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Maize Market Liberalisation in Benin: A Case of Hysteresis

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This article analyses the effect of 10 years of 'liberalisation' policy on price integration in the Beninese maize market. The comparison of price series for two periods, before *and* after *the policy change, shows that the expected positive effect is not confirmed by co-integration analysis. Though markets were, and are, integrated in the long run, the sluggish speed of adjustment has not improved, which is why observed price differences between market places are often larger than marketing costs can justify. We conclude that the liberalisation policies did not significantly affect maize market integration and that, therefore, more effective policy instruments are required in order to strengthen the competitive forces in the market.*

JEL classification: *R15, C32*

1. Introduction

During the last decade, extensive economic reforms have been undertaken in the agricultural sector in many sub-Saharan countries. Although some of the reforms have had a positive effect, generally the results have fallen short of expectations, and much remains to be done (Kherallah *et al.*, 2000 and 2002). In most countries, very little is known about the performance of the staple food market since the policy change. There are exceptions: Rashid (2004) deals with the performance of the maize market in Uganda and compares market integration for two periods after the reform, and

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Badiane (2000) compares the situation before and after the reform but gives only a global overview of the policy change and its impact on food markets for several countries. Having data, at our disposal, about the performance of the maize market in Benin just before the policy change (Lutz, 1994), we decided to carry out an additional survey (Adegbidi *et al.*, 2003) in order to be able to do a before and after comparison of market integration.

Since 1991, the Government of Benin has been implementing a structural adjustment programme intended to dramatically change the organisation of most economic sectors. The reforms in the food market involved the official liberalisation of marketing activities, the restructuring of the cereal marketing board (Office National d'Appui à la Sécurité Alimentaire, ONASA) and the establishment of a market information system (Badiane, 2000). This policy was expected to make the market more transparent, to strengthen competition and to improve market integration. Despite the policy change, the results of our recent survey indicate that no fundamental changes in the organisation of the maize market have emerged (Adegbidi *et al.*, 2003). In particular, the set of formal and informal rules and regulations co-ordinating exchange in the market places has hardly changed.

The reforms in the maize market in Benin have been mainly intended to encourage further development of the private market, whereas in most other African countries the policy change has focused on the removal of state interventions (Kherallah *et al.*, 2002). In Benin, the cereal marketing board only played a role at the fringe of the market, owing to budgetary and organisational problems (Badiane, 2000). Consequently, one might argue that the policy change in Benin was not dramatic, and that the reform could not be expected to bring about major changes. However, this belies the objectives the government set when launching the policy change. At that time, serious market imperfections were identified and it was expected that the new policies would bring about major improvements in market performance. Moreover, an important aspect of the policy change was the aim to create a more favourable institutional environment for entrepreneurs. This would indeed be a dramatic change, as in our study of the market before the reform, we concluded that in certain segments in the market, ambiguous government regulations were creating major obstacles to entry and discouraging competition (Lutz, 1994; Kuiper *et al.*, 1999).

Thorbecke (2000) made an important distinction between two types of liberalisation policies: those that focus on getting the prices right (devaluation, abolishment of tariffs, elimination of state commodity boards) and those that have a wider scope. The latter include trade liberalisation policies, together with a constellation of structural measures designed to improve the functioning of markets. The latter policies could be described as getting the prices and institutions right: 'a comprehensive and joint package of policy measures addressing institutions and other initial conditions to relax the constraining effects of binding elements is a sine qua non for the successful performance (...) of cash crops as well as of staple food markets. Pushing the price button in a setting where one or more complementary measures or conditions are absent is fruitless' (Thorbecke, 2000, p. 40).

We are particularly interested in ascertaining whether the degree of market integration has changed as a result of the liberalisation policy.[1] Ten years ago, we concluded that there were some major deficiencies and imperfections in the maize market. Although it was confirmed that markets were integrated in the long run, price adaptations were proven to be sluggish in the short term (Lutz *et al.*, 1995). This raises the question of whether evidence for improvement can be found in the present price series. Put differently, do the price series indicate that there has been a fundamental change in maize market price integration, or do they indicate that the policy measures have failed to reduce price sluggishness?[2]

The general hypothesis is in line with the assumed effects of the liberal policy, and postulates that market integration has improved during the last decade. The measures used to verify this hypothesis focus on two indicators: the long-run co-integrating vectors and the speed of adjustment (Rashid, 2004; Abdulai, 2000; Kuiper *et al.*, 1999). The contribution of this article is that we are able to test for changes in market integration for the same set of markets,

[1] Market integration refers to the co-movement of prices as a result of arbitrage. Although market integration is not synonymous for market efficiency, it may serve as an important indicator of market performance (see Section 3).

[2] The referees commented that it would be interesting to compare the seasonal price patterns and price volatility for the two periods under study. These are certainly important elements in the assessment of market performance. Regrettably, the length of the available price series (2 and 3 years) precludes such an analysis. As seasonal supply conditions differ between the years, a longer time series would be needed to analyse these features properly.

through a comparison of price series before and after the reform. Moreover, we perform an impulse response analysis, in order to shed some light on the short-run adjustment process.

The article is structured as follows. Section 2 portrays some features of the maize market and the changes in policy. Section 3 discusses the strengths and weaknesses of the co-integration approach and its appropriateness for analysing price adjustment processes in the Beninese maize market. Section 4 describes the method used for the price series analysis. The findings and conclusions are presented in Sections 5 and 6, respectively.

2. The maize market and the liberalisation policy

In general, Benin is self-sufficient in maize. As a result of different climatic conditions, there are important differences between the south and north of the country: in the south, maize is the staple food and is grown during two cropping seasons per year, but in the north it is a cash crop, grown in only one cropping season.

The distribution of maize as a commodity is regulated by a private market system, which is integrated into a larger international maize market. Traders operate within a spatial network of both *formal* (periodic spot markets) and *informal* market places.[3] Numerous petty traders and wholesalers are involved in the business. Some traders handle large volumes of maize to supply the urban centres and the feed industry. They are only incidentally involved in export transactions. However, because of seasonality in supply and the small-scale production systems, most wholesalers handle relatively small volumes (1000 kg per market day, see Lutz, 1994 and Adegbidi *et al.*, 2003).

The marketing costs involved in the transactions carried out between two formal market places are indicated in Table 1. The most regular type of transaction is taken into account: the wholesaler transfers 1000 kg per market day, pays all the formal taxes and generally involves a local assembler in the rural market and a broker in the urban market. The latter intermediary deals with the retailers. The figures confirm that after the devaluation in 1994, the nominal marketing costs increased significantly as a result of the high inflation rates immediately after the shock of devaluation.

[3] Informal markets have no official form of organisation, whereas formal markets are subject to regulations enforced by an official market authority.

Table 1: *Breakdown of Marketing Costs in Fcfa/kg: Azové–Cotonou (Short Distance: 144 km) and Nikki–Cotonou (Long Distance: 529 km)*

	Azove–Cotonou		Nikki–Cotonou	
	1987–9	1998–2001	1987–9	1998–2001
Wholesale trade				
Market tax	3.0	2.0	2.8	3.5
Freight (bag)	7.7	9.5	11.5	13.8
Transport trader	1.0	1.8	5.5	9.5
Assembler	1.2	–	–	2.0
Broker	2.9	3.9	2.9	3.9
Others	1.2	4.5	1.2	3.2
Gross margin	5.0	10.0	5.0	10.0
Retail trade				
Market tax	1.0	1.0	1.0	1.0
Gross margin	5.0	10.0	5.0	10.0
Total cost	28.0	42.7	34.9	56.9

Note: Wholesalers buy and sell approximately 1000 kg on the regional market per market day. In most market places they generally buy with the help of an assembler and sell with the mediation of a broker. Traders' gross margins (according to their own estimation) varied as a rule from 0 to 10 Fcfa per kg in the period 1987–9. Consequently, we imputed a gross margin of 5 Fcfa per kilogram as a normal mark-up. According to the traders, during the period 1998–2001 the margins doubled.
Source: Adegbidi *et al.*, 2003.

However, whereas the price of foreign exchange doubled, the marketing costs increased more slowly, because of a reduction in taxes and real freight costs.

The marketing system has to be flexible in order to accommodate seasonality in production. However, in general, in Benin, the trade relations between the rural centres with a maize surplus and the major towns are quite stable, as the urban deficits have to be continuously made good by the same set of rural areas with maize surpluses. The major maize granaries are located in the north of the country (Nikki): maize is a cash crop for most of the farmers here, and storage conditions are favourable. Table 2 confirms these relationships. The prices in the rural

Table 2: *Price Differences Between Rural and Urban Centres (Fcfa/kg)*

	Urban centers					
	Cotonou		Bohicon		Parakou	
	1987–9	1998–2001	1987–9	1998–2001	1987–9	1998–2001
Rural centers						
Azové						
Number of observations	140	282	140	282	140	282
Average $(P_i - P_j) > 0$	18.3	58.1	11.5	21.3	20.1	32.9
Average $\|(P_i - P_j) < 0\|$	3.5	–	5.9	6.5	13.1	3.4
$(P_i - P_j) < 0$ (%)	2	0	10	6	69	5
changes in sign	4	0	7	10	8	20
$\|P_i - P_j\| >$ marketing costs (%)	32	99	11	32	18	36
$\|P_i - P_j\| >$ transport costs (%)	88	100	86	95	51	71
Glazoué						
Number of observations	140	282	140	282	140	282
Average $(P_i - P_j) > 0$	15.9	53.6	9.7	17.6	15.1	29.4
Average $\|(P_i - P_j) < 0\|$	16.7	–	12.6	11.3	15.1	8.1
$(P_i - P_j) < 0$ (%)	9	0	21	8	75	8
changes in sign	16	0	16	17	11	12
$\|P_i - P_j\| >$ marketing costs (%)	18	91	7	23	34	44
$\|P_i - P_j\| >$ transport costs (%)	81	100	85	91	74	89
Kétou						
Number of observations	140	282	140	282	140	282
Average $(P_i - P_j) > 0$	22.6	51.4	16.1	24.4	18	27.6
Average $\|(P_i - P_j) < 0\|$	7.5	7	5.7	30	10.4	12.2
$(P_i - P_j) < 0$ (%)	1	1	8	22	57	10
changes in sign	2	4	12	13	26	10
$\|P_i - P_j\| >$ marketing costs (%)	43	86	21	33	13	19
$\|P_i - P_j\| >$ transport costs (%)	94	95	86	88	34	73

(*continued on next page*)

Table 2 (*continued*)

	Urban centers					
	Cotonou		Bohicon		Parakou	
	1987–9	1998–2001	1987–9	1998–2001	1987–9	1998–2001
Nikki						
Number of observations	140	282	140	282	140	282
Average $(P_i - P_j) > 0$	34.5	62.2	25.7	28	11.4	34.3
Average $\lvert(P_i - P_j) < 0\rvert$	3.7	9.8	6.3	15.3	3.1	15.2
$(P_i - P_j) < 0$ (%)	9	2	8	14	6	2
changes in sign	6	2	2	11	3	6
$\lvert P_i - P_j\rvert >$ marketing costs (%)	57	68	51	31	13	58
$\lvert P_i - P_j\rvert >$ transport costs (%)	83	98	85	71	91	98

Note: Prices are given in Fcfa/kg. 'P_i', price in urban centre and 'P_j', price in rural centre. 'Average $(P_i - P_j) > 0$' refers to the average value of positive price differences. 'Average $|(P_i - P_j) < 0|$' refers to the average value of negative price differences. '$(P_i - P_j) < 0$' indicates the number of negative price differences as a percentage of the total number of observations. Changes in sign calculates the number of times the price differential between two markets changes sign in the price series. '$|P_i - P_j| >$ marketing costs' calculates the number (percentage) of price differences that exceed the calculated marketing costs (see Table 1). '$|P_i - P_j| >$ transport costs' calculates the number (percentage) of price differences that exceed the freight costs (see Table 1). It is beyond the scope of this article to provide the details of this calculation. For further information, see Lutz (1994) and Adegbidi *et al.* (2003).
Source: Price series ONASA (1998, 1999, 2000, 2001) and data collected by the authors (1987–9).

markets are generally lower than those in the urban centres. The fact that the average price difference is appreciably larger than the direct freight costs indicates that trade relationships can be worthwhile. Only Parakou, a major urban centre in the north with excellent transport links with Nikki, shows a somewhat atypical pattern in the period 1987–9. At that time, the northern region was linked to the south by an unpaved road (300 km).

During the 1990s, major improvements in the road network caused transport costs to fall and therefore trade relationships between these regions intensified. Though the price differences vary, they do not allow for trade reversals except in some extraordinary circumstances. Any negative price differences observed are generally smaller than the transport costs.

There are alternative channels for the Beninese maize trade. Some traders integrate the function of the local assembler, the broker and/or retailer, whereas others prefer to operate on informal markets. This implies that even if the price differences are less than the calculated costs of marketing between formal market places (Table 1), some trade will be feasible through these alternative channels. Consequently, market integration may be sustained as long as the price differential exceeds the direct transport costs. Table 2 shows that for nearly all market pairs, the number of price differences smaller than transport costs is quite small, which indicates that trade relationships between markets are fairly continuous. The only relationship that may have been weak in the period before the policy change is the one between the rural markets in the south and Parakou.

We observed serious deficiencies and imperfections in the organisation of the maize market before the structural reforms were begun (Lutz, 1994; Kuiper *et al.*, 1999, 2003):

- Lack of information on market opportunities;
- Non-transparent enforcement of formal regulations;
- Entry barriers for non-residents or for persons not in local informal traders' organisations;
- Most transactions involve small volumes (thin markets) and take place in spot markets;
- Farmers are only passively involved in commercial activities in the marketing channel: farmers' organisations are not active in the food market and, therefore, only a minority of large-scale farmers are able to develop more profitable commercial strategies.

These problems inflate the costs of market exchange and have been used as an argument to explain the sluggish price adjustment process, in particular between markets that are far apart geographically. This is an important problem for Benin, as the major maize surplus regions are in the north, about 400 km from the towns in the south.

The general measures taken by the Benin government in the framework of the structural adjustment policy concern macroeconomic stabilisation measures: reduction of budget deficits, liberalisation of foreign exchange and privatisation and deregulation. A plan was made specifically for the agricultural sector, focusing on the reorganisation of the cotton industry. Some of the major institutional changes announced influenced the operations in the food market (Declaration of the Rural Development Policy in 1991, see Adegbidi *et al.*, 2003, p. 24):

- Construction and maintenance of feeder roads;
- Simplification of customs procedures and more systematic road checks;
- Improvement of the financing instruments and techniques that favour storage at the farm level and at all other levels in the food marketing channel;
- Regional integration of agricultural markets and their possible protection against subsidised imports;
- Establishment of a support system for exports and local traders and, in particular, a market information system concerning prices and international and domestic market conditions;
- A national policy to promote small- and medium-sized enterprises that are involved in the processing of agricultural commodities or supply of agricultural services to farmers.

As a consequence of these intentions, several policy measures were taken, the most dramatic of which were the restructuring of the national grain board (ONASA) and other state services (Badiane, 2000; Adegbidi *et al.*, 2003). However, the impact of the announced institutional changes has been questioned (Adegbidi *et al.*, 2003).

A rough indicator of the problem under study is provided in Table 2, which compares prices in two market places and presents the number of times the observed price difference exceeded the estimated marketing costs. From the table, it seems that the market in Bohicon is relatively well integrated, although in the period after the policy change, with one exception, the price differences vis-à-vis the rural centres in the hinterland (Azové and Glazoué) exceed the calculated marketing costs more frequently. The exception is the price difference compared with Nikki, which was more favourable in 1998–2001. The results for Parakou display a similar pattern although now also the relationship with Nikki, where the nearest major granary is situated, weakened.

The position of Cotonou, the country's major urban centre, is worrying. After the reforms, all the granaries in the study seem to have loosened their dependence on its market. From these results, it is clear that it would be worthwhile to analyse the process of price adjustment between market places in Benin.

3. Appropriateness of co-integration tests

Rashid (2004) discusses the importance of market integration for successful market reforms, and the strengths and weaknesses of the co-integration method. One of the weaknesses is that the method assumes continuous trade relations but that it cannot discern whether trade occurs between any two locations: 'suppose that two surplus markets, A and B, do not engage in trade because the price differential between them is less than transfer costs, but both markets supply a major urban location, C, with which price differentials are large enough to cover the transfer costs. Now, if price shocks in C are transmitted to A and B, all three markets are integrated. In this situation, the co-integration method will make the right diagnosis, but will fail to detect that there was no trade between A and B' (ibid., p. 110).

Several authors have criticised the use of co-integration tests as indicators for market efficiency. Rashid (2004) discusses the critique and focuses on two crucial assumptions: (i) the transaction costs are stationary (Barrett, 1996; McNew and Fackler, 1997; Barrett and Li, 2002) and (ii) there are no significant reversals or discontinuous trade flows across markets (McNew, 1996; Baulch, 1997a,b; Park *et al.*, 2002; Araujo *et al.*, 2005).

Barrett (1996) showed that a failure to observe co-integration may be rooted in non-stationary marketing costs. If the data are not deflated, this phenomenon will affect the test results — particularly in countries with high inflation rates. In Section 2, we noted that the price series we have examined are for two periods with low rates of inflation. The marketing costs between the two periods differ, because of the price shock following the devaluation of the Franc CFA in 1994. However, the consequences of this shock were absorbed in the price series during the first few years after the change (Badiane, 2000). Moreover, traders reported that transport costs were stable during the periods under study (Adegbidi *et al.*, 2003).

The second assumption — that there are no significant reversals or discontinuous trade flows across markets — is more relevant for Benin, as harvest seasons differ between the south and the north. The disparity in harvest seasons implies that bivariate long-run price relationships are non-linear and, subsequently, trade does not link all markets continuously (McNew, 1996). If this were true, straightforward co-integration analysis would be inappropriate, as it does not distinguish between periods characterised by different arbitrage conditions (autarky, efficient arbitrage, arbitrage failure). One could argue that this problem is less pertinent for Benin, as trade relationships with southern consumer centres are expected to be reasonably continuous: the north is the region with the main surpluses and with favourable storage conditions, whereas the large cities (permanent deficit regions) are located in the south. Moreover, although maize is the staple food in the south, it is of minor importance in the diets of farmers in the north. These farmers consider maize to be a cash crop and consequently are actively involved in spatial and temporal arbitrage. All this implies that ruptures in trade flows are not expected to persist for long.

In order to verify the argument presented earlier, it would be helpful to have data on trade flows to complement the information on price series (Baulch, 1997b). Regrettably, no such market data are available for the time series under study.[4] However, the price data give some information about trade opportunities and expected trade flows. For example, one could test whether price differences allow for trade reversals and discontinuous uni-directional trade flows. In Section 2, we showed that these rarely occur between some market pairs and certainly never occur for sustained periods of time. In the appendix of the next section, we apply a more formal procedure to check for linearity. The test verifies whether a three-regime band-threshold vector error-correction model (VECM) represents the relationship better than a linear VECM (i.e., the

[4] The literature provides some indications that there are seasonal trade flows (Gabre-Madhin *et al.*, 2001 and Lutz, 1994). We do agree that the flows reflect seasonality in a sense that the north transfers the largest volumes of maize during the lean season in the south. After the lean season, trade flows between the south and the north dwindle, but generally do not dry up completely because alternative channels (different types of petty traders) continue to operate (see also Section 2).

co-integration method). A three-regime band-threshold VECM distinguishes three possibilities for trade between two markets: $P_1 - P_2 > C$ (trade flows are expected to exist because price differences exceed marketing costs, where P_1 and P_2 are the prices of a homogeneous good traded on the spatially dispersed markets 1 and 2, respectively, and C are the marketing costs involved in trade between the two markets), $|P_1 - P_2| \leq C$ (no direct trade flows are expected, as price differences are lower than marketing costs) and $P_2 - P_1 > C$ (reversal in trade flows). A linear VECM exists only if the first or third option holds. It appears that the test results justify the assumption that trade relationships are continuous (linear) and reject the alternative of non-linearity of the band-threshold type. The empirical results of this procedure, together with the more qualitative information presented in Section 2, lead us to conclude that trade relationships between the rural surplus areas and the urban deficit areas are continuous and that we are justified in using the co-integration method for this study.

4. Method

This study analyses the price time series observed in seven market places, well distributed over the major maize consumption and/or production regions: three urban centres (Parakou, Bohicon and Cotonou) and four rural centres (Kétou, Glazoué, Azové and Nikki). In this article, we compare the period 1987–9 with the situation 10 years later as represented by the period 1998–2001 (see Table 2 for further details).[5]

Consider the time series of prices of a homogeneous commodity traded on p ($p > 1$) spatially dispersed markets. Let $\mathbf{P}_t$ be the ($p \times 1$) vector of these prices established at time t. For this moment, we assume that the prices have stochastic trends and, as a consequence of continuous market arbitrage at stationary marketing costs, these stochastic trends reduce to one common stochastic trend. In the appendix, we outline an econometric procedure to check for the presence of band-threshold effects allowing for trade reversals and periods of autarky.

[5] Most price series concern periodic markets with a cycle of 4 days. Two markets, Parakou and Glazoué, have a cycle of 7 days. In order to make these series compatible, we transformed them into a 4-day cycle by taking the last observed price as a proxy.

The simplest VECM that complies with our assumptions given earlier is:

$$\boldsymbol{\Delta}\mathbf{P}_t = \boldsymbol{\alpha\beta}'\mathbf{P}_{t-1} + \boldsymbol{\varepsilon}_t \quad (t = 1, \ldots, T; \quad \boldsymbol{\varepsilon}_0 = \mathbf{0}_p), \tag{1}$$

with fixed starting prices $\mathbf{P}_0$, $\boldsymbol{\Delta}\mathbf{P}_t = \mathbf{P}_t - \mathbf{P}_{t-1}$, $\boldsymbol{\alpha}$ is the $(p \times (p-1))$ matrix of adjustment parameters, $\boldsymbol{\beta}$ is a $(p \times (p-1))$ matrix containing the $(p-1)$ co-integrating vectors, $\boldsymbol{\varepsilon}_t$ is Gaussian white noise with covariance $\boldsymbol{\Omega}$, and $\mathbf{0}_p$ is a $(p \times 1)$ zero vector. The long-run model is identified exactly by specifying $\boldsymbol{\beta} = (\mathbf{H}_1\boldsymbol{\varphi}_1, \ldots, \mathbf{H}_{p-1}\boldsymbol{\varphi}_{p-1})$, where $\mathbf{H}_i$ $(i = 1, \ldots, p-1)$ is a $(p \times 2)$ matrix with elements $(i, 1)$ and $(p, 2)$ equal to one and all other elements equal to zero, and $\boldsymbol{\varphi}_i$ is a (2×1) vector. Using the relation (cf. Johansen, 1995, p. 39):

$$\boldsymbol{\beta}_\perp(\boldsymbol{\alpha}'_\perp\boldsymbol{\beta}_\perp)^{-1}\boldsymbol{\alpha}'_\perp + \boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}\boldsymbol{\beta}' = \mathbf{I}_p, \tag{2}$$

where $\boldsymbol{\alpha}_\perp$ and $\boldsymbol{\beta}_\perp$ are $(p \times 1)$ vectors such that $\boldsymbol{\alpha}'_\perp\boldsymbol{\alpha} = \boldsymbol{\beta}'_\perp\boldsymbol{\beta} = \mathbf{0}_{(1\times(p-1))}$ (cf. Johansen, 1995, p. 48) and $\mathbf{I}_p$ is a $(p \times p)$ identity matrix, we can derive from equation (1) that:

$$\begin{aligned}\mathbf{P}_t &= \boldsymbol{\beta}_\perp(\boldsymbol{\alpha}'_\perp\boldsymbol{\beta}_\perp)^{-1}\boldsymbol{\alpha}'_\perp\ \mathbf{P}_t + \boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}\boldsymbol{\beta}'\mathbf{P}_t \\ &= \boldsymbol{\beta}_\perp(\boldsymbol{\alpha}'_\perp\boldsymbol{\beta}_\perp)^{-1}\left(\boldsymbol{\alpha}'_\perp\mathbf{P}_0 + \boldsymbol{\alpha}'_\perp\sum_{i=0}^{t-1}\boldsymbol{\varepsilon}_{t-i}\right) \\ &\quad + \boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}\sum_{i=0}^{t-1}(\ \mathbf{I}_{p-1} + \boldsymbol{\beta}'\boldsymbol{\alpha})^i\boldsymbol{\beta}'\boldsymbol{\varepsilon}_{t-i},\end{aligned} \tag{3}$$

which is the common-trend representation (Stock and Watson, 1988). The common stochastic trend $(\boldsymbol{\alpha}'_\perp\mathbf{P}_t = \boldsymbol{\alpha}'_\perp\mathbf{P}_0 + \boldsymbol{\alpha}'_\perp\sum_{i=0}^{t-1}\boldsymbol{\varepsilon}_{t-i})$ is proportional to the permanent components $\boldsymbol{\beta}_\perp(\boldsymbol{\alpha}'_\perp\boldsymbol{\beta}_\perp)^{-1}$ $\boldsymbol{\alpha}'_\perp\mathbf{P}_t$ in the prices $\mathbf{P}_t$. The short-run dynamics $(\boldsymbol{\beta}'\mathbf{P}_t = \sum_{i=0}^{t-1}(\mathbf{I}_{p-1} + \boldsymbol{\beta}'\boldsymbol{\alpha})^i\boldsymbol{\beta}'\boldsymbol{\varepsilon}_{t-i})$ form the transitory components $\boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}\boldsymbol{\beta}'\mathbf{P}_t$ in the prices $\mathbf{P}_t$. If one of the prices in $\mathbf{P}_t$, say, price i $(i = 1, \ldots, p)$, does not display error-correcting behaviour, then this price is the common stochastic trend and hence $\boldsymbol{\alpha}_\perp$ is a zero vector, except for element i, which can be shown to be equal to one. However, if all prices are error-correcting, then it is of interest to compute impulse responses. The impulse response is the adjustment path a price follows as a consequence of a shock to the common stochastic trend. This shock could be, for example, the largest absolute value of $\boldsymbol{\alpha}'_\perp\boldsymbol{\varepsilon}_t$. Given that $\boldsymbol{\alpha}'_\perp\boldsymbol{\varepsilon}_T$ is not assigned decaying weights when h

($h = 1, 2, \ldots$) in $T + h$ increases, a shock in the common stochastic trend can also be considered to be a permanent shock in the prices. The short-run dynamics $\boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1} \sum_{i=0}^{t-1} (\mathbf{I}_{p-1} + \boldsymbol{\beta}'\boldsymbol{\alpha})^{i}\boldsymbol{\beta}'\boldsymbol{\varepsilon}_{t-i}$ are responsible for the fact that prices adjust gradually rather than immediately to the new long-run equilibrium created by the permanent shock. Recall, however, that if price i does not display error correction, then the ith row of $\boldsymbol{\alpha}$ is zero (and hence, $\boldsymbol{\alpha}_{\perp}$ is a zero vector, except for element i) showing that this price, which, in fact, is the common stochastic trend itself, does adjust instantly to the new long-run equilibrium.

The question now is how are the short-run dynamics triggered by the shock in the common stochastic trend? This analysis of how rapidly prices adjust towards their new equilibrium aims to reveal whether the maize markets in Benin show stronger price integration. In this respect, we follow Kuiper *et al.* (1999, p. 722) in assuming that the permanent shock is exogenous to the transitory shocks. If the homogeneous commodity under study is a crop being grown at certain places in the region, then the exogeneity assumption implies that an unexpected change in harvest is not caused by unexpected changes in transaction or transportation costs. Therefore, we use the simple linear relationships

$$\boldsymbol{\beta}'\boldsymbol{\varepsilon}_t = \boldsymbol{\delta}\boldsymbol{\alpha}'_{\perp}\boldsymbol{\varepsilon}_t + \mathbf{v}_t, \tag{4}$$

where $\boldsymbol{\delta}$ is a $((p-1)\times 1)$ vector of parameters to be estimated by equation-wise OLS and $\mathbf{v}_t$ is the $((p-1)\times 1)$ vector of residuals. Next, the vectors of impulse responses $\mathbf{R}_i$ ($i = 1, 2, \ldots$) are estimated as:

$$\mathbf{R}_i = [\boldsymbol{\beta}_{\perp}(\boldsymbol{\alpha}'_{\perp}\boldsymbol{\beta}_{\perp})^{-1} + \boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}(\mathbf{I}_{p-1} + \boldsymbol{\beta}'\boldsymbol{\alpha})^{i-1}\boldsymbol{\delta}]s, \tag{5}$$

where s could be chosen to be the largest absolute value of the T permanent shocks $\boldsymbol{\alpha}'_{\perp}\boldsymbol{\varepsilon}_t$ ($t = 1, \ldots, T$). Given that the elements (i, i) ($i = 1, \ldots, p-1$) in $\boldsymbol{\beta}$ are normalised to one, s is assigned the sign of the scalar $(\boldsymbol{\alpha}'_{\perp}\boldsymbol{\beta}_{\perp})^{-1}$ in order to generate a positive shock in the permanent components. Assuming the parameters $\boldsymbol{\alpha}$, $\boldsymbol{\beta}$ (and hence, $\boldsymbol{\alpha}_{\perp}$, $\boldsymbol{\beta}_{\perp}$), and $\boldsymbol{\delta}$ are known, and, as we have already assumed by applying OLS to equation (4), that $\mathbf{v}_t$ is orthogonal to $\boldsymbol{\alpha}'_{\perp}\boldsymbol{\varepsilon}_t$ and hence to all elements of $\boldsymbol{\varepsilon}_t$ if $\boldsymbol{\alpha}_{\perp}$ does not contain elements restricted to zero, the covariance matrices of the $\mathbf{R}_i$ ($i = 1, 2, \ldots$) are computed as:

$$\begin{aligned}\mathrm{cov}(\mathbf{R}_i) = \sum_{j=1}^{i} \Big\{ & [\boldsymbol{\beta}_\perp(\boldsymbol{\alpha}'_\perp\boldsymbol{\beta}_\perp)^{-1}\boldsymbol{\alpha}'_\perp + \boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1} \\ & \times (\mathbf{I}_{p-1} + \boldsymbol{\beta}'\boldsymbol{\alpha})^{j-1}\boldsymbol{\delta\alpha}'_\perp]\left[\sum_{t=1}^{T}(\boldsymbol{\varepsilon}_t\boldsymbol{\varepsilon}'_t)/T\right] \\ & \times [\boldsymbol{\beta}_\perp(\boldsymbol{\alpha}'_\perp\boldsymbol{\beta}_\perp)^{-1}\boldsymbol{\alpha}'_\perp + \boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1} \\ & \quad \times (\mathbf{I}_{p-1} + \boldsymbol{\beta}'\boldsymbol{\alpha})^{j-1}\boldsymbol{\delta\alpha}'_\perp]' \\ & + [\boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}(\mathbf{I}_{p-1} + \boldsymbol{\beta}'\boldsymbol{\alpha})^{j-1}]\left[\sum_{t=1}^{T}(\mathbf{v}_t\mathbf{v}'_t)/T\right] \\ & \times \left[\boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}(\mathbf{I}_{p-1} + \boldsymbol{\beta}'\boldsymbol{\alpha})^{j-1}\right]'\Big\}. \end{aligned} \tag{6}$$

The diagonal elements of $\mathrm{cov}(\mathbf{R}_i)$ are the variances of the impulse responses. These can be used to compute the confidence intervals at the common 95% confidence level in the case that we consider the largest absolute value of $\boldsymbol{\alpha}'_\perp\boldsymbol{\varepsilon}_t$ as the shock in the permanent component triggering the impulse responses. The fewer periods it takes before the impulse response of a price no longer differs significantly from its new permanent level after the shock in the permanent component, the more representative the price movements of this market are for the long-run common stochastic price trend. If other markets are much slower in this respect, then arbitrage will perhaps not function well.

To compare the speed of adjustment of the market prices between two sample periods, we may compute the transitory $\boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}(\mathbf{I}_{p-1} + \boldsymbol{\beta}'\boldsymbol{\alpha})^{i-1}\boldsymbol{\delta}s$ responses as obtained from equation (5) with variances to be obtained from $\boldsymbol{\delta\alpha}'_\perp\boldsymbol{\varepsilon}_t + \mathbf{v}_t$, see equation (4), to derive the number of periods that elapse before these responses are no longer significant. To compare the sample periods, we either insert the VECM estimates for the first sample period in the $\boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}(\mathbf{I}_{p-1} + \boldsymbol{\beta}'\boldsymbol{\alpha})^{i-1}$ part of the responses, or we insert the VECM estimates obtained for the second sample period. Note that $\boldsymbol{\delta}s$ and $\boldsymbol{\delta\alpha}'_\perp\boldsymbol{\varepsilon}_t + \mathbf{v}_t$ (and hence, the variances derived from $\boldsymbol{\delta\alpha}'_\perp\boldsymbol{\varepsilon}_t + \mathbf{v}_t$) are always VECM estimates for the second sample period.

The method outlined in this section deals with a VECM derived from a VAR of order one. Given that a VAR of a higher order can always be written in VAR(1) form, extension of the method to higher-order VECMs is straightforward. See, for example, Kuiper and Meulenberg (2004), who applied the permanent-transitory composition as derived in Bruneau and Jondeau (1999). In line with Granger (1969), this method is based on a precise definition of long-run non-causality in a multivariate context, according to which a variable x is not causal for another variable y in the long run if and only if the knowledge of the past of x does not improve the long-run prediction of y. According to this composition, the permanent component to be derived from a VECM based on a VAR of order l ($l = 2, 3, \ldots$)

$$\mathbf{\Delta P}_t = \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{P}_{t-1} + \sum_{k=1}^{l-1} \Gamma_k \mathbf{\Delta P}_{t-k} + \boldsymbol{\varepsilon}_t, \tag{7}$$

is given by $\boldsymbol{\beta}_\perp[\boldsymbol{\alpha}'_\perp(\mathbf{I} - \sum_{k=1}^{l-1}\Gamma_k)\boldsymbol{\beta}_\perp]^{-1}\boldsymbol{\alpha}'_\perp(\mathbf{P}_t - \sum_{k=1}^{l-1}\Gamma_k\mathbf{P}_{t-k})$, where the Γ_k are $(p \times p)$ parameter matrices such that, in contrast to the VAR(1) case, the common factor $\boldsymbol{\alpha}'_\perp(\mathbf{P}_t - \sum_{k=1}^{l-1}\Gamma_k\mathbf{P}_{t-k})$ does not necessarily reduce to the common factor $\boldsymbol{\alpha}'_\perp\mathbf{P}_t$ that was proposed in Gonzalo and Granger (1995) and used in, among others, Rashid (2004).

5. Findings

Let $\mathbf{P}_t = (P_{\mathrm{Co},t}, P_{\mathrm{Az},t}, P_{\mathrm{Bo},t}, P_{\mathrm{Ke},t}, P_{\mathrm{Gl},t}, P_{\mathrm{Pa},t}, P_{\mathrm{Ni},t})'$ be the vector of the prices of Cotonou (Co), Azové (Az), Bohicon (Bo), Kétou (Ke), Glazoué (Gl), Parakou (Pa) and Nikki (Ni). Hence, $p = 7$. First, the system is modelled as a VAR with constant terms included. Using 11 as the upper limit for the order of the VAR, we find that the order selection criteria (Akaike Information Criterion, Final Prediction Error criterion, Schwarz Criterion, and Hannan-Quinn criterion, see, e.g., Lütkepohl, 1991, pp. 128–138) all select a VAR(1) for both VARs: the VAR for the period 1987–1989 ($T = 140$) and the VAR for the period 1998–2001 ($T = 282$). Next, we determine the number of co-integrating vectors (i.e., long-run price relationships) using the *trace* test of the Johansen procedure as described in Johansen and Juselius (1990) and adjusted in Johansen (2002) to improve its small sample performance. The number of co-integrating vectors, say, r, is given by the rank of the matrix $\boldsymbol{\alpha\beta}'$ in eqaution (1). If $r = 0$, then there are no co-integrating vectors and if $r = p$, then the prices do not have a stochastic trend. However, if the prices do have a stochastic trend, then this trend should be common to all prices in order for $r = p - 1$, implying that all pairs of prices exhibit a long-run equilibrium.

Given that $r = p - 1$, we can test whether each price has a stochastic trend by generating a Gaussian white noise series η_t and using the *trace* statistic to test for $r = 1$ in a bivariate VECM with η_t and the considered price. We give η_t the same variance as the variance of the first differences of the price under consideration. The order selection criteria AIC, FPE, SC and HQ are used again to select the order of each bivariate VAR with constant terms included, again taking 11 as the upper limit and choosing the largest order selected. The *trace* test results are presented in Table 3 and clearly show that the restriction $r = 0$ (i.e., given that $p = 2, p - r = 2$) must be rejected, whereas $r = 1$ (i.e., $p - r \geq 1$ implying that the price is non-stationary as η_t is stationary by construction) cannot be rejected. Consequently, we

Table 3: *Testing for a Stochastic Trend in the Price Series in Bivariate ($p = 2$) VARs*

	95% Critical value	Co	Az	Bo	Ke	Gl	Pa	Ni
Sample period 1987–9								
Order VAR		1	2	1	2	1	1	2
trace statistic								
$p - r \geq 1$	4.13	0.98	2.44	1.15	1.01	0.96	0.77	0.36
$p - r = 2$	12.32	101[a]	63[a]	103[a]	62[a]	102[a]	105[a]	64[a]
Sample period 1998–2001								
Order VAR		1	1	1	1	1	1	1
trace statistic								
$p - r \geq 1$	4.13	0.03	0.28	0.07	0.37	0.76	0.00	1.56
$p - r = 2$	12.32	165[a]	167[a]	166[a]	168[a]	167[a]	165[a]	166[a]

Note: Critical values are obtained from Table 2 ($k = 0$) in MacKinnon *et al.* (1999). The *trace* statistics are corrected by the Bartlett correction factor along the lines of Johansen (2002) to improve their final sample properties. The VARs do not contain any deterministic terms as selected according to the procedure in Johansen (1992b), see also Johansen (1995, p. 98–100).

[a]Significant as the trace statistic is close to or greater than the 95% critical value.

Table 4: *Cointegration Test*

$p - r$	95% Critical value	Trace statistic		
		VAR(1) for 1987–9	VAR(1) for 1998–2001	VAR(1) with-out Kétou for 1998–2001
≥1	4.13	0.75	0.28	0.13
≥2	12.32	12.31[a]	2.92	12.21[a]
≥3	24.28	28.03[a]	16.85	37.48[a]
≥4	40.17	44.05[a]	43.94[a]	74.69[a]
≥5	60.06	78.82[a]	89.42[a]	124.87[a]
≥6	83.94	143.71[a]	165.69[a]	209.44[a]
=7	111.79	220.28[a]	252.49[a]	

Note: Critical values are obtained from Table 2 ($k = 0$) in MacKinnon *et al.* (1999). The *trace* statistics are corrected by the Bartlett correction factor along the lines of Johansen (2002) to improve their final sample properties. The VARs do not contain any deterministic terms as selected according to the procedure in Johansen (1992b), see also Johansen (1995, p. 98–100).
[a]Significant as the trace statistic is close to or greater than the 95% critical value.

conclude that all prices in both periods have a stochastic trend. Testing for a unit root in this way allows us to profit from the improved finite sample properties resulting from the small sample correction derived in Johansen (2002).

Table 4 presents the results of Johansen's (2002) *trace* test. According to these results, the markets are price-integrated in the period 1987–9. In the period 1998–2001, however, we find one common stochastic trend only if the price of Kétou is omitted from the VAR(1). Not omitting any price or omitting the price of a market other than Kétou leads to more than one stochastic trend. Consequently, we conclude that at the end of the 1990s Kétou was no longer part of the price integrated market system in Benin. Of relevance here is the observation made by Adegbidi *et al.* (2003, p. 56) that local traders in Kétou conduct much trade with Nigeria.[6] Consequently, the co-integration method shows that the markets

[6] Nigeria is a huge market, which is why the local Kétou market can behave somewhat independently.

are price-integrated in both periods, a result that would be very unlikely if trade between a pair of markets exhibited flow reversals or discontinuous uni-directional flows. Therefore, the *trace* test of the hypothesis that all spatially dispersed markets have the same stochastic price trend can also be considered as a test for the absence of trade reversals and/or periods of autarky.

In the VECM for 1987–9 as well as the VECM without the price of Kétou for 1998–2001, the equation-wise *F*-statistics testing for the absence of $\boldsymbol{\beta}'\mathbf{P}_{t-1}$ are all significant, implying that none of the rows of $\boldsymbol{\alpha}$ can be put equal to zero (Johansen, 1992a). Consequently, there is no market in the price-integrated market system that can be considered to function as price leader. To see how closely all markets drive the long-run price system, we perform the impulse response analysis described in the previous section.

The impulse responses to the largest absolute value in the error term $\boldsymbol{\alpha}'_{\perp}\boldsymbol{\varepsilon}_t$ of the random-walk model for the common stochastic trend $\boldsymbol{\alpha}'_{\perp}\mathbf{P}_t = \boldsymbol{\alpha}'_{\perp}\mathbf{P}_{t-1} + \boldsymbol{\alpha}'_{\perp}\boldsymbol{\varepsilon}_t$ (as can be derived by pre-multiplying equation (1) by $\boldsymbol{\alpha}'_{\perp}$) are displayed in Figure 1 for the sample 1987–9 and in Figure 2 for the sample 1998–2001. We can use Figures 1 and 2 to compare the speed of adjustment between both samples, because we obtain quite similar results for the two samples if we divide the change in the permanent level by the maximum price level for each individual market. The percentages obtained for the 1987–9 sample and, in brackets, for the 1998–2001 sample, were: 22 (23) for Co, 20 (18) for Az, 23 (22) for Bo, 19 (–) for Ke, 16 (19) for Gl, 19 (21) for Pa and 21 (16) for Ni. In both figures, the price of Azové quickly approaches the new permanent level: in 1987–9 within one period and in 1998–2001 after one period.

Table 5 presents the number of periods (of 4 days) needed for adjustment after a shock, that is, before the price level lies within the confidence interval of the new permanent price level. It appears that the northern markets, that is, Glazoué, Parakou and Nikki, are slowest in adjusting to the new equilibrium. This applies to both of the samples studied. In the 1998–2001 sample, after the trade reforms, the urban market of Cotonou is also among the slowest adjusters, taking as many as 36 days more when compared with 1987–9. Even though Nikki has considerably improved its speed of adjustment, there is no apparent overall improvement; the slow speed of adjustment of Cotonou and Parakou is particularly worrying.

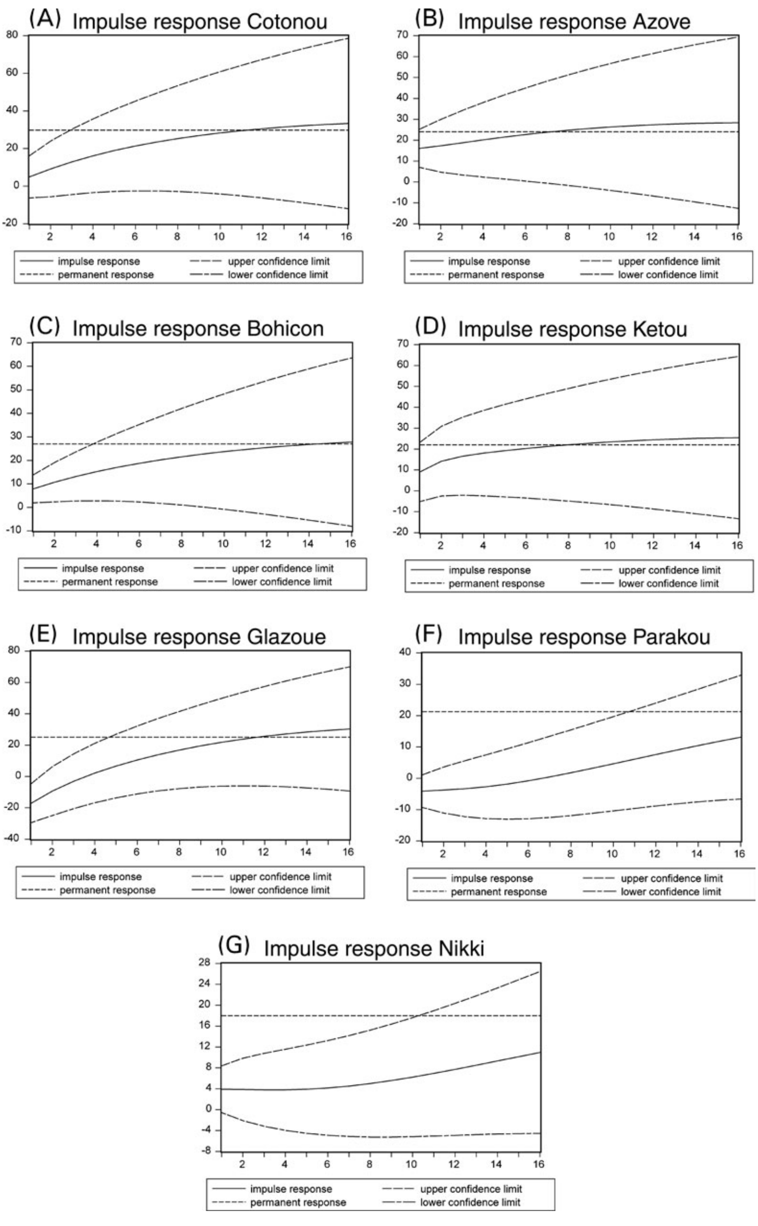

Figure 1: *Impulse Response to the Largest Shock in the Permanent Component 1987–9 (95% Confidence Interval, Number of Periods on X-axis, Difference vis-à-vis the Pre-Shock Price in Fcfa/kg on Y-axis).*

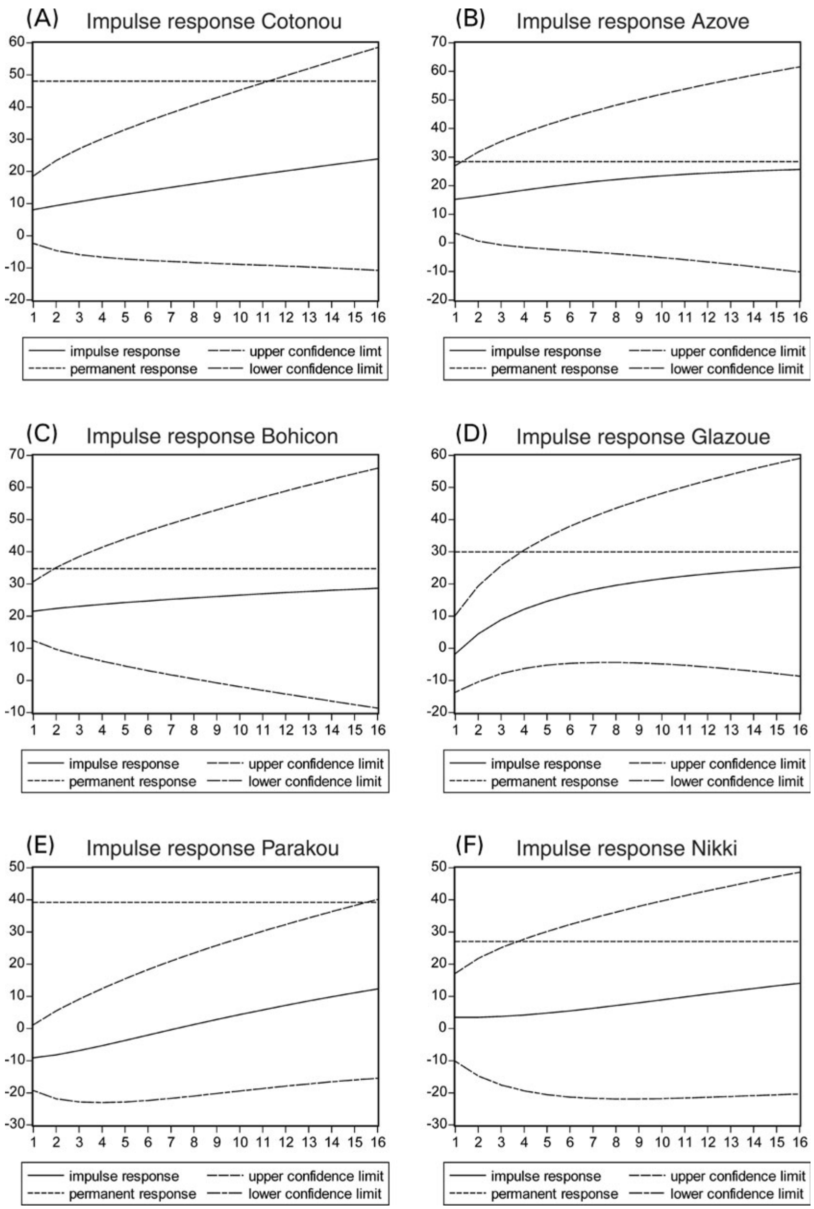

Figure 2: *Impulse Response to the Largest Shock in the Permanent Component 1998–2001 (95% Confidence Interval, Number of Periods on X-axis, Difference vis-à-vis the Pre-shock Price in Fcfa/kg on Y-axis).*

Table 5: *Number of Periods Before Impulse Response Does Not differ Significantly From the Permanent Level*

Market	Co	Az	Bo	Ke	Gl	Pa	Ni
1987–9	2	0	3	0	4	10	9
1998–2001	11	1	1	–	3	15	3

Note: One period is 4 days. Hence, j periods are $4 \times j$ days.

Table 6 shows the number of significant transitory $\boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}(\mathbf{I}_{p-1}+\boldsymbol{\beta}'\boldsymbol{\alpha})^{i-1}\boldsymbol{\delta}s$ responses for the 1998–2001 sample. The results, computed with $\boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}(\mathbf{I}_{p-1}+\boldsymbol{\beta}'\boldsymbol{\alpha})^{i-1}$ obtained from the 1987–9 sample VECM estimates, are compared with the results computed with $\boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}(\mathbf{I}_{p-1}+\boldsymbol{\beta}'\boldsymbol{\alpha})^{i-1}$ obtained from the 1998–2001 sample VECM estimates. Table 6 confirms the outcome presented in Table 5. The speed of adjustment based on $\boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}(\mathbf{I}_{p-1}+\boldsymbol{\beta}'\boldsymbol{\alpha})^{i-1}$ obtained from the 1998–2001 sample VECM estimates implies there is a further delay in adjustment.

6. Conclusions

The findings of this article confirm that the market reforms have not improved maize market integration. Though most markets

Table 6: *Number of Periods for the 1998–2001 Sample Before the Respective Transitory Responses* $\boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}(\mathbf{I}_{p-1}+\boldsymbol{\beta}'\boldsymbol{\alpha})^{i-1}\boldsymbol{\delta}s$ *Become Insignificant*

Market	Co	Az	Bo	Ke	Gl	Pa	Ni
$\boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}(\mathbf{I}_{p-1}+\boldsymbol{\beta}'\boldsymbol{\alpha})^{i-1}$ obtained from the VECM 1987–9 sample:							
	0	0	0	–	1	2	4
$\boldsymbol{\alpha}(\boldsymbol{\beta}'\boldsymbol{\alpha})^{-1}(\mathbf{I}_{p-1}+\boldsymbol{\beta}'\boldsymbol{\alpha})^{i-1}$ obtained from the VECM 1998–2001 sample:							
	5	1	3	–	2	5	2

Note: One period is 4 days. Hence, j periods are $4 \times j$ days.

ultimately become integrated, the price adjustment processes remain sluggish. The results for Cotonou, the major consumer market, are particularly disappointing, as are those for Kétou and Parakou. This finding was not anticipated, as the policy change, coupled with the current widespread use of mobile phones (an unknown technology in Benin 15 years ago) by wholesalers and the increased volume of maize trade (Cotonou is a fast-growing consumer market) were expected to intensify trade relationships. However, the findings underline the conclusions drawn by Adegbidi *et al.* (2003), that no fundamental changes in the organisation of the market have been brought about. To date, the maize market does not show two important characteristics of a liberalised market: open-ness to change and the development of new channel formats. It can also be concluded that it is too early to say that this market is halfway down the road to structural adjustment (Kherallah *et al.*, 2000).

Our results confirm the conclusions drawn some time ago by Badiane (2000, p. 153): 'Further progress in market reform will require not only further liberalisation, but a more concerted effort to go beyond the withdrawal of the public sector from agricultural marketing'. To date, the effects of the policy change on the functioning of the maize market have been disappointing. Put differently, the forces that resist change have generally prevailed. All this supports the conclusion that a crucial element in the improvement of the functioning of the market under study is the so-called issue of getting the institutions right.

The Beninese experiences are in line with a general observation made by Kherallah *et al.* (2002) and do indeed confirm that the establishment of an effective institutional environment is a complex, time-consuming process. Lutz (1994) and Adegbidi *et al.* (2003) propose four essential elements of this environment, which are expected to improve market integration. First, farmers' organisations should become more active and develop commercial strategies in order to produce a larger share of the value chain. The greater involvement of farmers will put pressure on traders to abide by market laws. Secondly, the reliability of the information disseminated by the market information service should be improved, in order to facilitate competitive forces in the market. Thirdly, the use of Grades and Standards, which was advocated almost 50 years ago by Abbott (1958), merits serious consideration (cf. Farina and Reardon, 2000), as it would reduce information asymmetries in the market and encourage the entry of non-resident wholesalers. Finally, there is a need for a well-equipped market authority that fosters competitive processes. In particular, the authority should guarantee access to markets for non-resident wholesalers, that is level the playing field. Though the 'Direction de la Concurrence et du Commerce Interne' is formally responsible for this task, that organisation seems to focus on the price levels of five commodities: pharmaceutical products, petrol, water, electricity, bread and office (and school) supplies. It has not undertaken activities in the food market, as these markets were considered to be 'completely deregulated' (Adegbidi *et al.*, 2003). Although the data presented in this article do not allow conclusions to be drawn about the causes of poor integration, we can conclude that the abovementioned elements do provide an interesting agenda for further research.

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Appendix. Checking for linear and band-threshold VECMs

Consider the prices of a homogeneous commodity traded on two spatially dispersed markets, market 1 and market 2. The prices are denoted as P_{1t} and P_{2t}. Let the transaction and transportation costs per unit of product be captured by the constant absolute price margin C ($C > 0$). Consequently, arbitrage occurs if $|P_{1t} - P_{2t}| > C$. Hence, there is no trade between the two markets if $|P_{1t} - P_{2t}| \leq C$. This situation can be described by the following three-regime band-threshold VECM of order one ($i = 1, 2$):

$$\Delta P_{it} = \begin{cases} \gamma_i(P_{1,t-1} - P_{2,t-1} - C) + \varepsilon_{it} & \text{if } (P_{1,t-1} - P_{2,t-1}) > C \\ \varepsilon_{it} & \text{if } |P_{1,t-1} - P_{2,t-1}| \leq C \\ \gamma_i(P_{1,t-1} - P_{2,t-1} + C) + \varepsilon_{it} & \text{if } (P_{2,t-1} - P_{1,t-1}) > C \end{cases} \quad \text{(A1)}$$

where trade flows are directed from market 2 (1) to market 1 (2) in the $(P_{1,t-1} - P_{2,t-1}) > C$ $((P_{2,t-1} - P_{1,t-1}) > C)$ regime. Consequently, given that $\gamma_1 < 0$ and/or $\gamma_2 > 0$, if there is trade, error correction occurs. In contrast, prices follow a random walk if there is no arbitrage. Among the interesting references on band-threshold co-integration are Balke and Fomby (1997), Prakash and Taylor (1997), Dercon and Van Campenhout (1998), Ejrnaes and Persson (2000), Lo and Zivot (2001), Goodwin and Piggot (2001), Sephton (2003) and Escobal (2005, Ch. 6).

If trade flows are predominantly directed from market 2 to market 1 and transaction and transportation costs are stationary around mean C, then the linear VECM may be most appropriate:

$$\Delta P_{it} = \gamma_i(P_{1,t-1} - P_{2,t-1} - C) + \varepsilon_{it} \quad \text{(A2)}$$

A model that nests (A1) and (A2) reads as:

$$\begin{aligned}\Delta P_{it} = \kappa_i\{ & P_{1,t-1} - P_{2,t-1} - \lambda|P_{1,t-1} - P_{2,t-1}|/(P_{1,t-1} - P_{2,t-1}) \\ & + \rho(P_{1,t-1} - P_{2,t-1} + |P_{1,t-1} - P_{2,t-1}|)/(P_{1,t-1} - P_{2,t-1}) \\ & - [\theta(P_{1,t-1} - P_{2,t-1}) - \tau|P_{1,t-1} - P_{2,t-1}|/(P_{1,t-1} - P_{2,t-1})] \\ & \times [1 - 0.5(P_{1,t-1} - P_{2,t-1} - C + |P_{1,t-1} - P_{2,t-1} - C|)/ \\ & \quad (P_{1,t-1} - P_{2,t-1} - C) - 0.5(P_{1,t-1} - P_{2,t-1} + C \\ & - |P_{1,t-1} - P_{2,t-1} + C|)/(P_{1,t-1} - P_{2,t-1} + C)]\} + \varepsilon_{it} \end{aligned} \tag{A3}$$

and reduces to (A1) if $\rho = 0$, $\theta = 1$ and $\tau = \lambda = C$ and becomes (A2) if $\lambda = \rho = -C$ and $\theta = \tau = 0$. Note that C is unknown and as a consequence of the discontinuity in the likelihood function, C cannot simply be estimated by maximum likelihood without the necessity of a grid search on the starting values of the optimisation algorithm. The parameters in (A3) can be estimated conditional on a grid of values for C to select that value of C, which maximises the likelihood. A more convenient procedure, however, is provided by the following iterative method. First, set C equal to the $(0.1T + 1)$th smallest value of $|P_{1t} - P_{2t}|$ (cf. Hansen, 1999), where $0.1T$ is rounded up to the next integer, and estimate the other parameters in (A3) conditional on this value. Secondly, if it appears that $\lambda \approx \rho < 0$, then (A1) may be rejected in favour of (A2). If, however, $\lambda \approx \tau > 0$, then set C equal to λ and reestimate (A3). Repeat until $\lambda = C$ and then check for evidence of the restrictions $\rho = 0$, $\theta = 1$ and $\tau = \lambda$.

We first apply our check for the bi-variate model presented in (A3), taking the rural surplus market Nikki as the reference market. The estimates are presented in Table A1. The values for C we started with for the 1987–9 sample (and in brackets for the 1998–2001 sample) are 7.00 (36.40), 4.00 (3.00), 6.61 (6.10), 3.60, 7.00 (3.40) and 5.53 (18.00) if P_1 is given by P_{Co}, P_{Az}, P_{Bo}, P_{Ke}, P_{Gl} and P_{Pa}. In the 1987–9 sample, we find evidence of a band-threshold only for Kétou, where the final estimate is 15.07. In the 1998–2001 sample, a band-threshold is found for Azové, although now the standard error of the estimate is large.

Table A1: *Estimates of the Parameters in (A3) to Check for Band-Thresholds*

P_1	P_2	Iteration	κ_1	κ_2	λ	ρ	θ	τ
Sample: 1987–9								
P_{Co}	P_{Ni}	1	−0.019 (0.030)	0.034 (0.015)	−33.71 (36.11)	−31.39 (17.67)	5.082 (9.958)	28.82 (46.32)
P_{Az}	P_{Ni}	1	−0.001 (0.024)	0.063 (0.021)	−3.246 (10.37)	−9.442 (4.784)	4.836 (18.10)	12.54 (41.93)
P_{Bo}	P_{Ni}	1	0.018 (0.027)	0.069 (0.021)	−14.10 (16.15)	−19.55 (8.093)	5.002 (13.48)	43.83 (61.57)
P_{Ke}	P_{Ni}	1	−0.309 (0.068)	0.059 (0.027)	12.19 (4.678)	−1.390 (2.539)	1.137 (9.376)	10.44 (13.17)
		2	−0.404 (0.091)	0.064 (0.035)	14.88 (4.180)	−1.501 (2.019)	0.456 (0.869)	10.09 (6.435)
		3	−0.412 (0.101)	0.067 (0.038)	15.41 (4.553)	−1.303 (1.970)	0.733 (0.501)	11.98 (5.783)
		4	−0.403 (0.101)	0.065 (0.037)	15.07 (4.734)	−1.320 (2.023)	0.769 (0.509)	11.77 (5.961)
P_{Gl}	P_{Ni}	1	−0.010 (0.020)	0.031 (0.019)	−32.10 (31.97)	−23.64 (12.83)	16.69 (13.56)	85.18 (49.36)
P_{Pa}	P_{Ni}	1	−0.040 (0.031)	0.047 (0.043)	−3.356 (44.04)	−5.841 (21.43)	5.254 (12.52)	7.867 (52.13)
Sample: 1998–2001								
P_{Co}	P_{Ni}	1	−0.005 (0.009)	0.043 (0.027)	−262.6 (220.6)	−164.7 (109.3)	−7.319 (7.130)	−187.6 (202.5)
P_{Az}	P_{Ni}	1	−0.015 (0.020)	0.132 (0.025)	10.35 (7.132)	−3.094 (4.071)	8.209 (12.47)	20.97 (25.67)
		2	−0.012 (0.021)	0.142 (0.026)	12.54 (7.421)	−2.693 (3.769)	0.587 (2.251)	10.10 (16.14)
		3	−0.012 (0.022)	0.149 (0.029)	14.25 (8.209)	−2.148 (3.835)	0.464 (1.582)	10.87 (14.30)
		4	−0.012 (0.022)	0.145 (0.031)	13.31 (9.055)	−2.498 (3.767)	0.686 (1.154)	11.05 (14.37)
		5	−0.012 (0.022)	0.144 (0.029)	12.98 (8.662)	−2.558 (3.986)	0.707 (1.595)	10.87 (14.74)
P_{Bo}	P_{Ni}	1	0.022 (0.018)	0.125 (0.040)	−7.895 (14.11)	−19.24 (7.121)	−1.384 (6.079)	20.51 (27.49)
P_{Gl}	P_{Ni}	1	−0.014 (0.023)	0.096 (0.025)	−3.435 (13.44)	−7.726 (6.366)	3.582 (17.00)	14.89 (39.28)
P_{Pa}	P_{Ni}	1	−0.003 (0.008)	0.010 (0.027)	−1346 (3789)	−673.4 (1837)	11.43 (37.21)	−202.6 (754.1)

Note: Standard errors are given in brackets.

Table A2: *Estimates of the Parameters in (A4) to Check for Band-Thresholds*

j	λ_j	ρ_j	θ_j	τ_j
Sample: 1987–9				
1	−45.46 (55.73)	−38.93 (29.40)	2.553 (14.48)	4.739 (67.07)
2	−33.35 (15.94)	−23.26 (12.22)	−5.712 (5.251)	−22.71 (18.53)
3	−17.99 (20.84)	−21.57 (11.65)	−0.148 (5.725)	−2.097 (26.27)
4	−6.734 (19.37)	−8.918 (8.969)	0.124 (0.775)	−1.115 (13.26)
5	−41.13 (26.80)	−28.43 (18.03)	−5.428 (8.937)	−25.42 (42.91)
6	−1.661 (25.90)	−5.460 (13.51)	1.625 (6.572)	6.663 (30.54)
Sample: 1998–2001				
1	−29.24 (40.19)	−47.78 (23.22)	2.616 (1.452)	98.06 (42.63)
2	2.840 (8.528)	−3.231 (7.063)	0.730 (0.584)	8.517 (5.046)
3	−18.56 (11.55)	−21.88 (8.079)	1.337 (2.174)	12.53 (10.45)
5	−4.394 (8.810)	−7.625 (8.041)	0.459 (3.507)	7.161 (9.366)
6	−95.46 (39.64)	−65.94 (21.63)	−2.821 (2.004)	−44.25 (32.35)

The version of (A3) that applies for more than two markets reads as:

$$\begin{aligned}\Delta P_{it} = \sum_{j=1}^{p-1}\{\kappa_{ij}\{&P_{j,t-1} - P_{p,t-1} - \lambda_j|P_{j,t-1} - P_{p,t-1}|/(P_{j,t-1} - P_{p,t-1})\\ &+ \rho_j(P_{j,t-1} - P_{p,t-1} + |P_{j,t-1} - P_{p,t-1}|)/(P_{j,t-1} - P_{p,t-1})\\ &- [\theta_j(P_{j,t-1} - P_{p,t-1}) - \tau_j|P_{j,t-1} - P_{p,t-1}|/(P_{j,t-1} - P_{p,t-1})]\\ &\times [1 - 0.5(P_{j,t-1} - P_{p,t-1} - C_j + |P_{j,t-1} - P_{p,t-1} - C_j|)/\\ &\quad (P_{j,t-1} - P_{p,t-1} - C_j) - 0.5(P_{j,t-1} - P_{p,t-1} + C_j\\ &- |P_{j,t-1} - P_{p,t-1} + C_j|)/(P_{j,t-1} - P_{p,t-1} + C_j)]\}\} + \varepsilon_{it}\end{aligned} \tag{A4}$$

where $i = 1, \ldots, p; j = 1, \ldots, p-1; P_1 = P_{\mathrm{Co}}; P_2 = P_{\mathrm{Az}}; P_3 = P_{\mathrm{Bo}}; P_4 = P_{\mathrm{Ke}}; P_5 = P_{\mathrm{Gl}}; P_6 = P_{\mathrm{Pa}}$; and $P_p = P_7 = P_{\mathrm{Ni}}$ is the reference market. Notice that in the 1998–2001 sample, we omit the Kétou price. The estimates for the parameters λ_j, ρ_j, θ_j and τ_j are presented in Table A2. The negative values (Nikki is a surplus market) and/or large standard errors for the estimates of the λ_j parameters, the tendency for $\lambda_j \approx \rho_j$, and the large standard errors of the θ_j and τ_j parameter estimates (such that we may expect these parameters to be insignificant) clearly indicate that the band-threshold VECM must be rejected in favour of the linear VECM. This conclusion is confirmed by rejecting the restrictions $\tau_j = \lambda_j$ ($\chi^2(5) = 19.41$, *P*-value $= 0.002$) and accepting $\lambda_j = \rho_j$ ($\chi^2(5) = 9.999$, *P*-value $= 0.075$) in (A4) for the 1998–2001 sample, provided that standard distribution theory holds.