Ethnicity, Gender and Human Capital Investments in Bolivia: Evaluation of an Old-Age Cash Transfer Program

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Abstract

This paper measures the impact of an old age unconditional cash transfer program in Bolivia (Bolivida) on children's human capital investments by ethnicity and gender of the recipient. Taking advantage of the regression discontinuity created by the program, I investigate whether this exogenous variation in income is allocated differently within indigenous, multiethnic, and non-indigenous families, conditional on having at least one eligible member and one school-age child living in the household.

I find that cultural factors play a key role in the human capital intrahousehold resource allocation process. Bolivida transfers to women and men both lead to increases in children's educational investments; however, women are more effective at promoting human capital accumulation. The pattern of allocation within ethnic groups is clear; non-indigenous and multiethnic beneficiaries have larger impacts on the expenditure in children's education than analogous indigenous beneficiaries. The children who benefited the most are those who were already enrolled in school before the program started.

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1 Introduction

The creation of an old age unconditional cash transfer program in Bolivia creates a quasi-experimental opportunity to study the intrahousehold resource allocation process upon children's human capital accumulation. As the probability of receiving the transfer changes discontinuously at the eligibility age, this sharp discontinuity in the program assignment mechanism can be used for identification of the conditional mean treatment effect under minimal assumptions (Hahn, Todd, and Van der Klaauw, 2001).

This paper presents evidence of the existence of cultural-based decision rules in the allocation of income on educational expenditure among children, and herein lies the main contribution of the paper. I focus on educational expenditure as this is the main form of human capital investment. The allocation of resources within indigenous and multiethnic families follow decision rules closely related to patriarchal family structures, where female have less power in the decision-making process. Conversely, non-indigenous families follow bargaining mechanisms in which the outcomes of women's allocations prove to be more autonomous or male-independent.

In order to parametrically evaluate the effect of the program, I use a reduced-form equation that links child-level educational expenditure with program participation, after controlling for total per capita income excluding the pension, socioeconomic and demographic characteristics of the household, and child-level attributes. The effect of the program is measured by comparing households that have at least one beneficiary member, with those that will have a beneficiary some time in the near future. The model is estimated using eligibility to avoid potential endogeneity in the pension receipt variable.

The core finding is that these cash transfers to both, women and men, lead to substantial improvements in children's educational expenditure. The effect, however, is stronger in households where the pension is received by women. The patterns of allocation with respect to the ethnicity of the recipient are unambiguous; non-indigenous beneficiaries have a stronger impact on human capital investments than their indigenous counterparts, conditional on program participation.

To the best of my knowledge, with the exception of Martínez (2004), estimates of the impact of this program are nonexistent. He evaluates the impact of the program on household food expenditure by geographic location and gender of the recipient. His paper finds that the pension has a positive effect on the consumption of household-produced agricultural products equivalent to one and a half times the value of the pension. This suggests that the program is promoting productive investments that ultimately create a multiplier effect of the transfer on household food consumption.

The intrahousehold allocation of old age pension programs has been analyzed in different contexts and settings. By far, the most widely studied pension program is the South African means-tested "Old Age Pension Program". Case and Deaton (1998) investigate the standards of living of those families who have a member that receives the pension and find that the Program is well targeted as far as it is reaching the poorest households. They also analyze the redistributive consequences of the transfer on food, clothing, housing, schooling, transport, health, and adult goods expenditure by gender of the head of the household. The paper finds that the pension is spent similarly to any other income but that there are gender specific expenditure patterns. Duflo (2003) analyzes the impact of this program on children's weight-for-height, and height-for-age. She finds that both variables are dramatically improved among girls when their grandmothers receive the transfer, but that their nutrition is entirely unaffected when the pension is received by their grandfathers. Her paper does not find any significant effect of the pension among boys. Finally, Edmonds (2006) analyzes the impact of the program on child labor and schooling responses. The paper finds large increases in schooling attendance and reductions in child labor when a member in the household becomes eligible for this cash transfer, particularly among black male recipients.

In the context of Latin America, de Carvalho Filho (2000) studies the effect of a Social Security Reform on child labor, and school enrollment in Brazil. He finds that girls and boys enroll more in school, and reduce their labor market participation, respectively, as a result of this exogenous variation in income.

2 Bolivia, Ethnicity and Cultural Diversity

Bolivia is the poorest country in South America and second only to Nicaragua in Latin America. In 1999, 41 percent of its population was living below the national poverty line (Barja, Monterrey, and Villarroel, 2004). Traditionally, the elderly and youngest have been the most vulnerable and unprotected segments; this is reflected in the poverty indicators by age that show that whereas no more than 37 percent of adults are poor, a 49 percent of children and elderly live in poverty (Udape and Unicef, 2005).

Within the Latin American countries, Bolivia has a complex and unusual multi-ethnic dimension, only comparable to Guatemala, Peru and Ecuador. Data from the last census in 2001 reveal that close to 62 percent of the population identify themselves as belonging to one of the many indigenous groups living in the country. This makes of Bolivia the country with the highest percent of indigenous population in the Region. This peculiar ethnic dimension is closely related to the income distribution; while close to 49 percent of indigenous people live below the poverty line, only 24 percent of the non-indigenous people is poor (Udape, 2006). However, cultural disparities among ethnic groups are important and go far beyond the income distribution. Nonetheless the most notorious observable differences might be race and language, there indeed exist profound differences in traditions, values and beliefs that might potentially derive in different decision rules in the allocation of resources within households. In particular, patriarchal family structures (more common among indigenous families) might limit the power of indigenous women to allocate their resources as compared to non-indigenous women.[1]

3 The Bolivida Old-Age Pension Program

In the middle of the 1990s, an array of unprecedented and still controversial economic and social structural reforms were implemented by the Bolivian Government with the aim of confronting a long-term stagnation of the economy. One of these reforms was the so called the "Capitalization", which sold half of the six largest state-owned companies (including the telecommunications, hydrocarbons, air transportation, railroad industries, electrical energy, and smelting companies) to private investors. Another of these reforms was the Pension System Reform, which eliminated the old publicly administered pay-as-you-go system, implemented the current privately managed system, and entitled *all* Bolivians aged 65 and over to receive a *flat* noncontributory and unconditional cash transfer. The Government targeted this segment of the population as traditionally it had been characterized for having no access to the system (coverage in Bolivia is one of the lowest in the Region; close to 80 percent of the Bolivians are not affiliated to the pension system, Yánez-Pagans and Landa (2007)).

However, the unique feature of the Bolivian pension reform was not this innovative unconditional cash transfer, but rather its association with the privatization process. As the Capitalization was expected to generate considerable revenues for the country, the Government determined that this pension would be financed with the dividends of the shares of the capitalized companies. The amount of the benefit was originally set up as an annuity of USD248 and, since then, the program has been used as a governmental mechanism to redistribute the gains of the Capitalization among the elderly Bolivians.

The Program was strategically introduced in 1997, before the elections, under the name of Bono Solidario (Bonosol). After the elections, a new administration took the office and had to deal with the fact that the transition costs of the reform were higher, and dividends of the capitalized companies lower, than expected. The liquidity problems were so serious that, in 1998, the program had to be put on hold for a couple of years as payments were financially untenable. In 2001, the amount of the pension was reduced to USD120 per year (approximately 50 percent of its original value), and the program resumed under the new

[1] For a description of family structures among ethnic groups in Bolivia see Stephenson (1999), and Larson, Harris, and Tandeter (1995).

current name of Bolivida (further discussion can be found in Rofman (2006), Barja and Urquiola (2003), and Dowers, Fassina, and Pettinato (2001)).

Bolivida is a large Program as far as the amount of money transferred to the households. von Gersdorff (1997) estimated that the amount of the transfers accounted for 50 and 85 percent of the annual income of the poor and extremely poor households, respectively. Table 1 shows the amount of the benefit by ethnic groups as a proportion of total per capita expenditure. Bolivida represents on average a 7 percent of the total per capita expenditure of indigenous families (ranging from 2 to 99 percent across the expenditure distribution), and a 4 percent of total per capita expenditure of non-indigenous families (ranging from 1 to 43 percent across the analogous distribution).

Table 1: Bolivida as proportion of total per capita expenditure[a]

Percentile	Mean PCE	Share Bolivida (%)	Mean PCE	Share Bolivida (%)
	Indigenous		Non-indigenous	
1	30	99	79	43
5	39	44	98	27
10	83	35	181	5
25	120	16	200	9
50	217	7	403	4
75	321	7	734	3
90	583	3	1,246	2
95	805	3	1,670	1
99	1,715	2	2,774	1

[a] Percentiles are based on total per capita expenditure. Total per capita expenditure (PCE) adjusted for equivalence scales.

Table 2: Share of age group receiving Bolivida[a]

	Indigenous	Non-indigenous
Panel A : Male		
55-59	0.000	0.000
60-64	0.000	0.000
65-plus	0.630***	0.809***
	(0.083)	(0.061)
Panel B : Female		
55-59	0.000	0.000
60-64	0.000	0.000
65-plus	0.768***	0.751***
	(0.051)	(0.069)

[a] Standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1.

An evaluation of who Bolivida is reaching has not been done; hence it is not clear who is benefiting from the program. Table 2 shows a simple estimation of who the beneficiaries of Bolivida are by age and ethnicity of the recipient. It can be seen that approximately 26 percent of the eligible elders are indeed not receiving the pension. This is not a trivial proportion and, indeed, this fact has important implications for the identification strategy that will be discussed later. It is reassuring though that, assuming that age is reported truthfully and without error, there are not non-eligible people benefiting from the pension.

4 Data, Sample, and Survey Design

The data consist of two nationally representative cross sectional Living Standards Measurement Study Surveys (LSMS) collected in 2000 and 2001 by the Bolivian National Institute of Statistics. The surveys include a comprehensive socioeconomic module with individual-level information on receipt of Bolivida, and a detailed expenditures module with individual-level information on educational expenditure for all members in the household aged 6 and over.

The sample comprises all school-age children (the school-age range considered is 6 to 13 years old as the minimum legal working age in Bolivia is 14) that live in households with at least one person in the age-range 55 to 74 years old (Bolivida eligibility age is 65). The sample excludes all households that do not have any member in the labor market, and those whose total reported income is missing. Non-relatives and domestic non-relative workers living in the household are also excluded. Observations that belong to the top one percent of the income and educational-expenditure distributions are not included in the analysis. The sample includes 1,410 school-age children, and 462 eligible members distributed among 950 households.

The surveys use a stratified two-stage sampling. The sampling frames for the Primary Sampling Units (PSU) and Ultimate Sampling Units (USU) are the lists of Census enumeration areas and dwellings, respectively. The geographical regions, and population agglomeration are used for the explicit stratification. The proportion of households classified as poor, and the average consumption expenditure are used for the implicit stratification. The sampling frame for the baseline survey was constructed on the basis of the 1992 Census enumeration areas list. Conversely, the follow-up survey uses an updated sampling frame that was constructed upon revised cartographic information compiled for the 2001 Census. The implications of having surveys coming from two different sampling frames are discussed later on in Subsection 6.2.1.

Selected summary statistics of the sample are reported in Table 3 for eligible (i.e. households that have at least one eligible member) and non-eligible families (i.e. households that do not have any eligible member) by ethnicity and gender of the potential recipient. The estimations correspond to unconditional means adjusted for sampling design. The table shows that, in general, characteristics of the sub-sample of non-eligible families are not too different from characteristics of families in the eligible sub-sample. The first panel presents the demographic characteristics of the households. Households with a woman eligible to receive

Table 3: Summary statistics for eligible and non-eligible households[a]

	ELEGIBLE						NON
	E_i^f	E_m^f	E_{ni}^f	E_i^m	E_m^m	E_{ni}^m	ELEGIBLE
Panel A : Demographics							
Number of children	2.16	1.99	1.91	2.38	2.30	2.69	2.22
Number of school-age children	1.42	1.44	1.51	1.71	1.65	1.71	1.58
Household size	5.36	5.23	5.24	5.96	6.03	6.67	5.88
Age of oldest person in the household	70.10	69.45	74.10	70.21	71.08	71.29	58.94
Age of head of household	55.63	60.53	59.65	67.10	67.94	62.09	54.62
Panel B : Years of education							
Head	3.51	8.09	6.56	2.72	7.04	5.16	5.57
Oldest member	0.17	6.70	3.54	2.13	6.66	3.38	4.29
Mother	4.39	8.29	9.77	2.60	7.94	7.39	5.28
Father	4.91	8.41	10.30	3.07	8.90	7.86	6.09
Panel C : Income, Expenditure and Bolivida (in Bs per month)							
Total per capita income excluding Bolivida	338.36	1172.94	916.76	360.77	686.76	644.80	493.68
Total per capita expenditure	299.65	901.33	1187.38	296.18	741.91	593.47	488.09
Children educational expenditure (ind level, def)	14.56	84.01	87.98	12.28	37.63	11.53	19.43
Children educational expenditure (ind level, not def)	36.02	204.51	215.55	30.19	92.66	29.18	47.73
Share children educational expenditure (hhold level)	0.04	0.07	0.05	0.05	0.04	0.03	0.04
Bolivida income[b] (eligible) at household level	72.99	68.00	83.50	93.27	81.78	91.19	0.00
Bolivida income (beneficiary) at household level	68.25	59.98	83.50	90.64	81.78	82.55	0.00
Panel D : Family structure							
Proportion of members 0-5	12.17	7.44	5.71	6.35	7.76	13.01	8.96
Proportion of members 6-13	28.38	29.51	30.54	30.74	27.19	27.15	27.94
Proportion of members 14-24	13.97	9.20	9.29	16.39	13.60	12.86	18.78
Proportion of members 25-64	22.66	30.42	29.22	15.6	25.40	23.36	44.33
Proportion of members 65-94	22.83	23.43	25.23	30.88	26.06	23.61	0.00

[a] For conciseness, standard errors are not reported. Total expenditure includes durable goods. All per capita variables are adjusted for equivalence scales. i=indigenous, ni=non-indigenous, m=multiethnic, f=female, m=male, def=deflacted.

[b] In 2001, the annual Bolivida pension was 816BS (68BS per month); 92 percent of the beneficiaries in the sample report receiving this amount and 8 percent report receiving just half of this (408BS). The average number of people per household receiving the pension is 1.36.

the pension are on average smaller, and live in households with less school-age children than their male counterparts. The age of the head reveals that most eligible women are not the heads of their households, but that most eligible men are in fact heads of household.

The second and third panels show that indigenous households are characterized by lower per capita income than their non-indigenous counterparts, and also by lower levels of education of their heads and oldest members. Interestingly, households with a multiethnic eligible member report having heads of household and oldest members with more years of education than household with a non-indigenous eligible member. Conversely, households with a non-indigenous eligible member report having children whose parents have more years of education than those with a multiethnic eligible member. It is also remarkable that the oldest member's years of education gender gap among indigenous is quite large. Last but not least, both eligible and non-eligible households allocate approximately 4.5 percent of their total income on children's human capital investments.

Finally, panel four reports information on family structure. The statistics show that the pension is received in households where approximately 28 percent of its members are school-age children.[2] Eligible households have on average a different family structure than non-eligible households. This is not surprising as non-eligible households have younger members among its members. Therefore, it is important that, besides the age and gender of the beneficiary, the model controls at least for household structure, household size, household income, and years of education of the oldest member, the head, and the parents.

5 The Bolivian Indigenous Population

Psacharopoulos and Patrinos (1994) point out that the identification of indigenous people in Bolivia is not simple. Social class and ethnic elements are very interrelated and, thus, are difficult to disentangle. In the context of this paper the definition of indigenous is critical for the analysis. Nonetheless many studies have used language spoken for classification, this condition is neither necessary nor sufficient to make a good categorization of indigenous people in the case of Bolivia (Ine and Maipo, 2003).

In general, information used to define indigenous people include a set of ethnolinguistic characteristics.[3] The surveys used in this paper include three questions aimed at identifying ethnic groups: (i) Do you consider yourself as belonging to an indigenous group?; (ii) What languages do you speak?; (iii) As

[2]Approximately 22 percent of the Bolivian children under the age of 14 live with a person that is eligible for receiving the pension.

[3]Ethnolinguistics refers to the study of language within ethnic groups or, more generally, to the relationship between language and culture.

a child, in what language did you first learn to speak?. The first two questions are only collected for household members that are at least 12 years old, and the third one is collected for all members in the household. The tabulation of these three questions for the sample is presented in Table 4 that shows that the percentage of indigenous population in Bolivia varies considerably upon the criterion selected for the classification of ethnic groups.

Table 4: Identification of the Bolivian indigenous population[a]

	Indigenous (%)	Non-indigenous (%)
Self identification	39.60	60.40
First language spoken	59.52	40.48
Language as a child	53.37	46.63

[a] Calculations include only people 12 years old and older.

Molina and Albó (2006) use the above three characteristics to construct an ethnolinguistic matrix for Bolivia. The criterion they use for classification is to weight self-perception more heavily among the three of them, as they consider this the most important one to determine ethnicity. I construct this matrix for my sample and use it to create an index of indigenism that classifies the population across 8 ethnolinguistic cohorts as Table 5 shows.

Table 5: Ethnolinguistic condition in Bolivia[a]

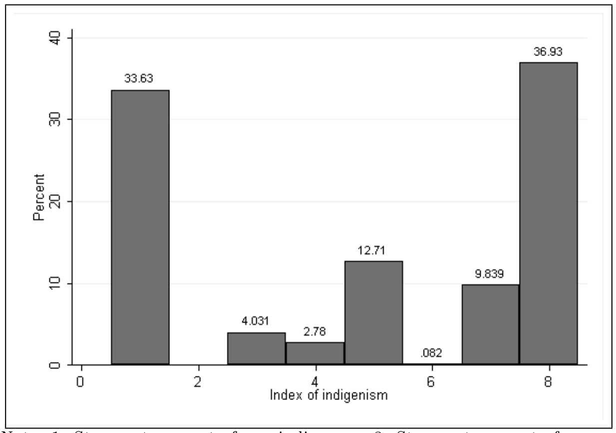
Index of indigenism	Self identification	First language spoken	Language as a child	%
1	no	no	no	33.7
2	no	no	yes	0.00
3	no	yes	no	4.04
4	no	yes	yes	2.74
5	yes	no	no	12.85
6	yes	no	yes	0.08
7	yes	yes	no	9.82
8	yes	yes	yes	36.77

[a] Calculations include only people 12 years old and older; 1=Strongest concept of non-indigenous; 8=Strongest concept of indigenous

The matrix is graphically shown in Figure 1, where it can be seen that nonetheless the Bolivian population is a multiethnic society, close to 70 percent of the population is located at either extremes of the distribution.

The approach adopted in the paper is to use the proposed index of indigenism to classify adult people among three exclusive ethnic groups as follows: indigenous (if the index is 1), non-indigenous (if the index is 8), and multiethnic (if the index lies between 2 and 7).

Figure 1: Ethnolinguistic condition in Bolivia



Note: 1=Strongest concept of non-indigenous; 8=Strongest concept of indigenous. Includes only people 12 years old and over.

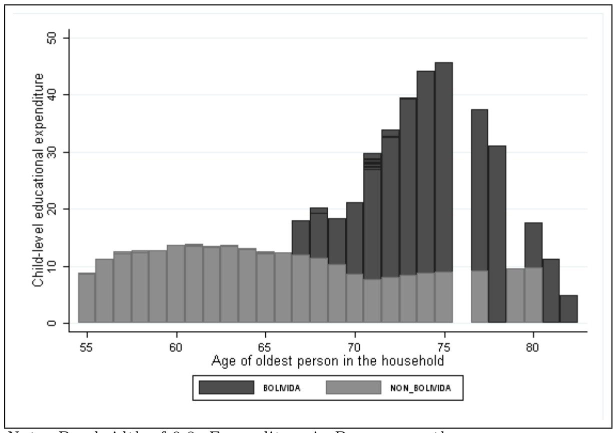
6 Estimating the effect of Bolivida on children's educational expenditure

6.1 Non-parametric exploratory analysis

I use locally weighted regressions to estimate the conditional mean of child-level educational expenditure as a function of the age of the head of the household, and the age of the oldest person living in the household. The smoothing is made separately for Bolivida families (i.e. households that have at least one recipient member) and non-Bolivida families (i.e. households that do not have any recipient member). These regressions are an attractive way to perform exploratory analysis as they allow for the data to determine the local shape of the conditional mean relationship without having to impose any restrictions regarding the underlying distribution of the errors.

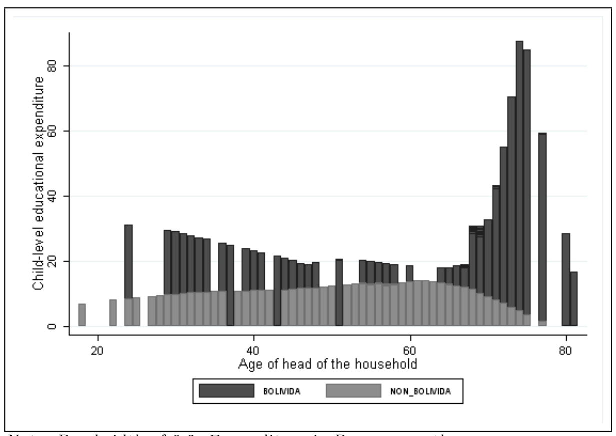
The estimation results are displayed in Figure 2 and Figure 3. Figure 2 shows that conditional on having at least one school-age child and one elder in the household, having a member that is at least 65 years old increases the pattern of expenditure on children's education if and only if the elder is a recipient of Bolivida. Figure 3 shows that, independently of the age of the head of the household, Bolivida families have always a relative higher conditional mean on children's educational expenditure. Moreover, it shows that when the head of the household reaches an age close to 65 years old, Bolivida households have a dramatic increase in the expenditure on children's education that is not observed in the non-Bolivida families.

Figure 2: Educational expenditure and oldest member age



Note: Bandwidth of 0.8. Expenditure in Bs. per month.

Figure 3: Educational expenditure and age of the head

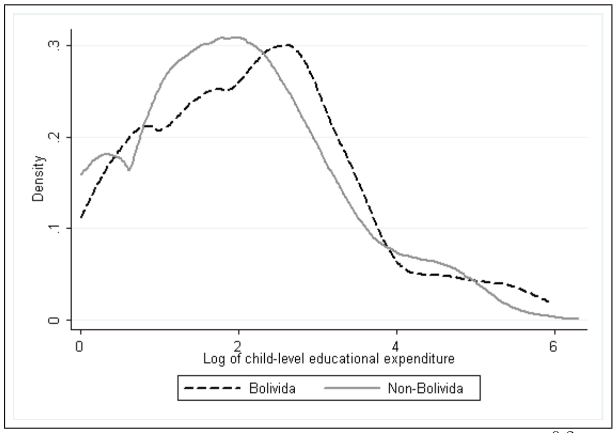


Note: Bandwidth of 0.8. Expenditure in Bs. per month.

As a final exercise, Figure 4 presents the child-level educational expenditure density functions estimated using a univariate Epanechnikov kernel.[4] It shows that the density of Bolivida families lies to the right of that of the non-Bolivida families. The Kolmogorov-Smirnov test of the equality of the distributions is rejected with a p-value of 0.003. In the next section these relations are parametrized to measure the effect of Bolivida on children's educational expenditure.

[4]Locally weighted regressions use the method of nearest neighbors for the estimation. Conversely, kernels use a well defined window width for the estimation.

Figure 4: Kernel density function estimation



Note: Bandwidth determined using Silverman's rule-of-thumb $0.9 * n^{0.2} * sd$.

6.2 Parametric approach

6.2.1 Identification strategy, model, and estimation

The introduction of Bolivida by the Government creates a discontinuity of who receives the pension at the age of 65. Hence, it can reasonably be assumed that women and men who are 64 years old (or are at the lower limit) are identical to women and men who are 65 years old (or are at the upper limit), except for the fact that the 65 year old receive the pension. Consequently, valid inference on the causal effects of Bolivida on children's educational expenditure can be obtained through the regression discontinuity design first introduced by Thistlethwaite and Campbell (1960). A limitation of the approach, however, is that only local effects are identified (i.e. treatment effects are identified only at the point at which the probability of receiving treatment changes discontinuously, Imbens and Angrist (1994)). Nevertheless, it is exactly this localized parameter the one of interest in this paper.

Two are the main assumptions made for the estimation. First, children who live with a Bolivida eligible member and a Bolivida near-future eligible member, do not differ in any time-variant or invariant unobservable way.[5] This implies that there are not child-level or household-level unobservable features affecting at the same time children educational expenditure and eligibility into the program. Second, household structure is exogenous in the model. This is a plausible assumption as Bolivida had just been resumed at the time of the analysis, and its continuation and fiscal sustainability were uncertain (Gamboa-Rivera, 2006), which implies that families had little time to change their structure (if

[5]The time-invariant assumption is required because panel data is not available.

ever at all) as a result of the program.

Under these assumptions, based on a Working-Leser Engel curve[6] (Working (1943) and Leser (1963)), the selected functional form restriction necessary to parametrically identify the effect of the Bolivia is the following reduced-form equation:

$$\begin{aligned} GEDUC_{ihc} = \beta_1 E_h^{if} + \beta_2 E_h^{im} + \beta_3 E_h^{mf} + \beta_4 E_h^{mm} + \beta_5 E_h^{nif} + \beta_6 E_h^{nim} \\ + \beta_7 Y_h + Z_h^{'}\Phi + X_i^{'}\Delta + \sum_{j=6}^{13} \Lambda_j 1_{j=c} + W_h^{'}\Theta + R^{'}\Omega + \upsilon_{ihc} \end{aligned} \tag{1}$$

where:

GEDUCihc=Log educational expenditure for child i, in household h, born in cohort c
E_h^{if}=Indicator for whether there is an indigenous eligible female in the household
E_h^{nif}=Indicator for whether there is a non-indigenous eligible female in the household
E_h^{mf}=Indicator for whether there is a multiethnic eligible female in the household
E_h^{im}=Indicator for whether there is an indigenous eligible male in the household
E_h^{nim}=Indicator for whether there is a non-indigenous eligible male in the household
E_h^{mm}=Indicator for whether there is a multiethnic eligible male in the household
Y_h=Log of total per capita income excluding the pension at household h
Z_h=Vector of household characteristics (explained in detail below)
X_i=Vector of child individual characteristics (explained in detail below)
1_c=Indicator variables for whether the child was born in cohort c
W_h=Vector of indicators for whether there is a member in the household aged k, where k goes from 55 to 74[7]
R=Interactions of departamental, province, and urban geographic divisions.

Child-level educational expenditure (GEDUC) includes registration fees, uniforms, texts, stationary supplies, copies, school board fees, transportation, contributions to school to pay teachers, and contributions to school to improve school infrastructure. The variable is deflacted by the educational consumer price index at departamental level (main geographic division, CPI[8], 1991=100)

[6]Original Working specification links household food expenditure to total expenditure; Leser presents a generalization of the function to all classes of goods. This functional form has been widely used in different contexts; see for instance Deaton (1997) for an augmented version with demographic characteristics of the household.

[7]Ideally, I would like to compare people on the age-range 64 to 65. However, this reduces the sample significantly. Therefore, following Duflo (2003), I use a larger interval on the age-range. An alternative specification using dummies on the age-range 60 to 69 yields similar results.

[8]Time series information of CPI disaggregated by type of expenditure is only available for 3 out of the 9 departmental divisions existing in Bolivia. Hence, departamental divisions with missing information are deflacted using a national average.

The Bolivida effects are captured by six eligibility dummies (E) based on ethnicity and gender of the recipient (more details are given further on in this Section). The excluded dummies in all cases are non-eligible members. The creation of mixed-ethnicity dummies allows to disentangle allocation of educational expenditure in multiethnic families, as close to 23 percent of the households in the sample are multiethnic (i.e. head of the household reports a different ethnic group than spouse).

The annual income and educational expenditures are adjusted to monthly averages as Bolivida payments were received throughout the year (SPVS, 2002), and information on the time of the receipt is not available in the survey. The income excluding the pension (Y) is additionally adjusted for economies of scale in the households.[9]

The vector of the household characteristics (Z) includes the logarithm of household size, the proportion of household members on age-ranges of 0-5, 6-13, 14-24, 25-64 and 65-94, head's age, head's years of education, and oldest member years of education. The vector of child-level individual characteristics (X) includes sex, mother's and father's ethnicity, mother's and father's years of education, indicator variables for the presence of the mother and/or the father, a dummy variable for whether the child was enrolled at school, and two dummy variables for whether the eligible member in the household is one of the parents of the child.[10]

The geographic interactions (R) are intended to capture time-invariant unobservables at departamental, province, and urban level. Children's age dummies control for child-cohort characteristics, and elderly dummies capture the presence of old members in the household.

The specification does not allow for non-linearity on per capita total income as kernel-weighted local polynomial smoothing evidences that the relation between the logarithm of children's educational expenditure and income excluding Bolivida is linear, conditional on having at least one school-age child and one elder living in the household (results available upon request).

In the applied microeconomics literature it is common to use expenditure, instead of income, to compute standard-of-living measures. Deaton and Zaidi (1999) give a good description of the theoretical and practical reasons why expenditure-based measures are preferable to income-based measures. Clearly, one of the main advantages of using expenditure variables is the smoother distribution that these variables have over time. Nevertheless, the reason why I use income instead of expenditure is to avoid potential pension endogeneity (i.e.

[9] Following World Bank (2003): Adult Equivalence Scales (AES)=1+0.7(adults -1) + 0.5 children.

[10] If neither the mother nor the father is present in the household, the information on ethnicity corresponds to the head of the household. An enrollment dummy is used (instead of an attendance dummy) to avoid potential measurement error in the variable as the survey was collected in November which coincides with the end of the school year.

expenditure is measured after the household receives the transfer and, hence, should not be included in the model).

The use of pension receipt dummies might be problematic as these variables could potentially be endogenous in the model. An alternative approach commonly used in this situation (see for instance Duflo (2003), and Hoddinott and Skoufias (2004)), is to estimate the impact of Bolivida conditional on having at least one eligible member. As mentioned in Section 3, there are not non-eligible members receiving the pension but there are indeed eligible members not benefiting from the pension.[11] It is not clear why differences between actual and estimated beneficiaries arise. It is possible, for instance, to devise a scenario in which deficiencies in the personal identification documentation system constrain eligible members to receive the pension. However, as the true reasons are unclear, using the eligible variable for the estimation is appealing as it avoids deriving in misleading inference if these unobservables are at the same time affecting educational expenditure and participation in the program.

It should be kept in mind, however, that this approach provides a measure of the intent to treat effect (i.e. the impact of having a potential beneficiary in the household) and not the effect of the treatment on the treated (i.e. effect of having an actual beneficiary). The advantage of using this methodology is that orthogonality between the eligible variable and the error term is assured by construction.[12] Nevertheless, it must be emphasized that the estimated coefficients are all biased downwards due to attenuation bias (i.e. the conditional mean of educational expenditure is calculated using households that indeed never received Bolivida).

Drawing on methodology of Hoddinott and Skoufias (2004) and Borooah and Iyer (2005), regressions are not estimated using a difference-in-differences (DiD) approach, but rather estimating the model year by year and comparing the conditional means. Two arguments support the use of this methodology. First, one of the main benefits of using DiD is to remove child-level time-invariant unobservables. However, as my data come from two cross sections, I cannot take advantage of this feature. Second, pooling my two cross sections for estimation is not straightforward since, as explained in Section 4, both surveys come from different sampling frames which implies that its comparison requires a reweighing scheme that is not possible to perform with the information publicly available (for discussions about sampling design see Pfefferman (1993), and Binder and Roberts (1993)).

[11] This would not be problematic if it will be the result of a random process, as the randomization will guarantee that there are not systematic differences in any other pre-treatment variables. However, if this responds to a non random process, inference would be contaminated beacuse the recipient variable would become endogenous in the model and create bias and inconsistency. For further details see Lee (2005)

[12] Furthermore, measuring the impact of Bolivida over beneficiaries is not straightforward, as I still would need to make sure that eligible-non-beneficiary members are an appropriate control group for recipient members.

6.2.2 Results

The results of the estimation of Equation 1 by OLS are shown in Table 6. The eligibility dummies by gender and ethnicity measure the impact of having at least one potential beneficiary in the household (as opposed to having at least one potential beneficiary in the near future) on children's educational expenditure, after controlling for child-level and household characteristics.

Columns (1) and (2) show that, before Bolivida, all conditional means are small (even negative for multiethnic females) and not statistically significant at the 5 percent level. Conversely, after the program, all eligibility coefficients become large and significant at the 1 percent level. Column (3) shows the differences between the conditional means of the eligibility dummies after and before the program. While the significance of these differences cannot be assessed (models cannot be nested as data for each of the years come from a different sampling frames), disparities in the allocations of pension income across gender and ethnic groups are evident. Undoubtedly, eligible women allocate more of their pension income in the accumulation of children's human capital investments as compared to eligible men. For instance, in 2000, having an indigenous female eligible member in the household has a statistically non-significant impact of 10.1 percent over children's educational expenditure. In 2001, having an indigenous female eligible member has a statistically significant impact (at the 1 percent level) of 99.8 percent on children's educational expenditure. Hence, assuming that all the assumptions in Section 6.2.1 hold, the difference attributed to Bolivia is 89.7 percent increase in the educational expenditure. Therefore, having an eligible female in the household increases educational expenditure by 90 and 124 percent on average within the indigenous and non-indigenous cohorts, respectively. Conversely, having an eligible male only increases these investments by 16 and 55 percent within the indigenous and multiethnic cohorts, respectively.

This suggests that Bolivida transfers to both, women and men, lead to improvements on children's educational expenditure, whereas the effect is stronger in households where the pension is received by women. Regarding to the ethnicity of the recipient, patterns of allocation are unambiguous; conditional on program participation, non-indigenous and multiethnic beneficiaries have a stronger impact on educational expenditure than their indigenous counterparts.

To this point, I assumed that the sex of the child affected the conditional mean of educational expenditure exclusively as a covariate. Next, I estimate the average effects of eligibility for girls and boys, and run a DiD for elibility and sex-child cohorts. Columns (4) to (7) show the baseline specification estimated separately for the sub-samples of girls and boys (note that approximately 53 percent of school-age children in the sample are boys). These coefficients provide indirect estimates of the interactions between the eligibility dummies, and sex of the child. Before the program, the point estimates are never statistically

Table 6: Effect of Bolivida on children's educational expenditure: OLS Regressions[a]

	Year 2000	Year 2001	Diff	Year 2000	Year 2001	Year 2000	Year 2001	Year 2000	Year 2001
	All			Girls		Boys		DiD (1=boy)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
E_i_f	0.1012	0.9981***	0.8969	-0.5502	0.3432	0.3254	1.2591**	0.8756	0.9160*
	(0.3847)	(0.3806)		(0.5394)	(0.4158)	(0.4416)	(0.5029)	(0.6762)	(0.5287)
E_i_m	0.7467*	0.9090***	0.1623	-0.3239	0.4521	0.6982*	1.3278***	1.0221	0.8757*
	(0.3886)	(0.3265)		(0.6028)	(0.3652)	(0.3957)	(0.3739)	(0.6609)	(0.4832)
E_ni_f	0.1423	1.3792***	1.2369	-0.5457	0.9446**	0.0846	1.5613***	0.6303	0.6167
	(0.5012)	(0.3471)		(0.6360)	(0.3769)	(0.5562)	(0.4653)	(0.8114)	(0.5389)
E_ni_m	0.5328	0.9685**	0.4357	-0.1812	0.0694	0.5092	1.3718***	0.6904	1.3024**
	(0.4558)	(0.4188)		(0.6324)	(0.4801)	(0.4462)	(0.4639)	(0.7541)	(0.6147)
E_m_f	-0.0696	1.1492***	1.2188	-0.5050	0.5795	-0.0082	1.0957**	0.4968	0.5161
	(0.5120)	(0.3978)		(0.7960)	(0.5134)	(0.4560)	(0.4489)	(0.8734)	(0.6329)
E_m_m	0.5219	1.0686***	0.5467	-0.5748	0.6905*	0.6433	1.4252***	1.2181	0.7346
	(0.4377)	(0.3214)		(0.7210)	(0.3760)	(0.4088)	(0.4396)	(0.7787)	(0.5200)
lytsinbono_i	0.3140***	0.3331***		0.4931***	0.3434***	0.2594***	0.3696***	0.3142***	0.3344***
	(0.0504)	(0.0530)		(0.0722)	(0.0678)	(0.0559)	(0.0618)	(0.0506)	(0.0524)
sex of child	-0.0331	0.2227***							
	(0.0911)	(0.0755)							
Observations	557	776		268	362	289	414	557	776
R^2	0.5458	0.4847		0.6134	0.5573	0.6236	0.5047	0.6184	0.5335
E_i_f=E_i_m (p-value)	0.083	0.598		0.483	0.699	0.749	0.924	0.174	0.222
E_m_f=E_m_m (p-value)	0.060	0.877		0.263	0.770	0.552	0.695	0.172	0.840
E_ni_f=E_ni_m (p-value)	0.409	0.073		0.428	0.027	0.762	0.212	0.385	0.025

[a] Standard errors in parentheses robust to correlation of residuals within households and heteroskedasticity; i=indigenous, ni=non indigenous, m=multiethnic; f=female, m=male; lytsinbono_i=log of per capita income net of Bolivida; *** p<0.01, ** p<0.05, * p<0.1. Columns (8) and (9) eligibility dummies correspond to interactions with sex of the child; girls are the excluded dummy.

significant (except for indigenous males), and report negative for the girls sub-sample and very small for the boys sub-sample. Conversely, after the program, the coefficients become all large and highly significant but only for the boys sub-sample. Therefore, the highly significant and large coefficients reported in column (2) are mainly driven by educational investments on boys. This also implies that the educational expenditure allocation is dependent on the sex of the child.

Columns (8) and (9) present the estimations of an alternative specification which includes interactions of sex of the child and all the independent variables on equation 1 with its corresponding main effects.[13] This specification corresponds to a DiD and allows to directly estimate the difference in educational expenditure among girls and boys in eligible households, versus the difference in educational expenditure among girls and boys in non-eligible households. Results show that there exists discrimination against human capital investments on girls, and that this is particularly pronounced within indigenous families, and non-indigenous males.

More interestingly, however, is the fact that there seem to exist culture-based decision rules in the allocation of human capital investments. In particular, women in indigenous and multiethnic families have less power in the allocation of Bolivida resources upon children's educational expenditure as compared to women in non-indigenous families. Columns (2) and (9) show that, conditional on program eligibility, human capital investment decision rules within indigenous and multiethnic families are very similar across gender groups; however, this is not the case within non-indigenous families. This implies that similar resources are being allocated almost identically by women and men within indigenous and multiethnic families, and that this allocation is specific to the gender of the recipient within non-indigenous families.

Consequently, my claim is that cultural factors play a key role in the allocation of resources upon children's human capital investments as they derive in different decision-making rules. More specifically, indigenous and multiethnic families follow decision rules closely related to patriarchal family structures (that limit the power of women to allocate their resources as compared to non patriarchal family structures), and non-indigenous families follow a bargaining decision process. This claim can be formally tested using a Wald test for simple linear hypothesis under the null that the marginal propensity to use money from the pension (captured by the eligibility dummies) differs within ethnic groups depending on the gender of the recipient. The p-values associated with the statistics are shown in the lower part of Table 6. Columns (2) and (9) confirm that, conditional on program eligibility, only point estimates for the non-indigenous cohort are statistically different (at the 7 percent level). Col-

[13]Initially, I estimated an specification that introduced only the interactions of sex of the child and eligibility dummies with its corresponding main effects. However, results showed that the slope of the relation between educational expenditure and the control variables needed to vary across sex-child cohorts.

umn (7) shows that investments on boys' education, within ethnic groups, is gender-independent (i.e. eligible women and men allocate their investments in a very similar way). Conversely, column (5) shows that the investments on girls' education, within ethnic groups, is gender-dependent; in particular, while indigenous women and men allocate their investments in a very similar way, non-indigenous eligible women and men allocate their investments in a very dissimilar way (with non-indigenous women prioritizing investments on girls' education).

Finally, it would be interesting to run the regressions by child-ethnic cohorts. However, this is unfeasible since, as mentioned in Section 6, only one of the three variables used to create the individual-level indigenous variable was collected for children 12 years old and under (i.e. first childhood language). Hence, using this variable to classify children across ethnic groups contradicts the criteria previously used to categorize adult ethnic groups. On the other hand, imputing the ethnicity of one of the parents (or even of the head of the household) as children's ethnicity might be conflictive and misleading, as not a trivial number of parents categorized as indigenous report having children whose first childhood language is not an indigenous language.

6.2.3 Extensions

Table 7 presents alternative specifications and estimation methods to assess the robustness of the previous results, and extend some of the previous findings. First, potential endogeneity of per capita income excluding the pension is addressed by excluding this variable as control in the original specification. However, if per capita income net of Bolivida is a relevant variable to include in the model, this last specification leads to omitted variable bias. Second, instrumental variables are used as a potential method to deal with endogeneity, and possible measurement error in the income variable. Third, a censored model is estimated to account for the fact that not a trivial proportion of school-age children in the sample report zero educational expenditures. Fourth, the marginal tobit coefficients are broken down using a McDonald Moffitt decomposition (McDonald and Moffitt, 1980) which provides important insight as to whether Bolivida is impacting children that were already enrolled in school, or children that were not enrolled in school, before the program started. For conciseness, only estimates for the year 2001 are reported (results for the year 2000 are available upon request).

To this point, reported parameter estimates have been based on a specification that assumes that the logarithm of the per capita income excluding the pension is exogenous. Nevertheless, if errors are not orthogonal to this covariate, then previously estimated coefficients are inconsistent and biased. Column (1) presents the original specification to be used as the benchmark model. Column (2) shows the estimated coefficients excluding this variable as cofounder. It can

Table 7: Effect of Bolivida on children's educational expenditure: Extensions[a]

	Original specification (1)	Excluding income net of Bolivida (2)	IV for income net of Bolivida (3)	Tobit coefficient (4)	Tobit marginal effect for all observations (5)	Tobit marginal effect for observations above the limit (6)	Tobit marginal effect for observations at the limit (7)
E_i_f	0.998**	1.039**	1.025**	1.012*	0.766	0.609	0.145
	(2.622)	(2.639)	(2.684)	(2.440)			
E_i_m	0.909**	0.848*	0.869*	0.966**	0.645	0.512	0.122
	(2.784)	(2.435)	(2.473)	(2.807)			
E_ni_f	1.379**	1.396**	1.390**	1.421**	1.057	0.839	0.2000
	(3.974)	(3.814)	(3.910)	(3.862)			
E_ni_m	0.9689*	0.959*	0.962*	1.003*	0.616	0.490	0.1167
	(2.313)	(2.325)	(2.327)	(2.288)			
E_m_f	1.149**	1.316**	1.258**	1.190**	0.706	0.560	0.133
	(2.889)	(3.293)	(3.222)	(2.779)			
E_m_m	1.069**	0.949**	0.991**	1.103**	0.708	0.562	0.134
	(3.325)	(2.704)	(2.694)	(3.288)			
lytsinbono_i	0.333**			0.362**			
	(6.289)			(6.061)			
sex of child	0.2227***	0.1990**	0.2072**	0.2351***			
	(0.0755)	(0.0797)	(0.0803)	(0.0812)			
Observations	776	776	776	776	776	776	776
R^2	0.485	0.430	0.461	-	-	-	-
E_i_f=E_i_m (p-value)	0.598	0.170	0.350	0.584			
E_m_f=E_m_m (p-value)	0.877	0.630	0.822	0.994			
E_ni_f=E_ni_m (p-value)	0.073	0.069	0.068	0.075			

[a] t-statistics in parentheses robust to correlation of residuals within households and heteroskedasticity; i=indigenous, ni=non-indigenous, m=multiethnic; f=female, m=male. lytsinbono_i=log of per capita income net of Bolivida; *** p<0.01, ** p<0.05, * p<0.1.

be seen that the basic results remain unchanged.

An alternative approach is to use instrumental variables (IV), which also allows to deal with the bias generated if the income variable is measured with error. Column (3) reports the IV estimates when the total per capita income excluding Bolivida is instrumented using a vector of variables that includes head's occupation, a dummy for whether the head is in the labor market, and a dummy for whether the head holds a secondary job.[14] I am well aware that these instruments might be problematic. However, it is not easy to find convincing instruments for this variable, as it is hard to argue for the exclusion restrictions (i.e. instruments should have no direct impact on children's educational expenditure). Nevertheless, it is reassuring that basic results remain unchanged, and that the tests of the validity, and relevance of instruments cannot reject the null.[15]

Among the school-age children included in the sample, 11 percent report having zero educational expenditures. This implies that, if the dependent variable is censored, then OLS might not be the best method to use for the estimation. Column (4) reports the eligibility coefficients estimated using a tobit regression model. The reported t-statistics are bootstrapped versions as non normality in errors is a specially difficult problem in this setting (tobit model is only identified if the assumption of normality is fulfilled, Segelman and Langche (1999)). The estimates closely approximate the previous findings.

Next, I use the McDonald Moffitt technique for censored models with lower bounds to break down the impact of Bolivida. This methodology allows to decompose the marginal effects of the tobit coefficients into the portion caused by observations above the limit (children that before the Bolivida were already enrolled in school) and observations at the limit (children that before the Bolivida were not enrolled in school). Results are shown in columns (6) and (7) where it is evidenced that the program is mainly affecting the educational investments of children that were already enrolled in school before the program started. The impact of the pension on children at the limit is minimal, as only few of them actually get to move across the limit as a result of the program.

Finally, the Wald tests show that, at the 8 percent, the previous linear hypothesis testing results hold, i.e. conditional on program participation, only the non-indigenous cohort has gender-specific outcomes in the pension allocation decision-making process.

[14]The occupational dummies correspond to the head, if appropriate, and to the spouse and/or the oldest member in the household otherwise.

[15]Test of overidentifying restrictions (test of instruments validity, i.e. uncorrelated with the error term and correctly excluded from the estimated equation) in the presence of heteroskedasticity is 6.01 with p-value 0.05. Anderson canonical correlation LR test (test of instrument relevance, i.e. equation is identified) is 36.03 with p-value of 0.00. The R^2 statistic for the instruments in the first stage (test of weak instruments) is 0.4421. For details see Baum, Schaffer, and Stillman (2003).

7 Conclusions

The Bolivida old age pension program led to an increase on children's human capital investments. Having at least one eligible elder in the household increases children's educational expenditure between 16 and 124 percent, depending on the gender and ethnicity of the recipient. However, statistical significance of these estimations cannot be assessed as data come from different sampling frames.

Perhaps unsurprisingly, the paper finds that female recipients are more effective at promoting children's human capital investments than their male counterparts. The increases in educational expenditure caused by female recipients are around 90 and 125 percent; and only around 16 and 55 percent among male recipients. More interestingly, patterns of allocation of human capital investments by ethnicity confirm that pension income in the hands of indigenous has a smaller impact on educational expenditure than analogous income in the hands of non-indigenous and multiethnic recipients.

The main contribution of the paper, however, is the evidence found that cultural factors play a key role in the intrahousehold children's human capital resource allocation process. Decision rules determining patterns of educational expenditure in Bolivia seem to be culture-dependent. In particular, whereas indigenous and multiethnic families allocate human capital investments following a unitary or dictatorial decision-making process, non-indigenous families use bargaining decision-making processes. The exact allocation mechanisms across ethnic cohorts, however, have not been analyzed in the paper and remain open to future research.

The paper finds significant differences between educational expenditure gender gaps for eligible and non-eligible households within indigenous families; this suggests that among indigenous there is a tendency to prioritize educational investments on boys as compared to the one on girls.

A decomposition of the marginal effects of Bolivida on children's educational expenditure reveals that the program is mainly affecting the human capital investments of children that were already enrolled in school before the program started. Conversely, the impact of the pension on children not previously enrolled in school is very limited, as this source of income in their households might be being allocated to other priority expenses if binding budget constraints in their household are the reason why they are not enrolled in school.

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