The Educational Consequences of Adolescent Childbearing and Union Formation in Brazil

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Abstract

While Brazil has high rates of adolescent fertility for its below-replacement total fertility rate, we know little about the causal effects of adolescent childbearing and adolescent union formation for women's education. In this article, we examine unique data from the 2013 School-to-Work Transitions Survey to address the consequences of adolescent childbearing and adolescent union formation on educational outcomes of Brazilian young women. We apply several analytical strategies to address the endogeneity between adolescent childbearing and educational outcomes. Our findings suggest that childbearing during the teenage years is detrimental to the educational attainment of Brazilian women, and that educational disadvantages persist once we take into account mother's selection into adolescent childbearing. The penalty for adolescent mothers ranges from -1.66 to -1.80 fewer years of schooling and from a 41 to 35 percent difference in the probabilities of graduating from high school. Additional findings show that marital unions among adolescent mothers have a compounding role at further hindering women's educational progress. Combined, our findings suggest that young mothers, particularly those in a marital union, face additional layers of disadvantages, demonstrating that early family formation is a meaningful *stratifier* for women in an already highly-stratified society.

INTRODUCTION

The consequences of adolescent childbearing have been a longstanding issue of interest in the social sciences. Globally, women whose first birth occurred in their adolescent years tend to exhibit worse social and economic outcomes than those who delayed childbearing or remained childless (Hoffman, Foster, and Furstenberg 1993; Fergusson and Woodward 2000; Lee 2010; Kane, Morgan, Harris, and Guilkey 2013; Diaz and Fiel 2016). Yet, scholars acknowledge that these differences may not reflect a causal effect of an adolescent birth. That is, because maternal age at first birth is not randomly distributed in the population, adolescent childbearing and the most salient social outcomes in young adult life are intertwined. While scholars have made significant progress toward understanding the effects of teenage childbearing in industrialized countries, the lack of data identifying the timing of events has hindered the appropriate examination of this question in low and middle-income countries.

Importantly, a different array of intervening elements emerges when examining the causal effects of adolescent motherhood on educational outcomes in low- and middle-income

societies, particularly highly-stratified contexts. One of the most salient is the role of early union formation. In such contexts, where social safety nets are limited, the presence of a partner is potentially crucial in explaining women's socioeconomic outcomes. Further, where the impact of early unions on women's educational outcomes has been largely overlooked, the effect of a union for adolescent mothers is not straightforward. That is, conceptually, we can think of mechanisms by which early unions could be either detrimental or positive to further the educational attainment of adolescent mothers, particularly, in contexts of disadvantage. In this article, we posit that marital unions are important to our understanding of the impact of adolescent childbearing on education in the specific context we analyze, Brazil.

The purpose of this article is two-fold. First, we seek to evaluate the causal effects of adolescent childbearing on young women's educational attainment in Brazil. We know a lot about the effects of education on fertility, but very little about how fertility impinges or advances education, particularly in low- and middle-income countries. We focus on two educational outcomes—complete years of schooling and graduation from high school—both of which are critical elements of human capital formation. Our second goal is to explore the impact of unions on the educational attainment of adolescent mothers. We use data from the School-to-Work Transition Survey (SWTS), a unique, nationally representative data source collected in 2013 by the International Labor Organization (ILO). The SWTS includes rich information on early socioeconomic background, based on retrospective questions, therefore before childbearing during adolescence. It is worth noting that the DHS—Demographic and Health Survey, the main dataset used for the study of fertility and education in middle- and low-income countries—does not have such information. This allows us to implement a number of strategies for addressing our research questions while taking selectivity into account. The SWTS is particularly unique given the lack of data on individuals' life trajectory in non-industrialized contexts.

Our study fills important gaps in the literature about the consequences of demographic events during adolescence in at least three important ways. First, while most studies have examined only one aspect of family formation at a time, we consider how both childbearing and union formation impact women's education. Second, we address the issue of selectivity into early childbearing by using a set of strategies suitable for exploiting retrospective data. While with their own set of limitations, the methodologies we use allow us to explore the extent to which the negative associations between early childbearing and educational outcomes hold after selection is accounted for—unmasking whether young family formation has a causal effect on education. Third, we use nationally representative data to study this question in the Brazilian context. The few studies that have examined the consequences of early demographic events in low and middle-income countries have used samples with small geographic representation or have not dealt with selectivity issues.

Brazil offers a compelling case for the examination of the causal consequences of adolescent childbearing and adolescent union formation on young women's education. On the one hand, the country has relatively high adolescent fertility rates combined with below-replacement total fertility rates (TFRs), a potential consequence of Brazil's high levels of inequality. In most low-fertility countries, a delay in childbearing occurred as fertility rates

declined, shifting the fertility curve to high age groups. In contrast, during Brazil's fertility transition there was a rejuvenation of the fertility schedule. That is, the fertility of young women had a *growing* relative contribution to the TFR in the country (Verona 2018), with declines in fertility occurring mainly through a pattern of stopping rather than spacing births throughout the reproductive life (Alves and Cavenaghi 2008). Likewise, despite below-replacement fertility levels, age at first union remains young. According to UNFPA, Brazil persists as one of the 41 countries in the world where the prevalence of early marriage is 30 percent or more (United Nations Population Fund 2012). Finally, the country remains one of the most unequal in the world, despite recent improvements, and such high inequality likely fuels the fertility patterns we examine. This overall scenario raises questions on whether early family formation translates into further educational and social stratification among Brazilian women.

The article is organized as follows. The first section summarizes current explanations for the impact of adolescent childbearing, with a focus on educational outcomes. It is followed by a discussion of the potential mechanisms driving an influence of a marital union on the educational attainment of adolescent mothers. We next discuss the issue of selection into adolescent childbearing. The fourth section introduces the data and methods, paying particular attention to the strategies we use in order to address the endogeneity of teenage childbearing with educational attainment. We then present our results, and subsequently discuss conclusions and implications of the study.

Adolescent Childbearing and Women's Education

The period in women's life course when they give birth has important consequences for their educational trajectories (Powell et al. 2006). Of particular importance is whether girls give birth during adolescence, when they have not yet finished a successful educational path and have not established a career (Lloyd et al. 2005). A large body of literature has established at least three major mechanisms explaining a detrimental effect of adolescent childbearing to women's educational outcomes.

First, considering a purely resource-allocation perspective, young mothers are at a clear disadvantage vis-à-vis older mothers because they have fewer autonomous resources available to invest in their children while also continuing in school (McLanahan 2004). Women traditionally have the primary responsibility of child care and childrearing, and parenthood for them often coincides with a shrinking of opportunities and reduced scope for independent action (Lloyd et al. 2005). If working, adolescent mothers are more often in low-paid jobs, which yield fewer resources and oftentimes, require demanding hours.

Second, another stream of research suggests that young mothers likely have fewer psychological and emotional tools than older mothers to successfully undertake the major responsibilities involved in motherhood (Berlin et al. 2002). When conceptualizing the consequences of adolescent childbearing, researchers often use an underlying definition of the teenage years as a period of exploration and social-emotional development, combined with human capital formation (Arnett 2000). Within this framework, adolescent childbearing marks an interruption of such processes, and is clearly detrimental to adolescent well-being and social-emotional growth. Adolescents are therefore less likely to be ready for

parenthood's demanding obligations, with preferences and tastes that are not conducive for combining the roles of mother and student.

Third, adolescent childbearing is also associated with health disadvantages for both mothers and their newborn, primarily because adolescents' physiological development is incomplete. In fact, pregnancy and childbirth complications are the second cause of death among adolescents globally (LeGrand and Mbacké 1993). Newborns born to adolescent mothers are also more likely to have low birth weight (Alam 2000), with the risk of long-term effects (Bacci, et al. 1993; LeGrand and Mbacké 1993).

Contrary to the notion described above, several scholars have cast doubt on the detrimental consequences of adolescent childbearing for all women. This body of research emphasizes the critical importance of context in explaining the consequences of adolescent childbearing. Low-income African American women in the US, for example, face barriers to bearing and raising children at an older age to the point that adolescent childbearing is understood as an adaptive strategy (Geronimus, Korenman, and Hillemeier 1994; Geronimus 2003). Empirical studies have confirmed that low-income African American women face a smaller penalty from adolescent childbearing than white women (Furstenberg et. al. 1987; Heard 2007). Combined, this research suggests that within disadvantaged populations, early childbearing might not be deviant; rather, a pattern of early fertility may be responsive to social-structural constraints and limited opportunities. That is, the actual consequences of teenage childbearing in women's lives may be small or even negligible in contexts where socioeconomic disadvantage is the main driver of negative outcomes.

Within this line of thinking, the extended family may also play a role to ameliorate disadvantages associated with adolescent childbearing. When adolescent mothers transfer childcare responsibilities to the extended family and grandparents, young mothers can remain enrolled in school (Johnson-Hanks 2002; Kaufman, de Wet, and Stadler 2001; Marteleteo and Dondero 2013). It is likely that the costs associated with early childbearing might be reduced because a large proportion of adolescent mothers lives in extended households, mainly with their parents or in-laws (Marteleteo and Dondero 2013).

Brazil offers a different context from that of developed countries within which adolescents frame their decisions and gather support. A middle-income country with low overall levels of education, adolescent childbearing is more normative in Brazil than in other contexts with the same low overall fertility rate. While fertility has declined dramatically in the last few decades, from a total rate of 6.3 in 1960 to current below-replacement levels (1.8) (UNFPA 2010), childbearing ages remain relatively young. Indeed, Brazil witnessed an uptick in teenage fertility rates in the 1980s and 1990s, a time period in which fertility rates for all other age groups fell. The combination of below-replacement total fertility, high rates of adolescent fertility and a rejuvenated fertility schedule underscore the importance of examining the consequences of adolescent childbearing for the formation of human capital for Brazil's young women. Further, as in other contexts, young childbearing is notoriously more prevalent among women of low socioeconomic status (SES) (Cavenaghi 2013).

Another important feature of adolescent childbearing in Brazil is that it often takes place within or leads to a marital union (Cabral and Heilborn 2005). A study of three metropolitan areas reports that the birth of a child led to a marital union for 31.3 percent of the young mothers; another 21.8 percent were already in a union at the time of the birth (Dias and Aquino 2006). In our data, as much as 57 percent of adolescent mothers also started marital unions at age 18 or younger. This pattern sharply departs from the context of adolescent childbearing in the US and other developed countries, where adolescent births rarely take place within or lead to a marital union (Martin et al. 2010). Given the prevalence of marital births among teenage child-bearers in Brazil, we further investigate the impact of these unions in the educational outcomes of adolescent mothers.

Adolescent Childbearing, Marital Union, and Women's Education

The second goal of this article is to investigate whether adolescent marital unions shape the association between adolescent childbearing and education among Brazilian women. Research on how adolescent union and adolescent childbearing intertwine to shape educational (dis)advantages is scarce and the potential effect of adolescent unions on the education of adolescent mothers remains unclear, particularly in the context of a developing country.

As with adolescent childbearing, adolescent marital unions mark an interruption of the exploration and social-emotional development that are typical of this age, likely conferring detrimental consequences to young women (Arnett 2000). Marital unions during adolescence could be consequential for adolescent mothers through a number of mechanisms. First, adolescent unions tend to be more fragile than later unions. They are commonly less stable, and of lower quality (Lyngstad & Jalovaara, 2010). That is, even if early unions do have a protective effect for certain groups, this impact is probably short-lived.

Second, according to the specialization perspective, household formation leads to a gender-based division of labor, a result of a joint strategy to maximize household well-being (Becker 1991). Women's time spent in household chores tends to increase after marriage, while men's decrease (Hersch and Stratton 2000; Gupta 1999). The specialization perspective would in turn predict a compounding detrimental effect of marital unions among young mothers, through the mechanism of lowering the incentives for accumulating human capital. Thus, scholars have shown that gender continues to be a critical factor organizing household interactions, where mothers are still mostly responsible for childcare and managing the home (Killewald and Gough 2013). A greater investment in children and household chores has plausibly a negative impact on women's earnings (Noonan, 2001).

Marriage also exacerbates the motherhood penalty, which further suggests that marital status has an additional impact in the proportion of time and energy that mothers allocate to their market-related activities versus house chores (Budig and England 2001; Glauber 2007). It is reasonable to expect that a similar pattern applies for educational outcomes. Studies on time use in Latin America confirm a strong gender disparity in unpaid domestic work. This evidence has demonstrated the marked prevalence of traditional gender norms within families, with marital unions strengthening the social roles of male provider and female

caregiver (Dias and Aquino 2006). In fact, Brazilian women tend to increase their time in unpaid-domestic work with each additional child, while fathers maintain theirs at the same level regardless of their number of children (Pinheiro et al. 2016).

Finally, adolescent mothers in marital unions may lose valuable support from the extended kin. Extended families may provide not only financial but childcare support, allowing adolescent mothers to stay in school. In the US, teenage mothers living in extended families are more likely to stay in school and/or work versus those who are in a marital union (Trent and Harlan 1994). In contrast, research suggest that for adolescent mothers who live with the father of their child, father's support was provided within the traditional gender roles of breadwinners and only backup child providers (Molborn and Jacobs 2015). The role of the extended family in buffering a potential negative penalty in the schooling of adolescent mothers is particularly important in contexts where the extended family provides ample support to mothers; but where such support is greater for single adolescent mothers than for those in a marital union (for Brazil, Marteleteo and Dondero 2013).

Contrary to the ideas discussed above, if we conceptually locate early union formation in contexts of disadvantage, the link between adolescent marital unions and negative outcomes becomes less straightforward. Adolescent mothers in unions come disproportionately from more disadvantaged families of origin less capable of supporting their socioemotional development and human capital formation (Amato and Booth 1997; Orbuch, Veroff, Hassan and Horrocks 2002). As individuals from disadvantaged backgrounds are more likely to experience instability, lack of supervision, or low-quality family ties in their families of origin, transitions to childbearing and marital unions may result on a positive change. Teenage mothers from such contexts might see family formation as a rational strategy intended to improve their quality of life.

Furthermore, the sizeable strain of research on the impact of (adult) marriage and cohabitation often finds beneficial effects of marital unions for a set of outcomes (for a review, Waite and Gallagher 2000). Following a new economics perspective, individuals are rational agents, and as such, they would start a marital union if they are better-off married than single, basing their decision on an analysis of opportunities, risks, and rewards (Becker, 1981). In a context of disadvantage, a marital partner might be beneficial to adolescents by providing emotional support, financial independence, and an exit from a (more) disadvantaged social location. In this view, having a child and forming a marital union might not be intrinsically opposed to the formation of human capital of adolescent mothers. While indirectly, some research points to this direction suggesting that adolescents in a context of overall disadvantage benefit from a marital union (Booth, Rustenbach, and McHale 2008; Bosick 2015).

Selection into Adolescent Childbearing

The literature on the educational consequences of adolescent childbearing can be divided in two generations based on methodological grounds. Initially, studies reported a strong and mostly negative correlation between adolescent childbearing and a number of disadvantageous outcomes in both high- (Mott and Marsiglio 1985; Hofferth 1987) and low-

income countries (Buvinic 1998; Gupta and Leite 1999). In these early contributions, fertility was modeled as exogenous to education.

A second generation of studies implemented methodological strategies to account for the endogeneity of fertility and social outcomes, yielding mixed results. These studies acknowledged that, since teenage mothers come disproportionately from disadvantaged backgrounds, failure to consider such disadvantages would likely result in an overestimation of the negative impact of adolescent childbearing. Social context is key to this line of argument. Adolescent mothers would fare worse than older mothers (or childless women) because disadvantages contribute both to selection into adolescent motherhood and to worse educational outcomes (Geronimus, Korenman, and Hillemeier 1994). Most of these studies report smaller, although still negative effects of adolescent childbearing on educational outcomes. The wide range of estimates in the literature can be a result of the many different methodological techniques used (Kane et al. 2013).

Methodologically, this second generation of studies focused on defining the appropriate counterfactual group. One approach is the use of siblings and cousins fixed effects¹ (Hoffman, Foster, and Furstenberg 1993; Levine, Emery, and Pollack 2007) following the work of Geronimus and Korenman (1992; 1993). Other studies have relied in a quasi-natural approach, comparing young mothers with older mothers who miscarried as adolescents (Hotz, McElroy, and Sanders 2005; Levine, Emery, and Pollack 2007; Marteleteo and Dondero 2013). Most miscarriages are random and result in a postponement of childbearing, therefore allowing for a more appropriate comparison group.

A third approach to address the endogeneity issue has been to use pre-fertility characteristics through strategies that reweight the data using propensity score matching, for example (Assini-Meytin and Greene 2015; Branson, Ardington, and Leibbrandt 2011; Ranchhod, Lam, Leibbrandt, and Marteleteo 2011). A fourth approach, when data is available, is to compare several methodologies and adjudicate (Kane et al. 2013).

The findings from this body of literature are mixed. In a highly-influential series of studies working with data from sisters and cousins, Geronimus and Korenman (1992, 1993) find that educational penalties for American teenage mothers become almost null once unobserved family background is accounted for. Yet, replicating Geronimus and Korenman's strategy with a different data source, Hoffman and colleagues (1993) conclude that disadvantages to education remain among young mothers. Further, using miscarriages as an instrumental variable, Hotz and colleagues (2005) find remaining small penalties in the likelihood of teenage mothers to graduate from high school. Assini and colleagues (2015), using matching procedures, find substantial remaining effects for adolescent fertility after controlling for unobserved factors. They evaluate the likelihood to depend on welfare and earned a GED or a high school diploma versus finishing college. Similarly, Kane and colleagues (2013) tested different analytical approaches and found a significant detrimental effect, ranging from 0.7 to 1.9 fewer years of education. All noted approaches rely on data

¹Sisters are more alike than unrelated women; therefore, comparing sisters who differ in their fertility timing allows for fewer unmeasured variables that correlate with adolescent childbearing. A limitation of this approach is that sample sizes are usually small and data availability is scarce, particularly if applied to developing countries.

that is rarely available for low- and middle-income countries, which can explain the dearth of studies on the causal educational consequences of adolescent childbearing in these countries.

Only a handful of studies using data on Latin American countries have investigated the impacts of teenage childbearing on education taking into account the role of pre-existing disadvantages, and most are not published in peer-review journals. Similarly to the overall literature described above, these studies have produced mixed results. Using miscarriages as an instrument, Azevedo et al. (2012) found no educational penalty for Mexican women who were teenage mothers, as well as a higher likelihood to be employed and a greater dependence on social welfare. A higher likelihood to be employed is not a clear positive indicator in developing countries, given the prevalence of precarious work among low-income women (ILO 2012).² Arceo-Gomez and Campos-Vasquez (2014) use a propensity score matching strategy and find a detrimental impact of teenage childbearing of 0.6–0.8 fewer years of schooling among Mexican women. Further, Kruger, Berthelon, and Navia (2009) used a set of instrumental variables tapping into the social acceptance of early fertility in specific geographical areas, and found that Chilean adolescent mothers were between 21 and 31 percent less likely to attend school or to complete high school versus their peers who were not adolescent mothers. Berthelon and Kruger (2014) apply both a propensity score matching strategy and fixed-effects using a sample of sisters in Chile. While their results using fixed-effects show a smaller penalty compared to those using a matching strategy, their study finds a remaining sizeable effect, with sisters who had experienced an early birth being 17 percent less likely to graduate from high school and 28 percent less likely to transition to tertiary education than their sisters who did not have a teenage birth.

We investigate the impact of teenage childbearing on educational outcomes of women in Brazil. In our own data, as much as 26 percent of women had a first child at age 18 or younger, a high prevalence even within the context of the Latin American region. A number of additional factors could either mitigate or exacerbate the educational consequences of teenage childbearing in Brazil. Adolescent fertility is more normative in Brazil, which would lead to a lower penalty. Additionally, the extended family and grandparents often have a present role in children's lives in Brazil. Social norms that favor kin cooperation at raising children could mitigate the educational penalty for adolescent mothers.

On the other hand, institutionalized support to mothers is scarce in Brazil (Marteleteo and Dondero 2013). Further, while the educational system in the country has experienced an unprecedented expansion in its recent history, access to secondary school remains far from universal (Marteleteo, Marschner, and Carvalhaes 2016). These factors, in turn, would predict a strong disadvantage for adolescent mothers. Equally important, adolescent mothers in Brazil are more likely to engage in marital unions than their peers in industrialized countries. In our data, over 50 percent of adolescent mothers were in a marital union at age 18 or

²For instance, descriptive work for Peru finds no difference in labor force participation of women who were early mothers versus those who are either childless or had children later in life (Alcazar and Lovaton 2006).

younger. Given its high prevalence, adolescent unions are an important piece in the analysis of the educational outcomes of teenage mothers.

Data and Methods

Data—We use data from the School-to-Work Transitions Survey (SWTS), a nationally-representative survey collected in Brazil in 2013 by the International Labor Organization (ILO). The SWTS contains valuable details on background characteristics of respondents while growing up, as well as retrospective questions on their family histories, such as first birth and first marital union, unique information rarely available in surveys for Latin America. The SWTS gathered information from 3,288 respondents ages 15–29. The set of questions referring to the social and economic background of respondents during their childhood allows us to create a plausible counterfactual group for analyzing later educational outcomes of women who were adolescent mothers.

The final analytical sample is comprised of 1,080 women ages 19 and older³; 90 percent of this sample has no missing values in any variable of interest (974 women). In an effort to include all relevant cases, we implement a multiple imputation procedure (Acocck 2005), with 10 imputations. We tested our models using different numbers of imputations, with negligible variation in the results. Moreover, estimates from models with imputed values are similar to estimates from models with missing values.⁴ In our final analytical sample, 28 percent of women became mothers at age 18 or younger, while 26 percent report they started their first marital union as teenagers. As noted above, among women who were adolescent mothers (N=302), about 57 percent also started a marital union while teenagers (N=171), though it is not possible to determine precisely what event happened before because for age at first marital union we have information on age and not date.

Measures—We run models with two dependent variables: completed *years of education* and *graduation from high school*.⁵ We define an adolescent marital union as occurring at ages 18 and younger, where “marital unions” include both marriage and cohabitation. The variables we used to measure socioeconomic background are *parental education*, *parental occupation*, *race*, and perceived *social class* during childhood. Parental education is the highest level of educational attainment among parents using three categories: elementary or less, high school education, and some tertiary and higher. Our measure for *parental occupation* refers to the occupation with the highest status among the parents, and has six categories: rural workers; manual, non-skilled workers and owners with no employees; manual, skilled workers; supervisors of manual work; routine, non-manual workers; and professional or managerial workers (Ribeiro 2006). *Race* is measured with a dichotomous variable (white/non-white), while our measure of *social class* during childhood⁶ has four categories: poor, low-middle, middle, and upper-middle and higher⁷. We use this set of

³We also removed the few individuals who did not report educational attainment or age at first child, as well as Asians (a total of 25 observations excluded).

⁴We use the Stata package for multivariate imputation by chained equations (MICE). The imputed models are not presented, but they are available upon request.

⁵Our data does not provide information on the exact date respondents left school. However, our educational outcomes are the result of fluid processes of school enrollment, school drop-out and grade retention. These educational markers such as the ones we use account for such fluid nature of educational attainment and minimize—though not eliminate completely—the potential issue of reverse causation.

measures as “pre-treatment variables,” affecting both *treatment* status (teenage childbearing) and our outcomes of interest. Additionally, our models include controls for age (continuous, from 19 to 29), *region* (5 categories), and bivariate measures for urban residence and whether the respondent was an internal migrant at age 15.

Methods

Adolescent Childbearing and Educational Outcomes: We first conduct Ordinary Least Square (OLS) regressions with years of schooling as outcome and logit models for predicting high school completion. We then implement four procedures that address selection into adolescent childbearing—Regression Adjustment (RA), Inverse-Probability-Weighted Regression Adjustment (IPWRA), Augmented Inverse-Probability-Weighting (AIPW)⁸, and a matching procedure, Coarsened Exact Matching (CEM). All are suitable for using retrospective data to estimate the causal effect of adolescent childbearing under the assumption of non-randomness in the distribution of women’s age at first birth. We show results using IPWRA and AIPW in the main text, since they provide the most-conservative estimators (lowest adolescent motherhood penalty). Details for the remaining strategies are presented in Appendix A. All results are consistent across procedures and findings remain stable.

We rely on the conceptual framework known as the *counterfactual model* (Rubin 1974; Wooldridge 2002, 2012). We simulate a comparison between the outcome of a woman who was an adolescent mother and the outcome of the same woman if not treated—if she did not have a child during adolescence. Clearly, this situation is, by definition, not-observable (Cerulli 2015). The procedures employed require the assumption of Conditional Mean Independence (CMI), also known as selection on *observables*. The CMI assumption is, plausibly, the most-salient limitation of this approach. That is, we assume we observe the “pre-treatment” covariates driving the non-random assignment into teenage childbearing. When $\mathbf{x}$ represents enough information observable to the analyst and determining the treatment, the CMI assumption states that the potential outcomes for *treated* and *un-treated* groups might be independent of treatment, conditional on $\mathbf{x}$ (Wooldridge 2012). Therefore, once these variables are accounted for, it is possible to estimate Average Treatment Effects on the Population (ATE) and Average Treatment Effects on the Treated (ATET). This strategy also requires assuming that our “pre-treatment” variables—parental education, parental occupation, race, and social class during childhood—are the main determinants of differential selection into early childbearing.

Conditional Independence Assumption:

⁶While the data we use are unique in that it provides information on social advantages during childhood, such information is collected retrospectively. Because this information is retrospective, as opposed to measured during childhood with longitudinal data, it might incur in recall bias. At the same time, our dataset is unique in that it has information rarely available for middle- and low-income countries (and none for Brazil), that is, social class during childhood and parental education and occupation—rarely available for adult respondents through nationally-representative household surveys such as the National Household Survey (PNAD). Parental education and social class during childhood are statistically correlated yet represent slightly different mechanisms in the intergenerational transmission of inequality.

⁷The specific question we used asked respondents their perception on what was the social class of their parents when the respondent was a child (or the family that raised him/her).

⁸We use the package *teffects* (Stata) for implementing RA, IPWRA and AIPW. A command for implementing CEM is also offered in Stata (*cem*).

$$(Y_1; Y_0) \perp D, \mathbf{x} \quad (1)$$

Conditional Mean Independence Assumption (CMI):

$$E(Y_1 | \mathbf{x}, D) = E(Y_1 | \mathbf{x}) \quad (2)$$

$$E(Y_0 | \mathbf{x}, D) = E(Y_0 | \mathbf{x}) \quad (3)$$

Where D represents the treatment (0/1), $\mathbf{x}$ is a row of covariates driving the treatment assignment, Y_1 represents the outcome when the individual is treated and Y_0 is the outcome of the same individual if not treated. Equation 1 states that, conditional on controlling for $\mathbf{x}$, Y_0 and Y_1 are probabilistically independent of D . Equations 2 and 3 represent the restriction of the independence only on the mean (Cerulli 2015).

Our strategies also require the non-violation of the *overlap assumption*, for each observation we should have a positive probability of each treatment level (Rosenbaum and Rubin 1983). Given the high proportion of women who were teen mothers in our sample, we have sufficient observations in each treatment status.

Double-Robust Estimators: IPWRA and AIPW: *Inverse-Probability-Weighted Regression Adjustment (IPWRA)* estimators allow to specify both the probability of treatment (or propensity score) and a model for the outcomes under analyses (years of education and graduation from high school). The treatment model is conducted first, and as a result Inverse Probability Weights to treatment are calculated. These are then used to perform an adjusted regression with the outcome of interest. Thus, this method is a combination of Inverse-Probability-Weighting (IPW) and Regression Adjustment (Wooldridge, 2010). On the other hand, *Augmented Inverse Probability Weighting (AIPW)* also allows for modeling both treatment and outcome of interest. The AIPW estimators use IPW as well, but include an augmentation term in the outcome model as a form of bias correction.

Both IPWRA and AIPW rely on the estimation of treatment assignment in addition to the outcome interest. Importantly, both have a doubly robust property, that is, they are expected to remain consistent even if one of these two models is not correctly specified (Wooldridge 2012; Słoczyński and Wooldridge, forthcoming).

Balance Checks: As noted, both IPWRA and AIPW rely on a specification modeling treatment assignment, both working with a “weighted sample.” It is also expected that this procedure will balance the covariates in our analyses. Results for our logit model predicting treatment status (or propensity score model) are presented in Appendix B. On the overall,

coefficients show the expected direction, with (lower) parental education and race (non-white) being important predictors of adolescent childbearing.

Further, we check for balance after weighting using the overidentification test developed by Imai and Ratkovic (2014).⁹ We also produced an unbalanced diagnosis comparing raw (original) and weighted data in terms of standardized differences and the variance ratio. All tests conducted point out to a substantial improvement of balance after the weighting process.

Sensitivity Analyses: While our data provide retrospective information rarely available for low- and middle-income countries, unmeasured factors may play a role at explaining our results. We conducted *deterministic* and *probabilistic* sensitivity analyses¹⁰ to explore the impact of this potential source of bias (Lash et al. 2009; Orsini et al. 2008), using high school incomplete (0/1) as the outcome of interest.

We conducted a *deterministic* sensitivity analysis to evaluate the effect of unmeasured confounding due to family structure while growing up. Family structure is associated with both educational outcomes and the risk of teenage pregnancy (Kane et al. 2013; Fergusson and Woodward 2000; among others), and has been found to be associated with teenage pregnancy in Brazil (Chagas de Almeida and Aquino 2009).

We obtained information on family structure at age 15 from the Pesquisa Nacional por Amostra de Domicílios (PNAD) conducted in 2014. While the PNAD design is cross-sectional, the 2014 wave added a module on social mobility with retrospective questions, including family characteristics of respondents at age 15. Our measure for family structure is binary where 1 identifies women who lived with either one parent or without a parent at age 15 (details available upon request).

Though imperfect, this strategy allows us to conduct a sensitivity analysis using information from the most widely-used source of socioeconomic microdata for Brazil. As mentioned, data sources including retrospective information are extremely rare in Brazil, and longitudinal datasets are non-existent. We followed the same age restriction with the PNAD 2014 data. By using only information from female household heads or partners ages 19–29, we minimize the number of women with children out of the household. We find that 38 percent of women in our PNAD 2014 sample lived with only one parent or no parent at age 15, and have a 36 percent increase in the risk of not completing high school (odds ratio [OR]=1.72; 95% confidence interval [CI]= 1.51–1.95).

We next implemented *probabilistic* sensitivity analyses using Monte Carlo simulations (20,000 simulations). Probabilistic sensitivity analyses are aimed to capture uncertainty about the specific parameters of a potential unmeasured confounder (Orsini et al. 2008). We assigned two trapezoidal independent distributions of an unmeasured confounder U with minimum value, low mode, upper mode, and maximum values of 0.4, 0.45, 0.55, and 0.6 for women who were teen mothers (exposed) and of 0.25, 0.30, 0.38, and 0.42 for women who

⁹Available in Stata: `tebalance overid`.

¹⁰All sensitivity analyses were conducted using the Stata version 14.1 (Orsini et al. 2008).

were not teen mothers (unexposed). We also specified a trapezoidal distribution for the effect of U on high school incompleteness RR, with critical values of 1, 1.2, 1.8, and 2. Details of these analyses are shown in Appendix C.

Adolescent Motherhood and Adolescent Unions—Our second goal is to evaluate a potential interaction between adolescent childbearing and adolescent marital unions at explaining educational outcomes among Brazilian women. We ran models for each educational outcome of interest including a binary by binary interaction (adolescent mother=0/1, adolescent union=0/1), using our original main analytical sample. We present full results of these models as well as predicted probabilities to illustrate our results.

Results

We start by presenting a summary of our set of pre-treatment variables by *treatment* level (first birth at 18 or younger=1 versus first birth at 18 or younger=0), shown in columns B and C in Table 1. As expected, women who became mothers as teens show a more unfavorable profile for all these covariates. Thus, while 64 percent in our control group is non-white, 72 percent of women who have been teen mothers identified themselves as such. Similarly, whereas only three percent of women who were adolescent mothers had parents with tertiary education, this proportion is nine percent in our control group. This suggests that adolescent mothers come disproportionately from disadvantaged households in Brazil, as in other places (Diaz and Aquino 2006, among others), confirming the need to address the issue of self-selection into early childbearing.

Balance Checks

Columns D–G in Table 1 present results for an unbalanced diagnosis comparing raw (original) and weighted data in terms of standardized differences and the variance ratio. Results show that our model for treatment improved substantially the balance for all pre-treatment covariates—for all cases, the weighted standardized differences are close to zero, while the variance ratio shows values that are close to 1. We also conducted a balance check after weighting using the overidentification test developed by Imai and Ratkovic (2014).¹¹ The resultant chi-square coefficient indicates we cannot reject the null hypothesis that our treatment model achieves balance for the covariates used.¹²

Descriptive Results

Column B in Table 2 shows descriptive statistics for the variables included in our models. Twenty-seven percent of the women in our sample had a first child at age 18 or younger and 25 percent started a marital union at age 18 or younger. A substantial proportion of respondents identify as non-white (66 percent) and a quarter had migrated internally at some point in their life (26 percent). About two-thirds of the women in our sample (65 percent) had parents with elementary schooling or less. About a third of the respondents had a parent with either a rural or unskilled occupation (15 and 18 percent), while respondents whose

¹¹Available in Stata: `tebalance overid`

¹² $\text{Prob} > \chi^2 = 0.8314$

parents had a skilled manual job make up for 35 percent. Only 13 percent report a parent with a professional or managerial type of work. Further, 38 percent considered themselves as poor while growing up.

Table 2 show the mean years of schooling and the proportions completing high school by selected covariates. Respondents who had a first birth at age 18 or younger had on average 8.42 years of schooling, while those who postponed childbearing to age 19 or later (or remained childless) had on average 10.6 years of schooling. Women who started a marital union before age 18 had on average 8.34 years of schooling while their counterparts who were not engaged in an adolescent union had 10.55 years of schooling. Table 2 also shows that disparities in educational attainment follow race and socioeconomic status lines. Non-white respondents had fewer years of schooling and lower proportions completed high school compared to white respondents. Likewise, respondents with elementary-school educated parents had almost 4 fewer years of schooling than respondents with parents with some tertiary education or more (9.24 versus 13.22). Similarly, a markedly lower proportion of daughters of low-educated parents finished high school than daughters of parents with some tertiary education or higher (47 percent versus 97 percent). Respondents with at least one parent in professional/managerial occupations had 12.33 years of schooling compared to only 8.52 among those with a rural worker parent.

Multivariate Results

Effect of Teenage Childbearing on Educational Outcomes—Tables 3 and 4 present a summary of results from OLS and logistic regression models in addition to results from the strategies for estimating treatment effects (counterfactual approach). Table 3 shows results with years of schooling as outcome of interest, while Table 4 reports results for graduation from high school. Summaries include results for the ATE, or the expected effect on a randomly drawn person from the population under analysis.¹³ For both outcomes of interest, we first run “null” models with only adolescent childbearing as predictor of educational attainment. Results for these models are shown in the first row of both tables, and correspond to a baseline for the interpretation of additional approaches. We obtain a “baseline” penalty of 2.28 fewer years of education for adolescent mothers, as well as a 0.42 reduction in the predicted probabilities of graduating from high school.

The second column for both tables indicates the estimated difference in educational attainment across the two groups, adolescent mothers versus not, or ATE. Thus, in Table 3, our OLS model for years of education after including socioeconomic controls, predicts 1.79 fewer years of schooling for women who were adolescent mothers versus those who did not have a child at early age, holding all other variables constant.

In Table 4, for ease of interpretation, we include the predicted probabilities of high school completion at each treatment level, calculated after implementing logit models. Similarly, for the IPWRA and AIPW strategies, this column presents the Potential Outcome Means (POMs), or predicted mean of graduation at each specific treatment level. Once again, we use the group of women who were not teen mothers as reference.¹⁴

¹³We include results for the ATET—those who were adolescent mothers—in Appendix F.

After including socioeconomic controls, our logit model indicates a disadvantage of 0.41 in the predicted probabilities of graduation from high school among adolescent mothers.¹⁵ Accordingly, whereas teenage mothers have a predicted probability of 0.30 of completing high school, those who either postponed motherhood or remained childless have a predicted probability of 0.71 (holding all other covariates constant). The full set of results for the OLS and logistic models are presented in Appendix D.

Our findings show that the negative coefficients representing adolescent childbearing we found in the OLS and logistic regressions remain large and significant for both outcomes in the models accounting for selection. In Table 3, Augmented IPW (AIPW) applied to years of education yields −1.69 fewer years of education attributable to adolescent childbearing, reduced from −2.28 in our baseline model, while the IPWRA results in a penalty of −1.66 years of education.

Further, Table 4 shows that both Augmented IPW (AIPW) and IPWRA estimates indicate an average reduction of 0.35 in the probability of completing high school due to early childbearing, reduced from 0.42 in our baseline model. Thus, according to estimates produced by an AIPW procedure, adolescent mothers have a potential mean of high school graduation of 0.316, in contrast to 0.664 for women who became mothers later in life (or remained childless). Other procedures yield similar results, with reductions in the potential means of high school graduation between the two groups ranging from 0.35 to 0.38 (Appendix E).

Across the strategies we implement, based on selection observables, we find a stable educational penalty associated to adolescent childbearing. While the coefficients are slightly reduced if compared with coefficients from OLS and logit models, penalties after accounting for selection persist as large and significant. Combined, these results suggest that differential selection into adolescent childbearing does not explain the lower educational outcomes of adolescent mothers vis-à-vis older mothers (or childless women).

Results for our ATET estimations are shown in Appendix F. Recent work has pointed out to the importance of plausible heterogeneity in the effects of teenage childbearing across different groups of women (Diaz and Fiel 2016). Yet, on the overall, our analyses do not show salient differences in the effects for the average teenager versus a typical woman who started childbearing at early age (for instance, IPWRA estimates a penalty of −1.66 years of schooling as ATE and −1.81 for ATET). In fact, this picture is consistent with those reported by Diaz and Fiel for the US (Diaz and Fiel), who also find homogeneous negative effects in high school completion.

Sensitivity Analysis—The original (observed) Odds Ratio between teenage childbearing and high school not-complete was 5.68, and further adjustment for other covariates did not substantially change the association. Using PNAD 2014, our *deterministic* sensitivity

¹⁴Complete results for all models are available upon request.

¹⁵Full set of results for OLS and Logistic Models (with coefficients for all covariates) is included in Appendix D.

analysis estimated that failure to adjust for family structure only biased our results by approximately five percent, with an externally adjusted OR 5.4.

Moreover, our *probabilistic* sensitivity analysis yielded a median bias-adjusted OR of 5.31, with 2.5th and 97.5th percentiles of 3.88 and 7.33 (a reduction of about seven percent). Taken as a whole, our sensitivity analyses suggest that unmeasured confounding is not a major factor affecting our overall findings.

Adolescent Childbearing, Marital Unions and Educational Outcomes—Finally, Table 5 shows results for models predicting years of schooling and high school graduation with our full set of covariates and an interaction term between adolescent childbearing and adolescent marital union. The interaction term is significant in the models for years of schooling ($b=0.779$, $p<0.05$), although it fails to statistical significance for high school graduation as an outcome.

While the main effect of early union is clearly negative for both outcomes, the positive sign of both interaction terms makes the interpretation of results less straightforward. Thus, to illustrate the effects, we calculated predicted probabilities at the means of covariates for all groups by teenage childbearing and teenage union status, separately for each outcome of interest—years of education and high school graduation. We present these results in Figure 1 and Figure 2. Figure 1, using years of education as outcome of interest, shows a gap of 0.8 fewer years of schooling for the group of women who were adolescent mothers and also engaged in a marital union during adolescence versus those mothers who remained single until age 19 or later. Where our interaction term for high school graduation remains not-significant, predicted margins obtained from this equation point out to a similar pattern of disadvantage affecting women who were adolescent mothers and also started unions early in life (Figure 2). Although our data does not allow us to disentangle whether a union preceeded childbearing in some cases (age), our findings suggest that early marital unions do not mitigate but rather intensify the penalty in years of schooling associated with teenage childbearing.

Discussion and Conclusions

The educational consequences of adolescent childbearing have long been of interest to social scientists. A recurring limitation of this body of research is the lack of appropriate data to examine the causal effects of adolescent childbearing in low- and middle-income countries while at the same time addressing the possibility that early social disadvantages might drive both educational penalties and early childbearing. In this article, we use a unique data source for Brazil, with information suitable for assessing the educational impact of adolescent childbearing using retrospective information. The article further addresses the role of adolescent marital unions, particularly common in Brazil, in explaining educational disadvantages among adolescent mothers.

Our study makes two main contributions. First, we find a large negative effect of adolescent childbearing in two key educational outcomes even after addressing differential selection into early motherhood. In models taking selection into account, the penalty for adolescent

mothers ranges from -1.66 to -1.80 fewer years of schooling and about 35 percent difference in the probabilities of graduating from high school. Our estimates indicate a larger penalty in Brazil when compared with industrialized contexts (for US, Kane et al. 2013). The evaluation of these results should consider that adolescent mothers account for about 28 percent of women in our data, by far a higher proportion than in most industrialized societies. Combined, a high educational penalty and a high prevalence of adolescent childbearing suggests that adolescent mothers face unparalleled educational disadvantages in this middle-income country and, consequently, further social stratification. These findings also suggest that the role of the extended family in buffering the negative effects of adolescent childbearing has been overstated in these contexts.

Further, our findings suggest that a marital union is associated with an additional layer of disadvantage for adolescent mothers. The role of adolescent marital unions for women's educational outcomes has been largely overlooked in the literature addressing adolescent childbearing. Our findings point to a negative effect of a partner early in life on women's accumulation of human capital in Brazil, suggesting that a partner's presence reinforces traditional gender roles. Our findings highlight the importance of exploring the context of family formation more broadly when analyzing the educational outcomes of adolescent mothers. Within this framework, early family formation impinges on women's education, therefore furthering social stratification in Brazil.

The current study presents several avenues for future research. One particularly important question refers to the mechanisms driving the additional disadvantage of early marital unions for adolescent mothers. Further research should shed light into understanding whether and how family structure and marital unions contribute to further adolescent mothers' educational disadvantages.

It is important to note that our dataset, while providing unique information key to our causal analyses, has limitations. Our total sample size is comparatively small, and therefore we are unable to explore differences across subgroups of adolescent mothers (by race for example). A small sample could also lower the quality of matching procedures; yet, using the CEM strategy, we were able to match the majority of our sample. Furthermore, while uncommon for developing countries, the set of pre-adulthood variables available in our data is certainly small. Information on women's family structure while growing up, or age at sexual debut would have been particularly useful, given their relevance in the literature (Kane et al. 2013, among others). Circumstances such as family size and number of children may also influence both adolescent fertility and educational outcomes, through the mechanism of fewer parental resources allocated for each child. Similarly, we lack measures of academic ability, such as test scores (Lou and Thomas 2015; Shearer et al. 2002, among others), as well as on non-cognitive abilities (Heckman et. al. 2006). Both of these factors have been found to be associated with teenage childbearing and educational achievement.

At the same time, the measures on social class we use are key characteristics shaping individuals' trajectories. Further, scholars have consistently pointed out that class markers during childhood have larger consequences in Latin American countries versus others with similar levels of development, given the persistent high rates of inequality in the region

(Torche 2014), suggesting that the factors we are able to measure may be even more consequential for Brazil than for other contexts.

While contributing to fill a gap in the literature on the causal effects of fertility and education, our analyses rely on the Conditional Mean Assumption. That is, we cannot completely rule out the endogeneity of early childbearing and educational outcomes and unobserved factors may still affect women's selection into adolescent childbearing. Our sensitivity analysis suggests that our results still hold even when we explore the potential impact of unmeasured confounders, though their role cannot be completely ruled out.

This study advances the literature on the consequences of adolescent childbearing and union formation on women's education in a middle-income country with a distinctive pattern of early fertility, below-replacement birth rates, and a high prevalence of marital unions among adolescent mothers. In fact, the proportion of female adolescents who are either mothers or pregnant remains high in Brazil, even if compared to other Latin American countries. Furthermore, while Brazil made considerable progress towards educational expansion, the country remains highly unequal (Torche and Ribeiro 2010), with educational opportunity playing a fundamental role as a mechanism for social stratification. Brazil, and Latin America as a whole, is characterized by a polarized distribution of educational attainment with a historically large premium for those with secondary and postsecondary schooling (Torche 2014). The outcomes we examine—years of education and graduation from high school—represent key assets in the lives of contemporary Brazilians (Marteleteo et al. 2016). Particularly for those from disadvantaged backgrounds, the culmination of high school is the first step in pushing through the bottleneck of social mobility in a highly unequal society. That young mothers, particularly those in a marital union, face additional layers of educational disadvantages demonstrates the importance of early family formation as a *stratifier* in an already highly-stratified society.

Additional Materials

Appendix A Other Procedures using a Counterfactual Model

Regression Adjustment (RA) is usually described as a general approach for estimating ATE, based in a logic of imputation. The procedure includes two steps. First, two separate regression models are fit of the outcome of interest on the selected covariates, for each treatment status (teen mother=1, teen mother=0). Each observation in the control group is “imputed” with its predicted outcome if they had received the treatment, and vice versa. Once all observations are imputed, the difference between values under the two conditions is calculated. The average of these differences over treated subjects produces the ATE (Cerulli 2015).

Coarsened Exact Matching (CEM). With the aim of applying a matching strategy to our study, we opt for implementing a CEM—Coarsened Exact Matching procedure (Iacus, King and Porro 2012). A number of contributions have compared different matching methods available (including those based on propensity scores) and concluded that CEM offers better balance properties (Blackwell et al 2009). As stated, the goal of this exercise is to compare

women with a similar probability of having had a birth in their teenage years, and therefore to remove endogeneity. Given our *pre-treatment* variables, we match individuals by race (two groups), parental educational attainment (three groups), parental occupation (six groups) and social class during childhood (four groups). The CEM procedure implies that the researcher “coarsens” each variable selected for the matching into meaningful categories (i.e.: age into five categories); distinctly assigning each observation into a strata in which all observations have identical values on the “coarsened” pre-treatment variables. Then, a weight is assigned to all observations, corresponding to their block. Observations in any stratum that do not have at least one observation in each treatment and control groups are discarded (Blackwell, Iacus, King, and Porro, 2009). We were able to conduct *exact* matching, meaning that we did not coarsen the variables used for the matching; our treated and control groups have values identical to the original pre-treatment covariates selected. We next re-estimate our equations with the balanced sample.¹⁶ We were able to match above 90 percent of observations in our analytical sample. Our statistic for unbalance is reduced to values close to zero after the CEM procedure is implemented.¹⁷ The table below presents additional details of the balancing procedure. We further conducted a *one-to-one* CEM matching using the same number of treated and control observations within each stratum. Using this approach yields to substantially similar results and we report them alongside with results from other strategies.

Coarsened Exact Matching (CEM) Matching Summary

	Original	After CEM
Unbalanced	0.3150	2.913e-16
Sample Size	1,080	1,012
Un-matched percent: 7.3		

Appendix B Logit model for predicting treatment status (propensity score model)

Non-white	0.289 ⁺ (0.156)
Parental education. Ref.: Elementary & lower high school	-0.543 ^{**} (0.187)
Some tertiary and higher	-0.996 [*] (0.442)
Parental occupation. Ref.: Rural workers Manual, unskilled workers	-0.0329

¹⁶With CEM (as with IPWRA and AIPW), the obtained weights function similarly to sample weights in subsequent regression models.

¹⁷The Blackwell et al. package for CEM implementation allows to measure the sample imbalance in the original data, as a baseline, in order to compare it with the resulting imbalance after the matching is conducted.

	(0.248)
Manual, skilled	0.288
	(0.214)
Supervisors of manual work	−0.0909
	(0.301)
Routine, non-manual workers	−0.110
	(0.342)
Professionals and managers	−0.473
	(0.376)
Social class during childhood. Ref.: Poor	
Low-middle	−0.138
	(0.163)
Middle	−0.308
	(0.209)
Upper-middle and higher	0.509
	(0.375)
Constant	−0.896***
	(0.223)
N	1,080

Standard errors in parentheses.

+ p<0.1

* p<0.05

** p<0.01.

Imputed Sample.

SOURCE: Own elaboration based on 2013 Brazil School-to-Work Transition Survey (ILO).

Appendix C Bias analysis for the potential impact of unmeasured confounding

Panel A: Deterministic Bias Analysis

	Association btw 1 parent/no parent and high school incomplete (RR)	Prevalence of 1 parent/no parent among females with 1st birth<=18 years old	Prevalence of 1 parent/no parent among females with 1st birth>18 yrs old/childless	External Adjusted Odds Ratio	Percent Bias
Data from PNAD 2014	1.36	0.5	0.34	5.4	5%
Trying Different Scenarios (not with real data)	Association between Confounder U and High School Incomplete (RR)	Prevalence of Confounder U among females with 1st birth<=18 years old	Prevalence of Confounder U among females with 1st birth>18 years old/childless	External Adjusted Odds Ratio	
Scenario 1	1.46	0.5	0.25	5.15	10%

Scenario 2	1.56	0.55	0.25	4.95	15%
Scenario 3	1.66	0.45	0.17	4.87	17%
Scenario 4	1.86	0.4	0.15	4.77	19%

Panel B: Probabilistic Bias Analysis

	Minimum Value	Mode: Lower Value	Mode: Upper Value	Maximum Value	Median U-adjusted teen pregnancy-high school incomplete OR
Prevalence of confounder U if teen birth=1	40	45	55	60	
Prevalence of confounder U if teen birth=0	25	30	38	42	
Association between confounder U and High School Incomplete (RR)	1	1.2	1.8	2	5.31

NOTE: Conventional (unadjusted) association: OR=5.68 [C.I.: 4.16–7.74].

Appendix D Full Set of Results. OLS and Logistic Models predicting Years of Education and Graduation from High School. Only Women 19 and older, Brazil 2013

	Years of Education	Graduation from HS
1st Birth at 18 and Younger	−1.79*** (0.163)	−1.751*** (0.170)
Non-white	−0.312 ⁺ (0.166)	−0.137 (0.167)
Parental education. Ref.: Elementary and lower high school	0.967*** (0.182)	0.858*** (0.181)
Some tertiary or higher	2.422*** (0.357)	2.511*** (0.679)
Parental occupation. Ref.: Rural workers		
Manual, unskilled workers	0.730** (0.269)	0.559* (0.258)
Manual, skilled	0.868*** (0.235)	0.687** (0.233)
Supervisors of manual work	0.957** (0.305)	0.742* (0.318)
Routine, non-manual workers	0.944** (0.351)	0.624 ⁺ (0.369)

	Years of Education	Graduation from HS
Professionals and managers	1.729*** (0.346)	1.306*** (0.366)
Social class during childhood. Ref.: Poor		
Low-middle	0.141 (0.171)	0.146 (0.169)
Middle	0.739*** (0.208)	0.505* (0.215)
Upper-middle or higher	0.504 (0.392)	0.0690 (0.468)
Migrated	−0.123 (0.166)	0.00650 (0.168)
Urban	0.412 ⁺ (0.224)	0.273 (0.211)
Constant	8.655*** (0.676)	−1.491* (0.672)
N	1,080	1,080

NOTES: Standard errors in parentheses.

⁺
p<0.1

*
p<0.05

**
p<0.01.

SOURCE: Own elaboration based on 2013 Brazil School-to-Work Transition Survey (ILO). Models include age and region. Imputed Sample.

Appendix E Complete Summary of Analysis

Panel A. Summary of Analysis predicting Years of Education

Method	Difference/ATE (teen mother 1 vs. teen mother 0)	S.E.	p	[95% Confidence Interval]		N
				<i>Lower Bound</i>	<i>Upper Bound</i>	
OLS, null model	−2.28	0.18	0.00	−2.63	−1.93	1,080
OLS, all covariates	−1.79	0.16	0.00	−2.11	−1.47	1,080
Regression Adjustment	−1.70	0.18	0.00	−2.04	−1.35	1,080
Inverse-Probability-Weighted Regression Adjustment (IPWRA)	−1.66	0.17	0.00	−2.01	−1.32	1,080
Augmented Inverse-Probability Weighting (AIPW)	−1.69	0.18	0.00	−2.04	−1.34	1,080
Matching: CEM	−1.80	0.17	0.00	−2.14	−1.47	1,080
CEM, one-to-one	−1.71	0.20	0.00	−2.10	−1.33	1,080

SOURCE: Own elaboration based on the School-to-Work Transition Survey (ILO, 2013). Imputed Sample.

Panel B. Summary of Analysis Predicting Graduation from High School.

Method	Difference/ATE (teen mother 1 vs. teen mother 0)	Predicted Prob. & POMs	S.E.	Significance	[95% Confidence Interval]		N
					<i>Lower Bound</i>	<i>Upper bound</i>	
Logit, null model							
Teen mother	−0.42	0.27	0.03	0.00	0.22	0.32	1,080
Non-teen Mother		0.69	0.02	0.00	0.65	0.72	
Logit, all covariates							
Teen mother	−0.41	0.30	0.03	0.00	0.24	0.36	1,080
Non-teen Mother		0.71	0.02	0.00	0.67	0.75	
Regression adjustment							
Teen mother	−0.35	0.32	0.03	0.00	0.27	0.37	1,080
Non-teen Mother		0.67	0.02	0.00	0.64	0.70	
Inverse-Probability- Weighted Regression Adjustment (IPWRA)							
Teen mother	−0.35	0.32	0.03	0.00	0.27	0.37	1,080
Non-teen Mother		0.67	0.02	0.00	0.64	0.70	
Augmented Inverse- Probability Weightning (AIPW)							
Teen mother	−0.35	0.32	0.03	0.00	0.26	0.37	1,080
Non-teen Mother		0.66	0.02	0.00	0.63	0.70	
Matching: CEM							
Teen mother	−0.36	0.26	0.03	0.00	0.21	0.31	1,080
Non-teen Mother		0.63	0.02	0.00	0.58	0.67	
CEM, one-to-one							
Teen mother	−0.38	0.26	0.03	0.00	0.21	0.31	1,080
Non-teen Mother		0.64	0.03	0.00	0.59	0.69	

SOURCE: Own elaboration based on the School-to-Work Transition Survey (ILO, 2013). Imputed Sample.

Appendix F Summary of analyses predicting years of education and high school completion. Average Treatment Effect on the Treated (ATET)

A. Years of Education and Adolescent Childbearing							
Method	Difference	S.E.	<i>p</i>	[95% Conf. Interval]		N	
Regression Adjustment	−1.82	0.18	0.00	−2.17	−1.47	1,080	
IPWRA	−1.81	0.18	0.00	−2.16	−1.46	1,080	
B. Graduation from High School and Adolescent Childbearing							
Method	Difference	Predicted Probabilities	S.E.	<i>p</i>	[95% Conf. Interval]		N
Regression Adjustment	−0.36		0.03	0.00	−0.42	−0.30	1,080

Teen mother	0.27						
Non-Teen Mother	0.63						
IPWRA	-0.36	0.03	0.00	-0.42	-0.30	1,080	
Teen mother	0.27						
Non-Teen Mother	0.63						

SOURCE: Own elaboration based on the School-to-Work Transition Survey (ILO, 2013). Imputed data. Sampling weights applied.

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FIGURE 1—.
Predicted margins for years of education by early childbearing and early union status (4 combinations), Brazil 2013

FIGURE 2.

Predicted probabilities of graduation from high school by early childbearing and early union status (4 combinations), Brazil 2013

TABLE 1

Balance information

	<i>Column B</i>	<i>Column C</i>	<i>Column D</i>	<i>Column E</i>	<i>Column F</i>	<i>Column G</i>
	<u>Means</u>		<u>Unbalance Diagnosis: Raw and Weighted Data</u>			
			<u>Standardized Differences</u>		<u>Variance Ratio</u>	
Covariate	Teen Birth=0	Teen Birth=1	Raw	Weighted	Raw	Weighted
Non-white	0.64	0.72	0.172	0.014	0.878	0.990
Parental Education						
Elementary or less	0.60	0.77				
High school	0.31	0.19	−0.275	−0.014	0.718	0.985
Tertiary	0.09	0.03	−0.227	0.034	0.430	1.109
Parental Occupation						
Rural worker	0.14	0.16				
Manual, low-skilled	0.17	0.19	0.025	−0.013	1.045	0.977
Manual, skilled workers	0.33	0.43	0.211	−0.022	1.115	0.986
Supervisors of manual work	0.11	0.09	−0.053	0.021	0.873	1.054
Routine, non-manual work	0.09	0.07	−0.095	−0.008	0.740	0.977
Professional, managerial	0.16	0.07	−0.287	0.031	0.467	1.069
Social Class during Childhood						
Poor	0.36	0.44				
Low-middle	0.37	0.37	−0.010	0.002	0.997	1.001
Middle	0.22	0.15	−0.208	0.028	0.711	1.041
Upper-middle and higher	0.04	0.05	0.042	−0.014	1.204	0.937
	711.00	263.00				
<u>Detail for Number of Observations</u>						
	Raw	Weighted				
Treated	263	486.30				
Control	711	487.70				
Total	974	974.00				

SOURCE: Own elaboration based on 2013 Brazil School-to-Work Transition Survey (ILO). Non-imputed sample.

TABLE 2

Mean years of schooling and proportion graduating from high school by selected covariates. Females, ages 19–29. Brazil, 2013

	<i>Column B</i>	<i>Column C</i>	<i>Column D</i>	<i>Column E</i>
	<u>Sample</u>	<u>Years Schooling.</u>		<u>Graduated from H.S.</u>
	Means and Proportions	Mean	S.D.	Proportions
Years of Education	10.00 (2.77)			
Graduated from High School	0.58			
Age	24.05 (3.09)			
Teenage Childbearing				
1st Child at 18/Younger	0.27	8.42	2.75	0.29
1st Child at 19/Older/Childless	0.73	10.59	2.54	0.70
Teenage Union				
1st Union at 18/Younger	0.25	8.34	2.72	0.30
1st Union at 19/Older/Never in a union	0.75	10.55	2.56	0.68
Race				
Non-white	0.66	9.71	2.68	0.56
White	0.34	10.56	2.85	0.64
Parental Education				
Elementary or less	0.65	9.24	2.65	0.47
High school	0.28	10.88	2.26	0.74
Tertiary	0.08	13.22	2.12	0.97
Parental Occupation				
Rural worker	0.15	8.52	3.08	0.37
Manual, low-skilled	0.18	9.60	2.27	0.53
Manual, skilled workers	0.35	9.70	2.47	0.55
Supervisors of manual work	0.11	10.27	2.56	0.65
Routine, non-manual work	0.08	10.61	2.22	0.67
Professional, managerial	0.13	12.33	2.75	0.88
Social Class during Childhood				
Poor	0.38	9.28	2.72	0.49
Low-middle	0.37	9.93	2.67	0.59
Middle	0.20	11.11	2.45	0.73
Upper-middle and higher	0.04	11.65	3.20	0.70
Migration				
Migrated	0.26	9.89	2.92	0.57
Non-migrated	0.74	10.04	2.71	0.59
Are of Residence				
Rural	0.13	8.89	3.30	0.43
Urban	0.87	10.16	2.64	0.61

	<i>Column B</i>	<i>Column C</i>	<i>Column D</i>	<i>Column E</i>
	<u>Sample</u>	<u>Years Schooling.</u>		<u>Graduated from H.S.</u>
	Means and Proportions	Mean	S.D.	Proportions
Region				
North	0.09	9.47	2.76	0.54
Mid-West	0.07	11.00	2.58	0.71
North-East	0.30	9.49	2.91	0.54
South	0.13	9.58	2.69	0.46
South-East	0.41	10.45	2.60	0.64
<i>N</i>	974			

SOURCE: Own elaboration based on 2013 Brazil School-to-Work Transition Survey (ILO). Non-imputed sample.

TABLE 3

Summary of analysis predicting years of education

Method	Difference/ATE (teen mother 1 vs.0)	S.E.	p	[95% Confidence Interval]		N
				<i>Lower Bound</i>	<i>Upper Bound</i>	
OLS, null model	−2.28	0.18	0.00	−2.63	−1.93	1,080
OLS, all covariates	−1.79	0.16	0.00	−2.11	−1.47	1,080
Inverse-Probability-Weighted Regression Adjustment (IPWRA)	−1.66	0.17	0.00	−2.01	−1.32	1,080
Augmented Inverse-Probability Weighting (AIPW)	−1.69	0.18	0.00	−2.04	−1.34	1,080

SOURCE: Own elaboration based on the School-to-Work Transition Survey (ILO, 2013). Imputed Sample.

TABLE 4

Summary of analysis predicting graduation from high school

Method	Difference/ATE	Predicted Prob. & POMs	S.E.	Significance	[95% Confidence Interval]		N
	(teen mother 1 vs.0)				Lower Bound	Upper bound	
Logit null model							
Teen mother	−0.42	0.27	0.03	0.00	0.22	0.32	1,080
Non-teen Mother		0.69	0.02	0.00	0.65	0.72	
Logit, all covariates							
Teen mother	−0.41	0.30	0.03	0.00	0.24	0.36	1,080
Non-teen Mother		0.71	0.02	0.00	0.67	0.75	
Inverse-Probability-Weighted Regression Adjustment (IPWRA)							
Teen mother	−0.35	0.32	0.03	0.00	0.27	0.37	1,080
Non-teen Mother		0.67	0.02	0.00	0.64	0.70	
Augmented Inverse-Probability Weightning (AIPW)							
Teen mother	−0.35	0.32	0.03	0.00	0.26	0.37	1,080
Non-teen Mother		0.66	0.02	0.00	0.63	0.70	

SOURCE: Own elaboration based on the School-to-Work Transition Survey (ILO, 2013). Imputed Sample.

TABLE 5

Results from OLS and logit models including an interaction between early childbearing and early union—
Brazil, 2013

	Years of Education	Graduated from H.S.
First child at 18/ younger	-1.603*** (0.222)	-1.602*** (0.228)
First union at 18/ younger	-1.586*** (0.240)	-1.242*** (0.240)
First child at 18/younger * First union at 18/younger	0.779 * (0.357)	0.547 (0.378)
Non-white	-0.330 * (0.162)	-0.177 (0.172)
Parental education. Ref.: Elementary and lower high school	0.834*** (0.179)	0.764*** (0.186)
Some tertiary & higher	2.248*** (0.350)	2.356*** (0.657)
Parental occupation. Ref.: Rural workers		
Manual, unskilled workers	0.593 * (0.262)	0.474 + (0.269)
Manual, skilled	0.693 ** (0.232)	0.570 * (0.244)
Supervisors of manual work	0.880 ** (0.300)	0.694 * (0.327)
Routine, non-manual workers	0.781 * (0.345)	0.513 (0.379)
Professionals and managers	1.554*** (0.342)	1.228 ** (0.375)
Social class during childhood. Ref.: Poor		
Low-middle	0.158 (0.167)	0.166 (0.173)
Middle	0.784*** (0.204)	0.576 ** (0.219)
Upper-middle and higher	0.541 (0.384)	0.140 (0.489)
Migrated	-0.0748 (0.163)	0.0472 (0.171)
Urban	0.402 + (0.220)	0.283 (0.213)
Constant	8.981***	-1.286 +

	Years of Education	Graduated from H.S.
	(0.662)	(0.702)
N	1,080	1,080

NOTES: Standard errors in parentheses.

⁺
p<0.1

^{*}
p<0.05

^{**}
p<0.01.

Imputed Sample.

SOURCE: Own elaboration based on 2013 Brazil School-to-Work Transition Survey (ILO). Models include age and region.