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Transmission of global wheat prices to domestic markets in Kenya: A cointegration approach

Jonathan Makau Nzuma*
Department of Agricultural Economics, University of Nairobi, Nairobi, Kenya. E-mail: jnzuma@gmail.com

Patrick Kipruto Kirui
Department of Agricultural Economics, University of Nairobi, Nairobi, Kenya. E-mail: kiruip@gmail.com

* Corresponding author

Abstract

This paper evaluates the extent to which changes in international wheat prices are transmitted to domestic markets in Kenya using an error correction model (ECM) that employs monthly producer price data for the period 2002 to 2020. Domestic wheat markets in Kenya were found to be strongly integrated while, international wheat markets were cointegrated with domestic prices at the port of Mombasa. The long-run elasticity of price transmission was estimated at 0.91, which implies that 91% of the changes in international wheat prices are transmitted to domestic markets in Kenya. The speed of adjustment was estimated at -0.069, which implies that it takes about 14 months for the changes in the international wheat price to be fully transmitted to the Kenyan domestic market. Wheat farmers in Kenya seem to be insulated from international price shocks given the long period of time it takes for domestic markets to adjust to international price changes. Even though not explicitly analysed, government border policies, market and infrastructure impediments seem to be underlying causes of the incomplete price pass-through, along with the low speeds of adjustments. Our analysis suggests that the main constraint to a complete pass-through is the existence of price-setting power at the producer level of the wheat market in Kenya. Investments in infrastructure development and the promotion of liberal trade policies can improve the transmission of international wheat price signals to domestic markets in Kenya.

Key words: cointegration; error correction model; spatial price transmission; wheat

1. Introduction

Over the recent past, empirical research on agricultural price transmission analysis has attracted renewed interest. Interest in this topic unquestionably increased after the so-called "food price crisis" of 2007/2008, in which international agricultural markets were shocked by increased price volatility (Listorti & Esposti 2012). These price increases have been attributed to supply shortages, increased biofuel production, reduced stock-to-use ratios, export bans by major grain exporters and panic buying by some major importers (Gilbert 2010; Garrido *et al*. 2016). Commodity prices rose rapidly again from 2010 to 2014. This recent turbulence in agricultural markets strongly revitalised empirical research on price transmission (Esposti & Listorti 2013).

Understanding the sources of domestic food price volatility in developing countries, and the extent to which it is transmitted from international to domestic markets, is critical to help design better global, regional and domestic policies to cope with excessive food price volatility and to protect the most vulnerable groups (Ceballos *et al*. 2017). The extent to which world commodity prices are transmitted to domestic markets in developing countries has serious food security implications for smallholder farmers and low-income consumers, especially in Sub-Saharan Africa (SSA), where the capacity to respond to food crises is limited. The perceived vulnerability of SSA countries to changes in world food prices justifies the need to model the behaviour of prices during price shocks (Abidoye & Labuschagne 2014).

Yet empirical evidence on the degree to which world market shocks are transmitted to domestic markets in SSA is limited. Most of the available studies in SSA examine the degree of price transmission between markets within a country (Abdulai 2000; Kuiper *et al*. 2003; Rashid 2004; Negassa & Meyers 2007; Van Campenhout 2007; Moser *et al*. 2009; Myers 2013). These studies used cointegration approaches in the form of error correction models (ECM) to find evidence of market integration and price transmission within markets for the same commodity in one country. The results indicate that both distance and an international border between the markets reduce the probability that the prices will be cointegrated and slow the speed of adjustment if they are cointegrated (Ceballos *et al*. 2017)

Fewer studies have examined the transmission of staple food prices from world markets to local markets in SSA (Conforti 2004; Baquedano *et al*. 2011; Minot 2011; Abidoye & Labuschagne 2014; Baquedano & Liefert 2014; Ceballos *et al*. 2017; Hatzenbuehler *et al*. 2017). Conforti (2004) found low levels of transmission of wheat prices in Ethiopia and Ghana from an ECM that used a mix of monthly and annual retail prices. In general, he found the degree of price transmission in SSA countries to be lower than that in Latin America and Asian countries. Minot (2011) used monthly consumer prices from 13 countries in SSA and found evidence of transmission in one fifth of the considered domestic price series. Overall, Minot (2011) found rice prices to be more integrated than maize and wheat prices. Abidoye and Labuschagne (2014) applied a threshold cointegration on monthly maize producer prices to find that only large long-run deviations in world prices are transmitted to South African markets. Baquedano *et al*. (2011) compared market integration for export and imported crops in Mali (cotton and rice) and Nicaragua (coffee and rice) using a generalised ECM and found Nicaraguan agriculture to be more integrated into world markets than that of Mali. Using a single equation ECM that applies monthly consumer prices, Baquedano and Liefert (2014) found that 84% of consumer markets in developing countries are cointegrated with world markets.

Ceballos *et al*. (2017) applied a multivariate GARCH (generalised autoregressive conditionally heteroscedastic) model to examine grain price and volatility transmission from world markets to local markets in 27 developing countries in Latin America, Africa and Asia. The authors observed significant interactions from international to domestic markets in only a few cases. Hatzenbuehler *et al*. (2017) used an ECM to find limited evidence of international price transmission in Nigerian food security crop markets and conclude that tradability matters for price transmission.

Overall, the studies examining the transmission of world food prices to domestic markets in SSA employ dynamic regression models based on the vector autoregressive model (VAR) to report incomplete price transmission from world to domestic markets in the region, with variations across countries and crops. According to Kalkuhl (2016), the rather slow rates of international price transmission could be attributed to the use of aggregate food price indices that mask the heterogeneity across countries and commodities, along with other policy and infrastructural impediments. In almost all cases, these studies use a mix of monthly and annual consumer price data, with little exploration of producer prices. Moreover, it is not clear whether these studies analysed the prices of raw or processed commodities and at what level of the market.

This paper examines the transmission of world wheat prices to domestic markets in Kenya using monthly producer prices for unmilled grain wheat at the wholesale level. To the best of our knowledge, the transmission of world wheat prices to domestic markets in Kenya has not been studied extensively. Mundlak and Larson (1992) used annual data of the FAO from 1968 to 1978 and a static model to estimate a long-run price transmission elasticity of 78% for Kenya's wheat. The static model was criticised by later authors, putting in doubt the reliability of its estimates. Unlike the earlier studies, which used either aggregate annual price data or price indices at the retail level, this paper employs a cointegration approach on monthly wholesale producer price data for 18 years to evaluate the transmission of world wheat prices to three domestic markets in Kenya (Mombasa, Nakuru and Eldoret). Nakuru and Eldoret are key wheat-producing zones, while Mombasa is the sea port of entry and a major consumption market, with a relatively well-developed milling industry.

Given Kenya's high dependence on wheat import in the face of vulnerable smallholder producers and low-income consumers, this study is timely to inform policy on the influences of international food price shocks on food security. It therefore would be useful to know the degree to which world wheat price shocks influence wheat prices in Kenya and what the effects are on producers. The paper seeks to answer four main questions: (1) Are domestic wheat markets in Kenya integrated? (2) Are world wheat markets integrated with domestic markets in Kenya? (3) To what extent are world wheat price changes transmitted to domestic markets in Kenya? and (4) How long does it take for domestic prices in Kenya to adjust to changes in world wheat prices? We used a cointegration model to evaluate domestic wheat market integration in Kenya and to estimate international price transmission elasticities.

Wheat is the second most important food crop in Kenya after maize in terms of production and consumption (Nyangito *et al*. 2002). Eighty-five percent of production occurs on large and medium-scale farms, using capital-intensive technologies (Monroy *et al*. 2013). The majority of the large-scale wheat farmers are members of the Cereal Growers Association (CGA), a powerful wheat farmers' lobby group that sets producer prices in collaboration with the Government of Kenya (GOK) and wheat millers. Annual per capita consumption is estimated at 0.028 metric tonnes (MT), as compared to 0.103MT for maize (Abate *et al*. 2015).

Domestic production falls short of wheat demand, creating a need for importation. Kenya is a net importer of wheat, bringing in two-thirds of its requirement to meet the annual consumption of 900 000 tonnes, against local production of 350 000 tonnes (USDA 2019). Kenya imports hard wheat, consisting of unmilled (grain) durum wheat and meslin (Musyoka 2009). The imported hard wheat is of high quality relative to domestically produced soft wheat and is used in baking. Millers blend imported hard wheat varieties with soft wheat in a 40:60 ratio to produce a flour quality that meets Kenyan market demands (Monroy *et al*. 2013). This implies that a quality adjustment needs to be made when comparing Kenya's wheat prices with those of imported wheat. The bulk of Kenya's wheat imports are sourced from Russia, Argentina, Ukraine, Canada, Latvia and the United States of America. In the recent past, more than half of Kenya's wheat imports have been sourced from Russia and Ukraine, with Ukraine playing a dominant role after Russia imposed an import ban in 2010 (Gitau *et al*. 2013). US wheat exports to Kenya are currently hindered by Kenya's long-standing restriction of Pacific Northwest (PNW) wheat due to the lack of an export certification protocol for *flag smut* between Kenya and the United States.

Currently, Kenya's wheat grain imports are accessed at 10% ad valorem for registered millers; otherwise, the Eastern Africa Community (EAC) common external tariff (CET) of 35% applies (USDA 2019). In addition, GOK maintains an understanding with the Cereal Millers Association (CMA), a key industry association, that CMA members must provide a mop-up plan for local produce before they can be granted import licences. There are about 103 registered grain millers in Kenya, with an installed capacity of 1.6 million MT per year (USDA 2019). Nineteen of these millers account for about 85% of the market and, although this does not necessarily imply market power in grain importation and milling, it indicates a concentration of activity in relatively fewer firms than those registered. Demand for wheat and wheat products in Kenya has increased over time as a result of a growing economy and rapid urbanisation. Wheat demand is fuelled by the considerable expansion in home and industrial baking. In addition to the traditional bakeries, most leading supermarket chains have opened baking units within their stores.

While trade policies and other factors (transaction costs, exchange rate risks, information imperfections, expectations, etc.) affect price transmission, this study does not explicitly consider the sources of these distortions to the law of one price (LOP), the underlying theoretical framework of price transmission. Instead, the potential influences of these distortions on the price transmission estimates are discussed in Section 2, while the results are interpreted with the potential effects of these factors in mind. This study adopts a non-structural approach that treats factors affecting transmission as external prior information, rather than as the outcome of a theoretical framework to be confirmed by the results, and focuses its attention on the dynamics of the transmission process.

We find evidence of strong domestic wheat market integration, coupled with high transmission rates of international shocks to domestic markets, but low speeds of adjustment of domestic markets to international price shocks. The low speeds of adjustment to international price shocks suggest that wheat farmers in Kenya are insulated from international price shocks owing to a number of factors that are explored in the paper. The remainder of this paper is organised as follows: Section 2 reviews the factors that influence the transmission of international price shocks to domestic markets. The third section presents the methodology and data used to measure the integration of domestic markets and the transmission of international price shocks to domestic markets. Section 4 presents a discussion of the empirical results, while the conclusions and policy implications are provided in Section 5.

2. Factors influencing international market integration

Market integration is affected by several factors, including border policies, price support mechanisms, transfer costs, the exchange rate and market structure (Zorya *et al*. 2014). Another potential source of inaccuracy in transmission estimation emanates from the lack of substitutability between domestic and foreign goods. Transmission-impeding policies include systems of managed (or fixed) prices, state trading enterprises, import quotas and trade-prohibitive tariffs (Liefert & Persaud 2009). Border policy instruments, such as import tariffs and export subsidies, export bans, export taxes and non-tariff barriers (NTBs) insulate domestic markets from international markets and impede the complete transmission of price signals coming from world prices (Ozturk 2020). If a high import tariff is imposed, the world price changes would pass through to the domestic prices partly, if at all, which would cause the international and domestic prices to move independently of each other (Rapsomanikis *et al*. 2003). Moreover, the implementation of price support policies, such as floor prices, deficiency payments or any other supports, can cause domestic and international prices to move independently.

Transfer costs may also cause weak or no cointegration at all (Ozturk 2020). Transfer costs can generate domestic transport and transaction costs (TT) so high that they preclude any trade in a product, as well as any transmission between the good's border and domestic prices (Balcombe *et al*. 2007). In developing countries in particular, transportation costs are high due to poor infrastructure. Weak infrastructure (physical, commercial and institutional) increases transaction costs and impedes the flow of price and other key market information from borders to interior regions (Barrett 2001; Fackler & Goodwin 2001; Barrett & Li 2002). Exchange rate changes can also retard transmission, where the change can result from either a macroeconomic policy or economy-wide developments within a system of floating exchange rates (Baquedano *et al*. 2011). When local currency appreciates (depreciates) against the US dollar, an increase in the commodity price in the local currency would be less (more) than an increase in the world price in dollars. Hence, prices deviate from each other.

Market structure may also hinder market integration. Market structure and conditions that can affect price and exchange rate transmission include domestic market power and weak infrastructure (Baquedano & Liefert 2014). Market power gives domestic producers price-setting potential such that changes in border prices are not transmitted completely to domestic prices. In non-competitive markets, for example, high prices in the world market may not be transmitted to the domestic consumer or to producer prices. Lastly, domestic products and their foreign counterparts might not be complete substitutes (homogenous) (Baquedano & Liefert 2014). If the imported good is of higher (lower) quality than its domestic analogue, then the estimate of transmission elasticity is understated (overstated) (Baquedano *et al*. 2011). Consequently, the transmission of world (foreign) to domestic prices for these goods will probably be incomplete.

While we do not explicitly model the effects of these factors on wheat market integration in Kenya, we note their potential effects on our price transmission elasticities. The existing state policies should allow a complete transmission between border and domestic prices (assuming it is not prohibitively high and no non-policy market conditions impede transmission). However, the existence of price-setting power – at the farm level by the Cereal Growers Association (CGA) and at the processing level by the Cereal Millers Association (CMA) – may hinder the full transmission of world wheat prices to domestic markets in Kenya. In addition, the existing weak market infrastructure, which imposes high transport and transaction costs, may impede a complete pass-through of changes in world wheat prices.

3. Methodology

The law of one price (LOP) forms the theoretical basis for spatial price transmission analysis. Assuming no policy distortions, prices adjusted for transport cost for a homogeneous commodity in spatially separated markets should be equal (Houck 1986). The causal relationship between the world price and domestic prices for traded commodities can be examined in an error-correction framework. The ECM model allows for the disentanglement of long-run and short-run dynamics in price interdependence (Listorti & Esposti 2012). The relationship between the domestic price and the world price can be derived from a static price transmission model, represented as follows:

$$P_{it}^{d} = \beta P_{it}^{w} + u_t, \tag{1}$$

where P_{it}^{d} and P_{it}^{w} are the natural log of the domestic price and the world price respectively, in nominal terms for commodity i at time t, and u_t are the residuals The parameter β is the price transmission elasticity of a change in P_{it}^{w} to P_{it}^{d} .

If the residuals, u_t, in Equation (1) are stationary, then there is a long-run equilibrium between the series (Engle & Granger 1987). Thus, the domestic prices and world prices are cointegrated. If two price series are cointegrated, they lend themselves to an error correction specification which is defined in the second stage of estimation. The model in Equation (1) represents a general form of the augmented distributed lag model (ADL). To deal with possible nonstationary issues within the price series and to evaluate the adjustment process for price transmission, we adopted the generalised ECM proposed by Hendry *et al*. (1984). The ECM has a general autoregressive distributed lag structure and, as such, is a direct transformation of an ADL model (Banerjee *et al*. 1990).

To correct the deviation from the equilibrium in the last period, the residuals of Equation (1) are inserted into a dynamic ECM and specified as follows:

$$\Delta p_t^{d} = \alpha + \gamma \Delta p_{t-1}^{w} + \delta\left(p_{t-1}^{d} - \beta p_{t-1}^{w}\right) + \mathcal{E}_{\mathrm{t}}, \tag{2}$$

where Δp_t^d and Δp_{t-1}^w are the first difference of the natural log of the domestic and world prices in US dollars of commodity *j* at time *t*. Given the ADL structure, β is the short-run transmission elasticity of changes in the border price to the domestic producer/consumer price. δ measures the speed of convergence toward the long-run relationship between the border price and the domestic price, and γ represents the long-run price transmission elasticity of a change in border prices to domestic consumer/producer prices. Finally, $\mathcal{E}_{\mathrm{t}}$ is an identical independently distributed (*iid*) error term.

This general form of Equation (2) can be characterised as an ADL(1,1) process, meaning that one lag of the dependent and one of the independent variable are considered as regressors (Baquedano & Liefert 2014). Equation (2) can be estimated by OLS as long as the error term represents a stable stationary process, i.e. $\mathcal{E}_{\mathrm{t}}$ is *I(0)* (Lütkepohl 2005). Equation (2) is used in this study to develop an ECM that characterises the data generation process. The two-step ECM has been used widely to deal with nonstationary issues in price series and to evaluate market integration and price transmission. Equation (2) suggests that changes in the domestic prices stem from two sources. First, changes in world price and, second, changes in the error correction term (ECT). Equation (2) may contain more than one lag of changes in domestic and world prices. Thus, an optimal lag length must be determined before the estimation of the ECM.

3.1 Estimation strategy

The estimation procedure of the price transmission elasticities proceeded in three steps: First, we tested for unit roots. Second, we tested for cointegration and, finally, we estimated the price transmission elasticities in an ECM, in which it was assumed that the underlying price series were nonstationary and cointegrated. To determine appropriate lag lengths, we relied on the Akaike information criterion (AIC) proposed by Johansen (1988). This criterion involves choosing the lowest AIC value that gives the model's best fit. In addition, the 2007/2008 food prices, coupled with Kenya's post-election crisis, might have caused a breakpoint in the data. To test for structural price breaks, we used a Chow test to divide the data into two subsamples based on a specified breakpoint (Chow 1960). It uses an *F*-test to determine whether differences exist between these two models.

3.1.1 Unit root and cointegration tests

The Augmented Dickey-Fuller (ADF; 1979) test and the Phillips-Perron test (Phillips & Perron 1988) are the most frequently used tests for unit root in the literature. An ADF test can be specified as:

$$\Delta Y_t = \mu + \eta t + \tau * Y_{t-1} + \sum_{j=1}^{p} \phi_j \Delta Y_{t-j} + e_t, \tag{3}$$

where Y_t is a random variable, possibly with non-zero mean, μ is a constant, t is a time trend and e_t is the error term that is independently and identically distributed (i.e. *iid* ~ $[0, \sigma^2]$). The null hypothesis of the unit root ($\tau^* = 1$) is tested against the alternative of stationarity. However, the estimated τ^* does not have a standard *t*-distribution and hence the critical values provided by Dickey and Fuller (1979) are used.

Once the order of integration of the variables is established, the series are tested for cointegration. This is undertaken using Johansen's (1988) maximum likelihood (MLE) procedure, which is based on an unrestricted VAR model that can be specified following Johansen and Juselius (1990), as:

$$\Delta Y_{t_i} = \mu + \Pi Y_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-1} + V_i, \tag{4}$$

where Y_t is an *n* x 1 vector of variables that are integrated of order one, ~I(*1*); *Π* is the rank of the matrix; V_t is a vector of random errors; while μ and Γ are parameter to be estimated. The Johansen MLE test evaluates the rank (*r*) of the matrix Π, which is defined as the product of two matrices (i.e. $\Pi = \alpha\beta'$). If $r = 0$, all variables are I(*1*) and thus non-stationary. In the case of $0 < r < n$, there exist *r* cointegrating vectors. In the third case, if $r = n$, all the variables are I(*0*) and thus stationary.

Johansen's methodology provides two test statistics for the number of cointegrating vectors: the trace and maximum eigenvalue tests, specified below (Johansen 1995) as:

$$J_{trace} = -T\sum_{i=r+1}^{n} \ln(1-\lambda_i^2) \tag{5}$$

$$J_{max} = (r, r+1) = -Tln(1-\lambda_{r+1}), \tag{6}$$

where *T* is the sample size and $\hat{\lambda}_i$ is the *i*th largest canonical correlation of Δy_t with y_{t-1} after correcting for lagged differences and deterministic variables, when present. The trace statistic tests the null hypothesis that the cointegration rank is equal to r against the alternative that it is equal to *n*. Conversely, the maximum eigenvalue statistic tests the null hypothesis that the cointegration rank is equal to *r* against the alternative that it is equal to r +1 (Hjalmarsson & Österholm 2010).

3.1.2 Price transmission analysis

Once a cointegrating relationship has been established among the variables, the next step is to analyse for price transmission. In this study, the estimated model tests for the effect of world wheat prices on domestic wheat prices in Kenya, since Kenya is a small net importer of wheat and so unlikely to influence world prices. Following Minot (2011), the ECM for a small importing country can be derived from Equation (2) and specified as follows:

$$\Delta p_t^d = \alpha + \theta\left(p_{t-1}^d - \beta p_{t-1}^w\right) + \delta\Delta p_{t-1}^w + \rho\Delta p_{t-1}^d + \mathcal{E}_{\mathrm{t}}, \tag{7}$$

where p_t^d is the log of the domestic wheat price in Kenya converted into US dollars; p_t^w is the log of world wheat price (Ukraine export) in US dollars; (p_{t-1}^d - $\beta_1 p_{t-1}^w$) is the error correction term; *Δ* is the difference operator, so $\Delta p_t = p_t - p_{t-1}$; α, θ, β, δ and ρ are parameters to be estimated; and ε_t is the error term. Since all price series are logged before estimation, the cointegration factor (*β*) is the long-run elasticity of price transmission. The expected value of *β* for imported commodities is $0 < \beta < 1$. If $\beta = 0.5$, this implies that 50% of the change in the international price will be transmitted to the domestic price in the long run.

The error correction coefficient (*θ*) reflects the speed of adjustment and falls in the range of $-1 < \theta < 0$. The coefficient (*δ*) is the short-run elasticity of the domestic price relative to the world price. In this case, it represents the percentage adjustment of domestic price one period after a 1% shock in international price. α captures the trend in the error correction term of the ECM. The coefficient on the lagged change in the domestic price (*ρ*) is the autoregressive term, reflecting the effect of each change in the domestic price on the change in domestic price in the next period. The expected value is $-1 < \rho < 1$.

3.2 Data sources

The study employed monthly wheat producer price data at the wholesale level for the period 2002 to 2020. The domestic wheat prices are based on monthly observations of wholesale prices compiled from Kenya's Ministry of Agriculture, Livestock, Fisheries and Cooperatives for the following markets: Mombasa, Nakuru and Eldoret. The price of wheat at Mombasa is used as the domestic reference price for Kenya. International wheat prices were obtained from the Global Information and Early Warning Systems (GIEWS) database of the Food and Agriculture Organization of the United Nations (FAO). The Ukraine milling wheat export price in US dollars was used as the international reference price. The domestic prices in Kenyan shillings were converted to US dollars. Exchange rate data was obtained from the Central Bank of Kenya (CBK). All prices are monthly per metric tonne (MT), and were normalised and converted to natural logarithms before estimation. The data was analysed using the Eviews7 statistical package.

4. Results and discussion

Table 1 presents the results of the Augmented Dickey Fuller (ADF) and the Phillips-Perron (PP) tests for unit roots at level and first differences. All variables were transformed into logarithms, while the models were estimated with a constant and a trend variable included. The hypothesis of a unit root in the level series cannot be rejected for the four price series at the 5% level for both the ADF and PP tests (Table 1). However, the hypothesis of a unit root on the first differenced series is rejected for all the four price series (Table 1). These findings imply that wheat prices are characterised by non-stationary trends of order one [I(*1*)]. Thus, the domestic and international wheat price series used in this study are non-stationary and integrated of order one. Since the wheat price series are integrated of order one, it is expected that they are jointly determined and might be cointegrated.

Table 1: Unit root test results for wheat prices in Kenya and Ukraine

Series	Level series		Lags	First differences		I(*d*)
	ADF test statistic	PP test statistic		ADF test statistic	PP test statistic	
Mombasa	-2.38	-3.01	1	-18.96	-19.11	I(1)
Nakuru	-2.37	-2.94	1	-11.50	-22.33	I(1)
Eldoret	-2.56	-2.79	1	-12.85	-12.77	I(1)
Ukraine	-3.11	-2.64	1	-6.90	-10.31	I(1)
5% critical values	-3.44	-3.44		-3.44	-3.44	

Table 2 presents the Johansen's cointegration test results for the three domestic markets in Kenya (Mombasa, Nakuru and Eldoret). Both the trace test statistic and the maximum eigenvalue test reject the null hypothesis of at least one cointegrating vector at the 5% level (Table 2). These findings suggest the existence of one cointegrating vector among the domestic wheat market prices in Kenya. It therefore can be concluded that wheat markets in Kenya have a stable long-run equilibrium relationship to which the variables in the system have a tendency to return and can be interpreted as being integrated. Thus, the three domestic wheat markets analysed in this study are strongly integrated.

Table 2: Johansen's cointegration test results for domestic market integration

No. of cointegrating equations	Trace statistic	5% critical value	Max. eigenvalue	5% critical value
None **	52.84	29.68	41.89	20.97
At most 1	10.95	15.41	6.89	14.07
At most 2	3.76	3.76	3.76	3.76

Note: The critical values are adopted from Osterwald-Lenum (1992). * (**) denotes rejection of the null at the 5% and 1% level

Table 3 presents the results of the bivariate cointegration test between domestic wheat markets in Kenya and the international market. Both the trace statistic and the eigenvalue test reject the null hypothesis of at least one cointegrating vector at the 5% level for Mombasa and Ukraine (Table 3). The existence of one cointegrating vector between the wheat price at Kenya's port of Mombasa and that of Ukraine implies that there is a stable long-run equilibrium relationship between wheat prices in Kenya and the Ukraine export price, since the Mombasa price is used as the domestic reference price. It therefore can be concluded that world wheat markets are integrated with domestic markets in Kenya.

However, the trace test and maximum eigenvalue test results fail to reject the hypothesis of at least one cointegrating vector between the domestic wheat prices in both Nakuru and Eldoret and the Ukraine prices at the 5% level (Table 3). These findings suggest that there is not a cointegrating relationship between the domestic market prices in Nakuru and Eldoret and the international market prices, and therefore there is no long-run relationship between these markets and the international market. The integration between the international prices and the prices in Mombasa is plausible, given that Mombasa is the port of entry, while Nakuru and Eldoret are hinterland producer markets. A probable explanation for the lack of integration between the international prices and prices in Nakuru and Eldoret could be that producer prices in these markets are predetermined before the importation of wheat by the CGA and government, based on estimated production costs (and not international prices, to protect large-scale wheat farmers). This explanation is validated by the fact that domestic wheat prices in these two markets are generally higher than imported wheat prices, accounting for all other factors.

Table 3: Bivariate cointegration test results for wheat market integration

Hypothesised number of CE(s)	Trace statistic	5% critical value	Maximum eigen statistic	5% critical value
Mombasa and Ukraine prices				
r = 0*	18.05	15.41	14.91	14.07
r ≤ 1	3.14	3.76	3.14	3.76
Nakuru and Ukraine prices				
r = 0	11.21	15.41	7.89	14.07
r ≤ 1	3.33	3.76	3.33	3.76
Eldoret and Ukraine prices				
r = 0	12.12	15.41	7.80	14.07
r ≤ 1	4.32	3.76	4.32	3.76

Note: The critical values are taken from Osterwald-Lenum (1992). * denotes the rejection of the hypothesis of no cointegration at the 5% level

Since the domestic wheat prices in Kenya and the international prices are cointegrated, the long-run and short-run dynamics between international and domestic wheat markets were estimated using an ECM framework. Before estimation of an ECM, it is important to determine the number of lags that will be incorporated into the estimation. Table 4 presents the optimal lag-length results based on a number of tests, which include the Akaike information criterion (AIC), the Schwarz information criterion (SBIC), the final prediction error (FPE), the sequential modified LR test statistic, and the Hannan-Quinn information criterion (HQIC). All five tests select the equivalence of a two-lagged model. It therefore can be concluded that a two-lagged specification is appropriate for all variables in the model. Thus, an optimal lag length of two was used in this study

Table 4: Test results for optimal lag length of the VAR

Lag	LogL	LR	FPE	AIC	SBIC	HQIC
0	-49.513	NA	0.005	0.468	0.499	0.481
1	578.254	1 238.412	1.89e-05	-5.202	-5.110	-5.165
2	601.120	44.693*	1.59e-05*	-5.374*	-5.220*	-5.312*
3	602.138	1.971	1.63e-05	-5.347	-5.131	-5.259

* indicates lag order selected by the criterion; N = 220

To test for the possibility of a permanent shift in the long-run relationship between domestic wheat prices in Kenya and world wheat prices, a Chow test of structural break was undertaken. We postulated that the 2007/2008 food price crisis, coupled with Kenya's 2007 post-election crisis, might have caused a structural break in prices and used December 2007 as the break point. The null hypothesis of no structural break was rejected at the 1% level (Table 5). Given this evidence of a structural break, we divided the estimation period for the relationship between the world wheat prices and producer prices in Kenya into the two periods of January 2002 to December 2007 and January 2008 to December 2020. We then estimated the long-run relationships for the two sub-samples and compared the results with those of the pooled data from January 2002 to December 2020.

Table 5: Chow test results for a structural break

Chow breakpoint test: 2008M01 Equation sample: 2002M01 2020M12			
F-statistic	29.521	Prob. F(1,226)	0.000
Log likelihood ratio	27.992	Prob. chi-square (1)	0.000
Wald statistic	29.522	Prob. chi-square (1)	0.000

The long-run relationship between world wheat prices and domestic producer prices in Kenya is presented in Table 6. The ECM results shows that international wheat prices are transmitted to the Kenyan domestic market in the long run but not in the short run, since the long-run adjustment parameter (β) is significant at the 5% level, while the short-run adjustment parameter (δ) is not (Table 6).

Table 6: Price transmission elasticities for wheat in Kenya (2002 to 2020)

Variable	**Estimation period**		
	January 2002 to December 2007	**January 2008 to December 2020**	**January 2002 to December 2020**
Constant	0.005 (0.662)	0.001 (0.172)	0.003 (0.745)
Long-run PTE ($\boldsymbol{\beta}$)	0.878 (4.115)	0.962 (3.208)	0.909 (7.634)
Short-run PTE (δ)	0.020 (0.085)	0.041 (0.578)	0.057 (0.864)
Speed of adjustment ($\boldsymbol{\theta}$)	-0.064 (-2.364)	-0.043 (-1.908)	-0.069 (-3.020)
Adjusted R^2	0.083	0.148	0.138
F-statistic	2.229	6.297	8.162
Log likelihood	97.396	208.518	308.186
Sample size	69	153	225

The figures in parenthesis are t-statistics

The long-run elasticity of price transmission from the world wheat markets to domestic markets in Kenya for the pooled sample was estimated at 0.91 (Table 6). A one percentage increase in the price of wheat in the international market will lead to an increase in the domestic price by 91% (Table 6). This finding suggests that 91% of wheat price changes in the markets are transmitted to the domestic markets in Kenya. Similarly, statistically significant long-run price transmission elasticity of 0.88 and 0.96 was found for the 2002 to 2007 and the 2008 to 2020 sub-samples respectively (Table 6). The high price transmission elasticities are not surprising, given that Kenya is a net importer of wheat, a commodity that is highly tradable.

The short-run price transmission elasticities (δ) are not significant across all sub-samples, which suggests that, in the short run, changes in wheat prices in the international markets do not have any significant influence on domestic prices in Kenya (Table 6). The coefficient of the speed of adjustment (θ) to the long-run equilibrium relationship was negative and significant at the 1% level for the period from 2002 to 2007 and the pooled sample. However the speed of adjustment for the 2008 to 2020 sub-sample was only significant at the 10% level. The ECT was estimated at -0.069 for the pooled sample and -0.064 for the 2002 to 2007 period (Table 6). The negative and statistically significant coefficient of the speed of adjustment (θ) for the domestic wheat market confirms the validity of the model. It also confirms our earlier finding that world wheat markets are integrated with domestic markets in Kenya. The coefficient indicates the time it takes for wheat prices in Kenya to return to equilibrium after a change in the world markets. The result indicates that it takes 14 months for wheat prices in Kenya to adjust to changes in international prices. The impulse response function (IRF) in Figure 1 corroborates our finding on the speed of adjustment. The IRF hits a plateau in the 14th month (Figure 1), suggesting that domestic markets take about 14 months to adjust to world shocks. The period from 2002 to 2007 had a lower speed of adjustment than the pooled sample. The adjusted R-square value (0.14) suggests that 14% of the variability of Kenya's wheat market is explained by the variability in the world wheat price.

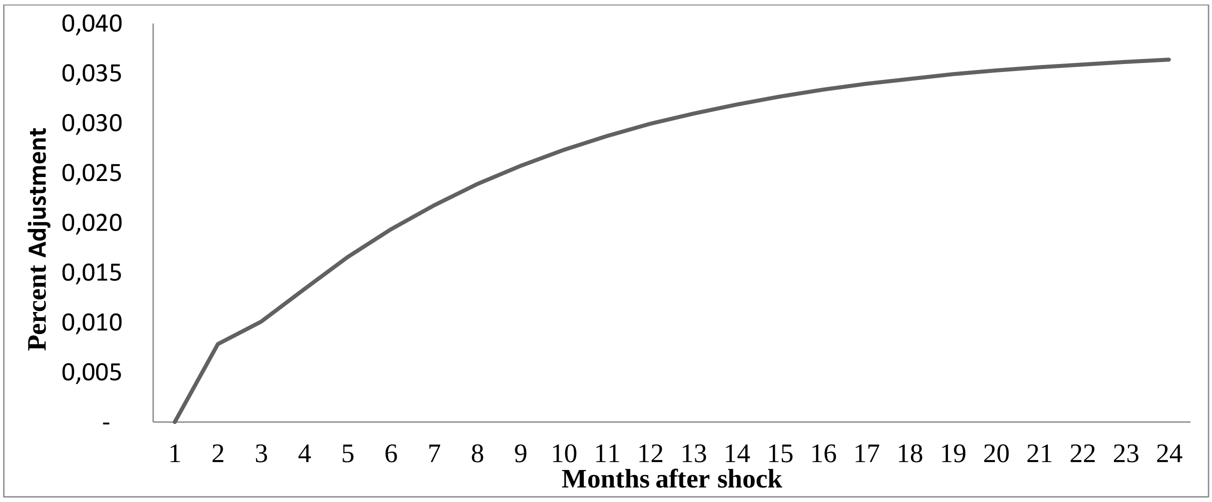

Figure 1: Kenya's price adjustments for wheat production to a 1% shock in the world price.
Source: Authors calculations from regression results.

The results of this study tend to agree with those of previous studies, which found relatively high degrees of price transmissions for food staples from international to domestic markets in SSA (Abidoye & Labuschagne 2014; Ceballos *et al*. 2017; Hatzenboehler *et al*. 2017). Abidoye and Labuschagne (2014) estimated a long-run price transmission elasticity of 98% for maize in South Africa, while Hatzenboehler *et al*. (2017) reported greater than full price transmission of world prices into several maize and rice markets in Nigeria. These earlier estimates compare favourably with our price transmission elasticity estimate of 91% for wheat in Kenya. Ceballos *et al*. (2017) attributed these high transmission elasticities to government policy, which may guide local prices to follow international prices even with minimal trade. This finding of slow speeds of adjustment is consistent with findings from earlier studies (Conforti 2004; Minot 2011; Abidoye & Labuschagne 2014; Baquedano & Liefert 2014; Ceballos e*t al*. 2017). These slow speeds of adjustments have been attributed to market failure/market distortion factors, such as transport and transaction costs, market power, increasing returns to scale in production, product homogeneity and differentiation, exchange rates, border policies, information asymmetry and domestic policies (Rashid & Minot 2010; Baquedano *et al*. 2011; Acharya *et al*. 2012; Abidoye & Labuschagne 2014).

5. Conclusions and policy implications

This study evaluated the transmission of world wheat prices into domestic markets in Kenya using a cointegration approach. All wheat price series were non-stationary and integrated of order one. One cointegrating relationship was established among the three domestic price series (Mombasa, Nakuru and Eldoret), implying that domestic wheat markets in Kenya are integrated. Moreover, world wheat prices are integrated with domestic wheat markets in Kenya, as represented by the domestic wheat reference price at the port of Mombasa. The long run of price transmission elasticity of wheat in Kenya was estimated at 0.91, while the speed of adjustment was estimated at -0.069. The low speed of adjustment of domestic prices to changes in international prices in the face of a high price transmission elasticity suggests the presence of some infrastructural or policy impediments. Even though not analysed explicitly, government border policies, market and infrastructure impediments seem to be underlying causes of the incomplete price pass-through, along with the low speeds of adjustment. Our analysis suggests that the main constraint to a complete pass-through is the existence of price-setting power at the producer level of the wheat market in Kenya.

References

Abate T, Mugo S, De Groote H & Regassa MW, 2015. Maize in Kenya: Chance for getting back to former glory? DT Maize: A Quarterly Bulletin of the Drought Tolerant Maize for Africa Project 4(3): 1–3. Available at https://repository.cimmyt.org/xmlui/bitstream/handle/10883/4477/57029-2015%20v4(3).pdf?sequence=1

Abdulai A, 2000. Spatial price transmission and asymmetry in the Ghanaian maize market. Journal of Development Economics 63(2): 327–49.

Abidoye BO & Labuschagne M, 2014. The transmission of world maize price to South African maize market: A threshold cointegration approach. Agricultural Economics 45(4): 501–12.

Acharya SS, Chand PR, Birthal SK & Negi DS, 2012. Market integration and price transmission in India: A case of rice and wheat with special reference to the world food crisis of 2007/08. Rome: Food and Agriculture Organization.

Balcombe K, Bailey A & Brooks J, 2007. Threshold effects in price transmission: The case of Brazilian wheat, maize, and soya prices. American Journal of Agricultural Economics 89(2): 308–23.

Banerjee A, Galbraith J & Dolado J, 1990. Dynamic specification and linear transformations of the autoregressive-distributed lag model. Oxford Bulletin of Economics and Statistics 52(1): 95–104.

Barrett CB, 2001. Measuring integration and efficiency in international agricultural markets. Review of Agricultural Economics 23(1): 19–32.

Barrett CB & Li JR, 2002. Distinguishing between equilibrium and integration in spatial price analysis. American Journal of Agricultural Economics 84(2): 292–307.

Baquedano FG & Liefert W, 2014. Market integration and price transmission in consumer markets of developing countries. Food Policy 44: 103–14.

Baquedano FG, Liefert W & Shapouri S, 2011. World market integration for export and food crops in developing countries: A case study for Mali and Nicaragua. Agricultural Economics 42(5): 619–30.

Ceballos F, Hernandez MA, Minot N & Robles M, 2017. Grain price and volatility transmission from international to domestic markets in developing countries. World Development 94: 305–20.

Chow GC, 1960. Tests of equality between sets of coefficients in two linear regressions. Econometrica 28: 591–605.

Conforti P, 2004. Price transmission in selected agricultural markets. FAO Commodity and Trade Policy Research, Working Paper No. 7, Commodities and Trade Division of the Food and Agricultural Organization of the United Nations, Rome, Italy.

Dickey DA & Fuller WA, 1979. Distribution of the estimators for autoregressive time series with a unit root. Journal of the American Statistical Association 74(366): 427–31.

Engle R & Granger C, 1987. Co-integration and error correction: Representation, estimation, and testing. Econometrica 55: 251–76.

Esposti R & Listorti G, 2013. Agricultural price transmission across space and commodities during price bubbles. Agricultural Economics 44: 125–39.

Fackler PL & Goodwin BK, 2001. Spatial price analysis. In Gardner BL & Rausser GC (eds.), Handbook of agriculture economics. Amsterdam: Elsevier.

Garrido A, Brümmer B, M'Barek R, Meuwissen MPM & Morales-Opaz C (eds.), 2016. Agricultural markets instability: Revisiting the recent food crises. Oxon UK: Routledge Press.

Gilbert CL, 2010. How to understand high food prices. Journal of Agricultural Economics 61: 398–425.

Gitau R, Mburu S, Mathenge MK & Smale M, 2013. Trade and agricultural competitiveness for growth, food security and poverty reduction: A case of wheat and rice production in Kenya. Tegemeo Institute of Agricultural Policy and Development WPS 45/2011, Nairobi, Kenya.

Hatzenbuehler PL, Abbott PC & Abdoulaye T, 2017. Price transmission in Nigerian food security crop markets. Journal of Agricultural Economics 68(1): 143–63.

Hendry DF, Pagan AR & Sargan JD, 1984. Dynamic specification. In Griliches Z & Intriligator M (eds.), Handbook of econometrics, Volume II. Amsterdam: Elsevier.

Hjalmarsson E & Österholm P, 2010. Testing for cointegration using the Johansen methodology when variables are near-integrated: Size distortions and partial remedies. Empirical Economics 39: 51–76. https://doi.org/10.1007/s00181-009-0294-6

Houck JP, 1986. Elements of agricultural trade policies. New York NY: Macmillan Publishing Company.

Johansen S, 1988. Statistical analysis of cointegration vectors. Journal of Economic Dynamics and Control 12(2): 231–54.

Johansen S, 1995. Likelihood-based inference in cointegrated vector autoregressive models. New York: Oxford University Press.

Johansen S & Juselius K, 1990. Maximum likelihood estimation and inference on cointegration – With applications to the demand for money. Oxford Bulletin of Economics and Statistics 52(2): 169–210.

Kalkuhl M, 2016. How strong do global commodity prices influence domestic food prices in developing countries? A global price transmission and vulnerability mapping analysis. In Kalkuhl M, Von Braun J & Torero M (eds.), Food price volatility and its implications for food security and policy. Cham: Springer.

Kuiper WE, Lutz C & Van Tilburg A, 2003. Vertical price leadership on local maize markets in Benin. Journal of Development Economics 71: 417–33.

Lütkepohl H, 2005. New introduction to multiple time series analysis. New York: Springer.

Liefert W & Persaud S, 2009. The transmission of exchange rate changes to agricultural prices. Economic Research Report No 76, Economic Research Service, United States Department of Agriculture Washington DC.

Listorti G & Esposti R, 2012. Horizontal price transmission in agricultural markets: Fundamental concepts and open empirical issues. Bio-Based and Applied Economics 1(1): 86–96.

Minot N, 2011. Transmission of world food price changes to markets in Sub-Saharan Africa. IFPRI Discussion Paper No. 1059, International Food Policy Research Institute, Washington DC.

Moser C, Barret C & Minten B, 2009. Spatial integration at multiple scale: Rice markets in Madagascar. Agricultural Economics 40: 281–94.

Mundlak Y & Larson DF, 1992. On the transmission of world agricultural prices. World Bank Economic Review 6(3): 399–422.

Musyoka MM, 2009. Wheat import demand and welfare effects of import controls in Kenya. KIPPRA Discussion Paper No. 100, Kenya Institute for Public Policy Research and Analysis, Nairobi, Kenya.

Monroy L, Mulinge W & Witwer M, 2013. Analysis of incentives and disincentives for wheat in Kenya. MAFAP technical notes series, FAO, Rome.

Myers RJ, 2013. Evaluating the effectiveness of inter-regional trade and storage in Malawi's private sector maize markets. Food Policy 41: 75–84. https://doi.org/10.1016/j.foodpol.2013.04.007

Negassa A & Meyers R, 2008. Estimating policy effects on spatial market efficiency: An extension to parity bound models. American Journal of Agricultural Economics 89: 338–52.

Nyangito H, Ikiara MM & Ronge EE, 2002. Performance of Kenya's wheat industry and prospects for regional trade in wheat products. KIPPRA Discussion Paper No. 17, Kenya Institute for Public Policy Research and Analysis, Nairobi, Kenya.

Osterwald-Lenum M, 1992. A note with quantiles of the asymptotic distribution of the maximum likelihood cointegration rank test statistics. Oxford Bulletin of Economics and Statistics 54: 461–71.

Ozturk O, 2020. Market integration and spatial price transmission in grain markets of Turkey. Applied Economics 52(18): 1936–48. https://doi.org/10.1080/00036846.2020.1726862

Phillips PC & Perron P, 1988. Testing for a unit root in time series regression. Biometrika 75(2): 335–46.

Rapsomanikis G, Hallam D & Conforti P, 2003. Market integration and price transmission in selected food and cash crop markets of developing countries: Review and applications. In Commodity market review 2003–2004. Rome: FAO.

Rashid S, 2004. Spatial integration of maize markets in post-liberalized Uganda. MTID Discussion Paper No. 721, International Food Policy Research Institute, Washington DC.

Rashid S & Minot N, 2010. Are staple food markets in Africa efficient? Spatial price analyses and beyond. Paper presented at the COMESA policy seminar on Food Price Variability: Causes, Consequences, and Policy Options, 25–26 January, Maputo, Mozambique.

USDA, 2019. Kenya: Grain and feed annual. Nairobi, Kenya: United States Department of Agriculture (USDA), Foreign Agriculture Service.

Van Campenhout B, 2008. Measuring trends in food market integration: Method and application to Tanzanian maize markets. Food Policy 32: 112–27.

Zorya S, Von Cramon-Taubadel S, Greb F, Jamora N, Mengel C & Würriehausen N, 2014. Price transmission from world to local grain markets in developing countries. Why it matters, how it works, and how it should be enhanced. In Gillson I & Fouad A (eds.), Trade policy and food security: Improving access to food in developing countries in the wake of high world prices. Washington DC: World Bank.