Is it Better to be a Boy? A Disaggregated Outlay Equivalent Analysis of Gender Bias in Papua New Guinea

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The demographic evidence of gender bias in many countries has provided an impetus for finding ways to study the status of women in developing countries. Because of the lack of accurate intra-household data, Deaton [1989] introduced a method for using household expenditure data to infer discrimination in the allocation of goods between boys and girls. Few studies of discrimination using the method, however, have detected bias even though alternative indicators suggest it is a serious problem. In this paper, we study the case of Papua New Guinea, a country in which there are many indicators of severe gender bias. Discrimination in the allocation of goods between boys and girls within households in Papua New Guinea is examined using Deaton's outlay-equivalent ratio method. Adding a boy to the household reduces expenditure on adult goods by as much as would a nine-tenths reduction in total outlay per member, but girls have no effect on adult goods expenditure. The hypothesis of Haddad and Reardon [1993] that gender bias is inversely related to the importance of female labour in agricultural production is not supported. Sensitivity analysis shows that bias in rural areas occurs equally regardless of the age of the household head, while bias against girls may be less in regions of the country that have ethnic groups which practice matrilineal descent.

I. INTRODUCTION

The question of unequal divisions between men and women and boys and girls inside the household is the topic of much recent research. Unequal allocations cause poverty and inequality to be understated when we use

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The Journal of Development Studies, Vol.40, No.4, April 2004, pp.115–136
ISSN 0022-0388 print
DOI: 10.1080/00220380410001673210

measures that assume that every household member is treated evenly [*Haddad and Kanbur, 1990*]. The demographic evidence of 'missing women' in Bangladesh, China, India and Pakistan provides further impetus for this line of research [*Dreze and Sen, 1989*]. However, the difficulty of observing the inner workings of households restricts many studies to 'externally' observable outcomes like health and nutrition [*Sen, 1984*]. Studies of individual consumption within the household, such as Pitt, Rosenzweig and Hassan [*1990*] on food in Bangladesh, are less common because they need data that are difficult and expensive to obtain in a reliable manner.

Because of the lack of accurate intra-household data, inferential methods of detecting gender bias using just standard household expenditure are attractive. Deaton [*1989*] introduced a method for using household expenditure data to infer discrimination in the allocation of goods between boys and girls. The method starts, somewhat paradoxically, with expenditures on goods that are not consumed by children – for example beer, gambling, and tobacco products. The addition of a child to the family is treated as a *pure* (negative) income effect on the demand for these goods, since the child does not contribute to family income and also is unlikely to participate in the consumption of these goods. Evidence for gender bias can then be found by (statistically) answering the question: Is the reduction in spending on these 'adult goods' larger when the additional child is a boy rather than a girl?

Despite the cleverness of Deaton's inferential method, the results have been disappointing. Curiously, in its few applications, the method sometimes finds bias against girls in locations where it is not expected, but no bias in places where other evidence strongly suggests that males are favoured. For example, Ahmad and Morduch [*1993*] find no evidence of gender bias in Bangladeshi household expenditure data, whereas the imbalance in sex ratios suggests considerable discrimination against girls. In Indian states where excess mortality and low literacy point to bias against girls, Subramanian [*1994*] finds no bias in household expenditures against girls; however, in Maharashtra, a state in which social indicators suggest the status of girls is relatively better, the expenditure method indicates there is a bias. Deaton [*1997*] lists other examples from Pakistan and Taiwan where the expenditure method finds no bias despite other studies indicating a preference for boys in these countries.

Several explanations have been advanced for the failure of the adult goods method to detect bias. One econometric reason may be sample truncation in cases where girls have been so discriminated against that they have died [*Udry, 1997*]. Alternatively, girls may have higher needs than boys, so that equal allocations still adversely affect girls [*Ahmad and Morduch, 1993*]. Critical interventions might not be made for girls when they are made for boys. A further reason for the failure of the method is that it has been used in some places that are not likely candidates for bias, such as Côte d'Ivoire [*Deaton, 1989*] and Burkina Faso [*Haddad and Reardon, 1993*]. In these countries, women are economically productive so girls are not seen as a burden on their parents. Haddad and Reardon advance this argument further by searching for differences in the degree of gender bias across agro-ecological zones. Their hypothesis, derived from Boserup [*1970*] and Rosenzweig and Schultz [*1982*], is that discrimination against girls will be less as the economic opportunities for women increase (such as in urban areas or in agricultural settings with high income potential), suggesting that the greatest discrimination will be in rural areas where the potential income from agriculture is low. According to Haddad and Reardon [*1993*], discrimination in households is expected to decline as the study area moves to higher potential rural areas; declining even further when the study examines urban areas.

In this paper we provide an example in which the adult goods method detects gender bias in a place, Papua New Guinea (PNG), where, *ex ante*, we expect there to be bias. As will be argued below, there are several reasons for our strong results. First, a number of economic indicators and the weight of other social science work suggest that males are favoured in PNG. Second, adult goods consumption, for a variety of reasons, is important in this setting (more than 12 per cent of the budget), and its high level may make the results more easily detectable. Third, the enormous heterogeneity in the sample data may facilitate identification of these effects; we have a nationally representative sample in a country that is almost certainly home to the most diverse set of ethnic and social groups of humans in the world. Fourth, unlike previous studies (to the best of our knowledge), we designed and managed the collection of the data ourselves, planning from the start to undertake adult goods analysis. As such, the questionnaire contains a set of well-defined categories for adult goods, and enumerators were taught to exercise extra care when collecting these data. Finally, unlike some previous analyses, we use econometric techniques and statistical tests that control for the survey design effects in our data.

To examine discrimination in PNG, the rest of the paper is organised as follows. Section II briefly examines gender bias in PNG. The two next sections review the methodology and describe the data and estimation approach. Section V contains the basic results. In this section, we look first at gender discrimination, in general, and perform sensitivity analysis with the age groups of children. Section VI presents disaggregated results, examining gender discrimination by agro-ecological region and between urban and rural areas, by the age group of the household head, and according to the dominant descent rule.

II. GENDER BIAS IN PNG

A preponderance of evidence has been amassed by social scientists that boys are highly favoured in PNG. The male share of the population is 51.6 per cent, a demographic imbalance similar to India and Pakistan [*UN, 1996*]. This is consistent with reports of a higher under-five mortality rate among females than among males; according to WHO [*1998*] the female under-five mortality rate is 10 per cent higher than for males.[1] This differential is unusual because higher mortality rates amongst girls are observed in only a few other countries, such as Bangladesh, China and India, where male preference is prevalent [*WHO, 1998*].

The literacy rate for adult females, at only 44 per cent, is 18 percentage points lower than for males [*Gibson, 2001*] and 49 per cent of women have never attended school, which is nearly twice the rate for men [*World Bank, 1999*]. Among those Papua New Guineans aged 25 and over, men have had about 1.3 years of schooling, yet this already-low figure is still almost double the attainment for women [*UNDP, 1994*]. In addition, a substantial gender gap in school enrolment is apparent at all schooling levels [*UNESCO, 1999*]. Women and girls in PNG also appear to be disadvantaged in health and nutrition [*Groos and Garner, 1988*], as well as in many other aspects of life [*Brouwer et al., 1998*].

Many other social scientists also find the PNG society is discriminatory against women. For example, Modjeska [*1982*] found that 'big men' manipulated economic activities (especially control over land and other communal resources) to dominate the social system, which allowed them not only to rise above other men in social status and power in PNG's highly stratified society, but also to establish a dominance over women. According to Fox [*1999*], such a strong masculine bias originates from the deeply embedded social belief that women are subordinate to men, and this leads to a lack of recognition and public acknowledgement of women's role in PNG society [*Cox and Aitsi, 1988*]. Consequently, insufficient attention and resource commitment, both by households and by the government, are paid to women's matters [*Overfield, 1995; Grossman, 1998*]. For example, it has been found that households in PNG invest more in the health of boys than of girls and that the boys take precedence over the girls when the family is economically constrained in child schooling decisions [*Fox, 1999*]. Such discriminatory practices against the girl child in turn result in excess mortality among girls and lower educational attainment for women.

Nevertheless, this bias is interesting because it exists in a country where women are economically productive – female labour is often considered to be a scarce factor in agricultural production in relative land-rich PNG [*Overfield, 1998*].[2] Positive bride-prices and the prevalence of polygamy are symptoms of this scarcity. A number of researchers have documented the value of women in agricultural production. For example, Johnson [*1988*] found that the two strongest explanations of success in coffee growing areas were the number of wives in a household and the number of non-wife females in the household. Overfield and Irog [*1992*] found a positive and statistically significant association between availability of female labour and success in cash cropping. Thus, PNG may be a counter-example to the view advanced by Haddad and Reardon [*1993*] that bias against girls may be 'rational' in the sense that it can be explained by factor endowments and production relations in agriculture.

The importance of adult goods in PNG household budgets and the concentration of cash earnings in the hands of adult males may mean that the adult goods method gives an unusually clear view of any tensions over intra-household consumption allocations. Adult goods potentially account for 12–13 per cent of the household budget in PNG, a figure that is robust across agro-ecological zones (Table 1, row 1). In comparison, in one country where adult goods analysis has previously been used – Burkina Faso – the budget share of *candidate* adult goods is reported as less than 10 per cent, but is probably even lower, since many of these goods do not meet the statistical criteria for true adult goods [*Haddad and Reardon, 1993*]. Budget shares for a similar array of adult goods in poor areas of other countries, such as China, are also significantly lower (below 8 per cent) [*Zhang, 1999*].

While true in many countries, the cash for purchasing both adult and other goods is unequally concentrated in the hands of male adults in PNG, even though women do play an important role in many economic activities. For example, women make up 87 per cent of the participants in sweet potato production, the dominant staple crop, and 83 per cent for other food crops [*World Bank, 1999*]. However, men more frequently perform cash earning jobs (comprising over 80 per cent of the participants), and undertake more rewarding cash cropping activities (comprising well over half of the participants). Even within cash cropping activities there appears to be discrimination against women; the returns to men's own-account coffee production are up to twice as high as the returns to women [*Overfield, 1998*].

Costs associated with raising children in PNG, especially those who are school-aged (typically from age seven onwards), may exacerbate tensions over the budgetary needs of children. Although school and health fees at one time were highly subsidised by the government, by the time of the survey, 1995, most of these had been eliminated through fiscal cuts, structural adjustment shifts, and bureaucratic neglect [*Jarrett and Anderson, 1989*]. For example, the average outlay of a poor family for a single secondary school student is equivalent to 36 per cent of the *household's* non-food expenditure [*World Bank, 1999*]. With a devaluing currency, children's clothing and shoes, most of which are imported, have become more expensive. In short, children, as every parent knows, can be expensive to raise, and almost certainly mean sacrifice in other areas. Hence, given these higher costs, in a society with a revealed preference towards boys (as shown by other indicators), with a high propensity to consume adult goods, and with a large part of the cash in the hands of males, if the adult goods analysis can be statistically implemented, we should expect to find it revealing gender bias in PNG. However, in areas in which an investment in a girl can raise her future income earning potential, the extent of the discrimination against girls may perhaps be attenuated.[3]

III. METHODS

The first step is to identify the set of adult goods. Following Deaton, Ruiz-Castillo and Thomas [*1989*] one can use the linear model:

$$p_i q_i = b_{0i} + b_{1i} x_G + \sum c_{ij} n_j + \mathrm{d}_i \cdot \mathbf{z} + v_i \tag{1}$$

where $p_i q_i$ is expenditure on candidate adult good i, x_G is total expenditure on adult goods, n_j is the number of people in each of eight age and gender classes, $\mathbf{z}$ is a vector of control variables, and v_i is a random error. Children affect only the total expenditure allocated to adult goods, not the allocation within the adult goods group once total group expenditure is given. Therefore, if good i is a genuine adult good, the age and gender of children should play no part in equation (1), which can be tested with an F-test on the relevant estimated c_{ij} coefficients (that is, those coefficients on the variables associated with the age categories for children).

An alternative test of whether a list of candidate goods contains only *genuine* adult goods uses the concept of 'outlay-equivalent ratios'. For any normal commodity i and demographic category r, the outlay-equivalent ratio π_{ir} is:

$$\pi_{ir} = \frac{\partial(p_i q_i)/\partial n_r}{\partial(p_i q_i)/\partial x} \cdot \frac{n}{x} \tag{2}$$

where π_{ir} is an expression that measures the effect of an additional person of type r on the demand for good i, measured as the percentage change in outlay (expenditure) per person that would have been necessary to produce the same effect on demand. The effect of an additional child should be that adult goods expenditure falls, given the extra needs that the child brings. Because this effect is exactly like a reduction in income, the reduced expenditure on each individual adult good ought to be in proportion to the marginal propensities to spend on each good [*Deaton, 1989*]. Hence, the π-ratios for any particular type of child should be the *same* across all of the adult goods, so this implication provides another test for whether the candidate goods are indeed adult goods.

The outlay equivalent ratios can be calculated from the coefficients produced for any estimated Engel curve. Following Deaton [*1989*], the Engel curve used here is specified as:

$$w_i = \frac{p_i q_i}{x} = \alpha_i + \beta_i \, ln\left(\frac{x}{n}\right) + \eta_i \, lnn + \sum_{j=1}^{J-1} \gamma_{ij} \left(\frac{n_j}{n}\right) + \delta_i \cdot \mathbf{z} + u_i \tag{3}$$

where w_i is the budget share for the ith adult good, x is the value of total household consumption, n is household size, n_j is the number of people in the jth demographic group, and z is a vector of control variables. The estimated parameters of equation (3) can then be used to calculate:

$$\pi_{ir} = \frac{(\eta_i - \beta_i) + \gamma_{ir} - \sum_{j=1}^{J} \gamma_{ij}(n_j/n)}{\beta_i + w_i} \tag{4}$$

for $r = 1,\ldots, J$, where γ_{iJ} is defined to be zero. Sample means can be used for the w_i and the n_j/n ratios. The calculated π-ratios can then be used to test the null hypothesis of even treatment for boys and girls of a certain age class:

$$H_0 : \pi_{ij} = \pi_{ik} \tag{5}$$

for all i adult goods, where j refers to boys and k refers to girls.[4] In addition to this test across demographic categories, the implication of equality of π-ratios across adult goods and within (child) demographic categories can also be tested:

$$H_0 : \pi_{ir} = \pi_{jr} \tag{6}$$

for all i and j that refer to adult goods and for all r demographic categories. General procedures for deriving the test statistics and the standard errors of the π-ratios are described by Deaton, Ruiz-Castillo, and Thomas [*1989*].

IV. DATA AND ESTIMATION ISSUES

Data used in this paper come from the Papua New Guinea Household Survey (PNGHS), the first nationwide household consumption survey ever conducted in PNG. The survey design and enumeration, which were carried out by the authors in 1995 and 1996, covered a random sample of 1,200 households, residing in 120 rural and urban communities ('clusters'), who were interviewed between January and December 1996. The survey team selected clusters from the enumeration areas of the 1990 census, stratifying the sample by sector (urban and rural), by environmental conditions (elevation and rainfall), and by the level of agricultural development.[5] The stratification allows us to create sub-samples that correspond to the agro-ecological zones used by Haddad and Reardon [*1993*]. The division of the sample into urban, and rural zones with high and low income-earning potential in agriculture facilitates the study's ability to answer hypotheses regarding income effects. We also generate a set of household weights generated on the basis of variation between the 1990 census estimates of the size of each cluster and the actual size found in 1996, and on the deviation of the actual number of households surveyed in each cluster from the target number. All results presented below take account of the clustered, weighted and stratified nature of the sample.

Enumerators interviewed each household twice, with the start of the two-week consumption recall period signalled by the first interview. Expenditure data were collected on all food (36 categories) and other frequent expenses (20 categories) during the recall period. The expenditure estimates include the imputed value of own-production,[6] net gifts received, and stock changes, so they should be a good measure of consumption during the recall period. An annual recall covered 31 categories of infrequent expenses. An inventory of durable assets was used to estimate the value of the flow of services from these assets, including rental services from owner-occupied dwellings. We also broke out six categories designed especially to capture adult goods. Enumerators were trained to be sure to explain to respondents the scope of the goods included in each commodity category, especially those for adult goods.

The data required for our analysis, the estimation of equations (1) and (3), are available for 1,144 households. The data set includes six commodities that may plausibly be candidates for adult goods: adult clothing, alcohol, betelnut (a mild narcotic similar to *pan*), gambling and lotteries, meals eaten away from home, and tobacco and cigarettes. All except for betelnut and gambling have been included as adult goods by Deaton [*1989*]. Eight demographic groups were created: the number of males and females in each of four age groups, 0–6 years, 7–14 years, 15–50 years, and over 50 years. To control for other household traits that could affect household expenditures, our specification includes a dummy variable for whether the household head was a primary school graduate (six years of schooling), three dummies for the main income source of the head (formal business or wages, tree crop agriculture, food crops and livestock), and four regional dummies.

Table 1 contains budget shares for the candidate adult goods along with means for the other Engel curve regressors, all disaggregated by agroecological zone. As discussed above, the total budget share of the candidate adult goods is rather constant across the zones, at approximately 12 per cent. The largest shares for the candidate goods are betelnut and tobacco products, while alcohol shows the greatest variation in budget shares across the zones. Comparing the urban and two rural zones demonstrates a pattern of per capita expenditures, education of the household head, and involvement in the formal sector that are consistent with those expected with changing income levels.

Consistent with the aggregate evidence of females having a smaller share of the PNG population [*UN, 1996*], the household data show that at each age group, females have a smaller demographic share than males. The gap between male and female demographic shares is especially apparent for children of age 7–14 years in all regions, and in the rural zone with low income potential. The uneven sex ratios are least apparent in the urban sector, and in the youngest age group (0–6 years).

TABLE 1
DESCRIPTION OF THE DATA (MEANS)

Variable	Papua New Guinea	Urban Zone	Rural High Zone	Rural Low Zone
Total candidate adult goods budget share	0.123	0.115	0.124	0.126
Adult clothing budget share	0.015	0.013	0.016	0.013
Alcohol budget share	0.017	0.032	0.017	0.007
Betelnut budget share	0.034	0.021	0.037	0.037
Gambling and lotteries budget share	0.011	0.010	0.010	0.016
Meals eaten out of home budget share	0.021	0.025	0.020	0.021
Tobacco and cigarettes budget share	0.025	0.015	0.025	0.033
Log per capita total expenditure	7.105	7.911	7.051	6.782
Log of household size	1.620	1.716	1.632	1.531
Share of males, $0 \leqslant 6$ years, in household	0.105	0.099	0.102	0.116
Share of females, $0 \leqslant 6$ years, in household	0.092	0.105	0.091	0.089
Share of males, 7–14 years, in household	0.103	0.092	0.105	0.103
Share of females, 7–14 years, in household	0.087	0.086	0.088	0.083
Share of males, 15–50 years, in household	0.270	0.323	0.261	0.264
Share of females, 15–50 years, in household	0.256	0.248	0.255	0.264
Share of males, > 50 years, in household	0.051	0.031	0.055	0.052
Share of females, > 50 years, in household	0.036	0.015	0.042	0.029
1 = head completed primary school, 0 = other	0.412	0.762	0.398	0.245
1 = main income of head from formal sector	0.262	0.757	0.197	0.150
1 = main income of head from tree crops	0.387	0.086	0.457	0.368
1 = main income of head from food crops	0.172	0.020	0.195	0.197
Number of Observations	1144	314	645	185
(percentage of national household numbers)	(100)	(13)	(64)	(23)

Note: Means are calculated using household sampling weights.

One possible econometric complication for the estimation of equation (1) arises when there are only a few adult goods; a regression for one adult good on the total would be rather like regressing something on itself [*Deaton, 1989*]. Observations with a large unexplained residual expenditure on a particular adult good could have a large total adult goods expenditure, with bias arising from the correlation between the equation error and the explanatory variable. To guard against this, we could adopt an instrumental variables approach, using total expenditure, x, as an instrument for x_G. Of course, if the bias is not statistically important, there would be a loss of efficiency when using instrumental variables, affecting the test of whether good i is a genuine adult good. To resolve the trade-off between potential bias and loss of efficiency, we use either OLS or IV to estimate equation (1), depending on the outcome of a Durbin–Wu–Hausman test.

There is also a complication in testing the hypothesis that the π-ratios for a particular type of child are equal across adult goods (equation 6) when the data come from a clustered survey, as they do here. Because the π-ratios are non-linear transformations of the OLS parameter vector $\mathbf{b}_i$, Deaton, Ruiz-Castillo and Thomas [*1989*] use the delta method to calculate the variances both within and across equations, using:

$$\{\mathbf{V}(\hat{\pi}_r)\}_{ij} = \mathbf{J}'_{ir}(\mathbf{X}'\mathbf{X})^{-1}\mathbf{J}_{ir}\sigma_{ij}$$

where $\mathbf{J}_{ir}$ is the $1 \times k$ Jacobian matrix of the transformation from the $\mathbf{b}_i$s into the scalar π_{ir}. $\mathbf{X}$ is the matrix of explanatory variables which are common in each adult goods equation, and σ_{ij} is the residual covariance between the ith and jth equations, estimated from:

$$\hat{\sigma}_{ij} = (n-k)^{-1}\mathbf{e}'_i\mathbf{e}_j$$

where $\mathbf{e}_i$ is the vector of residuals from the ith equation and $(n-k)$ is the degrees of freedom in each regression. The problem with clustered surveys is that the observations are not sampled independently, so the covariance matrix of the cross-equation parameters, $E\lfloor(\hat{\mathbf{b}}_i - \mathbf{b}_i)(\hat{\mathbf{b}}_j - \mathbf{b}_j)\rfloor$ is not simply the $\sigma_{ij}\,(\mathbf{X}'\mathbf{X})^{-1}$ that is used in the procedure outlined by Deaton, Ruiz-Castillo and Thomas [*1989*].

To address this feature of our data, we took several steps to modify the existing testing procedures. First, we note that the robust (or 'Huber/White/Sandwich') estimator of the variance is appropriate for single equation estimation with clustered survey data [*StataCorp, 1999*]. Hence, a statistically appropriate covariance matrix for testing the equality of the π-ratios across adult goods can be obtained by 'stacking' the equations for each adult good and using a fully interacted dummy variable model (that is, a dummy variable specified for each type of adult good is interacted with all of the variables in the stacked model).[7] To the extent that this modification of the testing procedure allows us to control for the survey design effects, and in particular the reduced precision that clustered samples bring, we will be less likely to reject null hypotheses (both within and across equations) than will previous studies that have ignored survey design effects. In other words, the failure of previous outlay-equivalent studies to detect gender bias becomes more surprising because most studies appear to have ignored survey design effects, making them more likely to reject the null of no gender bias.

V. BASIC RESULTS

Table 2 contains the results of testing the candidate adult goods, using the linear expenditure model described in equation (1). The instrumental variables estimator was used for all of the expenditure items except for betelnut and gambling, where this choice reflected the outcome of the Durbin–Wu–Hausman tests. The hypothesis that the age and gender of children plays no part in explaining the allocation of expenditures within the adult goods group is accepted for all candidate goods, both individually and jointly. This result is consistent with evidence from other countries, although gambling has typically not been tested as a potential adult good while betelnut is a specialised item for PNG (although the related *pan* has been used as an adult good in South Asia).

The Engel curves for these six adult goods and for the aggregate adult goods budget share explain between 3.3 per cent (tobacco products) and 13 per cent (meals out of the home) of the variation in budget shares. The Engel

TABLE 2
IDENTIFYING ADULT GOODS BY TESTING FOR EXCLUSION OF CHILD DEMOGRAPHICS

	t-test for excluding each age–sex group				
Adult good	Male 0–6	Male 7–14	Female 0–6	Female 7–14	Joint test for excluding child demographics
Adult clothing	1.29	1.60	0.18	0.68	1.01
Alcohol	0.78	0.20	0.71	1.56	1.11
Betelnut	1.80	0.12	0.65	1.22	1.48
Gambling	0.27	1.17	0.96	1.42	0.93
Meals out	1.23	1.34	0.26	0.60	0.65
Tobacco	0.90	0.35	0.90	0.21	0.40

Note: Test statistics are corrected for clustering, sampling weights and stratification. The *t*-test is distributed t_d where d is the number of clusters minus the number of strata (105) and the joint test is an adjusted Wald test, distributed as $F_{(k,\ d-k+1)}$ where k is the number of restrictions (4). The critical values at the 5 per cent level for t_d are 1.98 and for $F_{(k,\ d-k+1)}$ are 2.46.

curve for the aggregate adult goods group explains 7.8 per cent of the variation in budget shares. In all cases, F-tests indicate the overall significance of the variables in the regression, even when adopting the conservative approach of using degrees of freedom based on the number of clusters rather than the number of households.[8] To save space, these Engel curves are not presented because interest is not so much in the individual coefficients but in the combinations of the coefficients given by equation (4).

Table 3 contains the estimated outlay-equivalent ratios for each of the six adult goods and for the aggregate good formed from the sum of expenditures on the six individual goods. To interpret these ratios, note that the π-ratio of -0.49 for the effect of young boys on adult clothing means that the addition of a boy of age 0–6 years to the household has the same effect on adult clothing expenditure as would a 49 per cent reduction in total outlay per household member. The first four rows of Table 3 give the π-ratios for children. These should be negative, which they are in 23 of 28 cases. The

TABLE 3
OUTLAY EQUIVALENT RATIOS FOR ADULT GOODS

Gender and age	Adult clothing	Alcohol	Betelnut	Gambling	Meals out	Tobacco	All adult goods
Children				π-ratios			
Male 0–6	−0.49	−0.33	−0.46	0.79	−0.64	−0.13	−0.26
Male 7–14	−0.46	−1.06	−0.27	−1.80	−1.51	−0.52	−0.90
Female 0–6	−0.70	−0.50	−0.38	−0.11	0.20	−0.59	−0.34
Female 7–14	0.06	−0.43	0.76	−0.68	−0.74	0.68	−0.02
Adults							
Male 15–50	0.87	0.21	−0.50	0.20	−0.15	0.27	0.03
Male > 50	−0.83	−0.63	0.30	−0.07	1.71	0.38	0.17
Female 15–50	0.20	−0.53	0.07	−0.16	−1.28	0.43	−0.27
Female > 50	1.84	−0.86	−0.61	−0.82	−2.22	−1.20	−0.87
Children				Standard errors			
Male 0–6	0.53	0.37	0.29	0.74	0.38	0.54	0.18
Male 7–14	0.43	0.29	0.33	0.48	0.52	0.47	0.17
Female 0–6	0.51	0.34	0.49	0.48	0.59	0.83	0.20
Female 7–14	0.58	0.45	0.40	0.62	0.51	0.68	0.17
Adults							
Male 15–50	0.55	0.37	0.25	0.33	0.32	0.41	0.16
Male > 50	0.68	0.28	0.51	0.65	1.56	0.83	0.33
Female 15–50	0.47	0.35	0.42	0.45	0.56	0.66	0.20
Female > 50	1.08	0.28	0.60	0.39	1.20	0.52	0.25
	p-values for test of equal π-ratios for males and females						
Infants	0.69	0.70	0.86	0.31	0.29	0.61	0.74
Adolescents	0.56	0.30	0.08	0.02	0.10	0.17	0.00
Prime adults	0.43	0.18	0.33	0.55	0.05	0.86	0.28
Old adults	0.06	0.51	0.31	0.31	0.13	0.17	0.04

Note: Standard errors and p-values are corrected for clustering, sampling weights and stratification.

only anomalies are for the effect of young boys on gambling expenditures, the effect of young girls on meals consumed out of the home, and the effect of older girls on the consumption of adult clothing, betelnut, and tobacco products. In all five cases of positive π-ratios for children, the point estimates are surrounded by wide standard errors and the π-ratios are not statistically different from zero.

Although these anomalies occur with a similar low frequency to that in the original studies by Deaton [*1989*], they raise the question of whether all of the items included in Table 3 are genuine adult goods.[9] The results of the alternative test (that is, the test that the π-ratios are equal across adult goods categories within age and gender groups – or, along the rows in Table 3) are reported in the top panel of Table 4 (rows 1 to 4). While the null hypothesis is never rejected at conventional levels, lending support to this choice of goods, there is a potentially worrying aspect of the test results.[10] For 7–14 year-old boys, the test is on the borderline of rejection, with $p = 0.06$.[11]

On closer inspection of our results, however, there appears to be a way to allay this concern. The good that appears to contribute most to the inequality in π-ratios for 7–14 year old children is betelnut. It is possible that some older children in this group actually do consume some betelnut. In PNG, there is no legal age limit on its consumption, and in some communities there are no real norms against its use by children. In contrast, even in PNG, there are minimum age requirements and community norms in most areas against alcohol and tobacco use by children. The results in the bottom panel of Table 4 (rows 5 to 8) show that the equality of the π-ratios is more difficult to reject

TABLE 4
ADJUSTED WALD TESTS FOR EQUALITY OF π-RATIOS ACROSS ADULT GOODS

	Children			Adults	
Gender and age	Test	*p*-value		Test	*p*-value
Males 0–6	0.66	0.65	Males 15–50	1.54	0.18
Males 7–14	2.12	0.06	Males > 50	1.18	0.32
Females 0–6	0.36	0.88	Females 15–50	1.31	0.26
Females 7–14	1.78	0.15	Females > 50	1.63	
		Excluding Betelnut			
Males 0–6	0.79	0.53	Males 15–50	0.67	0.61
Males 7–14	1.60	0.17	Males > 50	1.03	0.39
Females 0–6	0.45	0.77	Females 15–50	1.48	0.21
Females 7–14	0.90	0.47	Females > 50	2.00	0.09

Note: The test statistics are adjusted Wald (W) tests, $(d\text{-}k+1)W/kd$ where d is the number of clusters minus the number of strata in the stacked model and k is the number of restrictions being tested. The tests are approximately F distributed with k and $d\text{-}k+1$ degrees of freedom under the null hypothesis that the π-ratios are the same across all candidate adult goods (with betelnut included $F_{(5,\ 630)}$ and with betelnut excluded $F_{(4,\ 525)}$).

once betelnut is removed from the list of adult goods. Hence, to ensure the robustness of our results, we present the tests for boy–girl discrimination both with and without betelnut included in the list of adult goods.

If boys are favoured over girls, the π-ratio for a given age category should be a bigger negative number for boys than for girls. It is apparent by looking down the columns of Table 3 (comparing rows 2 and 4) that for the older age group (7–14 year olds) this pattern holds for all six individual adult goods, as well as for the aggregate group. The point estimate for the aggregate adult goods group suggests that the addition of an older boy to the household reduces adult goods consumption by as much as would a 90 per cent reduction in total outlay per member. In contrast, the addition of a similarly aged girl to the household is equivalent to only a two per cent reduction in per capita outlay. The difference in these π-ratios for the aggregate good is statistically significant ($p < 0.01$), although differences for the individual adult goods are less precisely measured (row 18, columns 1 to 6).[12] If betelnut is excluded from the list of adult goods, the same conclusion of bias in favour of boys is reached: the π-ratio for the aggregate adult goods group is -1.11 for 7–14 year-old boys (standard error of 0.21) and -0.28 for similarly aged girls (standard error of 0.27) and the hypothesis that these two are the same is still rejected, at the $p < 0.02$ level.

In contrast to older children, for the 0–6 year-olds, only two individual adult goods (betelnut and meals out of the home) generate the pattern of a more negative π-ratio for boys than for girls. Moreover, inspection of the standard errors shows that the differences in π-ratios for boys and girls in the 0–6 years age group are all statistically insignificant. Hence, if there is any gender bias amongst the young children, it does not appear to be detectable using the adult good method.

The π-ratios in Table 3 also suggest that differences in the effects that adult men and women have on adult goods consumption are usually statistically insignificant.[13] Even if they were not, conclusions about gender bias for adults cannot be made on the basis of these results because they may just reflect gender differences in preferences (this explanation can be ruled out for children, who should exert only income effects on adult goods demand).

Sensitivity Analysis

How robust are these results to the definition of the children's age groups? Since it is possible that children in PNG by the age of 14 are already in the work force and bringing income into the household (which would violate one of the assumptions of the Deaton method), in this section we check whether the results are sensitive to the age groupings that we have chosen. To do so, the π-ratios were recalculated, using the Engel curve for the aggregate adult good, and (alternatively) age groups of 7–13, 7–12, 7–11, and 6–11 years for the older children (Table 5). These changes also required the starting age for the prime-age adult group to change, and the ending age for the younger child group to fall to five years in one case. While the changed age groups result in some slight changes in the π-ratios, the main findings from Tables 3 and 4 remain unchanged: the π-ratio for older boys is statistically significantly lower than it is for older girls, while there are no statistical differences in the π-ratios for boys and girls in the younger age group. The apparent robustness of the bias against girls in the older age group is important because, while it is plausible that 14-year old girls might either consume the adult good or contribute income to the household (both of which could make the π-ratio positive) this is much less plausible for 11-year old girls.[14]

VI. DISAGGREGATED RESULTS

The basic results suggest a bias in intra-household allocations in favour of older boys, which is consistent with a variety of other social science research in PNG [*Overfield, 1995*]. However, where the adult goods method can potentially provide new information is in studying how the degree of bias differs for different types of households. Identifying such differences, if they exist, may help to untangle the causes of the gender bias. Therefore, in this section of the paper, disaggregated evidence on the π-ratios is presented, following the example set by Haddad and Reardon [*1993*].

Disaggregating by Agro-ecological Zone

The pattern of gender discrimination that is found in the full sample is also apparent in the rural zones of PNG, when using the outlay equivalent ratios for the adult goods aggregate. In contrast, a different pattern is found for urban households (Table 6, columns 1 to 3). The results suggest that parents

TABLE 5
VARIATION IN OUTLAY EQUIVALENT RATIOS FOR THE AGGREGATE ADULT GROUP WITH DIFFERENT CHILD AGE GROUPS

Age groups	Infants: adolescents:	0–6 7–14	0–6 7–13	0–6 7–12	0–6 7–11	0–5 6–11
				π-ratios		
Male infants		– 0.26	– 0.33	– 0.34	– 0.34	– 0.31
Male adolescents		– 0.90	– 0.82	– 0.77	– 0.81	– 0.78
Female infants		– 0.34	– 0.40	– 0.41	– 0.41	– 0.43
Female adolescents		– 0.02	0.04	– 0.11	– 0.02	– 0.08
		p-values for test of equal π-ratios for males and females				
Infants		0.74	0.77	0.77	0.79	0.64
Adolescents		0.00	0.01	0.04	0.04	0.05

Note: The *p*-values are corrected for clustering, sampling weights and stratification.

in the rural sector favour boys over girls in the 7–14 year old age category in the high-income potential areas ($p < 0.01$), and also – although less precisely measured – in the low-income potential areas ($p < 0.06$ – row 10). In contrast, we cannot reject the hypothesis that the infant girl and infant boy π-ratios are the same in either of the rural zones (row 9). In urban areas, however, there is no sign of gender discrimination when examining either infant or adolescent children (column 1, rows 9 and 10).

This pattern of results across different areas appears to provide only partial support to the hypothesis of Haddad and Reardon [*1993*] that discrimination against girls will be less detectable as the economic opportunities for women increase. While the lack of apparent bias in urban areas may be consistent with there being greater economic opportunities in urban areas, it is puzzling that gender bias does not appear to vary across the two rural zones. Most observers would expect women's labour in agricultural production to be more valuable in the rural zone with high agricultural income potential and therefore the costs to households of discriminating against girls to be much greater. Moreover, the hypothesis that the π-ratios are equal across the urban–rural and agro-ecological zones is rejected (at the $p = 0.03$ level) only for *boys* (Table 5, column 4). Further tests show that this significant result is driven by the difference between the urban sector and each of the two rural zones, the

TABLE 6
OUTLAY EQUIVALENT RATIOS FOR THE AGGREGATE ADULT GROUP BY AGRO-ECOLOGICAL ZONE

Gender and age	Urban zone	Rural zones: High income potential	Rural zones: Low income potential	Test for equality across zones (p-value)
Children		π-ratios		
Male 0–6	– 0.70	– 0.20	– 0.54	0.66 (0.52)
Male 7–14	– 0.00	– 0.84	– 1.24	3.61 (0.03)
Female 0–6	– 0.47	– 0.27	– 0.57	0.21 (0.81)
Female 7–14	– 0.71	0.22	– 0.20	1.49 (0.23)
Adults				
Male 15–50	0.77	0.04	– 0.74	4.46 (0.01)
Male > 50	0.37	0.38	– 0.63	0.85 (0.43)
Female 15–50	– 0.43	– 0.46	0.29	1.30 (0.28)
Female > 50	– 0.80	– 0.84	– 1.32	0.27 (0.76)
	p-values for test of equal π-ratios for males and females			
Infants	0.70	0.81	0.96	
Adolescents	0.22	0.00	0.06	
Prime Adults	0.00	0.09	0.18	
Old Adults	0.34	0.02	0.49	

Note: The test for equality of the π-ratios across zones is an adjusted Wald test, distributed as $F_{(2,105)}$ under the null hypothesis. The p-values for the hypothesis tests are corrected for sample design effects.

comparison between the high- and low-income rural zones brings an insignificant result ($p = 0.24$).[15] Thus, while parents in both high- and low-income rural zones of PNG significantly reduce their consumption of adult goods when they have an additional male adolescent, urban parents do not seem to make similar adjustments in their consumption allocations. Because bias in favour of boys seems to be restricted to rural areas, which are presumably more tradition-bound, it might be construed from these findings that in PNG gender bias arises more from cultural rather than strictly economic factors.

Other Disaggregations

To further explore the hypothesis that cultural factors affect the degree of bias, two other disaggregations of the π-ratios are presented: by the age group of the household head and according to the dominant descent rule. The lifecycle position of the household, as reflected by the age of the head, may proxy for the strength of traditional attitudes, although age may also reflect the severity of credit constraints and the higher shadow value of child agricultural labour (because of the smaller size of households headed by young people). Regardless of the reasoning, when the sample is divided into 'Young' (40 years and under) and 'Old' (above 40) sub-samples based on the age of the household head, no significant differences in the π-ratios occur between the Young and Old groups (Table 7, column 3). Within each household head age group, the π-ratios for boys are more negative than they are for similarly aged girls, although the difference is not statistically significant for the 0–6 year old children (Table 7, row 5). Amongst the older children, the gender gap in the π-ratios is largest for households with an Old head, although the statistical significance of the difference is slightly higher for households with a Young head (Table 7, row 6). Thus, there are no clear differences in gender bias according to the age group of the household head.

Variation in descent rules may be an important determinant of cultural perceptions about gender and the role of women, which in turn may affect the degree of gender bias recorded by the outlay equivalent method. Within PNG there are a number of matrilineal societies, in which children are part of their mother's kin group and where only female children can pass kin identity on to their offspring.[16] In contrast to patrilineal and mixed descent systems, women in matrilineal systems may have greater status and more autonomy, so bias against girls may be less apparent.[17] To test this conjecture, the sample is divided into 'Matrilineal' and 'Other' sub-samples, based on the predominant descent system in the area where each cluster of surveyed households was located. Whereas the π-ratio for older (7–14 years) girls was positive in the Other sub-sample, it was negative in the Matrilineal sub-sample (Table 7, row 4, column 5), suggesting less bias against girls when

TABLE 7
OUTLAY EQUIVALENT RATIOS FOR THE AGGREGATE ADULT GOODS GROUP BY AGE GROUP OF THE HOUSEHOLD HEAD AND DESCENT SYSTEM

	Age of household head			Descent system		
	π-ratios		*p*-value for equal π	π-ratios		*p*-value for equal π
	Young	Old		Matrilineal	Other	
Male 0–6	− 0.18	− 0.74	0.46	− 0.15	− 0.27	0.78
Male 7–14	− 0.81	− 0.73	0.52	− 0.94	− 1.00	0.92
Female 0–6	− 0.11	− 0.52	0.37	− 0.02	− 0.42	0.34
Female 7–14	− 0.03	0.13	0.16	− 0.18	0.25	0.37
p-values for test of equal π-ratios for males and females						
Infants	0.77	0.58		0.82	0.61	
Adolescents	0.04	0.07		0.31	0.01	

Note: The *p*-values are corrected for clustering, sampling weights and stratification.

descent goes through the female line (although the difference is not statistically significant). Moreover, within the Matrilineal sub-sample, the hypothesis of the π-ratio for older girls being equal to the ratio for older boys is not rejected ($p=0.31$), whereas it is strongly rejected in the Other sub-sample ($p=0.01$).

The importance of different types of descent rules for the status of women and for bias against girls may also partially explain why gender bias was not found in urban areas. The main asset transferred to descendents in PNG is access rights to customary land. In urban areas these rights are likely to be less valuable (because one has to return to the rural area to exercise them), so even for urban households that originally came from patrilineal societies, there may be less reason to favour boys.

VII. CONCLUSIONS

In our paper we have tested for discrimination in the allocation of goods between boys and girls within households in Papua New Guinea using Deaton's [*1989*] outlay–equivalent ratio method. Adding an adolescent boy (7–14 years old) to the household reduces expenditure on adult goods by as much as would a nine-tenths reduction in total outlay per member, but girls have no effect on adult goods expenditure. The hypothesis of Haddad and Reardon [*1993*] that gender bias is inversely related to the importance of female labour in agricultural production is not supported. There is no evidence of bias against girls in the urban sector, nor of bias against younger girls in the 0–6 year old age group. These results are robust to the definition of the children's age groups used in the analysis and also do not vary by the age group of the household head.

Perhaps the most surprising aspect of our research is that it worked so well. In the past, adult goods-based approaches have not yielded results consistent with the findings of other social scientists and economic studies based on alternative approaches. Consequently, there has been some question about whether expenditure-based methods can be used as a reliable tool for investigating the nature and extent of gender discrimination [*Deaton, 1997*]. Although our results cannot explain why the adult goods method has failed to work in other countries, the successful application in this setting at least provides some grounds for believing that the method can be a useful tool. Of course, it may simply be the combination of factors in PNG that make bias greater and easier to pick up although we also believe that the higher quality of our data, which were collected with this purpose in mind, may have helped provide more precise measures of the adult goods consumption behaviour of poor households.

If our results have identified the existence of a deep-seated cultural–social bias against girls, educational and other development policies countering gender discrimination could lead to both gains to social welfare and more rapid, equitable growth. For example, interventions through PNG's widespread (though floundering) health system should target young girls. A one dollar expenditure on females should have a greater effect on welfare than if it were spent by the household on the boy. Moreover, this bias may mean that the incidence of poverty is understated because average household data will miss the fact that females in some household may have access to sufficiently fewer resources such that their individual consumption level falls below the poverty line. However, further work is needed in PNG to explain why bias in favour of boys is restricted to the rural areas and is also restricted to children in the older age groups. One plausible hypothesis, which we have not tested, is that the needs of younger children – primarily food and care – are met using resources primarily controlled by women. In contrast, the needs of older children – particularly school fees and clothing – require cash, which is controlled primarily by men who may favour boys over girls.

NOTES

1. There appears to be some ambiguity in the statistics, because the 1996 *Demographic and Health Survey* reports an under-five mortality rate of 108.2 for males and 91.3 for females, which is the opposite to the pattern reported in the WHO statistics [*NSO, 1997*].
2. In fact, according to the UNDP's gender-related human development index, the position of women in PNG while not good, is better than a number of other countries, such as Burkina Faso, Bangladesh and Côte d'Ivoire. While this seems to contradict the arguments of the social scientists that claim there is a lot of discrimination, it may be because the index weights some of the opportunities that PNG women face in a way that makes the measure look lower relative to other countries. It could also be that the index is not a very good

measure of human development. Case and Deaton [*2002*], for example, note that the UN's gender-related index misses a number of the deprivations that are associated with poverty.

3. In other words, adult goods consumption in the present may be affected both by current income and by expected lifetime income. If having girls in the family leads to an increase in expected lifetime income due to, say, bride price receipts, it may lead to a positive impact on adult goods consumption which is independent of any impacts due to discrimination. The outlay-equivalent literature does not consider the fact that the demographic structure of the household may affect expected lifetime income, with the result that the demographic variables may pick up omitted effects in addition to any discrimination that is present.
4. Unobserved household-level heterogeneity that affects consumption patterns should not affect the inferred level of bias, unless such heterogeneity affects the demographic coefficients for boys differently than for girls.
5. This classification was established from an agricultural mapping project [*Allen, Bourke and Hide, 1995*].
6. The monetary values for self-produced foods were the values used by respondents. Estimates of average expenditure are unchanged if these respondent-reported unit values are replaced by either cluster medians of the unit values or cluster averages of market prices [*Gibson and Rozelle, 1998*].
7. We are grateful to Roberto Gutierrez of Stata Corporation for suggesting this procedure. Because repeating the same observations induces a correlation in the disturbances of the stacked model, the clusters also have to be redefined as the interaction of the original clusters with the dummy variable for each replication of the dataset.
8. Specifically, we used an adjusted Wald (W) test for zero slopes: $(d-k+1/kd)$ $W \sim F(k, d-k+1)$, where d is the number of clusters minus the number of strata (105), and k the number of slope variables [*StataCorp, 1999*].
9. Three of 32 π-ratios were positive in the results for Côte d'Ivoire and 8 of 48 in the results for Thailand, which included some items such as 'men's and boy's clothing' which were known to include child goods.
10. Actually, there may be a second, although in our opinion, more minor concern. In contrast to the results in Deaton [*1989*], the equality of the π-ratios across goods for each of the *adult* demographic groups is also not rejected. While this equality is less informative for adults than it is for children, the failure to reject may indicate low test power, although once again, if the survey design effects are ignored the equality across the goods is rejected at the 0.02 level, for at least two of the adult groups.
11. If the survey design effects are ignored, as they seem to have been in some previous studies, the p-value for 7–14 year old girls would also fall into the rejection zone, at $p = 0.02$ (from 0.12 with survey design effects included).
12. The importance of controlling for survey design effects can be seen by noting that if the test of equal π-ratios for the effect of older boys and older girls on betelnut (alcohol) consumption is carried out with the clustering and stratification ignored, the p-values fall from 0.08 to 0.02 (0.30 to 0.13).
13. The p-value for the difference between π-ratios for prime age males and females in the aggregate adult goods group drops from 0.28 to 0.06 if betelnut is excluded, probably because adult women are equal participants in consumption of betelnut in contrast to alcohol and tobacco which are predominantly consumed by men.
14. In fact, child labour for cash incomes is quite rare in Papua New Guinea. Children between the ages of 11 and 18 may work in family-related enterprises if they have parental permission, a medical clearance and a work permit from a labour office. But according to the 2001 Report of Human Rights Practices, from the US Department of State, such employment is uncommon in PNG.
15. This equality across the two rural zones appears to be robust to the definitions of high and low income areas. If the lowest of the clusters in the high agricultural income potential zone are reallocated to the low income zone, and the outlay equivalent ratios re-estimated, the hypothesis of equality of π-ratios across the two rural zones is still not rejected ($p < 0.45$).

16. These include the Trobriands, the Tolai, the Kavieng, and various groups in Milne Bay and Bougainville. Within the household survey, these areas are identified as coming from the Esa'ala, Kavieng, Kokopo, Losuia and Rabaul districts.
17. However, Quisumbing and Otsuka [*2001*] report that in matrilineal areas of Sumatra, daughters are still disadvantaged in terms of schooling.

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