Educational Gap and Family Structure in Uruguay

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Abstract
In this paper we use household survey data to study the determinants of children's educational achievement in Uruguay. As an indicator of this educational achievement, we build the "educational gap" which is the difference between expected years of schooling of a child and actual years of schooling. Among the determinants, we introduce indicators of family environment, focusing on the impact of the parents' marital status on their children educational attainment. In particular, the results suggest positive infuence of having married parents on daughter's educational outcomes, after controlling for household background variables such as parents' education, income percapita, wealth and number of children.

Resumen
En este trabajo, utilizamos datos provenientes de la encuesta de hogares para estudiar los determinantes de los logros educativos de los niños en Uruguay. Como un indicador de esos logros educativos, construimos un índice de rezago escolar que consiste en la diferencia entre los años de educación que un niño se espera que tenga de acuerdo a su edad y los años de educación que realmente tiene. Entre los determinantes, introducimos indicadores del medioambiente familiar, centrándonos en el impacto del status marital de los padres sobre los logros educativos de los hijos. En particular, los resultados sugieren una influencia positiva de tener padres casados sobre los resultados educativos de las hijas, luego de controlar por variables características del hogar tales como la educación de los padres, el ingreso per cápita, la riqueza y el número de hijos.

JEL classification: J12, J13, C14, C34, I21
Keywords: censored data, treatment evaluation, education, family instability, cohabitation

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1. Introduction

Previous investigations analyse the possible determinants of schooling gap [1] -a censored variable- but methodological problems arise. When data are censored, OLS regression can provide misleading estimates. But also using traditional maximum likelihood methods for censored models, like Tobit, could not be appropriate because of the lack of normality of the error terms or due to the existence of heteroskedasticity. In order to overcome these requirements of strong distributional assumptions, many semiparametric alternatives were developed. To examine these methodological issues, this paper applies and compares different methods to estimate the determinants of the educational gap.
There are several papers that examine the possible relationship between family structure and child well-being (Axinn and Thornton, 1993; Brown, 2004; Manning and Lichter, 1996; McLanahan, 1985; Raley and Wildsmith, 2004). Brown (2004) provides an extensive summary on the emerging literature on the effects of cohabiting families on children residing with them, and suggests children's academic performance is negatively associated with cohabitation relative to marriage. Brown sets the hypotheses that both the impermanence of cohabiting unions and their incomplete institutionalization (unclear family roles, rights, and obligations) set the stage for a family environment that may undermine child development.

We use data of 2001 from the Continuous Households Survey (Encuesta Continua de Hogares, ECH) and employ the schooling gap as one proxy of child well-being. In the case of two biological-parent houses in Uruguay, the fact of having married parents seems to be a significant determinant of the daughters' schooling outcomes, while it does not impact on son's achievements. In fact, having married parents contributes to the decrease of the girls' educational gap. For example, taking into account the results of the Symmetrically Censored Least Squares Estimator, the fact of having married parents reduces the educational gap of the daughters by 0.123. The ECH for year 2001 was selected because of two reasons: it contains recent data and it is previous to the year 2002 in which Uruguay experienced one of the greatest adverse economic shocks of its history.

2. Background

In this section, we discuss the previous literature. We observe some main trends in estimating the determinants of school achievement. One trend is to estimate the educational gap using OLS. Take for example the case of Behrman, Birdsall and Székely (2000) who estimate the effects of parental education and household income on the schooling gap of their children. With respect to Brazil, they find that maternal education has a slightly stronger effect than paternal education in 1995. But the authors use OLS, despite the fact that the gap measure is censored. Also Andersen (2001) estimates the schooling gap of the teenagers by OLS. He finds that schooling gap is negatively related with the household income per capita, with the maximum of father's and mother's education, with the age of the head of household at the birth of the teenager and with the fact of living in urban areas. On the contrary, the schooling gap is positively related with the presence of a younger sister or brother, and with the fact of being a teenager that is not the son or the daughter of the head of the household.

Other approach is to use Probit models. For instance, Fishback and Baskin (1991) compare the level of educational achievement between black and white children in the first part of the twentieth century. They assume that child's literacy (a variable which was valued at one if household members said the child could write and zero if (s)he could not write) is a function of school inputs, household educational inputs, the time devoted to learning, and characteristics like de child's age and sex. Lacking direct measures of income and wealth, they include a variety of indirect measures: the age of the head and the spouse to measure their position in the life cycle; home ownership, and an index of occupational status. They estimate the effect of each determinant using maximum likelihood probit analysis. The analysis shows that the largest contributor to the black-white literacy gap was the difference between the educations of black and white parents. The estimation results also show that the length of the school term was a key school input for developing basic literacy, and the higher parents occupational

1. Schooling gap is the relative lag behind the age-appropriate schooling level. In other words, schooling gap is defined as the difference between expected years of schooling (number of years of schooling under assumption of an initial enrollment age of 6 and completing one grade per year without grade repetition) and actual years of schooling, as a proportion of expected years of education.

status contributed to the child's literacy too.

Another approach has been to use a Tobit procedure. For example, Psacharopoulos and Arriagada (1989) estimate educational attainment among 7 to 14 years old employing a Tobit model. They find that maternal education has a stronger effect than paternal education on boys and girls taken together. However, boys and girls are pooled, so it is not possible to make any further gender comparisons. Margo (1987) specifies a model of school attendance and constructs an equation as the outcome of a household utility maximization.

Parents derive utility from consumption of markets goods and home production by household members and from their children's schooling. How frequently a child attends school depends on the characteristics of the parents and the child; on the availability of schooling, quantity and quality; and on the returns to schooling compared with other uses of the child's time, which may vary with the household location. The dependent variable is the number of months of school attended in the census year. Because many children did not attend school the dependent variable is censored at zero, and Tobit analysis is used. Margo's results show that the presence of a child under age 5 in black families lowered school attendance among older siblings, presumably by increasing parental demands on the sibling's home time. Margo also finds that longer school terms and smaller class sizes encourage children of both races to attend school more frequently, but the effect was larger among blacks.

In addition better-trained teachers also increase attendance in the black schools. Finally Margo observes that urban children of both races attend school significantly longer than rural children, and the effect is larger among blacks. Saha (2005) also use a Tobit model, focus on one age cohort and restricts the analysis to children who turn 15 during the survey years. The explanatory variables used by Saha are: number of younger siblings, household size, parental age, rural dummy, income, mother's education, father's education. Saha found that maternal and parental education, and the household income were positively correlated with the educational attainment, while the household size, the fact of living in rural areas and the number of younger siblings were negatively correlated with the school outcomes. Also, Saha found that the maternal effects differed by household type: maternal education in two-parents household widened the gap between son and daughters educational attainment, and, in sharp contrast, maternal education in female-headed households contributed to the decrease in the gender gap.

Finally, consider the treatment of endogeneity in previous literature. Case and Deaton (1999) examine the relationship between school inputs, particularly the pupil-teacher ratio, and various measures of educational outcomes, including educational attainment, enrolment rates, the reason for not being in school, educational expenditures, and test scores.

They present the results of a series of regressions in which the pupil-teacher ratio (or the presence of other facilities) is an explanatory variable. Among the other controls are age, urbanization, sex, and various measures of family background, such as whether the household is headed by a woman, household size, the educational attainment of the head, and the logarithm of total expenditure per capita. They think of head's education as both a direct input into the educational process and a measure of household resources. In one of the regressions (estimated using OLS and Two Stage Least Squares with robust standard errors), the dependent variable is educational attainment measured as years completed. The reason to introduce Two Stage Least Squares is this: the pupil-teacher ratio for Blacks may be affected by household characteristics. Thus the estimated effects of the pupil-teacher ratio may be coming from the influence of unmeasured household characteristics. They consider possible instrumentation using the racial composition of magisterial districts (they checked that pupil-teacher ratios can be predicted by racial composition). They found that the TSLS results were very similar to the OLS results. Case and Deaton show that gender and household characteristics have important effects in the regressions. Black female students have on average about half a year of educational attainment more than the Black male students, and among Black students there are the expected positive effects of household resources and of education of the household head. Head's education is a strong predictor of educational attainment among both Blacks and Whites. Controlling for household background variables, they find strong and significant effects of pupil-teacher ratios on enrollment, on educational achievement, and on test scores for numeracy. Bjorklund et al (2005) focus on children who live with both biological parents and analyze whether marriage -in comparison with cohabitation- confers any educational advantages to children, for the case of Sweden. They use a natural experiment, namely the marriage boom in Sweden in the last two months of 1989, created by the reform of the widow's pension system (those who were married by the end of 1989, would be entitled to widow's pension if their husband died), to identify the causal effect of marriage on child outcomes. This experiment enables the authors to compare educational outcomes for children whose parents married in November and December 1989 to those of children whose parents were already married and to those of children whose parents continued to cohabit. They find that children whose parents married in the end of 1989 had similar educational outcomes than children of cohabiting parents which suggest some doubts on the direct causation of legal marriage on children educational outcomes.

For comprehensive handling of the problem of endogeneity, we should refer to Francesconi et al. (2006). They analyse the impact on schooling outcomes of growing up in a family headed by a single mother. They test the hypothesis that a family without father and/or mother in Germany is associated with worse educational outcomes, and employ propensity score matching models, mother fixed effects and quasi experimental models, and models based on comparisons between individuals whose fathers died, divorced, or remained married. The principal schooling outcome analysed is whether an individual has educational qualifications to university entrance level. They find that although almost all the point estimates indicate that non-intactness of family structure has an adverse effect on schooling outcomes, confidence intervals for estimated effects are wide so the data are consistent with the impact of family structure being zero as well as adverse. About the possible presence of endogeneity, they argue that there's disagreement about whether the family structure is causal: lone parenthood may be correlated with other socioeconomic disadvantages, and so inferior outcomes may arise from (potentially unobserved) factors other than a parent's absence.

The following section gives details about the data used in the empirical estimation. In section 4 different methods for censored regression models are defined, while Section 5 presents the results and makes comparisons between the different methods. The final section discusses conclusions, limitations of the approach adopted here and points out some issues for further research.

3. Data

We use cross-sectional data of the year 2001 from Continuous Households Survey (Encuesta Continua de Hogares - ECH) which includes socio-economic information of people and households. ECH is conducted by the National Institute of Statistics (Instituto Nacional de Estadística - INE) of Uruguay and is an urban representative sample with a total sample size of 57394 observations. We take into account only sons and daughters with ages which fall in the interval [8,14] and who live with both biological parents (a sample size of 4067 observations). We focus in the interval [8,14] because -as observed in Table 1- the proportion of children with positive schooling gap is nearly zero for children of 6 and 7 years (the initial enrolment age in Uruguay is usually 6), and children with 14 years old or above are considered to be part of the labor force by the ECH.

Table 1

age	percentage of children with education-gap = 0
6	99.50
7	99.17
8	67.77
9	63.49
10	64.76
11	64.44
12	59.64
13	60.67
14	54.96

3.1 The dependent variable

The dependent variable, educational gap of the sons and daughters, indicates the relative lag behind the age-appropriate schooling level. It is computed as (under the assumption of an initial enrolment age of 6):

$$\text{educ\_gap} = \left(\frac{age - 6 - years\_of\_schooling}{age - 6} \right)$$

In other words, educational gap is defined as the difference between expected years of schooling (number of years of schooling [2]) a child would have under assumption of an initial enrolment age of 6 and completing one grade per year without grade repetition) and actual years of schooling, as a proportion of expected years of education.

3.2 Family structure as a regressor

As stated in Section 1, previous research for other countries suggests some linkage between family structure and children school engagement. For this reason, this empirical application introduces -as a regressor for children educational gap- parents' marital status: a binary indicator variable which takes the value one if the parents are married, and zero in the case of cohabitation. This paper concentrates in these two types of family structure because of the increasing rate of cohabitation during the last thirty years (Brown, 2004; Raley and Wildsmith, 2004; Cid, Presno and Viana, 2004; Manning and Lichter, 1996). As an example of this trend, consider that, in Uruguay, the proportion of informal unions in the total of couples rose from 7.65 percent in 1963 census to 16.45 percent in 1996 census and this augmentation occurred basically in the younger age groups. For example, for the 15-19 age group this ratio is multiplied by more than three times in this period.

Introducing family explanatory variables pretends to stimulate further research on this topic which could be fruitful to improve our knowledge of the causes of the low educational achievements in our country relative to developed countries. Filgueira, Filgueira and Fuentes (2003) states that Latin

2 The variable "years of schooling" is measured as years completed both in primary and secondary school plus one. The reason to add the value "one" is that the survey (ECH) used does not provide information about the child's birthday and this is a problem in order to estimate the "schooling gap". In our country, a child is able to start primary school if (s)he is at least 6 years old before the 10th of May. Take for example that one child with age 7 could claim in the survey that she has 0 year completed of schooling (thus a schooling gap of 100 percent). But her birthday is the 20th. of May so she started primary school at 6 years old (as early as she was able to) but the survey was executed on August when she is 7 and ECH says that she has 0 year completed and in fact her educational gap is zero. To sum up, adding the value "one" to the years completed at school, we are able to guarantee that every child with an educational gap greater than zero really has a gap. It is important because, precisely, we wish to analyse the determinants of this gap.

American countries have invested considerable economic resources in order to improve their educational supply, particularly in terms of school infrastructure, human and material resources, and innovative strategies to make schools more appealing to students. However, children academic performance remains a daunting challenge because of great drop-out rates, low grade completion and low schooling rates.

Filgueira et al. observe that the key to this failure seems to be not on the supply side but on the demand side: little is known regarding how and why the targeted population behaves as it does, and thus, the primary focus of diagnosis and policy should go from supply to demand. And, precisely, in the demand side of education, the family could play a crucial role.

3.3 Other regressors for the educational gap

The explanatory variables also included in this paper are:

log household income per capita, entered linearly and quadratically: Brown (2004) states that poverty is closely linked to different features of child well-being like school outcome. Saha (2005) believes that, in the presence of credit constraints, poorer families are less able to pay for the direct costs of education, such as books and transportation; and poorer families are also more likely to send children into the work force to supplement family income;

subjob: indicates if the principal job of the mother and/or the father is an informal one (e.g., without social security in case of being ill or unemployed). An informal employment could imply job instability and thus could create worse household environment for children school engagement.

mother inactive and father inactive: these dummy regressors indicate if the mother or the father are not employed and not seeking for an employment (for example, the mother spends her time studying in order to complete her undergraduate degree, and looking after the children and the house). With these variables we intend to measure the closeness of the parent-child relationship. Datcher-Loury (1988) observes that greater child care time of highly educated but not of less well-educated mothers significantly raises offspring years of study.

mother education and father education, entered linearly and quadratically: each one shows the number of completed and approved years of education since Primary School. It is expected that, for example, children whose parents have a university degree are more engaged with school than those whose parents only have few completed years of Primary School (see Brown, 2004).

quantity of children with age below 15: socioeconomic literature (see, for instance, Becker [1988], Simon [1998] or Saha [2005]) suggests a negative relationship in the short run between number of children and parents resources per capita which could imply worse school engagement.

quantity of people with age above 59: the presence of grandparents in household composition could have a positive effect on children's school outcomes because of the greater guidance and supervision or the spill over effects of more contact with the adults. In the same sense, this research included home-aid: a binary regressor coded one for the presence of an additional adult at home which helps with homecare (laundry and meal preparation, etc.).

private children education: a proxy of education quality. Heckman and Rubinstein (2001) quote the conjectures that the decline in discipline in some public schools could be a major source of their failure on children's school engagement, and that the greater effectiveness of some private schools could come in producing more motivated and selfdisciplined students.

scholarship: a binary regressor with the value one in the case of a child with income from a scholarship. It could be expected that someone with a grant should show better academic performance.

public job: a dummy variable coded one if the mother and/or the father have a public job, and zero otherwise. The hypothesis is that parents public job could be an indicator of economic stability, thus it can in‡uence positively children education.

remittances: this regressor pretends to capture the conjecture that child human capital decisions could be positively related with the fact of having a family member working abroad. McKenzie and Rapoport (2005) observe that previous research has suggested the potential of remittance income to improve access to education of the poor. They also state that a new literature has emphasized a possible link between expectation of future migration and current schooling decisions: education is needed to migrate, and since income abroad is much larger than at home, this raises the potential returns to schooling.

number of people with income at home: the hypothesis is that the larger number of individuals at home with a personal income (salary, profits, pensions, etc.), the greater the closeness of children to real world: the offspring experiment the need of being educated to cope with the market.

absolute wealth: the ECH provides information about thirteen goods that show welfare and each household could have: hot water heater, electric tea kettle, refrigerator, color television, cable TV service, VCR player/recorder, washing machine, dishwasher, microwave, computer, internet connection, automobile for personal use, telephone service. These goods could show different levels of wealth. For each good i, we have constructed a dummy variable d_i which takes value 1 if the house has this good or service, and 0 otherwise. Then we have developed the index

"wealth"= $\frac{1}{13}\sum_{i=1}^{i=13} d_i$

relative wealth: besides the previous wealth index which is an absolute indicator of wellbeing, we have built also an index of relative wealth using the goods information of the ECH. For each good i, we have constructed a dummy variable d_i which takes value 1 if the house has this good or service, and 0 otherwise. Thus, we have developed this indicator in two steps:

1st) the sample mean of each d_i is calculated;

2nd) "relative wealth index" $= \frac{\sum_{i=1}^{i=13}\left[1-mean(d_i)\right] d_i}{\sum_{i=1}^{i=13}\left[1-mean(d_i)\right]}$

(therefore, as an indicator of relative welfare, the formula above shows that greater average of people in the sample having a particular good implies less relative welfare).

Besides quadratic and interactive forms of these explanatory variables, we also included among the regressors dummies with the purpose of controlling potential effects of population density and economic situation of the region of residence, or the possible incidence of the sector of the economy in which the parents are employed.

3.4 Summary Statistics

The Continuous Household Survey (ECH) of the year 2001 provides information of 6.384 children in the interval of age [8,14]. Among them, 4.067 are children living with both biological parents (so they represent a 64 percent of the children of this interval). Other 1.479 children live with his/her biological father or mother (alone or with a step-father/mother).
Other 665 children claim to live in a household where the grandfather/grandmother is the person withmore authority in the house (the "chief" in terms of the ECH). Other 114 children claim to be only "other relatives" while 59 children describe themselves as no relatives at all.

Table 2

Descriptive statistics for daughters living with both biological parents in the interval of age [8,14] * means are statistically different at 1%; ** at 5%; * at 10%**

	Cohabit	Married		
	(258 obs)	(1797 o.)	Difference	p-value
father age	43.09	43.13	-0.04	0.936
mother age	38.45	39.75	-1.3***	0.008
child age	10.60	11.09	-0.49***	0.000
child educational gap	0.147	0.072	0.075***	0.000
father education	6.86	10.02	-3.16***	0.000
mother education	6.98	10.37	-3.39***	0.000
people living at home	5.85	5.13	0.72***	0.000
n. of children age $<$ 15	2.77	2.09	0.68***	0.000
n. of people age $>$ 59	0.08	0.09	0.01	0.779

Note: This table includes the results of t-tests on the equality of means allowing the variances to be unequal. “Cohabit” column contains the daughters who live with cohabiting parents; “Married” column contains the daughters who live with married parents.

Table 3

Descriptive statistics for sons living with both biological parents in the interval of age [8,14] * means are statistically different at 1%; ** at 5%; * at 10%.**

	Cohabit	Married		
	(235 obs)	(1777 o.)	Di¤erence	p-value
father age	41.74	43.23	-1.48**	0.023
mother age	37.04	39.95	-2.91***	0.000
child age	10.49	11.19	-0.70***	0.000
child educational gap	0.155	0.094	0.06***	0.000
father education	7.01	9.92	-2.91***	0.000
mother education	7.44	10.22	-2.78***	0.000
people living at home	6.01	5.08	0.93***	0.000
n. of children age $<$ 15	2.94	2.05	0.89***	0.000
n. of people age $>$ 59	0.12	0.09	0.03	0.265

Note: This table includes the results of t-tests on the equality of means allowing the variances to be unequal. “Cohabit” column contains the sons who live with cohabiting parents; “Married” column contains the sons who live with married parents.

The tables 2 and 3 show the means of individual and household characteristics by parental marital

status and by child gender. The cause of presenting different tables for boys and girls is that in developing countries (Saha, 2005), older children, usually girls, are oftenresponsible for home production and care of younger siblings. And these tasks could mean less time to devote to school work and, then, worse academic performance.

Descriptive factors to note are the statistically significant differences between twobiological cohabiting parents and married parents. Cohabiting parents are younger and have less completed years of schooling. Their children are younger but have greater schooling gap.

Another feature to mark is that cohabiting households have bigger family sizes and a larger number of younger siblings. In spite of these differences between the children who live with married parents and those who live with cohabiting parents, we have to bear in mind that in order to asses properly the determinants of the different educational gap, we ought to execute econometric analysis (as we do in the next section).

Table 4
Educational gap - children with age among [8,14]

	Bio Paren. Cohab		Bio Paren. Marr	
	Girls (258)	Boys (235)	Girls (1797)	Boys (1777)
Median	0.111	0.125	0	0
Mean	0.147	0.155	0.072	0.094
Std. Dev.	0.166	0.172	0.127	0.145
Variance	0.028	0.029	0.016	0.021
Skewness	0.799	0.810	2.121	1.855
Kurtosis	2.718	2.816	8.543	6.925

The plots below show that the educational gap is skewed right for all the children with age among [8,14] and it is more marked for the children who live with married parents.

Also, in reference to the Kurtosis analysis, the figures below show that the peakedness is more pronounced for the children who live with married parents because the proportion of children with a educational gap near 0 is greater among the children living with married parents.

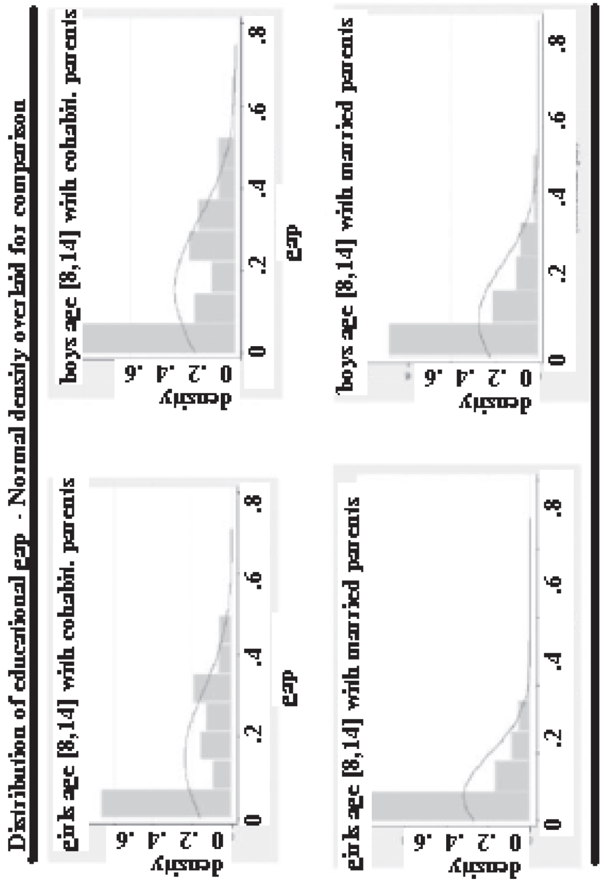

Figure 1

4. Methods of Estimation

In order to look for more robustness in the estimation of the determinants of the educational gap, we employ -besides OLS- three parametric approaches and two semiparametric methods. We discuss some advantages and drawbacks of each method below.

Binary Probit Model

One possibility is to define the educational gap, y_i, as a binary response variable, taking on the values zero when the actual grade attainment does not lag behind the ageappropriate schooling level, and one otherwise. But allowing only binary response, we lose information about the relative lags and their possible determinants.

Multinomial Ordered Models

Also we employ Multinomial Ordered Models to test the determinants of the absolute schooling gap (take into account that considering only the absolute gap, that dependent variable could take only the integer values from 0 to 8 because children ages are in the interval [8, 14]). We use in this paper an Ordered Logit Model, which contains the limitation of the IIA (independence from irrelevant alternatives) assumption.

Tobit Model

The educational gap, y_i, is a doubly censored variable which takes on the value zero and one with positive probability. In other words, the dependent variable suffers from interval censoring: the values of the true dependent variable, $y_i{*}$, are observed only if they fall within the interval [0,1].
Algebraically,

$$y_i^* = x_i'\beta + u_i, \qquad u_i|x_i \backsim Normal(0,\, \sigma^2)$$

$$y_i = 0 \quad \text{if} \quad y_i^* \leq 0$$

$$y_i = y_i^* \quad \text{if} \quad 0 < y_i^* < 1$$

$$y_i = 1 \quad \text{if} \quad y_i^* \geq 1$$

where x_i is a K x 1 vector of observed regressors, β is a K x 1 vector of unknown regression coefficients to be estimated, u_i is an unobserved error.

Tobit assumptions

Heteroskedasticity and nonnormality result in the Tobit estimator $\hat{\beta}$ being inconsistent for β, and entirely changes the functional forms for $E(y|x,\ 0 < y_i{*} < 1)$ and $E(y|x)$. Wooldrige (2002) observes that $y_i{*}$ should have a homoskedastic normal distribution and the variable y should be (roughly) continuous when $y > 0$, Thus the Tobit model is not appropriate for ordered responses. In the empirical application of this paper, we do a Tobit analysis with robust standard errors to cope with the possible existence of heteroskedasticity.

Normality was also tested using various procedures. Kernel density estimators were used to approxi-

3. In the Tobit model, the dependent variable is educational gap and the regressors are the household characteristics (family structure, income, wealth, parents' education, quantity of children at home,...) and dummies controlling population density, the economic situation of the region of residence and the sector of the economy in which the parents are employed.

mate the density f (*residuals of robust* TOBIT) [3] and a Normal density was overlaid for comparison. We tested normality in two ways: (1) a test based on a combination of a test on skewess and a test on kurtosis [4]; (2) the Shapiro-Francia test [5]. The results suggest that we can reject that the error term is normally distributed. Thus the distribution of the residuals (the estimation analogous to the error term) could be subject to nonnormality. If so, the Tobit estimators will not provide a consistent estimate. Therefore the common practice in previous literature of employing Tobit estimators for estimating educational attainment should be checked through to avoid inappropriate conclusions.

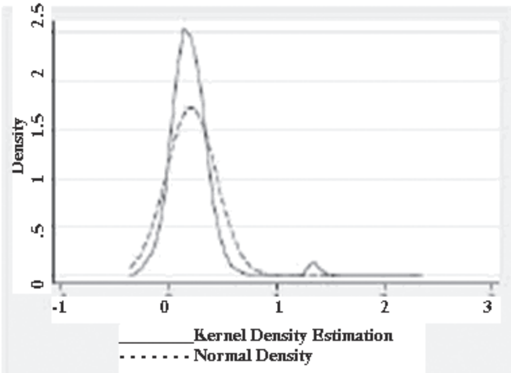

Figure 5
Kernel Density estimation of the residuals of the Tobit model for education-gap (variables and results in Table 5) - Normal density overlaid for comparison - Only

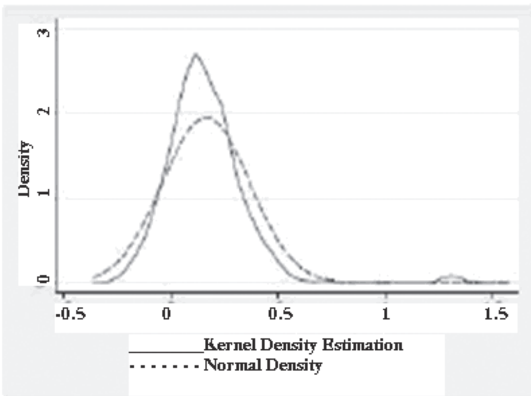

Figure 6
Kernel Density estimation of the residuals of the Tobit model for education-gap (variables and results in Table 6) - Normal density overlaid for comparison - Only sons with age [8,14]

Figure 2

4. Tested with "sktest" command of STATA
5. Tested with "sfrancia" command of STATA.

Thus, relaxing distributional assumptions on the error terms and seeking for models which succeed with those weaker distributional assumptions is mandatory to obtain accurate results.

Semiparametric Censored Regression Models

As we have seen in the previous sections, Tobit models require some specfications of the error distribution: normality and homoskedasticity. In order to relax these requirements, the semiparametric approach has been proposed in the recent economic literature to provide consistent estimates for censored data. Thus one of the advantages of the semiparametric approaches for censored models is that estimators are consistent under weaker distributional assumptions. The attribute "semiparametric" in this model comes from the fact that the distribution of the errors u_i given the explanatory variables does not have a known parametric form.

This paper uses two semiparametric estimators for censored regression models: the censored least absolute deviations (CLAD, Powell, 1984) and the symmetrically censored least squares (SCLS, Powell, 1986) (for a summary, see Chay and Powell, 2001, or Cameron and Trivedi, 2005).

Censored Least Absolute Deviations Estimator

The key distributional assumption of CLAD estimator is that ujx has median zero, and this means weaker distributional assumptions than the Tobit model which need normal errors. CLAD estimator is a generalization of least absolute deviations estimation for the standard linear model. Thus, the CLAD estimator minimizes the sum of absolute deviations of y_i over all β:

$$S_T(\beta) = \left(\tfrac{1}{T}\right) \sum_{t=1}^{T} |y_t - y_t^*|$$

where

$y_t^* = 1$ if $x_i'\beta \geq 1$

$y_t^* = x_i'\beta$ if $0 < x_i'\beta < 1$

$y_t^* = 0$ if $x_i'\beta \cdot\ 0$

Powell (1984) shows that CLAD $\hat{\beta}$ estimation is consistent, asymptotically normal and its asymptotic covariance matrix can be consistently estimated. Thus, tests of hypotheses concerning the unknown regression coefficient can be constructed, which are valid in large samples (precisely, in this paper we work with more than 4.000 observations: it could be seen as a "large sample"). Unlike estimation methods based on the assumption of Gaussian distributed errors terms, the CLAD estimator is consistent and asymptotically normal for a wide class of error distributions, and is also robust to heteroskedasticity.

Symmetrically Censored Least Squares Estimator

This estimator is based on the assumption that errors are symmetrically (and independently) distributed around zero, so is less restrictive than Tobit requirements (normally distributed and homoskedastic errors). The SCLS estimators are consistent and asymptotically normal for a wide class of symmetric error distributions with heteroskedasticity of unknown form. But the assumption of SCLS that errors are symmetrically and independently distributed around zero is stronger than the zero median restriction of the CLAD estimator.

Powell (1986) states that if the underlying error terms were symmetrically distributed about zero, and if the latent dependent variables were observable, classical least squares estimation would yield consistent estimates of the parameter vector β. But due to the censoring, the observed dependent variable y has an asymmetric distribution. Powell's approach consists in symmetrically censoring the dependent variable y (it is usually known as a "symmetric trimmed" method) so that symmetry can be restored, and then the regression coefcients can be estimated by least squares. Symmetric censoring of the dependent variable implies that observations with values above the censoring point are dropped, and this means that there could be a loss of efficiency due to the information dropped in those observations. However this problem is reduced in the present paper because a relative large sample is used.

Treatment Evaluation and Parents' Marital Status

The typical dilemma in treatment evaluation involves the inference of a causal association between the treatment and the outcome. In this paper, we pay particular attention to the effects of parent's marital status on the educational attainment of their children.

Thus, we observe (y_i, x_i, D_i), $i = 1; \ldots, N$, where y_i is the educational gap, x_i represents the regressors, and D_i is the treatment variable and takes the value 1 if the treatment is applied (cohabiting parents) and is 0 otherwise (married parents). The impact of a hypothetical change in D on y, holding x constant, is of interest. But no individual is simultaneously observed in both states: with the data available, it is not possible to view the same child both with married parents and with cohabiting ones. Moreover, the sample does not come from a randomized social experiment: it comes from observational data and the assignment of individuals to the treatment and control groups is not random. Hence, we estimate the treatment effects based on propensity score: this approach is a way to reduce the bias performing comparisons of outcomes using treated and control individuals who are as similar as possible (Becker and Ichino 2002). The propensity score is defined as the conditional probability of receiving a treatment given pre-treatment characteristics:

$$p(X) \equiv \Pr\{ D - 1 \ X \} - E\{D|X\}$$

where $D = \{0, 1\}$ is the indicator of exposure to treatment and X is the vector of pretreatment characteristics.

The propensity score was estimated in this application using a logit model [6]. Due to the probability of observing two units with exactly the same value of the propensity score is in principle zero since $p(X)$ is a continuous variable, various methods have been developed (for a summary, see Cameron et al. 2005) to match comparison units sufficiently close to the treated units. So, after estimating $p(X)$ we employed Kernel Matching method [7].

5. Empirical Results

Results

Tables 5 and 6 present the results of these estimations for girls and boys respectively. In most cases, the signs of the significant regressors come to be the expected ones (see Section 3). The number of children at home has operated in the hypothesized direction: this variable seems to worse children's school outcomes. On the other hand, according to the previous tables, family's wealth, parents' education (especially mother's education) and, in the case of daughters, the fact of having married parents have positive and significant effects on offspring school engagement. Maternal education seems to have a greater positive effect than father's education on the children educational attainment. This fact is consistent with the suggestions of the literature.

A possible explanation (see Saha, 2005) is that mothers tend to spend more time directly assisting

6. Applied with the Stata ado file "pscore" developed by Becker and Ichino (2002).

7. This matching method was applied using the Stata ado files psmatch2 developed by E. Leuven and B. Sianesi (2003) "PSMATCH2: Stata module to perform full Mahalanobis and propensity score matching, common support graphing, and covariate imbalance testing".

children with school work. The tables below show, considering CLAD results, each additional year of mother education reduces educational gap of sons by 0.021 while each additional year of father education reduces educational gap of sons only by 0.008. One exception in the signs theoretically predicted seems to be the positive sign of quantity_of_people_with_income_at_home, perhaps suggesting that more members of the family on the market could mean smaller child care time, and thus worse children's educational outcomes. Regarding the family structure issue, the results suggest that girls living with two-biological married parents experience better outcomes on educational attainment.

Considering CLAD results, the fact of having married parents reduces educational gap of daughters by 0.094. In the case of the sons, though the sign of the coefficient is the same, it is not significant in any estimation method used. Thus, negative cohabiting effects on educational attainment seem to be more pronounced against daughters. A possible explanation is that instability of the cohabitation unions (Brown, 2004) has a deeper influence on daughters with ages among [8,14] because of the different psychological characteristics of boys and girls at those ages. There's a rising literature in the psychological field which discusses gender-specific learning differences. For instance, Sax (2005) asserts that the brain of girls and boys develops differently; the brain is wired differently; girls hear better; and girls and boys respond to stress differently: Sax argues that stress enhances learning in males and the same stress impairs learning in females. This last fact could be related with the girls' worse school outcome than the boys', as a consequence of the instable environment of cohabitation.

Table 5

Estimates of educational gap (all heteroskedasticity-robust); only girls among [8,14].

Number of observations: 2055 - (estimated standard errors in parentheses) - *** significant at 1%; ** at 5%; * at 10%

Dependent Variable: educational gap

	OLS	PROBIT	TOBIT	OLOGIT	CLAD	SCLS
married parents	**-.069 (.031)****	-.289 (.318)	-.100 (.068)	**-1.03 (.551)***	**-.094 (.046)***	**-.123 (.099)***
quantity children age<15	.013 (.002)***	.073 (.026)***	.022 (.005)***	.154 (.043)***	.034 (.017)*	.037 (.015)*
quantity people age>59	-.010 (.008)	-.091 (.111)	-.025 (.025)	-.120 (.172)		.001 (.141)
quantity people with income	.010 (.004)**	.113 (.045)**	.023 (.010)**	.217 (.070)***	.039 (.020)*	.018 (.025)
log income per capita	-.035 (.065)	-.178 (.718)	-.024 (.162)	-.550 (1.27)	-.057 (.024)*	-.030 (.935)
(log income per capita)^2	.001 (.003)	-.006 (.045)		.021 (.079)		-.001 (.067)
father's education	-.006 (.004)	-.103 (.042)**	-.022 (.009)**	-.196 (.072)***	-.003 (.004)	.041 (.075)
(father's education)^2	.000 (.000)	-.001 (.002)		-.001 (.004)		-.003 (.006)
mother's education	-.018 (.005)***	-.120 (.046)***	-.032 (.010)***	-.237 (.082)***	-.021 (.011)*	-.074 (.074)*
(mother's education)^2	.000 (.000)	.001 (.002)		.004 (.004)		.002 (.008)
(father_educ)x(mother_educ)	.000 (.000)	.006 (.003)*	.001 (.001)	.007 (.006)		
(married_p)x(father_educ)	.004 (.003)	.071 (.033)**	.014 (.007)*	.129 (.059)**		
(married_p)x(mother_educ)	.002 (.004)	-.041 (.040)		-.021 (.069)		
parents' public job	.021 (.011)*	.088 (.128)	.047 (.032)	.140 (.122)		.168 (.132)
home_aid	.009 (.032)	.076 (.515)	.038 (.129)	-.059 (.801)		
mother_inactive	-.001 (.010)	.092 (.104)	.007 (.025)	.130 (.132)		-.071 (.104)
father_inactive	.034 (.021)	.282 (.241)	.077 (.055)	.041 (.313)		.118 (.132)
children private education	.010 (.008)	.186 (.110)*	.049 (.027)*	.193 (.174)		
parents' subjob	.008 (.007)	.001 (.076)	.010 (.018)	.006 (.118)		-.043 (.117)
children scholarship	.051 (.051)		.009 (.096)	1.12 (.611)*		-.003 (.163)
remittances	.226 (.163)		.409 (.205)**	1.58 (2.32)		.277 (.607)
absolute wealth	-.214 (.072)***	-1.04 (.686)	-.338 (.156)**	-2.55 (1.15)**		-.257 (.678)
relative wealth	.158 (.061)***	.803 (.643)	.240 (.147)	1.62 (1.06)		-1.01 (1.06)
constant	.427(.263)	2.35 (2.78)	.513 (.620)		.528 (.214)*	.656 (3.30)
R-squared	.219					
Pseudo R-squared		.148		.096		
Sigma			.262			

Table 6
Estimates of educational gap (all heteroskedasticity-robust); only boys among [8,14].
Number of observations: 2012 - (estimated standard errors in parentheses) - *** significant at 1%; ** at 5%; * at 10%
Dependent Variable: educational gap

	OLS	PROBIT	TOBIT	OLOGIT	CLAD	SCLS
married parents	-.045 (.033)	-.187 (.303)	-.072 (.062)	-.547 (.452)	-.002 (.033)	-.007 (.042)
quantity children age<15	.014 (.002)***	.081 (.026)***	.023 (.005)***	.103 (.038)***	.029 (.008)*	.021 (.012)*
quantity people age>59	.001 (.011)	.069 (.104)	.011 (.022)	-.099 (.163)	.015 (.030)	-.006 (.083)
quantity people with income	.013 (.004)***	.164 (.043)***	.029 (.008)***	.270 (.061)***	.028 (.008)*	.021 (.021)
log income per capita	-.036 (.059)	.159 (.633)	.062 (.136)	.015 (.922)	-.054 (.018)*	.170 (.805)
(log income per capita)^2	.001 (.003)	-.031 (.039)		-.026 (.058)		-.013 (.056)
father's education	-.010 (.004)**	-.074 (.041)*	-.019 (.008)**	-.136 (.066)**	-.008 (.001)	-.024 (.018)
(father's education)^2	.000 (.000)	.001 (.002)		.005 (.003)		.001 (.002)
mother's education	-.024 (.005)***	-.220 (.048)***	-.048 (.010)***	-.358 (.080)***	-.021 (.004)*	-.039 (.040)
(mother's education)^2	.001 (.000)***	.007 (.002)***	.001 (.000)***	.012 (.003)***		.001 (.004)
(father_educ)x(mother_educ)	.000 (.000)	-.001 (.003)	.000 (.000)	-.003 (.004)		
(married_p)x(father_educ)	.002 (.003)	.035 (.031)	.007 (.006)	.037 (.052)		
(married_p)x(mother_educ)	.003 (.004)	.001 (.038)	.003 (.008)	.037 (.068)		
parents' public job	.008 (.0129)	-.097 (.142)	-.001 (.034)	.157 (.119)		.037 (.076)
home_aid	-.001 (.039)	.023 (.413)	.017 (.112)	-.249 (.904)		
mother_inactive	.002 (.011)	-.035 (.105)	.002 (.022)	.065 (.116)		.007 (.036)
father_inactive	.031 (.028)	.309 (.241)	.064 (.054)	.277 (.315)		.039 (.121)
children private education	.007 (.009)	.150 (.114)	.039 (.027)	.155 (.182)		
parents' subjob	.005 (.007)	.025 (.078)	.007 (.017)	-.089 (.110)		.018 (.053)
children scholarship	-.124 (.023)***	-1.21 (.701)*	-.311 (.127)**	-1.28 (.895)		-.045 (.088)
remittances	-.028 (.030)		-1.04 (.101)***			
absolute wealth	-.194 (.079)**	-1.39 (.713)**	-.330 (.148)**	-2.23 (1.05)**		-.187 (.377)
relative wealth	.138 (.066)**	1.28 (.657)*	.242 (.137)*	1.44 (.972)	.036 (.131)	.065 (.439)
constant	.547 (.237)**	1.94 (2.45)	.376 (.523)		.554 (.135)*	-.168 (2.92)
R-squared	.221					
Pseudo R-squared		.155		.095		
Sigma			.255			

Robustness Check

Also we introduced and tested two suggestions of Berlinski et al. (2007). Firstly, these authors study the determinants of the levels of completed education among individuals aged 7-15 in Uruguay. Children can enroll in the first grade of primary education if they become 6 before the 10th. of May. Since the ECH Survey gives no information on birth date, they restrict the sample to the months of January to April. Secondly, Berlinski et al. study the effect of pre-primary education on children's subsequent school outcomes and they suggest a positive relationship. Thus, in this paper, we also introduced the binary regressor pre-primary education and restricted the sample to the months of January to April. But the new regressor has no significative impact on the educational gap and the results are similar to the tables 5 and 6 (see tables 10 and 11 in the Appendix)

Testing Endogeneity

One way in which endogeneity could arise is from the "omitted variables problem" and it might have appeared in the applied part of this paper because of the possible linkage between the variable "parents' marital status" and the unobserved "parents' irresponsibility".
With the intention of eliminating, or at least mitigating, the possible omitted variable bias, we introduced proxy variables.

Proxy binary variables for unobserved "parents' irresponsibility" (takes value one in case of parents' irresponsibility):

a) The survey asks the parents who have a job and didn't work last week for the reasons of this attitude. If they answer: "because of bad weather or not too much work to do", then "parents' irresponsibility" takes value one.

b) The survey asks the parents who have a job if they would like to work more hours. If they answer: "yes, but I did nothing to work more hours" or "yes, but I am not searching for other job", then "parents' irresponsibility" takes value one.

c) The survey asks the unemployed parents if they did anything to find a job last week. If they answer: "nothing", then "parents' irresponsibility" takes value one.

In this paper, these different proxy variables were aggregated in one dummy variable which takes value one if any of the dummies above is different from zero. We tested its significance using Tobit, CLAD and SCLS models, for boys and girls separately, with "educational gap" as the dependent variable. In no one of these models, the coefficients of this proxy variable of "parents' irresponsibility" were significantly different from zero (see Table 7). Thus, we did this exploratory exercise but we were not able to find a good proxy of parents' irresponsibility. The variables employed as proxy could be disputed but they were selected due to the restriction of variables available in the ECH survey.

Table 7
Searching for a proxy of parents' irresponsibility
Tobit(MLE) (heteroskedasticity-robust), CLAD and SCLS estimates of educational gap.
Dependent Variable: educational gap - (estimated standard errors in parentheses)
*** significant at 1%; ** at 5%; * at 10%

	girls among [8,14] - 2055 observations			boys among [8,14] - 2012 observations		
	TOBIT	CLAD	SCLS	TOBIT	CLAD	SCLS
parents' irresponsibility	.001 (.025)	-.024 (.026)	-.011 (.104)	.022 (.023)	.001 (.026)	.007 (.123)
married parents	-.100 (.068)	-.080 (.050)	-.121 (.108)	-.070 (.062)	-.002 (.034)	-.006 (.041)
quantity children age<15	.022 (.005)***	.036 (.019)*	.037 (.018)*	.023 (.005)***	.029 (.007)*	.021 (.015)
quantity people age>59	-.025 (.025)		.005 (.188)	.011 (.022)	.014 (.030)	-.008 (.064)
quantity people with income	.023 (.010)**	.035 (.018)*	.018 (.028)	.028 (.008)***	.028 (.009)*	.021 (.024)
log income per capita	-.023 (.162)	-.057 (.029)*	-.035 (1.08)	.062 (.136)	-.053 (.020)	.186 (1.23)
(log income per capita)^2			-.001 (.079)			-.014 (.086)
father's education	-.022 (.009)**	-.001 (.004)	.042 (.079)*	-.019 (.008)**	-.008 (.004)*	-.024 (.023)*
(father's education)^2			-.003 (.006)			.001 (.001)
mother's education	-.032 (.010)***	-.022 (.010)	-.072 (.053)	-.047 (.010)***	-.021 (.005)*	-.038 (.033)
(mother's education)^2			.002 (.006)	.001 (.000)***		.001 (.002)
(father_educ)x(mother_educ)	.001 (.001)			.000 (.000)		
(married_p)x(father_educ)	.014 (.007)*			.007 (.006)		
(married_p)x(mother_educ)				.002 (.008)		
parents' public job	.047 (.032)		.166 (.143)	-.002 (.034)		.038 (.103)
home_aid	.038 (.129)			.018 (.112)		
mother_inactive	.006 (.033)		-.063 (.121)	-.015 (.029)		
father_inactive	.076 (.055)		.119 (.146)	.058 (.054)		.038 (.116)
children private education	.049 (.027)*			.039 (.027)		
parents' subjob	.010 (.018)		-.040 (.140)	.005 (.018)		.019 (.066)
children scholarship	.010 (.097)		-.012 (.247)	-.307 (.127)**		-.044 (.134)
remittances	.410 (.205)**		.269 (.662)	-1.04 (.101)***		
absolute wealth	-.338 (.156)**		-.259 (.742)	-.339 (.149)**		-.178 (.482)
relative wealth	.240 (.147)		-1.00 (1.17)	.250 (.138)*	.037 (.121)	.046 (.522)
constant	.513 (.620)	.529 (.260)	.663 (3.76)	.376 (.523)	.551 (.151)*	-.231 (4.44)
Sigma	.262			.255		

Treatment Evaluation

Table 8
Average Effect (on Educational Gap) of Treatment (Cohabiting Parents) on the Treated (ATT)

	Girls	Boys
ATT	**.0214**	**.0175**
n. treat	258	230
n. contr.	1,797	1,777
Treated	.1467	.1512
Controls	.1253	.1336
S.E.	.0116	.0123
T-stat	1.84	1.42

Note: estimation with the Kernel Matching method

The point estimates indicate that having cohabiting parents increases the educational gap both for girls and boys, and the effect of the "treatment" (having cohabiting parents) is greater in the case of daughters. The ATT is significantly different from zero at the 5 percent level in the daughters' case but only at the 10 percent in the sons' case. Thus, using the propensity score and the Kernel matching method, there's some evidence to support the positive influence of having married parents on their children's educational attainment, especially on girls'. In order to evaluate the goodness of the matching, we should bear in mind that the matching method intends to make comparisons between treated and control individuals who are as similar as possible. This similarity between the treated and control individuals can be seen in the mean comparison test (t-test) shown on the Table 9: there's no statistically significant difference in the characteristics of the treated and control matched individuals. This fact denotes that the matching is fine.

Table 9
Descriptive Statistics for the Treated, not Treated and Matched groups

Household environment of girls with age [8,14]

Variable	Sample	Mean Treated	Mean Control	t-test t	t-test p>t
log income per capita	Unmatched	7.2913	8.0068	-13.27	0.000
	Matched	7.2913	7.326	-0.55	0.581
any parent unemployed	Unmatched	.23643	.14524	3.78	0.000
	Matched	.23643	.23406	0.06	0.949
mother's education	Unmatched	6.9806	10.378	-12.69	0.000
	Matched	6.9806	7.1729	-0.75	0.451
father's education	Unmatched	6.8605	10.023	-11.51	0.000
	Matched	6.8605	7.1362	-0.98	0.326
relative wealth	Unmatched	.16771	.37738	-13.60	0.000
	Matched	.16771	.17927	-0.83	0.409
household ownership	Unmatched	.45736	.68893	-7.44	0.000
	Matched	.45736	.46816	-0.25	0.806
more than one family at home	Unmatched	.04651	.0089	4.90	0.000
	Matched	.04651	.0471	-0.03	0.975
number of individuals per room	Unmatched	3.0303	2.1911	11.30	0.000
	Matched	3.0303	3.0049	0.17	0.863

Household environment of boys with age [8,14]

Variable	Sample	Mean Treated	Mean Control	t-test t	t-test p>t
log income per capita	Unmatched	7.3224	7.9726	-11.29	0.000
	Matched	7.3279	7.3736	-0.69	0.491
any parent unemployed	Unmatched	.22553	.13506	3.71	0.000
	Matched	.22174	.21871	0.08	0.938
mother's education	Unmatched	7.4447	10.224	-10.15	0.000
	Matched	7.5022	7.7107	-0.75	0.456
father's education	Unmatched	7.0128	9.9226	-10.29	0.000
	Matched	7.087	7.3678	-0.89	0.374
household ownership	Unmatched	.47234	.67136	-6.06	0.000
	Matched	.47826	.48647	-0.18	0.861
more than one family at home	Unmatched	.0383	.00788	4.14	0.000
	Matched	.01739	.02216	-0.37	0.714
sum of parents age	Unmatched	78.783	83.18	-4.74	0.000
	Matched	78.961	78.787	0.13	0.897

6. Conclusions

There's a growing body of research on the determinants of children's school performance and not few methodological problems appear in previous investigations about the determinants of educational gap. This paper has extended prior research considering -besides the possible existence of endogeneity- censored regression models -such as Tobit Model- and semiparametric alternative approaches -such as the Censored Least Absolute Deviations Estimator and the Symmetrically Censored Least Squares Estimator. Drawbacks and advantages of the different estimation methods have been discussed. In the empirical application, this study introduces indicators of family environment and focuses on the impact of the parents' marital status on their children's educational attainment. In particular, the results suggest positive influence of having married parents on daughter's educational outcomes, after controlling for household background variables such as parents' education, income per capita, wealth and number of children. This finding is consistent with previous investigations and with the theoretical hypotheses that both the impermanence of cohabiting unions and their incomplete institutionalization (unclear family roles, rights, and obligations) set the stage for a family environment that may undermine child development. Finally, this paper includes an application of the propensity score approach for treatment evaluation of parents' marital status: all the point estimates indicate that having married parents has a positive effect on children's schooling outcomes, especially on girls'. This present study contributes to the economic literature in this field by applying more suitable estimation methods and by checking through the possible faults or omissions of methods used in previous investigations.

For further research, four considerations about the empirical application: First, a significant shortcoming of the survey used in this paper, is that it does not have longitudinal data or cohort information[8]: there's no information available about the marriage history information of the biological parents [9]. Thus, one drawback of Continuous Households Sur vey (ECH) is that it does not provide measurement of the duration of the different family structures or the number of different family transitions that children have experienced (so long term or cumulative effects of family structure can't be observed). Second, besides taking into account data from all the available years of the Continuous Households Survey, in order to contribute to unravel the complexities of family issues, it could be useful to wide the range of family structures and also test the different incidence of, for instance, the two-biological-parent families, stepfamilies and female-headed households over the children education attainments. Moreover, it could be interesting to evaluate also for the other years of the ECH survey if cohabiting effects on educational attainment could be biased against daughters -a kind of unwelcome discrimination- as it is suggested in this paper. Third, this investigation could be completed testing also not only children school engagement but also other behavioral and emotional effects. Fourth, one major problem with the data used for the empirical application is that there is no measure of the children ability which should be positive correlated with school performance.

On the theoretical field of estimation methods for censored regression models, other semiparametric alternatives for censored models could be evaluated.

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8. Brown (2004) quotes previous research which using longitudinal data also suggests a positive relationship between two-biological married parents and child well-being.
9. Longitudinal data with individual life trajectories would allow us to observe, for instance, how parents' attitudes about cohabitation influence the child's subsequent marital and cohabitation experience (Axinn

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Table 10
Estimates of educational gap (all heteroskedasticity-robust).
Sample restricted to the months of January to April; only girls among [8,14]
Number of observations: 687 - (estimated standard errors in parentheses) - *** significant at 1%; ** at 5%; * at 10%
Dependent Variable: educational gap

	OLS	PROBIT	TOBIT	OLOGIT
married parents	**-.153 (.068)****	-.788 (.733)	**-.445 (.206)****	-1.131 (1.139)
preschool attendance	-.021(.020)	.005 (.197)	-.035 (.060)	.025 (.284)
quantity children age<15	.025 (.006)***	.161 (.062)***	.062 (.018)***	.208 (.080)**
quantity people age>59		-.056 (.251)	.016 (.076)	-.109 (.374)
quantity people with income	.015 (.012)	.154 (.099)	.050 (.033)	.242 (.146)*
log income per capita	-.020 (.090)	.018 (.888)	-.008 (.286)	-.590 (1.177)
(log income per capita)^2	.001 (.005)	-.004 (.060)		.056 (.080)
father's education	.013 (.009)	.030 (.095)	.017 (.029)	.134 (.145)
(father's education)^2		-.003 (.006)		-.010 (.008)
mother's education	-.038 (.011)	-.281 (.091)***	-.088 (.028)***	-.480 (.169)***
(mother's education)^2		.002 (.005)	.001 (.001)	.008 (.006)
(father_educ)x(mother_educ)			-.002 (.002)	-.001 (.011)
(married_p)x(father_educ)	-.005 (.007)	.020 (.077)	.007 (.023)	-.008 (.106)
(married_p)x(mother_educ)	.020 (.010)**	.105 (.079)	.052 (.024)**	.152 (.138)
parents' public job	.034 (.024)	.332 (.348)	.135 (.114)	.288 (.286)
home_aid	.048 (.061)	1.712 (.782)**	.536 (.267)**	1.211 (1.207)
mother_inactive	-.019 (.025)	.520 (.235)**	.077 (.080)	.065 (.264)
father_inactive	-.027 (.038)	-.032 (.464)	-.074 (.145)	-.151 (.514)
children private education		.236 (.282)	.078 (.099)	.218 (.438)
parents' subjob	.006 (.015)	-.027 (.179)		.286 (.252)
remittances	.256 (.201)		.610 (.241)**	1.843 (1.879)
absolute wealth	-.358 (.156)**	-4.198 (1.470)***	-1.206 (.471)**	-5.182 (2.060)**
relative wealth	.265 (.129)**	2.839 (1.439)**	.744 (.458)	3.010 (2.130)
constant	.411 (.371)	1.300 (3.410)	.451 (1.096)	
R-squared	.340			
Pseudo R-squared		.308		.156
Sigma			.355	

Table 11
Estimates of educational gap (all heteroskedasticity-robust)
Sample restricted to the months of January to April; only boys among [8,14]
Number of observations: 642 - (estimated standard errors in parentheses) - *** significant at 1%; ** at 5%; * at 10%
Dependent Variable: educational gap

	OLS	PROBIT	TOBIT	OLOGIT
married parents	-.086 (.055)	-.815 (.606)	**-.308 (.136)****	**-2.223 (.975)****
preschool attendance	-.007 (.019)	.176 (.179)	.011 (.046)	.060 (.227)
quantity children age<15	.007 (.005)	.080 (.060)	.023 (.014)*	.002 (.072)
quantity people age>59	-.029 (.015)*	-.020 (.193)	-.017 (.043)	-.030 (.279)
quantity people with income	.027 (.008)***	.327 (.084)***	.064 (.016)***	.384 (.104)***
log income per capita	-.145 (.117)	.401 (1.554)	.027 (.049)	-.069 (1.950)
(log income per capita)^2	.006 (.006)	-.054 (.098)		-.032 (.123)
father's education	.004 (.009)	.023 (.091)	.015 (.019)	-.157 (.128)
(father's education)^2		.001 (.005)		.011 (.006)*
mother's education	-.044 (.011)***	-.462 (.103)***	-.113 (.021)***	-.643 (.182)***
(mother's education)^2	.001 (.000)***	.015 (.004)***		.018 (.004)***
(father_educ)x(mother_educ)		-.010 (.006)	-.002 (.001)*	-.017 (.007)**
(married_p)x(father_educ)	-.008 (.006)		-.005 (.016)	.034 (.119)
(married_p)x(mother_educ)	.021 (.008)**	.143 (.086)*	.057 (.018)***	.343 (.170)**
parents' public job	.009 (.020)	-.137 (.375)	.026 (.093)	-.152 (.246)
home_aid	-.033 (.032)		-1.265 (.194)***	-32.61 (.574)***
mother_inactive	-.003 (.025)	-.115 (.219)	-.025 (.049)	.088 (.228)
father_inactive	-.005 (.061)	.025 (.480)	.031 (.135)	-.250 (.663)
children private education	.012 (.019)	.353 (.267)	.124 (.067)*	.211 (.354)
parents' subjob	.013 (.016)	-.075 (.177)	-.009 (.041)	-.094 (.216)
absolute wealth	-.130 (.161)	-1.206 (1.762)	-.209 (.393)	-1.529 (2.068)
relative wealth	.129 (.137)	1.233 (1.568)	.191 (.364)	1.412 (1.888)
constant	1.044 (.472)**	1.923 (5.992)		
R-squared	.362			
Pseudo R-squared		.309		.129
Sigma			.279	

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“Los profesores del Máster son de primer nivel y la apuesta a la excelencia es total. La exigencia y rigurosidad técnica que experimenté en la UM me ayudó a dar el salto para hacer el Doctorado en Economía que estoy preparando en el exterior”.

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