Nutrition and economic growth in South Africa: A momentum threshold autoregressive (MTAR) approach

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NUTRITION AND ECONOMIC GROWTH IN SOUTH AFRICA: A MOMENTUM THRESHOLD AUTORERGESSIVE (M-TAR) APPROACH

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ABSTRACT:

Purpose: This purpose of our paper is to examine asymmetric co-integration effects between nutrition and economic growth for annual South African data from the period 1961-2013.

Design/methodology/approach: We deviate from the conventional assumption of linear co-integration and pragmatically incorporate asymmetric effects in the framework through a fusion of the momentum threshold autoregressive and threshold error correction (MTAR-TEC) model approaches, which essentially combines the adjustment asymmetry model of Enders and Silkos (2001); with causality analysis as introduced by Granger (1969); all encompassed by/within the threshold autoregressive (TAR) framework, a la Hansen (2000).

Findings: The findings obtained from our study uncover a number of interesting phenomena for the South Africa economy. Firstly, in coherence with previous studies conducted for developing economies, we establish a positive relationship between nutrition and economic growth with an estimated income elasticity of nutritional intake of 0.15. Secondly, we find bi-direction causality between nutrition and economic growth with a stronger causal effect running from nutrition to economic growth. Lastly, we find that in the face of equilibrium

shocks to the variables, policymakers are slow to responding to deviations of the variables from their co-integrated long run steady state equilibrium.

Originality/value: In our study, we make a novel contribution to the literature by exploring asymmetric modelling in the correlation between nutrition intake and economic growth for the exclusive case of South Africa.

1. INTRODUCTION

Economic growth can essentially be described as the increase in the quality and quantity of goods and services, resulting from multitudes of entrepreneurs hiring more workers, introducing technological innovations, and improving worker productivity (Khamfula, 2005). In practice, the concept of economic growth is used to describe the process of allocating production factors to productive use and the allocation of such resources is subject to the constraints of an economy's infrastructure. Of recent, it has become increasingly recognized that in order to improve economic performance, qualitative aspects of economic growth are probably more important than growth itself (Arora and Vamvakidis, 2005). In this context, the effects of nutrition as well as the channels through which it affects or is affected by economic growth has recently received considerable attention in the academic realm. Even though the current empirical findings on the relationship between nutrition and economic growth have failed to produce anything but a weak consensus concerning their co-integration properties, the general outlook, nevertheless, remains optimistic of a 'technically-determined' correlation existing between the two variables.

Given such a universal and unchallenged belief of a correlation between nutrition and economic growth, one would imagine that such a strongly held conviction would be readily demonstrable with reference to well-established empirical evidence. The unanimity of these views, however, seems to be embodied in re-iterated political and editorial statements rather than in the academic literature. Examples of inadequacies existing in the empirical literature are not difficult to come across. Take for instance, Fan and Pandya-Lorch (2011) who argue upon how existing research fails to answer the question facing several developing economies on how to set priorities and sequence interventions in maximizing the benefits arising from the dynamic and nonlinear relationship between nutrition and economic growth. Furthermore, Thomas and Frankenberg (2002) have highlighted on how the current literature has failed to identify a deterministic causal relationship between the two variables thus warranting further research on the subject matter. Thereby motivated, to a large extent, by the empirical hiatus existing in the standard empirical literature; which ranges from failure to take into account for nonlinearities at a macroeconomic level (Vecchi and Coppola, 2003); to a failure to observe causal effects between the two variables (Neeliah and Shankar, 2008), our study seeks to develop a general econometric framework for asymmetric modelling that circumvents these issues in a coherent manner. We pragmatically address these issues through a fusion of the momentum threshold autoregressive and threshold error correction (MTAR-TEC) model approaches, which essentially combines the adjustment asymmetry model of Enders and Silkos (2001); with causality analysis as introduced by Granger (1969); all encompassed by/within the threshold autoregressive (TAR) framework, a la Hansen (2000).

The motivation behind the choice of our empirical approach can be rationalized as follows. Engle and Granger (1987) introduced the concept of establishing cointegration effects amongst time-series variables as a means of ensuring that the variables of interest follow a common long-run trend and the general estimation of the correlation between the variables will, thereafter, not yield spurious results. In the standard literature, it is a common and well-accepted practice for researchers to investigate the effects between nutrition and economic growth under the implicit assumption of linear cointegration and causality effects. However, as noted by Shimokawa (2010), the assumption of a symmetric adjustment mechanism may be too restrictive in accounting for the dynamic and nonlinear effects in the co-relationship between the two variables. Such a contention of an asymmetric relationship between the variables may be plausible, for a variety of reasons, of which on the forefront of these reasons, is that a number of empirical studies have found that economic growth, at least, evolves as a nonlinear process over the business cycle (see Beechey and Osterholm, 2008; and Shelly and Wallace, 2011 for examples). Thereby, in ignoring nonlinearities when investigating the macroeconomic effects of nutrition on economic growth, researchers are prone to ignoring underlying cointegration asymmetries in the microeconomic foundations of business cycle theory connecting the two variables (Vecchi and Coppola, 2003). This, in turn, can rise up a hypothetical contention of a possible nonlinear adjustment process of nutrition and output growth towards their long-run steady-state equilibrium.

The cohesiveness of the selected MTAR-TEC approach renders it a noble candidate for extending the current literature, particularly for an emerging economy like South Africa.

One of the principle advantages with this empirical approach is that, unlike other methods commonly employed in the literature, the MTAR-TEC model, on account of being derived from Hansen’s (2000) TAR framework, can accommodate for asymmetric unit root testing, asymmetric cointegration analysis as well as asymmetric causality analysis within a singular framework. By mapping our obtained results towards applicable policy implications, we are then able to extract or isolate a variation of applicable policy interventions dependent upon different states of the business cycle. In particular, we are able to evaluate as to whether nutrition and economic growth evolve as nonlinear processes over the business cycle, and if so, to what extent are they asymmetrically cointegrated in a general equilibrium sense. For instance, we are able to establish that positive and negative shocks to the time series produce different adjustment effects of the variables to long-run their steady-state equilibrium. These results are not only intriguing for researchers but also to offer a broader perspective for policymaking in its “*never-ending*” challenge to eradicate poverty through improved nutrition and economic growth.

The remainder of our manuscript, hereafter, is presented as follows. The following section provides a brief review of previous literature and summarizes a range of statistical techniques used to examine the correlation between nutrition on economic growth. Section three outlines the MTAR-TEC model framework whilst section 4 presents the main empirical results. We conclude our study in section 5 of our manuscript by deriving the policy-relevance of the study’s findings.

2. LITERATURE REVIEW

Even though it is generally accepted that economic growth is at least a necessary precondition for reducing levels of poverty, very little is known about the relationship between economic growth and nutrition, and, hence, very little can be deduced on how economic policies can be geared towards improving nutritional intake. Isolating the effects of improved nutrition on productivity and economic growth, at large, remains a novel field of study in the academic paradigm and quantifying the effects of nutrition on economic growth has, over time, proved to be quite a challenging task for research academics and practitioners, alike. The available empirical research up to date has focused on nutrition-productivity growth linkages and this field of research has been primarily carried out by nutritionists and medical doctors, although an increasing number of economists have taken a keen interest on the subject matter (Strauss, 2004). As conveniently noted by Salois et. al. (2010), two types of empirical approaches been adopted so far in the existing literature; namely, macroeconomic and microeconomic approaches. Microeconomic studies mainly make use of data from household surveys and focus on the impact of nutrition intake upon health of individuals, whereas the macro or aggregate alternative typically investigate gains/losses from nutrition/malnutrition in terms of growth in national income (Karlsen and Rikardson, 2007). Most macroeconomic studies estimate the panel or single country data between gross domestic product (GDP) (or some other closely-related measure of income) and nutritional intake and tend to reveal a significantly positive relationship between the two variables (see Cole, 1971; Bouis and Haddad, 1992; and Arcand, 2001 for illustrations). On the other hand, microeconomic studies treat nutrition as being a unique dimension of human capital and tend to find that the availability of household income does not necessarily make access to the food available (see Deolalikar, 1988; Vecchi and Coppola, 2004; and Karlsen and Rikardson, 2007).

From a chronological perspective, the empirical frameworks used in examining the relationship between nutrition and economic growth have undergone a variety of radical experimental phases in the literature. In earlier empirical studies, the investigation into the relationship between nutrition and economic growth was mainly conducted through the specification of linear growth models, which were typically estimated using the ordinary least squares (OLS) technique. The resulting empirical evidences presented in these earlier empirical studies were indicative of a positive correlation between the two variables, although some empirical works (e.g. Behrman and Deolalikar, 1987) provide little or no evidence to validate this notion. A popular citation among this earlier literature is a study conducted by Correa and Cummins (1970), who found that for Latin American economies, increased calorie intake accounts for approximately 5 percent of the GDP growth, or alternatively stated, has a 0.05 estimate of income elasticity for calorie demand, whereas for industrialized economies an increased calorie consumption seemed to show a negligible effect on economic growth. Similarly, Reutlinger and Selowsky (1976) come up with a positive correlation between nutrition and income growth solely for a pool of developing economies. However, differing from study of Correa and Cummins (1970), the authors estimate relatively higher income elasticities of calories which ranged from 0.15 to 0.30. Another empirical study worth taking note of is that presented by Strauss (1986), who reported significant impacts of increased calorie consumption on farm output and wages in

Sierra Leone. The author particularly finds that an increase of 1 percent in calorie intake increased productivity output by approximately 1.6 percent but this effect ceases to exist at sufficiently high levels calorie consumption. In a attempt to account for the varying elasticity estimates obtained in these previous studies, Behrman and Deolalikar (1989) conclude from their study that as food budgets increase from very low levels, there is a very pronounced increase in the for food variety more specifically for developing or emerging economies. An important implication for this finding is that since the elasticity of substitution is higher amoung poor households, any increase in food prices will cause the poor to curtail their food consumption more dramatically than the rich.

Unfortunately, the cumulative evidence of a positive correlation between nutritional status and output productivity found in the early literature has been prematurely misinterpreted as been indicative of causal effects existing among the variables. As a consequence, a number of spurious and misleading policy implications have being drawn from these empirical findings. However, through the consolidation of appropriate statistical tools into the empirical literature, it has been possible for more current research studies to formally probe into cointegration and causality effects between the time series variables. Recently, some scholars have considered the use of vector autoregression (VAR) models and various cointegration techniques, for a more widespread interpretation of the regression results, in the sense of adding another dimension of policy implications derived from the empirical analysis. This cluster of studies appears to be solely responsible for reviving the academic interest on the subject matter and, as a consequence, has resulted in an extension of the current knowledge of nutrition-economic growth relationship in a dynamic manner. Take for instance, Neeliah and Shankar (2008), who employ the Johansen's (1991) cointegration technique as well as Granger (1969) causality tests to nutritional intake and economic growth data for the Mauritian economy. The authors find that even though both time series variables are first difference stationary (i.e. a preliminary indication of cointegration amoung the variables), formal cointegration or causality tests performed on the data reveal that the variables are neither cointegrated nor are there any causality effects among them and, consequentially, any estimated regression between the two variables growth will be prove to be spurious. Contrary to these findings, Taniguchi and Wang (2003) consider running granger causality tests for Sub-Saharan, Latin American and Asian countries and report that causality runs in both directions, even though the impact of economic growth on nutritional intake is more significant than that of nutritional intake on economic growth. In an even more recent study, Halicioglu (2011) employs the ADRL cointegration technique to Turkish data and estimates an income elasticity of calorie intake of 0.22 whereby causality is established to run uni-directional from income growth to calorie intake. Ogundari (2011) extends upon the study of Halicioglu (2011) by utilizing a vector-autoregressive-error-correction-model (VAR-ECM) approach to Nigerian data and establishes cointegration between nutritional demand and GDP output growth with a 1 percent increase in GDP resulting in an increase in nutritional demand of between 0.059 and 0.073 percent. Furthermore, the short-run dynamics associated with the estimated error correction model (ECM) show that a shock to GDP will result in a speed of adjustment of nutrition to the long-run steady state of approximately 26 percent to 29 percent. Notably, the authors do not perform formal granger causality tests, but conclude that the estimated impulse response functions of the VAR-ECM system lend support of output growth leading to increases in nutritional demand.

As previously highlighted, the empirical results drawn from studies using cointegration techniques and causality analysis can be used to draw out several useful policy implications. For example, if causality is established to run from nutrition intake to economic growth, one can conclude that improvements in nutrition intake at the household level will result in an improvement in overall output productivity. Such a causal relationship is illustrated in a variety of microeconomic models depicting the dynamic relationship between the two variables (see Fogel (2004) and Meng et. al. (2004)) and discourages the strict pursuit of development strategies aimed at improving economic growth or national income since these policies are seen to be inefficient at alleviating hunger. On the other hand, if causality is found to run from economic growth to nutritional intake, then this indicative that policymakers should be more concerned with improving and distributing output productivity in a manner as to influence the nutrition intake of an economy's inhabitants. In this instance, the overriding issue of poverty alleviation is relative to the level of national income and the issue should be addressed and initiated at a macroeconomic policy level. However, as is clearly evident from our sample review of previous studies presented so far, even with the improved calibre of statistical methods applied in the literature, the empirical evidence still remains inconclusive of the extent of cointegration and causal effects between nutritional intake and productivity growth. This conclusion is further reiterated in a recent study of Ogundari and Abdulai (2013) who, by employing meta-regression analysis to a sample of 40 studies, find that even though the positive correlation commonly found between nutrition and income is significant, there, however, exists a publication bias of the obtained calorie-income elasticities. One highly justifiable reason for the aforementioned inconclusiveness, as point out by Shimokawa (2010), is that the majority of current empirical literature tends to ignore the possibilities of asymmetries existing in the relationship between these two variables. Surprisingly, a handful of studies have either supported the notion or have been equally indicative of existing asymmetrical effects between the time series variables and yet little empirical work has been formally conducted to verify this phenomenon. Despite the current quantity of publication on the subject matter being quite limited in volume, this new wave of empirical literature typically provides evidence of the income elasticities of nutrition varying across a range of different economic conditions (Mondal et. al., 2005), different time periods (Skoufias, 2009), different genders (Shimokowa, 2010) and different incomes (Salois et. al., 2010). Generally, these studies hypothesize on the estimated income-calorie relationship being described as a curve as opposed to straight line and conduct their estimations based on spline functions (Skoufias et. al., 2009), quantile regressions (Salois et. al., 2010) and non-parametric estimators (Shimokowa, 2010). Notwithstanding the efforts put by these authors into reaching a general consensus of asymmetric behaviour governing the whole correlation between nutrition and economic growth, this new wave of empirical literature has yet to explore the possibility of asymmetric effects from a cointegration and causality perspective.

3. THEORETICAL AND EMPIRICAL FRAMEWORK

In spite of the existing literature, the exploration of the co-movement between nutritional intake and economic growth in South Africa remains unknown and requires formal investigation. From a theoretical perspective, two strands of literature are commonly concerned with modelling the correlation between nutrition and income. The first strand of literature, makes use of the nutrition based efficiency wage model as introduced by Leibenstein (1957) and further developed by Mirrlees (1976) and Stiglitz (1976), and this theory depicts that higher wage rates allow workers to improve nutritional intake, which, as an important component of human capital, enhances productivity within an economy. This, in turn, enables individuals to enhance their accumulation of income or wealth. According to the nutrition-based efficiency wage model, productivity depends on nutrition and this relation can be depicted in the following function:

$$gdp = f(des) \tag{1}$$

Where gdp is representative of national income and des represents a measure of nutritional intake. The second strand of theoretical literature relies more on the Engel curve which, in its functional form, takes the demand for food calories to be dependent upon income. Initially, Engel curves where used to describe how household expenditure on a particular commodity varied with the household budget, of which later modifications to this initial specification, replaced household expenditure with demand for food calories. The functional form for the nutrition-based Engel curve can therefore take the following specification:

$$des = f(gdp) \tag{2}$$

For the simple fact that the actual causal relationship between economic growth and nutrition is unknown a prior for South Africa, we begin our empirical analysis by specifying two bivariate regression equations. In the first regression, we place GDP growth as the dependent variable (i.e. nutrition-based efficiency wage model) as in Correa and Cummins (1970) and Taniguchi and Wang (2003):

$$gdp = \psi_{10} + \psi_{11} des + \xi_{t1} \tag{3}$$

Whereas in the second regression, we follow Strauss and Thomas (1998), Meng et. al. (2004) and Skoufias et. al. (2009) by assuming that the annual percentage change in nutrition is dependent upon economic growth in the regression equation (i.e. nutrition-based Engel curve):

$$des = \psi_{20} + \psi_{21} gdp + \xi_{t2} \tag{4}$$

From the above long-run regressions gdp is output growth rate, des is the year-on-year percentage change in the dietary energy supplies and ψ_i are the associated regression coefficients. In introducing asymmetric adjustment between the observed time series variables, we follow Enders and Siklos (2001) and allow the residual deviations from the long-run equilibrium to behave as a TAR process. Formally, we model the residuals obtained from regressions (3) and (4) as follows:

$$\Delta\xi_{ti} = I_t\rho_1\xi_{t-1} + (1 - I_t)\rho_2\xi_{t-1} + \sum_{i=1}^{p}\beta_i\,\Delta\xi_{t-i} + \varepsilon_t \qquad (5)$$

From equation (3) asymmetric adjustment is implied by different values of ρ_1 and ρ_2. If ξ_{t-1} is found to be stationary, then the least squares (LS) estimates of ρ_1 and ρ_2 will have an asymptotic multivariate normal distribution for any given value of a consistently estimated threshold. Enders and Silkos (2001) demonstrate that a sufficient condition for stationary of ξ_{t-1} is that ρ_1, $\rho_2 < 0$ and $(1-\rho_1)(1-\rho_2) < 1$. Enders and Dibooglu (2001) suggest a more formal test of the null hypothesis of no cointegration (i.e. $\rho_1 = \rho_2 = 0$) against the alternative of cointegration (i.e. $\rho_1 \neq \rho_2 \neq 0$). If the null hypothesis of no cointegration is rejected, then we can proceed to test for the null hypothesis of symmetric adjustment (i.e. $\rho_1 = \rho_2$) against the alternative of asymmetric adjustment (i.e. $\rho_1 \neq \rho_2$). The co-integration tests are then evaluated using standard F-test statistics. Concerning the asymmetric modelling of the error terms of our growth regression (3) and (4), we opt to estimate each of our regression specifications using four types of asymmetric cointegration relations, namely; TAR with a zero threshold; consistent-TAR with a nonzero threshold; MTAR with a zero threshold; and consistent-MTAR with a nonzero threshold. For our TAR model with a zero threshold, we use the following indicator function:

$$I_t = \begin{cases} 1, if\,\xi_{t-1} \geq 0 \\ 0, if\,\xi_{t-1} < 0 \end{cases} \qquad (6)$$

And for our c-TAR model with a nonzero threshold, I_t, is set according to:

$$I_t = \begin{cases} 1, if\,\xi_{t-1} \geq \tau \\ 0, if\,\xi_{t-1} < \tau \end{cases} \qquad (7)$$

The threshold variable governing asymmetric behaviour is denoted by τ and the estimated threshold value is denoted as τ. Enders and Silkos (2001) suggest the use of a grid search procedure to derive a consistent estimate of the threshold. Our choice of nonzero threshold estimate follows the same procedure as that used for estimating the TAR models as described in Hansen (1999). The TAR cointegration models, as derived by combining equation (5) with equations (6) and (7) are designed to capture potential asymmetric deep movements in the residuals if, for example, positive deviations are more prolonged than negative deviations. Enders and Granger (1998) and Caner and Hansen (2001) suggest that by permitting the Heaviside indicator function, I_t, to rely on the first differences of the residuals, $\Delta\xi_{t-1}$, A MTAR version of the residual modelled in equation (5) can hence be developed. The implication of the MTAR model is that correction mechanism dynamic since by using $\Delta\xi_{t-1}$, it is possible to access if the momentum of the series is larger in a given direction relative to the direction in the alternative direction. Given such a scenario, the MTAR model can effectively capture large and smooth changes in a series. Unlike the TAR model which shows the "depth" of the swings in equilibrium relationship, the MTAR can capture spiky adjustments in the equilibrium relationship since it permits decay in the relationship to be captured by $\Delta\xi_{t-1}$ instead of ξ_{t-1}. TAR and MTAR models allow the residuals to exhibit different degrees of autoregressive decay depending on the behaviour of the lagged residual and its first difference respectively. In the MTAR model with a zero threshold, I_t, is set as:

$$I_t = \begin{cases} 1, if\ \Delta\xi_{t-1} \geq 0 \\ 0, if\ \Delta\xi_{t-1} < 0 \end{cases} \tag{8}$$

Whereas for the c-MTAR model with a nonzero threshold, I_t, is set as:

$$I_t = \begin{cases} 1, if\ \Delta\xi_{t-1} \geq \tau \\ 0, if\ \Delta\xi_{t-1} < \tau \end{cases} \tag{9}$$

According to the granger representation theorem, an error correction model can be estimated once a pair of time series variables is found to be cointegrated. When the presence of threshold cointegration is validated, the error correction model can be modified to take into account asymmetries as in Blake and Fombly (1997). The asymmetric error-correction model also can exist between a pair of time series variables of Δdes_t and Δgdp_t when they are formed in an asymmetric cointegration relationship. The TAR-VEC model for a zero threshold can be expressed as:

$$\begin{pmatrix}\Delta des_t \\ \Delta gdp_t\end{pmatrix} = c + \begin{cases} \lambda^{+}\xi_{t-1}^{+} + \sum_{i=1}^{p} \alpha_k^{des+}\,\Delta des_{t-k}^{+} + \sum_{i=1}^{p} \beta_k^{gdp+}\,\Delta gdp_{t-k}^{+}\,, if\ \xi_{t-1} < 0 \\ \lambda^{-}\xi_{t-1}^{-} + \sum_{i=1}^{p} \alpha_k^{des-}\,\Delta des_{t-k}^{-} + \sum_{i=1}^{p} \beta_k^{gdp-}\,\Delta gdp_{t-k}^{-}\,, if\ \xi_{t-1} \geq 0 \end{cases} \tag{10}$$

Whereas the c-TAR-TEC model with a nonzero threshold is given as:

$$\begin{pmatrix}\Delta des_t \\ \Delta gdp_t\end{pmatrix} = c + \begin{cases} \lambda^{+}\xi_{t-1}^{+} + \sum_{i=1}^{p} \alpha_k^{des+}\,\Delta des_{t-k}^{+} + \sum_{i=1}^{p} \beta_k^{gdp+}\,\Delta gdp_{t-k}^{+}\,, if\ \xi_{t-1} < \tau \\ \lambda^{-}\xi_{t-1}^{-} + \sum_{i=1}^{p} \alpha_k^{des-}\,\Delta des_{t-k}^{-} + \sum_{i=1}^{p} \beta_k^{gdp-}\,\Delta gdp_{t-k}^{-}\,, if\ \xi_{t-1} \geq \tau \end{cases} \tag{11}$$

The MTAR-TEC model with a zero threshold is specified as:

$$\begin{pmatrix}\Delta des_t \\ \Delta gdp_t\end{pmatrix} = c + \begin{cases} \lambda^{+}\xi_{t-1}^{+} + \sum_{i=1}^{p} \alpha_k^{des+}\,\Delta des_{t-k}^{+} + \sum_{i=1}^{p} \beta_k^{des+}\,\Delta gdp_{t-k}^{+}\,, if\ \Delta\xi_{t-1} < 0 \\ \lambda^{-}\xi_{t-1}^{-} + \sum_{i=1}^{p} \alpha_k^{des-}\,\Delta des_{t-k}^{-} + \sum_{i=1}^{p} \beta_k^{gdp-}\,\Delta gdp_{t-k}^{-}\,, if\ \Delta\xi_{t-1} \geq 0 \end{cases} \tag{12}$$

Whereas, the c-MTAR-TEC with a nonzero threshold is given by:

$$\begin{pmatrix}\Delta des_t \\ \Delta gdp_t\end{pmatrix} = c + \begin{cases} \lambda^{+}\xi_{t-1}^{+} + \sum_{i=1}^{p} \alpha_k^{des+}\,\Delta des_{t-k}^{+} + \sum_{i=1}^{p} \beta_k^{gdp+}\,\Delta gdp_{t-k}^{+}\,, if\ \Delta\xi_{t-1} < \tau \\ \lambda^{-}\xi_{t-1}^{-} + \sum_{i=1}^{p} \alpha_k^{des-}\,\Delta des_{t-k}^{-} + \sum_{i=1}^{p} \beta_k^{gdp-}\,\Delta gdp_{t-k}^{-}\,, if\ \Delta\xi_{t-1} \geq \tau \end{cases} \tag{13}$$

Through the above described systems of error correction models, the presence of asymmetries between the variables could initially be examined by examining the signs on the coefficients of the error correction terms. Furthermore, three types of joint hypotheses can be formed from the specified TEC models. Firstly, granger causality tests can be implemented by testing whether all Δdes_t and Δgdp_t are statistically different from zero based on a standard F-test and if the λ coefficients of the error correction are also significant. The null hypothesis that Δdes_t dose not lead to Δgdp_t can be denoted as: H_{03}: $\alpha_k = 0$, i=1, ..., k and the null hypothesis that Δgdp_t does not lead to Δdes_t is: H_{04}: $\beta_k = 0$, i=1, ..., k. The second type of hypothesis would be the cumulative symmetric effects which is relatively a long-run test for asymmetry. The final hypothesis tests whether it is possible to get back to equilibrium after a shock, and if it is the case, how long will it take. Since the causality tests are sensitive to the selection of the lag length, we determine the lag lengths using the AIC criterion.

4. EMPIRICAL ANALYSIS

4.1 UNIVARIATE TIME SERIES ANALYSIS

Considering the nature of our research, the data used in the empirical analysis consists of the annual percentage change in gross domestic product (GDP) which is gathered from the South African Reserve Bank (SARB) website whereas nutrition is measured by calorie intake expressed in calories/capita/day was collected from various food balance sheets (FAO, 2010). The empirical analysis uses annually adjusted data obtained for the periods extending from 1960 to 2009. Our choice of sample period and periodicity reflects the limitations in the availability of the time-series data on nutrition and economic growth for South Africa. We also take into consideration the discussions of Hodge (2009) concerning the volatility complexities associated with South African data and advocate on the use of filtering techniques in order to smooth the data In particular, we apply the Hodrick-Prescott (HP) filter to the time series variables prior to being incorporated into the econometric analysis.

Prior to testing for unit roots within the individual time series, we begin our analysis by estimating self-exciting threshold autoregressive (SETAR) processes for the time series variables as means of evaluating whether nonlinear behaviour exists amoung the observed variables. In particular, we apply Hansen's (2000) conditional least squares (CLS) technique which entails performing a grid-search over a predetermined range of threshold variable estimates belonging to a set $\Psi = [\underline{\gamma}, \overline{\gamma},]$ with the optimal estimates $\hat{\gamma}$ chosen by minimizing the following objective functions $\hat{\gamma} = \text{argmin}_{\gamma \epsilon \Psi}\, Q_T(\gamma)$. Once we obtain an estimate of $\hat{\gamma}$, which maximizes the explanatory power of the SETAR regressions, the corresponding slope coefficients and residual errors of the SETAR regressions are estimated via backward substitution. As a means of validating the threshold effects, Hansen (2000) suggests the use of a likelihood ratio ($LR(\lambda)$) statistic which tests the null hypothesis of no threshold effects against the alternative of threshold effects. In our empirical analysis, we obtain the following estimates of the SETAR process for Δdes_t:

$$des_t = \underset{(0.08)^{*}}{1.08} des_{t-1} \, I.(des \leq \lambda_{des}) + \underset{(0.07)^{*}}{0.36} des_{t-1} \, I.(des > \lambda_{des}) \tag{14}$$

$$\lambda_{des} = 0.47; \; LR(\lambda_{des}) = \underset{(0.00)^{***}}{12.42}; \; R^2 = 0.78; AIC = -349; MAPE = 48.7\%$$

Whereas for Δgdp we estimate the following asymmetric data generating process:

$$gdp_t = \underset{(0.01)^{**}}{0.59} gdp_{t-1} \, I.\left(gdp \leq \lambda_{gdp}\right) + \underset{(0.02)^{*}}{1.64} des_{t-1} \, I.\left(gdp > \lambda_{gdp}\right) \tag{15}$$

$$\lambda_{gdp} = 4.7; \; LR(\lambda_{gdp}) = \underset{(0.00)^{***}}{10.67}; \; R^2 = 0.76; AIC = -214; MAPE = 3.205\%$$

Our estimation results, as reported above, reveal that both des and gdp reject the null hypothesis of no thresholds and are therefore rendered as SETAR (1,1) processes. Based upon Hansen's (2000) SETAR modelling procedure, we obtain a threshold estimate of -1.08% for des, whereas a threshold of 4.7% is estimated for gdp. In deciding the lag period of the model, we apply the AIC and BIC rule to select the number of lag's to include and the estimation results show that a lag of 1 period is optimal for the SETAR model. From the coefficients of the lagged values in both the upper and lower regime, we find that des behaves in a persistent manner and seems to contain a unit root above its threshold level, whereas below this level des is not persistent and seems to be stationary. Conversely,gdp tends to evolve as a stationary, non-persistent process above its threshold level and exhibits unit root behaviour at rates of above 4.7 percent.

Distinguishing between nonlinearity and unit roots in the time series variables is considered important since they render different dynamics over the business cycle. Unit roots, on one hand, imply that a shock to either unemployment or output growth would lead to a new natural rate in the long-run. On the other side of the spectrum, asymmetric behaviour in the individual time series may be a result of hysteresis within the data generating process of the time series. Enders and Granger (1998) and Enders and Siklos (2001) demonstrate the problem of low power associated with traditional unit root tests when the underlying data generating process of time series is found to be asymmetric. Therefore, in order to formally validate our preliminary evidence of persistence amoung the variables, we proceed to apply the nonlinear unit root test of Bec, Salem and Carrasco (2004) (BBC hereafter) in order to test for the presence of unit roots against the null hypothesis of a stationary nonlinear SETAR process. The unit root test of BBC (2004) is based upon Dickey-Fuller's representation of Hansen's (2000) TAR model and can be specified as follows:

$$\Delta y_t = \sum_{r=r_1}^{r_2} \delta_r y_{t-1} y_{t-d}^r + \sum_{j=1}^{p} \alpha_j \Delta y_{t-j} + \varepsilon_t \tag{16}$$

Where ε_t~i.i.d.(0, σ^2). The unit root test is based upon the statistical significance of the parameters (δ_{r1},...,δ_{r2}). BBC set r_1=1, r_2=2 and derive the distribution from Supremum-based tests statistics on the Wald, Lagrange Multiplier (LM) and LR statistics in testing $\delta_1 = \delta_2 = 0$ against the null of $\delta_1 \neq 0$ or $\delta_2 \neq 0$ when the actual data generating process is a unit root and d=1. The distribution of the test statistics are nuisance free. The BBC unit root tests are performed on time series variables of gdp and des with the results reported below in Table 1.

$Table$ 1: BBC $Nonlinear$ $Unit$ $Root$ $Tests$

	Critical Values			des_t	gdp_t
τ_1	10%	5%	1%		
W_{sup}^{BBC}	16.18	18.4	23.01	128.28 (16.98)***	553.17 (41.74)
LM_{sup}^{BBC}	15.59	17.63	21.76	34.39 (12.41)***	43.31 (20.88)*
LR_{sup}^{BBC}	15.77	17.89	22.23	61.86 (14.46)***	119.71 (19.70)*

Significance level codes: '***", '**' and '*' denote the 1%, 5% and 10% significance levels respectively. Tests statistics for the first differences of the variables i.e. Δdes_t and Δgdp_t, are given in parenthesis.

From the results reported in table 1, we find that the all test statistics cannot reject the null hypothesis of a unit root for both des and gdp at their levels. However, in their first differences, we find that all test statistics cannot reject alternative hypothesis of a stationary TAR process, all except for the LM statistic obtained for gdp. We therefore conclude that that all series are characterized by a unit root in their levels, whereas their first differences reverted to a stationary TAR processes. All in all, we observe that the time series appear to be both nonlinear and non-stationary and as a result, we may (as a pre-speculation), assume that the time-series variables tends to asymmetrically move more or less together over time, a phenomenon that is later confirmed via formal co-integration analysis.

4.2 Threshold Cointegration Model Estimates

Given evidence of all series being integrated of order I(1), we then proceed to test for long-run equilibrium by employing the Ender and Silkos (2001) asymmetric cointegration methodology. Table 2 below presents the results of the threshold cointegration analysis performed for the nutrition and economic growth employing the TAR, c-TAR, MTAR and c-MTAR model specifications.

$Table\ 2$: $Threshold\ Cointegration\ Regressions$

$dependent\ variable$	des_t				gdp_t			
$model\ type$	tar	$c-tar$	$mtar$	$c-mtar$	tar	$c-tar$	$mtar$	$c-mtar$
ψ_{i0}	-0.21 (0.00)***	-0.21 (0.00)***	-0.21 (0.00)***	-0.21 (0.00)***	1.84 (0.00)***	1.84 (0.00)***	1.84 (0.00)***	1.84 (0.00)***
ψ_{i1}	0.15 (0.00)***	0.15 (0.00)***	0.15 (0.00)***	0.15 (0.00)***	5.25 (0.00)***	5.25 (0.00)***	5.25 (0.00)***	5.25 (0.00)***
$threshold\ variabe/value(\tau)$	0	0.038	0	0.015	0	0.07	0	-0.074
$\rho_1\xi_{t-1}$	0.009 (0.00)***	0.009 (0.00)***	-0.011 (0.00)***	-0.02 (0.00)***	0.007 (0.00)***	-0.007 (0.00)***	-0.003 (0.103)*	-0.007 (0.00)***
$\rho_2\xi_{t-1}$	-0.013 (0.00)***	-0.013 (0.00)***	0.009 (0.00)***	-0.01 (0.00)***	-0.008 (0.00)***	-0.008 (0.00)***	-0.011 (0.00)***	-0.004 (0.506)
$\beta_l\Delta\xi_{t-1}$	1.00 (0.00)***	1.00 (0.00)***	1.00 (0.00)***	1.02 (0.00)***	1.00 (0.00)***	1.00 (0.00)***	1.01 (0.00)***	0.99 (0.00)***
R^2	0.9895	0.9896	0.9899	0.9904	0.9851	0.9851	0.9873	0.9852
$H_0: \rho_1 = \rho_2 = 0$	22.31 (0.00)***	22.51 (0.00)***	20.80 (0.00)***	23.19 (0.00)***	12.585 (0.00)***	12.57 (0.00)***	18.495 (0.00)***	12.699 (0.00)***
$H_0: \rho_1 = \rho_2$	1.95 (0.169)	2.16 (0.148)	0.39 (0.535)	2.87 (0.098)*	0.21 (0.65)	0.19 (0.67)	7.76 (0.008)	0.35 (0.56)
aic	-495.72	-495.94	-494.10	-496.65	-336.414	-336.393	-343.828	-336.569
bic	-488.32	-488.54	-486.70	-489.25	-328.014	-328.992	-336.428	-329.168
$number\ of\ observations$	49	49	49	49	49	49	49	49

Asterisk (*) denotes 10% significance levels. Tests statistics for the coefficients from the threshold cointegration model and the p-values for the hypothesis testing are all given in parenthesis.

The estimation results depict that there is indeed a long-run relationship between the two variables as the null hypothesis of no cointegration is rejected in favour of the alternative hypothesis of threshold cointegration for all the estimated models. This implies that the long-run regression estimates can be interpreted with non-spurious interpretations. We find that for the gdp model regression, a one percent increase in the rate of nutrition intake is associated with an increase in gdp of roughly 5.2 percent. In terms of the nutrition-based efficiency wage hypothesis, this result indicates that an improvement in nutritional intake will lead to an improvement in productivity output, in terms of improved human capital input. Similar interpretations are deduced from the des regression, in which we find that a percentage increase in gdp results in an increase of nutritional intake of roughly 0.15 percent which indicates an income elasticity of nutrient of 0.15 for the observed data which is significantly different from zero. In translating this obtained result to the nutrition-based Engel curve, this implies that improved productivity will lead to improved nutritional status of the economy. In this sense, the above-described evidence leads to support of both the efficiency wage hypothesis and the Engel curve for South African data. Overall, our long-run regression elasticity estimates are in coherence with those obtained other studies like Bouis and Haddad (1992) for the Phillipines; Babatunde (2008) for Nigeria and Reutlinger and Selowsky (1976) for other emerging economies. It is also worth noting that all regression results have a strong explanatory power with a general coefficient of determination (R^2) of 0.98 being observed. Furthermore, all regressions passed the diagnostic tests such as the Durbin Watson (DW) for autocorrelation.

Subsequent to estimating our long-run regression, we model the TAR and MTAR variations of the residuals obtained from the des and gdp regressions (1) and (2). Using the conditional least squares (CLS) method as described in Hansen (1999), we obtain threshold values of 0.038 and 0.015, for the c-TAR and c-MTAR model when des is employed as a dependent variable, respectively. On the other hand, when gdp is used as the dependent variable, the estimated thresholds for the c-TAR and c-MTAR models are 0.07 and -0.074, respectively. As previously mentioned, a sufficient condition for validating asymmetric cointegration amoung the time series variables is that the residuals (i.e. ε_t) from equation (3) must be stationary i.e. ρ_1, $\rho_2 < 0$ and $(1-\rho_1)(1-\rho_2) < 1$. From table 2, we observe that only the c-MTAR model using des as a dependent variable and the c-TAR, MTAR and c-MTAR model for gdp as a dependent variable satisfy this condition. In the aforementioned models where the condition of stationary residuals is satisfied, we find that the speed of adjustment towards equilibrium is faster in the case of a shock to ε_t. We also find that the absolute parameter ρ_1 is higher compared to the estimated ρ_2 coefficient, for all estimated models with exception of the c-MTAR models for both des and gdp regressions.

In turning to more formal asymmetric cointegration tests, we find that all estimated models cannot reject the null hypothesis of symmetric cointegration in favour of the alternative of asymmetric cointegration effects, with exception for the c-MTAR model with des as a dependent variable. Therefore, as is based upon the presented empirical evidence, the c-MTAR specification is deemed to provide the most adequate description of asymmetric behaviour between the time series variables in contrast to the estimated TAR models. Following this evidence, an asymmetric co-integration relationship between nutrition and economic growth is validated and thus warrants the estimation of corresponding error correction mechanism (ECM) with long-run asymmetric equilibrium.

4.3 Threshold Error Correction Model Estimates

Having provided evidence supporting asymmetric adjustment, an asymmetric error correction model can be used to investigate the movement of the time-series variables towards their long-run equilibrium relationship. Table 3 reports the estimates of threshold error correction (TEC) model as given by equations (8) to (11). For each of the model specifications, the lag length was selected using the AIC information criterion. Key statistics are reported in table 3, including the null hypothesis of granger causality tests, cumulative asymmetric tests as well as symmetric momentum equilibrium adjustment path. It should be noted that the estimates of our threshold error correction models for the TAR, c-TAR, MTAR and c-MTAR models of both des and gdp as dependent variables are presented in table 3 inorder to provide a comparison between the obtained results.

$Table\ 3$: $(M)TAR - TEC\ RESULTS$

$dependent\ variable$	Δdes_t				Δgdp_t			
$model\ type$	$tar - tec$	$mtar - tec$	$c - tar - tec$	$c - mtar - tec$	$tar - tec$	$mtar - tec$	$c - tar - tec$	$c - mtar - tec$
$\lambda^{-}\xi^{-}_{t-1}$	-0.036 (0.00)***	-0.036 (0.00)***	-0.0361 (0.00)***	-0.0383 (0.00)***	-0.056 (0.00)***	-0.057 (0.00)***	-0.0603 (0.00)***	-0.065 (0.00)***
Δdes^{-}_{t-k}	0.8191 (0.00)***	0.8102 (0.00)***	0.8250 (0.00)***	0.7771 (0.00)***	0.2697 (0.04)*	0.5236 (0.00)***	0.2269 (0.09)*	0.3656 (0.01)*
Δgdp^{-}_{t-k}	-0.0051 (0.84)	0.072 (0.00)***	0.0683 (0.00)***	0.083 (0.00)***	1.0332 (0.00)***	0.9657 (0.00)***	1.0474 (0.00)***	0.9983 (0.00)***
$\lambda^{+}\xi^{+}_{t-1}$	-0.0363 (0.02)*	-0.037 (0.00)***	-0.0326 (0.03)*	-0.0272 (0.00)***	-0.1319 (0.00)***	-0.038 (0.02)*	-0.1512 (0.00)***	-0.061 (0.00)**
Δdes^{+}_{t-k}	0.9163 (0.00)***	0.9144 (0.00)***	0.9336 (0.00)***	0.8676 (0.00)***	0.3059 (0.08)*	0.6761 (0.00)***	0.2147 (0.20)	0.6224 (0.00)***
Δgdp^{+}_{t-k}	-0.0051 (0.84)	-0.003 (0.78)	-0.0116 (0.653)	-0.0029 (0.78)	0.9414 (0.00)***	0.7869 (0.00)***	0.9753 (0.00)***	0.8229 (0.00)***
R^2	0.9994	0.9994	0.9994	0.9994	0.9998	0.9999	0.9999	0.9998
$number\ of\ observations$	49	49	49	49	49	49	49	49
DW	0.741 (0.00)***	0.75 (0.00)***	0.745 (0.00)***	0.776 (0.00)***	0.582 (0.00)***	0.65 (0.00)***	0.580 (0.00)***	0.550 (0.00)***
H_1	0.000 (0.985)	0.118 (0.73)	0.066 (0.798)	3.569 (0.07)*	4.30 (0.45)**	6.13 (0.01)**	8.167 (0.00)***	0.000 (0.986)
H_{20}	19.783 (0.00)***	29.361 (0.00)***	20.877 (0.00)***	30.965 (0.00)***	396.38 (0.00)***	426.01 (0.00)***	429.99 (0.00)***	473.625 (0.00)***
H_{21}	114.276 (0.00)***	348.71 (0.00)***	114.864 (0.00)***	206.986 (0.00)***	2.189 (0.13)*	37.67 (0.00)***	1.54 (0.23)	20.526 (0.00)***
H_{30}	15.028 (0.00)***	58.31 (0.00)***	18.56 (0.00)***	61.909 (0.00)***	3.892 (0.05)*	59.26 (0.00)***	2.879 (0.09)*	37.127 (0.00)***
H_{31}	3.079 (0.087)*	7.76 (0.08)*	4.267 (0.045)**	8.148 (0.00)***	0.075 (0.79)	3.03 (0.09)*	0.010 (0.920)	9.536 (0.00)***
H_{40}	15.028 (0.00)***	58.31 (0.00)***	18.56 (0.00)***	61.909 (0.00)***	3.89 (0.06)*	59.26 (0.00)***	2.879 (0.09)*	37.127 (0.00)***
H_{41}	3.079 (0.087)*	7.76 (0.08)*	4.267 (0.045)**	8.148 (0.00)***	0.075 (0.79)	3.03 (0.09)*	0.010 (0.920)	9.536 (0.00)***

Significance level codes: '***", '**' and '*' denote the 1%, 5% and 10% significance levels respectively. P-values are reported in parenthesis.

From the results presented in Table 3, the estimates of all error correction terms are found to be significant for both the long-run and short-run regression estimates and the predictive power of the asymmetric error-correction models as measured by the R^2 statistic are found to be encouragingly high. The sign and trend of deviations are important in determining how quickly policymakers are likely to respond to deviations from equilibrium. In particular, the value of the adjustment parameters determines the speed of reversion back to steady-state equilibrium when nutrition and economic growth temporarily depart from their underlying equilibrium relationship following either a negative or positive shock to the variables. For instance, the p-values for the estimates of $\lambda^{-}\xi_{t-1}^{-}$ and $\lambda^{+}\xi_{t-1}^{+}$ indicate that for all econometric models, with the exception of the c-MTAR-TEC model, a positive shock to des results in a quicker adjustment back to its long-run equilibrium in comparison to the effect of a negative shock. Conversely, we find that for both TAR-TEC and C-TAR-TEC models, a positive shock to gdp will result in a much quicker reversion back to steady-state equilibrium in contrast to a negative shock, whereas on the other hand, the M-TAR-TEC model responds slightly stronger to negative shocks when compared to positive shocks. In further taking into consideration the asymmetric cointegration tests on the TEC models, we find that the F-statistic rejects that the null hypothesis of symmetric cointegration adjustment (i.e. the coefficients λ^{-} and λ^{+} are equal) for the TAR-TEC, MTAR-TEC and c-TAR-TEC models on the Δgdp regression whereas the same hypothesis can only be rejected for the c-TAR-TEC model on the Δdes regression. Moreover, it should be noted that the adjustment coefficients between the various models, and, do not appear noticeably different.

Moreover, we applied the Granger causality tests based on the TEC models to examine causal relations between nutrition and economic growth. The hypotheses of granger causality between nutrition and economic growth are assessed with F-tests. Generally speaking, causality between the two time series variables is found to run bi-directional, with the exception for the c-TAR-TEC model with Δgdp as a dependent variable where causality is found to run from gdp to des. These results provide overwhelming evidence in favour of nutritional-intake having a two-way co-relationship with wealth and income. This, in conjunction with the significant estimates obtained from the cointegration results presented in Table 2, strongly advocates for the existence of both the efficiency wage hypothesis and Engel curve for the case of South Africa. However, nutrition appears to have a stronger causal effect on gdp compared to the impact of economic growth on nutritional intake, a finding which is similar to that obtained in Tiffin and Dawson (2002) yet contrary to that obtained in Wang and Taniguchi (2001). The threshold co-integration tests reveal significant asymmetric co-integration for the c-MTAR-TEC with Δdes as a dependent variable and for the TAR, M-TAR and c-TAR models with Δgdp as a dependent variable. The p-values obtained from the hypothesis testing short-run dynamics are found to be significant for all model specifications with the exception of the TAR-TEC and MTAR-TEC models with

Δgdp as a dependent variable, in which only the short-run dynamics of the coefficients associated with nutrition are found to be significant. Furthermore, the cumulative asymmetric effects are also examined. We find strong evidence of asymmetric cumulative effects both upwards and downwards for all estimated models. The final type of asymmetry examined is the momentum equilibrium adjustment path asymmetries, in which all estimated models, with the exception of the TAR models on the Δgdp regression.

Several interesting facts emerge regarding the overall estimation of our asymmetric cointegration and error correction model. In general, the results of the asymmetric co-integration test for the TEC models as reported in table 3 are similar to those performed for the co-integration models in table 2, in the sense of the c-MTAR model with Δdes as a dependent variable, being the only regression which cannot reject alternative hypothesis of asymmetric co-integration among the time series variables. Therefore our results indicate that the des (and not gdp) is responsible for asymmetric cointegration adjustments between the two variables. In contrasting the four estimated models, the MTAR with consistent threshold estimate is clearly the best model based upon the cointegration tests. According to the presented empirical evidence, MTAR model has a better explanatory ability of asymmetric cointegration between nutrition and economic growth in South Africa in comparison to their counterpart TAR specifications. We, therefore, conclude our empirical analysis by declaring that a smooth adjustment co-integration relation exists between nutritional intake and economic growth in South Africa.

5. CONCLUSION OF THE STUDY

The United Nations Millennium Development Goals (MDG) deem that income poverty and malnutrition, as indicators of poverty, should be halved by the year 2015 whereas the Accelerated and Shared Growth of South Africa (ASGISA) have a set an economic growth target of 6 percent as planned to be achieved by 2014. By analysing the asymmetric cointegration effects between nutrition and economic growth for annual South African data, our study presents a number of intriguing policy implications. In particular, we find strong empirical evidence in support of both the efficiency wage hypothesis and the Engel curve for South African data. We take this finding to be of considerable importance since it draws the implication that the health status of the South African economy, through its nutritional intake, bears a two-way relationship with productivity and ultimately income wealth. While the common yet sole use of the income-elasticity of calorie intake, as obtained from the Engel curve, may reveal how nutritional intake is affected income, it infers little to policymakers on how productivity output affects the diet consumption within the economy. This result may produce limitations on the efficiency of policy formulation, as it places policymakers under the impression that the sole reliance on development strategies through nutritional programs aimed at improving economic growth may be sufficient for overall economic development.

Our study therefore adheres that whatever the implied individual merits of implemented developments policies are, they are consequentially of limited value if policymakers do not directly address nutritional issues within economic development programs. With specific reference to policymakers in developing or emerging economies and for international aid agencies, an important conclusion which can be drawn from our study is that all policies- including food aid – which enhance food security and reduce undernourishment in developing countries can account for improvements in economic growth. In advocating for improved policies which strengthen food security by focusing on their humanitarian benefits, the implications drawn from our study may serve as a reminder that the direct focus on nutrition policies should neither be ignored or be planned in isolation but should be implemented in conjunction with economic growth policies. On the other end of the spectrum, relying solely upon labour markets interventions and economic growth strategies is not sufficient enough to eradicate current poverty problems. We conclude that the economic returns to investing in nutritional programs far outweigh their costs and policy reforms supporting productivity growth need to be accompanied by strategic investment programs aimed at tackling the overriding problem of poverty via nutrition- specific development programs.

While we are able to establish a significant, positive correlation between nutrition and economic growth we, however, interpret our overall findings with extreme caution as the asymmetries found in the cointegration relation between nutritional intake and economic growth in South Africa present reservations with regards to interpreting our obtained empirical results. Specifically, we find that the cointegration asymmetries in the nutrition-productivity co-relation are a result of slow adjustments back to the long-run steady-state equilibrium in the face of negative and positive shocks to both nutrition and productivity output. In other words, policy-induced shocks to either nutritional intake or productivity output, as implemented through various development policies, will result in slow reversion or responses to the counter variable as they deviate from cointegrated long-run steady state equilibrium. A plausible explanation for such a long delay in the equilibrium adjustment process following a shock to the variables, may be that policymakers have not yet identified and thus explored other possible "avenues" which directly link nutritional programs towards improved economic growth and vice-versa. Identifying and exploring the use these intermediary channels between nutrition intake and economic growth may serve well as a candidate for potential future research.

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