Gender Bias in Orphanhood: Ethiopian Adults Consumption Evidence

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Abstract

This paper examines the evidence for parental gender discrimination against (mostly HIV) orphans in terms of gender from the application of a model of intrahousehold inequality based on goods consumed solely by adult, for instance tobacco. The model predicts that the expenditure share on adult goods is likely to be higher with a ceteris paribus increase in the proportion of girls. Applied to a 2004-2005 Ethiopian household expenditure survey, the results indicate that a. the bias against girls is the dominant effect across different categories of biological and orphan children; b. parents treat girls differently from boys with similar genetic relationship to them, contrary to the prediction of the hypothesis of genetic relatedness. The results suggest that the common findings in the literature that there is no gender dimension to parental treatment of orphan children in Africa have missed the gender bias evidence because they are based on solely on welfare outcomes in education and health; not explicitly obtained from a model of intahousehold inequality.

Keywords: Child Gender; Adult Goods; Intrahousehold Inequality; HIV-orphans; Africa; Ethiopia.

JEL Codes: D12; D13; J16; O55.

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1—Introduction

The HIV/AIDS epidemic has had a devastating impact in Africa, and Sub Saharan Africa (SSA) is its epicenter. Its prevalence rates in the region are the world's highest: above 10% in 16 countries (World Bank 2002). Ethiopia, the country whose orphan problems we examine in this paper from with reference to child gender, has an HIV/AIDS prevalence rate for women aged 15-24 that ranks eighth among the most affected African countries. Ethiopia is also a country that has experienced famine, and war; historically poverty and hunger have been important causes of orphanhood. The age structure of adult mortality can provide some idea about the relative impact of HIV infection on orphanhood as compared to its other causes. Since most adult HIV/AIDS mortality is accounted for by prime-age adult death, and middle-aged adults tend to have older children, many studies found that orphans tend on average to be older than non-orphaned children and have poorer welfare outcomes, for example, a greater likelihood of lacking basic education, e.g. Case et. al. (2006). The indirect evidence on age of orphans and pattern of parental mortality in this study, presented in table 1 below, suggest the HIV infection has been the dominant cause of the rising tide of orphans in Ethiopia in recent decades. However, anecdotal reports also indicate that gender bias against non-orphan or orphan girls is a pervasive practice in Ethiopia. Erulkar et. al. (2011) found that 64% of the sample children aged 10-19 from in the slums of Addis Ababa were girls, 36% just aged 10-14. The study also found 79% of surveyed girls had never been to school, and 24% were employed as domestic workers. Koohi-Kamali (2008) employed the same Ethiopian survey used in the current study, see section 4, and reports the ratio of the ratio of primary school enrollment of girls to boys as 0.806, similar to Pakistan (Aslam and Kingdon 2008, table 2); the population sex ratio of female to male as 0.895, considerably outside the 0.93-0.96 range of naturally occurring ratio (Parazzini et al. 1998). The latter is taken as a broad indicator of female health and mortality, Sen (1990). Despite the evidence of gender disparities in welfare outcomes, there is no matching evidence of discrimination in household internal resource allocation that can explain such outcomes across non-orphan and orphan children.

The present paper examines this puzzle by addressing to two features common to the studies on African orphans. First, although the differences in outcomes are attributed to parental preferences regarding treatment of non-orphaned/orphaned children, that interpretation is not based on any explicit model of intrahousehold inequality, see next section. We shall employ the Rothbarth (1943) model, which attempts to isolate the effects of children on adults' consumption by assuming that a change in the number of children or their gender leaves adults' consumption preference unchanged; that is, such demographic changes have income effects but no substitution effects. The model then measures those effects by changes in the level of household expenditures on goods consumed solely by adults as a proxy for the exclusively adult component of household budget. However, we argue that some modifications to the specification of that model is necessary and briefly examine the factors that have been disregarded in many previous applications of the model, see Koohi-Kamali (2008) for more extensive discussion; that can account for the failure to obtain evidence of child gender bias (see Deaton 1997). The commonly employed specification based on the age of children lacks empirical support and is not justified by any hypothesis. We propose a child gender specification based on the number of children that draws theoretical plausibility from fertility behavior of parents: parents with a pro-boy preference who fail to have male children in early cycles are hence more likely to have a larger number of children; for example (see Clark 2000) for evidence. Furthermore, the sample employed should be consistent with the Rothbarth model presupposition of a nuclear household context (see Deaton & Muellbauer 1986). A typical household survey from a developing country contains a high proportion of nonnuclear households, whereas the model does not propose to predict consumption behavior of such households.

A second feature common to the reported test results on African orphans is the finding that while there are differences in treatment received by orphaned and fostered children compared to non-orphans, those differences are not based on gender[1]. We provide evidence of differences in treatment by different categories of orphan children by adding child gender (proportion of girls) as a variable to the Rothbarth model of intrahousehold inequality. We demonstrate that such an explicit model of internal inequality can uncover evidence of gender discrimination against orphans or fostered children that may well have been missed in the approaches not based on a hypothesis of intrafamily resource allocation.

We employ the Household Consumption and Income Expenditure Survey (HICES) for 2004-2005 conducted by the Ethiopian Statistical Authority to test the predictions of the model regarding orphanhood. Contrary to the common findings that there is no gender dimension to orphan child discrimination in Africa, we report evidence for gender effects of orphanhood on adults' consumption patterns by testing a few proposed mechanisms of child gender discrimination examined.

Section 2 discusses the three different hypotheses employed in the literature on African orphans. In section 3, we discuss the models employed in this paper and present its specifications; section 4 examines the data. Section 5 presents the main findings of this research on testing each of the three hypotheses on orphanhood by the Rothmarth model, indicating that not only does the model effectively detect different treatment of children in terms of their relationship to household head/spouse, but also, more importantly, it can uncover evidence of discrimination by child gender missed in the earlier studies. A final section contains some concluding remarks.

2— Literature

The existing literature on orphanhood in SSA deals mainly with disparities in welfare outcomes with respect to children's health and especially their education. Three approaches can be discerned in these studies. The most common is to conduct an analysis of parental treatment of orphans by comparing it with treatment of biological with non-orphan children (Ainsworth *et. al.* 2005; Evans & Miguel 2007). The second approach is guided by the hypothesis that orphans suffer greater loss of welfare from losing a mother than from losing a father (see for example Evans & Miguel 2007).[2] Case *et al.* (2004) and Case and Ardington (2006) test this hypothesis for differences in parental investment in children's human capital (years of schooling) and found evidence for greater loss of child welfare when orphanhood is due to the death of a mother rather than a father; in addition to education, Beegle *et al.* (2007, 2009) provide extensive evidence for differences in investment in child health (by height of children), and come to a similar conclusion. The third approach is based on the Hamiltonian hypothesis from evolutionary biology (Hamilton 1964a, b), according to which altruistic behavior between any two individuals is an increasing function of the degree of genetic relatedness between them; thus, for instance, one's own children are treated more favorably than one's grandchildren, who in turn receive better treatment than more distant relatives, the latter being better treated than non-relatives, and so forth. Note that this hypothesis is largely silent on the gender dimension of genetic relatedness; the predictive implication of the hypothesis is that differences in child treatment occur across different child categories, not *within* each category, for example between orphans and non-orphans, rather than between orphan boys and orphan girls.[3] This hypothesis was first tested by Case *et al.* (2000) in terms of differences in welfare outcomes among biological, orphaned and fostered[4] children and more supporting evidence was obtained for it by Case *et al.* (2004) and by Case and Ardington (2006). The absence of gender effects on consumption is hence a notable feature of the literature; the exception being by the category of educational expenditure as “child goods”. Finding of gender bias by spending on child education is suggestive; for example Gong et. al. (2005), or Kingdon (2005), and Aslam and Kingdon (2008). However, the basis for the interpretation of the evidence by educational expenditure as measuring internal household inequality is not clear. The focus of the model is how the parents allocate resources differently to boys and girls by making changes in their own, as opposed to their household, living standard based on consumption of goods known a prior to exclude consumption by children, see Muellbauer (1987). Since children have no decision-making power in that regard; changes in children’s living standard are not as informative; though such changes can provide supporting evidence in terms of welfare consequences of intrahousehold inequality by education or health. Jensen (2002) proposes an alternative approach that relies mainly on increased fertility to explain parental gender bias in internal educational allocation that leads to some interesting insights; however, on the limitation of explaining gender bias solely by demographic change, see Leung (1991).

3—Model and Specifications

The paradox is that the application of the model, first proposed in Deaton (1987), has produced little of interest despite strong evidence of bias against girls by welfare outcome in many countries over the last two decades. Koohi-Kamali (2008, 2016) discusses the model’s gender specification and sample truncation problems, and demonstrates how the suitable modification of the model, as discussed in section 3, can reveal evidence of discrimination against girls.

However, in extending the model to child gender bias with respect to orphan children, one would have to deal with a new potential source of endogeneity, since taking in an orphan child is a deliberate decision made by the household. A cross-section context is not an ideal one to deal with this question; however, some idea can be obtained from modelling whether a household has an orphan by the estimation of a probit function with an orphan indicator as the dependent variable (see also Gong *et al.* 2005, who take a similar approach, using Chinese cross-section consumption data to model whether a family's decision to have a second child is based on the gender of the first child). We take up this question in section 5.

First, we deal with the truncation issue by confining the samples employed to strictly nuclear households; for example both Gong *et al.* (2005) in tests of child gender bias, and Nelson (1988) in estimates of equivalence scale and cost of children, employ similarly truncated samples. Second, we examine the effect of child gender using an index based on the number of children in the household. However, we control separately for the age of children. To do so, we employ the version of the Working-Leser functional form on which almost all previous studies are based, on the influence of child age on gender bias, see section 5 below. Since there is considerable evidence that expansion of the Working-Leser specification by a quadratic term for logarithm of per capita total expenditure considerably improves the functional form flexibility and fit of the budget share equations (see for example Banks *et al.* 1997), we expand the budget share equations with a quadratic term for the logarithm of per capita expenditure to capture the curvature effect on the estimates.

Provided one takes into account these issues, the Rothbarth model remains rather attractive if the focus of analysis is on child welfare or inequality among children. The model is well understood, its key assumptions have good empirical support, see Deaton and Muellauer (1986), and there is broad consensus on its effectiveness for estimating the cost of a child; Gronau (1988), in a cogent critique of equivalence scales models, regards it as “the only” theoretically consistent model among them. Indeed, the case for its application to child gender bias is even stronger, since it would not involve obtaining compensating income estimates for the presence of children. In this study, we apply the Rothbarth model detection of gender bias between biological and non-biological children.

We adopt three types of specification of the model; each specification enters the budget share equations in two versions: with separate indices for each category of children, once with and once without gender distinction. So

$$w_j = \alpha_0 + \alpha_1 \ln(\frac{x_h}{n_h}) + \alpha_2 [\ln(\frac{x_h}{n_h})]^2 + \eta \ln n_c + \gamma_{orph} \frac{n_{orph}}{n_c} + \sum_{a=1}^{4} \delta_a \frac{n_a}{n_c} + \delta z + \varepsilon_j \qquad (1)$$

where w_j denotes budget share of adult goods j, x_h total expenditure for household h, n_h is household size, n_a / n_c indicates the proportions of children in different age groups a over total number of children n_c, n_{orph} / n_c is the proportion of orphans in the household and constitutes our index of child category in (1), z is the vector of additional variables: regional community-level dummies, variables specific to the analysis of child gender bias, such as dummies for the mother’s age and her educational level, their interactive terms, and so on; ε_j a normally distributed error term. In two respects, (1) is different from the usual semi-log budget share equation employed in earlier applications of the Rothbarth model to child gender bias. With the sample restricted to nuclear households, each observation contains two adults, the head and the spouse, so the proportion of adults does not appear above since it is no longer a variable: variation in household size comes entirely from the change in the total number of children across households, and thus demographic proportions are defined over n_c rather than n_h. Single-parent households are excluded here (see however the evidence in Koohi-Kamali (2008) for female-headed households). Note that equation (1) employs separate demographic proportions by child age and child number (which in earlier studies were combined into the same indices), with child categories defined over the number of children, both for gender-dependent and gender-independent indices. The rationale for doing this is that it allows comparison with gender-distinct estimates that should be defined relative to the number of children[5]. We employ five different age groups of children, two years and under, three to six, seven to nine, and 10 to 12 years, omitting those aged 13 to 15 (maximum age) as the omitted category group, for sensibility to child age, see below. Similarly, there are two categories of proportions of children, for biological children of the household head and spouse, and for orphan children; the one included above is the proportion of orphans, the proportion of children with two biological parents is the reference category. An orphan child is defined as one with at least one dead parent (the child *may* be living with the other parent; see below). Thus, the coefficient estimate on n_{orph}/n_c indicates the amount by which average expenditure share on an adult good is expected to change if a child with biological parents is replaced with an orphan child, keeping age structure of the children constant. More specifically, if adults are less willing to cut their standard of living—as indicated by their consumption of adult goods—when there are more orphans in the household, that is, when n_{orph}/n_c increases, then one would expect this proportion to have a positive effect on the budget share of each adult good, that is the higher n_{orph}/n_c, relative to that for the biological children of the head/spouse, the greater the expenditure on adult good j. Therefore, bias against orphans should result in a *positive* sign for this index. In all the equations estimates below, the following variables were also included, although we have not reported the estimates here in the interest of saving space. The vector of variables in z controls for urban/rural, and for the latter, whether the household resides in a rural area under the safety net system run by the government, and if so, whether in addition the area has a free school feeding program. We also include controls for regions and an extensive list of community-specific fixed effects, for example instances of HIV/AIDS death or sickness reported by survey participants, distances to the nearest primary school, health center, marketplace, sources of water supply, fire wood collection, household-level controls such as parents age and education, and so on; see Appendix on the list of control variables employed. Finally, we note that we employ all the sample observations, both positive and zero that combines the extensive and intensive margins of gender differences in expenditure on each single group of adult goods, as in all previous applications of the model[6].

We write the gender-distinct version of (1) as

$$w_j = \alpha_0 + \alpha_1 \ln\left(\frac{x_h}{n_h}\right) + \alpha_2\left[\ln\left(\frac{x_h}{n_h}\right)\right]^2 + \eta \ln n_c + \frac{1}{n_c}\left(\gamma_{biog}\, n_{biog} + \gamma_{orphg}\, n_{orphg} + \gamma_{orphb}\, n_{orphb}\right)$$
$$+ \sum_{a=1}^{4} \delta_a \frac{n_a}{n_c} + \delta z + \varepsilon \qquad (2)$$

where n_{biog}, n_{orphg}, and n_{orphb} are, respectively, the number of girls with two biological parents, number of orphan girls, and number of orphan boys in the household, which enter (2) as proportions of n_c, leaving the proportion of boys with biological parents as the reference child category. Thus, the gender coefficients indicate the amount by which average expenditure share on an adult good is expected to change if a boy with biological parents is replaced by a girl/boy in each of the other three child categories specified in (2), holding constant the age structure of the children. The child gender coefficient estimates can be similarly interpreted: bias against each girls/boys category in (2) should result in a *positive* sign for the index of that category. We provide least square estimates by equations (1) and (2) for each i=4 adult good in this study. All expenditures are calculated on an annual basis, and the definition of total expenditure employed includes imputed rent, but excludes investment in house ownership.

(1) and (2) should also be modified to test the hypotheses underlying the other two approaches to child welfare in orphanhood discussed in section 1. Regarding the differences in treatment of paternal and maternal orphans, we have

$$w_j = \alpha_0 + \alpha_1 \ln\left(\frac{x_h}{n_h}\right) + \alpha_2\left[\ln\left(\frac{x_h}{n_h}\right)\right]^2 + \eta \ln n_c + \frac{1}{n_c}\left(\gamma_{orphp}\, n_{orphp} + \gamma_{orphm}\, n_{orphm} + \gamma_{orphp*m}\, n_{orphp*m}\right) + \sum_{a=1}^{4} \delta_a \frac{n_a}{n_c} + \delta z + \varepsilon \quad (3)$$

where n_{orphp}, n_{orphm}, and $n_{orphp*m}$ are, respectively, number of paternal orphans, number of maternal orphans that enter (3) as proportions of n_c, and an interactive term between the two in order to take into account the additional effect of double orphans on the budget share of adult goods; the proportion of children with biological parents is the reference child category. As paternal and maternal orphans are not mutually exclusive categories, the effect of being a two-parent orphan (double orphan) consists of the sum of coefficient estimates on paternal and maternal indices above plus the interaction term of the two; a significantly positive interaction term would indicate greater adult discrimination against orphans than that suggested from the sum of the one-parent orphan (single orphan) coefficient estimates, while a significantly negative interaction term would signal a smaller impact than the sum of the estimates obtained from the two single-orphan coefficients (see Ainsworth *et al.* 2005). The gender-distinct version of (3) can be similarly formulated as

$$w_j = \alpha_0 + \alpha_1 \ln\left(\frac{x_h}{n_h}\right) + \alpha_2\left[\ln\left(\frac{x_h}{n_h}\right)\right]^2 + \eta \ln n_c + \frac{1}{n_c}\left(\gamma_{aliveg}\, n_{aliveg} + \gamma_{orphpg}\, n_{orphpg} + \gamma_{orphpb}\, n_{orphpb} + \gamma_{orphmg}\, n_{orphmg} + \gamma_{orphmb}\, n_{orphmb}\right) + \sum_{a=1}^{4} \delta_a \frac{n_a}{n_c} + \delta z + \varepsilon_j \quad (4)$$

where n_{aliveg}, n_{orphpg}, n_{orphpb}, n_{orphmg}, n_{orphmb} are the number of girls with both parents alive, number of paternal orphans girls, number of paternal orphan boys, number of maternal orphan girls, and number of maternal orphan boys, the reference category being the proportions of boys with both parents alive[7].

Finally, in order to test the hypothesis of the impact of the genetic relatedness to adults' treatment of children, we write the share equation without gender distinction as

$$w_j = \alpha_0 + \alpha_1 \ln(\frac{x_h}{n_h}) + \alpha_2[\ln(\frac{x_h}{n_h})]^2 + \eta \ln n_c + \frac{1}{n_c}(\gamma_{adopt}\, n_{adopt} + \gamma_{fost}\, n_{fost}) + \sum_{a=1}^{4}\delta_a \frac{n_a}{n_c} + \delta z + \varepsilon_j \qquad (5)$$

where n_{adopt}, n_{fost} are the number of adopted and fostered children and enter (5) as proportions of n_c, with the proportion of children with biological parents as the reference child category. In this study, we define adopted children as those with at least one parent dead, while the fostered are those with both parents alive, but at least one of them not living with the child.[8]

The gendered version of (5) is written as

$$w_j = \alpha_0 + \alpha_1 \ln(\frac{x_h}{n_h}) + \alpha_2[\ln(\frac{x_h}{n_h})]^2 + \eta \ln n_c + \frac{1}{n_c}(\gamma_{biog}\, n_{biog} + \gamma_{adoptg}\, n_{adoptg} + \gamma_{adoptb}\, n_{adoptb} + \gamma_{fostg}\, n_{fostg} + \gamma_{fostb}\, n_{fostb}) + \sum_{a=1}^{4}\delta_a \frac{n_a}{n_c} + \delta z + \varepsilon_j \qquad (6)$$

where n_{biog}, n_{adoptg}, n_{adoptb}, n_{fostg}, n_{fostb} are, respectively, the number of girls with biological parents, number of adopted girls, number of adopted boys, number of fostered girls, and number of fostered boys, all of which enter (6) as proportions of n_c; the reference child category is the proportion of boys with biological parents. However, if a household contains both orphan and fostered children with the same degree of relatedness to the household head, then this additional effect should be independently controlled in order to obtain clearer evidence for the hypothesis. I deal with this issue by adding an interactive tem ($n_{adopt*}n_{fost}$) to (5) and (6), and discuss the outcome in section 5. Once again, each of the child category coefficient estimates in (5)-(6) measures the difference between the reference category and the other child category proportions by determining the amount of change they induce in the budget share of an adult good. I examine the estimates obtained from the applications of (1)-(6) in three separate parts of section 5 below.

4—Data

The data for this study are based on the 2004-2005 HICES conducted by the Ethiopian government's Statistical Authority, consisting of over 21,299 households and covering the entire country. Table 1 show the samples employed in this study. Nuclear households with at least one child account for 52% of the total; about a third (17% of all the households surveyed) with non-biological children of the head or spouse. Moreover, the HICES 2004-2005 suggests that Ethiopia has a significant orphan problem. The survey defines these as children with dead parents or children living in households with absent biological parents; respectively 12% and 31%. The household is the observation unit throughout this study; descriptive statistics referring to various individual-level averages are obtained from the number of individuals residing in the sample households.

Table 1

Average Years of Age by Child Category (standard deviations in brackets)		
	Head/Spouse Biolog. Children	Orphan Children
All Children	5.69 (4.191)	9.71 (3.909)
Boys	5.72 (4.226)	9.65 (3.860)
Girls	5.65 (4.151)	9.75 (3.956)

The notably higher average age of orphan children given in table 1 blow is consistent with

HIV/AIDS as the main cause of orphanhood in Ethiopia; the disease hits mainly the middle age adults with older children. This pattern of orphan age is hard to explain in terms of other historical causes of the Ethiopian orphanhood such as war or famine. The living arrangement of children provides an important context for different approaches to the welfare analysis of orphanhood. Tales 2 and 3 provide some clues as to the dominant cause of orphanhood. First, table 2 shows the average orphan child is nearly twice as old as the average non-orphan child. Second, table 3 reveals another related and noteworthy feature of the sample, namely that it contains substantially more orphans with a dead father than with a dead mother. Taken together with table 1, the evidence is consistent with HIV/AIDS as the main cause of Ethiopia's orphanhood because men have a higher mortality rate, and because women tend to marry older men (see Ainsworth 2005). The other causes of orphanhood such as famine, hunger or non-HIV diseases are less likely to result in age structures of this type [9].

We employ four categories of expenditure that appear to be reasonable candidates for adult goods in Ethiopia: tobacco (only a few observations available for alcohol), coffee, adult clothing[10] and adult personal services and personal effects and items; the latter category consists of shaving-related goods and services, in addition to perfumes, handbags and wallets, walking sticks and wigs, and provide the strongest evidence of gender bias. As with most adult goods, consumption by children cannot be ruled out. However, the above items are selected on the robustness to age definition of a child reported in Koohi-Kamali (2008). That study employed the current survey and the same list of adult goods based on samples with a child definition age of 15 and 16, and show that test outcomes remain nearly unchanged, and insensitive to the precise child age definition. The survey incidence of positive purchase ranges from just under a quarter to well over a third of the households. As such, the percentages of non-purchase are not unlike those expected of some common adult goods in India (Subramanian & Deaton 1991). The observations are inclusive of zero expenditure as discussed in section 3.

Table 2

Living Arrangement of Children Aged 0-15 Years by Parental Vital Status (row % in brackets)					
	Maternal Orphans	Paternal Orphans	Double Orphans	Biological/ Non-Orphan	Total
All Children	523 (2.44)	1286 (6.04)	261 (1.21)	19356 (90.30)	21426 (100)
Boys	253 (2.32)	619(5.67)	136 (1.25)	9915 (90.77)	10923 (100)
Girls	270 (2.57)	667 (6.35)	125 (1.19)	9441 (89.89)	10503 (100)

Table 3

Living Arrangement of Fostered Children Aged 0-15 Years by Parental Presence/Absence (row % in brackets)					
	Mother Not Living In	Father Not Living In	Both Parents Absent	Both Parents at Home	Total
All Children	323 (1.70)	1920 (9.20)	1122 (5.8)	15982 (82.61)	19348 (100)
Boys	173 (1.75)	907 (9.15)	449 (4.53)	8382 (84.57)	9911 (100)
Girls	150 (1.59)	1013 (10.73)	673 (7.13)	7601 (80.55)	9437 (100)

5.Estimation and Tests Outcomes

The results of applying the above method will be discussed from three perspectives, focusing on the differences: in parental treatment between orphan and non-orphan children, between children with living and dead mother/father, and between biological, orphan and fostered children. Section 5.1 presents the estimated gender effects, while section 5.2 examines the joint hypothesis of no gender effects on adult consumption by testing the versions of the model with and without child gender against each other based on the estimates reported in section 5.1.

5.1—Estimation of child gender effects on consumption

However, first we must discuss the question of gender endogeneity of orphan variables in (1)-(6); endogeneity poses a problem for this and other similar cross-section studies of orphanhood (for example Case *et al.* 2000). Nonetheless, some insight into this issue can be gained from modeling the probability of the decision to take in one or more orphans conditional on the household's demographic, gender and socio-economic features (for a similar attempt, see also Gong *et. al.* 2005.) Table 4 reports the outcome for a probit model with an indicator dependent variable of one if there is at least one orphan child in the household, zero otherwise. Model 1 (Columns 2 and 3) of the table shows the results with the proportion of girls and of children in each age group as independent variables, with the age group 13-15 years as the reference, and logarithms of household size and total expenditure and its square; with additional controls, see section 3. as discussed in section.

Note the probability of taking in an orphan is statistically significant and inversely related to the proportion of girls in the household. More interesting is how the age of children is also negatively related to the probability of having orphans in the family. These negative age coefficient estimates tend to decline in size and in statistical significance as child age increases: it is largest for young children, while negative effects disappear for the oldest child aged 10-12. To explore further the relationship between the probability of having an orphan child and child gender and age, model 2 shows the probit estimates with the proportion of girls and the proportions of children in each of the four child age groups as interactive terms. Once again, the interactive variables for the youngest children are negatively significant and tend to decline in size and significance for the older group (7-9 years), while the negative effect turns positive,

Table 4. Probit Estimates for Probability of Having Orphan(s) in the Household, Conditional on Household Demographic and Socio-economic Characteristics, Ethiopian HICES 2004-2005

	Model 1		**Model 2**	
Variable	**Coefficient**	***z*-value**	**Coefficient**	***z*-ratio**
Log(pcexpend)	.91877	1.32	0.69629	1.01
Log(pcexpend)^2	-0.07083	-1.42	-0.05276	-1.06
Log(hhold size)	0.03267	1.48	0.026188	1.18
Girl_proportion	-0.11483	-1.98	-	-
Child_0-2	-1.76875	-6.27	-	-
Child_3-6	-0.74377	-8.07	-	-
Child_7-9	-0.34294	-3.69	-	-
Child_10-12	0.00039	0.00	-	-
*Girl*0-2*	-	-	-0.41138	-2.52
*Girl*3-6*	-	-	-0.48619	-4.56
*Girl*7-9*	-	-	-0.10803	-1.03
*Girl*10-12*	-	-	0.11557	1.11
Constant	-4.08920	-1.68	-3.57388	-1.48
Log Likelihood	-2387.22		-2429.24	
Pseudo R^2	0.3212		0.3093	
%correct predic	0.86		0.86	
Sample size	8289			

Notes: Dependent variable: 1 if the household contains only biological kid(s); 0 if at contains orphan(s). Reference group: Addis Ababa for region, 13-15 years for child age; proportion of boys for child gender.

though not significantly so, for the oldest child age group of 10-12. This suggests that a possible motive for taking an orphan into the household is the benefits derived from the labor of that child: a very young child imposes additional costs on the household without such benefits; contrary to the expectation that if adopted at a very young age, fostered orphans are also more likely to be treated like biological children (Beegle *et al.* 2007). The probability of adopting a child conditional on the child's being a girl is significantly negative but declines for older girls. This outcome is also consistent with the large difference between non-biological and biological children given in table 2; the former being on average nearly twice as old regardless of gender.

We employ a larger list of variables to examine intrahousehold inequality than that in table 4, given the importance of extensive controls in budget share analysis[1]. Turning to the orphan effects on adult consumption, Table 5, top panel, suggests that there is positively significant evidence of discrimination against orphans by two categories (adult items and tobacco). However, the lower panel reveals that seven out of 12 positively significant instances are accounted for by the girls' effects on consumption; only one instance (coffee) by the boys' effect. Thus, the gender effects dominate orphan effects.

Table 5

Effects of Orphans on the Budget Share of Adult Goods
(estimates have heteroskedasticity-consistent standard errors; *t*-ratios in brackets)

Adult Goods	Ad. Person. Items	Coffee	Tobacco	Adult Clothes
	row estimated by (2), reference group: biological children			
Orphans	0.00206 (2.05)	0.00214 (1.23)	0.00156 (1.85)	0.00034 (0.52)
	rows estimated by (4), reference group: biological boys			
Biological Girls	0.00199 (3.87)	0.00757 (2.65)	0.00155 (1.69)	0.00057 (0.97)
Orphan Girls	0.00395 (3.84)	0.00891 (2.56)	0.00311 (2.83)	0.00226 (2.16)
Orphan Boys	0.00185 (1.51)	0.00992 (2.83)	0.00067 (0.68)	0.00004 (0.05)

Table 6, upper panel, indicates three significantly positive instances of discrimination out of 12; note the insignificant double-orphan effects for all expenditure categories. The lower panel shows 10 out of 20 positively significant instances of discrimination are against girls, both maternal and paternal. Note the absence of positively significant boy effects on adult consumption; once gain the gender effects dominate.

[1] Ensuring non-missing variables across this extensive list of control variables results in somewhat smaller samples than that employed in table 4.

Table 6

Effects of Proportions of Maternal/Paternal Orphans on the Budget Share of Adult Goods (estimates have heteroskedasticity-consistent standard errors; *t*-ratios in brackets)*****				
Adult Goods	Ad Person. Items	Coffee	Tobacco	Adult Clothes
	rows estimated by (3), reference group: biological children			
Paternal Orphans	0.00027 (2.69)	0.00180 (0.84)	0.00124 (1.17)	0.00133 (2.36)
Maternal Orphans	0.00056 (0.56)	0.00780 (2.07)	0.00181 (1.44)	0.00089 (0.91)
Maternal*Paternal	-0.00096 (0.39)	-0.00511 (1.25)	-0.00144 (0.73)	-0.00036 (0.30)
	rows estimated by (4), reference group: boys with parents alive			
Girls/ ParentsAlive	0.00193 (3.66)	0.00370 (2.78)	0.00160 (1.76)	0.00051 (0.83)
Girls/FatherDead	0.00454 (3.58)	0.00281 (1.13)	0.00260 (2.10)	0.00257 (1.68)
Boys/FatherAlive	0.00211 (1.28)	0.00110 (0.35)	-0.00016 (0.16)	0.00140 (1.35)
Girls/MotherDead	0.00231 (1.81)	0.00989 (2.02)	0.00287 (2.37)	0.00243 (1.64)
Boys/MotherDead	-0.00109 (1.20)	0.00364 (0.72)	-0.00035 (0.21)	-0.00049 (0.51)

Table 7, upper pane, indicates fostered children have larger positive effects on the consumption of adult items and clothes consumption compared to adopted children; apparently supporting more favorable treatment of the genetically related children. However, the lower panel reveals the impacts of both adopted and fostered girls are mostly positive and statistically significant. By contrast, there is only one positive instance for boy effects (on adult items; a negative one for tobacco). Note also that, except for coffee, there are larger positive effects for biological girls than adopted boys. Hence, the gender effects dominate even when the degree of genetic relatedness is held constant, an outcome inconsistent with the genetic hypothesis. We have obtained similar results to those reported here for a weaker version of the hypothesis neutral with respect to the within-category characteristics of children, such as their age or gender. That is, testing whether biological girls are expected to be treated better than orphan and/or fostered boys. The test outcomes reported here remain robust (results available from the author.)

Two further issues of possible relevance remain. One important type of nuclear household in the context of high HIV are the grandparent-headed households looking after their grandchildren, see Duflo (2003). However, the HICE 2004-05 does not provide sufficient information to distinguish between grandparents and other types of related care-takers, hence table 3

Table 7

Effects of Proportions of Adopted & Fostered Children on the Budget Share of Adult Goods (estimates have heteroskedasticity-consistent standard errors; *t*-ratios in brackets)*****				
Adult Goods	Ad. Person. Items	Coffee	Tobacco	Adult Clothes
	rows estimated by (5), reference group: biological children			
Adopted Children	0.00105 (1.82)	-0.00021 (0.13)	0.00065 (0.79)	-0.00048 (0.37)
Fostered Children	0.00177 (2.34)	0.00360 (1.66)	-0.00044 (0.53)	0.00165 (1.47)
	rows estimated by (6), reference group: boys with parents living in the same family			
Girls+par. live-in	0.00181 (3.34)	0.00284 (2.16)	0.00145 (1.70)	0.00025 (0.41)
Adopted Girls	0.00248 (2.41)	0.00208 (1.08)	0.00281 (2.50)	0.00021 (0.14)
Adopted Boys	0.00127 (1.46)	0.00106 (0.41)	-0.00002 (0.02)	-0.00097 (0.66)
Fostered Girls	0.00269 (2.98)	0.00899 (3.33)	0.00198 (1.82)	0.00244 (1.67)
Fostered Boys	0.00253 (2.19)	0.00026 (0.09)	-0.00191 (1.65)	0.00067 (0.59)

provides a somewhat limited view of the living arrangement of the Ethiopian orphans although the nuclear family sample restriction imposed does not remove any information in that regard.

The second issue is whether the presence of the alive parent of maternal or paternal orphans in the household, or the financial contributions to the household by the fostered child parents (Case *et.al.* 2006), affect the above results. We have examined this question by adding orphan child's parent living-in and financial receipt variables and their interactives with various orphan child categories. The results (available from the author) remain very similar to those reported here; the additional variables are predominantly statistically insignificant.

5.1—Test gender effects on consumption

However, these single-parameter tests provide evidence for gender effects of each child category with and without gender distinction whereas the principle outcome of interest is whether such effects represent joint statistically significant differences between boys and girls. This section tests the joint restrictions for gender effects in (1) *vs.* (2), (3) *vs.* (4), and (5) *vs.* (6).

Table 8 shows the *chi*-square values for Wald tests of the above joint gender parameter restrictions for each of the four adult goods of this study; each row tests the prediction of each hypothesis regarding the treatment of orphans. We note that in 10 out of the 12 cases in table 8, the hypotheses of no child gender effect are rejected. The strongest evidence is for adult personal items, the weakest for adult clothes. For the first row, even the *chi*-square value for adult clothes rejects the hypothesis of no gender effect between orphan/non-orphan boys and girls. The estimates for the lower two rows reject the hypothesis of no gender bias: three of the four adult goods values given in each row, corresponding to each one of the three types of orphanhood, are above their relevant *chi*-square critical values.

The above can be contrasted with the gender effects estimated in Koohi-Kamali (2008). In that study, households with non-biological children of the household head were excluded from the samples employed, and clear evidence of child gender bias in Ethiopia by consumption of adult goods was reported, indicating that gender effects remain dominant regardless of the orphan child presence in the household.

The question is whether the gender of orphans has an impact of its own on adults' consumption, and if so, to what extent the test statistics reported in table 8 can be attributed specifically to gender bias related solely to orphanhood. Table 9 shows the results of testing the joint gender coefficient restriction confined to gender restriction tests for the coefficient estimates on children who are *not* biological offspring of the household head and his/her spouse. The *chi*-square values are smaller, indicating that in some instances, the child gender bias does not have an additional impact for orphans beyond that captured by the biological child-parent link. However, it is still the case that in 7 out of 12 instances, the joint restrictions hypothesis that orphan gender has no impact is rejected. Since for each child category in tables 8 and 9, the degree of relatedness for boys and girls is held constant, the jointly statistically significant differences between boys and girls render further evidence against the Hamiltonian hypothesis.

Table 8

Chi-square Values for Wald Test of Joint Restrictions of Child Gender Effects on Consumption of Adult Goods#

Goods	Adult Items	Coffee	Tobacco	Adult Clothes
(1) vs. (2) restriction tested: proportion of biological girls=0 & proportion of orphan girls=proportion of orphan boys				
Orphan	17.71	7.31	6.96	6.01
(3) vs. (4) restriction tested: proportion of biological girls=0 & proportion of paternal orphan girls=proportion of maternal orphan boys & proportion of maternal orphan girls=proportion of paternal orphan boys				
Dead Parents	19.16	10.21	8.01	3.17
(4) vs. (6) restriction tested: proportion of biological girls=0 & proportion of adopted girls=proportion of adopted boys & proportion of fostered girls=proportion of fostered boys				
Adopt/Foster	13.56	8.50	14.19	2.36

First and second rows based on eq. (1)-(4) with the orphan sample (6745); third row based on eq. (5)-(6) with the fostered sample (7102), see Appendix A.
1st row: critical *chi*-square for d.f=2 at 5% confidence level: 5.99
2nd row: coefficient; critical *chi*-square for d.f=3 at 5% confidence level: 7.81
3rd row: critical *chi*-square for d.f=3 at 5% confidence level: 7.81

Table 9

Chi-square Values for Wald Test of Restrictions of Child Gender Effects across Orphan Child Categories on Consumption of Adult Goods#

Goods	Adult Items	Coffee	Tobacco	Adult Clothes
(1) *vs.* (2) restriction tested: *proportion of orphan girls=proportion of orphan boys*				
Orphan	4.06	0.12	4.67	1.81
(2) *vs.* (4) restriction tested: *proportion of paternal orphan girls=proportion of paternal orphan boys & proportion of maternal orphan girls=proportion of maternal orphan boys*				
Dead Parents	7.94	6.11	2.92	0.84
(5) *vs.* (6) restriction tested: *proportion of adopted girls=proportion of adopted boys & proportion of fostered girls=proportion of fostered boys*				
Adopt/Foster	1.81	2.33	12.80	6.46

First and second rows based on eq. (1)-(4) with the orphan sample (6745); third row based on eq. (5)-(6) with the fostered sample (7102), see Appendix A.
1st row: critical *chi*-square for d.f=1 at 5% confidence level: 3.84
2nd row: critical *chi*-square for d.f=2 at 5% confidence level: 5.99
3rd row: critical *chi*-square for d.f=2 at 5% confidence level: 5.9

6—Conclusion

Although testing the predictions of the Rothbarth model employed requires only cross-section data, the orphan effects obtained from this model may prove to be endogenous when analyzed with longitudinal survey data. However, we have employed an extensive set of observable controls for individual and community-level differences in order to mitigate such sources of estimation bias.

Subject to this limitation, the evidence of this study highlights the statistically significant impact of child gender on adult consumption; the effects are present to varying degrees in all four categories of adult goods examined in this study. Regarding maternal orphan, the weight of evidence in this paper tends to lend *some* support to the hypothesis that maternal orphans girls are more poorly treated than maternal orphan boys, that is the loss of a mother is more critical for a girl. Moreover, most of the evidence reported in this paper suggests statistically significant differences between boys and girls who stand in the same degree of relatedness to the household head, and that is contrary to the Hamiltonian rule. The joint restrictions on all female/male child coefficients in the budget-share equations are widely rejected; similar gender tests confined to children not biologically linked to the head are for the most part also rejected. This suggests that the earlier approaches not based on an explicit model of intrahousehold resource distribution may well have missed the evidence on the effects of orphanhood on consumption.

A good deal of attention is paid to the public policy on the adoption of Ethiopian children from countries outside Ethiopia; little to the welfare consequences of child adoption inside Ethiopia. The public policy implications of our finding point to benefits of child labour as a likely motive for child adoption within Ethiopia; contrary to the belief that children adopted at young age are likely to be better treated. Hence there is need for a policy outcome focused on the age and gender of the adopted children. Broad policies that support families but lack this focus are likely to achieve little. At a more general level, public policy designed to lower the incidence of HIV directly will substantially contribute to the resolution of the orphan problem in Ethiopia.

Notes

[1]For example, for Tanzania, Beegle *et al.* (2007:10) found "that the impact of orphanshood had no gender dimension—the impact was never significantly sex-dependent." Similarly, for South Africa, Case and Ardington (2006:417) reported "no evidence that female orphans are at particular risk"; although the multi-country cross-section study by (Case *et al.* 2004) did find that girls face a significantly greater risk of being kept out of school in several African countries, they also state that "female orphans are at no particular risk" in that regard (Case *et al.* 2004:502). A possible exception is Ainsworth et al. (2005:429-430), which found a large decline in number of hours of school attendance among maternal orphan girls, but found no similar gender effect regarding enrollment: "Girls are no less likely to attend than boys, at all ages."

[2] The approach is not explicitly stated as a hypothesis, but the theoretical and empirical evidence that child welfare improves when the mother controls the division of household resources provides indirect support for this hypothesis (Lundberg and Pollack 1993; Hoddinott and Hadad 1995).

[3] A weaker version of the hypothesis may be proposed with the degree of relatedness neutral with respect to the within-category characteristics of children, such as their age or gender. In this version within-category differences are permitted but the hypothesis predicts that across-category differences should depend on the degree of relatedness *regardless* of child characteristics. That is, biological girls are expected to be treated better than orphan and/or fostered boys; see section 5 on the evidence for the weaker version of the hypothesis.

[4] Orphan children have at least one parent dead, while fostered children are commonly defined as those whose biological parents are alive, but who live in households with non-biological parents; see section 3. The non-orphan fostered child may be better off than an orphan if absent parents give financial support to the child, choose better care-providers, and monitor them (Case *et al.* 2006).

[5] Defining gender indices by age with reference to the number children is the standard, for example Deaton (1987), or Kingdon (2005); the critical difference here is that the gender index is defined regardless of child age as proportion of children.

[6] As two-step decision approach has had a mixed record of successful identification of the incidence of zero consumption for applications to expenditure data. For example, from estimates on educational expenditure in Kingdon (2005), the evidence is mixed on whether the two-step method provides significant additional insight on gender bias beyond that offered by single-equation school enrollment estimation. The identification becomes a particularly acute problem with application to some key adult goods such as tobacco and alcohol, see Atkinson *el. at* (1984), because the zero consumption observations are a combination of non-consumption and measurement error zeros due to infrequency of purchase left out by the survey adopted cycle of consumption, see Deaton and Irish (1986).

[7] It turns out that the estimates from (3) suggest that the interactive term for double orphans, to be discussed in section 5 below, affects the budget shares insignificantly.

[8] There are thus no orphans in the fostered category, which contains the smaller group of "not-related" fostered children, for example domestic child servants in urban households.

[9] The HICE 2004-05 questionnaire does not provide direct information on the incidence of HIV in the sampled households.

[10] The HICES does provide separate codes for adult (older than 15 years) expenditure but only for clothing items (with survey item code class of 202); not for footwear.

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APPENDIX A: descriptive data

Table 2a-Mean and St. Dev. of the principle Variables				
Variables:	Mean	Stand. Deviation	Mean	Stand. Deviation
	'Orphan' Sample		'Fostered' Sample	
Ad. Personal bud. share	0.0040358	0.0153129	0.0040787	0.0153364
Adult Cloths bud. share	0.0019591	0.0140839	0.0020831	0.0154075
Coffee budget share	0.0130555	0.0453582	0.0130603	0.0449826
Tobacco budget share	0.0062068	0.0222804	0.0061619	0.0219525
Log. total expenditure	7.2204597	0.5689646	7.2324212	0.5709257
Log. total expendit. Sq.	52.4587112	8.4171823	52.6338269	8.4691739
Log. Household size	0.8591529	0.5775614	0.8321878	0.5842834
Prop. Children 0_2	0.3062517	0.3396914	0.2939220	0.3376579
Prop. Children 3_6	0.2800469	0.2833801	0.2763165	0.2872991
Prop. Children 7_9	0.1728955	0.2277303	0.1772947	0.2390554
Prop. Children 10_12	0.1275377	0.2139519	0.1313210	0.2234462
Prop. Biolog. Children	0.4504973	0.4975804	0.4500846	0.4975373
Prop. Biolog. Boys	0.2093403	0.4068679	0.2068431	0.4050706
Prop. Biolog. Girls	0.4121572	0.4922596	0.4114334	0.49211281
Prop. Orphans	0.1393842	0.3191967	-	-
Prop. Orphan Gilrs	0.0720472	0.2189816	-	-
Prop. Orphan Boys	0.0673370	0.2133353	-	-
child, both parent alive	0.8584759	0.3204032	-	-
girl, both parents alive	0.4153319	0.3576660	-	-
boy, both parents alive	0.4431440	0.3627068	-	-
Prop. Paternal orphans	0.0766887	0.2206977	-	-
Prop. Pat. orphan girl	0.0444887	0.1754644	-	-
Prop. Pat. orphan boy	0.409014	0.1679614	-	-
Prop. maternal orphans	0.0426425	0.1702500	-	-
Prop. mat. orphan girl	0.0159164	0.1057538	-	-
Prop. mat. orphan boy	0.0176545	0.1104682	-	-
Prop. adopted child	-	-	0.0840762	0.2608274
Prop. adopted girl	-	-	0.0415544	0.1693306
Prop. adopted boy	-	-	0.0425219	0.1705838
Prop. fostered child	-	-	0.1128995	0.2811302
Prop. fostered girl	-	-	0.0681847	0.2158767
Prop. fostered boy	-	-	0.0447148	0.1794433
Sample size#	6745		7102	

Orphan sample exclusive of households with fostered children; Fostered sample inclusive of households with fostered children.

List of additional controls and their interactive terms (retained only if statistically significant)**.**

Individual and household level: husband and spouse age, education, employment and marital status, asset ownership, including land, cattle and durables; household amenities, including cooking, drinking water, sanitation amenity.

Community-level: availability of public work scheme, school feeding program, distance to the nearest school, hospital, marketplace, and to the sources of water and fire-wood.

Country-wide: Rural-urban dummy; and 10 regional controls with Addis Abba as the reference category.

Intercatives: household-level interactives for age, education, employment and marital status; community-level and household-level interactives with public works, school feeding program, sources of water and fire-wood; and country-wide and household-level interactives with urban-rural and regional dummies.