Consumer Behaviour in India

An Application of the Linear Expenditure System

N Bhattacharya

In two articles in 'The Economic Weekly' Ashok Rudra (November 7, 1964) and Ashok Rudra and Pushpam Paul (November 28, 1964) applied the simplest form of linear expenditure system to some time-series data on domestic consumption in India. Income elasticities obtained thus were very different from the usually accepted cross-section elasticities from the National Sample Survey (NSS); and the elasticities of necessaries (like foodgrains) and of luxuries were much closer to unity.

The fitted model was utilised by Ashok Rudra in a very important discussion of the consequences of the proposed relative rates of growth of agriculture and industry.

An attempt is made here to re-examine the Paul-Rudra estimates of elasticities. It is found that owing to limitations of data the estimates have very wide margins of error. Further, the most plausible "point" estimates are closer to the cross-section elasticities than Rudra and Paul seemed to suggest.

IN THE present state of time-series statistics on consumption and prices in India, analyses would not give firm estimates of elasticities. But so far as the income elasticities are concerned, there is considerable scope for improving the cross-section estimates (till now estimates by extremely naive methods) to make them more nearly applicable for projection of consumer demand over time.

The Model

Many extensions of the linear expenditure system are presented by Stone (1964); but for various reasons, such as the short time-span covered in our analysis, the slow change in level or pattern of consumption in India (see reference 6, pp 4-9), and the comparative insignificance of durable goods in India, the simplest model introduced by Stone (1954) seems to be the only one suited to this analysis. This can be written as

(1) $p_{it} q_{it} = p_{it} \bar{q}_i + b_i (C_t - \Sigma_j p_{jt} \bar{q}_j)$

$i = 1, 2, \ldots, n;\ t = 1, 2 \ldots T$

Here, P_{it} is the price of the i-th commodity and q_{it} is the corresponding quantity consumed by the average consumer in period t, $\bar{q}_i$ the "committed" quantity which the average consumer tends to purchase irrespective of prices, $C_t = \Sigma_j p_{jt} q_{jt}$ the total consumer expenditure on all commodities, and b_i the fraction of the non-committed amount, $C_t - \Sigma_j p_{jt} \bar{q}_j$, allotted to commodity i.

The model shows expenditure on each commodity as a homogeneous function of C_t and all the prices. In order that expected commodity-wise expenditures add up to C_t identically, it is necessary and sufficient that $\Sigma b_i = 1$. The model satisfies a third criterion arising out of pure theory, viz, the Slutzky condition of symmetry of the substitution effect. In fact, Stone (1954) started from the general linear homogenous model.

(2) $p_{it} q_{it} = b_i C_t + \Sigma_j q_{jt} p_{jt}$

$i = 1, 2, \ldots n;\ t = 1, 2 \ldots T$

Where the adding-up criterion is satisfied if $\Sigma b_i = 1$ and $\Sigma_i a_{ij} = 0$ for all j. The imposition of the Slutzky condition throws the a_{ij} into the form implied in (1) and considerably reduces the number of parameters at the cost of interlocking the equations for different commodities and making estimation considerably difficult.[1,2]

Since real expenditure rises very slowly over the period, linearity in C_t is not a serious assumption but linearity in prices is, since relative prices do alter appreciably. A serious difficulty is that the model cannot handle inferior goods or complementaries, which must be remembered when using a fine classification of commodities. But the real limitation of the model in the present analysis is that all the price-quantity demand curves are necessarily inelastic, which may be unrealistic for commodity-groups like clothing, milk and milk products introduced in our investigation.

A two-stage iterative method of estimation was employed by Stone (1954). It minimises the unweighted sum of squares

$$S(b, \bar{q}) = \Sigma_i \Sigma_t \left\{ p_{it} q_{it} - p_{it} \bar{q}_i - b_i (C_t - \Sigma_j p_{jt} \bar{q}_j) \right\}^2$$

over variations in b and $\bar{q}$ where b and $\bar{q}$ are the vectors $(b_1, b_2 \ldots b_n)$ and $(\bar{q}_1, \bar{q}_2 \ldots \bar{q}_n)$ respectively. Stone recognised the arbitrariness of this criterion. Residuals for larger commodity groups may receive undue weightage like those for periods with higher price levels.[3] Also, no account is taken of the correlations between residuals for different commodities, nor of autocorrelations of residuals over time.

Nevertheless, the idea is to start from an initial vector $b^0 = (b_1^0, b_2^0, \ldots b_n^0)$ to estimate $\bar{q}^1 = \bar{q}_1^1, \bar{q}_2^1, \ldots \bar{q}_n^1)$ by minimising $S(b^0, q)$ for variations in $\bar{q}$, next to estimate b^1 by minimising $S(b, \bar{q}^1)$ for variations in b; and so on. At each step, the ordinary least squares method is applicable. We show below that the process converges under very mild conditions. Although it seems plausible, there is no proof that the solution reached is a relative minimum when all arguments of $S(b, \bar{q})$ are allowed to vary simultaneously, instead of the b_i's alone or the $\bar{q}_i$'s alone. What is more important, it is uncertain whether, in general, the solution gives an absolute minimum independent of the initial b^0. In the present case, we satisfied ourselves on these points by carrying out some trial calculations starting from different initial values.

That the Stone method of estimation converges can be proved thus, following Stone, Aitchison and Brown (1955): Let $\bar{q}^r$ be the estimate of $\bar{q}$ starting from b^{r-1} and b^r the estimate of b starting from $\bar{q}^r$. (Incidentally, $\Sigma b_i{}^r = 1$ necessarily.) When minimising $S(b, \bar{q}^r)$ for variations in b, b^{r-1} is available among the choices. Hence
$S(b^r, \bar{q}^r) \leqslant S(b^{r-1}, \bar{q}^r)$

Similarly, $S(b^r, \bar{q}^{r+1}) \leqslant S(b^r, \bar{q}^r)$

Hence the values of S(b, q) obtained at each step form a monotone non-increasing sequence. Since the sequence is bounded below at zero, the sequence converges to a limit and from this it can be shown that the sequence b^r converges.

Given b^r, the estimate $\bar{q}^{r+1}$ is only a rational function of the observations and b^r. Hence if b^r converges, $\bar{q}^{r+1}$ also converges under very general conditions. It can be shown that in the estimation of $\bar{q}^{r+1}$ multicollinearity does not arise if all the price series do not move in the same proportions over time. Needless to add, the whole approach assumes the constancy of demand functions, so that the identification problem does not arise.

The Material

Rudra (1964) and Paul and Rudra (1964) examined two sets of data relating to rural and urban India, one distinguishing between three commodity-groups and the other, between six. From the use made in the earlier article, it is clear that the three-commodity case was much more important. The values of p_{it} q_{it} were obtained from eight rounds of the NSS, numbered 8th to 15th (covering the period July 1954 to June 1960) as weighted averages of the corresponding rural and urban estimates. The p_{it}'s were calculated by Paul and Rudra from official wholesale price series, by averaging commodity-specific price relatives using weights based on NSS family budgets.

The material analysed here includes the Paul-Rudra data, but many other sets having the same lay-out were analysed. Separate calculations were done for rural India, urban India and all-India, and in each case, for the half-samplewise and combined estimates of p_{it} q_{it} from the NSS. Unfortunately, the same wholesale price indices had to be used in all these cases. However, much greater emphasis was given to the three-commodity case, and here, retail price indices of cereals obtained from unpublished NSS sources were sometimes used for the foodgrains group, along with wholesale price indices for the other two groups. These retail price indices are available separately for rural and urban areas, but they are somewhat provisional, particularly those for the 10th round.

Considerations of space prevent us from presenting the basic material, viz (i) the p_{it} q_{it}'s for rural and urban India obtained from published or draft NSS reports, (ii) the approximate weightages of the rural and the urban sectors used for obtaining rural-plus-urban estimates, and (iii) the wholesale/retail price indices employed as described earlier. The Paul-Rudra figures for p_{it} q_{it} have since been revised for the 15th round; also they were slightly inaccurate in obtaining the figures for the 11th round. Most of our calculations are based on the revised data, but the Paul-Rudra data were sometimes used to see the effect of these differences in data.

It may be stated here that the wholesale and retail indices for foodgrains do not agree very well, and this will have important consequences in the sequel. It may also be stated that halfsample divergences are on the whole larger for the urban sector than for the rural sector. This must be partly due to the frequently smaller sample sizes in urban areas. Figures for the 8th round (urban) are particularly erratic. However, figures for foodgrains are consistent even for the urban sector.

For the rural-plus-urban results for three commodities it was found that the iterations converge or tend to converge to the same point whatever the initial vector b^0 in the plausible region. We did not continue the iteration sufficiently long to actually reach the same point starting from each chosen initial vector, but the diagrammatic indications seem to be fairly clear. It is also apparent that the residual sum of squares tends to reach an absolute minimum. It may be stated that these tendencies are absolutely clear for halfsample 2, where the convergence is quite rapid; for half sample 1 the convergence is very slow and the tendencies are not quite clear; the position is intermediate for the 'combined' data.

In this case of three commodities, it might have been more convenient to have a programme for only estimating $\bar{q}^r$ from b^r, and to map out the residual sum of squares as a function of b_1 and b_2 by trying a large number of (b_1, b_2)-points. This would enable one to approach the absolute minimum by trial and error methods.

The best point estimates of the parameters are shown in Table 1.

Table 1: Rural-plus-Urban Estimates: Three Commodity-Groups

			b_1	b_2	b_3	$\bar{q}_1$ (Rs)	$\bar{q}_2$ (Rs)	$\bar{q}_3$ (Rs)
Wholesale prices for all commodity-groups	half-sample	1	0.19	0.45	0.36	4.5	-0.6	1.5
		2	0.24	0.30	0.46	6.5	4.4	4.5
	combined		0.22	0.36	0.42	5.4	2.8	2.6
NSS price for foodgrains only	half-sample	1	0.14	0.47	0.39	6.3	-1.4	0.7
		2	0.16	0.35	0.49	7.3	4.6	5.0
Probable range	combined		0.14	0.40	0.46	6.3	1.9	2.2
			0.1-0.3	0.25-0.5	0.3-0.55	4.8	-3-6	0-5.5

It should be noted that $\bar{q}_1$, $\bar{q}_2$, $\bar{q}_3$ are quantities evaluated at base (i.e 1952-53) prices.

We must explain the figures in the last row which, though partly subjective, are very important. In arriving at these intervals we remembered the appreciable halfsample divergence (especially for commodity-groups 2 and 3) and the effect of the choice of price indices for foodgrains. We remembered yet another point. The Stone model does not allow the estimation of standard errors, but one can get very rough notions by observing the variation of the residual sum of squares around the optimum point in the $(b_1\ b_2)$ diagram. Surely, if one draws a contour map with the residual sum of squares 20 per cent above its minimum value, there would be little to choose between any two points inside that closed curve: the available data cannot possibly discriminate between two points within this region.

Regarding results for the individual rural and urban sectors for three-commodity groups, we did not experiment with many initial vectors since the results obtained with the plausible initial vectors formed a consistent pattern with the rural-plus-urban results. For the rural sector, the situation is similar to rural-plus-urban regarding speed of convergence, and regarding half-sample divergence or effect of substituting NSS price for wholesale price for foodgrains. The probable ranges of the parameters are as follows:

Rural Estimates

b_1	b_2	b_3
0.08-0.32	0.28-0.44	0.37-0.53
$\bar{q}_1$ (Rs)	$\bar{q}_2$ (Rs)	$\bar{q}_3$ (Rs)
4.5-7.5	—2.6-4	—3-4

The corresponding ranges for the urban parameters are:

Urban Estimates

b_1	b_2	b_3
0.04-0.19	0.20-038	0.48-0.73
$\bar{q}_1$ (Rs)	$\bar{q}_2$ (Rs)	$\bar{q}_3$ (Rs)
2.5-6	0-6.6	—4-5

The results for six-commodity groups are similar for rural and rural-plus-urban. For both these sectors, the results are generally similar to corresponding results for three-commodity groups in the sense that subtotals of b_i's and $\bar{q}_i$'s obtained in the former case are close to the b_i and $\bar{q}_i$ of the broader groups in the latter case. For the urban sector, however, the $\bar{q}_i$'s tend to be more often negative in the six-commodity case. The estimates for foodgrains as well as for fuel and light remain very stable.

Since computations in this case are far less extensive than for the three-commodity case, we shall not venture to give probable intervals for the true parameter values. The ranges indicated by halfsample estimates have been used for obtaining the elasticity ranges in Table 3, but they may not contain the true values. Thus, we have no idea of the possible effect of using some retail price series instead of the wholesale ones. Also, with only 100 cycles completed in all cases, the approach to the optimal solutions is usually less close.

Consider now the results for the Paul-Rudra data in Table 1. The estimates given in Rudra (1964) are as follows:

	b_1	b_2	b_3
"quantity" fit	0.355	0.302	0.343
"price" fit	0.364	0.297	0.339
	$\bar{q}_1$	$\bar{q}_2$	$\bar{q}_3$
"quantity" fit	6.00	4.34	4.98
"price" fit	5.85	4.23	4.86

It is necessary to explain the terms "quantity" fit and "price" fit. Both fits followed the Stone method in all respects, except that in the very beginning equation (1) was divided throughout by p_{it} for the "quantity" fit, and by q_{it} for the "price" fit. The former has the advantage of using a residual sum of squares where residuals for periods with higher price levels do not receive greater weightage. In our calculations we did not introduce this modification because, on the whole, the later rounds with higher price levels deserved greater weights in virtue of larger sample sizes. The price fit is not of direct interest here, and where fits are so close as here, both should give very similar estimates.

In one of our exercises we chose an initial $b^* = (0.35, 0.30, 0.35)$ approximating the Paul-Rudra estimates. The iteration led somewhere near $b = (0.25, 0.35, 0.40)$ with appreciable reduction in the sum of squares. We examined the goodness of fit, and our fit as judged by the residual sum of squares was equally superior to the Paul-Rudra "quantity" fit in terms of both quantities and expenditures: the difference in the residual sum of squares was roughly about 20 per cent. Presumably, Paul-Rudra stopped after a few cycles, having been misled by the agreement between results of successive cycles. But on closer examination we felt — we emphasise the subjective element since no probability model is available — that the Paul-Rudra fit did not look significantly inferior: it was clear that the Paul Rudra estimates cannot be ruled out. The correlation coefficients between observed and expected quantities are as shown in Table 2.

This gave us our rule of thumbs (used in an intuitive way) that the minimum residual sum of squares could easily go up by 20 per cent without there being serious increase in the values of the correlation coefficients.

In this and other cases we found that the residual sum of squares often changes very slowly in particular directions near the optimal solution. The Hessian in one case had

Table 2: Correlation Coefficients between Observed and Expected Quantities

	Commodities		
	Foodgrains	Food excluding Foodgrains	Non-Food
Paul-Rudra quantity fit	0.905	0.922	0.961
our fit	0.918	0.945	0.952

Table 3: Magnitude of Elasticities of Total Expenditure Estimated by the Linear Expenditure System for All-India

	3-Commodity Groups			6-Commodity Groups					
	Foodgrains	Food Excluding Foodgrains	Non-Food	Milk etc	Other Food	Clothing	Fuel & Light	Other Non-Food
Rural	0.2-0.8	0.95-1.6	1.15-1.7	1.7-2.2	0.7-1.2	1.2-1.9	0.45-0.6	1.4-1.7
Urban	0.15-0.6	0.55-1.1	1.15-1.8	1.4-1.6	0.5-0.9	1.2-2.1	0.5-0.65	1.4-1.6
Rural-plus-Urban	0.3-0.85	0.85-1.7	0.9-1.6	1.8-2.1	0.7-1.2	1.05-1.8	0.6-0.7	1.4-1.5

one latent root very close to unity. It is too much to expect precise estimates from the data of such nature and volume.

Similar results were obtained for the six-commodity case. In both cases, data differences produced some differences in estimates.

The calculations omitting one or two rounds indicate that the estimates are not very sensitive to such actions. In this sense, the short length of the series is not a very serious limitation.

We shall first consider the estimates $\bar{q}_i$. The appearance of some negative $\bar{q}_i$'s probably means that these committed quantities are small and nonsignificant. (A negative $\bar{q}$ *is* not quite absurd, however.) For the rural-plus-urban three-commodity case, this happens for group 2, i e, food excluding foodgrains. Even for non-food, the $\bar{q}$ is not definitely significant. Very broadly speaking, only the committed quantity for foodgrains is definitely positive in all three sectors, while those for the other groups are unreliable and probably small. But perhaps, $\bar{q}_2$ for food excluding foodgrains may also be taken to be positive in urban areas, but not in rural. These results are sensible, but the estimates are too erratic.

In the 6-commodity case, $\bar{q}_1$ (for foodgrains) is positive for all sectors, but larger for the rural sector than for the urban; while all other $\bar{q}_i$'s seem to be positive for the rural sector, most of them except for those for fuel and light show at least one negative half-sample estimate for the urban sector. On the whole, the results are acceptable but erratic, as in the three commodity case.

Our intervals for the b_i's give the following intervals of elasticities with respect to total expenditure. As stated earlier, the intervals may be too narrow for the 6-commodity situation.

Actually, the elasticities vary a little over rounds, but we have considered the average elasticity by dividing the b_i's by average proportion of budget spent on the commodity. The most plausible estimates should be around the midpoints of the intervals shown in Table 3.

In the Stone model, elasticities with respect to C_t are the same for q_{it} and p_{it} q_{it}, but these elasticities are more nearly comparable with quantity elasticities from cross-section data than with the value elasticities, since the difference between these is mainly due to quality variation not reflected in time-series data. But quantity elasticities are not available for many commodities. Anyway, the following gives the magnitude of some quantity and value elasticities from cross-section analysis (Bhattacharya, 1964). These are to be compared with the midpoints of the intervals given above. Quantity elasticities are not known for most commodities, but it is known that quality variation is more pronounced in urban areas (Iyengar, 1963).

It can be seen that, on the whole, the estimates obtained in the analysis cannot be regarded as "significantly" different from those in Table 4; the agreement is rather good except for foodgrains in the urban sector.

Conclusion from Cross-section Analysis

On the whole, the time-series estimates have very wide margins of uncertainty. The series could be somewhat lengthened, but the uncertainty arises mostly out of the unreliability of the NSS estimates of p_i, q_{it} (as indicated by half-sample divergence) and the lack of information regarding price move-

Table 4: Magnitude of Quantity and Value Elasticities of Total Consumer Expenditure estimated from NSS Family Budgets for All-India

	Rural		Urban	
Commodity-Group	Quantity	Value	Quantity	Value
Foodgrains	0.3-0.5	0.4-0.6	0.1-0.2	0.3-0.4
Milk and Milk Products		1.5-2.0		1.4-1.8
Clothing		1.4-1.7		1.6-1.8
Fuel and Light		0.5-0.7		0.7-0.8

ments (as shown by the effect of using retail instead of wholesale prices for foodgrains). The data also suffer probably from multi-collinearity. The correlation between C_i's and p_{it}'s are very high. The correlation coefficients between p_{it} q_{it} and C_i are 0.96, 0.97 and 0.90 for the three broad groups using rural-plus-urban, combined data.

The Paul-Rudra estimates of income elasticities are within bounds of possibility, but the best point estimates are quite different and not far from the cross-section estimates based on the NSS[4]. The Paul-Rudra criticisms of cross-section elasticities cannot, certainly, be ignored. As yet, the cross section estimates have been obtained by almost primitive methods (Bhattacharya, 1964), by fitting some curve to the regression of per capita commodity consumption of a household (y) on the per capita total consumer expenditure on all commodities (x). It is not merely a question of trying different algebraic forms and selecting the best-fitting one from them (see Roy and Dhar, 1960; Sinha, 1966). One has also to take explicit account of the effects of household size and composition. Published tables show negative correlation between household size and x; and unpublished tables show that the proportion of adults is considerably higher at the higher levels of x. The methods of Prais and Houthakker (1955) seem to be too detailed for the purpose. Following Brown (1954), one might estimate separate elasticities for different size-composition classes of households, and average these to get one overall elasticity. Or, matters might be simplified by the use of some notional scale of equivalence, or some scale obtained by simple methods (see Nicholson, 1964, postscript). One has to try also to eliminate the effects of, say, education or occupation or size of landholding, which influence preferences and which may be correlated with x in the cross-section data. This may be done by introducing them in the regression equation or by having separate analyses for separate strata. In short, the income elasticity should show the influence of income, keeping all other factors fixed as far as possible. How far the current estimates approach this ideal is anybody's guess, and it may be risky to use them for forecasting demand.[5] Finally, efforts should be made to estimate quantity elasticities in addition to value elasticities for as many commodity groups as possible.

One subtle problem seems to be unnoticed in the literature on cross-section analysis.[6] The NSS enquiries employ a moving reference period, so households interviewed on different dates furnish accounts for different periods of 30 days (preceding the date of interview). As the date of interview is spread over the survey period, which may be a few months or one year, seasonal variation is superimposed on true variation between households. The size distribution of x exaggerates the true extent of inequality, and what is more important, the Engel curves may be distorted by the seasonality of x and y. The recent NSS rounds have subrounds of shorter duration, covering representative subsamples of the round sample. We can calculate subround-wise (or seasonwise) elasticities or round-based estimates mixing all subsamples together. But is there any rigorous method of estimating the elasticity that would obtain if all households furnished information for the same annual reference period? This latter elasticity seems to be the one required for most applications. (See Table 4).

That such problems are important is shown by two pieces of evidence. First, we note the sudden shift in cross-section elasticities estimated for clothing. Up to the 6th round of the NSS, the elasticities were around 0.9 for rural India and 1 for urban, but from the 7th round onwards the values were consistently higher, of the order of 1.5 for the rural sector and 1.7 for the urban sector (Krishnan, 1964). In the opinion of the present author this must be due to the change in reference period from the 7th round. Up to 6th round, the reference period was 'last 30 days' for all commodities, without distinction.

Second, while the first Agricultural Labour Enquiry (ALE) employing an annual reference period showed that the percentages of income spent on food and other groups did not change rapidly with income, the later ALEs carried out on the NSS pattern showed the usual shifts in such proportions, and in particular, a decline, with rise in income, in the percentage of income spent on foodgrains.[7] Seasonality might have been a disturbing factor in the later enquiries. (One might recall how in Friedman's theory of the consumption function (Friedman, 1957) the presence of transitory elements in income leads to the under-estimation of the marginal propensity to consume. The analogy with the present case is not complete, since the transitory (here seasonal) components of x and y are correlated. But essentially the trouble is the same in usual cross-section analyses of NSS data, viz, that the classifying character x is not permanent consumption but is affected by seasonality.)

If foodgrains consumption is much less seasonal than consumption of other commodities like clothing, then if households are ranked in ascending order of x, those with higher percentage spent on foodgrains (z) will tend to be those with lower values of x, just because they gave information for months in which they spent very little on other than foodgrains: and the reverse will happen for seasonal commodities. (This makes x and z interdependent, rather than x determining z.) It is quite likely that, because of the co-variation between seasons within households, the elasticities of many necessaries are being underestimated and those of many luxuries overestimated in the usual process of estimation, and the Paul-Rudra estimates may not be far from truth. It is important to re-examine the matter by carrying out type studies, by analysing NSS 1st round data relating to the annual reference period, and by using a better measure of permanent consumption by excluding unusual medical or ceremonial expenses, say, from total expenditure on all commodities.

For estimation of own-price and substitution elasticities — the latter might be ignored if only broad commodity groups are considered — the only promising line of approach may be to utilise the already numerous subroundwise estimates from the later rounds of the NSS, along with corresponding retail prices built up from NSS or other sources (see Brown, 1958). But the effects of seasonality must be somehow eliminated; perhaps the 'between seasons' income elasticity should be used to correct for income changes.

When the income-elasticity is known, it may not be absurd to guess the magnitude of the own-price elasticity using one's notions about the scope for substitution. As Shupack (1962, pp 571-573) showed, time-series analysis may not give any firm estimates of own-price elasticities even when data of very superior quality are available. Indeed, there are few really successful time-series analyses except for simple markets for particular commodities.

Before concluding, we might mention the very impressive attempt by Krishnan (1964) to estimate both types of elasticities by an integration of cross-section and time-series approaches. Krishnan used the NSS p_{it} q_{it}'s for households in different ranges of total consumer expenditure and in different rounds of the NSS. Again, lack of price data was the main hurdle. The same wholesale indices were used for all the expenditure levels, which is very unrealistic as shown by extensive unpublished tabulations of NSS material (see Iyengar and Bhattacharya, 1965). That Krishnan's results were generally sensible seems to be partly a matter of luck.

References

[1] Bhattacharya, N (1964): A survey of work on demand projection done at the Indian Statistical Institute, Calcutta. Mimeographed paper I S I (Calcutta). Planning Division Series.

[2] Brown, J A C (1954): The Consumption of Food in Relation to Household Composition and Income. Econometrica, 22, 444-460.

[3] Brown, J A C (1958): 'Seasonality and Elasticity of the Demand For Food in Great Britain, since De-rationing', Journal of Agricultural Economics, December 1958, 1-12.

[4] Friedman, M (1957): 'A Theory of the Consumption Function', Princeton University Press.

[5] Government of India, Ministry of Labour (1951): Report on the First Agricultural Labour Enquiry.

[6] Government of India, National Sample Survey (1960): Tables with Notes on Budgets of Agricultural Labourers, 11th-12th Rounds.

[7] Government of India, Planning Commission (1964): Report of the Committee on Distribution of Income and Levels of Living, Part I.

[8] Iyengar, N S (1963): Estimation of Quality Elasticities etc. Sankhya, Series B, 25, 15-22.

[9] Iyengar, N S and Bhattacharya, N (1965): On the effect of differentials in consumer prices index on measures of inequality. Sankhya, Series B, 27, 47-56.

[10] Krishnan, T N (1964): The role of Agriculture in Economic Development, Unpublished MIT thesis.

[11] Kuh, E and Mayer, J R (1957): How Extraneous Are Extraneous Estimates? Review of Economics and Statistics, 39, 380-393.

[12] Nicholson, J L (1964): Redistribution of Income in the United Kingdom in 1959, 1957 and 1953. In International Association for Research in Income and Wealth: Series X on Income Redistribution and the Statistical Foundations of Economic Policy.

[13] Paul, Pushpam and Rudra, Ashok (1964): 'Demand Elasticity for Foodgrains', The Economic Weekly, November 28, 1964, pp 1893-4.

[14] Prais, S J and Houthakker, H S (1955): 'The Analysis of Family Budgets', Cambridge University Press.

[15] Roy, J and Dhar, S K (1960): in 'Studies on Consumer Behaviour', Indian Statistical Series No 6, Asia Publishing House, Bombay.

[16] Rudra, Ashok (1964): 'Relative Rates of Growth: Agriculture and Industry', The Economic Weekly, November 7, 1964, pp 1773-87.

[17] Shupack, M B (1962): 'The Predictive Accuracy of Empirical Demand Analysis', Economic Journal, LXXII, 550-575.

[18] Sinha, R P (1966): 'An Analysis of Food Expenditure in India'. Journal of Farm Economics, 48, 113-123.

[19] Stone, J R N (1954): Linear Expenditure Systems and Demand Analysis: An Application to the Pattern of British Demand. The Economic Journal, LXIV, 511-27.

[20] Stone, J R N, (1964): Models for Demand Projections. In 'Essays on Econometrics and Planning' edited by C R Rao, Pergamon Press.

[21] Stone, Richard, Aitchison, J and Brown, J A C (1955): Some Estimation Problems in Demand Analysis. The Incorporated Statistician, 5, 1-13.

Notes

[1] Theory shows that substitution effect with respect to own price is negative for any commodity. This implies $0 < b_i < 1$ for all i. These and the constraint $C_t - \Sigma_i p_{it} \bar{q}_i > 0$ for all t are not easily utilised in any method of estimation. In the present case, these were automatically satisfied.

[2] We did some regression analysis using model (2); the results will be reported separately.

[3] Thus, according to Table 5 of Paul-Rudra (1964) the fits are closer for the three larger commodity-groups among the six. In the three commodity cases, the groups are fortunately nearly equal.

[4] Some differences could, in fact, be expected. For discussions on the conceptual distinction between cross-section elasticities and time-series elasticities see Kuh and Meyer (1957).

[5] Another point worth investigation is brought out in Sinha (1966), viz, the distinction between elasticities of consumption of home-produced commodities and of consumption out of purchase. This raises basic questions regarding the validity of the model.

[6] The problem exists in many other countries, but we shall discuss it in terms of the Indian situation.

[7] See references 4 and 5.