CAUSALITY AND ASYMMETRY IN PRICE TRANSMISSION MECHANISM: THE CASE OF NAMIBIA MUTTON MARKET

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Abstract

This study examines the price transmission mechanisms in the Namibian mutton market. It further estimates the causality links between the consumer and producer prices proxied with their indices. The traditional (Engle-Granger) and standardized (Enders & Siklos) Augmented Dickey-Fuller procedures was used to test for cointegration and asymmetry in price transmission. The following results were found: asymmetric price transmission between producer and consumer prices, direction of causation runs from producer to consumer price. In general, results suggested that actors in the market respond more quickly to shocks that stretch their market margin than those that squeeze them.

Key words: price transmission, asymmetry, Namibia.

1. Introduction

Agriculture in Namibia contributes around 5.5% of the national Gross Domestic Product (GDP) though 25% to 40% of Namibians depend on subsistence agriculture and herding. Approximately 70% of the country’s population is directly or indirectly dependent on agriculture to stand a living. The harsh climate and landscape make farmers in Namibia to practice free-ranging livestock production and it produces high quality meat for both national and international markets. Livestock farming consists of cattle, sheep, goats, and pigs, and contributes 2.7% to the national GDP (van Wyk, 2011). Namibia is a surplus producer of mutton and lamb, and has been exporting live sheep and mutton long before the country gained independence from South Africa (Sarker and Oyewumi, 2015). The overall state of affairs in agriculture in Namibia shows that farmers are facing an uphill struggle against not only climatic conditions which are far from optimal, but also an economic environment that has become intensely globalized and competitive.

A concern that is frequently raised by farmers is the relationship between the producer- and retail prices, or the variation in the price margin between the carcass price that the farmer or feedlot receive and the retail price that the consumer pays (Mare and Ogundeji, 2013). This calls for an understanding of the price transmission mechanism in the food value chain. Price transmission from producer to consumer prices is a key characteristic of the supply chain and has received a lot of attention in the literature. In agricultural markets we often observe that an increase of producer prices is transmitted more fully and faster to consumer prices while producer price decrease is passed through the supply chain to consumer prices incompletely and at a lower speed (Vavra and Goodwin, 2005, Rajcaniova and Pokrivcak, 2013). Despite its relative importance for policy making, a review of literature shows that price transmission study has not been given much attention in Namibia. Therefore the objective of this study is to evaluate the price transmission mechanism (symmetric or asymmetric) in the Namibian mutton market. In addition to this the direction of causality between the producer and consumer prices will also be established. To the best of the author’s knowledge, this paper is the first to evaluate price transmission in the mutton value chain of Namibia, thereby contributing to bridging the gap in literature and stimulate debate in this regard.
The Namibian mutton production value chain is presented in Figure 1. The Figure shows that are various markets and linkages between the various role players in the mutton industry.

![Figure 1: Mutton value chain map in Namibia](image)

Source: Adapted from van Wyk (2011).

The Namibian sheep value chain is structured around four main components, namely on-farm production, primary processing (slaughtering), secondary processing (de-boning) and marketing (wholesale and retail). Feedlots as in South Africa and elsewhere are not viable or competitive within Namibia, this was attributed to the scarcity and unreliability of feed (grain) production and the high transport cost. Sheep are transported over long distances to meat processing plants and other markets. The extended transport distances have a negative influence on quality because of bruising and stress (van Wyk, 2011).

2. Materials and methods

2.1. Data Description

Monthly producer prices of mutton for Namibia (from January 2002 to December 2013) were sourced from the Meat Board of Namibia. The consumer prices were not available at the time this analysis was done but the Consumer Price Index (CPI) for mutton were sourced from the Namibian Statistic Agency. The Producer Price Index (PPI) was calculated from the producer prices using December 2012 as the base. These were done to make both prices indexes. Therefore, the CPI was used as a proxy for the
consumer prices (retail price) and PPI as proxy for the producer prices (farm price). The summary statistics of the data is presented in Table 1.

Table 1: Descriptive statistics of mutton price series

<table>
<thead>
<tr>
<th>Variable</th>
<th>Obs.</th>
<th>Mean</th>
<th>Std. Dev.</th>
<th>Min</th>
<th>Max</th>
</tr>
</thead>
<tbody>
<tr>
<td>PPI (Producer Price)</td>
<td>144</td>
<td>71.50</td>
<td>20.85</td>
<td>40.73</td>
<td>121.36</td>
</tr>
<tr>
<td>CPI (Consumer Price)</td>
<td>144</td>
<td>63.67</td>
<td>21.54</td>
<td>34.32</td>
<td>103.27</td>
</tr>
</tbody>
</table>

A logarithm transformation of the variables is applied, such that results may be interpreted in percentage change terms.

2.2. Estimation approach

In this study, three (3) competing models, namely, Engle-Granger, Threshold Autoregressive (TAR), Momentum Threshold Autoregressive (M-TAR), and Momentum Consistent TAR models. The Engle-Granger methodology is based on the traditional Dickey-Fuller test. The remaining three models on the contrary apply standardized Dickey-Fuller technique to allow asymmetric adjustments from disequilibrium positions. Critical values developed by Enders & Siklos (2001) for such models.

The approach used by Alemu and Ogundeji (2010) was followed in this study. It starts by testing the statistical properties of variables. Once statistical properties of variables are confirmed, cointegration tests are carried out. Two approaches are commonly used to test long-run relationships among variables - the Johansen (1996) and the Engle-Granger (1987). If conducted using the Johansen methodology, the test will be about establishing the rank of \( \pi \) in [1].

\[
\Delta x_t = \pi x_{t-1} + \sum_{i=1}^{k} \Delta x_{t-i} + \psi_t
\]

Where \( x_t = [x_{1t}, x_{2t}] \)' is a matrix consisting of the two price variables which are I(1), \( \pi \) is a matrix of coefficients, \( \psi_t = [\psi_{1t}, \psi_{2t}] \) is a matrix of disturbance terms, and \( k \) is lag length to be determined based on the Akaike information criterion (AIC). Cointegration (long-run relationships) between variables is confirmed when the rank of \( \pi \) is different from zero i.e. \( \text{rank}(\pi) \neq 0 \).

The Engle-Granger method on the other hand is conducted in two steps. In step one, an equation given by a long-run equation given by [2] is fitted

\[
x_{it} = \phi_0 + \sum_{i=1}^{n} \phi_i x_{it} + \varepsilon_t
\]

Where \( x_{it} \)'s are individual variables, the \( \phi_i \)'s are parameter estimates, and the \( \varepsilon_t \)'s are the residuals.

In step two, [3] is fitted to check for the stationarity of the variable \( \varepsilon_t \) in [2] by testing the null hypothesis of no cointegration i.e. \( \rho = 0 \) against its alternative hypothesis of cointegration i.e. \( -2 < \rho < 0 \).

\[
\Delta \varepsilon_t = \rho \varepsilon_{t-1} + \sum_{i=1}^{k} \delta_i \Delta \varepsilon_{t-i} + \nu_t.
\]
Once cointegration is established, error-correction representation of the form given by [4] for Johansen

\[ \Delta x_t = \pi \Delta x_{t-1} + \sum_{i=1}^{k} \Delta x_{t-i} + \psi, \]

and by [5] for Engle-Granger are estimated

\[ \Delta x_{tr} = \sum_{i=1}^{n} \phi_i \Delta x_{tr} + \phi_{tr+1} \varepsilon_{tr-1} + v_t. \]

One major weakness of the above methodologies is their implicit assumption that the system makes symmetric adjustment toward equilibrium. If adjustment is asymmetric, the estimated equations could be misspecified.

Enders & Siklos (2001) extended the Dickey-Fuller equation to test for the cointegration, with asymmetric adjustment toward equilibrium being made part of the alternative hypothesis. They suggested estimating [6]

\[ \Delta \varepsilon_t = \rho_1 I_t \varepsilon_{t-1} + \rho_2 (1 - I_t) \varepsilon_{t-1} + \sum_{i=1}^{k} \Delta \varepsilon_{t-i} + \nu_t. \]

Where \( \varepsilon_t \) is as defined before, \( \rho_1 \) & \( \rho_2 \) are adjustment coefficients, and \( I_t \) is Heaviside indicator given by [7] below. The necessary and sufficient conditions for the stationarity of \( \varepsilon_t \) requires that \( \rho_1 \) & \( \rho_2 \) be less than zero and \( (1 + \rho_1) (1 + \rho_2) < 1 \) for any value of \( \tau \) (Petrucelli and Woolford, 1984).

\[ I_t = \begin{cases} 1 & \text{if } \varepsilon_{t-1} \geq \tau \\ 0 & \text{if } \varepsilon_{t-1} < \tau \end{cases} \quad \text{Where } \tau \text{ is the threshold.} \]

The threshold in [7] could take a value of 0 - the case of TAR and M-TAR models. Where \( \tau \) is assumed unknown (the case of momentum consistent TAR), a grid search is conducted to search for its value over the potential threshold variable. The consistent estimate of \( \tau \) is given by that which yields the lowest residual sum of squares in [1] (Chan, 1993). The TAR model differs from the M-TAR model on how the Heaviside indicator is created. In TAR models, it is created using \( \varepsilon_t \) in [5]. But in M-TAR models, first difference of \( \varepsilon_{t-1} \) i.e. \( \Delta \varepsilon_{t-1} \) is used instead. M-TAR models are preferred to TAR models when deviations from equilibrium are believed to exhibit more momentum in one direction than the other (Enders and Siklos, 2001).

Two types of tests are performed on estimates from [6]. Firstly, to check for cointegration by jointly testing for the null hypothesis that \( \rho_1 = \rho_2 = 0 \). Secondly, to test for symmetry in adjustment by checking whether \( \rho_1 = \rho_2 \). A non-standard testing procedure is recommended to conduct the first test as parameters are only identified in the alternative null hypothesis. Enders & Siklos (2001) developed critical values by running a Monte Carlo experiment. The critical values are compared against a value from F-distribution to conduct the test for cointegration. They labeled the F-statistic \( \Phi \) for TAR and M-TAR models and its analog \( \Phi^* \) for the Momentum consistent TAR model. According to Enders & Siklos (2001), the \( \Phi \) and \( \Phi^* \) must be used only when \( \rho_1 \) and \( \rho_2 \) satisfy convergence conditions (the necessary and sufficient conditions discussed above).

After cointegration is confirmed and the true values of \( \tau \), \( \rho_1 \) & \( \rho_2 \) are known, a standard F-test could be invoked to test for the symmetric null hypothesis \( (\rho_1 = \rho_2) \) against its alternative of asymmetric adjustment \( (\rho_1 \neq \rho_2) \) because OLS estimates for \( \rho_1 \) & \( \rho_2 \) have an asymptotic multivariate normal distribution (Tong, 1990).
Equation [4] can further be expanded to analyze the error correction model as follows:

\[ \Delta x_{1t} = \alpha_{11} + \sum_{i=1}^{k} \varphi_{1i} \Delta x_{1t-i} + \sum_{i=0}^{k} \beta_{1i} \Delta x_{2t-i} + \mu_1 \rho_1 \varepsilon_{r-1} + (1-I_t) \rho_{12} \varepsilon_{r-1} + v_{1t} \]

\[ \Delta x_{2t} = \alpha_{21} + \sum_{i=1}^{k} \varphi_{2i} \Delta x_{1t-i} + \sum_{i=0}^{k} \beta_{2i} \Delta x_{2t-i} + \mu_2 \rho_1 \varepsilon_{r-1} + (1-I_t) \rho_{22} \varepsilon_{r-1} + v_{2t} \]

Where the \( \alpha \)'s, \( \varphi \)'s, \( \beta \)'s and \( \rho \)'s are parameter estimates and the \( \varepsilon \)'s are independently and normally distributed residuals. Parameter estimates such as \( \varphi_{21} \) and \( \beta_{12} \) give information about the direction of causality but instead results from a pair wise causality test is presented and analyzed.

3. Results and discussion

The usage of non-stationary time series in regression analysis can lead to spurious regression. Therefore, augmented Dickey Fuller (ADF) test was used to ascertain the stationarity of the price series. Test results indicated that both variables were stationary at first difference, that is they are integrated of order one I(1). Therefore differencing them only once was sufficient to make them stationary. This is followed by estimation of equation [2] to determine long-run relationships between producer and consumer prices which gave the following results:

\[ x_{1t} = -0.2482 + 1.1043x_{2t} + \varepsilon_t \]

Where \( x_{1t} \) is the logarithm of the CPI and \( x_{2t} \) the logarithm of the PPI. The two estimated parameters are significant at 1% level of significance.

The residuals from the long run equation were used to test for cointegration (long run relationship) between \( \bar{X}_1 \) & \( \bar{X}_2 \) using [6] and [7] for four competing models - Engle-Granger, TAR, M-TAR and Momentum Consistent TAR models. In all the four cases, the value of \( k \) (lag length) in [6] was set at 2 using the AIC criteria. Results are given in Table 2.

According to Table 2, the no-cointegration null hypothesis could not be rejected for the Engle-Granger equation at conventional significance levels (column 2). The value of \( \rho_1 \) is -0.1817 and the t-statistics for the null hypothesis \( \rho_1 =0 \) is -4.5138. The critical values for EG test are -4.07, -3.37 and -3.03 at 1%, 5% and 10% levels of significance, respectively. Therefore, according to EG cointegration test, the variables are cointegrated at 1%, 5% and 10% levels of significance. Note that this is the approach which makes an implicit assumption that the process exhibits symmetric adjustment toward equilibrium.

<table>
<thead>
<tr>
<th>Model</th>
<th>Engle-Granger</th>
<th>Threshold Autoregressive (TAR)</th>
<th>Momentum-Threshold Autoregressive (M-TAR)</th>
<th>Momentum-consistent Threshold Autoregressive</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \rho_1 )</td>
<td>-0.1817 (-4.5134)</td>
<td>-0.2886 (-3.5033)</td>
<td>-0.4807 (-7.1187)</td>
<td>-0.7601 (-6.7320)</td>
</tr>
<tr>
<td>( \rho_2 )</td>
<td>NA</td>
<td>-0.1561 (-3.5796)</td>
<td>-0.0731 (-1.7356)</td>
<td>-0.0917 (-2.5147)</td>
</tr>
<tr>
<td>AIC</td>
<td>-4.9385</td>
<td>-4.9403</td>
<td>-5.1097</td>
<td>-5.1187</td>
</tr>
<tr>
<td>( \phi )</td>
<td>NA</td>
<td>11.3776</td>
<td>26.1258</td>
<td>25.7084</td>
</tr>
</tbody>
</table>
Following the work by Enders and Siklos (2001), Alemu and Ogundeji (2010), and Alemu 2011. Three other competing models (column 2 through 5) were analyzed. These models are different in terms of the type of threshold variable they use. In TAR models, the potential threshold variable is given by the residual from the long-run equation. In M-TAR and Momentum consistent TAR models, the threshold variable is given by the first difference of the residuals from the long-run equation. Their similarity lies on the assumption they make that there could be asymmetric adjustment toward equilibrium (Alemu and Ogundeji, 2010). Results from these models are presented in Table 2. Rows 2 and 3 give the ρ1 and ρ2 values for the TAR (-0.2886, -0.1561), M-TAR (-0.4807, -0.0731) and MC-TAR (-0.7601, -0.0917). Row 5 gives the F-statistics ϕ and ϕ* useful to test the null hypothesis of ρ1 = ρ2 = 0, which are 11.38, 26.13 and 25.71. These values were compared against the critical values (bottom of Table 2). Cointegration was confirmed at 1%, 5% and 10% for TAR, M-TAR and MC-TAR models. Diagnostic tests on residual were conducted for autocorrelation and heteroskedasticity were performed. The null hypothesis of no autocorrelation and no heteroskedasticity are accepted at conventional level of significance.

Results in Table 2 can further be analyzed to choose the model that best fits the data. Following Enders & Siklos (2001), who used the AIC, the Momentum consistent TAR model is selected because it has the lowest AIC value. Therefore, further analysis of the results will be on the adjustment parameters from the best model, that is, the MC-TAR model. Petruccelli and Woolford (1984) defined the necessary and sufficient condition that must be met for long run relationship to be confirmed as ρ1 < 0, ρ2 < 0 and ρ1 ρ2 < 1 for any value of the threshold. Enders and Siklos (2001) rewrote the conditions as ρ1 < 0, ρ2 < 0 and (1+ρ1)(1+ρ2) < 1. The other property of the test is that convergence to equilibrium is faster when ρ1 and ρ2 are both negative. Table 2 (column 5, rows 2 and 3) provides the values of the adjustment parameters ρ1 and ρ2 for the MC-TAR model. The two parameters are negative (-0.7601, -0.0917) and the operation (1+ρ1)(1+ρ2) gives a value of 0.02, which is smaller than 1. With these condition satisfied, Enders and Siklos (2001) critical values was used to test the no cointegration null hypothesis of ρ1 = ρ2 = 0. As recorded in Table 2, the computed F-statistic of 25.71 is higher than the corresponding critical value at 1% level of significance. Therefore, results are again in favour of the presence of long run relationship among the price series.

With the confirmation of cointegration, the adjustment parameter can be tested for symmetric null hypothesis that ρ1 = ρ2. This test is to confirm whether actors respond to negative and positive shocks differently. We found that the point estimates for ρ1 and ρ2 in [6] are different from each other (Table

<table>
<thead>
<tr>
<th>ρ1, ρ2</th>
<th>NA</th>
<th>2.2063 (0.1397)</th>
<th>27.9087 (0.0000)</th>
<th>31.8462 (0.0000)</th>
</tr>
</thead>
<tbody>
<tr>
<td>γ</td>
<td>NA</td>
<td>0</td>
<td>0</td>
<td>0.0396</td>
</tr>
<tr>
<td>Lag length (k)</td>
<td>2</td>
<td>2</td>
<td>2</td>
<td>2</td>
</tr>
<tr>
<td>err +γ</td>
<td>-</td>
<td>-</td>
<td>-</td>
<td>-0.2834 (-7.1614)</td>
</tr>
<tr>
<td>err -γ</td>
<td>-</td>
<td>-</td>
<td>-</td>
<td>-0.0585 (-4.3431)</td>
</tr>
</tbody>
</table>

* entries are estimated value of ρ1, ρ2 with t-statistic in parentheses
* entries in this row are the sample values of ϕ and ϕ*. Critical values for two variables case and one-lagged for ϕ are 4.99, 6.01 and 8.30 at 10%, 5%, and 1%, respectively. The corresponding values for ϕ* are 6.02, 7.08, and 9.51, see Enders and Siklos (2001).

* entries in this row are the sample F-statistic for the null hypothesis that the adjustment coefficients are equal. Significance levels are in parenthesis.

Q(p) is the p-value for the autocorrelation test of the first p residuals. It is based on Ljung-Box statistic

Test for first order ARCH residuals. The numbers in parenthesis represents p-values.
2, column 5, Row 6), suggesting that adjustments toward equilibrium are asymmetric. They give information about the degree of persistence of discrepancies from equilibrium. For example, the values and $\rho_1 = -0.7601$ and $\rho_2 = -0.0917$ imply that on average 23% of the positive and about 91% of the negative shocks are rolled over to the next month. This implies that retailer and producer prices are characterized by negative rather than positive asymmetric price transmission. This suggests that deviations from equilibrium persist when they are caused by negative shocks (an increase in producer prices) than in the case of positive shocks (a decrease in producer prices). This means that in the case of mutton market actors in the market react faster to shocks that stretch their margin than shocks that squeeze the margin. In other words, consumer prices of mutton react more fully or rapidly to a decrease in producer prices than to an increase in producer prices.

The point estimate of the coefficient of the error correction terms are -0.2834 for the positive error correction term and -0.0585 for the negative one. Both estimates have correct sign and are significant at 1% significance level. This suggests that positive and negative disequilibria are corrected. However, the response is the strongest for positive shocks. According to the result found, every month, retailers correct about 28% of the positive and 5% of the negative disequilibria.

Next, the direction of causality between the CPI for mutton and PPI for mutton is confirmed using the Granger causality tests. The null hypothesis and the test results is presented in Table 3.

Table 3. Test results on causality

<table>
<thead>
<tr>
<th>Granger Causality Tests</th>
<th>F-Statistic</th>
<th>Probabilities</th>
</tr>
</thead>
<tbody>
<tr>
<td>CPI does not Granger Cause PPI</td>
<td>0.60366</td>
<td>0.4385</td>
</tr>
<tr>
<td>PPI does not Granger Cause CPI</td>
<td>50.0225</td>
<td>0.0000</td>
</tr>
</tbody>
</table>

According to the results found, producer inflation does granger cause consumer price inflation. The null hypothesis that producer inflation does not cause consumer inflation is rejected at 1% level of significance. However, the reverse is not true. The test cannot reject the null hypothesis that a consumer price does not Granger cause producer prices at conventional level of significance. This means that the direction of causality is unidirectional and runs from Producer price to the consumer price.

4. Conclusion

This study analyzed price transmission in the Namibian mutton market. The following results were found. Cointegration test revealed that the price series are cointegrated. That is, there exist long run relationships among the price series. Secondly, we found unidirectional causation running from producer to consumer prices but not vice versa. Thirdly, the procedures developed by Enders & Silkos (2001) were used to test the null hypothesis of symmetric adjustment against its alternative hypothesis of asymmetric adjustment. A strong evidence of asymmetry was confirmed and the symmetric null hypothesis was rejected. It can be concluded that in the case of mutton market in Namibia, actors in the market react faster to shocks that stretch their margin than shocks that squeeze the margin. In other words, consumer prices of mutton react more fully or rapidly to a decrease in producer prices than to an increase in producer prices.
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