationship between p_{rt} and p_{wt}: $\delta(x_t - c_{rt}) = p_{rt} - (1 + \delta)p_{wt}$. Notice that each time the disequilibrium error consists of a linear combination of c_{rt} and x_t. Hence, testing for cointegration can be seen as a check of our assumption that c_{rt} and x_t are stationary.

The long-run equilibrium implies a common stochastic trend in the prices p_{rt} and p_{wt}. To study the econometric implications of this trend, let $\alpha_\perp = [\alpha_{r\perp}, \alpha_{w\perp}]'$ be a (2×1) vector of parameters of full rank such that $\alpha'\alpha_\perp = 0$. Premultiplying (14) by $\alpha_\perp'$ shows that $\alpha_\perp'\Delta\mathbf{X}_t$ does not respond to the disequilibrium errors, while its solution, $\alpha_\perp'\mathbf{X}_t$, is driven by the stochastic trend, $\alpha_\perp'\sum_{i=0}^{t}\varepsilon_i$. Further define $\beta_\perp = [\beta_{r\perp}, \beta_{w\perp}]'$ as a (2×1) vector of full rank such that $\beta'\beta_\perp = 0$. Using the relation $\beta_\perp(\alpha_\perp'\beta_\perp)^{-1}\alpha_\perp' + \alpha(\beta'\alpha)^{-1}\beta' = \mathbf{I}$ it can be seen that:

$$\mathbf{X}_t = \beta_\perp(\alpha_\perp'\beta_\perp)^{-1}\alpha_\perp'\mathbf{X}_t + \alpha(\beta'\alpha)^{-1}\beta'\mathbf{X}_t\,. \tag{16}$$

Because β does not contain zero elements, neither does $\beta_\perp$. Moreover, $\alpha_\perp'\beta_\perp$ is a scalar unequal to zero. Consequently, the stochastic trend introduced by $\alpha_\perp'\mathbf{X}_t$ is the *common* stochastic trend in the prices p_{rt} and p_{wt}. This is why Gonzalo and Granger (1995) define $\alpha_\perp'\mathbf{X}_t$ as the common factor. Note that if $\alpha_{r\perp} = 0$ ($\alpha_{w\perp} = 0$), then $\alpha_w = 0$ ($\alpha_r = 0$). The implication is that p_{wt} (p_{rt}) captures the common factor, whereas p_{rt} (p_{wt}) does all the correction to eliminate any deviation from long-run equilibrium, see also (16). In turn, this is equivalent to saying that there is long-run causality running from p_{wt} to p_{rt} (p_{rt} to p_{wt}). See Hall and Milne (1994), Granger and Lin (1995) and Gonzalo and Granger (1995) for the concept of long-run causality and see, for example, Tiffin and Dawson (1996) and Dawson and Tiffin (1998) for empirical applications.

Using the econometric concepts introduced above, we can now derive the testable implications that discriminate between our three strategic channel pricing models: *Model* 1, given by (9a) and (9b) and implying price leadership of the wholesaler; *Model* 2, formed by (9a) and (10) and implying that the retailer dominates since the wholesaler is only allowed to set its price on the basis of the farm gate price; and *Model* 3, composed of (9b) and (12) to capture the fact that the retailer buys directly from the farmer while the wholesaler is involved in market arbitrage. Given that $c_{wt} + p_{ft}$ introduces the stochastic trend, while a linear combination of the stationary variables c_{rt} and x_t represents the disequilibrium error, it follows that in *Model* 1 the common factor is captured by a linear combination of both prices, p_{rt} and p_{wt}, see (8b). Consequently, both p_{rt} and p_{wt} will display error correction. In *Model* 2 p_{wt} is not error correcting, see (10), but p_{rt} is: (11) includes c_{rt} and x_t. Lastly, in *Model* 3, p_{rt} does not show error-correcting behavior, see (12), but p_{wt} does, see (9b).

Notice that these results are counter-intuitive when compared with the literature on spatial (i.e., horizontal) price integration, where it is the price of the reference (i.e., dominant) market that should not show error-correcting behavior (for example, Silvapulle and Jayasuriya, 1994; or Dercon, 1995). On the contrary, in our first two channel (i.e., vertical) pricing models, *Model* 1 and *Model* 2, the price of the leader does respond to the error-correction term. This shows that it is important to assign the common stochastic trend and the disequilibrium errors to the respective variables in the theoretical model (in our framework the stochastic trend is generated by $c_{wt} + p_{ft}$ and the disequilibrium errors are introduced by a linear combination of x_t and c_{rt}), before formulating hypotheses on price dominance in terms of exclusion of the error-correction term.

4. Empirical analysis

All five markets are periodic and are held in a four-day cycle. Daily wholesale and retail prices are available for all five markets: for 190 market days at Cotonou (4 September 1987 to 29 September 1989), 160 market days at Bohicon (1 January 1988 to 29 September 1989), 184 market days at Azové (28 September 1987 to 29 September 1989), 174 market days at Dassa (3 November 1987 to 25 September 1989) and, lastly, 144 market days at Kétou (3 March 1988 to 25 September 1989). For each market the time series of the retail prices and wholesale prices are displayed in one figure, see Figure 1 (data available from authors on request). See Lutz (1994) or Lutz et al. (1995) for a description of the elaborate method used to collect these market prices. Annual inflation was only 2 to 3 per cent during the sample period and can be ignored when compared to the stochastic trend fluctuations in the prices. Hence, the price series were not deflated.

Figure 1 nicely shows the coherence between the wholesale and retail prices. Moreover, the considerable price fall in 1988 clearly marks the end of the lean season after the relatively bad harvest in 1987, and the start of a new promising harvest. Nevertheless, in spite of this large price shock, the estimated breakpoint test statistic of Zivot and Andrews (1992) does not reject $I(1)$ against the alternative of single structural breakpoint stationarity, see Table 1, although the test statistic for the wholesale price series of Bohicon is just a bit smaller than the 5% critical value.

For each market we considered the retail and wholesale prices and estimated bivariate VARs of order k = 1,..., 11. All computations were performed in EViews, Version 3.1. To determine the appropriate order, the commonly used Akaike criterion (AIC) was computed (see, for example, Lütkepohl, 1991). The estimate for k, denoted k^*, was chosen so that AIC was minimized. The results are presented in Table 2 and show that all k^* are much smaller than the maximum lag length fixed at 11, suggesting that it is unnecessary to conclude that it is more fruitful to increase the information set by adding new variables (and hence, equations) to the VAR rather than to automatically increase the lag length.

Next, we applied the Johansen procedure (Johansen, 1992) to jointly test for cointegration and deterministic components, see also Harris (1995, p. 97). We found the wholesale price and the retail price to be cointegrated for each market; this supports our assumption that the unit retailing cost c_t and the exogenous demand shifts captured by x_t can be considered to be stationary. The results are presented in Table 3 and are based on (13) and (14). It appears that none of the selected models contain deterministic terms. Comparing the trace statistics with their critical values shows that for each market $r = 0$ (i.e., no cointegration) must be rejected (which was true for all models for the deterministic components) while $r \leq 1$ (i.e., cointegration) cannot be rejected. In contrast to the widely used Engle-Granger two-step method for cointegration testing (Engle and Granger, 1987), the Johansen procedure is invari-

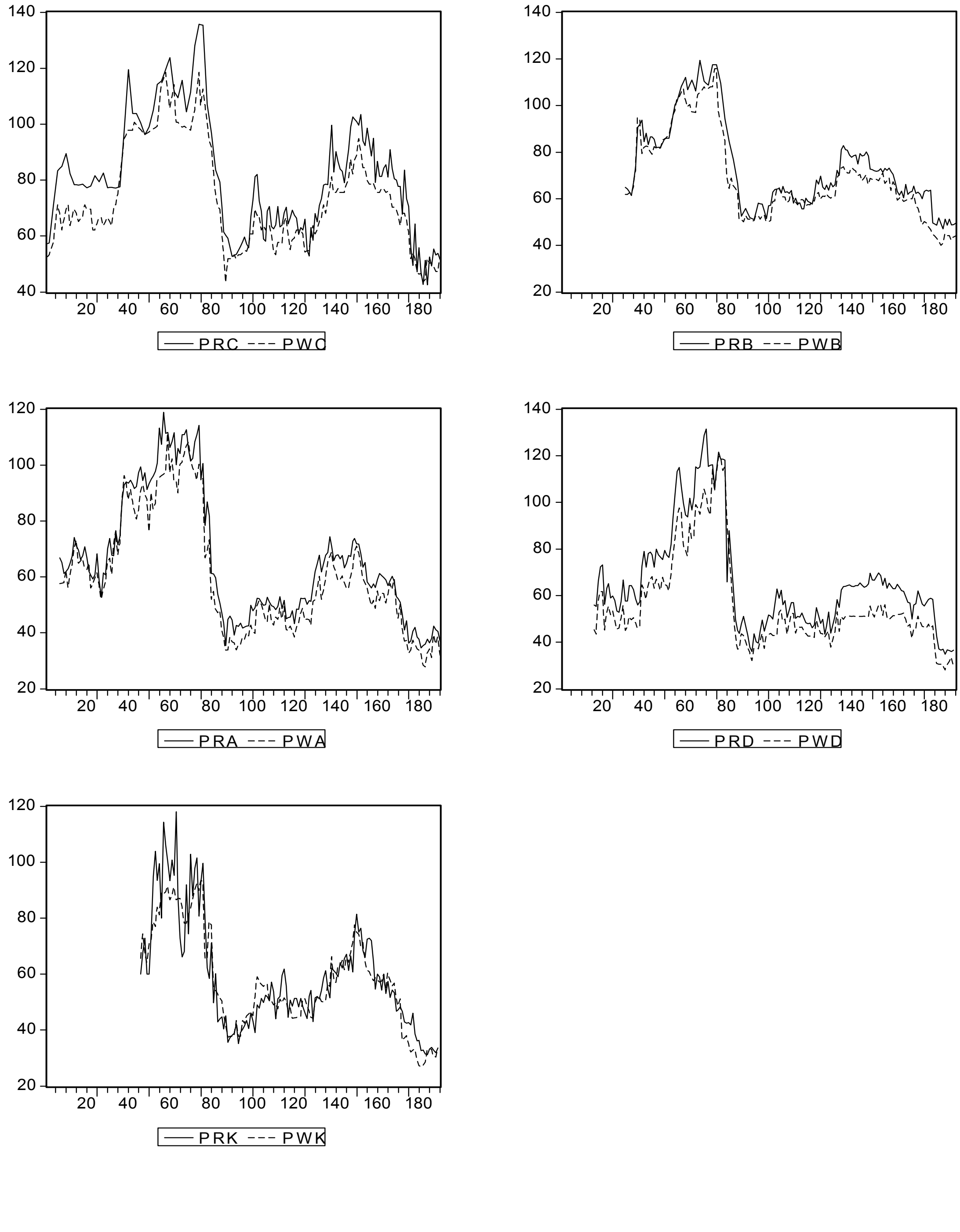

Figure 1 Prices in Fcfa per kilogram per day. Period between each daily observation is three days (markets are held in a four-day cycle). Observation 1 represents date 9/4/87; observation 20 is date 11/19/87; observation 40 is date 2/7/88; observation 60 is date 4/27/88; observation 80 is date 7/16/88; observation 100 is date 10/4/88; observation 120 is date 12/23/88; observation 140 is date 3/13/89; observation 160 is date 6/1/89; observation 180 is

date 8/20/89. PRC: retail price Cotonou; PWC: wholesale price Cotonou; PRB: retail price Bohicon; PWB: wholesale price Bohicon; PRA: retail price Azové; PWA: wholesale price Azové; PRD: retail price Dassa; PWD: wholesale price Dassa; PRK: retail price Kétou; PWK: wholesale price Kétou.

Table 1 Minimum *t* values for Model (A) with $\bar{k} = 8$ obtained by applying the procedure of Zivot and Andrews (1992) to test for *I*(1) against the alternative of single structural breakpoint *I*(0).

Series	minimum *t* value	date of breakpoint
retail price Cotonou	–3.42	6/30/88
wholesale price Cotonou	–3.69	6/30/88
retail price Bohicon	–3.96	7/04/88
wholesale price Bohicon	–4.83*	6/26/88
retail price Azové	–3.97	6/22/88
wholesale price Azové	–4.17	6/30/88
retail price Dassa	–4.30	7/12/88
wholesale price Dassa	–4.07	7/12/88
retail price Kétou	–4.24	6/30/88
wholesale price Kétou	–4.04	6/30/88

* and ** indicate that the *I*(1) hypothesis is rejected at the 0.05 and 0.01 levels, respectively. Critical *t* values are –4.80 (0.05 level) and –5.34 (0.01 level), see Zivot and Andrews (1992, Table 2).

ant to the choice of the variable selected for normalization (Hamilton, 1994). In our presentation we chose the retail price to be the left-hand variable of the cointegrating relationship (e_{rt} is the disequilibrium error, see Table 3). For each market place we found the parameter of the wholesale price to be highly significant (using asymptotic *t* values). If we took the wholesale price as the left-hand variable and estimated the parameter of the retail price, we found all parameters to be significant as well. Consequently, both prices must be *I*(1) and their relationship is a real cointegrating relationship, complying with our assumption that x_t and c_{rt} are stationary.

Table 2 Order determination of bivariate VARs consisting of the retail price series and the wholesale price series of the corresponding market.

Market	AIC at *k* =										
	1	2	3	4	5	6	7	8	9	10	11
Cotonou	11.63	11.59*	11.60	11.62	11.59	11.62	11.64	11.67	11.71	11.75	11.77
Bohicon	10.08	10.08	10.04*	10.09	10.14	10.15	10.17	10.20	10.22	10.28	10.32
Azové	11.48	11.48	11.46	11.45*	11.46	11.50	11.52	11.51	11.53	11.55	11.58
Dassa	11.96	11.93	11.92*	11.96	11.97	11.99	12.02	12.01	12.02	12.05	12.04
Kétou	12.51	12.50	12.48	12.48	12.42*	12.48	12.52	12.56	12.61	12.58	12.57

* indicates lowest value. AIC is Akaike information criterion. Maximum lag length of each VAR is set at 11.

From the cointegrating relationships in Table 3 we can also deduce that the parameter of the wholesale price is significantly greater than one in all markets except Kétou. In Kétou the parameter can be restricted to one, confirming our observation that most consumers in Kétou buy directly from the farmers or are farmers themselves. In the other markets, retailers' long-run margin behavior is characterized by charging a percentage mark-up (cf. Von Ungern-Sternberg, 1994).

Table 3 Testing for cointegration among the retail price series and wholesale price series of the corresponding market.

Market	Intercept	$r \leq$	Trace	Critical Values 5%	1%	Cointegrating Relationship (standard error in parentheses)
Cotonou	absent in	0	30.31	12.32	16.36	$p_{rt} = 1.11\, p_{wt} + e_{rt}$
	VECM	1	0.20	4.13	6.94	(0.01)
Bohicon	absent in	0	24.69	12.32	16.36	$p_{rt} = 1.07\, p_{wt} + e_{rt}$
	VECM	1	0.58	4.13	6.94	(0.01)
Azové	absent in	0	29.45	12.32	16.36	$p_{rt} = 1.09\, p_{wt} + e_{rt}$
	VECM	1	0.61	4.13	6.94	(0.01)
Dassa	absent in	0	22.43	12.32	16.36	$p_{rt} = 1.17\, p_{wt} + e_{rt}$
	VECM	1	0.85	4.13	6.94	(0.02)
Kétou	absent in	0	13.74	12.32	16.36	$p_{rt} = 1.039\, p_{wt} + e_{rt}$
	VECM	1	1.03	4.13	6.94	(0.022)

The critical values are obtained from MacKinnon et al. (1999, Case I, $k = 0$), see also Osterwald-Lenum (1992, Table 0). e_{rt} is the residual of the cointegrating relationship.

Based on the long-run parameter estimates presented in Table 3, we estimated the short-run parameters α and Γ_j (this time, including an intercept). The α parameters, which can be interpreted as adjustment parameters, are of particular interest, because they were used to test our models. The estimates of the adjustment parameters are presented in Table 4 and appear to be in favor of our hypotheses, in particular if one compares the t values with the Dickey-Fuller critical values (cf. Schotman, 1989) which, in absolute terms, are larger than the critical values of the standard t distribution so that, by way of approximation, we applied a one-sided t test at the 0.01 level instead of the 0.05 one. *Model* 2, implying that retailers dominate the wholesalers and are able to exercise some monopolistic behavior, applies to Cotonou, Bohicon and Kétou, because p_{rt} is error correcting (α_r is significant and lies within -2 and 0) and p_{wt} is not (α_w is insignificant). The results for Azové and Dassa according to which p_{wt} is error correcting (α_w is significant and lies within 0 and 2) and p_{rt} is not (α_r is insignificant), comply with *Model* 3, indicating that there is a direct link between the retail price and the farm gate price, while the wholesalers are able to be involved in market arbitrage, leaving them some leeway to influence wholesale prices.

5. Conclusions

Recall that we proposed a method for empirically testing whether or not wholesalers have some price setting power vis-à-vis the retailers. The method was applied to three models that were considered as possible candidates for describing the vertical price relationships in the marketing channels of local maize markets in Benin. A salient feature of our method is that the common stochastic trend and the deviations from the long-run vertical price equilibrium must be assigned to the variables in each model being considered. Doing this for the application in this paper, we found that the exclusion restrictions on the error-correction structure led to testable implications discriminating between the three models.

Table 4 Testing for long-run causality between the retail price series and the wholesale price series of the corresponding market.

Market	Effective sample size	Estimate α_r	t value	Estimate α_w	t value
Cotonou	188	–0.30*	–4.85	0.09	1.69
Bohicon	157	–0.33*	–4.90	–0.07	–0.82
Azové	180	–0.16	–1.29	0.39*	3.04
Dassa	164	0.20	1.96	0.34*	4.26
Kétou	134	–0.37*	–2.77	0.07	0.88

* indicates significantly different from zero (one-sided test at the 0.01 level). α_r is the coefficient of $e_{r,\,t-1}$, that is, the error correction term (see Table 3), in the equation for Δp_{rt} and α_w is the coefficient of $e_{r,\,t-1}$ in the equation for Δp_{wt}.

As far as our limited evidence goes, we conclude that retailers do not allow wholesalers to behave as vertical price leaders in the sense of Stackelberg leadership, unless the wholesalers are involved in market arbitrage. In fact, in the towns wholesalers do not have alternative market opportunities and retailers dominate the local market price formation process. In Kétou, the few retailers that exist seem to be able to exploit some opportunities for monopolistic competition. In Dassa and Azové wholesalers dominate: retail prices are stuck to the stochastic trend, while wholesalers have alternative arbitrage opportunities, giving them some freedom to influence prices.

Our empirical results indicate that relations between wholesalers and retailers vary between market places. In contrast to common assumption, retailers play a crucial role in the price formation process. Local market conditions are decisive for the distribution of market power among retailers and wholesalers. Consequently, the statement that ‘the retail market segment is dominated by the wholesale segment’ needs to be tested before it is imposed as an assumption on a model.

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