Measuring Integration and Efficiency in Maize Grain Markets: The Case of South Africa and Mozambique

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Abstract:
Price transmission between the South African market and other regional markets is not as straightforward, despite South Africa's role of a surplus producer for the region. There appears to be a host of local factors that must be taken into account in order to anticipate the likely level of regional food prices. This article assesses the degree of market integration and the speed of price adjustment to spatial price differentials between the SAFEX maize price in South Africa and maize grain and maize meal prices in Maputo, Mozambique. The findings of this study indicate that under certain trading regimes, there is no evidence of a long-run relationship between Mozambican and South African maize grain prices. This implies that any large deviations, within these regimes, which exceed transaction costs, could continue to grow with no tendency towards equilibrium. However, the trade volume data indicates maize grain exports from South Africa into Mozambique in every month except for three within the sample set. Hence, the empirical findings of this paper are unexpected given a simple arbitrage argument. Possible reasons for these findings are highlighted in the article. It is interesting to note that when the same empirical analysis is undertaken for the SAFEX maize prices and maize meal prices in Maputo then there is in fact evidence of a long-run relationship between these prices in a high import regime. These findings are not surprising and are what we would expect since two of the largest milling companies, located in Maputo are responsible for the majority of the volume of maize grain imported into the country from South Africa.

Keywords: price transmission, market integration, cointegration, trade regimes

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1. INTRODUCTION

Since 2006, events within global food, energy and financial markets have raised legitimate concerns about food security, particularly for developing countries within the southern African region. Recent experiences from 2007 and 2008 have shown that there exists a complex relationship between South African food prices and those within southern Africa. For instance, domestic maize grain prices in South Africa switched from trading at import parity in 2007 (due to a severe drought) to trading at export parity in 2008 due to a bumper crop that was harvested in 2008. Yet, at the same time world grain prices soared, which implied that the South African export parity prices increased dramatically. Hence, the net effect was that apart from two months South African maize prices mostly traded sideways until the last quarter of 2008 when prices actually started trading softer on the back of much lower international prices. The SAFEX maize price was still following export parity levels at that time. However, over the same period, maize prices in interior areas of southern Africa continued to rise to historically unprecedented levels despite falling global and South African maize grain prices. For example, in December 2008, SAFEX maize grain prices were quoted at $167/MT while prices in Maputo, Mozambique reached a record high of $546/MT. While the direction of the relationship between world prices and those within South Africa change over time due to the switch in trade regimes, there is an undeniable relationship between world prices and those within the South African market purely because of the fact that world prices together with the exchange rate set the import-export parity price band within which the South African grain prices are discovered. Yet, the price transmission between the South African market and other regional markets is not as straightforward, despite South Africa's role of surplus producer for the region. There appears to be a host of local factors that must be taken into account in order to anticipate the likely level of regional food prices.

To this end, this study assesses the degree of market integration and the speed of price adjustment to spatial price differentials between the SAFEX maize price in South Africa and maize grain and maize meal prices in Maputo, Mozambique. Because maize has been consistently exported from South Africa to Maputo almost every month since January 1990, this is a particularly relevant trade route to test for price transmission in the region.

An innovative aspect of our analysis is that it allows for the degree of price transmission and speed of adjustment to vary according to the volume of trade flows between the two markets as might be plausibly expected. Other things equal, consistent and high-volume trade flows might be expected to lead to greater price transmission and shorter times for prices in one market to adjust to changes in prices in the other market, at least in the absence of trade barriers or non-competitive behavior. Moreover, our model allows for the endogenous determination of the monthly trade volumes at which the transmission and co-integration price relationships change.

The paper is organized as follows. Section 2 presents an overview of South Africa's maize grain trade within the region, with particular focus on South Africa-to-southern Mozambique trade routes. Section 3 reviews the relevant literature on price transmission modeling and empirical findings from Africa. Section 4 presents the model used in this study. Sections 5 and 6 describe the data used in the analysis and present the main findings of the paper. The concluding Section 7 summarizes the main findings, discusses the main policy implications, and suggests avenues for future research in the area of price transmission analysis.

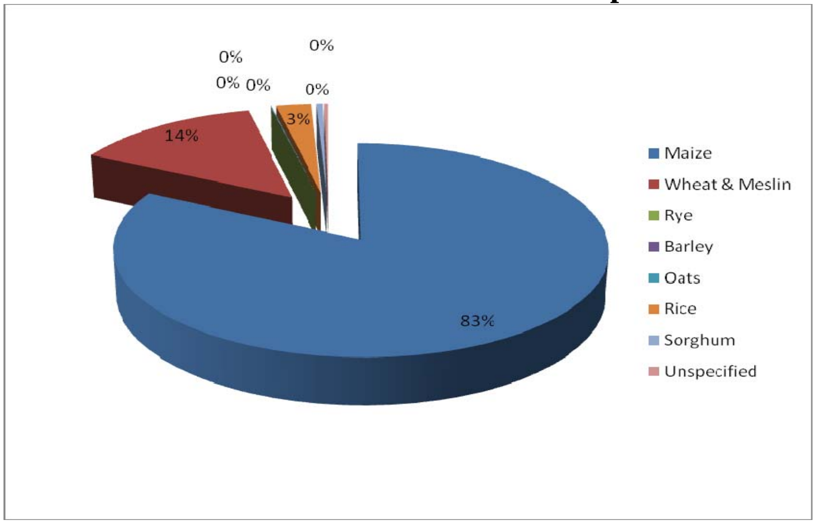
2. REGIONAL GRAIN TRADE: RSA TO SOUTHERN MOZAMBIQUE

2.1. Regional Grain Trade

International trade has taken on an increasingly important role in the South African economy. Over the past two decades, both imports and exports have grown faster than the overall economy. For instance, between 2005 and 2006, Exports' and Imports' percentage share of GDP increased from 26.8% to 29.1% and 28.3% to 33.0%; respectively (World Bank, 2007). This growth rate in trade occurred, despite the overall decline in GDP between 2005 and 2006. In terms of overall trade, between 1992 and 2006, agricultural commodities and goods accounted for 4.4% and 2.3% of total exports and imports; respectively (DTI, 2007).

Of this share, the maize industry is the largest contributing subsector. Figure 1 below illustrates the disaggregated percentage share of total cereal export values. When disaggregated, the maize subsector, between 2007 and 2009 contributed approximately 83% to the total value of cereal exports. This amounted to approximately R8.1 billion, in nominal terms.

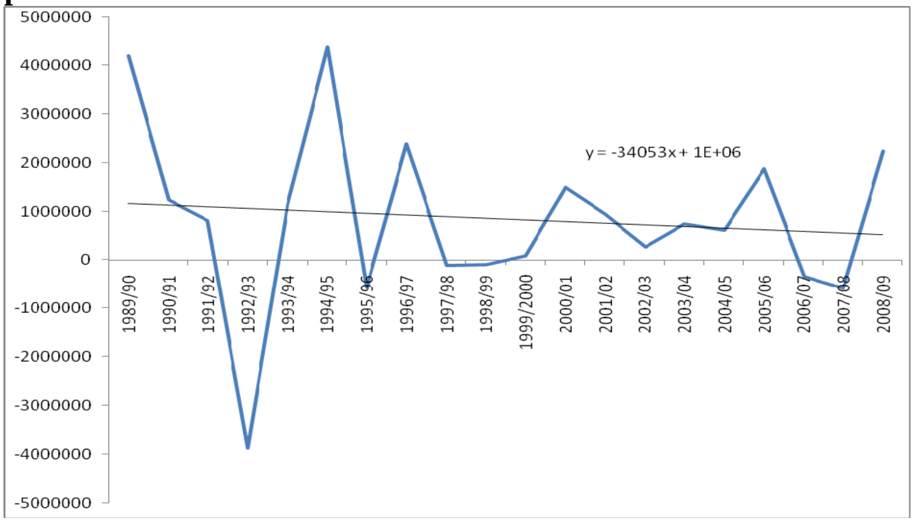
Figure 1: Maize Grain Contribution to Total Cereal Export Values: 2007 to 2009



Source: Department of Trade and Industry, http://www.thedti.gov.za

In general, South Africa's maize grain sector generates a trade surplus in terms of maize grain and its products (see Figure 2 below). It is only in years of drought, that a maize deficit occurs (marketing years 92/93, 95/96 and 2006/07). However, despite maintaining a trade surplus, net export volumes have been decreasing; when a linear trend line is added to the data, the net exports decline at an average rate of about 34,053 metric tons a year.

Figure 2. Net Exports of Maize Grain and Maize Meal: South Africa: 1989/90 to 2008/09 ('000 MT)

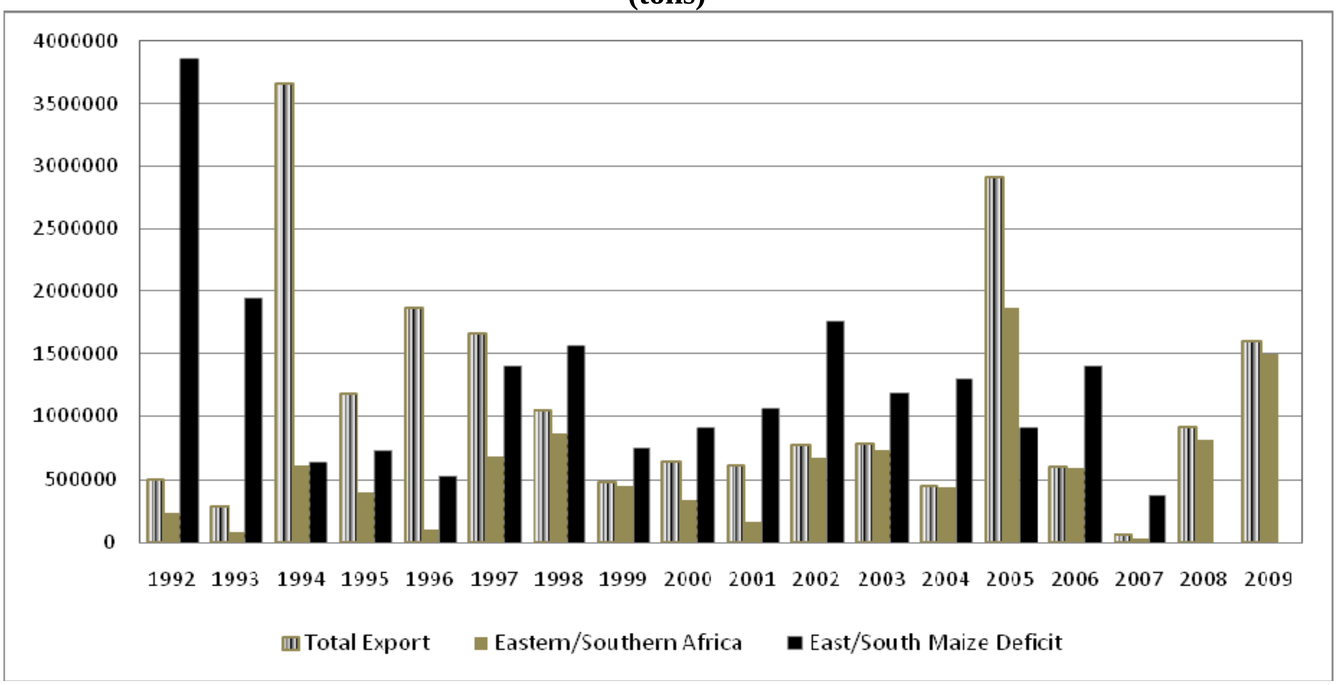


Source: SAGIS, http://www.sagis.org.za/

According to a scenario exercise that was undertaken by the Bureau of Food and Agricultural Policy (BFAP), total exports of white maize could remain above 1 million tons of grain, with a downward trend of approximately 5000 tons of maize grain per year between 2010 and 2017. However, should a drought within the region occur, trade volumes are expected to immediately drop to zero in 2010 and then increase to approximately 1.3 million tons in 2011. Under this scenario, within the context of the current global uncertainty, any regional shock, such as a drought, would have adverse impact on regional self-sufficiency. Under this scenario, the South African market would not be able to supply regional markets with staple maize grain in the short-term. The exact impact of this possible downward trend on regional trade flows is difficult to say without further empirical analysis.

In general, surplus maize grain and meal is exported mainly to BLNS countries (Botswana, Lesotho, Namibia, and Swaziland), Harare in Zimbabwe, Kenya, Mozambique, Zambia, and Mauritius. Figure 3 below depicts South Africa's total volume of maize grain exports disaggregated by destination region.

Figure 3: South African Maize Grain Export Volume and Regional Maize Deficit: 1992 to 2009 (tons)



Source: South African Revenue Services (SARS)

Between 1992 and 1997, less than one-half of total maize grain exports were destined for eastern and southern Africa. However between 1998 and 2009, the percentage share of maize grain exports destined for the region increased significantly. In 1992, this share amounted to approximately 45% of total maize grain exports versus 89% share of total export bound for the region in 2008. Clearly this trend indicates the growing importance of the eastern and southern African region as an output market for South African maize grain, as well as a growing reliance of the region on the South African maize market. In fact, the proportion of the regions grain requirements (measured as total imports minus exports) supplied by the South African market trends upward except for periods of regional drought; (1992, and 2001 marketing years) and during periods of unfavorable pricing conditions. In 2005/06 production season SA plantings decreased by 43% due to the extremely low maize prices in 2005. For many farmers it was not profitable to plant. South Africa went from export parity in 2005 close to import parity in 2006. With the price recovery of 2006 the area recovered completely in the 2006/07 season; however a severe drought kept prices at import parity for 2007. Hence, maize prices remained relatively close to import parity for two seasons and South Africa only serviced long-term clients in across the borders.

In general, the change in the make-up of export markets can be attributed to several factors. These include; the removal of sanctions against South Africa within the southern African region, as well as South Africa's involvement in regional and continental agreements such as the New Economic Partnerships for African Development (NEPAD), African Union (AU), and SADC.

However, it should be noted that despite the growing importance of the interregional trade, South African maize grain traders face several constraints to efficiency. These include (Crichton, 2008):

1. Uncertainty caused by unpredictable export bans, import tariffs, state importation and/or stock releases. For example, during the 2005/06 marketing year, the Zambian government imposed an import duty on maize given its assumption of a high carry over-stock from the previous marketing season due to export bans imposed within the 2004/05 marketing year on maize grain (Fews Net, 2005).
2. Lack of suitable storage facilities within export markets.
3. Lack of sufficient funding on part of regional consumers.
4. Poor quality of maize grain originating within regional markets.
5. Non-tariff trade barriers in terms of non-GMO requirement for white maize. For example, Zambia prohibits GMO maize, while countries such as Zimbabwe, Malawi, and Angola will only allow the importation of milled GMO maize products. Currently, 45% of South African white maize is GMO free. However, given the methods of monitoring the two streams at the silos[4], there exists potential for cross-contamination.

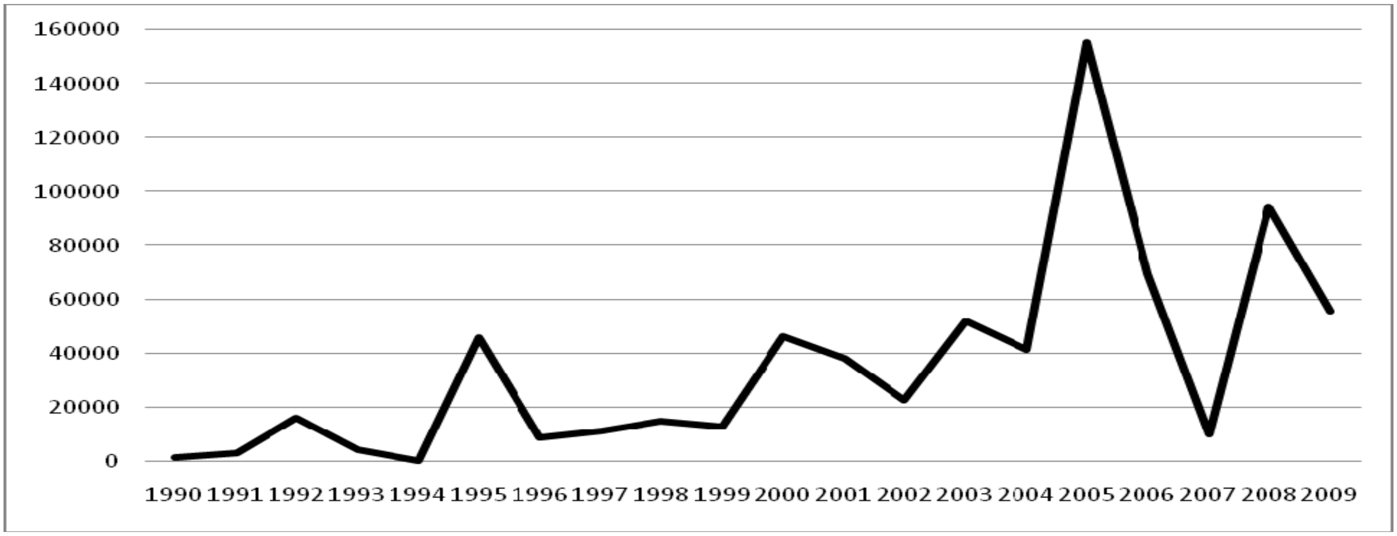
Overall, the extensive trade reforms and market restructuring through most of the 1990's have had a positive impact on the balance of trade in terms of maize grain and products.

2.2 RSA Trade with Mozambique

In the case of Mozambique, the South African market serves as an important source of maize grain, particularly for the southern, deficit producing region of the country. Between 1990 and 2009, maize grain imports from South Africa occurred in nearly every month except for three, indicating a steady flow of grain between the two countries. In general, the share of total RSA maize exports to the region accounted for by Mozambique can range from as little as 0.1% in 1994 to 33% in 2007. Figure 4 below depicts the annual volume of white maize trade flows between RSA and Mozambique.

Figure 4: South Africa to Mozambique Maize Grain Flow: 1990 to 2009 (tons)

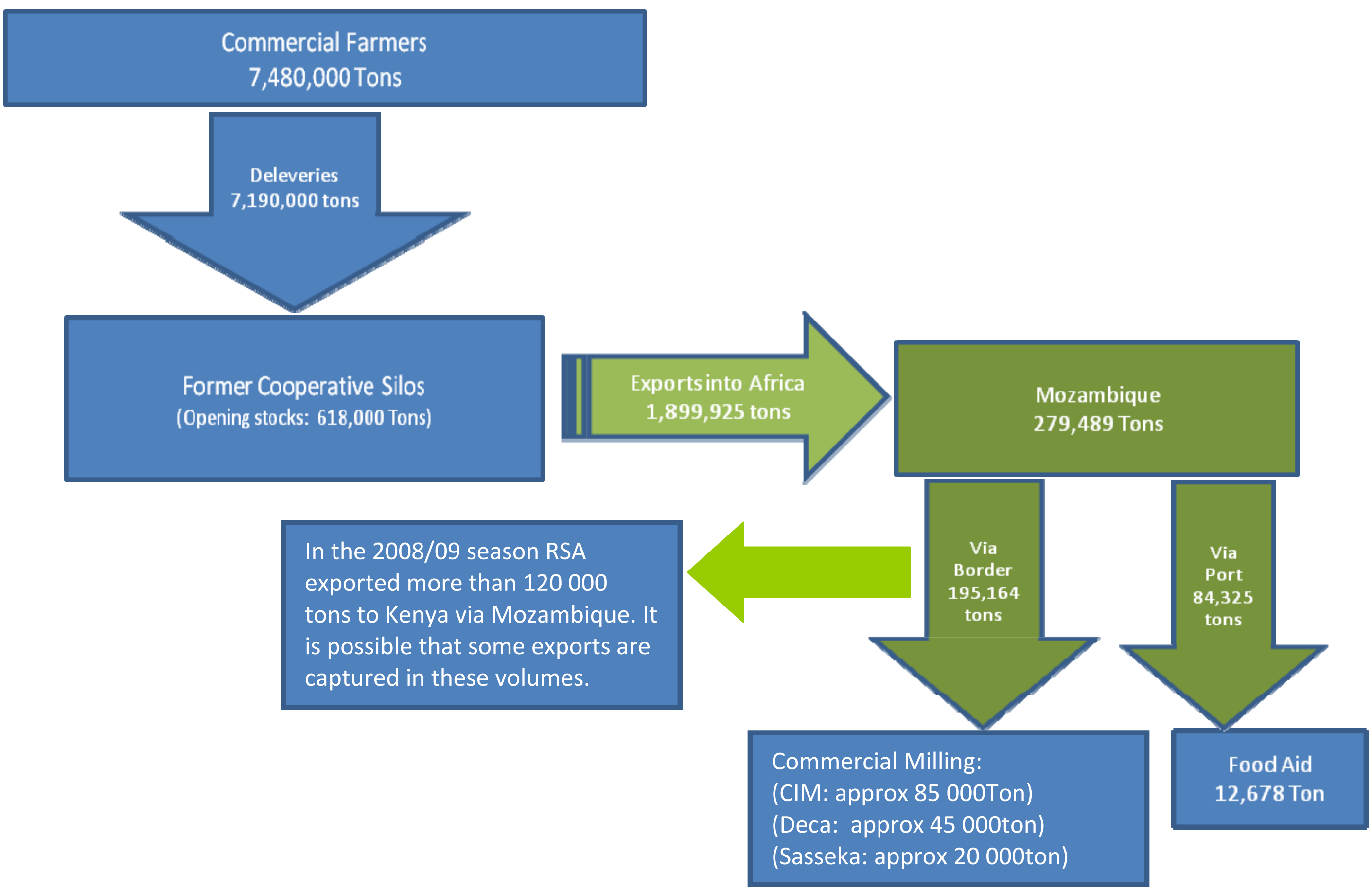
[4] Silo operators, given time restraints, rely on the verbal assurance of the farmer on the GMO status of maize grain delivered rather than testing every truck load of maize delivered (Meyer, 2008)



Source: SARS

On the whole, the total volume of white maize being imported into Mozambique from South Africa has been increasing between 1990 and 2009; with the mean volume of grain traded amounting to approximately 35,240 tons. Figure 5 below illustrates a flow diagram that depicts the flow of maize grain from South Africa into Mozambican market.

Figure 5: White Maize Flow Diagram: RSA to Mozambique: 2008/2009[5]



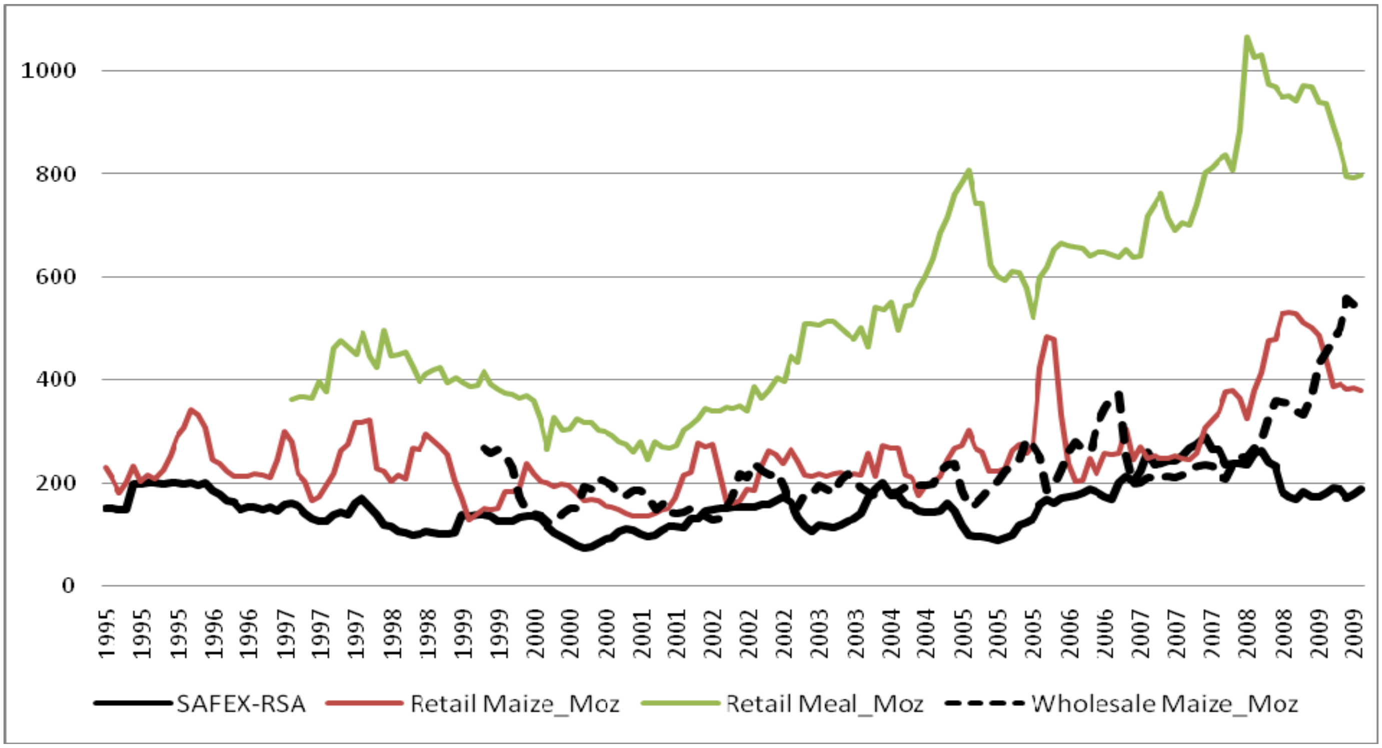
[5] Deliveries refer to maize grain received directly from the farm (not from a commercial storage) of any producer (small and/or large scale) on any premise (silos of cooperatives, processors, etc).

Large grain traders within South Africa dominate the grain export market into the region. These include multi-national companies such as Cargill and Louis Dreyfus as well as former cooperatives such as AFGRI (AFGRI, 2009). In the case of AFGRI exports of maize grain into Mozambique requires cash before delivery and travels via rail and/or ocean. Typically, grain destined for southern Mozambique, specifically Maputo, goes by rail and can be loaded from AFGRI's grain silos located in Wonderfontein, Midddlegurg, Bethal, Ermelo, and Carolina. At the time of this study, average freight rates were estimated at R400/MT (approximately USD 53.00) while administration fees were approximately R20/MT (approximately USD 2.67/MT). Trade administration fees include filing of phytosanitary certification, non-GMO or GMO testing certification, importation license, and payment of a 17% VAT (AFGRI, 2009; CIM, 2009). Imports, from time of loading to off-loading at the final destination, can take from 14 to 21 days (CIM, 2009).

The structure of the milling industry in and around the Maputo area strongly influences the maize grain trade between RSA and southern Mozambique (Tschirley and Abdula, 2007). In 2007 Tschirley and Abdula found that the retail maize meal market was dominated by two large-scale milling enterprises; Companhia Industrial de Matola (CIM) and MEREC (Sasseka) Industries. Together these mills held almost 100% share of the retail maize meal market in Maputo and were responsible for a significant portion of total maize imports into southern Mozambique originating in South Africa (Tschirley and Abdula, 2007). In the 2008/2009 marketing year, it is estimated; assuming full-capacity operation, CIM's grain requirement alone would require 45% of total grain imports into southern Mozambique (CIM, 2009). Moving outside of the Maputo region, a third large-scale mill (Deca) operates in Shimoio which has an estimated capacity of 40 000 tons. It is important to note that these three large-scale mills have the ability to reclaim the import tax (EVA) of 17% on maize grain. Yet this process is complex, takes time and seems to require significant lobbying by the large role players to eventually receive the import tax back from government. This can be regarded as a serious trade barrier for smaller traders that do not have the lobbying power or the financial means to carry the costs of the import tax over a longer period of time. In general, within the Maputo market, maize grain at the retail level is thinly traded and is entirely of domestic origin (Tschirely and Abdula, 2007). The imported maize from South Africa is directly sourced by the large companies and processed for maize meal. Quality issues are frequently encountered with the maize that is produced locally, mainly due to a general lack of infrastructure like storage and efficient means of transport.

Hence, it becomes clear that it is unlikely for maize grain prices in Mozambique to follow the SAFEX prices in South Africa, despite of the fact that Mozambique is a net importer of maize from South Africa. The apparent disjoint between the maize grain and meal price trends, lends further evidence to the disjoint between wholesale markets within the two countries. Figure 6, below, illustrates monthly SAFEX and Maputo wholesale/retail maize grain and meal prices between January 1995 and September 2009.

Figure 6. Maize Grain and Meal Prices: 1995 to 2009
(Nominal USD/MT)



From the price trends it is clear that there is very little co-movement between RSA (as measured by SAFEX) and Maputo wholesale maize grain prices. There are periods where wholesale prices in Maputo increased while SAFEX prices declined; specifically 2000/2001 and 2003. When correlation coefficients are calculated for the various price series we find little evidence for a linear relationship between wholesale and retail maize grain in Maputo and SAFEX, whereas some indication in the case of retail maize meal prices in Maputo and SAFEX grain prices. Table 1 below summarizes the coefficient measures:

Table 1. RSA and Mozambican Maize Grain and Meal Coefficient Measures

	SAFEX
SAFEX	1.0
Retail Maize	0.475963
Retail Meal	0.673249
Wholesale Maize	0.348163

3. REVIEW OF TRANSMISSION LITERATURE

There is an extensive literature exploring the degree to which shocks are transmitted among spatially or vertically separated markets via price adjustment[6]. Some early studies estimated an asymmetric adjustment function (e.g., Wolfram, 1971; Houck, 1977; and Ward, 1982) of the form:

$$\Delta P_t^{out} = \alpha + \alpha_1^+ \sum_{t=1}^{T} D^+ \Delta P_t^{in} + \alpha_1^- \sum_{t=1}^{T} D^- \Delta P_t^{in} + \varepsilon_t \qquad (1)$$

where P_t^{out} represents prices at time t in the destination market, P_t^{in} prices in the source market; D^+ and D^- are dummy variables with $D^+ = 1$ if $P_t^{in} \geq P_{t-1}^{in}$ and $D^+ = 0$ otherwise; $D^- = 1$ if $P_t^{in} < P_{t-1}^{in}$ and $D^- = 0$ otherwise. This model allows for two price adjustment coefficients to be estimated; α_1^+ and α_1^- for periods of rising and

[6] See Meyer and von Cramon-Taubadel (2004); Fackler and Goodwin (2002) for a comprehensive survey of the literature.

decreasing input prices, respectively. The null hypothesis of symmetric price transmission between the two markets is rejected if $\alpha_1^+ \neq \alpha_1^-$.

One important requirement for the first difference model (1) to be valid is that the price series be nonstationary but not cointegrated. Commodity price series often display evidence of unit roots (nonstationarity) but also feature a long-run equilibrium relationship that connects them in the long-run. This leads to a situation where unit root tests indicate that P^{in} and P^{out} are nonstationary but a linear combination of them is stationary, in which case the prices are cointegrated. In this case regressions using the first difference form (1) will lead to incorrect inference because they ignore the cointegrating relationship (von Cramon-Taubadel and Fahlbusch, 1994; and von Cramon-Taubadel, 1998).

The existence of a long-run cointegrating relationship between P^{in} and P^{out} implies the need for an error correction model (ECM) that incorporates asymmetric adjustment terms (Engle and Granger, 1987; Granger and Lee, 1989; von Cramon-Taubadel, 1998, von Cramon-Taubadel and Loy, 1996). The asymmetric ECM model can then be represented by the following equation;

$$\Delta P_t^{out} = \alpha_0 + \sum_{j=1}^{k} (\alpha_j^+ D^+ \Delta P_{t-j+1}^{in}) + \sum_{j=1}^{L} (\alpha_j^- D^- \Delta P_{t-j+1}^{in}) + \phi^+ ECT_{t-1}^+ + \phi^- ECT_{t-1}^- + \upsilon_t \quad (2)$$

Here the ECT term represents the residual from the linear estimated relationship between market prices and transactions costs over time and can be expressed as;

$$ECT_t = P_t^{out} - \hat{\beta}_0 - \hat{\beta}_1 P_t^{in} - \hat{\beta}_2 \tau_t \quad (3)$$

where τ_t represents the cost to transfer goods between the source and destination markets, assuming continuous unidirectional trade the equilibrium relationship between P^{out} and P^{in}. The residual, which represents an I(0) random variable, measures the short-term deviations from equilibrium caused by shocks to price and/or transaction costs. Including the lagged residual term in the ECM allows for dynamic inter-market price adjustment to their long-run equilibrium state.

However, this model assumes linear error correction (i.e. constant parameters ϕ^+ and ϕ^-) whereby any deviation from long-run equilibrium, regardless of its size, will be corrected at the same rate. This model specification is appropriate only if we assume a single trade regime (continuous, unidirectional trade). If there are shifts between trade and non-trade regimes over time then small shocks to the long-run equilibrium may have different dynamic adjustment effects than large shocks, because small shocks may not induce a change in trade regimes while large shocks do.

Consequently, threshold cointegration models [7] have been used to test spatial market equilibrium allowing for different dynamic adjustments depending on whether shocks to equilibrium fall over or under a threshold level. Equation 4 illustrates how this type of model can be specified (Meyer and von Cramon-Taubadel, 2004):

[7] The threshold approach was first introduced by Tong (1983) and later modeled and tested by Balke and Fomby (1997), Tsay (1989), Goodwin and Holt (1999), Goodwin and Harper (2000); and Goodwin and Piggott (2001).

$$\Delta P_t^{out} = \begin{cases} \alpha_0 + \sum(\alpha_{1j}\Delta P_{t-j+1}^{in}) + \phi_1 ECT_{t-1} + v_t & if \quad ECT_{t-1} < C_1 \\ \alpha_0 + \sum_{j=1}^{k}(\alpha_{2j}\Delta P_{t-j+1}^{in}) + \phi_2 ECT_{t-1} + v_t & if \quad C_1 \leq ECT_{t-1} \leq C_2 \\ \alpha_0 + \sum_{j=1}^{k}(\alpha_{3j}\Delta P_{t-j+1}^{in}) + \phi 3 ECT_{t-1} + v_t & if \quad ECT_{t-1} > C_2 \end{cases} \qquad (4)$$

In this case the thresholds to be estimated are given by C_1 and C_2, and whenever ECT falls within the $[C_1, C_2]$ interval, the deviations from the long-run equilibrium, which, when compared to adjustment costs, are so small they will not lead to a price adjustment (Goodwin and Piggott, 2001). This interval is known as the neutral band. In a spatial context, transactions costs, associated with trade between two different markets, may be used to estimate the thresholds. In this case, the thresholds may arise when the price differential between the source and destination markets greatly exceed transaction costs (Meyer and von Cramon-Taubadel, 2004).

However, there are two points that should be made with regards to this model. First, data limitations on transaction costs can make it very difficult to get accurate estimates of a plausible threshold that corresponds to the minimum incentives required to encourage price adjustment (i.e. trade between the source and destination market). Secondly, given the model formulation, all parameters can potentially vary across regimes with the exception of the β coefficients in the long-run equilibrium relationship. This assumption of linearity in the ECT does not take into consideration spatial equilibrium theory that suggests that different long-run equilibrium conditions may potentially exist between trading versus non-trading regimes.

The development of a method to test for threshold effects in the cointegration relationship itself by Gonzalo and Pitarakis (2006) enabled Stephen et. al. (2008) to specify a switching error correction model (SECM) that allowed for potential differences in long-run price relationships between markets under different trading regimes. The model specification is as follows:

$$\begin{aligned} \Delta P_t^{out} &= \left(\sum_{j=1}^{L} (\alpha_{1j}^{trade} \Delta P_{t-j+1}^{in}) + \phi^{trade} ECT_{t-1}^{trade} \right) I_t^{trade} \\ &+ \left(\sum_{j=1}^{H} (\alpha_j^{non-trade} \Delta P_{t-j+1}^{in}) + \phi^{non-trade} ECT_{t-1}^{non-trade} \right) I_t^{non-trade} + v_t \end{aligned} \qquad (5)$$

Here $ECT_{t-1}^k = P_{t-1}^{out} - \hat{\beta}_0^k - \hat{\beta}_1^k P_{t-1}^{in} - \hat{\beta}_2^k \tau_{t-1}, \quad k \in \{trade, non-trade\}$ and I_t^{trade} and $I_t^{non-trade}$ are indicator functions for specific trade regime into which each time period, *t*, falls. The usefulness of the SECM model is that it will allow for complete variation of all parameters across trading regimes. This allows not only tests for differences between the speed of the market's response to shocks in each regime (the ϕ^{κ} parameters) but also, to determine if different mechanisms, besides trade, are responsible for price adjustments between spatially separated markets ($\phi^{non\text{-}trade}$ parameter). The key assumption in the literature is that physical arbitrage is the mechanism by which spatially separated markets return to equilibrium. Therefore, any evidence of cointegration between markets in non-trade periods would suggest that other factors (for example, information flow) besides physical trade flows may play a role in bringing about spatial competitive equilibrium through price adjustments. Given this then the null hypothesis is that $\phi^{non\text{-}trade} = 0$, which implies that any price adjustment within the destination market occurs only under trade. Also, under the null, $\beta^{trade} = \beta^{non\text{-}trade}$, i.e. the cointegrating vector is the same regardless of trading regimes.

4. METHODOLOGY

The SECM model, as implemented by Stephen et. al. (2008) uses exogenous sample separation in order to test for spatial price adjustment under differing trade regimes rather than estimation of threshold parameters. However, undertaking sample separation at zero trade is somewhat arbitrary since the price adjustment process back to the long-run equilibrium may in fact vary across periods of high versus low trade. For instance, in periods of high trade we might expect the price difference to better reflect transfer costs, therefore resulting in a stronger price transmission between to the markets. Whereas, in low-import regimes, the price connection between the two markets may be broken; fundamentally altering the rate and degree of price transmission. In order to account for price transmission regime changes at a trade threshold the model specification used in this study is represented by:

$$\begin{aligned}\Delta P_t^{out} &= \left(\sum_{j=1}^{L}(\alpha_{1j}^{low-trade}\Delta P_{t-j+1}^{in}) + \phi^{low-trade}ECT_{t-1}^{low-trade}\right)I_t^{low-trade} \\ &+ \left(\sum_{j=1}^{H}(\alpha_{j}^{high-trade}\Delta P_{t-j+1}^{in}) + \phi^{high-trade}ECT_{t-1}^{high-trade}\right)I_t^{high-trade} + v_t\end{aligned} \qquad (6)$$

All variable definitions remain the same, except now the two regimes defined are low versus high trade periods where;

$$\begin{aligned}I_t^{low-trade} &= 1 \quad \text{if} \quad q_t \le C \quad \text{zero otherwise} \\ I_t^{high-trade} &= 1 \quad \text{if} \quad q_t > C \quad \text{zero otherwise}\end{aligned} \qquad (7)$$

Here, q_t represents the quantity of imports from South Africa; C is the import threshold at which the price transmission process shifts. If q_t is exogenous, then equation (6) can be estimated as a conventional threshold model using a grid search procedure to locate the appropriate threshold that minimizes the sum of squared residuals over the full sample. In the case of South Africa and Mozambique, exogeneity of q_t is a reasonable assumption when using monthly trade data since it can take at least a month to arrange and transport imports (see Section 2). This implies that the decision to have imports flow at *t* was instigated at *t-1* and cannot respond to contemporaneous observations of RSA and/or Mozambican prices. One slight complication is that it may be appropriate to impose different stationarity and/or cointegration restrictions in the two regimes. For example, it may be that variables are nonstationary in both regimes but only cointegrated in the high import regime. This possibility can be accounted for by imposing the appropriate restrictions on (6) and then applying the appropriate estimation procedure given the restrictions in each regime.

The form of the appropriate estimation equation depends on the stochastic properties of the underlying data being used. In the case of South Africa and Mozambique, over the entire sample period, there is little statistical evidence of cointegration between the price series; however this does not allow us to infer no cointegration within different trade regimes. Given this uncertainty we estimate multiple forms of the transmission model and the results checked for robustness across different identification assumptions. The three assumptions we allow for are cointegration, stationarity, and partial cointegration.

4.1 Cointegration

If we assume there is at least one cointegrating relationship between the variables within our system and that this cointegrating vector is given by the long-run spatial equilibrium relationship between domestic Mozambican and South African maize grain prices we could use a Vector Error Correction (VEC) representation and estimate the model using the standard Engle-Granger two-step approach or Johansen maximum likelihood methods of estimation. However, following Phillips and Loretan (1991) we use a single equation error correction model (SEECM) which has the advantage of requiring a single-equation estimation for each trade regime. Given our assumptions, our model takes the form:

$$\Delta M_t = \alpha_0 + \beta_1 \Delta S_t + \beta_2 \Delta D + \beta_3 \Delta T_t + \lambda\left(M_{t-1} - \beta_{1j} S_{t-1} - \beta_2 D_{t-1} - \beta_3 T_{t-1}\right) + \sum_{i=1}^{n} A_i (\Delta M_{t-i} - \beta_1 \Delta S_{t-i} - \beta_2 \Delta D_{t-i} - \beta_3 T_{t-1}) \\ + \sum_{i=1}^{n} b_i \Delta S_{t-i} + \sum_{i=1}^{n} c_i \Delta D_{t-i} + \sum_{i=1}^{n} e_i \Delta T_{t-1} + \varepsilon_{pt} \tag{8}$$

where: M = wholesale or retail maize prices in Maputo, S = SAFEX spot prices, D=diesel prices, T = 17% ad valorem tariff, and α_0 is a constant term. The β parameters represent the long-run price transmission effect (the long-run equilibrium relationship between prices and transfer costs), while λ is the "speed of adjustment" parameter which determines how quickly the long-run equilibrium reasserts itself after a shock. Optimal asymptotic Gaussian inference on (8) can be achieved by choosing *n* to eliminate residual autocorrelation and applying nonlinear least squares (NLS); see Phillips and Loretan (1991).

Although there is no generally accepted criterion for what constitutes "rapid" adjustment, one commonly reported summary measure used for comparative purposes is the "half-life" (the number of periods it takes for half of the full adjustment back to long-run equilibrium to occur). The half-life is often computed using the formula $hl = \ln(0.5)/\ln(1+\lambda)$ but this formula assumes shocks decay at a constant rate, which may not be true when (4) contains higher order autoregressive terms (see Murray and Papell, 2004). In this case half-lives can be computed using simulation.

4.2. Stationarity

If we assume that all variables are stationary and that our explanatory variables, S, D, and T are exogenous then a convenient single-equation form for the cointegrating relationship under stationarity can be given by:

$$\Delta M_t = \alpha_0 + \beta_1 \Delta S_t + \beta_2 \Delta D_t + \beta_3 \Delta T_t + \lambda\left(M_{t-1} - \beta_{1j} S_{t-1} - \beta_2 D_{t-1} - \beta_3 T_t\right) \\ + \sum_{i=1}^{n} A_i (\Delta M_{t-i} - \beta_1 \Delta S_{t-i} - \beta_2 \Delta D_{t-i} - \beta_3 \Delta T_t) \\ + b_0 S_{t-1} + \sum_{i=1}^{n} b_i \Delta S_{t-i} + c_0 D_{t-1} + \sum_{i=1}^{n} c_i \Delta D_{t-i} + e_0 T_{t-1} + \sum_{i=1}^{n} e_i \Delta T_t + \varepsilon_{pt} \tag{9}$$

Early work on spatial price transmission was often undertaken in this framework (see references). Optimal asymptotic Gaussian inference on (5) can again be obtained by choosing *n* to eliminate residual autocorrelation and applying NLS.[8] The long-run transmission parameters β and speed of adjustment parameter λ have the same interpretation as in the cointegration model. Furthermore, these parameters are identified and NLS provides optimal Gaussian inference, but only under the assumption that S_t, D_t, and T_t are exogenous. If the exogeneity assumption does not hold, then β estimates will contain simultaneous equations bias under the stationarity assumption, unless an instrumental variables procedure is used. However, exogeneity is a reasonable assumption in the current context. There is unlikely to be contemporaneous feedback from Mozambique prices to RSA prices because Mozambique markets are small relative to the RSA market and trade only flows one way out of RSA.

[8] Actually, the stationary model is linear and could be estimated using ordinary least squares. However, using NLS allows direct estimation of the long-run transmission parameters β and the speed of adjustment parameter λ, which facilitates comparison with other model forms.

4.3. Partial Cointegration

Another set of assumptions we could make are that S_t, and D_t, are nonstationary and cointegrated while our price series are a stationary linear combination of these cointegrating vectors. In other words, we assume a partial cointegration and that S_t is exogenous but allow for contemporaneous feedback between diesel prices and local maize prices. The model form under these assumptions is as follows:

$$\begin{aligned}\Delta M_t &= \alpha_0 + \beta_1 \Delta S_t + \beta_2 \Delta D_t + \beta_3 \Delta T_t + \lambda\left(M_{t-1} - \beta_{1j} S_{t-1} - \beta_2 D_{t-1} - \beta_3 T_t\right) \\ &+ \sum_{i=1}^{n} A_i (\Delta M_{t-i} - \beta_1 \Delta S_{t-i} - \beta_2 \Delta D_{t-i} - \beta_3 \Delta T_t) \\ &+ b_0 [S_{t-1} + (\frac{\beta_2}{\beta_1}) D_{t-1}] + \sum_{i=1}^{n} b_i [\Delta S_{t-i} + (\frac{\beta_2}{\beta_1}) \Delta D_{t-i}] + \sum_{i=1}^{n} c_i \Delta D_{t-i} + \sum_{i=1}^{n} e_i \Delta T_t + \varepsilon_{pt}\end{aligned} \tag{10}$$

As with the Cointegrating and Stationary version of the model, optimal asymptotic Gaussian inference on (11) can be obtained by choosing n to eliminate residual autocorrelation and applying NLS. In this case, β_1 has the same long-run price transmission interpretation as before (the long-run change in p_t resulting from a permanent one unit change in S_t, holding D_t and T_t constant). But in this case it is important to note that for S_t to change permanently without a corresponding permanent change in D_t requires a permanent change in the long-run equilibrium relationship between S_t and D_t. Hence, in this case β_1 really represents the long-run response of M_t to a permanent unit change in the equilibrium relationship between S_t and D_t, while λ still characterizes the speed with which the resulting adjustment takes place.

4.4 Summary

Each of the estimation equations (8)-(10) provides a slightly different single equation model for estimating both the extent of long-run price transmission and the speed of adjustment. Each form is derived under a different set of identification assumptions. In the cointegration model (8) all variables are assumed nonstationary with a single cointegrating vector characterizing the long-run spatial price relationship. No exogeneity assumptions are required. In the stationary model (9) all variables are assumed stationary and S_t and D_t exogenous. In the partial cointegration model (11) S_t and D_t are nonstationary and cointegrated but M_t is stationary. In this case S_t is also assumed exogenous.

Alternative identification assumptions are also possible. For example, it could be assumed that M_t and S_t are nonstationary and cointegrated but D_t is stationary ($\beta_1 = 1$, $\beta_2 = 0$, and $|\gamma_1| < 1$, with transfer costs also exogenous. Or that M_t and D_t are nonstationary and cointegrated but S_t is stationary ($|\beta_1| < 1$, $\beta_2 = 0$, and $\gamma_1 = 1$, with S_t also exogenous). These alternative identifications would lead to slightly different versions of the estimation equations (6). However, estimation equations for these identifications are not provided explicitly here because these situations are unlikely to be relevant to the particular application studied here (see the preliminary estimation results below). It is also possible that both prices and transfer costs are all nonstationary and there are no cointegrating relationships among them, leading to an estimation form that includes only first differences of variables. In this case, however, no equilibrium relationship exists and there is no-long run price transmission. So estimating this form of the model would not be helpful in drawing inferences about the extent of price transmission (zero transmission is a maintained assumption).

5. DATA AND DESCRIPTIVE RESULTS

5.1 Data

The models specified in this study are estimated using monthly data on retail maize meal, maize grain as well as wholesale maize grain prices within Maputo, Mozambique; transportation costs; tariffs; and trade flows of white maize grain between Randfontein, South Africa and Maputo.

The monthly maize price data were converted from local currency into U.S. Dollars (USD) using monthly exchange rates. The data were acquired from the national statistical agencies within the respective countries. The specific sources in Mozambique include the Ministry of Agriculture, Agricultural Market Information Center (SIMA), and the National Statistics Institute (INE) of Mozambique. In the case of South Africa, SAFEX white maize spot prices were used, as provided by the South Africa Grain Information System (SAGIS), while exchange rate data was sourced from the Statistical Agency of South Africa. The sample period ranged from March 1997 to September 2009 for retail maize meal prices; January 1995 to September 2009 for retail maize grain prices, and October 1998 to January 2009 for Wholesale maize grain prices.

Trade volume and value data was provided by the South African Revenue Services (SARS) and contains monthly exports of maize grain from South Africa into Mozambique. The volumes recorded include grain intended for commercial, private, and/or food aid use. The import value data was used to calculate the 17% value-added tax on grain imported into Mozambique from South Africa. Although this tax is intended for maize grain imports that will not require further processing, it is applied to all maize grain imports; with reimbursements being distributed to grain processors. However, according to local processors, there is a lag, which could range from 6 to 12 months, between paying the tax and receiving the reimbursement. For this reason we include the tariff as a possible cost to cross-border trade.

In light of the fact that transportation costs are not publicly available, monthly wholesale diesel prices within South Africa were used as a proxy instead.

5.2. Unit Root and Cointegration Tests

In conducting our analysis of cointegration, we begin by testing for the presence of a unit root for each price series. Table 2 presents the p-values for augmented Dickey-Fuller and Phillips-Perron test which were computed using Mackinnon's asymptotically valid distribution. The tests were first applied to the full sample and the results support the evidence of nonstationarity for all variables except for retail maize grain prices in Mozambique and tariffs. This implies that within our sample, there is not cointegrating relationship between SAFEX and domestic retail maize grain prices; therefore in this case, the partial cointegration model (11) may be the most appropriate for the full sample estimation.

Since we are interested in determining whether or not the cointegrating relationship changes as we move from one trade regime to another, we divided the sample set into two sub-samples; "Low" and "High" imports from South Africa, using the median level as a divisor. The results of nonstationarity continue to be supported in both low and high import trade regimes for SAFEX, Transportation, domestic wholesale maize and retail maize meal prices, while Tariffs remain stationary within each regime. However, for retail maize grain, the results change suggesting nonstationarity in both individual regimes. This result suggests that when a threshold is used to separate the sample the evidence supports use of the cointegration model rather than the partial cointegration model.

Table 2: Probability Values for Unit Root Tests

Test	Mozambique Prices			RSA Price	Diesel	Tariff
	P_{rmz}	P_{rmm}	P_{wmz}	SAFEX		
Full Sample						
-ADF	0.0235	0.8190	0.7723	0.2331	0.5935	0.000
-PP	0.0690	0.7943	0.9590	0.2546	0.5385	0.000
Low Imports						
-ADF	0.7521	0.9893	0.9234	0.5454	0.9869	0.000
-PP	0.8220	0.9919	0.9905	0.5686	0.9650	0.000
High Imports						
-ADF	0.0833	0.3364	0.8333	0.1890	0.1917	0.000
-PP	0.1669	0.3536	0.9432	0.2820	0.0921	0.000

In general, if two or more time series are themselves non-stationary (i.e. they exhibit a unit root), but a linear combination of them is stationary, then the series are said to be cointegrated and any deviation from the mean between the price series, is not expected to continue. In a spatial context, the spread will have the tendency to return to the long-run equilibrium relationship that satisfies the LOP. To test for cointegration between our price series we used two tests, the Engle-Granger two-step method and the Johansen procedure. These tests were applied to three models in which retail maize grain, retail maize meal and wholesale maize grain prices in Mozambique were the dependent variables. The results are presented in Table 3 below.

Table 3. Unit Root and Cointegration Tests for the Entire Sample Period

Test	Test Statistics	Critical Value[9]
Engle-Granger Two-Step Method		
Retail Maize – no linear time trend		
Lagged residual	-3.40*	-3.34
Retail Maize Meal-no linear time trend		
Lagged residual	-2.81	-3.34
Wholesale Maize-no linear time trend		
Lagged residual	-0.58	-3.34
Johansen method (r=0)		
Retail Maize		
Maximum eigenvalue test statistic	113.53	27.07
Trace test statistic	164.19	47.21
Retail Maize Meal		
Maximum eigenvalue test statistic	117.25	27.07
Trace test statistic	144.9	47.21
Wholesale Maize		
Maximum eigenvalue test statistic	98.79	27.07
Trace test statistic	123.16	47.21

The results of the Engle-Granger Two-step methods indicate that regardless of whether we are modeling wholesale grain prices or retail maize meal prices, there is no evidence of cointegration between Mozambican, SAFEX, Diesel and Tariff price series. The test statistics on the lagged residuals are above the critical value of -3.34 at the 5% level of significance. In other words, there is no long-run relationship between M and S, within our sample period; i.e. the price differential between these two markets could become very large with no tendency for them to come back together, while controlling for transportation costs. Based on a simple arbitrage argument, this result is surprising since within the sample period, every month, except for 3 (9/96,

[9] ADF and Engle-Granger test at 5% level of significance, Johansen at the 95% level of significance.

10/96 and 8/02), maize grain was exported into Mozambique from South Africa; indicating sufficient incentive within the market for traders to engage in trade and possibly move the price differential back to some equilibrium value.

In order to ensure robustness of our cointegration testing, we used the Johansen procedure to test the three models for cointegration. In all three cases, the Johansen's maximum eigenvalue statistics as well as the trace statistics confirms the finding of at least one cointegrating vector; which is what we would expect. The null hypothesis of Johansen trace test statistics is that there are at most r (0 < r < k) cointegrating vectors and thus (k-r) common stochastic trends. The null is rejected if the test statistics exceed their critical values. In the case of retail maize meal prices, because the trace statistic at r = 0 of 117.25 exceeds its critical value of 27.07, we reject the null hypothesis of no cointegrating equations. In contrast, the trace statistic at r = 1 of 27.65 is less than its critical value of 29.68 for retail maize meal prices, and r = 2 of 10.03 is less than its critical value of 15.41 for retail maize grain prices, we cannot reject the null hypothesis that there are one or fewer cointegrating equations and two or fewer cointegrating equations within retail maize meal and maize grain models; respectively. For the wholesale maize grain we found evidence of at least one cointegrating vector within the model.

6. PRICE TRANSMISSION RESULTS FOR SOUTH AFRICA - MOZAMBIQUE TRADE

Price transmission results for both the wholesale and retail prices are presented in the tables below. Given that our main interest is in the long-run price transmission between the dependent variables and the explanatory variables (β_1 and β_2); as well as with the speed of adjustment (λ), only the estimation results for the relevant coefficients are presented. The parameters not reported determine the autocorrelation structure of the data and it is important to note that *n*=1 lags was sufficient to eliminate autocorrelation in the residual of all model forms.

Goodness of fit results are also included in the tables. Model R^2 is reported but should be interpreted with care because all models are estimated with the dependent variable in difference form to facilitate direct estimation of the speed of adjustment parameter. Each of the models has an alternative representation with the dependent variable in levels form, which would increase the reported R^2 considerably (even though the two representations are formally equivalent). The *F* column in the tables is the *p*-value for testing the null hypothesis that the associated regression model has no explanatory power. Results show that all forms of the regression model have significant explanatory power.[10]

6.1 Wholesale Maize Grain Prices

Table 3. Price Transmission Result for Wholesale Maize Grain Prices: 10/1998 to 01/2009

Model	β_1	β_2	β_3	λ	Half-life	No. Obs.	R^2	F-statistic
Full Sample								
Cointegration	0.419	68.896	0.163	-0.128	5.06	122	0.2494	0.0006
	(0.50)	(23.11)*	(0.25)	(0.04)*				
Partial Cointegration	0.066	10.876	-0.352	-0.128	5.06	122	0.2494	0.0006
	(0.16)	(28.74)	(0.531)	(0.04)*				
Stationary	0.066	-5.469	-0.022	-0.128	5.06	122	0.2494	0.0006
	(0.16)	(7.39)	(0.04)	(0.04)*				

[10] Notice that for a given sample or subsample all of the goodness of fit statistics, and the estimated speed of adjustment parameter and its standard error, are identical. This is because the different model forms (cointegration, stationary, partial cointegration, and stationary transfer costs) use different just identifying assumptions to estimate the β s but each model has the same explanatory power for changes in the domestic maize price.

Imports <=2633								
Cointegration	-0.002	73.369	0.309	-0.166	3.82	68	0.4808	0.0000
	(0.45)	(23.06)*	(0.30)	(0.05)*				
Partial Cointegration	-0.205	1299.735	-0.397	-0.166	3.82	68	0.4808	0.0000
	(0.15)	(49920.65)	(0.32)	(0.05)*				
Stationary	-0.205	8.423	-0.022	-0.166	3.82	68	0.4808	0.0000
	(0.15)	(6.39)	(0.03)	(0.05)*				
Imports >2633								
Cointegration	-0.645	27.94	392.49	-0.306	1.90	54	0.3541	0.0253
	(0.65)	(17.79)	(1.16)*	0.08)*				
Partial Cointegration	0.485	-19.151	6.072	-0.265	2.25	54	0.3924	0.0175
	(0.35)	(29.83)	(3.04)	(0.08)*				
Stationary	0.485	-36.735	1.934	-0.265	2.25	54	0.3924	0.0175
	(0.35)	(19.01)	(1.15)	(0.08)*				

***=5% level of significance**

The F-statistics p-value indicates that all forms of the regression model have significant explanatory power. The first part of Table 3 shows results for the full sample model assuming no import threshold. The model is estimated under cointegration, stationarity, and partial cointegration assumptions. In all three cases there is no evidence of long-run price transmission between either RSA prices or tariff costs to domestic wholesale maize prices in Mozambique. However, there is some evidence of a statistically significant long-run relationship between transport costs and Mozambican prices under the cointegration assumption. The coefficient measure indicates that a permanent $1 increase in diesel prices within South Africa will result in a $68 per metric ton increase in the price of Wholesale maize grain prices in the Maputo market. The speed of adjustment parameter is statistically significant at conventional levels and displays an estimated half-life of 5.06 months. This suggests that any shock to Mozambican maize prices dies out relatively quickly (half of the adjustment takes place within five months), but that prices adjust back to their mean (under stationarity) or mean rate of change (under nonstationarity) rather than back to some long-run equilibrium relationship with RSA prices.

When we allow for an import threshold which causes a regime shift in the price transmission relationship there appears to be strong support for the existence of a threshold effect[11]; a standard *F*-test of no threshold against the alternative of an optimal threshold provided a *p*-value of 0.0006. Nevertheless, in a low-import regime, i.e. periods where import volumes were below 2633 metric tons of grain, there continues to be no strong evidence of long-run price transmission between RSA and Mozambican wholesale maize grain prices. However, in high import regime, under the cointegration assumption, there is evidence of a long-run price relationship between tariff costs and domestic maize grain prices while under the stationarity and partial cointegration assumptions the evidence does not support statistically significant long-run price transmission. The speed of adjustment parameter, under both regimes, does support a slightly faster adjustment to shocks in the high import regime (half-life of 2.25. months versus 5 months for the full sample) implying then, that shocks do dissipate more quickly in the high import regime.

Given these results, the implication is that there exists no long-run relationship between RSA and Mozambican wholesale maize grain prices within our sample period; i.e. the price differential between these two markets could become very large with no tendency for them to come back together, while controlling for transportation and tariff costs. Based on a simple arbitrage argument, this result is not what we would expect since within the sample period every month, except for 3 (9/96, 10/96 and 8/02), maize grain was exported into Mozambique from South Africa; indicating sufficient incentive within the market for traders to engage in trade and possibly move the price differential back to some equilibrium value.

[11] All model forms provide the same threshold and test statistic because the assumptions used in each are just identifying. However, this test needs to be interpreted with care because the grid search procedure used to identify the optimal threshold introduces a nuisance parameter into the test for no threshold effects (see references). A standard F-test using the optimal threshold does not take this into account but the bootstrap procedure developed by Hansen and Seo (2002) to solve this problem is not directly applicable here because it applies to linear models while the models in this paper are nonlinear. Nevertheless, the standard F-test for structural change using the optimal threshold does strongly support a threshold effect.

One possible explanation for the lack transmission between these two markets is that South African maize grain is thinly traded within the Maputo wholesale grain market and the majority of all imports destined for Maputo are intended for domestic maize grain processors, some of which source their entire input requirements solely from South Africa (Nel, 2009). For example; within the 2008/2009 marketing year the large-scale Mill, Companhia Industrial Da Matola (CIM) located in Maputo, sourced their entire grain requirements directly from AFGRI, a South African based Agri-business firm involved in, among other things, grain trading. In fact, all grain silos located within towns that border Mozambique (i.e., Wonderfontein, Middleburg, Bethal, Ermelo, and Carolina) are owned by AFGRI; and according to AFGRI, one of their largest customers within Mozambique is CIM (van Rensburg, 2009).

6.2 Retail Maize Grain Prices

Table 4. Price Transmission Result for Retail Maize Grain Prices: 01/1995 to 09/2009

Model	β_1	β_2	β_3	λ	Half-life	No. Obs.	R^2	F-statistic
Full Sample								
Cointegration	0.353	59.244	0.052	-0.165	3.84	175	0.2074	0.0001
	(0.40)	(19.71)*	(0.27)	(0.04)*				
Partial Cointegration	0.285	47.763	-0.038	-0.165	3.84	175	0.2074	0.0001
	(0.19)	(68.23)	(0.19)	(0.04)*				
Stationary	0.285	-13.037	-0.003	-0.165	3.84	175	0.2074	0.0001
	(0.19)	(9.46)	(0.02)	(0.04)*				

***=5% level of significance**

The results for the full sample model assuming no import threshold indicates that there is no evidence of long-run price transmission between either RSA prices or tariff costs to retail maize prices in Mozambique. However, there is some evidence of a statistically significant long-run relationship between transport costs and Mozambican prices under the cointegration assumption. The coefficient measure indicates that a permanent $1 increase in diesel prices within South Africa will result in a $59 per metric ton increase in the price of Wholesale maize grain prices in the Maputo market.

The speed of adjustment parameter is statistically significant at conventional levels and displays an estimated half-life of 3.84 months. This suggests that any shock to Mozambican maize prices dies out relatively quickly (half of the adjustment takes place in less than four months), but that prices adjust back to their mean (under stationarity) or mean rate of change (under nonstationarity) rather than back to some long-run equilibrium relationship with RSA prices.

6.3 Retail Maize Meal Prices

Table 5. Price Transmission Result for Retail Maize Meal Prices: 03/1997 to 09/2009

Model	β_1	β_2	β_3	λ	Half-life	No. Obs.	R^2	F-statistic
Full Sample								
Cointegration	3.669	70.997	0.158	-0.066	10.15	149	0.2262	0.0002
	(1.78)*	(68.19)	(0.98)	(0.03)*				
Partial Cointegration	0.259	5.01	0.230	-0.066	10.15	149	0.2262	0.0002
	(0.22)	(7.60)	(1.40)	(0.03)*				
Stationary	0.259	36.52	-0.047	-0.066	10.15	149	0.2262	0.0002
	(0.22)	(10.49)*	(0.05)	(0.03)*				
Imports <= 82								

Cointegration	4.267	-71.623	0.354	-0.066	4.09	21	0.7658	0.0752
	(4.20)	(262.19)	(0.19)	(0.03)*				
Partial Cointegration	0.888	-14.899	3.304	-0.156	4.09	21	0.7658	0.0752
	(0.79)	(34.48)	(4.50)	(0.19)				
Stationary	0.888	-120.67	0.016	-0.156	4.09	21	0.7658	0.0752
	(0.79)	(43.51)*	(0.06)	(0.19)				
Imports > 82								
Cointegration	3.192	81.70	-0.946	-0.068	9.84	128	0.2838	0.0000
	(1.63)*	(64.30)	(2.29)	(0.03)*				
Partial Cointegration	0.202	5.182	-0.860	-0.068	9.84	128	0.2838	0.0000
	(0.23)	(7.98)	(1.76)	(0.03)*				
Stationary	0.202	48.89	-0.061	-0.068	9.84	128	0.2838	0.0000
	(0.23)	(10.80)*	(0.08)	(0.03)*				

***=5% level of significance**

The first part of Table 5 shows results for the full sample model assuming no import threshold. The model is estimated under cointegration, stationarity, and partial cointegration assumptions. In all three cases there is no evidence of long-run price transmission between Tariffs and domestic retail maize meal prices, while there is some evidence of a long-run relationship between transportation costs and domestic prices under the stationarity assumptions. Under cointegration there is evidence of a statistically significant long-run relationship between RSA and Mozambican prices.[12] The coefficient estimate indicates that a permanent $1 increase in SAFEX maize grain prices will result in a corresponding $3.70 increase in retail maize meal price within the Maputo domestic market. This finding is not surprising and is what we would expect since two of the largest milling companies, located in Maputo are responsible for the majority of the volume of maize grain imported into the country from South Africa (Nel, 2009).

The speed of adjustment parameter within the full sample model is statistically significant at conventional levels and displays an estimated half-life of 10.15 months. This suggests that any shock to Mozambican maize meal prices dies within 10 months and prices adjust back to their long-run equilibrium relationship with RSA prices under cointegration assumptions, but back to their mean (under stationarity) or mean rate of change (under partial cointegration).

The second part of Table 5 shows results allowing for an import threshold which causes a regime shift in the price transmission relationship. The optimal estimated threshold was 82 metric tons of imports and a standard F-test of no threshold against the alternative of an optimal threshold provided a p-value of 0.002, suggesting strong support for the existence of a threshold effect. Nevertheless, in the low import regime there is no strong evidence of long-run price transmission between RSA prices, transport costs, and/or tariffs and Mozambican prices. However, in a high import regime, under cointegration, there is evidence of a long-run price relationship between SAFEX and domestic maize meal prices while under the stationarity assumption there is some evidence of a statistically significant long-run relationship between transport costs and Mozambican prices. In partial cointegration cases the evidence does not support statistically significant long-run price transmission. The speed of adjustment parameter, although not significant in most cases under the low trade regime, does support a slightly faster adjustment to shocks in the high import regime (half-life of 9.84. months versus 10.15 for the full sample). Hence, shocks do dissipate more quickly in the high import regime.

7. CONCLUSION AND SUGGESTIONS FOR FURTHER RESEARCH

[12] In the cointegration model the null hypotheses of $\beta_1 = 0$ and $\beta_2 = 0$ cannot be tested formally because this would violate the maintained assumption of cointegration. However, the standard errors for these parameters reported in Table 5 are asymptotically valid for testing other hypotheses about β_1 and β_2 (see Phillips and Loretan, 1991) so it is clear that there is some evidence of a long run connection between Mozambican and RSA prices.

Neo-classical economic price theory tells us that prices are flexible and are responsible for efficient allocation of productive resources within markets. According to LOP, in an efficient market there must be only one price. In a spatial context, allowing for positive transaction costs, this implies that the difference in price between two markets should be exactly equal to cost of moving the commodity from one market to another. In such a case, the markets are said to have arrived at a spatial equilibrium, which is *Pareto* efficient. In a market with arbitrage and trade, any violations of LOP will be transitory and forces within the market will act to restore the long-run equilibrium condition when it has been subject to a shock. How long violations can persist depends on the state of information technology, whether markets operate with inventories and market structures. For instance, in an commodity market, with free information, inventories and no barriers to entry or exit, marketing agents can be expected to tolerate only short and transitory violations of LOP.

A convenient econometric way of analyzing the nature of LOP as an equilibrium condition is an error-correction model. In such a model an equilibrium law of one price is estimated using an error correction term. Given the existence of a long-run or equilibrium price relationship between markets, any violation or shock will be corrected for so that the equilibrium price difference is restored. If markets are not well integrated one cannot establish or estimate the fundamental law of one price identity. In this case, there would be no long-run relationship between the various markets price series and any regression involving these series would be spurious and tell us nothing meaningful.

The findings of this study indicate that under certain trading regimes, there is no evidence of a long-run relationship between Mozambican and South African maize grain prices. This implies then that any large deviations, within these regimes, which exceed transaction costs, could continue to grow with no tendency towards equilibrium. This finding is unexpected given a simple arbitrage argument. The trade volume data indicates maize grain exports from South Africa into Mozambique in every month except for three within the sample set. Some possible reasons for this finding could include:

1. Market Power: The existence of market power could limit the extent of arbitrage and allow for price differentials to remain well above the pareto efficient level
2. Inventories: inventory management can be an important element of a firm's adjustment to exogenous shocks therefore inaccurate crop estimates may contribute to the inability of mark actors to appropriately respond economic pricing signals.
3. Policy Intervention: ad hoc policy interventions, such as export bans or government buying programs, both domestically and/or regionally can result in increased risk and uncertainty for grain traders.
4. Asymmetric Information: Distorted market information on prices and/or crop projections could result in market actors not engaging in profit-maximizing behavior

References

Abdulai, A. (2000) "Spatial Price Transmission and asymmetry in the Ghanaian Maize Market", Journal of Development Economics, Vol. 63, pp. 327-349.

Barrett, C.B. (2001). "Measuring Integration and Efficiency in International Agricultural Markets." Review of Agricultural Economics, 23(1):19-32.

Balke, N.S. and T.B. Fomby (1997). "Threshold cointegration". International Economic Review, Vol. 38:627-645.

Engle, R.F. and C.W.J. Granger (1987). "Cointegration and Error-Correction: Representation, Estimation and Testing." Econometrica, 55(2):251-276.

Fackler, P.L. and Goodwin, B.K. (2002) "Spatial Price Analysis", In Gardner, B.L. and Rausser, G.C., editors, Handbook of Agricultural Economics, pp 972-1024, Elsevier Science.

Gonzalo, J. and Pitarakis, J.Y.. (2006) "Threshold effects in cointegrating relationships". Oxford Bulletin of Economics and Statistics, 68(s1): 813-833.

Goodwin, B.K., T.J. Grennes, and M.K. Wohlgenant (1990). "A Revised Test of the Law of One Price using Rational Price Expectations", American Journal of Agricultural Economics, Vol. 72(8).

Goodwin, B.K. and D.C. Harper (2000). "Price Transmission, Threshold Behavior and Asymmetric Adjustment in the U.S. Pork Sector', Journal of Agricultural & Applied Economics, Vol. 32:543-553.

Goodwin, B.K. and M.T. Holt (1999). "Asymmetric Adjustment and Price Transmission in the U.S. Beef Sector", American Journal of Agricultural Economics, Vol. 81:630-637.

Goodwin, B.K. and N.E. Piggott (2001). "Spatial Market Integration in the Presence of Threshold Effects", American Journal of Agricultural Economics, Vol.8392):302-317.

Granger, C.W.J. and T.H. Lee (1989). "Investigation of Production, Sales and Inventory Relationships using Multicointegrations and non-symmetric Error Correction Models", Journal of Applied Econometrics, Vol. 4, pp. 135-159.

Hansen, B.E. and Seo, B. (2002) "Testing for Two-Regime Threshold Cointegration in Vector Error correction Models", Journal of Econometrics, Vol. 110, pp. 307-316.

Houck, J.P. (1977) "An Approach to specifying and estimating nonreversible Functions", American Journal of Agricultural Economics, Vol. 59, pp.570-572.

Jayne, T.S., Chapoto, A., Minde, I. and Donovan, C. (2008). The 2008/09 Food Price and Food Security Situation in Eastern and Southern Africa: Implications for Immediate and Long-run Responses. International Development Working Paper, Michigan State University.

Klovland, J.T. (2005) "Commodity market integration 1850-1913: Evidence from Britain and Germany", European Review of Economic History, Vol 9, pp 163-197.

Meyer, J. and Von Cramon-Taubadel, S. (2004) “Asymmetric Price Transmission: A Survey”, Journal of Agricultural Economics, Vol. 55 (3), pp581-611.

Nel, Tinus (2009). Managing Director, Companhia Industrial Da Matola (CIM). Personal Interview. Maputo, Mozambique.

Reagan, P.B. and Weitzman, M.L. (1982) “Asymmetries in Price and Quantity Adjustments by the competitive Firm”, Journal of Economic Theory, Vol. 27, pp. 410-420.

Sexton, R., C. Kling, and H. Carmen (1991). “Market Integration, Efficiency of Arbitrage, and Imperfect Competition: Methodology and Application to U.S. Celery Markets”, American Journal of Agricultural Economics, Vol. 73(8).

Stephans, E., Mabaya, E., Von Cramon-Taubadel, S., and Barrett, S. (2008) “Spatial Price Adjustment with and without Trade”, paper presented at the AAEA Annual Meeting, Orlando, Florida.

Takayama, T. and G.G. Judge (1971). “Spatial and Temporal Price and Allocation Models. Amsterdam: North Holland.

Tong, H. (1983). “Threshold Models in Non-Linear Time Series Analysis”, New York, Springer Verlag.

Tsay, R.S. (1989). “Testing and Modeling Trheshold Autoregressive Processes”, Journal of the American Statistical Assoiciation, Vol. 84: 231-240.

Van Rensburg, Carla (2009). Junior Trader, AFGRI. Personal Interview. Centurion, South Africa.

Von Cramon-Taubadel, S. (1998) “Estimating asymmetric Price Transmission with the Error Correction Representations: An Application to the German Pork Market”, European Review of Agricultural Economics, Vol. 25, pp. 1-18.

Von Cramon-Taubadel, S. and Fahlbusch, S. (1994) “Identifying asymmetric price transmission with error correction models”, Poster Session EAAE European Seminar in Reading.

Von Cramon-Taubadel, S. and Loy, J.P. (1996) “Price Asymmetry in the international Wheat Market: Comment”, Canadian Journal of Agricultural Economics, Vol. 44, pp. 311-317.

Ward, R.W. (1982) “Asymmetry in Retail, wholesale and shipping Pint Pricing for fresh Vegetables”, American Journal of Agricultural Economics, Vol. 62, pp. 205-212.

Wolffram, R. (1971) “Positivistic Measures of aggregate Supply Elasticities: Some new Approaches – some critical Notes”, American Journal of Agricultural Economics, Vol. 53, pp. 356-359.