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An Examination of Calorie Demand Relationship in Pakistan

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Abstract: This paper has examined the long-run relationship between daily per capita calorie intake, per capita income and food prices for Pakistan using aggregate data 1960-2001. Cointegration analysis yields the income-calorie elasticity of 0.21, while food-price elasticity is insignificant. Thus, economic growth, as measured by increasing per capita income, has significantly improved calorie intake in Pakistan; future economic growth can alleviate further inadequate calorie intake. However, significant improvements in calorie intake cannot be made directly by food subsidies. Nevertheless, the policies that lower food prices also have the indirect effect of increasing real incomes via the income effect. So it is in this sense that food-subsidy policies may have a role in improving calorie intakes. In the context of access to food, it would be important to identify the food insecure people, which are financially poor and are unable to acquire sufficient food, even if the overall supply of food in the country is sufficient. Further, causality tests indicate a bidirectional relationship from income to calorie intake; and from calorie intake to income.

Key words: Per capita calorie, cointegration, Granger-causality, Pakistan

Introduction

In Pakistan per capita calories intake has grown from 1940 (kilo) calories per day in 1960 to 2457 in 2001 with an average annual growth rate of 0.60 per cent (Govt. of Pakistan, 2003). Notwithstanding this increase, food security remains an unfulfilled dream for currently about 42 million people (UN, 2001). The fact that about one third of the population does not have access to food needed for adequate nutrition is manifested by the incidence of malnutrition. Among the 174 nations covered by the latest survey using the criteria of Human Development Index (HDI), Pakistan was ranked as number 135 (UN, 2001). Access to food is mainly related to per capita income, and for the last four decades Pakistan has been trying to increase the per capita income. Food poverty (calorie based) incidence showed that about one-third of the households are living below the food poverty line (consuming calories below the recommended level) and they are not meeting their nutritional requirements. The incidence of food poverty is higher in rural areas (35%), than in urban areas (26%) (UN, 2001).

The relationship between calorie intake and income therefore is crucial and much of the literature on malnutrition has focused on this relationship but estimated calorie-income elasticity vary considerably (Bouis, 1994, summarizes this literature). There are two lines of inquiry: first, does calorie intake rise with income, and second, is income generation affected by calorie intake? The former focuses on the estimation of calorie-demand relationship, while the latter is at the centre of the efficiency wage hypothesis. It is therefore clear that causality in the calorie-income relationship can run in either or both directions. It is evident from the literature on the efficiency wage hypothesis where there is concern about the endogeneity of income (see for example, Strauss 1986; Sahn and Alderman, 1988; Haddad and Bouis, 1991 and Behrman *et al.*, 1997). Recently, Dawson and Tiffin (1998); Dawson (2002) examined the long-run calorie-income relationship for India and Pakistan and the estimated elasticities are 0.34 and 0.19 respectively. However, notwithstanding the fact that food subsidies are common in LDCs and the estimates with respect to food prices in general, tend to be significant.

Our aim here is to re-examine the long-run relationship between per capita calorie intake, per capita income, and food prices using aggregate time series data and cointegration analysis for Pakistan and to test for the direction of causality between calories and income. The remainder of the paper is organized as follows: Section 2 discusses our empirical methodology, Section 3 discusses the data and results, and Section 4 summarizes and concludes.

Materials and Methods

Many time series are non-stationary and in general OLS regressions between non-stationary data are spurious. The presence of unit roots in the autoregressive representation of a time series leads to non-stationarily, and such series, referred to as being integrated of order one [I (1)], must be first-difference to render them stationary (or integrated of order zero). Where [I (1)] series move together and their linear combination is stationary, the series are cointegrated and the problem of spurious regression does not arise. Cointegration implies the existence of a meaningful long-run equilibrium (Granger, 1998). Since a cointegrating relationship cannot exist between two variables which are integrated of a different order, we first test for the order of integration of the variables.

In testing for the presence of unit roots in the individual time series using the augmented Dickey-Fuller (ADF) test (Dickey and Fuller, 1981; Said and Dickey, 1984), both with and without a deterministic trend, we follow the sequential procedure of Dickey and Pantula (1987): the null of the largest plausible number of unit roots, assumed to be three, is tested and, if rejected, that of two unit roots is tested and so on until the null is not rejected. The number of lags in the ADF-equation is chosen to ensure that serial correlation is absent using the Breusch-Godfrey statistic (Greene, 2000, p.541). If they are integrated of the same order, Johansen's (1988) procedure can then be used to test for the presence of a cointegrating vector between calorie intake, income and food prices. The procedure is based on maximum likelihood estimation of the error correction model:

$$\triangle z_t = \delta + \Gamma_1 \triangle z_{t-1} + \Gamma_2 \triangle z_{t-2} + \ldots.. \Gamma_{p-1} \triangle z_{t-p+1} + \pi z_{t-p} + \mu_t \qquad (1)$$

Where $Z_t = [C_t, Y_t, P_t]$, C_t is per capita calorie intake, Y_t is income, and P_t is the price of food, $\alpha z_t = z_t - z_{t-1}$ and π and Γ_i are (nxn) matrices of parameters with $\Gamma_i = -(I - A_1 - A_2 --- ... A_i)$, $(I = 1, \ldots k - 1)$, and $\pi = I - \pi_1 - \pi_2 --- \pi_k$. The term πz_{t-p} provides information about the long-run equilibrium relationship between the variables in z_t. Information about the number of cointegrating relationships among the variables in z_t is given by the rank of the π-matrix: if π is of reduced rank, the model is subject to a unit root; and if $0<r<n$, where r is the rank of π, π can be decomposed into two (n x r) matrices α and β, such that $\pi = \alpha\beta'$ where $\beta' z_t$ is stationary. Here, α is the error correction term and measures the speed of adjustment in Δzt and β contains r distinct cointegrating vectors, that is the cointegrating relationships between the non-stationary variables.

The Johansen procedure estimates (1), and trace statistics are used to test the null hypothesis of at most r cointegrating vectors against the alternative that it is greater than r. If cointegration is established, then Engle and Granger (1987) error correction specification can be used to test for Granger causality. If the series Ct and Yt are both I (1) and cointegrated, then the ECM model is represented by the following equations.

$$\Delta C = \alpha_0 + \sum_{I=1}^{n} \beta_I \Delta C_{t-I} + \sum_{I=1}^{n} \beta_J \Delta Y_{t-I} + \delta ECT_{t-I} + \mu_t \qquad (2)$$

$$\Delta Y = \Phi_0 + \sum_{I=1}^{n} \sigma_I \Delta Y_{t-I} + \sum_{I=1}^{n} \sigma_J \Delta C_{t-I} \lambda ECT_{t-I} + \varepsilon_t \qquad (3)$$

Where Δ is the difference operator μ_t and ϵ_t are the white noise error terms, ECT_{t-I} is the error correction term derived from the long-run cointegrating relationship., while *n* is the optimal lag length orders of the variables which are determined by using the general-to-specific modeling procedure (Hendry and Ericson, 1991). Our null hypotheses are as follows. *Y* will Granger cause C if $\beta_j \neq 0$. Similarly, C will Granger cause Y if $\sigma_i \neq 0$. There will be bidirectional causality if $\beta_j \neq 0$ and $\sigma_i \neq 0$. To implement the Grnger - causality test, F-statistics are calculated under the null hypothesis that in Equ (2) and (3) all the coefficients of β_j, $\sigma_i = 0$.

Results and Discussion

The annual time series data relate to Pakistan for 1960-01. Calorie intake is average per capita energy (calorie) intake per day derived from national food balance sheets (FAO, 2000; GOP, 2002-03). It has trended erratically upward and has grown from 1940 (kilo) calories in 1960 to 2457 in 2001 with an average annual growth rate of 0.60 per cent. Real per capita gross domestic product (GDP) (2000 rupees) is calculated as (nominal per capita GDP) × (100/CPI) (GOP, 2002-03). Real per capita GDP has trended upward at a relatively constant average annual growth rate of 2.44 per cent, ranging between 8628 rupees in 1960 and 23483 rupees in 2001. The real food price index (2000 rupees) is the corresponding nominal index deflated by the consumer price index.

Augmented Dickey-Fuller (ADF) tests are used to test for unit roots in the series in logarithms. Lags are added so that the Breusch-Godfrey LM-statistic (Greene, 2000, p.541) rejects serial correlation up to fourth order. Table 1 presents the results of the ADF-tests performed with and without a linear trend. In the non-trended model, unit roots appear in C_t, Y_t and P_t. However, in the trended model, Ct series appear to be a stationary series. Trend is also significant, but by using Ø3-test we cannot reject the null of the unit root and of no trend. Therefore, we conclude on balance that all series are I (1), that is, they are stationary after first difference. Therefore cointegrating relationship between them is now sought.

We now test for cointegration in our model following Johansen (1988). First, to determine the order of the vector autoregressive (VAR), we carried out adjusted LR-tests on the VAR with a maximum of four lags. The results indicate that the LR-test statistics rejected order zero (0.000), but do not reject the VAR with order one. So we choose one as the order of VAR for our model.

The second step in the Johansen procedure is to test for the presence and number of cointegrating vectors among the series in the model. The results are presented in Table 2, which imply that trace statistics strongly rejects the null hypothesis that there is no cointegration between Ct, Yt, and Pt (i.e. r = 0), but do not reject the hypothesis that there is one cointegrating relationship (i.e r = 1). It is therefore concluded that our model has one cointegrating vector (i.e., a unique long-run equilibrium relationship exists).

Table 1: DF test for Unit Root Results

Variables	Test statistics for Non-trended Model	Test statistics for Trended Model	Trend	Ø3
C_t	- 2.22	- 3.98	2.87	5.75
Y_t	- 0.79	- 2.67	2.46	3.89
P_t	- 0.57	- 1.90	-1.92	2.00
CV*	- 2.93	- 3.60	2.85	6.73

*95% of confidence level.

Table 2: Co integration Results --- Trace Statistics

Equation Tested	H0	H1	Statistics
	r=0	$r \geq 1$	52.22 (34.87)
	$r \leq 1$	$r \geq 2$	14.36 (20.18)
C_t Y_t P_t	$r \leq 2$	$r \geq 3$	4.45 (9.16)

Critical values (95% confidence level) in parentheses.

Table 3: Johansen Normalized Estimates

Variables	Normalized Estimates
C_t	-1.00
Y_t	0.21(2.1)
P_t	0.01(0.17)

Note: t-statistics in parentheses

Table 4: Granger-causality test (using error correction approach)

Causality	Lags	F-Statistics	P-Value
LY $\Rightarrow$ LC	1,1	4.56	0.01
LC $\Rightarrow$ LY	1,1	2.27	0.10

Table 3 shows the Johansen normalized estimates. Because all variables are defined in logarithms, the estimated coefficients in each cointegrating vector associated with Y_t and P_t, when normalized on C_t, are the income and food-price long-run elasticities of demand for calorie intake. In our results, the income elasticity of calorie demand is 0.21 which implies that a 1% increase in per capita income increases per capita calorie demand by 0.21%. The results are significant the only reason that coefficient is small is that as income increases, individuals may diversify their diets from a taste perspective as they substitute more expensive sources of calories for less expensive ones. Further substitution may occur by consuming complements to good nutrition, such as clean water, good sanitation or women's time in child care. However, we also find that coefficient of food prices bears a positive sign but has a insignificant effect on calorie intake. It is possible that limited variability in food prices, combined with our small sample, has prevented us from identifying a significant impact.

Since the series C_t and Y_t are found to be cointegrated, we test for the direction of Granger causality following Granger (1969) and results are presented in Table 4. Results show that Y_t Granger-causes Ct at 1% level of significance, and C_t Granger-causes Y $_t$at 10% level of significance. Causality tests reveal the existence of bidirectional causality between the series i.e., from income to calorie intake; and from calorie intake to income. It is clear that by increasing the per capita income and improving the per capita calorie intake we can make better the food security situation.

Conclusion: This study has examined the long-run relationship between daily per capita calorie intake, per capita income and food prices for Pakistan using cointegration analysis. Using annual data for 1961-2001, there is evidence that such a relationship exists, and that a 1% increase in real per capita income raises the daily per capita calorie intake by 0.21%. The results were found statistically significant. However, the coefficient of food prices bears the positive sign but is statistically insignificant. Further, causality tests indicate a bidirectional relationship from income to calorie intake; and from calorie intake to income.

The implications of our results for development policies that seek to alleviate inadequate calorie intake are clear. First, in the long run the estimate of the calorie-income elasticity supports the conventional wisdom that income growth can alleviate inadequate calorie intake. Because as income increases, individuals may diversify their diets from a taste perspective as they substitute more expensive sources of calories for less expensive ones. Further substitution may occur by consuming complements to good nutrition, such as clean water, good sanitation or women's time in child care. Second, we also find that food prices have an insignificant effect on calorie intake. However, it is possible that limited variability in food prices, combined, with our small sample, has prevented us from identifying a significant impact. Given this caveat, the implication of our result is that policies that lower food prices also have the indirect effect of increasing real incomes via the income effect. So it is in this sense that food-subsidy policies may have a role in improving calorie intakes. In the context of access to food, it would be important to identify the food insecure people, which are financially poor and are unable to acquire sufficient food, even if the overall supply of food in the country is sufficient.

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