# THE EFFECT OF INCOME AND FOOD PRICE CHANGES ON THE ACQUISITION OF FOOD BY LOW-INCOME HOUSEHOLDS

by

Harold Alderman

Report submitted to:

Agency for International Development
Office of Nutrition

Under DAN-1274-C-00-3077-00

International Food Policy Research Institute
1776 Massachusetts Avenue, N.W.
Washington, D.C. 20036

October, 1984

# TABLE OF CONTENTS

Page

VOLUME I. A REVIEW OF THE EVIDENCE

ACKNOWLEDGMENTS . . . . . v
SUMMARY . . . . . vi
1. INTRODUCTION: THE IMPACT OF FOOD POLICY, INCOME, AND PRICE CHANGES ON HUMAN NUTRITION . . . . . 1
2. POTENTIAL USES OF PARAMETERS OF CONSUMER RESPONSE IN MODELING FOOD POLICY . . . . . 9
Modeling Impacts of Income Growth . . . . . 9
Leakages Out of Transfer Programs . . . . . 11
Estimating Linkages Between Sectors . . . . . 15
Uses of Price Parameters on a Disaggregate Level . . . . . 19
The Use of MPCs to Identify At-Risk Groups . . . . . 22
3. EMPIRICAL PATTERNS OF PARAMETERS AND THEIR RELATIONSHIP TO POLICY: A REVIEW OF PRINCIPAL STUDIES . . . . . 26
Application of Frisch Technique . . . . . 26
Estimates Using Cross-Sectional Data . . . . . 30
United States Data . . . . . 30
Indonesia . . . . . 32
Thailand . . . . . 33
Brazil . . . . . 34
Bangladesh . . . . . 36
Sri Lanka . . . . . 38
Estimates Combining Time Series and Cross-Section Data . . 38
Philippines . . . . . 38
India . . . . . 39
Other Estimates Using System of Demand Approach . . . . . 42
4. GENERALIZATION OF PATTERNS . . . . . 45
Patterns in Income Elasticities . . . . . 46
Patterns in Own-Price Elasticities . . . . . 52
Magnitude . . . . . 52
Relationship of the Elasticities of Poor and Rich Consumers . . . . . 55
Conclusion . . . . . 72
Recommendations for Application of Parameters . . . . . 74
APPENDICES
Table 1 -- Meat Products Elasticities . . . . . 76
Table 2 -- Rice Elasticities . . . . . 77
Table 3 -- Wheat and Bread Elasticities . . . . . 79
Table 4 -- Coarse Grain Elasticities . . . . . 80
Table 5 -- Root Crop Elasticities . . . . . 81
Table 6 -- Milk Product Elasticities . . . . . 82

TABLE OF CONTENTS
(continued)

Page

## Figures

Figure 1 -- Price and Income Elasticities for Lower Income Consumers . . . . . 3
Figure 2 -- Price and Income Elasticities for Upper Income Consumers . . . . . 4
Figure 3 -- Plot of Standardized Price Elasticities on Standardized Income . . . . . 5

## Tables

Table 3.1 -- Studies Presenting Price Elasticities by Income Group . . . . . 27
Table 4.1 -- Calorie Expenditure Elasticities for Low-Income Families . . . . . 49
Table 4.2 -- Food Expenditure Elasticities for Low-Income Families . . . . . 50
Table 4.3 -- Cross-Tabulation of Price Elasticities and Real Income Effects of Price Changes . . . . . 54
Table 4.4 -- Parameters of Regression of Hicksian Elasticities on Log of Income . . . . . 57
Table 4.5 -- Regressions of Standardized Substitution Elasticities on Log of Income . . . . . 60
Table 4.6 -- Regressions of Standardized Substitution Elasticities on Income Share . . . . . 61
Table 4.7 -- Regressions of Standardized Substitution Elasticities on Population Percentile . . . . . 62
Table 4.8 -- Regressions of Standardized Uncompensated Elasticities on Population Percentile . . . . . 65
Table 4.9 -- Regressions of Standardized Uncompensated Elasticities on Income Share . . . . . 66

TABLE OF CONTENTS
(continued)

Page

VOLUME II. THEORETICAL AND METHODOLOGICAL CONCERNS FOR MEASUREMENT OF DEMAND PARAMETERS FOR LOW-INCOME CONSUMERS

INTRODUCTION . . . . . 83
1. RESTRICTIONS OF PARAMETERS FOR ESTIMATING CONSUMER RESPONSE WITH DISAGGREGATED COMMODITY GROUPINGS . . . . . 86
Systems of Demand . . . . . 86
Testing of Restrictions for Demand Analysis . . . . . 91
Restrictions by Judicious Choice . . . . . 97
Choice of Functional Form . . . . . 100
2. ESTIMATION CONCERNS ASSOCIATED WITH CROSS-SECTIONAL DATA . . 107
Interpretation of Parameters . . . . . 107
Case Study of Time Series . . . . . 116
Non-Consumption . . . . . 120
Case Studies on Non-Consumption . . . . . 122
3. CRITERIA FOR DISTINGUISHING GROUPS FOR PARAMETER ESTIMATION. 132
Case Study: Sample Truncation . . . . . 136
Real Income and Household Demographics . . . . . 140
4. OTHER ISSUES . . . . . 148
Variation of Parameter by Source of Income . . . . . 148
Issue of Control of Family Income . . . . . 154
Case Study: Effects of Women's Earnings on Marginal Calorie Purchases . . . . . 160
Distribution of Household Purchases . . . . . 164
5. CONCLUSIONS AND RESEARCH RECOMMENDATIONS . . . . . 168
BIBLIOGRAPHY . . . . . 171

Figure

Figure 1 -- Truncation of Normally Distributed Calorie Variable . . . . . 134

TABLE OF CONTENTS
(continued)

## Tables

Table 2.1 -- Income and Price Elasticities - Mindanao . . . . . 118
Table 2.2 -- Income and Price Elasticities - Visayas . . . . . 119
Table 2.3 -- Alternative Estimates for Grain in Mindanao . . . . 124
Table 2.4 -- Alternative Estimates for Grain in Visayas . . . . 125
Table 2.5 -- Alternative Estimates for Food in Sri Lanka . . . . 126
Table 3.1 -- Table Results of Monte Carlo Experiment on Sample Groupings . . . . . 138
Table 4.1 -- Models Used to Estimate the Marginal Propensity to Consume Calories . . . . . 162
Table 4.2 -- Marginal Propensities to Consume Calories for Men and Women . . . . . 163

## ACKNOWLEDGMENTS

Many people contributed to this study. Howarth Bouis, Michael Demousis, Per Pinstrup-Andersen, and David Sahn all had a hand in the conceptualization and refinements of the project. Bruce Cogill assisted in making the estimations for the case studies. Patricia O'Brien-Place and Roberta van Haeften of the Nutrition Economics Group, Office of International Cooperation and Development, United States Department of Agriculture, provided support and advice on many aspects of the report. The report and my continuing education greatly benefitted from many insightful comments and suggestions made by Angus Deaton on an earlier draft.

Publication of this report on a revised version as an IFPRI research report is currently under consideration. I am grateful to the United States Agency for International Development, Bureau of Science, Office of Nutrition, for funding under DAN-1274-C-00-3077-00.

## SUMMARY

Planners and government officials regularly employ measurements of a population's or subpopulation's response to changes in incomes and prices in their formulation of food policy. While it has been recognized for some time that the poor are more likely than wealthier consumers to adjust their purchases of food when their income changes, only in recent years have analysts attempted to determine whether the poor are similarly more responsive to price changes. This report reviews the evidence of those attempts and discusses the econometric problems raised by the study of the food demand of groups within a population.

Fifteen studies that have estimated price response by income group are compared in this report. They cover 11 countries and differ in their methods of estimation, in that assumptions underlying the choice of functional form, and in the periods over which their data was collected. Some general patterns emerge. The disaggregate price elasticities of the studies are frequently large in absolute value, exceeding 1 in a third of the estimates. This implies that consumers readily substitute other foods and nonfoods for another food item when that item's price rises. The poor are statistically more likely to make such substitutions than the well-off. Such substitution is in addition to the changes in food consumption that the poor make following a price rise that is attributable to reduction in real income.

A number of generalized relationships are identified. For example, it is estimated that a 10 percent increase in income leads

to declines in substitution elasticities of between 1 and 3 percent. Similarly, it is estimated that the price elasticity for rice of the poorest 10 percent of a population will be 80 percent higher than the mean value for the whole population. Similar relationships were estimated for meat, wheat and coarse grains, root crops, and milk. In all cases, the pattern of declining price responsiveness with increased income (or ranking in income) was statistically significant.

Responses to income changes were indicated by estimations of income elasticities for food expenditures and calories. Food expenditures for families that consume 1,750-2,000 calories per day increase by 8.2 percent with a 10 percent increase of income. Calorie intakes increase, however, by only 4.8 percent, as a portion of the increased expenditure goes to increasing the perceived quality of the diet. More of the variation of calorie intake is not explained by income than there is variation of food expenditures.

The policy implications and uses of such parameters include estimations of leakage of subsidies to nonfood items. Increases in income of the poor do lead to increases in calorie consumption greater than those of the overall population but transfers cannot be expected to be used solely to raise calorie intake. However, with high propensities to spend on food it is also unlikely that in-kind transfers will reduce demand for agricultural products. With larger price responses, the poor are more likely to increase their food consumption following the introduction of price subsidies or decrease it when domestic food prices rise. It is possible, however, to identify food items consumed mainly by the poor -- for example, coarse grains -- for

which the poor are price-responsive, and in so doing to target pricing policies. It is important, however, to note that an increase in the food consumption of one item due to price changes is partially offset by decreases in the consumption of others. While the net effect can be low in some countries, it is frequently appreciable even for the poor.

The estimation of price elasticities by income group is hindered by the scarcity of data with sufficient variation in prices. Time-series data generally do not report consumption by groups, and therefore most, though not all, estimation has been done with cross-sectional data. Restrictions on parameters derived from theory and from assumptions about the nature of consumer utility make it possible to estimate reductions of parameters. This computational advantage comes at the price of unrealistic behavioral models when dealing with a number of food items and is not recommended for food policy analysis.

Cross-sectional studies may generate price parameters that differ from time series, with the parameters from the former generally being larger. Evidence for this is mixed, however, and there is a need to obtain panel or other disaggregate data for several periods to explore this pattern further. Cross-sectional studies have the additional problem of including nonconsumers. This can lead to serious estimation biases that can be handled by techniques that study the probability of a purchase as well as the amount purchased.

If one distinguishes subsamples of a population to be studied by levels of calories consumed rather than by income, serious biases are also possible. The theoretical prediction that income elasticities

will be underestimated when the sample is subdivided by intake is verified by 10 statistical experiments reported here.

Another concern for policy is whether sources of household income, distinguished by gender or by type of employment, affect income and price parameters. While there is empirical evidence that household demographies affect demand and a theoretical foundation for their inclusion in demand studies, there is less experience with sources of income. In a study presented to illustrate some of the methodological problems of such an investigation, it was found that the propensity of households in Sri Lanka to increase calories was higher following an increase in income by women than following an increase in earnings by men. Increases in male earnings, however, led to higher total food expenditures than increases in women's earnings. These differences were statistically significant but were not large, and therefore have limited policy implications. They are, however, of importance for the understanding of the nature of interhousehold behavior and for the application of demand theory to practical problems.

# IMPACT OF INCOME AND FOOD PRICE CHANGES IN FOOD ACQUISITION BY LOW-INCOME HOUSEHOLDS

## VOL. I. A REVIEW OF THE EVIDENCE

by

Harold Alderman

Report submitted to:

Agency for International Development
Office of Nutrition

Under DAN-1274-C=00-3077-00

International Food Policy Research Institute
1776 Massachusetts Avenue, N.W.
Washington, D.C. 20036

October, 1984

# 1. INTRODUCTION: THE IMPACT OF FOOD POLICY, INCOME, AND PRICE CHANGES ON HUMAN NUTRITION

A complex web of interactions contributes to malnutrition and, therefore, a government may influence the nutritional status of its population through a variety of actions. Some improvements in nutritional status may be achieved through programs that do not appreciably change the economic environment or a family's physical endowment -- for example, through health care and education programs. However, prices and incomes are principal determinants of a family's food intake and, hence, its nutrition. Consequently, a government's food policies become a major influence on the nutritional status of a population, whether or not such involvement is designed with nutrition as a primary goal.

In most societies where considerable malnutrition exists, the households with malnourished members responded to changes in income and prices differently than society as a whole. While knowledge of the average price response of a population may be sufficient for determining a market-clearing price, it will not enable a planner to know how different parts of a population will change their consumption following price changes. A related point is that the absence of information on commodity substitution limits a planner's ability to influence the consumption of specific groups through price mechanisms without introducing disincentives, costly general subsidies, or

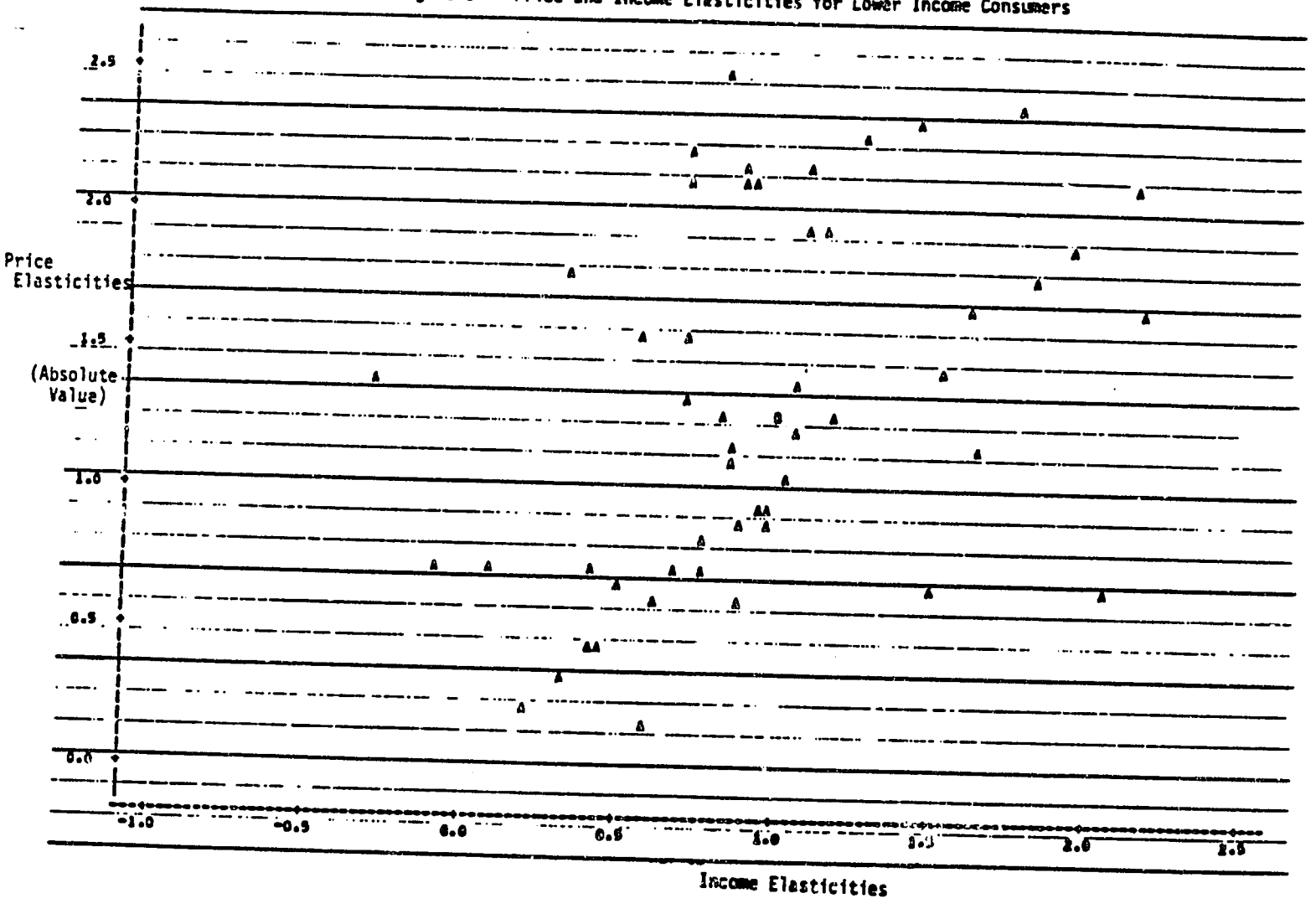
targeting schemes that would strain the administrative resources of a country.

Reliable estimates of the effects that changes in prices and incomes have on groups of households with malnourished members are scarce, however, and their use in designing food policy has been limited. Only in the last ten years have efforts been made to distinguish the price response of poor households from the response of the whole population.

There is no consensus on whether it is necessary or feasible to make empirical studies of the size of readjustment of the purchases by poor households following price changes. While few studies have focused on how much demand for food by low-income households is influenced by changes in the economic environment, those that have generally indicate that the poor do responded differently than the general population. This is illustrated graphically by Figures 1-3.

Figure 1 illustrates a set of disaggregate price elasticities and income elasticities for the poor for a number of food items. The elasticities for upper income consumers are shown in Figure 2. These points are derived from surveys and estimates discussed at length later in the report. The data points illustrate the percentage change of consumption of various food items following either a percentage increase in income or a percentage decrease in the price of the food item. While the data points are scattered, reflecting differences in areas and commodities covered, it is clear that the points for the low-income consumers are, in general, higher and further to the right

Figure 1 -- Price and Income Elasticities for Lower Income Consumers




A = One observation
B = Two observations

Figure 2 -- Price and Income Elasticities for Upper Income Consumers

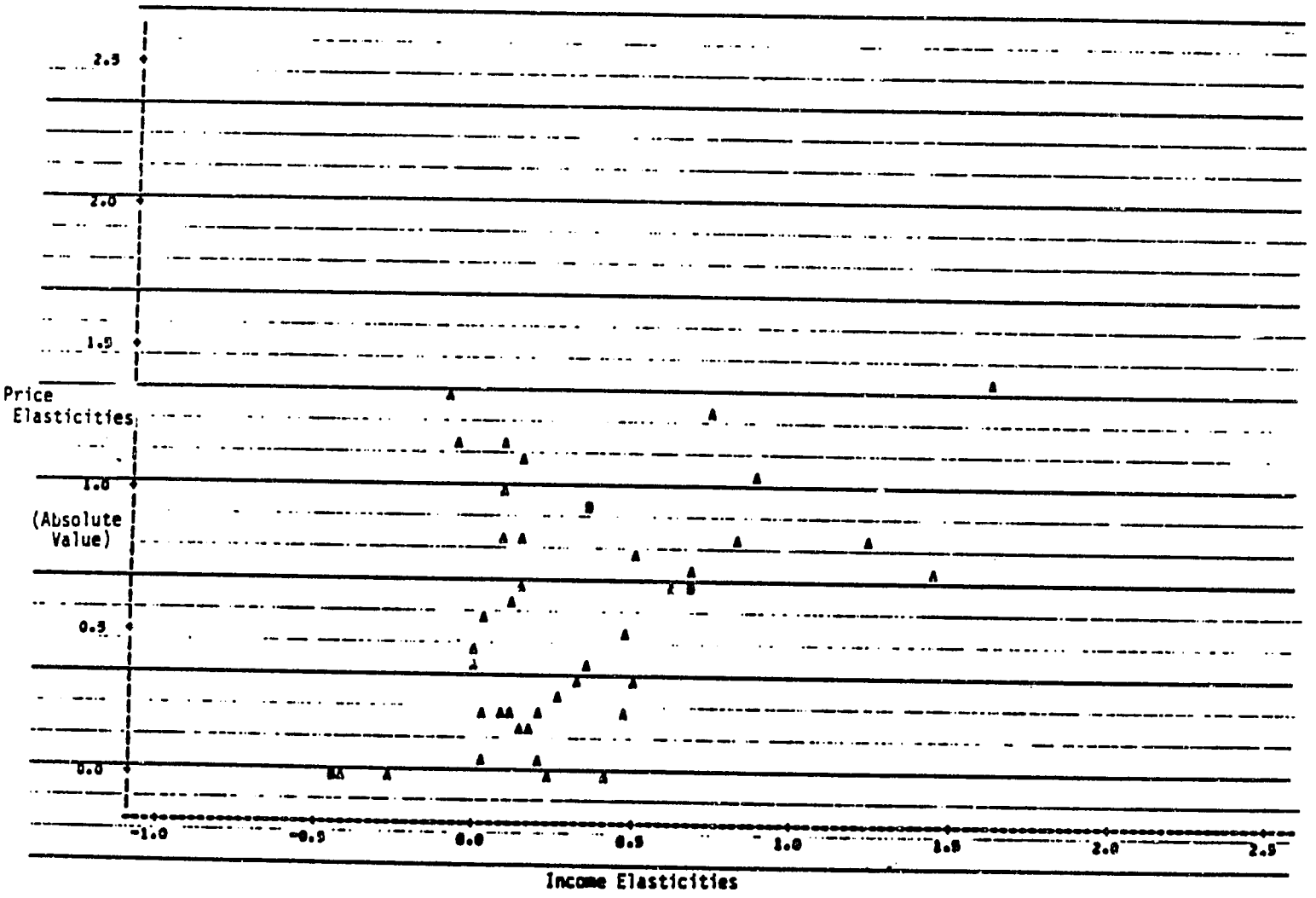
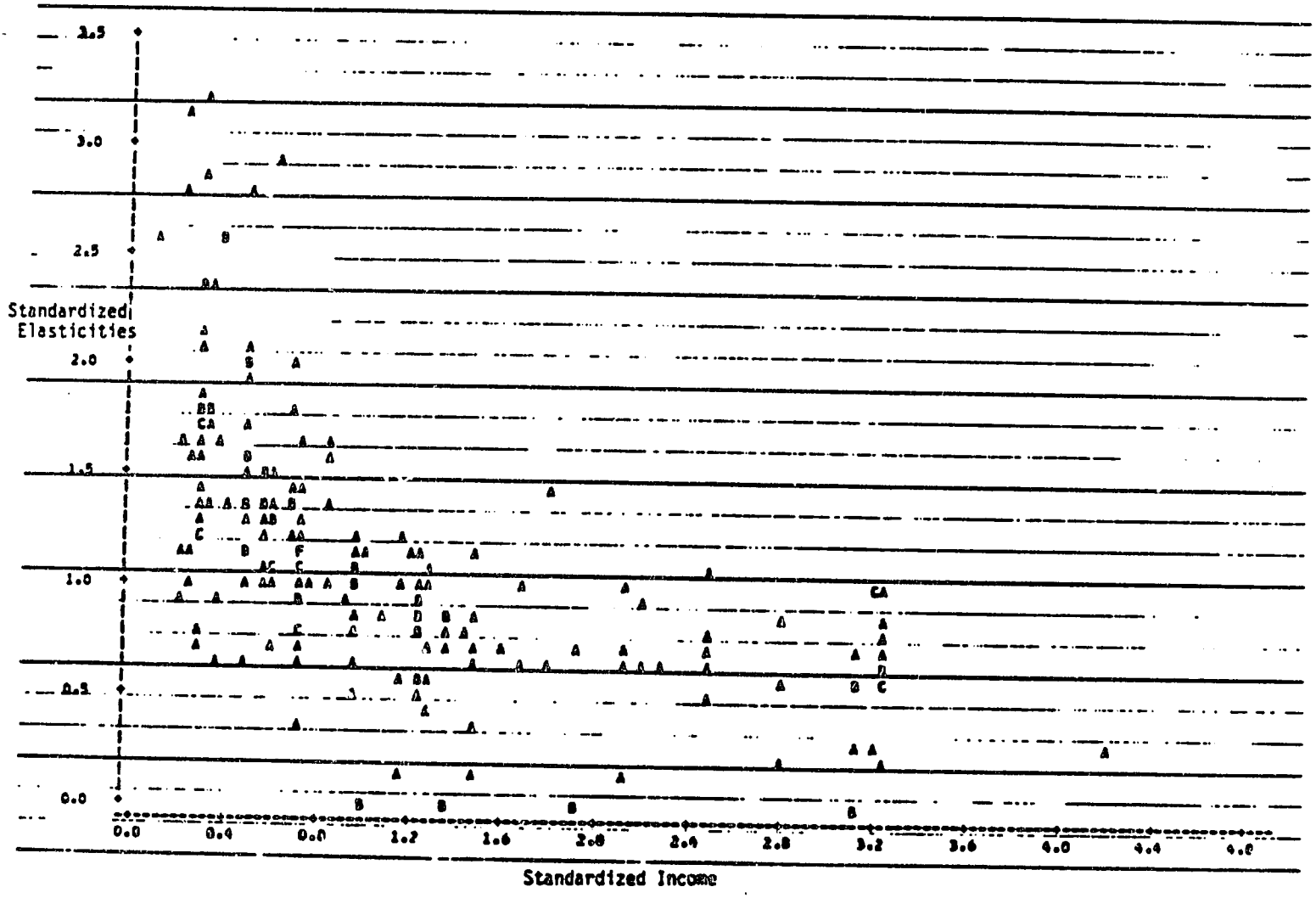


Figure 3 -- Plot of Standardized Price Elasticities on Standardized Income




Elasticities are standardized for each commodity by dividing the income group's elasticity in each country by that country's mean elasticity for the commodity.

Incomes are standardized in a similar manner.

than the points for the upper-income consumers. This implies that both income and price elasticities are larger for the poor.

Figure 3 smooths out much of the intracountry variance by standardizing the price elasticities of different income groups in each country by dividing them by the mean value for that commodity in the country. Similarly, the horizontal axis represents incomes that are standardized by dividing the group's mean value by the sample mean income. Virtually all points that represent price responses greater than the mean (standardized elasticities greater than one) are found in the portion of the graph corresponding to groups with incomes below the national average.

The purpose of this study is to analyze the patterns in food purchasing behavior illustrated in these graphs. The review assesses the current state of knowledge revealed by existing studies covering different economic environments. Any generalizable findings from these studies that may be useful for planners will be identified. In part, it is hoped that such patterns will reduce uncertainty when policy must be made with limited information. In addition, current knowledge should be useful for guiding future research on the determinants of food acquisition by the poor.

This is not merely an issue of improving the precision of measurements. The attractiveness of certain policy measures may hinge on knowledge of how different groups would respond to them. At times the effect of a policy or the need for compensating measures are not apparent from theoretical or qualitative knowledge and can only be known from empirical results.

The next chapter of this volume includes a discussion of some of these applications of demand parameters in food policy work. This is to illustrate why there is an interest in measuring the responses of the poor, as well as how much they currently consume.

The response of the poor to food price changes is a special part of welfare analysis. Most economists argue that real income is the main measure of household welfare and that changes in quantities of various goods consumed are, in part, welfare-preserving adjustments that maximize well-being with a given income. The specific concern with food consumption diverges from pure consumer sovereignty and reflects the social process of transfer programs.[1] It also reflects the externalities of nutrition in health and human capital development and the possibility that the maximization of the welfare of an individual and a household might differ. This paper will focus on the specific nutritional concerns, although the economists' concept of welfare should not be forgotten.

The interest in the responses of the poor is not new. What is new is the possibility of applying econometric techniques to new data sources. The scarcity of the type of data necessary to conduct such studies has been the principal barrier to directly estimating parameters disaggregated by income groups.

Cross-sectional data sets may provide an acceptable basis for estimating income parameters and, therefore, the income effects of price changes. However, they may not provide reliable estimates of

---

[1] See, for example, J. Tobin, "On Limiting the Domain of Inequality," Journal of Law and Economics XIII, 1970, pp. 263-277; and L. Thurow, "Cash Versus In-Kind Transfers," American Economic Review 64, 1974, pp. 190-195.

price parameters unless they refer to several points in time or several geographical locations and thus provide for sufficient and relevant price variation. Time series that have data by income strata are even scarcer than broad cross-sectional data sets.

The costs of establishing such a data base are high and the time needed frequently exceeds what planners have. An alternative is methodological innovation. Using economic theory and advances in computerized techniques, a number of researchers have explored ways to estimate demand parameters with limited price variations. The third chapter of this report discusses 15 studies which report such estimates using a variety of techniques and data sets from 11 countries.

It is recognized that policy planners may not always have the data necessary for such estimates. For this reason, the available evidence is analyzed in Chapter 4 and patterns from those studies are illustrated using statistical evidence.

As this section is intended to assist policy planners, efforts have been made to minimize the technical language and algebra in the text. But since much of the innovations and much of the debate about the meaning of the results involve econometrics, this review includes a second volume on methodology. Some readers may prefer that it precede the discussion of results, while for others it may be too technical for their tastes. No attempt has been made to produce a state-of-the-art consumer theory. Rather, Volume II discusses particular problems and methodologies that are important when disaggregating by income groups and by commodities. It also discusses particular features of cross-sectional data used to estimate price parameters.

## 2. POTENTIAL USES OF PARAMETERS OF CONSUMER RESPONSE IN MODELING FOOD POLICY

### Modeling Impacts of Income Growth

Human nutrition may be affected by food policy through changes in 1) family incomes; 2) the prices a household has to pay, which in turn reflects the availability of production and imports; 3) fluctuations in these prices and incomes; 4) the composition of household income between cash, credit and in-kind earnings; 5) intra-household income distribution and budget control; 6) the opportunity cost of time for family members, which affects not only incomes through wage potentials but allocation in nonmarket activities as well. Age, reproduction status, and level of activity necessary for employment are also relevant to nutrition status.

Governments have several instruments that they can use to bring about changes in the determinants of household food consumption. They may, for example, influence price through marketing boards, price floors and ceilings, or by changes in imports. The first two, however, also affect income policies while the changes in imports contribute to and reflect a variety of foreign exchange quantity and rate regulations. Indeed, as the understanding of the intricacies of the interaction of macro- and microeconomic factors has improved, it has become increasingly clear that the degrees of freedom a government has to determine instruments are far fewer than the number of goals a planner may hope to achieve. It is the interdependence of the

instruments of policy that stimulates the interest in developing finer tools for policy planning. This, in turn, stimulates an interest in greater precision in measuring the reaction of groups of consumers and producers to changes in their endowments and in prices.

This chapter presents a number of examples of such uses of estimates of consumer response for policy planning. It is clear that in the long run policies that promote equitable growth will eliminate most forms of malnutrition caused by underconsumption. Furthermore, as is frequently pointed out, government funds used to raise current consumption are unavailable for capital investment and, therefore, may delay that growth. While to some degree this tradeoff is not strictly an either/or choice, as current consumption may increase human capital through improved health and education, it is still useful to be able to estimate the time frame of the impacts of income growth on nutrient consumption. If planners view the speed at which development goals alleviate underconsumption as too slow, they can choose to either redirect development programs or to introduce programs that focus on the nutrition of specific groups while general growth strategies are being implemented.

Reutlinger and Selowsky's approach to projecting the number of undernourished people from 1965 to 1990 using alternative growth rates was influential. They found that the absolute number of people at risk of undernutrition would increase in that period, although the percentage and the average calorie deficit would decline. Their study,

---

[1] S. Reutlinger and M. Selowsky, Malnutrition and Poverty (John Hopkins Press, 1976).

however, used an assumed income elasticity in order to generate projections worldwide. The elasticity chosen for calories was 0.15 for intake equal to the calorie requirement using a semi-logarithmic form. Two similar studies used income elasticities derived from household budget data from seven countries to make projections under alternative scenarios that included changes in growth and exogenously determined price changes to make the same point.[1] Berg presented his results under a chapter with a revealing heading, "The Short-Run Inadequacy of the Growth Process." He notes that under historical rates of income growth and no price changes, it could take between 22 (high-growth scenario, Morocco) and 235 years (low-growth scenario, Bangladesh) to eliminate the food energy deficit in the lowest decile of the population in the countries studied. If the assumption of neutrally distributed income growth was varied, Berg, and Knudsen and Scandizzo as well, found that the results could be quite different. Both studies also found that the calorie income elasticities of the poor (variously defined) differed greatly between countries, indicating that for some populations the effect of income growth on nutrition would be more pronounced than in others.

## Leakages Out of Transfer Programs

Recognizing that income growth may be insufficient to deal with

---

[1] A. Berg, Malnourished People: A Policy View, Poverty and Basic Needs Series (Washington, D.C.: World Bank, June, 1981); and O. Knudsen and P. Scandizzo, "The Demand for Calories in Developing Countries," American Journal of Agricultural Economics 64, 1982, pp. 80-86. These studies overlapped in generation and data, using four data sets in common.

underconsumption within a planning horizon or that pockets of underconsumption may exist within an affluent society raises the possibility of targeting nutritional policies on specific groups. It is for such a purpose that disaggregated demand parameters are most commonly used. Assume that the government has distinguished groups for which the raising of food intake is a particular goal of food policy. These may have been identified by nutritional indicators, income, or any of a number of other characteristics. Having identified the target groups, one can look at either the effect of specific programs designed to increase the food consumption of the groups or one can investigate the effects of other policies that do not have explicit nutritional goals but may, nevertheless, have a large effect on food acquisition. Even before modeling price parameters, marginal propensities to consume (MPC) or income elasticities can be used to provide information about the targeting and the spillovers of programs and policies under consideration.

For example, the absolute benefits to the poor of a subsidy on particular commodity will be highest in per capita terms if the income elasticity is negative, and the relative benefits to the poor (as a share of total income) will be higher if the elasticity is less than 1.[1] This information is, however, more readily available by looking

---

[1] Note that there is a formal relationship between income elasticities and the distribution of a commodity that has been elucidated by Kakwani, who has used Lorenz curves to estimate elasticities. See N. Kakwani, "On the Estimation of Engel Elasticities from Grouped Observations with Application to Indonesian Data," Journal of Econometrics 6, 1977, pp. 1-19. See also P. Pinstrup-Andersen, "Decision Making in Food and Agricultural Research Policy: The Distribution of Benefits from New Agricultural Technology Among Consumer Income Strata," Agricultural Administration 14, 1977, pp. 13-25.

at group means than by estimating elasticities with statistical techniques. The elasticities provide a summary representation.

Other information is revealed more clearly by estimates of parameter. For example, consider the concept of leakage which pertains not to the share of subsidies going to nontarget groups but to the percentage of an income transfer going to nontarget items. In any income transfer program focused on food (including programs embodied in a commodity subsidy available in inframarginal amounts), leakages can be defined as the portion of income not spent on food. This, then, is reflected in the MPC for nonfood items or (1-MPC) for food. The fiscal cost of raising the food consumption of the target group is inversely proportional to the MPC for food; the higher the MPC the lower the cost. For example, suppose the MPC for food is 0.4, then for each additional dollar to be spent on food, $2.50 needs to be transferred. While such leakages reflect consumer preferences, they dilute the nutritional effect of transfer programs.

Selowsky has used such information to explore the effectiveness of transfer programs.[1] For example, if resale of in-kind transfers is possible and costless, then it would be expected that transfers in excess of previously consumed amounts will be resold if they exceed 1/(1-MPC) times the amount consumed initially. This pertains to the actual sale of the commodity or food stamps; substitution of the transfer commodity for other purchases would occur at all levels. The

---

[1] M. Selowsky, "Target Group-Oriented Food Programs: Cost Effectiveness Comparisons," American Journal of Agricultural Economics 61, 1979, pp. 988-999.

effectiveness of such programs, then, may be the same as a cash transfer. If, on the other hand, the possibility of resale is eliminated, for example, by on-site feeding, then the nutrition program will be more cost-effective in raising consumption than an income transfer, provided that the transfer exceeds the original consumption/(1-MPC). Selowsky also investigates the cost effectiveness of subsidy programs, food stamp programs with payment in stamps, and programs that target food to children within a family.[1] Evaluating the effectiveness of the former requires knowledge of price as well as income parameters.

Studying consumption within the family requires information on household shares of food. In the case of a school feeding program, suppose that children consume 20 percent of the family's food; then, if one wants to raise children's consumption by an amount greater than 0.2X (Average Propensity to Consume in Household)/[1-0.2 MPC]), one must feed the child the entire amount of food deemed appropriate.[2] That is, the program must be able to compensate for 100 percent substitution for food previously eaten at home. In practice, the effectiveness of such programs involve a variety of institutional arrangements not captured in household parameters. Nevertheless, such analyses of theoretical efficiencies are useful guides for project expectations.

---

1 These are also discussed at length in Reutlinger and Selowsky, Malnutrition and Poverty.

2 Selowsky uses MPC in the numerator, though APC seems more appropriate.

Another use of MPC in studying leakages comes when one is calculating the potential disincentives to producers from food aid. It is frequently argued that concessional aid can partially offset the production disincentive by raising the demand for food to match the increased supply. Income can be transferred either implicitly -- with concessionally priced foods available in inframarginal amounts -- or explicitly, with a cash transfer possibly funded by revenue generated by selling food, obtained at concessional rates, to the general public at commercial rates. If one knows the MPC of the target group, one can calculate the accompanied transfer (or wage component of a food-for-work program) that will allow local producer incentives to be constant.

## Estimating Linkages Between Sectors

Another example in which marginal spending patterns indicated by MPCs are employed is in the study of linkages from one type of development program to other sectors of the economy. In one variant of this concept, low-income consumers are viewed as consuming products that have large share of labor and low shares of capital, foreign exchange, or both. Growth of the income of low-income consumers, then, leads to even more equitable growth patterns as labor-intensive goods are purchased. Knowledge of the MPC of various income groups, along with the factor shares of various products, allows this hypothesis to be tested. King and Byerlee cite studies from Pakistan and Turkey supporting this view and use their own estimates from Sierra

Leone as further evidence.[1]

Another variant of this theme is in terms of locational linkages. One form of this approach looks at the foreign exchange component of expenditure patterns and, therefore, is similar to the approach that concentrates on scarce factors. Another form looks at leakages from rural areas into urban. This latter type of linkage was also studied by King and Byerlee, who found that the propensity to consume goods produced in rural areas did not vary by income. Recently, Hazell and Roell have found rather different linkage patterns in Malaysia than in Nigeria.[2] They found that the shares of expenditures going to locally produced goods declined as income rose in Malaysia, but that they were more constant in Nigeria. The results of such regional studies may depend on the definition of linkage (a rural leakage is an urban linkage) and of tradeables and nontradeables, and they may be sensitive to where regional boundaries are set. Such results are useful, however, if one is interested in regional or sectoral equity.

Both factor share and regional linkage studies have moderate data requirements. They generally use household surveys to estimate MPC and do not need estimates of price parameters. Mellor has taken such a linkage approach closer to food pricing policies.[3] Since MPCs for

---

1 R. King and D. Byerlee, "Factor Intensity and Locational Linkages of Rural Consumption Patterns in Sierra Leone," American Journal of Agricultural Economics 60, 1978, pp. 197-208.

2 P. Hazell and A. Roell, Rural Growth Linkages: Household Expenditure Patterns in Malaysia and Nigeria, IFPRI Research Report No. 41 (Washington, D.C., 1983).

3 J. Mellor, The New Economics of Growth (Cornell University Press, 1976); and "Food Price Policy and Income Distribution in Low-Income Countries," Economic Development and Cultural Change 27, 1978, pp. 1-26.

grains and other foods are generally high for low-income consumers, Mellor notes that income growth among this group will increase demand for these fairly labor-intensive commodities. He also notes that a poverty-oriented growth strategy could become short-circuited if such linkages put pressure on food prices to rise or increase the amount of foreign exchange allocated to food imports. As new agricultural technology allows for reductions in the unit costs of producing food, the technological changes that have occurred in recent decades have offered the possibility of sustained poverty-oriented growth in countries such as India, growth that was not possible in earlier decades. The new technologies, then, can promote such growth in two ways: by increasing income in groups with high MPCs for labor-intensive groups that generate more employment, and by lowering the unit costs of producing food, which lowers costs of principal wage goods, and again, generates employment.

In actuality, the data necessary to estimate linkages from propensities to spend on different goods differ from the data ideally suited to investigate the changes of real income attributed to changes in production technology. The latter can be assessed accurately only if substitution parameters are known. For a first approximation, however, it is possible to use the real income effect of a price change. In Mellor's study, the effect of a price change was calculated by multiplying the budget share of foodgrains by the income elasticities of other goods. A standard Slutsky decomposition shows that this overestimates the effect for substitute goods. This approach, then, is suited for illustrating patterns of demand and the

derived demand for factors of production but is not designed to project quantities demanded following a price change.

More complex general equilibrium models have been developed to look at the growth and distribution of investment and consumption under alternative pricing policies.[1] McCarthy and Taylor observed that "typical in models of this type, most of the macro-adjustment occurs via price-induced change in real income."[2] This effect was illustrated by Taylor with a model of the Egyptian economy in which a reduction of food subsidies had a contractionary effect on the entire economy. While it is beyond the scope of this review to explore the fine points of such macroeconomic models, they illustrate one use of demand parameters for modeling linkages with the general economy. The linkages come through a number of loops in which factor demand, household income and savings, and investment are made endogenous. Demand parameters obviously affect equilibrium prices and, through iterative loops, income and savings. The demand parameters used in the consumption loop of such a model commonly come from a linear expenditure system (LES) as the linearity aids in the computations. While the assumptions used in the LES can be criticized on empirical grounds (see Volume II), there is little evidence about how sensitive the policy conclusions from such a model are to the source of demand

---

1 F. D. McCarthy and L. Taylor, "Macro Food Policy Planning: A General Equilibrium Model for Pakistan," Review of Economics and Statistics, 1979, pp. 107-121. See also L. Taylor, Macro Models for Developing Countries (McGraw-Hill, 1979).

2 F. D. McCarthy and L. Taylor, "Macro Food Policy Planning," op. cit., p. 116.

parameters. Similarly, there is little evidence from which one can gauge the optimal level of commodity aggregation to model nutritional impacts without overburdening the computations. McCarthy and Taylor disaggregated food into rice, wheat and sugar, as well as rice, wheat, and sugar products.

## Uses of Price Parameters on a Disaggregate Level

In the discussion above, the focus was on the use of models that study sectorial flows and macroeconomic performance. A planner is often not concerned with changes of the price of food as an aggregate index but with changes in the prices of individual commodities and for specific groups and locales. Often, in such cases, the complexity of the analysis may be increased by including more classes of consumers and goods and at the same time the task may be simplified by focusing on partial rather than general equilibrium.

Such an approach is particularly applicable to analysis of such consumer-oriented food policies as price subsidies. For example, Reutlinger and Selowsky[1] illustrated the cost effectiveness of a general food price subsidy under a variety of supply assumptions. The cost of achieving a target consumption for the poor will obviously decrease the larger their price elasticity is. Reutlinger and Selowsky noted that the cost is affected both by the elasticity of the poor and the relative sizes of the elasticities of the target group and the general population. Using formulas such as those derived by

---

[1] Op. cit.

Reutlinger and Selowsky, the fiscal cost of the subsidy can be calculated and the effectiveness of such an outlay relative to other programs can be evaluated. For example, they compare a general price subsidy to a targeted price subsidy that might be available at a fair price shop. They also consider income transfer programs as discussed above, including a food stamp program that requires purchase of stamps.

Such comparisons rarely include the costs of identifying the target group and assuring that they and only they are the recipients of the price subsidy. As these costs may be considerable, the cost effectiveness of a targeted scheme (including targeted income transfers) relative to a general price subsidy may be much less than estimated.

As an alternative to targeting by family, some researchers have proposed targeting price subsidies by commodity.[1] The cost-effectiveness of such a form of targeting is, again, inversely related to the price elasticity.[2] Also, as discussed above, the subsidy is more skewed to the poor if the commodity is an inferior good.

Indeed, analysis of such commodity targeting differs little from analysis of a general food price subsidy, except that in the former, $E_{total}$ is likely to be smaller relative to $E_{poor}$. However, the data

---

1 C. P. Timmer, "Food Prices and Food Policy Analysis in LDC's," Food Policy 5, 1980, pp. 188-199; also C. P. Timmer, W. Falcon, and S. Pearson, Food Policy Analysis (Baltimore: Johns Hopkins University Press, 1983).

2 If more than one item is subsidized or taxed, their inverse relationship is only approximate.

requirements for such analyses do differ. If one is calculating the effect of a price change of a single food item rather than the price change of an aggregation of food, information on substitutions between commodities is required, as the net calorie change from a food subsidy on one good will be reduced by decreased consumption of goods that are substitutes for the targeted commodity. The total response to such price changes is indeterminate, hence the need to estimate cross-price parameters for the target group.

The use of price parameters in nutrition-oriented food policy is not restricted to analysis of food subsidies and transfer programs. As an increase of the local supply of a commodity will affect the price, the distributional and nutritional effects of a change in agricultural production can be calculated with the aid of such parameters. This has been a subject of the studies by Mellor, cited above, and by Pinstrup-Andersen, et al.[1] Such studies also require information on the supply elasticities of the commodities in question and, if research is the instrument by which local supply is to be increased, information on the expected returns to research is useful. The study by Pinstrup-Andersen et al. found that the increase of the supply of a commodity was not a good predictor of nutritional effects even though such an approach is commonly used. The reasons for this included 1) differences in the shares of any commodity going to nutrient-

---

[1] P. Pinstrup-Andersen, N. Ruiz de Londoño and E. Hoover, "The Impact of Increasing Food Supply on Human Nutrition: Implications for Commodity Priorities in Agricultural Research and Policy," American Journal of Agricultural Economics 58, 1976, pp. 131-142.

deficient groups; 2) the size of the substitution between goods; and 3) only moderate correlation between relative supply increases and real income effects (see below).

## The Use of MPCs to Identify At-Risk Groups

Researchers have recognized a number of difficulties in establishing calorie requirements that can be applied to a general population. While the current debate about the extent of variation of an individual's calorie requirements and of the possibility of individual homeostatic regulation of requirements is far from being resolved, it is widely established that calorie requirements vary by activity level.[1] For this reason alone, it is difficult to determine the calorie deficiency of individuals from their food intake. A share of the controversy over the amount of malnutrition, then, reflects the normative assessment of the patterns of desired activity as well as requirements for basic biological processes.

As an alternative to using predetermined calorie requirements to indicate caloric deprivation, it has been proposed that dietary adequacy be derived from individual behavior. Since it is known that as income rises, individuals shift their food expenditures toward goods perceived as having higher quality, including embodied services for transportation and preparation, attempts have been made to use

---

[1] For an excellent discussion of some of the controversies around the use of calorie requirements, see G. Beaton, "Energy in Human Nutrition: Perspectives and Problems," W. O. Attwater Memorial Lecture, Nutrition Reviews 41, 1983, pp. 325-340.

responses to changes in income as indicators of nutritional status.[1] Edirisinghe and Poleman proposed that the point at which households begin to purchase calories from costlier but more preferred staples would provide a threshold indicative of household dietary adequacy.[2] Defining this point as the point at which the income elasticity for inferior staples became negative, they found the threshold between 1,250 and 1,900 kilo calories per capita per day in regions of four countries. The lower cutoff was for Indonesia as a whole (although rural Indonesia, comprising 80 percent of the population, was reported to have a threshold of 1,560). The next lowest threshold observed was 1,530 kilo calories in Northeast Brazil. Excluding the Indonesian threshold on the ground that it is inconsistent with the rural data, the thresholds were between 68 and 88 percent of reported calorie requirements.

A more extreme use of thresholds is made by Lipton.[3] His category of the "ultra-poor" has expenditure shares for food that rise as expenditures do. This implies that they have income elasticities for

---

1 Thirty years ago, Bennet stated what is known as Bennet's Law: that the share of starchy staples in the diet declines as income rises. See M. Bennet, The World's Food (New York: Harper, 1954). See also F. D. McCarthy, "Food Consumption, Income Distribution and Quality Effects," Food Policy 2, 1977, pp. 79-82.

2 N. Edirisinghe and T. Poleman, "Behavior Thresholds as Indicators of Perceived Dietary Adequacy and Inadequacy," International Agricultural Economic Study #17, Cornell University, 1983.

3 M. Lipton, Poverty, Undernutrition, and Hunger, World Bank Staff Working Paper 597 (Washington, D.C.: World Bank, 1983); see also V. Rao, "Measurement of Deprivation and Poverty Based on the Proportion Spent on Food: An Exploratory Exercise," World Development 9, 1981, pp. 337-354.

food greater than 1. He suggests that these ultra-poor will spend an average of 75 to 80 percent of their budgets on food, with a small portion of their budgets committed to other basic goods, including housing. By this criterion, 10 to 30 percent of rural and 5 to 20 percent of urban India was found to be ultra-poor. Lipton also reports studies from Nigeria, Brazil, Bangladesh, and Indonesia in which the ratio of food outlays to total outlays was observed to rise as the income of the lowest income groups rose. These families were also below Edirisinghe's and Poleman's threshold in that they were not observed to change the composition of their diets as expenditures rose.

Such threshold analysis poses two problems. Firstly, from an econometric standpoint, it is not clear that a functional form which best fits an entire data set on an OLS or other criterion is always the most accurate estimator of a turning point. Threshold analysis basically looks at second derivatives of a functional form, and there is little evidence about how sensitive threshold estimates are to alternative approaches. More important, there is a conceptual issue about the most appropriate cutoff. If the cutoff point is given in terms of budget shares of food, lower budget shares, say a value of 65 percent in an urban area, may indicate that relatively high expenditures on housing are needed to maintain health and safety. Similarly, when looking at income elasticities, one can argue that a value may be less than 1 merely because housing is further below a minimum (perhaps unknown) basic requirement than is food, rather than because food consumption has reached a point where taste rather than

need predominates. Nevertheless, while such behavioral approaches may underestimate the target group that would be chosen with a criterion of functional consequences of underconsumption, it does identify the portion of the population for which the cost effectiveness of a nutrition-oriented program is probably highest.

## 3. EMPIRICAL PATTERNS OF PARAMETERS AND THEIR RELATIONSHIP TO POLICY: A REVIEW OF PRINCIPAL STUDIES

A number of studies have reported price responses disaggregated by income class (see Table 3.1). The principal ones are discussed below. The studies differ in their countries data and in methodology. Each approach has its advantages and drawbacks, which are briefly discussed. However, as this is not a manual of demand estimation, the main concern is the general patterns of income-specific price responses.

### Application of Frisch Technique

One of the first attempts to identify the price response of low-income consumers for use in food policy analysis was made by Pinstrup-Andersen and colleagues.[1] The econometric technique employed in this series of studies was developed by Frisch in 1959 (see Vol. II for a discussion of methodologies) but apparently not applied to food policy work before these studies. The method allows for the estimation of a

---

[1] P. Pinstrup-Andersen, N. Ruiz de Londoño and E. Hoover, "The Impact of Increasing Food Supply on Human Nutrition: Implications for Commodity Priorities in Agricultural Research and Policy," American Journal of Agricultural Economics 58, 1975, pp. 131-142; also P. Pinstrup-Andersen, "Decision Making in Food and Agricultural Research Policy: The Distribution of Benefits from New Agricultural Technology Among Consumer Income Strata," Agricultural Administration 14, 1977, pp. 13-25; and P. Pinstrup-Andersen and E. Caicedo, "The Potential Impact of Change in Income Development in Food Demand and Human Nutrition," American Journal of Agricultural Economics 60, 1978, pp. 402-412.

Table 3.1 -- Studies Presenting Price Elasticities by Income Group

| Study | Country or Region | Systems Approach | Aggregation | Time Series | Commodities |
|---|---|---|---|---|---|
| 1) Pinstrup-Andersen, et al. (1976)a/ | Cali, Colombia | Frisch | 1479 households | no | 22 goods |
| 2) Rosenberger (1978) | USA | no | 7500 households | no | fluid milk |
| 3) Timmer and Alderman (1979) | Indonesia | no | 1800 cells | 3 rounds | rice, cassava |
| 4) Thomson (1979) | USA | Frisch, also single equations | 5482 households | no | 28 goods |
| 5) Ahmed (1981) | Rural Bangladesh | LES | 234 cells | no | 10 goods |
| 6) Murty and Radhakrishna (1981) | India | modified LES | 288 cells | 1950/1-1970/1 | 9 goods |
| 7) Trairatvorakul (1981) | Thailand | no | 11,450 households | monthly 1 year | 16 goods |
| 8) Dixon (1982) | Java, Indonesia | no | unspecified cells | 3 rounds | rice, cassava, gaplek |
| 9) Bouis (1982) | Philippines | no | unspecified cells | 1974-77 16 qtrs. | rice, maize, wheat |
| 10) Williamson-Gray (1982) | Brazil | no | 343 cells | no | 16 goods |
| 11) Pitt (1983) | Rural Bangladesh | no | 5750 households | 4 qtrs. | 9 goods |
| 12) Strauss (1983) | Sierra Leone | QES | 138 households | no | 6 goods + leisure |

Table 3.1 -- Studies Presenting Price Elasticities by Income Group (Continued)

| Study | Country or Region | Approach | Aggregation | Time Series | Commodities |
|---|---|---|---|---|---|
| 13) Pinstrup-Andersen, et al. (1983) | Urban Sudan | LES | 4930 households | monthly, 1 year | 7 goods |
| 14) Murty (1983) | India | LES | 288 cells | 1950/1-1970/1 | 9 goods |
| 15) Sahn (1984) | Sri Lanka | no | 4753 households | no | 12 goods |

a/ Complete citations are in the text.

complete matrix of own- and cross-price parameters with limited data, primarily cross-section data without price variation. The importance of the study, then, is not so much the econometric results, which can be criticized because of the severe restriction imposed by the methodology, but their application.

Having derived the matrix, Pinstrup-Andersen and colleagues investigated the direct and indirect effects on calorie consumption of a hypothetical shift in the supply curve that might result from agricultural research on 22 commodities. The exercise was carried out with three sets of assumptions about the costs of research and two sets of supply elasticities. By illustration, they found that the commodities given priority in order to meet goals of reducing calorie deficits under the assumptions that equal percentage increases had equal costs and that supply elasticities were zero were rice, sugarcane, maize, and cassava. If equal quantity increases had identical costs, the priorities would shift to oil seeds, sugarcane, maize, and rice. Other crops would be given priority if increasing protein was the objective.

Pinstrup-Andersen also used these parameters to estimate the gains in consumer surplus which would follow a shift in supply. He found, for example, that for the poor the greatest gains when supply elasticities were zero came from increases in rice production, but when they were unity the greatest gains came from increases in the supply of beef. With different commodity bundles consumed by the upper income groups, changes in beef supply was the largest contributor to consumer surplus under both assumptions while rice was, at

best, only the fifth largest. Pinstrup-Andersen and Caicedo used these parameters in yet another context. They estimated the nutritional effects of changes in income distribution under fixed and expanding food supplies. In the former scenario, prices rose more rapidly, virtually negating the nutritional effect of the distribution-mediated income growth. Under the alternative, redistribution led to a large improvement in nutrient intake.

The studies of Pinstrup-Andersen and colleagues were criticized by Brandt and Goodwin who correctly noted limitations of the Frisch methodology.[1] Whether the research priorities or effects of different distribution scenarios would vary qualitatively with different means of estimation has not been directly investigated. These studies, however, indicate the potential contribution to food policy analysis that can come from understanding price responses by selected income groups.

## Estimates Using Cross-Sectional Data

### United States Data

Rosenberger's study was directly influenced by the study of Pinstrup-Andersen et al. and attempted to estimate income class-specific price parameters directly without the restrictive assumptions

---

[1] J. Brandt and J. Goodwin, "The Impact of Increasing Food Supply on Human Nutrition - Implications for Commodity Priorities in Agricultural Research and Policy: Comment," American Journal of Agricultural Economics 62, 1980, pp. 588-591.

of the Frisch methodology.[1] She studied consumption of fluid milk using cross-sectional data from the United States. Since the Federal Milk Order System establishes prices for milk on a regional basis that is unlikely to be correlated with any significant determinant of demand, the cross-sectional prices exhibited variances that were relatively free of identification problems that can occur with other causes of regional price variation. Using a semi-log demand equation and dummy variables in interaction with price terms [log P x income class dummy variables] she estimated the differences between the price parameter of each of 13 income groups and the price parameter of the highest income bracket. With one exception, the price elasticity declined in absolute value from the poorest income group to the wealthiest; from -2.1 to -0.22, with a weighted mean of -0.82. Rosenberger then used such parameters to estimate the effect on milk and calcium consumption that would follow the phased deregulation of the fluid milk market.

Thomson's study was designed to compare estimation methodologies.[2] Using the same data as Rosenberger, she compared elasticities derived from the Frisch methodology with direct single-equation estimates for 20 commodities. Rather than using dummy variables, she ran separate regressions for each of four income classes. For all but two of these commodity groups (baked goods and tea and coffee), the lowest income

---

1 A. Rosenberger, "The Impact of the Federal Milk Marketing Order System on Dairy Product Consumption." Unpublished M.A. Thesis (Cornell University, 1978).

2 A. Thomson, "Nutrition, Food Demand and Policy." Unpublished Ph.D. Thesis (Stanford University, 1982).

stratum exhibited higher absolute values of price parameters. Thomson did not report compensated price elasticities but, in view of the low income elasticities (only those for alcoholic beverages were greater than 0.5) and the high degree of disaggregation that would lead to small budget shares, it is reasonably certain that these patterns indicate that the poor have a greater propensity to substitute between food items.

Indonesia

Similar approaches were used by Timmer and Alderman[1] and Dixon.[2] These studies used a log-log quadratic equation to estimate income terms and interaction terms with dummy variables to estimate price terms. Timmer and Alderman focused on basic staple commodities in order to evaluate the possibility of mitigating the effect of rising rice price on nutrition with policies affecting maize and cassava in Indonesia. Using cell means from regions and income groups in three separate rounds of a household survey, they found that price elasticities for rice and cassava declined in absolute terms as incomes increased, as did cross-price elasticities of cassava with rice prices. Estimates of rice quantity which changed with cassava prices were not significant, nor were similar estimates with maize. Dixon

---

1 P. Timmer and H. Alderman, "Estimating Consumption Parameters for Food Policy Analysis," American Journal of Agricultural Economics 61, 1979, pp. 982-997.

2 J. Dixon, Food Consumption Patterns and Related Demand Parameters in Indonesia: A Review of Available Evidence, Rice Policies in Southeast Asia Project Working Paper #6 (Washington, D.C., IFPRI, 1982).

focused this methodology on Java and added dried cassava and sweet potato to the analysis, but found no clear patterns in the elasticities for these two commodities. This may be due, in part, to the aggregation of consumers and nonconsumers and the wide confidence intervals in the resulting estimates.

Thailand

Trairatvorakul's study resembles the above in that it uses a single-equation methodology (log-log quadratic) and primarily cross-sectional data.[1] He ran separate regressions for income strata and also for subsamples stratified by calorie intakes. His study differs from the others in that it introduced price exogenously using Ministry of Commerce regional monthly data and therefore avoided the possibility of price and income being correlated due to quality effects. For 10 of the 16 commodities, the poor proved more price responsive. The difference was statistically significant for 8 of these commodities. These included nonglutinous rice, beef, chicken, saltwater fish and eggs, but not pork, freshwater fish, sugar, or prepared noodle plates. Trairatvorakul also estimated cross-price effects, finding chicken and pork significant gross substitutes for rice for middle- and upper-income consumers but not for the poorest quartile. Pork and chicken also appeared to be substitutes for each other in the higher income brackets. A number of the cross-price effects estimated were insigni-

---

[1] P. Trairatvorakul, "Food Demand and the Structure of the Thai Food System." Ph.D. Thesis (Harvard University, 1981).

ficant and a few had unexpected signs. This difficulty was also reported by Timmer and Alderman and by Thomson.

Trairatvorakul used these parameter estimates in conjunction with supply elasticities and information on income and land holdings to project the short-run effects on consumption that would follow a change in rice prices. Among his findings was that the net gains to the rural poor would be small, as many purchase rice rather than sell it. For the urban poor, the calculated reduction in calorie consumption in the short run was large.

Brazil

An additional study using a methodology similar to Timmer and Alderman is a study done by Williamson-Gray using 198 cell means from a survey of 55,000 Brazilian families.[1] The price variation in this study was regional; for all income groups the average price paid by the lowest income group in a given subregion was used as the independent variable. As in the studies mentioned above, Williamson-Gray found that the poor were more responsive to prices. Out of 16 commodities or commodity groupings, the compensated and uncompensated price elasticities of the poorest 30 percent were not more negative than that of the middle 50 percent of the population only for dairy products. The pattern is similar with a comparison of the estimates for the middle-income group and the upper income bracket. Dairy price

---

[1] C. Williamson-Gray, "Food Consumption Parameters for Brazil and Their Applications to Food Policy," IFPRI Research Report 32 (Washington, D.C., 1982).

elasticities again are an exception to the basic trend and, in addition, the elasticity for legumes from the higher income group is virtually identical to that of the poor (-0.61, -0.46, -0.63 for the three income groups).

The differences in compensated price elasticities were frequently pronounced. For example, the elasticity for the cereal group was -0.74 for the poor, -0.16 for the middle, and not significantly different from zero for the rich. The corresponding parameters for roots were -1.39, -0.76, -0.23 while there were -0.50, -0.11, -0.10 for meat and fish. Williamson-Gray also estimated cross-price responses. While she observed the general pattern of low significance for many estimates, a number were significant. For example, the uncompensated cross-price elasticity for the quantity of rice with the price of wheat was 1.71, 1.09 and 0.79 for the three groups. Conversely, the elasticities with the quantity of wheat and the price of rice were 1.03, 1.07, and 0.29. Rice and cassava were substitutes while maize was a gross complement with both rice and legumes. The latter were substitutes with the estimates for meat and fish with the price of legumes being 1.10, 0.83, and 0.65.

Williamson-Gray used these results to explore the efficiency of using food substitutes to increase the calorie consumption of the poor. She modeled options with equal fiscal costs of subsidies on wheat products, milk, and rice. The results indicated that subsidies on wheat would actually add to the problem of calorie deficiencies because the cross-price effects were so pronounced for the poor. Subsidies on milk, on the other hand, would lead to net increases in

calorie consumption, but the greatest share of the benefits would accrue to the wealthy. Hence the subsidies are inefficient tools for meeting nutritional goals. Rice subsidies, however, have positive effects mainly for the poor, reflecting the large estimated own-price elasticity for that group.

Bangladesh

Pitt's study of rural Bangladesh differs from the above studies in a number of respects.[1] The price variation in his study was due both to seasonal supply shifts and the spatial variation of the 800 villages surveyed in a country with poor transportation. As the estimates were made using household data with 5,750 observations, Pitt used a methodology designed for data with a significant percentage of non-consumers during the survey. The method predicts market entry and allows for decomposition of response into entry and consumption after entry. Unlike most of the studies above, he did not estimate price parameters for different groups by separating the sample or by introducing dummy variables. Instead the price parameters are of the form $\beta ij = \beta ij^{0} + \beta\gamma ij * \log$ (expenditures). This has the advantage that price parameters vary continuously rather than discretely by income but constrains the change in price parameters to be monotonic. Pitt's form, then, modifies the semi-log function which in its basic presentation has elasticities that vary inversely with quantities. If a good is normal, then, the elasticities would decline in absolute value

---

1 M. Pitt, "Food Preferences and Nutrition in Rural Bangladesh," Review of Economics and Statistics LXV, 1983, pp. 105-114.

as income rose even if $\beta\gamma_{ij}$ were zero. However, he found that for the majority of the 9 commodities, $\beta\gamma_{ij}$ was positive so that own-price responsiveness decreased with income. For example, for the poorest 10th percentile, the uncompensated price elasticity for rice was -1.30 while it was -0.83 for the 75th percentile. The compensated elasticities were -0.57 and -0.26 respectively. Similarly, the pure substitution elasticities for wheat were -0.73 and -0.09, respectively. Only for fish, mustard oil, and onions were the compensated own-price elasticities larger in absolute value for the higher income group than for the poor. Pitt also established cross-price parameters (without symmetry restrictions). Unlike the situation with own-price parameters, no clear pattern with relation to expenditure was noted.

Since Pitt's study covers commodities that make up 95 percent of calorie intake he was able to establish nutrient price elasticities that are the weighted sum of individual commodity elasticities weighted by their share in total nutrient intake. Despite the differences in own-price response for rice mentioned above, cross-price effects make the net uncompensated calorie elasticity for changes in the price of that grain greater in absolute value for the well-off (-0.52 compared to -0.48 for the poor). A different pattern was observed for wheat, with the poor having a 2 percent increase in calory intake with a 10 percent decline in wheat prices compared to the middle classes increase of 0.3 percent in calorie intake. Wheat and potatoes were the only commodities for which the nutrient elasticities of the poor for protein, calories, 2 minerals and 3 vitamins were all negative. For other commodities, the cross-price effects offset the own-price response for at least one nutrient.

Sri Lanka

Sahn's analysis of the 1980/81 household expenditure survey for Sri Lanka includes a technique to account for entry into the market for commodities such as meat, milk, and bread.[1] As there are few families that do not consume such goods as rice, fish, sugar, vegetables, spices, and coconuts, estimates for these goods were obtained using OLS. Price variation was generally spatial, although the survey was conducted in 2 rounds over 7 months. Covariance techniques were used to control for differences between the rural areas, urban areas, and estates.

Of the 13 commodities studied, 8 had significantly smaller absolute price elasticities for the upper income groups. This pattern was particularly pronounced for fish, milk, and pulses. In addition to these 8, the uncompensated own-price elasticity for rice declined slightly with higher incomes. However, as both the budget share of rice and income elasticity declined sharply with income, the compensated substitution elasticity for rice was lower in absolute value for the lower income groups. Furthermore, unlike in many of the other studies, there was no significant differences in the price elasticities for bread or wheat of specific income groups.

Estimates Combining Time Series and Cross-Section Data

Philippines

In the studies reported above, large and significant price

[1] Sahn, D. "Analysis of Food Consumption Patterns and Parameters in Sri Lanka," IFPRI Research Report in preparation.

responses were obtained from a household survey undertaken in less than one year. An appropriate question, then, is whether such patterns can be duplicated using time series data or combinations of time series and cross-sections. A study by Bouis provides some evidence to answer that question.[1] He used quarterly data consisting of 1,000 interviews from over 15 quarters from 1974 to 1977. The estimates, which combined the time series and cross-sectional data, were run separately for 4 income groups in each of 12 regions. The commodities studied were rice, corn grits, and wheat. While the absolute value of the own-price elasticities for rice generally declined as income rose, in some maize-consuming regions the middle-income groups were more responsive to prices than either the poor or the well-off. Such a pattern was observed for both rice and maize.

An additional noteworthy point of Bouis' study is that the measured price responses were frequently large -- in the range of -2 to -1. While such magnitudes are common in cross-sectional studies, Bouis' estimates have much the flavor of time series, with spatial differences minimized by the separate regional regressions.

## India

With the exception of the work of Pinstrup-Andersen and colleagues, the studies discussed to this point employed single-equation techniques with no restrictions derived from demand theory.

---

[1] H. Bouis, "Rice Policy in the Philippines." Unpublished Ph.D. Thesis (Stanford University, 1982).

While such an approach is important for its flexibility, it will be worthwhile to see if the general patterns are also found in studies that use a systems approach to estimation (see Volume II for a discussion of the merits and shortcomings of such an approach). Murty's studies of demand for food in India are a unique attempt to estimate income-specific price parameters using time series data.[1] Both studies estimate separate systems of demand for five urban and five rural income groups using cell means from rounds 2-25 of the Indian National Sample Survey collected between 1950/51 and 1970/71. Each round of this survey contained 12 or 13 cell means. The estimations, then, combine some of the properties of time series and of cross-sectional techniques.

In the study coauthored with Radhakrishna, demand was modeled using a modification of the linear expenditure system (LES) proposed by Nasse,[2] which allows the underlying additivity assumption to be relaxed. Six food groups and three nonfood groups were included. The pattern of declining absolute values of own-price elasticities with rising income was observed in these estimates, with values for the meat, milk, and sugar groups exceeding unity for many income

---

1 K. Murty, "Consumption and Nutrition Patterns of ICRISAT Mandate Crops in India," Economics Program Progress Report #53, International Crops Research Institute for the Semi-Arid Tropics, 1983; and K. Murty and R. Radhakrishna, "Agricultural Prices, Income Distribution and Demand Patterns in a Low Income Country," paper presented at the First International Working Conference on Computer Applications in Food Production and Agricultural Engineering, Havana, Cuba, October, 1981.

2 P. Nasse, "Analyse des Effects de Substitutions dans un System Complet de Functions de Demand," Annals de L'INSEE 5, 1970, pp. 88-100.

groups. The full matrix of price and income elasticities from this study was used to model the effects of income transfer schemes. As in the study by Pinstrup-Andersen and Caicedo, the effects were, to a degree, sensitive to the assumptions of the approach to redistribution but also indicative of the pressures on cereal prices that such a policy would introduce. The study also illustrated the importance of food price stabilization for other sectors, as the real income effects of a cereal price change leads to significant changes in the demand for other goods.

Murty's study, which employed the standard LES, concentrated on the demand for the five cereals that made up the aggregate group in the earlier study. In addition, it included three other food groups and one nonfood group. As in the Bouis' study that used time series and cross-section, the pure substitution direct-price elasticities were frequently between -2.5 and -1. For example, the estimated price responses for the poorest rural expenditure class were -1.39 for superior cereals and -2.29 for sorghum. The corresponding elasticities for the highest rural expenditure bracket were -0.39 and -0.96. Similar patterns were estimated with cross-price terms. Murty then used the own- and cross-price elasticities to calculate the calorie intake elasticities for each group. Since superior cereals include rice and wheat and, therefore, have the largest share of total calories, price changes from that aggregate were most influential in the changes on total calorie intake. As discussed in Volume II, some caution needs to be taken when employing these estimates because the LES excludes negative income elasticities. For example, the estimated

elasticities for sorghum and millet were positive, although the data presented (for 1973/74, outside the actual period of estimation) indicated that those grains were clearly inferior goods for all expenditure groups in urban areas.

Other Estimates Using System of Demand Approach

Two other studies within the set reported here used the LES system with separate estimates for different expenditure groups. Ahmed used cell means from the data set analyzed by Pitt.[1] As both the commodity aggregation and the means of reporting results by income group differ, a direct comparison of results is not possible. Nevertheless, one gets the general impression that the price elasticities in the two studies differ by a greater degree than do the income elasticities. For example, Pitt reports rice and wheat separately and finds that the 75th expenditure percentile had own-price elasticities of -0.83 and -0.06. As the budget share to rice is five times that of wheat, an average grain elasticity weighted for expenditures would be around -0.70. An elasticity weighted for calories or quantities would be approximately -0.65 while the LES own-price estimate for foodgrains for the richest half of the population was -0.38. Similarly, Pitt reported the price elasticity for pulses to be -0.51 compared to the LES estimate for an aggregation of pulses and vegetables of -0.32. Price elasticity for fish and milk from Pitt's study weighted for

---

1 R. Ahmed, "Agricultural Price Policies under Complex Socio-Economic and Natural Constraints: The Case of Bangladesh," IFPRI Research Report 27 (Washington, D.C., 1981).

expenditures would be -0.8 while an elasticity weighted for calories would be -0.5. The LES estimates for an aggregation of meat, fish, milk and eggs was -0.35. With account taken for the fact that Ahmed reports elasticities for the poor as represented by the poorest half of the population and Pitt represents the poor by the 10th percentile, the impression is that the LES expenditures are less price responsive for the poor just as they are for the wealthy. There is no independent means of ascertaining where the "true" elasticities fit between these reports.

The study of price response in urban Sudan used data from 4,930 households to estimate elasticities for 7 commodity groups.[1] Separate LES were estimated for 5 expenditure groups. The own-price elasticity for bread estimated for the poorest 12.5 percent of the population was -0.83. The corresponding elasticity for the wealthiest half of the population was -0.22. On the basis of such estimates, the effect of an uncompensated change in the price of bread (subsidized at the time of the analysis) on nutrition was analyzed. It was predicted that a net reduction of 175 calories in the daily diet of the poor would follow a 50 percent increase in bread prices.

One further use of a system approach is included in this review. Strauss employed a quadratic expenditure system (QES) to study demand

---

[1] P. Pinstrup-Andersen, J. von Braun, T. Uy, and W. Floro, "Impact of Changes in Income and Food Prices on Food Consumption by Low-Income Households in Urban Khartoum with Emphasis on the Effect of Changes in Wheat Bread Prices" (Washington, D.C.: IFPRI, April 1983), Mimeograph.

for 7 commodities in rural Sierra Leone.[1] His study differs from the others reported in two major respects. First, the QES is more flexible than the LES, although it retains the potential advantage of employing restrictions derived from demand theory. Second, as household labor supply is included in the study, the labor-leisure tradeoff was measured.

In this study a single system was estimated from a sample of 138 households. Differences in price elasticities by income group came from the quadratic nature of the system. The variation of the price parameters, then, is continuous with expenditures. Again, one notes compensated own-price elasticities declining with expenditures, although not for all food commodities. They do so for the two largest food groups (in expenditure terms) -- rice and fish and animal products. The uncompensated elasticities for roots and for oil increase slightly with expenditures. In this community, the large budget share of rice leads to negative gross substitution with roots when the rice price increases. Few of the compensated cross-price elasticities for food reported by Strauss are appreciable, as the majority are less than 0.1. Little substitution between goods and leisure was observed; the effect of a wage change was largely the income effect.

---

[1] J. Strauss, "Determinants of Food Consumption in Rural Sierra Leone," Journal of Development Economics 11, 1983, pp. 327-354.

## 4. GENERALIZATIONS OF PATTERNS

In a study published in 1971, Weisskopf presented regression results which he used to argue that structuralist models of development are incorrect in assuming that substitution between food and other aggregate groups is low.[1] A decade later Ray felt the need to test the relationship of budget shares to prices.[2] Similarly, though convinced that consumers are responsive to prices, at least when dealing with disaggregated foods, Timmer found it worthwhile to ask whether substitution elasticities are constant across income groups.[3]

One of the goals of this review is to discover patterns in existing studies that address such broad questions. It may be possible to use those patterns to establish a general rule. To be sure, there will always be exceptions to such rules, but it may be possible to use the general pattern as a reference that can be used to explain the exceptions.

---

1 R. Weisskopf, "Demand Elasticities for a Developing Economy," in Studies in Development Planning, ed. H. Chenery (Harvard University Press, 1971).

2 R. Ray, "The Testing and Estimation of Complete Demand Systems on Household Budget Surveys," European Economic Review 17, 1982, pp. 349-69.

3 C. P. Timmer, "Is There 'Curvature' in the Slutsky Matrix?" Review of Economics and Statistics LXIII, 1981, pp. 395-402.

Planners may not have sufficient data with which to estimate parameters that can be used to fix the precision of projections and to verify or refute other patterns. In such circumstances it may be useful to have general rules of thumb to narrow the uncertainty of any assumptions that must be used.

Patterns in Income Elasticities

Basic guidelines already exist about income elasticities. Engel's law and, to a lesser degree, Bennet's law have achieved the type of veneration given in other subcultures to the wisdom of the elders. Similarly, it is widely, if not universally, acknowledged that income elasticities for food items decline with income. While there may be some questions about the rate at which they decline, any data set sufficient to estimate income elasticities should be sufficient to estimate the decline as well. Finally, if one needs more guidance, broad reviews of Engel elasticities, such as those of Houthakker[1] and Weisskopf are available.

Houthakker's analysis of cross-sectional data confirmed Engel's law in that all elasticities for total food expenditures were less than one, although estimates from poorer countries were close to unity. Houthakker also found that the inclusion of variables for family size reduced the apparent income elasticity.[2] He reports 71

---

1 H. Houthakker, "An International Comparison of Household Expenditure Patterns Commemorating the Centenary of Engel's Law," Econometrica 25, 1957, pp. 532-551.

2 This issue is addressed in greater detail in Volume II.

estimates of elasticities from various samples and subsamples, most of them adjusted for family size. In 11 of those, generally samples from Southeast Asia and Ghana, as well as pre-oil boom Libya, the elasticity exceeded 0.8. In 34 of those, including a number of estimates from different urban regions of the United States in 1950, the elasticity was between 0.6 and 0.8, while only 5 samples yielded estimates below 0.4.

Weisskopf's estimates differed from Houthakker's in that they are from time series data, yet the results are similar. For 8 of the 16 countries he studied, the short-run results exceeded unity, while for another 7 the results exceeded 0.8. The estimates were below 0.4 for only one country, Israel.

These studies indicate that elasticities for food expenditure are appreciable. While the results reported from such studies do not allow for estimates of marginal propensities to consume (MPCs), they are consistent with estimates of MPC ranging from 0.3 to 0.6 in low-income countries, with 0.7 not being unlikely for low income consumers in rural areas.[1]

With the exceptions noted above, the studies also confirm Engel's law. In addition, they confirm a pattern so regular as to be almost a corollary of Engel's law, that elasticities for food, as well as budget shares, decline with income. For all this, regularities in food expenditures provide only rough guidance for nutrition components of food policy analysis.

---

[1] See, for example, P. Hazell and A. Röell, op. cit.

The main reason for this is that elasticities by items or by nutrients are far less regular than are those for total food expenditures. For example, in one of the few published studies that compare calorie elasticities, variations between countries appear to be quite large.[1] In that study, the poor in Sri Lanka were found to have elasticities for calories of only 0.18, while they had elasticities of 0.44 for India and 0.56 for Morocco.

It is, of course, expected that elasticities for calories will be less than those for total food expenditures. People want to increase the quality as well as the quantity of food purchased, especially as their physical capacity for food intake is saturated. The income elasticity for food expenditures can be decomposed to the sum of the income elasticity for calories plus the elasticity of the change in the price of calories (quality) with respect to income.[2] This latter term is surely positive, and should dominate the total term as incomes rise.

An impression of the range and size of calorie and expenditure elasticities can be obtained by looking at Tables 4.1 and 4.2. The elasticities reported were all calculated with the same functional form:

$$\left.\begin{matrix} LKcal \\ LFX \end{matrix}\right\} = a + \beta LTX + \beta_2 (LTX)^2 + \beta_3 Number \times LTX + \beta_4 Rural \times LTX + C_1 Rural + C_2 Number$$

---

1 O. Knudsen and P. Scandizzo, op. cit.

2 $\partial \ln(P*Kcal)/\partial \ln Y = \partial \ln Kcal/\partial \ln Y + \partial \ln P/\partial \ln Y$. If calories can be considered a measure of quantity, the second term on the right can be considered a quality elasticity.

Table 4.1 -- Calorie Elasticities for Low-Income Families*

| Country | Urban | Rural | Regression $R^2$ |
|---|---|---|---|
| Sri Lanka | .41 | .60 | .65 |
| Thailand | .26 | .29 | .28 |
| Egypt | .20 | .34 | .31 |
| Sudan | .30 | .33 | .42 |
| Indonesia | .55 | .61 | .63 |
| Nigeria (Funtua) | - | .94 | .72 |
| Nigeria (Gusau) | - | .94 | .84 |
| Malaysia (Muda) | - | .65 | .76 |
| Brazil** | .10 | .23 | .99 |
| Bangladesh** | .40 | .40 | .99 |
| Morocco** | .54 | .77 | .99 |

* Low income defined as the average income of families which consume 1750-2000 calories per capita per day.

** From aggregate data.

Sources: Sri Lanka: 1981 Labour Force and SocioEconomic Survey, Department of Census and Statistics

Thailand: 1975/76 Household Expenditure Survey, Central Bank

Egypt: 1981/82 Household Budget Survey, IFPRI-Institute of National Planning

Sudan: 1978/79 Household Expenditure Survey, Department of Statistics

Indonesia: 1976 SUSENAS Survey

Nigeria and Malaysia: World Bank Project Monitoring Studies

Brazil: 1974/75 National Household Expenditure Survey, Brazilian Geographical and Statistical Institute

Bangladesh: 1976/77 Household Expenditure Survey, Census Bureau

Morocco: 1971 Household Consumption Survey, Ministry of Planning. Data reported in Knudsen, G. and P. Scandizzo, Nutrition and Food Needs in Developing Countries. World Bank Staff Working Paper # 328, 1979. This analysis differs from the approach reported in that work.

Table 4.2 -- Food Expenditure Elasticities for Low Income Families*

| Country | Urban | Rural | Regression $R^2$ |
|---|---|---|---|
| Sri Lanka | .72 | .86 | .91 |
| Thailand | .62 | .65 | .77 |
| Egypt | .71 | .68 | .75 |
| Sudan | .74 | .84 | .80 |
| Indonesia | .88 | .98 | .87 |
| Nigeria (Funtua) | - | .89 | .95 |
| Nigeria (Gusau) | - | 1.04 | .96 |
| Malaysia (Muda) | - | .88 | .76 |
| Brazil** | .83 | .83 | .99 |
| Bangladesh** | 1.06 | 1.06 | .99 |

* See previous table.

** From aggregate data.

where:

LKcal = the logarithm of calories per capita per day;

LFX = the logarithm of food expenditures per capita per month;

LTX = the logarithm of total expenditures per capita per month;

Number = household size;

Rural = a variable defined as 1 if the household is from a rural area and 0 otherwise.

The elasticities estimated in this manner vary by income. The values reported in Tables 4.1 and 4.2 are calculated at the average per capita monthly income of households with calorie consumption of 1,750-2,000 calories per day. It should be readily apparent that the calorie elasticities are on average less than the food expenditure elasticities. Respective means are 0.48 $\pm$ 0.26 and 0.82 $\pm$ 0.12.[1] Also note that the rural population generally have both higher calorie and food expenditure elasticities than the urban at the incomes that correspond to equal consumption. It is of further interest that estimating equations for calories have lower $R^2$ terms, implying that they have less explanatory power than the food expenditure elasticities. Similarly, the standard deviation is less for food expenditure elasticities, which indicates that such elasticities vary less between countries than calorie elasticities do.

Implicitly, the average quality elasticity is 0.34 $\pm$ 0.22, which indicates that even if calories are consumed less than average requirements, there is a tendency to shift to higher-priced sources

---

[1] The smaller variation of the food expenditure elasticities implies a negative correlation of calorie and quality elasticities.

of calories. This quality elasticity, however, is on the average less than the calorie elasticity.

## Patterns in Own-Price Elasticities

### Magnitude

Theory offers few guidelines about the nature of changes of price parameters as incomes rise. The Slutsky decomposition shows that the total price effect can be decomposed into a pure substitution effect and a real income effect.

(4.1) $E_{ij} = e_{ij} - w_j n_i$

[Price elasticity] = [Substitution elasticity] - [budget share to goods with price change] x [income elasticity of good being measured]

When the income elasticity for this jth good is less than one, the budget share declines with income and, therefore, the real income effect of a price change declines with income, even if the income elasticity is constant over income groups. However, the substitution effect may be, and usually is, large relative to the real income effect. This is especially true with disaggregated commodities. For example, the real income effect exceeded 10 percent of the total price effect in less than one-third of the estimates for specific meat products in the data sets discussed in the previous section, and those were mainly in additive systems. In no case did the real income effect reach 25 percent of the total elasticity. The real income effect was higher for rice, as the budget share for rice was appreciable in a number of samples, particularly Bangladesh and Indonesia.

However, the real income effect for coarse grains was a minor share of the total price effect. As indicated in Table 4.3, the real income effect is virtually always less than 0.25 while most price elasticities greatly exceed that value. The use of the easily derived real income effect, then, would only be useful, if useful at all, for items that are a larger share of the budget, particularly aggregate groups. However, even with those groups it is likely to capture only a small share of the total price effect.

The sizes of the elasticities reported in the Appendix show that the poor are quite price responsive and that most elasticities decline with increased income. Often the absolute value of an elasticity for specific food items exceeds one, indicating that expenditures on the item decline following a price increase. Even the average elasticities for individual commodities across all income group are frequently greater than one (see Table 4.3). These values are perhaps higher than conventional wisdom implies they are. Do they, in fact, measure a response different than what is generally denoted by the term "price elasticity"? Timmer and Alderman suggest that the results measured from cross-sectional data are long-run responses that should exceed the short-run adjustment, perhaps by as much as a factor of 2. Indeed, Ray found price elasticities for food as an aggregate measured in cross-section to be -1.16 for rural India and -1.28 for urban India. The comparable time series results were -0.68 and -0.62.[1] The time series results of Murty and Radhakrisha, Murty, and Bouis,

---

[1] R. Ray, op. cit.

Table 4.3 -- Cross-Tabulation of Price Elasticities and Real Income Effects of Price Changes

| Number Observations with values: | | Price Elasticities (Absolute Value) <.25 | .25-.5 | .5-.75 | .7.5-1 | >1 | Total |
|---|---|---|---|---|---|---|---|
| Real Income Effects | <.25 | 33 | 28 | 37 | 28 | 64 | 190 |
| | .25-.5 | 0 | 0 | 4 | 1 | 6 | 11 |
| | .5-.75 | 0 | 0 | 0 | 3 | 1 | 4 |
| Total | | 33 | 28 | 41 | 32 | 71 | 205 |

however, also estimate some price responses that exceed those conventionally attributed to necessities. Furthermore, the studies of Williamson-Gray and Trairatvorakul, and the time-series studies from India as well, introduce prices exogenously and therefore are not attributing quality responses to price. Similarly, Rosenberger's study has a variable for price which, although derived from household budget surveys, is for a commodity that not only is homogenous but that also has a price determined by an exogenous government policy unlikely to be correlated with a regional price index.

## Relationship of the Elasticities of Poor and Rich Consumers

Timmer seeks a regular pattern in the change of compensated price elasticities with income, noting that his approach is pragmatic rather than theoretical.[1] It is sometimes argued that poor people, having fewer items in their budgets, have less opportunities for substitution, hence the absolute value of $e_{ij}$ should be smaller for the group.[2] On the other hand, one could find virtually infinite substitution between two inexpensive calorie sources if consumption by the poor resembled demand for commodity characteristics in the manner of programming models for least-cost diets. Almost all of the studies included in Table 1 include at least one inexpensive source of calories that can be a substitute for the main grain of that culture. Furthermore, it is

---

1 Timmer, 1981, op. cit.

2 L. Jackson, "Hierarchic Demand and the Engel Curve for Variety," Review of Economics and Statistics LXVI, 1984, pp. 8-15. Note, however, that some inferior goods may be consumed only by the poor.

possible, in principle, to disaggregate by quality so that substitution of different grades of a commodity can be studied, although not infrequently the prices of the different grades are highly collinear. While neither the magnitude nor the sign of the derivative of the substitution parameter with respect to income can, therefore, be conclusively predicted by theory, the functional form of an estimate may inadvertently constrain the relationship.

In Timmer's study the change of the absolute value at $e_{ij}$ with respect to income is explored by regressing values of $|e_{ij}|$ estimated in flexible functions (specifically, from the Rosenberger and Timmer-Alderman studies above) on income,

$$|e_{ij}| = b_{ij} + c \log\text{Income}. \tag{4.2}$$

The $b_{ij}$ terms for separate commodities were not significant, while the coefficient c had a value of -0.339 with a t-statistic of 8.3. It is useful to explore this relationship using a larger data set and with separate commodity groups. In this exercise it is important to exclude those studies in which the price parameters are generated from the Frisch or other additive systems, as the additivity assumption severely restricts the price parameter. Table 4.4 presents different c parameters from versions of equation (4.2) estimated separately for five commodity groups.

The fits of these equations are poor, reflecting in part the heterogeneity of the data, categories and methods of estimation, and the influence of some extreme values for meat in Thailand and rice in Brazil. In order to deal with the latter outrider, the rice equation was re-estimated. A dummy variable for aggregate commodities using

Table 4.4 -- Parameters of Regression of Hicksian Elasticities on Log of Income

| Commodity Groups | Meat | Rice | Wheat and Coarse Grains | Root Crops | Milk |
|---|---|---|---|---|---|
| Coefficient of LnIncome | -0.55 | -.08 | -.25 | -.25 | -.21 |
| t-Statistic | 1.66 | 0.60 | 2.10 | 1.67 | 4.14 |
| Regression $R^2$ | .14 | .01 | .24 | .15 | .38 |
| Degrees of Freedom | 17 | 38 | 14 | 15 | 28 |

estimates for aggregate cereals from Brazil was also included. As in a few other studies, rice was grouped with wheat and other grains. It was expected that the more aggregated a group, the lower the estimated elasticity. This is because substitutions within a group are masked. The equation was

$$(4.3) \qquad |e_{ii}| = 1.26 \underset{(4.29)}{-0.20} \text{ LnIncome } \underset{(3.62)}{-0.38} \text{ aggregate}; \quad R^2 = 0.46$$

Equation (4.3) indicates that with the modifications the estimates conform to the pattern observed with coarse grains, roots, and milk. The variable for aggregation included in the meat equation had the expected negative sign but was not significant. If the Thai data points are removed, the $R^2$ values improves to 0.4 and the aggregation coefficient becomes significant. The coefficient of log at income drops to -0.22 and the t-value rises to 1.86, which is significant only at the 10 percent level.

Considering these modifications, one notes that the income coefficients are generally between -0.2 and -0.25. Equation (4.4) generalizes equation (4.2) for all commodities,

$$(4.4) \qquad |e_{ii}| = 1.36 \underset{(4.31)}{-.17} \text{ LnIncome}; \qquad R^2 = 0.14.$$

This, then, supports Timmer's view of a generalizable pattern in compensated (Hicksian) elasticities, but the coefficient is somewhat smaller than he observed.[1] This is not an artifact of the adjustment of income for purchasing power parity. That adjustment was made in the

---

[1] Note that the coefficient in the aggregate equation is lower than those of individual commodities due to the absence of intercept terms for specific commodities.

current study but not by Timmer. If equation (4.4) is run using income in dollar equivalents, the coefficient of LnIncome is -0.15.

With price elasticities generally between 0.5 and 1.5, this equation implies that a 10 percent increase in income will lead to a 1 to 3 percent decline in the compensated price elasticity; the larger decline coming from the smaller elasticity.

The results in equation (4.4) as well as Table 4.4 are important for hypothesis testing as they reconfirm that there is a significant curvature in the Slutsky matrix for food commodities. To be sure, there are exceptions, as indicated by Bouis' results for Mindanao, by Strauss' estimates for root crops, and by Pitt's estimates for fish. Nevertheless, the general empirical pattern reported here should serve as a standard against which individual cases can be tested and, therefore, it improves upon a priori notions based on the assumption that the size of the substitution effect increases as the number of substitutes in a budget increases with income. The relationship in equation (4.4) has two limitations, however. Firstly, it offers no insight into the reasons for such a pattern and, therefore, no guidance about when exceptions can be expected. Secondly, although the relationship between income and the elasticities is statistically significant, the estimating equation is heterogenous and has low predictive power.

This second difficulty can be readily addressed. As mentioned above, one motive for this research is to provide guidelines for the use of policy parameters when data or time limitations rule out estimates of income elasticities for specific groups. Tables 4.5, 4.6, and 4.7 present three variations on regressions using the ratio of a

Table 4.5 -- Regressions of Standardized Substitution Elasticities on Log of Income

| Commodity | Intercept | Coefficient of LnIncome | $R^2$ |
|---|---|---|---|
| Meat (excluding Thai data) | 1.32 | -.106<br>(1.43) | .15 |
| Meat (including Thai data) | 2.48 | -.380<br>(1.42) | .11 |
| Rice | 1.25 | -.137<br>(1.38) | .05 |
| Wheat and Other Grains | 1.48 | -.126<br>(1.38) | .12 |
| Root Crops | 1.76 | -.258<br>(4.34) | .54 |
| Milk | 2.04 | -.243<br>(3.83) | .35 |

Table 4.6 -- Regressions of Standardized Substitution Elasticities on Income Share

| Commodity | Intercept | Coefficient of Income Share | $R^2$ |
|---|---|---|---|
| Meat (excluding Thai data) | 1.26 | -.265 (4.16) | .59 |
| Meat (including Thai data) | 1.82 | -.579 (1.80) | .16 |
| Rice | 1.10 | -.252 (2.26) | .12 |
| Wheat and Other Grains | 1.25 | -.214 (2.05) | .23 |
| Root Crops | 1.30 | -.256 (3.97) | .50 |
| Milk | 1.48 | -.356 (4.04) | .38 |

Table 4.7 -- Regressions of Standardized Substitution Elasticities on Population Percentile

| Commodity | Intercept | Coefficient of Percentile | $R^2$ |
|---|---|---|---|
| Meat (excluding Thai data) | 1.34 | -.078 (3.93) | .56 |
| Meat (including Thai data) | 2.21 | -.213 (2.56) | .28 |
| Rice | 1.29 | -.096 (3.35) | .23 |
| Wheat and Other Grains | 1.42 | -.081 (2.51) | .31 |
| Root Crops | 1.34 | -.063 (2.93) | .35 |
| Milk | 1.68 | -.117 (5.40) | .52 |

compensated elasticity for a specific group to the mean elasticity for the population as the dependent variable, $e_{stand}$ = compensated elasticity, of group/average elasticity.

With such regressions it may be possible to construct elasticities for specific groups that are consistent with estimated aggregate price elasticities obtained from time series. Table 4.5 reports results similar to Table 4.4. The standardization poses particular problems with nonsignificant price responses (as in the Thai meat regressions), as such values have been set at zero. This both exaggerates the differences between the upper income-group elasticities and the mean value and, by depressing the mean value aggregated from various groups, increases the ratio for the lower income groups. This contributes to the higher observed coefficients when the Thai data is included as well as the lower $R^2$ values.

Table 4.6 gives results from regressions of $e_{stand}$ on group income/average income per capita. These again indicate that as income goes up relative to the population mean, the elasticity falls. Note that as both incomes and elasticities are standardized by country mean values, there is little effect remaining for a specific country; the regressions are less cross-country regressions than within-country regressions. Table 4.6 shows improvement in the $R^2$s of all but one commodity regression. The fit of the figures in Table 4.7 is, in general, a further improvement. It shows a regression of $e_{stand}$ on a population arrayed by income (the poorest decile being 1 and richest decile being 10). As examples of how these equations can be used, suppose that time series estimation indicates that the countrywide

compensated-price elasticity for wheat is -0.5. If the mean income is $300, then the elasticity of a household with an income of $100 would be -0.59;[1] while a family with $500 would have an elasticity of -0.45. By the equation in Table 4.7, the elasticity of the household in the 1st decile of income would be -0.67. In this example, there is no guarantee that the calculations from Table 4.6 and Table 4.7 will be identical at any given income. Indeed, while there are some income distribution patterns that may generate identical estimates, such a result would be coincidental. The results in neither Table 4.6 nor 4.7 will predict a sign change with the range of data used here, although this is possible with the former and, considering the maximum allowable value for percentile, it is not possible with the latter. To a degree the formulation in Table 4.6 is the more logical in that it deals with relative purchasing power rather than a ranking according to purchasing power. However, the equations in Table 4.7 fit slightly better and seem adequate in guiding first approximations for policy projections.

Recall that these equations indicate changes in the pure substitution effect. For policy use uncompensated elasticities are usually required, and it is necessary to add the real income effect to these substitution parameters. Such information is frequently available. If, however, even this information is unavailable, the relationships in Tables 4.8 and 4.9 can be used. These tables parallel the previous two tables but use a standardized uncompensated elasticity as the

---

[1] That is, -0.5(1.25-0.214 x 0.33).

Table 4.8 -- Regressions of Standardized Uncompensated Elasticities on Population Percentiles

| Commodity | Intercept | Coefficient of Percentile | $R^2$ |
|---|---|---|---|
| Meat (excluding Thai data) | 1.43 | -.083<br>(3.82) | .55 |
| Rice | 2.06 | -.176<br>(6.63) | .54 |
| Wheat and Other Grains | 1.47 | -.085<br>(2.68) | .34 |
| Root Crops | 1.38 | -.064<br>(2.79) | .33 |
| Milk | 1.71 | -.119<br>(5.47) | .52 |

Table 4.9 -- Regressions of Standardized Uncompensated Elasticities on Income Share

| Commodity | Intercept | Coefficient of Income Share | $R^2$ |
|---|---|---|---|
| Meat (excluding Thai data) | 1.34 | -.288 (4.14) | .58 |
| Rice | 1.74 | -.502 (4.26) | .32 |
| Wheat and other Grains | 1.29 | -.229 (2.20) | .26 |
| Root Crops | 1.34 | -.260 (3.77) | .47 |
| Milk | 1.51 | -.363 (4.08) | .38 |

dependent variable ($E_{stand}$ = group uncompensated elasticity/average uncompensated elasticity).

Since the budget shares of most commodities other than grains are small, it is not surprising that there is little difference between the results in the two sets of tables. Furthermore, since the budget shares of most foods decrease with rising income, as do most income elasticities, it can also be expected that the slopes of the regressions for uncompensated elasticities will generally be steeper than their counterparts.[1] When, however, income elasticities and budget shares are known with some accuracy, it seems preferable to use Tables 4.6 and 4.7 to fix the unknown component.

Timmer mentions a possible theoretical explanation for the pattern observed in his study. If one assumes additivity and also that the marginal utility of money is roughly constant, then the ratio of the change of the income elasticity that comes with a change of income to the change of the price elasticity that comes with a change in income will give a measure of Frisch's money flexibility, the absolute value of which, according to Brown and Deaton, centers around 2.[2] Although these assumptions are strong, Timmer's results were consistent with such a value. The results here, however, have a lower but still plausible value for money flexibility. Regressing $\eta$ on

---

[1] The income elasticities themselves decline with income in a manner that is well known. If standardized income elasticities are regressed on population shares, the coefficients of population analogous to those in 4.7 and 4.9 would be -0.11 for meat, -0.22 for rice, -0.11 for wheat, -0.22 for coarse grains, -0.50 for roots and -0.12 for milk.

[2] J. Brown and A. Deaton. "Models of Consumer Behavior: A Survey." Economic Journal 82 (1972) pp. 1145-1236.

LogIncome for those commodities for which $\eta$ was estimated by subpopulation and, therefore, were not functionally constrained, gave a value of -0.27 ($t = 6.78$) for $\partial\eta/\partial$LnIncome. The ratio of the coefficient and that from equation (4.4) implies an absolute value for money flexibility of 1.58.[1] Such a relationship is implicitly behind many systems of demand estimates, primarily the Frisch methodology and the LES, and suffers from the same limitations that they do (see Volume II).

In an attempt to explain the curvature in terms of behavior, Bouis has hypothesized that the marginal utility of a food item can be decomposed into marginal utilities from increased calories, diet variety, and taste.[2] It is possible with equal generality to substitute marginal utility from nutrient content or satiety for the first of these. The second is intriguing because it is not a property of the good itself but of its relation to other goods; the more the diet is concentrated on one good, the higher the marginal utility from variety obtained from increases in other goods. Taste is a subjective element common to most, if not all, preference orderings. If food intake is low, the marginal utility of satiety is high relative to the other components. Consequently, the marginal rate of substitution between items of similar nutrient composition (two cereals, for

---

1 A similar estimate can be obtained by regressing $\eta$ on $|e_{ii}|$ for those studies in which both parameters were estimated in a flexible manner. This gives a value of 1.42 for the money flexibility. The $R^2$ of the relationship was 0.42.

2 H. Bouis, "If There is Curvature In the Slutsky Matrix, What Do the Curves Look Like?" (Washington, D.C.: IFPRI, 1983), Mimeograph.

instance) would be high. As the diet concentrates on a few inexpensive sources of nutrients, the utility of variety for other items increases and diet diversity takes a larger role. Furthermore, as diets become varied and the marginal utility of nutrients declines, the taste and status properties of various goods dominate the marginal utility, and less substitution between items may take place. The poor purchase 'food,' while the rich purchase 'particular foods.' Bouis presents examples that indicate that the pattern in the curvature of substitution parameters is indeterminate and not either monotonically increasing or decreasing. Much of his exploration centers on substitution between and within broad groups of foods and has not yet been fully developed for empirically testable hyphotheses. One implication can be checked on existing studies, however. Consistent with other models, the approach states that the marginal utility of calories dominates the total marginal utility of food items for low-income consumers, while variety and taste dominate for those with higher incomes. As nonfood items contribute to the latter (and possibly, also, to variety) but not to satiety, one would expect larger substitution within food groups but little between foods and non-foods at low incomes.

While few of the studies included include nonfood items in a nonadditive manner, the net nutrient elasticities can be calculated from some of the studies. If one presumes that the subjective quality of satiety is determined most by calories, one can focus on them and subsume the other nutrients under taste and variety. To conduct this inquiry, either a complete matrix of cross-price elasticities or a

direct estimate of $\partial Kcal/\partial P_i$ is needed.[1] So far, this discussion and analysis has not focused on cross-price terms, as few have been estimated satisfactorily. However, those reported are generally consistent with the pattern of declining absolute values with rising incomes.

Timmer reports compensated price elasticities for calories from rice, cassava and shelled corn in Indonesia.[2] These were -0.023 for the poor and -0.648, -0.643 and -0.583 for the other income groups. Pitt found small compensated changes in total calories with an increase of any food price despite appreciable substitution between items. This occurs for both income groups reported. Similarly, the compensated change in calories with a change in the price of superior cereals, estimated for rural India by Murty using an LES, is negligible for the lowest rural income groups and only -0.08 for the rich. It is about -0.1 for those in urban areas. The changes following a change in the price of other grains would also be small.

Such a result, however, is not universally reported. As Trairatvorakul points out, with rice dominating total calorie intake in Thailand, substitution for rice leads to substitution of calorie sources for other goods. Although it is not indicated in his study, these are most likely to be nonfood items. This was also observed in Sri Lanka. In Brazil, a 10 percent increase in rice prices would lead

---

1 Sahn, op. cit., reports results which indicate no differences in the two approaches. There is not, however, enough evidence to generalize on the relative efficiencies of the two approaches.

2 Timmer (1981), op. cit.

to a net decline of 4.5 percent of calories for the poorest 30 percent of the population and a decrease of 2 percent for the middle 50 percent of the population. Rice provided 19.6 percent of daily calories in Brazil in 1974/75. Changes in the price of cassava (15 percent of calories), sugar (13.5 percent of calories) or bread (8.3 percent) had no measurable effect on net calorie intake with direct estimates of total calorie intake on various prices. Using the matrix of cross-price elasticities provided by Williamson-Gray, one is able to make a similar calculation for the net effect of a change in the price of cereals.[1] Calculations done in this manner indicate that the compensated substitution elasticities for cereal prices are -0.4 for the poor and -0.16 for the middle-income group. Changes in calorie intakes following changes in the prices of other goods are negligible. Unlike the Thailand example, however, the large change in calories following a rice or cereal price change cannot be attributed to one commodity that dominates caloric consumption. Rather, cereals and both edible oils and sugar are highly complementary. A rise in the price of cereals would lead to reductions in the quantities of oil and sugar purchased. This effect, which declines in absolute value with increases in income, reinforces the direction of the change in the amount of calories in cereal consumption following a change in cereal prices.

---

[1] Williamson-Gray, op. cit., Tables 43-45.

## Conclusion

To summarize, then, while there is a fair amount of evidence indicating that absolute values for compensated own- and cross-price elasticities decline with increased incomes, there is no general pattern showing that the net compensated calorie elasticity following a change in any commodity price is negligible for the poor. To be sure, it is low in many cases and in others the general curvature is weak,[1] but there are other cases in which total calories purchased by the poor does decline with a price increase. This is also indicated by the number of examples that show large compensated elasticities for cereals taken as an aggregate. Again, such a case indicates that even the poor in many communities substitute calories for other attributes of foods (and even nonfoods).

Were this not true, one could estimate the total effect on calorie consumption with estimates of income elasticities and budget shares. However, as discussed above, this would underestimate the nutritional effect in a sufficiently large number of cases to reinforce the importance of estimating a matrix of price and cross-price

---

[1] This is indicated by a regression of the form of eq. 4.8 which includes a dummy variable which is one for commodities which are greater than .25 of budget (presumably also the bulk of calories) and an interaction term to shift the slope of LnIncome. The regression is

$$|e_{ii}| = 1.53 - \underset{(4.57)}{0.21}\ \text{LnIncome} - \underset{(1.75)}{0.55}\ \text{dummy} + \underset{(1.62)}{0.14}\ \text{dummy} * \text{LnIncome} \qquad r^2 = 0.16 \quad \text{d.f.} = 110$$

This indicates that for foods that comprise a large part of the budget, the absolute values of compensated elasticities are smaller and change less with changes in income.

elasticities for the poor. This could either be by direct estimates or by using the approximations discussed earlier based on the sample of the ratio of specific group elasticities to market elasticities with rising incomes.

On a list of unanswered questions is why the poor have a non-zero compensated substitution elasticity for calories. This question is analogous to the question of why the diet of even the poor does not conform with least-cost diets based on nutrient costs. While it is recognized that all consumers make budgetary allocations based on subjective notions of personal welfare that might be different from a planner's narrow concerns, it is plausible that at some low level of food intake, there will be virtually no amount of nonfood that would substitute for food. It has already been discussed that the inflection point in Engel curves may be a useful tool for showing who the ultra-poor are. Similarly, it may be useful to use inflection points for compensated price elasticities of calories as an indicator of groups for which quality does not substitute for quantity.

Another task that could aid food policy analysis is verification of whether direct estimates of $\partial Kcal/\partial P_j$ differ from direct of estimates of $\Sigma\ w_i \partial Q_i/\partial P_j$. In many cases it may be easier to estimate the former than the latter.

The difficulty in estimating cross-price terms, either in the aggregate or by groups, is basically a problem of econometrics and of data scarcity. Issues related to estimation techniques are discussed in the second volume of this report. Other research tasks are discussed in its conclusion.

Recommendations for Application of Parameters

Ideally, one would like to derive appropriate parameters designed to measure specifically responses to policy changes being considered. However, data is frequently not available and the time and expense of collection and analysis is, in many cases, prohibitive. Food expenditure changes following changes in income occur regularly enough that rules of thumb are probably sufficient. Marginal propensities of the poor to spend on food are in the range of 0.5-0.7, and elasticities are about 0.8. Calorie elasticities are approximately half that for the poor and are probably higher in rural areas.

Responses to price changes are, however, less regular. Price elasticities for meat and dairy products are frequently above one. There is evidence that in a number of cases, the price elasticities of staples are also around one. As such values are not universal, however, it is advisable in the absence of other data sources to estimate national or regional price responses from time-series data and to estimate group-specific values using Table 4.7 and estimates of income parameters and budget shares. If the latter are not available, a general relationship following Table 4.9 can be employed. These methods are preferable to approximations based on real income effects alone or from an assumption that the poor respond much as the general population does.

Attempts should be made to obtain cross-price parameters for major commodities for food policy analysis. Substitution between food commodities, including between major staples and coarse grains or root

crops, is likely and will generally moderate the total effect of a price change. However, unless there is a major food item which has a negative income elasticity, effects of income changes will generally reinforce each other rather than moderate a response.

Changes in both income and price policies can have important effects on nutrition. While price elasticities net of substitutions may be moderate, price changes following changes in subsidy or exchange rate policies may be greater in percentage terms than plausible income transfers or growth in the short-run. The consequences and potential to influence the effect on welfare can be best understood using price and income effect, disaggregated by commodities and income groups.

# APPENDIX

## Table 1 -- Meat Products Elasticities

| Country | Study Code* | Income | Income Elasticity | $\lvert E_{ii} \rvert$ | $\lvert e_{ii} \rvert$ |
|---|---|---|---|---|---|
| Colombia | 1 | 10.374 | 1.52 | 1.46 | 1.30 |
| | 1 | 18.564 | 1.35 | 1.30 | 1.16 |
| | 1 | 27.339 | 0.99 | 0.99 | 0.87 |
| | 1 | 53.196 | 0.67 | 0.69 | 0.63 |
| | 1 | 122.109 | 0.47 | 0.50 | 0.45 |
| | 1 | 10.374 | 1.94 | 1.89 | 1.83 |
| | 1 | 18.564 | 1.65 | 1.61 | 1.55 |
| | 1 | 27.339 | 1.12 | 1.12 | 1.08 |
| | 1 | 53.196 | 0.82 | 0.82 | 0.79 |
| | 1 | 122.109 | 0.69 | 0.70 | 0.68 |
| United States | 4 | 36.853 | 0.20 | 0.15 | 0.15 |
| | 4 | 100.832 | 0.21 | 0.10 | 0.10 |
| | 4 | 159.057 | 0.25 | 0.11 | 0.11 |
| | 4 | 292.632 | 0.20 | 0.07 | 0.07 |
| Bangladesh | 6 | 11.866 | 1.51 | 0.65 | 0.55 |
| | 6 | 20.876 | 1.14 | 0.48 | 0.37 |
| | 10 | 5.100 | 0.50 | 0.66 | 0.62 |
| | 10 | 25.500 | 1.02 | 0.97 | 0.90 |
| Thailand | 7 | 13.300 | 0.70 | 2.21 | 2.17 |
| | 7 | 28.000 | 0.58 | 0.00 | 0.00 |
| | 7 | 54.040 | 0.40 | 0.00 | 0.00 |
| | 7 | 13.300 | 0.42 | 7.18 | 7.16 |
| | 7 | 28.000 | 0.48 | 0.00 | 0.00 |
| | 7 | 54.040 | 0.23 | 2.25 | 2.24 |
| Brazil | 9 | 24.852 | 1.45 | 2.35 | 2.25 |
| | 9 | 53.656 | 0.74 | 1.29 | 1.25 |
| | 9 | 245.784 | 0.15 | 0.82 | 0.82 |
| Sierra Leone | 11 | 8.865 | 1.00 | 1.29 | 1.16 |
| | 11 | 14.400 | 0.92 | 0.92 | 0.81 |
| | 11 | 22.230 | 0.64 | 0.81 | 0.74 |
| Sudan | 13 | 15.479 | 1.03 | 1.04 | 0.90 |
| | 13 | 19.665 | 0.97 | 0.92 | 0.79 |
| | 13 | 28.405 | 0.21 | 0.34 | 0.32 |
| Sri Lanka | 15 | 13.508 | 2.17 | 1.68 | 1.58 |
| | 15 | 70.422 | .50 | .35 | .32 |

*Codes are the same as in Table 3.1.
Insignificant results are reported as zero.
Income adjusted for purchasing power parity.

# Appendix

## Table 2 -- Rice Elasticities

| Country | Study Code | Income | Income Elasticity | $\lvert E_{ii} \rvert$ | $\lvert e_{ii} \rvert$ |
|---|---|---|---|---|---|
| Colombia | 1 | 10.374 | 0.42 | 0.43 | 0.40 |
| | 1 | 18.564 | 0.39 | 0.40 | 0.38 |
| | 1 | 27.339 | 0.39 | 0.40 | 0.39 |
| | 1 | 53.196 | 0.26 | 0.26 | 0.25 |
| | 1 | 122.109 | 0.19 | 0.19 | 0.19 |
| Indonesia | 3 | 5.148 | 1.16 | 1.92 | 1.49 |
| | 3 | 8.352 | 0.92 | 1.47 | 1.12 |
| | 3 | 13.068 | 0.68 | 1.16 | 0.94 |
| | 3 | 30.222 | 0.28 | 0.74 | 0.68 |
| | 5 | 14.550 | 0.33 | 0.31 | 0.19 |
| | 5 | 36.420 | 0.11 | 0.56 | 0.53 |
| | 5 | 93.480 | -0.12 | 0.00 | 0.00 |
| | 5 | 8.944 | 0.83 | 1.28 | 1.03 |
| | 5 | 19.856 | 0.48 | 0.45 | 0.35 |
| | 5 | 44.688 | 0.13 | 0.18 | 0.16 |
| United States | 4 | 36.853 | 0.23 | 0.84 | 0.84 |
| | 4 | 100.832 | 0.03 | 0.09 | 0.09 |
| | 4 | 159.057 | 0.05 | 0.04 | 0.03 |
| | 4 | 292.632 | 0.02 | 0.03 | 0.03 |
| Bangladesh | 6 | 11.866 | 0.87 | 0.62 | 0.20 |
| | 6 | 20.876 | 0.57 | 0.38 | 0.16 |
| | 10 | 5.100 | 1.19 | 1.30 | 0.57 |
| | 10 | 20.876 | 0.94 | 0.83 | 0.26 |
| Thailand | 7 | 13.300 | 0.41 | 0.74 | 0.64 |
| | 7 | 28.000 | 0.08 | 0.71 | 0.70 |
| | 7 | 54.040 | 0.00 | 0.46 | 0.46 |
| Philippines | 8 | 6.096 | 0.10 | 0.73 | 0.11 |
| | 8 | 14.064 | 0.10 | 0.69 | 0.67 |
| | 8 | 22.656 | 0.20 | 0.68 | 0.64 |
| | 8 | 53.184 | 0.00 | 0.40 | 0.40 |
| Brazil | 9 | 24.852 | 0.85 | 4.31 | 4.25 |
| | 9 | 53.656 | 0.30 | 2.95 | 2.94 |
| | 9 | 245.784 | -0.22 | 1.15 | 1.15 |
| Sierra Leone | 11 | 8.865 | 0.88 | 2.16 | 1.93 |
| | 11 | 14.400 | 0.67 | 0.78 | 0.62 |
| | 11 | 22.230 | 0.08 | 0.45 | 0.43 |

Appendix

Table 2 -- Rice Elasticities
(Continued)

| Country | Study Code | Income | Income Elasticity | $\lvert E_{ii} \rvert$ | $\lvert e_{ii} \rvert$ |
|---|---|---|---|---|---|
| India | 12 | 5.680 | 1.05 | 1.39 | 1.05 |
| | 12 | 9.700 | 0.95 | 1.27 | 0.95 |
| | 12 | 14.780 | 0.81 | 0.69 | 0.42 |
| | 12 | 25.320 | 0.47 | 0.48 | 0.35 |
| | 12 | 65.060 | 0.34 | 0.39 | 0.32 |
| | 12 | 4.966 | 1.07 | 1.23 | 0.86 |
| | 12 | 9.048 | 0.84 | 1.14 | 0.87 |
| | 12 | 14.820 | 0.48 | 0.66 | 0.51 |
| | 12 | 25.428 | 0.20 | 0.29 | 0.24 |
| | 12 | 64.116 | 0.13 | 0.21 | 0.19 |
| | 14 | 5.680 | 0.97 | 0.92 | 0.33 |
| | 14 | 9.700 | 0.89 | 0.78 | 0.28 |
| | 14 | 14.780 | 0.77 | 0.54 | 0.15 |
| | 14 | 25.320 | 0.44 | 0.20 | 0.03 |
| | 14 | 65.060 | 0.32 | 0.33 | 0.25 |
| | 14 | 4.966 | 0.96 | 0.89 | 0.41 |
| | 14 | 9.048 | 0.75 | 0.73 | 0.39 |
| | 14 | 14.820 | 0.49 | 0.36 | 0.17 |
| | 14 | 25.428 | 0.14 | 0.15 | 0.11 |
| | 14 | 64.116 | 0.15 | 0.18 | 0.16 |
| Sri Lanka | 15 | 12.508 | 0.68 | 0.70 | 0.49 |
| | 15 | 70.420 | 0.15 | 0.67 | 0.65 |

# Appendix

## Table 3 -- Wheat and Bread Elasticities

| Country | Study Code | Income | Income Elasticity | $\lvert E_{ii} \rvert$ | $\lvert e_{ii} \rvert$ |
|---|---|---|---|---|---|
| United States | 4 | 36.853 | 0.10 | 0.16 | 0.16 |
| | 4 | 100.832 | 0.03 | 0.17 | 0.17 |
| | 4 | 159.057 | 0.05 | 0.27 | 0.27 |
| | 4 | 292.632 | 0.04 | 0.22 | 0.22 |
| Brazil | 9 | 24.852 | 1.09 | 1.96 | 1.91 |
| | 9 | 53.656 | 1.23 | 0.84 | 0.81 |
| | 9 | 245.784 | 1.45 | 0.73 | 0.73 |
| Bangladesh | 10 | 5.100 | -0.10 | 0.72 | 0.73 |
| | 10 | 25.500 | -0.24 | 0.06 | 0.06 |
| Sudan | 13 | 15.479 | 0.89 | 0.91 | 0.84 |
| | 13 | 19.665 | 0.78 | 0.74 | 0.69 |
| | 13 | 28.405 | 0.14 | 0.22 | 0.21 |
| Sri Lanka | 15 | 12.508 | 0.77 | 0.85 | 0.83 |
| | 15 | 70.420 | 0.09 | 0.85 | 0.85 |

# Appendix

## Table 4 -- Coarse Grain Elasticities

| Country | Study Code | Income | Income Elasticity | $\lvert E_{ii} \rvert$ | $\lvert e_{ii} \rvert$ |
|---|---|---|---|---|---|
| Colombia | 1 | 10.374 | 0.63 | 0.63 | 0.60 |
| | 1 | 18.564 | 0.54 | 0.55 | 0.53 |
| | 1 | 27.339 | 0.44 | 0.44 | 0.43 |
| | 1 | 53.196 | -0.26 | 0.00 | 0.00 |
| | 1 | 122.109 | -0.43 | 0.00 | 0.00 |
| Philippines | 8 | 6.096 | -0.30 | 1.37 | 1.40 |
| | 8 | 14.064 | -0.56 | 1.53 | 1.56 |
| | 8 | 22.656 | -0.50 | 1.47 | 1.49 |
| | 8 | 53.184 | -0.10 | 1.32 | 1.32 |
| Brazil | 9 | 24.852 | 0.33 | 1.77 | 1.77 |
| | 9 | 53.656 | 0.15 | 1.09 | 1.09 |
| | 9 | 245.784 | 0.04 | 0.58 | 0.58 |
| India | 12 | 5.680 | 1.25 | 2.29 | 2.19 |
| | 12 | 9.700 | 1.03 | 1.94 | 1.86 |
| | 12 | 14.780 | 0.40 | 0.68 | 0.66 |
| | 12 | 25.320 | 0.25 | 0.57 | 0.56 |
| | 12 | 65.060 | 0.36 | 0.96 | 0.95 |
| | 12 | 5.680 | 1.08 | 2.18 | 2.15 |
| | 12 | 9.700 | 1.22 | 2.44 | 2.41 |
| | 12 | 14.780 | 0.43 | 0.74 | 0.73 |
| | 12 | 25.320 | 0.59 | 1.28 | 1.27 |
| | 12 | 65.060 | 0.36 | 0.96 | 0.95 |
| | 12 | 4.966 | 0.70 | 1.54 | 1.49 |
| | 12 | 9.048 | 0.55 | 1.53 | 1.49 |
| | 12 | 14.820 | 0.53 | 1.95 | 1.93 |
| | 12 | 25.428 | 0.09 | 1.37 | 1.37 |
| | 12 | 64.116 | 0.10 | 1.16 | 1.16 |
| | 12 | 4.966 | 0.92 | 2.11 | 2.09 |
| | 12 | 9.048 | 0.72 | 2.10 | 2.09 |
| | 12 | 14.820 | 0.33 | 1.29 | 1.28 |
| | 12 | 25.428 | 0.07 | 1.00 | 1.00 |
| | 12 | 64.116 | 0.08 | 0.99 | 0.99 |
| Sudan | 13 | 15.479 | 0.43 | 0.45 | 0.43 |
| | 13 | 19.665 | 0.33 | 0.32 | 0.31 |
| | 13 | 28.405 | 0.18 | 0.29 | 0.29 |

# Appendix

## Table 5 -- Root Crop Elasticities

| Country | Study Code | Income | Income Elasticity | $\lvert E_{ii} \rvert$ | $\lvert e_{ii} \rvert$ |
|---|---|---|---|---|---|
| Colombia | 1 | 10.374 | 0.22 | 0.23 | 0.22 |
| | 1 | 18.564 | 0.27 | 0.28 | 0.27 |
| | 1 | 27.339 | 0.24 | 0.25 | 0.25 |
| | 1 | 53.196 | -0.41 | 0.00 | 0.00 |
| | 1 | 122.109 | -0.44 | 0.00 | 0.00 |
| Indonesia | 3 | 5.148 | 0.99 | 1.28 | 1.25 |
| | 3 | 8.316 | 0.68 | 0.82 | 0.81 |
| | 3 | 13.068 | 0.39 | 0.94 | 0.93 |
| | 3 | 30.222 | -0.08 | 0.78 | 0.78 |
| | 5 | 8.944 | 0.85 | 1.09 | 1.06 |
| | 5 | 19.856 | 0.86 | 0.82 | 0.80 |
| | 5 | 44.688 | -0.63 | 0.67 | 0.68 |
| | 5 | 8.944 | 0.83 | 2.49 | 2.46 |
| | 5 | 19.856 | -1.02 | 2.06 | 2.08 |
| | 5 | 44.688 | -2.90 | 1.00 | 1.04 |
| Brazil | 9 | 24.852 | 0.71 | 1.36 | 1.34 |
| | 9 | 53.656 | 0.50 | 0.76 | 0.76 |
| | 9 | 245.784 | 0.09 | 0.23 | 0.23 |
| Bangladesh | 10 | 5.100 | 1.61 | 1.68 | 1.66 |
| | 10 | 25.500 | 1.88 | 0.96 | 0.94 |
| Sierra Leone | 11 | 8.865 | 0.60 | 0.15 | 0.12 |
| | 11 | 14.400 | 1.00 | 0.26 | 0.20 |
| | 11 | 26.730 | 0.86 | 0.31 | 0.18 |
| Sri Lanka | 15 | 12.508 | 1.07 | 1.33 | 1.31 |
| | 15 | 70.420 | 1.27 | 1.24 | 1.23 |

# Appendix

## Table 6 -- Milk Product Elasticities

| Country | Study Code | Income | Income Elasticity | $\lvert E_{ii} \rvert$ | $\lvert e_{ii} \rvert$ |
|---|---|---|---|---|---|
| Colombia | 1 | 10.374 | 1.83 | 1.79 | 1.74 |
| | 1 | 18.564 | 1.65 | 1.62 | 1.57 |
| | 1 | 27.339 | 1.13 | 1.12 | 1.07 |
| | 1 | 53.196 | 0.63 | 0.64 | 0.62 |
| | 1 | 122.109 | 0.20 | 0.20 | 0.20 |
| United States | 2 | 11.508 | 0.87 | 2.10 | 2.08 |
| | 2 | 34.387 | 0.70 | 1.29 | 1.29 |
| | 2 | 57.540 | 0.66 | 1.14 | 1.14 |
| | 2 | 80.419 | 0.59 | 1.03 | 1.03 |
| | 2 | 103.435 | 0.54 | 0.84 | 0.84 |
| | 2 | 126.451 | 0.48 | 0.77 | 0.77 |
| | 2 | 149.467 | 0.47 | 0.59 | 0.69 |
| | 2 | 172.483 | 0.44 | 0.67 | 0.67 |
| | 2 | 195.499 | 0.45 | 0.65 | 0.65 |
| | 2 | 218.515 | 0.43 | 0.59 | 0.59 |
| | 2 | 247.148 | 0.46 | 0.55 | 0.55 |
| | 2 | 310.442 | 0.41 | 0.50 | 0.50 |
| | 2 | 459.909 | 0.46 | 0.00 | 0.00 |
| | 2 | 574.852 | 0.47 | 0.22 | 0.22 |
| | 4 | 36.853 | 0.41 | 1.58 | 1.58 |
| | 4 | 100.832 | 0.21 | 1.01 | 1.00 |
| | 4 | 159.057 | 0.16 | 0.52 | 0.52 |
| | 4 | 292.632 | 0.13 | 0.63 | 0.63 |
| Bangladesh | 10 | 5.100 | 2.52 | 1.08 | 1.02 |
| | 10 | 25.500 | 1.91 | 0.25 | 0.20 |
| Sudan | 13 | 15.479 | 0.94 | 0.96 | 0.90 |
| | 13 | 19.665 | 1.04 | 0.97 | 0.91 |
| | 13 | 28.40[illegible] | 0.35 | 0.56 | 0.54 |
| India | 14 | 5.680 | 1.75 | 2.40 | 2.35 |
| | 14 | 9.700 | 2.34 | 1.13 | 1.04 |
| | 14 | 14.780 | 1.89 | 0.88 | 0.79 |
| | 14 | 25.320 | 1.48 | 0.81 | 0.79 |
| | 14 | 65.060 | 0.67 | 0.69 | 0.66 |
| | 14 | 4.966 | 1.65 | 1.16 | 1.13 |
| | 14 | 9.048 | 2.06 | 0.66 | 0.60 |
| | 14 | 14.820 | 1.84 | 1.06 | 1.00 |
| | 14 | 25.428 | 1.45 | 1.12 | 1.00 |
| | 14 | 64.116 | 0.87 | 1.03 | 1.00 |
| Sri Lanka | 15 | 12.508 | 2.14 | 2.10 | 2.08 |
| | 15 | 70.420 | 1.62 | 1.39 | 1.36 |

82A

# IMPACT OF INCOME AND FOOD PRICE CHANGES ON FOOD ACQUISITION BY LOW-INCOME HOUSEHOLDS

## VOL. II. THEORETICAL AND METHODOLOGICAL CONCERNS FOR MEASUREMENT OF DEMAND PARAMETERS FOR LOW-INCOME CONSUMERS

by

Harold Alderman

Report submitted to:

Agency for International Development
Office of Nutrition

Under DAN-1274-C-00-3077-00

International Food Policy Research Institute
1776 Massachusetts Avenue, N.W.
Washington, D.C. 20036

October, 1984

## INTRODUCTION

In principle, the estimation of the response of low-income consumers to price and income changes is no different from a general study of consumer demand. There are no properties derived from demand theory unique to low-income households. Just as the techniques for measuring demand employed when a data set comes from the developed world differ little from those employed with a similar data set from low-income countries, the econometrics of consumer demand are largely unaffected by the differences of income within a country. Yet there are unique problems for this task. These stem from the goals of the research and the nature of the data available.

As the goals of food policy are diverse and many instruments to implement policy are for single commodities, food policy analysis is most useful when it focuses on a set of disaggregated commodities. A complete matrix of parameters for n commodities would have $n^2$ price and cross-price parameters and n income parameters. As an example, if one is dealing with aggregate groups -- say, food, other non-durables and durables -- there would be only 12 parameters to be estimated, whereas if one chooses to focus on, say, 5 food commodities as well as the other 2 aggregate groups, there are 56 parameters to ferret out. Even this modest task is seldom possible with time-series data, as such series generally have too few observations. Furthermore, the observed prices may be sufficiently correlated to render individual

parameters unstable. Using restrictions derived from either demand theory or assumptions about the nature of the underlying utility functions, however, one can reduce the number of parameters to be estimated. Such techniques are briefly discussed in the next section.[1]

While these techniques for estimating systems of demand were originally derived for use with time-series data, they can be used with cross-sectional data as well.[2] As time-series data is usually collected in aggregation, they seldom carry enough information to make it possible to extract price responses by income groups. Cross-sectional data, on the other hand, frequently are available in a disaggregated form, and furthermore, may contain sufficient degrees of freedom that the analyst need not set restrictions on the parameters. Nevertheless, reliance on cross-sectional data for the estimation of price parameters introduces problems of interpretation and of estimation not usually present when dealing with time-series estimation. Such issues are discussed in the second section below.

---

1 As such techniques are applied to general estimation tasks as well as the special tasks discussed in this paper, the discussion is necessarily brief. For further discussion, see J. Brown and A. Deaton, "Models of Consumer Behavior: A Survey," Economic Journal 82, 1972, pp. 1145-1236; and A. Deaton and J. Muellbauer (1980a), Economics and Consumer Behavior, (England: Cambridge University Press, 1980). Both reviews are excellent and accessible. See also, L. Phlips, Applied Consumption Analysis, (New York: North-Holland Press, 1974), and A. Powell, Empirical Analytics of Demand Systems (Lexington, Mass.: Heath Press, 1974).

2 C. Lluch, "Consumer Demand Functions for Spain 1958-64," European Economic Review 2, 1971; and L. Lau, W. Lin, and P. Yotopolous, "The Linear Logarithmic Expenditure System: An Application to Consumption-Leisure Choice, Econometrica 46, 1978, pp. 843-868.

A further issue that needs to be addressed is the criterion for sample separation. While groups of households for which parameters are to be estimated can be determined largely by the use to which the parameters will be employed, there are some concerns based on statistical as well as welfare theory that should also be considered. These are discussed in the third section.

A fourth section introduces problems related to the estimation of parameters for food policy use that are not central to this study. These include problems such as different sources of income producing different spending patterns, decisionmaking within a household, and distribution of household purchases. The paper concludes with an outline of research questions of relevance to the study of the acquisition of food by low-income households.

## 1. RESTRICTIONS OF PARAMETERS FOR ESTIMATING CONSUMER RESPONSE WITH DISAGGREGATED COMMODITY GROUPINGS

### Systems of Demand

A number of techniques for estimating complete systems of demand have been developed. Such systems, which estimate all income, price, and cross-price parameters for groups of commodities, do not necessarily include restrictions on parameters -- a number are quite flexible. However, use of complete demand systems is a prerequisite for applying many of the restrictions employed to reduce the number of parameters to be estimated.

Many such restrictions are general conditions derived from any arbitrary utility function ($u = f[x]$) and consumer maximization of utility subject to a budget constraint. For example, theory requires symmetry in the Slutsky matrix of compensated price effects. As the $n^2$ price parameters may be decomposed into compensated price effects and income effects, imposing this condition a priori reduces the number of parameters by $n/2\ (n-1)$.

Furthermore, if all expenditures, including savings, exhaust the budget, the adding-up criterion can be used to eliminate one equation.[1]

---

[1] This criterion can be further decomposed into the Engel and Cornot aggregations. The former states that the sum of all income elasticities weighted by the budget shares equals one ($\Sigma w_i \eta_i = 1$) while the latter states that the sum of all own- and cross-price Marshallian elasticities for the ith commodity weighted by the corresponding jth budget shares equals the negative of the budget share to the ith good ($\sum_j w_j E_{ji} + w_i = 0$).

In addition, one can impose the condition that each estimating equation is homogenous of degree zero. As this condition implies that increasing all prices and incomes by the same proportion leaves demand unchanged, it is also referred to as the absence of money illusion. If this restriction holds, one additional parameter in each equation is redundant and need not be estimated. Not all these restrictions are fully independent, and taken together they reduce the system to n/2 (n+1)-1 parameters to be estimated.

Note that these restrictions are not particularly severe. Indeed, the adding-up criterion is basically an accounting device. If the analyst derives the demand equations from any specified direct or indirect utility function, this condition will automatically hold. Such a starting point will also generate conditions under which symmetry and homogeneity can be imposed.[1] Alternatively, an analyst can begin by specifying a demand equation and imposing such restrictions on it, although the restrictions may then only hold locally.

It is also possible to impose n restrictions on the main diagonal of the Slutsky matrix (compensated own-price elasticities are known to be zero or negative as well as require the matrix to be of a negative semi-definite form. These provide an opportunity to use prior information in appropriate statistical techniques. They may also be used to

---

[1] Indirect utility functions have certain theoretical advantages as a starting point. See L. Lau, "Complete Systems of Consumer Demand Functions Through Duality," in M.D. Intrilligator, Frontiers of Quantitative Economics, Vol. III (New York: North Holland, 1977, pp. 59-85; and A. Deaton and J. Muellbauer (1980a) op. cit.

test the results of the estimation, but they do not reduce the number of parameters to be estimated.

Theory, then, helps to reduce the parameters by approximately half. If n represents more than a few aggregates of commodities, this is still a large number. Another set of restrictions that can be employed pertains to the separability of commodities. It is not derived from demand theory, but implies that the underlying utility function has special properties. These properties are discussed by Goldman and Uzawa and by Pollak[1]. The assumption referred to as weak separability implies that the ratio of marginal utilities of goods in one group is independent of goods not in the group.

$$(1) \qquad \partial[(\partial u/\partial q_{ri})/(\partial u/\partial q_{rj})]/\partial q_{sk} = 0$$

where r and s denote groups and i,j,k commodities within groups. This allows a decision tree in which expenditures are first allotted between groups and then among commodities within groups. The number of parameters to be estimated in any one demand equation will be $N_R + M - 1$, where M is the number of groups and $N_R$ is the number of commodities in the particular group that contains the item analyzed. The main difficulty with such estimates comes in defining appropriate price indexes for the M groups.

---

1 S. Goldman and H. Uzawa, "A Note on Separability in Demand Systems," Econometrica 32, 1964, pp. 387-98; R. Pollak, "Conditional Demand Functions and the Implication of Separable Utility," Southern Economic Journal 37, 1971, pp. 423-33.

A utility function of the form

(2) $u = f[u_r(q_i, q_j) + u_s(q_k) + \ldots]$

has the property of strong or additive separability. This implies that the marginal utility of any goods belonging to one group is independent of the quantities of goods in <u>any</u> other group. The commodities can then be rearranged into two subgroups:

(3) $u = f[u_1(q_{i,j}) + \{u_2(q_k) + \ldots\}]$.

With only two commodity blocks, only $N_R + 1$ parameters have to be estimated for each commodity. This is especially useful if food groups can be placed in a block strongly separable from other additively separable groups.[1]

A special case of strong separability in which every group consists of only one commodity is known as want independence or additivity, and allows for the estimation of a complete system of demand given only income elasticities and one price derivative. This procedure, due to Frisch, then reduces the number of estimated parameters to (n-1) income derivatives and one own-price derivative from which the compensated price derivatives can be derived.[2] The own-price deriva-

---

1 See P. S. George and G. King, <u>Consumer Demand for Food Commodities in the United States with Projections for 1980</u>, Giannani Foundation Monograph 26 (University of California at Davis, 1971); and A. DeJanvry, J. Bieri and A. Nunez, "Estimations of Demand Parameters Under Consumer Budgeting: An Application to Argentina," <u>American Journal of Agricultural Economics</u> 54, 1972, pp. 422-430.

2 R. Frisch, "A Complete Scheme for Computing All Direct and Gross Demand Elasticities in a Model with Many Sectors," <u>Econometrica</u> 27, 1959, pp. 177-96. In actuality, the Frisch method requires income elasticities, budget shares, and an estimate of the money flexibility parameter (elasticity of marginal utility of income with respect to income. Money flexibility is not directly observed but derived for one independently estimated price elasticity.)

tives can be calculated from the Cornot aggregation. This method has been used to estimate a complete system of demand for various income groups using cross-sectional data.[1]

Another system of demand which is widely employed, in part because of the computational advantage gained from the underlying additive utility and in part because it is easier to use in large general equilibrium models, is the linear expenditure system (LES). Less flexible than Frisch's method in that it restricts the underlying Engel curves to be linear, it has the advantage of estimating the price parameters (and by inference, money flexibility) internally rather than deriving them from an exogenous source. In some applications, the price variance has been derived from a single cross-sectional source.[2] If no price variation is reported, it is still possible to generate a parameter matrix using assumptions about base, or committed expenditures, or by making additional assumptions to modify the system into the extended linear expenditure system (ELES), which includes savings behavior. Similarly, one can use wage data and assumptions based on the theory of time allocation,[3] or demographic

---

1 P. Pinstrup-Andersen, N. de Londoño, E. Hoover, op. cit.

2 Lluch, 1971 op. cit.; J. Strauss, op. cit.; and R. Pollak and T. Wales, "Estimation of Complete Demand Systems from Household Budget Data: The Linear and Quadratic Expenditure Systems," American Economic Review 68, 1978, pp. 348-59. The authors use this system and the less restrictive quadratic expenditure system to estimate parameters from more than one linked cross-sectional surveys.

3 R. Betancourt, "The Estimation of Price Elasticities under Additive Preferences," International Economic Review 12, 1971, pp. 283-92.

data and a theory of measurement of household composition derived from Barten,[1] to derive a complete matrix of parameters without any observations of prices.

## Testing of Restrictions for Demand Analysis

Such convenience in estimation comes at a cost, however. The assumption of additivity is contrary to intuition for disaggregated food commodities. While it may be plausible that the satisfaction obtained from an increase in total food is independent of the amount of housing, it is less likely that the utility derived from an additional unit of rice is independent of the quantity of bread consumed. Furthermore, Frisch's method and the LES imply that all goods are substitutes and none are inferior goods. Finally, the labor or demographic theories referred to above are themselves restrictive and not without plausible alternatives.

The assumption of additivity has been tested empirically in a number of studies and has fared poorly. Statistical tests of the restrictions have generally led to its rejection.[2] Furthermore, comparisons of parameters derived from additive systems with other

---

1 A. Barten, "Family Composition Prices and Expenditure Patterns," in P. Hart, G. Mills, and J. Whitaker (eds.), Economic Analysis for National Economic Planning (Woborn, Mass.: Butterworth Press, 1964), pp. 277-92. See also H. Howe, "Cross-Sectional Application of Linear Expenditure Systems: Response to Sociodemographic Effects," American Journal of Agricultural Economics 59, 1977, pp. 141-148.

2 A. Barten, "The Systems of Consumer Demand Functions Approach: A Review," Econometrica 45, 1977, pp. 23-51; reports the results of a number of estimations. See also, A. Deaton, "Specification and Testing in Applied Demand Analysis," Economic Journal 88, 1978, pp. 524-36.

systems have indicated wide divergences in values.[1] Brandt and Goodwin found that the divergence was small for aggregate groups but noteworthy with disaggregated groups. King states (without giving his derivation) that the LES requires income elasticities to be larger than price elasticities, which must lie between 0 and -1. This is in contrast to what he states is "Wold's rule" -- that income elasticities of necessities are smaller than their price elasticities while the opposite occurs with luxuries. King apparently constrains the so-called committed expenditure in the LES to be greater than or equal to 0. This is not necessary and reflects a somewhat artificial interpretation of the model.[2] Deaton points out that Frisch's method implies that the price elasticities are approximately proportional to the income elasticities and that the approximation would be closer the smaller the budget share of that good.[3] Individually, each of the empirical tests and comparisons may be subject to reservations, as they may reflect problems with the functional forms employed rather than with additivity per se. They may reflect the difficulty in comparing time series with cross-sectional estimates. But collectively they point to the conclusion that additive systems should be used in disaggregate food policy analysis only with extreme caution.

---

1 A. Deaton, "The Measurement of Income and Price Elasticities," European Economic Review 6, 1975, pp. 261-74; A. Thomson, op. cit.; J. Brandt and J. Goodwin, "The Impact of Increasing Food Supply in Human Nutrition -- Implications for Commodity Priorities in Agricultural Research and Policy: Comment," American Journal of Agricultural Economics 62, 1980, pp. 588-91.

2 R. King, "Choices and Consequences," American Journal of Agricultural Economics 61, 1979, pp. 839-48.

3 A. Deaton, "A Reconsideration of the Empirical Implication of Additive Preferences," Economic Journal 84, 1974, pp. 338-48.

While additivity provides the most powerful set of restrictions to help in obtaining parameters for food policy analysis, it is worthwhile investigating tests of the other restrictions discussed above. Although these restrictions are intuitively more appealing than additivity and may be derived from demand theory, statistical tests frequently reject them.[1] As Christenson et al. have pointed out, the rejection of such restrictions logically precludes a test of additivity, although theoretical restrictions may be imposed, despite rejection, for further work.[2] While the negativity of the diagonal terms of the Slutsky matrix was not rejected in any of the tests reports, symmetry was in about half the cases. So was homogeneity (which unlike symmetry can be imposed on a single-equation basis). Deaton and Muellbauer report similar rejection of homogeneity and symmetry using a functional form different than any reviewed by Barten.[3] While Barten does not counsel abandonment of demand theory in guiding empirical work, noting that the alternative, agnosticism, is not attractive, these results are disturbing nevertheless.

Perhaps some behavioral factors have been ignored. For example, the relatively weak condition of homogeneity may have been rejected as money illusion, may be a short-run phenomenon, and would not appear in

---

1 A. Barten, 1977, op. cit.

2 L. Christenson, D. Jorgenson, L. Lau, "Transcendental Logarithmic Utility Functions," American Economic Review 65, 1975, pp. 367-83.

3 A. Deaton and J. Muellbauer, "An Almost Ideal Demand System," American Economic Review 70, 1980b, pp.312-26.

a more dynamic analysis. Deaton argues that money illusion occurs when inflation is accelerating, as the consumer does not face all prices simultaneously but compares current prices with recollections of other commodity prices.[1] Barten, however, cites the six-year separation of data points covered in Lluch's study as an example of a rejection of homogeneity in which there should have been ample time for dynamic adjustment.[2] Clearly, there are grounds for further work on money illusion and the dynamics of consumers' adjustments to changes.

A more basic explanation can be offered for the sobering results cited by Barten. Demand theory is based on the behavior of the individual consumer while most empirical studies, especially those involving time series, require the aggregation of individual data. (Aggregation of commodities also presents difficulties related to index number problems, the solution of which are intimately connected to issues of separability.)

If q is a vector of goods and P is a vector of prices, then, according to microeconomic theory,

(4) $q_n = f_n\ (y_n, P)$ where n denotes the individual.

To form an average,

(5) $\overline{q} = 1/N(\Sigma q_n)$ and $\overline{y} = 1/N(\Sigma y_n) = 1/N\ \Sigma P' q_n = P' q.$

One estimates

(6) $\overline{q} = f(\overline{y}, P)$

---

1 A. Deaton, "Involuntary Savings Through Unanticipated Inflation," American Economic Review 67, 1977, pp. 899-910.

2 C. Lluch, op. cit.

but in general (6) is not a behavioral equation. Note that from equations (4) and (5),

(7) $\bar{q} = 1/N \Sigma f_n (y_n, P)$

which does not necessarily equal (6).

An analogous problem can be observed if $q = f(y, P, Z)$, where Z is a vector or matrix of consumer characteristics, perhaps age or race. For example, if q is not independent of age or q is not a linear function of age, information on age distribution is required, unless, of course, average age and age distribution are constant over the time series. One has less theoretical guidance on Z, and the characteristics included are frequently discrete dummy variables, but the relevant issues are the same. Both the translog system and the almost ideal demand system can be aggregated consistently.[1] Aggregation, of course, is less of a problem if the sample is cross-sectional, although such data sets may contain only cell means for analysis.

Muellbauer discusses the conditions under which equations (6) and (7) are equivalents.[2] Obviously, if every consumer were the representative consumer often alluded to, there would be no problem. Also, estimates from grouped data will be consistent if the budget share allocated to each good is independent of y. This implies, of course, linear Engel curves with unitary elasticities. Actually, this condition of linear Engel curves can be relaxed. Intuitively, this

---

1 Christensen, et al., op. cit; Deaton and Muellbauer, 1980b, op. cit.

2 J. Muellbauer, "Aggregation, Income Distribution, and Consumer Demand," Review of Economic Studies 62, 1975, pp. 525-43.

should be apparent. If the demand function were either strictly concave or convex in income, then

(8) $\alpha f(y_1) + (1-\alpha)f(y_2) \lesseqgtr f[\alpha y_1 + (1-\alpha y_2)]$

which would not give consistent aggregation. Muellbauer suggests the use of $y_0$, a representative income not necessarily equal to $1/N\ \Sigma y_n$ as a representative income that takes account of income distribution. Unless income distribution is constant, $y_0$ will vary over time or regions even if average y does not. Muellbauer, however, offers no empirical evidence and the question of whether restrictions imposed on demand systems will fare better when $y_0$ or some other consideration of income distribution is taken into account is, as yet, unanswered.

Deaton indicates another possible reason for apparent rejection of homogeneity; he suggests that the consumer might maximize utility subject to a rationing constraint that requires the consumption of one or more goods to be fixed exogenously.[1] In the example investigated, rationing was defined broadly to include housing as fixed in the short-run. Therefore, the study resembles investigation of production functions with fixed factors. Deaton found that a rationed effect model with housing constrained rejected homogeneity for only one of the seven commodities tested, although the homogeneity was rejected for the total system. This was compared to four rejections out of eight for the same data without imposing the restriction on housing.

---

[1] A. Deaton, "Theoretical and Empirical Approaches to Consumer Demand Under Rationing," in A. Deaton (ed.), Essays in the Theory and Measurement of Consumer Behavior, in honor of Sir Richard Stone (England: Cambridge University Press, 1981), pp. 55-72.

Quantity restrictions, kinked budget constraints, and various other market imperfections, then, may contribute to the difficulty in applying restrictions derived from theory to observed behavior.

Restrictions by Judicious Choice

Frequently, instead of estimating a complete system of demand parameters, analysts choose to concentrate on income and own-price parameters and a subset of cross-price parameters chosen by the researcher as most likely candidates for substitution or complementation. This can be done with a single equation and is, in effect, equivalent to assuming that $\partial q_i/\partial P_k = 0$ for a subset of commodities. If the assumption is not true, then it leads to missing variable bias for the included parameters, provided that the omitted and included variables are correlated. However, in general, it will be inconsistent to assume that the omitted price elasticities are all zero if the Cornot aggregation and symmetry are to hold. For example, if only own prices are included in an equation, the Cornot aggregation and symmetry imply that the own-price elasticity is -1. Perhaps. But the analysis should not begin with such an assumption.[1] A similar argument can be made about the real income effect from the Slutsky decomposition, although it is possible but not likely for $E_{ij} = e_{ij} + w_j n_j = 0$ even when $w_j$ and $n_j$ do not equal zero. Exclusion of all cross-

---

[1] T. Young, "Single Equation Demand Estimation," European Economic Review 9, 1982, pp. 103-105; makes a similar argument using homogeneity, in which including only own-price response implies that the income elasticity equals the negative of the price elasticity. Homogeneity is perhaps a slightly stronger assumption than aggregation, which is basically bookkeeping.

price terms, then, must rest either on the possibility that they are not statistically different from zero, perhaps because the variance is small, and not actually zero, or that the prices are uncorrelated with the included price. This is not as unlikely as it may seem, for if the analysis is conducted with real prices and if the omitted class of goods is composed of many commodities, then the price index for the aggregate good may be uncorrelated even if some of the components are not. This is an empirical question that may be verified if the data is available.

Cross-price parameters may be difficult to estimate, but they are nevertheless important for food policy analysis. If a set of goods is excluded, then one has to be cautious in applying the parameters derived. For example, suppose that the demand for rice is being estimated and that the principal substitute, bread, has a fixed price. Statistically, [illegible] n, all derivatives with respect to the price of bread are unestimable. Therefore, there will be no missing variable bias for the own-price parameter for rice, but there may be one if rice prices are not included with bread estimates. If one is concerned with the total food or calorie consequence of a change in rice prices, the net demand for wheat must be included. In general, it is not unlikely that low-income consumers in many areas where so-called inferior grains are consumed are particularly likely to substitute between major sources, depending on their relative prices. The potential effect of a change in the price of a major staple on nutrition in such a community, then, could be either underestimated or overestimated if own-price elasticities are employed in isolation and avenues

for mitigating policy interventions overlooked. Pinstrup-Andersen et al. took the ratios of the total effects on the poor of a change in the price of a good. These ratios were 1.19 for rice, 1.33 for cassava, 1.39 for potato, and only 0.51 for milk and 0.55 for beef.[1] Williamson-Gray found that the net effect of a subsidy on wheat was an insignificant decrease in calorie consumption because of substitutions.[2] Yetley and Tun found the total effect of a change in the rice price was only 55 percent of the own-price effect alone.[3]

Similar differences in net and total effects were reported by Pitt.[4] These studies used different methodologies and different data bases, and collectively they remind the food policy analyst that, although they are difficult to estimate, cross-price parameters can have a major effect on food consumption.

Even if it is not possible to estimate all parameters in a system that has restrictions across equations, it may be useful to estimate the parameters for all commodities jointly. Single-equation estimates are inefficient in the sense that they are unable to use information contained in the estimates for other commodities. One way in which some of this information may be used is the method of seemingly unre-

---

1 Op. cit.

2 Op. cit., pp. 41-42.

3 M. Yetley and S. Tun, "Household Demand Analysis for Assessing Nutritional Impact of Development Programs," Staff Report AGES810806, Agricultural Development Branch, Economic Research Service (USDA, August 1981).

4 Op. cit.

lated regressions (SUR).[1] This method takes into account the correlations of residuals between equations, based on the assumption that even if all consumers have identical demand functions and each equation is specified without heteroskedasticity, an exogenous shock that affects actual purchases for the ith commodity in household h will affect demand for the jth commodity as well. Note, however, that SUR does not result in a gain in efficiency if the independent variables in each equation are identical or if the residuals are uncorrelated. The method may be used with any functional form, including local approximations that are not necessarily globally consistent with demand theory.

## Choice of Functional Form

Another type of restriction that has been used in conjunction with income-specific price elasticities is a functional relationship that is an interaction of price and income terms. While many of the studies reporting price parameters by income group have been made with covariance techniques, using interaction of dummy variables and price elasticities, or by dividing the sample into subsamples, it is possible to estimate equations of the form $Q = \beta(Y, P, Y \times P)$. This, intuitively, has an advantage over covariance techniques because price elasticities are disjunct in covariance methods; a small change of household income at the extreme of an arbitrary division of the sample can lead to a

---

[1] A. Zellner, "An Efficient Method of Estimating Seemingly Unrelated Regressions and Tests for Aggregation Bias," Journal of the American Statistical Association 57, 1962, pp. 348-368.

marked change of price elasticities for that household. If, on the other hand, the price parameter is specified with a continuous second derivation with respect to income, one can test whether $\partial^2 Q/\partial P \partial Y = 0$ with statistical techniques. If the null hypothesis is rejected, then the price elasticities vary continuously with income. By Young's theorem, however, the income elasticity will vary by price. An early test of such an interaction was designed not to test the hypothesis that price elasticities differ across income groups, but that quality effects would require that if $\partial \bar{P}/\partial Y > 0$, then positive income and price would interact when higher priced items tended to be luxuries and less expensive ones necessities.[1] Cramer's study found no effects in time-series analysis. He did, however, note a number of positive interactions using cell means for each of 40 quart-erly cross-sectional surveys. This is consistent with the quality hypothesis, but equally consistent with the general view that the absolute value of price elasticities are smaller for more affluent consumers.

Another problem with estimating a relationship that includes a price times income interaction is that by specifying the relationship, the rate of change of price elasticities is constrained as income rises to a specific form.[2] While this may be sufficient for generating the parameters for subsequent policy work, it cannot be used for specifying relationships as in Volume I.

---

1 J. S. Cramer, "Interaction of Income and Price in Consumer Demand," International Economic Review 14, 1973, pp. 351-63.

2 An example of such a study is Pitt, op. cit.

Functional forms may indicate variations of price elasticities by income groups that are constrained by the form or are artifacts of it. Various estimating equations generate elasticities that are functions of quantities consumed or budget shares.

A standard exercise in consumption theory illustrates that the derivative of the cost function with respect to price is the Hicksian (compensated) demand and that the derivative of this demand with respect to price is the compensated price parameter:

$$(9) \qquad \frac{\partial^2 C(u,p)}{\partial P_i \, \partial P_i} = \frac{\partial H(p,Y)}{\partial P} < 0.$$

To go one step further,

$$(10) \qquad \frac{\partial^3 C(U,P)}{\partial P_i \, \partial P_i \, \partial u} \quad \frac{\partial u}{\partial Y} = \frac{\partial H(P,Y)}{\partial P_i \partial Y}.$$

Although $\partial u/\partial Y$ is unambiguously positive, the sign of the expression is indeterminate and is subject to the formulation of the cost function.

As an illustration, consider the AIDS which is derived from a flexible approximation of an unknown cost function.[1] The own-price elasticity (uncompensated) from this model is

$$(11) \qquad \frac{\gamma ij - B_i \alpha i - B_i \Sigma \gamma ik \log P_k}{w_i} - 1$$

where the $\gamma$, B, and $\alpha$ are estimated parameters and $w_i$ is the budget

---

[1] This illustration is presented in elasticity terms which clearly differ from the Hicksian price parameter. The example could be constructed in derivative terms as well but the point should be clear with the simplier calculations here.

share of the ith good. Calling this expression

(12) $\frac{f'(p)}{w_i} - 1,$

then,

(13) $\partial \left[ \left( \frac{f'(p)}{w} - 1 \right) \right] / \partial y = -[f'(p)/w^2] \, [\partial w_i / \partial y].$

Generally, $f'(p)$ is invariant with respect to Y, but $w_i$ clearly is not. This result implies that with AIDS, the poor will be less price responsive (will have a less negative elasticity) than the rest of the population under either of the two conditions. The first is when $\partial w_i/\partial y > 0$ and $f'$ is positive. The former relationship describes a luxury good ($\eta > 1$) and, in this system, the latter denotes that the price elasticity lies between 0 and -1. The second is when $\partial w_i/\partial y < 0$ and $F'(p)$ is negative. This is a situation in which the commodity is a necessity or inferior good and $E_{ii}$ is less than -1.

These conditions do not always hold, as some studies that used the AIDS show.

This illustration is not meant to imply that AIDS is unique in its implications for the changes of price response by income groups. Similar illustrations could be derived for almost any estimation technique, whether derived directly from theory or based on a more pragmatic approach. The question is whether any such pattern that is built into an estimation conforms to reality. To answer it, one needs either to allow the parameter to vary in a flexible manner,[1] as it did

---

[1] AIDS is a second order expansion of an arbitrary indirect utility function. In principle, a third order expansion would give the necessary flexibility but is surely unestimable.

in many of the examples given in the previous section, or to estimate parameters for different subgroups and compare the results.

When such a formulation is used with covariance analysis, the sign of the interaction term of price times a dummy for low-income consumers may not necessarily indicate a change in absolute elasticities by income group; in some cases, it may indicate that the implicit pattern is moderated by the choice of functional form.

The reader will note that the main focus of this study is on the estimation of price parameters. This is not because the estimation of income parameters yields no issues of interest. Indeed, some of those are addressed below. In general, however, the estimation of income parameters for low-income consumers involve no new or more severe statistical problems not found in general population estimates.

The major issue involved in the estimation of income parameters is the choice of functional form. Similar to the discussion above, the choice of underlying Engel functions generally entails assumptions about changes in elasticities as incomes change. Theory provides little assistance in this matter -- mainly the Engel aggregation -- and even that is frequently made secondary to goodness of fit. Other considerations revolve around the expected curvature, the possibility that consumption will be saturated, and the likelihood that there is either a minimum amount of consumption or a minimum amount of expenditures for entry into the market. For example, linear relationships imply that the marginal propensity to consume is constant. Although such a relationship is not likely to conform to experience, the linear form underlies a number of demand studies. Log-linear relationships

are also widely used because the coefficient of an estimating equation yields elasticities directly. The linearity of the (transformed) function allows for easy manipulation in larger models. The form, however, requires that elasticities are constant for all amounts of consumption. Again, this is unlikely and, indeed, if a good is a luxury item -- that is, if $n_i$, the income elasticity, is greater than one -- the budget share would increase without limit until it accounted for all of income. However, the form is useful as a local approximation. Similarly, other common functional forms have implications for either the income elasticities or the marginal propensity to consume. The inverse form,

(14) $Q = a_0 - b/y$,

is asymptotic to $a_0$ and has a marginal propensity inversely proportional to the square of income (y). The semi-log form

(15) $Q = a_0 + \ln(y)$

is not asymptotic but has an elasticity inversely proportional to the amount of consumption and a marginal propensity to consume that varies inversely with income. A log inverse relationship

(16) $\mathrm{Ln}(Q) = a_0 - b/y$

is asymptotic and has an elasticity inversely proportional to income. Another form that is common, in part because it adds up properties, is the Leser's budget share equation:

(17) $w = a_0 + b \log Y$,

which gives an elasticity of $b/w + 1$.

There are, in addition, a number of three-parameter forms in general use, such as the log-log inverse or quadratic forms of

equation 17. These have the advantage of allowing a good to vary from necessity to inferior good at different incomes.

Each of these forms have their proponents and uses for which they are most appropriate. Prais and Houthakker have shown the implications of many of these forms which will usually diverge less at the mean of the sample than at the extremities.[1] Box-Cox transformations[2] and Fourier transformations[3] offer greater flexibility although they are difficult to interpret. Furthermore, as they are polynomial or trigometric expansions, they are difficult to estimate, especially when many variables are included.

---

1 See S. Prais and H. Houthakker, The Analysis of Family Budgets (England: Cambridge University Press, 1971).

2 J. Aaness and A. Rodseth, "Engel Curves and Systems of Demand Functions," European Economic Review 20, 1983, pp. 95-121; and J. Baylock and R. Green, "Analysis of Flexible Engel Functions," Agricultural Economics Research 32, 1980, pp. 12-20.

3 M. Wohlgennut, "Conceptual and Functional Form Issues in Estimating Demand Elasticities for Food," American Journal of Agricultural Economics 66, 1984, pp. 211-15.

## 2. ESTIMATION CONCERNS ASSOCIATED WITH CROSS-SECTIONAL DATA

### Interpretation of Parameters

Given the problems noted in restricting the number of parameters to be estimated for disaggregated demand responses, such work is generally performed with large cross-sectional data sets. The advantages of such sets are not only that they have greater degrees of freedom but also in the type of information they provide. Time-series data is seldom available disaggregated by income groups, although, as discussed below, a price- and income-interaction term may, in principle, allow one to circumvent such a deficiency. Furthermore, cross-sectional data frequently contain demographic information, the importance of which has been noted in a number of studies.[1] Even functions designed with consistent aggregation cannot be expected to be able to explore demographic effects as deeply as disaggregated data. Generally, the average and the distribution of a demographic variable, say, family size, changes little over time, but varies widely in a cross-section.

Estimates of price parameters with cross-sectional data using either systems or single-equation methodologies are increasingly com-

---

[1] See, for example, the review by R. Pollak and T. Wales, "Demographic Variables in Demand Analysis," Econometrica 49, 1981, pp. 1533-51. A less theoretical approach is reported by J. Adrian and R. Daniel, "Impact of Socioeconomic Factors on Consumption of Selected Food Nutrients in the United States," American Journal of Agricultural Economics 58, 1976, pp. 31-8.

mon. Nevertheless, it is important to recognize that the interpretation of the parameters generated differs from the more conventional time-series estimates.

That estimates from cross-sectional data differ from time series has been known for some time. Kuh observed that the coefficient for capital stock in investment functions generated by time series exceeded the coefficient from cross sections, while the coefficients for profits from cross sections were typically twice as large as those from time series.[1] Similarly, Peterson observed an agricultural supply response from an international cross-sectional data set that was much larger than those commonly observed from time-series data.[2] A comparison of demand estimates using time series and pooled cross sections (cell means) for 17 years of Indian data found time-series price elasticities for food less in absolute value than from cross section.[3]

The major reason proposed for these phenomenon is the differences between short-run and longer-run adjustments to a change. This is most well-known in supply response analyses, in which lagged variables are used to gain a perspective on the dynamics of adjustments. An example of demand models in which the previous period's consumption

---

1 E. Kuh, "The Validity of Cross-Sectionally Estimated Behavior Equations in Time Series Applications," Econometrica 27, 1959, pp. 197-214.

2 W. Peterson, "International Farm Prices and the Social Cost of Cheap Food Policies," American Journal of Agricultural Economics 1979, pp. 12-21.

3 R. Ray, op. cit.

is included as a regressor can be found in Pollak.[1] The positive coefficients generally observed for such lagged variables are interpreted as indicators of consumer habits. This interpretation, while plausible, does raise some problems. Note, for example, that even if habit formation did not exist, the effects of the indivisibility of durable goods and of liquidity constraints would influence the purchase of other commodities as well. Since purchases of durables are frequently negatively correlated with the stock of durables, itself a reflection of the previous period's purchases, adding up budget shares would imply that the coefficients for the lagged variables for some of the non-durables were positive.[2] The habit-formation hypothesis has an additional problem in that most welfare theories assume that tastes are exogenous, which is clearly not tenable in a habit-formation model.

Other constraints and disequilibriums may also influence the time-series analysis. These include constraints on labor-leisure choice, housing, and temporary or permanent shortages of goods in settings with price rigidities. There is a large literature on disequilibrium

---

1 R. Pollak, "Habit Formation and Dynamic Demand Functions," Journal of Political Economy, pp. 60-78, 1970. See also L. Blanciforti and R. Green, "An Almost Ideal Demand System Incorporating Habits: An Analysis of Expenditures on Food and Aggregate Commodity Groups," Review of Economics and Statistics LXV, 1983, pp. 511-15.

2 A. Deaton and J. Muellbauer, 1980 op. cit., Chapter 13 points out that dynamic time-series analysis is frequently based on static one-period utility maximization subject to one period budget constraints and discusses an alternative model based on aggregation and interhousehold social relationships. See also R. Pollak, "Habit Formation and Long-Run Utility Functions," Journal of Economic Theory 13, 1976, pp. 272-97.

and rationing, broadly defined.[1] One general conclusion that can be drawn from the literature parallels the more widely-known condition from production literature that short-run cost functions lie within the long-run envelope. Short-run responses with fixed (human or other) capital costs, then, are less than long-run substitution effects.[2] Compared to production studies, however, little has been done on the distribution and time frame of lagged adjustments in consumer behavior.

Time-series estimates may also underestimate actual responses by virtue of aggregation. For example, suppose that aggregate income grows but that the growth is concentrated in upper-income groups that typically have low elasticities for staple food. In such a case, parameters derived from a methodology that do not include income distribution will be misleading if they are extrapolated to the poor. Shares of total consumption of a normal good by the upper income groups will exceed their share in the population. Both aggregate income and price responses will reflect these quantity weights and will, therefore, reflect the presumably lower elasticities of the upper income brackets.

But, as Kuh has pointed out, differences between time-series and cross-section studies reflect the biases of both approaches. What, then, are the potential biases in cross-sectional analysis? An

---

1 See A. Deaton, 1981, op. cit.

2 See R. Pollak, "Conditional Demand Functions and Consumption Theory," Quarterly Journal of Economics 83, 1969, pp. 70-78.

important consideration in the answer to this question is the sources of price variation used for the estimates. Household budgets are often analyzed with the assumption that the prices faced by all individuals are the same at a given time. In single-period surveys, then, there would be no price variation and price parameters would not be estimable without strong assumptions. Prices are frequently not even collected or reported in such surveys. The assumption that price variation is insufficient is obviously not valid for the numerous cross-sectional surveys from which price parameters have been estimated. What are the sources of such price variations?

Some may be temporal, reflecting seasonal rounds of the survey. Such estimations, in effect, pool cross-sectional and time-series data, albeit over a short time span.

Generally, for such estimations the researcher includes covariance terms or more sophisticated variance-component models.[1] If income and prices are already included, however, dummy variables for rounds are proxies for unknown nonmarket effects. Of course, if the seasonal effect influences tastes, as in ice cream or festival foods, the interpretation is straightforward. If, however, one is dealing with a staple grain, then a significant variable for a round may reflect cash constraints or changes in labor, and through them, energy output. Also, it may reflect errors in the price variables or other market stickiness not fully in accord with the assumption underlying in cross-

---

[1] G. Maddala, Econometrics, Chapter 14 (McGraw Hill, 1977); and Y. Mundlak, "On the Pooling of Time Series and Cross-Section Data," Econometrica 46, 1978, pp. 69-85.

sectional data that there is a long-run equilibrium. Indeed, if real prices vary between rounds, then the long-run adjustment interpretation is diluted. Similarly, while seasonal fluctuations in income can be modeled with a permanent income hypothesis and expenditures can be used as proxies for permanent income, the adjustment to unanticipated changes in income is unlikely to be instantaneous.

Even when only a single round of a survey is conducted over a wide area, differences in climate may create differences in the cropping cycle. For example, in any given month, some regions will be closer to harvest than others and will, therefore, have different local prices. Similarly, even when trade is involved, differences in transportation networks will create for differences in regional prices. Such spatial price differences may reflect the real costs of marketing and may persist even in a competitive economy. For the purpose of demand estimates, however, competition need not be assumed; collusion among sellers or local monopoly behavior represents shifts of the supply curve and may not interfere with the identification of the demand curve.

An assumption that is critical, however, is that there are no unspecified regional differences in the demand curves. Otherwise, differences in prices caused by shifts in supply cannot be distinguished from those caused by shifts in demand, and a classic simultaneity bias may occur. In such a case, a positive bias in the price parameters can be expected. But with individual or cell means data from a number of towns or villages within a number of larger districts, a portion of the bias can be removed by covariance

analysis. For example, Timmer and Alderman used dummy variables for six regions for urban residence, and for the two rounds in their cross-section study of data aggregated by province.[1]

When spatial variations in prices reflect differences in markets, it is important to ascertain whether these differences are nominal or real. Cross-sectional studies, like analyses of time-series data, should deflate income and prices by cost of living or similar price indexes. Which price indexes to use and how they should be used have been major topics of discussion in consumption and welfare theory, and two related but distinct issues are important here. One use of price indexes is to allow systems to be estimated in real income so that symmetry restrictions can be imposed.[2] Some systems, such as AIDS, estimate the index internally and nonlinearly. However, substitution of an externally-derived price index not fully consistent with the parameter estimates allows for linear estimation and has been shown to impose little bias on the series estimations (no comparison has been made within cross-section estimations). A prerequisite for internally-estimated real-income deflators is the inclusion of all prices and commodity groups in the estimations.

This, then, involves a second use of price indexes. If one imposes homogeneity but does not impose symmetry restrictions on the compensated

---

1 op. cit.

2 See G. Swamy and H. Binswanger, "Flexible Consumer Demand Systems and Linear Estimations: Food in India," American Journal of Agricultural Economics 65, 1983, pp. 675-84; for a discussion of this and related issues.

substitution, then the Marshallian relationship can be estimated without a real income deflator, since under homogeneity, $f(Y,P) = f(Y\lambda, P\lambda)$ where $\lambda = 1/\text{Index}$. This, however, does not avoid the need for an index of nonfoods or some other aggregate of other commodities. Exclusion of such an index would create missing variable bias as discussed in the previous section. Exclusion, then, of regional cost-of-living deflators and/or regional price indexes for other goods poses a problem that is more severe the less homogenous the demand function and the greater the missing variable bias. Occasionally, the impact of changes in prices (regional differences) that are nearly uncorrelated with other prices is investigated. An example is Rosenberger's investigation of exogenously federally-determined milk prices. In such a case, the consequence of excluding a price index is probably small. This is not always the case, however, and when data is available the importance of regional price indexes should be investigated.

There are other reasons for differences in prices observed in cross-sectional data. Most well-known are differences in quality. Most cross-sectional data do not include direct observations on market prices of a standard or indicator commodity. They include, instead, individual prices or expenditures that include quality premiums. One can, therefore, expect a bias in estimating price responses. This problem is more complex than just commodity aggregation, for not only is one dealing with an aggregation of goods of various quality but also one is dealing with a price belonging to one component of the aggregation, but which component is not known. The problem can be

modeled as an error-in-variable problem. Suppose the true quantity model is

(18) $Q = a + \beta P^*$

where $P^*$ is the quantity price.

Quantity is observed correctly but $P = P^* + q$ only where q is a quality premium. The regression of Q on P, then, will give a biased coefficient of $\beta$, say $\beta'$. By standard-errors-in-variable procedures,

(19) $\beta' = \beta/(1 + \sigma_q^2/\sigma_{p^*}^2)$

where $\sigma_q^2$ and $\sigma_{p^*}^2$ are variances.

The error in measurement of own-price parameters, then, will make the response less negative. The problem does not end here, however, for not only are other price variables also measured with error, but the model excludes any substitution between quantity and quality. Expanding equation (18) to

(20) $Q = a + \beta P^* + Cq$

does not change the errors-in-variable part of the bias on $\beta$ but it does add a missing variable bias which, given the positive correlation between P and q, depends on the sign of C. If quality is a gross complement, then the second component of the bias will be negative, otherwise the substitution will make the observed $\beta'$ still less negative.

Another sort of bias comes when the price per unit is itself a function of the quantity purchased. If there are substantial discounts for large bulk purchases, then the association of high prices and small purchases may reflect a causality going from quantity

to price and not vice versa. This will exaggerate the own-price response.

Case Study of Time Series

As discussed, seldom is data available in a panel that would make it possible to construct time series for low-income groups within a larger population. One exception is the data used by Bouis, which is from 15 quarterly surveys made in several districts of the Philippines.[1] While not strictly a panel form because the same households are not included in each round, the data provides an opportunity to compare time-series and cross-sectional results from the same data set.

The approach here is somewhat artificial. The most efficient way to analyzing this data is to use the whole sample with or without covariance techniques for combining time-series and cross-sectional data. Such an approach was employed by Bouis. However, as one usually has either time-series or cross-sectional data, but not both, this data has been packaged in either of those forms.

For the cross-sectional analysis, regressions were run independently for each quarter. Separate log-linear regressions were run for each of four income quartiles and separately for Mindanao and Visayas Islands. For each class and region, then, there are 15 estimates of each elasticity. In order to derive the cross-sectional elasticities for each class and region, the different elasticities were averaged

---

[1] Bouis, op. cit.

and weighted by the inverse of their variance. This method gives higher weights to parameters obtained with greater precision.

Time-series estimates were also obtained separately for each income group and island. In this case, the means of each cell were calculated and included in separate time-series regressions with 15 observations each. This method differs little, then, from the standard time-series approach except that the poor are not averaged with the rich. To help handle the problem of collinearity, a principal components technique developed by Mundlak was used.[1] While the technique solved some estimation problems (and introduced others, as the package was designed for more variables than employed here), it did not introduce patterns not also apparent in the OLS forms of the estimates.

The results are presented in Tables 2.1 and 2.2. After making allowance for subdivisions of the data, which make the cells or cross-sections small (50-60 households in a cell), and, therefore, the parameters unstable, the results of both techniques seem plausible. But they are, however, dissimilar. The income elasticities from the time series are low, with rice in Visayas generally an inferior good and maize in both regions a strongly inferior good for the middle-income groups. Contrary to expectations, many of the own- and cross-price elasticities in the quarterly time-series estimations exceed the corresponding parameter from the cross-sectional regression. This

---

[1] Y. Mundlak, "On the Concept of Non-Significant Functions and Its Implications for Regression Analysis." Journal of Econometrics 18, 1981, pp. 139-49.

Table 2.1--Income and Price Elasticities - Mindanao

| | Cross-Sectional | | | | | | Time Series | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\eta_{rice}$ | $\eta_{maize}$ | $\varepsilon_{rr}$ | $\varepsilon_{mm}$ | $\varepsilon_{rm}$ | $\varepsilon_{mr}$ | $\eta_{rice}$ | $\eta_{maize}$ | $\varepsilon_{rr}$ | $\varepsilon_{mm}$ | $\varepsilon_{rm}$ | $\varepsilon_{mr}$ |
| Income Group 1 | .43 | -.09 | -2.33 | -.81 | 1.08 | 2.22 | .37 | -.26 | -.59 | -.48 | .66 | .86 |
| | (3.32) | (0.70) | (3.90) | (1.92) | (2.59) | (4.51) | (1.73) | (1.21) | (1.14) | (1.40) | (1.90) | (1.68) |
| Income Group 2 | .53 | -.70 | -1.29 | -1.01 | .66 | .01 | .76 | -2.22 | .29 | -.95 | .89 | -.19 |
| | (1.66) | (2.17) | (3.97) | (1.98) | (1.27) | (0.02) | (3.55) | (3.57) | (3.55) | (1.56) | (3.55) | (0.74) |
| Income Group 3 | .40 | -.41 | -.49 | -.54 | 1.05 | .85 | .45 | -2.09 | -.51 | .16 | -.09 | 2.37 |
| | (1.33) | (1.26) | (1.04) | (2.14) | (2.26) | (1.65) | (2.67) | (3.08) | (2.67) | (0.11) | (2.67) | (3.08) |
| Income Group 4 | .18 | -.02 | -.08 | -.40 | .25 | .21 | -.09 | -.18 | -.05 | -2.69 | .16 | 1.06 |
| | (1.54) | (0.15) | (0.18) | (0.96) | (0.65) | (0.44) | (1.44) | (0.24) | (1.44) | (2.05) | (1.44) | (0.71) |

t-values in parentheses

Source: Estimated as discussed in text using data from Bouis, op. cit.

Table 2.2--Income and Price Elasticities - Visayas

| | Cross-Sectional | | | | | | Time Series | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\eta_{rice}$ | $\eta_{maize}$ | $\varepsilon_{rr}$ | $\varepsilon_{mm}$ | $\varepsilon_{rm}$ | $\varepsilon_{mr}$ | $\eta_{rice}$ | $\eta_{maize}$ | $\varepsilon_{rr}$ | $\varepsilon_{mm}$ | $\varepsilon_{rm}$ | $\varepsilon_{mr}$ |
| Income Group 1 | .28 | -.40 | -2.18 | -.33 | .56 | 2.05 | .09 | -.12 | -.88 | -.07 | .06 | 1.10 |
| | (2.37) | (3.36) | (4.22) | (0.68) | (1.12) | (4.18) | (3.61) | (3.08) | (3.61) | (3.08) | (3.61) | (3.08) |
| Income Group 2 | .76 | -.45 | -1.65 | .59 | .00 | .24 | .07 | -1.52 | -1.39 | -1.78 | .07 | -1.06 |
| | (2.80) | (1.88) | (3.47) | (1.32) | (0.00) | (0.51) | (3.48) | (1.27) | (3.48) | (2.62) | (3.48) | (2.83) |
| Income Group 3 | .59 | -.60 | -1.18 | -.59 | .56 | .34 | -.18 | -2.07 | -.81 | -1.89 | .42 | 1.27 |
| | (2.88) | (3.01) | (2.22) | (1.19) | (0.94) | (0.64) | (1.30) | (2.54) | (3.11) | (2.03) | (1.30) | (2.76) |
| Income Group 4 | -.09 | -.09 | -.81 | -.86 | .03 | 1.55 | -.28 | -.22 | -.12 | -.55 | .48 | .89 |
| | (0.87) | (0.78) | (1.80) | (1.85) | (0.10) | (3.15) | (3.52) | (2.28) | (0.79) | (2.28) | (3.86) | (2.28) |

t-values in parentheses

Source: Estimated as discussed in text using data from Bouis, op. cit.

particular case study, then, does not lend support to the view that cross-sectional analysis measures a long-run adjustment that exceeds time-series results in absolute value.

In general, the pattern of larger absolute values of price elasticities for the poor is supported in the cross-sectional analysis, while the time-series results are mixed. Both approaches produce patterns that do not always decline strictly monotonically -- there is no reason that they should. The time series has a few erratic values that makes patterns hard to discern.

More evidence, then, would be useful to indicate whether time series using village or district data and covering quarters or monthly observations offer a better way to estimate price parameters by income group than cross-sectional data. Similarly, where repeated surveys are available, it is still an open question whether the data should be analyzed as a time series -- for example, with first-difference equations -- or with techniques that allow parameters to be derived from variations across households as well.

## Non-Consumption

One remaining problem typical of cross-sectional surveys is the issue of nonconsumption. Not all respondents will have consumed a given good during the period covered in the survey. Nonconsumption is more likely the more disaggregated the commodity groups and the shorter the recall period. If the data include observations on use, including stock draw-down or buildup or the flow of services from durables rather than purchase, then nonconsumption is less likely.

From the standpoint of estimation, nonconsumption or zero-value observations may lead to biased estimates. If the true model is

(21) $q = \beta x + u$ or $u = q - \beta x$,

then, because $q > 0$ for all cases, the distribution of u is nonnormal, which biases ordinary least squares estimation. If the regression is estimated only for the values for which $q > 0$, then only those observations for which $u > -\beta x$ are used and, therefore, $E(u) \neq 0$, which again violates a basic OLS assumption. Tobin originally stated the problem and devised an estimation procedure called Tobin's probit or Tobit to deal with it.[1] The procedure is a maximum likelihood technique with the likelihood function of

(22) $L = \prod_1 1/\sigma \; f(q_i - \beta x/\sigma) \prod_2 F(-\beta x/\sigma)$,

where f(.) is the standard normal density and F(.) is the cumulative normal density. The subscript 1 pertains to the portion of the sample for which $q_i > 0$ and 2 for those observations for which $q = 0$. Pitt has used such a scale model in an analysis of food demand in Bangladesh and has shown that it is inappropriate for models that have expenditures or budget shares as dependent variables.[2] McDonald and Moffit have shown that the parameters from a Tobit estimation can be decomposed into two components:[3]

---

1 J. Tobin, "Estimation of Relationships for Limited Dependent Variables," Econometrica 26, 1958, pp. 24-36.

2 Pitt, op. cit.

3 J. McDonald and R. Moffit, "The Uses of Tobit Analysis," The Review of Economics and Statistics 62, 1980, pp. 318-21.

(23) $\partial Q/\partial X = F(z)\ (\partial Q^*/\partial X) + E(Q^*)\ (\partial F(z)/\partial X)$.

The total change in Q is composed of the change in Q conditional upon Q being above the limit, weighted by the probability that it is above the limit, plus the change in that probability weighted by the expected value of Q, if it is above the limit.

Tobit estimations are constrained in that the same parameters must be used to explain both entry and quantity responses. If one is particularly interested in identifying the determinants of entry, one can employ a method introduced by Heckman for labor participation studies.[1] The method uses a probit which is designed for zero-one predictions for estimating entry, and an OLS modified to account for the selection procedure for response conditional upon entry. The two estimates can be combined according to equation 23.

## Case Studies on Non-Consumption

Tobit and probit estimations are nonlinear and relatively expensive. In addition, they are not generally available in packaged programs designed for large cross-sectional surveys. Nor can they be easily modified to include restrictions across equations or seemingly unrelated regressions. While the bias of using OLS when Tobit is

---

[1] J. Heckman, "Sample Selection Bias as a Specification Error," Econometrica 47, 1979, pp. 153-161. Another explanation of the technique is presented in Z. Griliches, B. Hall, and J. Hausman, "Missing Data and Self Selection in Large Panels," Annales de L'INSEE 30-31, 1978, pp. 137-176. For an application of Heckman's probit, or Hobbit, to food demand see H. Alderman and J. von Braun, The Impact of the Egyptian Food Ration and Subsidy System on Income Distribution and Welfare, IFPRI Research Report 45 (Washington, D.C., 1984).

appropriate can be calculated in terms of the true parameters and variance, Tobit itself involves some assumptions about the appropriate model. It is worthwhile, then, to investigate the degree of bias in the parameters of OLS relative to Tobit. Two data sets were used to illustrate this. One is the subsample of the Philippine data set discussed above. Although logarithmic regressions are not ideal for a Tobit procedure, they were used in this example to remain as close as possible to Bouis' original analysis.[1] The other data set was a subset of the 1980/81 Labour Force and SocioEconomic Survey by the Department of Census and Statistics, Government of Sri Lanka.[2] For this example, data for Colombo and Mattara districts were used (16 percent of the total sample). Tables 2.3 to 2.5 present coefficients from OLS, Tobit, and Probit/OLS combined.[3] The OLS estimates include the zero value observations. OLS estimates on only positive observations are not reported for the parameters and are misleading even if entry is random, as indicated by equation (23). If entry is random and $\partial f(z)/\partial x$ is zero, it would still be necessary to weight the OLS parameters by the probability of consumption to give the true popula-

---

1 H. Bouis, op. cit. While standard Tobit procedures cannot handle negative values that may occur if the logarithm of a small positive consumption is taken, it is always possible to rescale the units so that all values are either zero or greater (not equal) to 1. The logarithms will then be positive or undefined. The undefined logarithms can then be set to zero.

2 D. Sahn, op. cit.

3 Tobit derivatives are nonlinear even though the original specification was linear, and the coefficients reported here are, in actuality, the derivative of the Tobit at the mean of all independent variables.

Table 2.3 -- Alternative Estimates for Grain in Mindanao

| | Log RICE | | | | | Log CORN | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | TOBIT | Weighted PROBIT/OLS | | (Total of 3-4)* | OLS | TOBIT | Weighted PROBIT/ OLS | | (Total of 8-9)* |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
| Constant | .21 | -1.47 | -1.63 | 1.62 | (-0.01) | 3.11 | 2.47 | 1.72 | 2.02 | (3.74) |
| Log (Income) | .33<br>(3.68) | 0.44<br>(4.09) | .39<br>(3.90) | .01<br>(9.73) | (.39) | -.30<br>(3.47) | -0.33<br>(3.32) | -.33<br>(3.40) | -.03<br>(1.14) | (-.33) |
| Log (Rice Price) | -2.19<br>(5.46) | -2.36<br>(5.14) | -2.01<br>(4.42) | -.47<br>(4.44) | (-2.48) | 2.28<br>(5.88) | 2.57<br>(5.60) | 2.61<br>(5.72) | .17<br>(1.32) | (2.61) |
| Log (Corn Price) | .71<br>(2.43) | 0.82<br>(2.41) | .73<br>(2.23) | .17<br>(1.82) | (.90) | -.74<br>(2.63) | -0.71<br>(2.19) | -.77<br>(2.50) | -.14<br>(1.65) | (-.91) |
| Round | .80<br>(6.50) | 0.83<br>(6.10) | .82<br>(6.07) | 0.04<br>(1.44) | (.82) | -.78<br>(6.65) | -0.93<br>(6.47) | -.71<br>(5.83) | -.11<br>(2.70) | (-.82) |
| $R^2$ | .14 | | | .06 | | .14 | | | .04 | |
| Number | 779 | | | 412 | | 779 | | | 504 | |
| Mean | 1.58 | | | 2.99 | | 1.92 | | | 2.98 | |
| Number of Non-Zero Observations | | 412 | | | | | 504 | | | |

* Variables with a t value less than 1.645 ($p < .10$ cutoff) were not included in the total.

Table 2.4--Alternative Estimates for Grain in Visayas

| | Log RICE | | | | | Log CORN | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | TOBIT | Weighted PROBIT/OLS | | (Total of 3-4) | OLS | TOBIT | Weighted PROBIT/ OLS | | (Total of 8-9) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
| Constant | 1.09 | .214 | .18 | 1.81 | 1.99 | 2.78 | 1.67 | 1.18 | 1.55 | 2.73 |
| Log (Income) | .27<br>(2.90) | 0.30<br>(2.79) | .21<br>(2.33) | .05<br>(2.04) | .25 | -.35<br>(3.73) | -0.37<br>(3.67) | -.32<br>(3.38) | -.32<br>(1.06) | (-.33) |
| Log (Rice Price) | -1.73<br>(6.43) | -1.94<br>(6.07) | -1.71<br>(5.83) | -.20<br>(2.59) | -1.91 | 1.75<br>(6.48) | 1.90<br>(6.33) | 1.99<br>(6.35) | .02<br>(0.21) | 1.99 |
| Log (Corn Price) | .50<br>(1.47) | 0.61<br>(1.52) | .51<br>(1.50) | .03<br>(0.33) | .0 | -.36<br>(1.05) | -.36<br>(0.97) | -.33<br>(0.95) | -.04<br>(0.48) | .0 |
| Round | .19<br>(1.68) | 0.22<br>(1.65) | .15<br>(1.27) | .04<br>(1.35) | .0 | -.30<br>(2.66) | -.33<br>(2.52) | -.29<br>(2.50) | -.03<br>(0.83) | -.29 |
| $R^2$ | .06 | | | .03 | | .08 | | | .01 | |
| Number | 878 | | | 587 | | 878 | | | 430 | |
| Mean | 2.04 | | | 3.07 | | 2.90 | | | 2.98 | |
| Number of Non-Zero Observations | | 587 | | | | | 430 | | | |

Table 2.5--Alternative Estimates for Food in Sri Lanka

| | Q Bread | | | | | Q Milk | | | | | Q Meat | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | OLS | TOBIT | Weighted PROBIT/OLS | | (Total of 3-4) | OLS | TOBIT | Weighted PROBIT/ OLS | | (Total of 8-9) | OLS | TOBIT | Weighted PROBIT/ OLS | | (Total of 13-14) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (11) | (12) | (13) | (14) | (15) |
| Log (Expenditure) | 21.3 | 19.9 | 9.2 | 12.5 | 21.7 | 14.4 | 10.4 | 8.8 | 16.7 | 25.5 | 15.8 | 5.1 | 4.7 | 5.50 | 10.2 |
| | (5.17) | (4.58) | (12.59) | (3.36) | | (6.62) | (7.28) | (6.95) | (6.9) | | (13.58) | (8.79) | (5.6) | (8.8) | |
| Log (Price Dependent) | 12.9 | 4.2 | -24.0 | 22.0 | 0.0 | -48.3 | -24.4 | -6.4 | -34.9 | -41.35 | -16.1 | -6.7 | -9.4 | -3.8 | -12.8 |
| | (0.35) | (0.11) | (0.89) | (0.79) | | (19.97) | (16.59) | (4.25) | (16.95) | | (3.93) | (4.25) | (4.04) | (2.9) | |
| Log Price Dependent)* Poor Dummy | -28.8 | -38.3 | -13.9 | -13.1 | -13.9 | 2.0 | -1.9 | -.16 | -3.5 | -3.7 | 3.8 | 0.0 | 0.1 | 1.2 | 1.2 |
| | (3.90) | (4.80) | (2.80) | (1.26) | | (1.74) | (2.14) | (2.73) | (2.52) | | (5.84) | (0.01 | (0.03) | (1.79) | |
| Log (Rice Price) | 10.8 | 7.8 | 0.22 | 9.6 | 0.0 | Log Coconut Price 3.9 | 5.1 | 3.0 | 9.3 | 9.3 | Log Fish Price -1.3 | -.3 | -.6 | -.7 | 0.0 |
| | (1.1) | (0.93) | (0.03) | (1.23) | | (0.93) | (1.69) | (1.51) | (2.11) | | (1.04) | (0.44) | (0.98) | (0.74) | |
| Urban Dummy | 33.6 | 38.2 | 21.9 | 7.3 | 21.9 | 5.5 | 5.0 | 3.6 | 8.4 | 13.0 | 2.70 | 4.5 | 6.03 | .06 | 6.0 |
| | (8.3) | (8.93) | (7.98) | (0.76) | | (2.74) | (3.63) | (3.80) | (3.97) | | (2.46) | (6.15) | (6.71) | (0.[illegible]) | |
| $R^2$ | .28 | | | .18 | | .41 | | | .43 | | .29 | | | .42 | |
| Number | 786 | | | 620 | | 786 | | | 464 | | 786 | | | 174 | |
| Mean | 74 | | | 94 | | 12 | | | 20 | | 5.9 | | | 26.7 | |
| Number of Non-Zero Observations | | 620 | | | | | 464 | | | | | 174 | | | |

Note: Other household size and composition variables were also significant but not reported here.

tion parameters. Of course, for a double-log equation the elasticities would still be correct, by the process of corresponding errors, but the marginal propensities would be incorrect.

Results from the Philippine study, which includes only observations from the poorest quartile, show little difference between the OLS and the Tobit results or between these results and the weighted sum of the probit and conditional OLS (equation 23). Most differences are surely within acceptable margins of uncertainty. But there are some features of this data that rule against any generalization. Notice, for example, that the Tobit elasticities are the same as the entry elasticities from the probit weighted by average consumption if it is nonzero.[1] Virtually all the measured response is from entry.

This led to some problems with the estimates of OLS conditional upon entry. The use of the inverse of Mills' ratio in such a study works best when the determinants of entry differ from those of conditional response. When the determinants are the same or similar, the Mills' ratio term is apt to be collinear. In the regressions on Philippine data, the information contained in that term swamped the small amount of information in the response equation (note the low $R^2$ values) and the coefficients were implausible, probably reflecting only residual curvature partially accounted for by the Mills' ratio term is included. The Philippine results reported here do not include

---

[1] This result differs from McDonald and Moffit's use of a decomposition of a Tobit in that in their study the decomposition was constrained to follow the restrictive assumption of a Tobit while the current decomposition uses separate estimation to derive the results.

the Mills' ratio. The results for Sri Lanka do include it in two out of the three equations. In the third, the meat equation, the inclusion rendered almost all terms insignificant without an appreciable change in $R^2$. A regression with a $R^2$ of 0.42 and low T values is a sign of multicollinearity, which for the purposes here was dealt with by excluding the Mills' term.

One notes that the OLS (whole sample) and Tobit coefficients for the Sri Lanka data (using a semi-log function) differ more than in the Philippines data. For example, the bread own-price elasticity for the poor is a third higher in the Tobit estimate, and the overall milk price elasticity is half as large with the Tobit as the OLS and 40 percent as large in the meat equations. The latter estimate shows marked differences with regard to income response as well. While the Tobit has a theoretical advantage in a comparison with OLS, the Probit/OLS method has an equally valid theoretical basis. In the comparison mentioned of meat and milk, the two-step approach appears midway between the Tobit and OLS. It appears different than either approach in regard to the poor's bread response, mainly because the method of adding the responses used here set the OLS value to zero. But if it were taken at the observed value, the weighted sum would be close to the OLS coefficient.

Even with this small set of equations, no general pattern emerges. Clearly the researcher risks major differences in reported response if the number of zero values is large. (There is, of course, no gain to a Tobit if nearly every family consumes the good and, indeed, the algorithms may not even work then.) Tobits are, however, more expen-

sive than other methods and, as mentioned, difficult to use in systems. They are not advised for exploratory work.

One approach that may be worthwhile for single equation estimates is a comparison of a general OLS equation with a linear (but biased) entry equation plus a conditional OLS equation, combining the two as in equation (23). If these equations differ only slightly, one can use OLS on the whole sample. Otherwise, the choice is between Tobit or the two-step approach. The latter is particularly recommended if the analyst is interested in determinants of entry, especially if the determinants are exogenous market structure variables that are not included in the conditional response equation. It is not necessary that the probit use the same variables or functional form as the second step. It was done here mainly for illustration. In one use of the two-step approach, the method distinguished between fixed cost for market entry and variable costs upon entry.[1] The two-step approach provides more information but does not have the pedigree of the Tobit, being originally designed for labor questions, and more work on the behavior of the Mills' ratio term and the heteroskedasticity it may introduce may be useful.

The problem of zero-valued observations is avoided when the data includes cell means for districts or income groups. Such aggregations generally include a number of positive values in each cell. If a researcher's concern is with average response to changes in income and prices, such information is sufficient. If one has a particular

---

[1] H. Alderman and J. von Braun, op. cit. Note that this study looked at the response to market times as well as cash prices.

interest in the determinants of entry, cell means that include the percentage of non-zero observations within the cell are enough for the application of logit or probit models.

Aggregation in cross-section data is, of course, theoretically no different from aggregation in time series. It has the advantage of improving the apparent statistical fit by averaging out random shocks within a cell,[1] but can fall victim to aggregation bias as discussed above. Furthermore, when different cells contain different numbers of households, cell means can introduce heteroskedasticity. This form of heteroskedasticity is based on the statistical theorem that the variance of the mean is inversely proportional to the number of observations. One then needs to weight each observation by the square root of n for an efficient estimation. Whether dealing with cell means or individual observations, heteroskedasticity may also be present if the error terms vary by income group or size of purchase. Prais and Houthakker discuss one such form -- when the variance is proportional to the square of the predicted value -- but more general patterns may be discerned.[2]

Data may occasionally be available in a form that retains properties both of time series and cross-sectional disaggregation. In some

---

1 This is only apparent. Estimations from grouped data are always less efficient than for disaggregated data. See J. S. Cramer, "Efficient Grouping, Regression, and Correlation in Engel Curve Analysis," Journal of American Statistical Association, 1962.

2 For a discussion of heteroskedasticity, see G. Judge, W. Griffith, R. Hill, and T. Lee, The Theory and Practice of Econometrics, Chapter 4 (New York: Wiley and Sons, 1980). Also, H. White, "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity," Econometrica 48, 1980, pp. 819-838.

surveys, groups of consumers or panels are interviewed at regular intervals. With repeated observations on the same individuals over time, either changes in taste can be modelled endogenously or individual taste differences can be removed using covariance techniques for pooled cross-sections and time-series data, or by using first differences. Such data, however, are expensive to collect and are rarely available.

## 3. CRITERIA FOR DISTINGUISHING GROUPS FOR PARAMETER ESTIMATION

This review has been prompted by the need to identify the demand parameters of subsets of a population for use in food policy analysis. It is worthwhile, then, to consider how the criteria for subdividing the population are determined. Such criteria are needed regardless of the functional form; they are needed whether or not a systems approach is used, and whether the subdivisions will be applied with covariance analysis or with separate estimations on separate samples. Even when the parameters vary continuously rather than discretely, any tabular form of presentation or scenario used to present the findings to policymakers will report the parameters by subgroups. The division is, of course, influenced by the data availability and expected application, but frequently the choice is less clear-cut than a division by a discrete characteristic such as region, employment category, or sex of the household head.

As the goal here is to formulate nutritional policy, one logical choice is to subdivide the sample by calorie consumption and estimate separate parameters for severe or moderately calorie-deficient households. There are, however, statistical and practical limitations to this.

From a statistical standpoint, subdividing the sample using the dependent variable for truncation introduces the same nature of bias that stimulated the development of the Tobit model. Suppose that

Observed Kcal = True Kcal $\pm$ error. Presume this error is recording error and equal for 11 levels of Kcal, distributed normally with mean at zero. Suppose also that the analyst picks a cutoff at Kcal*, as in Figure 1. Given the normal distribution, if Kcal* is less than $\bar{K}y$, more cases in which Observed Kcal < Kcal* < True Kcal will be included in the subsample than cases in which Observed Kcal > Kcal* > True Kcal will be excluded. Under such conditions, the error in the truncated sample is not distributed normally and has a mean value less than zero. Furthermore, the distribution of the error will be heteroskedastic as it will vary by income. This is so because $\bar{K}y$ increases with income while Kcal* is a constant. The bias in the error due to truncation will vary by income groups even though, under the original assumptions, the error was unaffected by the level of Kcal.

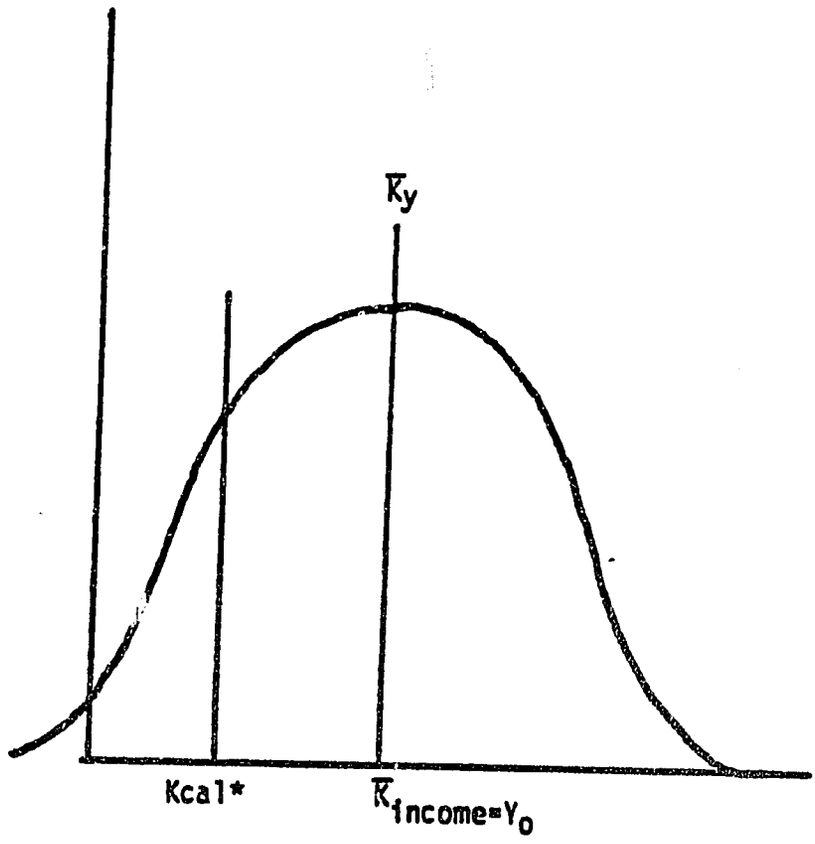
The discussion has been in terms of measurement error but, given the known importance of individual variation of calorie requirements, it can be restated.[1] A sample divided by calorie consumption, in effect, aggregates two types of consumers. Suppose the population has four subgroups, low-income consumers with and without high biological needs for food and high-income consumers with and without high biological needs for food. (Family aggregation complicates matters, but is tangential here. As changes in activity change requirements, the term is used here in the sense of what would be consumed if activity were

---

[1] The importance of individual variation is discussed by a number of authors. See G. Beaton, "Energy in Human Nutrition: Perspective and Problems," Nutrition Reviews, 1983, pp. 325-40. Beaton also makes the point that requirement is a normative concept.

Figure 1

Truncation of Normally Distributed Calorie Variable



not curtailed because too few calories were available). The low Kcal group will include individuals with high requirements for food but limited resources and individuals with low demand but adequate resources. The income elasticities generated reflect what happens to consumption in the group when income increases and will be an average of the low response of the low-requirement/high-income group and the average response of the bulk of the low-income group. This is, again, a missing variable bias, for a variable denoting requirements is missing. One can expect $\partial$Kcal/$\partial$requirements > 0 and, since the truncated sample includes mainly high-income/low-requirement individuals and low-income individuals with varied requirements, requirements in the truncated sample will be negatively correlated with income (but not in the overall sample). This will bias income elasticities downward.[1] An analogous point can be made with price responses in which the truncated sample includes both high- and low-requirement groups at high prices and only low-requirement groups with low prices. In such a case, requirements and price in the truncated sample will be positively correlated. While missing variable bias is a general problem with estimation, the essential point of this discussion is that truncation by the value of the dependent variable can exacerbate the problem by introducing correlation of the missing variable and the independent variable where none had existed before. The same kind of problem exists if, instead of separate regressions on a subsample, different parameters are estimated by covariance techniques.

---

[1] Note that this argument applies to a model of the form $Q = aY + \beta P + CReq$. If the requirements enter as an interaction as well, $Q = aY + \beta P + CR + a2Y*R + \beta^2 P*R$, the argument is more complex.

More commonly, the analysis is done on individual commodities rather than calories. The problem using a Kcal cutoff is, then, somewhat mitigated, but as Kcal is actually a linear transformation of the sum of the commodities, the error in a few commodities will dominate the error in Kcal and the truncation can introduce estimation errors.

Case Study: Sample Truncation

The studies by Williamson-Gray and Trairatvorakul give elasticities by income groups and by calorie groups.[1] Most income elasticities in the former are lower in the calorie-deficient group than in the same proportion of poor individuals, while the pattern with own-price responses is more mixed. The latter study does not report the estimates in a manner that is directly comparable, but the average elasticities taken from the calorie truncated sample are generally lower -- the ones for rice are disturbingly lower -- than in the sample divided by income groups. While such comparisons are in accord with the above discussion, they cannot indicate the degree of bias in that the true parameters are unknown. Accordingly, a monte carlo type of experiment was conducted in which calorie intakes were generated using known parameters plus a random error. After division of the sample, the parameters were estimated and compared to the known values. The sample consists of 200 observations with incomes equally divided in the integers from 1 to 10. Price was randomly distributed

---

[1] Op. cit.

about a mean of 1 with a standard deviation of 0.067 in one set of experiments and 0.15 in another. The calorie-generating equation was Kcal = 1,477 + 600Log Income - 800Log Price + u, if income was less than or equal to 3 (30 percent of sample) and Kcal = 1,477 + 600Log Income - 600Log Price + u otherwise. The term u was generated at random with a mean of zero and a standard deviation of 166.67. The results are presented in Table 2.6.

A number of points are noteworthy. First, all but one of the income parameters in the Kcal truncated sample differ from the true values at the 5 percent level, while none of the whole-sample estimates and only one of the income-truncated estimates are statistically different. Secondly, although the small price variation in the first experiment seems insufficient for reliable estimates, the distortions in the price parameters in the two subsamples differ with no obvious pattern. A direct comparison is not possible because the construction of the data, using income to determine the price parameter value, favors the income subsample. Both subsamples yield price parameters more useful for policy application than the complete sample with dummy variables. Using a dummy variable for either low income or calorie adequacy, one would conclude that there is no significant difference in population parameters in 3 of the last 5 experiments, when in fact a real different was constructed. Furthermore, if one ignores insignificant parameters for the dummy variable one has price parameters for the low-income group of -618, -671, -923, -1,012, and -514. Including even insignificant parameters would give -809, -920, -923, -1,012, and -670, which would be closer to the real values. The

Table 3.1--Table Results of Monte Carlo Experiment on Sample Groupings

| Parameter* | Subsample KCAL < 2000 | | | | Subsample Lowest 30% Income | | | Whole Sample With Dummy for Price * Income < 3 | | | | Whole Sample With Dummy for Price * Calorie Adequate | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | Income | Price | N | $R^2$ | Income | Price | $R^2$ | Income | Price | Dummy | $R^2$ | Income | Price | Dummy | $R^2$ |
| Standard Deviation of Price Index = .067 | | | | | | | | | | | | | | | |
| Experiment 1 | 506 | -897 | 42 | .74 | 607 | -1172 | .78 | 612 | -815 | -356 | .88 | 610 | -1330 | 226 | .88 |
| | (10.46) | (2.96) | | | (13.83) | (4.11) | | (36.36) | (4.00) | (0.97) | | (36.55) | (1.77) | (0.55) | |
| Experiment 2 | 411 | -360 | 42 | .53 | 567 | -446 | .68 | 602 | -544 | 125 | .87 | 601 | -455 | -28 | .87 |
| | (6.47) | (0.94) | | | (10.56) | (1.24) | | (34.52) | (2.44) | (0.32) | | (33.89) | (0.58) | (0.06) | |
| Experiment 3 | 504 | -596 | 32 | .70 | 700 | -991 | .80 | 582 | -594 | -510 | .85 | 578 | .2449 | 909 | .85 |
| | (8.14) | (0.93) | | | (14.72) | (2.52) | | (32.65) | (2.98) | (1.12) | | (32.02) | (1.72) | (1.24) | |
| Experiment 4 | 431 | -590 | 36 | .49 | 576 | -694 | .70 | 574 | -391 | -295 | .86 | 574 | -1068 | 314 | .86 |
| | (5.42) | (1.04) | | | (11.53) | (1.95) | | (34.20) | (1.81) | (0.75) | | (33.50) | (1.04) | (0.58) | |
| Experiment 5 | 383 | -817 | 36 | .64 | 595 | -955 | .73 | 618 | -631 | -317 | .85 | 616 | -1185 | 247 | .85 |
| | (7.53) | (2.01) | | | (11.72) | (2.71) | | (33.66) | (2.82) | (0.76) | | (33.67) | (1.09) | (0.43) | |
| Standard Deviation of Price Index = .15 | | | | | | | | | | | | | | | |
| Experiment 6 | 484 | -760 | 37 | .76 | 665 | -849 | .79 | 597 | -618 | -191 | .87 | 596 | -1043 | 201 | .86 |
| | (9.94) | (6.73) | | | (14.17) | (6.93) | | (35.21) | (7.10) | (1.28) | | (35.38) | (3.29) | (1.17) | |
| Experiment 7 | 413 | -755 | 39 | .65 | 576 | -941 | .80 | 601 | -671 | -249 | .87 | 598 | -1202 | 245 | .87 |
| | (7.75) | (4.45) | | | (11.74) | (6.45) | | (35.60) | (7.13) | (1.49) | | (35.48) | (2.93) | (1.12) | |
| Experiment 8 | 454 | -907 | 42 | .66 | 542 | -930 | .78 | 580 | -493 | -430 | .88 | 576 | -1546 | 515 | .88 |
| | (7.76) | (5.55) | | | (11.63) | (7.06) | | (34.71) | (6.33) | (2.83) | | (34.47) | (4.61) | (2.86) | |
| Experiment 9 | 434 | -1033 | 39 | .65 | 567 | -1009 | .75 | 612 | -472 | -540 | .89 | 608 | -2041 | 777 | .89 |
| | (7.50) | (5.74) | | | (12.11) | (6.54) | | (39.54) | (5.67) | (3.28) | | (40.07) | (5.62) | (4.01) | |
| Experiment 10 | 417 | -422 | 43 | .48 | 629 | -665 | .70 | 618 | -514 | -156 | .85 | 616 | -887 | 177 | .85 |
| | (6.01) | (1.69) | | | (11.11) | (3.86) | | (32.24) | (5.63) | (0.85) | | (616) | (1.94) | (0.74) | |

* T values in parenthesis

problem here is that the variations in price are small, although the variation is not markedly less than could be expected from cross-sectional data. The reliability of the parameters when the dummy variable indicates that calorie adequacy is even less. Using this data one would conclude that the price parameters for the malnourished were -1,043, -1,202, -1,546, -2,041, and -841, thereby, in 4 of the 5 experiments, seriously overestimating the potential impact of price changes. Conclusions about the general population would be similarly misleading, particularly if the common practice of assuming nonsignificant interaction parameters to be zero were followed.

One further point of minor interest: while on one hand the $R^2$ values mean little considering the constructed data, their range should put some perspective on the use of such numbers as assessment of regression fit. In this case the true model is known, the income parameter is homogenous for the entire population, the price parameter has little individual variation, and the only source of error is a random term with a standard deviation that is equivalent to less than the caloric value of two tablespoons of cooking oil. Nevertheless, the $R^2$s of some subgroups are below 0.50. Surely the analyst will not be nearly as fortunate in model specification or data accuracy as with this model, and should not be disappointed to observe $R^2$s of 0.2 or 0.3 in a cross-section analysis.

Two general conclusions follow from this small experiment, one stronger than the other. Firstly, truncation of the sample by calorie grouping is ill-advised. Secondly, in this small experiment, with its limited price variability but significant income variability, dividing

the sample by income appears to be a more reliable means of establishing price parameters than the use of covariance techniques. This conclusion may be true only for this case; with smaller income variance or less firmly defined income parameters such a division may make estimates of income parameters difficult. With a large enough sample, however, it may aid in assessing price parameters and may, in addition, minimize quality and other spurious variations in price.

## Real Income and Household Demographics

The most common alternative to subdivision of the sample by caloric intake is subdivision by income. Alternatively, one can subdivide by employment category, region, or sex of household head, depending, in part, on the expected use of the data. For example, if the data will be used for a program to lower the consumer price by improving or subsidizing the transport of food to a region in which post-harvest prices are remarkably high, then a regional breakdown may be sufficient, although the remaining cross-sectional price variability will be limited. More commonly, however, the targeting or planning refers to the "poor" defined either in terms of consumption of certain goods, especially food and shelter, or in terms of income that makes it possible to purchase a certain amount of goods. These two concepts are conceptually different.[1] The former assumes that different goods have social weights that reflect a concept of one-

---

[1] For a discussion of poverty indicators, see M. Lipton, Poverty, Undernutrition and Hunger, World Bank Staff Working Paper 597 (Washington, D.C., 1983). Also, A. K. Sen, Levels of Poverty: Policy and Change, World Bank Staff Working Paper 401 (Washington, D.C., 1980).

dimensional equity -- that is, the view that society considers equity on the consumption of a certain bundle of goods more important than equity overall. The latter view is founded more on the concept of consumer sovereignty, assuming basically that, since a household maximizes its utility subject to price and budget constraints, then it is sufficient to look at the size of income at which nutritional or similar consumption goals can be met by the household. As mentioned, using the former criterion in estimation is likely to create econometric biases. Even when it is administratively more convenient to identify low-consumption groups rather than low-income groups, it may be advisable to estimate income and price parameters by income group and apply them to the low-consumption groups.

If income is used to distinguish subdivisions of the population, then the question remains whether household income, income per capita, or some other standardized income term is more relevant. The potential for misclassification is readily apparent if household income is used; a large poor household may rank above a single individual with a moderate income. The alternative of per capita income, however, is not without problems concerning economies of scale and percentage of children in a household. If economies of scale exist for some commodities then a household with N individuals and M units of earning will be less well off than a household with 2N members and 2M units of earnings. Note that this is true even if there are no economies of scale for the consumption of foods, which are the central focus of the study. If some households have economies of scale in housing or transportation, then they have more disposable income for food items

at the same per capita income. The issue here is the possibility of welfare comparisons based on some normalized income and not the importance of demographic variables in predicting consumption of a given good. The latter use of demographics can be achieved ad hoc, without theoretical underpinnings, by including such variables in any demand equation. If, however, one wishes to go from commodity-specific interest0 to a method of comparing household income, some assumptions about consumer behavior are necessary.[1]

Adult equivalency scales based on relative caloric requirements are commonly used to standardize households. While this distinguishes the different requirements for children and has some justification in aggregate food demand equations, it does not allow for economies of scale (after all, two adult males are exactly twice one) and it has no foundation in utility theory. Utility is not likely to be a function of calorie requirements.

Another means of scaling household income can be based on Engel's law, which holds that the budget share of food declines monotonically with increases in real income. It can be assumed, then, that two households with identical budget shares to food have identical real incomes. This information can be used to compare the real incomes of families with different sizes and composition. More simply, one can subdivide the sample by the total budget share of food, but that risks the truncation bias discussed above. Or, to be more sophisticated, one can

---

[1] For reviews, see A. Deaton and J. Muellbauer, 1980, Chapter 8 and R. Pollak and T. Wales, op. cit.

create weights as done by Deaton.[1] If, for example,

(24) Budget share of food = $w = \beta_0 + \beta_1 \ln(Y/N) + \beta_2(\ln Y/N)^2 + \beta_3 Na + \beta_4 N_C + \beta_5 Na N_C + \beta_6 Na \ln(Y/N) + \beta_7 N_C \ln(Y/N)$

where Na is the total number of adults, $N_C$ is total number of children, and $N = Na + N_C$; then relative to a household reference standard of two children and two adults, standardized income ($Y^*$) can be derived by solving the following equation:

(25) $\beta_2 \ln(Y^*/4)^2 + (\beta_1 + 2\beta_6 + 2\beta_7) \ln Y^*/4 + 2\beta_3 + 2\beta_4 + 4\beta_5 - \beta_1 \ln (y_i/n_i) - \beta_2 (\ln y_i/n^i)^2 - \beta_3 n_a^i - \beta_4 n_c^i - \beta_5 n_a^i n_c^i - \beta_6 n_a^i (\ln y^i/n^i) - \beta_7 n_c^i \ln (y^i/n^i) = 0$

The equation is solved for $Y^*$ given the observed income $y^i$ and observed household members $n_a^i$ and $n_c^i$.[2]

Equation (24) is not necessarily the only estimating equation for such a technique. For example, some of the parameters may not be significantly different from zero and should be dropped, or other specifications of the budget share equation might be employed. Once $Y^*$ is derived one can get correction factors $(Y^*/4)/(y^i/n^i)$ for any number of adults, children, and income. Using this estimating equation, the correction factors will vary by each of the variables, including income, generating a three-dimensional matrix of such factors.

---

1 A. Deaton, "Inequality and Needs: Some Experimental Results from Sri Lanka," Population and Development Review 9, 1983, pp.35-49.

2 The equation is not as formidable as it appears or as it is to type. Once the parameters are identified, it reduces to a single quadratic equation. The larger of the two roots is the correct result.

Although Y* is derived from information about budget shares, it sidesteps the truncation bias because it predicts standardized income from the complete data set. The information can then be used to subdivide the sample for further work.[1] This method is, however, limited by a number of assumptions. Firstly, the household scaling is identified from the estimation equation for the budget share of food, which in actuality contains both the general income standardization and the specific budget reallocation towards or away from food. Both are necessary for accurate prediction but can be different scaling factors. The former standardization should, in principle, be a weighted average of all commodity effects. Simultaneous estimation of all weights, unfortunately, is underidentified even in a complete system of demand equations. The use of food expenditures for standardization of demographic weights, then, is based on the assumption that, since the budget shares of food are high in low-income areas, the family income deflator will be dominated by the food-specific effects. This, then, gives an acceptable approximation.

This method is only one of many that can be shown to lead to different conclusions.[2] Deaton compares the Engel curve weighting with a method that estimates the cost of children by comparing purchases of

---

1 This, of course, suggests a way to get around the bias by subdividing by calories. One could estimate Kcal using all observations and appropriate variables. This predicted Kcal could be used to subdivide the sample.

2 A. Deaton, Three Essays on A Sri Lanka Household Survey, Living Standards Measurement Study, Working Paper 11 (Washington, D.C.: World Bank, October, 1981).

two bundles of goods -- one that children share and another with goods consumed primarily by adults, such as tobacco and alcoholic beverages. It is assumed that the latter purchases serve as indicators of welfare, independent of the number of children and that differences in expenditures on these goods, controlled for nominal income, will reflect the costs of children. Note that this method does not handle the economies of scale of adult members and, as it reflects a partitioning of commodities, indirectly includes an assumption of additivity.

A conceptually different technique is based on the interpretation of the budget constraint $Y = \Sigma PQ$. If one assumes that real income equals the sum of expenditures on goods weighted by the commodity-specific scale, then $Y^* = \Sigma PQ^* = \Sigma\ PQi/(N_a + \lambda_i N_C)$. Furthermore, this can be reformulated as $Y^* = \Sigma P_i^* Q_i$, implying that changes in the household change effective unit costs. (Each child raises the cost of candy relative to beer.) Using such a formulation in nonlinear estimations, it is possible to estimate household scaling, providing that observed prices vary.[1] The approach based on the budget share of food, however, is clearly easier to estimate than a complete nonlinear system and should be sufficient for most purposes.

With scaled income, a cutoff point for the target group is still needed. Occasionally a poverty line is given or is defined in terms of either calories consumed or the ability to consume a certain

---

[1] Pollak and Wales, op. cit. They discuss a technique of varying committed expenditures as household composition varies as well.

amount. Alternatively, one arbitrarily determined percentile can be focused on the poorest quartile or decile, for instance. There is one technique that allows the data to determine the group classifications internally. Rather than focusing on, say, the ith percentile of scaled income, it is possible to allow the parameters to switch at an internally determined point. The basic model, from Quandt, is

$$(26) \qquad \begin{aligned} Q_i &= X\beta 1 + u_x & 0 < Y < Y^* \\ Q_i &= X\beta 2 + u2 & Y^* < Y < \infty \end{aligned}$$

where Y is some classification variable, for example, percentage of a poverty index.[1] The maximum likelihood algorithm will search for the best switching point. This will give disjointed rather than continuous changes of parameters. It cannot, unfortunately, rigorously test whether $\beta 1$ and $\beta 2$ are statistically different, although Akin et al. used such a method to determine a stratification and then made separate regressions on both strata. They then used a Chow test to determine statistical differences.

On the other hand, testing for statistical differences using covariance techniques, is fairly straightforward. One can create dummy variables for any group and use a multiplication interaction in the regression for dummy x price. The test whether the derivative with respect to price for that group differs from the general population will be the t-statistic for the parameter of the interaction variable. The total response of the group will be the sum of the

---

1 R. Quandt, "A New Approach to Estimating Switching Regressions," Journal of the American Statistical Association 67, 1972, pp. 306-16; for application to food policy, see J. Akin, D. Guilkey, and B. Popkin, op. cit.

general population response and the coefficient of the interaction term, while the variance will be the sum of the variance of the two parameters and twice their covariance. That is, $var(\beta_1 + \beta_2) = S^2(S_{11} + S_{22} + 2S_{12\mu}$, where $S^2$ is the estimate of the variance (the sum of squares/(n-k-1) when k is the number of parameters) and $S_{11}$, $S_{22}$, and $S_{12}$ are the elements of the covariance matrix of parameters. Alternatively, one can include interaction terms for all groups and then exclude an independent price term. The problem of collinearity when dummy variables for all subgroups are included will be avoided, as the intercept will not be a linear combination of the interactions. This gives the price parameter for each group; the difference between the price parameters can be tested, as above, with the standard error derived from the covariance of the parameters.

# 4. OTHER ISSUES

## Variation of Parameter by Source of Income

Frequently one is interested in differences in consumption patterns from different sources of income rather than by level of income. Estimates that distinguish the source of income may be used to analyze the effects of transfer programs,[1] patterns of consumption from earnings of different household members,[2] or marginal propensities to consume from cash as opposed to agricultural or in-kind earnings.[3] Not infrequently such tests fail to specify the constrained utility maximization that is used to generate the difference in the marginal propensity to consume (MPC) by source of income.[4] Consequently, with such an omission, the tests at best fail to improve our understanding

---

1 D. West and D. Price, "The Effects of Income, Assets, Food Programs, and Household Size on Food Consumption," American Journal of Agricultural Economics 58, 1976, pp. 725-730; S. Kumar, Impact of Subsidized Rice on Food Consumption and Nutrition in Kerala, IFPRI Research Report 5 (Washington, D.C., 1979).

2 J. Kinsey, "Working Wives and the Marginal Propensity to Consume Food Away from Home," American Journal of Agricultural Economics 65, 1983, pp. 10-19.

3 B. Massell, "Consistent Estimation of Expenditure Elasticities from Cross-Sectional Data on Households Producing Partly for Subsistence," Review of Economics and Statistics 51, 1969, pp. 136-42.

4 For example, one must be careful to avoid an implicit assumption of separable utility from items consumed from the two sources of income.

of the determinants of behavior. At worst, the tests reflect only statistical artifacts of misspecified functional forms.

Suppose that one source for the differences in MPC from different sources of income reflects imperfect knowledge about the households' budget constraint. For example, suppose one decisionmaker with the objective of maximizing the utility from three categories of goods

(27) $\max u = f(X_1, X_2, X_3)$

subject to $P_1X_1 + P_2X_2 = Y_1$, and $P_3X_3 = Y_2$; and where one can distinguish $X_2$ and $X_3$ conceptually but not in practice, so that analysis of demand investigates consumption of $q_1 = X_1$ and $q_2 = X_2 + X_3$. Such a situation may arise if $Y_2$ refers to on-farm production of foods also available on the market or to food stamps. Since $q_2$ can be purchased with $Y_1$ or $Y_2$ currency, the two currencies are convertible. The concern in this case is the effective exchange rate of $Y_2$ to $Y_1$. This includes transaction costs or rates of discounts of $Y_2$. The appropriate demand function is $q = f[(Y_1 + \lambda Y_2), P]$ where $\lambda$ is the discount rate. If the transaction costs are fixed costs of market entry rather than costs per unit of $Y_2$, then $q = f[(Y_1 + Y_2 - \lambda), P]$. The problem is compounded by the frequent lack of a direct observations of $Y_2$. If $Y_2$ is estimated by $P_3X_3$, the usual problem of errors in variables must be considered.[1]

A variant of this hypothesis would be when one source of income is perceived as more transient than the other. For example, if a

---

[1] Massell, op. cit.

transfer program is expected to terminate in a brief time.[1] The more transitory income would then be discounted relative to the base income for the purpose of consumption.

This is essentially a measurement problem and not a behavioral problem with which it is sometimes confused. Consider a condition in which the objective is

(28) $\max u = u_{(z)} (X_1)$

where $u_{(z)}$ indicates that preferences are endogenously determined by z, and subject to $P_1 X_1 = (Y_1 + Y_2)$. Consider, for example, tastes for opera, which are likely a function of past exposure. Now suppose that $z = z_1$ when $Y_2 = 0$ and

$z = z_2$ when $Y_2 > 0$ or a more general relationship, $z = f(Y_2)$. This may come about when transfers (food stamps, feeding programs) are accompanied by educational programs or when participation in a program heightens consumer awareness of certain commodities similar to the manner that name recognition and other advertising techniques are assumed to shift demand. Various forms of moral suasion and community pressure may be relevant here. The test for the effect of sources of income or demand in this case would to be to test if $\partial q_1{}^2/\partial(Y_1 + Y_2)\partial Y_2 = 0$.

Tests for differences of sources of income on MPC are sensitive to the specification of the Engel form. Sources of income and total income are frequently correlated because of, for example, eligibility

---

[1] J. Hausman and D. Wise, "Discontinuous Budget Constraints and Estimation: The Demand for Housing," Review of Economic Studies XLVII, 1980, pp. 75-96.

requirements for transfer programs or negative correlations of farm and non-farm income. Observed differences in MPC may, then, reflect only the misspecification of the Engel relationship and the correlation of one source of income and an omitted variable.

The linear Engel relationship is clearly a poor choice, as one expects MPC to decline for items such as food as income rises. Yet, as it is most amenable to algebraic manipulation, it will be used to illustrate some differences between the two models discussed above. In the former,

(29) $$q_2 = \alpha + \beta(Y_1 + \lambda Y_2) = \alpha + \beta(Y_1 + Y_2) + \gamma Y_2$$

where $\gamma = \beta(\lambda - 1)$. $\lambda$ is a discount factor, presumably between 0 and 1, hence $\gamma$ is presumed to be of opposite sign to $\beta$, and $|\gamma| < |\beta|$.

This constrasts with the case in which

(30) $$q_2 = \alpha + \beta(Y_1 + Y_2) + \gamma[Y_2/(Y_1 + Y_2)],$$

or, in terms of covariance analysis

(31) $$q_2 = \alpha + \beta(Y_1 + Y_2) + \gamma D(Y_1 + Y_2),$$

where D is a dummy variable defined as 1 if the household participates in the program presumed to change MPC and 0 otherwise. In this case, the hypothesis is that $\gamma > 0$.

There are a number of variants to this approach. One could assume that preferences change in a functional manner and introduce an interaction term. For example, D could be the proportion of income in-kind, in which case, from basic algebra, it becomes merely the introduction of in-kind earnings as the second regressor in the linear case of equation (31), or an interaction term in a logarithmic functional form. Alternatively and equivalently, one could formulate the problem as

testing linear restrictions. In the linear model this is simply a test of whether $\beta_1 = \beta_2$ in $Q = \beta_1 y_1 + \beta_2 y_2$, while in a quadratic model it is a tests of whether $\beta_1 = \beta_2$ and $\gamma_1 = \gamma_2 = \gamma_3$ in $Q = \beta_1 y_1 + \beta_2 Y_2 + \gamma_1 Y_1^2 + C_2 Y_1 Y_2 + C_3 Y^2$. This approach is not possible in a logarithmic model as $\beta \log(Y_1 + Y_2) = \beta \log Y_1 + \beta \log Y_2$. Furthermore, one has to be careful that the Engel relationship is not misspecified. One means of dealing with this is to specify the model with a flexible form with income defined as $(Y_1 + Y_2)$.

Having chosen the best form in this method, one can then test the restrictions on the parameters, either with linear constraints or with interaction terms. If the restrictions are accepted, the interpretation is straightforward -- there is no observed difference in MPC from source of income. If, however, the restrictions are rejected, caution need be taken not only to distinguish between alternative hypotheses but also to ascertain whether the sample is homogenous. In transfer programs recipients are frequently chosen on the basis of household characteristics. These characteristics may influence consumption decisions. The problem would be particularly worrisome if participation were voluntary or otherwise indirectly self-selecting. If z in equation (28) is a discontinuous variable, particularly a zero-one dummy indicating participation, it is not always possible to rule out the hypothesis that tastes differ before participation and not because of it. This issue is more relevant for transfer programs than it is for farm income. If z is a continuous function of some measure of degree of participation, years enrolled in the program, for example, then a causal interpretation is more plausible.

Note that neither equation (29) or (30) result in a variant of equation (30) in which $\lambda > 1$. In an example in which $Y_2$ represents farm in-kind earnings, it is not uncommon to assume that at equal income levels a farm household would consume more food than a non-farm household of identical tastes. Such, however, would not occur in the specification of equation (27), but could reflect instead the difficulty of evaluating the proper marginal prices or a form of equation (28).

Equation (29) is estimated, but the true relationship is

(32) $$q = \alpha + \beta(Y_1 + \lambda Y_2) + \gamma P,$$

then there is a missing variable bias. The estimate of the direction of the bias is straightforward. Suppose that self-sufficient farmers consider the marginal price to be the farmgate price while others pay the retail price. Note that the value of P would be the retail price if the marginal purchase of the farm household or of the household receiving in-kind transfers is in the higher priced market. In this situation, households with higher values of $Y_2$ are less likely to have their marginal price in the retail market and, therefore, $Y_2$ is negatively correlated with the missing price variable. As one can expect c to be negative in equation (32), the missing variable bias is positive; the observed value of $\gamma$ from equation (29) will be equal to the expected value of $\gamma + [(Y_2' Y_2)^{-1} Y_2' P]\gamma$.

If marketing margins are known and do not vary by the size of a transaction, they provide an estimate of the average value of $\lambda$. If one has observations of only retail prices, one could estimate

(33) $$q = \alpha + \beta(Y_1 + \lambda Y_2) + \gamma[P-(1-\lambda)P \times D],$$

where D is a dummy variable defined as 1 if the marginal purchase of

the commodity is in the market denoted by $Y_2$ income. Equation (33) or equation (29) could be estimated by a nonlinear technique. In actuality, of course, equation (29) does not require a nonlinear method, but a logarithm or semi-log variant of the Engel curve would.

Issue of Control of Family Income

Recently the assumption that a household can be characterized as having a single ordering of preferences has been challenged.[1] Even within a nuclear family, it is unlikely that all individuals have identical preference ordering. The simplifying assumption that a household has a single purse controlled by one individual does not eliminate the possibility of an internal dynamic that brings about compromise in budget allocations. Each member has the potential to influence the allocation process through combinations of various threats such as physical violence, pouts, tantrums, withholding of favors,[2] etc., and rewards. In order to predict the outcome of a change of food policy, it is not important how the final internal agreement is reached if it is stable.[3] However, the weighting of the preferences of household members in equilibrium may be influenced by changes directly or indirectly affected by policy instruments. If so,

---

1 B. Rogers, "The Internal Dynamics of Households: A Critical Factor in Development Policy," paper submitted to USAID, October 1983; C. Jones, "The Mobilization of Women's Labor for Cash Crop Production: A Game Theoretic Approach," American Journal of Agricultural Economics 65, 1983, pp. 1049-1054.

2 See, for example, Aristophanes' Lysistrata.

3 Stable weights are implied by P. Samuelson, "Social Indifference Curves," Quarterly Journal of Economics 60, 1956, pp. 1-22.

it is possible that demand parameters will vary according to factors that influence bargaining within the household.

This is a difficult issue to study. Only a few researchers have attempted to quantify the effects of one of the factors most commonly assumed to influence the weighting of preferences within a household, namely, the control of income by sexes. Roger's review and Guyer's study of income control in Africa show some of the major difficulties in modeling such behavior.[1] These include the difficulty of defining the household as a single functional unit rather than as a web of obligations and benefits; the difficulty in obtaining information on the income of household members and the transfers between them, and the difficulty of modeling exchanges of services (including refraining from delivering on a threat) that are not monetarized, but contribute to the utility of household members. It is also important to know how information flows between household members. To the degree that one member can conceal income, that member avoids compensating actions by other members. Furthermore, control of income by female members is frequently linked with other variables. Households headed by females are frequently, but not necessarily, poorer or have different dependency ratios than other households. If women in different households have different opportunity costs, either because of differences in education or other determinants of market wages or nonmarket productivity, then the price of time will differ. This, of course, will

---

[1] B. Rogers, op. cit.; J. Guyer, "Household Budgets and Women's Incomes," Working Paper 28, African Study Center (Boston University, 1980).

make goods which are time-intensive in consumption more costly for some households than others, even if cash prices are the same. It is difficult, then, to distinguish preferences from costs in such circumstances.[1]

In some contexts there are differences in the seasonality and lumpiness of female and male incomes which, in and of themselves, influence food purchases and are distinct from differences by sex in propensities to consume. Finally, a difficulty in generalizing about control of income comes because most observations are made for a single culture. No universal rule as widely applicable as Engel's law has yet appeared from the studies of control of income.

As an example of the complexity of such relationships, consider Guyer's observations in Southern Cameroon. At first glance it appears that the society she studies is a near-polar case. Males earn income from different occupations than women and, furthermore, women bear most of the responsibility for purchasing food. However, in this small sample (26 women), 40 percent of the cash income of the married women come from husbands and 3 percent from other males. Women who are not married obviously received no income from husbands but received 17 percent of their income from other males. To determine the effects of different food policies, it is necessary to understand what influences the amount transferred between household members.

---

[1] R. Pollak, and M. Wachter, "The Relevance of Household Production and Its Implications for the Allocation of Time," Journal of Political Economy 83, 1975, pp. 255-277; review the basic full (time included) model of a household and discuss the restriction conditions under which costs and preferences can be distinguished.

Presumably males have an idea of the use women put the money to and may increase or decrease the transfer according to their notions of what a household should consume. While it is appealing to simplify the household by assuming that most of the increments to women's incomes is spent on food, while most of men's earnings is spent on beer and radios,[1] the fact that food purchasing responsibilities are largely given to one sex is insufficient to establish the net effect of transfers. One can imagine a number of scenarios that could moderate the effect of increases of women's earnings on household food availability, even if it were established that women's MPC on food is higher than males. For example, if a male saw female earnings increase, he could decrease the amount transferred. Similarly, increases in male earnings could lead to a greater absolute or even proportional amount of his earnings going to household needs. When purchasing requirements are less distinct and where types of employment are less segregated by sex, this issue is even harder to measure.

In another region of Cameroon, Jones found that there was little difference in either the amount of rice retained by the household or the amount spent on sauce ingredients when females controlled the disposition of the crop or headed the household. Women, however, maximized their personal rather than household income by their labor allocation. If all sources of income contributed equally to the

---

[1] While there is some evidence for such behavior, there are also a number of studies with rather different conclusions. See Rogers, op. cit.

household well-being, this behavior would be inefficient. Even if the consumption effect is not measurable, this information supports the view that marginal propensities differ.

Leaving aside the difficulty of determining how allocative decisions are made within the household, there are econometric problems involved in measuring the net effect of women's earnings.

Suppose a model of the form

(34) $\max u(u_1, u_2)$

where $u1 = f(x_1, x_2)$, $u2 = g(x_3, x_4)$, and subject to $P_1 X_1 + P_2 X_2 = Y_m$

$P_3 X_3 + P_4 X_4 = Y_f$

If $X_1$ and $X_2$ are observationally distinct from $X_3$ and $X_4$ and if utility is additive, then there are no particular problems in measuring demand by source of income. Commonly, however, there is an overlap of commodities and the analyst observes only $q_2 = (X_2 + X_3)$ rather than the good distinguished by purchasing agent. Furthermore, $u = (u_1, u_2)$ is probably not additive and the true relationships are of the form $X_2 = q(Y_m, X_3)$ and $X_3 = q(Y_f, X_2)$. As it makes little intuitive sense to suppose that there are two separate and distinct budget constraints or two distinct margins, a functional form such as $Q = B_m Y_m + B_F Y_F$ has little theoretical foundation, although it may be predictive. More plausible is a model of the form

(35) $\max u(F)(X_1, X_2, X_3, X_4)$

subject to $\Sigma P_j X_j = Y_f + Y_m$, and

where u(F) indicates that the observed (or weighted) preference ordering for the household can be influenced by the percentage of income earned by the women. In principle, this could be tested using

a model that allows the income (and price parameters) to measure the MPC from total income but to vary by a function of women's income. For example,

(36) $Q = a + \beta(Y_f + Y_m) + \beta_2 (Y_f + Y_m) * \lambda,$

where $\lambda = (Y_f/Y_f+Y_m)$.

If the true model were linear, this might reduce to

(37) $Q = a + \beta_1(Y_f + Y_m) + \beta_2(Y_f).$

However, if the model is misspecified, $\beta_2(Y_f)$ may indicate only the misspecification. For example, suppose that the true model was quadratic.

(38) $Q = a + \beta_1(Y_f + Y_m) + C(Y_f + Y_m)^2.$

If the quadratic term is omitted and C is negative, then the observed $\beta 2$ will have the opposite sign of the correlation of $Y_f$ and total income; if women work mainly in poorer households, $\beta_2$ will appear to be positive. Such misleading results are potential in any model in which Engel curves have more curvature than is included in the model or when household composition effects are excluded.

Another approach to the problem is to assume that income is transferred between members but that the purchasing pattern, conditional upon a member's own and transferred income, reflects one individual's preferences. For example, suppose

(39) $Q_f = f(Y_f + \text{Transfer}),$

where Transfer $= g(Y_f + Y_m)$ with or without constraints on negative transfers.

This supposes that males transfer to females according to total, not personal, earnings. If Transfer is known, such a model would be recursive and simple to solve. Otherwise, it reduces to

(40) $Qf = f[Yf + \lambda(Yf + Ym)],$

where $\lambda$ is a constant for all commodities purchased by women. Many variants of such models can be proposed for various social structures. Jones proposes a bargaining model that differs from the above examples. Currently, however, there is little data available that can distinguish between alternative models.

## Case Study: Effects of Women's Earnings on Marginal Calorie Purchases

While a complete model of the dynamics of intrahousehold decisionmaking is beyond the scope of this study, it is worthwhile presenting an illustration of an approach to measuring the marginal impact of income earned by women. This example, adapted from Sahn, provides evidence on both the size and significance of the gender effect in one culture and illustrates some measurement problems encountered in such an analysis.[1]

The basic estimating equation used was of the form

(41) $$CAL = \alpha + b1 LnY + b2 (LnY)^2 + C1 (Y_F/Y)*LnY + d_L Z_L + e_K Z_K * LnY,$$

where:

$CAL$ = calories consumed per adult equivalency unit per day,

$LnY$ = log of income,

$Y_F$ = value of income earner by females,

$Y_F/Y$ = ratio of female to total income = RATIO,

$Z_L$ = the vector of other independent variables,

$d_L$ = the vector of coefficients to be estimated,

$Z_K$ = subset of $Z_L$,

$e_K$ = the vector of coefficients to be estimated.

---

[1] Sahn, op. cit.

Included among the other independent variables were household size (HHSIZE), household size times per capita expenditures (HHSX), the number of income earners (EARNERS), and the ratio of adult equivalency units to household size (AEURAT). In addition, data on the educational achievement of the mother and father, MOMED and DADED, were included because they differ for households in which different shares of income are earned by women. The age of the major income earner was included (AGE1) to control for life-cycle variations. The amount of farmland owned (FARMLAND), and a price index (PRINDEX) were also included.

The total effect of a change in women's income is $\partial Cal/\partial Y_F + \partial Cal/\partial Y * \partial Y/\partial Y_F$. Note that $\partial Y/\partial Y_F = 1$. Since a change in male income is the same as the change in household income, with female income held constant, the second term on the right is also the change in consumption following a change in male earnings. The first term is the effect of female earnings.

The MPC out of female earning, then, is

$$(42) \qquad \frac{\partial CAL}{\partial Y_F} = \frac{\partial CAL}{\partial Y}\bigg|_{Y_M = Y_{M_0}} = \frac{1}{Y}\cdot(b_1 + 2b_2 LnY + C_1 Y_F\left(1 - \frac{LnY}{Y}\right) + C_1 LnY + e_K Z_K$$

where $\partial CAL/\partial Y_F$ is the MPC out of female income,

while $$\frac{\partial CAL}{\partial Y_M} = \frac{\partial CAL}{\partial Y}\bigg|_{Y_F = Y_{F_0}} = \frac{1}{Y}\left(b_1 + 2b_2 LnY + C_1 Y_F\left(\frac{1-LnY}{Y}\right) + e_K Z_K\right).$$

The results of this estimation are presented in Table 4.1 and the elasticities are in Table 4.2. In this form, the difference in $MPC_M$ and $MPC_F$ decline with income (from 0.048 to 0.009) but are constant

Table 4.1 -- Models Used to Estimate the Marginal Propensity to Consume Calories

| Independent Variable | Dependent Variable: Calories/AEU | | | | | |
|---|---|---|---|---|---|---|
| | Model 1 | | Model 2 | | Model 3 | |
| INTERCEPT | -25521.6 | | 24520.6 | | -24570.4 | |
| LnY | 5994.01 | (17.03)[a] | 5741.47 | (16.19) | 5751.75 | (16.22) |
| $(LnY)^2$ | -238.03 | (11.32) | -223.30 | (10.59) | -224.06 | (10.63) |
| PRINDEX | -1744.00 | (26.70) | -1693.16 | (25.73) | -1692.77 | (25.73) |
| HHSIZE | 725.19 | (11.57) | 717.75 | (11.47) | 718.13 | (11.48) |
| HHSX | -95.70 | (11.99) | -95.28 | (11.95) | -95.20 | (11.94) |
| AGE1 | 3.33 | (5.10) | 3.65 | (5.39) | 3.55 | (5.24) |
| AEURAT | 2870.29 | (24.07) | 2873.53 | (24.16) | -2854.01 | (23.94) |
| EARNERS | 5.12 | (0.43) | 10.28 | (0.82) | 4.99 | (0.39) |
| MOMED | -11.77 | (1.27) | -5.61 | (0.61) | -6.93 | (0.46) |
| DADED | -28.07 | (3.34) | -26.80 | (3.18) | -24.90 | (2.87) |
| FARMLAND | 7.17 | (2.40) | 2.75 | (0.89) | 2.66 | (0.86) |
| URB | 2922.60 | (9.48) | 2880.80 | (9.37) | 2894.76 | (9.42) |
| EST | -2302.79 | (3.51) | -2252.67 | (3.44) | -2286.16 | (3.49) |
| URBX | -410.75 | (10.66) | -401.35 | (10.44) | -403.19 | (10.49) |
| ESTX | 327.93 | (3.92) | 322.75 | (3.87) | 326.10 | (3.91) |
| $(Y_F/Y*LnY)$ | 9.64 | (2.04) | 11.12 | (2.29) | 19.62 | (3.16) |
| $(Y_F/Y)^2$ | | | | | -39.73 | (2.19) |
| NMY[1] | | | -485.12 | (4.72) | -484.64 | (4.72) |
| AGY[2] | | | 185.56 | (4.49) | 187.94 | (4.55) |
| FSY[3] | | | -74.45 | (1.09) | -83.83 | (1.23) |
| CPY[4] | | | 126.19 | (1.31) | 125.03 | (0.19) |
| BPY[5] | | | 7.94 | (0.19) | 10.02 | (0.81) |
| $R^2$ | 0.644 | | 0.648 | | 0.648 | |

[a] t-statistics are in parentheses.

[1] Ratio of income from non-money sources (excluding home consumption) to total income
[2] Ratio of income from agricultural production, including home consumption and agricultural sales, to total income.
[3] Ratio of income from food stamps, pensions and remittances to total income.
[4] Ratio of income from periodic cash payments to total income.
[5] Ratio of income from business profits to total income.

SOURCE: Sahn, op. cit.

Table 4.2 -- Marginal Propensities to Consume Calories from Income Earned by Men ($MPC_M$) and Women ($MPC_F$)

| Derivative | Low Expenditure Group[a] | | | Average Expenditure Group[b] | | | High Expenditure Group[c] | | |
|---|---|---|---|---|---|---|---|---|---|
| | % Female Income | | | % Female Income | | | % Female Income | | |
| | 0 | 25 | 80 | 0 | 25 | 80 | 0 | 25 | 80 |
| MODEL 1: | | | | | | | | | |
| ∂CAL/∂RATIO | 71.06 | 71.06 | 71.06 | 75.76 | 75.76 | 75.76 | 86.98 | 86.98 | 86.98 |
| $MPC_M$ | 1.192 | 1.177 | 1.161 | 0.679 | 0.672 | 0.659 | 0.160 | 0.153 | 0.153 |
| $MPC_F$ | 1.237 | 1.222 | 1.206 | 0.708 | 0.701 | 0.688 | 0.171 | 0.169 | 0.164 |
| $MPC_F - MPC_M$ | 0.045 | 0.045 | 0.045 | 0.029 | 0.029 | 0.029 | 0.011 | 0.011 | 0.011 |
| MODEL 2: | | | | | | | | | |
| ∂CAL/∂RATIO | 81.94 | 81.94 | 81.94 | 87.37 | 87.37 | 87.37 | 100.12 | 100.12 | 100.12 |
| $MPC_M$ | 1.171 | 1.160 | 1.135 | 0.674 | 0.666 | 0.650 | 0.162 | 0.159 | 0.153 |
| $MPC_F$ | 1.223 | 1.212 | 1.187 | 0.708 | 0.700 | 0.684 | 0.174 | 0.171 | 0.165 |
| $MPC_F - MPC_M$ | 0.052 | 0.052 | 0.052 | 0.034 | 0.034 | 0.034 | 0.012 | 0.012 | 0.012 |
| MODEL 3: | | | | | | | | | |
| ∂CAL/∂RATIO | 144.44 | 124.59 | 8.92 | 153.99 | 134.14 | 90.47 | 176.79 | 156.94 | 115.39 |
| $MPC_M$ | 1.171 | 1.155 | 1.077 | 0.674 | 0.663 | 0.651 | 0.162 | 0.158 | 0.154 |
| $MPC_F$ | 1.262 | 1.233 | 1.128 | 0.734 | 0.715 | 0.686 | 0.183 | 0.177 | 0.167 |
| $MPC_F - MPC_M$ | 0.091 | 0.078 | 0.051 | 0.060 | 0.052 | 0.035 | 0.021 | 0.019 | 0.013 |

a Corresponds to 2nd decile of population where yearly per capita expenditures = Rs. 1587
b Corresponds to mean of population where yearly per capita expenditures = Rs. 2583
c Corresponds to 10th decile of population where yearly per capita expenditures = Rs. 8267

SOURCE: Sahn, op. cit.

regardless of the share of female income.[1]

A variant of the model above, model 2, includes variables for sources of income such as percent from nonagriculture in-kind (NMY), from agriculture (AGY), and from transfers (FSY). The result of this modification is to increase the size of the coefficient of FEMINCR and the difference in MPCs from male and female earnings.

Model 3 is a variation of equation (41) in which the variable $(Y_F/Y)$ is replaced by two terms $(Y_F/Y)*LnY$ and $(Y_F/Y)^2$. This model, which has a slightly better fit, generates values of $\partial Cal/\partial ratio$ that decline as share of female earnings increase, which is logical; going from 10 to 20 percent female earnings has a greater effect than going from 80 to 90 percent. In this form, also, the difference in male and female MPCs is greater when incomes are small than when they are large. Note that both the FEMINCRX and FEMINCR2 coefficients are significant.

## Distribution of Household Purchases

While a family probably has a concept of how food will be distributed when it decides how much food will be purchased, the measurement questions are different. The total quantity of food purchased by the household is not a strong predictor of the effect on nutrition if the distribution among family members is unknown and if it differs greatly between households.

---

[1] This does not imply $\partial Cal/\partial ratio = 0$, holding income constant. In fact, it is 76 calories, which is small considering that this implies a household with all earnings by women consumes 76 calories more than one with equal earnings by males only.

Food distribution within a household may be biased by the sex or age of the recipient. Studies of such biases are reviewed by Lipton and by Carloni.[1] The former concludes that sex-based food discrimination is rare for adults and slightly less rare for children. What evidence exists is mainly from South Asia. As Chen et al. noticed, the effect of sex biases on nutrition reflects differences in the rates of visits to clinics as well as in household food allocations.[2]

There are two approaches to measuring the allocation of food within a household. One is to measure total food consumption and include demographic variables as regressors. This approach, which is discussed in Section 3, allows the data to determine the average household weights for different individuals. A flaw in it is that individual commodity weightings are not always consistent with total income weightings, but the method has been used in various forms with some success.

A second approach measures individual family members' intake, either by recall or direct observation. This approach, obviously, has the advantage of using more information, especially if data on activity, illness, pregnancy, and lactation are recorded.[3]

---

1 M. Lipton, op. cit.; A. Carloni, "Sex Disparities in the Distribution of Food Within Rural Households," Food and Nutrition 7, 1981, pp. 3-12.

2 L. Chen, B. Huq, S. D'Souza, "Sex Bias in the Family Allocation of Food and Health Care in Rural Bangladesh," Population and Development Review 7, 1981, pp. 55-70.

3 R. Chaudhury, personal communication, has observed that apparent biases are less when differences in activities as opposed to average WHO energy requirements are used to standardize food requirements.

While both approaches are useful for food and nutrition policy, the issue relevant to this review is the interrelation with price and income parameters. First, as has already been discussed, such parameters are estimated with more accuracy when demographic information is included. Second, for the application of the parameters to policy, it is necessary to divide the MPC by the share of food going to the target groups. This, of course, will give lower MPCs (but will not affect elasticities), and these will be lower the more narrowly defined the target group.[1]

An important issue to be studied is whether food shares vary during different times of the year or because of some other factors. The most commonly proposed view is that the main male laborers will increase their share as activities increase in peak seasons of the agricultural years. If so, women and children, whose activities also increase at this time, will see a decline in their share of food. Moreover, each share in these periods may be part of a smaller whole because planting seasons generally coincide with periods of high food prices and low household stores.

A related issue is whether shares are constant no matter what the income is or whether increments in food following an increase in income or decline in prices are distributed to or away from already favored members. For example, if a household conceptualizes indivi-

---

[1] For a discussion of leakages in the specific context of supplementary feeding programs, see G. Beaton and H. Ghassemi, "Supplementary Feeding Programs for Young Children in Developing Countries," American Journal of Clinical Nutrition 34, 1982, pp. 864-916.

dual minimum needs and recognizes social aspects of food, it may make efforts to see that all members reach the minimum. After this is achieved, increments may go disproportionally to favored members until some point of diminishing utility, at which time the share to other members will again begin to rise. If shares vary by income group then $\partial Q_i/\partial Y = \partial Q_H/\partial Y * share_i + \partial share_i/\partial Y * Q_H$. As mentioned, $\partial share_i/\partial Y$ may follow a pattern that is complex and difficult to measure.

Finally, distribution may differ if the head of the household is a woman. This latter issue parallels the discussion in the previous subsection. It has, in addition, the problem of measuring distribution when most household demographics are not representative of the general population.

## 5. CONCLUSIONS AND RESEARCH RECOMMENDATIONS

A general pattern of declining absolute value of price elasticities has been observed for five commodity groups using a variety of estimation techniques. The data employed in these estimations is generally cross-section data, but similar patterns and sizes of results were observed using time-series data or combinations of time-series and cross-section data. Furthermore, a pattern of high income elasticities for food expenditures and lower elasticities for calories has been observed for a number of countries.

As a conclusion, it is useful to make two sets of recommendations. One set is intended for planners who intend to apply price and income parameters to food policy analysis. The other is intended for analysts involved in the estimation of such parameters.

There is a need to explore more fully the implications of price elasticities derived from cross-section data. Central to the estimation of price parameter from cross-section data is the problem of quality of commodities. It is not clear how to define or measure quality nor is the nature of the substitution between it and quantity clear. Until both theoretical and empirical work is done in this area, the meaning of cross-section estimates is controversial. A few countries, including India, Indonesia, Sri Lanka, Philippines, Bangladesh, and Egypt, have conducted national household expenditure surveys for a number of years. While this study reports on some

analyses of combinations of time-series and cross-section analysis of such data, the potential is largely untapped. This is, in part, because of the difficulty of getting access to the relevant surveys, particularly in disaggregate form, and because of the cost of analysis.

An alternative and related approach would be to collect panel data. This is expensive, although national coverage with tens of thousands of households need not be undertaken. What is needed, at least initially, are studies of whether responses measured in such a manner differ from cross-sectional or aggregate time-series analysis. This need not be done on a national scale. In this respect, first-difference techniques or techniques of varying parameters could be explored.

Most developing countries intervene in the food allocation process. Many of these interventions create dual price systems, or quantity restrictions and similar market disequilibria. In such situations, consumer demand cannot be characterized in terms of price and income responses alone. Both the econometric consequences and the policy implications of such type of markets are largely unexplored. Recognition of the characteristics of such markets and refinement of techniques to measure demand in such environments, then, should be a priority of applied research on consumer demand.

The returns to increased understanding of the distribution of food within the family and of the dynamics of budget allocation, especially in complex families, are surely high. This requires innovations in data collection and techniques to measure differences in

parameters by source of income or type of household. This is especially true for the new concern with the role of women but applies also to household size and the birth order and sex of infants. Similarly, studies of in-kind transfers are fragmentary and likely to benefit from increased attention.

It is particularly important to include cross-price responses in the analysis of food policy. As aggregation of commodities generally obscures the effect of a number of specific policy instruments, it is important to continue the process of exploring techniques that aid in estimating cross-prices. This includes analysis of flexible functional forms, perhaps using combined time-series and cross-section data and an exploration of Mundlak's technique of principal components mentioned above.

As the consequence of the collinearity of regressors depends in part on the variance of the dependent variable, it is useful to explore estimates of total calories regressed on a set of food prices. These results should be compared to the effect on calories from elasticities estimated from individual commodities.

Such explorations can enable practical policy-oriented research to be closely allied to theory with no sacrifice of applicability.

# BIBLIOGRAPHY

Aaness, J. and Rodseth, A. "Engel Curves and Systems of Demand Functions." European Economic Review 20 (1983), pp. 95-121.

Adrian, J. and Daniel, R. "Impact of Socioeconomic Factors on Consumption of Selected Food Nutrients in the United States." American Journal of Agricultural Economics 58 (1976), pp. 31-38.

Ahmed, R. Agricultural Price Policies under Complex Socio-Economic and Natural Constraints: The Case of Bangladesh. Research Report 27 (Washington, D.C.: International Food Policy Research Institute, 1981).

Akin, J., Guilky, D., and Popkin, B. "The School Lunch Program and Nutrient Intake: A Switching Regression Analysis." American Journal of Agricultural Economics 65 (1983), pp. 477-85.

Alderman, H. and von Braun, J. The Impact of the Egyptian Food Ration and Subsidy System on Income Distribution and Welfare. Research Report 45 (Washington, D.C.: IFPRI, 1984).

Barten, A. "Family Composition Prices and Expenditure Patterns," in P. Hart, G. Mills and J. Whitaker (eds.), Economic Analysis for National Economic Planning. (Woburn, Mass: Butterworth Publishers, Inc., 1964), pp. 277-92.

_____. "The Systems of Consumer Demand Function Approach: A Review." Econometrica 45 (1977), pp. 23-51.

Baylock, J. and Green, R. "Analysis of Flexible Engel Functions." Agricultural Economics Research 32 (1980), pp. 12-20.

Beaton, G. "Energy in Human Nutrition: Perspective and Problems." Nutrition Reviews 41 (1983), pp. 325-40.

Beaton, G. and Ghassemi, H. "Supplementary Feeding Programs for Young Children in Developing Countries." American Journal of Clinical Nutrition 34 (1982), pp. 864-916.

Bennet, M. The World's Food. (New York: Harper, 1954).

Berg, A. Malnourished People: A Policy View. Poverty and Basic Needs Series (Washington, D.C.: World Bank, 1981).

Betancourt, R. "The Estimation of Price Elasticities under Additive Preferences." International Economic Review 12 (1971), pp. 283-92.

Blanciforti, L., and Green, R. "An Almost Ideal Demand System Incorporating Habits: An Analysis of Expenditures on Food and Aggregate Commodity Groups." Review of Economics and Statistics LXV (1983), pp. 511-15.

Brandt, J. and Goodwin, J. "The Impact of Increasing Food Supply in Human Nutrition -- Implications for Commodity Priorities in Agricultural Research and Policy: Comment." Americal Journal of Agricultural Economics 62 (1980), pp. 588-91.

Brown, J. and Deaton, A. "Models of Consumer Behavior: A Survey." Economic Journal 82 (1972), pp. 1145-1236.

Bouis, H. "Rice Policy in the Philippines," Ph.D. Thesis (Stanford University, 1982).

_____. "If There is Curvature in the Slutsky Matrix, What do the Curves Look Like?" (Washington, D.C.: IFPRI, 1983), Mimeograph.

Carloni, A. "Sex Disparities in the Distribution of Food Within Rural Households." Food and Nutrition 7 (1981), pp. 3-12.

Chen, L., Huq, B., and D'Souza, S. "Sex Bias in the Family Allocation of Food and Health Care in Rural Bangladesh," Population and Development Review 7 (1981), pp. 55-70.

Christenson, L., Jorgensen, D., and Lau, L. "Transcendental Logarithmic Utility Functions." American Economic Review 65 (1975), pp. 367-83.

Cramer, J. S. "Efficient Grouping, Regression, and Correlation in Engel Curve Analysis." Journal of American Statistical Association 1962.

_____. "Interaction of Income and Price in Consumer Demand," International Economic Review 14 (1973), pp. 351-63.

Deaton, A. "Inequality and Needs: Some Experimental Results from Sri Lanka." Population and Development Review 9 (1983), pp. 35-49.

_____. "Involuntary Savings Through Unanticipated Inflation. American Economic Review 67 (1977), pp. 899-910.

_____. "A Reconsideration of the Empirical Implication of Additive Preferences." Economic Journal 84 (1974), pp. 338-48.

_____. "Specification and Testing in Applied Demand Analysis." Economic Journal 88 (1978), pp. 524-36.

_____. "The Measurement of Income and Price Elasticities." European Economic Review 6 (1975), pp. 261-74.

_____. "Theoretical and Empirical Approaches to Consumer Demand Under Rationing." in Deaton, A., (ed.) Essays on the Theory and Measurement of Consumer Behavior, in honor of Sir Richard Stone (New York: Cambridge University Press, 1981), pp. 55-72.

_____. Three Essays on A Sri Lanka Household Survey. Living Standards Measurement Study, Working Paper 11 (Washington, D.C.: World Bank, 1981).

Deaton, A. and Muellbauer, J. "An Almost Ideal Demand System." American Economic Review 70 (1980b), pp. 312-326.

_____. Economics and Consumer Behavior. (New York: Cambridge University Press, 1980).

De Janvry, A., Bieri, J., and Nunez, A. "Estimations of Demand Parameters Under Consumer Budgeting: An Application to Argentina." American Journal of Agricultural Economics 54 (1972), pp. 422-30.

Dixon, J. Food Consumption Patterns and Related Demand Parameters in Indonesia: A Review of Available Evidence. Rice Policies in Southeast Asian Project Working Paper # 6 (Washington, D.C.: IFPRI, 1982).

Edirisinghe, N. and Poleman, T. "Behavior Thresholds as Indicators of Perceived Dietary Adequacy and Inadequacy," International Agricultural Economics Study # 17 (Ithaca, N.Y.: Cornell University, 1983).

Frish, R. "A Complete Scheme for Computing all Direct and Gross Demand Elasticities in a Model with Many Sectors." Econometrica 27 (1959), pp. 177-96.

George, P. S. and King G. Consumer Demand for Food Commodities in the United States with Projections for 1980. Giannani Foundation Monograph # 26 (Davis, CA: University of California, 1971).

Goldman, S. and Uzawa, H. "A Note on Separability in Demand Systems." Econometrica 32 (1964), pp. 387-98.

Griliches, Z., Hall B., and Hausman, J. "Missing Data and Self-Selection on Large Panels." Annals de L'INSEE 30-31 (1978), pp. 137-176.

Guyer, J. "Household Budgets and Women's Incomes." Working Paper # 28, African Study Center, (Boston, Mass: Boston University, 1980).

Hausman, J. and Wise, D. "Discontinuous Budget Constraints and Estimation: The Demand for Housing." Review of Economic Studies XLVII (1980), pp. 75-96.

Hazell, P. and Roell, A. Rural Growth Linkages: Household Expenditure Patterns in Malaysia and Nigeria. Research Report 41, (Washington, D.C.: IFPRI, 1983).

Heckman, J. "Sample Selection Bias as a Specification Error," Econometrica 47 (1979), pp. 153-161.

Houthakker, H. "An International Comparison of Household Expenditure Patterns Commemorating the Centenary of Engel's Law." Econometrica 25 (1957), pp. 532-551.

Howe, H. "Cross-Sectional Application of Linear Expenditure Systems: Response to Sociodemographic Effects." American Journal of Agricultural Economics 59 (1977), pp. 141-148.

Jackson, L. "Hierarchic Demand and the Engel Curve for Variety." Review of Economics and Statistics LXVI (1984), pp. 8-15.

Jones, C. "The Mobilization of Women's Labor for Cash Crop Production: A Game Theoretic Approach." American Journal of Agricultural Economics 65 (1983), pp. 1049-59.

Judge, G., Griffith, W., Hill, R., and Lee, T. The Theory and Practice of Econometrics. (New York: John Wiley and Sons, 1980)

Kakwani, N. "On the Estimation of Engel Elasticities from Grouped Observations with Application to Indonesian Data." Journal of Econometrics 6 (1977), pp. 1-19.

King, R. "Choices and Consequences." American Journal of Agricultural Economics 61 (1979), pp. 839-48.

King, R. and Byerlee, D. "Factor Intensity and Locational Linkages of Rural Consumption Patterns in Sierra Leone." American Journal of Agricultural Economics 60 (1978), pp. 197-208.

Kinsey, J. "Working Wives and the Marginal Propensity to Consume Food Away from Home." American Journal of Agricultural Economics 65 (1983), pp. 10-19.

Knudsen, O. and Scandizzo, P. Nutrition and Food Needs in Developing Countries. Staff Working Paper 328 (Washington, D.C.: World Bank, 1979).

_____. "The Demand for Calories in Developing Countries." American Journal of Agricultural Economics 64 (1982), pp. 80-86.

Kuh, E. "The Validity of Cross-Sectionally Estimated Behavior Equations in Time Series Applications." Econometrica 27 (1959), pp. 197-214.

Kumar, S. Impact of Subsidized Price on Food Consumption and Nutrition in Kerala. Research Report # 5 (Washington, D.C.: IFPRI, 1979).

Lau, L. "Complete Systems of Consumer Demand Functions through Duality." in M. D. Intriligator, Frontiers of Quantitative Economics. Vol. III (New York: North-Holland, 1977), pp. 59-85.

Lau, L., Lin, W. and Yotopolous, P. "The Linear Logarithmic Expenditure System: An Application to Consumption-Leisure Choice." Econometrica 46 (1978), pp. 843-868.

Lipton, M. Poverty, Undernutrition and Hunger. Staff Working Paper 597 (Washington, D.C.: World Bank, 1983).

Lluch, C. "Consumer Demand Functions for Spain, 1958-64," European Economic Research 2 (1971).

Maddala, G. Econometrics. (New York: McGraw-Hill, 1977).

Massell, B. "Consistent Estimation of Expenditure Elasticities from Cross-Sectional Data on Households Producing Partly for Subsistence." Review of Economics and Statistics 51 (1969), pp. 136-42.

McCarthy, F. D. "Food Consumption, Income Distribution, and Quality Effect." Food Policy 2 (1977), pp. 79-82.

McCarthy, F. D., and Taylor, L. "Macro Food Policy Planning: A General Equilibrium Model for Pakistan." Review of Economics and Statistics (1979), pp. 107-121.

McDonald, S. and Moffit, R. "The Uses of Tobit Analysis." Review of Economics and Statistics 62 (1980), pp. 318-21.

Mellor, J. "Food Price Policy and Income Distribution in Low-Income Countries." Economic Development and Cultural Change 27 (1978).

_____. The New Economics of Growth. (Ithaca, N.Y.: Cornell University Press, 1976).

Muellbauer, J. "Aggregation, Income Distribution and Consumer Demand." Review of Economic Studies 62 (1975), pp. 525-43.

Mundlak, Y. "On the Concept of Non-Significant Functions and Its Implication for Regression Analysis." Journal of Econometrics 16 (1981), pp. 139-49.

_____. "On the Pooling of Time Series and Cross-Section Data." Econometrica 46 (1978), pp. 69-85.

Murty, K. Consumption and Nutrition Patterns of ICRISAT Mandate Crops in India. Economic Progress Report 53 (India: International Crops Research Institute for the Semi-Arid Tropics, 1983).

Murty, K. and Radhakrishna, R. "Agricultural Prices, Income Distribution and Demand Patterns in a Low-Income Country." Paper presented at the First International Working Conference on Computer Applications in Food Production and Agricultural Engineering, Havana, Cuba, October 1981.

Nasse, P. "Analyse des Effects de Substitutions dans un System Complet de Functions de Demand." Annals de L'INSEE 5 (1970), pp. 88-110.

Peterson, W. "International Farm Prices and the Social Cost of Cheap Food Policies." American Journal of Agricultural Economics (1979), pp. 12-21.

Phlips, L. Applied Consumption Analysis. (New York: North Holland, 1974).

Pinstrup-Andersen, P. "Decision Making in Food and Agricultural Research Policy: The Distribution of Benefits of New Agricultural Technology Among Consumer Income Strata." Agricultural Administration 14 (1977), pp. 13-25.

Pinstrup-Andersen, P. and Caicedo, E. "The Potential Impact of Change in Income Development in Food Demand and Human Nutrition." American Journal of Agricultural Economics 60 (1978), pp. 402-412.

Pinstrup-Andersen, P., de Londoño, N., and Hoover, E. "The Impact of Increasing Food Supply on Human Nutrition: Implications for Commodity Priorities in Agricultural Research and Policy." American Journal of Agricultural Economics 58 (1976), pp. 131-42.

Pinstrup-Andersen, P., von Braun, J., Uy, T., and Floro, W. "Impact of Changes in Income and Food Prices on Food Consumption by Low-Income Household in Urban Khartoum with Emphasis on the Effect of Changes in Wheat Bread Prices." (Washington, D.C.: IFPRI, 1983), Mimeograph.

Pitt, M. "Food Preferences and Nutrition in Rural Bangladesh." Review of Economics and Statistics LXV (1983), pp. 105-114.

Pollak, R. "Conditional Demand Functions and Consumption Theory." Quarterly Journal of Economics 83 (1969), pp. 76-78.

_____. "Conditional Demand Functions and the Implications of Separable Utility." Southern Economic Journal 37 (1971), pp. 423-33.

_____. "Habit Formation and Dynamic Demand Functions." Journal of Political Economy (1970), pp. 60-78.

_____. "Habit Formation and Long-Run Utility Functions." Journal of Economic Theory 13 (1976), pp. 272-97.

Pollak, R., and Wachter, M. "The Relevance of Household Production and Its Implications for the Allocation of Time." Journal of Political Economy 83 (1975), pp. 255-77.

Pollak, R. and Wales, T. "Demographic Variables in Demand Analysis." Econometrica 49 (1981), pp. 1533-1551.

_____. "Estimation of Complete Demand Systems from Household Budget Data: The Linear and Quadratic Expenditure Systems." American Economic Review 68 (1978), pp. 348-59.

Powell, A. Empirical Analytics of Demand Systems. (Lexington, Mass: D.C. Heath and Co., 1974).

Prais, S., and Houthakker, H. The Analysis of Family Budgets (New York: Cambridge University Press, 1971).

Quandt, R. "A New Approach to Estimating Switching Regressions." Journal of American Statistical Association 67 (1972), pp. 306-316

Rao, V. "Measurement of Deprivation and Poverty Based on the Proportion Spent on Food: An Exploratory Exercise." World Development 9 (1981), pp. 337-354.

Ray, R. "The Testing and Estimation of Complete Demand Systems on Household Budget Surveys." European Economic Review 17 (1982), pp. 349-69.

Reutlinger, S. and Selowsky, M. Malnutrition and Poverty. (Baltimore, Md.: Johns Hopkins University Press, 1976).

Rogers, B. "The Internal Dynamics of Households: A Critical Factor in Development Policy." Paper submitted to USAID, October 1983.

Rosenberger, A. "The Impact of the Federal Milk Marketing Order System on Dairy Product Consumption." M. A. Thesis (Cornell University, 1978).

Sahn, D. "An Analysis of Food Consumption Patterns and Parameters in Sri Lanka." Forthcoming Research Report (Washington, D.C.: IFPRI).

Samuelson, P. "Social Indifference Curves." Quarterly Journal of Economics 60 (1956), pp. 1-22.

Selowsky, M. "Target Group-Oriented Food Programs: Cost Effectiveness Comparisons." American Journal of Agricultural Economics 61 (1979), pp. 998-99.

Sen, A. K. Levels of Poverty: Policy and Change. Staff Working Paper 401 (Washington, D.C.: World Bank, 1980).

Strauss, J. "Determinants of Food Consumption in Rural Sierra Leone." Journal of Development Economics 11 (1983), pp. 327-54.

Swamy, G. and Binswanger, H. "Flexible Consumer Demand Systems and Linear Estimation: Food in India." American Journal of Agricultural Economics 65 (1983), pp. 675-84.

Taylor, L. Macro Models for Developing Countries. (New York: McGraw-Hill, 1979).

Timmer, C. P. "Food Prices and Food Policy Analysis in LDCs." Food Policy 5 (1980), pp. 188-99.

Timmer, C. P. "Is there 'Curvature' in the Slutsky Matrix?" Review of Economics and Statistics LXIII (1981), pp. 395-402.

Timmer, C. P., and Alderman, H. "Estimating Consumption Parameters for Food Policy Analysis." American Journal of Agricultural Economics 61 (1979), pp. 982-87.

Timmer, C. P., Falcon, W., and Pearson, S. Food Policy Analysis. (Baltimore, Md.: Johns Hopkins University Press, 1983.

Thomson, A. "Nutrition, Food Demand and Policy." Ph.D. Thesis (Stanford University, 1982).

Thurow, L. "Cash Versus In-Kind Transfers." American Economic Review 64 (1974), pp. 190-95.

Tobin, J. "Estimation of Relationships for Limited Dependent Variables." Econometrica 26 (1958), pp. 24-36.

Trairatvorakul, P. "Food Demand and the Structure of the Thai Food System." Ph.D. Thesis (Harvard University, 1981).

Weisskopf, R. "Demand Elasticities for a Developing Economy." in Studies in Development Planning, H. Chenery (ed.), (Cambridge, Mass: Harvard University Press, 1971).

West, D. and Price, D. "The Effects of Income, Assets, Food Programs and Household Size on Food Consumption." American Journal of Agricultural Economics 58 (1976), pp. 725-30.

White, H. "A Heteroskodasticity-Consistent Covariance Matrix Estimation and a Direct Test for Heteroskodasticity." Econometrica 48 (1980), pp. 819-38.

Wohlgennut, M. "Conceptual and Functional Form Issues in Estimating Demand Elasticities for Food." American Journal of Agricultural Economics 66 (1984), pp. 211-215.

Williamson-Gray, C. Food Consumption Parameters for Brazil and Their Application to Food Policy. Research Report 32, (Washington, D.C.: IFPRI, 1982).

Yetley, M., and Tun, S. "Household Demand Analysis for Assessing Nutritional Impact of Development Programs." Staff Report AGES810806, Agricultural Development Branch, Economic Research Service, U.S. Department of Agriculture (Washington, D.C.: USDA, 1981).

Young, T. "Single Equation Demand Estimation." European Economic Review 9 (1982), pp. 103-105.

Zellner, A. "An Efficient Method of Estimating Seemingly Unrelated Regressions and Tests for Aggregation Bias." Journal of the American Statistical Association 57 (1962), pp. 348-68.