Journal of Applied Economics, Vol. III, No. 2 (Nov 2000), 353-386

DOES GENDER AND BIRTH ORDER MATTER WHEN PARENTS SPECIALIZE IN CHILD'S NUTRITION? EVIDENCE FROM CHILE

LUIS RUBALCAVA*

Centro de Investigación y Docencia Económicas

and

DANTE CONTRERAS**

Universidad de Chile

Using household survey data from Chile the current paper presents evidence of how the nutritional status of the child reflects differences in parental preferences and child rearing technology within an intra-household allocation approach that includes a health production function. From the household optimization problem we estimate the nutritional status of the child conditional on a set of child, family and community covariates that reflect parental preferences and parental child rearing technology. We test directly whether birth-order in the family and whether being a son or being a daughter reflect how parents allocate the resources, given that the Chilean family is often linked to a machismo sentiment in the division of household chores. Logit estimates of the nutritional status of the child show

---

* Luis Rubalcava is current professor in the Department of Economics at the *Centro de Investigación y Docencia Económicas, A.C* . in Mexico City. E-mail: lrubalca@dis1.cide.mx.

** Dante Contreras is current professor in the Department of Economics at the *Universidad de Chile,* in Santiago de Chile. E-mail: dcontrer@decon.facea.uchile.cl.

We wish to thank Janet Currie, Andrew Davenport, Dean Hyslop, Arnold Harberger, Arleen Leibowitz, Emundo Murrugara, Keunkwan Ryu, Wei-Yin Hu, Graciela Teruel and two anonymous referees. The paper was also benefited from comments received at the Population Association of America Meetings, DC (1997), the Labor and Population Raw Seminar at UCLA, and from the Center for the Study of Women Congress, UCLA (1997). We specially wish to thank Duncan Thomas for his extraordinarily helpful comments and support along the course of this study. Luis Rubalcava is thankful to the Tinker Research Grand (Center for Latin American Studies, UCLA). All remaining errors are ours.

gender specialization on child rearing: mothers give more resources to their daughters and fathers to their sons. This gender polarity is significant for non-oldest daughters and non-oldest sons, reflecting perhaps infant-order experience in child-care specialization. We also find that father's education is less important than mother's education. Nevertheless, mothers with higher education levels than their spouse seem to assign less family resources to their children than those who are relatively less educated.

JEL classification codes: I12, D1
Key words: child's nutrition, intrahousehold allocation of resources

I. Introduction

Parental decisions have a profound effect on a child's human capital development. Given the family's endowment, the way parents decide how to allocate household resources has a direct impact on the child's health and education. These decisions, in turn, may affect not only the productivity of the children once they have grown up, but also impact their life expectancy. It is in this context that the present paper emphasizes the impact of family resources and parental preferences on the provision of child health within the household.

We explore child nutritional status and parental resources within households in Chile. Unlike the traditional family literature that conceives the household as a single decision-maker, we adopt an intrahousehold allocation approach, relax the unitary preferences assumption, and introduce a health production function to disentangle how parental preferences and differences in parental child-rearing technology affect the nutritional status of the child. In particular, we test whether there is any gender or birth-order differentiation by parents that could be captured through the nutritional status of the child, conditional on each parent's characteristics. The gender and birth-order analysis is based on the machismo sentiment in both sexes that is often encountered in the Chilean family (Raczynski and Serrano, 1986). In addition, the birth-order hypothesis allows one to capture any parental apprenticeship in child rearing or differences in predilection among children according to birth order. Parents may gain more experience in taking care of latter children. Likewise, older children in the household could be seen differently by their parents and consequently be treated dissimilarly relative to their younger siblings.

From the family maximization problem, we derive the household's demand for child health, conditional on a set of child, family and community covariates that reflect parental preferences and child-rearing technology in the allocation of health among children. We employ the Chilean Household Survey Casen 92. The Chilean National Ministry of Planning conducted this national, cross-sectional survey.

The second section of the paper gives a brief overview of the intrahousehold model literature. Part III defines the theoretical model. Section IV describes the data. Part V outlines the empirical strategy and presents the results. Conclusions are found at the end of the paper.

II. Household Behavior Models

Traditionally, household decisions have been modeled on the assumption that the household is a single decision-maker unit that maximizes a sole utility function subject to a set of constraints dictated by the household budget and available technology (Becker, 1964). These models assume either that all household members have common preferences, or that there is one household member, a dictator or representative agent, who determines the allocations of all household members in either an altruistic or selfish manner. Under this assumption the important factors for the family maximization problem are household aggregates rather than individual resources, where the optimal demands depend on aggregated household resources and not on each family member's income. This is *called the income pooling hypothesis*, and can be tested empirically.

While the common preferences approach has been shown to be a useful framework in many circumstances, taking the unitary representation of the household as a benchmark is sometimes questionable. From the theoretical point of view, individualism lies at the foundation of microeconomic theory. The fact that individualism requires one to allow that different individuals may have different preferences poses questions as to the way household behavior should be modeled. Should we rely on common preference analysis when characterizing family behavior when designing targeting public policy?[1] How far wrong can one go by simply presuming that intrahousehold inequality does not exist, when in fact it does? Are similar individuals treated differently in the allocation of resources within the family? If the family itself allows disparities among its members, should that be an issue of assessment when analyzing household demands and designing public policies?[2] These and other questions challenge the common preference model, especially when this framework fails to explain intrahousehold allocation decisions in terms of differences in tastes and bargaining power among family members.

Several empirical analyses have highlighted these issues. Sen (1984) summarizes a number of studies, which claim that girls are less favored than

[1] The impact of a direct subsidy to targeted households may differ depending on who receives the check and the relative bargaining power of the family member receiving the allotment. Lundberg, Pollak and Wales (1997) utilize a natural experiment framework provided by a shift in the United Kingdom welfare system in the late 1970s to test the pooling hypothesis upon household behavior. Prior to 1977 child allowance was directed to the father, however over the period 1977 through 1979 the program was replaced by a check paid to the mother. This represented a substantial redistribution of income in terms of the mother's bargaining power. The authors document an increase in spending on women's and children's as opposed to men's followed after the policy change.

[2] Sen (1984) claims that characterizing the family as a single decision-making unit could lead to misleading conclusions when evaluating standards of living on the basis of market data if disparities within the family are not taken into account. Market demands could well reflect the relative importance of different items as seen by the decision markers and not necessarily from each family member welfare.

boys in terms of food division within the household for the case of India.[3] Haddad and Kanbur (1990) have analyzed the consequences of ignoring intrahousehold inequality within a targeted public policy framework. Schultz (1990) finds that, in Thailand, resources in the hands of women reduce fertility more than resources held by men, and that the impact of non-labor income on labor supply outcomes depends on who in the household controls that income. Thomas (1990) reports that child health (survival probabilities, height for age and weight for height) along with household nutrient intakes tend to rise more, if additional (non-labor) income is in the hand of women rather than men. Using the same data, Thomas (1993a) reports that income in the hands of women is associated with increases in the share of the household budget spent on health, education and housing. Evidence in India indicates that children are more likely to attend school and receive medical attention if the mother has more assets (Duraisamy, 1992; Duraisamy and Malathy, 1991). Quisumbing (1994) finds that among Philippine households, bequests to children are a function of resource control, with son inheriting more land than daughters when fathers have more education or when the household has more land. This literature is still recent, but the results overwhelmingly suggest that the effect of resources on welfare of individuals within the household may depend on who control the resources.

Recognizing that the assumptions of the unitary model of the household do not necessarily hold, several innovative and theoretically more appealing models have been proposed in recent years. These models characterize the intrahousehold allocation mechanism while relaxing the common preference assumption. Some models suggest that household allocation decisions may be characterized as the result of a *bargaining* process, where each member seeks to allocate family resources over which she or he has control in relation

[3] Undernutrition among children within the family is shown to be higher for females than males in West Bengal. Male to female calorie and protein intake ratios are marked by male predominance among young children in Bangladesh. In addition, female morbidity relative to male's is found to be greater within families of Greater Calcutta.

to her (his) preferences and bargaining power. Horney and McElroy (1981) have proposed a Nash equilibrium framework, in which each family member cooperates with each other in order to raise her (his) individual utility level above a certain threshold point. Bourguignon and Chiappori (1992); Bourguignon and Chiappori (1994);[4] and Browning, Bourguignon, Chiappori and Lechene (1994), have proposed a more general (collective) framework, where the common preference model may be tested. In this more general *collective model*, the only assumption made is the one of Pareto efficiency. The household behavior mechanism is viewed as an example of a repeated "game" where each person knows the preferences of the other members in the household. This symmetry of information and the fact that the game is repeated, supports the idea that agents find mechanism to allocate efficient outcomes. Within this setting, the household utility function is comprised of a weighted sum of family member's utility functions, which in turn are maximized subject to total household resources. The weighting scheme that leads to the Pareto efficient allocation is provided by a sharing rule or welfare weights, which are function of each household member's individual bargaining variables. Those individuals with more bargaining power within the family get bigger weights to their utility function, making the household welfare function to reflect more their tastes. The resulting household demand functions are sensitive to the welfare weights, and consequently to the variables that reflect these weights and the individuals' bargaining power within the family. In this framework, if individual incomes are taken as proxies for bargaining

[4] Focusing on price-variation analysis, the authors show that under the collective household modeling, household demands need not satisfy the usual symmetry condition on the Slutsky matrix that the individual theory of demand (and common preference model) predicts. The Slutsky matrix is found to be equal to a symmetry matrix plus a rank-one matrix. The authors proceed to test this main result on single-head households and married-couple households. The Slutsky symmetry condition is not rejected for singles but it is for couples. Finally, the derived predictions of the collective setting are not rejected on couples' data, which according to the authors, provides support for the collective model as a viable alternative to the unitary model.

power measures, and members in the family do not share the same preferences, the optimal household demands are functions of household *individual* incomes and not of simply total household pooled income as suggested by the common preference model. This is called the pooling alternative hypothesis and it can be tested empirically by examining (i.e., in a regression framework) the impact of individual incomes on household demands. If individual incomes have significantly different effects on the analyzed household demand(s), this rejects the income pooling hypothesis and the common preference model.[5]

III. The Model

In this section we propose a general model within the intrahousehold allocation framework, that provides an empirical test of how the nutritional status of children may reflect differences in parental preferences and child rearing technology in the provision of health, conditional on parent's age, education and income, and on the characteristics of the child. In particular, we will focus our attention on whether there is parental differentiation in terms of child's birth-order and gender within households in Chile. It becomes important to incorporate gender and birth-order differentiation in the model for several reasons. Parents may rely on their oldest son to look after them when old, while daughters assist their husband's families. In such cases one might expect to see parents invest more in a son rather than in a daughter. The participation of girls and boys in different household tasks can be another

[5] One must be careful when making this statement, since the reverse logic does not apply: If the income coefficients are not different from each other, one cannot conclude that the family behaves as a single decision unit. Other variables -unobservable to the econometrician- may be the true bargaining factors. We will address this point in the paper by testing the common preference assumption focusing on *differences* in parental education attainment as bargaining factors, as well as parental individual nonlabor income. Conditional on education, age and individual current incomes, differences in education may modify the way each parent sees her (his) partner's long term income profile and consequently, reflect some bargaining power in the family.

reason for cost differentiation and human capital formation. Unlike their brothers, young girls may not go to school in order to provide child care support.[6] Likewise, child's human capital investments might differ by birth order, either because of biological factors or due to behavioral influences. Parents may learn from the experience of their older children and be more efficient at raising later ones. Finally, there might be differences in resource allocations simply because of tastes, which in turn may reflect social and cultural norms. These differences are relevant in the Chilean society, where the division of household tasks and probably child rearing itself, are influenced by a machismo sentiment in both sexes (Gissi J., 1984). Household chores are mainly described as a mother's issue (Raczynski and Serrano, 1986; Aranda, 1986).

Following Bourguignon and Chiappori (1992) we relax the common preference assumption and allow the household family welfare (W) to be a weighted function of each parent, mother's and father's, utility (U^m and U^f, respectively). The weighting rule (Ω) allows us to capture the influence of bargaining factors within the family that may reflect the negotiation process in the allocation of resources. Among these power variables, we focus our attention on mother's and father's individual income to test the common preference model via the pooling hypothesis. Conditional on parental age and education levels differences in parental education are also considered as bargaining variables in our model. The idea behind this is that differences in parental levels of education may reflect dissimilarities in the way individuals see their potential long-term incomes, and act as bargaining variables. Both variables, individual incomes and differences in parental education, enter in our model as bargaining factors through the welfare weight (Ω). As we discuss latter in the model, *levels* of parental educational are likely to reflect parental ability and knowledge in the procurement of her/ his children's health, and therefore, are key elements in the literature of the health home production

[6] In agricultural communities, the argument can be reversed if sons provide help on the farm.

functions. However, human capital investments also reflect the person's permanent income, and possible bargaining power within the household when differences in education between parents are significant. Hence, it strikes us as very unlikely that differences in parental education would have a substantial impact on a couple's home production function rather than reflecting a bargaining power issue, once we control for each parent levels of education. Consequently, the pooling hypothesis in the model is also tested by looking at the effect of parental differences in education on each child nutritional status, once we control with parental levels human capital attainment.

In addition, we specify each parent's preferences as dependent upon parental (observed and unobserved) characteristics, and on *all* household member's private and public consumption. The latter allows us not only to explain any altruistic behavior and externalities in consumption, but also to capture any other preference interactions that are essential in modeling why parents allocate resources in the provision of health to their children:

$$W = \left[\, \Omega; U^{m}\left(X, H, u_m, u_f, e_m, e_f\right), U^{f}\left(X, H, u_m, u_f, e_m, e_f\right)\right]$$

$$0 \le \Omega\left[Y_m, Y_f; (Diff.Edu.)\right] \le 1 \qquad (1)$$

X represents a vector of household market commodities, including leisure; *H* stands for all nonmarket goods produced at home, such as child's health investment in terms of parental rearing; u_m and u_f denote mother's and father's observed background characteristics such as age and education; and e_m and e_f correspond to vectors of parental unobservable characteristics, such as tastes reflecting child-gender and child birth-order predilection.

The household welfare function is maximized, subject to the family budget constraint;[7]

[7] For simplicity, we assume parental incomes to be the only source of family monetary resources.

$$PX = Y_m + Y_f\,,$$

$$Y_i = E_j + n\,y_j \qquad \text{for} \quad i = w, f\,; \tag{2}$$

P is a vector of market prices excluding the price of leisure; and Y_m and Y_f stand for mother's and father's total income. We further assume parental total income be linear combination of parental earnings (E) and nonlabor income (n y_i). Earnings depend as usual, on individual's wage and on a time constraint For exposition purposes, parental nonlabor income is assumed to be predetermined in our static model.[8] Later in the estimation process, we consider the possibility of measurement error in income.

The health of the children in the family does not depend merely on the parents' preferences in the allocation of resources. Other variables such as child biological factors, community characteristics and each parent's specific technology in raising children become important elements in determining the health status of the child. Therefore, we introduce to the model a nonmarket commodity production function that enables us to capture any parental child-rearing technology in the procurement of health:

$$H = H\left(X, X_n, \boldsymbol{\theta}, \boldsymbol{\eta}_p, \boldsymbol{\eta}_c\right) \tag{3}$$

We allow the nonmarket commodity production function (H) to depend on any market purchased (X) and nonmarket (X_n) inputs that are related to the health status of the child, such as food intake, health services and breast-feeding respectively. We also incorporate a vector of child's characteristics (θ), such as age, gender and birth-order, that controls for biological factors influencing the child's health outcome. In addition, we introduce a vector of parental-specific characteristics (η_p), that reflect each parent child-rearing

[8] The exogeneity assumption should be taken with caution when interpreting the regression results, in the event that nonlabor income reflects previous labor supply decisions that are correlated to current household's unobserved characteristics.

technology. η_p can be thought as mother's and father's age and human capital; child-rearing experience in terms of birth-order, and as any other parental child-rearing specific ability in terms of parent-son and parent-daughter gender matching. These characteristics may also account for the fact that parents learn from the experience of their older children, and be more efficient at raising later ones. Finally, the nonmarket commodity production function depends on regional and community characteristics (η_c) that capture characteristics related to the environment.

The maximization process leads to aggregate market and nonmarket household commodity demands for each element of *X* and *H*, which includes child's health investment:

$$X^* = G_x\left(P, y_m, y_f; u_m, u_f, e_m, e_f; q, h_p, h_c\right) \tag{4}$$

$$H^* = G_z\left(P, y_m, y_f; u_m, u_f, e_m, e_f; q, h_p, h_c\right)$$

These optimal demands depend on a vector *P* of commodity prices, and on the set of observed and unobserved household characteristic and community characteristics that reflect parental preferences and child-rearing technology in the allocation of resources within the household. Section III of the paper deals with the empirical strategy that estimates the child's health as a component of H^*.[9]

We turn next to the description of the data.

IV. The Data

The Chilean Household Survey CASEN92 (*Encuesta de Caracterización Nacional* 1992) was carried out by the Chilean National Ministry of Planning in collaboration with the University of Chile. The 1992 survey consists of a

[9] In this context, H^* corresponds to a vector of household aggregate health demand, that includes individual household member health status, such as son's and daughter's nutritional levels.

nationwide cross-sectional random sample of 143,459 individuals and 35,948 households. It encapsulates detailed socioeconomic and demographic information at an individual and household level, about labor and nonlabor income, dwelling characteristics, gender, age, levels of education and the nutritional status of the children, among other variables. This information becomes essential to the empirical implementation of our model. Having detailed demographic information at an individual level for parents living in the household facilitates the estimation of the health status of the child as an intrahousehold resource allocation outcome and child-rearing technology. Similarly, information about each child's gender, age and consequently birth-order, permits one to analyze how child characteristics relate to parental demographics in determining the allocation of resources on child health investment.

The Survey provides data about child nutritional status in terms of biomedical risk for those children with five years or less of age. This biomedical hazard is defined under five ordered categorical variables (*normal or eutrofic, over-weighted, biomedical risk, moderate, and accentuated malnutrition*), which capture the overall healthiness of the child relative to Chilean national health standards. We should say that the entire Chilean population is entitled to basic public health services since a preventive health system was established. This means that for families to be eligible for governmental subsidies, each child has to be subject to medical controls on a periodical basis. The regular visits to medical clinics provide each child with a health record, measured by professionals physicians, that serves as source for the nutritional status information in this survey.[10] This allows one to have a multidimensional health indicator for the child, while embracing an objective classification criteria for the empirical analysis.

For estimation purposes, we select a subsample of 11,702 observations for all children less than six years of age living in the household. For these children the nutritional biomedical hazard is defined, and at least one parent

[10] Visits vary from a quarterly to a monthly basis depending on the child's nutritional status.

is present in the household either on married status or cohabiting.[11] We aggregate the nutritional five-ordered categorical variable into a dichotomous variable. The categories: “biomedical risk”, “moderate” and “accentuated malnutrition” were redefined as malnourished, while “normal or eutrofic” and “over-weighted or obese” categories were collapse into the now called well-nourished category. Although, in general terms the over-weighted category may reflect bad nutrition habits, the nature of an over-weighted child compared to a malnourished infant, makes both categories two separate health problems. The Chilean Ministry of Health through regular medical evaluation identifies whether the over-weighted condition of the child, is accompanied with low or high levels of proteins intake. When a child is over-weighted but, at the same time, presents lower levels of proteins intake, she or he is classified as a child with malnutrition problems.[12]

Table 1 shows that under the new definition, 83 percent of the children were found to be well-nourished in our sample, while 17 percent lay under the malnourished category. The distribution of child’s age is largely even across ages, with a mode at four years old. With respect to household

[11] The reader may be concern about potential selectivity bias in our sample since we are focusing on married or cohabiting households with infants. However, the purpose of the article is to examine how parental characteristics reflect the allocation of resources towards investments in their children’s health. It is not our intention to estimate either a health production function for household children, or the estimation of a health conditional demand for all Chilean infants. Therefore, conditional on the targeted population of our interest we consider our sample faces no selection problems. Moreover, since our targeted population is couple households, and because the marital status of a person is a matter of choice, and therefore endogenous, we do not differentiate between married and cohabiting families.

[12] In a previous version of this paper we estimated an ordered logit model following each five biomedical categories. However, given our sample small variability in several categories, we do not obtain precise estimates. (Children classified in biomedical risk, moderate and accentuated malnutrition represent only 6.03%, 4.84% and 6.32% of the sample, respectively). Therefore, in order to gain efficiency in our estimates and because the nutritional-status distribution mode laid at the “normal or eutrofic” category (82.81%), we decided to regroup the nutritional level of the children into two basic categories: well nourished and malnourished. We believe no valuable information is lost.

composition we can see that most of the families have either one or two children under six years old.

Mothers are, on average, slightly younger, presenting lower earnings and have lower nonlabor income compared to fathers (Table 2).[13]

Table 1. Child Characteristics

Child's Nutritional Status		
		Relative Frequency
Mal-nourished		17.19
Well-nourished		82.81
Child's Age Distribution (years)		
Zero		14.45
One		15.29
Two		17.21
Three		17.95
Four		18.20
Five		16.89
Children Less than Six Years Old in the Household		
One child		54.12
Two children		37.09
Three children		7.90
Four children		0.85
Five children		0.04
Sample size	11,702	100.00

[13] 90 percent of the mothers report zero nonlabor income, yet only 0.4 percent of the fathers are found to have zero non-earned income.

Table 2. Parental Characteristics

Parental Age and Income

	Mean	Standard Error
Mother's age	29.807	(0.058)
Father's age	33.424	(0.071)
Mother's labor income	19,909.870*	(719.241)
Father's labor income	125,333.300*	(2,139.365)
Mother's nonlabor income	1,427.040*	(0.001)
Father's nonlabor income	11,103.200*	(0.003)

Parental Education Distribution

	Mother	Father
None	247	249
	(2.11%)	(2.13%)
1-8 years	5,196	5,082
	(44.40%)	(43.43%)
9-12 years	5,026	4,904
	(42.95%)	(41.91%)
13-18 years	1,233	1,467
	(10.54%)	(12.54%)

Parental Relative Education and Age Distribution

Mothers	With lower education than their spouse	With higher education than their spouse	Total
Younger than	8,262	1,398	9,660
their spouse	(70.60%)	(11.95%)	(82.55%)
Older than their spouse	1,707	335	2,042
	(14.59%)	(2.86%)	(17.45%)
Total	9,969	1,733	11,702
	(85.19%)	(14.81%)	(100.00%)

*1992 Chilean Pesos.

The great majority of the parents have at least some degree of education, displaying both similar distributions with modes at nine to twelve years of education. However, from a family perspective, only 15 percent of the mothers show higher levels of education than their spouse. This may represent differences in mother's and father's long-term earnings profiles, and consequently, reflect some bargaining power in the family. We take advantage of this fact to test the common preference assumption using differences in educational levels, as additional source of power.[14]

V. Empirical Strategy and Results

The empirical strategy focuses on the nutritional status of the child as an indicator reflecting the household child-health investment decisions represented in equation (4). According to our model, the core of our analysis is to regress the nutritional status of the child (h_{ij})[15] on parental characteristics such as individual nonlabor incomes, age and levels of education to test differences in parental preferences and child-rearing technology. The hypothesis of gender differentiation is tested by allowing the core model be fully interacted with a child-gender dummy variable (GENDER), while the birth-order differentiation hypothesis is carried on by fully interacting the model with a birth-order dummy variable (B/ORDER). In what follows we explain the core model (A), using the gender hypothesis specification:

Gender Hypothesis Specification

$$h_{ij} = 1\{A + A \times GENDER + e_{ij} > 0\} \qquad \text{i = child, j = household} \qquad (5)$$

[14] According to the pooling hypothesis, only aggregate resources determine household behavior, thus the effect of differences in educational levels on the child nourishment should be zero in order not to reject the common preference model.

[15] h_{ij} is an indicator that takes the value of one if well nourished, and zero otherwise.

where

$$A = a + b_m n y_m + b_f n y_f + Z_m g_m + Z_f g_f + d_{1c} C + \qquad (6)$$

$$+ d_G GENDER + d_B B/ORDER + D_r j_r$$

In the core model, mother's (Z_m) an father's (Z_f) characteristics such as parental age and education are include in the regression as important covariates in determining either child-rearing technology or parental preferences in the provision of their children's health.[16] Both characteristics enter our model as variables affecting the family welfare function and the nonmarket child-rearing production function. Consequently, any differences captured by parental age and education *levels* will reflect differences in child-rearing technology and parental taste differentiation.

In terms of child-rearing technology, one would like to think that the more educated the parent is, the more efficient he or she becomes in procuring his (her) children's health. Different age levels may also reflect differences in parental energy effort in terms of child rearing. To control for parental education, we use three categorical variables for each parent: education between 1-8 years, education between 9-12 years, and education between 13-18 years.[17] Additionally, mothers and fathers age enter in our regressions as a second order polynomial to capture age nonlinearities on rearing technology.[18]

---

[16] Despite fact the theoretical framework derives demand equations that depend on wages, we avoid the problem of predicting wages for non-workers (81 percent of women). This of course means that the coefficients on education will partly capture wage effects as mentioned in the discussion.

[17] 18 years of education is the highest level attained by either parent in our sample. Less than one year of education is the left-out category.

[18] $(age_m + age_f)^2$.

We also control for mother's (ny_m) and father's (ny_f) nonlabor income,[19] to test the common preference assumption in child health procurement, by examining the income pooling hypothesis.[20] In this framework, individual incomes are taken as bargaining power variables affecting the nutritional status of the child through the welfare weights. To find significantly different effects of mother's and father's nonlabor income on the nutritional status of the child, would contradict the common preference result which states that household optimal allocation of resources depends only on all members pooled income. This would suggest that mothers and fathers share different tastes in the procurement of their children's health. Additionally, we include the age of the child (C),[21] together with a gender (GENDER) and birth-order (B/ORDER) dummy to control for biological factors that influence the health development of the infant. GENDER takes the value of one if the child is a son and zero if she is a daughter. B/ORDER is equal to one if the infant is the oldest child living in the household at the time of the interview, and zero otherwise. Regional and rural-urban categorical variables (D_r) are employed to account for community heterogeneity, such as climate and economic conditions.[22] In order to test whether there is any gender differentiation in parental preferences or child-rearing technology we allow the core model (A) to be fully interacted with the child GENDER dummy.

[19] Parental unearned income also enters the model as a second order polynomial, i.e., $(ny_m + ny_f)^2$.

[20] Labor income reflects the decision on labor supply and is part of the household behavior, therefore, we exclude it in the logit estimation regressors. Logit specifications, were also tested using total labor income with a conditional logit (Chamberlain, 1980). However, small variability of the dependent variable within families prevented us to successfully identify the model.

[21] We also include a quadratic term with respect to child's age in the empirical estimation.

[22] People may be concerned about the potential endogeneity of these variables due to migration issues, for instance. Nevertheless, in view of the omitted-variable bias that one could incur when neglecting them, we have resolved to include them. The survey does not provide migration information.

Table 3 presents logit estimates under the gender (son/daughter) hypothesis.[23] The age of the mother has a greater impact on the nutritional status of the daughters rather than on the nutritional status of her sons. Likewise, in terms of parental age, fathers seem to direct more resources to the provision of their son's health than to their daughters'. The same pattern can be seen with respect to parental education. At higher levels of education (13 - 18 years), mothers have a stronger effect on daughters, while fathers show a greater effect on sons. The fact that this gender differentiation only takes place at higher levels of education may reflect the role of parental human capital in allowing both parents to specialize in the allocation of resources when procuring the health of their daughters and sons.

A general concern is the problem of measurement error in unearned income that is often encountered in household surveys. However, the fact that measurement error in paternal (or maternal) nonlabor income does not differ across siblings in the same household indicates that any bias transmitted to the estimates is common across siblings. Therefore, we exploit within-household variation in child gender (and birth-order) to test the income pooling hypothesis in a difference-in-difference framework that eliminates the measurement error bias. We test whether mother's and father's differential income effect with respect to gender (or birth/order) is equal.[24]

The p-value of [0.37] for the difference-in-difference effect of nonlabor income on child's nourishment based on gender predilection, does not allow one to reject the common preference model.

---

[23] Since we are not estimating a conditional demand and its elasticity, but only interested in how parental characteristics individually reflect the allocation of resources, Logit estimates are reported in coefficient terms as opposed to marginal probabilities.

[24] The Appendix presents the intuition of the test.

Table 3. Logit Parental Effect on Child Nourishment Gender Fully Interacted Model

	Son	Daughter	Difference
Mother age	-0.008	0.020	-0.028
	[0.443]	[0.056]	[0.058]
Father age	0.040	0.003	0.037
	[0.000]	[0.744]	[0.005]
Mother education (1-8 years)	0.162	0.293	-0.131
	[0.562]	[0.220]	[0.722]
Mother education (9-12 years)	0.253	0.634	-0.381
	[0.382]	[0.012]	[0.322]
Mother education (13-18 years)	0.385	1.292	-0.907
	[0.288]	[0.000]	[0.070]
Father education (1-8 years)	0.560	0.373	0.187
	[0.021]	[0.123]	[0.585]
Father education (9-12 years)	0.889	0.446	0.443
	[0.001]	[0.079]	[.0220]
Father education (13-18 years)	1.554	0.434	1.120
	[0.000]	[0.173]	[0.016]
Mother income	-1.355	0.256	-1.611
	[0.264]	[0.810]	[0.317]
Father income	0.654	0.719	-0.065
	[0.074]	[0.051]	[0.901]

Table 3. (Continue) Logit Parental Effect on Child Nourishment Gender Fully Interacted Model

	Difference in difference
Income pooling	-1.546
(Diff.-Diff.)	[0.367]

Notes: The sample size corresponds to 11,702 children; the Pseudo R^2: is 0.034. Logit estimates are reported in coefficient terms as opposed to marginal probabilities. Results for age and income correspond to second-degree polynomials, and represent the total effect of the variable calculated at the sample mean. Therefore, P-values are reported in [parenthesis]. Parental education enters the model nonparametrically via marginal categorical variables; the omitted category is zero years of schooling. p-values are shown under the difference-in-difference income-pooling hypothesis. In addition to the covariates of interest shown in the table, each regression includes the age of the infant and a set of categorical variables denoting the child's gender and birth-order, as well as, the region of residence.

We now replace the GENDER dummy with the B/ORDER dummy and proceed to test differences in birth-order in the same way.

Birth-Order Hypothesis Specification

$$h_{ij} = \{A + A \times B/ORDER + e_{ij} > 0\} \tag{7}$$

According to the birth-order hypothesis (Table 4), we find larger effects of mother's age on oldest children. The characteristics of the father reflect no birth-order differentiation. However, neither can we reject common preference with respect to birth-order health status based on the income's difference-in-difference [p-value of 0.52].

Table 4. Logit Parental Effect on Child Nourishment
Birth-Order Fully Interacted Model

	Oldest	Latter	Difference
Mother age	0.027	-0.007	0.034
	[0.042]	[0.432]	[0.034]
Father age	0.009	0.025	-0.017
	[0.462]	[0.001]	[0.247]
Mother education	0.221	0.232	-0.011
(1-8 years)	[0.646]	[0.237]	[0.983]
Mother education	0.374	0.459	-0.085
(9-12 years)	[0.445]	[0.028]	[0.873]
Mother education	0.243	1.261	-1.018
(13-18 years)	[0.652]	[0.000]	[0.102]
Father education	0.541	0.417	0.124
(1-8 years)	[0.161]	[0.029]	[0.774]
Father education	0.838	0.577	0.261
(9-12 years)	[0.035]	[0.004]	[0.558]
Father education	1.148	0.887	0.261
(13-18 years)	[0.013]	[0.001]	[0.627]
Mother income	-2.015	-0.340	-1.675
	[0.253]	[0.663]	[0.385]
Father income	0.510	0.885	-0.345
	[0.335]	[0.007]	[0.575]

Table 4. (Continue) Logit Parental Effect on Child Nourishment Birth-Order Fully Interacted Model

	Difference in difference
Income pooling	-1.330
(Diff.-Diff.)	[0.524]

See notes Table 3. Pseudo R^2: 0.035

Parental Levels of Education and Preferences

The estimates presented in Tables 3 and 4 do not allow us to differentiate in terms of education, those effects coming via preference from those coming via child-rearing technology. In an effort to analyze the preference effect in child nutrition with respect to parental human capital, a new categorical variable (D_E) is introduced to the core model:

$$A_1 = A + d_E D_E \tag{8}$$

Now extended core model (A_1) includes a dummy variable (D_E) equal to one for those families where the mother reports higher levels of education than her spouse. Thus, *conditional* on each parent level of education, age and income, the interaction of A_1 with GENDER and B/ORDER allows one to test the common preference model through the education bargaining power effect in terms of gender and birth-order predilection.

Gender Hypothesis Specification

$$h_{ij} = 1\{A_1 + A_1 \times GENDER + e_{ij} > 0\} \tag{9}$$

Table 5 shows that, after controlling for the parental education categorical

Table 5. Logit Parental Effect on Child Nourishment Gender Fully Interacted Model

	Son	Daughter	Difference
Mother age	-0.025	0.019	-0.043
	[0.069]	[0.165]	[0.023]
Father age	0.054	0.004	0.050
	[0.000]	[0.731]	[0.003]
Mother education	0.282	0.386	-0.104
(1-8 years)	[0.316]	[0.108]	[0.778]
Mother education	0.572	0.935	-0.363
(9-12 years)	[0.076]	[0.001]	[0.393]
Mother education	0.873	1.739	-0.866
(13-18 years)	[0.039]	[0.000]	[0.131]
Father education	0.327	0.076	0.251
(1-8 years)	[0.224]	[0.779]	[0.511]
Father education	0.499	-0.022	0.521
(9-12 years)	[0.124]	[0.945]	[0.251]
Father education	1.035	-0.162	1.197
(13-18 years)	[0.016]	[0.686]	[0.041]
Education difference	-0.363	-0.443	0.080
as bargaining	[0.045]	[0.015]	[0.755]

Table 5. (Continue) Logit Parental Effect on Child Nourishment Gender Fully Interacted Model

	Son	Daughter	Difference
Mother income	-1.348	0.226	-1.575
	[0.267]	[0.830]	[0.328]
Father income	0.646	0.682	-0.036
	[0.077]	[0.063]	[0.945]
			Difference in difference
Income pooling (Diff.-Diff.)			1.539
			[0.368]

See notes Table 3. Pseudo R^2: 0.036.

variable (D_E), the effect of parental age on child nourishment is larger and is consistent with the same gender bias pattern observed in Table 3. Mother's age is found to be larger for daughters, while fathers continue to direct more resources to their sons for high levels of education (13-18 years). The common preference model, through the education bargaining power dummy cannot be rejected in terms of gender differentiation, [p-value of 0.76]. However, negative and significant estimates of the education dummy variable on the child's nutritional status show that mothers with higher education attainment relative to their spouse channel less resources to their children than those mothers who are relatively less educated. This may reflect high child rearing opportunity costs in terms of mothers' household chore decisions.

We now proceed to test birth-order predilection with education as power factor.

Birth-Order Hypothesis Specification

$$h_{ij} = 1\{A_1 + A_1 \times B / ORDER + e_{ij} > 0\} \tag{10}$$

Logit estimates presented in Table 6 show no evidence of child birth-order differentiation with respect to parental child-rearing technology (e.g., levels of age and education), and with respect to parental predilection (e.g., differences in nonlabor income and human capital). This fact leads us to the next issue that is that child-rearing technology and parental preferences may alternate in their effects and take different directions depending on the gender and birth-order of a child.

Table 6. Logit Parental Effect on Child Nourishment Birth-Order Fully Interacted Model

	Oldest	Latter	Difference
Mother age	0.011	-0.011	0.022
	[0.516]	[0.336]	[0.278]
Father age	0.022	0.029	-0.007
	[0.134]	[0.004]	[0.693]
Mother education	0.345	0.328	0.017
(1-8 years)	[0.472]	[0.097]	[0.974]
Mother education	0.708	0.756	-0.048
(9-12 years)	[0.171]	[0.001]	[0.933]
Mother education	0.767	1.696	-0.929
(13-18 years)	[0.207]	[0.000]	[0.184]
Father education	0.302	0.143	0.159
(1-8 years)	[0.462]	[0.510]	[0.732]

Table 6. (Continue) Logit Parental Effect on Child Nourishment Birth-Order Fully Interacted Model

	Oldest	Latter	Difference
Father education (9-12 years)	0.421 [0.371]	0.144 [0.581]	0.277 [0.607]
Father education (13-18 years)	0.590 [0.302]	0.340 [0.324]	0.250 [0.708]
Education difference as bargaining	-0.376 [0.098]	-0.410 [0.009]	0.034 [0.902]
Mother income	-2.106 [0.232]	-0.334 [0.668]	-1.772 [0.358]
Father income	0.467 [0.376]	0.829 [0.008]	-0.362 [0.555]
			Difference in difference
Income pooling (Diff.-Diff.)			1.410 [0.499]

See notes Table 3. Pseudo R^2: 0.036

Gender-Birth/Order Hypotheses Specification

$$h_{ij} = 1\{A_1 + A_1[(GENDER) + (B/ORDER) + \\ +(GENDER \times B/ORDER)] + e_{ij} > 0\} \tag{11}$$

To analyze this possibility, we fully interact the core model (A1) with

GENDER and B/ORDER simultaneously. Table 7 presents the results. We find gender differentiation with respect to parental age. Mothers continue to direct more resource to their daughters, while fathers to their sons. However, this gender polarity is significant only for latter-born daughters and latter-born sons, respectively (columns [(B) minus (D)]). These results provide evidence of how parenting experience may lead to specialization in childcare. Additionally, education of the mother appears to be more important than father's education in providing nourishment of the child, (columns B and D). This issue seems plausible considering that Chilean mothers spend relatively more time with their children than fathers. We also obtain weak evidence of gender differentiation with respect to birth-order. If we focus on the birth-order hypothesis, contrary to the previous results, we find evidence of birth-order differentiation by gender. Looking at high levels of education (13-18 years), the mother assigns less resources to the oldest child if he is a son (columns [(A) minus (B)]), but makes no differentiation in terms of birth-order among daughters, (columns [(C) minus (D)]).

Finally, the common preference hypothesis, with respect to gender and birth-order predilection, cannot be rejected using education differences and nonlabor income as bargaining factors. Nevertheless, the negative estimates of the education bargaining dummy variable indicate that mothers with relative more education than their spouse, *conditional to the level of education of each spouse*, direct fewer resources to their latter children than those who are relatively less educated. This may reflect high child-rearing opportunity costs in terms of the mother's out-of-home activities due to the high correlation between education and potential earnings.[25]

[25] Ideally one would like to estimate the model for two separate household samples: working and non-working mothers. This, as opposed to interacting a categorical variable for labor force participation, given that leisure decisions are endogenous to our model. Unfortunately, we cannot drive our empirical results into that step: only 10 percent of the women in our sample report positive nonlabor income, and from them, only 19 percent report to have entered the labor force [see footnotes 13 and 16]. This leaves us with very few observations to estimate the interacted models, especially when 80 percent of our sample infants lie on the well-nourished category [see footnote 12].

Table 7. Logit Parental Effect on Child Nourishment
Gender and Birth-Order Fully Interacted Model

	Son		Daughter		Gender Diff.		Birth-order diff.	
	Oldest	Latter	Oldest	Latter	Oldest	Latter	Son	Daughter
	(A)	(B)	(C)	(D)	(A) – (C)	(B) – (D)	(A) – (B)	(C) - (D)
Mother age	-0.010	-0.034	0.016	-0.012	-0.026	-0.046	0.024	0.004
	[0.669]	[0.04]	[0.491]	[0.468]	[0.433]	[0.050]	[0.421]	[0.895]
Father age	0.038	0.054	0.025	0.003	0.014	0.051	-0.015	0.021
	[0.076]	[0.000]	[0.275]	[0.827]	[0.656]	[0.014]	[0.554]	[0.554]
Mother education (1-8 years)	-0.370	0.407	0.565	0.312	-0.936	0.095	-0.778	0.253
	[0.666]	[0.177]	[0.356]	[0.238]	[0.375]	[0.813]	[0.393]	[0.704]
Mother education (9-12 years)	-0.225	0.718	1.068	0.843	-1.293	-0.124	-0.943	0.225
	[0.809]	[0.103]	[0.103]	[0.008]	[0.255]	[0.793]	[0.342]	[0.757]
Mother education (13-18 years)	-0.679	1.403	1.374	2.099	-2.053	-0.696	-2.081	-0.724
	[0.530]	[0.004]	[0.074]	[0.000]	[0.122]	[0.325]	[0.080]	[0.432]
Father education (1-8 years)	0.675	0.265	0.288	0.075	0.387	0.190	0.410	0.213
	[0.240]	[0.393]	[0.644]	[0.807]	[0.648]	[0.664]	[0.530]	[0.760]
Father education (9-12 years)	0.974	0.420	0.296	-0.075	0.678	0.494	0.554	0.371
	[0.157]	[0.263]	[0.668]	[0.840]	[0.487]	[0.347]	[0.479]	[0.636]
Father education (13-18 years)	1.912	0.773	-0.036	-0.004	1.949	0.778	1.139	-0.032
	[0.035]	[0.118]	[0.964]	[0.993]	[0.107]	[0.262]	[0.270]	[0.973]

Table 7. (Continued) Logit Parental Effect on Child Nourishment Gender and Birth-Order Fully Interacted Model

	Son		Daughter		Gender diff.		Birth-order diff.	
	Oldest	Latter	Oldest	Latter	Oldest	Latter	Son	Daughter
	(A)	(B)	(C)	(D)	(A) – (C)	(B) – (D)	(A) – (B)	(C) - (D)
Education as bargaining	-0.028 [0.543]	-0.389 [0.074]	-0.442 [0.163]	-0.402 [0.077]	0.235 [0.615]	0.013 [0.966]	0.181 [0.654]	-0.040 [0.919]
Mother income	-1.027 [0.724]	-0.859 [0.546]	-0.875 [0.793]	1.661 [0.292]	-0.152 [0.973]	-2.519 [0.235]	-0.168 [0.959]	-2.536 [0.492]
Father income	1.584 [0.143]	0.625 [0.128]	0.468 [0.519]	1.160 [0.018]	1.116 [0.392]	-0.535 [0.402]	0.959 [0.407]	-0.692 [0.429]

	Gender Diff-in-diff		Birth–order Diff-in-diff	
	Oldest	Latter	Son	Daughter
Income pooling (Diff.-Diff.)	1.268 [0.786]	1.984 [0.373]	1.128 [0.749]	1.844 [0.626]

See notes Table 3. Pseudo R^2: 0.040

VI. Conclusions

Household decisions have been traditionally modeled by treating the household as the elementary decision unit. However, this approach provides no information about how family resources are allocated within the household. This is important because household behavior could well reflect the decision marker's welfare but not necessarily the other family members' well being. We believe more research has to be done in the interest of economic modeling to improve the understanding of intrahousehold allocation.

This paper examines the nutritional status of Chilean children, in a context of family resources, where mother and father characteristics reflect differences in child-rearing technology and parental preferences. Mother and father incomes, and differences in education are taken as bargaining variables reflecting tastes. Levels of education and parental age enter our model as child-rearing technology factors.

We find gender specialization in child-rearing: mothers direct more family resources towards their daughters, while fathers channel more to their sons. This gender polarity is significant for parental age and high levels of education. Additionally, the education of the father becomes less important than mother's education in attending the nourishment of the children. This supports that household chores are essentially a woman's task in the majority of Chilean families.

Although it is not possible to conclude if the income-pooling hypothesis holds for gender and birth-order predilection, the evidence shows how important is examining household models for policy analysis, when resource allocation is also a function of child-rearing parental abilities. The evidence also shows that, holding each parent schooling constant, mothers with a higher education level than their spouse direct less resources to their children, than those who are relatively less educated. This may reflect the increase in child-rearing opportunity cost when mothers are better educated to perform out-of-home activities. If this is the correct interpretation, then the mother's decision to perform other activities different from home chores, should not be viewed exclusively in terms of her individual opportunity cost, but also on the basis of her children's welfare. Along these lines, a larger supply of public childcare services could help offset the associated loss in children's welfare.

Appendix

The following is an informal approach to stimulate the intuition behind the difference-in-difference pooling hypothesis testing, contingent to measurement error in current unearned income. Without loss of generality, we will focus on the gender preference hypothesis:

Let the true fully-interacted model be represented by

$$h_2 = b_0 + X\, q + b_1\, y_m + b_2\, y_f + m_s \qquad \text{(m = mother; f = father)}$$

$$h_d = g_0 + X\, q + g_1\, Y_m + g_2\, Y_f + m_d \qquad \text{(s = son; d = daughter)} \qquad (12)$$

Let the bias on income coefficients (caused by the measurement error) be independent to the gender predilection of the mother and the father, but different across parents:

$$p\, lim\hat{b}_1 = b_1 + \Theta_m \qquad\qquad p\, lim\hat{g}_1 = g_2 + \Theta_m$$

$$p\, lim\hat{b}_2 = b_2 + \Theta_f \qquad\qquad p\, lim\hat{g}_2 = g_2 + \Theta_f \qquad (13)$$

Testing the gender common preference assumption implies:

$$\mathrm{H_o}: \quad \mathrm{b_1 = b_2} \ \text{ and } \ \gamma_1 = \gamma_2 \qquad (14)$$

Nevertheless, testing H_0 with (13) may cause to reject H_0 even when it is true. However, the common preference assumptions can consistently be tested using a difference-in-difference approach:

$$H_0 : (b_1 - g_1) - (b_2 - g_2) = 0.$$

Since

$$p \lim \left[\left(\hat{b}_1 - \hat{g}_1 \right) - \left(\hat{b}_2 - \hat{g}_2 \right) \right] = (b_1 - g_1) - (b_2 - g_2) \qquad (15)$$

References

Aranda, X. (1986), "Las Families Pobres Rurales en Areas de Pequeña Propiedad, 71-88" En Búsqueda de la Familia Chilena, Ed. Ediciones Universidad Católica de Chile, pp. 224.

Becker, G.S. (1964), "Human Capital. New York", Columbia University Press.

Bourguignon, F. and Chiappori, P. (1992), "Collective Models of Household Bahvior". *European Economic Review* 36, 355-364.

Bourguignon, F., Browning M., Chiappori, P. and Lechene, V. (1993), "Intra Household Allocation of Consumption: A Model and Some Evidende from French Data", *Annales de Economie et de Statistique* 29,137-156.

______ (1994), "Income and Outcomes: A Structural Model of Intrahousehold Allocation". *Journal of Political Economy* 102 nº. 6, 1067-1096.

Chamberlain, G. (1980), "Analysis of Covariance with Qualitative Data", *Review of Economic Studies* 47, 225-238.

Chiappori, P. A. (1992), "Collective Labor Supply and Welfare", *Journal of Political Economy* 100, nº. 31, 437-467.

Chiappori, P. and Lechene, V. (1993), "Intra Household Allocation of Consumption: A Model and Some Evidence from French Data", Annales de Economie et de Statistique 29, 137-156.

Duraisamy, P. (1992), "Gender, Intrafamily Allocations of Resources and Child Schooling in South India", Economic Growth Center, Discussion paper No. 667, Yale University.

Duraisamy, P. and Malathy, R. (1991), "Impact of Public Programs on Fertility and Gender Specific Investment in Human Capital of Children in Rural India: Cross Sectional and Time Series Analyses", in T.P. Schultz (ed.), Research in Population Economics, Volume 7, Greenwich, CT: JAI Press, 157-187.

Gissi, J. (1984), "El Machismo en los Dos Sexos", 63-98. *La Estructura Familiar en los Sectores Populares Urbanos. El Machismo en los Dos Sexos*, Ed. Cenpafal, pp. 98.

Haddad, L. and Kanbur, R. (1990), “How Serious is the Neglect of Intrahousehold Inequality?” *Economic Journal.*

Horney, M. J., and McElroy, M.B. (1981), “Nash-bargained Household Decisions: Toward a Generalization of the Theory of Demand”, *International Economic Review* 22 nº. 2, 333-349.

Lundberg, S. J., Pollak, R. A. and Wales, T. J. (1997), “Do Husbands and Wives Pool their Resources? Evidence from the U.K. Child Benefit”, Journal of Human Resources, 32.3: 463-480.

Quisumbing, A. (1994), “Intergenerational Transfers in Phillipine Rice Villages: Gender Differences in Traditional Inheritance Customs”, *Journal of Development Economics,* 43.2: 167-196.

Raczynski, D. and Serrano, C. (1996), “Mujer y Familia”, 37-70. *En Búsqueda de la Familia Chilena*, Ed. Ediciones Universidad Católica de Chile, pp. 224.

Schultz, T.P. (1990), “Testing the Neoclassical Model of Family Labor Supply and Fertility”, Journal of Human Resources XXV 4, 599-634.

Sen, A. K. (1984), “Family and Food: Sex Bias in Poverty”, 346-368. *Resources, Values, and Development,* Ed. Amartya Sen, pp.546.

Thomas, D. (1994), “Like Father, Like Son; Like Mother, Like Daughter”, *Journal of Human Resources* XXIX 4, 950-988.

______ (1990), “Intrahousehold Resource Allocation, An Inferential Approach”, *Journal of Human Resources* XXV 4, 635-657.