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Misallocated Talent? Teen Pregnancy, Education and Job Sorting in Colombia

Jorge M. Agüero[1]

[1]Department of Economics.
University of Connecticut
jorge.aguero@uconn.edu

Policy makers and international organizations often argue that teenage pregnancy affects girls' life trajectories by, for example, limiting their employment opportunities. These concerns are amplified in regions with high teen pregnancy rates such as Latin America. We use a unique dataset from Colombia that allows us to instrument for early motherhood with the age at menarche. We find that teen pregnancy reduces school attainment and increases the number of children ever born. However, when considering eight indicators of labor supply, including labor force participation, type of job and occupation while accounting for multiple hypothesis testing, we find that much (if not all) of the negative effects on labor supply attributed to teen motherhood are due to selection. Our findings weaken the claim that early motherhood leads to a path of low-quality employment or a misallocation of talent due to job sorting. We discuss the role that family network and co-residence plays as a mechanism to buffer the effects of early motherhood on labor supply.

KEYWORDS
adolescent pregnancy, early motherhood, female labor force, family network

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¿Talento mal distribuido? Embarazo adolescente, educación y la selección de puestos de trabajo en Colombia

Jorge M. Agüero[1]

[1]Departamento de Economía.
Universidad de Connecticut.
jorge.aguero@uconn.edu

Los responsables políticos y las organizaciones internacionales suelen argumentar que el embarazo en la adolescencia afecta a la trayectoria vital de las niñas, por ejemplo, limitando sus oportunidades de empleo. Estas preocupaciones se amplifican en regiones con altas tasas de embarazo adolescente como América Latina. Utilizamos un conjunto de datos único de Colombia que nos permite instrumentar la maternidad temprana con la edad de la menarquia. Encontramos que el embarazo en la adolescencia reduce el rendimiento escolar y aumenta el número de hijos nacidos. Sin embargo, al considerar ocho indicadores de la oferta de trabajo, incluyendo la participación en la fuerza laboral, el tipo de trabajo y la ocupación, y teniendo en cuenta las pruebas de hipótesis múltiples, encontramos que gran parte (si no todos) de los efectos negativos sobre la oferta de trabajo atribuidos a la maternidad adolescente se deben a la selección. Nuestros resultados debilitan la afirmación de que la maternidad temprana conduce a una trayectoria de empleo de baja calidad o a una mala asignación del talento debido a la selección del empleo. Discutimos el papel que desempeña la red familiar y la corresidencia como mecanismo para amortiguar los efectos de la maternidad temprana en la oferta de trabajo.

KEYWORDS
embarazo adolescente, maternidad temprana, fuerza laboral femenina, redes familiares

1 | INTRODUCTION

Latin America's high adolescent fertility rate is often seen as a major public policy challenge (PAHO et al, 2017). For example, the region's ability to take advantage of the demographic dividend would be compromised if the youth face reproductive and productive roles that compete against each other. In fact, De Hoyos et al (2016) suggest that teenage pregnancy is a leading factor behind the large number of so-called "ninis;" young people out of school and out of work. Even if early childbearing does not affect labor force participation, it could lessen the quality of jobs women can pursue. These young mothers could be inefficiently engaged in low-quality jobs, creating a misallocation within the labor force and affecting overall productivity.

However, it has been difficult to identify whether the observed negative outcomes represent a true causal effect of teen childbearing or if teen mothers are negatively selected so that they will have poorer outcomes irrespective of their early childbearing. Data limitations in Latin America have also prevented the analysis of medium- to longer-term effects of early childbearing (Azevedo et al, 2012). The goal of our paper is to address these issues by measuring the longer-term effects of teen pregnancy, particularly on the human capital accumulation and labor supply of women. We focus on the case of Colombia, a country with an adolescent fertility rate close to the regional average and where a unique dataset allows us to overcome existing limitations while expanding on previous work about the causal impact of teenage motherhood in the region.

Most of the papers exploring the causal impact of early childbearing have focused on developed countries. For example, many papers have exploited variation in the timing of pregnancies among sisters using either complex longitudinal studies (e.g., Geronimus and Korenman, 1992; Groger and Bronars, 1993) or administrative records (e.g., Aizer, Devereux and Salvanes, 2019) to estimate longer-term impacts on education and labor-market outcomes.[1] Others have used miscarriages as an instrument.[2] However, recent papers have challenged this strategy because miscarriages are not random, are subjected to non-classical measurement error, and can be predicted by the frequency of abortions in the individual's neighborhood (Fletcher and Wolfe, 2009; Ashcraft, Fernandez-Val and Lang, 2013).

Our paper advances this literature by applying an alternative instrumental variable (IV) that is unaffected by the previous criticism and is available in a recent nationwide and large dataset for Colombia. Following work by Ribar (1994), Klepinger, Lundberg, and Plotnick (1999) and Field and Ambrus (2008), we instrument teen pregnancy with age at menarche. As shown by these authors and confirmed below for the case of Colombia (see section 3 below), women who have their first menstrual period at a younger age are more likely to become teen mothers. This is of course due to biology, as conception requires the onset of puberty. In the context of high teen pregnancy rates, this biological feature acts as a binding constraint on exposure to early childbearing. In particular, natural variation in the timing of first menstruation produces quasi-random differences in the earliest age at which adolescents are at risk of becoming mothers.

The validity of the IV strategy depends on whether unobserved differences in family background shape both adolescent maturation and later outcomes. Field and Ambrus (2008) conducted an extended review of the medical literature and concluded that most of the variation in the age of menarche is due to genetic factors. For instance, Kaprio et al.

[1] Berthelon, Kruger and Eberhard (2017) focus on Chile and had to limit their analysis to sisters who cohabitate as adults, which are not a random sample of the population.

[2] See Hotz, McElroy and Sanders (2005) for a study about the United States and Goodman et al (2004) for Britain. Azevedo, Lopez-Calva and Perova (2012) applied the same method to the case of Mexico.

(1995)'s landmark study shows that the correlation among monozygotic twins in the age of onset is nearly three times larger than the correlation observed in dizygotic twins. This has been further validated in recent work from a variety of contexts including developing countries (e.g., Jahanfar, Lye and Krishnarajah, 2013; Ameade and Garti 2016). This suggests a minimal role of environmental influences, including family background, on maturation.

We provide further evidence in favor of the validity of the exclusion restriction by showing that, in our sample, the onset of menarche is not related to adolescents' health status measured by height. Height has been consistently linked to early health status and nutrition and it is largely driven by prepubescent growth (e.g., Thomas and Strauss, 1997; Vogl, 2014). Thus, the null link between the age at menarche and this key indicator of health, allows us to rule malnutrition as a possible confounder for the instrument and the outcomes of interest.

Employing this identification strategy, we find that becoming a mother during adolescence has a positive effect on longer-term fertility and reduces school attainment. We also find that it increases marriage (though, less precisely estimated). However, when considering eight measures of labor supply –including different indicators of labor force participation, type of job and occupation– while also accounting for multiple hypothesis testing, we find that much of the negative effects of teen motherhood on labor supply is due to selection into motherhood.

This is a new and important result. Previous work by Agüero and Marks (2008, 2011) and Aaronson (2017) have shown that in countries with low and middle levels of economic development, either contemporaneously or historically, motherhood does not impose a penalty on the employment of women. Knowing that this pattern is also found for the *timing* of motherhood represents an important contribution to the literature. Thus, our findings weaken the claim that early motherhood, overall, leads to a path of low-quality employment or a misallocation of talent due to job sorting. However, heterogenous analysis reveal that early motherhood negatively impacts the labor supply of older women and as well as indigenous groups and minorities.

We also uncover a new mechanism that helps explain the weaker effect on labor supply. Women who were teenage mothers live in households where more adults are present and in particular, other adult women. We also show that this higher presence of other adult women provides the caring for the additional children that teen mothers have, allowing them to buffer the effect on labor supply.

2 | DATA

The main dataset for the analysis is the 2015-16 Colombian Demographic and Health Survey (CDHS). This is a nationwide household survey that collects information about women of reproductive age: 15-49 at the time of the survey.[3] The 2015-16 CDHS collects information on two sets of variables relevant for our study. First, as in nearly all the DHS datasets, women in the sample report the number of children, their marital status, schooling (years of schooling and levels of school attainment) and their labor market involvement. For the labor market we consider eight variables covering labor participation and type of employment and sector: currently working, worked in last 12 months, ever worked, works all year, paid worker, paid in cash, works in sales and works as a professional. We use all these variables to assess the impact of teenage motherhood. Second, and this is a unique feature of the 2015-16 CDHS, for all women in the sample the age at menarche is also recorded. This

[3]This DHS is part of a USAID global project collecting information on women and their children in over 65 countries. This dataset as well as the other are free to download from https://dhsprogram.com

variable permits us to instrument for teenage motherhood.[4]

Following Field and Ambrus (2008) and others, we restrict the sample to women aged 25-49 (born between 1965 and 1991). This lets us focus on longer-term effects and on women who were mainly done with their education to fully measure the impact on this outcome.[5] This restriction allows us to expand and complement the work by Arceo-Gómez and Campos-Vázquez (2014) on short term impacts. They applied an individual fixed-effect model combined with propensity score matching to a panel in Mexico and focus on childless adolescents at baseline who became mothers when observed three years later.[6]

We further restrict the sample to women whose onset of menarche started no earlier than 12. As shown in appendix Figures A.1 and A.2, this allows us to focus on the sample for which the instrument has a direct impact on the probability of teen motherhood. This pattern is consistent with the findings of Field and Ambrus (2008) for the case of Bangladesh. Our final sample has nearly 20,000 observations (N=19,891).[7] This sample size is an important expansion compared to previous work in the region.

Appendix Table A.1 summarizes the main features of the sample (column 1). More than 40% of the sample had a child before 20 and the average age at first birth is 20.8 for the full sample. The mean age at menarche is 13.5.[8] The average number of children is 2.4. Two-thirds of the sample works (at the time of the survey) and over 90% worked for pay. Most of the working women are employed in sales (58%) and 1 in 8 women work as professionals (jobs requiring a college degree). Regarding marital status and household composition, 69% of women are married or cohabitate, 20% live with at least one of their parents and in households with 1.3 other adult women (excluding the respondent).

This table also offers a preview of our main results. When we compare the mean values by teenage motherhood status (columns 2 and 3), we find significant differences in the pair-wise comparison for all the outcomes (at the 1%) level. For instance, teen mothers are less likely to work compared to their non-teen mother counterparts (0.63 vs. 0.68, respectively). Comparing these columns is analogous to an OLS estimation (without controls). However, when using age at menarche many of the differences disappear. Splitting the sample into those with a late menarche (at 14 or later, column 4) and women with an earlier menarche (column 5) is equivalent to a simplified reduced-form regression when using age at menarche as an instrument. The differences in employment probability are no longer relevant (0.67 vs. 0.66) and lack statistical significance. These findings are our first indication that the selection into teen motherhood is a mayor driver of the correlates between early childbearing and labor supply of women in our context.

In the next section we describe the methodology employed to estimate the effects of teen motherhood on the outcomes described above.

3 | METHODS

Our econometric model uses Y_i as the outcome of interest. This includes fertility education, marital status and employment variables. Our variable of interest is early motherhood. This

[4]No other DHS in Latin America includes this variable. A previous survey (2010 CDHS) collects age at menarche only for women aged between 10 and 17 as part of the anthropometric module. It is this survey (2010 CDHS) that allows us --as well as Urdinola and Ospino (2015)-- to test for the association between height and the onset of puberty. Unfortunately, there are no anthropometric measures recorded in the 2015-16 CDHS.

[5]According to the 2005 Colombian Population Census, approximately only 10% of women were still in school by the age of 25.

[6]For their longer-term impacts they used a different (cross-sectional) survey with propensity score matching.

[7]We also eliminated 36 observations where the age at menarche is recorded as happening after the age at first birth.

[8]See Appendix Figure A.3 for the distribution of this variable.

would be measure by T_i and is equal to one if individual i became a mother as a teenager (at 19 or before) and zero otherwise. As a robustness check, we will use age at first birth instead of the binary indicator. Also, A_i represents woman i's age at menarche and is the instrument used to identify the first-stage equation. Our 2SLS model is given by:

$$Y_i = \beta_0 + \beta_1 E\left[T_i|A_i\right] + \beta_i X_i + \varepsilon_i \quad (1)$$

$$T_i = \pi_0 + \pi_1 A_i + \pi_2 X_i + \upsilon_i \quad (2)$$

The set of controls (X_i) includes fixed effects by month and year of birth, ethnicity, region and urban location. It also contains interactions of region and urban/rural to account for additional geographical variation. Standard errors are clustered at the primary sample unit. Results are robust to clustering at the individual level.

The validity of our identification strategy relies on two important assumptions about the instrumental variable. First, the age a menarche should strongly alter the probability that a woman becomes a teenage mother. This is expected to hold due to biology, as conception requires the onset of puberty. As Figure 1 confirms, the probability of becoming a mother before age 20 decreases for women with a later onset of menarche. For instance, this probability goes from 47% when the age at menarche is 12 to just 8% when the onset of menarche occurs at 18 (or later). As a complement, we observe a positive association between age at menarche and the age at first birth (Figure 2).

Table 1 shows the regression analog of these graphs and displays the results of estimating Equation (2). In column (1), we find that each additional year delaying the onset of menarche is associated with a 12% decline (=-0.048/0.413) in the probability of being a teen mother. This is significant at the 1% level and unaffected by the inclusion of controls (column 2). In column (3), we consider a more flexible specification, using age-specific fixed effects (plus controls). Again, we observe that a later onset of the menarche decreases teen motherhood. Thus, in the paper we consider two 2SLS models, as a way to show the robustness of our identification strategy. In one, we employ age at menarche as an instrument leading to a just-identified model. Second, we use age-specific dummies for the onset of the menarche as instruments. This constitutes our over-identified models. In both cases, our first stage is very strong. The F-statistic for the first model is 374.7 and 98.3 for the over-identified model. When the variable of interest if the age at first birth, the corresponding F-statistics are 327.6 and 64.4, respectively.

The second assumption to validate our identification strategy refers to the exclusion restriction. This requires that the age at menarche affects the outcomes of interest only through its effect on the probability of motherhood before age 20 (or age at first birth). While a formal test is not possible, we provide many arguments in favor of this assumption.

As mentioned in the introduction, Field and Ambrus (2008) conducted an extensive review of the medical literature and concluded that most of the variation in the age of menarche is due to genetic factors. More recent work has continued to validate this claim (e.g., Jahanfar, Lye and Krishnarajah, 2013; Szwed et al., 2013; Ameade and Garti, 2016). The landmark study by Kaprio et al. (1995) provides a clear example. This study shows that the correlation among monozygotic twins in the age of onset is nearly three times larger than the correlation observed in dizygotic twins. This implies a very minimal role of environmental influences, including family background, on maturation.

Field and Ambrus (2008) also identify additional possible factors known from field and laboratory experiments that could affect menarche: geography and climate; strenuous physical activity or stress, and the sex composition of peer group. We do not expect any of these factors to confound our analysis for several reason. First, accounting for geographic and climate fixed effects in our estimates did not change our findings (comparing columns 1 and 2 in Table 1). Second, while hard labor in developing countries is associated with poverty, in Colombia time use surveys show that adolescent girls rarely engage in work outside the home and instead tend to allocate their time towards home production (Torres and Agüero, 2017). Finally, there is no reason to suspect that sex composition of peer groups is distributed across individuals in a way that correlates with adult outcomes. However, in developing countries there could be room for additional factors due to, for example, acute malnutrition.

To our advantage, malnutrition in Colombia is lower than in other Andean and Latin American countries.[9] For example, the proportion of adult women shorter than 145cm, a common threshold for short women, is just 3.5 percent in Colombia, compared to 9.5 percent in Bolivia and Peru. A pattern is found in child stunting among these countries. Furthermore, Colombia's nutritional status is much better than what is observed the poorest countries in the region, such as Haiti and Guatemala. Nonetheless, these smaller malnutrition rates could still represent a problem for our analysis if differences in nutrition within our sample are large enough to exert an influence on the endocrine systems that regulate menarche.[10]

We provide evidence that rejects this possibility by exploring the relationship between menarche and height. Height has been consistently linked to early health status and nutrition and it is largely driven by prepubescent growth (e.g., Thomas and Strauss, 1997; Vogl, 2014). Using data from the 2010 CDHS, Urdinola and Ospino (2015) find that in a regression of height on age at menarche, the estimated coefficient is very close to zero and statistically insignificant. We conduct a similar analysis and report the results in Table 2.

We find no association between a delay in the age of menarche and height. Regressing height against age a menarche, an including controls for altitude of the area as well as district, race and age fixed effects, results in an estimate of 0.05, which represents a 0.03% of the mean. This tiny effect is not statistically different from zero (column 1). In column (2), we consider a more flexible specification and include age-specific dummies for the onset of menarche. Again, these estimates are neither economically nor statistically significant. Furthermore, we cannot reject the null hypothesis that all these dummies are different from zero (p-value: 0.242).

In Figure 3, we present additional evidence that reinforces this null effect. Using the same dataset, we estimate the kernel density of height for multiple menarcheal ages (11-17). This exercise reveals that the population distributions, in addition to the averages, are remarkably similar across all subsamples. This finding is consistent with the work of Field and Ambrus (2008) where the density estimates also overlap in the case of Bangladesh. Altogether, the evidence argues in favor of the validity of using age of menarche as an instrument for teen motherhood.[11]

To estimate the impact of early childbearing on a set of indicators for schooling and labor market outcomes, we employ two ways to account for the multiple number of outcomes

[9] This feature should be heavily considered when using this instrument in settings where malnutrition is acute, as in some areas of Sub-Sharan Africa. See Burger et al (2019) for a discussion.

[10] See Field (2007) for an extensive review of the links between nutrition and menarche based on the medical literature.

[11] An alternative approach for the case of Colombia could come from the availability of family planning programs using data from the expansion of *Profamilia* reported by Miller (2010). However, two features impede us from using such approach. First, the CDHS does not record information about place of residence during adolescence (or at birth). Second, Miller focuses on older women in order to have a both a treatment and a control group. Women included in the 2015-16 CDHS are too young so all the groups are treated as adolescents using Miller's design, preventing us from using that additional test. However, as described above, we added an alternative robustness check by using age-specific dummies for the onset of the menarche. That allows us to have two 2SLS models: one that is just-identified and another one that is over-identified.

considered. First, we create indices for each family of outcomes following Kling, Liebman, and Katz (2007). These indices combine all the variables in each family of outcomes. To construct the indices, we define each outcome so that higher values correspond with better outcomes. Given that almost all our variables are binary, we set them equal to one to signal a better outcome. Then we standardize each outcome into a Z-score by subtracting the mean and dividing by the standard deviation. After that, we average all the Z-scores and estimate the effect of teen motherhood on these standardized outcome indices.

Second, we correct for the potential issue of simultaneous inference using multiple hypothesis testing. Based on Benjamini and Hochberg (1995), we apply the concept of a false discovery rate (FDR) to allow inference when conducting many tests. Intuitively, FDR allows a researcher to tolerate a certain number of tests to be incorrectly discovered. An FDR adjusted q- value of 0.05 implies that 5 percent of *significant tests* result in false positives compared with an unadjusted p-value of 0.05 that implies 5 percent of *all tests* result in false positives. In the regression tables, we show standard errors based on unadjusted p-values as well as FDR adjusted q-values that address the multiple hypotheses being tested in a given family of outcomes. The results of applying this methodology are described in the next section.

4 | RESULTS

We start by exploring the effect of teen motherhood on the total number of children. One possibility is that early motherhood accelerates the arrival of the first born without affecting the complete fertility rate. Our results reject this possibility (see Table 3). We find that teenage motherhood is associated with a higher number of children, after controlling for age (and the other covariates described above). This is observed even when instrumenting for early motherhood as in columns (2) and (3) of Table 3, for the just-identified and the over-identified models, respectively, where teen mothers have between 1.2-1.5 more children.

In Table 4 we focus on the schooling effects. Again, we find that early motherhood has a detrimental impact on schooling. For example, after instrumenting, we find that teen mothers have 1.72 fewer years of schooling, representing a decline of 19% with respect to the mean. This impact seems to be driven by a reduction in dropping out of school during adolescence. These effects are robust to the adjustment for simultaneous inference using multiple hypothesis testing (shown in brackets) and when using an index that includes all schooling variables. For the index, we estimate a decline of around 40 percent of a standard deviation.

Turning our attention to the marriage market, Table 5, we find that some of the impacts are driven by selection: the 2SLS estimates tend to be smaller than the OLS. However, the 2SLS parameters are less precise, as expected.

These three findings, an increase in the number of children, a decrease in schooling and a possible increase in marriage, are consistent with previous findings in Colombia and the Latin American region. For example, Miller (2010) focuses on an older cohort (born between 1911 and 1955) and shows that the introduction of family planning programs in Colombia is associated with an increasing in schooling, a reduction in fertility and marriage. Arceo-Gómez and Campos-Vázquez (2014), use a fixed-effects model in Mexico and find that teen motherhood has similar impacts in the short-run, though they report a smaller effect on schooling (0.6-0.8 years of schooling). Thus, a consensus seems to emerge about the impact of teen motherhood on schooling (negative), on fertility (positive) and marriage (also positive, smaller, but less precise).

We now explore the impacts on labor supply. Table 6 shows the estimated effects when considering eight different indicators of employment. Comparing OLS estimates (column 1) against the 2SLS (columns 2 and 3) suggests that much of the negative effect on labor supply attributed to teen motherhood is due to selection. The 2SLS estimates are a fraction of the OLS even when including the index that summarizes all eight indicators. Overall, these estimates reveal a very small effect, if not a null impact, on labor supply.

This pattern is consistent with the results found for advanced countries reviewed in the introduction. This is also aligned with previous work on the role of children on their mothers' employment in lower-income countries. Agüero and Marks (2008, 2011) and Aaronson (2017) have shown that for low- and middle-income countries, either contemporaneously or historically, motherhood does not impose a penalty on the employment of women. Knowing that this pattern is also found for the timing of motherhood represents an important contribution to the literature.

Furthermore, Agüero and Marks (2011) find that children negatively affect the *quality* of the job their mothers pursue. On Table 6, when considering whether a woman is paid (at all) or paid in cash only, we find that the OLS and 2SLS are not very different (but as expected, the latter are less precise). However, a different conclusion is obtained when considering occupations. While the OLS results continue to indicate a possible job sorting, with early motherhood increasing the probability of working in sales and reducing the chances of working in professional occupations, the 2SLS reveal very clear null impacts. Thus, we find very weak evidence that early childbearing affects the quality of the jobs these mothers pursue.

All these previous findings for fertility, education, marriage market and labor supply are confirmed when estimating the impact of the age at first birth as an alternative measure for early motherhood. These findings are summarized in Table 7. Again, delaying having a child is positively associated with higher fertility and lower schooling. These findings are not driven by selection as shown by the 2SLS in columns (2) and (3). For the marriage market, the 2SLS have a similar point estimate than the OLS but are less precise. As before, the impacts on the labor supply seem to be driven by selection and our 2SLS suggest much modest impacts that are not statistically different from zero. Thus, there is much weaker evidence that there is a true tradeoff between the reproductive and productive roles of young women.

We now explore heterogenous effects. We focus on race and age as two variables that are available in the data and cannot be affected by early childbearing themselves.[12] These exercises are reported in Table 8. We start by exploring differences by age cohorts and divide the sample into older (born between 1965-1999) and younger women (born between 1980 and 1991) using the median age as the cutoff. Panel A (older women) and B (younger group) show that the effects for the overall sample in terms on fertility and education remain in these two sub-groups. However, we observe a negative and larger impact on the labor market index for older women. When dividing the sample by ethnicity/race, the impacts are larger for indigenous and other minorities groups (although less precisely estimated) while negligible for the rest of the population (Panels C and D).

Finally, we provide new evidence of a mechanism explaining the weaker effect of early childbearing on labor market outcomes, despite the increase in fertility and the reduction in human capital related outcomes. Previous work has postulated that the prevalence of an informal economy could explain the limited effects of motherhood on labor supply in developing countries and in particular, in Latin America (e.g., Miller, 2009; Berniell et al, 2019). They argue that an informal economy provides a flexible work environment for women that have more children. Here we consider an alternative, but

[12] The 2015-16 Colombian DHS does not contain socio-economic variables about the parents of the women interviewed limiting the variables that could be considered for this analysis.

complementary, hypothesis: women who became mothers during adolescence have different living arrangements that facilitate child rearing buffering the employment impact of early motherhood. Table 9 provides evidence in favor of this novel mechanism. Teen mothers are more likely to live with other adult women (Panel A, columns 1-4), with an overall increase of 15 percent of standard deviation in the index. The results are more precise when considering the effect of age at first birth as the variable of interest (see Panel B). In Table 10, we provide evidence that these living arrangements facilitate the caring of children teen mothers have. The DHS asked women with young children about who tends for them. Limiting the analysis to this smaller sample, we show that teen mothers are more likely to have another person take care of their children (column 1, Panels A and B), relative to women who become mothers later. This help in childcaring responsibilities buffers the impact of early motherhood on labor market outcomes.

5 | CONCLUSIONS

We present a new methodology to study the longer-term effect of teen motherhood on the schooling and labor supply outcomes of women in the context of Latin America. We focus on the case of Colombia, where the availability of a unique dataset, allows us to account for selection into teen motherhood by instrumenting using age at menarche. We provided strong evidence about the validity of this instrument, ruling out possible violations of the exclusion restrictions.

Our findings on fertility (positive), education (negative) and marital status (positive, but less precise) are aligned with previous work on the region, even though these studies have used alternatives methodologies and, in some cases, centered on shorter terms.

Yet, in terms of the effects on labor supply, we conclude that much (if not all) of the negative effect on labor supply attributed to teen motherhood is due to selection. We find weak evidence that early motherhood leads to job sorting when considering type of job (paid and paid in cash) and a null effect when exploring sorting into different occupations. Thus, we do not find enough evidence to conclude that there is a true trade off between the reproductive and productive roles of young women. While we find negative effects for older women and racial minorities, overall our findings do not support the claim that early motherhood leads to a path of low-quality employment or a misallocation of talent.

We uncover a novel mechanism for this limited effect on labor supply. We find strong evidence showing that household composition differs for women who had an early motherhood as they now live in homes where there are more adults and in particular more adult women. These adult women buffer the effect of teen motherhood by providing the caring for the additional children and we present evidence in favor of this argument. Thus, while there is weak evidence that becoming a teen mother has a negative causal effect on the future labor outcomes of these mothers, as documented in our study, it is possible that the altered living arrangements limit the labor supply of the co-residing women who help with the caring responsibilities of the extra children. This is an interesting line of research that should be further explored.

6 | TABLES

TABLE 1 First stage: age at menarche and teen motherhood

	Dependent variable: Mother before age 20 (=1)		
	(1)	(2)	(3)
Age at menarche	-0.048***	-0.048***	
	(0.003)	(0.003)	
Age at menarche:			
13			-0.022**
			(0,009)
14			-0.044***
			(0,010)
15			-0.129***
			(0,012)
16			-0.198***
			(0,017)
17			-0.257***
			(0,023)
18			-0.394***
			(0,021)
Observations	19.891	19.891	19.891
R-squared	0.016	0.044	0.046
Controls	N	Y	Y
Mean dep. var.	0.413	0.413	0.413

Notes: Robust standard errors in parentheses clustered at the primary sampling unit. Controls include controls for fixed effect in month and year of birth, region, urban and ethnicity. For column (3), the omitted category is menarche at age 12. Data source: 2015 Colombian DHS. *** p<0.01, ** p<0.05, * p<0.1

TABLE 2 Validity of the instrument: age at menarche and height

	Dependent variable: height (cm.)	
	(1)	(2)
Age at menarche	0.0507	
	(0.080)	
Age at menarche:		
12		-0.126
		(0.168)
13		0.129
		(0.208)
14		0.166
		(0.245)
+15		-0.227
		(0.456)
Observations	8.212	8.212
R-squared	0.171	0.172
Joint F-test		1.406
Prob > F		0.242

Notes: Robust standard errors in parentheses clustered at the survey sampling area. All regressions include controls for altitude of the area as well as district, race and age fixed effects. In column (2), the omitted category is menarche at age 11. The null hypothesis for the joint F-test is that all coefficients of age at menarche are zero. *** p<0.01, ** p<0.05, * p<0.1. *Source:* 2010 Colombia DHS.

TABLE 3 Teen motherhood and number of children

Dependent variable: Number of children			
	OLS	2SLS	
		just-identified	over-identified
	(1)	(2)	(3)
Teenage motherhood	1.461***	1.219***	1.473***
	(0.021)	(0.160)	(0.151)
Observations	19.891	19.891	19.891
R-squared	0.385	0.381	
Mean dep. var.	2.427	2.427	2.427

Notes: Robust standard errors in parentheses clustered at the survey sampling area. All regressions include controls for fixed effect in month and year of birth, region, urban and ethnicity. In column (2), teen motherhood is instrument with the age at menarche (F-stat: 374.7). In column (3) the instruments are fixed effects for the age at menarche (F-stat: 98.3), where the omitted category is menarche at age 11. Stars denote statistical significance at 1, 5, and 10% levels based on unadjusted p-values. *Source:* 2015-16 Colombian DHS.

TABLE 4 Teenage motherhood and schooling

Dependent variable	OLS	2SLS		Obs.	Mean of dep. var.
		just-identified	over-identified		
	(1)	(2)	(3)		
Years of schooling	-2.510***	-1.724***	-1.656***	19,891	9.093
	(0.055)	(0.424)	(0.400)		
	[0.00000]	[0.00005]	[0.00004]		
Primary or more	-0.122***	-0.250***	-0.198***	19,891	0.839
	(0.005)	(0.039)	(0.036)		
	[0.00000]	[0.00000]	[0.00000]		
Secondary or more	-0.278***	-0.189***	-0.189***	19,891	0.541
	(0.007)	(0.050)	(0.049)		
	[0.00000]	[0.00018]	[0.00010]		
Tertiary	-0.236***	-0.006	-0.058	19,891	0.298
	(0.006)	(0.049)	(0.048)		
	[0.00000]	[0.90720]	[0.23363]		
Schooling index	-0.530***	-0.409***	-0.391***	19,891	
	(0.012)	(0.088)	(0.083)		

Notes: In columns (1)-(3) each cell shows the coefficient of regressing each outcome on teenage motherhood. All regressions include controls for fixed effect in month and year of birth, region, urban and ethnicity. In column (2), teen motherhood is instrument with the age at menarche (F-stat: 374.7). In column (3) the instruments are fixed effects for the age at menarche (F-stat: 98.3), where the omitted category is menarche at age 11. Robust standard errors in parentheses clustered at the survey sampling area. Stars denote statistical significance at 1, 5, and 10% levels based on unadjusted p-values. FDR q-values are computed over all four outcomes (within a column) and are shown in square brackets. FDR q- values indicate the probability of false positives among significant tests. *Source:* 2015-16 Colombian DHS.

TABLE 5 Teen motherhood and marital status

Dependent variable: Currently married or cohabiting (=1)			
	OLS	2SLS	
		Just-identified	Over-identified
	(1)	(2)	(3)
Teenage motherhood	0.083***	0.044	0.076
	(0.007)	(0.052)	(0.050)
Observations	19.891	19.891	19.891
R-squared	0.041	0.039	
Mean dep. var.	0.685	0.685	0.685

Notes: Robust standard errors in parentheses clustered at the survey sampling area. All regressions include controls for fixed effect in month and year of birth, region, urban and ethnicity. In column (2), teen motherhood is instrument with the age at menarche (F-stat: 374.7). In column (3) the instruments are fixed effects for the age at menarche (F-stat: 98.3), where the omitted category is menarche at age 11. Stars denote statistical significance at 1, 5, and 10% levels based on unadjusted p-values. *Source:* 2015-16 Colombian DHS.

TABLE 6 Teenage motherhood and labor market outcomes

Dependent variable	OLS (1)	2SLS Just-identified (2)	2SLS Over-identified (3)	Obs.	Mean of dep.var.
Currently works	-0.042***	0,000	0,012	19.891	0,66
	(0,007)	(0,052)	(0,050)		
	[0.00000]	[0.99619]	[0.80678]		
Worked last 12 months	-0.023***	-0,062	-0,052	19.891	0,774
	(0,006)	(0,046)	(0,043)		
	[0.00021]	[0.18398]	[0.22810]		
Ever worked	-0.015***	-0.051*	-0,035	19.891	0,93
	(0,004)	(0,028)	(0,026)		
	[0.00007]	[0.06799]	[0.18243]		
Works all year	-0.055***	-0,038	-0,029	19.891	0,644
	(0,007)	(0,053)	(0,050)		
	[0.00000]	[0.47461]	[0.55772]		
Paid worker	-0.020***	-0,046	-0,046	19.891	0,904
	(0,004)	(0,032)	(0,029)		
	[0.00000]	[0.15388]	[0.11694]		
Paid in cash	-0.032***	-0,023	-0,045	19.891	0,853
	(0,005)	(0,039)	(0,036)		
	[0.00000]	[0.54327]	[0.20378]		
Works in sales	0.128***	0,006	0,025	19.891	0,576
	(0,007)	(0,055)	(0,052)		
	[0.00000]	[0.91539]	[0.62822]		
Works as professional	-0.111***	0,023	0,000	19.891	0,125
	(0,005)	(0,036)	(0,036)		
	[0.00000]	[0.53460]	[0.99799]		
Labor market index	-0.111***	-0,060	-0,059	19.891	
	(0,008)	(0,061)	(0,057)		

Notes: In columns (1)-(3) each cell shows the coefficient of regressing each outcome on teenage motherhood. All regressions include controls for fixed effect in month and year of birth, region, urban and ethnicity. In column (2), teen motherhood is instrument with the age at menarche (F-stat: 374.7). In column (3) the instruments are fixed effects for the age at menarche (F-stat: 98.3), where the omitted category is menarche at age 11. Robust standard errors in parentheses clustered at the survey sampling area. Stars denote statistical significance at 1, 5, and 10% levels based on unadjusted p-values. FDR q-values are computed over all eight outcomes (within a column) and are shown in square brackets. FDR q- values indicate the probability of false positives among significant tests. *Source:* 2015-16 Colombian DHS.

TABLE 7 The effect of age at first birth

Effect of age at first birth by method					
	OLS	2SLS			
Dependent variable:		Just-identified	Over-identified	Obs.	Mean of dep.var.
	(1)	(2)	(3)		
Number of children	-0.165***	-0.130***	-0.161***	17,934	2,692
	(0,002)	(0,017)	(0,016)		
Schooling index	0.065***	0.047***	0.046***	17,934	
	(0,001)	(0,010)	(0,010)		
Married or cohabitating	-0.004***	-0,005	-0,007	17,934	0,724
	(0,001)	(0,006)	(0,006)		
Labor market index	0.014***	0,008	0,008	17,934	
	(0,001)	(0,007)	(0,007)		

Notes: In columns (1)-(3) each cell shows the coefficient of regressing each outcome on teenage motherhood. All regressions include controls for fixed effect in month and year of birth, region, urban and ethnicity. In column (2), age at first birth is instrument with the age at menarche (F-stat: 327.6). In column (3) the instruments are fixed effects for the age at menarche (F-stat: 64.4), where the omitted category is menarche at age 11. Robust standard errors in parentheses clustered at the survey sampling area. Stars denote statistical significance at 1, 5, and 10% levels based on unadjusted p-values. *Source:* 2015-16 Colombian DHS.

TABLE 8 Heterogenous effects (2SLS)

	Dependent variable:		
	Number of children	Education index	Work index
	(1)	(2)	(3)
Panel A: Older women (born before 1979, N=9940, F-stat=229)			
Teenage motherhood	1.051***	-0.364***	-0.199**
	(0.244)	(0.121)	(0.078)
Panel B: Younger women (born in 1979 or later, N=9951, F-stat=152)			
Teenage motherhood	1.426***	-0.481***	0.102
	(0.179)	(0.123)	(0.096)
Panel C: Indigenous and other minorities (N=4419, F-stat=93.19)			
Teenage motherhood	1.426***	-0.651***	-0.206*
	(0.371)	(0.178)	(0.117)
Panel D: Other ethnic groups (N=15472, F-stat=286)			
Teenage motherhood	1.161***	-0.346***	-0.03
	(0.169)	(0.101)	(0.069)

Notes: Each cell shows the coefficient of regressing each outcome on teenage motherhood. All regressions include controls for fixed effect in month and year of birth, region, urban and ethnicity. Teenage motherhood is instrument with the age at menarche (just-identified model). F-stat refers to the first stage F-statistic. Robust standard errors in parentheses clustered at the survey sampling area. Stars denote statistical significance at 1, 5, and 10% levels based on unadjusted p-values. *Source:* 2015-16 Colombian DHS.

TABLE 9 Effect on living arrangements (2SLS)

	Number of adults in the household		Lives with (=1)				Index
	All	Women	Any adult	Women	Mother	Father	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Effect of teen motherhood (N=19,891)							
Teen motherhood	0.249*	0,147	0.066*	0.095*	0,025	0,014	0.149*
	(0,150)	(0,099)	(0,035)	(0,055)	(0,042)	(0,033)	(0,083)
	[0.09652]	[0.13906]	[0.05702]	[0.08142]	[0.55401]	[0.67820]	
Panel B: Effect of age at first birth (N=17,934)							
Age at first birth	-0.044***	-0.019*	-0.008**	-0.010*	-0,001	-0,004	-0.020**
	(0,017)	(0,010)	(0,004)	(0,006)	(0,004)	(0,003)	(0,009)
	[0.00845]	[0.07446]	[0.04172]	[0.08708]	[0.88847]	[0.21009]	
Mean of dep. variable	1,72	0,570	0,898	0,403	0,180	0,095	

Notes: Each cell shows the coefficient of regressing each outcome on teenage motherhood (panel A) or age at first birth (panel B) using a 2SLS model. All regressions include controls for fixed effect in month and year of birth, region, urban and ethnicity. Robust standard errors in parentheses clustered at the survey sampling area. Stars denote statistical significance at 1, 5, and 10% levels based on unadjusted p-values. FDR q-values are computed over all outcomes (within a row) and are shown in square brackets. FDR q- values indicate the probability of false positives among significant tests.

TABLE 10 Effect on childcare responsabilities (2SLS)

	Not the mother	Grandparent	Other family member	Any family member	Index
	(1)	(2)	(3)	(4)	(5)
Panel A: Effect of teen motherhood (N=5,574)					
Teen mother	0.160*	0,029	0,018	0,047	0,155
(F-stat: 89.19)	(0,096)	(0,110)	(0,067)	(0,114)	(0,163)
	[0.09741]	[0.79017]	[0.79275]	[0.68086]	
Panel B: Effect of age at first birth (N=5,574)					
Age at first birth	-0.018*	-0,003	-0,002	-0,005	-0,018
(F-stat: 67.99)	(0,011)	(0,013)	(0,008)	(0,013)	(0,019)
	[0.09822]	[0.79046]	[0.79288]	[0.68167]	
Mean of dep. Variable	0,805	0,34	0,093	0,433	

Notes: Each cell shows the coefficient of regressing each outcome on teenage motherhood (panel A) or age at first birth (panel B) using a 2SLS model. All regressions include controls for fixed effect in month and year of birth, region, urban and ethnicity. Robust standard errors in parentheses clustered at the survey sampling area. Stars denote statistical significance at 1, 5, and 10% levels based on unadjusted p-values. FDR q-values are computed over all outcomes (within a row) and are shown in square brackets. FDR q- values indicate the probability of false positives among significant tests.

7 | FIGURES

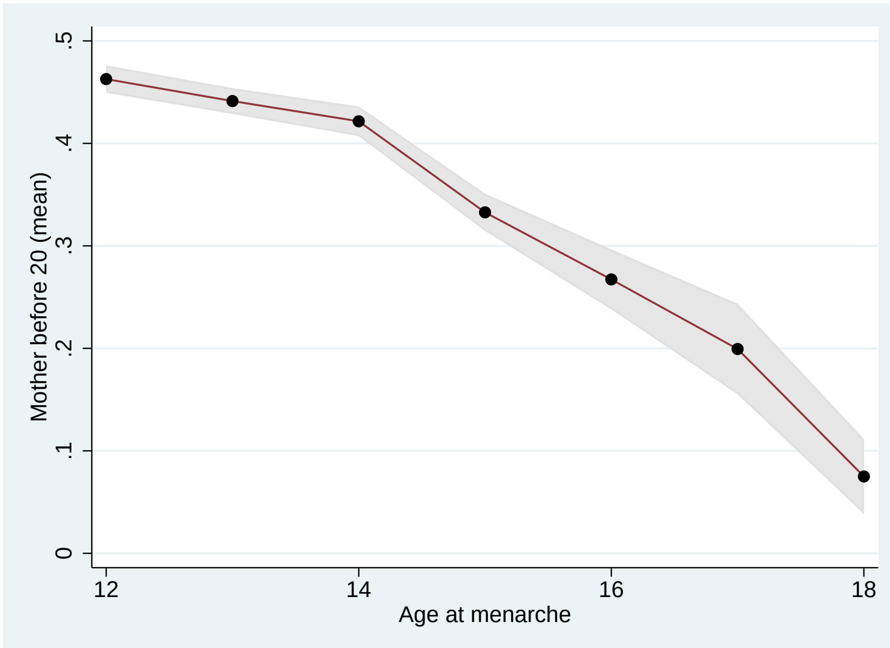

FIGURE 1 Age at menarche and teen motherhood. *Notes:* The sample includes women between 25 and 49 who reached menarche between 12 and 18. For less than 0.01 % of women, menarche arrived at age 19 or later. They are grouped in the 18 category. Shaded areas reflect the 95% confidence interval. *Source:* 2015-16 Colombian DHS.

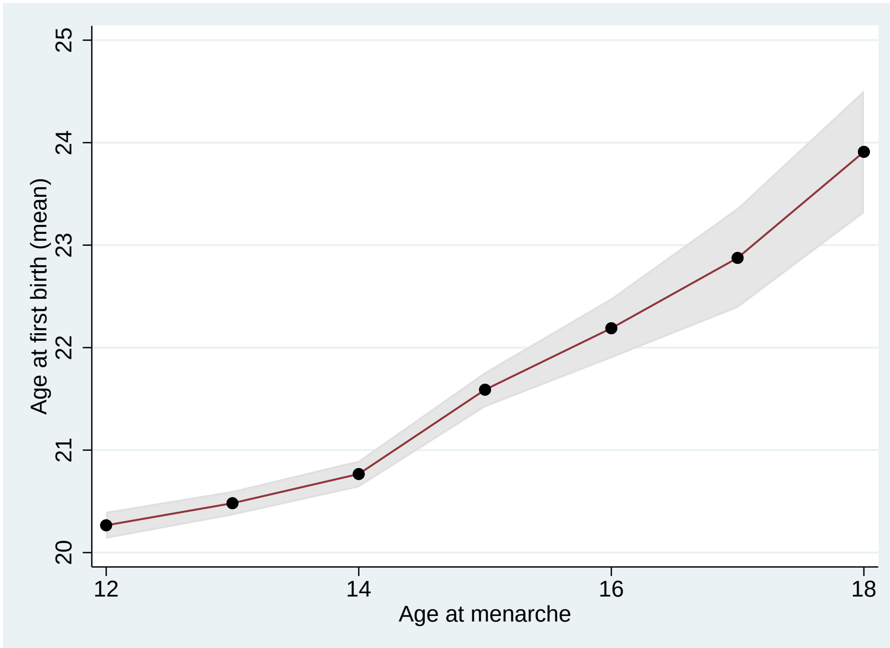

FIGURE 2 Age at menarche and age at first birth. *Notes:* The sample includes women between 25 and 49 who reached menarche between 12 and 18. For less than 0.01% of women, menarche arrived at age 19 or later. They are grouped in the 18 category. Shaded areas reflect the 95% confidence interval. *Source:* 2015-16 Colombian DHS.

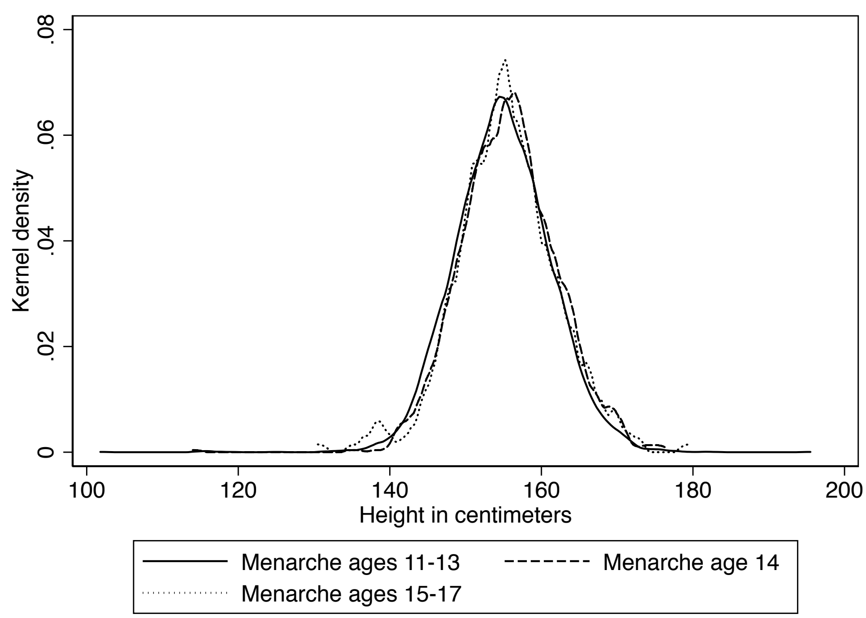

FIGURE 3 Height distributions by age at menarche. *Notes:* The sample includes women between 11 and 17 who reached menarche between 11 and 17. Epanechnikov kernel, bandwidth=0.8. *Source:* 2010 Colombian DHS.

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8 | APPENDIX

| Tables

TABLE A.1 Summary statistics

Variable	All women (1)	Teen mother No (2)	Teen mother Yes (3)		Early menarche No (4)	Early menarche Yes (5)	
Sample size	19.891	11.673	8.218		8.493	11.398	
Age at menarche	13,46	13,61	13,26	*	14,72	12,53	*
Age at first birth	20,78	23,82	17,18	*	21,31	20,38	*
Teen mother	0,413	0	1		0,362	0,451	*
Number of children	2,427	1,803	3,314	*	2,447	2,413	
Currently works	0,66	0,683	0,628	*	0,666	0,655	
Worked last 12m	0,774	0,789	0,754	*	0,78	0,77	
Ever worked	0,93	0,94	0,916	*	0,937	0,925	+
Works all year	0,644	0,676	0,599	*	0,648	0,641	
Paid worker	0,904	0,919	0,883	*	0,908	0,901	
Paid in cash	0,853	0,874	0,824	*	0,853	0,854	
Works in sales	0,576	0,534	0,636	*	0,575	0,577	
Works as professional	0,125	0,172	0,057	*	0,119	0,129	°
Years of schooling	9,093	10,25	7,45	*	9,008	9,156	°
Primary or more	0,839	0,899	0,754	*	0,843	0,835	
High-school or more	0,541	0,667	0,362	*	0,532	0,548	°
Tertiary	0,298	0,402	0,15	*	0,282	0,309	*
Married or cohabitating	0,685	0,644	0,744	*	0,683	0,687	
Lives with mother	0,18	0,226	0,115	*	0,17	0,188	*
Lives with father	0,095	0,121	0,057	*	0,09	0,099	°
Lives with a parent	0,2	0,248	0,132	*	0,187	0,21	*
Adults in household	4,288	4,367	4,175	*	4,257	4,311	
Adult women	3,288	3,368	3,175	*	3,257	3,312	
Share of adult women	0,723	0,726	0,718	*	0,721	0,724	

Notes: The sample includes women between 25 and 49 who reached menarche between 12 and 18. Early menarche is defined as an onset before age 14. Statistical differences between columns (2) and (3), as well as (4) and (5) at the 10%, 5%, and 1% are denoted by °, +, and *, respectively.

Figures

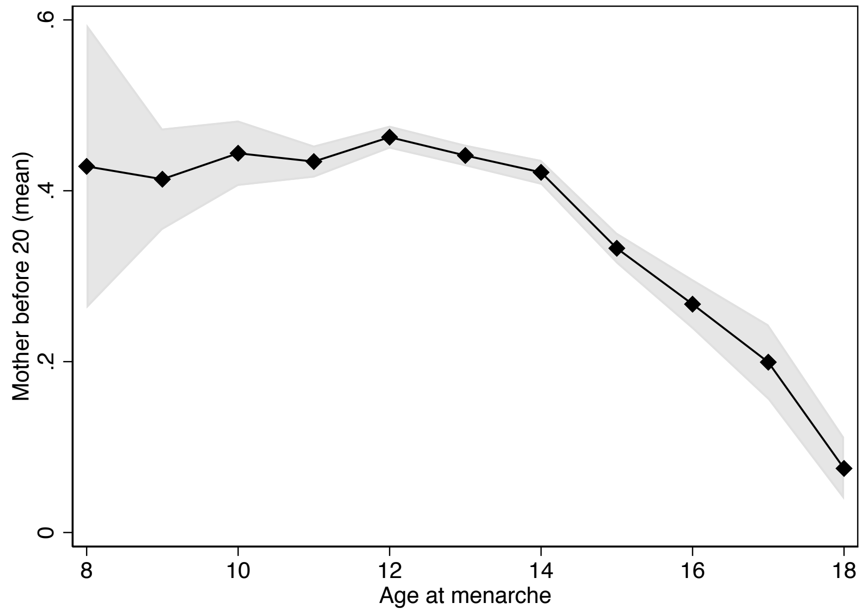

FIGURE A.1 Age at menarche and teen motherhood: full sample. *Notes:* The sample includes women between 25 and 49. Shaded areas reflect the 95% confidence interval. *Source:* 2015-16 Colombian DHS.

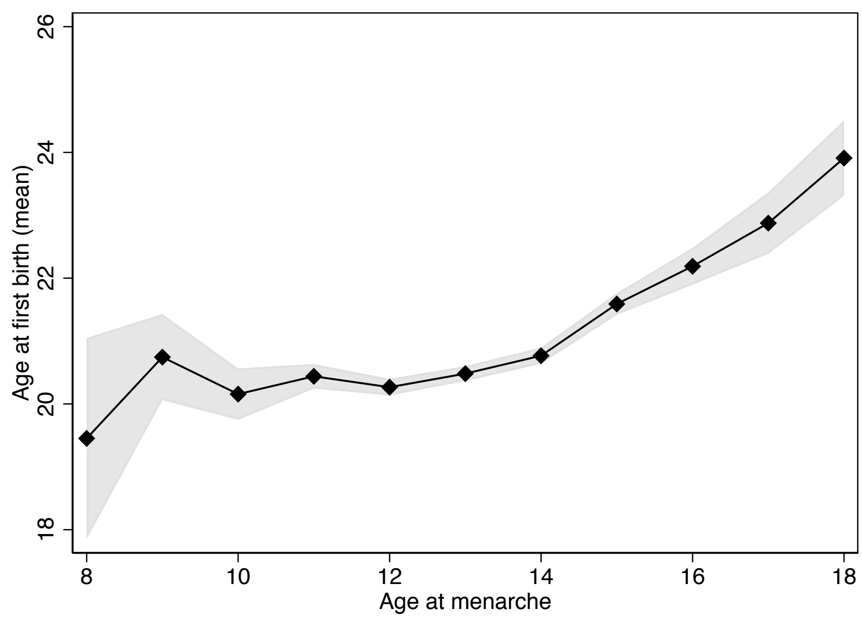

FIGURE A.2 Age at menarche and age at first birth (full sample). *Notes:* The sample includes women between 25 and 49. Shaded areas reflect the 95% confidence interval. *Source:* 2015-16 Colombian DHS.

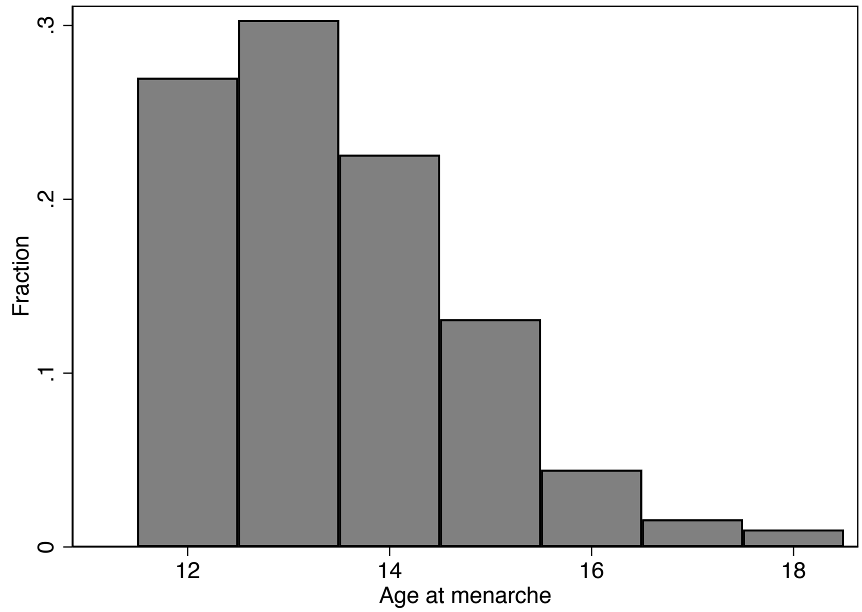

FIGURE A.3 Distribution of age at menarche. *Notes:* The sample includes women between 25 and 49 who reached menarche between 12 and 18. For less than 0.01% of women, menarche arrived at age 19 or later. They are grouped in the 18 category. *Source:* 2015-16 Colombian DHS.