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GENDER-BIAS IN INDIA: THE IMPORTANCE OF HOUSEHOLD FIXED-EFFECTS

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ABSTRACT:

This paper presents an analysis of gender patterns in intra-household allocation of resources based on household level consumption data. Invoking the assumption that households seek to equalize the marginal utility of wealth when they allocate resources over the life-cycle, the paper provides a rationale for parental behaviour pertaining to the intertemporal allocation of goods among children. Estimation results based on panel data from India show that controlling for the unobserved marginal utility (household fixed) effect is crucial. Once allowance is made for fixed effects, the results indicate that there is no longer any gender-differential in the allocation of resources.

JEL Classification: D10, J16, O12

Keywords: intra-household allocation, marginal utility of wealth, wealth-shock.

1. Introduction

The allocation of resources within the household has become an important research issue in recent years. Any analysis of the distribution of income across households could present a wrong picture if there are inequalities in the distribution of resources within households (see Haddad and Kanbur (1990)). There is considerable evidence that resources are not allocated randomly within households, and that the distribution of resources, pecuniary as well as non-pecuniary, is unequal within the family in many developing countries[1], particularly in Asia. Such an unequal distribution of goods usually takes the form of a bias against females or children. In India, for example, evidence based on mortality rates and human capital investments (spending on health, nutrition and education) indicate that there is a bias against girls in the intra-household allocation of resources (Bardhan (1984); Behrman (1988); Harriss (1990); Rosenzweig and Schultz (1982); Sen (1984); Sen and Sengupta (1983))[2].

While the evaluation of the standard of living of a household as a unit can be done on the basis of household-level data, the measurement of the welfare of each individual within a household poses a difficult problem since household surveys do not usually provide any information on the consumption of each individual. Further it is problematic to assign a particular level of

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1. See the references on India cited below; Deolalikar (1993) for Indonesia, Parish and Willis (1993) for Taiwan, Schultz (1993) on investments in health and education in many groups of developing countries. Thomas (1990) finds some evidence for such non-random allocation of resources in Brazilian households. For an overview of the literature on within-household resource allocation, see Haddad, Hoddinott and Alderman (1994). The evidence provided by Deaton (1987) and Svedberg (1990) shows that there is relatively little gender bias in Africa when compared with Asian countries.

2. In the Indian case, North-South distinction seems to be important in explaining gender-bias in intra-household allocation of food and other resources. See Harriss (1990).

consumption to each individual since many commodities are consumed jointly within the household. In the case of food, there are many nutrition surveys[3] that collect data on individual intakes based on 24-hour recall information. Such surveys, however, do not collect data on other goods at an individual level. Almost all the studies on gender-bias in intra-household allocation are based on nutrition surveys; only a few studies (see Deaton (1987,1989), Haddad and Hoddinott (1991) and Subramanian and Deaton (1991)) look at household level consumption expenditure data to examine the presence, or absence, of inequalities within the household.

Existing studies of the intra-household allocation of goods use models in which the allocation is determined in one of the following four ways: (i) parents allocate resources based on the differential labour market returns to boys and girls (Rosenzweig and Schultz(1982)), (ii) parents allocate resources according to their own utility, which depends on the well-being of their children (an approach due to Behrman, Pollak and Taubman(1982), Behrman(1988)); (iii) households allocate resources based on the productivity of individual members (Pitt, Rosenzweig and Hassan(1990)); (iv) resources are allocated according to the relative bargaining power of the family members (Haddad and Hoddinott (1991) and Thomas(1990)). The demand theory approach by Deaton (1987,1989) can be categorized under (ii). This paper belongs to (i) above, in that parents allocate resources to male and female children in a pure investment sense, but it uses household demand data.

The goals of this paper are two-fold. First, to determine the pattern of intra-household allocation by looking only at household level consumption data.

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3. The Indian data set used in this paper also has a nutrition survey (see Behrman (1988)). Pitt, Rosenzweig and Hassan (1990) utilize a nutrition survey from Bangladesh.

Second, to look for a rationale as to why parents may discriminate between male and female children in the intra-household allocation of resources, in some countries or societies. This paper is also an attempt to take up the hitherto neglected role of dynamic optimization in intra-household allocation problems in a simple way. For this purpose, I use a unique panel of consumption data from India.

There is plenty of anthropological and other evidence from India that parents spend a considerable amount of resources on the marriages of female children (see Bardhan (1984), Harriss (1990), Miller (1980) and Rao (1993)). The practice of giving dowry for the marriages of female children has long been in existence in many parts of India. I hypothesize that such high costs of raising female children may be an important determinant of (differential) allocation of resources within the household. In a life-cycle context, the birth of a girl may have the same impact as a negative shock on life-time household wealth. For example, if the child born is a girl, then parents may start a life-cycle savings plan in order to meet marriage expenses at a later date. Browning and Subramaniam (1995) provide evidence on marriage costs for the households in the data set used here. They find that cash dowry, without including in-kind dowry which is also substantial, and marriage expenses amount, on average, to a year's income for the households. In such cases, the presence of an additional girl in the household may lead to a reduction in total household expenditure and also to a reduction in its major component, food expenditure[4].

My approach to test this hypothesis is to estimate several Engel curves to test for the effects of gender on consumption patterns. In a typical Engel curve

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4. Srinivas (1984) observes that paying large sums of dowries cause considerable economic hardship to households with daughters to be married.

relationship, total household expenditure is held constant. Under this method, it is possible to test for the impact of gender only on the composition of demand. For example, if we take two households with the same total outlay to spend, then only the composition of demand can change with the gender-composition of the household (since total expenditure, or outlay, is held constant, spending less on one good means more expenditure on some other good). Therefore I estimate equations in which total household income is included as a major explanatory variable instead of total household expenditures. This could capture any gender-bias in intra-household allocation that comes from parental consumption and saving behaviour following the birth of a child.

The cross-section Engel curve equations with household income as the main explanatory variable indicate that male children receive more resources than female children in the case of food, with the result more pronounced for young male children. These cross-section equations are useful only for determining the pattern of intra-household allocation. In order to model parental behaviour, as outlined above, pertaining to the intertemporal allocation of resources within the household, I control for household specific fixed effects utilizing the panel nature of the data.

The following section provides an outline of the framework and the estimation procedure. In Section 3, I provide a description of the data used in this paper and present the estimation results. Section 4 summarizes the findings.

2. Framework and Estimation Procedure

Modern theories of intertemporal allocation assume that individuals attempt to keep their marginal utility of expenditures constant over time (see Browning

(1992) or Browning, Deaton and Irish (1985)). This assumption could offer an intuitive and plausible explanation for parental economic behaviour in allocating resources over time, as children come into and leave the household[5]. Parents may favour male children over female children in some social and cultural settings, India for example, because of several factors: economic contributions made by female children may be considerably less than that made by male children to the household, or returns to investments in girls may be considerably lower than for boys.

In many countries, labour market opportunities are limited for women[6]. Yet another important factor could be marriage-related financial commitments that daughters impose on their parents without any tangible returns. In such contexts parents may optimally allocate the limited resources available, based on a cost-benefit calculation. Or, to put it differently, parents may view daughters as a drain on their wealth. If that is the case, a preponderance of female children will take the form of a wealth-effect: girls will increase parental marginal utility of lifetime wealth, because they may potentially reduce their parents' wealth as a result of the above factors. Then, in a utility maximizing framework, parents may reduce household expenditures following the birth of a female child, because it has the same effect as a negative shock on life-time wealth. They may allocate more resources to boys (because the marginal utility of wealth is lower than for girls) and fewer to girls, since they want to keep

5. Browning and Subramaniam (1995) present a model of life-cycle intra-household allocation and derive testable predictions for three competing hypotheses: differential allocation due to wealth-effect or parental preferences or differential returns.

6. Rosenzweig and Schultz (1982) provide evidence that differential returns in the labour market play a crucial role in explaining male-female children survival differentials.

their marginal utility of expenditures constant over time. I assume here that parents agree on the allocation of resources within the household. However, as Thomas (1990) shows, incomes under the control of mother and father may have different effects on the allocation of resources. Haddad, Hoddinott and Alderman (1994) discuss the different policy implications of unitary versus non-unitary models of within-household allocation of resources.

To analyze the relationship between the gender composition and consumption patterns of households, I start from the general equation for an Engel curve

$$C_{ht} = f(Y_{ht}, a_{ht}, Z_h, \alpha_h, u_{ht}) \qquad \textbf{(1)}$$

where C_{ht} is household consumption expenditure at time t, Y_{ht} is total household income, a_{ht} is a vector of household demographic composition variables, Z_h is a vector of time-invariant control variables, u_{ht} is the standard error term, and α_h is an unobserved household-specific fixed effect. The vector a_{ht} may include the numbers of males and females in different age groups, and age variables. The vector Z_h may contain variables that control for spatial and temporal effects[7].

If households are assumed to keep marginal utility constant, the specification above shows a simple way to formalize the process of intra-household allocation. The household fixed effect, α_h, can be visualized as the unobserved marginal utility of household expenditures. Let the stream of marginal utilities over time (from Period 1 to Period T) of household h be given by: $(\alpha_{h,1}, \alpha_{h,2}, \ldots, \alpha_{h,t-1}, \alpha_{h,t}, \alpha_{h,t+1}, \ldots, \alpha_{h,T})$. The constant marginal utility assumption states that each period's marginal utility will be equal to some fixed

7. The list of variables in these two vectors that are included in the estimation equations is provided in Section 3.1 below.

α_h. Then if there is a change in the gender composition of children in household h at time t the marginal utilities can be ranked as follows:

(1) No wealth-shock: $\alpha_{h,t+1} = \alpha_{h,t-1}$

(2) Wealth-shock: if the household has one more girl, then $\alpha_{h,t+1} > \alpha_{h,t-1}$; if the household has one more boy, then $\alpha_{h,t+1} < \alpha_{h,t-1}$.

The following figure illustrates the wealth-shock idea:

INSERT FIGURE 1 HERE

Under such a specification, without controlling for the household specific fixed effect, one would expect to see differential gender effects on household consumption expenditures. Once the unobserved effect is controlled for, there should not be any differences in parental allocative behaviour towards male and female children. But it should be emphasized that the household fixed effect may represent other time invariant variables as well, besides the unobserved marginal utility of wealth[8]. For example, in the fixed effects version, the impact of time invariant characteristics such as caste or village effects cannot be estimated but they may be included in the fixed effect.

In order to test the above hypothesis, I estimate a levels-version and a fixed-effects version of equation (1). To be more specific, I estimate equations of the following form:

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8. Browning and Subramaniam (1995) formalize the ideas presented here and look at the issue of identification: their model shows that the differences in marginal utility of wealth, which households attempt to keep constant over time (treated as a fixed effect), determine parental behaviour in the allocation of resources between male and female children. They also disentangle this type of wealth effect from parental preferences for bias.

$$\log C_{ht} = \beta_o + \beta_M M_{ht} + \beta_F F_{ht} + \beta_Y Y_{ht} + \beta_z Z_h + u_{ht} \qquad (2)$$

where log C_{ht} is the log of (annual) consumption expenditure for household h, M_{ht} and F_{ht} are the numbers of males and females in three age groups (adults, young and old children) in the household, and Y_{ht} is the log of household income. Testing for gender-effects on consumption patterns in this framework would involve a simple F-test on the equality of the β_M and β_F coefficients. An inequality such as $\beta_M > \beta_F$ has been interpreted in the literature as evidence that boys are favoured over girls (see Subramanian and Deaton (1991)[9]). I find that the same result holds for the rural Indian sample used. However, in the framework of the model presented here, as I have pointed out above, this result may not indicate that boys are getting more resources, but instead that households with predominantly more girls are "poorer" and must adjust their consumption patterns over the life-cycle accordingly. Thus the finding that the coefficients for males and females are unequal may be a consequence of an omitted wealth effect that is correlated with the gender structure of the household. I use the within-estimator approach to control for the unobserved (fixed) wealth-effect in the Engel curves: parental response to a female birth takes the form of reducing current consumption (in order to keep marginal utility constant). In the fixed effects version, one can no longer reject the equality of the gender coefficients, suggesting that the levels-version result favouring males is a consequence of the unobserved marginal utility of wealth effect.

9. Subramanian and Deaton estimate budget share equations for different categories of goods using the National Sample Survey data from the state of Maharashtra in India.

I also estimate a variant of equation (2) in which I include dummy variables for the presence of male and female children in the household, in addition to the numbers of male and female children in the household. This would allow us to test if there is any differential (gender) effect based on the birth-order of children. Harriss (1990) presents some evidence that within-household discrimination is more likely to occur against higher order female children. The results in the present paper show that having a female child, holding everything else constant, reduces total expenditures of the ICRISAT households considerably.

Endogeneity of Income: In the specification above in equation (2), I have treated household income as an exogenous variable on the premise that households respond to a wealth-shock by changing total expenditures, holding income constant. However, this need not be the case, since, for a given (negative) wealth-shock, households could respond in two ways: (i) by reducing the consumption of goods, in order to save; and/or (ii) by reducing the consumption of leisure, in order to increase their income. If (ii) holds, then income could no longer be treated as exogenous in the specification. One way to address this problem is to instrument income using variables that are correlated with income but not with the unobserved fixed effect (marginal utility of wealth). I construct a village by landholding group by year median income variable for each household in each of the four landholding groups (landless, and three cultivating groups based on land-size: small, medium and large) for each year, and use this variable to instrument income. For example, for household 1 in the Landless group in Village 1 in Year 1976, this variable would be the median of the incomes of all the <u>other</u> (that is, excluding household 1) households in the Landless group in Village 1 in Year 1976. This variable has two advantages over using some other instrument like household assets, for example. Household assets (or wealth) may not be completely uncorrelated with the unobserved marginal utility of wealth effect[10]. The variable I construct also captures an important aspect of household economies in the study villages: that incomes of households within a landholding group in a village may be correlated, given their dependence on weather[11].

Therefore, the levels estimates are obtained using two-stage least squares. The fixed effects model is estimated using ordinary least squares, since the marginal utility effect (and thus the correlation between income and wealth shock) is purged by the inclusion of the fixed effect, and the random effects estimates, which are not presented, are obtained using GLS[12].

3. Data

The data used in the paper come from the ICRISAT VLS (International Crops Research Institute for the Semi-Arid Tropics Village Level Studies) Panel Data set. These data are from a survey initiated by ICRISAT in 1975 in three distinct agroclimatic regions of semi-arid tropical India. The data set contains information on consumption, income and production for 120 continuous households

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10. Browning and Subramaniam (1995) find that wealthier households in the ICRISAT villages spend a considerably higher amount (about 130 per cent of a year's income) on the marriages of daughters.

11. This variable also explains the variation in household incomes over time a lot better than a set of village-by-year interaction dummies, which is another way of capturing the impact of weather and other spatial factors on household incomes.

12. It should be noted that household income is not the only endogenous right hand side variable. For example, household composition could be endogenous (if not the sex of children); timing of fertility could be endogenous to the consumption (demand) process. The reader should bear in mind that these potential sources of endogeneity are not investigated in this paper.

from three villages (two from the State of Maharashtra and one from the State of Andhra Pradesh) for a period of ten years from 1975-84. Within each village, a random sample of 40 households (ten labour households and thirty cultivating households) was chosen for the panel. ICRISAT also conducted four rounds of a nutrition survey in the study villages during 1976-78. I use the data pertaining to the period 1976-81, because consumption data[13] were not collected in detail at the beginning in 1975 and at the end after 1981.

Table 1 presents some descriptive statistics for the variables used in estimation. I use the food price index to deflate nominal food expenditures and the general price index to obtain real income and real total expenditures. The share of food in total expenditure is around 70% for all the three villages. The 6-year standard deviation in household food expenditures is proportionately less than that in household incomes. The average household size is 6, and the average total number of children (age 0-15) is 2.5.

3.1. Results and Analysis

3.1.1 *Discussion of Results from Food Expenditure Equations*

In this section, results from the estimation of equation (2) are presented. Here, apart from the numbers of adult males and females (age > 15), numbers of male and female children in two age groups (0 to 4 years old, and 5 to 15 years old) and household income in period t, I also include the mean ages and the

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13. Of the total 120 households from the three villages, data are available for 104 households for all the years from 1976 to 1981. Five of these households were dropped from the analysis in this paper: two single person households were not included, and three other households in which all the children were reported to be living outside the household units for the entire length of the panel were also dropped. In addition, income data were missing for five of the households (one year for each household). In the analysis, I have dropped these years rather than dropping the households altogether. Thus the sample size is 589 in the tables presented below.

square root of the mean ages of male and female children. These variables account for the changing gender and age composition within households over time. The mean age variables are included in order to account for the discontinuities that arise from including only the numbers of children in the two age bands[14]. The vector Z_h contains a number of village(2), household caste(2) and time(5) dummies.

Table 2 presents the estimation results for 6 equations. I first analyze the levels-version results: all the levels equations are estimated using two-stage least squares to correct for the endogeneity of income. In equation (2.1) there are four children groups and equation (2.2) is a restricted version of (2.1). In 2.3, I include two dummy variables to indicate the presence of a male or female child in the household, along with the numbers of male and female children. The most important thing to note in these two equations is the significant positive effect for all the male groups. Female children in the age group 0-4 do not have any significant effect on food expenditures in (2.1), and in (2.2) when female children are aggregated, they do not have any effect. Table 4 presents the F-test results for equality of gender effects: in (2.1) there is a significant difference between male and female children in the 0-4 group [$F(1,568)=7.80$, $p=0.005$], but not for the 5-15 age group. However, the overall difference, for the 0-15 groups, is significant in (2.2) [$F(1,570)=7.79$, $p=0.005$].

The coefficients on the gender-dummy variables in 2.3 show that having a female child, ceteris paribus, reduces household food expenditure significantly,

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14. For example, the impact on food expenditures of a 6 year old is likely to be more similar to that of a 5 year old than it is to that of a 14 year old. See Browning (1992) for a detailed discussion of various ways of including children variables in consumption equations. See also Banks and Johnson (1993).

by about 16 per cent; but having a male child does not have any significant effect on household food expenditure[15]. However, there is no significant difference between the dummy coefficients themselves. The first male child increases (though not significantly) household food expenditure by about 7 per cent and the first female child decreases (significantly with $F(1,568)=4.34$, $p=0.037$) it by about 12.5 per cent (F-test value for this difference is $F(1,568)=4.01$, $p=0.046$). The marginal effect of the first child is the sum of the estimated coefficients for the child-dummy and the number of children variables. Note that the specifications here do not pick up any possible non-linearities in the marginal effects of children on consumption. In terms of gender-birth order effects, the F-test result [$F(1,568)=5.01$, $p=0.026$] in Table 4 for equation 2.3 shows that higher order male children have a significant effect on expenditures.

The results from 2.1 and 2.2, based on household level expenditure data, are not directly comparable with the results in the literature based on intra-household data: Behrman (1988) uses the ICRISAT nutrition data set to look at the distribution of individual nutrient intakes (and of individual health outcomes) of sibling children and finds that female children on average get less food than male children during lean seasons. Based on the same survey, Behrman and Deolalikar (1990) find that females have lower price elasticities of demand for nutrients than males. But the results above from the levels-equations do not imply that girls get fewer nutrients than boys. They suggest a cross-sectional response on the part of the households, in the sense that households with more girls have lower food expenditure. Or, households with predominantly more girls

15. Variables that further disaggregated the categories of children based on birth order were not significant in the equations, indicating that the first vs higher order distinction is sufficient.

are impoverished. However, who bears the burden of the lower food expenditures within these households requires a separate analysis. Behrman's (1988) finding that boys are favoured may arise due to the wealth-effect of female children; however, he does not find any differential allocation in favour of boys during surplus seasons.

Another important point made in the gender-bias literature is that second born or higher order female children are more likely to discriminated against in the intra-household allocation of resources (Harriss (1990), Das Gupta (1984)). Equation 2.3 provides a test of this point, in a limited way, by including a dummy variable for the presence of a child to distinguish between first born and higher order children. The results show that households with a female child have a much lower expenditure on food, and that the first born female child has a big negative effect on food expenditure. The marginal response of having one more girl, is not significant whereas there is a significant positive marginal response of having one more boy. These results are not in line with the above cited evidence for birth-order discrimination: they indicate that the first female child has a much bigger wealth shock than higher order female children.

In all three equations, the difference between adult males and adult females is significant (see the F-test results in Table 4). While it is not clear why adult gender-composition exhibits a significant difference here, these results may support the finding that men consume more nutrients than women in the ICRISAT households (see Walker and Ryan (1990)). Most of the other variables are significant: the income elasticity is around 0.50. Of the age variables, the male children age variables are significant. Both the village dummies and two of the time dummies are also significant (see Table 6). Income variability within the villages and over time is one possible reason for the significance of these variables. The six year standard deviation was higher than the mean income for the representative household in Aurepalle, the reference village in the equations.

In the first three equations of Table 2, there is no control for the unobserved household fixed effect, α_h, which is characterized as the parents' marginal utility of wealth. Columns 4 to 6 of Table 2 present the fixed effects estimates which control for the unobserved effects by estimating the model in deviation form (or, the within-estimator approach). In equations (2.4) to (2.6), no male children group has any significant effect on food consumption, with the only significant effect being that of female children in the 5-15 group. The F-test results presented in Table 4 show that there is no difference between the effects of male and female children on food expenditures, with controls for fixed effects: the bias in favour of male children, which is present in the levels equations, vanishes with control for the unobserved household fixed effect. If the household fixed effect is not controlled for, one would conclude that male children are associated with higher household food expenditures. The children dummy variables are also not significant here (equation 2.6), indicating that having a female child does not lead to a reduction in household food expenditures, as in the levels equations. Thus these results clearly suggest that, with appropriate controls for the wealth-adjustments that parents make, there is no gender-bias in the intra-household allocation of food over and above the impact of such an intertemporal behaviour. There is no significant difference between adult male and female groups. The male age coefficients are significant and income is still significant; however, the effects of the time-invariant variables cannot be observed in the fixed-effects version.

The last three rows of Table 2 present some test statistics for comparing across the levels, random effects (GLS) and the fixed effects models. The F-test results presented in the first of these rows clearly favour the hypothesis that the household fixed effects are highly significant in the food expenditure equations. These results were obtained by estimating the fixed effects equations with 98 household dummies. The second row presents the results to test for the hypothesis that there are no error components (random effects vs no effects), this is a Lagrange-Multiplier test: these are also highly significant. The random effects estimator is a GLS estimator which accounts for the fact that the households are observed six times during the program. However, I have presented only the levels (2SLS) results, which treat the data in a cross-section framework ignoring the panel feature, for purposes of comparison with the cross-section results available in the literature which suggest that there are gender differentials in intra-household allocation. The last row in Table 2 presents the Hausman test statistics to compare the fixed effects and the random effects results: the test values, which are significant at at least the 0.0005 level, are clearly in favour of the fixed effects model.

3.1.2 *Discussion of Results from Total Expenditure Equations*

In order to test the hypothesis that parents may be forced to reduce total household expenditures (hence, food expenditures), I have estimated the above equations with total expenditure as the dependent variable. The estimation results are presented in Table 3, F-test results are presented in Table 5 and the coefficients on the time-invariant control variables are reported in Table 6. There is statistically strong support for this hypothesis, at least for young children and for the presence of children (as captured by the dummy variables). Male children in 0-4 age group significantly increase total expenditures whereas female children do not have any effect and there is a significant difference between them [F(1,568)=8.05, p=0.005]. However in the case of children in the 5-15 age group, a similar difference is not found. There is a significant difference between male and female children in the overall (0-15) group in both the equations.

The coefficients on the gender-dummy variables in equation 3.3 show that having a female child reduces household total expenditure significantly and considerably, in support of the wealth-shock hypothesis, by about 18 per cent; having a male child has a positive effect on total expenditure, increasing it by 6 per cent, but not significantly. There is also a significant difference between these dummy coefficients [F(1,568)=5.19, p=0.023]. The first male child <u>increases</u> (though not significantly) household total expenditure by about 13 per cent and the first female child <u>decreases</u> (significantly with F(1,568)=5.16, p=0.023) it by about 15 per cent. The F-test value for this difference between the impact of the first male vs first female child is F(1,568)=7.01, p=0.008. In terms of gender-birth order effects, there is no significant difference among higher male and female children. An important difference here, when compared with the results from food expenditure equations, is that the adult gender composition does not have any significant effect on total expenditures in the levels equations. The income elasticity is higher, at around 0.56, and many of the time-invariant variables are significant.

Qualitatively, the fixed effects equations for household total expenditures are very similar to the results presented above for food expenditures. Child gender-composition variables do not have significant impact on total expenditures, with the exception of female children in the 5-15 age group (equation 3.4 in Table 3). Male children, in either age group or in the aggregate, are no longer significantly different from young female children, with control for fixed effects. Another important point to note is that the children dummy variables are no longer significant. However, with control for fixed effects, adult gender- composition variables have a significant effect on total expenditures as shown in the first row of Table 3. Both adult males and females increase total expenditures, but with no significant difference between them. It looks like with control for the marginal utility effect, in moving from the levels to the fixed effects equations, the effect of the child gender-composition variables is transferred to the adult gender-composition variables. The F, LM and the Hausman test results presented in the last three rows of Table 3 are strongly in favour of the fixed effects model.

The cross-sectional (or the levels) and the panel results presented above are in support of the wealth-shock hypothesis: households with at least one female child are poorer in the sense that their total expenditures are lower by about 18 per cent, and food expenditures by about 16 per cent. The marginal utility of wealth hypothesis provides a plausible motive for parental behaviour in Indian rural households, that parents initiate a life-cycle savings program upon the birth of a female child by reducing current consumption.

4. **Summary and Conclusions**

In this paper, I examine the issue of gender-bias in intra-household allocation of resources from the perspective of the modern theories of intertemporal allocation. The empirical analysis is based on a panel of household level consumption data from India. Several Engel curve relationships are estimated, with household income as the first explanatory variable, to determine gender patterns in consumption. The levels-version results of this paper are in line with the findings in the literature that parents favour boys over girls in the allocation of resources. However, given the household nature of the data used, these results should be interpreted in a cross-sectional sense, that female children are associated with lower current household expenditures rather than from an intra-household point of view.

Based on the assumption that households seek to equalize the marginal utility of expenditures when they allocate resources, I model the parental marginal utility of wealth as the unobserved household-specific fixed effect that stays constant over time. If parents optimally allocate resources over the life-cycle, then they would reduce current consumption expenditures in order to increase savings following the birth of a female child, because female children increase the marginal utility of wealth, by reducing life-time wealth due to high marriage costs. The fixed effects estimation results, based on this rationale, indicate that once the unobserved household fixed effect is controlled there is no evidence of any significant gender-bias in the intrahousehold allocation of food and other resources. These results point out that the differential effects of gender-composition, against female children, in the allocation of resources derive entirely from the presence of the unobserved wealth effect. Browning and Subramaniam (1995) provide supporting evidence that the wealth-adjustments over time, in the form of increased savings and reduced consumption in households following a female birth, match closely with the marriage costs of daughters in the ICRISAT households. While the results in this paper do not directly imply that female children receive fewer resources within the household, as the bulk of the intra-household literature suggests, female children may receive fewer resources as a result of their negative wealth effect.

The results in this paper also have a number of important policy implications. They suggest that households with more female children have lower consumption than other households. Miller (1980) and Harriss (1990) provide anthropological evidence that dowries and marriage costs are significantly lower in some parts of India than other parts. For example, the prevalence of dowry is much higher in general in the North than in the South. In some Southern parts of the country where wet-irrigation is practised, with more earnings opportunities for women, there is a system of reverse-dowry or bride-price in existence. Thus, to the extent that female children contribute in the form of earnings to the household pool of income, one would expect the impact of the wealth-effect on household consumption to be lower. Given that the legislative attempts through the Dowry Prohibition Act of 1961 have been largely unsuccessful, policies to promote labour market opportunities for women are the only plausible means to solve this problem with its roots in the culture of the country.

Table 1 : Descriptive Statistics for ICRISAT Households, 1976-81[a]

Variables	Median	SD
Food Expenditure	3166.00	1523.00
Log of Food Expenditure	7.95	0.46
Total Expenditure	4617.00	2346.00
Log of Total Expenditure	8.32	0.47
Total Household Income	6188.33	5754.93
Number of adult males	1.77	0.98
Number of adult females	1.75	0.96
Number of male children, 0-4	0.39	0.67
Number of female children, 0-4	0.30	0.56
Number of male children, 5-15	0.99	0.92
Number of female children, 5-15	0.79	0.89
Household size	5.99	2.78
Mean age of male children	5.98	4.69
Mean age of female children	5.09	4.89

a. Number of observations:594; all expenditure values are in 1983 rupees.

Notes:
(1) Food also includes edible oils and fats. Non-food expenditures include: narcotics, tea, coffee, tobacco, pan, alcohol, clothing, tailoring expenses, chappals and footwear, medicines, cosmetics, soap and barber services, travel and entertainment, electricity, water charges, cooking fuel, labour expenses for domestic work and other expenses.

(2) Total household income is the sum of: net trade income, net crop income, net livestock income, net land rent, salary and wage incomes.

Table 2 : Food Expenditure Equations - Levels (Two Stage Least Squares) and Fixed Effects Estimates for ICRISAT Data:1976-81 (Dependent Variable - Log of real food expenditure)[a,b]

	Levels Equations - 2SLS Estimates			Fixed effects Equations		
Variables	2.1	2.2	2.3	2.4	2.5	2.6
Males (>15)	0.067 (0.032)	0.068 (0.022)	0.067 (0.022)	0.084 (0.031)	0.084 (0.030)	0.086 (0.029)
Females (>15)	0.009 (0.021)	0.009 (0.020)	0.012 (0.020)	0.087 (0.024)	0.085 (0.023)	0.073 (0.023)
Males (0-4)	0.135 (0.032)	--	--	0.035 (0.026)	--	--
Females (0-4)	0.017 (0.031)	--	--	0.027 (0.032)	--	--
Males (5-15)	0.051 (0.021)	--	--	0.037 (0.030)	--	--
Femls (5-15)	0.031 (0.023)	--	--	0.066 (0.031)	--	--
Males (0-15)	--	0.089 (0.019)	0.089 (0.018)	--	0.029 (0.020)	0.035 (0.020)
Femls (0-15)	--	0.027 (0.016)	0.029 (0.019)	--	0.041 (0.025)	0.041 (0.026)
Male child dummy	--	--	-0.019 (0.084)		--	-0.004 (0.025)
Female child dummy	--	--	-0.157 (0.064)		--	-0.007 (0.022)
Log real income[c]	0.505 (0.081)	0.501 (0.079)	0.485 (0.075)	0.163 (0.023)	0.163 (0.023)	0.188 (0.024)
Mean age of Male Children	0.044 (0.014)	0.032 (0.012)	0.025 (0.015)	0.030 (0.016)	0.028 (0.015)	0.029 (0.015)
Sqrt(Mean age of Male Chn)	-0.147 (0.053)	-0.124 (0.047)	-0.091 (0.073)	-0.103 (0.061)	-0.091 (0.060)	-0.099 (0.061)
Mean age of Fem Children	-0.005 (0.014)	-0.005 (0.012)	-0.021 (0.014)	0.022 (0.017)	0.026 (0.016)	0.024 (0.016)

Sqrt(Mean age of Fem Chn)	0.016 (0.049)	0.025 (0.046)	0.112 (0.059)	-0.087 (0.063)	-0.087 (0.062)	-0.079 (0.062)
Intercept	3.221 (0.649)	3.256 (0.621)	3.406 (0.602)	--	--	--
R-squared	0.604	0.606	0.621	0.193	0.192	0.211
Fixed vs No Effects (F)	--	--	--	F(98,479) 5.20	F(98,481) 5.29	F(98,479) 5.05
Random vs No Effects (LM)[d]	Chsq(1) 92.42	Chsq(1) 91.76	Chsq(1) 87.07	--	--	--
Random vs Fixed Effects[d]	--	--	--	Chsq(11) 44.39	Chsq(9) 74.66	Chsq(11) 31.96

a. Sample size is 589 in all the columns; Standard errors in parentheses; b. All time-invariant controls are reported in Table 6. c. Endogenous variable in levels equations, see text for instruments used. d. Lagrange-Multiplier and Hausman tests, respectively.

Table 3 : Total Expenditure Equations - Levels (Two Stage Least Squares) and Fixed Effects Estimates for ICRISAT Data:1976-81 (Dependent Variable - Log of real total expenditure)[a,b]

	Levels Equations - 2SLS Estimates			Fixed effects Equations		
Variables	3.1	3.2	3.3	3.4	3.5	3.6
Males (>15)	0.034 (0.024)	0.039 (0.023)	0.037 (0.023)	0.064 (0.030)	0.069 (0.029)	0.071 (0.029)
Females (>15)	0.009 (0.022)	0.009 (0.021)	0.015 (0.021)	0.093 (0.023)	0.091 (0.022)	0.079 (0.022)
Males (0-4)	0.149 (0.034)	--	--	0.034 (0.025)	--	--
Females (0-4)	0.021 (0.033)	--	--	0.024 (0.031)	--	--
Males (5-15)	0.019 (0.022)	--	--	0.001 (0.029)	--	--
Femls (5-15)	0.036 (0.024)	--	--	0.055 (0.030)	--	--
Males (0-15)	--	0.076 (0.019)	0.071 (0.019)	--	0.011 (0.019)	0.016 (0.019)
Feml (0-15)	--	0.015 (0.019)	0.036 (0.020)	--	0.035 (0.024)	0.035 (0.025)
Male child dummy	--	--	0.057 (0.088)	--	--	-0.001 (0.023)
Female child dummy	--	--	-0.183 (0.067)	--	--	-0.010 (0.021)
Log real income[c]	0.576 (0.086)	0.566 (0.083)	0.546 (0.079)	0.137 (0.022)	0.134 (0.022)	0.152 (0.023)
Mean age of Male Children	0.040 (0.015)	0.021 (0.012)	0.021 (0.016)	0.021 (0.015)	0.013 (0.014)	0.015 (0.015)
Sqrt(Mean age of Male Chn)	-0.125 (0.056)	-0.086 (0.050)	-0.097 (0.077)	-0.063 (0.059)	-0.039 (0.057)	-0.047 (0.059)

Mean age of Female Chn	-0.002 (0.014)	-0.003 (0.013)	-0.019 (0.015)	0.013 (0.016)	0.018 (0.016)	0.015 (0.015)
Sqrt(Mean age of Fem Chn)	0.001 (0.052)	0.015 (0.049)	0.106 (0.062)	-0.051 (0.061)	-0.055 (0.059)	-0.046 (0.059)
Intercept	3.047 (0.688)	3.138 (0.658)	3.316 (0.633)	--	--	--
R-squared	0.586	0.586	0.608	0.166	0.164	0.180
Fixed vs No Effects (F)	--	--	--	F(98,479) 8.14	F(98,481) 8.55	F(98,479) 7.86
Random vs No Effects (LM)[d]	Chsq(1) 117.76	Chsq(1) 123.64	Chsq(1) 115.76	--	--	--
Random vs Fixed Effects[d]	--	--	--	Chsq(11) 129.63	Chsq(9) 89.93	Chsq(11) 160.22

a. Sample size is 589 in all the columns; Standard errors in parentheses; b. All time-invariant controls are reported in Table 6. c. Endogenous variable in levels equations, see text for instruments used. d. Lagrange-Multiplier and Hausman tests, respectively.

Table 4 : F-test Results for Gender Equality in Food Expenditure Equations[a]

	Levels Equations - 2SLS Estimates			Fixed Effects Equations		
Variable	2.1	2.2	2.3	2.4	2.5	2.6
Adults	4.690 (0.030)	4.920 (0.027)	4.260 (0.039)	0.000 (0.913)	0.000 (0.937)	0.090 (0.768)
Children,0-4	7.800 (0.005)	--	--	0.040 (0.827)	--	--
Children,5-15	0.046 (0.498)	--	--	0.038 (0.547)	--	--
All Children	3.900 (0.020)	7.790 (0.005)	--	0.370 (0.689)	0.110 (0.734)	--
Adults and Children	4.280 (0.005)	6.740 (0.001)	--	0.260 (0.857)	0.060 (0.943)	--
First Child	--	--	4.010 (0.046)	--	--	0.000 (0.967)
Higher Order Children	--	--	5.01 (0.026)	--	--	0.020 (0.888)

Table 5 : F-test Results for Gender Equality in Total Expenditure Equations[a]

	Levels Equations - 2SLS Estimates			Fixed Effects Equations		
Variable	3.1	3.2	3.3	3.4	3.5	3.6
Adults	0.780 (0.379)	1.230 (0.270)	0.620 (0.432)	0.390 (0.541)	0.250 (0.626)	0.040 (0.852)
Children,0-4	8.050 (0.005)	--	--	0.060 (0.795)	--	--
Children,5-15	0.260 (0.610)	--	--	1.410 (0.233)	--	--
All Children	4.770 (0.009)	4.370 (0.037)	--	1.230 (0.294)	0.490 (0.493)	--
Adults and Children	3.550 (0.014)	2.940 (0.053)	--	0.840 (0.476)	0.300 (0.744)	--
First Child	--	--	7.010 (0.008)	--	--	0.050 (0.827)
Higher Order Children	--	--	1.51 (0.220)	--	--	0.290 (0.587)

a. Numbers in parentheses are p-values in both the tables.

Table 6 : Food and Total Expenditure Equations - Levels Estimates for Time Invariant Control Variables[a]

	Food Exp. Equations			Total Exp. Equations		
Variables	2.1	2.2	2.3	3.1	3.2	3.3
Dummy for Village 3[b]	0.323 (0.034)	0.323 (0.034)	0.316 (0.032)	0.401 (0.036)	0.397 (0.035)	0.386 (0.034)
Dummy for Village 5	0.074 (0.039)	0.088 (0.037)	0.087 (0.037)	0.175 (0.041)	0.196 (0.039)	0.191 (0.039)
Dummy (1976)[c]	0.018 (0.044)	0.019 (0.043)	0.014 (0.043)	-0.000 (0.046)	0.001 (0.046)	-0.004 (0.045)
Dummy (1977)	-0.058 (0.045)	-0.057 (0.045)	-0.063 (0.044)	-0.143 (0.048)	-0.143 (0.048)	-0.148 (0.047)
Dummy (1978)	-0.039 (0.044)	-0.037 (0.044)	-0.042 (0.043)	-0.104 (0.046)	-0.102 (0.046)	-0.106 (0.045)
Dummy (1979)	0.131 (0.043)	0.136 (0.043)	0.136 (0.042)	0.073 (0.046)	0.079 (0.045)	0.080 (0.044)
Dummy (1980)	0.195 (0.043)	0.195 (0.042)	0.189 (0.042)	0.138 (0.045)	0.138 (0.045)	0.131 (0.044)
Low Caste Dummy[d]	0.151 (0.076)	0.134 (0.070)	0.125 (0.069)	0.157 (0.080)	0.126 (0.075)	0.118 (0.073)
Medium Caste Dummy	0.066 (0.045)	0.066 (0.043)	0.055 (0.043)	0.064 (0.047)	0.058 (0.046)	0.048 (0.045)

a. Standard errors in parentheses; b.Village 1 is dropped;c.Year 1981 is dropped;d.High caste is dropped.

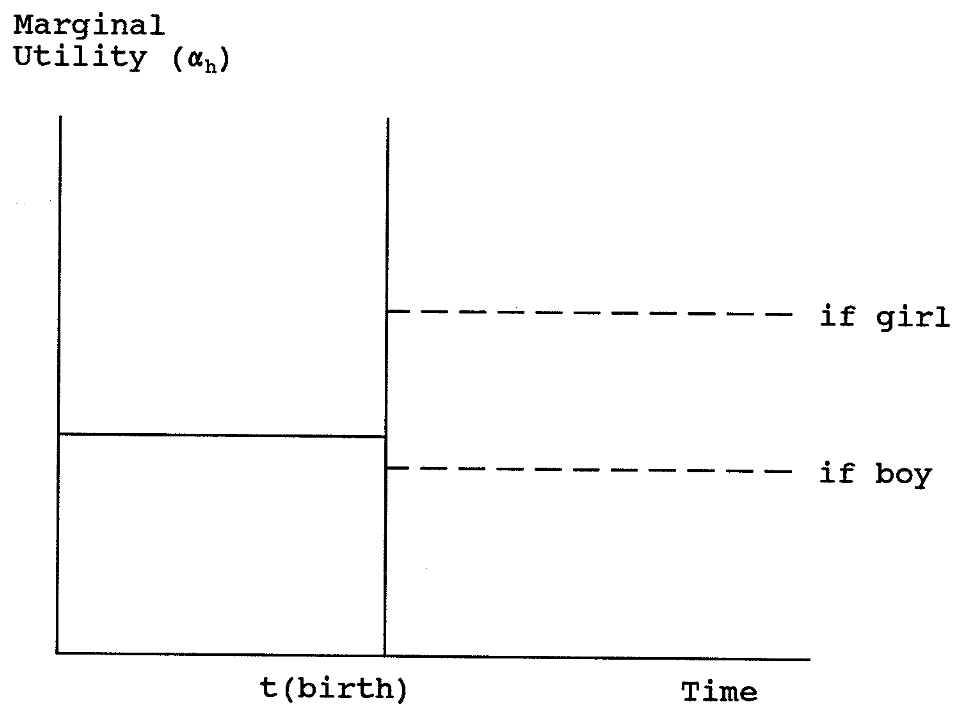

Figure 1 : Marginal Utility of Wealth at Birth

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