INTRAHOUSEHOLD RESOURCE ALLOCATION IN RURAL PAKISTAN: A SEMIPARAMETRIC ANALYSIS

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SUMMARY

We estimate semiparametric Engel curves for rural Pakistan using a large household survey. This allows us to obtain consistent estimates of the effects of household size and composition on consumption patterns even when these demographic variables are correlated with an unknown function of income. The coefficients on the household composition variables are used to infer patterns of intrahousehold allocation. While there is little evidence of gender differences among children, adult males appear to consume more than adult females. Amongst males, workers consume more than dependents. There is no evidence of differential treatment of the elderly or of higher birth-order children. We identify substantial economies of size in food consumption. We also find that Engel curves for food, adult goods and child goods are quadratic logarithmic, which implies that some commonly used demand models are inappropriate. © 1998 John Wiley & Sons, Ltd.

1. INTRODUCTION

We investigate household expenditures in rural Pakistan using semiparametric methods. The work is motivated by an interest both in establishing the shape of the Engel curve and in identifying age–gender patterns in consumption. The first has implications for the analysis of demand and taxation. The second informs the literature concerned with gender bias in South Asia and is of potential use in making welfare comparisons across heterogeneous households.

An important question that arises in devising development strategies in low-income countries pertains to the impact of income growth on the welfare of the poor. If foodshare is used as an (inverse) indicator of welfare, then we are interested in the rate at which a rise in income generates a decline in foodshare. This depends upon the shape of the food Engel curve. Existing studies have tended to assume that foodshare is linear in the logarithm of per capita expenditure.[1] This is questionable in the context of a sample that includes very poor households since it is plausible that, at the lower end of the income distribution, foodshare either does not decline with income or it declines more slowly than at higher incomes. Identification of the curvature of the Engel curve has the further advantage that it permits us to infer the rank of the underlying demand system (Gorman, 1981). Sound policy analyses can depend upon finding the correct specification.[2] This paper therefore determines the relation of income and foodshare non-parametrically for a given parametric specification of the effects of household size and structure on consumption. Controls

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[1] Income is particularly volatile in agrarian economies. We shall follow the convention of using expenditure instead of income since the former is the smoother of the two variables and more likely to be log normal.
[2] See, for example, Banks, Blundell and Lewbel (1997) in the context of estimation of the welfare costs of indirect taxation in the UK.

for compositional heterogeneity across households, as we shall see, are of independent interest. An advantage of semiparametric estimation is that it offers consistent estimates of the coefficients on these compositional variables when the functional dependence of foodshare on income is of unknown form.

There is growing interest in intrahousehold allocation since the evidence has been accumulating against the unitary model (e.g. Browning *et al.*, 1994). Any implications for intrahousehold inequality are of particular interest in poor societies, where relative neglect can have harsh consequences. Data from several low income countries reveal that morbidity and mortality rates are higher for females than for males (see Section 2). We investigate whether there are gender differences in consumption that underlie this phenomenon. Since differences in consumption levels may reflect differences in needs rather than any biases in intrahousehold food allocation, we control for the work status of individuals. Since gender differences in consumption may appear at some ages and not at others, we use a finer specification of age effects than the literature on gender bias has so far permitted. Age effects on consumption are also interesting with regard to the status of the elderly in the large integrated families that distinguish developing countries. With the progress of the demographic transition and the associated increase in the proportion of the elderly in society, it is unclear how viable continued family support for the elderly will be. Age may have further effects via birth order. There is considerable evidence from poor societies of first-born children being treated differently than their siblings. We therefore investigate relative neglect of the elderly and of higher birth-order children. We do not attempt to determine the motivations underlying intrahousehold allocation but only to detect the consumption patterns that they give rise to.

2. THE LITERATURE

Fine accounts of recent developments in non-parametric techniques are provided by Silverman (1986) and Härdle (1990), for example. There is a recent but growing literature that applies these techniques to the estimation of Engel curves. For instance, Banks *et al.* (1997), Lewbel (1991) and Delgado and Miles (1996) study expenditure patterns in the UK, the USA and Spain respectively. Strauss and Thomas (1990) and Subramanian and Deaton (1996) plot non-parametric calorie–expenditure curves for Brazil and India respectively. Estimation of semiparametric Engel curves which incorporate demographic variables is relatively uncommon. However, it is very useful if one is interested in the effects on budget shares of household size and structure and if these (parametrically specified) variables are correlated with the unspecified function of total expenditure per capita.[3]

There is considerable evidence of excess female mortality and morbidity in South Asia, especially amongst children (e.g. D'Souza and Chen, 1980; Das Gupta, 1987). This has been attributed to discrimination against females in the intrahousehold allocation of food and health care (e.g. Basu, 1989; Behrman, 1988; Chen, *et al.*, 1981; Harriss, 1990). The fact that girls also tend to have lower completed schooling than boys (e.g. Strauss and Beegle, 1995) suggests that the intrahousehold allocation of resources among children is not guided by inequality aversion or by needs alone. The data appear consistent with parental preference of boys, as also with the maximization of returns to investment in regions where investments in boys produce higher

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[3] There is, for instance, widespread evidence of a negative relation of per capita expenditure and household size (e.g. Lipton and Ravallion, 1994, section 4·2).

returns. Son preference has been documented in infertility studies where women report a desire for sons over daughters (e.g. Rukannudin, 1982, for rural Pakistan, Das Gupta, 1987, for India). Rosenzweig and Schultz (1982) argue that gender differences in child survival in India can be explained in terms of gender differences in expected earnings.

While the evidence on *outcomes* like mortality and literacy is contained in census data, the evidence on *inputs* like nutrients and hospital admissions, which policy can attempt to target, has largely relied upon scattered small-scale investigations. Large-sample studies have tended to take the indirect approach of inferring intrahousehold distribution from data on household expenditure and household composition since these are often available in national surveys.[4] These studies estimate Working–Leser Engel curves (Working, 1943; Leser, 1963) for food or adult goods. Interestingly, more often than not, they find no 'gender bias'. Where gender differences in consumption patterns do emerge, the existing literature does not attempt to distinguish differential needs for allocational bias.

While there are indications from small-sample studies that girls of higher birth order receive less care (e.g. Das Gupta, 1987), large-sample investigations of birth-order effects are scarce. With the exception of Kochar (1996), who studies medical expenditures on the elderly in Pakistan, there is similarly little work concerned with how the elderly fare in the intrahousehold distribution of resources.

We have generalized the investigation of intrahousehold inequalities in consumption beyond gender differences. This paper also differs from earlier studies in employing a richer, less restrictive specification of age–gender effects, in distinguishing dependents from workers, and in allowing the expenditure function to be determined non-parametrically.

3. SPECIFICATION

3.1 The Semiparametric Model

For the reasons put forward in Section 1, we permit the functional dependence, F, of budget share (ω) on log per capita expenditure (y) to be determined by the data. This gives the semiparametric model

$$\omega_i = \beta^T x_i + F(y_i) + v_i \tag{1}$$

where subscript i denotes household, $F(\ )$ is unknown, and x_i is a vector of J variables representing household size and composition and other relevant covariates. Estimates are obtained following the semiparametric procedure outlined in Robinson (1988). This involves non-parametric regressions for which we have used the Fast Fourier Transform algorithm of Härdle (1987).

3.2 The Empirical Model

The data are from the Household Income and Expenditure Survey of Pakistan conducted by the Federal Bureau of Statistics in 1987–8 on a stratified random sample of about 18,000 households. The investigation is confined to rural households with no more than 20 members, of which there

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[4] See Browning and Subramaniam (1995), Subramaniam (1996) and Subramanian and Deaton (1991) on India, Ahmad and Morduch (1993) on Bangladesh, Burgess and Zhuang (1997) on China, and Deaton (1989) on the Côte d'Ivoire and Thailand.

are 9740 in the sample. These households exhibit considerable substantial heterogeneity. The estimated model is

$$\omega_i = F(y_i) + \alpha \ln N_i + \Sigma_k \gamma_k (N_{ki}/N_i) + \varphi^T z_i + v_i \qquad (2)$$

ω is the share of the household budget that is allocated to food. Food expenditure includes an imputed value for home-grown produce. y is the logarithm of total expenditure per capita. Leaving $F(\ )$ unrestricted permits a commodity that is initially a luxury to become a necessity at higher income levels. N is household size and $\alpha < 0$ implies size economies. These may arise as a result of indivisibilities in perishable foods, which result in less waste in larger families. Alternatively, larger families may be better placed to exploit discounts on bulk purchases. (N_{ki}/N_i) are the proportions of household members in K age–gender–work status groups,[5] z denotes birth-order, region and season effects, and v reflects stochastic variations in tastes and other unobservables. Though fertility may be endogenous to the standard of living in the household, we treat household size and structure as exogenous. One may therefore think of the equation as reflecting a long-run reduced form.

Existing studies that use expenditure data to investigate intrahousehold inequality in consumption (see footnote 4) have specified *household composition* in terms of a small number of age groups such as 0–4, 5–14, 15–59 and 60-plus. This paper includes a variable for every age in the range 0 to 14 (the data are in integers).[6] This can be important since there may be gender differentials in consumption at certain ages and not at others. Also, this information might enable the researcher to determine the seriousness of any neglect. For instance, Das Gupta (1987) finds that girls in India are most vulnerable to shocks to their family incomes in the age range 0–2. Adults are grouped into the categories 15–24 (young), 25–59 (prime-age) and 60-plus (elderly). So as to control for differences in food needs arising from different levels of activity, allocations to individuals are permitted to depend upon their work status. It is not unusual for children 8 years and older to work.[7] Unpaid family workers are regarded as working if they contribute at least 15 hours a week to household production. Since the market is often underdeveloped in rural economies, it makes sense to count non-market work.

Birth-order dummies for male and female children allow for the possibility that the first child has a different impact on budget shares as compared to later children of the same sex. For instance, certain fixed costs may be incurred for the first-born alone; parents may favour the first-born son since it is customary that they will spend their old age with him; or second and further daughters may be particularly undesirable when higher-order births reflect an effort to achieve the desired number of sons. We include dummies to control for which of Pakistan's four *provinces* a household is located in. These allow regional price variation but households within a region are assumed to face the same prices. We also include dummies for the quarter in which a household was interviewed, which we expect will proxy for *season*. Agrarian economies experience large seasonal fluctuations in activity and income. Food requirements can therefore vary dramatically

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[5] If the coefficient on the kth variable is γ_k, then the change in foodshare upon replacing a person in the suppressed group (suppressed on account of collinearity) with a person from the kth group is γ_k/N.
[6] Age 0 refers to children less than a year old. We adopt the convention of defining people 15 years of age or older as adults.
[7] Using the Pakistan Integrated household Survey to study child labour in rural Pakistan, Bhalotra (1998) estimates that about 7% of children aged 10–14 work outside the household, another 20% are employed on the household farm or enterprise, and an even larger fraction are engaged in domestic work.

over the agricultural cycle for agricultural workers (e.g. Payne, 1985), and there are indications that females bear the brunt (e.g. Dercon and Krishnan, 1997).

4. RESULTS: FOOD ENGEL CURVES

4.1 The Shape of the Engel Curve

Figure 1(a) displays the semiparametric estimate of the *food* Engel curve. The vertical axis refers to the budget share adjusted for the effects of household size and composition. Consistent with Engel's law, the curve slopes downwards. Food is a broad aggregate of 82 commodities which, on average, consumes 52% of the household budget in rural Pakistan. For illustrative purposes, we estimate an Engel curve for the more homogeneous category of *milk and milk products* (Figure 1(b)). These are relatively expensive items, which constitute 25% of the average food budget. The inverted U-shape of this Engel curve implies that milk is a luxury for many households in the sample. Quadratic curves fitted by OLS on the same data appear to provide good approximations to the actual functional forms. Figure A1 in the Appendix displays 95% confidence intervals around the semiparametric curves, which clearly encompass the quadratic fits.[8] Table I presents parametric (OLS) estimates of the Engel curves which show that the quadratic in log per capita expenditure is highly significant. It also reports a test (see Härdle and Mammen, 1993) of the semiparametric against the parametric curves which shows that we cannot reject the parametric quadratic model against the semiparametric model.

A quadratic food Engel curve is an interesting result. Food Engel curves are linear in the richer populations of Spain (Delgado and Miles, 1996), the UK (Banks *et al.*, 1997) and the USA (Lewbel, 1991), suggesting that the quadratic is a feature of the income levels in our sample and the stage of economic development.[9] In our data, as expenditure rises, foodshare appears to fall less rapidly for poorer households than for richer ones. This is consistent with the food basket of the poor being deficient in either quantity or quality. A further possible explanation of the curvature is that indivisibilities in alterative (non-food) commodities are binding at low income levels.[10] Another is that the distribution of household expenditure is conditioned by the array of goods and services that the household may choose from, and poorer households tend to live in remote villages where choices are more limited.

Quadratic Engel curves imply a demand system of rank three (Gorman, 1981; Lewbel, 1991). This means that rank two Piglog demand systems such as the Almost Ideal Demand System of Deaton and Muellbauer (1980), the Log-Translog model, and the Linear Expenditure System of Stone (1954) are inappropriate for analyses of demand in rural Pakistan.

Might the non-linearity of the Engel curve reflect specification error? While we have permitted the demographic and labour market characteristics of individuals to shift the Engel curve, we have restricted the slope coefficients to be the same across fairly heterogeneous groups. This pooling of data across heterogeneous groups may give rise to a spurious curvature.[11] Consider households with different *numbers of earners*. Suppose that the underlying Engel relation is linear

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[8] The confidence intervals were generated using 500 replications of the wild bootstrap (see Härdle and Mammen, 1993).
[9] Working (1943) found non-linearities at low levels of total expenditure in his sample of US households in the mid-1930s, and remarked that 'the laws of expenditure applicable in one culture and in one epoch may not be applicable in another culture or in another epoch'.
[10] In other words, soon after food needs are met, the household may be able to divert expenditures towards items like clothing and glass bangles. However, it may not be able to buy items like a bicycle or a stove. At higher levels of income, these items become affordable and foodshare falls more rapidly than before.
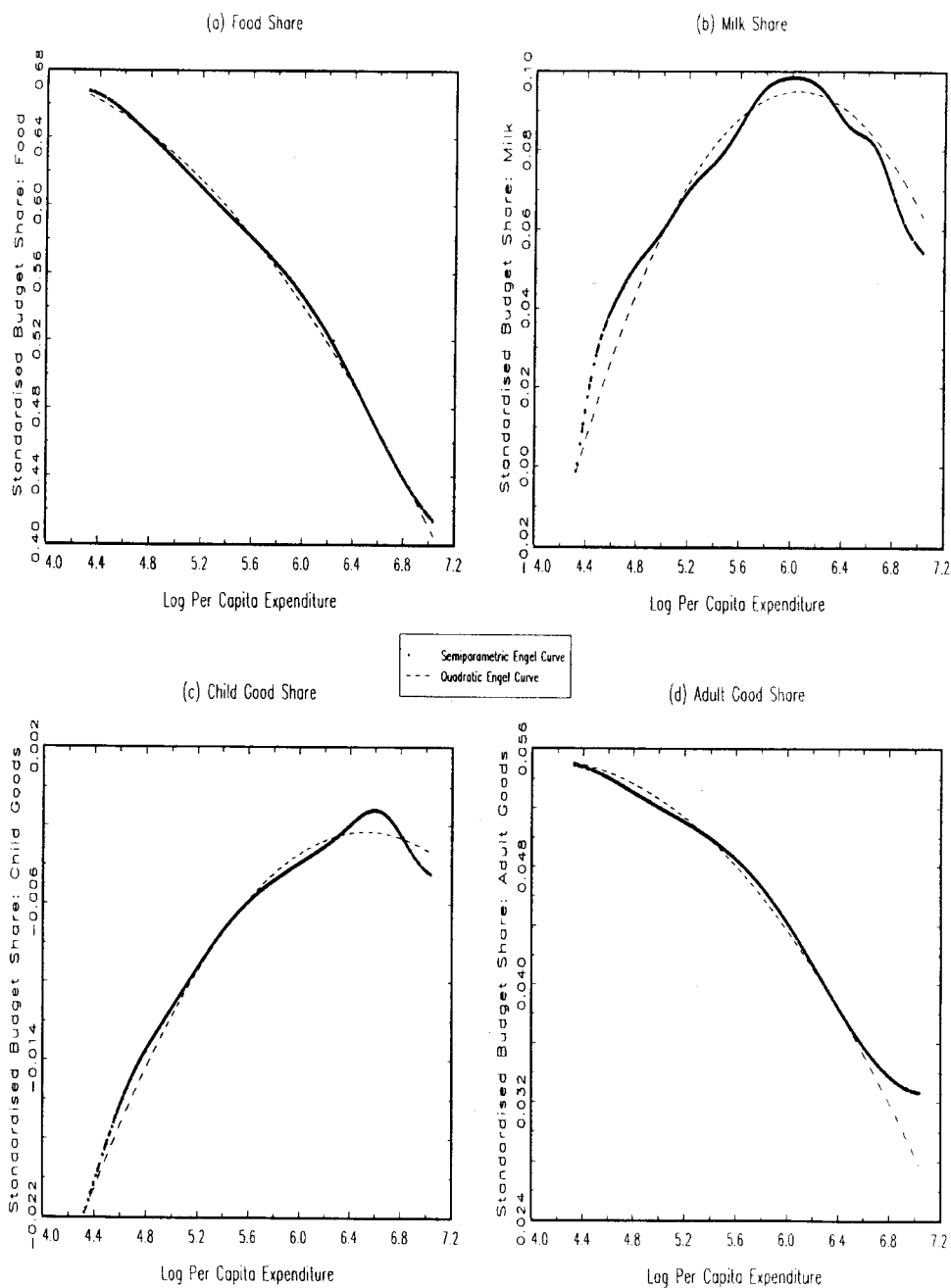[11] We are grateful to an anonymous referee for pointing this out.

Figure 1. Semiparametric and quadratic Engel curves

Table I. Semiparametric and parametric Engel curves

Dependent variable	Foodshare		Milkshare		Adult goods share		Child goods share	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Variable	Semiparametric	Parametric	Semiparametric	Parametric	Semiparametric	Parametric	Semiparametric	Parametric
ln Y		0·16 (4·5)		0·36 (15·6)		0·015 (1·4)		0·044 (7·9)
(ln Y)^2		−0·023 (7·1)		−0·030 (15·0)		−0·002 (2·4)		−0·003 (6·9)
ln size	−0·014 (4·9)	−0·016 (5·8)	0·014 (6·1)	0·012 (5·5)	−0·002 (1·8)	−0·002 (2·5)	0·006 (11·0)	0·006 (10·6)
Female birth order	−0·001 (0·2)	−0·001 (0·1)	0·003 (1·0)	0·003 (0·9)	−0·000 (0·3)	−0·000 (0·2)	−0·000 (0·5)	0·000 (0·4)
Male birth order	−0·005 (1·4)	−0·004 (1·0)	0·003 (0·9)	0·003 (1·1)	−0·001 (0·7)	−0·001 (0·4)	0·001 (1·4)	0·001 (1·3)
χ^2_5 demographics	309·7*	339·2*	7702*	5486*	3908*	3153*	2215*	2524*
χ^2_3 regions	328·8*	338·6*	432·6*	435·5*	433·9*	443·4*	265·7*	269·3*
χ^2_3 seasons	10·7*	9·2	0·75	0·93	4·1	5·4	1·9	1·6
Adjusted R^2		0·22		0·071		0·117		0·194
Härdle–Mammen	0·12 (0·22)		0·12 (0·23)		0·003 (0·005)		0·001 (0·01)	
Mean of dep. var.	0·52	0·52	0·13	0·13	0·053	0·053	0·017	0·017
Y elasticity	Figure 3(a)	−0·19	Figure 3(b)	0·19	Figure 3(d)	−0·14	Figure 3(c)	1·6
Size elasticity	−0·027	−0·031	0·11	0·092	−0·038	−0·038	0·35	0·35

Notes:

Number of observations = 9643 (after 1% trimmed at low densities). Standard errors are Newey–West heteroscedasticity-consistent. Absolute t-ratios are in parentheses. χ^2_k are Wald tests of joint significance, k = degrees of freedom, a star indicates significance at 1%. Y is per capita expenditure, dep. var. is dependent variable. Demographics are size, birth-order and composition (age–gender–work status) variables. Härdle–Mammen is a test of the semiparametric against the parametric model; the figure in parentheses is the 5% critical value, generated by 500 replications of the wild bootstrap. The elasticities are for commodity *shares*. Figure 3 presents elasticities for commodity *expenditures*, which are equivalent to the share elasticities plus unity. The Y-elasticities are evaluated at the mean level of per capita expenditure (=5·6 in logs).

for each group of households, but steeper for multiple-earner households. If, in addition, the per capita expenditure of multiple-earner households is greater than that of one-earner households, then the Engel curve for multiple-earner households will lie mostly to the right of the Engel curve for one-earner households. Pooling across one and multiple earner households will then tend to produce a non-linear Engel curve. Similarly, pooling data across *regions* may generate spurious non-linearity if the Engel curves for the different regions have different slopes, given that the regions have different levels of expenditure per capita. Figure 2(a) plots semiparametric Engel curves for three subsamples of the data, consisting of households with one, two, and three or more earners.[12] Their parametric counterparts confirm that, for every subsample, the relation of foodshare and log per capita expenditure is a quadratic.[13] Semiparametric curves for each of the four provinces of Pakistan are displayed in Figure 2(b). Parametric regressions show that the Engel curves for Sind and Baluchistan are linear. However, 75% of the sample households live in Punjab and NWFP which support quadratic logarithmic curves.

A further specification issue pertains to the potential endogeneity of per capita expenditure (y). To allow for this, we followed the procedure in Blundell and Duncan (1997), whereby expenditure is regressed (parametrically) on income (the instrument) and the residual from this regression is included as an additional variable in the semiparametric model. The quadratic shape of the Engel curve persists although it is attenuated. The attenuation leaves unchanged the estimates of intrahousehold consumption patterns. (Results are not reported here but are available upon request.)

4.2 Non-parametric Elasticities

For reasons indicated in Section 1, we are interested in finding expenditure elasticities for the full range of incomes in the data. Moreover, if there is a common element of measurement error in food expenditure and total expenditure (such as may arise from imputing value to home-grown produce) and if this can be localized to a part of the expenditure distribution (e.g. if only rich landowning households have home-grown produce), then it is useful to have non-parametric estimates of elasticities at every point of the Engel curve. This is because the elasticity at any point depends only upon the data in the bandwidth around it. In contrast, the parametric estimate of the elasticity at some middle point in the distribution will tend to be biased because the erroneous data bias the slope of the entire Engel curve. The computation of nonparametric elasticities is straightforward once an estimate of the smooth, differentiable Engel function has been obtained (e.g. Deaton, 1997, pp. 195–6).[14] Figure 3(a) presents nonparametric estimates of elasticities with respect to per capita expenditure (Y) of *food expenditure* rather than foodshare. These are given by $[1 + m'_\omega(y)/m_\omega(y)]$, where $m_\omega(y)$ is the non-parametric regression estimate of the expectation of foodshare (ω) conditional on $y(=\ln Y)$.

Since the food Engel curve slopes downwards, the expenditure elasticities are less than unity. For a range of low incomes (Rs 100–260 per capita per month), the elasticity remains fairly

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[12] In rural Pakistan, 54% of households have one earner, 23% have two earners and 16% have three or more earners. The remaining 6% have no earners. Average household size increases with the number of earners but the proportion of earners is nevertheless increasing in the number of earners. *Yet, contrary to expectation, average per capita expenditure is (slightly) decreasing in the number of earners.* This is consistent with poverty compelling women and children to work.
[13] The OLS expenditure coefficients are as follows, *t*-statistics in parentheses: One-earner: $0.25(6.3)y - 0.03(8.9)y^2$; Two-earner: $0.24(3.8)y - 0.03(5.4)y^2$; Three or more earners: $0.30(3.3)y - 0.04(4.6)y^2$.
[14] Ideally, we want the elasticities from the semiparametric model but plots (available on request) of the non-parametric and semiparametric Engel curves are very similar in scale and shape.

(a) Labour Market Status of Household
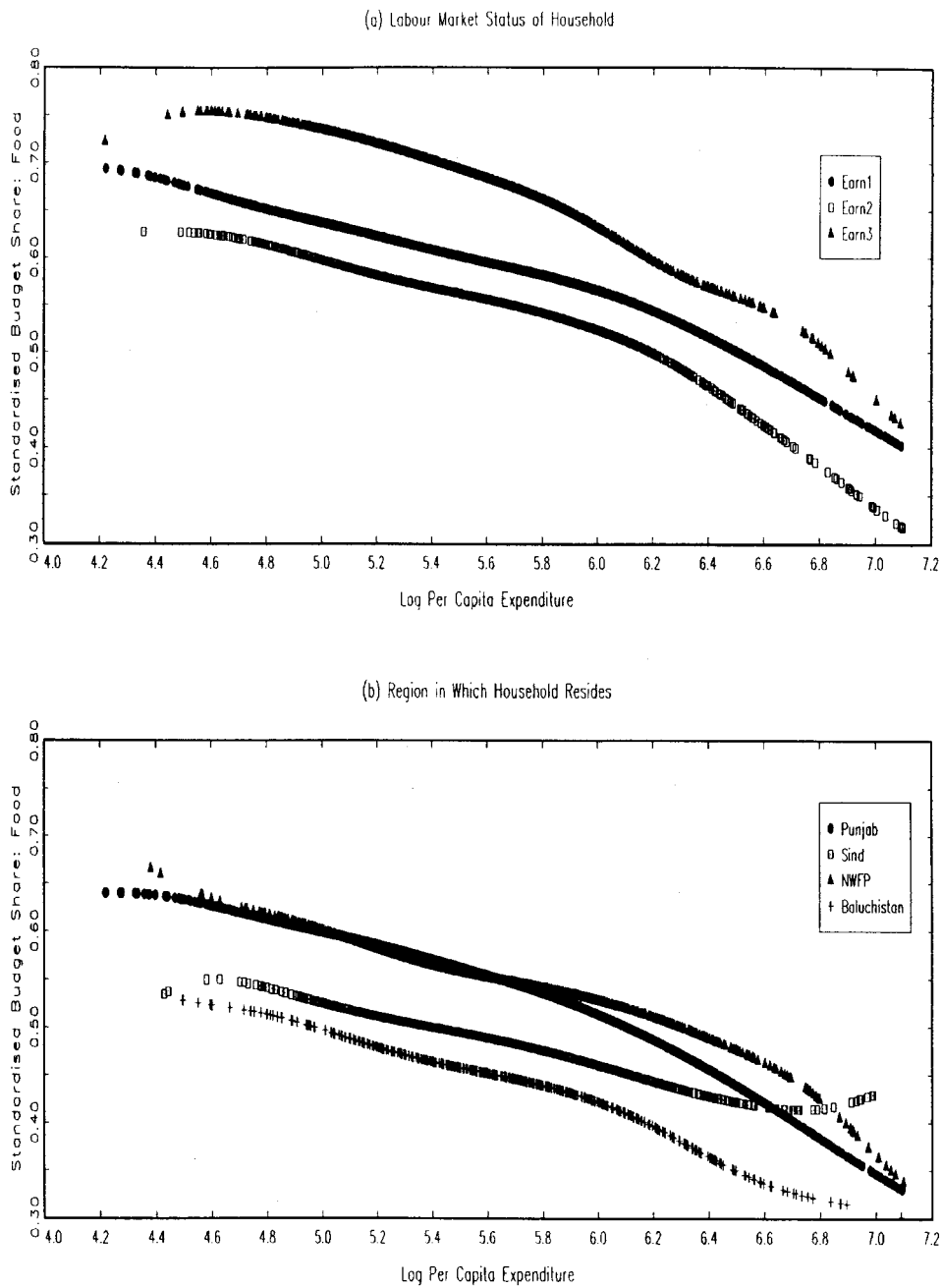
(b) Region in Which Household Resides

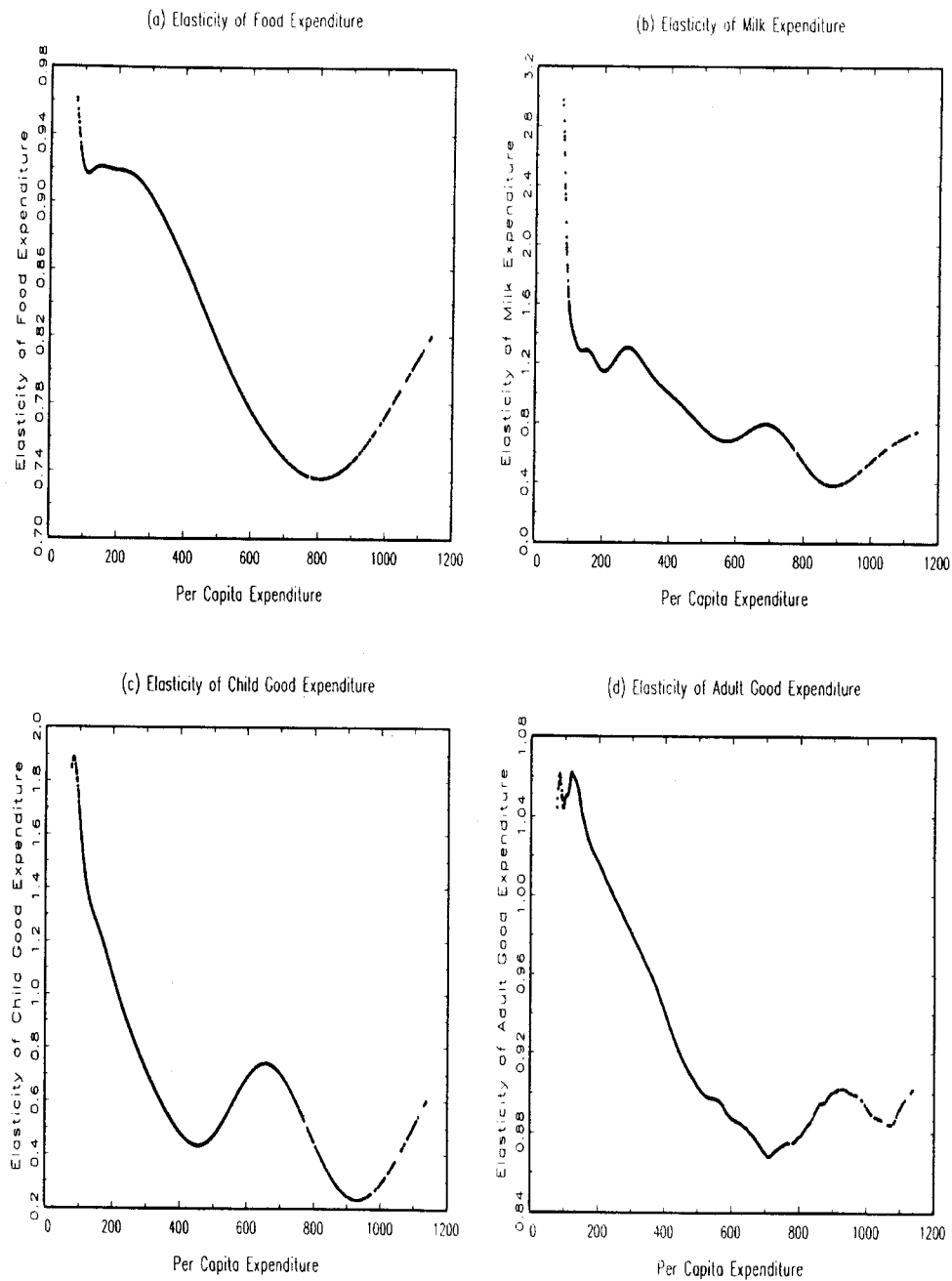Figure 2. Semiparametric food Engel curves: investigating heterogeneity

Figure 3. Expenditure elasticities from non-parametric Engel curves

constant at about 0·92. At levels of Y above the sample mean (Rs 260), it begins to fall and for relatively well-off households (Y of Rs 800), it is about 0·73. These large food expenditure elasticities are consistent with most households in the sample being poor. In comparison, the elasticity at the sample mean is 0·75 in rural Maharashtra in India (Subramanian and Deaton, 1991) and about 0·3 in the UK (based on estimates in Banks *et al.*, 1997). There is a general tendency for the elasticity of milk expenditure to decline with rising incomes, though less rapidly than in the case of food expenditure (Figure 3(b)). Milk and milk products are luxuries (elasticity > 1) for households with Y less than about Rs 400 and necessities after.

4.3 Scale Effects

Milk consumption reflects size diseconomies. The *food* aggregate, however, displays significant economies of size. Note that this is in spite of controlling for household composition, so larger households are not merely 'saving' on the basis of having a high proportion of children, who consume less than adults. The size coefficient of $-0·016$ implies that a doubling of household size (holding per capita expenditure constant) would decrease foodshare by 1 percentage point. If size were constant, then a household with the mean level of per capita expenditure would require an increase in *per capita* expenditure of 13% in order to achieve this same decrease in foodshare.[15] Our estimates of Engel curves for subsamples of the data defined by number of earners reveal that size economies are about twice as large for one-earner families and they diminish as household size expands. In any case, what we find is an impressive economy on food, which is a rival good. It suggests that, once we have taken account of more likely size economies on shared goods (cooking stove, house, etc.), the economic gains to household expansion or integration may be quite substantial. It also implies that, where foodshare is used as a measure of household welfare, per capita expenditure understates the welfare of relatively large households.

4.4 Household Composition Effects

There are no *birth-order* effects in consumption. There is thus little support here for the hypothesis that parents favour older children.

When the proportion of workers is included in a food Engel curve with age–gender effects alone, it emerges with a positive and significant coefficient of 0·04. Given that total expenditure per capita is being held constant, this supports the interpretation that workers, on average, have greater food *needs*. Existing studies of food and adult food expenditures in the 'gender bias' literature have made no attempt to control for needs when studying intrahousehold allocations.

Composition is denoted by a vector measuring the proportion of household members in different age, gender and work-status groups. For children, we have the proportion of every age (in years) in the range 0–14. Adults are grouped into three age categories, 15–24, 25–59 and 60-plus. Interaction with gender dummies yields 36 age–gender variables. No children under the age of 8 work. For males and females 8 years or older, the age–gender variable is interacted with a

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[15] We have used the semiparametric estimates underlying Figure 1(a) to perform this calculation. The parametric estimates in column 2 of Table I yield a figure of 12%. This is close, as we would expect since the parametric and semiparametric curves coincide in the neighbourhood of mean expenditure (see Figure 1(a)).

work-status dummy. This gives 56 compositional variables. Wald tests of their joint significance are in Table I (χ^2 demographics). The following results are based on Wald tests of the equality of coefficients at individual ages. These are available from the authors upon request.

- *Children:* There are no significant *gender* differences in the consumption patterns of children,[16] though there are *work-status* effects for boys. Working boys consume significantly more than dependent boys (and dependent girls).[17] However, working girls do not systematically get more than dependent girls.

- *Adults:* When composition is restricted to age–gender, we find that the presence of male adults in the household increases the food budget to a greater extent than does the presence of female adults, and this is true in each of the three age bands. Distinguishing individuals by their work-status, we find that a systematic *gender* difference in food is restricted to working individuals.[18] This suggests that the difference might pertain to differences in the sorts of work performed by men and women. There are *work-status* effects among adults similar to those found for children. Working males raise foodshare (and milkshare) significantly more than dependent males do, especially at the ages 15–25. In contrast, the consumption patterns of women are not systematically influenced by their work-status. In Section 1, we argued that there may be *age* 'discrimination' in favour of or against *the elderly*. Tests of coefficient equality between the elderly (over 60), prime-age adults (15–60) and young adults (15–25) do not reveal any consumption differences between the elderly and other adults. Rather, we find that when young adults are not working, they have a significantly smaller effect on foodshare than older adults.

4.5 Region and season Effects

The Engel curves include province and quarter dummies. Wald tests of their joint significance are presented in Table I. The season effects are generally small and only marginally significant. The region fixed effects are highly significant.

5. NON-FOOD EXPENDITURES

5.1 Motivation and Definitions

Since the sum of expenditure per capita is held constant in estimation of the food Engel curve, any drop in foodshare must be compensated by a rise in the share of other goods. Suppose that, when a boy in the household is 'replaced' by a girl, the budget share of food falls. This might suggest that boys are better fed than girls. However, even as girls induce lower spending on food, they may induce higher spending on other goods. It is therefore difficult to use the food Engel curve to draw conclusions about gender differences in overall spending.

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[16] At the 10% level of significance, the age-specific Wald tests are consistent with boys consuming more than girls at the ages of 3 (food) and 7 (food and milk). We might put this down to sampling variation.

[17] We are looking at boys aged 8–14 at every age. Workers consume more food than non-workers at ages 8, 10, 13 and 14; and more milk at ages 10, 12, 13 and 14.

[18] The result that working men consume more than working women does not apply to the sub-category of *milk*. Among dependent adults, females appear to consume more *food* than males in the age range 15–25, whereas in the age band 25–60, this tendency is reversed. In *milk* consumption alone, dependent females aged 15–60 do significantly better than males. This might be related to child-bearing.

An alternative is to study the allocation of the sum of expenditures on children using the adult goods method of Rothbarth (1943), discussed in Deaton *et al.* (1989). Adult goods are those on which the presence of children has only income effects. We define adult goods to include tea and coffee, alcohol and drugs, tobacco and tobacco products, men's footwear and women's footwear. If parents are driven to drink more because their children wear them out or to smoke less to protect their children, then alcohol and tobacco, for example, will not be valid adult goods. If the chosen composite is valid then the coefficients on the proportions of children in the household will be negative. We also investigate consumption patterns in child goods, defined to include children's furniture and nursery items, toys, children's footwear, pocket money and icecream. While this is interesting, it is assigned secondary importance since many child goods are likely to be luxuries in rural Pakistan and our main concern is with basic necessities like food.

5.2 Estimates of Engel Curves for Adult and Child Goods

Semiparametric and parametric estimates of Engel curves for child and adult goods are in Figures 1(c) and 1(d). Adult goods appear to be necessities. Child goods are luxuries until the very edge of the income range in our sample. As before, the dotted lines in the figures are the quadratic Engel curves fitted on the same data. In both cases, they afford a good approximation (Härdle–Mammen statistics in Table I). Non-parametric estimates of expenditure elasticities are plotted in Figures 3(c) and 3(d).[19] The coefficients on the logarithm of household size suggest economies in adult goods and diseconomies in child goods expenditures. The latter may be a reflection of sibling rivalry that bolsters children's demands for toys and the like, though an alternative explanation is that the purchase of a child good is better justified the more children there are in the family to share it.

- *Adult goods:* Most though not all of the compositional variables for children are, as expected, negative. Their relative coefficients reveal no significant *gender* or *work-status* differences in the allocation of non-adult expenditures among *children*. Among *adults*, their allocation favours males aged 25–60, irrespective of work status. Among 15–25-year olds, working males get more than working females and more than dependent males. Amongst the elderly, there is no gender difference but workers get more than dependents.

- *Child goods:* In the child goods equation, most of the child variables have positive coefficients. For under-8s and working *children*, there are no differences. However, among children aged 10–14 years, if neither is working, there appears to be a *gender* preference in favour of boys. There is also a curious *work-status* effect among boys, whereby the share of child goods is larger in households with dependent boys than in households with working boys, *ceteris paribus*. Might this be because working boys have no time for toys? The composite nature of child goods makes it difficult to speculate. The *adult* coefficients make interesting suggestions. The share of the budget devoted to child goods would register a significant increase if a working male were replaced by a working female. The latter is consistent with the views that women gain greater control over the allocation of the budget if they work, and that they have different preferences

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[19] Unlike in the case of food and milk, the non-parametric Engel curves for adult and child goods are not similar to the semiparametric curves: unsurprisingly, controlling for demographics alters the manner in which expenditures on these goods vary with living standards. Therefore Figures 3(c) and (d) which are based upon non-parametric estimates are not directly comparable with the semiparametric curves in Figures 1(c) and (d). For instance, Figure 1(c) implies that child goods are luxuries for most households but Figure 3(c) suggests that they become necessities at about the mean level of per capita expenditure.

from men. Child goods' share would also rise if a working adult were replaced by a dependent adult of the same gender. Since expenditure per capita is held constant, this would appear to reflect a complementarity between child goods and adult leisure.

6. GENDER BIAS

Our investigations do not reveal any systematic gender differential in the intrahousehold allocation of expenditures on *children*. We do, however, find that working boys tend to have a significantly larger effect on foodshare than dependent boys and dependent girls.[20] We also report some evidence to suggest that, on average, workers' have greater food needs than non-workers (Section 4.4). Yet we cannot rule out gender bias because we cannot rule out the possibility that boys are over compensated for working. In any case, our investigations do not provide any overwhelming evidence of bias.

Yet the evidence of excess female mortality and morbidity among children in South Asia (Section 2) may be consistent with our findings for any of the following reasons.

(1) Girls may be neglected in ways other than food denial. For instance, parents may take sons to a medical clinic more readily than they take daughters (e.g. Basu, 1989; Das Gupta, 1987). Alternatively, parental care may be allocated relatively favourably to sons and this can be critical when the child is ill. Rose (1996) finds evidence to suggest that parents in rural India spend more time at home when they have sons than when they have daughters.

(2) If gender bias characterises only a fraction of households in the sample and the selection rule that distinguishes these households from the rest is unknown, then any bias that is present may be obscured. Morduch and Stern (1997) estimate a mixture model on data from Bangladesh to deal with this problem and find that controlling for heterogeneity reveals a pro-male bias in health outcomes.

(3) Girls will fare worse than boys in the population if they belong disproportionately to poorer households. This, in turn, is possible in the following circumstances. If there is son preference in fertility (see Section 2) and parents only control family size after a son is born, then larger families will have relatively more daughters and the proposed correlation arises because larger families tend to have lower per capita expenditure (e.g. Ahmad and Morduch, 1990). Alternatively, if marriage costs arise for daughters but not for sons then households maximizing intertemporal utility will begin a savings programme when a daughter is born (e.g. Browning and Subramaniam, 1995). Less is consumed by all children in such households but this effect is predominant in households with more daughters.[21] While these mechanisms result in girls being worse off than boys, they will not show up in our investigation since we hold per capita expenditure constant.

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[20] Recall that while allocations to working girls are insignificantly different from allocations to working boys, working girls do no better than dependent girls.
[21] Instead of cutting consumption, the household could increase income by increasing labour supply (Rose, 1996). However, while food expenditures may thus be maintained, child health may suffer. If it is the mother who joins the labour market (or increases hours on the market), then the children may suffer from less time and care. If, instead, it is the children who go out to work then, at given food expenditure, they may have worse health.

7. CONCLUSIONS

It is common practice to *assume* a functional form for demand models. Our semiparametric estimates of Engel curves suggest that the popularly used Piglog class of demand models is inappropriate for rural Pakistan. The data favour a quadratic logarithmic specification of the Engel curve which, in the case of food, contrasts with linear Engel curves obtained for the USA, UK and Spain. However, it is consistent with our sample of rural households representing a range of incomes that includes some at which food needs remain unmet. The non-parametric estimate of the elasticity of food expenditure with respect to total expenditure is almost 0·90 at the mean, an indication of how poor these households are.

We have obtained estimates of the effects on budget shares of household size and composition, allowing these variables to be correlated with an unknown function of household expenditure. These estimates are potentially useful not only in the long run prediction of the effects on demand of demographic change but also in the comparison of welfare levels across heterogeneous households.

We find economies of size in the consumption of food and adult goods and diseconomies of size in consumption of milk and child goods. Food is by far the most important commodity, consuming 52% of the budget on average. Size economies in food consumption suggest that a doubling of household size is equivalent to a 13% increase in *per capita* expenditure for a household with the mean level of per capita expenditure, holding constant composition, region and season. Together with the more evident size economies that obtain from durables like housing and cooking stoves, this contributes to an economic rationale for the integrated family that characterises poor societies.

A central objective of this research was to investigate whether the distribution of resources within families tends to favour males over females, older over younger children, or prime-age over elderly adults. Rather more attention was paid to children because they are most vulnerable to neglect, the consequences of which can be lasting. Identification of intrahousehold inequality may, for instance, assist poverty targeting. There is no evidence of birth-order effects. We are unable to identify systematic gender differences in consumption among children. However, among working adults, men have a significantly larger impact on foodshare than do women. We also find, in common across children and adults, that male workers command more of the household food budget than do male dependents. There is no corresponding difference between female workers and female dependents. The elderly do not appear to consume any less than younger adults, irrespective of whether they work.
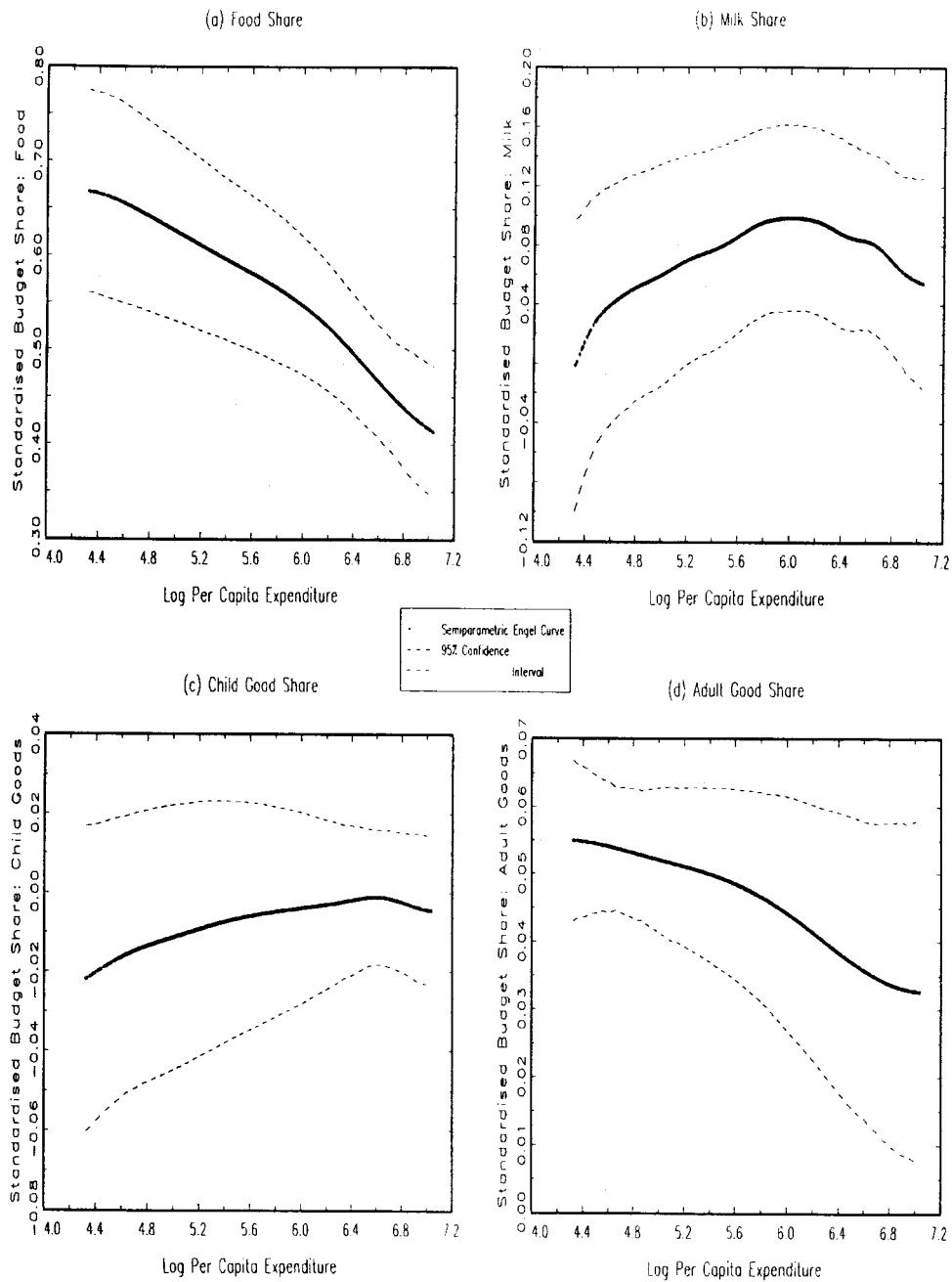
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Figure A1. Semiparametric Engel curves with confidence bands

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