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Nutrition in Pakistan: Estimating the Economic Demand for Calories

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Abstract: In the last four decades, per capita calorie intake in Pakistan has grown from 1750-2450 (kilo)calories with an average annual growth rate of 0.90%. Nevertheless, 20% of Pakistan's population is still undernourished.This paper has examined the long-run relationship between daily per capita calorie intake and per capita income for Pakistan using cointegration analysis. Using annual data for 1961-1998, there is strong evidence taht such a relationship exists, and that a 1% increase in real per capita income raises the daily per capita calorie intake by 0.19 per cent. Further, causality tests indicate a unidirectional relationship from income to calorie intake; we find no evidence of causality in the opposite direction. This result substantiates Engel's law and provides no support for the hypothesis that income generation is constrained by calorie intake. There are two caveats to the results. First, data limitations restrict the number of observations to 38. Whilst it is not uncommon to find such small samples used in analyses of this type, some caution is necessary in interpreting the results as a consequence of the low power exhibited in some of the tests employed. Second, an aggregation problem arises from the use of national data since we are implicitly adding-up across non-linear relationships at the micro, household level; further, distributional changes in income have not been accounted for in the model. The implications of our results for development policies which seek to alleviate inadequate calorie intake in Pakistan are clear. First, the estimate of the calorie-income elasticity albeit low supports the conventional wisdom that income growth can alleviate inadequate calorie intake. However, nutritional status, measured in terms of nutrient deficiency, may not improve: 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.

Key words: Calories, per capita calorie consumption, nutrition in Pakistan.

Introduction

Substantial increases in per capita calorie consumption in Pakistan have taken place in the last 40 years, rising from 1753 (kilo)calories in 1961 to 2447 in 1998. (FAO, 2000). Notwithstanding this increase, 20% of the total population is still undernourished (United Nations, 2001). During the 1970s, such undernutrition was generally thought to reflect a lack of protein but by 1980, there was broad agreement on its cause: "Serious and extensive nutritional deficiencies occur in virtually all developing countries, though they are worst in low-income countries. They are usually caused by undernourishment - a shortage of food ... Malnutrition is largely a reflection of poverty: people do not have enough income for food. Given the slow income growth that is likely for the poorest people in the foreseeable future, large numbers will remain malnourished for decades to come. ... The most effective long-term policies are those that raise the incomes of the poor..." (World Development Report 1980, p.59).

Engel's law is derived from the relationship between the demand for calories and income: the proportion of income spent on food diminishes as income increases. This law is summarised in the income elasticity of calorie demand, ηM, which is the percentage change in calorie demand brought about by a 1% increase in income. It is expected that $0<\eta M<1$ and conventional wisdom (World Bank, 1986, pp.v and 10) is that calorie-income elasticities, while not equal to unity, are assumed to be substantially greater than zero. Much of the empirical literature on nutrition in LDCs has focused the calorie-income relationship but estimated calorie-income elasticities vary considerably. For example, Bouis and Haddad (1992) for the Philippines, and Ravallion (1990) for Indonesia provide estimates that are either close to, and/or insignificantly different from zero, whereas Behrman and Deolalilar (1987) for India and Strauss (1984) for Sierra Leone produce estimates of around 0.82 (Bouis, 1994, summarises this literature).

An alternative approach concerns the 'efficiency wages hypothesis' where income generation is affected by calorie intake. Again results are conflicting: for example, Strauss (1986) finds a significant relationship between farm productivity and calorie intake in Sierra Leone, while Deolaliker (1988) for India finds no evidence that nutrition determines wages (Bliss and Stern, 1978, survey this literature).

It is clear that causality in the calorie-income relationship can run in either (or both) direction(s). Recently, Dawson and Tiffin (1998) examine the long-run calorie-income relationship. Using annual data for India, results show that calorie intake is caused by income and the calorie-income elasticity is 0.34.

The focus here is to estimate the long-run relationship between per capita calorie intake and per capita income using aggregate annual data for Pakistan and to test for the direction of causality between calories and income. The remainder of the paper is organised as follows: Section 2 examines the empirical methodology, Section 3 discusses the data and results, and Section 4 summarises and concludes.

Materials and Methods

The method adopted to investigate the calorie-income relationship using time series data does not require the specification of causality prior to estimation. The vector autoregressive (VAR) model is the basis for this analysis and is expressed as:

$$\begin{bmatrix} C_t \\ M_t \end{bmatrix} = \begin{bmatrix} \mu_1 \\ \mu_2 \end{bmatrix} + \sum_{i=1}^{k} \begin{bmatrix} a_{k,11} & a_{k,21} \\ a_{k,21} & a_{k,22} \end{bmatrix} \begin{bmatrix} C_{t-i} \\ M_{t-i} \end{bmatrix} + \begin{bmatrix} \epsilon_{1t} \\ \epsilon_{2t} \end{bmatrix} \quad (1)$$

where C_t is calories, M_t is income, k is the lag length, μ and A_i are matrices of parameters to be estimated, and ϵ_t are error

Table 1: Unit root tests

Variable	ADF-Test Non-Trended Model	Trended Model	KPSS-Test Non-Trended	Trended Model Model
C_t	-2.26	-4.11	2.57	0.27
M_t	-1.66	-1.36	3.75	0.23
Critical value*	-2.93	-3.50	0.46	0.15

Note * 95% confidence level

Table 2: Trace Statistics

H_o (H_1)	Model 1	Model 2
r=0(r=1)	50.85(15.41)	20.56(15.41)
r≤(r=2)	11.40(3.76)	1.98(3.76)

Note:95 % confidence level in parentheses (Osterwald-Lenum, 1992)

Table 3: Normalised cointegrating vectors

Variable	β	α	β	α
C_t	-1.00	-0.57	-1.00	0.17
M_t	0.17	-0.13	0.19	0
		1.51		

Note: t-statistics in parentheses

terms.

As Harris (1995, pp.14-25) explains, many economic time series are non-stationary and that in general ordinary least squares regressions between non-stationary data are spurious. The presence of unit roots in a time series leads to non-stationarity. In such cases, each series must be first-differenced to make it stationary and it is referred to as integrated, or I(1). Therefore, the first step is to test for the presence of unit roots using the augmented Dickey-Fuller (ADF) test (Dickey and Fuller, 1981, and Said and Dickey, (1984): the null hypothesis is of a unit root while the alternative is stationarity. Often to substantiate conclusions from the ADF-tests, the KPSS-test (Kwiatkowski *et al.*, 1992) is used where the null hypothesis is of stationarity while the alternative is a unit root.

Where integrated series move together and their linear combination is stationary, the series are cointegrated and the problem of spurious regression is absent. Cointegration implies the existence of a meaningful long-run equilibrium (Granger, 1988); the Johansen (1988) procedure tests for cointegration and is based on the estimation of the VAR model in (1) transformed into its vector error correction model (VECM) form:

$$\Delta x_t = \mu + \sum_{i=1}^{k-1} \Gamma_i x_{t-i} + \pi x_{t-k} + \varepsilon_t \quad (2)$$

where $x_t = [C_t, M_t]'$, $\Delta x_t = x_t - x_{t-1}$, μ and Γ_i ($\Gamma_i = -[I - A_1 - ... - A_i]$ for $i = 1,...,k-1$) are (2×1) and (2×2) matrices of parameters respectively, π ($\pi = -[I - A_1 - ... - A_k]$) is a (2×2) matrix of parameters and ε_t is a (2×1) vector of white noise errors. When the model is subject to unit roots, π is of reduced rank (r) and when $0 < r < 2$, π can be decomposed into $\pi = \alpha\beta'$. Where $r = 1$, (2) can be rewritten in full as:

$$\begin{bmatrix}\Delta C_t\\ \Delta M_t\end{bmatrix} = \begin{bmatrix}\mu_1\\ \mu_2\end{bmatrix} + \sum_{i=1}^{k-1}\begin{bmatrix}\Gamma_{i,11} & \Gamma_{i,12}\\ \Gamma_{i,21} & \Gamma_{i,22}\end{bmatrix}\begin{bmatrix}\Delta C_{t-i}\\ \Delta M_{t-i}\end{bmatrix} + \begin{bmatrix}\alpha_1\\ \alpha_2\end{bmatrix}\begin{bmatrix}\beta_1 & \beta_2\end{bmatrix}\begin{bmatrix}C_{t-k}\\ M_{t-k}\end{bmatrix} + \begin{bmatrix}\epsilon_{1t}\\ \epsilon_{2t}\end{bmatrix} \quad (3)$$

The Granger representation theorem (Engle and Granger, 1987) shows that $\beta' x_t$ is stationary implying that x_t is cointegrated with r distinct cointegrating vectors given by the columns of β. Johansen's (1988) procedure estimates (3); trace statistics are used to determine the rank of π which can then be decomposed to give the cointegrating vector, β.

Two possible models are admitted in (4). Model 1 is where there are no linear trends in the levels of the variables and the first-differenced series have a zero mean; here the intercept is restricted to the cointegration space. Model 2 is where there are linear trends in the levels of the variables and there is an intercept in the short-run model only. To test between these models, the Pantula principle (Harris, 1995, p.97) is used to test the joint hypothesis of both rank and the deterministic components.

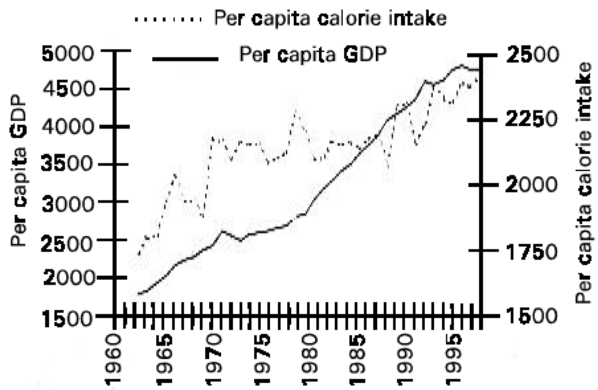

Fig. 1: Per capita GDP and calorie intake in Pakistan (1961-1968

Equation (3) can be used to test the direction of causality (Granger, 1969) between calories and income. Following Hall and Milne's (1994) definition of (long-run) causality, C_t does not cause M_t if $\alpha_2 = 0$ in (3). Similarly, M_t does not cause C_t if $\alpha_1 = 0$ in (3). Bidirectional causality between M_t and C_t is also possible and is implied by $\alpha_1 \neq 0$ and $\alpha_2 \neq 0$.

Results and Discussion

The annual data relate to Pakistan for 1961-98 and are shown in Figure 1. Calorie intake is average per capita energy (calorie) intake per day, calculated on the basis of per capita dietary energy supply derived from national food balance sheets (source: FAO, 2000). It has trended erratically upwards and has varied between 1753 (kilo)calories in 1961 and 2447 in 1998 with an average annual growth rate of 0.90%. Real per capita GDP (in 1981 constant rupees) is real GDP/population (sources: World Bank, 2001). Real per capita GDP has trended upwards at a relatively constant average annual growth rate of 2.68%, varying between 1823 rupees in 1961 and 4970 rupees in 1996.

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 both C_t and M_t; this conclusion is substantiated in the trended model for M_t but not for C_t. The KPSS-tests indicate that both C_t and M_t have unit roots irrespective of whether or not the model contains a trend. Therefore, we conclude on balance that C_t and M_t are I(1), that is, they are stationary after first-differencing. Therefore a cointegrating relationship between them is now sought.

The first step of the Johansen procedure is to select the order of the VAR. The LR-statistic, adjusted for small samples (Sims, 1980), is used to test the null hypothesis that the order of the VAR is k against the alternative that it is four where k=0,1,...,4, and results show that k=1. The Johansen procedure and trace statistics are used to test between Models 1 and 2 and to test for the presence of a cointegrating vector using the Pantula principle. From the results in Table 2, Model 2 is chosen with one cointegrating vector (r = 1).

Using the preferred Model 2, the cointegrating vector for the unrestricted model, normalised on C_t, is shown in Table 3 and implies that $C_t = 1.17GDP_t$. To test the direction of causality between C_t and M_t, the significance of the α-coefficients is tested. Testing $\alpha_1 = 0$, $\chi^2 = 2.01$ (p-value = 0.16) and the null is not rejected; testing $\alpha_2 = 0$, $\chi^2 = 13.70$ (p-value = 0.00) and the null is rejected. Both results are supported by the t-statistics shown in Table 3. Thus, the direction of causality is unidirectional from M_t to C_t, that is, changes in income lead to changes in calorie intake and there is no support for the hypothesis that income generation is determined by calorie intake. Also shown in Table 3 is the normalised cointegrating vector under the restriction that $\alpha_2 = 0$. The implied relationship is: $C_t = 0.19\ M_t$; the income elasticity of calorie demand is 0.19 which implies that a 1% increase in per capita income increases per capita calorie demand by 0.19%; Engel's law is valid in this case.

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