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EL ROL DE LOS ESTUDIOS DE POBLACIÓN TRAS LA PANDEMIA DE COVID-19 Y EL DESAFÍO DE LA IGUALDAD EN AMÉRICA LATINA Y EL CARIBE

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Gender Differences in High-School Dropout: Vulnerability and Adolescent Fertility in Chile

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Introduction

Adolescence is a period of rapid changes and preparation for adulthood. The initial concerns about the consequences of adolescent fertility assumed that pregnancy was a turning point which altered the trajectory teens were following, placing them on a path of vulnerability, i.e., facing a lack of resources from which they might not recover in adulthood. It was thought that pregnancy during adolescence would hinder proper school progress, the accumulation of human capital and participation in the labor force. However, several years of research have shown that teenagers who become pregnant are not a random sample of the population, but a selective sample that is more likely to have limited socioeconomic resources and to belong to disadvantaged groups, either because of their rural residence, their ethnic minority status, or due to having grown up in non-intact families, that is to say, they were a vulnerable group to begin with. This paper studies the association between adolescent fertility and the high-school dropout rate in Chile, taking into account this selectivity and inquiring into gender differences in these outcomes. Many efforts have been expended to handle the selectivity of vulnerable teenagers into pregnancy or motherhood, mainly in U.S.-based research. These efforts, which are scarce in Latin America in general, and in Chile in particular, have considerably improved our knowledge of the educational consequences of adolescent fertility. However, there is little literature about the educational consequences of early childbearing among men, as well as about gender differences in the educational outcomes of early mother or fatherhood. The main distinguishing feature of this study is to contribute by generating estimates regarding the educational setbacks of adolescent fertility in a Latin American country, as well as to invite men into the discussion about the effect of adolescent fertility on education.

In the following pages, we will review previous studies about educational outcomes and adolescent fertility, paying special attention to selectivity issues and the approaches that have been followed to deal with this aspect of the research. We synthetize the main findings about gender differences in educational attainment in the presence of early childbearing. We also introduce the Chilean case. Next, we describe the data we used and our analytical strategy. We present the results of our analyses and finally we highlight our main findings and discuss their implications.

Previous Research

Adolescent Fertility and Education

From a life course perspective (Elder, 1998), adolescence is a critical stage, because what happens during this period may impact the trajectories people will follow in different life domains. Adolescents' behavior reflects their own trajectories of physiological and psychological development, but also the influence of the social institutions they belong to, among which family and school are key environments. Adolescents experience continuity and change in their interpersonal relations and in the roles they play in these key institutions, all of which are embedded in the larger structures of society, such as gender or the socioeconomic stratification systems, and they are conditioned by the peculiarities of the historic time in which they live (Crosnoe & Johnson, 2011). In this vein, Buhr et al. (2014), considering the multilevel structure of the life course, distinguish between two types of determinants of fertility decisions: external conditions and internal conditions. The former refers to "structural, economic and institutional circumstances at the societal or regional level, as well as at the level of social environments and social relationships". The second group, internal conditions, "comprise physiological aptitudes, personality traits, dispositions, values, and aspirations" (Buhr & Huinink, 2014, p. 2). Other authors have made similar distinctions, referring to macro, meso and micro determinants of fertility in general (Balbo, Billari, & Mills, 2013) or of adolescent fertility (Pantelides, 2004). Pantelides' framework includes among the macrosocial factors the systems of stratification of a society (on the basis of socioeconomic status, gender, or race/ethnicity) and the social policies available to adolescents. At a meso level, key determinants are their urban/rural status, family structure, and the influence of peers. At a micro-level, adolescent fertility depends on variables such as age, race/ethnicity and educational attainment, but also on more subjective features, such as individual perceptions, attitudes and knowledge about gender roles, sexual behaviors and contraception (Pantelides, 2004). Macrosocial determinants of teen fertility have hardly been studied, and anyway, there is not much to say about them when studying a single country (as is the case in this study). By way of contrast there is plenty of research about meso and micro level determinants of adolescent fertility. In that line of investigation, we know that teens living in rural areas are more likely to get pregnant than teens in urban areas (Rodríguez Vignoli, 2014; Santos, 2009), the same as teens who have grown up in non-intact families (Kane, Morgan, Harris, & Guilkey, 2013; Pantelides, 2004, 2004); and we know that as teens age, the likelihood of transitioning to parenthood increases. Adolescent fertility is also more probable among ethnic minorities and teens with a lower educational attainment (Almeida, Aquino, & Barros, 2006; Buhr & Huinink, 2014; Shahidul, 2015). Among the most proximate determinants of fertility, an earlier age at the first sexual encounter increases the time exposed to the risk of getting pregnant, and is therefore positively related to adolescent fertility; and the use of contraceptives is negatively related to adolescent fertility (Di Cesare & Rodriguez Vignoli, 2006; Santelli & Melinikas, 2010).

Besides the determinants of adolescent fertility, there is a considerable body of research delving into the consequences of early childbearing in relation to the accumulation of human capital, which links to the multidimensionality of the life course. Because the different life domains are mutually interdependent, fertility experiences have an influence on other domains. In particular, spillover effects (Schiemann, Glavin, & Milkie, 2009), which are a type of interdependence in regard to outcomes (Buhr & Huinink, 2014), are likely to be observed when an early childbirth and the subsequent childcare responsibilities hinder the educational trajectories of adolescents. Starting from the basic idea that postponing childbirth may contribute to obtaining a higher educational attainment, which is not only more human capital for women, but also a resource for their children (Schultz, 2007), the initial research on adolescent fertility and education assumed that the former was the cause of worse educational outcomes. Broadly speaking, there are two potential mechanisms by which adolescent fertility could decrease educational attainment: the opportunity costs and the effects of the stress associated with becoming a mother at an early age. In the first case, teens who become mothers would greatly decrease their ability to accumulate human capital because of the child care tasks they must assume. Teen mothers have less available time and less energy, hence their deficient educational outcomes and reduced employability (Becker, 1981). Regarding stress, the rapid transition to the role of mother would have psychological consequences that would hinder the subsequent accumulation of human capital (Coleman, 2006; Hagestad, 1990).

However, these causal analyses of the relation between adolescent fertility and education have received many critiques, related to endogeneity. In a given model, an endogenous variable is a variable that is correlated with the error term. Predictors included in a model (a vector of variables, X) should explain variations in the outcome (Y), but all of the factors which are not explicitly included in the model are included the error term. For instance, if variable (Z) explains changes in Y, but we are not able to observe it, that variable Z is included in the error term. If Z can explain any of the X variables in the model, the error term is correlated with X, so that X is endogenous. Endogeneity appears when Z is correlated to Y, but also to X, and not included in the model (Antonakis, Bendahan, Jacquart, & Lalive, 2010). In the case of adolescent fertility and educational outcomes, typically there are common unobserved variables that predict both teen childbearing and educational outcomes. One of the sources of endogeneity is selection bias. Adopting terminology from a causal approach, given that selection into the groups of “treatment” (teen childbearing) and “control” (not teen childbearing) is not random – because adolescent fertility is more frequent among the most socially disadvantaged, who are more likely to have poor educational outcomes even in the absence of adolescent fertility -- the estimates of treatment effect (teen childbearing) on educational outcomes that do not deal with this problem are typically biased (Kane et al., 2013). In other words, it may be that socioeconomic vulnerability and other vulnerabilities, instead of fertility itself, may be the cause of a low educational attainment (Geronimus & Korenman, 1993; Hotz, McElroy, & Sanders, 2005).

Because the initial research on the association between adolescent fertility and educational outcomes did not deal with this endogeneity due to treatment selection, it overestimated the effects of fertility on educational outcomes (Diaz & Fiel, 2016). Several strategies have been used to properly estimate the effect of adolescent fertility on educational outcomes. These strategies address endogeneity differently and impose different assumptions, dealing either with observed or unobserved variables, and they lead to different estimates of the effect of adolescent fertility and educational outcomes (Kane et al., 2013). In general, though, these studies estimate that the causal effect of adolescent fertility on educational outcomes is lower than was originally thought. For instance, in the U.S. the reduction in the number of years of schooling due to adolescent fertility would be between zero and two years (Fletcher & Wolfe, 2009; Kane et al., 2013; Lee, 2010).

Among the techniques used to properly estimate the effect of adolescent fertility on educational outcomes are fixed-effect models, instrumental variables, and propensity score matching. Fixed-effect models have typically used samples of siblings to include the effect of unobserved variables previous to the pregnancy, which determine both adolescent fertility and educational attainment, and that are shared between sisters. This strategy requires samples in which one girl got pregnant during adolescence whereas one of her sisters did not. Instrumental variable models involve a two-stage least square estimation. In the first stage, several exogenous predictors and at least one instrumental variable are used to model the endogenous variables (teen pregnancy). Miscarriages and spontaneous abortion have been used as instruments. In the second stage, predicted values from the first model are used, instead of pregnancy itself, to model educational attainment, controlling for other exogenous variables. Finally, propensity score matching is a technique with fewer parametric assumptions, whose data requirements are less complex than the others. To deal with the selection of certain types of adolescents into pregnancy, a two-step analysis is implemented. First, the probability that an adolescent will get pregnant is predicted by means of a regression for binary outcomes, including a set of preexisting predictors of pregnancy (the propensity score). Second, matched “pairs” of treated and control cases are formed using their predicted propensity score. The effect of the treatment on educational outcomes is estimated by averaging differences within pairs (Diaz & Fiel, 2016), that is to say, comparing the attainment of similar adolescents who did and did not got pregnant. Studies using this approach have estimated that in the U.S., teen mothers are 40 percent less likely to finish college than are childless teens (Lee, 2010). In Mexico, Arceo-Gomez and Campos-Vasquez (2014) found that teen pregnancy decreased educational attainment by 1-1.2 years of schooling. Notice that even though propensity score matching does not solve the problem of selectivity due to unobserved variables – something that the instrumental variables approach does, which is why this strategy is more appropriate when embracing a causal analysis -- it reduces the bias in the estimation of the treatment effect due to the observed variables that are included in the estimation (Stuart, 2010).

Gender differences in adolescent fertility and educational outcomes

Past research about variations in life course outcomes among teen mothers and fathers is scarce. Most of the research on the particular topic of adolescent fertility and educational outcomes has examined only women[1]. As an exception, Marsiglio (1987), when analyzing fathers in the U.S., found that Black teenage fathers were more likely to get a high-school degree than were non-poor Whites. There are a few studies comparing men and women. In Cameroon, where high-school completion is more likely among men than women, a life-table based analysis concluded that pregnancy in secondary education is responsible to a great degree, but not for all of the gender gap in the school dropout rate (Eloundou-Enyegue, 2004). In the U.S., one study found worse educational results among teen fathers than teen mothers (Pirog & Magee, 1997), and a more recent study finds no differences in the probability of high-school completion between men and women who became parents during adolescence, even though it also finds that gender is a significant moderator of the association between parenting responsibilities and educational outcomes: primary caregiving responsibilities are associated with lower chances of graduating from high school for mothers, but not for fathers, and working is related to lower chances of finishing high school for fathers, but not for mothers (Mollborn, 2010). These findings suggest that teen parents who conform to the gender-stereotypical parenting responsibilities (that is so say, mothers as primary caregivers and fathers as breadwinners) are at highest risk for dropping out of high school. In Brazil, one study finds that teen pregnancy is a frequent predecessor of becoming a school dropout both for men and women, but the reasons for justifying dropping out of school are gender-specific. The main reason for dropping out that mothers report is pregnancy, whereas the main reason that fathers report is work (Almeida et al., 2006). Just as is the case in Mollborn's investigation, this study suggests that even though it is not extensively studied, early parenthood

[1] Research on the relationship between adolescent fertility and educational outcomes is not common in European-based research, given that adolescent fertility tends to be low in countries that are advanced in their fertility transition, but adolescent fertility is still relatively high in the U.S., and most of the research on this topic is produced in this country.

probably impacts the educational trajectories of men, perhaps not pushing them into the household for caregiving tasks, but pushing them into the labor market at an early age.

The parenting responsibilities (childcare and breadwinning) are likely to have a traditional gender distribution, but both men and women who experience teen parenthood are likely to see an effect of these responsibilities on their educational trajectories. Given that we know so little about gender differences in the educational outcomes of men and women who become parents during their adolescence, it is not clear whether the burden of an early birth is greater for women or men. On the one side, “the biological burden of mothering and the typically greater intensity of parenting responsibilities for mothers than for fathers would suggest that the negative educational consequences of teenage parenthood may be more severe for women than for men” (Mollborn, 2010, p. 153). On the other hand, one of the few studies in this area actually finds that the burden was worse for men (Pirog & Magee, 1997). Still, neither of these studies deals with the issue of endogeneity in the relation between teenage childbearing and educational outcomes. In addition, these findings can be context-specific, given that social norms about early childbearing and the social support coming either from the young parent’s family of origin or the State, vary in different societies.

Adolescent Fertility and Educational Outcomes in Chile

Regarding the specificity of the Chilean context, adolescent fertility is relatively high. In 2016, there were 32.3 births per 1,000 women in the 15-19 age interval (DEIS, n.d.), as compared to 20 in the U.S, 13 in the U.K. or 3 in Switzerland in 2017 (United Nations Population Division, n.d.). Since the fertility decline began in the 1960s, the total fertility rate has been continually decreasing, but adolescent fertility has not. It has certainly declined from its 1960s values, but the decline has been slower than the decline of the next age groups. As seen in Figure 1, during the nineties adolescent fertility actually increased. It declined again between 1999 and 2004; it increased again between 2005 and 2009, and since then it has declined continuously, which means that this last trend of decline has extended for about a decade (DEIS, n.d.; Rodríguez Vignoli, 2014; Rodríguez Vignoli, Di Cesare, & Páez, 2017). Chile shares with other countries in the region the pattern of a relatively high adolescent fertility and a relatively slow fertility decline among teens. Indeed, Latin America has the second highest adolescent fertility rate, after Sub Saharan Africa (Rodríguez Vignoli, 2014; Rodríguez Vignoli et al., 2017).

[FIGURE 1 HERE]

The association of adolescent fertility with socioeconomic status (SES) in Chile is well known. For instance, in 2006 an adolescent from the lowest SES group had 17 times more chances of becoming a mother than an adolescent from the highest SES group (Rodriguez Vignoli & Di Cesare, 2010). One aspect that has been highlighted in Chile is that fertility under age 15 has also followed a non-monotonic trend of decline. Instead, there are periods in which it has grown-- and the socioeconomic bias of fertility is even more evident under age 15 (Molina, Molina, & González A, 2007; Rodríguez Vignoli, 2011).

Regarding educational outcomes, secondary education has been compulsory since 2006. According to the 2017 census, 75 percent of the population of 25 years of age or older has a high-school diploma or further studies (Instituto Nacional de Estadísticas, 2018). School dropout rates are relatively low as compared to other Latin American countries (CEPAL, 2002). According to estimates by the Ministry of Education (2013), dropping out rates are between 2.4 and 3.8 percent. Dropping out is more likely during secondary school, with the highest rates in the 9^{th} and 11^{th} grades (Centro de Estudios de la Niñez, 2014). It is also more common among poor students and/or students living in rural areas (Santos, 2009), and it is more frequent among men than women (Bonomelli, Castillo & Croquevielle, 2020). Even though Chile has advanced substantially in expanding educational access, it has done less about shortening the gap in the quality of education offered to students of different SES groups, ending up with] a school system that has been qualified as hyper-segregated (Bellei, 2013). Hence, the association between adolescent fertility and education is quite relevant in Chile, considering that adolescents with higher chances of experiencing a pregnancy are likely to be receiving an education of inferior quality.

Some previous research has looked at the association between adolescent fertility and educational outcomes in Chile. Most of this research uses cross-sectional data, which is an obstacle to assessing the effect of pregnancy. Some of them use clinical, non-random samples, and, therefore, it is not possible to generalize their results to the total population. From this literature, we highlight Rodríguez's study (2005). He used census data, and found that teen mothers are less likely to be full-time students than nulliparous teens. Tough, there has been some improvement over time --in 1982, teen mothers had zero chances of attending school, while in 2002 their chances were 20 percent. In a clinical study, using a retrospective analysis of teen mothers in a maternity ward, Molina et al. (2004) found that a sizeable proportion of the sample had dropped out of school before getting pregnant, because of financial difficulties or problems at home.

There are two published studies addressing issues of endogeneity in the relation between adolescent fertility and educational outcomes, one of them using a fixed-effect approach and the other using propensity score matching. In the former, Berthelon, Kruger and Eberhard (2017) estimate the causal effect of teen motherhood on several educational outcomes. They pool data from ten rounds of the most important (cross sectional) household survey in Chile (Encuesta Nacional de Caracterización Económica, CASEN), covering the 1990-2011 period, to perform fixed-effect models in a sample of siblings. The authors find that becoming a teen mother reduces educational outcomes. More specifically, among women 20-24, teen motherhood reduces the probability of completing high school by 16 percent. Although the data the study uses is not very recent, this investigation makes a valuable contribution in terms of measuring the precise effect of teen motherhood on educational outcomes. However, the generalizability of these findings is limited. Even though 10 rounds of the CASEN survey are used to deal with the problem of reduced samples that sibling studies typically face, the analysis is limited to young siblings (women 20-24 years old) living with their parents, in families in which one sister has a child during her adolescence, whereas the other did not. It is likely that by doing so they are excluding young women who do not live with their parents (perhaps because they became teen mothers), as well as women from families in which both (or all) sisters had a child during adolescence. The other study (Berthelon & Kruger, 2017), by using propensity score matching, overcomes the problem of generalizability. This work was carried out by two of the same authors, and it uses the same data source, although they add one more round of CASEN (2013). The analysis is similar, in terms of the included covariates and the observed outcomes, even though the analytical technique the authors use is different. They find that adolescent mothers' odds of completing high school were 23 percent lower than the chances of those who did not bear a child during their teen years. Note that by using propensity score matching techniques the authors are not able to account for unobservables that affect both adolescent motherhood and educational outcomes, and therefore, this effect should not be interpreted as causal.

There are a few Chilean studies exploring the life course trajectories of teen fathers, as compared with teen mothers. They suggest that adolescent fertility has a greater impact on the school trajectories of girls than boys. Madrid (2006), using a cross-sectional National Survey of Youth shows that 50 percent of young men who become fathers during their teens recognize that they stopped their educational career in order to work, whereas 60 percent of women who became mothers in their teens recognize they stopped studying in order to look after their babies. In the same vein, Molina et al. (2004), using a nationally representative household survey, conclude that motherhood is the main reason why women drop out of school, whereas the main reason for men is economic hardship. Qualitative evidence supports these findings, in as much as adolescent fathers report they associate being a father with the need to financially provide for their babies, more than they associate it with assuming the daily tasks of caregiving, whereas teen mothers relate motherhood to domestic labor and childcare, which hinders the continuation of their educational careers (Sadler & Aguayo, 2006). Consistently, women who become mothers in their teens in Chile tend to have low labor force participation rates (Molina et al., 2004; Rodríguez Vignoli, 2005). If they work, they are likely to hold lower quality jobs (Molina et al., 2004). Madrid's estimates indicate that young men and women (15 to 29 years old) who became parents in their teens are more likely to work than those who postponed childbearing. Qualitative research highlights that this traditional gender division of labor occurs at times when there is an increasing movement towards more involvement of men in childcare, an image that is strongly diffused by the mass media (Miller, 2011; Olavarría, 2001; Valdés, 2009), even though men would experience this demand for higher involvement ambivalently, because they feel that the need to be a close and loving father competes with the need to provide for and protect their families (Herrera & Pavicevic, 2016; Olavarría, 2017; Valdes & Godoy, 2008).

On basis of this previous research, this study asks about gender differences in the association of adolescent parenthood and educational outcomes in Chile. In particular, we analyze the high-school dropout rate of teen men and women, taking into consideration their socioeconomic status, sociodemographic characteristics, and characteristics of their sexual initiation.

Data & Methods

Data comes from the VIII Chilean Survey of Youth, carried out in 2015 (September to December). The sample design is probabilistic and multistage. This is a nationally representative survey[2], whose target population is totally composed of non-institutionalized men and women, ages 15-29. The sample size was 9,393 people. The main questionnaire was applied face-to-face, but there was a self-administered booklet with questions about sensitive topics, including sexual and reproductive health. The main questionnaire incorporated questions about school attainment, family background and sociodemographic characteristics, among other factors.

Our dependent variable is high-school dropout status, a dichotomous variable which identifies young people whose educational attainment is incomplete secondary school or less, and who are currently not enrolled in school. People who are enrolled or whose educational attainment indicates having completed secondary school or more are the reference category. Our treatment is teen parenthood, which is a dummy variable, identifying people who had a baby before age 20. The main questionnaire asked about the number of children the respondent has and the age of the eldest child. Using these questions, in addition to the age of the respondent, we were able to identify people who became parents during their teenage years. People who had previously

[2] The sample is also representative at the regional level (regions are the main geographic units into which Chile is divided) and at the urban/rural level.

reported they had never had sex were classified as not having a teen unplanned pregnancy. Therefore, people were classified as having experienced adolescent fertility either if they became parents in their adolescence or if they reported an unplanned teen pregnancy.

To measure the socioeconomic status (SES) of young people before fertility–related events or dropping out of school, we include a categorical variable that classifies the type of elementary school the respondents attended[3]. In Chile, there are three types of schools, classified according to their funding: public schools, semi-public schools, and private schools. Public schools are fully funded by per student subsidies granted by the State (vouchers) and they are administered by municipalities. Semi-public (or private-subsidized) schools also receive vouchers, but they are run by the private sector, and many of them charge an additional fee. Private schools are fully funded by fees paid by the students' families and are administered privately (Contreras, 2001). In the highly segregated Chilean school system, generally children from the better-off families attend private schools, the poorest attend public schools and semi-public schools receive students from mid-income families. For instance, in 2015, 24 percent of the children attending public elementary schools lived in poor households, a percentage that reached only 14 percent among children attending semi-public schools and less than two percent among children attending private schools[4]. We use the type of elementary school people attended as a proxy of socioeconomic status of origin. The reference category is public schools.

Considering our research questions, we included gender (male/female) and the respondent's year of birth. The rest of our covariates are categorical indicators of ethnicity (self-identification with any indigenous group, versus no identification), urban/rural status (a dichotomous indicator, which depends on the size of the population of the municipality where the respondent lives), their geographical area of residence (Northern, Central, Southern[5], with Southern as the reference group), migrant status (proxied by respondent's nationality; people with dual

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[3]Alas, our data do not include more direct indicators of the socioeconomic status of the respondents' families of origin, such as their fathers' educational attainment or occupation. The alternative of limiting the sample only to young people living with at least one of their parents (given that the data includes the education and occupation of every household member) implies losing 43% of the sample, which is why we prefer to use a retrospective question about the type of elementary school the respondent attended as a proxy.

[4] These percentages are computed using CASEN 2015, and are available upon request. They correspond to poverty defined as insufficiency of income. An alternative measure is multidimentional poverty (inability to satisfy needs in five dimensions). The figures follow the same pattern: 29 percent of children attending public elementary schools, 19 percent of children attending elementary semipublic school and 5 percent of children attending elementary private schools are poor.

[5]At the time of the data collection, Chile was geographically divided into 15 regions. The regions are numbered I to IV, plus XV (which is the farthest Northern region, but was created more recently, hence its number) and are classified as Northern. Regions V to VIII plus Region XIII, where the Capital city is located, are classified as Central. Regions IX to XIV are classified as Southern.

nationality are classified as immigrants), and disability status (self-reported as one or more of the following conditions: vision impairment, deaf or hard of hearing, speech disorders, physical disability, intellectual or mental disabilities, and psychiatric disabilities). We also include two variables related to sexual debut: age at first sexual encounter and the use of contraceptives at first intercourse, as a dichotomous indicator (used any versus did not use). All of these covariates can be considered prior to treatment assignment, perhaps with the exception of area of residence, although residential mobility tends to be low in Chile.

Regarding missing data, our treatment indicator has missing information for 105 respondents (1.2 percent of the total sample). We excluded these cases because the statistical procedure we followed required no missing values in the treatment indicator. Therefore, our analytic sample is made up of 9,298 people. Only three of the covariates we considered have missing values, namely, type of school (7.2 percent), the sexual debut related variables, age at first sexual encounter (34.9 percent) and contraceptive use at first sexual encounter (33.9 percent). These last two are especially large, because the corresponding questions were not applied to those who have not initiated their sexual lives, but the statistical approach we followed deals with missing variables in the covariates, as explained below.

After a sample description according to gender, our analysis starts with a bivariate description of the variables we are studying according to the treatment/control classification. We assess the significance of the differences between men and women or treated and control subjects, using tests of differences in means. We then use a propensity score analysis to estimate the effect of teen parenthood on high school dropout status. As mentioned earlier, for efficient inference and accurate estimation, we would like to have the treatment and control groups be as similar as possible. Randomized experiments use randomized assignment mechanisms to ensure this similarity, with the resulting groups being only randomly different from each other, either in observed or unobserved variables. Given the difficulties of implementing randomized experiments in social sciences (undoubtedly in the area of our concern, adolescent fertility and educational outcomes), propensity score analysis is an often-used approach with observational data. Propensity score analysis assumes strong ignorable treatment assignment, which implies that the treatment is independent from the outcome, *given the covariates*, and that all observed cases have a positive probability of receiving the treatment (Rosembaum, 2010).

Most of the applications using propensity scores are in matching, but there are potential drawbacks in this type of analysis, given that in some cases a very large number of subjects may be needed in the control group, and for some matching techniques, the use of a caliper and the imposition of a number of subjects being matched in the control group may result in an important sample reduction, since many subjects in the control group are not used. Alternatively, treated subjects may be discarded if a match cannot be found. Besides matching, propensity scores may be used for weighting the data and running different types of models, using all the subjects in the sample, which maintains the statistical power to detect treatment effects. Using the propensity scores as weights, we downsize the importance of cases depending on their propensity scores (Olmos & Govindasamy, 2015). We adopt this strategy for estimating the effect of teen parenthood on high-school dropout status for men and women in Chile.

Our strategy encompasses obtaining the propensity scores and then using them to create weights that are applied to a generalized linear probability (GLM) model of high-school dropout status, using the identity link function[6]. First, the probability of exposure to treatment (the propensity score) is estimated using a generalized boosted regression model, which has been argued to be superior to logistic regression as an estimation model for propensity scores (Ridgeway, McCaffrey, Morral, Burgette, & Griffin, 2017). We use the covariates we listed above (SES of origin, year of birth, urban/rural status, region, race/ethnicity, migrant status, disability status, age at first sexual encounter, and contraceptive use at first intercourse). If there are missing values in the covariates, this procedure constructs weights that also balance missingness in the treatment and control groups (Ridgeway et al., 2017). We then assess balance, defined as the similarity of the distributions of the full set of covariates between the treatment and control groups (Stuart, 2010). We compare standardized effect sizes (the difference in means of each covariate between the treatment and the control group, divided by the standard deviation of the full treated group)[7]. The balance of covariates with missing values is assessed by itself. We compare our results to the balance obtained using other strategies (nearest neighbor with different calipers, pair matching, full matching), verifying that our strategy has the best balance. The propensity scores estimated in this stage are used to obtain weights. The weight of each treatment case is 1 and the weight of each control case is $w_i = 1/(1-p(x_i))$. Therefore, control cases with features that are dissimilar to treatment cases are given small weights, whereas control cases with features that are very similar to treatment cases are given greater weights. (Ridgeway et al., 2017).

The last part of our analysis consists of modelling the probability of high-school dropout status. We use GLM (identity link), estimating six specifications. The first specification regresses high-school dropout status only based on the treatment (teen parenthood). The second specification

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[6]A GLM model with identity link function is equivalent to a linear probability model. The developers of the package we use in R (twang) use the GLM terminology, so we maintain their terms for consistency.

[7] We also check the Kolmogorov-Smirnov (KS) statistics. Results are available upon request.

incorporates the sociodemographic covariates. The third specification adds the two indicators of sexual debut. The fourth to sixth specifications are the same as these first three, but restricting the sample to people who are 20 years old or more, considering that our sample is age-heterogeneous and that those respondents who are still in their teen years may experience either parenthood or dropping out of school during their adolescence after the data collection. All specifications employ the weights from the propensity score analysis, that is to say, we use a combination of propensity score weighting and covariate adjustment. The covariate adjustment in our models accounts for small residual biases and increases efficiency in the estimates (Stuart & Rubin 2007), whereas the propensity score weighting helps to account for any nonlinearities or functional form problems in the regression controls (Curtis et at, 2007).

We estimate the average treatment effect on the treated (ATT), which corresponds to the effect of adolescent parenthood on high-school dropout status among those teens who experienced adolescent parenthood. Analyses were run for men and women separately. We did not use sample weights in any of our analyses, although we repeated our descriptive analysis using survey weights and stratification and found similar results. These results are available upon request. All the analyses were performed using the statistical software R. We used the package `twang` (Ridgeway et al, 2017) for the propensity weighting approach.

Results

Table 1 summarizes the characteristics of the sample. Slightly more than half of the sample are women. Ten percent of the respondents dropped out of school, and even though the proportion is a bit higher among women than men, this difference does not reach statistical significance at the bivariate level. Differences between men and women are significant and large, when it comes to adolescent parenthood, which reaches 20 percent among women but only eight percent among men. The large difference between male and female teen parenthood raises the question of men under-reporting fatherhood in this survey. We cannot directly answer that question, but we can use birth records, from the vital statistics system (DEIS, s.f), to compute the male and female age-specific fertility rates, for the relevant age interval, as a benchmark. In 2015, the female fertility rate among women 15-19 years old was 39.3 per thousand women, whereas for men in the same age interval it was 16 per thousand. In a similar vein, a recent press release by the National Institute of Statistics indicates that in 2018 there were 47 births to men under 15 years old and 472 births to women in the same age interval; in the next age-interval (15-19), there were 7,028 births to teen fathers, which is about half the births registered to teen mothers (16,897) (Instituto Nacional de Estadísticas, 2020). In addition, note that non-response for father's age in the 2015 births records reaches 9.2 percent, whereas for mother's age it is only .02 percent. Hence, even though some teen fatherhood may be underreported in birth records --registered as non-response--, the difference between male and female parenthood both in ENAJU and in the birth records suggest that women partner with older men[8].

The mean age at sexual debut is 16 and about three quarters of the sample declared they used some type of contraception at their first intercourse. A small proportion of the respondents attended private elementary schools, which we consider a proxy of high SES. The percentage living in rural areas is similar to the percentage self-identifying as indigenous, about 12-13 percent. About half of the respondents live in the Central area of the country, less than three percent are immigrants, and seven percent report one or more disabilities.

[TABLE 1 HERE]

Table 2 explores the association between teen parenthood and all of the variables we considered. Starting from our outcome, dropout status is more than four times more likely among people who became parents in their teenage years than among people who went through adolescence childless. Most of the people who become parents during their adolescence are women. As expected, people who became teen parents are more likely to have a socioeconomically disadvantaged background, as proxied by the type of elementary school they attended. Youngsters who became parents in their adolescence are more likely to live in the Northern area of Chile than youngsters who go through their adolescence childless, and youngsters who did not have children in their teen years are more likely to live in the Central area of the country than teen parents. The proportion of teen parents is higher among migrants and slightly lower among people with disabilities. As expected, the mean age at first sexual encounter is earlier for people who became parents as compared to people who did not have children during their adolescence, and the rate of contraceptive use at the first sexual encounter is lower.

[TABLE 2 HERE]

[8] Because the male age-specific (15-19) fertility rate is not the same as the percentage of young men who became teen fathers, we used another data source to get a figure we could directly compare to the 8 percent teen fatherhood we find with ENAJU. We used the Survey of Social Protection (Encuesta de Protección Social, EPS), which also had a 2015 round]. This is a nationally representative survey, covering the population 18 years old and older. It is the only nationally representative survey including a history of all the births men and women have in their lives. We selected men 18-29 (the most similar to the ENAJU sample, which is 15-29) and used their dates of birth in addition to the dates of birth of their children to identify among young men who became fathers in their adolescence. The percentage of teen fatherhood we obtained is 11 percent, a bit higher than the eight percent we obtain with ENAJU. Note that ENAJU includes younger men (under age 18) who are less likely to become fathers, which could partially explain the difference.

Next, we move to our propensity score analysis. Table 3 shows the balance of the covariates before and after the implementation of the generalized boosted model to generate the propensity, that is to say, it summarizes the effect of matching and reweighting on the covariates. The first panel shows the results for women and the second half the results for men. As can be seen, there were significant differences between the treatment and control groups in several of the variables before obtaining the propensity scores, and there were more significant differences in the sample of women than in the sample of men. However, none of those differences are significant after reweighting to balance the propensity scores, neither in the male or female samples. Figure 2 displays the balance obtained on the observed covariates after our propensity score analysis, making clear that the procedure made the statistically significant differences we observed before obtaining the propensity scores insignificant.

[TABLE 3 HERE]

We next obtained the ATT of teen parenthood on high-school dropout status, using different covariate specifications. These results are summarized in Table 4 (the complete results of these models are included in the Appendix, Table A1). Given that these are GLM using the identity link function (which is the same as a linear probability model), the coefficients should be interpreted as marginal probabilities. We start from a model with no covariates, which shows the effects of teen parenthood on high-school dropout status after dealing with the observed selectivity of teens into parenthood, and we next adjust for sociodemographic and sexual debut variables. Across the first three specifications, the estimated effect on high school-dropout status for women decreases from 19.9 percent to 17.5 percent (a decline of about two percentage points), but it does not lose statistical significance. For men, the decrease is greater, going from 16.2 percent to 9.9 percent, and it also remains significant (a decline of about six percentage points). In practice our best estimate -- after adjusting for all the controls we were able to include-- indicates that we would observe 18 percent less dropping out among the young women who became mothers in their teens, if they had not became mothers then. The corresponding effect for young men who became fathers is 10 percent less dropping out. The results are similar when we exclude from the sample subjects who are still in their adolescence, although they are slightly reduced for women. The estimated ATT, including all covariates, within this restricted sample, is shown in specification 6, and it reaches 16 percent for women and 10 percent for men.

[TABLE 4 HERE]

Discussion

In this study, we use a combination of propensity score techniques and regression techniques to investigate the association between high-school dropout status and teen parenthood for men and women. Most of our estimates indicate that, after correcting for observed selectivity into parenthood, the dropout rate of teen mothers is 18-16 percent higher than that of women who did not bear children in their teen years. The corresponding effect for men is 10 percent. To our knowledge, this is the first time this type of analysis has been applied to study the association between teen parenthood and educational outcomes in both men and women in Chile. One previous study (Berthelon, Kruger & Eberhard, 2017), on women only, using a fixed-effect approach on a sample of sisters, found that teen motherhood reduces the probability of completing high school by 16 percent, a result that is similar to ours, even though the samples are not directly comparable. The other Chilean study on this topic, also analyzing only women (Berthelon & Kruger, 2017), also finds that adolescent mothers are less likely to complete high school. More specifically, the authors report an OR=0.23, which is not directly comparable to our results, not only because we use a different model, but we also focus on women in a different age range, and use a different covariates adjustment[9].

Methodologically, we have relied on propensity score analysis to deal with the selectivity issue. The main advantage of propensity score techniques is that they reduce the large covariate bias between treatment and control groups, by making them comparable in all observable variables used in the analysis. By using propensity score weighting, instead of the most popular propensity score matching application, we avoid a major drawback of matching, which is the loss of statistical power that may prevent us from detecting treatment effects –loss of power that derives from the sample reduction that happens when subjects in the treatment groups are discarded because a match cannot be found or subjects in the control group are not used (Littnerova et al, 2013). We combine propensity score weighting with regression modeling, in which the covariates can further reduce small residual biases, increasing the efficiency in the estimation (Stuart & Rubin 2007). We believe that this strategy gives us an advantage in terms of the generalizability of the results—given that we use a recent nationally representative survey of the youth-- , an issue that may become a drawback when using other strategies, such as fixed-effects models on sibling samples or an instrumental variable approach.

[9] Berthelon & Kruger study several outcomes in their 2017 article. We have reported here only the one that is closest to our research question, and only for the analytical sample that resembles ours the most, which is 24-year-old women. They included as covariates the educational attainment of the respondents' mothers and fathers, the respondents' year of birth, the rural location of their households, and their region of residence.

The propensity score techniques, though, cannot handle the selectivity due to unobserved variables, and there certainly are unobservables in the data we use which are critical to teen parenthood. For instance, we do not have access to variables that measure religious values prior to treatment, the power-imbalance in terms of gender in adolescent romantic relationships, or the role of friends. We acknowledge that these variables are important, but since they are unobservable, we cannot assess the magnitude of the bias of excluding them from our estimations. That is why our results make no causality claims, and we recognize the superiority of techniques that deal with unobservables when causality is the goal. Our goal in this study is not to establish a precise causal effect of teen parenthood on high-school dropout status, but to study the association between teen parenthood on high-school dropout status, obtaining results that may be generalized to the total population of young Chileans, because we believe this type of results may have important policy implications. However, because the problem of unobservables is critical when studying teen parenthood and educational outcomes, we tested how large the effect on unobserved heterogeneity would have to be in order to nullify the propensity score estimates. Our results do not suggest that there are unobserved variables which would make our estimate of the ATT null for women or men (even though the evidence is stronger for the sample of women).

In terms of the magnitude of our results, we find that parenthood is associated with a greater setback in the educational trajectories of women than men (recall that the estimated effect of teen parenthood on dropout status is 18-16 percent for women; 10 percent for men). Considering that life domains are mutually interdependent, as the life course perspective highlights, it is likely that the greater educational setback for women will have a greater impact on their employment trajectories, and consequently, on their financial autonomy.

Even though men are not in the most prejudicial position, one of this study's major contribution is exploring the relation between high-school dropout status and teen parenthood for men and not only for women. Even though the magnitude of our results is smaller for men than women, they still suggest that teen fatherhood is associated with a setback in young men's educational careers --perhaps other life trajectories are also affected by this event. Giving visibility to the fact that a not negligible proportion of men who experience teen fatherhood reach a low educational attainment is important, because it calls for policies targeted toward teen fathers in the educational realm, policies which in the case of Chile are underdeveloped.

In terms of public policies and interventions, we believe that these interventions may be most helpful if they consider both genders. We need policies oriented toward decreasing teen parenthood, for instance, promotion of more effective sexual education or access to sexual and reproductive health services for adolescents, which in Chile are currently not very widespread. But we also need remedial policies, operating after the teens have already become parents, and which probably will be more effective in taking into consideration the traditional parenting norms*,* according to which mothers are the primary caregivers and fathers are the breadwinners, because the interventions carried out are likely to yield different results for teen mothers than for teen fathers.

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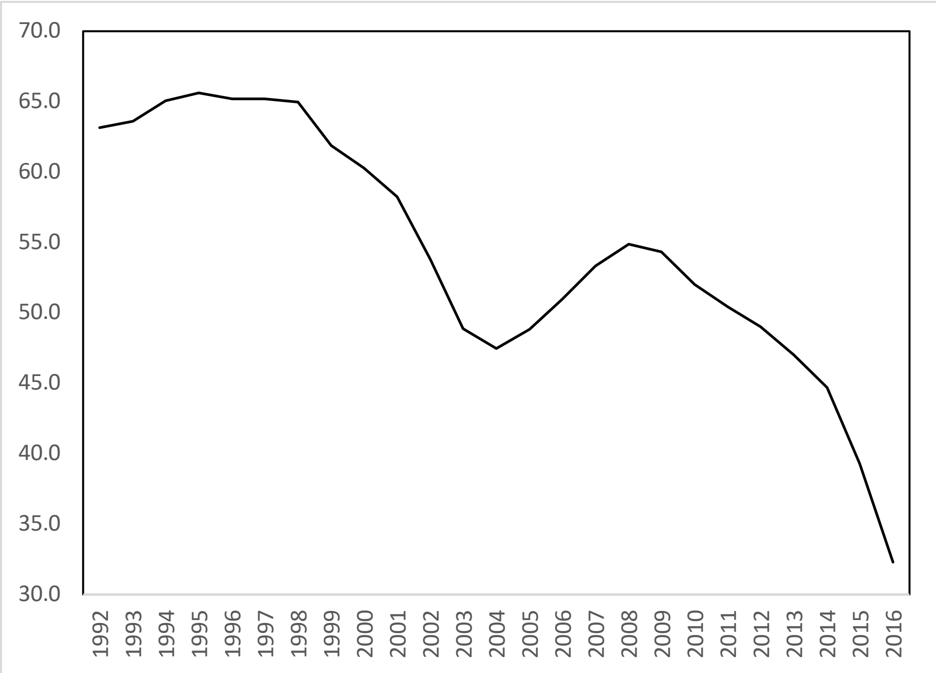
Figure 1: Number of Births per 1,000 Women 15-19,

Chile, 1992-2016



Source: DEIS (Departamento de Estadísticas e Información en Salud, Department of Information and Statistics in Health), Chile

Table 1: Sample Description, according to Gender (n=9, 285)

	Men (n=4,322)	Women (n=4,951)	Total
High school dropout	9.82	11.06	10.48
Teen parenthood ***	7.97	20.24	14.52
Mean year of birth ***	1994.11	1993.45	1993.76
Type of school			
Private***	5.56	3.82	4.63
Semi-public	37.73	36.74	37.20
Public*	56.70	59.44	58.17
Rural status *	12.30	13.89	13.15
Area of residence			
Northern **	20.29	22.79	21.62
Central **	58.48	55.25	56.76
Southern	21.23	21.96	21.62
Ethnicity	12.19	13.55	12.91
Migrant status*	2.19	2.98	2.62
Disability status	6.35	7.02	6.70
Mean age at sexual debut ***	16.17	16.71	16.46
Used contraception at first intercourse ***	74.16	72.76	73.40

**** p<0.01, ** p<0.05, * p<0.1*

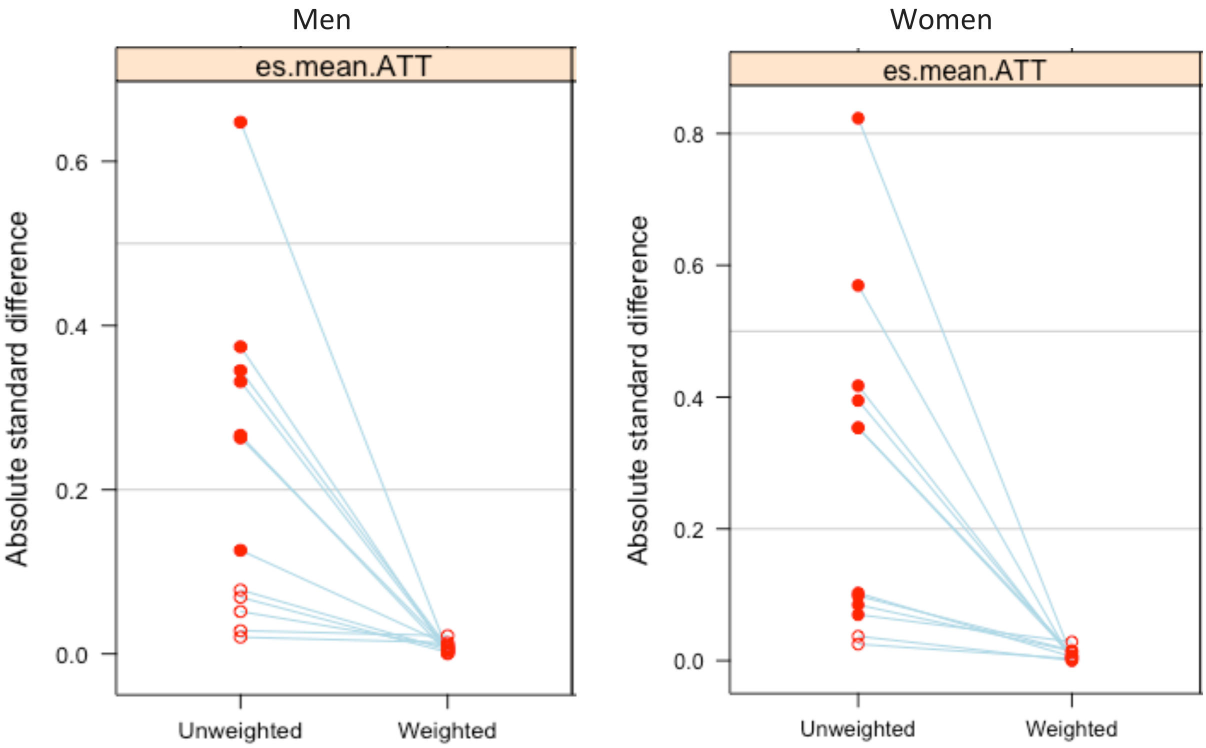
Table 2: Association between Teen Parenthood, High School Dropout Status and other Covariates (n=9, 285)

	Treated (teen parent)	Control (not teen parent)
High school dropout ***	29.99	7.16
Female ***	74.40	49.80
Mean Year of Birth, (Std. Dev)	1991.85	1994.09
Type of school		
Private***	2.83	4.91
Semi-public ***	25.68	39.14
Public***	71.47	55.93
Rural status	13.72	13.06
Area of residence		
Northern ***	29.08	20.36
Central ***	49.55	57.98
Southern	21.36	21.66
Ethnicity	13.35	12.85
Migrant status *	3.70	2.43
Disability status**	5.11	6.97
Mean age at sexual debut (std. dev) ***	15.62	16.64
Used contraceptive at first intercourse ***	60.42	76.24

**** p<0.01, ** p<0.05, * p<0.1*

Figure 2: Comparison of Absolute Standardized Bias

Before and after Propensity Score Weighting



For each variable considered in the modeling, lines connect dots for the same variable with and without weighting. Solid dots denote statistically significant differences (*p*<0.05)

Table 3: Covariate balance with propensity score weighting, women and men

	Unadjusted			Adjusted		
	C	T	Std effect size	C	T	Std effect size
Sample of Women						
Year of birth	1993.878	1991.780	0.569***	1991.770	1991.780	0.003
Type of school						
Private	0.039	0.023	0.107**	0.023	0.025	0.011
Semi-public	0.373	0.229	0.345***	0.251	0.249	0.005
Public	0.528	0.665	0.298***	0.665	0.665	0.000
Area of residence						
Northern	0.209	0.303	0.205***	0.300	0.303	0.006
Central	0.570	0.483	0.175***	0.483	0.483	0.002
Southern	0.221	0.214	0.016	0.216	0.214	0.005
Rural status	0.136	0.150	0.037	0.150	0.150	0.000
Ethnicity	0.134	0.143	0.025	0.142	0.143	0.003
Migrant status	0.027	0.041	0.070 **	0.035	0.041	0.028
Disability status	0.074	0.055	0.085	0.058	0.055	0.014
Use of contraception at first intercourse	0.771	0.597	0.354***	0.603	0.597	0.012
Age at first intercourse	17.055	15.685	0.823***	15.675	15.685	0.006
Age at first intercourse missing	0.379	0.181	0.417***	0.180	15.685	0.003
Use of contraception at first intercourse missing	0.369	0.183	0.395***	0.181	0.181	0.006
Type of school missing	0.060	0.084	0.088**	0.079	0.084	0.016

(Continues)

(Continued)

	Unadjusted			Adjusted		
	C	T	Std effect size	C	T	Std effect size
Sample of Men						
Year of birth	1994.291	1992.052	0. 647***	1992.063	1992.052	0.003
Type of school						
Private	0.053	0.035	0.097*	0.033	0.035	0.012
Semi-public	0.356	0.255	0.232***	0.258	0.255	0.007
Public	0.514	0.626	0.272***	0.631	0.626	0.009
Area of residence						
Northern	0.198	0.255	0.130 **	0.254	0.255	0.002
Central	0.212	0.212	0.002	0.537	0.533	0.008
Southern	0.589	0.533	0.112 **	0.208	0.212	0.009
Rural status	0.101	0.125	0.078	0.105	0.101	0.013
Ethnicity	0.123	0.107	0.052	0.112	0.107	0.016
Migrant status	0.022	0.026	0.028	0.025	0.026	0.009
Disability status	0.066	0.041	0.126	0.040	0.041	0.004
Use of contraception at first intercourse	0.755	0.625	0.755***	0.628	0.625	0.004
Age at first intercourse	16.244	15.454	0.374***	15.480	15.454	0.012
Age at first intercourse missing	0.374	0.209	0.345 ***	0.211	0.209	0.004
Use of contraception at first intercourse missing	0.361	0.203	0.332***	0.198	0.203	0.010
Type of school dependency missing	0.077	0.084	0.026	0.078	0.084	0.020

**** p<0.01, ** p<0.05, * p<0.1*

Table 4. Estimates of average treatment effect on the treated (ATT)

Description of specification		Women	Men
Specification 1:	coeff.	0.199***	0.163 ***
Propensity score weights, no covariates	std. error	(0.017)	(0.026)
Specification 2:	coeff.	0.196 ***	0.142***
Propensity score weights, sociodemographic covariates	std. error	(0.017)	(0.025)
Specification 3:	coeff.	0.175***	0.099***
Propensity score weights, all covariates	std. error	(0.018)	(0.026)
Specification 4:	coeff.	0.182***	0.161***
Propensity score weights, no covariates, sample restriction	std. error	(0.019)	(0.028)
Specification 5:	coeff.	0.184***	0.134***
Propensity score weights, sociodemographic covariates, sample restriction	std. error	(0.018)	(0.027)
Specification 6:	coeff.	0.160***	0.101 ***
Propensity score weights, all covariates sample restriction	std. error	(0.020)	(0.029)

****p ≤ 0.001; **p ≤ 0.01; *p ≤ 0.05; (*) p ≤ 0.1.*

Propensity score weights were obtained from generalized boosted regression models.

Specifications 2 and 5 exclude age at first sexual encounter and use of contraception at first intercourse

APPENDIX

Table A1: GLM results for different specifications, including covariates coefficients.

(Teen parenthood as treatment)

	Specification 2	Specification 3	Specification 5	Specification 6
Sample of Women				
Intercept	0.312 (4.434)	2.470 (5.073)	6.037 (6.198)	8.090 (6.694)
Treatment (teen parent)	0.196 (0.017) ***	0.175 (0.018) ***	0.184 (0.018) ***	0.160 (0.020) ***
Year of Birth	-0.001 (0.002)	-0.001 (0.002)	-0.002 (0.003)	-0.004 (0.003)
Type of school				
Private	-0.093 (0.043) *	-0.104 (0.040) **	-0.118 (0.043) **	-0.144 (0.032) ***
Semi-public	-0.134 (0.016) ***	-0.111 (0.017) ***	-0.120 (0.019) ***	-0.097 (0.020) ***
Rural status	0.070 (0.026) **	0.077 (0.029) **	0.065 (0.029) *	0.071 (0.033) *
Area of residence				
Northern	0.038 (0.023) (*)	0.026 (0.025)	0.028 (0.026)	0.010 (0.028)
Central	0.052 (0.021) *	0.052 (0.0023) *	0.047 (0.024) (*)	0.048 (0.026) (*)
Ethnicity	0.012 (0.024)	0.026 (0.027)	0.005 (0.027)	0.024 (0.030)
Migrant status	0.005 (0.045)	-0.037 (0.047)	-0.022 (0.045)	-0.075 (0.044) (*)
Disability status	-0.052 (0.024) (*)	-0.049 (0.033)	-0.035 (0.034)	-0.028 (0.037)
Age at first sex		-0.025 (0.005) ***		-0.026 (0.006) ***
Used contraceptive at first intercourse		-0.046 (0.019) *		-0.053 (0.0213) *

(Continues)

(Continued)

	Specification 2	Specification 3	Specification 5	Specification 6
Sample of Men				
Intercept	-9.141 (6.614)	-8.870 (7.180)	-14.563 (8.828)	-14.130 (9.282)
Treatment (teen parent)	0.142 (0.024) ***	0.099 (0.026) ***	0.134 (0.027) ***	0.102 (0.029) ***
Year of Birth	0.005 (0.003)	0.005 (0.004)	0.007 (0.004) (*)	0.007 (0.004)
Type of school				
Private	-0.168 (0.044) ***	-0.147 (0.027) ***	-0.162 (0.054) **	-0.144 (0.030) ***
Semi-public	-0.182 (0.022) ***	-0.146 (0.024) ***	-0.182 (0.024) ***	-0.149 (0.026) ***
Rural status	0.112 (0.046) *	0.098 (0.049) (*)	0.097 (0.050) (*)	0.076 (0.053)
Area of residence				
Northern	0.037 (0.038)	0.009 (0.041)	0.012 (0.042)	-0.017 (0.047)
Central	0.017 (0.032)	-0.006 (0.035)	0.014 (0.037)	-0.022 (0.040)
Ethnicity	0.072 (0.045)	0.064 (0.048)	0.104 (0.052) *	0.082 (0.056)
Migrant status	-0.053 (0.070)	-0.116 (0.036) **	-0.007 (0.082)	-0.079 (0.040) (*)
Disability status	-0.026 (0.059)	-0.035 (0.061)	-0.063 (0.067)	-0.107 (0.046) *
Age at first sex		-0.014 (0.006) *		-0.019 (0.008) *
Used contraceptive at first intercourse		-0.072 (0.029) *		-0.057 (0.031) (*)

****p ≤ 0.001; **p ≤ 0.01; *p ≤ 0.05; (*) p ≤ 0.1.*

Standard errors in parentheses

Table A2: Relative influence of the covariates on the estimated propensity score

(Teen parenthood as treatment)

	Women sample	Men sample
Year of Birth	28.534	44.618
Type of school		
Private	0.296	0.357
Semi-public	0.176	0.577
Public	5.661	6.748
Rural status	0.295	0.820
Area of residence		
Northern	1.382	1.167
Central	0.719	1.100
Southern	0.015	0.875
Ethnicity	0.453	0.063
Migrant status	0.400	0.203
Disability status	0.290	1.022
Age at first sex	48.981	30.022
Used contraceptive at first intercourse	10.271	12.610