EARLY CHILDBEARING, UNION STATUS, AND WOMEN'S HEALTH AT MIDLIFE*

Kristi Williams[a]

Sharon Sassler[b]

Fenaba Addo[b]

Elizabeth Cooksey[a]

Adrianne Frech[c]

[a]Department of Sociology, The Ohio State University

[b]Department of Policy Analysis and Management, Cornell University

[c]Department of Sociology, The University of Akron

Corresponding Author: Kristi Williams, Department of Sociology, 238 Townshend Hall, The Ohio State University, Columbus OH 43210; williams.2339@osu.edu; Phone: 614-688-3207; Fax: 614-292-6687

EARLY CHILDBEARING, UNION STATUS, AND WOMEN'S HEALTH AT MIDLIFE

ABSTRACT

Despite the prevalence of early and nonmarital childbearing, little is known about their long-term consequences for women's mental and physical health over the life course or about the role that subsequent marriage might play in ameliorating or perhaps exacerbating any negative effects. We use 29 years of panel data from the National Longitudinal Survey of Youth 1979 to examine differences in age-40 self-assessed health between women who had an adolescent birth (prior to age 20) or young adult first birth (age 20-23) and those whose first birth occurred at later ages. We then differentiate women who gave birth during adolescence or young adulthood by their marital status at birth and after the birth to estimate the effect of marriage on the midlife self-assessed health of young mothers. To partially address selection bias, we employ multivariate propensity score matching. Results suggest few negative long-term health consequences of early childbearing except among black women who have their first birth in young adulthood (age 20-23). Among those who have an early first birth, marriage at birth appears beneficial for the midlife health of white women, while marriage after a nonmarital early first birth may pose health risks for black and Hispanic women.

EARLY CHILDBEARING, UNION STATUS, AND WOMEN'S HEALTH AT MIDLIFE

Although the adolescent birth rate in the United States has declined in recent years, approximately 10% of all births (20% of first-births) occur to women age 19 or younger—a higher rate than any other developed country (Martin et al. 2010). An additional 25% percent of births (30% of first-births) in the U.S. occur during early adulthood (between the ages of 20 and 24). Moreover, the majority of these "early" births (prior to age 24) occur to unmarried women. Despite the prevalence of early and nonmarital childbearing, little is known about their long-term consequences for women's mental and physical health over the life course or about the role that subsequent marriage might play in ameliorating or perhaps exacerbating any negative effects.

We use 29 years of panel data from the National Longitudinal Survey of Youth 1979 to examine differences in age-40 self-assessed health between women who had an adolescent birth (prior to age 20) or young adult first birth (age 20-23) and those whose first birth occurred at later ages. We then differentiate women who gave birth during adolescence or young adulthood by their marital status at birth and after the birth to estimate the effect of marriage on the midlife self-assessed health of young mothers. To partially address selection bias, we employ multivariate propensity score matching.

Early Childbearing, Race/Ethnicity, and Women's Health

Several U.S. studies find that, at least up until about age 30, earlier ages at first birth are associated with worse short-term and long-term health outcomes across numerous domains, including physical impairment (Mirowsky 2002), serious chronic conditions (Mirowsky 2002), mortality (Henretta 2007; Mirowsky 2005) and an index of health problems (including self-assessed health) (Mirowsky 2005). Research on European populations is generally consistent with this pattern (Doblhammer 2000; Grundy and Tomassini 2005). Although biological

processes may be implicated in linking childbearing patterns to later life health, the general pattern of findings to date are more suggestive of a role of social processes. Strictly biodevelopmental perspectives suggest that childbearing that occurs early in the life course, when the organism is young and biologically resilient, should produce *better* long term health outcomes (Godsen and Rutherford 1995; Mirowsky 2005), but the evidence to date has failed to support such a hypothesis. In contrast, social processes associated with early childbearing clearly suggest negative outcomes. Early childbearing is linked to lower levels of educational and socioeconomic attainment, higher rates of subsequent marital and family instability, and increased stress throughout the life course. All of these factors, in turn, can have cumulative negative consequences for health and well-being later in life.

Despite the evidence that early childbearing is linked to poorer health outcomes among women, extant research has been limited in a number of ways. The greatest limitation is a lack of attention to the threat of selection in inferring a *causal* role of early childbearing on health. This is an important policy question as efforts aimed at reducing early childbearing are premised on the view that it causes a range of negative individual and society outcomes. However, it is well-established that women who begin childbearing in their adolescent years differ substantially on a number of background characteristics that are themselves strongly associated with health. As a result, it may be the risk factors that predispose women to adolescent or early childbearing, rather than the fertility timing itself, that are responsible for the negative association of early childbearing with health later in life. Our data allow us to control for an unprecedented range of background characteristics predictive of entry into early childbearing, including socioeconomic and family background and cognitive ability. We also employ multivariate propensity score analysis to determine whether significant observed associations of early childbearing with

midlife health persist when women who had a early first birth are matched with those who have a similar estimated propensity of having an early first birth, based on their background characteristics.

A second limitation of prior research is a lack of attention to race/ethnic differences in the effects of early childbearing on women's health. This is an important consideration in the U.S. where rates of early and nonmarital childbearing are substantially higher for black, compared to white and Hispanic women. However, substantial evidence casts doubt on the dominant cultural belief that *early* childbearing has widespread individual and social costs for low income African American women (Geronimus 2003). As Geronimus has argued (1996), early fertility may be an adapative strategy for low income urban African-American women vulnerable to "weathering"-- accelerated declines in health that pose substantial challenges to bearing and raising children at older ages. On the other hand, evidence that African Americans have more difficulty accessing high-quality health care services than whites (Smedley et al. 2003) could mean that black women may be especially vulnerable to any long-term negative health consequences of early childbearing. The consequences for Hispanic women may also differ. Although no prior research has examined the relative effects of early childbearing among Hispanic compared to white women, Williams et al. (2011) find that the negative long-term health consequences associated with *nonmarital* childbearing observed among white and black women do not extend to Hispanic women. Although this countervailing evidence does not support strong directional hypotheses, both demographic and ecological factors warrant consideration of potential race/ethnic differences.

A third limitation of prior research on early childbearing is an exclusive focus on the consequences of *teen childbearing* for women's health, with no attention given to the impact of

childbearing that occurs in the young adult years. Using data from the British Household Panel Survey, Francesconi (2008) recently found that negative outcomes for *children* associated with birth timing are not limited to those born to teen mothers but also extend to those born to mothers in their early 20s. Thus, "public policies targeting only teenage parenthood may not be enough to address the social exclusion and disadvantages of all young parents and their families (Francesconi 2008: 94-95)." It is unknown whether similar patterns exist in the U.S. where the average age at first birth (25) is substantially lower than in the U.K. (29.1) or whether *mothers'* health (as opposed to children's outcomes) is negatively affected by young adult childbearing in any context. Addressing this question is especially important in a U.S. context given the demographics of birth timing. In the United States, the highest first birth rates occur in the 20-24 age group. Our analyses considers both the long term health consequences of *adolescent* childbearing (births that occur to women aged 19 or younger) and *early adult* childbearing (births that occur to women aged 20-23). For brevity, we refer to the combination of both groups as *early childbearing*.

Early Childbearing, Subsequent Union Status, and Women's Health

An important consideration in attempting to assess a causal impact of early childbearing with later life health is the ability to unravel the unique effects of age and marital status at birth. Early first births are more likely than later births to occur outside of a marital union and nonmarital fertility has been linked to poor health outcomes among women in U.S. (Henretta 2007; Williams et al. 2011) and European populations. Although little research has directly examined the mechanisms linking nonmarital fertility with poor health, subsequent biosocial processes of social disadvantage, family instability and chronic stress exposure (similar to those theorized to be relevant in early childbearing) likely play important roles. Determining the extent

to which any negative health consequences of early childbearing operate through their effects on nonmarital fertility has substantial relevance for family and public health policy. We address this question in two ways. First, we estimate the extent to which any negative long-term health consequences of early childbearing are explained by the association of early- and nonmarital childbearing. Second, we limit our analysis to women who had an early first-birth and consider whether those who are married at birth or those who are unmarried at birth but later marry have better outcomes than their counterparts who are married at the time of their early first birth, giving attention to potential race/ethnic differences and employing multivariate propensity score analysis to partially address selection processes.

DATA AND MEASURES

Data

The 1979 National Longitudinal Survey of Youth (NLSY79) data are uniquely suited to investigating long-term consequences of early childbearing. The NLSY79 is an ongoing survey of a nationally representative sample of 12,686 young men and women ages 14-22 in 1979; it originally included oversamples of black, Hispanic, military and poor white respondents. Although the military and economically disadvantaged white oversamples were dropped prior to 1991, the remaining respondents were interviewed annually through 1994 and biennially since. By 2008, data on detailed union histories had been continuously collected over a 29 year period, and measures of health and well-being ascertained at age 40, a time when health problems begin to emerge. As of 2006, all NLSY79 mothers were over age 40.

A total of 3,986 women had given birth prior to age 40, and 3,471 (89.3%) of them completed the age 40 health assessment. We further limit this sample to the 3,896 women whose first birth took place prior to age 36 and who therefore had sufficient time to enter a union prior

7

to reaching age 40; of this sub-sample, 3,391 (89.1%) completed the self-assessed health measure. We exclude 491 (14.48%) respondents who were missing data on additional variables included in the analysis The amount of data missing was most sizable for maternal years of education, poverty status, and AFQT score. Separate models are estimated for each of the three categories of race/ethnicity: (1) Non-Hispanic, non-black (hereafter described as "White") (n = 1,413) (2) Black (n = 901) and (3) Hispanic (n = 591).Propensity score matching was implemented using the psmatch2 program created for STATA by Leuven and Sianesi (2003). Bootstrapped standard errors were calculated for all matching estimations.

Measures

Self-assessed Health is measured at age 40 with a single question: "In general, would you say your health is excellent, very good, good, fair, or poor?" Responses are coded 1-5 with higher values indicating better health. Self-assessed health is highly predictive of subsequent morbidity and mortality (Idler and Benjamini 1997) and is widely recognized as a valid and reliable indicator of health status in the general population (Ferraro and Farmer 1999).

Age at First Birth. Dummy variables distinguish women whose first birth occurred: (a) during adolescence (prior to age 19), (b) in early adulthood (age 19-23) and (c) age 24 or older (up to age 36). The reference category is age 24-36.

Marital History. Dummy variables differentiate women by their marital status at the birth of the first child and their subsequent marital history: (a) unmarried at first birth and remained unmarried through age 40, (b) unmarried at birth and married by age 40 and (c) married at first birth (reference category).

Control Variables include US nativity (1=foreign born), cognitive ability (1980 AFQT score, which ranges from 1-99), poverty status at baseline (1 = in poverty, based on whether the

family income was below the federal poverty income guidelines for the particular family size in 1978), family composition at age 14 (1=lived with both biological parents), and the respondent's mother's years of education as a proxy for the socioeconomic status of her family of origin.

Although the number of biological children a woman has by age 40 and her marital status at age 40 likely influence health, they are consequences rather than causes of nonmarital fertility and subsequent union history. Controlling for these potential mechanisms would underestimate the true total effect of single motherhood or subsequent union history on midlife health and they are, therefore, excluded from the models presented. However, supplementary analyses revealed that their inclusion did not change the overall pattern of findings, suggesting their role as potential mediators is negligible.

Analysis

Our analyses address two basic questions: (1) Does adolescent/early parenthood undermine mid-life health and to what extent is this explained by the association of adolescent/early parenthood with nonmarital childbearing? And (2) Among women who had an adolescent or early first birth, how does marital status at the child's birth and subsequent marital history affect midlife health? We use OLS, ordered probit regression, and multivariate propensity score matching to address these questions.[1]

Selection into early childbearing and subsequently forming unions is a nonrandom process. In order to identify the effect of early childbearing and later union transitions on mother's health, we minimize the potential bias due to nonrandom selection, by conditioning on pre-treatment observable characteristics (Rosenbaum and Rubin 1983). We match women in our sample based on the predicted probability they had an early first birth or experienced a particular union history (the propensity score) as a non-parametric feature of these pre-treatment (or

predetermined) observed characteristics (Dehejia and Wahba 2002). PSM models thus estimate the average effect of having an early first birth (or, in later models, union history) among women similarly likely to engage in such behaviors. Thus, in the second stage of each part of our analysis, we determine whether significant associations identified in the first-stage OLS models are robust to a propensity score matching specification. As Morgan and Harding (2006) argue, the use of matching techniques provides a sizable advance over merely conceding that selection bias may be present in some form and speculating on the sign of the bias.

We use one-to-one nearest neighbor matching with replacement (Morgan and Harding 2006; Rosenbaum and Rubin 1983). Nearest neighbor matching (with and without replacement) and other approaches (interval matching, kernel matching) are closely related. There are not yet standardized guidelines for choosing a matching procedure (Gangl 2010; Morgan and Winship 2007), though nearest neighbor matching estimates have been shown to produce lower standard errors, compared with caliper matching, and are preferred when there is a large overlap in propensity scores between treatment and control groups (Black & Smith, 2004; Dehejia & Wahba, 2002). Matched observations from the treatment and control groups are used to compute differences in health at age 40, and the average difference is then computed across all matches. Where necessary, we trim the sample to ensure common support (i.e., that the range of propensity scores is the same for treatment and control groups). We use bootstrapped standard errors ($n = 50$) to ensure the most robust outcomes for our PSM results.

Descriptives

Table 1 shows descriptive statistics (means and standard deviations [in parentheses] or percentages) for all variables by race/ethnicity.

(Tables 1 and 2 about here)

RESULTS

Midlife Health of Women with an Adolescent / Early First Birth

Table 2 presents results from OLS regression models comparing the health at age 40 of

women who had an adolescent or early first birth to that of their counterparts who had a first

birth between the ages of 24 and 36 by race/ethnicity.[2] For each race/ethnic group, the first

models show the estimated effect of early childbearing before conditioning on background

characteristics or marital status at birth. As shown in Models 1 (white women), 2 (black women)

and 3 (Hispanic women), prior to conditioning on a wide range of background characteristics,

having a first birth in adolescence or young adulthood is associated with worse midlife health for

white and black, but not Hispanic, women. Controlling for background characteristics reduces

these estimated effects but they remain statistically significant for black and white women.

Notably, among white women, background characteristics appear to explain approximately 67%

of the initial association of adolescent childbearing and 47% of the association of childbearing in

young adulthood with midlife health. A smaller proportion of the estimated effect is explained

for white women (28.5% of the effect of adolescent childbearing and 11.8% of the effect of

childbearing in young adulthood).

The final set of models for each race/ethnic group (Models 3, 6, and 9) enter a control for

marital status at birth to determine whether the consequences of early childbearing with midlife

health are partially mediated by marital status at birth (i.e., the fact that young mothers are more

likely to be single mothers). The only group for whom this is relevant is white women with an

adolescent first birth. The estimated effect of adolescent childbearing among white women is

reduced by 50% to nonsignificance after marital status at birth is controlled. Marital status at

birth contributes less to the estimated effect of adolescent childbearing among black women

(27.95%) and to the estimated effect of childbearing in young adulthood among white (8.1%) and black (15.7%) women, and these associations remain statistically significant after it is controlled. For Hispanic women, neither adolescent / early first births nor nonmarital childbearing is negatively associated with health at midlife in any of the three models. In separate analyses (not shown) we entered interaction terms to test whether the estimated effect of age at first birth on midlife health depended on marital status at first birth or subsequent union history. None of the coefficients for these interaction terns were significantly different from zero. Although our aim is not to identify mechanisms through which early childbearing produces long-term health consequences, substantial evidence of a positive association of parity with health problems in later life (Grundy et al.) lead us to consider whether the consequences of early childbearing were mediated by the number of children born to the respondent by age 40. However, we found no evidence to support this hypothesis for any race/ethnic group in our analysis.

(Table 3 about here)

We next employ multivariate propensity score matching to determine whether the significant associations presented in Table 3 are robust to an approach that better accounts for the differential selection of women into adolescent / early parenthood.   This analysis estimates the likelihood of experiencing an adolescent or early first birth for the total sample of women who became mothers, by conditioning on pre-treatment observable characteristics (Rosenbaum and Rubin 1983). Covariates used to estimate the propensity for adolescent / early first birth include, family structure at age 14, cognitive abilities, poverty status at baseline, respondent's mother's years of education, contraceptive use prior to first birth, baseline educational expectations (1 = expect a college degree), frequency of religious attendance, health limitations that began prior to

the first birth, and traditional gender roles. Significant predictors of adolescent or early first births varied across models but generally included mothers' AFQT score, family structure at age 14, contraceptive use at first birth, nativity, and educational expectations

Table 3 presents the average treatment effects for the treated, estimated from the PSM models. In the first panel, the treatment refers to experiencing a first birth under age 20 (compared to age 24-35). The results for white women are consistent with the OLS results as neither set of analyses indicates that having an early first birth is associated with the midlife health of white women. The results for black women, however, are inconsistent with the OLS results which suggested that black women whose first birth occurred prior to age 20 have worse midlife health than their counterparts whose first birth occurred age24-35. Here, the ATT estimate (-.183) is substantially smaller than the OLS coefficient (-.351) and is not significantly different from 0. Thus, after accounting for differential selection into early childbearing, having an adolescent first birth is not associated with worse midlife health for black women. In the second panel, the treatment refers to experiencing a first birth between the ages of 20-23 (compared to age 24-35) and the ATT is estimated separately for black and white women. The OLS results in Table 2 indicated that first births occurring between the ages of 20-23 are negatively associated with midlife health for both white (-.124*) and black (-.220** women). The PSM models in Table 3 indicate that, for white women, this difference is substantially smaller (-.098) and no longer statistically significant whereas for black women the size of the coefficient increases slightly (.298*) and remains significant. In sum, the PSM analyses leads to the following conclusions. First, OLS estimates of the negative consequences of adolescent first births among black women and first births in young adulthood (age 20-23) are largely due to differential selection into adolescent childbearing. However, even among black women with a

similar estimated propensity to have an early first birth in young adulthood (age 20 to 23), those who actually have early first birth have significantly lower levels of age-40 self-assessed health than their counterparts whose first birth occurs between the ages of 24 and 35. In sum, adolescent first births are not significantly associated with worse health at midlife for white, black, or Hispanic women and early first births in young adulthood (age 20-23) are negatively associated with midlife health only among black women. We find no evidence of negative health consequences of adolescent / early childbearing for Hispanic women. [1]

Nonmarital Childbearing, Subsequent Union History and Health at Midlife

In the second part of the analysis, we limit the sample to women who had an early (adolescent or young adult) first birth (prior to the age of 24) and use OLS regression and, in supplementary analyses (available upon request), ordered probit to examine whether marital status at birth and marital history after birth are significantly associated with midlife health. The central question addressed here is whether marriage at the time of birth or subsequently is associated with better (or perhaps worse) midlife health outcomes for women who have an adolescent / early first birth. Results are shown in Table 4 separately for white, black, and Hispanic women. In the first model for each race/ethnic group (odd-numbered models), the "unmarried" coefficient shows the estimated difference in age-40 self-assessed health of women who had an adolescent early/first birth while unmarried compared to their counterparts who had an adolescent / early first birth while married. For this sample of women who had an early first birth, being unmarried at the first child's birth is significantly associated with worse midlife health than is being married for white women. Marital status at birth appears to have no consequences for the long term midlife health of black or Hispanic women with early first births.

---

[1] Consistent with the OLS results, propensity score models for Hispanic women (not shown) indicated no significant association of adolescent or young adult childbearing with the midlife health of Hispanic women.

In the second model for each race/ethnic group (even-numbered models) we further disaggregate women who were unmarried at the first birth to differentiate those who subsequently married from those who remained unmarried. The aim is to consider whether subsequent marriage ameliorates (or perhaps exacerbate) any negative consequences associated with being unmarried at birth for the midlife health of women who experienced adolescent/early parenthood. Results indicate that, among white women with early first births, the negative health consequences of being unmarried at birth continue to be observed both among those who subsequently marry and those who remain unmarried through age 40. Examining the size of the coefficients for the disaggregated marital status dummy variable suggests that white young single mothers who subsequently marry may have worse health than those who remain never-married but supplementary analysis (not shown) indicates that this difference is not statistically significant. Nevertheless, we explore this contrast further in the propensity score models that follow.

In contrast to the results for white women, Models 4 and 6 suggest that the negative consequences of being unmarried at birth for black and Hispanic young mothers depend on subsequent union history. Interestingly, only those black and Hispanic young single mothers who subsequently marry have worse age-40 health than their counterparts who were married at birth. Single mothers in these groups who remain single through age 40 have no worse health at midlife than their counterparts who had an early first birth while married. Subsequent analysis (not shown) confirms that young black and Hispanic single mothers who subsequently marry have worse health at midlife than their counterparts who are single at birth and never marry. In sum, the OLS analyses reveal that for white young mothers, being unmarried at birth undermines midlife health and subsequent union history does little to alter this pattern. However,

for black and white young mothers, being unmarried at birth is only associated with worse health outcomes for those who subsequently marry. It appears that for Hispanic and black young mothers, marriage after the birth of the child poses long term health risks.

(Table 4 about here)

(Table 5 about here)

In the final set of analyses, we employ multivariate propensity score models to determine whether these significant associations of marital status at birth and marital history with the midlife health of young mothers shown in Table 4 are robust to a consideration of differential selection into marriage (at birth or later). We first estimate the likelihood of occupying each of the marital history categories that we consider separately for white, black and Hispanic women, employing 3 contrasts: (1) "unmarried at birth but later married" compared to "married at birth," (2) "unmarried at birth and never married by age 40" compared to "married at birth," and (3) "unmarried at birth and later married" compared to "unmarried at birth and never married by age 40." The covariates used to predict marital history include family structure at age 14, cognitive abilities, poverty status at baseline, respondent's mother's years of education, contraceptive use prior to first birth, baseline educational expectations (1 = expect a college degree), frequency of religious attendance, health limitations that began prior to the first birth, and traditional gender roles.

Table 5 presents the average treatment effects among the treated from the PSM models. Results for black women are shown in the first column. Recall that the OLS results in Table 4 indicated that, for black women who had an early first birth, those who were unmarried at birth and later married report lower levels of midlife self-assessed health than those who were married at birth (-.215*). The results from the PSM analyses in Table 5 strengthen these findings and

16

indicate that they persist and increase in magnitude (-.474***) after accounting for differential selection into an early first birth. Marriage after the first birth, however, continues to be unrelated to midlife health for black women with nonmarital early first births, compared to those who were married at their first child's birth (-.080 in OLS models; -.027 in PSM models) and compared to those who were unmarried at birth and never-married by age 40 (-.126). In other words, the negative effects of having a nonmarital first birth among black young mothers are limited to those who subsequently marry.

Estimating similar PSM models for white and Hispanic women is complicated by the small number of women who were unmarried at birth but later married. Therefore, we are only able to estimate PSM models using a combined white and Hispanic subsample for contrasts involving this group. Recall that the OLS models indicated that being unmarried at the child's birth was negatively associated with self-assessed health for white young mothers regardless of their subsequent union history (-.455* among white women who later marry; -.239* among white women who never marry by age 40). The PSM models support this pattern, although we are only able to estimate the effect of being unmarried and later marrying (compared to having a marital first birth) in the combined sample of white and Hispanic women (-.578*). Our PSM estimate of the effect of being unmarried at birth and never-marrying compared to being married at birth separately for white women support the OLS findings. White women with early first births who never marry have worse age-40 health than those whose first birth occurred in marriage (-.239*** OLS; -.244+ PSM), although the difference is only significant at the .10 level in the PSM models.

The results for Hispanic women parallel those for black women. The OLS models indicated that the negative effects of nonmarital fertility on midlife health among Hispanic young

mothers are limited to those who subsequently marry (-.424* OLS among those who subsequently marry; .128 OLS among those who never marry). The PSM models are consistent with these findings although we are limited to considering the combined group of Hispanic and white women (-.578* PSM among those who subsequently marry). Similar to the OLS models for Hispanic women and the OLS and PSM models for black women, there appears to be no difference in the midlife health of Hispanic women with a marital compared to nonmarital first birth as long as those with a nonmarital first birth never marry (.128 OLS; .282 PSM).

DISCUSSION

Considered together, our results suggest substantial race/ethnic heterogeneity in the estimated long term health consequences of (a) having an early first birth and (b) nonmarital childbearing and union history among those who have an early first birth. For white women, although early first births themselves are not associated with long term negative health outcomes, those who have an early first birth report better health if they are married at the birth than if they are unmarried, and subsequent marital history among the unmarried does little to alter this pattern.

For black and Hispanic women, adolescent childbearing has no long-term negative health consequences (I found a study that uses instrumental variables and arrives at a similar conclusion but didn't look at race diffs or the 20-23 age category). In fact, the only negative health consequences of early fertility were observed among black women who have their first birth as young adults (age 20-23). Moreover, among black women and Hispanic women who have an early first birth, any negative consequences of nonmarital fertility are limited to those who subsequently marry. In other words, being unmarried at birth is not a health risk factor for black or Hispanic young mothers as long as they remain unmarried. Subsequent marriage,

18

however, may pose risks for this group. Black and Hispanic young single mothers who subsequently marry (but not those who remain never-married) have worse midlife health than their similar young mothers who were married at birth and, more importantly, worse health than those who, by age 40, have never-married.

Implications and limitations of these findings will be discussed further in the completed version of this paper.

---

[1] Because we obtained similar results with both techniques, we present the OLS results for ease of interpretation. Results of the ordered probit analyses are available upon request.

Table 1: Descriptive Statistics

	White			Black			Hispanic		
Age 1st birth	<20	20-23	24-36	<20	20-23	24-36	<20	20-23	24-36
Age 40 self-assessed health	3.53	3.63	3.88	3.27	3.33	3.63	3.40	3.42	3.54
Nonmarital 1st birth (0 = marital)	32.28%	12.81%	5.62%	84.70%	64.54%	40.09%	42.40%	24.71%	15.69%
Foreign-born mother (0 = native)	1.90%	3.54%	1.92%	0.91%	0.80%	6.60%	17.97%	28.82%	24.02%
Mother's education (years)	10.54	11.43	12.38	9.96	11.08	11.05	7.20	7.39	8.37
Lived with both parents age 14	71.52%	72.75%	85.48%	40.87%	50.20%	59.91%	58.06%	65.88%	75.98%
AFQT	38.13	48.23	60.51	17.91	21.87	29.16	21.12	23.96	35.86
In poverty 1979 (0 = not poverty)	20.25%	11.72%	6.03%	60.96%	40.24%	37.26%	43.78%	39.11%	26.47%
Observations	316	367	730	438	251	212	217	170	204

Table 2: OLS Regression Coefficients Predicting Age 40-Self Assessed Health: U.S. Women Who Had a First Birth Prior to Age 34

					Age 40 Self-Assessed Health				
		White			Black			Hispanic	
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Age at 1st birth									
<19	-0.358***	-0.118*	-0.059	-0.351***	-0.251***	-0.181*	-0.138	0.058	0.068
	(0.067)	(0.071)	(0.073)	(0.085)	(0.089)	(0.094)	(0.103)	(0.106)	(0.108)
20-23	-0.257***	-0.135**	-0.124*	-0.296***	-0.261***	-0.220**	-0.122	0.026	0.029
	(0.063)	(0.064)	(0.064)	(0.095)	(0.096)	(0.097)	(0.109)	(0.109)	(0.109)
(Ref = 24 -35)									
Unmarried at	---	---	-0.295***	---	---	-0.179**	---	---	-0.047
1st birth	---	---	(0.081)	---	---	(0.079)	---	---	(0.098)
Mother foreign	---	0.187	0.178	---	0.040	0.012	---	0.123	0.120
born	---	(0.173)	(0.172)	---	(0.231)	(0.231)	---	(0.103)	(0.103)
Mother's education	---	0.035***	0.034***	---	0.029**	0.029**	---	0.003	0.003
(years)	---	(0.012)	(0.012)	---	(0.014)	(0.014)	---	(0.012)	(0.012)
Lived with both	---	0.021	0.003	---	-0.162**	-0.165**	---	-0.066	-0.070
parents age 14	---	(0.065)	(0.065)	---	(0.070)	(0.070)	---	(0.092)	(0.092)
Cognitive ability	---	0.006***	0.006***	---	0.006***	0.006***	---	0.011***	0.011***
	---	(0.001)	(0.001)	---	(0.002)	(0.002)	---	(0.002)	(0.002)
In poverty 1979	---	-0.275***	-0.233***	---	-0.105	-0.082	---	-0.161*	-0.157*
	---	(0.087)	(0.087)	---	(0.074)	(0.074)	---	(0.093)	(0.094)
Constant	3.884***	3.087***	3.147***	3.623***	3.251***	3.324***	3.539***	3.163***	3.177***
	(0.037)	(0.157)	(0.157)	(0.070)	(0.186)	(0.188)	(0.074)	(0.161)	(0.164)
Observations	1,413	1,413	1,413	901	901	901	591	591	591
R-squared	0.024	0.072	0.081	0.019	0.050	0.056	0.004	0.070	0.071

Table 3. Propensity score matching results (Average Treatment Effect on the Treated) estimating the effect of early childbearing on age 40 self-assessed health among women with a first birth prior to age 36

	Age 40 Self-Assessed Health	
	White Women	Black Women
	ATT	ATT
First Birth Under Age 20 Compared to Age 24-35		
First Birth Under Age 20	-.185	-.183
(0 = first birth age 24-35)	(.127)	(.131)
Treatment observations	298	401
Control observations	734	212
Total n	1,032	613
First Birth Age 20-23 Compared to Age 24-35		
First birth age 20-23	-.098	-.298*
(0 = first birth age 24-35)	(.087)	(.121)
Treatment observations	356	241
Control observations	734	212
Total n	1,090	453

Notes: +p<.10; *p < .05; **p < .01; ***p < .001 (two-tailed tests); Standard errors in parentheses; ATT = average treatment effect for the treated.

Table 4: Age 40 Self-Assessed Health Regressed on Union Status at Birth and Subsequent Union History among Women Who Had a First Birth Prior to Age 24

| | Age 40 Self-Assessed Health | | | | | |
| | White | | Black | | Hispanic | |
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
Marital status at birth						
Unmarried (Ref = married)	-0.253***	---	-0.151	---	0.015	---
	(0.096)	---	(0.101)	---	(0.110)	---
Marital status at birth and later						
Unmarried at birth, later married	---	-0.455*	---	-0.215*	---	-0.424*
	---	(0.257)	---	(0.113)	---	(0.206)
Unmarried at birth, never married	---	-0.239**	---	-0.080	---	0.128
	---	(0.099)	---	(0.102)	---	(0.116)
(Ref = married at birth)	---	---	---	---	---	---
First birth at age 20-23	-0.084	-0.086	-0.019	-0.002	-0.029	-0.005
	(0.079)	(0.079)	(0.088)	(0.088)	(0.106)	(0.105)
Mother foreign born	0.443*	0.438*	0.020	-0.009	0.063	0.050
	(0.232)	(0.232)	(0.427)	(0.428)	(0.128)	(0.127)
R's mother's education (years)	0.063***	0.062***	0.021	0.021	0.009	0.011
	(0.018)	(0.018)	(0.017)	(0.017)	(0.015)	(0.015)
Lived with both parents age 14	0.022	0.019	-0.142*	-0.143*	0.002	-0.018
	(0.086)	(0.086)	(0.083)	(0.083)	(0.108)	(0.107)
AFQT	0.005***	0.005***	0.007***	0.007**	0.012***	0.011***
	(0.002)	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)
In poverty 1979 (Ref = no)	-0.275**	-0.272**	-0.029	-0.023	-0.125	-0.116
	(0.109)	(0.109)	(0.088)	(0.088)	(0.109)	(0.109)
Constant	2.780***	2.794***	3.138***	3.126***	3.118***	3.097***
	(0.204)	(0.204)	(0.210)	(0.208)	(0.173)	(0.173)
Observations	682	682	682	682	387	387
R-squared	0.086	0.088	0.026	0.028	0.062	0.077

Notes: +p<.10; *p <.05; **p <.01; ***p < .001 (two-tailed tests); Standard errors in parentheses;

Table 5. Propensity score matching results (ATT) estimating the effect of marital status at birth and marital history on age 40 self-assessed health among women with an adolescent / early first birth

	Age 40 Self-Assessed Health			
	Black Women	White Women	Hispanic Women	White & Hispanic
	ATT	ATT	ATT	ATT
Unmarried at birth but later married compared to married at birth				
Unmarried at birth and married by age 40	-.474***	---	---	-.578*
(0 = married at birth)	(.142)	---	---	(.270)
Treatment observations	176	---	---	38
Control observations	149	---	---	755
Total n	325	---	---	793
Unmarried at birth and never married compared to married at birth				
Unmarried at birth and never married	-.027	-.244+	.282	---
(0 = married at birth)	(.134)	(.131)	(.188)	---
Treatment observations	317	127	101	---
Control observations	151	507	245	---
Total n	468	634	346	---

Notes: +p<.10; *p < .05; **p <.01; ***p < .001 (two-tailed tests); Standard errors in parentheses;
ATT = average treatment effect for the treated.