# The Demand for Calories in Turkey

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2011

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## Abstract

This research examines dynamic causal relationships between per capita calorie intake, per capita income and food prices using time series data for Turkey during 1965-2007.
ARDL cointegration analysis yields an income elasticity of calorie intake of 0.22, while the food-price elasticity is insignificant. The results suggest that economic growth in Turkey has improved calorie intake; future income growth can alleviate further inadequate nutrition. This result confirms Engel's law too.
An augmented form of Granger causality analysis is conducted amongst the variables. The short-run causality testing reveals the existence of only one causality which is running from income to calorie intake. The post-sample variance decompositions indicate that income is the main cause of the increased calorie intake in the long-run. The estimated long-run model appears to have stable parameters.

**Keywords:** calorie intake, cointegration, Granger-causality, Turkey

**JEL Classifications:** C22, D10, E20, F10

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## 1. Introduction

Malnutrition is mainly a reflection of poverty: people are deprived of food since they do not have enough income. The adverse impact of malnutrition on economic growth was highlighted initially in Correa (1970) and subsequently was improved by Cole (1971). Malnutrition may impede the economic growth of developing countries in many ways, such as: by reducing the life expectancy which reduces the productivity years to be expected from newly born children; by reducing resistance to disease which increases absence from work; and by inhibiting the mental and physical development of children's growth which decreases their potential productivity. Conventional wisdom in development economics suggests that malnutrition will disappear only with improvements in income that accompany the development process in the long-run. Therefore, traditionally, income policies have pursued an improvement in the availability of nutrition with a view to improving human development. This orthodox view in development economics has been challenged in some empirical studies, suggesting that increases in income will not result in substantial improvements in nutrition intakes. The economic policies designed to reduce poverty do not imply the elimination of malnutrition, as some empirical evidence on the calorie-income elasticity is found to be close zero. Thus, the nature of the relationship between calorie intake and income is crucial and much of the literature on malnutrition analyzes this relationship empirically. Empirical studies concerning calorie intake and income are broadly to test two hypotheses: is the calorie intake determined by income or vice versa? The former is based on the estimation of calorie-demand relationships, while the latter is related to the efficiency wage hypothesis. The motivation of this study is two fold: as far as this study is concerned, there exists no calorie demand estimation for Turkey using time series econometric techniques and the adopted econometric methodology. The bounds testing approach to cointegration has not been utilized previously in the literature of the demand for calories. This study focuses on the calorie demand income relationship at aggregate level rather than the institutional and structural aspects of changes in food consumption, which requires microeconomic aspects of household income.

The objectives of this study are as follows: i) to estimate the calorie-income elasticities both in the short-run and long-run using time-series econometric techniques; ii) to establish the direction of causal relationships between calorie intake, income, and price within and out of the sample period; and iii) to implement parameter stability tests to ascertain stability or instability in the calorie demand function.

The remainder of this paper is organized as follows. Section 2 presents a short literature review. Section 3 describes the study's model and methodology. Section 4 discusses the empirical results, and finally Section 5 concludes.

**2. A Brief Literature Review**

In development economics, there are two theories concerning the relationship between calorie consumption and income. The first strand of inquiry is based on the efficiency wage hypothesis that was initiated by Leibenstein (1957) and then was theorized by Stiglitz (1976). According to this hypothesis, the efficiency of workers depends on their wages through their nutrition that their income allows them to purchase. The empirical results for the efficiency wage hypothesis are mixed. There are some empirical studies to support this hypothesis, see for example, Strauss (1986) for Sierra Leone; Shan and Alderman (1988) for Sri Lanka; Van Den Boom *et al*. (1996) for Ghana; Behrman *et al*. (1997) for Pakistan; Weinberger (2004) for India; while Deolalikar (1988) finds no evidence that nutrition determines wages in India. Korjenek (1992) provides detailed summary results of the empirical studies concerning the efficiency wage hypothesis. The critical review of this approach and further theoretical suggestions are found in Strauss and Thomas (1998).

The second line of inquiry relates to calorie consumption and income. This approach also allows for testing the Engel's curve hypothesis which states that the proportion of income spent on food diminishes as income rises. Much of the literature on the calorie-income nexus tests the existence of the Engel's curve hypothesis. As Dawson and Tiffin (1998) summarizes, there are two approaches in testing the Engel's curve hypothesis. The direct approach estimates a reduced form Engel equation of the demand for calories. The indirect approach estimates food demand/expenditure systems for a small number of food groups and then converts the resulting food-income elasticities using constant calorie-to-food conversion factors. Reutlinger and Selowsky (1976) presented the first econometric result on the existence of the Engel's curve using cross-sectional and cross-country data. The key parameter in the studies of calorie-demand is the elasticity of calorie demand with respect to income. Table 1 displays the summary results of some empirical evidence on calorie-income elasticities. As can be seen from Table 1, the magnitude of calorie-income elasticities vary substantially due to food group's aggregation, data frequency, variable definition, method of approach and specification, and estimation method. However, many studies are in support of the Engel's curve hypothesis. A comprehensive review and details of the calorie-demand estimations in the 1980s are presented in Bouis and Haddad (1992) arguing that much of the variation in the calorie-income elasticity estimates can be explained very simply by the particular calorie and income variables that are used in the regression analysis. The reliability of the calorie-income elasticities were also questioned in Bouis (1994) in the selected countries such as Bangladesh, India, Indonesia, Pakistan, Sri Lanka and Thailand, suggesting that the survey techniques developed by nutritionists give more reasonable values more often than the survey techniques used by economists, in the long-run.

**Table 1. Calorie-Income Elasticity Estimates**

| Author(s) | Elasticity | Data | Method | Country | Causality |
|---|---|---|---|---|---|
| Reutlinger and Selowsky (1976) | 0.17 | CS | OLS | Developing countries | NA |
| Strauss (1984) | 0.82 | CS | OLS | Sierra Leona | NA |
| Dawson (1997) | 0.07 | CS | OLS | 41 Developing countries | NA |
| Dawson and Tiffin (1998) | 0.34 | TS | JJ | India | Y→C |
| Angulo *et al*. (2001) | various | TS | FIML | 14 EU plus Norway | NA |
| Dawson (2002) | 0.19 | TS | JJ | Pakistan | Y→C |
| Tiffin and Dawson (2002) | 0.31 | TS | JJ | Zimbabwe | Y↔C |
| Mushtaq *et al*. (2007) | 0.21 | TS | JJ | Pakistan | Y→C |
| Neeliah and Shankar (2008) | NA | TS | JJ | Mauritius | No causality |
| Dawson and Sanjuan (2011) | 0.25 | PD | PEM | 41 Developing countries | Y→C |
| Ogundari (2011) | 0.06 | TS | JJ | Nigeria | Y→C |

Keys: CS (Cross-Section), TS (Time Series), PD (Panel Data), FIML (Full Information Maximum Likelihood), OLS (Ordinary Least Squares), JJ (Johansen-Juselius cointegration method), PEM (Panel Econometric Methods), NA (Not available/applied), Y→C indicates that the causality runs from income (Y) to calorie intake (C), Y↔C indicates bi-directional causality.

The empirical evidence obtained for the calorie-income relationship using household data is outlined in Gibson and Rozelle (2002). Examples of this approach includes Strauss (1986) for Sierra Leone (1986); Behrman and Deolalikar (1987) for India; Subramanian and Deaton (1996) for India and Behrman *et al*. (1997) for Pakistan. The results provide a wide range of calorie-income elasticities.

## 3. Model and Methodology

Considering the empirical literature on calorie demand, this study adopts the following long-run relationship between calorie intake, income and food prices in double linear logarithmic form as:

$$c_t = a_0 + a_1 y_t + a_2 p_t + \varepsilon_t , \qquad (1)$$

where $c_t$ is the logarithm of calorie intake per capita in a day ; $y_t$ is the logarithm of real per capita income; $p_t$ is the logarithm of real food prices; and $\varepsilon_t$ is the classical error term.

The short-run dynamic adjustment process of the long-run relationship in equation (1) may provide useful policy recommendations. It is possible to incorporate the short-run dynamics into equation (1) by expressing it in an error-correction model, as suggested in Pesaran *et al.* (2001).

$$\Delta c_t = \beta_0 + \sum_{i=1}^{n1} \beta_{1i} \Delta c_{t-i} + \sum_{i=0}^{n2} \beta_{2i} \Delta y_{t-i} + \sum_{i=0}^{n3} \beta_{3i} \Delta p_{t-i} + \beta_4 c_{t-1} + \beta_5 y_{t-1} + \beta_6 p_{t-1} + v_t \qquad (2)$$

This approach, also known as autoregressive-distributed lag (ARDL)[1], provides the short-run and long-run estimates simultaneously. Short-run effects are reflected by the estimates of the coefficients attached to all first-differenced variables. The long-run effects of the explanatory variables on the dependent variable are obtained by the estimates of $\beta_5$-$\beta_6$ that are normalized on $\beta_4$. The inclusion of the lagged-level variables in equation (2) is verified through the bounds testing procedure, which is based on the Fisher (F) or Wald (W)-statistics. This procedure is considered as the first stage of the ARDL cointegration method. Accordingly, a joint significance test that implies no cointegration hypothesis, ($H_0$: all $\beta_4$ to $\beta_6 = 0$), against the alternative hypothesis, ($H_1$: at least one of $\beta_4$ to $\beta_6 \neq 0$) should be performed for equation (2). The F/W test used for this procedure has a non-standard distribution. Thus, Pesaran *et al.* (2001) compute two sets of critical values for a given significance level with and without a time trend. One set assumes that all variables are $I(0)$ and the other set assumes they are all $I(1)$. If the computed F/W-statistic exceeds the upper critical bounds value, then the $H_0$ is rejected, implying cointegration. In order to determine whether the adjustment of variables is toward their long-run equilibrium values, estimates of $\beta_4$-$\beta_6$ are used to construct an error-correction term (EC). Then lagged-level variables in equation (2) are replaced by $EC_{t-1}$ forming a modified version of equation (2) as follows:

$$\Delta c_t = \beta_0 + \sum_{i=1}^{n1} \beta_{1i} \Delta c_{t-i} + \sum_{i=0}^{n2} \beta_{2i} \Delta y_{t-i} + \sum_{i=0}^{n3} \beta_{3i} \Delta p_{t-i} + \lambda EC_{t-1} + \mu_t \qquad (3)$$

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[1] Some recent applications of the ARDL approach to cointegration can be found in following studies: Halicioglu (2011, 2010, and 2007), Yavuz *et al.* (2010), Ay and Yardımcı (2007), Erbaykal and Sürekçi (2006), and Kasman *et al.* (2005).

Equation (3) is re-estimated one more time using the same lags previously. A negative and statistically significant estimation of $\lambda$ not only represents the speed of adjustment but also provides an alternative means of supporting cointegration between the variables. Pesaran *et al.* (2001) cointegration approach has some methodological advantages in comparison to other single cointegration procedures. Reasons for the ARDL are: i) endogeneity problems and inability to test hypotheses on the estimated coefficients in the long-run associated with the Engle-Granger (1987) method are avoided; ii) the long and short-run coefficients of the model in question are estimated simultaneously; iii) the ARDL approach to testing for the existence of a long-run relationship between the variables in levels is applicable irrespective of whether the underlying regressors are purely stationary $I(0)$, purely non-stationary $I(1)$, or mutually cointegrated; iv) the small sample properties of the bounds testing approach are far superior to that of multivariate cointegration, as argued in Narayan (2005).

The Granger representation theorem suggests that there will be Granger causality in at least one direction if there exists a cointegration relationship among the variables in equation (1), providing that they are integrated order of one. Engle and Granger (1987) caution that the Granger causality test, which is conducted in the first-differenced variables by means of a VAR, will be misleading in the presence of cointegration. Therefore, inclusion of an additional variable to the VAR system, such as the error correction term would help us to capture the long-run relationship. To this end, an augmented form of the Granger causality test involving the error correction term is formulated in a multivariate $p$th order vector error correction model.

$$(1-L)\begin{bmatrix} c_t \\ y_t \\ p_t \end{bmatrix} = \begin{bmatrix} \theta_1 \\ \theta_2 \\ \theta_3 \end{bmatrix} + \sum_{i=1}^{p}(1-L)\begin{bmatrix} \phi_{11i}, \phi_{12i}, \phi_{13i} \\ \phi_{21i}, \phi_{22i}, \phi_{23i} \\ \phi_{31i}, \phi_{32i} \phi_{33i} \end{bmatrix}\begin{bmatrix} c_{t-i} \\ y_{t-i} \\ p_{t-i} \end{bmatrix} + \begin{bmatrix} \delta_1 \\ \delta_2 \\ \delta_3 \end{bmatrix}\left[EC_{t-1}\right] + \begin{bmatrix} \omega_{1t} \\ \omega_{2t} \\ \omega_{3t} \end{bmatrix} \quad (4)$$

$L$ is the lag operator. $EC_{t-1}$ is the error correction term, which is obtained from the long-run relationship described in equation (1), and it is not included in equation (4) if one finds no cointegration amongst the vector in question. The Granger causality test may be applied to equation (4) as follows: i) by checking statistical significance of the lagged differences of the variables for each vector; this is a measure of short-run causality; and ii) by examining statistical significance of the error-correction term for the vector that there exists a long-run relationship. As a passing note, one should reveal that equation (3) and (4) do not represent competing error-correction models because equation (3) may result in different lag structures on each regressors at the actual estimation stage; see Pesaran *et al*. (2001) for details and its mathematical derivation. All error-correction vectors in equation (4) are estimated with the same lag structure that is determined in unrestricted VAR framework.

Establishing Granger causality is restricted to essentially within sample tests, which are useful in distinguishing the plausible Granger exogeneity or endogenity of the dependent variable in the sample period, but are unable to deduce the degree of exogenity of the variables beyond the sample period. To examine this issue, the decomposition of variance of the variables may be used. The variance decompositions (VDCs) measure the percentage of a variable's forecast error variance that occurs as the result of a shock (or an innovation) from a variable in the system. Sims (1980) notes that if a variable is truly exogenous with respect to the other variables in the system, its own innovations will explain the entire variable's forecast error variance (i.e., almost 100%). By looking at VDCs, policy makers gather additional insight as to what percentage (of the forecast error variance) of each variable is explained by its determinant.

## 4. Results

Annual data over the period 1965-2007 were used to estimate equation (2) and (3) by the ARDL cointegration procedure of Pesaran *et al*. (2001). Variable definition and sources of data are cited in the Appendix.

To implement the Pesaran *et al*. (2001) procedure, one has to ensure that none of the explanatory variables in equation (1) is above $I(1)$. Three tests were used to test unit roots in the variables: Augmented Dickey-Fuller (henceforth, ADF) (1979, 1981), Phillips-Perron (henceforth, PP) (1988), and Elliott-Rothenberg-Stock (henceforth, ERS) (1996). Unit root tests results are displayed in Table 2 which warrant implementing the ARDL approach to cointegration, as the variables are in the combination of $I(0)$ and $I(1)$. To account for a possible endogenous structural break in the series, Zivot and Andrews (ZA) (1992) test was carried out. A brief methodological explanation of this test is placed in the Appendix. The results of the ZA are demonstrated in Table 3. The ZA results are in line with the unit root tests.

**Table 2. Unit Root Results**

| Variables | ADF | PP | ERS |
|---|---|---|---|
| $c_t$ | 4.97* | 4.01* | 4.60* |
| $y_t$ | 2.07 | 1.37 | 1.62 |
| $p_t$ | 2.41 | 2.63 | 1.53 |
| $\Delta c_t$ | 8.80* | 12.7* | 4.72* |
| $\Delta y_t$ | 3.61* | 6.92* | 3.65* |
| $\Delta p_t$ | 3.94* | 7.34* | 3.45* |

Notes: The sample level unit root regressions include a constant and a trend. The differenced level unit root regressions are with a constant and without a trend. All test statistics are expressed in absolute terms for convenience. Rejection of unit root hypothesis is indicated with an asterisk. Δ stands for first difference.

**Table 3. Zivot-Andrews Unit Root Test**

| Variables | Intercept | | | Trend | | | Both | | |
|---|---|---|---|---|---|---|---|---|---|
| | *k* | *t* | *TB* | *k* | *t* | *TB* | *k* | *t* | *TB* |
| $c_t$ | 0 | 5.29* | 1997 | 0 | 6.95* | 1991 | 0 | 6.91* | 1987 |
| $y_t$ | 4 | 3.41 | 2001 | 4 | 4.15 | 1998 | 4 | 4.24 | 1997 |
| $p_t$ | 0 | 4.32 | 1999 | 0 | 4.31 | 1996 | 0 | 4.36 | 1994 |

Notes: Estimation with 0.15 trimmed. Maximum lag length (*k*) is determined by AIC. *t* is the *t*-test statistic calculated. All test statistics are expressed in absolute terms for convenience. *TB* is the time of break. Rejection of unit root with a structrual break hypothesis is indicated with an asterisk.

The presence of long-run relationship was established applying a bounds test to equation (2). Considering that this study is utilizing annual data with a small sample size including only 43 observations, the maximum lag length in the ARDL model was set equal to 2. The results of the bounds testing are reported in Table 4. Table 4 illustrates that the computed F/W statistics are above the upper bound values at all level of significance confirming the existence of a cointegration relationship between the variables of equation (1).

**Table 4. The results of F and W tests for cointegration**.

| The assumed long-run relationship; $F/W(c \mid y, p)$ | | | | |
|---|---|---|---|---|
| F-statistic | 95% LB | 95% UB | 90% LB | 90% UB |
| 14.71 | 4.07 | 5.17 | 3.33 | 4.31 |
| W-statistic | | | | |
| 44.15 | 12.23 | 15.53 | 9.99 | 12.95 |

If the test statistic lies between the bounds, the test is inconclusive. If it is above the upper bound (UB), the null hypothesis of no level effect is rejected. If it is below the lower bound (LB), the null hypothesis of no level effect cannot be rejected.

On establishing a long-run cointegration relationship amongst the variables of equation (1), a two-step procedure to estimate the ARDL model was carried out. First, in search of the optimal lag length of the differenced variables of the short-run coefficients, Adjusted R-bar Squared Criterion ($\overline{R}^2$) was utilized and in the second step and then the ARDL model was estimated. The results of $\overline{R}^2$ based ARDL model are displayed in Panel A, B, and C of Table 5. The results of long-run coefficients are presented in Panel A of Table 5, whereas the short-run estimates are reported in Panel B of Table 5. Finally, Panel C of Table 5 demonstrates the short-run diagnostic test results. The overall regression results are satisfactory in terms of diagnostic tests. The short-run diagnostics obtained from the estimation of equation (2) suggest that the estimated model is free from a series of econometric problems such as serial correlation, functional form, normality, and heteroscedasticity. The long-run elasticity of calorie demand, with respect to income, is 0.2202 suggesting that for each 1% increase in the per capita income, the per capita daily calorie intake will rise by about 0.22%. Thus, the policies aiming at increasing per capita income in order to alleviate malnutrition will not result in substantial improvements in average daily per capita calorie consumption. The nutrition deficiency, however, may not improve if people diversify their diets when their income rises, as they may subsitute more expensive calories in place of less expensive calories. Moreover, people may resort to further substitution by consuming complements to good nutrition, such as healthy foods, sanitation or medical services. The impact of food prices on daily calorie intake is insignificant. The speed of adjustment parameter is –0.6475, suggesting that when the calorie demand equation is above or below its equilibrium level, it adjusts by 65% within the first year. The full convergence to its equilibrium level takes a little less than two years.

**Table 5. ARDL Cointegration Results**

| Panel A : Long-run estimates<br>Dependent variable $c_t$ | | | Panel B: Error correction representation results<br>Dependent variable $\Delta c_t$ | | |
|---|---|---|---|---|---|
| Regressor | Coefficient | T-ratio | Regressor | Coefficient | T-ratio |
| $y_t$ | 0.2202* | 5.9175 | $\Delta y_t$ | 0.0346 | 0.0495 |
| $p_t$ | 0.1E-3 | 0.0993 | $\Delta p_t$ | 0.0014 | 0.1914 |
| Constant | 2.6798* | 2.8660 | $EC_{t-1}$ | -0.6475* | 5.6442 |

| Panel C: Diagnostic test results | | | | | | | |
|---|---|---|---|---|---|---|---|
| $\overline{R}^2$ | 0.78 | F-statistic | 24.1* | $\chi^2_{SC}(1)$ | 0.005 | $\chi^2_{FF}(1)$ | 3.063 |
| RSS | 0.01 | DW-statistic | 2.01 | $\chi^2_{N}(2)$ | 0.497 | $\chi^2_{H}(1)$ | 1.056 |

*, **, and, *** indicate, 1%, 5%, and 10% significance levels respectively. RSS stands for residual sum of squares. T-ratios are in absolute values. $\chi^2_{SC}$, $\chi^2_{FF}$, $\chi^2_{N}$, and $\chi^2_{H}$ are Lagrange multiplier statistics for tests of residual correlation, functional form mis-specification, non-normal errors and heteroskedasticity, respectively. These statistics are distributed as Chi-squared variates with degrees of freedom in parentheses. The critical values for $\chi^2(1) = 3.84$ and $\chi^2(2) = 5.99$ are at 5% significance level.

**Table 6. Results of Granger Causality**

| | *F*-statistics (probability) | | | |
|---|---|---|---|---|
| Dependent Variable | $\Delta c_t$ | $\Delta y_t$ | $\Delta p_t$ | $EC_{t-1}$ (*t*-statistic) |
| $\Delta c_t$ | - | 2.91** (0.06) | 1.96 (0.15) | -0.14 (1.30) |
| $\Delta y_t$ | 0.23 (0.28) | - | 0.58 (0.56) | |
| $\Delta p_t$ | 0.53 (0.58) | 2.12 (0.13) | - | |
| Causality inference : $y \rightarrow c$. | | | | |

* and ** indicate 5 % and 10 % significance levels, respectively. The probability values are in brackets. The optimal lag length is 2 and is based on SBC.

As can be seen in Table 6, the Augmented Granger causality tests suggest the non-existence of a long-run causality amongst the variables. However, there is a short-run casusality running from income to the calorie intake.

**Table 7. Decomposition of Variance**

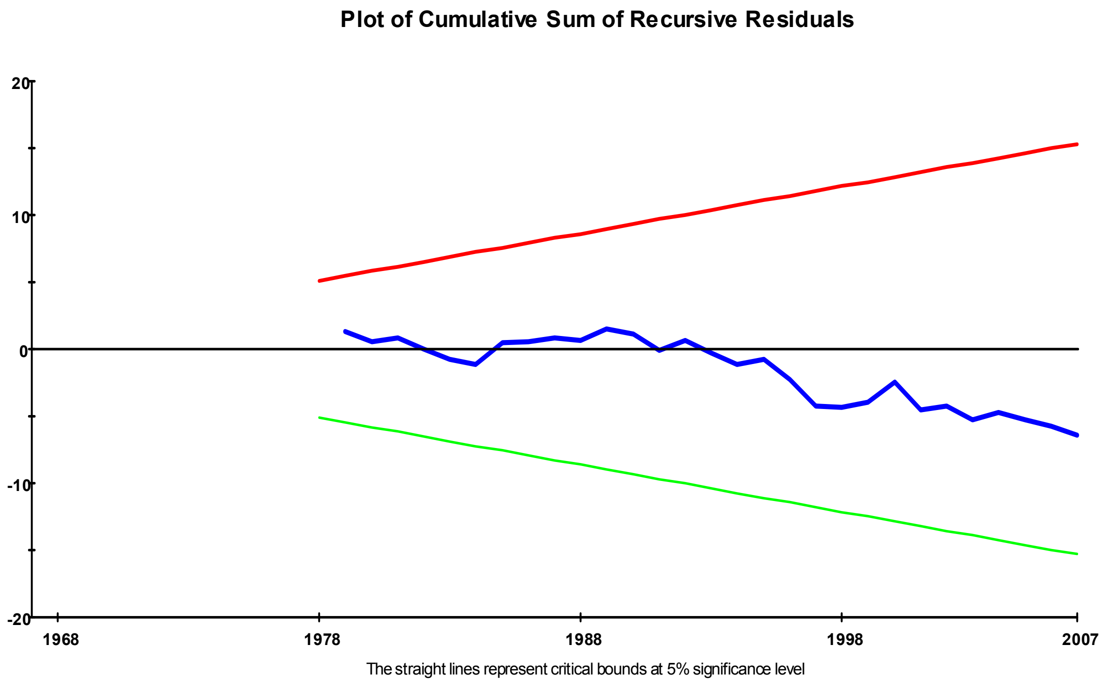
| Percentage of forecast variance explained by innovations in: | | | |
|---|---|---|---|
| Years | Calorie intake | Income | Price |
| 1 | 100.00 | 0.00 | 0.00 |
| 2 | 80.34 | 9.58 | 10.06 |
| 3 | 65.77 | 23.17 | 11.05 |
| 4 | 55.16 | 34.41 | 10.41 |
| 5 | 47.00 | 43.74 | 9.25 |
| 10 | 23.00 | 72.31 | 4.59 |

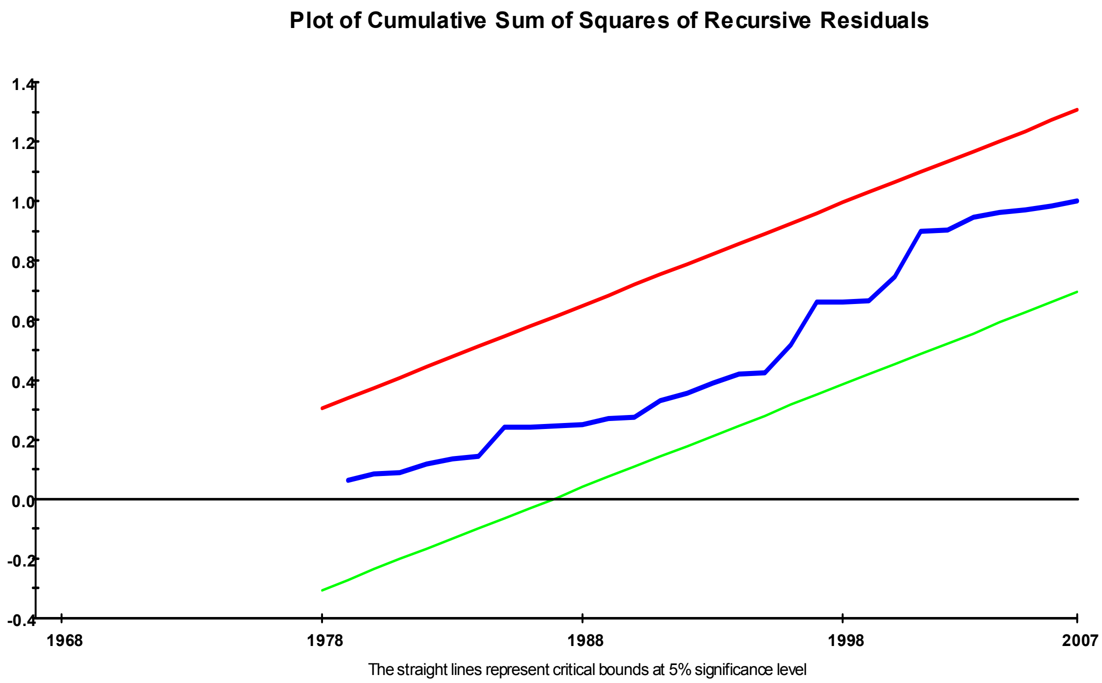
Notes: Figures in the first column refer to horizons (i.e., number of years). All figures are rounded to two decimal places. The covariances matrices of errors from all the VECMs appeared to be very small and approaching zero suggesting that the combinations of all the variables in these models are linear. Therefore, the ortohogonal case for the variance decompositions are applied.

Table 7 provides the summary results for the VDCs. As for the VDCs, a substantial portion of the variance of calorie intake (80.34%) is explained by its own innovations in the short-run, for example, at the two-year horizon. In the long-run, for example, at the ten-year horizon, the portion of the variance of calorie intake substantially decreases to 23% implying that other variables explain about 77% of the shocks in the calorie intake. The post-sample VDCs also indicate that 72.31% of the shocks in the calorie intake is due to innovations in income at the ten year-horizon, emphasing the fact that income is the main cause of the calorie intake.

The stability of the coefficients estimates in equation (3) was checked through the CUSUM and CUSUMSQ tests of Brown *et al.* (1975). The results are presented in Figures 1 and 2 respectively. The graphical representation of these tests demonstrate that the estimated coefficients are stable in the $\overline{R}^2$ based error-correction model as the plots of CUSUM and CUSUMSQ statistics are within the critical bounds. Therefore, the policy makers may use the regression coefficients to draw policy recommendations.

## Figure 1 and 2 : CUSUM and CUSUMSQ Plots for Stability Tests



## 5. Conclusions

This study has tested the long-run relationship between daily per capita calorie intake, per capita income and food prices for Turkey using the bounds testing approach to cointegration.

The findings of this article are in line with the previous empirical evidences for other developing countries. The results also support the Engle curve hypothesis.

The results demonstrate the existence of a statistically significant long-run relationship between calorie intake and income, indicating that a 1% rise in real per capita income increases the daily per capita calorie intake by 0.22%. Therefore, the impact of income increases on calorie intake is limited. The impact of food prices on the daily calorie intake, however, is insignificant. The augmented Granger causality and variance decomposition analysis confirm the unidirectional relationship from income to calorie intake. This study finds no reverse causality or feedback. Causality from income to calorie intake supports the conventional wisdom that income growth can alleviate and eventually eliminate inadequate calorie intake. The stability of the estimated parameters in the calorie demand equation allows the decision makers to design long-term policy frameworks.

The policy recommendation of this research is limited due to fact that it uses aggregate data and it is inevitable that institutional and structural details causing changes in calorie demand are probably lost. However, the Ministry of Health and those institutions concerned with the issues of public health should prepare more detailed dietary guidelines to curb the diseases eminating from malnutrition. These policies should be developed to target the different segments of the public so that efficiency of these measurements could be maximized. To this extent, the public should be encouraged to consume less saturated fat, reduced cholesterol intake and decreased intake of refined sugars. Educational and fiscal policy tools, such as the teaching of comprehensive healthy diets at schools, using the media to draw attention to the consequences of malnutrition, designing selective taxation policy for the foods with mal-nutritious contents and incentives for producing healthy foods, should lead to the calorie demand with healthy nutrition in the long-run.

## Appendix

*Structural Break Tests*

Zivot and Andrews (1992) proposed a structrual break test which allows endogenously determined breakpoints in the intercept, trend function, or in both. This test requires running the following regression for all potential break points, $T_B$, ($1<T_B<T$):

$$\Delta y_t = \mu + \beta t + \theta_1 DU_t + \gamma_t DT_t + \alpha y_{t-1} + \sum_{i=1}^{k} \phi_i \Delta y_{t-i} + \varepsilon_t \qquad (5)$$

where $DU_t$ and $DT_t$ are break dummy-variables that are define as follows:

$$DU_t = \begin{cases} 1 \;\; if \;\; t > T_B \\ 0 \;\; otherwise \end{cases} \qquad \text{and} \qquad DT_t = \begin{cases} t - T_B \quad if \; t > T_B \\ 0 \quad otherwise \end{cases}$$

where $k$ is the number of lags determined for each possible breakpoint by one of the information criteria such as AIC. Equation (5) is sequentially estimated and $T_B$ is chosen in order to minimize the one-sided $t$-statistics of the hypothesis $\alpha$=0. Thus, the break point is the point least favorable to the null hypothesis of unit root process with a drift and excludes any structural points.

*Data definition and sources*

Data are collected from three different sources, namely; Food Balance Sheets of Food and Agriculture Organization of the United Nations (FAO), International Financial Statistics of the International Monetary Fund (IMF) and Main Economic Indicators of OECD.

$c$ : is the daily per capita (kilo) calorie availability in logarithm. Source: FAO.
$y$ : is the real per capita gross domestic product (GDP) in logarithm. The GDP is divided by population and is also deflated by the consumer price index (CPI). Source: IMF.
$p$ : is the real food price index in logarithm. It is deflated by the CPI. Source: food prices comes from OECD and CPI comes from IMF.

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