Eat Drink Man Woman:
Testing for Gender Bias in China Using Individual Nutrient Intake Data*

Albert Park
Department of Economics, University of Michigan

Pungpond Rukumnuaykit
Department of Economics, Michigan State University

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Abstract

We present unitary and sharing rule models of the household that explicitly account for three parental concerns that may lead to gender bias in the allocation of resources to children—equity, efficiency, and preferences. Deaton's test of the effect of household composition on adult good expenditures is employed using data on fathers' and mothers' nutrient intake from the China Health and Nutrition Survey. We find that rural fathers, especially less educated ones, favor sons while rural mothers do not. Parental differences in gender bias are statistically significant, a result which is inconsistent with the unitary model and equity bias explanations of gender bias.

JEL codes: D12, I12, J13, J16, O12, O53

Keywords: household, children, nutrition, gender, China

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Correspondence author: Albert Park, Department of Economics, University of Michigan, Ann Arbor, MI 48109-1220, email: alpark@umich.edu.

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1. Introduction

It is now well-established that the unitary household model overlooks aspects of intra-household resource allocation that carry important implications for the welfare of individual family members (Haddad and Kanbur, 1990; Haddad, Hoddinott, and Alderman, 1997). A growing body of empirical evidence finds that in many developing countries women and girls, in particular, receive fewer household resources than men and boys, which may lead to poorer relative health and nutritional status (World Bank, 2001).

In this paper, we define gender bias broadly to be *any* difference in household resource allocation associated with characteristics that differ by gender. In allocating resources to sons and daughters, three main concerns of parents are equity, efficiency, and preferences. The equity concern refers to the desire of parents that children be equally well-off. If the needs of boys and girls differ, whether in terms of nutrition, other inputs such as clothing, parental time, etc., then observed gender bias may be due to "equity bias." For example, boys could have poorer health endowments than girls and so require greater nutritional inputs. The efficiency concern refers to the economic returns to investments in boys versus girls. Parents may benefit directly from the labor returns of children, whether through home goods production, income generation, or from future returns through remittances or cohabitation in old age. If these returns depend on child investments and differ for boys and girls, gender bias can arise from "efficiency bias." So defined, efficiency bias does not imply that allocations are efficient in an economic sense since parents maximize their selfish returns rather than the welfare of their children.

Rather, it recognizes that differential treatment of boys and girls may be consistent with rational optimizing behavior by parents, independent of the needs of children or any inherent preference favoring the welfare of boys versus girls. Finally, gender bias may arise from discriminatory attitudes of parents, or "preference bias." For example, in many Asian societies where sons carry on the family name, families often are thought to exhibit son preference, which could lead to resource allocation decisions favoring boys over girls.[1]

A standard approach for testing for gender bias in household expenditures is to measure whether the effect of an additional son on the consumption of adults differs from the effect of an additional daughter (Deaton, 1987, 1989, 1997; Deaton, Ruiz-Castillo, and Thomas, 1989). The advantage of this method is that it can be implemented using household expenditure and demographic data commonly available in large-scale household surveys. Although it has been applied to data from developing countries around the world, the test has not proven to be very robust in practice.[2] Part of the problem may be that the tests frequently rely on expenditure data for assignable adult goods such as tobacco and alcohol that constitute a relatively small share of expenditures. In contrast, we focus on adult food consumption, which accounts for the majority of household expenditures in many developing countries.[3] In 1993, the last year of our data,

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[1] One could argue that preference bias is similar to efficiency bias if boys provide greater "psychic" income to parents, but we prefer to draw a distinction between material and "psychic" benefits.

[2] Few studies found statistically significant evidence of gender bias even in countries where there is evidence that such bias is likely to exist. Gibson and Rozelle (forthcoming) provide a summary of applications of the test to data from Bangladesh, India, Pakistan, Taiwan, Cote d'Ivoire, and Burkina Faso (see also Deaton, 1997) and show that in contrast to other countries, there is strong evidence of gender bias in Papa New Guinea.

[3] The consumption of goods with small expenditure shares could be less income elastic, the noise-signal ratio for such goods could be greater, or many households could report zero consumption of such goods. On the other hand, demand for food or calories could be less income elastic than luxury goods, especially at low levels of income.

food consumption accounted for 58 and 50 percent of household expenditures in rural and urban China, respectively (National Statistical Bureau, 1994). An additional limitation of Deaton's method as commonly implemented is that it cannot distinguish among the three types of gender bias described above.

In this paper, we develop tests of gender bias that exploit individual nutritional intake data. We treat fathers and mothers consumption of calories, protein, and fat as adult goods in implementing tests à la Deaton. Such data make possible a richer characterization of intra-household resource allocation than is possible using household-level expenditure data. Specifically, we can look separately at the consumption of each parent to test for differences in the gender bias of fathers and mothers. We argue that such differences are inconsistent with the unitary model and cannot be explained by equity bias.

To support this claim and to motivate other empirical predictions, we first present unitary and sharing rule models of parental consumption that explicitly account for the three sources of gender bias, and which yield theoretical insights of independent interest. We then conduct the empirical analysis using data from the 1991 and 1993 waves of the China Health and Nutrition Survey, which collected detailed information on individual nutritional intake from rural and urban households in eight Chinese provinces. Our main finding is that rural fathers, especially those that are poorly educated, exhibit gender bias while rural mothers do not, with the difference in gender bias being statistically significant. Urban parents exhibit gender bias only toward adolescent children (age 12 to 15) but there is no evidence of differences in the gender bias of urban fathers and mothers.

The paper is organized as follows. First, we review the relevant literature, including existing evidence of son preference in China. In section 2, we present unitary and sharing rule models of parental consumption and derive comparative statics for the effects of equity, efficiency, and preference bias. In section 3, we describe the empirical methodology. The data and some descriptive results are discussed in section 4. The main empirical findings are presented in section 5. Section 6 concludes.

Previous Literature

A large body of literature examines gender bias in household resource allocation in developing countries. The strongest evidence for gender bias is from South and Southeast Asia. For India, in particular, researchers have found that girls have higher infant mortality rates (D'Souza and Chen, 1980; Rosenzweig and Schultz, 1982), poorer nutrition as measured by anthropometric indicators (Sen, 1984; Sen and Sengupta, 1983; Behrman and Deolalikar, 1990), and receive fewer household resources (Subramanian and Deaton, 1990). Research has also found evidence that boys receive more nutrients than girls in Bangladesh (Chen, Huq, and D'Souza, 1981) and the Philippines (Evenson et al., 1980; Senauer et al., 1988). Income and price elasticities of demand for girls is greater than for boys for health care in Pakistan (Alderman and Gertler, 1997), and for education and health care in other developing countries (World Bank, 2001).

A few studies have attempted to distinguish among the different types of gender bias. Behrman, Pollak, and Taubman (1982) find that in the U.S., parent educational investments in children are more equitable than would be expected in a pure investment model of such investments. Rosenzweig and Schultz (1982) present evidence of efficiency bias in household resource allocation in India by showing a connection between future productivity and survival probability. They also suggest that equity concerns increase with wealth. Behrman (1988) presents evidence of pro-male bias in nutrient investments in the lean season in India, and also finds that equity concerns matter, but not as much as in the U.S. Pitt, Rosenzweig, and Hassan (1990) find that the higher level and variance of men's consumption in Bangladesh in part reflects the greater sensitivity of their productivity to health status, but that households also care about equity. Kremer et al. (1997) find that even after accounting for different needs based on energy expenditure, women in Bangladesh consistently received less of their energy requirements than their children or husbands.

Previous work has found that the way in which fathers and mothers allocate resources, especially food, is different. It is widely perceived that men spend a higher share of their income on personal consumption than women (Haddad, Hoddinott, Alderman, 1997). Von Braun (1988) found a positive relationship between the amount of cereal production under women's control and household calorie consumption in the Gambia. Garcia (1990) found that in the Philippines, households in which mothers had a higher share of income consumed greater amounts of calories and protein. Haddad and Hoddinott (1994) found that increasing the women's share of cash income in the Ivory Coast raised the budget share of food and lowered the budget share of alcohol and cigarettes. Similarly, Thomas (1990) found that women's nonlabor income increases the household's demand for calories and protein, and increases children's survival rates and weight-for-height. The gender preferences of mothers and fathers may also be different. Using household survey data form the United States, Brazil, and Ghana, Thomas (1994) found that "in all three countries, the education of the mother has a bigger effect on her daughter's height; paternal education, in contrast, has a bigger impact on his son's height."

In China, a common saying is that "a married daughter is like water spilled on the ground." Lee and Wang (1999) describe the long history of son preference in China:

> "Son preference [in China] dates back to the origins of ancestral worship in the second and third millennia B.C. and was reinforced by a patrilineal and patrilocal familial system, supported by the imperial and especially late imperial state, which systematically discriminated against daughters (Bray 1997). Only sons could sacrifice to the family spirits. Only sons could carry the family name. Only sons, with rare exceptions, could inherit the family patrimony (Bernhardt, 1995). Not only did patrilocal marriage customs require daughters to marry out, but also hypergamous marriage patterns required upper class families to provide a dowry to accompany them. Daughters, therefore, were not only culturally considered inferior; they were also perceived by most families as a net economic and emotional loss" (p. 47).[4]

The most widely cited evidence of gender bias in contemporary China is the high sex ratio at birth, which reportedly rose from 113 males per 100 females during 1984-1989 to 115 during 1990-1994. This large gender imbalance may be due to sex-selective abortion, higher female infant mortality rates, including female infanticide, and unreported or "missing" females (Coale and Banister, 1994; Gu and Roy, 1995; Lee and Wang, 1999). In some poor areas, the neonatal and infant female excess mortality is several times higher than the national average (Lavely, Mason, and Li, 1996). Missing females are less likely to receive schooling and hospitalization services (Zeng et al., 1993). Girls in China have lower school enrollment rates, especially in poor areas

[4] Thus the common saying that "a married daughter is like water spilled on the ground," a resource that one cannot retrieve.

(Brown and Park, 2002; Hannum and Xie, 1994). We are unaware of previous studies of the intrahousehold allocation of food or other resources in China.

2. Modeling Parental Consumption

Unitary Household Model

Consider a family consisting of a father (f), mother (m), and child (c). We begin by assuming a unitary household model in which family members have common preferences over resource allocation decisions, so that the distribution of individual incomes does not affect allocation decisions. The household solves a one-period maximization problem by allocating available income to consumption by each individual (X_f, X_m, and X_c):

$$\underset{X_f, X_m, X_c}{Max}\, U(X_f, X_m, X_c) = U_f(X_f) + \beta_m U_m(X_m) + \beta_c U_c(X_c) \tag{1}$$

$$\text{s.t.} \quad X_f + X_m + X_c = Y(X_f, X_m, X_c)$$

This simple model illustrates the effect of equity, efficiency, and preference biases on resource allocation. Equity effects on consumption are captured by differences in the individual utility functions U_f, U_m, and U_c. Efficiency concerns are incorporated by allowing income (Y) to be endogenous to individual consumption levels. Differences in the marginal return to consumption by different individuals can affect allocation decisions (i.e., income earners are better fed). Finally, preferences are captured by different weights placed on the utility of different household members (β_m, β_c).

We solve the model explicitly assuming individual log utility functions with utility parameters u_i and fixed marginal returns to individual consumption (a_i):

$$U(X_f, X_m, X_c) = u_f \log(X_f) + \beta_m u_m \log(X_m) + \beta_c u_c \log(X_c) \qquad (2)$$

$$\text{s.t. } Y = Y_0 + a_f X_f + a_m X_m + a_c X_c$$

Here, $a_i<1$ is a necessary condition for a bounded solution. The resulting model leads to an intuitive mapping of equity, efficiency, and preference concerns to parameters u_i, a_i, and β_i. Gender bias in the treatment of children arises if the parameters u_c, a_c, or β_c differ for boys and girls.

It is straightforward to solve the household's maximization problem described in (2). Consumption by household member j, whether defined as the amount or share of consumption, is increasing in u_j, a_j, and β_j and decreasing in u_i, a_i, and β_i ($i\neq j$). We can derive a closed form solution for the relative consumption of mothers (subscript *m*) versus fathers (subscript *f*), where W_i is the share of household expenditures consumed by person i:

$$\frac{W_m}{W_f} = \frac{X_m}{X_f} = \frac{u_m(1-a_f)\beta_m}{u_f(1-a_m)} \qquad (3)$$

Differences in the gender bias of fathers and mothers occur if $\frac{W_m}{W_f}$ differ for boys and girls, which by definition must stem from differences in u_c, a_c, and β_c. However, it is clear from (3) that under the unitary model, the relative consumption shares of mothers and fathers do not depend on child parameters, and so cannot be affected by the gender of the child. Differences in the equity, efficiency, or preference parameters of boys versus girls affect parental consumption but do so in equal proportion for fathers and mothers. Thus, the unitary model cannot explain differences in the gender bias of fathers and mothers.

Sharing Rule Model

The sharing rule model assumes that household resource allocation is Pareto-efficient, and so is cooperative, but is sufficiently general to accommodate different assumptions about underlying bargaining processes and preferences (Chiappori, 1988 and 1992; Browning et al., 1994; Browning and Chiappori, 1998). The unitary model can be considered a special case of the sharing rule model. The sharing rule model assumes that the consumption of father and mother can be modeled as if the couple first splits household income according to a sharing rule, which can depend on the distribution of individual incomes, bargaining, etc. Then, each individual makes consumption decisions that selfishly maximize his or her own utility. We adapt the model to allow fathers and mothers to purchase consumption for their child, which is treated as a public good.

The father solves the following maximization problem:

$$\underset{X_f, X_{fc}}{Max}\, U(X_f, X_{fc}) = U_f(X_f) + \beta_{fc} U_c(X_{fc} + X_{mc}) \tag{4}$$

$$\text{s.t.} \quad X_f + X_{fc} = Y - \theta(Y_f(X_f), Y_m(X_m), Y_c(X_{fc} + X_{mc})) \,.$$

Here, X_{fc} is the amount of income the father spends on the child's consumption, X_{mc} is the amount of income the mother spends on the child's consumption, β_{fc} measures how much the father values consumption by the child, and θ is the mother's income based on the sharing rule. θ is a function of individual incomes, so that, for example, fathers may receive a greater share of their own income than of their wives' income. For simplicity, we also assume that children's income accrues to their parents through the sharing rule. The mother solves an analogous problem to the father, choosing X_m and X_{mc}. Her income is θ rather than Y-θ and her preference weight for her child's consumption is β_{mc}.

The father's problem and the mother's problem each yields a reaction function in which the parent's spending on the child is a negative function of the other parent's spending on the child. This is because additional spending by the other parent reduces the marginal utility of additional child consumption. The two reaction functions can be solved to find a unique solution. It is then straightforward to solve for the optimal consumption of the two parents. For the functional forms for utility and income in (2), we can solve explicitly for X_m and X_f. Not surprisingly, the signs of the comparative statics for parental consumption with respect to the u_i, a_i, and β_i are the same as in the unitary case. However, the expression for the relative consumption of mothers and fathers is somewhat different (derived in the appendix):

$$\frac{W_m}{W_f} = \frac{X_m}{X_f} = \frac{\beta_{fc} u_m (1 - a_f (1 - \theta_f))(1 - a_c \theta_c)}{\beta_{mc} u_f (1 - a_m \theta_m)(1 - a_c (1 - \theta_c))} \tag{5}$$

Note first that just as for the unitary model, the utility parameter for the child (u_c) drops out, so equity bias cannot explain differences in parents' consumption. However, in contrast to the unitary model, the other parameters related to the child can affect the relative consumption of mothers versus fathers. First, if the preferences of mothers and fathers towards sons and daughters differ, the relative consumption levels will differ if the couple has a son as opposed to a daughter. A simple example would be that mothers treat sons and daughters equally (β_{mc} does not change with the child's gender) but fathers have a higher β_{fc} when the child is a son. Second, if the sharing rule does not divide children's income equally between fathers and mothers ($\theta_c \neq 0.5$), differences in the returns to consumption of children can lead to different relative consumption levels of parents. For example, if fathers control the income of children, then $\theta_c=0$ and higher economic returns to children's consumption will reduce the father's share of consumption as he invests more in children while the mother invests less. Third, if the sharing rule for allocating children's income (θ_c) is different for boys and girls (e.g., the mother controls more of the girl's income and the father controls more of the boy's income), this could also induce different relative consumption levels when the child is a boy versus a girl. While this seems unlikely for child labor returns, more abstract interpretations of this effect may be more plausible. For example, relative to fathers, mother's may receive more future utility from healthy girls than from healthy boys (e.g., emotional support, remittances, etc.) or they may place higher value on the work that girls do around the house.

The sharing rule model thus helps illustrate the possible explanations for differential gender bias of fathers and mothers. Although empirically we cannot distinguish among the three above explanations, a finding of differential gender bias is inconsistent with the unitary model and with equity bias explanations of gender bias.

3. Methodology

We modify the outlay equivalence method proposed by Deaton (1989) to test for the effect of additional children of different genders on parental consumption. When an additional child becomes part of a household, the child places new demands on household resources. If total expenditures remain constant, other household members must sacrifice their own consumption for the sake of the child. Deaton's test examines whether the reduction in expenditures on "adult goods" (e.g. makeup, tobacco, liquor, jewelry, and adult clothing) differs when children are boys rather than girls. In this paper, we exploit the availability of consumption data of individual household members by defining nutrient intake of fathers and mothers as adult goods.

For any good consumed by parent i and a demographic category r, the outlay equivalence ratio, $\pi_{ir,}$ is defined as follows:

$$\pi_{ir} = \frac{\partial(q_i)/\partial n_r}{\partial(q_i)/\partial x}\frac{n}{x} \tag{6}$$

Here, q_i is parent i's consumption, n_r is the number of persons in age-gender cohort r, n is household size, and x is total household consumption. The outlay equivalence ratio is the fractional change in per capita consumption that would have the same effect on adult good consumption as an additional person in the specific age-gender group.

In contrast to conventional practice, we define q_i and x in nutritional units (calories, grams of protein, grams of fat) rather than by expenditures.[5] Our tests therefore ignore changes in adult non-food consumption or in the price (or quality) of nutrients consumed. It would be possible in theory for a parent whose share of calories reacts similarly to the presence of a son or daughter (i.e., no evidence of gender bias) to still exhibit gender bias in his or her shares of food or non-food expenditures. Nonetheless, using nutrients rather than expenditures remains attractive because nutrients are large and essential consumption items and can be measured separately for each household member. Given a reasonable degree of separability between food and other consumption items in the utility function, it is hard to imagine preferences that would lead to an opposite gender bias in the allocation of other resources that would offset or reverse a positive finding of gender bias in nutrient allocation.

Following Deaton (1989), π_{ir} can be calculated from parameters estimated from the following equation.

$$w_i = \frac{q_i}{x} = \alpha_i + \beta_i \ln\left(\frac{x}{n}\right) + \eta_i \ln n + \sum_{1}^{J-1} \gamma_{ir}\left(\frac{n_r}{n}\right) + \delta_i z + \mu_i \qquad (7)$$

Here w_i is the father or mother's consumption share, z is a vector of controls, in our case dummy variables for whether mothers and fathers have greater than primary education, and J is the total number of age-gender cohorts. Age groups have the following age ranges: 0-3, 4-7, 8-11, 12-15, 16-25, 26-40, 41-60, and 61 and older.

The effects of an additional person of age-gender cohort r on consumption of parent i, can then be calculated as

[5] The CHNS does not provide individual food expenditure data. It also has only limited information on non-food expenditures.

$$\pi_{ir} = \frac{(\eta_i - \beta_i) + \gamma_{ir} - \sum_{1}^{J-1} \gamma_{ij}\left({n_j}/{n} \right)}{\beta_i + w_i}. \tag{8}$$

Following Deaton (1989), we report standard errors for π_{ir} using the delta method. The variance of the estimated $\hat{\pi}_{ir}$ is defined as follows:

$$Var(\hat{\pi}_{ir}) = \hat{\sigma}_i^2 J_{ir}^{'} (X'X)^{-1} J_{ir}. \tag{9}$$

Here, $\hat{\sigma}_i^2 = (N-k)^{-1} \hat{u}_i^{'} \hat{u}_i$, where N is the number of observations, k is the number of estimated parameters in (7), and $\hat{u}_i$ are the estimated residuals from (7). J_{ir} is the Jacobian matrix evaluated at the parameter estimates and sample means, and X contains the regressors in (7).

By calculating π_{ir} for each parent's nutrient consumption with respect to both boys and girls, we are able to test the following null hypotheses:

1. The consumption of fathers does not exhibit gender bias ($\pi_{fb}=\pi_{fg}$).
2. The consumption of mothers does not exhibit gender bias ($\pi_{mb}=\pi_{mg}$).
3. Sons affect the consumption of fathers as much as the consumption of mothers ($\pi_{fb}=\pi_{mb}$).
4. Daughters affect the consumption of fathers as much as the consumption of mothers ($\pi_{fg}=\pi_{mg}$).
5. The gender bias of fathers is the same as the gender bias of mothers ($\pi_{fb}-\pi_{fg}=\pi_{mb}-\pi_{mg}$)

If there exist data on the consumption of fathers and mothers from the same set of families, estimation of (7) is more efficient if the consumption share equations for fathers and mothers are estimated simultaneously using a seemingly unrelated regression (SUR) model. Then, the above hypotheses can be directly tested by imposing restrictions on the γ_{ir}. For example, to test hypothesis 1, whether the effect of boys (b) on the consumption of fathers (f) is the same as that of girls (g), one can conduct an F-test for the restriction $\gamma_{fb}=\gamma_{fg}$. Similarly, to test hypothesis 3, whether the effect of boys (b) on the consumption of fathers (f) differs from their effect on the consumption of mothers (m), one can impose the cross-equation restriction $\gamma_{fb}=\gamma_{mb}$ and conduct an F-test.[6]

Given panel data on fathers and mothers consumption, one can pool the data and identify (7) from the cross-sectional variation, or try to exploit the panel data to estimate how consumption shares in the same family adjust to changes in family composition. Unfortunately, given the availability of only two years of data separated by only two years, there are too few households with changes in family composition to identify a fixed effects model. We thus follow the former strategy and pool the data from the two years. We also estimate (7) allowing for clustering of the errors for observations from the same households over time.[7] It turns out that if we estimate (7) separately for each year (1991 and 1993), many of the key results are the same and statistically significant.

One might suspect that the outlay equivalence ratios and hypothesis test results may differ for different household types. For instance, rural and urban communities could differ systematically in their cultural norms or in typical types of adult and child work activity (e.g., manual farm labor in rural areas versus administrative or service work in urban areas). Such differences could systematically alter the nutrient demand of

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[6] The formula for the variance of $\hat{\pi}_{ir}$ does not include the standard errors of the estimated γ_{ir} so there is not a direct correspondence between the hypothesis test results and the reported standard errors of $\hat{\pi}_{ir}$.

[7] We compare the results for single-equation estimation of fathers and mothers consumption share with and without clustering. The standard errors of the coefficients on the demographic group indicators change very little and none of the test results for hypotheses 1 and 2 change when clustering is allowed. Reported results ignore clustering to facilitate SUR estimation of the two consumption share equations.

family members. Values and work activities might also vary with the educational attainment of parents. We thus conduct the analysis separately for rural and urban sub-samples, and for groups defined by the education level of fathers and mothers.

4. Data and Sample Description

We utilize data from the 1991 and 1993 waves of the China Health and Nutrition Survey (CHNS) administered by the Population Center at the University of North Carolina at Chapel Hill.[8] A unique feature of the CHNS is the collection of individual nutrient consumption data for each member of surveyed households based on intensive 3-day food intake surveys in which trained health personnel visited households daily to monitor food preparation, consumption, and wastage by conducting detailed 24-hour recall surveys of each individual, observing food preparation equipment and serving instruments, and measuring changes in food stocks. The 1991 and 1993 survey samples included both rural and urban households from seven provinces in different parts of the country: Guangxi, Henan, Hubei, Hunan, Jiangsu, Liaoning, and Shandong.[9]

To reduce bias from unobserved heterogeneity associated with differences in family composition, we restrict our attention to the sample of nuclear families with a father, mother, and children less than 16 years old. We exclude households with no children or with members aged 16 and older who are not the father or mother (e.g., adult

[8] The 1989 wave did not collect nutritional intake data for all household members, and contains numerous missing values. Nutritional intake data from subsequent waves has not been released publicly.
[9] Four counties in each province were selected randomly. The sampling units consisted of 32 urban neighborhoods, 30 suburban neighborhoods, 32 towns, and 96 villages.

children, grandparents).[10] This leaves us with a pooled sample of 1737 observations, including 1310 rural household observations and 427 urban household observations.

Table 1 describes the sibling composition of sample households. Among sample households, 40.93 percent have one child, 44.68 percent have two children, and 14.39 percent have three or more children. There are more one-child households with sons (23.20 percent) than daughters (17.73 percent) and more two-son, no-daughter households (9.90 percent) than two-daughter, no-son households (7.43 percent). This suggests that households are more likely to stop having children if they have a son, but not necessarily that sons are favored within households.[11] Overall, 51.69 percent of children are boys, suggesting that there is not strong selection bias due to selective abortion or higher mortality rates among girls.

Table 2 displays summary statistics on the average daily nutrient consumption by gender for young children (age 7 and younger), older children (age 8 to 15), and parents (age 16 to 60). For all nutrients and demographic groups (except fat for children age seven and below), mean male consumption is higher than mean female consumption. The male to female ratio of nutrient consumption increases with age in both urban and rural areas and for all three nutrients. The male-female differences for all nutrients and in both rural and urban areas are statistically significantly different from zero for older children (age 8-15), and adults (age 16 to 60), but not for young children (age 0-7). The adult male to female consumption ratio is higher for calories than for protein and fat, especially in urban areas. Children's consumption is higher in urban areas irrespective of

[10] Most of the main results are robust to including all households, or households with at least one child below age 16 (allowing other adult family members).
[11] Differential stopping behavior could lead to average welfare differences between boys and girls if girls tend to be in larger households with fewer resources per capita (Jensen, 2002).

nutrient (calories, protein, and fat), gender, or age (younger versus older children), while adult consumption is higher in rural areas, likely reflecting greater physical exertion in work.

5. Results

The paper's main results are presented in Tables 3A, 3B, and 3C, which report estimated outlay equivalence ratios and hypothesis test results for calories, protein, and fat, respectively. Each table reports results separately for the rural and urban samples. We divide boys and girls into four age categories: 0-3, 4-7, 8-11, and 12-15, enabling us to test separately for gender bias in food allocations to children in different age groups. Reassuringly, we find that all outlay equivalence ratios are negative, and that they tend to increase with children's age as would be expected if older children consume more. Urban outlay equivalence ratios tend to be smaller in magnitude than rural outlay equivalence ratios.

Looking first at the outlay equivalence ratios for calorie consumption (Table 3A), we examine our five hypotheses about gender bias. First, we find strong evidence of gender bias in fathers' consumption, especially among rural households. For the rural sample, the difference in fathers' outlay equivalence ratios with respect to boys versus girls increases with child age. The effect on consumption shares of an additional boy is greater than that of an additional girl for all age groups except the youngest age group, and the difference is statistically significantly different than zero at the 99 percent confidence level. The change in fathers' calorie consumption when the household has an additional boy aged 4 to 7 is equivalent to the change associated with an income reduction of 73.7 percent, compared to 63.5 percent for a similarly aged girl, a difference of 10.1 percent. The same income percent difference is 9.7 percent for boys and girls aged 8 to 11, and 15.7 percent for those aged 12 to 15. For urban households, the fathers' outlay equivalence ratio is greater in magnitude for boys than girls for three of the four age groups and is statistically significant at the 95 percent confidence level for children aged 8 to 11 (and nearly significant at the 90 percent level for the 12 to 15 age group). The difference in outlay equivalence ratios for boys and girls aged 8 to 11 is similar in magnitude to the rural sample (0.115) and is somewhat smaller for children aged 12 to 15 (0.104).

In contrast, we find no strong evidence of gender bias in mothers' calorie consumption for the rural sample. Interestingly, we do find evidence that, similar to urban fathers, urban mothers favor boys in the older child age groups. With regards to hypotheses 3 and 4, we do not find differences in fathers' and mothers' responses to children to be statistically significantly different than zero. However, for the rural sample, the estimated magnitudes of fathers and mothers outlay equivalence ratios appear similar for boys but differ systematically for girls. This provides suggestive evidence that parents sacrifice calorie consumption equally for boys but that fathers sacrifice less for girls.

The most interesting results are for hypothesis 5. We find strong evidence in the rural sample that the gender bias in fathers' calorie consumption differs from the gender bias in mothers' calorie consumption (statistically significant at the 95 percent confidence level for children age 4-7 and at the 90 percent confidence level for the two older age groups). As noted earlier, this result is inconsistent with the unitary model and simple explanations of gender bias associated with the different needs of boys and girls, or equity bias.

Next, we turn to the results for protein and fat. For protein, the pattern of results are very similar to that for calories. The consumption of rural fathers exhibits gender bias, that of rural mothers does not, and the difference in gender bias is statistically significant at the 95 percent confidence level for the two older child age groups. The magnitude of rural father's gender bias toward older children (age 8-15), measured by the difference in outlay equivalence ratios for boys and girls, is greater than for calories. There also is evidence that urban fathers and mothers both exhibit gender bias for older children, but that differences in gender bias are not statistically significantly different than zero.[12] Although still not statistically significant at conventional confidence levels, the differences between fathers and mothers consumption reductions in response to additional children is larger for protein than for calories (in both urban and rural areas). In other words, mothers sacrifice even more protein than calories for children relative to fathers. For girls in rural areas, these differences are nearly statistically significant.

For fat, the results for the most part are consistent with the earlier findings. Focusing on the hypothesis tests, for the rural sample, fathers exhibit gender bias, but now only for the two older age groups, while mothers do not for any group, and the difference in gender bias is now only statistically significantly different from zero for the oldest age group. For children age 12 to 15, father's gender bias is even larger in magnitude than for calories or protein. In the urban sample, only mothers exhibit gender bias, and only for older children; differences in gender bias are not significantly different

[12] In contrast to the results for calories, there is evidence that fathers' and mothers' gender bias differs for children age 0-3 (statistically significant at the 90 percent confidence level).

than zero. For the rural sample, mothers' fat consumption is much more responsive to the presence of additional children than fathers' fat consumption. The difference in magnitudes of the outlay equivalence ratios for mothers and fathers ranges from 0.288 and 0.477 and are statistically significant at the 95 percent confidence level for both boys and girls and across age groups.

To investigate whether gender bias is related to the education of parents, we report outlay equivalence ratios and hypothesis test results for calorie consumption in four sub-samples of the rural sample: households with fathers with more than primary education, households with fathers with primary or less education, households with mothers with more than primary education, and households with mothers with primary or less education (Table 4).[13] The main finding is that gender bias in fathers' consumption is much greater in magnitude among fathers with low levels of education than among those who are better-educated. Differences in outlay equivalence ratios for boys and girls aged 4 to 7, 8 to 11, and 12 to 15, are 0.195, 0.165, and 0.173 for poorly educated rural fathers, compared to 0.101, 0.097, and 0.157 for all rural fathers. For mothers, in contrast, education level does not systematically affect the extent of gender bias. The exact same pattern of results are found for protein and fat (not reported), with the differences associated with fathers' education even more pronounced for fat.

6. Conclusion

In this paper, we exploit the availability of individual nutrient intake data from China to examine whether fathers and mothers exhibit gender bias in the intrahousehold allocation of nutrients by testing for the effects of household composition on fathers' and mothers' nutrient consumption. Our main empirical finding is that in rural areas, fathers exhibit significant gender bias while mothers do not, and that the difference in gender bias of fathers versus mothers is statistically significant. Fathers' gender bias is not statistically significant for infants (age 0 to 3) but increases with age. The pattern is similar whether one looks at calories, proteins, or fats. This is consistent with studies cited earlier that find that fathers and mothers allocate resources differently. We also find that poorly educated rural fathers exhibit much greater gender bias than better educated rural fathers. In urban areas, there is less evidence of gender bias, and no evidence that the gender bias of fathers and mothers is different. Both fathers and mothers exhibit gender bias in the allocation of proteins and fats to older children.

By explicitly modeling how gender bias can result from equity bias, efficiency bias, and preference bias in unitary and sharing rule models of the household, we show how our empirical results can help distinguish among different potential explanations of gender bias. Differences in the gender bias of fathers and mothers are inconsistent with the unitary household model or equity bias explanations. They are consistent with explanations in which fathers benefit differentially from boys and girls, while mothers do not, for example if there is efficiency bias and fathers control child income. They are also consistent with different preferences of fathers and mothers. Our findings suggest that to reduce gender bias, policies that focus on changing parental gender attitudes should focus on fathers, and that income-generating programs or empowerment efforts should focus on mothers.

[13] The urban sub-sample was too small to disaggregate in this way.

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Table 1: Son and Daughter Composition of Sample Households (in percent)

	Number of Sons				
Number of Daughters	0	1	2	3	Total
0	0.00	23.20	9.90	1.04	34.14
1	17.73	27.35	4.49	0.12	49.68
2	7.43	5.58	0.75	0.00	13.76
3	1.15	0.58	0.17	0.00	1.90
4	0.06	0.23	0.06	0.00	0.35
5	0.17	0.00	0.00	0.00	0.17
Total	26.54	56.94	15.37	1.15%	100.00

Table 2: Average Daily Nutrient Intakes of Sample Household Members

	Rural			Urban		
Ages	0~7	8~15	16~60	0~7	8~15	16~60
Males						
N	645	634	1308	155	150	427
Calories	1471	2161	2974	1576	2237	2736
	(541)	(602)	(749)	(494)	(586)	(627)
Protein (grams)	44.6	63.9	88.1	52.0	72.3	87.6
	(18.6)	(21.6)	(27.1)	(17.9)	(23.1)	(23.6)
Fat (grams)	36.1	50.8	69.9	50.7	69.7	85.5
	(23.3)	(29.1)	(36.4)	(27.3)	(34.8)	(36.9)
Females						
N	596	593	1310	94	197	427
Calories	1440	2007	2571	1545	2019	2336
	(492)	(499)	(591)	(452)	(515)	(562)
Protein (grams)	43.6	59.8	76.5	49.5	65.4	75.4
	(17.0)	(18.9)	(22.3)	(14.9)	(18.3)	(20.7)
Fat (grams)	36.7	47.0	60.6	52.6	62.3	74.0
	(23.2)	(26.4)	(30.6)	(27.3)	(30.4)	(31.0)
Male/Female Ratio						
Calories	1.022	1.077**	1.157**	1.020	1.108**	1.171**
Protein	1.021	1.068**	1.152**	1.051	1.105**	1.161**
Fat	0.982	1.080*	1.153**	0.965	1.119*	1.155**

Note: * and ** denote that means for males and females are statistically different at the 95 and 99 percent confidence levels based on 2-sided t-tests that allow for unequal variances by group.

Table 3A: Outlay Equivalence Ratios and Gender Bias Tests for Calorie Consumption of Fathers and Mothers

	Rural (N=1310)				Urban (N=427)			
Ages	0~3	4~7	8~11	12~15	0~3	4~7	8~11	12~15
Outlay Equivalence Ratios (standard errors in parentheses)								
π_{fb}	-0.533	-0.737	-0.914	-1.074	-0.381	-0.551	-0.802	-0.962
	(0.065)	(0.058)	(0.059)	(0.065)	(0.135)	(0.118)	(0.121)	(0.123)
π_{fg}	-0.461	-0.635	-0.818	-0.917	-0.274	-0.570	-0.687	-0.857
	(0.065)	(0.059)	(0.062)	(0.065)	(0.137)	(0.121)	(0.122)	(0.120)
π_{mb}	-0.552	-0.774	-0.961	-1.060	-0.411	-0.646	-0.949	-0.858
	(0.062)	(0.056)	(0.056)	(0.062)	(0.169)	(0.154)	(0.157)	(0.161)
π_{mg}	-0.557	-0.762	-0.944	-1.012	-0.461	-0.612	-0.733	-0.817
	(0.062)	(0.056)	(0.059)	(0.062)	(0.177)	(0.161)	(0.159)	(0.156)
Gender Bias Tests (p-values in parentheses*)								
$\|\pi_{fb}\|-\|\pi_{fg}\|$	0.072	0.101	0.097	0.157	0.108	-0.019	0.115	0.104
	(0.175)	(0.003)	(0.010)	(0.002)	(0.261)	(0.748)	(0.032)	(0.110)
$\|\pi_{mb}\|-\|\pi_{mg}\|$	-0.005	0.012	0.017	0.048	-0.050	0.035	0.215	0.041
	(0.917)	(0.706)	(0.632)	(0.323)	(0.193)	(0.604)	(0.049)	(0.045)
$\|\pi_{fb}\|-\|\pi_{mb}\|$	-0.019	-0.037	-0.047	0.014	-0.030	-0.095	-0.147	0.104
	(0.441)	(0.466)	(0.497)	(0.621)	(0.285)	(0.263)	(0.281)	(0.319)
$\|\pi_{fg}\|-\|\pi_{mg}\|$	-0.096	-0.126	-0.127	-0.094	-0.188	-0.041	-0.046	0.041
	(0.344)	(0.349)	(0.391)	(0.449)	(0.214)	(0.278)	(0.275)	(0.324)
$(\|\pi_{fb}\|-\|\pi_{fg}\|)-(\|\pi_{mb}\|-\|\pi_{mg}\|)$	0.077	0.089	0.080	0.108	0.158	-0.054	-0.100	0.063
	(0.277)	(0.045)	(0.097)	(0.086)	(0.119)	(0.593)	(0.770)	(0.916)

Data source: 1991 and 1993 waves of the China Health and Nutrition Survey
*Reported p-values are from tests of restrictions on the coefficients of boy and girl demographic groups in the fathers' and mothers' consumption share estimation models.

Table 3B: Outlay Equivalence Ratios and Gender Bias Tests for Protein Consumption of Fathers and Mothers

	Rural (N=1310)				Urban (N=427)			
Ages	0~3	4~7	8~11	12~15	0~3	4~7	8~11	12~15
Outlay Equivalence Ratios (standard errors in parentheses)								
π_{fb}	-0.506	-0.716	-0.896	-1.061	-0.389	-0.527	-0.766	-0.932
	(0.071)	(0.063)	(0.064)	(0.071)	(0.153)	(0.133)	(0.137)	(0.139)
π_{fg}	-0.426	-0.624	-0.765	-0.864	-0.228	-0.552	-0.665	-0.825
	(0.071)	(0.064)	(0.067)	(0.071)	(0.155)	(0.137)	(0.138)	(0.136)
π_{mb}	-0.648	-0.826	-1.001	-1.092	-0.498	-0.798	-1.055	-1.045
	(0.065)	(0.058)	(0.059)	(0.065)	(0.156)	(0.137)	(0.141)	(0.142)
π_{mg}	-0.635	-0.808	-0.993	-1.043	-0.646	-0.753	-0.905	-0.920
	(0.065)	(0.059)	(0.062)	(0.066)	(0.158)	(0.141)	(0.141)	(0.139)
Gender Bias Tests (p-values in parentheses*)								
$\|\pi_{fb}\|-\|\pi_{fg}\|$	0.080	0.092	0.131	0.197	0.161	-0.025	0.101	0.106
	(0.162)	(0.013)	(0.001)	(0.000)	(0.137)	(0.703)	(0.094)	(0.149)
$\|\pi_{mb}\|-\|\pi_{mg}\|$	0.013	0.018	0.007	0.049	-0.148	0.044	0.150	0.125
	(0.807)	(0.592)	(0.845)	(0.340)	(0.179)	(0.514)	(0.015)	(0.097)
$\|\pi_{fb}\|-\|\pi_{mb}\|$	-0.141	-0.110	-0.105	-0.031	-0.108	-0.271	-0.289	-0.113
	(0.150)	(0.186)	(0.208)	(0.286)	(0.321)	(0.280)	(0.285)	(0.340)
$\|\pi_{fg}\|-\|\pi_{mg}\|$	-0.209	-0.184	-0.229	-0.179	-0.418	-0.201	-0.240	-0.095
	(0.113)	(0.139)	(0.139)	(0.167)	(0.235)	(0.299)	(0.296)	(0.341)
$(\|\pi_{fb}\|-\|\pi_{fg}\|)-(\|\pi_{mb}\|-\|\pi_{mg}\|)$	0.067	0.074	0.123	0.148	0.309	-0.070	-0.049	-0.018
	(0.361)	(0.120)	(0.021)	(0.035)	(0.077)	(0.526)	(0.758)	(0.993)

Data source: 1991 and 1993 waves of the China Health and Nutrition Survey

*Reported p-values are from tests of restrictions on the coefficients of boy and girl demographic groups in the fathers' and mothers' consumption share estimation models.

Table 3C: Outlay Equivalence Ratios and Gender Bias Tests for Fat Consumption of Fathers and Mothers

	Rural (N=1310)				Urban (N=427)			
Ages	0~3	4~7	8~11	12~15	0~3	4~7	8~11	12~15
Outlay Equivalence Ratios (standard errors in parentheses)								
π_{fb}	-0.567	-0.701	-0.878	-1.036	-0.432	-0.631	-0.765	-0.884
	(0.096)	(0.086)	(0.087)	(0.096)	(0.190)	(0.165)	(0.170)	(0.172)
π_{fg}	-0.453	-0.667	-0.772	-0.795	-0.343	-0.714	-0.708	-0.827
	(0.095)	(0.087)	(0.091)	(0.096)	(0.192)	(0.170)	(0.171)	(0.168)
π_{mb}	-0.871	-1.070	-1.179	-1.324	-0.411	-0.646	-0.949	-0.858
	(0.087)	(0.079)	(0.079)	(0.087)	(0.169)	(0.154)	(0.157)	(0.161)
π_{mg}	-0.813	-1.014	-1.182	-1.272	-0.461	-0.612	-0.733	-0.817
	(0.087)	(0.079)	(0.083)	(0.087)	(0.177)	(0.161)	(0.159)	(0.156)
Gender Bias Tests (p-values in parentheses*)								
$\|\pi_{fb}\|-\|\pi_{fg}\|$	0.114	0.034	0.107	0.241	0.089	-0.084	0.057	0.057
	(0.139)	(0.496)	(0.052)	(0.001)	(0.507)	(0.312)	(0.447)	(0.533)
$\|\pi_{mb}\|-\|\pi_{mg}\|$	0.057	0.057	-0.003	0.052	-0.050	0.035	0.215	0.041
	(0.411)	(0.210)	(0.948)	(0.440)	(0.677)	(0.285)	(0.047)	(0.051)
$\|\pi_{fb}\|-\|\pi_{mb}\|$	-0.304	-0.369	-0.300	-0.288	0.021	-0.015	-0.183	0.026
	(0.048)	(0.041)	(0.055)	(0.064)	(0.326)	(0.314)	(0.301)	(0.325)
$\|\pi_{fg}\|-\|\pi_{mg}\|$	-0.361	-0.347	-0.410	-0.477	-0.118	0.103	-0.025	0.010
	(0.038)	(0.044)	(0.040)	(0.033)	(0.289)	(0.358)	(0.322)	(0.354)
$(\|\pi_{fb}\|-\|\pi_{fg}\|)-(\|\pi_{mb}\|-\|\pi_{mg}\|)$	0.057	-0.023	0.110	0.189	0.139	-0.118	-0.158	0.016
	(0.532)	(0.837)	(0.143)	(0.054)	(0.495)	(0.195)	(0.503)	(0.460)

Data source: 1991 and 1993 waves of the China Health and Nutrition Survey
*Reported p-values are from tests of restrictions on the coefficients of boy and girl demographic groups in the fathers' and mothers' consumption share estimation models.

Table 4: Outlay Equivalence Ratios and Gender Bias Tests for Calorie Consumption by Rural Parents, by Parents' Education Level

Ages	0~3	4~7	8~11	12~15	0~3	4~7	8~11	12~15
	Father's Education Low (N=404)				Father's Education High (N=906)			
Outlay Equivalence Ratios								
π_{fb}	-0.393	-0.691	-0.876	-0.944	-0.592	-0.765	-0.922	-1.158
π_{fg}	-0.388	-0.495	-0.711	-0.771	-0.510	-0.694	-0.866	-1.014
π_{mb}	-0.610	-0.883	-1.045	-1.132	-0.537	-0.745	-0.934	-1.043
π_{mg}	-0.491	-0.855	-0.947	-1.075	-0.564	-0.740	-0.957	-0.990
Gender Bias Tests (p-value in parentheses*)								
$\|\pi_{fb}\|-\|\pi_{fg}\|$	0.005	0.195	0.165	0.173	-0.026	0.005	-0.023	0.052
	(0.965)	(0.009)	(0.011)	(0.028)	(0.161)	(0.062)	(0.216)	(0.033)
$\|\pi_{mb}\|-\|\pi_{mg}\|$	0.119	0.028	0.098	0.057	-0.026	0.005	-0.023	0.052
	(0.327)	(0.698)	(0.118)	(0.456)	(0.640)	(0.899)	(0.593)	(0.418)
$\|\pi_{fb}\|-\|\pi_{mb}\|$	-0.217	-0.192	-0.169	-0.188	0.054	0.021	-0.011	0.116
	(0.184)	(0.227)	(0.258)	(0.259)	(0.767)	(0.764)	(0.761)	(0.956)
$\|\pi_{fg}\|-\|\pi_{mg}\|$	-0.103	-0.360	-0.236	-0.304	-0.054	-0.046	-0.091	0.023
	(0.235)	(0.145)	(0.214)	(0.190)	(0.635)	(0.678)	(0.667)	(0.821)
$(\|\pi_{fb}\|-\|\pi_{fg}\|)-$	-0.114	0.167	0.067	0.117	0.108	0.066	0.079	0.092
$(\|\pi_{mb}\|-\|\pi_{mg}\|)$	(0.563)	(0.058)	(0.272)	(0.181)	(0.194)	(0.198)	(0.223)	(0.278)
	Mother's Education Low (N=653)				Mother's Education High (N=657)			
Outlay Equivalence Ratios								
π_{fb}	-0.541	-0.794	-0.980	-1.122	-0.574	-0.735	-0.892	-1.059
π_{fg}	-0.473	-0.637	-0.901	-0.982	-0.495	-0.666	-0.786	-0.892
π_{mb}	-0.591	-0.797	-0.981	-1.092	-0.575	-0.816	-1.015	-1.089
π_{mg}	-0.578	-0.789	-0.969	-1.045	-0.612	-0.810	-0.998	-1.047
Gender Bias Tests (p-value in parentheses*)								
$\|\pi_{fb}\|-\|\pi_{fg}\|$	0.067	0.157	0.079	0.140	0.080	0.069	0.107	0.168
	(0.404)	(0.006)	(0.137)	(0.031)	(0.248)	(0.097)	(0.041)	(0.047)
$\|\pi_{mb}\|-\|\pi_{mg}\|$	0.013	0.007	0.012	0.047	-0.037	0.006	0.017	0.042
	(0.870)	(0.894)	(0.816)	(0.448)	(0.572)	(0.871)	(0.731)	(0.607)
$\|\pi_{fb}\|-\|\pi_{mb}\|$	-0.050	-0.003	-0.001	0.030	0.000	-0.081	-0.122	-0.029
	(0.578)	(0.700)	(0.753)	(0.834)	(0.362)	(0.332)	(0.328)	(0.397)
$\|\pi_{fg}\|-\|\pi_{mg}\|$	-0.105	-0.152	-0.068	-0.063	-0.117	-0.144	-0.212	-0.155
	(0.502)	(0.493)	(0.655)	(0.682)	(0.293)	(0.295)	(0.278)	(0.317)
$(\|\pi_{fb}\|-\|\pi_{fg}\|)-$	0.055	0.150	0.067	0.093	0.116	0.063	0.089	0.126
$(\|\pi_{mb}\|-\|\pi_{mg}\|)$	(0.569)	(0.034)	(0.293)	(0.206)	(0.254)	(0.277)	(0.209)	(0.265)

Data source: 1991 and 1993 waves of the China Health and Nutrition Survey

*Reported p-values are from tests of restrictions on the coefficients of boy and girl demographic groups in the fathers' and mothers' consumption share estimation models (equation xx).

Appendix
Derivation of Relative Consumption of Mother and Father in Sharing Rule Model

Following our solution to the unitary model, we assume individual log utility, fixed marginal returns to individual consumption (a_i), and a sharing rule with fixed proportions of each individual's income going to the mother (θ_i). We can rewrite the father's optimization problem in equation (4) as follows:

$$\underset{X_f, X_{fc}}{Max}\, U(X_f, X_{fc}) = u_f \log X_f + \beta_{fc} u_c \log(X_{fc} + X_{mc}) \qquad \text{(A-1)}$$

$$\text{s.t.} \quad X_f + X_{fc} = (1-\theta_f) a_f X_f + (1-\theta_m) a_m X_m + (1-\theta_c) a_c (X_{fc} + X_{mc})$$

If we set up this problem as a Lagrangian and take first order conditions with respect to X_f and X_{fc}, we can derive the following expression for the optimal X_f:

$$X_f^* = \frac{u_f (1 - a_c (1-\theta_c))}{\beta_{fc} u_c (1 - a_f (1-\theta_f))} (X_{fc}^* + X_{mc}^*) \qquad \text{(A-2)}$$

Similarly, we can write the mother's optimization problem as follows:

$$\underset{X_m, X_{mc}}{Max}\, U(X_m, X_{mc}) = u_m \log X_m + \beta_{mc} u_c \log(X_{fc} + X_{mc}) \qquad \text{(A-3)}$$

$$\text{s.t.} \quad X_m + X_{mc} = \theta_f a_f X_f + \theta_m a_m X_m + \theta_c a_c (X_{fc} + X_{mc})$$

Solving for optimal X_m:

$$X_m^* = \frac{u_m (1 - a_c \theta_c)}{\beta_{mc} u_c (1 - a_m \theta_m)} (X_{fc}^* + X_{mc}^*) \qquad \text{(A-4)}$$

Equation (5) in the text follows directly from the expressions (A-2) and (A-4).