Like Teenage Mother, Like Daughter? Black-White Differences in Maternal Age Effects on Teenage Childbearing

Abstract

Black women are at much higher risk of teenage childbearing. In this paper, we examine the extent to which this higher risk can be attributed to black-white differences in maternal age distributions and effects. Using a high-quality administrative dataset on 35,544 first births in North Carolina, we find that the increased risk is primarily due to weaker maternal age effects, so that black women born to adult mothers are at much higher risk than their white counterparts. (White and black women born to teenage mothers, on the other hand, are at similar risk of teenage childbearing.) We apply an order-invariant decomposition method and show that two-thirds of the black-white gap in maternal age effects can be explained by weaker associations between paternal characteristics and teenage fertility outcomes among blacks.

While previous work has focused on the role of intergenerational transmission, our analysis suggests that successive generations of black women continue to face higher risks of falling into disadvantage even after one generation escapes it. Our evidence also highlights the potential impacts of weaker marital prospects and expectations among black women on teenage fertility behavior, through reducing family resources at younger ages and reducing gains to postponing childbearing at older ages.

Introduction

The U.S. teenage birth rate is at a historic low, falling from 61.8 births per 1000 women aged 15 to 19 in 1991 to 29.4 in 2012. The decline has been especially sharp for black teenage birth rates, which fell by 60% over this time period (from 114.8 to 44.0, vs. from 52.6 to 27.4 for whites) (Martin et al. 2013). Nevertheless, the risk of teenage childbearing remains substantially higher for black women than for their white counterparts, and the U.S. teenage birth rate remains high compared to those in other developed nations, which range from 6 in France and 5 in Japan, to 14 in Canada and 26 in the U.K. (World Bank 2014).

There are multiple causes of the black-white gap in teenage birth rates, including demographic, economic and social differences. In this paper, we focus on two demographic differences: lower maternal ages at first birth among blacks, and black-white differences in the effects of maternal age on teenage fertility outcomes. Black women tend to be younger than white women at the time of first birth, with an average age of 22.7 in 2006 compared to 26.0 for white women (Matthews and Hamilton 2009). There are a number of reasons why having lower maternal ages at first birth may increase black women's risk of teenage childbearing. First, women who were younger at first birth may complete less education and create poorer family environments for learning, reducing their daughters' educational expectations and opportunity costs of early childbearing (Barber 2001; Manlove 1997). Second, younger mothers are less likely to be married and more likely to divorce and experience multiple partnership changes, exposing their daughters to financial hardship and emotional stress (Barber 2001; Manlove 1997; McLanahan 2009). Moreover, mothers who were themselves unmarried at the time of birth may be less likely to emphasize traditional values and be more accepting of early sexual activity and non-marital childbearing (Abrahamse et al. 1988; Barber 2001; Kahn and Anderson 1992), lowering their daughters' expectations and preferences for marriage before childbearing. Third, women who begin childbearing at a younger age tend to have higher completed fertility, which may increase their daughters' ideal family size and reduce their ideal age at first birth (Barber 2001; Kahn and Anderson 1992; Manlove 1997). Fourth, younger mothers may not only provide poorer family environments for learning, but also be more likely to live in poorer neighborhoods and to send their daughters to worse schools with more negative peer influences (Kahn and Anderson 1992; Manlove 1997) , further reducing their likelihood of being less engaged and motivated in school (Manlove 1997; Manlove 1998).

Understanding the process by which women born to younger mothers face higher risk of teenage motherhood is valuable for gaining insight into population dynamics and the intergenerational transmission of disadvantage. However, there have been relatively few studies of this process, as not many datasets contain information on multigenerational fertility outcomes. Most of the previously used datasets contain relatively dated survey data on teenagers in the 1950s to 1980s (an exception is Meade et al. (2008) who use data on teenagers in the early 2000s), and have relatively small sample sizes, ranging from around 6,000 to smaller than 500 (Barber 2001; Furstenberg et al. 1990; Hardy et al. 1998; Kahn and Anderson 1992; Manlove 1997; Meade et al. 2008). While some use nationally representative data (Kahn and Anderson 1992; Manlove 1997), others focus on select samples, some of which are mostly white (Barber 2001) or mostly black (Furstenberg et al. 1990). The studies universally find that women born to teenage mothers face considerably higher risk of teenage childbearing. At the same time, Kahn and Anderson (1992) – who use the 1998 National Survey of Family Growth and is the only study which compares the results for blacks and whites – find that the effects of maternal age are much weaker for blacks, so that black women born to younger and older mothers have more similar teenage childbearing outcomes.

In contrast to previous work, this paper is based on a high-quality administrative dataset on 35,544 first births in North Carolina born in 1987-1989. Like Kahn and Anderson (1992), we find that maternal age effects are much weaker for blacks. We show for the first time that these weak maternal age effects are largely due to a much higher probability of teenage birth among black women born to adult mothers compared to their white counterparts. (Black and white women born to teenage mothers are about equally likely to have a teenage birth.) This difference alone accounts for around 40% of the black-white gap in teenage birth probabilities, with the difference in maternal age distributions accounting for another one-third. Hence, while previous literature has highlighted the role of intergenerational transmission and shared social positioning, leading to a cycle of poverty (e.g., Furstenberg et al. 1990), our analysis suggests that successive generations of black women continue to face higher risks of falling into disadvantage even after one generation escapes it.

We also explore why black women born to older mothers are at much higher risk of teenage childbearing. While previous studies use sequential adding of regression covariates to tease out the relative importance of individual mediators, we apply a decomposition method suggested by Gelbach (2009) which produces estimates that are invariant to the order in which the covariates are introduced. We find, consistent with previous work, that family-level characteristics, especially maternal education, are the most important mediators of maternal age effects (Kahn and Anderson 1992; Manlove 1997). Unlike previous work, however, which has paid relatively little attention to the potential effects of paternal characteristics on daughters' teenage fertility outcomes, we find that paternal education is also an important mediator of maternal age effects even after accounting for maternal education. Finally, we show that two-thirds of the black-white difference in teenage childbearing risk among those born to older mothers can be explained by weaker associations between paternal characteristics and teenage fertility outcomes among blacks. Our results suggest that an important reason why women born to older mothers are less likely to become teenage mothers is that they tend to have more educated fathers, and that this effect is much stronger for whites than for blacks.

The two main findings in this paper are that a) the black-white gap in the probability of teenage childbearing is primarily due to the much higher probability of teenage birth among black women born to adult mothers, compared to their white counterparts; and that b) two-thirds of this increased probability of teenage birth among black women born to adult mothers is due to the weaker associations between paternal characteristics and teenage fertility outcomes among blacks. To our knowledge, neither of the findings has been previously shown. Our analysis does not allow us to make the causal claim that paternal education reduces daughters' probability of teenage childbearing, since it could instead be the case that women who provide better family environments for their children also tend to choose more educated and involved partners. Nevertheless, our evidence is consistent with the story of weaker marital prospects and expectations among black women,[1] increasing black women's risk of teenage childbearing through reducing family resources at younger ages and reducing gains to postponing childbearing at older (teenage) ages.

[1] McLanahan (2009) finds that unmarried black mothers have much lower expectations of marriage, despite stronger pro-marriage attitudes (although they are also more likely to believe that having a partner is not necessary for raising a child).

Methods

There are two main research objectives in this paper. The first objective is to explore the relative contributions of black-white differences in maternal age distributions and maternal age effects to the gap in teenage birth probabilities. To do so, we use the following identity:

$$\mathrm{P_j}(T = 1) = \sum_x \mathrm{P_j}(MA = x)\mathrm{P_j}(T = 1 | MA = x)\,, \tag{1}$$

where $\mathrm{P_j}(T = 1)$ is the probability of teenage childbearing for group *j*, and $\mathrm{P_j}(MA = x)$ is the probability that maternal age *MA* is equal to *x* for group *j*, where *x* represents three possible values: 19 or below, 20-25, and 26 or above. Hence, the first component of the identity represents the maternal age distribution for group *j*, while the second component represents the probabilities of teenage childbearing for each maternal age category for group *j*. We compute the components of the identity separately for whites and blacks, and use standardization to illustrate the relative importance of each component to the gap in teenage birth probabilities.

The second objective is to examine potential explanations for black-white differences in maternal age effects. To do so, we first model an individual's teenage childbearing outcome, represented by a dichotomous indicator T_i (1 = individual *i* gave birth as a teenager), as a linear function of her maternal age,[2] represented by a 2x1 row vector $\boldsymbol{MA_i}$ of dichotomous indicators for the last two of the above maternal age categories (where the first is omitted). Algebraically,

[2] There are two drawbacks to using linear probability models compared to other discrete choice models such as probit or logistic models. First, unlike the case where the dependent variable is continuous, the estimates are not efficient; however, in this paper, power is unlikely to be an issue due to the large sample size. Second, the predicted values of $\mathrm{P}(T_i = 1)$ may fall outside the

$$P(T_i = 1) = \alpha_0 + \boldsymbol{MA}_\mathbf{i}\boldsymbol{\alpha}_\mathbf{1} + \mathbf{1}\boldsymbol{C}_\mathbf{i}. \tag{2}$$

where $\boldsymbol{\alpha}_\mathbf{1}$ is a column vector representing the model coefficients, which give the marginal differences in teenage childbearing probabilities relative to the omitted class, and $\boldsymbol{C}_\mathbf{i}$ is a column vector of indicators for cohort fixed effects. The model is estimated using least squares.

Next, we expand the model to include other variables which may be associated with both individual *i*'s maternal age and teenage childbearing outcome, including: a) her family-level characteristics, namely her maternal education, paternal characteristics, parental marital status, and family size, b) her neighborhood-level characteristics, namely the student poverty and crime rates and the proportions of students passing end-of-grade tests and living with single parents in her school and school district, and c) her early educational outcomes,[3] namely her end-of-grade test scores in third to fifth grade and whether she was held back in school by fifth grade. These additional variables are represented by a row vector $\boldsymbol{X}_\mathbf{i}$ of *k* dichotomous indicators (see Table 1). Algebraically,

$$P(T_i = 1|\boldsymbol{X}_\mathbf{i}) = \beta_0 + \boldsymbol{MA}_\mathbf{i}\boldsymbol{\beta}_\mathbf{1} + \boldsymbol{X}_\mathbf{i}\boldsymbol{\beta}_\mathbf{2} + \mathbf{1}\boldsymbol{C}_\mathbf{i}, \tag{3}$$

where $\boldsymbol{\beta}_\mathbf{1}$ and $\boldsymbol{\beta}_\mathbf{2}$ are 2x1 column vectors representing the model coefficients. The model is similarly estimated using least squares.

---

range of zero to one. Despite the drawbacks, we prefer this specification due to its compatibility with the Gelbach decomposition method.

[3] While data on later outcomes, including test scores in eighth grade, are available, we focus on early outcomes in order to avoid estimation issues stemming from simultaneous causality.

To examine black-white differences in maternal age effects $\boldsymbol{\alpha_1}$, we explore the extent to which $\boldsymbol{X_\mathrm{i}}$ accounts for these maternal age effects. While previous work uses sequential adding of covariates to determine the relative importance of each variable, we apply an improved method which produces estimates that are invariant to the order in which the covariates are introduced. Gelbach (2009) shows that the difference between $\widehat{\boldsymbol{\alpha}}_{\mathbf{1}}$ and $\widehat{\boldsymbol{\beta}}_{\mathbf{1}}$, hereafter denoted as the 2x1 column vector $\widehat{\boldsymbol{\delta}}$, can be decomposed using the following identity:

$$\widehat{\boldsymbol{\delta}} = \widehat{\boldsymbol{\tau}}\widehat{\boldsymbol{\beta}}_{2}, \tag{4}$$

where $\widehat{\boldsymbol{\tau}}$ is a 2xk matrix, with each of the two rows representing the coefficients for each of the two elements of $\boldsymbol{MA_\mathrm{i}}$ when it is regressed on each of the k $\boldsymbol{X_\mathrm{i}}$s. This decomposition allows us to compute the individual contributions of each of the k predictors to each component of $\widehat{\boldsymbol{\delta}}$.[4] We estimate the above coefficients and use standardization to illustrate the relative importance of each component of $\boldsymbol{X_\mathrm{i}}$ to the black-white gap in teenage birth probabilities.

[4] Intuitively, the change in the estimated effects of maternal age on teenage childbearing once $\boldsymbol{X_\mathrm{i}}$ is included may be thought of as the product of a) the association between maternal age and $\boldsymbol{X_\mathrm{i}}$, represented by $\widehat{\boldsymbol{\tau}}$, and b) the association between $\boldsymbol{X_\mathrm{i}}$ and teenage childbearing, represented by $\widehat{\boldsymbol{\beta}}_{\mathbf{2}}$. Here, the component $\widehat{\boldsymbol{\tau}}$ reflects the fact that $\widehat{\boldsymbol{\alpha}}_{\mathbf{1}}$ is computed with $P(T_\mathrm{i} = 1)$ rather than $P(T_\mathrm{i} = 1|\boldsymbol{X_\mathrm{i}})$ as the dependent variable, so that $\boldsymbol{X_\mathrm{i}}$ is not held constant.

Data

The dataset in this paper follows three North Carolina birth cohorts (the "Sample") and consists of three components. The first component is each Sample individual's birth certificate, which provides information about her family background in terms of maternal race, age, educational attainment and marital status at the time of birth.[5] We also estimate her family size by summing the number of North Carolina birth certificates to her mother within 20-22 years of her mother's first birth. The second component is each Sample individual's public school administrative records, which provide information about her school and school district environments in terms of student poverty and crime rates and the proportions of students passing end-of-grade tests and living with single parents,[6] her end-of-grade test scores between third and fifth grade, and her ages at registration for these grades. The third component is any North Carolina birth certificate which lists a Sample individual as the mother, which provides

[5] Maternal race is self-reported on birth certificates, with the following possible values: white, black, American Indian, Asian, other non-white, and unknown. We define Sample individuals' race using maternal race, which, unlike paternal race, is almost always reported.

[6] While most of the sample attended third to fifth grade between 1996 and 2000, estimates of school poverty and crime rates and the proportion of school students who pass end-of-grade tests are based on 2005, 2001 and 2004 data respectively, the earliest years when the data are available; similarly, estimates of the school district poverty rate and the proportions of school district students who pass end-of-grade tests and live with single parents are based on 2004, 2002 and 2004 data respectively. Around 2.3% and 0.7% of observations have missing values for school and school district characteristics respectively; for these observations, the missing values are imputed using other school and school district characteristics.

information about whether she had a teenage birth and her exact age at the time of first birth. The dataset is de-identified and was obtained from the North Carolina Education Research Data Center, which performed all data linkages at the individual level, with permission from the NC Department of Health and Human Services.

The three Sample birth cohorts were born between 1987 and 1989, where 1987 is the first year when linked birth certificates are available from the North Carolina Education Research Data Center, and 1989 is the last cohort for which there are complete teenage childbearing histories (since data for the third component are available only up to and including 2009). Since maternal age at first birth is not reported on birth certificates, we restrict our sample to first births.

Between 1987 and 1989, 61,033 female singleton first births were born in North Carolina to state residents. This sample excludes out-of-state births to North Carolina residents (4.1% of all births), who are slightly more likely to be born to younger, less-educated and unmarried white mothers (see Appendix A1). Of this original sample, we drop 1,881 observations (3.1%) who are neither non-Hispanic white nor non-Hispanic black (data on Hispanic ethnicity were not available in 1987 so all white and black individuals were included in that year), and another 105 observations (0.2%) with missing data on maternal age or education or parental marital status. Observations with missing data for paternal characteristics, on the other hand, are not dropped since they constitute a fairly large proportion of the sample (see Table 1); instead, we include dichotomous indicators for missing data on father's age or education. Most significantly, we drop 23,503 observations (38.5%) who did not attend public school in third grade, eighth grade or at age 15, the final age before individuals are legally allowed to drop out of school in North Carolina. (A small number of observations who attended charter schools were also dropped due to lack of data on test scores and school/school district characteristics.) Hence, the final sample excludes individuals who did not survive to teenage years (including 560 observations with infant death certificates) as well as individuals who moved out of state or attended private schools during these ages. These individuals are excluded not just to obtain a dataset with more complete educational data, but also because the remaining individuals are less likely to have moved out of state, so that any teenage births they have had are more likely to be captured by the third component of the dataset. Relative to the original sample of female singleton first births in North Carolina, the final sample over-represents women from disadvantaged family backgrounds, with a higher proportion born to black (30.6% vs. 27.6%) or unmarried mothers (33.8% vs. 30.6%), and hence disproportionately includes the population at higher risk of teenage childbearing. The final sample size is 35,544.

The dataset used in this paper has several important advantages for studying the black-white difference in probabilities of teenage motherhood. First, it tracks the teenage childbearing experiences of recent birth cohorts who were teenagers in the mid to late 2000s, rather than in the early 2000s (Meade et al. 2008) or between the late 1950s and late 1980s (Barber 2001; Furstenberg et al. 1990; Hardy et al. 1998; Kahn and Anderson 1992; Manlove 1997). This is potentially important given the dramatic decline in teenage birth rates in the past two decades, which may reflect changes in selection into early childbearing. Second, it is considerably larger than datasets used in previous studies, with a sample size of 35,544 compared to sample sizes ranging from 6,084 (Kahn and Anderson 1992) to smaller than 500 (Barber 2001; Furstenberg et al. 1990). Third, it is based on high-quality administrative records rather than longitudinal or retrospective survey responses. While survey datasets such as the National Survey of Family Growth and the National Longitudinal Survey of Youth (used in Kahn and Anderson 1992 and

Meade et al. 2008 respectively) tend to have rich individual-level detail, including items such as ideal age at childbearing and educational expectations, administrative data are likely to be more accurate, especially in the case of school test scores. In this dataset, all North Carolina public school students in the same grade and year take the same end-of-grade test, and each individual's test score is represented by her Z-score relative to the test results of all students (including male and other female students not included in the sample) who took the test.[7] Around 1.0% and 0.9% of third to fifth reading and math grade test scores are missing; for these observations, test scores are imputed using the individual's test scores in the other subject, the number of grades she eventually completed and her age at school exit.

Table 1 provides some summary statistics. Of the women born between 1987 and 1989, those who became teenage mothers were much more likely to be born to young unmarried mothers with no high school degree, to have fathers with no high school degree or missing characteristics, and to have at least three siblings. They also tended to attend schools and live in school districts where the students were relatively likely to come from poor or single-parent households, commit crimes, and fail end-of-grade tests. Similar to their peers, they were doing much more poorly in reading and math as early as third to fifth grade, and were more likely to be held back by fifth grade.

[7] Since individual test scores are represented by their Z-scores relative to the test results of all students who took the test, the average test score should be around zero if the sample is representative of North Carolina public school students. In this sample, the average test scores are 0.16 and 0.05 for reading and math in third to fifth grade, where the reading scores in particular are higher than zero since only non-Hispanic white and black women born in North Carolina, who are unlikely to be taking English as a second language, are included.

Results

Decomposition of the Black-White Gap in Probability of Teenage Childbearing

Of the 35,544 women in the sample, 2.5% gave birth before age 17 and another 14.0% had their first birth between ages 17 and 19, so that the total probability of teenage motherhood is 16.5%. The vast majority of the teenage mothers were unmarried at the time of birth (97.2% among those who gave birth before age 17 and 87.7% among older teenage mothers). Table 2 shows that black women are almost twice as likely to experience teenage motherhood, and that these births are more likely to be non-marital. Compared to teenage women in the late 1950s to late 1980s, recent cohorts are substantially less likely to become adolescent mothers but much more likely to be unmarried at the time of birth (see Table 2), possibly due to the increasing social acceptability of non-marital childbearing even as eventual marriage remains an ideal (Furstenberg et al. 1990; Lichter, Batson and Brown 2004).

Figure 1 shows the teenage motherhood survival curves for this sample by race and whether they were born to teenage mothers. It illustrates two points about the black-white gap in probabilities of teenage motherhood. First, consistent with Meade et al. (2008), we find that the risks of childbearing are universally very low until around age 15, when the curves begin to diverge and continue to do so throughout the teenage years. This suggests that most of the black-white differences in teenage fertility behavior occur at older, rather than younger, teenage ages (see Appendix A2). Second, while the survival curves for women born to white and black teenage mothers lie fairly close to each other, those for women born to white and black older mothers are quite far apart, with a substantially higher probability of childbearing among the latter group. This suggests that maternal age effects are weaker for blacks than for whites, not because of weaker intergenerational transmission of teenage motherhood – in fact, women born to black teenage mothers appear to be at slightly higher risk than their white counterparts – but because being born to an adult mother is less protective for blacks.

Figure 2 also illustrates the second point, this time by showing the probabilities of teenage motherhood by race at all maternal ages. While the curves for whites and blacks both start out at probabilities of around 0.40, the latter curve is substantially flatter, so that the black-white gap in teenage childbearing probabilities increases with maternal age (see Fig. 2d). Figure 2 also provides an additional insight: while the literature has focused almost exclusively on the differences between women born to teenage and adult mothers (an exception is Hardy et al. 1998, who compare women born to mothers aged 20-24 and mothers aged 25 or above), the evidence suggests that the association between maternal age and teenage childbearing in the next generation is continuous, so that while women born to teenage mothers face the highest risk of all, women born to mothers in their early 20s are also at elevated risk compared to women born to mothers in their late 20s or older.

Thus far, the graphical evidence suggests that maternal age effects are weaker for blacks than for whites, due to higher probabilities of teenage childbearing among black women born to older mothers. We now turn to standardization to illustrate the relative importance of maternal age distributions and effects on the black-white gap in probabilities of teenage motherhood. The top half of Table 3 shows that black women are much more likely to be born to teenage mothers than white women, and that their risks of teenage motherhood are higher at all maternal ages, resulting in a total gap of around 11 percentage points in the probability of teenage motherhood (0.239 vs. 0.132 for whites). The bottom half of the table shows that if black women had exactly the same distribution of maternal ages as their white counterparts, the gap would narrow by a third to 7 percentage points (0.202 vs. 0.132); alternatively, if black women faced the same risks of teenage motherhood at all maternal ages as their white counterparts, the gap would narrow by more than half to 5 percentage points (0.180 vs. 0.132). The results suggest that weaker maternal age effects is the primary reason why black women are more likely to become teenage mothers, with the higher risk among women born to adult mothers alone accounting for 40% of the gap (see Table 3).

The weaker maternal age effects among blacks also have some interesting implications for population dynamics. Instead of estimating the proportion of daughters born to teenage mothers who become teenage mothers themselves, we could estimate the proportion of teenage mothers whose mothers also gave first birth as teenagers. In our sample, 49.7% of women who became teenage mothers also had mothers who gave first birth as teenagers, with a higher proportion among black teenage mothers (57.1% vs. 43.9% for whites). However, once we standardize maternal age distributions using the values for whites (see Appendix A3), black teenage mothers are actually *less* likely to be born to mothers who gave first birth as teenagers (31.2% vs. 43.9%), even though the intergenerational transmission of teenage childbearing is slightly stronger for blacks.

Mediators of Maternal Age Effects

Next, we explore why black women born to older mothers are at higher risk of teenage childbearing than their white counterparts. To do so, we compare the contributions of six potential mediators to maternal age effects for whites and blacks: maternal education, paternal age and education, parental marital status, family size, neighborhood characteristics and early educational outcomes. After controlling for these variables, maternal age effects continue to be statistically significant for whites but not for blacks (see Table 4).

One key finding from previous literature is that family-level characteristics, especially maternal education, are the most important mediators of maternal age effects (Kahn and Anderson 1992; Manlove 1997). However, relatively little attention has been given to the potential effects of paternal characteristics on daughters' teenage fertility outcomes (an exception is Manlove (1997), who controls for father's occupational class). We find that the effects of paternal education are large and highly significant even after accounting for maternal education. Nevertheless, these effects are somewhat weaker than those of maternal education, potentially due to union dissolution, lower involvement of fathers relative to mothers, or stronger associations between child outcomes and characteristics of same-sex parents.[8] On the other hand, similar to Mollborn and Lovegrove (2010), we find little evidence of paternal age effects. One potential explanation is that the paternal age data used in this paper may not necessarily reflect fathers' age at first birth; alternatively, paternal inputs may depend less on paternal age, with Mollborn and Lovegrove (2010) finding that teenage and older fathers have similar levels of involvement and attitudes towards parenting.

While maternal education is even more predictive for black women than for their white counterparts, paternal characteristics are more weakly associated with teenage childbearing outcomes (though still statistically significant). This result may be due to lower union stability in both married and unmarried black families (McLanahan 2009), or due to lower involvement of

[8] Thomas (1994) argues that parental characteristics affect household bargaining power, which disproportionately benefits same-sex children due to the parental preferences or technological advantages.

black fathers even in intact families (Harris et al. 1998), which has fallen further in recent generations.[9] We argue that the result may explain why previous papers, which do not control for paternal characteristics, find that co-residence with both biological parents is more strongly associated with teenage childbearing outcomes for white women (Kahn and Anderson 1992; Moore et al. 1998), since co-residence is likely to be associated with higher paternal age and education. (Our evidence, on the other hand, suggests that being born to married mothers is slightly more predictive of teenage childbearing outcomes for blacks after controlling for paternal characteristics.)

We also find that the teenage fertility outcomes of black women are more strongly associated with having at least three siblings. One potential explanation is that the intergenerational transmission of family size is stronger in black families; alternatively, the result may reflect the increased risk associated with more frequent maternal partnership changes, since they increase the probability of having half-siblings, especially if the mother is unmarried (McLanahan 2009). There is some evidence that adolescent girls who live with non-biological fathers are more likely to exhibit risky behaviors (Amato and Rivera 1999), possibly because of lower paternal investment (McLanahan 2009) or lower maternal ability to discipline or monitor their children's behavior due to financial hardship and lower social and emotional support (Barber 2001; Kahn and Anderson 1992; McLanahan 2009).

Unlike family-level characteristics, neighborhood-level characteristics are generally weakly associated with teenage fertility outcomes. The two most predictive variables are school

[9] In a sample of disadvantaged black women living in Baltimore, 21% of unmarried fathers provide child support and 55% have no contact with their children, compared to 41% and 38% in the previous generation (Furstenberg et al. 1990).

poverty rates and academic performance, which are only marginally significant for both whites and blacks. Our findings are consistent with findings from the Moving to Opportunity experiment (Ludwig et al. 2013), as well from cross-sectional studies, including Abrahamse et al. (1988), who use data from the High School and Beyond study, and Manlove (1998) and Moore et al. (1998), who use nationally representative data from the National Education Longitudinal Study, which also find that neighborhood and school characteristics are not strongly predictive of teenage childbearing outcomes.

Finally, again consistent with previous studies (Abrahamse et al. 1988; Manlove 1998; Moore et al. 1998), we find that better early educational outcomes are associated with lower probability of teenage motherhood for both whites and blacks, with slightly larger effects for the latter. While the associations are weak compared to those of family-level characteristics, we suspect that academic performance at older ages is likely to be a stronger mediator of maternal age and education effects by more closely reflecting college expectations (Moore et al. 1998), especially for black women (Abrahamse et al. 1988).

Decomposition of Maternal Age Effects

While the above regression analysis highlights the importance of family-level characteristics for teenage childbearing outcomes for both whites and blacks, it does not offer a breakdown of the individual contribution of each mediator to the observed maternal age effects, or how this breakdown differs for whites and blacks. To do so, we apply the order-invariant decomposition method proposed by Gelbach (2009) and present our results in Table 5. For both white and black women, maternal education explains the largest proportions of maternal age effects. This is particularly true for women born to young adult mothers (aged 20-25) compared to women born to teenage mothers, where maternal education alone accounts for 45% of the difference in probability of teenage motherhood. The reduced risk of teenage childbearing among women born to even older mothers (aged 26 or above) compared to women born to teenage mothers, on the other hand, is less solely driven by maternal education (which accounts for 34% of the difference in probability of teenage motherhood) and more due to a combination of maternal education, paternal characteristics and residual age effects.

Table 5 shows that maternal age effects are smaller for blacks than for whites by around 4 percentage points, and that this difference is due to large part to the smaller contribution of paternal characteristics to maternal age effects among blacks. This result is, in turn, due either to a) weaker associations between maternal age and paternal characteristics (represented by $\hat{\boldsymbol{\tau}}$ in Eq. 4), or b) weaker associations between paternal characteristics and daughters' teenage childbearing outcomes (represented by $\widehat{\boldsymbol{\beta}}_2$ in Eq. 4) (see Appendix A4). Turning once again to standardization to illustrate the relative importance of the two potential explanations, we find that if the associations between maternal age and paternal characteristics for black women were replaced by those for white women, the black-white gap in probability of teenage childbearing would fall to around 3 percentage points; alternatively, if the associations between paternal characteristics and daughters' teenage childbearing outcomes were replaced instead, the gap would fall to around 1.5 percentage points. Hence, around two-thirds of the black-white difference in maternal age effects can be attributed to the lower importance of paternal characteristics for daughters' teenage childbearing outcomes in black families, potentially due to more frequent maternal partnership changes (McLanahan 2009) or lower paternal involvement (Harris et al. 1998).

Our results are broadly consistent with previous studies in finding that maternal education and parental co-residence are important mediators of maternal age effects. Similar to Manlove (1997), we find that school performance accounts for around 8% of maternal age effects. However, our estimates of the contribution of maternal education to maternal age effects are larger than in previous papers (Kahn and Anderson 1992; Manlove 1997), even though the latter's use of sequential adding of covariates should lead to overestimation since maternal education is introduced before most of the other covariates. In part due to this difference, as well as to our ability to account for the effects of paternal characteristics, our analysis also explains a larger proportion of maternal age effects than previous papers (around 80%, compared to less than half in Kahn and Anderson 1992 and Manlove 1997). The remaining 20% of maternal age effects may reflect the role of other variables not available in our administrative dataset, such as household income, maternal marital history or daughters' marital and educational aspirations.

Discussion

Teenage childbearing in the United States has long been of great concern to policymakers due to the view that having a child early in life leads to negative consequences: for the mother herself, in terms of lower future job and marital prospects; for the child, in terms of worse health and environments and hence poorer life chances (Abrahamson et al. 1988; Meade et al. 2008); and for the rest of society, in terms of higher welfare program costs and foregone tax revenues (U.S. Department of Health and Human Services 2014; Graefe and Lichter 2002), estimated to be on the order of $32 billion per year (Abrahamson et al. 1988).[10] While some recent literature

[10] The original figure in Abrahamson et al. (1988) is $16 billion, which we update to 2014 dollars using inflation data from the Bureau of Labor Statistics.

shows that the consequences of teenage childbearing for the mother are less negative than previously believed (Geronimus and Korenman 1992; Hotz et al. 2005), the latest evidence suggests that early motherhood does lead to worse life outcomes (Ashcraft and Lang 2006; Fletcher and Wolfe 2009; Hoffman 2008; Kane et al. 2013), consistent with the observation that most teenage births are unwanted or mistimed (Kissin et al. 2008).

The two main findings in this paper are that a) the black-white gap in the probability of teenage childbearing are primarily due to the much higher probability of teenage birth among black women born to adult mothers, compared to their white counterparts; and that b) two-thirds of this increased probability of teenage birth among black women born to adult mothers is due to the weaker associations between paternal characteristics and teenage fertility outcomes among blacks. While our analysis does not allow us to make the claim that the black-white gap in the probability of teenage childbearing is caused by lower paternal involvement, the evidence is consistent with the story of weaker marital prospects and expectations among black women, increasing black women's risk of teenage childbearing by reducing family resources at younger ages and reducing gains to postponing childbearing at older (teenage) ages. We also note that there are other potential explanations for the black-white gap in the probability of teenage childbearing, including cultural, economic and social differences such as greater value placed on taking on nurturing roles and adult responsibilities among black adolescent girls, higher health risks associated with later childbearing among black women and greater support from black kinship networks for child rearing (Meade et al. 2008). The final point is consistent with Manlove (1998) and the dataset used in this paper, which find that black teenage mothers complete more schooling than white teenage mothers (see Appendix A5), despite having poorer test scores in third to fifth grade (not shown).

The evidence presented in this paper has some implications for policy. Previous literature has highlighted the role of intergenerational transmission and shared social positioning, leading to a cycle of poverty (e.g., Furstenberg et al. 1990). Our analysis, on the other hand, suggests that successive generations of young black women continue to face higher risk of teenage motherhood even after one generation escapes it, and that this increased risk may be due in part to lower marital prospects and stability among black women, possibly due to higher poverty and incarceration rates among black men. At the same time, while we find that it is harder to break out of disadvantage than previously suggested, we also find that strong early educational performance is associated with substantially lower risk, and we suspect that school performance at older ages and college expectations are even more important. Our evidence suggests that policies which increase black families' stability and access to resources through improving educational and market opportunities for low-income women and men is key to bridging the black-white gap in teenage birth rates and promoting the well-being of the next generation. On the other hand, we find weak evidence that peer characteristics at the school and school district levels matter, suggesting that policies aimed at improving broader environments are likely to have minimal spillover effects on local teenage birth rates.

Conclusion

There are several potentially important limitations to this paper. One of our central findings is that the associations between paternal characteristics and teenage fertility outcomes are much weaker for black women, which may explain why black women born to adult mothers are at much higher risk of teenage childbearing than their white counterparts. While our administrative data on paternal characteristics are useful, we cannot distinguish between unmarried cohabiting fathers and non-resident fathers, where the latter group is less likely to provide financial support and less interested in being involved in childrearing (McLanahan 2009). Furthermore, since our data on family-level characteristics are collected at the time of birth, they may not reflect family circumstances at teenage ages.

Second, our measure of sibship size may underestimate actual family size, since it is based on mother's completed fertility and does not include paternal stepsiblings or adopted children. This measurement issue is likely to be particularly relevant for black families, who are at higher risk of union dissolution.

Third, the administrative data do not allow us to look further into the channels through which the mediators of maternal age effects operate. For example, we are unable to investigate whether parental education matters due to its "direct" or "active" effects on parenting style, communication or monitoring, or due to more "indirect" or "passive" channels such as higher household income or different ideological values.[11] In the case of paternal education, it would be especially interesting to see whether "direct" effects are especially weak since fathers generally interact less with their daughters than with their sons (Harris et al. 1998; Harris and Morgan

[11] While there is a strong association between household income and teenage fertility outcomes (Abrahamse et al. 1988; Hardy et al. 1998; Meade et al. 2008), there is mixed evidence that maternal parenting style, communication and involvement matter for teenage childbearing outcomes (Kahn and Anderson 1992; Manlove 1998; Meade et al. 2008), although maternal interest in daughters' educational performance and communication about how pregnancy occurs appear to be significant factors (Kahn and Anderson 1992; Manlove 1997). Similarly, there is mixed evidence that paternal attitudes, monitoring and discipline matter, possibly due to varied measurements of paternal attitudes and involvement (Guilamo-Ramos et al. 2012).

1991), or whether they are stronger if paternal education also increases the propensity of fathers to spend time with their daughters.

Fourth, while we find little evidence that school and school district characteristics are associated with teenage fertility outcomes, we note that local access to contraception or abortion for adolescents is likely to matter (Kahn and Anderson 1992). Furthermore, while peer characteristics at the school or broader levels may be less important, those of more immediate peers may be more influential, with some evidence that school classmates' delinquency and substance use are associated with higher risk of teenage motherhood (Meade et al. 2008).

Appendix A1 Maternal Characteristics in North Carolina, 1987-1989

	Total	% white	% below 20 years old	% with less than high school degree	% unmarried
			Vital Statistics		
1987	93,501	69.3	15.7	23.2	24.9
1988	97,579	67.6	16.0	23.2	26.3
1989	102,105	66.8	16.4	22.9	27.7
			Birth certificates		
1987	89,704	69.0	15.2	22.7	24.1
1988	93,507	67.2	15.4	22.7	25.5
1989	97,996	66.5	15.9	22.3	27.0

Notes: Vital Statistics data include out-of-state births to mothers residing in North Carolina, while birth certificate data include only in-state births to mothers residing in North Carolina. Vital Statistics data were obtained online from the Centers for Disease Control and Prevention at http://www.cdc.gov/nchs/data_access/vitalstats/VitalStats_Births.htm. Birth certificate data were obtained from the North Carolina Education Research Data Center. All calculations are made by the authors.

Appendix A2 Decomposition of Maternal Age Effects into Effects on First Birth at Ages Below 17 and First Birth at Ages 17-19

The goal of this exercise is to decompose maternal age effects into differences in the probabilities of first birth a) by age 17 and b) between ages 17 and 19. One complication is that very early childbearing reduces the risk of first birth at older teenage ages, so that maternal age effects on the two probabilities do not add up to the total effects on probability of teenage motherhood. We use an approximation which decomposes this interaction effect.

Let P(*x*, *y*) denote the probability of giving first birth between ages *x* and *y*, then

$$P(0, 19) = 1 - [1 - P(0, 17)]x[1 - P(17, 19)] = P(0, 17) + P(17, 19) - P(0, 17)xP(17, 19),$$

where the third term is negative since a higher probability of giving birth by age 17 reduces the risk of giving first birth between ages 17 and 19. For two maternal age ranges *m1* and *m2*, we would like to estimate the two terms on the right side of the following equation:

$$P_{m1}(0, 19) - P_{m2}(0, 19) = [P_{m1}(0, 17) - P_{m2}(0, 17)] + [P_{m1}(17, 19) - P_{m2}(17, 19)].$$

To do so, we use the following approximation for differences in the negative third term:

$$P_{m1}(0, 17)*P_{m1}(17, 19) - P_{m2}(0, 17)xP_{m2}(17, 19) = [P_{m1}(0, 17) - P_{m2}(0, 17)]x[P_{m1}(17, 19) + P_{m2}(17, 19)]/2$$
$$+ [P_{m1}(17, 19) - P_{m2}(17, 19)]x[P_{m1}(0, 17) + P_{m2}(0, 17)]/2,$$

where $[P_{m1}(17, 19) + P_{m2}(17, 19)]/2$ and $[P_{m1}(0, 17) + P_{m2}(0, 17)]/2$ act as weights. The approximation thus yields:

$$P_{m1}(0, 17) - P_{m2}(0, 17) = [P_{m1}(0, 17) - P_{m2}(0, 17)] + [P_{m1}(0, 17) - P_{m2}(0, 17)]x[P_{m1}(17, 19)$$
$$+ P_{m2}(17, 19)]/2,$$

and

$$P_{m1}(17, 19) - P_{m2}(17, 19) = [P_{m1}(17, 19) - P_{m2}(17, 19)] + [P_{m1}(0, 17) - P_{m2}(0, 17)]x[P_{m1}(17, 19)$$
$$+ P_{m2}(17, 19)]/2.$$

The above approximation, when applied to the calculation of $P_{m1}(0, 19) - P_{m2}(0, 19)$ for the sample in this paper, yields estimates which are accurate up to at least 6 decimal places. The results are presented below:

Marginal Differences in Probability of Teenage Childbearing by Maternal Age

	All	Whites	Blacks
	Probability of teenage childbearing by maternal age		
First birth before age 20			
Mother was aged 19 or below	0.286	0.268	0.305
Mother was aged 20-25	0.160	0.139	0.213
Mother was aged 26 or above	0.063	0.048	0.131
First birth before age 17			
Mother was aged 19 or below	0.050	0.040	0.062
Mother was aged 20-25	0.021	0.014	0.038
Mother was aged 26 or above	0.008	0.005	0.021
First birth at ages 17-19 (conditional)			
Mother was aged 19 or below	0.248	0.238	0.260
Mother was aged 20-25	0.142	0.127	0.182
Mother was aged 26 or above	0.055	0.043	0.113
	Marginal differences in probabilities		
First birth before age 20			
Mother was aged 19 or below	-	-	-
Mother was aged 20-25	-0.126	-0.129	-0.092
Mother was aged 26 or above	-0.098	-0.091	-0.082
First birth before age 17			
Mother was aged 19 or below	-	-	-
Mother was aged 20-25	-0.023	-0.021	-0.019
Mother was aged 26 or above	-0.012	-0.008	-0.015
First birth at ages 17-19			
Mother was aged 19 or below	-	-	-
Mother was aged 20-25	0.102	-0.109	-0.074
Mother was aged 26 or above	-0.086	-0.083	-0.067
Number of observations	35,544	24,678	10,866

Notes: The conditional probabilities of first birth at ages 17-19 are calculated using the sample of women who did not give first birth before age 17. Marginal differences in probabilities are relative to the previous maternal age category and are calculated by the authors.

Appendix A3 Standardization for Proportion Born to Mothers Aged 20 or Below

To standardize the proportion born to mothers aged 20 or below for black women using the values for white women, we apply the following reweighting for black women:

$$w_i = (y_i - x_i)/(n - b)/(x_i/b) ,$$

where n is the total sample size,

b is the number of black women.

If i was born to a mother aged 20 or below, then

y_i is the number of women born to mothers aged 20 or below,

x_i is the number of black women born to mothers aged 20 or below.

If i was born to a mother aged above 20, then

y_i is the number of women born to mothers aged above 20,

x_i is the number of black women born to mothers aged above 20.

Hence, for black women born to mothers aged 20 or below, w_i is the proportion of white women born to mothers aged 20 or below, divided by the proportion of black women born to mothers aged 20 or below. These women are assigned a reweighting of less than one, since the numerator is smaller than the denominator. Similarly, for black women born to mothers aged above 20, w_i is the proportion of white women born to mothers aged above 20, divided by the proportion of black women born to mothers aged above 20. These women are assigned a reweighting of more than one, since the numerator is larger than the denominator.

Appendix A4 Decomposition of Maternal Age Effects on Teenage Childbearing into Effects Associated with Maternal Education, Paternal Characteristics, Parental Marital Status, Family Size, Neighborhood Characteristics and Early Educational Outcomes: Detailed Breakdown

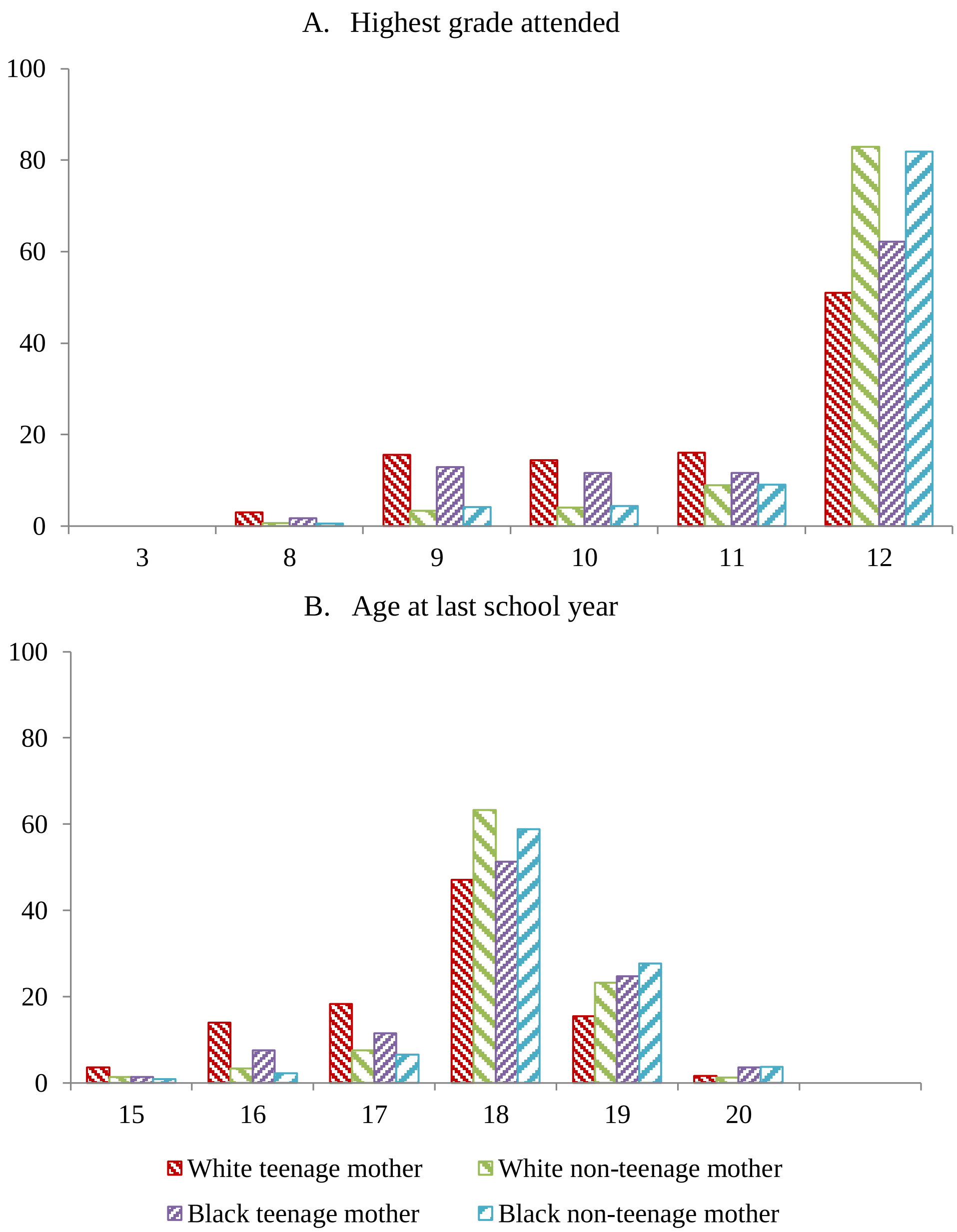
	All		Whites		Blacks	
	$\hat{\tau}$	$\hat{\beta}_2$	$\hat{\tau}$	$\hat{\beta}_2$	$\hat{\tau}$	$\hat{\beta}_2$
Maternal age = 20-25						
Maternal education						
Less than high school		-		-		-
High school	0.280***	-0.081***	0.256***	-0.080***	0.327***	-0.087***
More than high school	0.284***	-0.119***	0.285***	-0.107***	0.283***	-0.162***
Paternal age						
19 or below		-		-		-
20-25	0.153***	0.011	0.052***	0.015	0.202***	0.007
Over 25	0.295***	0.009	0.324***	0.016	0.184***	0.004
Age unknown	-0.270***	0.042**	-0.130***	0.070**	-0.281***	0.025†
Paternal education						
Less than high school		-		-		-
High school	0.191***	-0.065***	0.145***	-0.075***	0.184***	-0.035**
More than high school	0.178***	-0.088***	0.193***	-0.099***	0.116***	-0.062
Education unknown	-0.294***	-0.055***	-0.155***	-0.046*	-0.289***	-0.045†
Parental marital status						
Married	0.387***	-0.040***	0.290***	-0.024*	0.272***	-0.034**
Number of younger siblings						
None		-		-		-
One or two	0.025***	0.014**	0.002	0.012*	0.044***	0.019†
More than two	-0.091***	0.053***	-0.063***	0.038***	-0.117***	0.079***
School characteristics						
Student poverty rate > 55%	-0.111***	0.014**	-0.074***	0.009†	-0.030**	0.018†
Student passing rate > 80%	0.117***	-0.013**	0.067***	-0.010†	0.048***	-0.014
Crimes per 100 students > 0	-0.027***	0.005	-0.028***	0.006	-0.004	0.002
School district characteristics						
Child poverty rate > 17%	-0.036***	0.004	-0.005	0.003	-0.010	0.008
Student passing rate > 83%	0.053***	0.002	-0.002	0.004	0.020*	-0.001
Students w/one parent > 23%	-0.066***	0.005	0.022**	0.002	-0.029**	0.007
Early educational outcomes						
Reading Z-score < -1	-0.084***	0.035***	-0.061***	0.041***	-0.057***	0.028*
Reading Z-score > 1	0.092***	-0.033***	0.099***	-0.029***	0.021***	-0.055
Math Z-score < -1	-0.082***	0.014†	-0.054***	0.006	-0.044***	0.016***
Math Z-score > 1	0.073***	-0.023***	0.079***	-0.020***	0.013***	-0.045*
Retained by 5th grade	-0.079***	0.001	-0.073***	0.003	-0.069***	-0.004
Residual age effect	-0.018*		-0.035***		0.010	
Total age effect	**-0.126**		**-0.130**		**-0.093**	

Maternal age = 26 or above						
Maternal education						
Less than high school		-		-		-
High school	0.026***	-0.081***	0.004	-0.080***	0.098***	-0.087***
More than high school	0.619***	-0.119***	0.636***	-0.107***	0.544***	-0.162***
Paternal age						
19 or below		-		-		-
20-25	-0.241***	0.011	-0.390***	0.015	-0.068***	0.007
Over 25	0.823***	0.009	0.830***	0.016	0.686***	0.004
Age unknown	-0.379***	0.042**	-0.167***	0.070**	-0.492***	0.025†
Paternal education						
Less than high school		-		-		-
High school	0.089***	-0.065***	-0.004	-0.075***	0.205***	-0.035**
More than high school	0.489***	-0.088***	0.510***	-0.099***	0.321***	-0.062
Education unknown	-0.414***	-0.055***	-0.198***	-0.046*	-0.519***	-0.045†
Parental marital status						
Married	0.575***	-0.040***	0.375***	-0.024*	0.587***	-0.034**
Number of younger siblings						
None		-		-		-
One or two	-0.039***	0.014**	-0.054***	0.012*	-0.066***	0.019†
More than two	-0.130***	0.053***	-0.099***	0.038***	-0.167***	0.079***
School characteristics						
Student poverty rate > 55%	-0.229***	0.014**	-0.163***	0.009†	-0.090***	0.018†
Student passing rate > 80%	0.205***	-0.013**	0.119***	-0.010†	0.103***	-0.014
Crimes per 100 students > 0	-0.065***	0.005	-0.058***	0.006	-0.043***	0.002
School district characteristics						
Child poverty rate > 17%	-0.113***	0.004	-0.069***	0.003	-0.049***	0.008
Student passing rate > 83%	0.082***	0.002	-0.001	0.004	0.033**	-0.001
Student w/one parent > 23%	-0.060***	0.005	0.076***	0.002	-0.029*	0.007
Early educational outcomes						
Reading Z-score < -1	0.141***	0.035***	-0.100***	0.041***	-0.112***	0.028*
Reading Z-score > 1	0.249***	-0.033***	0.257***	-0.029***	0.070***	-0.055
Math Z-score < -1	-0.147***	0.014†	-0.097***	0.006	-0.104***	0.016***
Math Z-score > 1	0.210***	-0.023***	0.220***	-0.020***	0.047***	-0.045*
Retained by 5th grade	-0.116***	0.001	-0.100***	0.003	-0.123***	-0.004
Residual age effect	-0.053***		-0.074***		-0.015	
Total age effect	**-0.224**		**-0.220**		**-0.174**	

Notes: Values within each box refer to the estimated contribution of each component to maternal age effects (relative to maternal age = 19 or below). Coefficients are estimated using the Gelbach decomposition method, where $\hat{\boldsymbol{\tau}}$ represents differences in each mediator at each maternal age category, and $\widehat{\boldsymbol{\beta}}_2$ represents the estimated effect of each mediator on the probability of teenage motherhood.

†p <.10; *p < .05; **p < .01; ***p < .001

Appendix A5 Highest Grade Attended and Age at School Leaving by Teenage Motherhood Outcomes



Notes: Number of grades attended and age at last school are based on individuals' administrative school records up to age 23.

References

Abrahamse, A. F., Morrison, P. A., & Waite, L. J. (1988). Beyond Stereotypes: Who Becomes a Single Teenage Mother? *Rand Report Series*, 3489.

Amato, P. R., & Rivera, F. (1999). Paternal Involvement and Children's Behavior Problems. *Journal of Marriage and Family*, *61* (2), 375-384.

Ashcraft, A., & Lang, K. (2006). The Consequences of Teenage Childbearing. *NBER Working Paper*, 12485.

Barber, J. S. (2001). The Intergenerational Transmission of Age at First Birth among Married and Unmarried Men and Women. *Social Science Research, 30*: 219-247.

Bureau of Labor Statistics. (2014). *CPI Inflation Calculator*. http://www.bls.gov/data/inflation_calculator.htm. Accessed 15 July 2014.

Fletcher, J. M., & Wolfe, B. L. (2009). Education and Labor Market Consequences of Teenage Childbearing: Evidence Using the Timing of Pregnancy Outcomes and Community Fixed Effects. *Journal of Human Resources*, *44* (2): 303-325.

Furstenberg, F. F. Jr., Levine, J. A., & Brooks-Gunn, J. (1990). The Children of Teenage Mothers: Patterns of Early Childbearing in Two Generations. *Family Planning Perspectives*, *22* (2): 54-61.

Gelbach, J. B. (2009). When Do Covariates Matter? And Which Ones, and How Much? *University of Arizona Economics Department Working Paper*, 09-07.

Graefe, D. R., & Lichter, D. T. (2002). Marriage Among Unwed Mothers: Whites, Blacks and Hispanics Compared. *Perspectives on Sexual and Reproductive Health*, *34* (6): 286-293.

Geronimus, A. T., & Korenman, S. (1992). The Socioeconomic Consequences of Teen Childbearing Reconsidered. *Quarterly Journal of Economics*, *107* (4): 1187-1214.

Guilamo-Ramos, V., Bouris, A., Lee, J., McCarthy, K., Michael, S. L. Pitt-Barnes, S., & Dittus, P. 2012. Paternal Influences on Adolescent Sexual Behaviors: A Structured Literature Review. *Pediatrics*, *130* (5): 1313-1325.

Hardy, J. B., Astone, A. M., Brooks-Gunn, J., Shapiro, S., & Miller, T. L. (1998). Like Mother, Like Child: Intergenerational Patterns of Age at First Birth and Associations with Childhood and Adolescent Characteristics and Adult Outcomes in the Second Generation. *Developmental Psychology*, *34* (6): 1220-1232.

Harris, K. M., Furstenberg, F. F. Jr., & Marmer, J. K. (1998). Paternal Involvement with Adolescents in Intact Families: The Influence of Fathers over the Life Course. *Demography*, *35* (2): 201-216.

Harris, K. M., & Morgan, S. P. (1991). Fathers, Sons and Daughters: Differential Paternal Involvement in Parenting. *Journal of Marriage and the Family*, *53* (3): 531-544.

Hoffman, S. D. (2008). Updated Estimates of the Consequences of Teen Childbearing for Mothers. In S. D. Hoffman and R. A. Maynard (Ed.), *Kids Having Kids: Economic Costs and Social Consequences of Teen Pregnancy (second edition)* (pp. 74-92). Washington D.C.: Urban Institute Press.

Hotz, V. J., McElroy, S. W., & Sanders, S. G. (2005). Teenage Childbearing and Its Life Cycle Consequences. *Journal of Human Resources*, *40* (3): 683-715.

Kahn, J. R., & Anderson, K. E. (1992). Intergenerational Patterns of Teenage Fertility. *Demography*, *29* (1): 39-57.

Kane, J. B., Morgan, S. P., Harris, K. M., & Guilkey, D. K. (2013). The Educational Consequences of Teen Childbearing. *Demography*, *50*: 2129-2150.

Kissin, D., Anderson, K. E., Kraft, J. M., Warner, L., & Jamieson, D. J. (2008). Is There a Trend of Increased Unwanted Childbearing Among Young Women in the United States? *Journal of Adolescent Health*, *43*: 364-371.

Lichter, D. T., Batson, C. D., & Brown, J. B. (2004). Welfare Reform and Marriage Promotion: The Marital Expectations and Desires of Single and Cohabitating Mothers. *Social Service Review*, *78* (1): 2-25.

Ludwig, J., Duncan, G. J., Gennetian, L. A., Katz, L. J., Kessler, R. C., Kling, J. R., & Sanbonmatsu, L. (2013). Long-Term Neighborhood Effects on Low-Income Families: Evidence from Moving to Opportunity. *NBER Working Paper*, 18772.

Manlove, J. (1998). The Influence of High School Dropout and School Disengagement on the Risk of School-Age Pregnancy. *Journal of Research on Adolescence*, *8* (2): 187-220.

Manlove, J. (1997). Early Motherhood in an Intergenerational Perspective: The Experiences of a British Cohort. *Journal of Family and Marriage*, *59* (2): 263-279.

Martin, J. A., Hamilton, B. E., Osterman, M. J. K., Curtin, S. C., & Matthews, T. J. 2013. "Births: Final Data for 2012." *National Vital Statistics Reports*, 62 (9).

Matthews, T. J., & Hamilton, B. E. (2009). Delayed Childbearing: More Women are Having Their First Child Later in Life. *National Center for Health Statistics Data Brief*, 21.

McLanahan, S. (2009). Fragile Families and the Reproduction of Poverty. *Annuals, American Academy of Political and Social Science*, *621*: 111-131.

Meade, C. S., Kershaw, T. S., & Ickovics, J. S. (2008). The Intergenerational Cycle of Teenage Motherhood: An Ecological Approach. *Health Psychology*, *27* (4): 419-429.

Mollborn, S., & Lovegrove, P. J. (2010). How Teenage Fathers Matter for Children: Evidence from the ECLS-B. *Journal of Family Issues*, *32* (3): 3-30.

Moore, K. A., Manlove, J., Glei, D. A., & Morrison, D. R. (1998). Non-Marital School-Age Motherhood: Family, Individual, and School Characteristics. *Journal of Adolescent Research*, *13*: 433-457.

North Carolina Education Research Data Center. http://www.childandfamilypolicy.duke.edu/project_detail.php?id=35.

Thomas, D. (1994). Like Father, Like Son; Like Mother, Like Daughter: Parental Resources and Child Height. *Journal of Human Resources*, *29* (4): 950-988.

U.S. Department of Health and Human Services, Office of Adolescent Health. *Overview: Teenage Pregnancy and Childbearing*. http://www.hhs.gov/ash/oah/adolescent-health-topics/reproductive-health/teen-pregnancy/. Accessed 15 March 2014.

World Bank. *Adolescent Fertility Rate (Births per 1,000 Women Ages 15-19)*. http://data.worldbank.org/indicator/SP.ADO.TFRT. Accessed July 15 2014.

Table 1 Characteristics and Educational Outcomes of Women Born in North Carolina in 1987-1989, by Teenage Childbearing Outcome

	All	% of sample Teenage mother	Non-teenage mother
Family-level characteristics			
Maternal race			
Non-Hispanic white	69.43	55.67	72.15
Non-Hispanic black	30.57	44.33	27.85
Maternal age at birth			
19 or below	28.70	49.74	24.54
20-25	39.26	38.12	39.48
26 or above	32.04	12.15	35.97
Maternal education at birth			
Less than high school	25.48	47.78	21.06
High school	41.36	40.91	41.45
More than high school	33.17	11.31	37.49
Paternal age at birth			
19 or below	6.91	10.92	6.12
20-25	29.73	31.39	29.41
26 or above	42.99	22.15	47.11
Age unknown	20.37	35.54	17.37
Paternal education at birth			
Less than high school	15.87	24.76	14.11
High school	35.17	29.43	36.31
More than high school	26.65	7.51	30.43
Education unknown	22.31	38.30	19.14
Parental marital status at birth			
Married	66.22	44.83	70.45
Number of younger siblings (born to mother)			
None	17.07	12.81	17.91
One or two	43.66	42.24	43.94
More than two	39.27	44.95	38.14
Neighborhood-level characteristics			
School			
Student poverty rate > 55%	48.13	59.31	45.92
Student poverty rate unknown	2.27	2.78	2.17
Proportion of students passing tests > 80%	56.55	45.87	58.66
Proportion of students passing tests unknown	0.57	0.56	0.57
Number of crimes per 100 students > 0	18.21	21.11	17.64
Number of crimes unknown	1.30	1.52	1.26
School district			
Child poverty rate > 17%	39.91	45.07	38.88
Child poverty rate unknown	0.69	0.95	0.64
Proportion of students passing tests > 83%	43.79	38.98	44.74
Proportion of students passing tests unknown	0.00	0.00	0.00
Proportion of students with one parent > 23%	47.88	52.51	46.96
Proportion of students with one parent unknown	0.69	0.95	0.64

Educational outcomes in 3rd to 5th grade			
Test scores			
Reading Z-score < -1	11.27	19.42	9.66
Reading Z-score > 1	17.88	5.32	20.37
Reading Z-score unknown	1.13	1.31	1.09
Math Z-score < -1	12.81	21.13	11.16
Math Z-score > 1	15.15	4.50	17.25
Math Z-score unknown	1.11	1.28	1.07
Grade progression			
Retained by 5th grade	11.42	16.75	10.36
Number of observations	35,544	5,869	29,675

Notes: Data for family-level characteristics are obtained from linked North Carolina birth certificates, while neighborhood-level characteristics and educational outcomes in third to fifth grade are obtained from North Carolina public school administrative records, with permission from the North Carolina Department of Health and Human Services and North Carolina Research Data Center. The values for neighborhood-level characteristics and educational outcomes in third to fifth grade include imputed values.

Table 2 Probability of Teenage Motherhood and Being Married Conditional on Teenage Birth, 1950s-1980s and 2002-2009

	All	Whites	Blacks
Probability of teenage motherhood			
1950s-1980s	-	0.193	0.418
2002-2009	0.165	0.132	0.240
Probability of being married (conditional on teenage birth)			
1950s-1980s	-	0.772	0.220
2002-2009	0.109	0.182	0.165

Notes: Data for 1950s-1980s are taken from Kahn and Anderson (1992).

Table 3 Probability of Teenage Motherhood, Decomposed and Standardized by Maternal Age

	Decomposition					
	All		Whites		Blacks	
	%	Prob.	%	Prob.	%	Prob.
Mother was aged 19 or below	28.7	0.286	21.6	0.268	44.7	0.305
Mother was aged 20-25	39.3	0.160	40.3	0.139	36.9	0.213
Mother was aged 26 or above	32.0	0.063	38.1	0.048	18.3	0.131
Total probability	0.165		0.132		0.239	
	Standardization (using values for whites)					
	Same distribution of maternal ages		Same probabilities of teenage motherhood		Same probabilities of teenage motherhood among women born to adult mothers	
			Blacks			
	%	Prob.	%	Prob.	%	Prob.
Mother was aged 19 or below	21.6	0.305	44.7	0.268	44.7	0.305
Mother was aged 20-25	40.3	0.213	36.9	0.139	36.9	0.139
Mother was aged 26 or above	38.1	0.131	18.3	0.048	18.3	0.048
Total probability	0.202		0.180		0.197	

Table 4 Mediators of Maternal Age Effects on Probability of Teenage Motherhood (dependent variable = giving birth before age 20)

	All		Whites		Blacks	
	1	2	3	4	5	6
Family-level characteristics						
Maternal age at birth						
19 or below	-	-	-	-	-	-
20-25	-0.126***	-0.018*	-0.130***	-0.035***	-0.093***	0.010
26 or above	-0.224***	-0.053***	-0.220***	-0.074***	-0.174***	-0.015
Maternal education at birth						
Less than high school		-		-		-
High school		-0.081***		-0.080***		-0.087***
More than high school		-0.119***		-0.107***		-0.162***
Paternal age at birth						
19 or below		-		-		-
20-25		0.011		0.015		0.007
26 or above		0.009		0.016		0.004
Age unknown		0.042**		0.070**		0.025
Paternal education at birth						
Less than high school		-		-		-
High school		-0.065***		-0.075***		-0.035†
More than high school		-0.088***		-0.099***		-0.062**
Education unknown		-0.055***		-0.046*		-0.045†
Parental marital status at birth						
Married		-0.040***		-0.024*		-0.034**
Number of younger siblings						
None		-		-		-
One or two		0.014**		0.012*		0.019†
More than two		0.053***		0.038***		0.079***
Neighborhood-level characteristics						
School						
Student poverty rate > 55%		0.014**		0.009†		0.018†
Student passing rate > 80%		-0.013**		-0.010†		-0.014
Crimes per 100 students > 0		0.005		0.006		0.002
School district						
Child poverty rate > 17%		0.004		0.003		0.008
Student passing rate > 83%		0.002		0.004		-0.001
Students w/one parent > 23%		0.005		0.002		0.007
Educational outcomes for 3rd to 5th grade						
Test scores						
Reading Z-score < -1		0.035***		0.041***		0.028*
Reading Z-score > 1		-0.033***		-0.029***		-0.055***
Math Z-score < -1		0.014†		0.006		0.016
Math Z-score > 1		-0.023***		-0.020***		-0.045*
Grade progression						
Retained by 5th grade		0.001		0.003		-0.004
Constant	0.290	0.282	0.271	0.287	0.312	0.252
R^2	0.055	0.095	0.059	0.100	0.024	0.053
Number of observations	35,544		24,678		10,866	

Notes: Coefficients are marginal effects relative to the omitted class, estimated using a linear regression model. All regressions include cohort fixed effects.

†p <.10; *p < .05; **p < .01; ***p < .001

Table 5 Decomposition of Maternal Age Effects on Teenage Childbearing into Effects Associated with Maternal Education, Paternal Characteristics, Parental Marital Status, Family Size, Neighborhood Characteristics and Early Educational Outcomes

	All	Whites	Blacks
Maternal age = 20-25			
Maternal education	-0.057***	-0.051***	-0.074***
Paternal characteristics	-0.019***	-0.026***	-0.006
Parental marital status	-0.016***	-0.007*	-0.009**
Number of younger siblings	-0.004***	-0.002***	-0.008***
Neighborhood characteristics	-0.003***	-0.001***	-0.001**
Educational outcomes	-0.009***	-0.007***	-0.004***
Residual age effect	-0.018*	-0.035***	0.010
Total age effect	**-0.126**	**-0.130**	**-0.093**
Maternal age = 26 or above			
Maternal education	-0.076***	-0.068***	-0.097***
Paternal characteristics	-0.037***	-0.045***	-0.014
Parental marital status	-0.023***	-0.009*	-0.020**
Number of younger siblings	-0.007***	-0.004***	-0.014***
Neighborhood characteristics	-0.007***	-0.003**	-0.004***
Educational outcomes	-0.020***	-0.017***	-0.010***
Residual age effect	-0.053***	-0.074***	-0.015
Total age effect	**-0.224**	**-0.220**	**-0.174**

Notes: Values within each box refer to the estimated contribution of each component to maternal age effects (relative to maternal age = 19 or below). Coefficients are estimated using the Gelbach decomposition method.
*p < .05; **p < .01; ***p < .001

Fig. 1 Teenage Motherhood Survival Curves

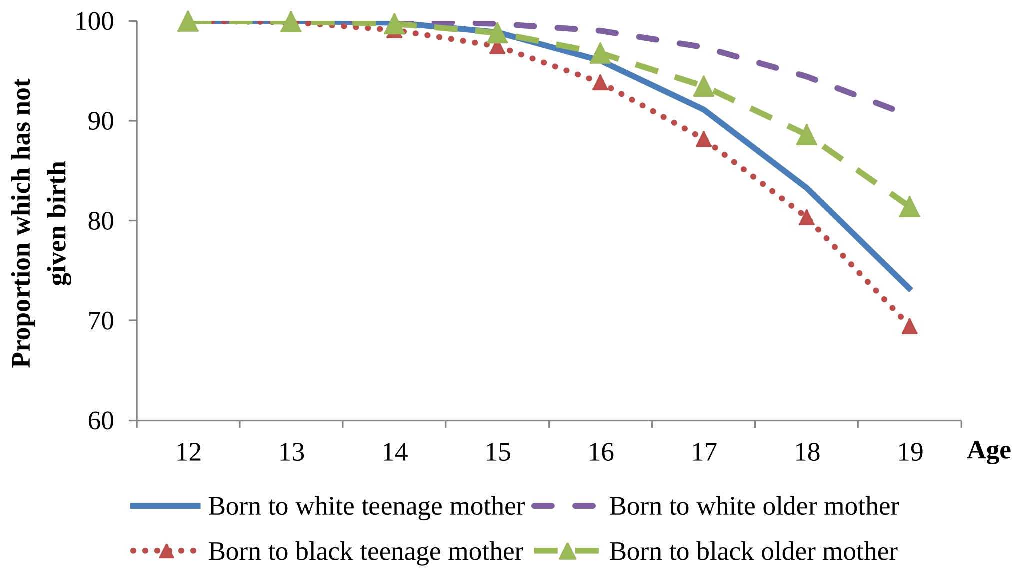
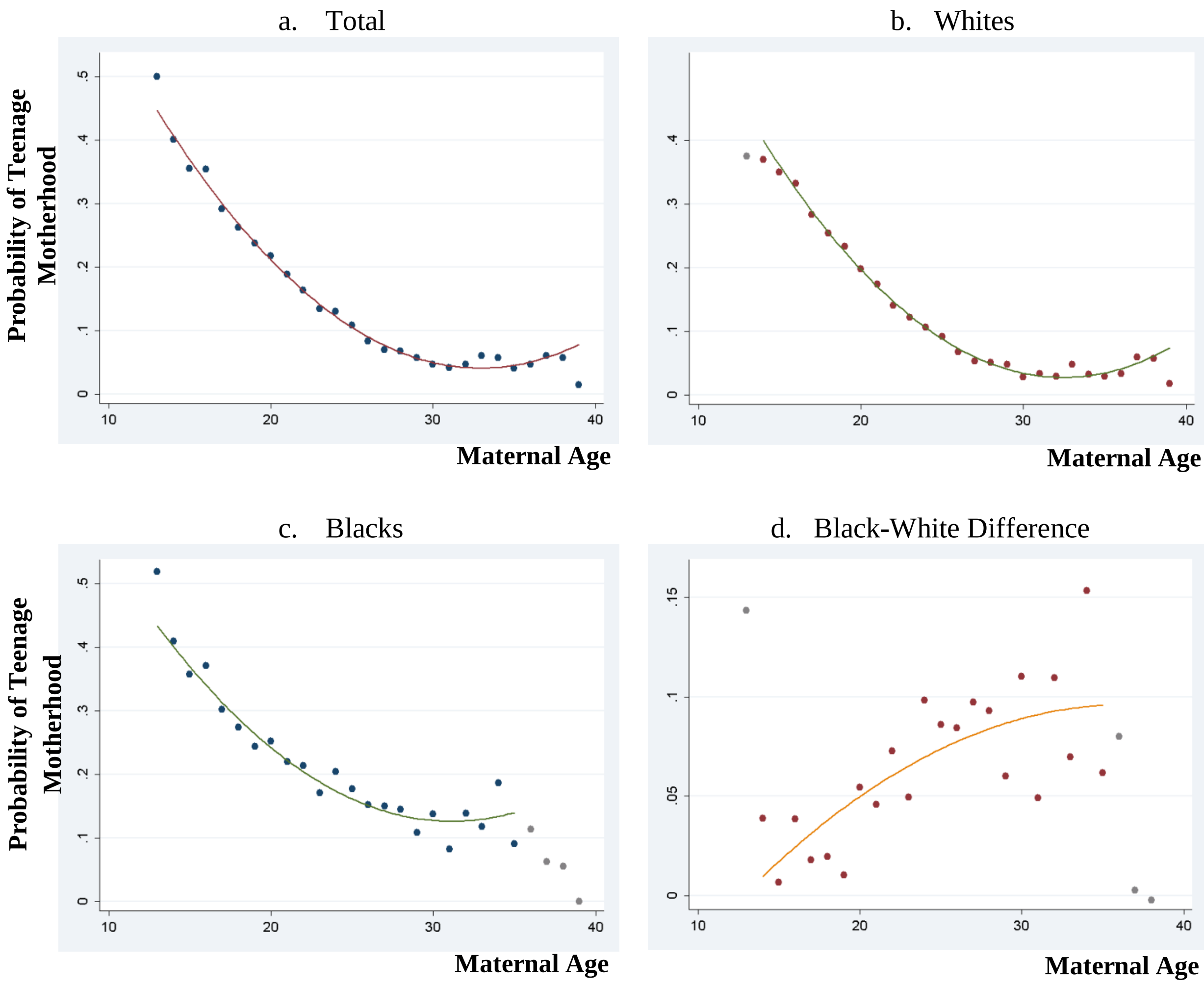


Fig. 2 Probability of Teenage Motherhood by Maternal Age



Notes: The gray data points represent fewer than 50 observations. The curves are fitted using a quadratic functional form which assigns equal weight to each data point.