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African-American/White Differences in the Age of Menarche: Accounting for the Difference

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Abstract

Lifetime health disparity between African-American and white females begins with lower birthweight and higher rates of childhood overweight. In adolescence, African-American girls experience earlier menarche. Understanding the origins of these health disparities is a national priority. There is growing literature suggesting that the life course health development model is a useful framework for studying disparities. The purpose of this study was to quantify the influence of explanatory factors from key developmental stages on the age of menarche and to determine how much of the overall race difference in age of menarche they could explain. The factors were maternal age of menarche, birthweight, poverty during early childhood (age 0 through 5 years), and child BMI z-scores at 6 years. The sample, drawn from the US National Longitudinal Surveys of Youth Child-Mother file, consisted of 2337 girls born between 1978 and 1998. Mean age of menarche in months was 144 for African-American girls and 150 for whites.

An instrumental variable approach was used to estimate a causal effect of child BMI z-score on age of menarche. The instrumental variables were prepregnancy BMI, high gestational weight gain and smoking during pregnancy. We found strong effects of maternal age of menarche, birthweight, and child BMI z-score (-5.23 , 95% CI $[-7.35, -3.12]$) for both African-Americans and whites. Age of menarche declined with increases in exposure to poverty during early childhood for whites. There was no effect of poverty for African-Americans. We used Oaxaca decomposition techniques to determine how much of the overall race difference in age of menarche was attributable to race differences in observable factors and how much was due to race dependent responses. The African-American/white difference in childhood BMI explained about 18% of the overall difference in age of menarche and birthweight differences explained another 11%.

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Keywords

USA; race disparities; menarche; childhood BMI; instrumental variables; age; girls

Introduction

African-American females are less healthy than whites throughout their lives. As neonates they have lower birthweight (Blumenshine et al., 2010; Whitehead & Helms, 2010); in childhood they are more likely to be obese (Ogden et al., 2006); and in adulthood they experience higher rates of chronic diseases such as cardiovascular disease and depression (Baker et al., 2010). Understanding the origins and pathways of these health disparities is a high national priority and there is a growing literature suggesting that a life course health development model is a useful framework for studying disparities (Braveman & Barclay, 2009; Halfon & Hochstein, 2002; Lu & Halfon, 2003). Within the life course model, health development is seen as the result of biological processes unfolding during defined developmental stages, interacting with physical, social, economic, and psychological environments. Health disparities are thought to originate in exposures to disadvantage at critical developmental stages and/or because of compounding and cumulative effects with continual exposure to disadvantage. This framework has the benefit of being able to consider proximal factors as well as distal risks from previous development stages. In this study, we were interested in understanding the racial difference in age of menarche as a function of factors from developmental periods (i.e., intrauterine, early childhood, middle childhood, adolescence) known to be associated with age of menarche and to have significant variation across racial group.

Contemporary African-American girls experience a younger age of menarche than do their white counterparts (Huang et al. 2009), a finding replicated in several large population studies. For example, the Pediatric Research in Office Settings data revealed that mean age of menarche was 12.2 years for African-American girls compared to 12.9 years for whites (Herman-Giddens et al., 1997); the National Health and Nutrition Examination Survey (NHANES III) data indicated mean ages of 12.1 and 12.7 years for the two groups respectively (Wu et al., 2002); and the Bogalusa Heart Study reported mean age of 12.1 years for African-American girls and 12.5 years for whites (Freedman et al., 2002a). Although the race differences in age of menarche were statistically significant in all studies, its explanation remains elusive.

Research has traced differences in age of menarche to early compromise in the intrauterine environment. Size at birth, whether measured by birthweight, weight for gestational age, or a combination of birthweight and birth length, has been used as a summary measure of the quality of the intrauterine environment. There is no consensus in the literature about the effect of birth size on age of menarche. However, it does appear to be sensitive to controls for growth in infancy and childhood. Adair (2001) reported that long and light infants had earlier menarche and the effect was more pronounced controlling for post natal growth in infancy. Tam and colleagues (2006) found that long and light infants had an earlier menarche, but only after controlling for adiposity at age 8 years. Sloboda and colleagues (2007) found independent effects of low birth weight and weight gain through age 8 years, with earliest menarche occurring in girls with low birth weight and high BMI. In contrast, Terry et al. (2009) found that a higher birthweight was associated with an earlier age of menarche, but only after controlling for weight gain until age 7 years. Dos Santos Silva et al. (2002) found that higher birthweight was associated with lower age of menarche after controlling for growth in infancy, but the effect disappeared when controls for growth

through age 7 are added. Direct comparison of these studies is difficult as there is variation in the study populations, in the growth measures, and the analytic approaches.

Stress levels related to factors in the child's early environment, including material resources, environmental toxins, harsh/positive parenting styles, and family structure, have been linked to age of menarche (Belsky et al., 2010; Hulanicka et al., 2001). Exposures to nicotine in utero and during childhood have been associated with later age of menarche (Ferris et al., 2010; Windham et al., 2008). It is widely acknowledged that African-American families experience increased material hardship and levels of stress (Logan, 2002), thereby introducing a mechanism through which possible race differences in the age of menarche may originate. Lower economic status at age 7 and declining economic status between birth and age 7 have been associated with a younger age of menarche in a racially mixed population (James-Todd et al., 2010). However, others have reported different race effects of exposure to poverty on age of menarche. Compared to girls of the same race above the poverty line, African-American girls living in poverty had later menarche whereas white girls living in poverty had earlier menarche (Braithwaite et al., 2009).

A girl's BMI has been the most widely researched factor in understanding race differences in age of menarche. Overweight is currently accepted as the primary explanation for differences in age of menarche across all races (Anderson et al., 2003) and in particular the African-American/white difference (Freedman et al., 2002b). Increased adiposity has been linked to early menarche, with numerous mechanisms hypothesized to be in play. Frisch and McArthur (1974) hypothesized that a minimum weight or body fat percent was required for menarche. Current theories focus on hormonal mechanisms that may be altered in girls with higher BMIs, including the influence of leptin activation on the gonadotropin-releasing hormone pulse generator, but definitive paths are not known and considerable research is ongoing. Furthermore, because the obesity epidemic has disproportionately impacted African-American girls (Strauss & Pollack, 2001) many have inferred that the earlier age of menarche experienced by African-American girls is due to increased BMI (Kaplowitz et al., 2006; Lee et al., 2007; Wang, 2002).

Only a few studies have explicitly explored the reasons for the African-American/white difference in age of menarche and the results are far from conclusive. Kaplowitz et al. (2001) found an association between increased weight and earlier puberty for both white and African-American girls, but the effect was stronger for whites, leading them to conclude that other factors, perhaps genetics or environmental factors, may also be operating. Anderson et al. (2003) concluded that increased BMI leads to earlier menarche, but that race differences in age of menarche cannot be explained entirely by race differences in BMI.

While this substantial body of work highlights key factors associated with age of menarche, there remain gaps in our knowledge. First, prior approaches using standard methods like OLS may have resulted in biased estimates of the effects of child BMI because BMI and age of menarche are likely to be influenced by some of the same unobserved factors. The second gap is linked to understanding the racial difference in age of menarche. Few studies have explicitly examined why racial differences in age of menarche occur. The purpose of this study was to identify and quantify the influence of explanatory factors from key developmental stages on the age of menarche. This study is novel in the sense that we use an instrumental variable (IV) approach, i.e., two-stage least squares, to overcome the problem of endogenous child BMI and to obtain causal effects of the explanatory factors on age of menarche. The study also assesses the extent to which each explanatory factor contributes to the race difference in age of menarche by implementing a Blinder-Oaxaca decomposition analysis.

Research Methods

Sample

Data were drawn from the 1979–2010 waves of the U.S. based National Longitudinal Surveys of Youth (NLS) Child-Mother file, a widely used, publicly available data set that includes a representative sample of individuals born between 1957 and 1964 as well as an over-sample of African-Americans (CHRR, 2000). Surveys were conducted annually between 1979 and 1994 and biennially thereafter. Information on birth outcomes for all biological children born to the female respondents were collected at each interview. Beginning in 1986, additional child assessments, including measures of height and weight, were conducted biennially. The study sample was drawn from girls born between 1978 and 1998 to non-Hispanic Caucasian (white) mothers from the cross-section sample and African-American mothers from either the cross-section sample or the oversample.

The potential sample (N=3030) excluded girls who reported not having reached menarche in the 2010 survey. The sample size was reduced by missing data for the following reasons: (a) girls who were not interviewed in 2010 and had no information about age of menarche in prior interviews (n=205); (b) girls who reported age of menarche in years but not months (n=166); (c) girls who lacked height and weight data during a prepubertal period defined as age 48–95 months and at least 24 months prior to menarche (n=72); (d) girls who were missing other variables used in the analysis (n=250). The study sample consisted of 2337 girls (77% of the eligible sample): 957 African-American girls born to 681 mothers and 1380 white girls born to 1056 mothers. The Ohio State University Institutional Review Board approved the study.

Measurement

Endogenous Variables

Pre-pubertal BMI z-score: The prepubertal period was defined by age and age relative to menarche. Girls needed to be observed between the ages of 48–95 months and at least 24 months prior to menarche. If multiple interviews occurred during this period, the interview that was closest to 6 years of age (72 months) was selected. Mothers were given the option to report their children's heights or have the interviewer complete the measurement. Interviewers were trained by the National Opinion Research Center at the University of Chicago to conduct height and weight measurements, with clothes on but without shoes. Approximately 78% of all heights and weights were measured. Measured height was slightly higher than mother-reported height (approximately 2 cm), although there was no difference in interviewer measured and mother reported weight. Mother reported heights were adjusted based on linear regression coefficients for age categories. BMI z-scores were calculated using the CDC program designed for this purpose because research has shown they correspond more accurately to body fatness than percentile scores (Field et al., 2003).

Age of Menarche: Mothers of girls at least 8 and less than 14 years of age were asked whether they had reached menarche. If so, the mother reported the year and month of menarche. Girls 14 and older were asked these questions directly. This information was combined with the girls' year and month of birth to determine their age of menarche in months. Over 60 percent reported within one year of menarche.

Exogenous variables

Race: The girls' race was determined by maternal race as established by the interviewer during the initial 1979 interview. Mixed race girls are categorized by race of the mother.

Maternal age of menarche: Age of menarche for mothers was collected in 1984 and 1985 when they were 19 to 28 years of age. About 75% of the mothers reported age in years and 25% reported age in months. Age in months was used when available and age in years was converted to age in months for comparability with their daughter's age of menarche. Maternal age of menarche is a proxy for partially shared genetic and environmental factors across generations.

Birthweight (kilograms): Birthweight is mother reported during the first interview following the child's birth. Pounds and ounces have been converted to kilograms.

Percent time in poverty age 0–5 years: Based on all sources of information about family composition and income in the year prior to the NLS interview, the survey creates an indicator for whether the child was living in poverty. The percent time in poverty is the average time each girl spent in poverty from birth through 5 years of age. For example, a girl who spent 3 years in poverty and 2 years above poverty would have a value of 0.6 because she had spent 60% of her early childhood living in poverty.

Instrumental variables for childhood BMIz

Smoking during pregnancy: The NLS collected data on whether or not the mother smoked during pregnancy in the first interview following the birth of each child. The variable was coded as 0/1.

Maternal Prepregnancy BMI: The NLS collected data on maternal height in 1981, 1982, 1984 and 1985. It was also collected during the first interview after the respondent was 40 years of age. When multiple reports were recorded we used the following algorithm based on Huang and colleagues (2009): (i) if reported heights followed a non-decreasing pattern, the last reported height was selected; (ii) when reported heights fluctuated within a 2-inch range, the median height was selected; (iii) if reported heights fluctuated more than 4 inches from the median, the outliers were discarded and the procedures described in (i) and (ii) were followed; (iv) if height was not reported at least once during the earlier years we used data from the health at 40 module. Data on prepregnancy weight was collected during the first interview after the birth of the child. BMI was computed as weight in kilograms divided by height squared in meters.

High gestational weight gain: The Institute of Medicine has published guidelines to categorize maternal gestational weight gain based on pre-pregnancy BMI; below, within and above recommended weight (Rasmussen et al., 2009). The NLS collected data on weight gain during pregnancy during the first interview following the birth of the child. The Institute of Medicine guidelines were used in conjunction with weight gain and pre-pregnancy BMI to construct an indicator variable for whether weight gain was high, i.e. above the recommendations.

Sample selection bias

To assess potential sample selection bias within race, we tested for differences in variable means for the complete case and excluded samples. T-tests and chi-square tests were used for the continuous and discrete variables respectively. Within each racial group there was no difference between the variable means for the included and excluded sample, providing some assurance that biases due to sample selection have not been introduced.

Analysis

Analysis of the complete case sample began by testing for race differences in the variable means. We then examined race differences in the response of age of menarche to each of the explanatory variables to inform model specification about important interactions with race.

Age of menarche is specified to be a linear function of childhood BMI z-score (*BMIz*) birthweight, maternal age of menarche, the percent of time spent in poverty during the first 5 years of life, and an indicator for girls who were African-American to capture otherwise unmeasured race differences in age of menarche. Although BMI z-scores are widely used in research to predict age at menarche, childhood *BMIz* is a potential endogenous regressor. Unmeasured factors, such as diet, exercise and metabolism, could affect both age of menarche and *BMIz*. If this is the case, OLS will produce inconsistent estimates. In particular the estimated effect of *BMIz* will be biased and cannot be interpreted as a causal effect of child *BMIz* on age of menarche.

The method of instrumental variables (IV) has been suggested as a solution to the problem of endogenous explanatory variables when instruments can be found that satisfy a stringent set of conditions. An instrument for *BMIz* must be correlated with *BMIz* after netting out the effects of the other explanatory variables and it must be uncorrelated with the error term in the age of menarche equation (Hernán & Robins, 2006; Wooldridge, 2010). Hernán and Robins (2006) described three conditions required for an instrument. We paraphrase them as follows: (i) the instrument must have a causal effect on *BMIz*, (ii) it must affect age of menarche only through its effect on *BMIz*, and (iii) it must not share common causes with age of menarche, i.e., no confounding for the effect of the instrument on age of menarche. Further, an instrument that satisfies condition (i) must not be a weak instrument, i.e. it must have a sufficiently strong partial effect on *BMIz*; otherwise estimates will be biased towards the OLS estimates (Bound et al., 1995). If conditions (ii) and (iii) are satisfied there is no correlation between the instrument and the error term in the structural age of menarche equation. The instruments for *BMIz* used in this application are maternal prepregnancy BMI, high maternal weight gain and smoking during pregnancy. The estimated effect of *BMIz* on age of menarche is the average effect for those girls whose *BMIz* is itself affected by the particular instruments (Angrist & Imbens, 1995).

Two-stage least squares (2SLS) is a type of IV estimator which is efficient in the class of all linear IV estimators if conditions (i)–(iii) are satisfied and the errors are homoskedastic. The 2SLS estimator operates by first estimating a reduced form equation for *BMIz*, i.e. an OLS equation of *BMIz* as a function of all of the potential instruments and the explanatory variables in the age of menarche equation. The predicted value of *BMIz* is by construction uncorrelated with the error term in the age of menarche equation. In the second stage of 2SLS, consistent estimates of the effect of *BMIz* on age of menarche are obtained by using the predicted value of *BMIz* in the OLS age of menarche equation.

A test for weak instruments based on the minimum eigenvalue statistic was implemented based on the widely used standard proposed by Stock and Yogo (2005). The instruments proposed in this application cannot be justified on theoretical grounds alone. Thus we rely on a Sargan test to provide statistical evidence that the instruments are not correlated with the error term in the structural age of menarche equation. The test does not guarantee that conditions (ii) and (iii) are satisfied, it only provides probabilistic support that they are satisfied. In the absence of multiple instruments this test cannot be performed.

A Hausman test for exogeneity is also implemented to determine whether *BMIz* is in fact endogenous. It compares the OLS and 2SLS estimators of parameters from age of menarche equation. If *BMIz* is uncorrelated with ϵ , the two estimators should differ only by sampling

error. Since the tests for overidentifying restrictions and exogeneity are valid under the homoskedasticity assumption, that too is tested. Finally, the potential endogeneity of birthweight as well as *BMIz* was examined. All analyses were conducted using Stata 12. Robust standard errors of the reported IV estimates are corrected for clustering between siblings.

The final part of the analysis decomposed the total difference in mean age of menarche between the races into two additive elements: the fraction attributable to (1) race differences in means of observable characteristics and (2) differences across groups in the responses to the observable characteristics, where responses are measured by group-specific regression coefficients. This method reveals how much of the inequality in menarcheal age can be explained by inequalities in birthweight, child *BMIz*, maternal age of menarche and exposure to poverty in early childhood. The method requires that the coefficients estimates have a causal interpretation; therefore the predicted value of *BMIz* from the reduced form is used in place of observed *BMIz* in linear regression in the race specific regressions underlying the estimates.

Decomposition techniques, first used by Oaxaca (1973) and Blinder (1973) to study race and gender differences in wages, exploit the property that the linear regression line goes through the means of the variables and hence the observed race difference in an outcome is the difference between the coefficients from race specific regressions evaluated at the race specific variable means. Theoretically, there is no “correct” set of “reference” coefficients. Neumark (1988) suggested using coefficients from a regression that pools the races. An alternative is to select one of the race specific regression coefficients as the reference set. Their interpretation varies slightly.

The decompositions require first estimating an age of menarche equation on three different samples- African-Americans, whites, and the pooled sample denoted respectively by $g=\{A,W,P\}$. For each of the samples age of menarche is regressed against a vector of explanatory variables including a constant, predicted *BMIz*, birthweight, mother’s age of menarche and fraction of time spent in poverty during early childhood. Let the coefficient estimates of the three regressions be represented by Γ_g . To economize on notation, let $\mathbf{M}_g$ denote race specific mean age of menarche and $\mathbf{X}_g$ denotes the transpose of a vector of race specific means including a constant. The Neumark decomposition based on coefficients from the pooled regression is

$$\mathbf{M}_A - \mathbf{M}_W = \mathbf{X}_A \Gamma_A - \mathbf{X}_W \Gamma_W = (\mathbf{X}_A - \mathbf{X}_W) \Gamma_P + [\mathbf{X}_A(\Gamma_A - \Gamma_P) - \mathbf{X}_W(\Gamma_W - \Gamma_P)].$$

The first term on the far right hand side, $(\mathbf{X}_A - \mathbf{X}_W) \Gamma_P$, can be interpreted as the part of race difference in age of menarche due to differences in characteristics. This is referred to as the explained difference. For example if the coefficient on an explanatory variable is 2 and the race difference in the variable means is 3, then that variable would be said to “explain” a 6 month difference in age of menarche. The second term, $[\mathbf{X}_A(\Gamma_A - \Gamma_P) - \mathbf{X}_W(\Gamma_W - \Gamma_P)]$, is the part of the overall race difference in age of menarche due to differences in the responses of the two groups, where response is measured by differences in coefficients. It is referred to as “unexplained differences” because there is no theory about why the coefficients from race specific regressions should differ from the coefficients from the pooled regression. Note that if the coefficients from the race specific regressions are not different, and therefore are not different from the pooled regression coefficients, then this second term is close to zero. Estimates for the Oaxaca decomposition, were generated using the Oaxaca command in Stata, with the pooled option.

Results

Table 1 presents evidence on race differences in variable means. African-American girls reached menarche about 6.54 months before whites. Childhood *BMIz* was about one-quarter of a standard deviation higher for African-American girls than whites. The mean difference in birthweight was 240 grams, with African-American neonates being smaller than whites. By the time they turned 6 years of age, African-American girls on average lived in poverty almost half of the time, more than triple the time spent by white girls. African-American mothers were one BMI point heavier prepregnancy than whites (6.3 pounds for a 5'7" woman). No race difference was found in high gestational weight gain. Maternal age of menarche was not different by race when each mother was counted as one observation (reported in Table 1). However, African-American mothers with earlier menarche relative to whites had more daughters in the sample. There was a race difference of about three months in maternal age of menarche when examining that variable in a sample based on the number of daughters (not reported in Table 1). African-American mothers smoked significantly less during pregnancy than their white counterparts.

A series of race-specific univariate regressions with age of menarche as the dependent variables were estimated for each of the explanatory variables. With the exception of fraction of time spent in poverty between age 0 and 5 years, we could not reject that the coefficients were the same for both races. However, there was a negative, statistically significant association between age of menarche and time spent in poverty for whites, but a positive, statistically insignificant association for African-Americans. Figure 1 presents graphical results of these relationships. Based on this finding, we estimated models that fully interacted race with percent of time in poverty.

Table 2 column one presents results from an OLS model with age of menarche as the dependent variable. The OLS results suggest that a one standard deviation increase in *BMIz* is associated with a decrease in age of menarche by 1.31 months. The estimates suggest that white girls who are always in poverty experience menarche 5.48 months earlier than whites never in poverty; in sharp contrast there was no statistically significant effect of poverty on age of menarche for African-American girls. The coefficients represent the effect of poverty for each race separately. Girls whose mothers experienced menarche one month later have menarche 0.14 months later. After controlling for the observable characteristics, including poverty, African American girls experienced menarche 7.21 months earlier.

The estimates from the first stage reduced form childhood *BMIz* equation are reported in column 3. *BMIz* increased with birthweight, prepregnancy BMI, high gestational weight gain and smoking during pregnancy. It was higher for African-American girls than white girls. Before describing results of the 2-SLS estimation we report on the specification tests performed. The Breusch-Pagan test of the null hypothesis of constant variance, distributed $\chi^2(1)$ under the null, had a value of 0.45. We could not reject that the errors were homoskedastic ($p=0.50$). The F-statistic ($F(3,2328)$) for the test of the null hypothesis that the instruments were weak was 36.16 ($p<0.01$). Thus we have confidence that the instruments have explanatory power in the reduced form *BMIz* equation. Under the null hypotheses that the instrumental variables are exogenous in the age of menarche equation, the Sargan test would be distributed $\chi^2(2)$ since we have 2 overidentifying restrictions. Based on the value of the test statistic, 2.86, we could not reject that the instruments are uncorrelated with the error term in the structural age of menarche equation ($p=0.24$). Finally the Hausman test ($F(1, 2329)$) of the null hypothesis that *BMIz* is exogenous was 15.41 ($p<0.01$), providing evidence that *BMIz* is endogenous and that difference between the OLS estimates and the 2-SLS estimates is statistically significant. These tests indicate that this is an appropriate application of instrumental variables.

Table 2 column 2 presents results from the 2-SLS age of menarche equation. The 2-SLS estimates suggest that a one standard deviation increase in childhood *BMIz* led to a 5.23 months decline in age of menarche, almost 4 times larger than the OLS estimate. The coefficient on birthweight in kilograms was 3.00, which implies that on average an increase in birthweight of 100 grams increased age of menarche by about 10 days. The effects of poverty and maternal age of menarche were comparable to those found in OLS.

Table 3 summarizes results from the Oaxaca decomposition of the total sample difference in age of menarche between African-Americans and whites. That difference was 6.54 months. Columns 1 and 2 report on the decomposition based on use of coefficients from the pooled sample as reference group. Column 1 reports the difference in age of menarche explained by mean differences in each of the explanatory variables. Based on coefficients from a pooled regression, higher average *BMIz* accounted for 1.15 months of the earlier age of menarche experience by African-American girls. Lower birthweight accounted for 0.70 month of the earlier age of menarche. Differences in means of other explanatory variables could not explain statistically significant differences in age of menarche. Column 2 reports on unexplained differences in age of menarche attributable to differences between the race-specific coefficients evaluated at the race specific variable means. Differences in the coefficients on poverty suggest that, given race specific means of poverty rates, African-American girls would have menarche 1.81 months later than whites. Race differences in the coefficients on *BMIz* suggest that they accounts for an unexplained difference of 0.06 months in age of menarche.

Discussion

The over-arching purpose of this study was to explain racial difference in age of menarche as a function of prior health outcomes and risk factors. A life course health development framework was used to link outcomes and risks during earlier developmental periods with later health outcomes and explore how race differences in these exposures contributed to race differences in age of menarche. We found strong effects of maternal age of menarche, birthweight, and child BMI z-score on age of menarche for both African-Americans and whites in the multivariate analysis as previously reported in the literature. Age of menarche declined with increases in exposure to poverty during early childhood only for whites. There was no effect of poverty for African-Americans. These race differences in the effects of poverty were also reported by Braithwaite and colleagues (2002).

We provided causal estimates of the negative effect of child BMI z-score on age of menarche by using instrumental variables to obtain consistent estimates of the effect sizes. Since the BMI z-score is not linear in weight for a given baseline height, girls with higher childhood BMI z-score, holding constant height, would have to lose more weight to increase predicted age of menarche by a fixed number of months. For example, a 72-month old girl with height of 3 feet 9 inches, at the mean of the empirical height-for-age distribution, weighing 44 pounds has a BMI z-score of zero. If, however, a girl weighed 49 pounds, at this baseline height, her BMI z-score would be one; and if she weighed 59 pounds her BMI z-score would be two. Our IV estimates of predicted that, *ceteris paribus*, the heaviest girl (59 pounds) would have menarche 5 months earlier than the middle weight girl (49 pounds) who in turn would have menarche 5 months earlier than the girl weighing 44 pounds. The clinical implication is that girls with higher BMI z-scores would need to lose more weight to affect any given increase in age of menarche.

The instrumental variable results highlight potential biases in OLS estimates of the effect of child BMI on age of menarche. The biases are attributable to correlation between the error term in the age of menarche equations and child BMI z-score arising from common

unobserved factors. The IV results suggested that the effect of BMI z-score on age of menarche was four times greater than the OLS results. The instrumental variables for child BMI z-score were maternal smoking during pregnancy, maternal prepregnancy BMI and high maternal weight gain during pregnancy. Typically the public health literature favors instruments drawn from a random controlled trial or a valid natural experiment (Hernán & Robins, 2006). We are not aware of suitable natural experiments to address the effect of over nutrition. The primary concern with instruments of the sort used in this paper is that they may be correlated with the error term in the structural age of menarche equation. While our instruments cannot be justified on theoretical grounds, they did satisfy exogeneity tests. The test, however, is statistical and does not “guarantee” exogeneity of the instruments in either this sample or in others. It merely suggests that the probability that the instruments are invalid is sufficiently low that one can have some degree of confidence in the validity of the instruments and interpretations of the estimated effect sizes as causal estimates.

Other threats to the validity of the instruments come from the potential for weak instruments which leads to estimates biased towards those of ordinary least squares (Bound et al., 1995). Again statistical tests suggested that the explanatory power of the instruments for child BMI z-score was sufficiently great that the instruments were not weak. But again this is a statistical test, not a guarantee. It is also known that there is a limit on the strength of instruments placed by the amount of confounding between the BMI z-score and the error term in the age of menarche equation (Martens et al., 2006). It is likely that confounding between age of menarche and child BMI is high and the potential strength of the instruments is relatively low. In addition the R-squared of the BMI z-score was .07. Finally, instrumental variables are always biased in finite samples, although they are consistent as the sample size goes to infinity.

In this study, increasing birthweight was associated with a later age of menarche. But because we do not have an adequate measure of weight gain during infancy we could not tease apart a birthweight effect from what may be related to an effect of accelerated growth in infancy. An additional concern with the estimated effect of birthweight on age of menarche was its potential bias due to possible endogeneity, i.e. correlation with the error term in the age of menarche equation. Since we had three instrumental variables we could use them to test for endogeneity simultaneously in birthweight and child BMI z-score. Tests showed that the instruments were valid and not weak for both child BMI z-score and birthweight. But we found we could not reject the hypothesis that birthweight was exogenous, giving us confidence in the 2SLS that finds that a 100 gram increase in birthweight increases age of menarche by about 1 week in this specification.

The effect of early childhood poverty on age of menarche was different by race. There was a social gradient in age of menarche for whites but not African-Americans. The finding is important because it sits at the intersection of literatures on economic, race and gender disparities in health. First, there is a growing body of evidence suggesting that the effect of low socio-economic status on development of overweight is gender-specific (Robinson et al., 2009). Second, the social gradient in overweight for girls is race specific. White adolescent girls with higher socio-economic status have lower prevalence of overweight, but high socio-economic African-American girls have higher prevalence of overweight (Wang & Zhang, 2006). If factors driving early menarche are also driving overweight, we would expect to observe comparable social gradients in age of menarche and overweight. Whites display similar socio economic gradients in overweight and age of menarche, but African-Americans have an inverse social gradient in overweight in conjunction with a weak or largely absent social gradient in age of menarche. These contrasts may provide a unique opportunity to disentangle the effects of socio-economic status on overweight and maturation.

Poverty is only one measure of socio-economic status and a particular measure may be more relevant for a given racial group. To explore the robustness of our finding about the race difference in the socio-economic gradient in age of menarche we re-estimated the model using maternal education and father absence separately and in combination with our poverty variable. Father absent was measured as the fraction of time the father was not in the household from birth through age 5 years, to be comparable with our measure of child poverty. Maternal education was categorized according to whether the mother was a high school dropout. Neither variable had a statistically significant effect, individually or in combination with child poverty, on age of menarche for African-American girls. However, for white girls having a father who was absent more during childhood and/or a mother who was a high school dropout was associated with a lower age of menarche. Thus, the question appears to be why relative advantage does not confer greater protective health for African-American girls. This important topic needs to be explored in future research.

The Oaxaca decomposition analysis parsed and quantified the African-American/white difference in age of menarche into an explained component (39%) and an unexplained component (61%). Of the 2.52 months of difference explained by the 2SLS model, BMIz accounted for 1.15 months (18%) of the overall difference of 6.54 months) and birthweight accounted for 0.70 months (11%).

Theoretically, there is no “correct” set of “reference” coefficients in an Oaxaca analysis. To test the sensitivity of using coefficients from the pooled sample we also ran alternative Oaxaca analyses using race specific regression coefficients as the reference category (Daymont & Andrisani, 1984). Results using each race as the reference group were qualitatively similar to the results reported here based on the racially pooled sample.

There are limitations of the study. First is the assumption that there is a linear relation between age of menarche and the explanatory variables, in particular BMI z-score. However, the effect of a given increase in child BMI may depend on the child’s BMI. The effect of child BMI may be nonlinear or more generally heterogeneous across children. Second, the study relies on a point-in-time measure of BMI and does not consider the entire trajectory of child growth. Third, the analysis was performed on the complete case sample, although results from conditional mean imputation were similar. We were particularly concerned about loss of observations due to missing values of instruments. Therefore, the IV models were re-estimated using combinations of two out of three of the instruments on a larger sample population. The results were qualitatively similar to the ones reported in Table 2 suggesting that the results are not sensitive to missing data in the instruments. Fourth, maternal measures of height and weight, and age of menarche were self-reported, and not all the heights and weights on the children were measured. There were not physical examinations to confirm pubertal stage, only maternal or self-reported age of menarche. Consequently, self-reported data may have led to measurement error in explanatory variables, which would bias estimates towards zero. Fifth, the study sample consisted of girls born to a representative sample of mothers in the 1957–1964 birth cohorts. Thus, results may not generalize to other samples.

The current study has several design and methodological strengths. These include an intergenerational sample with extensive data on both mothers and daughters, a longitudinal cohort of girls born between 1978 and 1998, an instrumental variables approach to account for endogeneity between child BMI and age of menarche, a large sample of African-American and white females, and a cumulative measure of poverty. In conclusion, this study provides evidence that race differences in birthweight and childhood BMI do contribute to racial differences in the age of menarche. However, race differences in childhood BMI explained less than 20 percent of the overall race difference in age of menarche for girls in

the NLS study sample. This conclusion is similar to that of Anderson and colleagues (2003) who concluded that racial differences in age of menarche cannot be explained entirely by differences in BMI.

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References

- Adair LS. Size at birth predicts age at menarche. *Pediatrics*. 2001; 107:E59. [PubMed: 11335780]
- Anderson SE, Dallal GE, Must A. Relative weight and race influence average age at menarche: results from two nationally representative surveys of US girls studied 25 years apart. *Pediatrics*. 2003; 111:844–850. [PubMed: 12671122]
- Angrist J, Imbens G. Two-stage least squares estimates of average causal effects in models with variable treatment intensity. *Journal of the American Statistical Association*. 1995; 90:431–442.
- Baker TA, Buchanan NT, Spencer TR. Disparities and social inequities: is the health of African American women still in peril? *Ethnicity & Disease*. 2010; 20:304–309. [PubMed: 20828107]
- Belsky J, Steinberg L, Houts RM, Halpern-Felsher BL. The development of reproductive strategy in females: early maternal harshness --> earlier menarche --> increased sexual risk taking. *Developmental Psychology*. 2010; 46:120–128. [PubMed: 20053011]
- Blinder A. Wage Discrimination: Reduced Form and Structural Estimates. *The Journal of Human Resources*. 1973; 8:436–455.
- Blumenshine P, Egarter S, Barclay CJ, Cubbin C, Braveman PA. Socioeconomic disparities in adverse birth outcomes: a systematic review. *American Journal of Preventive Medicine*. 2010; 39:263–272. [PubMed: 20709259]
- Bound J, Jaeger D, Baker R. Problems with instrumental variables estimation when the correlation the instruments and the endogenous explanatory variables is weak. *Journal of the American Statistical Association*. 1995; 90:433–450.
- Braithwaite D, Moore DH, Lustig RH, Epel ES, Ong KK, Rehkopf DH, et al. Socioeconomic status in relation to early menarche among black and white girls. *Cancer Causes Control*. 2009; 20:713–720. [PubMed: 19107561]
- Braveman P, Barclay C. Health disparities beginning in childhood: a life-course perspective. *Pediatrics*. 2009; 124(Suppl 3):S163–S175. [PubMed: 19861467]
- CHRR. NLSY79 Child and Young Adult Data: User Guide. Columbus, Ohio: Center for Human Resource Research; 2000.
- Daymont T, Andrisani P. Job Preferences, College Major, and the Gender Gap in Earnings. *Journal of Human Resources*. 1984; 19:408–428.
- dos Santos Silva I, De Stavola BL, Mann V, Kuh D, Hardy R, Wadsworth ME. Prenatal factors, childhood growth trajectories and age at menarche. *International Journal of Epidemiology*. 2002; 31:405–412. [PubMed: 11980804]
- Ferris JS, Flom JD, Tehranifar P, Mayne ST, Terry MB. Prenatal and childhood environmental tobacco smoke exposure and age at menarche. *Paediatric and Perinatal Epidemiology*. 2010; 24:515–523. [PubMed: 20955229]
- Field AE, Laird N, Steinberg E, Fallon E, Semega-Janneh M, Yanovski JA. Which metric of relative weight best captures body fatness in children? *Obesity Research*. 2003; 11:1345–1352. [PubMed: 14627755]
- Freedman DS, Khan LK, Serdula MK, Dietz WH, Srinivasan SR, Berenson GS. Relation of age at menarche to race, time period, and anthropometric dimensions: the Bogalusa Heart Study. *Pediatrics*. 2002a; 110:e43. [PubMed: 12359816]
- Freedman DS, Khan LK, Serdula MK, Dietz WH, Srinivasan SR, Berenson GS. Relation of age at menarche to race, time period, and anthropometric dimensions: the Bogalusa Heart Study. *Pediatrics*. 2002b; 110:e43. [PubMed: 12359816]

- Frisch RE, McArthur JW. Menstrual cycles: fatness as a determinant of minimum weight for height necessary for their maintenance or onset. *Science*. 1974; 185:949–951. [PubMed: 4469672]
- Halfon N, Hochstein M. Life course health development: an integrated framework for developing health, policy, and research. *Milbank Quarterly*. 2002; 80:433–479. iii. [PubMed: 12233246]
- Herman-Giddens ME, Slora EJ, Wasserman RC, Bourdony CJ, Bhapkar MV, Koch GG, et al. Secondary sexual characteristics and menses in young girls seen in office practice: a study from the Pediatric Research in Office Settings network. *Pediatrics*. 1997; 99:505–512. [PubMed: 9093289]
- Hernán M, Robins J. Instruments for causal inference: An epidemiologists dream? *Epidemiology*. 2006; 17:360–372. [PubMed: 16755261]
- Huang B, Biro FM, Dorn LD. Determination of relative timing of pubertal maturation through ordinal logistic modeling: evaluation of growth and timing parameters. *Journal of Adolescent Health*. 2009; 45:383–388. [PubMed: 19766943]
- Hulanicka B, Gronkiewicz L, Koniarek J. Effect of familial distress on growth and maturation of girls: a longitudinal study. *American Journal of Human Biology*. 2001; 13:771–776. [PubMed: 11748816]
- James-Todd T, Tehranifar P, Rich-Edwards J, Titievsky L, Terry MB. The impact of socioeconomic status across early life on age at menarche among a racially diverse population of girls. *Annals of Epidemiology*. 2010; 20:836–842. [PubMed: 20933190]
- Kaplowitz P. Pubertal development in girls: secular trends. *Current Opinions in Obstetrics and Gynecology*. 2006; 18:487–491.
- Kaplowitz PB, Slora EJ, Wasserman RC, Pedlow SE, Herman-Giddens ME. Earlier onset of puberty in girls: relation to increased body mass index and race. *Pediatrics*. 2001; 108:347–353. [PubMed: 11483799]
- Lee JM, Appugliese D, Kaciroti N, Corwyn RF, Bradley RH, Lumeng JC. Weight status in young girls and the onset of puberty. *Pediatrics*. 2007; 119:e624–e630. [PubMed: 17332182]
- Logan, J. *Separate and Unequal: The neighborhood gap for blacks and Hispanics in metropolitan areas*. Albany: University of Albany; 2002.
- Lu MC, Halfon N. Racial and ethnic disparities in birth outcomes: a life-course perspective. *Maternal and Child Health Journal*. 2003; 7:13–30. [PubMed: 12710797]
- Martens EP, Pestman WR, de Boer A, Belitser SV, Klungel OH. Instrumental variables: Application and limitations. *Epidemiology*. 2006; 17:260–267. [PubMed: 16617274]
- Neumark D. Employers' Discriminatory Behavior and the Estimation of Wage Discrimination. *The Journal of Human Resources*. 1988; 23:289–295.
- Oaxaca R. Male-Female Wage Differentials in Urban Labor Markets. *International Economic Review*. 1973; 14:693–709.
- Ogden CL, Carroll MD, Curtin LR, McDowell MA, Tabak CJ, Flegal KM. Prevalence of overweight and obesity in the United States, 1999–2004. *JAMA*. 2006; 295:1549–1555. [PubMed: 16595758]
- Rasmussen KM, Catalano PM, Yaktine AL. New guidelines for weight gain during pregnancy: what obstetrician/gynecologists should know. *Current Opinions in Obstetrics and Gynecology*. 2009; 21:521–526.
- Robinson WR, Gordon-Larsen P, Kaufman JS, Suchindran CM, Stevens J. The female-male disparity in obesity prevalence among black American young adults: contributions of sociodemographic characteristics of the childhood family. *American Journal of Clinical Nutrition*. 2009; 89:1204–1212. [PubMed: 19190069]
- Sloboda DM, Hart R, Doherty DA, Pennell CE, Hickey M. Age at menarche: Influences of prenatal and postnatal growth. *Journal of Clinical Endocrinology and Metabolism*. 2007; 92:46–50. [PubMed: 17062767]
- Stock, JH.; Yogo, M. Testing for weak instruments in IV regressions. In: Andrews, D.; Stock, J., editors. *Identification and inference for econometric models: a festschrift in honor of Thomas Rothenberg*. Cambridge: Univeristy Press; 2005. p. 80–108.
- Strauss RS, Pollack HA. Epidemic increase in childhood overweight, 1986–1998. *JAMA*. 2001; 286:2845–2848. [PubMed: 11735760]

- Tam CS, de Zegher F, Garnett SP, Baur LA, Cowell CT. Opposing influences of prenatal and postnatal growth on the timing of menarche. *Journal of Clinical Endocrinology and Metabolism*. 2006; 91:4369–4373. [PubMed: 16926251]
- Terry MB, Ferris JS, Tehranifar P, Wei Y, Flom JD. Birth weight, postnatal growth, and age at menarche. *American Journal of Epidemiology*. 2009; 170:72–79. [PubMed: 19439580]
- Wang Y. Is obesity associated with early sexual maturation? A comparison of the association in American boys versus girls. *Pediatrics*. 2002; 110:903–910. [PubMed: 12415028]
- Wang Y, Zhang Q. Are American children and adolescents of low socioeconomic status at increased risk of obesity? Changes in the association between overweight and family income between 1971 and 2002. *American Journal of Clinical Nutrition*. 2006; 84:707–716. [PubMed: 17023695]
- Whitehead N, Helms K. Racial and ethnic differences in preterm delivery among low-risk women. *Ethnicity and Disease*. 2010; 20:261–266. [PubMed: 20828100]
- Windham GC, Zhang L, Longnecker MP, Klebanoff M. Maternal smoking, demographic and lifestyle factors in relation to daughter's age at menarche. *Paediatric and Perinatal Epidemiology*. 2008; 22:551–561. [PubMed: 19000293]
- Wooldridge, J. *Econometric Analysis of Cross-Sectional and Panel Data*. Cambridge, Massachusetts: The MIT Press; 2010.
- Wu T, Mendola P, Buck GM. Ethnic differences in the presence of secondary sex characteristics and menarche among US girls: the Third National Health and Nutrition Examination Survey, 1988–1994. *Pediatrics*. 2002; 110:752–757. [PubMed: 12359790]

- Childhood BMIz is shown to be causally related to timing of menarche
- A one standard deviation increase in childhood BMIz causes age of menarche to decline by 5.2 months
- Race differences in BMI explain 18% of the difference between African – American and white girls in menarcheal age

Figure 1.
Univariate Regression Lines

Table 1

Variable means by race, p-value for test of difference in means (standard deviation of continuous variables in parentheses)

	African American	White	p
Age of menarche (months)	143.55 (0.52)	150.09 (0.40)	<0.01
Childhood BMI z-score	0.10 (0.05)	−0.11 (0.04)	<0.01
Age months when BMI was recorded	73.56 (0.36)	72.53 (0.29)	0.02
Birthweight (kilograms)	3.11 (0.02)	3.34 (0.01)	<0.01
Percent time in poverty age 0–5 years	0.48 (0.01)	0.15 (0.01)	<0.01
Maternal pre-pregnancy BMI	23.69 (0.16)	22.70 (0.12)	<0.01
High pregnancy weight gain	0.35 (0.02)	0.38 (0.01)	0.15
Maternal age of menarche (months)	153.88 (0.78)	154.10 (0.56)	0.82
Smoke during pregnancy	0.30 (0.01)	0.36 (0.01)	<0.01
Sample Size			
Girls	957	1380	
Mothers	681	1056	

Test statistics for difference in means: t-tests for continuous variables and Pearson chi-sq tests for discrete variables
Standard deviation of continuous variables in parentheses

Table 2

Regression results [95% confidence intervals in square brackets]

	(1) OLS Age of menarche	(2) 2-stage least squares Age of menarche	(3) Reduced form OLS Childhood BMIz
Childhood BMI z-score	-1.31 *** [-1.72,-0.89]	-5.23 *** [-7.35,-3.12]	
Birthweight (kilograms)	1.80 ** [0.63,2.97]	3.00 *** [1.53,4.48]	0.27 *** [0.16,0.38]
Percent poverty (0–5 years of age) *White	-5.48 *** [-8.45,-2.50]	-5.38 ** [-8.68,-2.08]	-0.11 [-0.42,0.19]
Percent poverty *African-American	1.81 [-0.74,4.37]	1.04 [-1.74,3.82]	-0.11 [-0.35,0.14]
Maternal age of menarche (months)	0.14 *** [0.10,0.17]	0.12 *** [0.08,0.16]	-0.00 [-0.01,0.01]
African-American	-7.21 *** [-9.08,-5.35]	-5.93 *** [-8.12,-3.74]	0.27 ** [0.09,0.44]
Prepregnancy BMI			0.06 *** [0.05,0.07]
High gestational weight gain			0.17 ** [0.05,0.30]
Smoke during pregnancy			0.21 ** [0.08,0.34]
<i>N</i>	2337	2337	2337
<i>R</i> ²	0.10	.	0.07

95% confidence intervals in brackets

*
 $p < 0.05$,**
 $p < 0.01$,***
 $p < 0.001$

Table 3

Results from Oaxaca Decomposition

	Explained	Unexplained
Totals	-2.52	-4.02
Predicted BMI z-score	-1.15 ***	0.06 *
Birthweight	-0.70 ***	-3.99
% time in poverty, 0-5 yrs	-0.53	1.81 **
Maternal age of menarche	-0.14	-7.30

Note: Total difference in mean age of menarche between African-American, whites is 6.54 months.

*
 $p < 0.05$,

**
 $p < 0.01$,

 $p < 0.001$