INTRAHOUSEHOLD RESOURCE ALLOCATION IN RURAL PAKISTAN: A Semiparametric Analysis

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Abstract

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 bias amongst children, adult males appear to get more than adult females. There is a tendency among males for workers to get more than dependents. There is no evidence of differential treatment of the elderly and 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 nonlinear, which suggests that the PIGLOG class of demand models is inappropriate.

Keywords: semiparametric estimation; intrahousehold resource allocation; gender bias; Engel curves.

JEL Nos.: C14, D12, I12.

Intrahousehold Resource Allocation in Rural Pakistan: A Semiparametric Analysis

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1. Introduction

We investigate household food 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 welfare. The second informs the literature concerned with gender bias in South Asia and is of potential use in the calculation of equivalence scales.

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. Since the majority of the rural population in Pakistan spend more than two-thirds of their income on food, it seems reasonable to employ the convention of using foodshare as an (inverse) indicator of their welfare. The rate at which a rise in income generates a decline in foodshare depends upon the shape of the 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 some 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

1 We 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. Income is particularly volatile in agrarian economies.

more slowly than at higher incomes. Once we have identified the curvature of the Engel curve, we can directly 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 nonparametrically for a given parametric specification of the effects of household size and structure on consumption. Controls for compositional heterogeneity across households, as we shall now see, are of independent interest. An advantage of semiparametric estimation is that it offers precise estimates of coefficients on these compositional variables when the functional dependence of foodshare on income is of unknown dimension.

Data from many low income countries reveal that morbidity and mortality rates are higher for females than for males, especially amongst children (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 (note that it is not unusual for children to work). Work-status is expected to proxy needs, albeit imperfectly. In an alternative specification, we control for earnings status. A role for earnings status is consistent with any of the following. Intrahousehold allocation may maximise returns to investment (e.g., Rosenzweig and Schultz, 1982), it may reflect individual bargaining powers (e.g., Manser and Brown, 1980), or it may reward members according to their contribution to household income. Since gender differences in consumption may appear at some ages

[2] See, for example, Banks, Blundell and Lewbel (1994) in the context of estimation of the welfare costs of indirect taxation in the U.K.

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 (see Deaton and Paxson 1992). 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 nonparametric techniques are provided by Silverman (1986), Hardle (1990), Stoker (1991) and Green and Silverman (1994). There is a recent but growing literature that applies these techniques to the estimation of Engel curves. For instance, Banks *et al* (1994), Lewbel (1991) and Delgado and Miles (1996) study expenditure patterns in the UK, the US and Spain respectively. Strauss and Thomas (1990) and Subramanian and Deaton (1996) plot nonparametric 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 the parametrically specified household size and structure 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[4] and this has been attributed to discrimination against females in the intrahousehold allocation of food and health care[5]. The fact that girls also have lower completed schooling (e.g., Strauss and Beegle, 1995) suggests that this might reflect preferences and not differences in needs and endowments. Son preference has been documented in India, where women surveyed in fertility studies report a desire for sons over daughters (e.g., DasGupta, 1987). The generally lower status of females has been put down, alternatively, to cultural factors (e.g., DasGupta, 1987), the exclusion of females from holding immovable property including land (Miller, 1981), dowry practices (e.g., Harriss, 1990) and the relatively low participation of women in agricultural work in areas where they appear most disadvantaged (Bardhan, 1974, 1982).

While the evidence on outcomes like morbidity and literacy is contained in census data, the evidence on inputs like nutrients and hospital admission, which policy can attempt to target, relies upon scattered small-scale investigations. Large sample studies have tended to take the more indirect approach of inferring intrahousehold distribution from data on household expenditure and household composition since this is what is

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[3] There is, for instance, considerable evidence of a negative relation of per capita expenditure and household size (e.g. Lipton and Ravallion, 1994, section 4.2).

[4] For example, D'Souza and Chen (1980), Simmons *et al* (1982), DasGupta (1987), Rosenzweig and Schultz (1982), Sen (1989), Government of India (1988), Coale (1991).

[5] See Basu (1989), Sen and Sengupta (1983), Kynch and Sen (1983), Behrman (1988), Behrman and Deolalikar (1991), Deolalikar (1991), Chen, Huq and D'Souza (1981), and Kielmann *et al* (1990), for example.

available in national surveys[6]. These studies estimate Working-Leser Engel curves (Working (1943), Leser (1963)) for food or adult-goods. Interestingly, more often than not, they find no evidence of gender bias. Where gender differences in consumption patterns do emerge, the existing literature does not attempt to distinguish differential needs from allocational bias. This paper differs from earlier studies in employing a richer, less restrictive specification of age-gender effects, in distinguishing dependents from workers/earners, and in allowing the expenditure function to be determined nonparametrically[7]. While there are indications from small sample studies that girls of higher birth order receive less care (e.g., DasGupta, 1987), large sample investigations of birth order effects are scarce. There is similarly little work that is concerned with how the elderly fare in the intrahousehold distribution of resources but Kochar (1996) presents evidence for Pakistan that adults older than 42 suffer more days of illness than younger adults but do not account for greater medical expenditures.

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[6] See, Browning and Subramaniam (1995), Subramaniam (1996) and Subramanian and Deaton (1991) on India, Ahmad and Morduch (1993) on Bangladesh, Burgess and Juzhong (1997) on China, Rudd (1993) on Taiwan, and Deaton (1989a) on the Cote d'Ivoire and Thailand.

[7] Restricting the form of the foodshare-expenditure relation can bias estimates of household composition effects and thereby alter inferences on the prevalence of gender differences in consumption. For example, Murthi (1994) finds that equivalence scales are sensitive to the functional form of the food Engel curve.

3. Semiparametric Estimation

3.1. The Semiparametric Engel Curve

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

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

where subscript i denotes household, F( ) is unknown, and x_i^T is a vector of J variables representing household size and composition and other relevant covariates. Following Robinson (1988), the first step is to run nonparametric regressions of ω_i and of each of the J elements of x_i on y_i, and generate estimates of the conditional expectations,

$$M_\omega(y) = E(\omega \mid y) \tag{2}$$

$$M_{x(j)}(y) = E(x_j \mid y) \quad \text{for all j} \tag{3}$$

The estimates, m_ω, $m_{x(j)}$ are obtained using the Nadaraya-Watson kernel regression procedure (Nadaraya (1964), Watson (1964)). A quartic kernel is used and the bandwidth is chosen by least-squares cross validation to minimise the integrated mean square error (Hardle, 1990: chapter 5). The second step is to regress the residuals from (2) on the residuals from (3) using ordinary least squares:

$$\omega_i - m_\omega(y_i) = \beta^T (x_i - m_x(y_i)) + v_i \tag{4}$$

where $m_x^T = (m_{x(1)}, m_{x(2)}, ..., m_{x(J)})$. Robinson (1988) shows that the semiparametric estimator, b, of the vector β is $\sqrt{n}$ consistent and asymptotically normally distributed. Define

$$f(y) = m_\omega(y) - b^T m_x(y) \tag{5}$$

Then our estimate of the function F(y) is obtained by graphing f(y) against y.

To obtain the variance-covariance matrix of the estimators, define W as the (nx1) vector of observations on $(\omega_i - m_\omega(y_i))$ for i=1,..., n where n is the sample size. Define X as the (nxJ) matrix of observations on $(x_i - m_x(y_i))^T$. Let $S_{AB} = n^{-1}A^TB$. Then

$$\text{(Estimated) } \mathrm{Cov}\{n^{1/2}(b - \beta)\} = s^2 S^{-1}{}_{XX} \tag{6}$$

where $s^2 = S_{WW} + 2S_{WX}b + b^T S_{XX}b$. To test hypotheses involving combinations of coefficients of the vector β, such as the equality of male and female coefficients at a given age (and work-status), we can use conventional Wald tests since the distribution of $n^{1/2}(b - \beta)$ is asymptotically normal.

3.2. Nonparametric Estimates of Engel Curve Elasticities

It is interesting to find out how expenditure elasticities vary with income. 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 that this can be localised to a part of the expenditure distribution (e.g. if only rich landowning households have home-grown produce), then having nonparametric estimates of elasticities at every point of the Engel curve is especially useful. This is because the elasticity at any point depends only upon the data in the bandwidth around it (see, for example, Deaton, 1996, p. 202). 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.

Ideally, we want to obtain the elasticity of expenditure from the semiparametric model, (2), but this requires estimates of $b^T m_x'(y)$ (the derivative of the second term in (5), where x is a vector). As a first approximation, we assume that these terms are close to zero and obtain elasticities from the nonparametric regression of ω_i on y_i. This seems reasonable since plots (available on request) of the nonparametric and semiparametric Engel curves are very similar in scale and shape. The computation of nonparametric elasticities is straightforward once an estimate of the smooth, differentiable Engel function has been obtained (e.g., Deaton, 1996, pp. 195-6). If there are n data points (ω_i, y_i), and K(z) is the kernel, then

$$m_\omega(y) = (nh)^{-1} \sum_i \omega_i K [(y-y_i)/h] / g(y) \qquad (7)$$

where h is the bandwidth, $m_\omega(y)$ is our estimate of $E(\omega|y)$, and g(y) is the kernel estimate of the density at y, that is,

$$g(y) = (nh)^{-1} \sum_i K [(y-y_i)/h] \qquad (8)$$

The derivative of (7) is

$$m_\omega'(y) = [\, n^{-1}h^{-2} \Sigma_i\, \omega_i\, K'\,[(y\text{-}y_i)/h] - m_\omega(y)g'(y)\,] \,/\, g(y) \qquad (9)$$

$K'(z)$ is the derivative of the kernel. The derivative of the density estimator in (8) is

$$g'(y) = n^{-1}h^{-2}\, \Sigma_i\, K'\,[(y\text{-}y_i)/h] \qquad (10)$$

Recall that Y is per capita total expenditure, y is its logarithm, and ω is foodshare. Then $\partial \ln\omega/\partial \ln Y = \omega^{-1}(\partial\omega/\partial y) = \omega^{-1}(m_\omega'(y))$. Therefore the elasticity of *food expenditure* ($=\omega Y$) with respect to per capita total expenditure is $[1 + m_\omega'(y) / m_\omega(y)]$.

4. The Empirical Specification

We use the Household Income and Expenditure Survey of Pakistan conducted by the Federal Bureau of Statistics in 1987-88 on a stratified random sample of about 18000 households. The investigation is confined to rural households, of which there are 9760 in the sample. The survey collects data for up to 40 members of a household. However, since 99.8% of households in the sample are no larger than 20, we have deleted observations pertaining to the 19 households with more than 20 members. These households exhibit considerable compositional heterogeneity.

Letting i index the household, the estimated model is

$$\omega_i = F(y_i) + \alpha \ln N_i + \Sigma_k\, \gamma_k(N_{ki}/N_i) + \varphi^T z_i + v_i \qquad (11)$$

ω is the share of the household budget that is allocated to food, which is an aggregate of 82 commodities including milk, tobacco, beverages, and alcohol[8]. 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 $\alpha<0$ suggests economies of household size. These may arise as a result of indivisibilities in perishable foods since larger families will then waste less. Alternatively, larger families are 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[9], z denotes birth-order, region and season effects, and v reflects stocastic variations in tastes and other unobservables. Though fertility may be endogenous to the standard of living of the household, we treat household size and structure as exogenous. One may therefore think of the equation as reflecting a long run reduced form. Defining y in per capita rather than in adult equivalent terms is fairly innocuous since log size as well as household composition variables are in the model.

Existing work (refer footnote 6) has 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)[10]. This can be important since there may be gender differentials in consumption at certain ages and not at others. Also, this

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[8] Pakistan is a predominantly Muslim country. Yet the national survey does record some nonzero expenditures on alcohol.

[9] 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.

[10] Age 0 refers to children less than a year old. We follow the convention of defining people 15 years of age or older as adults.

information might enable the researcher to determine the seriousness of any neglect. For instance, Rose (1994) and DasGupta (1987) find 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). While a stepwise constant specification of adult age is less than satisfactory, ours would seem to be a significant improvement on existing work, especially to the extent that the focus is on allocations to children. For reasons mentioned in Section 1, allocations to individuals are permitted to depend upon their *earning status*. In a variant model, we distinguish individuals by their *work status* in an attempt to control for different food needs arising from different levels of activity. It is not unusual for Pakistani children 8 years and older to work/earn. About 12% of boys and 8% of girls aged 8-14 work, while 70% of men and 19% of women work. The distinction between "workers" and "earners" is that "landlords, pensioners, and the like" belong to the latter but not to the former category. It is important to note that women and other unpaid family workers are regarded as "working" *and* "earning" if they contribute at least 15 hours a week to household production[11]. Since the market is often underdeveloped in rural economies, it makes sense to count nonmarket 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

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[11] Only 2.6% of adult females are unpaid family workers compared with 16% of adult males. Amongst children, 0.55% of girls and 5.3% of boys fall into this category. As this is probably contrary to expectation, it is worth stressing that unpaid family work includes domestic tasks, farming on family land, and employment in family enterprises. The latter are not female dominated.

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 over the agricultural cycle for agricultural workers (Payne, 1985), and there are indications that females bear the brunt[12].

5. Results: Food Engel Curves

5.1. The Shape of the Engel Curve

Figure 1(a) shows the semiparametric estimate of the *food* Engel curve, (5). The vertical axis refers to the budget share corrected or standardised for the effects of household size and structure. Consistent with Engel's law, the curve slopes downwards. A quadratic curve fitted by OLS on the same data appears to provide a good approximation to the actual functional form. This is reinforced by the relative fit statistic (Ait-Sahalia *et al*, 1994) being fairly close to zero (Table 1). This would appear to be a feature of the income levels in our sample and the stage of economic development since food Engel curves are linear in the richer populations of Spain (Delgado and Miles, 1996), the UK (Banks, Blundell and Lewbel,

[12] See, for example, Behrman (1988) and Rose (1994) on rural South India and Dercon and Krishnan (1997) on rural Ethiopia.

1994) and the US (Lewbel, 1991)[13]. Foodshare falls less rapidly for poorer households, presumably because their incomes constrain either the quantity or the quality of their food basket. A further possible explanation of the curvature is that indivisibilities in alternative (non-food) commodities are binding at low income levels[14]. 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.

Food is a broad aggregate 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) and Table 1). These are relatively expensive items, which constitute 25% of the average food budget. The inverted U-shape of this Engel curve suggests that milk is a luxury for many households in the sample.

Finding that the food and milk Engel curves are quadratic in log expenditure implies a rank three demand system (Gorman, 1981, Lewbel, 1991). This means that rank two PIGLOG demand systems (Muellbauer, 1975) such as the Almost Ideal Demand System of Deaton and Muellbauer (1980), the log Translog model of Jorgenson, Lau and Stoker (1982) and the Linear Expenditure System of Stone (1954) are inappropriate for analyses of demand in rural Pakistan.

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[13] Working (1943) found nonlinearities in his US mid-1930s sample at low levels of total expenditure 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".

[14] 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.

5.2. Nonparametric Elasticities of Food Expenditure

Nonparametric estimates of the elasticity of food and milk expenditure with respect to total per capita expenditure (Y) can be read off Figures 2(a) and 2(b). Since the Engel curve for foodshare is downward sloping, the expenditure elasticities are less than unity. A striking feature of 2(a) is that the elasticity remains fairly constant, at about 0.90, over a range of low incomes. It is only at levels of Y above the sample mean (Rs. 260 per capita per month) that it begins to fall. For households with Y of Rs. 600 per month, towards the upper end of the distribution, it is about 0.73. The size of the food expenditure elasticities is consistent with the fact that households in the sample are 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*, 1994). Turning to Figure 2(b), there is a general tendency for the elasticity of milk expenditure to decline with rising incomes. Milk and milk products are luxuries for households with Y less than Rs. 400 p.m. and necessities after.

5.3. Scale Effects

The logarithm of household size has a negative coefficient in the foodshare equation, indicating economies of size in *food* consumption. Note that this is in spite of controlling for household composition. The coefficient of -0.015 implies that a doubling of household size would decrease foodshare by 1 percentage point. If size were constant, this decrease in foodshare would require an increase in per capita income of 13%. Thus, where foodshare is used as a measure of household welfare, per capita expenditure understates the welfare of relatively large households. Size economies are likely to be appear for nonrival goods like durables as well, suggesting economic gains to integrated families which the poor may have to take into account. This is pertinent now that the process of economic development is tending to encourage nuclearisation of the family, with young adults leaving their parental village either for work or for greater independence. *Milk* consumption reflects size diseconomies.

5.4. Household Composition Effects

Birth Order Effects

In the foodshare equation, the female birth order variable is insignificant but the male birth order dummy is negative and significant at 10%. There are no birth order effects in milk consumption. There is thus little support here for the hypothesis that parents favour older children.

Gender and Work-Status Effects by Age

This section refers to food Engel curves, there being no significant gender differences in milk consumption (see Table 1, column 3).

Age-gender effects

We permit a full set of age variables for children and group adults into three age categories, 15-24, 25-59 and 60-plus. This is both because we are somewhat more interested in children and because consumption patterns are expected to be more stable for adults than for children. Interaction with gender dummies yields 36 age-gender variables. The compositional variables are jointly significant (Wald test in Table 1).

The presence of male adults in the household appears to increase 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. This might reflect gender bias but it might, instead, reflect male-female differences in tastes or needs. There are no significant gender differences in the consumption patterns of children[15]. Figure 3(a) depicts age effects on foodshare. The loci of points for boys and girls overlap, the fluctuations dominating any level difference. While this is consistent with finding insignificant gender differences in consumption, the spikiness of the age plots suggests that the extent of sampling variation may be obscuring the true pattern.

Distinguishing workers/earners from dependents

When the proportion of workers is included in the model, it emerges with a positive and significant coefficient of 0.04 (elasticity at mean of 0.02), which is interesting given that total expenditure per capita is being held constant. Interaction of age-gender variables with the work status dummy gives 72 variables. Since the semiparametric and parametric models in Table 1 have very similar coefficients on variables like size, it appears that the quadratic is an adequate fit. Moreover, any correlation of the demographic variables with expenditure appears weak. The parametric models in Table 2 show that the expenditure coefficients are insensitive to the precise specification of the demographic variables. Therefore, to avoid unnecessary computation, the following discussion refers to OLS estimates of a (quadratic) parametric model (column 2, Table 2). Appendix Table A

[15] The following are the Wald tests of the equality of the male and female coefficients for the different categories: young adults (χ^2_1=4.7; p=0.03*), prime-age adults (χ^2_1= 23.8; p=0.0*), elderly adults (χ^2_1=5.7; p=0.02*), all adults (χ^2_3=15.8; p=0.0*), and all children (χ^2_{15} =16.1; p=0.37). Each of the 15 age-specific tests for children shows a rejection of the null.

presents Wald tests of the equality of coefficients for (i) males and females, and (ii) workers and dependents.

A gender difference in favour of boys appears for working children though the age-specific tests reveal that it is only at age 8 that the boy coefficient significantly exceeds the girl coefficient, and that the converse is the case at age 9. Amongst adults, the tendency for males to consume more than females persists for 25-59 year olds ("prime-age adults"), whether we are looking at workers or non-workers. However, for 15-24 ("young") and 60-plus ("elderly") adults, the larger male coefficient is only apparent amongst workers. Indeed, for dependents aged 15-24, the female coefficient is significantly larger than the male coefficient. When earnings status replaces work status, the one change is that the gender difference observed amongst prime-age adults becomes restricted to earners. Amongst boys, girls, young males and elderly males, workers and earners appear to get a larger food allocation than dependents. However, amongst adult females, there is no difference between workers and dependents.

The elderly vs. prime-age and young adults

Appendix Table B concentrates on age differences amongst adults rather than on gender differences. The motivating question pertains to the treatment of the elderly. The estimates indicate that food allocations to the elderly are insignificantly different from those to prime-age adults. However, when young adults are not working, they have a significantly smaller effect on foodshare than older adults.

5.5. Region and Season Effects

The Engel curves include province and quarter dummies. The region fixed effects are significant in every case. The seasonal effects are generally small and only occasionally marginally significant.

6. Nonfood Expenditures

6.1. Motivation and Definitions

Since the sum of expenditures 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 suggests 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 son-preference in general.

One option is to estimate Engel curves for all expenditures that are thought to impact upon well-being such as food, health and education. The difficulty with this is that health and education are often subsidised so that private expenditures do not adequately reflect the use of these services. At the same time, the costs of transport and of parental time may constrain their use, introducing room for parents to discriminate between children but in a manner that is not recorded in available data. An alternative is to study the allocation of the sum of expenditures on children (including food). This alternative is the adult-goods method of Rothbarth (1943), discussed in Deaton and Muellbauer (1986) and Deaton *et al* (1989b). For this method to be effective, adult goods consumption must not be influenced by the presence of children in the household. We define adult goods to include tea and coffee, alcohol and drugs, tobacco and tobacco products, men's footwear and women's footwear. However 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, are not valid adult goods. If the chosen composite is in fact an adult good then the coefficients on the child variables will be negative. Relying on a similar notion of separability, we specify child goods 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.

6.2. Estimates of Engel Curves for Adult and Child Goods

Semiparametric Engel curves for child and adult goods are in Figures 1(c) and 1(d) and Table 3. There is a considerable range of incomes over which consumption of adult goods is inelastic, after which it drops. Child goods are luxuries till the very edge of the income range in our sample. As before, the solid lines in the figures are the quadratic Engel curves fitted on the same data. In both cases, they afford a good approximation (relative fit statistics in Table 3). Nonparametric estimates of the expenditure elasticities are plotted in Figures 2(c) and 2(d). Interestingly, adult goods become necessities at incomes above the mean. The coefficients on the logarithm of household size suggest economies in adult goods consumption and diseconomies in child goods. 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. The male birth order variable is positive and significant for child goods, suggesting that, *ceteris paribus*, the first male child costs more than his male siblings.

On account of space constraints, the Wald tests of the gender neutrality of coefficients at different ages are not reported but are available upon request. In the distribution of child goods, boys seem to be favoured over girls once they reach the age of ten (this is clearly seen in Figure 3(c)). The adult goods equation reveals no significant gender differences in the allocation of non-adult expenditures among children, though their allocation amongst adults favours males (see Figure 3(d)).

However, since the coefficients on the child variables are not all negative, we seem to have a composite adult good for which the assumption that children exert pure income effects is untenable. An alternative possibility is that there is gender bias in permanent income. If parents perceive their permanent income as being higher because they have a son rather than a daughter, then they might increase spending on everything including adult goods. This will conflict with the tendency for the coefficients on the boy variables to fall on account of a substitution from adult to nonadult goods in response to the child's presence. A further problem is that many of the adult goods in our composite are addictive in nature, which will limit changes in their consumption (Strauss and Beegle, 1995). Added to this is the fact that their budget shares are often very small (the composite in our sample claims 5% of the budget).

7. Is There No Gender Bias?

Our investigations do not reveal any systematic gender differential in the intrahousehold allocation of food to children. Yet, there is evidence of excess female mortality and morbidity amongst children. There is no necessary contradiction but this underlines the fact that a complete investigation of gender bias would have to go further than this. Our results, like most in the literature, are no more than indicative. The observations on consumption and health can be reconciled in any of the following ways.

(i) 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), DasGupta (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.

(ii) Girls may have greater needs than boys so that, with equal food allocations, they are worse off. This could be the case if, amongst the rural poor, the average boy goes to school while the average girl does a lot of housework, including fetching and carrying water and fuelwood (evidence of hte latter is in Dasgupta (1993), for example).

(iii) 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 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)[16]. Alternatively, if marriage costs arise for daughters but not for sons then households maximising intertemporal utility will begin a

[16] There is evidence of son preference in fertility in rural Pakistan (Rukannudin, 1982), India (Arnold *et al*, 1996) and Taiwan (Parish and Willis, 1992), for instance.

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[17]. While any such mechanisms will tell on the well-being of girls, they will not show up in our investigation since we hold per capita expenditure constant.

(iv) If gender bias characterises only a fraction of households in the sample and we do not incorporate a rule for the selection of households into this group, 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.

8. Conclusions

It is fairly common practice to assume a functional form for demand models. Our semiparametric estimates of Engel curves for rural Pakistan suggest that the popularly used PIGLOG class of demand models is inappropriate. The data favour a quadratic logarithmic specification. In the case of food, the result for Pakistan stands in contrast to that for the US, UK and Spain, all of which have Engel curves linear in the logarithm of expenditure. 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 nonparametric estimate of the elasticity of food

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[17] 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.

expenditure with respect to total expenditure is almost 0.90 at the mean, an indication of how poor these households are.

We find economies of size in the consumption of food and adult goods and diseconomies of size in the consumption of milk and child goods. Food is by far the most important commodity, consuming 52% of the budget on average. The size economies in food consumption suggest that a doubling of household size is equivalent to a 13% increase in per capita income, holding constant composition, region and season. 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. An insight into intrahousehold distribution is useful for policy design to the extent that policy measures excite reinforcing or counteracting responses on the part of household decision makers. For instance, a change in earnings opportunities for females may alter the intrahousehold distribution of nutrition or leisure. We are unable to identify systematic gender differences in consumption amongst children. There is also no robust evidence of birth-order effects. Our estimates suggest that household food consumption would register a statistically significant decline if an adult male were replaced by an adult female, even if both work and earn. We also find that, whether they are children or adults, male workers tend to 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 prime-age adults, irrespective of whether they work.

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Table 1
Semiparametric and Parametric Engel Curves for Food and Milk

Variable	Foodshare (1) Semiparametric	Foodshare (2) Parametric	Milkshare (3) Semiparametric	Milkshare (4) Parametric
constant		-0.44 (3.5)		-1.10 (12.3)
ln pcE		0.14 (3.2)		0.39 (12.7)
$(\ln pcE)^2$		-0.021 (5.3)		-0.033 (11.9)
ln size	-0.015 (5.4)	-0.016 (5.7)	0.012 (5.2)	0.012 (5.2)
female birth order	-0.001 (0.2)	-0.001 (0.3)	0.003 (1.0)	0.002 (0.8)
male birth order	-0.006 (1.8)	-0.006 (1.7)	0.002 (0.7)	0.002 (0.7)
(age)x(gender)	yes	yes	yes	yes
province dummies	yes	yes	yes	yes
season dummies	yes	yes	yes	yes
R^2		0.18		0.065
log likelihood		9776.86		11347.66
relative fit statistic		0.031		0.042
χ^2_{35} (age)(gender)	75.8 (p=0.0)	76.6 (p=0.0)	50.3 (p=0.05)	49.4 (p=0.05)
mean of dep var	0.52	0.52	0.13	0.13
expen. elasticity	see Figure 2(a)	0.84	see Figure 2(b)	1.14
size elasticity	-0.029	-0.031	0.023	0.023
gender patterns	For 15-60 year olds alone, male coefficient exceeds female coefficient	as in column (1)	At age 7 alone, male coefficient exceeds female coefficient.	as in column (3)

Notes: pcE=per capita expenditure, sample size after 1% is trimmed at low densities=9643. Standard errors are Newey-West, corrected for heteroskedasticity. Absolute t-ratios are in parentheses. The relative fit statistic (Ait-Sahalia *et al*, 1994) compares the fit of the semiparametric curve with the parametric least squares curve. It consists of the mean square difference of the predicted budget share from the semiparametric and parametric models divided by the sample variance of the semiparametric prediction. The closer the statistic is to zero, the better is the fit. The age-gender variables denoting household structure are the gender-specific proportions of children of age 0,1,..,14 and of adults aged 15-25, 25-60 and 60^+. "Gender patterns" are based on a χ^2 test of the null hypothesis that the male coefficient equals the female coefficient for every age/age-group. Elasticities are at means, mean (ln pcE)=5.625, mean (ln size)=1.697. The expenditure and size elasticities refer to responses of food expenditure, not foodshare. These are the foodshare elasticities plus one.

Table 2
Introducing Controls for Work Status
Food Engel Curves : OLS Estimates

Variable	(1) Age x Gender	(2) Age x Gender x Work-status
constant	0.391 (3.7)	0.377 (3.6)
ln pcE	0.163 (4.5)	0.163 (4.5)
$(\ln pcE)^2$	-0.023 (7.2)	-0.023 (7.2)
ln size	-0.018 (6.4)	-0.016 (5.8)
Household composition:		
(age)x(gender)	yes	
(age)x(gender)x(work status)		yes
female birth order dummy	-0.001 (0.2)	-0.000 (0.1)
male birth order dummy	-0.005 (1.4)	-0.004 (1.0)
Province dummies:		
Punjab	-0.056 (16.1)	-0.054 (15.6)
NWFP	-0.026 (6.2)	-0.023 (5.5)
Sind	-0.032 (8.7)	-0.032 (8.8)
Season dummies:		
quarter 1	0.004 (1.7)	0.004 (1.4)
quarter 2	-0.001 (0.4)	-0.001 (0.6)
quarter 3	-0.004 (1.4)	-0.004 (1.6)
Adj.R^2	0.211	0.217
N	9741	9741
χ^2 (hh structure)	χ_{38}^2 =232.14 (p=0.0)	χ_{58}^2 =339.24 (p=0.0)

Notes: See Notes to Table 1. The sample size is larger in the absence of trimming. This is why column 1 here is not identical to column 2 in Table 1. The dummies for Baluchistan province and for the fourth quarter are suppressed. χ^2 (household structure) is a Wald test of the joint significance of the variables describing household structure and size.

Table 3
Semiparametric and Parametric Engel Curves for Adult and Child Goods

	Adult goods share		Child goods share	
Variable	(1) Semiparametric	(2) Parametric	(3) Semiparametric	(4) Parametric
constant		0.002 (0.1)		-0.174 (7.7)
ln pcE		0.026 (1.7)		0.053 (6.5)
$(\ln pcE)^2$		-0.003 (2.4)		-0.004 (5.6)
ln size	-0.002 (2.0)	-0.002 (2.5)	0.007 (12.6)	0.007 (12.6)
female birth order	-0.000 (0.3)	-0.000 (0.3)	-0.000 (0.4)	0.000 (0.4)
male birth order	-0.001 (0.9)	-0.001 (0.8)	0.002 (2.4)	0.002 (2.4)
(age)x(gender)	yes	yes	yes	yes
province dummies	yes	yes	yes	yes
season dummies	yes	yes	yes	yes
R^2		0.113		0.169
log likelihood		20429.41		23878.58
relative fit statistic		0.019		0.0046
χ^2_{35} (age)(gender)	276.6 (p=0.0)	266.1 (p=0.0)	397.4 (p=0.0)	393.8 (p=0.0)
mean of dependent variable	0.053	0.053	0.017	0.017
expenditure elasticity	see Figure 2(c)	0.853	see Figure 2(d)	0.532
size elasticity	-0.038	-0.038	0.412	0.412
gender patterns	For 15-60 year olds, male coefficient exceeds female coefficient. At age 3 alone, the female coefficient is larger.	as in column (1)	Above age 10, the male coefficient exceeds the female coefficient.	as in column (3)

Notes: N=9643. Also see Notes to Table 1.

Appendix Table A
Gender and Work-Status Effects on Foodshare
Wald tests of coefficient equality

(1) ADULTS

working males vs working females
Joint : $\chi^2_3=53.0;\ p=0.0^*$
Young: $\chi^2_1=6.7$; $p=0.0^*$
Prime-age: $\chi^2_1=18.4$; $p=0.0^*$
Elderly: $\chi^2_1=2.9$; $p=0.10$

dependent males vs dependent females
Joint: $\chi^2_3=19.7;\ p=0.03^*$
Young: $\chi^2_1=8.8$; $p=0.00^*$
Prime-age: $\chi^2_1=4.3$; $p=0.04^*$
Elderly: $\chi^2_1=0.22$; $p=0.64$

male workers vs male dependents
Joint : $\chi^2_3=47.5;\ p=0.0^*$
Young: $\chi^2_1=37.6$; $p=0.0^*$
Prime-age: $\chi^2_1=0.65$; $p=0.42$
Elderly: $\chi^2_1=7.5$; $p=0.0^*$

female workers vs female dependents
Joint : $\chi^2_3=17.6;\ p=0.06$
Young: $\chi^2_1=0.90$; $p=0.34$
Prime-age: $\chi^2_1=0.98$; $p=0.32$
Elderly: $\chi^2_1=1.06$; $p=0.30$

Summary: The male coefficient exceeds the female coefficient for workers and for prime-age dependents. The coefficient on workers is larger than on dependents amongst young and elderly males.

(2) CHILDREN

working boys vs working girls
Joint (8-14 years): $\chi^2_7=19.5;\ p=0.01^*$
Age 8: $\chi^2_1=13.6$; $p=0.0^*$
Age 9: $\chi^2_1=8.3$; $p=0.0^*$

dependent boys vs dependent girls
Joint (8-14 years): $\chi^2_7=5.6;\ p=0.59$
Joint (0-7 years): $\chi^2_8=10.1;\ p=0.26$
(Children≤7 years are all dependents)

boy workers vs boy dependents
Joint (8-14 years): $\chi^2_7=30.2;\ p=0.0^*$
Age 8: $\chi^2_1=12.0$; $p=0.00^*$
Age 13: $\chi^2_1=3.4$; $p=0.07$
Age 14: $\chi^2_1=5.3$; $p=0.02^*$

girl workers vs girl dependents
Joint (8-14 years): $\chi^2_7=18.4;\ p=0.01^*$
Age 8: $\chi^2_1=9.8$; $p=0.0^*$
Age 9: $\chi^2_1=12.6$; $p=0.0^*$

No gender difference for dependent children. For working children, the male coefficient is larger. Workers and dependents have different coefficients in both gender groups.

Notes: These tests are based on the coefficients on household structure in Table 2. *Children* are defined as 0-14 year olds. For adults, *young*=15-25, *prime-age*=25-60, *elderly*=60 and older. A * indicates that the null hypothesis of coefficient equality is rejected at the 5% level of significance. Except for 9 year old child workers and young adult dependents, the null is rejected in favour of the male coefficient exceeding the female coefficient. Similarly, except for 8-year old girls, the null is rejected in favour of the coefficient on workers exceeding that on dependents. Age-specific tests are not shown when the coefficients are equal. Let *w*=worker, *d*=dependent. Then the results imply the following broad patterns. For young adults, *wm>wf=df>dm.* For prime-age adults, *wm=dm>wf=df*. For the elderly, *wm>wf=df=dm.* If *workers* is replaced by *earners*, the only change is that there ceases to be any significant gender difference amongst prime-age dependents

Appendix Table B
Elderly vs. Prime-Age vs. Young Adults
Wald tests of coefficient equality
***Format*: Hypothesis (Wald statistic; p-value)**

(1)
Age x Gender

all persons
elderly=prime age (χ^2_2=0.53; p=0.77)
prime-age=young (χ^2_2=7.2; p=0.03)*
elderly=young (χ^2_2=4.0; p=0.14)

males:
elderly=prime-age (χ^2_1=0.55; p=0.81)
prime-age=young (χ^2_1=7.14; p=0.01)*
elderly=young (χ^2_1=3.35; p=0.07)
females:
elderly=prime-age (χ^2_1=0.53;p=0.45)
prime-age=young (χ^2_1=0.47;p=0.49)
elderly=young (χ^2_1=0.66;p=0.80)

Summary: Prime age males do better than young males. The other age differences considered are insignificant.

(2)
Age x Gender x Work-status

workers:
elderly=prime age (χ^2_2=0.47; p=0.49)
prime-age=young (χ^2_2=0.27; p=0.60)
elderly=young (χ^2_2=1.7; p=0.43)
dependents:
elderly=prime age (χ^2_2=0.21; p=0.65)
prime-age=young (χ^2_2=11.3; p=0.0)*
elderly=young (χ^2_2=11.6; p=0.0)*

male workers:
elderly=prime-age (χ^2_1=0.74; p=0.39)
prime-age=young (χ^2_1=0.32; p=0.57)
elderly=young (χ^2_1=1.28; p=0.26)
male dependents:
elderly=prime-age (χ^2_1=0.82; p=0.37)
prime-age=young (χ^2_1=13.7; p=0.0)*
elderly=young (χ^2_1=10.7; p=0.0)*
female workers
elderly=prime-age (χ^2_1=0.21; p=0.65)
prime-age=young (χ^2_1=0.14; p=0.71)
elderly=young (χ^2_1=0.42; p=0.52)
female dependents
elderly=prime-age (χ^2_1=0.43; p=0.51)
prime-age=young (χ^2_1=0.32; p=0.86)
elderly=young (χ^2_1=0.27; p=0.61)

There is a tendency for elderly and prime-age adults to do better than young adults but this is only amongst male dependents. Any other differences are insignificant.

Notes: The tests use parameter estimates on the household structure variables in the models reported in Table 2. A * indicates that the null hypothesis of coefficient equality is rejected at 5%. *Elderly*=60 and older, *prime-age*=25-59, *young*=15-24 years.

Figure 1. Semiparametric and Quadratic Engel Curves

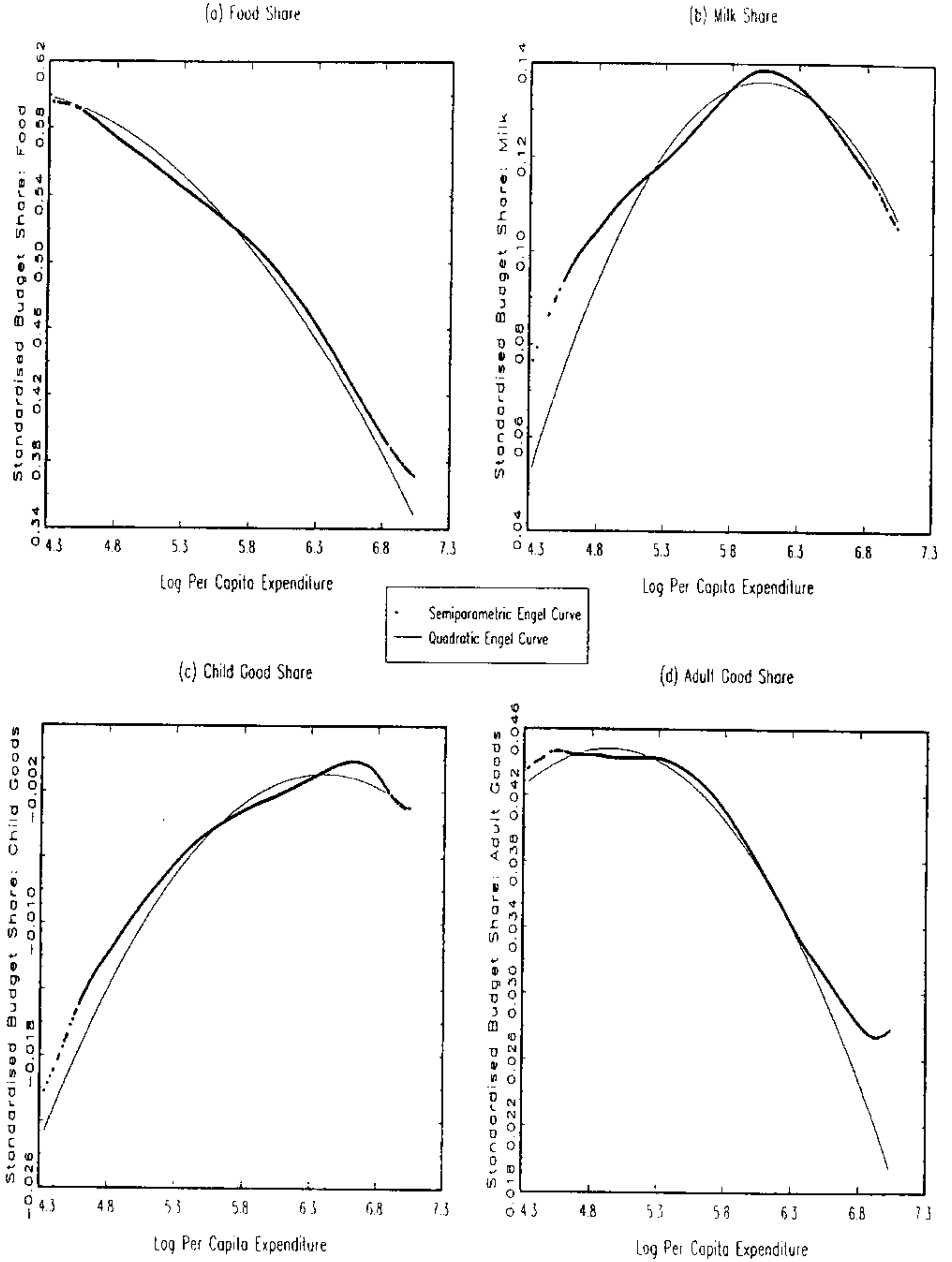

Figure 2. Expenditure Elasticities from Nonparametric Engel Curves

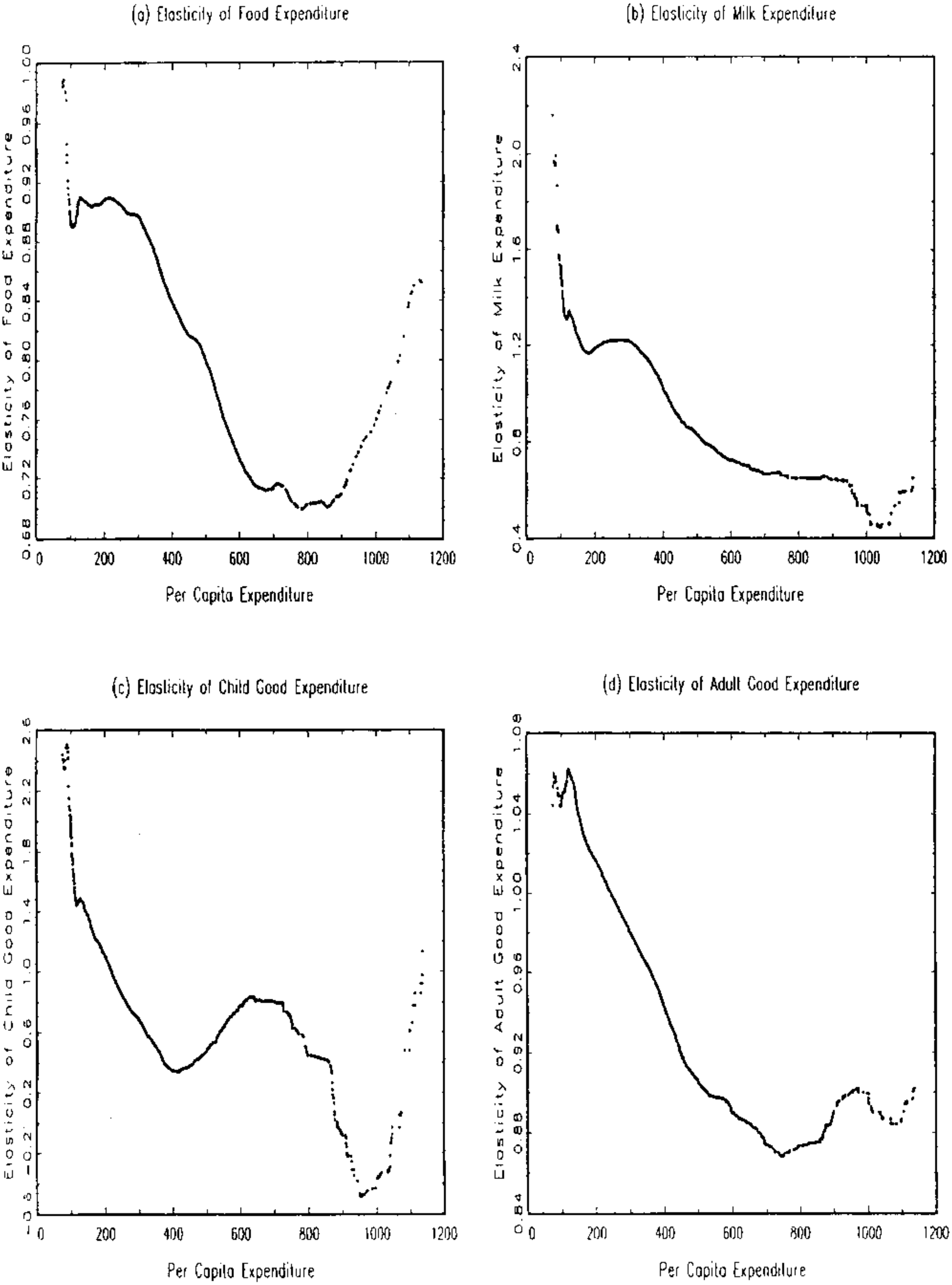

Figure 3. Estimated Coefficients on Gender/Age Dummies from Semiparametric Estimation