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RESEARCH PAPER

The effects of teenage childbearing on education, physical health, and mental distress: evidence from Mexico

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Abstract
This paper estimates the effects of teenage childbearing on education, working, physical and mental health, and physical activity of young girls in Mexico using two waves of the nationally representative Mexican Family Life Survey. We employ a propensity score matching model that accounts for a rich set of baseline covariates that predict teenage childbearing to attempt to reduce the bias due to confounding variables associated with teenage childbearing. The results demonstrate that teenage childbearing is associated with an increase in the probability of being overweight, and reductions in physical activity and the probability of high school completion. Moreover, the results are consistent when we employ sibling fixed effects to account for unobservable family background.

Key words: bounding; early pregnancy; health; Mexico; propensity score

JEL classification: I1; J13

1. Introduction

Teenage childbearing is of great policy concern, especially in developing countries where high adolescent childbearing rates persist.[1] Teenage childbearing is commonly associated with adverse socioeconomic outcomes for both mothers and their children and thus received considerable academic and policy interest. However, teenage childbearing might be a marker of disadvantage, rather than a cause of adverse outcomes. In order to design an effective poverty reduction policy, it is important to understand whether teenage childbearing is a cause of poor outcomes or a consequence of socioeconomic disadvantages. Moreover, while most research has focused on examining the education and labor market consequences of teenage childbearing, less is known about health consequences. This paper estimates the effects of teenage childbearing on the education and health of young girls in Mexico.

[1]Africa has the highest adolescent fertility rates in the world, followed by Latin America and the Caribbean [United Nations (2013)].

Estimating the impact of teenage childbearing is challenging because fertility decisions are not random. Unobserved heterogeneity between early and later child-bearers might bias the results as young mothers might be a select group that would experience adverse outcomes even in the absence of a child. To attempt to account for selection bias, recent rigorous studies, predominately in the United States (US), have used within-family estimations using data on pairs of sisters and twins [Geronimus and Korenman (1992), Holmlund (2005), Webbink et al. (2008), Gunes (2016)], instrumental variables using twin births and miscarriage as instruments [Bronars and Grogger (1994), Klepinger *et al.* (1999), Hotz *et al.* (2005), Ashcraft and Lang (2006), Fletcher and Wolfe (2009), Fletcher (2012), Ashcraft *et al.* (2013)], and propensity score matching [Chevalier and Viitanen (2003), Levine and Painter (2003), Kane *et al.* (2013)]. While the results generally suggest a negative relationship between teenage childbearing and socioeconomic outcomes, credible causal research for the US points to small effects [for a survey of the literature, see Kearney and Levine (2012)].

There is a lack of rigorous empirical evidence on the adverse consequences of teenage childbearing in the context of developing countries.[2] Moreover, while most rigorous research so far in both developed and developing countries has focused on examining education and labor market outcomes, the evidence on how the health and health behaviors of young mothers are affected is scarce. In the US, Gunes (2016) finds no effect of teenage childbearing on health outcomes and modest effects on preventive health behaviors, while Fletcher (2012) finds no effects on risky health behaviors. In Australia, Webbink *et al.* (2008) find that teenage childbearing negatively affects health behaviors. In South Africa, Ardington *et al.* (2015) find that teenage childbearing is associated with a higher mortality risk before the age of 30, largely due to AIDS-related deaths.

This paper adds to the few rigorous studies exploring the consequences of teenage childbearing in the context of developing countries. We estimate the effects of teenage childbearing on the education and health outcomes and health behaviors of young women, using data from the rich nationally representative longitudinal Mexican Family Life Survey conducted in 2005–2006 and 2009–2012. Specifically, using a propensity score matching approach, we explore the effects of teenage childbearing (first birth at age 18 or earlier) on high school completion, working, being overweight and anemic, self-reported health status, mental health, and physical activity. To this end, this paper employs a propensity score matching approach and estimates inverse probability weighted regression models. The results suggest that teenage childbearing significantly reduces the probability of high school completion, increases the probability of being overweight and anemic, and reduces physical activity.

The rich dataset permits employing propensity score matching, among other empirical approaches, in order to attempt to overcome endogeneity concerns associated with selection bias. In particular, the dataset contains a rich set of pre-treatment controls which can be employed as baseline characteristics in calculating inverse propensity score weights. Using various balancing tests, we find

[2]For example, in the context of developing countries, Ranchhod *et al.* (2011) and Ardington *et al.* (2015) show that teenage childbearing reduces education in South Africa using propensity score matching approach. Branson and Byker (2018) study a reproductive health intervention in South Africa and find that delaying teen childbearing increases schooling.

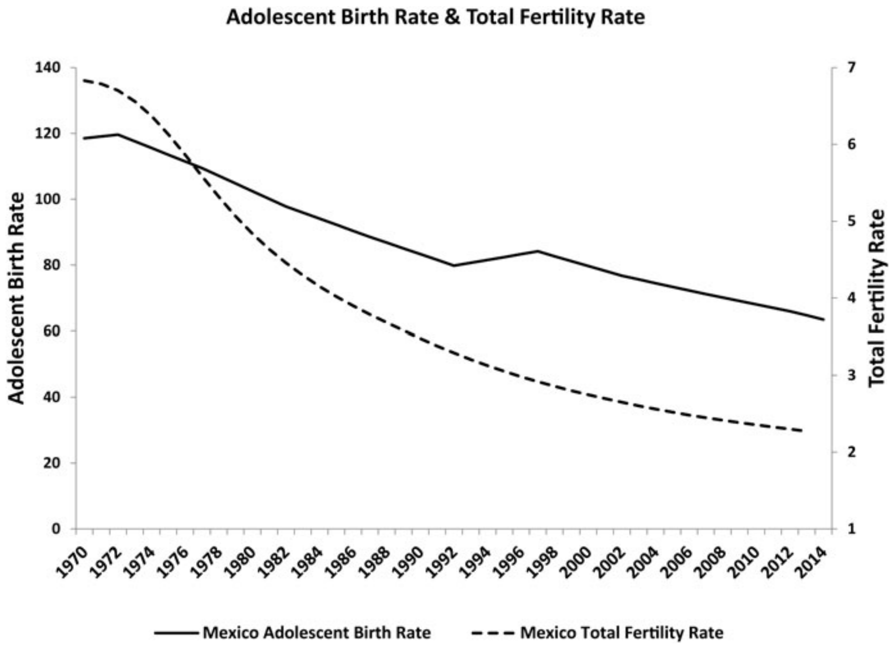

Figure 1. Adolescent birth rate and total fertility rate.

no significant differences in the observable characteristics of treated (teenage mothers) and control groups after using inverse propensity score weights. Second, we perform robustness checks using alternative matching techniques, including radius matching and entropy balance. Finally, we estimate the effects of teenage childbearing using an entirely different approach—employing sibling fixed-effects models. The results are consistent using the alternative sibling fixed-effects approach. Furthermore, we check the sensitivity of these results to various sample restrictions and alternative measures of teenage childbearing.

In Mexico, teenage childbearing remains high despite substantial declines in overall fertility rates (Figure 1). While high levels of teenage childbearing contribute to high levels of total fertility in most developing countries, such as in the majority of sub-Saharan African countries, there is not a strong link between the two rates in Mexico. In Mexico, the demographic transition has been mostly driven by the declines in fertility among older women. This paper therefore explores the effects of giving first birth as an adolescent in a context of high teenage childbearing rates where policies that reduced overall fertility have not been successful in reducing teenage childbearing.

The remainder of this paper is organized as follows: section 2 provides background on fertility trends in Mexico, section 3 presents a conceptual framework, section 4 describes the data and sample construction, section 5 presents the empirical methodology, section 6 discusses the results, and section 7 concludes.

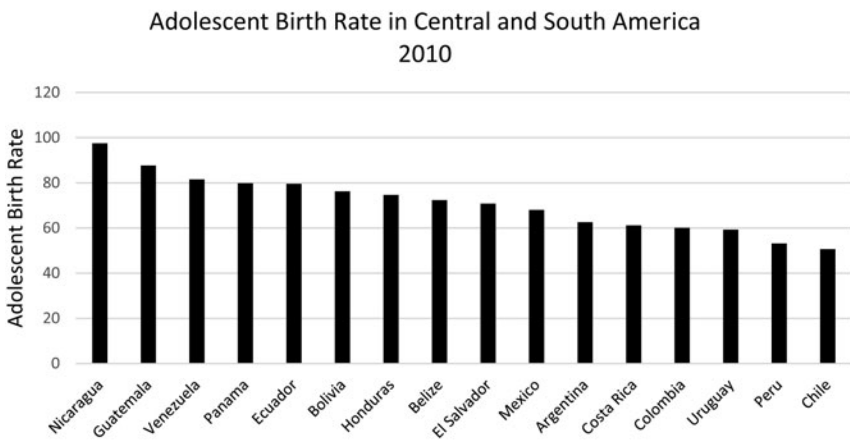

Figure 2. Comparison of adolescent birth rates in 2010.
Source: World Development Indicators, http://data.worldbank.org/data-catalog/world-development-indicators

2. Context

In many countries in Latin America and the Caribbean, including Mexico, fertility has decreased substantially over the past decades. In the 1970s, Mexico launched a policy to reduce population growth. The policy had a strong rural focus as few rural residents had access to contraceptives and health care. New clinics and hospitals were built, more supplies were provided, and health workers were sent out to recruit women to use contraceptives [Potter (1999)]. As a result, the average number of children per woman (i.e., total fertility rate) in Mexico dropped from 6.83 in 1970 to 2.52 in 2005 (Figure 1). Adolescent birth rate also decreased from about 119 to 74 over the same period. Yet, teenage childbearing remains quite high. In 2010, adolescent birth rate in Mexico (68.1 births per 1,000 women aged 15–19) was lower than most Central American countries but higher than many South American countries (Figure 2). Also, it was significantly higher than the average rate in OECD countries (27.1 births per 1,000 women) and the US (33.9 births per 1,000 women) [WDI (2017)].

One explanation for the high adolescent childbearing rates despite substantial reductions in total fertility rates could be the strong emphasis of Mexican population policies on postpartum long-run contraception, such as intrauterine devices (IUDs) and sterilization. Specifically, while the rates of sterilization doubled between 1981 and 1995 (almost 40% of all women using a family planning method were sterilized) and the share of women using IUD increased from 8.8% to 23.5%, the share of women using hormonal methods (the pill), which are more prevalent among younger women, decreased from 35.1% to 10.9% [Potter (1999)]. In 2009, while 67% of married women aged 15–49 reported using a modern method of contraception, only 53% of married young women between the ages of 15 and 24 reported using one [Juarez *et al.* (2013)]. Sixty-six percent of never-married young women who were sexually active in the past month used a contraceptive method, while only 17% of young women who were sexually experienced but did not have sex in the past month used a contraceptive method. This suggests that both married and unmarried young women are at a high risk of

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unplanned pregnancies. However, ethnographic accounts of teenage sexuality in Mexico suggest that adolescent childbearing, especially in rural areas, is often not a random, unexpected event resulting from risky behavior but rather a result of social norms about early coupling and motherhood [Stern (2007)]. Therefore, understanding the costs of adolescent childbearing could be important for reducing teenage childbearing either through increasing demand for contraception or changing social norms.

3. Conceptual framework

The links between adolescent childbearing and educational and economic outcomes have been extensively explored, such as increased opportunity cost of education and labor force participation due to childbearing [Ribar (1994), Klepinger *et al.* (1999), Ribar (1999) among many others]. Because the links are well established for educational and labor market outcomes, while only a few rigorous studies explore the relationship between teenage childbearing and health, this section presents a brief conceptual framework guiding the estimation of the effect of first birth as a teenager on health.

Suppose women derive utility from consumption, C, and health, H.[3] They choose consumption and investment in health to maximize utility subject to a budget constraint and a health production function:

$$\max_{C,\, I} U(C, H) \text{ s.t.} p_c C + p_h I = Y \text{ and } H = h(I; \epsilon,\ E,\ \mu),$$

where p_c and p_h represent the marginal cost of the consumption good and the health investment good, I, respectively.[4] Health (H) is a function of investments in health which build on a health endowment, ϵ. The effectiveness of investments may depend on the level of education, E, as well as various community characteristics that determine access to and quality of healthcare services, μ.

The reduced-form demand functions for the consumption and investment goods can then be presented as the functions of prices and income: $C = f(p_h, p_c, Y)$ and $I = f(p_h, p_c, Y)$. Therefore, the reduced-form demand function for health takes the form: $H = h(p_h, p_c, Y; \epsilon, E, \mu)$. This simple model illustrates that teenage childbearing could affect the health of the mother in a number of ways.

Teenage childbearing could reduce education, which might affect health directly by reducing health knowledge and efficacy of investments in producing positive health outcomes [Grossman (2006), Cutler and Lleras-Muney (2010), Soares (2015)]. Reduced education could also lead to worse labor market outcomes including lower income and labor force participation, which could adversely affect health. Specifically, unemployment and lower earnings will likely reduce the demand for health inputs, such as preventive health goods, curative healthcare, and nutritious foods.

Additionally, teenage childbearing could be associated with worse socioeconomic outcomes, such as labor market participation due to childcare responsibilities and

[3]Health in this paper is a latent variable, measured by being overweight, being anemic, reporting good health, and a mental health score (more details in section 4).

[4]Health investments can include both monetary and time investments. The price associated with time investments is the opportunity cost of time and the psychological cost of effort. Measures of health investments in this paper include the probability and frequency of physical activity.

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lack of childcare or high childcare costs, which would in turn negatively affect the demand for health inputs and outcomes. Greater childcare responsibilities may also reduce time investments in health, such as time spent exercising [Gunes (2016)].

Teenage childbearing could also affect the probability of marriage. If teenage childbearing increases the likelihood of marriage, then it could lead to better health inputs and outcomes through greater household income. However, even if teen fertility is associated with higher marriage rates, quality of marriage may not be high. For example, in the US, Lichter and Graefe (2011) find that teen mothers are less likely to get married and also that their unions are more unstable. Similarly, Gunes (2016) and Webbink *et al.* (2008) find that teen mothers in the US and Australia match with a lower quality spouse. On the other hand, if teenage childbearing increases the likelihood of being single then it could reduce household income and in turn adversely affect health.

Teenage childbearing could worsen health directly through changing the health endowment of a woman due to increased stress or other biological channels. While Gunes (2016) finds no effect on mental health, Liao (2003) finds that early childbearing is associated with worse mental health soon after birth and Henretta *et al.* (2008) find that negative mental health effects persist even in midlife. On the contrary, teenage childbearing could improve health if women begin to value health more and reduce the demand for unhealthy behaviors such as smoking. Yet, Webbink *et al.* (2008) find that teenage childbearing is associated with a higher probability of smoking in Australia. On the other hand, Fletcher (2012) finds no effect on smoking and a possible reduction in drug use and binge drinking in the US.

However, teenage childbearing is not a random shock to health and education outcomes. If women decide the timing of their childbearing based on a cost-benefit calculation, then early childbearing may not have any effect on health (or education). More specifically, only women with poor health, education, or labor market prospects might choose to have a child as a teenager because the benefits of delaying childbearing would be low and likely not exceed the costs of foregone utility from teenage childbearing.

Next, we discuss the outcomes explored in this paper and then present the empirical methodology used to estimate the relationship between teenage childbearing and socioeconomic outcomes, accounting for potential selection into teenage childbearing.

4. Data

The data used in this paper are from the Mexican Family Life Survey (MxFLS), which follows 35,000 individuals from 8,400 households in 150 communities throughout the country.[5] The MxFLS is the first nationally representative longitudinal survey in Mexico. The first wave was conducted in 2002 (MxFLS-1), and the second (MxFLS-2) and third (MxFLS-3) waves were conducted during 2005–2006 and 2009–2012, respectively. Because the analysis requires two consecutive waves, and wave 1 is missing key baseline characteristics and is incomplete in various other respects, we are restricted to waves 2 and 3.[6]

[5] http://www.ennvih-mxfls.org/.

[6] Due to missing variables, wave 1 is subject to sample selection bias and we find that the balancing tests are unsupported using wave 1. Using a single set of waves is also advantageous as it reduces concerns with potential changes in the environment over time and changes in the implementation of the survey between waves.

4.1 Education and health outcomes

We examine the effect of having first child as a teenager on education as measured by the probability of dropping out of school prior to high school completion. Human capital can also be accumulated through work. Therefore, we examine employment status as a measure of human capital accumulation in addition to using it as a labor market outcome.

For health outcomes, we group the outcome variables into three categories: "physical health," "mental health," and "health behaviors." Physical health outcomes explored include anemia, being overweight, and self-reported health status. The binary indicator for being overweight (BMI of 25 or higher) is calculated using height and weight measured by trained health workers during the survey. Teenage childbearing has been associated with an increased risk of being overweight or obese later on in life [Chang *et al.* (2013)]. It is important to know whether young mothers are more likely to have a higher BMI because obesity is associated with a higher incidence of other diseases, such as type 2 diabetes, high blood pressure, and coronary heart disease and is responsible for 6–11% of deaths caused by major non-communicable diseases [Lee *et al.* (2012)].

While being overweight may be a side effect of teenage childbearing, another negative health consequence may be the increased risk of anemia. Adolescents between the ages of 10 and 19 have high iron needs and anemia is a cause of concern because it may have long-run negative consequences on cognitive development and growth [PAHO/WHO (2008)]. Teenage childbearing further exacerbates the risk of anemia and it may be associated with health complications for the young mother as well as the infant. Anemic mothers are more likely to have pre-term, low-weight babies, and the mother's iron deficiency may have long-run consequences for the infant's physical and mental health [Viteri (1994)]. The MxFLS data provide information on hemoglobin levels in blood samples taken during the survey and thus allow studying the incidence of anemia (hemoglobin of <12 g/100 ml).

Self-reported health status (an overall measure of health and well-being) has been shown to be a strong determinant of mortality and morbidity [see Idler and Benyamini (1997) for a survey of the literature].[7] The outcome is a dummy variable equal to 1 if the respondent reports very good or good health, and 0 otherwise (regular, bad, or very bad health).

In addition to physical health outcomes, we also examine mental health outcomes. Mental health is of interest because poor mental health may affect socioeconomic outcomes and thus short-run shocks to mental health may have long-term consequences [Haushofer and Fehr (2014), Cornaglia et al. (2015), Lybbert and Wydick (2015)]. Moreover, poor mental health negatively affects the birth weight of the infants, which has adverse long-term consequences [Conway and Kennedy (2004)]. Mental health outcomes explored in this study are survey-based and measured based on a 20-question assessment of depressive symptoms developed by Calderon (1997) for the Mexican population.[8] The assessment includes questions

[7]While two out of twenty-seven studies surveyed by Idler and Benyamini (1997) include young people in their analysis, none of the studies focus on the accuracy of adolescent reporting in particular. There are, however, several papers that document the reliability of self-reported health measures among adolescents [e.g., Fosse and Haas (2009) and Allen *et al.* (2016)].

[8]The Calderon scale is generally similar to the CES-D scale, which has been used extensively in the US. Both are based on a similar set of questions and use identical aggregation methods (the Calderon scale

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such as "In the last 4 weeks, have you felt sad or depressed?" and "In the last 4 weeks, have you lost interest in things?" Respondents can answer each question by "No," "Yes, sometimes," "Yes, lots of times," and "Yes, all the time" and each question is given a score between 1 (for no symptoms) and 4 (for symptoms present all the time). Thus, the overall assessment values range from 20 to 80 with higher values signifying greater mental health impairment. Clinical experience has shown that scores in the range of 20 through 35 are normal, scores between 36 and 45 indicate some anxiety, scores between 46 and 65 indicate a moderate level of depression and anxiety, and scores between 66 and 80 indicate severe depressive symptoms [Calderon (1997)]. Using this assessment, we measure mental health both as a continuous variable and a dummy variable indicating whether the scores are in the normal range.

The health behaviors explored relate to physical activity. Physical activity is measured as a dummy variable indicating whether the respondent routinely exercises during the weekdays, as well as the number of days the respondent exercises in a given week. Forty-one percent of 15–18 year olds in Mexico did not meet the physical activity recommendation of 60 min of moderate-to-vigorous physical activity per day in 2012, and sports practice was not a common type of physical activity in women of childbearing age in 1999 [Hernández *et al.* (2003), Lee *et al.* (2012), del Martinez *et al.* (2014)]. Physical inactivity could be associated with obesity as well as poor mental health outcomes.[9]

4.2 Sample of analysis and teenage childbearing

Our main "treatment" variable is defined as teenage childbearing, which is a dummy variable indicating whether the women had their first child at age 18 or earlier. Since the data do not contain retrospective information on the adolescence of older women, we are unable to model the risk factors for teenage pregnancy of all women. Thus, the main sample of analysis includes girls at the ages of 15–18 in 2005 (MxFLS-2, "baseline" hereafter) that have had no pregnancy prior to 2005 and complete a follow-up survey in 2009–2012 (MxFLS-3). We also exclude girls younger than 18 at endline (to avoid the problem of censoring). This yields a total of 805 girls who have non-missing baseline information used in the calculation of the propensity score and are also followed-up in MxFLS-3. Ten percent (82) of these girls give their first birth by the time they are 19.[10]

Table 1 compares the baseline characteristics of the girls that experienced teenage birth and those who did not. A lower proportion of girls that subsequently had teenage childbearing were in school at baseline (0.52 vs. 0.62) although the difference is only significant at the 10% level and there are no significant differences in cognitive test scores (measured by Raven's progressive matrices assessment) between the two groups. Girls that experienced teenage childbearing were less likely to be single (0.87 vs. 0.95, p-value = 0.001) and more likely to have ever had sex (0.22 vs. 0.06, p-value<0.001). A significantly higher proportion of the girls with teenage childbearing lived in small towns (less than 2,500 inhabitants) compared to the girls that did not have teen childbearing, their households appear to be poorer (based on

starts at 20, whereas the CES-D scale starts at 0, but both have a range of 60 points). The estimates using the Calderon scale are thus directly comparable to the estimates using the CES-D scale.

[9]See Soares (2015) for a detailed discussion on complementarities between health and health behaviors.

[10]This is similar to the prevalence of teen childbearing in Mexico [Challenges (2007)].

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Table 1. Baseline characteristics of girls prior to any birth

	No teenage childbearing	Teenage childbearing	p value
Individual characteristics			
Age			
Age 15	0.24	0.51	0.000
Age 16	0.28	0.28	0.975
Age 17	0.24	0.17	0.189
Age 18	0.25	0.04	0.000
Currently in school	0.63	0.52	0.068
Cognitive test score	7.44	7.3	0.676
Single	0.96	0.87	0.000
Age at menarche	11.93	11.89	0.757
Ever had sex	0.06	0.22	0.000
Risk averse	0.06	0.07	0.701
Impatient	0.41	0.39	0.720
Normal range of mental health	0.93	0.93	0.967
Household characteristics			
Household size	5.8	5.24	0.032
Mother away	0.09	0.17	0.013
Father away	0.2	0.38	0.000
Household owns a house	0.86	0.78	0.044
Household owns a car	0.44	0.28	0.007
Household owns a washing machine	0.89	0.87	0.551
Household owns domestic appliances	0.89	0.83	0.099
Household lives in locality with more than 100,000 inhabitants	0.33	0.24	0.123
Household lives in locality with 15,000–100,000 inhabitants	0.09	0.09	0.957
Household lives in locality with 2,500–15,000 inhabitants	0.15	0.09	0.123
Household lives in locality with less than 2,500 inhabitants	0.44	0.59	0.011
Observations	723	82	

Note: Mean (std) for continuous variables; proportions for discrete variables; p-values calculated using a t-test.

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the availability of several household assets), and they are more likely to report having a mother or father who is away.[11]

4.3 Attrition and non-response

There are 1,328 girls who meet the age restriction at baseline and have had no prior pregnancy ("target sample"). The sample of analysis includes 805 girls (61% of the target sample) since there is attrition between the survey waves and non-response to the relevant fertility modules of the questionnaires at baseline. We compare the characteristics of the girls excluded from the analysis to the girls in the sample of analysis in Appendix Table 1. Women who were lost to follow-up or did not respond to the survey questions were generally older than our sample of women and thus less likely to be single or still attending school and more likely to have ever had sex at baseline.[12]

5. Methodology

If teenage childbearing (treatment) was random, then an OLS model comparing the average outcomes of girls who had their first birth as a teenager, TF, and those who did not would yield unbiased estimates of the treatment effect, β_0:

$$Y = \alpha + \beta_0 \mathrm{TF} + W_1\beta_1 + \epsilon \qquad (1)$$

where Y is the outcome of interest, W_1 is a vector of observed individual, household and community characteristics, and ϵ is a random error term.

However, teenage childbearing is likely not random and girls that did not have their first birth as a teenager may not provide the appropriate counterfactual for what would have happened had the teenagers not become mothers. For example, if girls with teenage childbearing were more likely to drop out of school or have unhealthy lifestyles even without having a child, then the effect of childbearing would be overestimated. Alternatively, if only girls who were in a good physical or mental health state were willing to become mothers, then any negative effects of teenage childbearing may be underestimated. In other words, estimating model (1) may yield biased results due to selection into teenage childbearing.

5.1 Propensity score matching

In order to account for observable determinants of teenage childbearing, and thereby attempt to reduce concerns about possible selection bias, we employ an inverse

[11]Note that it is unsurprising that the difference in age 18 at baseline between the two groups is significant as the proportion of girls aged 18 at baseline that are included in the treatment group is quite small by definition.

[12]We test whether selection bias may affect our main results using a simple test of selection bias and did not find evidence for selection bias for any outcome, except for dropping out of school. Specifically, we calculate the inverse mills ratio from a regression of a dummy variable indicating inclusion in the empirical analysis on baseline characteristics and include this term as an additional control in our main model (similar to Heckman selection model). While selection bias may be a concern in estimating the effect of teenage childbearing on schooling, the inclusion of the additional control increases the effect of teenage childbearing, suggesting that the reported coefficient is a lower bound estimate.

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propensity score weighting (IPW) approach.[13] We use a logit model and a rich set of individual characteristics to predict the probability of first birth at age 18 or earlier (TF):

$$\mathrm{TF}_{i,t+1} = \alpha_0 + X_{i,t}\gamma_1 + \nu_{i,t+1}, \tag{2}$$

where $X_{i,t}$ is a rich set of baseline characteristics presented in Table 1. More specifically, we use age fixed effects, school attendance, cognitive test scores, marital status, age at menarche, ever having had sex, being risk averse, being impatient, having a normal range of mental health scores, household size, size of locality where household lives, mother away, father away, and household ownership of various assets (a house, a car, a washing machine, domestic appliances).[14] We also control for survey year and state fixed effects.[15]

As mentioned, teenage childbearing is not random, and the general aim of propensity score matching is to mimic randomization by creating treatment and control groups which, after weighting, are balanced on observables. Interpreting the effects as causal requires the so-called conditional independence assumption, which entails that factors that affect teenage childbearing decisions and outcomes related to teenage childbearing (treatment-specific outcomes) are observable. To this end, we leverage the rich set of pre-treatment characteristics to remove the dependence between teenage childbearing and treatment-specific outcomes by conditioning on these covariates. If the assumption that observables account for confounding factors holds, then we can interpret our findings as causal.

The IPW approach uses the propensity scores estimated from equation (2) to calculate weights that are included in the estimation of the effects of teenage childbearing on the outcomes of interest. If $p(X_{i,t})$ is the propensity score, then the weight is $1/p(X_{i,t})$ for girls with teenage childbearing and $1/(1 - p(X_{i,t}))$ for girls without teenage childbearing. The use of the weights equalizes the distribution of the confounders in the two groups and the average treatment effect using an IPW approach would then be estimated as the difference in the weighted average outcome of each group. We use regression-adjusted IPW analysis to account for differences in characteristics at endline. Overall, we include the following regression controls: age fixed effects, household asset ownership (house, car, washing machine, domestic

[13]Several papers studying the effects of adolescent childbearing on socioeconomic outcomes have used propensity score matching [Chevalier and Viitanen (2003), Levine and Painter (2003), Kane *et al.* (2013)]. Moreover, recent studies use propensity score matching methods to explore various relationships in developing countries. For example, Woode *et al.* (2017) employ a propensity score matching method to explore the role of health insurance in the relationship between parental health shocks and child work in Rwanda.

[14]In the IPW regressions, risk aversion and impatience are not significant determinants of teenage pregnancy. While insignificance is not necessarily a reason to exclude these variables, the results are similar when we exclude these variables from the IPW regressions. The results are available upon request.

[15]The estimated propensity scores for the full sample range from 0 to 0.9 both for girls with and without teenage childbearing (Appendix Figure 1), indicating a good predictive power of the model (we also provide a figure with common support by score in the Online Appendix). OLS and IPW have the same number of observations as there are no individuals that fail the common support restriction (for a given value of the propensity score, there should be both a treated and untreated observation). Moreover, a figure showing the bias reduction across covariates is provided in the Online Appendix. A separate propensity score matching regression is estimated for each outcome, accounting for missing values.

Table 2. Internal balancing test—baseline characteristics after weighting by the inverse propensity score

	Difference (Teenage childbearing—Not)	SE	*t*-Statistic
Individual characteristics			
Age			
Age 15	0.096	0.078	1.241
Age 16	−0.08	0.056	−1.448
Age 17	0.111	0.1	1.108
Age 18	−0.127**	0.063	−2.004
Currently in school	−0.014	0.09	−0.154
Cognitive test score	−0.381	0.578	−0.659
Single	0.01	0.017	0.572
Age at menarche	0.073	0.173	0.423
Ever had sex	−0.016	0.023	−0.699
Risk averse	−0.018	0.025	−0.729
Impatient	0.06	0.092	0.651
Normal range of mental health	0.002	0.036	0.050
Household characteristics			
Household size	−0.101	0.291	−0.348
Mother away	−0.014	0.031	−0.465
Father away	0.069	0.086	0.807
Household owns a house	0.025	0.04	0.636
Household owns a car	−0.008	0.096	−0.079
Household owns a washing machine	0.005	0.047	0.114
Household owns domestic appliances	0.004	0.048	0.089
Household lives in locality with more than 100,000 inhabitants	−0.002	0.088	−0.023
Household lives in locality with 15,000–100,000 inhabitants	−0.01	0.037	−0.283
Household lives in locality with 2,500–15,000 inhabitants	−0.061	0.041	−1.482
Household lives in locality with less than 2,500 inhabitants	0.073	0.091	0.805
Observations	805		

Note: (1) Internal balancing test based on a regression model weighted with the inverse propensity score weights. Each number is from a separate regression of the variable on an indicator for early pregnancy. (2) Variables included are those used in the propensity score. Additional variables in the propensity score model not presented here are State Fixed Effects and Year Fixed Effects.

appliances), household size, size of area of residence, state fixed effects, and year of survey fixed effects.

We perform various tests to verify that observable characteristics are similar across treatment and control groups after weighting by the inverse propensity scores. First, we examine whether the means of the observable baseline characteristics are balanced. The results in Table 2 suggest that there is no significant difference in the means of the baseline characteristics between girls that experienced teenage childbearing and those that did not once the means are weighted using the inverse propensity scores, with the exception of age 18 at baseline. To account for age effects, we control for age fixed effects in the estimations. We also perform a regression-based balancing check, based on Smith and Todd (2005), as an additional test. In particular, we estimate the following regression for every j covariate included in the propensity score estimation:

$$X_{i,j,t} = \mu_0 + \mu_1 p(X_{i,t}) + \mu_2 p(X_{i,t})^2 + \mu_3 p(X_{i,t})^3 + \eta_0 \mathrm{TF}_{i,t+1} + \eta_1 \mathrm{TF}_{i,t+1} p(X_{i,t}) + \eta_2 \mathrm{TF}_{i,t+1}\, p(X_{i,t})^2 + \eta_3 \mathrm{TF}_{i,t+1}\, p(X_{i,t})^3 + u_{i,j,t}.$$

We test for joint significance of the coefficients denoted with η for each covariate (presented in Appendix Table 2). If the propensity score satisfies the balancing assumption, then teenage childbearing should not provide any additional information about the covariate X. Consistent with the means-based balancing test, the results of the regression-based balancing tests demonstrate that propensity-score-weighted teenage childbearing is not significantly correlated with the variables, except in two cases (dummy variables for age 18 and school attendance at baseline). We control for the only variable (age 18 at baseline) that fails both tests in the estimations.

Employing IPW has several advantages. First, IPW is computationally parsimonious, and is asymptotically efficient, at least practically [Huber *et al.* (2013)]. Second, compared to other matching approaches, it does not require choosing a tuning parameter, such as the choice of the number of matched neighbors in the nearest neighbor model and the choice of distance in the radius matching model [Huber *et al.* (2013)]. Moreover, the degree of bias and efficiency depends on the selection of the tuning parameter and data structure [Caliendo and Kopeinig (2008)]. Finally, while direct matching methods usually reduce sample size due to discarding unmatched individuals, IPW retains sample size as only observations outside the range of common support are discarded, which in turn reduces potential bias and increases the precision of the estimates [Garrido *et al.* (2014)]. The primary drawback to IPW, and PSM more generally, is that it requires that all factors related to selection are observable, or at least are correlated to observables, and that average treatment effects are based on the matched sample [Kane *et al.* (2013)]. Nevertheless, while IPW is our main estimation method, we also provide robustness checks to alternative matching algorithms. Specifically, we employ radius matching and entropy balance methods, which are discussed in more detail in section 6.2.

As discussed above, the validity of the IPW approach hinges on the assumption that observable baseline characteristics account for selection into teenage childbearing. While the baseline characteristics include a rich set of covariates, there remains a concern that unobservable characteristics might be related to both teenage childbearing and the outcomes of interest. For example, characteristics such as intrinsic motivation, which might not be reflected in cognitive test scores or other observables (e.g., risk-aversion or patience) might be related to the chances of becoming pregnant as a teenager and various SES outcomes. In this case, because motivation plausibly reduces the likelihood of teenage childbearing and improves SES outcomes, the IPW estimates would overstate the adverse impacts of teenage childbearing. While teenage pregnancy is typically unplanned, it is possible that risk preferences might interact with fertility and family preferences, which in turn might be correlated with teenage pregnancy and later adult outcomes. Thus, we emphasize that selection bias might remain a concern as the observable baseline characteristics might not account for all of the relevant factors in selection into teenage childbearing.

5.2 Sibling fixed effects

While the main analysis uses matching, we also perform robustness checks using sibling fixed effects (FE). We estimate the following regression model:

$$Y_{i,g} = \alpha + \gamma_0 \mathrm{TF}_i + W_1\gamma_1 + \kappa_g + \nu_{i,g}, \quad (3)$$

where $Y_{i,g}$ is the outcome of interest for woman i in sibling group g, κ_g represents the sibling fixed effects, and the other variables are defined as before.[16] The sibling FE model eliminates bias due to unobservable characteristics that are constant within families such as family background, community characteristics, or genetic predisposition to certain health conditions and health behaviors that may also be correlated with teenage childbearing. We estimate equation (3) using samples consisting of women aged (i) 30 and below and (ii) 35 and below, and restricting the sample to siblings with age differences within 5 years (or less) to reduce heterogeneity due to unobserved parental background and investments. To circumvent the problem of censored teenage childbearing, the youngest women in the sibling FE analysis are aged 19 at endline.

While sibling FE models account for confounding factors common for siblings, many of the potential concerns regarding the IPW estimates are also concerns with confounding factors within families. For example, differences in intrinsic motivation or preferences within siblings might be related to both the probability of becoming pregnant as a teenager and later adult outcomes. The caveat regarding selection bias remaining a concern is therefore also relevant to the FE estimates as well.

6. Results

6.1 Main results

Panel A of Table 3 presents the results from an OLS estimation of the effect of teenage childbearing, not accounting for any selection into teenage childbearing. The results suggest that teenage childbearing is associated with a significantly higher probability of dropping out of school. Teenage childbearing is positively correlated with being overweight and anemic, and negatively correlated with physical activity. Poor health, lack of physical exercise, and the stress of child-bearing and child-rearing could

[16]Sibling groups are based on observations from all three waves of the MxFLS, and thus, sibling groups include siblings who lived in the same household any time between 2002 and 2009.

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Table 3. Effect of teenage childbearing on educational and health outcomes

	Education and labor market outcomes		Physical health outcomes			Mental health outcomes		Health behaviors	
	Dropped out of school	Working	Overweight (BMI ≥25)	Anemic	Reports good health	Mental distress score	Normal range of mental health	Exercises	Exercise frequency
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Panel A: OLS									
Teenage childbearing	0.2976***	0.0074	0.1361**	0.1687***	−0.0659	−0.1956	−0.0065	−0.1447***	−0.5987***
	(0.0537)	(0.0561)	(0.0653)	(0.0510)	(0.0595)	(0.9461)	(0.0355)	(0.0324)	(0.1287)
Observations	805	805	641	603	680	680	680	680	680
Panel B: augmented OLS									
Teenage childbearing	0.1910***	−0.0121	0.1387**	0.1481***	−0.0593	−0.0732	−0.0143	−0.1569***	−0.6342***
	(0.0499)	(0.0556)	(0.0678)	(0.0511)	(0.0618)	(0.9622)	(0.0353)	(0.0344)	(0.1386)
Observations	805	805	641	603	680	680	680	680	680
Panel C: IPW									
Teenage childbearing	0.2809***	−0.0848	0.1150*	0.1272**	−0.0018	0.7131	−0.0588	−0.1733***	−0.6559***
	(0.0591)	(0.0525)	(0.0628)	(0.0607)	(0.0503)	(0.8507)	(0.0446)	(0.0227)	(0.0906)
Observations	805	805	641	603	680	680	680	680	680
Mean outcome girls without teen childbearing	0.45	0.29	0.4	0.09	0.72	25.62	0.92	0.18	0.69

Note: (1) All models control for the following individual and household characteristics: age fixed effects, household asset ownership (house, car, washing machine, domestic appliances), household size, size of area of residence, state fixed effects, years of survey fixed effects. Panel B additionally controls for the baseline characteristics included in the matching model. (2) Robust standard errors in parentheses. (3) *Significance at the 10% level, **significance at the 5% level, ***significance at the 1% level.

increase mental distress. Yet, teenage childbearing does not have any significant effect on mental health. We cannot rule out that this finding might be due to a measurement problem or the fact that more than 90% of girls have mental health scores that are in the normal range. While the estimated effect on self-reported health is negative, it is not statistically significant. While the results of the OLS estimations are informative, they are likely subject to selection bias.[17]

In an attempt to control for potential confounding factors and reduce selection bias, analysis in Panel B conditions the OLS estimates on the baseline characteristics included in the propensity score model. Results between the two models are similar.

Finally, Panel C of Table 3 presents the results of the effects of teenage childbearing using inverse probability weights. The estimated effects using IPW are consistent with the OLS estimates, suggesting that teenage childbearing negatively affects education, physical health (being overweight and anemic), and physical activity. Specifically, teenage childbearing increases the probability of dropping out of school by 28 percentage points, similar to the effect size found in Ardington *et al.* (2015) in South Africa, where teenage mothers are 26 percentage points more likely to drop out of school by age 20. In the US, Hotz *et al.* (2005) find that teenage childbearing decreases the probability of obtaining a High School diploma by 5–12%. Teenage childbearing is also associated with 12 and 13 percentage points lower probability of being overweight and anemic, respectively. While there are no comparable studies examining anemia prevalence, our estimate for being overweight is similar to the effect found by Webbink *et al.* (2008) for Australian teenagers. They find that teenage childbearing in Australia is associated with 19 percentage points higher probability of being overweight. The higher effect found in Australia could be due to differences in institutions, and social and cultural norms. We further test one potential mechanism which could explain overweight risk in Mexico—physical activity. We find that teenage childbearing reduces the probability of exercising by 17 percentage points and exercise frequency by about 1 day a week. Gunes (2016) similarly finds that teenage childbearing reduces physical activity in the US. Lastly, consistent with Gunes (2016), we do not find significant effects on mental health.

Comparing the OLS with IPW results in Panels A and C of Table 3, we find that using the propensity score matching approach to account for selection on observable characteristics reduces the magnitude of the coefficients on dropping out of school before completing high school, being overweight, and being anemic by 6%, 16%, and 25%, respectively. Thus, girls who are more likely to leave school before graduating, be overweight, have anemia are also more likely to give birth as a teenager. On the other hand, employing IPW increases the magnitude of the coefficients on the probability of exercising and exercise frequency by 20% and 10%, respectively. Thus, girls who are more likely to exercise are less likely to give birth as a teenager. Consistent with the OLS estimates, we find no significant effect on self-reported

[17]While the OLS estimates are potentially subject to selection bias, the treatment effect can be bound using a bounding-approach proposed by Altonji *et al.* (2005) and Oster (2019). In particular, following Altonji *et al.* (2005), we estimate the bounds of the treatment effects using a proportionality factor of one, and find that the bounding sets exclude zero for all outcomes. Moreover, we calculate the proportionality factor needed such that the treatment effect is equal to zero, and find that the proportionality factors are always greater than one, implying that the OLS estimates are consistent with significant treatment effects. More details regarding the bounding approach and the results are presented in the Online Appendix.

health and mental health. Overall, we find relatively small differences between OLS and IPW, suggesting that either OLS is less biased than expected or that there are unobserved confounding factors that are not related to the baseline observable characteristics.

Our IPW method uses observed characteristics of 15–18 years old girls at the time of the initial survey as the determinants of teenage childbearing. Because first births as a teenager take place within 1–3 years of the baseline survey, observed characteristics at baseline are good proxies for conditions at the time of first birth as a teenager (at age 18 or earlier). To investigate the extent to which the estimates are affected by the length of time in which the observable characteristics precede the treatment, we employ a robustness check. In particular, the estimations restrict the treated group to only women with first births within 2 years of the baseline survey. In the Online Appendix, we show that the results are consistent with the baseline results.[18]

We also re-estimate the effects using an alternative measure of teenage childbearing —probability of first birth at age 19 or earlier—and find that the results are consistent with the main results.[19] Specifically, the probability of first birth by age 20 increases the probability of dropping out of school and being overweight by 24 and 16 percentage points, while it reduces the probability of exercising by 11 percentage points and exercise frequency by half a day.[20]

Finally, if overweight girls are more likely to engage in risky behavior, then the estimated effects might be biased. For example, Averett *et al.* (2013) find that overweight or obese girls in the US are more likely than their recommended-weight peers to have ever had anal intercourse and are exposed to the same risks from vaginal intercourse as their peers once they have had vaginal intercourse. We control for being overweight at baseline in the estimations as a robustness check and find that the effects are consistent.[21]

6.2 Alternative matching models

In this section, we provide robustness checks to alternative matching algorithms, including entropy and radius matching approaches. While each matching method has its own limitation, it would be reassuring that there are significant adverse effects of teenage childbearing when we use different methods.

Following Hainmueller (2012), we employ an entropy matching approach that reweights the dataset by adjusting the sample to pre-specified moments of the covariate distribution to create balanced samples. We also employ a radius matching approach that matches each individual from a treatment group with individuals from a control group within a pre-specified distance around the propensity score. Panels A and B of Table 4 show the results using entropy and radius matching approaches,

[18]We also employ another robustness check by restricting the sample of analysis to include only women that give birth. The results (provided in the Online Appendix) are generally consistent with the main results, but there is some loss of precision due to smaller sample sizes.

[19]We use girls at the ages of 15–19 in 2005 that have had no pregnancy prior to 2005 and complete a follow-up survey in 2009–2012 (MxFLS-3). We also exclude girls younger than 19 at endline to avoid the problem of censoring.

[20]Full set of results are available upon request.

[21]Results are provided in the Online Appendix. Furthermore, using the change in the overweight status between baseline and follow-up surveys as an outcome, we find that teenage childbearing is positively associated with becoming overweight (it increases becoming overweight by 12 percentage points).

Table 4. Effects of teenage childbearing using different matching algorithms

	Education and labor market outcomes		Physical health outcomes			Mental health outcomes		Health behaviors	
	Dropped out of school	Working	Overweight (BMI ≥25)	Anemic	Reports good health	Mental distress score	Normal range of mental health	Exercises	Exercise frequency
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Panel A: entropy match									
Teenage childbearing	0.173**	−0.0313	0.138**	0.133**	−0.0473	0.254	−0.0188	−0.122**	−0.508**
	(0.0560)	(0.0614)	(0.0676)	(0.0530)	(0.0630)	(1.109)	(0.0366)	(0.0352)	(0.127)
Panel B: radius matching									
Teenage childbearing	0.213**	−0.044	0.169**	0.157**	−0.027	−0.17	−0.008	−0.148**	−0.589**
	(0.06)	(0.0663)	(0.0787)	(0.0619)	(0.0692)	(1.1226)	(0.0396)	(0.0431)	(0.1821)

Note: (1) Entropy matching approach based on Hainmueller (2012). (2) Radius matching uses caliper of 0.1. (3) All models control for the following individual and household characteristics: age fixed effects, household asset ownership (house, car, washing machine, domestic appliances), household size, size of area of residence, state fixed effects, years of survey fixed effects. (4) Robust standard errors in parentheses. (5) *Significance at the 10% level, **significance at the 5% level, ***significance at the 1% level.

Table 5. Effects of teenage childbearing using sibling fixed effects

	Education and labor market outcomes		Physical health outcomes			Mental health outcomes		Health behaviors	
	Dropped out of school	Working	Overweight (BMI ≥25)	Anemic	Reports good health	Mental distress score	Normal range of mental health	Exercises	Exercise frequency
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Panel A: sample of women aged 30 and younger									
Teenage childbearing	0.2208***	−0.0483	0.1822***	0.0592	−0.0028	0.3435	−0.0132	−0.0601	−0.3479*
	(0.0486)	(0.0522)	(0.0593)	(0.0472)	(0.0681)	(0.7876)	(0.0311)	(0.0461)	(0.1920)
Number of women	1,229	1,229	860	816	1,004	1,004	1,004	1,004	1,004
Number of sibling groups	556	556	395	376	460	460	460	460	460
Panel B: sample of women aged 35 and younger									
Teenage childbearing	0.1989***	−0.0512	0.1602***	0.0592	0.0207	0.399	−0.0101	−0.0607	−0.3230*
	(0.0487)	(0.0508)	(0.0571)	(0.0456)	(0.0678)	(0.7718)	(0.0309)	(0.0449)	(0.1832)
Number of women	1,373	1,373	953	901	1,108	1,108	1,108	1,108	1,108
Number of sibling groups	615	615	433	410	502	502	502	502	502

Note: (1) Sample includes siblings with an age difference of 5 years or less. Sample based on siblings in the second wave of survey (MxFLS-2). Sample restricted to women 19 years old or older. (2) All models control for the following individual and household characteristics: age fixed effects, household asset ownership (house, car, washing machine, domestic appliances), household size, size of area of residence, state fixed effects, years of survey fixed effects. (3) Robust standard errors in parentheses. (4) *Significance at the 10% level, **significance at the 5% level, ***significance at the 1% level.

Table 6. Effects of first births at older ages using sibling fixed effects

	Education and labor market outcomes		Physical health outcomes			Mental health outcomes		Health behaviors	
	Dropped out of school	Working	Overweight (BMI ≥25)	Anemic	Reports good health	Mental distress score	Normal range of mental health	Exercises	Exercise frequency
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Panel A: Sample of women aged 25–30									
Birth between the ages of 20–24	0.029	0.0581	−0.0189	0.0055	0.0384	1.7672	0.0241	−0.0135	0.1849
	(0.0934)	(0.1119)	(0.0960)	(0.0770)	(0.1483)	(2.2150)	(0.0873)	(0.1188)	(0.4839)
Number of women	211	211	140	134	158	158	158	158	158
Number of sibling groups	100	100	65	64	75	75	75	75	75
Panel B: Sample of women aged 25–35									
Birth between the ages of 20–24	0.0439	−0.0173	0.0113	0.0631	0.019	0.9882	0.0572	−0.0425	0.0116
	(0.0708)	(0.0962)	(0.0836)	(0.0841)	(0.1103)	(1.5441)	(0.0641)	(0.0848)	(0.3419)
Number of women	343	343	218	207	253	253	253	253	253
Number of sibling groups	161	161	101	98	119	119	119	119	119

Note: (1) Sample includes siblings with an age difference of 5 years or less. Sample based on siblings in the second wave of survey (MxFLS-2). (2) All models control for the following individual and household characteristics: age fixed effects, household asset ownership (house, car, washing machine, domestic appliances), household size, size of area of residence, state fixed effects, years of survey fixed effects. (3) Robust standard errors in parentheses. (4) *Significance at the 10% level, **significance at the 5% level, ***significance at the 1% level.

respectively. The results confirm the adverse negative effects of teenage childbearing on education and health outcomes. Moreover, the estimated effects are not statistically different from the estimated effects using IPW.

6.3 Sibling FE

Table 5 presents the results of employing sibling FE using different samples, and shows that the results corroborate the main effects. The sibling FE estimates in Panel A (Panel B) suggest that teenage childbearing increases the probability of dropping out of school by 22 (20) percentage points and the probability of being overweight by 18 (16) percentage points, while it reduces exercise frequency by about 0.35 days (0.32) in a week. While the estimated effects on the probability of being anemic and exercising are statistically insignificant, the imprecision of the estimates might be due to little variation in the outcome variables or teenage childbearing across siblings, which is a well-known concern in within-sibling approaches.

We also explore whether first births at older ages have similar effects. Because we do not have the requisite information for older women at the time of their first birth, we cannot employ an analogous propensity score matching approach to explore the effects of later childbearing. Alternatively, we employ the sibling FE approach. In particular, we restrict the sample to female siblings aged 25–30 (and 25–35) in the second wave, and estimate the effect of the probability of first birth between the ages of 20–24 using sibling fixed effects (control group is first birth after 25 or no first births). The results in Table 6 suggest insignificant effects of giving birth at older ages on educational and health outcomes and the estimated effects are much smaller than the effects of teenage childbearing in general, particularly for the probability of being overweight and dropping out of school.

7. Conclusion

Despite the reduction in total fertility rates in Mexico, little progress has been made in reducing teenage childbearing. In order to design effective poverty-reduction policies, it is important to understand the costs associated with teenage childbearing. There is a vast literature on the effects of teenage childbearing on the socioeconomic outcomes of mothers and their children in the developed world but there is little rigorous evidence from developing countries. In addition, while most research has focused on studying the education and labor market outcomes of young mothers or the education and health of their children, the evidence on health outcomes and behaviors of young mothers is limited. Health of teen mothers could deteriorate as a consequence of teenage childbearing for a number of reasons, including disruptions in schooling and employment [Ribar (1994), Ribar (1999), Klepinger *et al.* (1999)], which might lead to reduced health knowledge and investments in health [Grossman (2003)], reductions in the probability and quality of marriage [Webbink *et al.* (2008)], and increased stress [Liao (2003)].

This paper explores the effect of teenage childbearing on socioeconomic outcomes of young girls in Mexico in terms of education and health. To attempt to reduce the bias associated with confounding factors related to teenage childbearing, we employ an inverse probability weighting approach that accounts for a set of baseline covariates that are often not available to researchers. The results suggest that teenage childbearing negatively affects education, working, and health and health behavior of young girls in the short run. Specifically, teenage childbearing increases the probability of dropping out of school before completing high school and being overweight, and reduces physical activity. Moreover, the results are robust to employing alternative matching methods, including entropy and radius matching models, using sibling fixed-effects approach that accounts for unobservable family background, various sample restrictions, and using an alternative measure of teenage childbearing. Using sibling fixed-effects approach, we also find that giving birth at older ages is not associated with the educational and health outcomes.

Overall, this paper aims to bring more attention to the relevance of socioeconomic consequences, particularly health consequences, of early childbearing in designing policies to address poverty. While young mothers could potentially overcome the negative consequences of early childbearing on education through more training over time and catch up to their peers, deterioration in health may have serious short-term and long-term consequences, which might not be easily overcome.

Supplementary material. The supplementary material for this article can be found at https://doi.org/10.1017/dem.2020.2.

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