

RETURNS TO BIRTHWEIGHT

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Abstract—We use data on monozygotic twins to obtain improved estimates of the effect of intrauterine nutrient intake on adult health and earnings and thus to evaluate the efficacy of programs aimed at increasing birthweight. We use the results to evaluate the bias in cross-sectional estimates and to assess the proposition that health conditions play a major role in determining the world distribution of income. We show that there is considerable variation in the incidence of low birthweight across countries, and our estimates suggest that there are real payoffs to increasing body weight at birth. Increasing birthweight increases adult schooling attainment and adult height for babies at most levels of birthweight, but has no effect on adult body mass. The effect of increasing birthweight on schooling, moreover, is underestimated by 50% if there is no control for genetic and family background endowments as in cross-sectional estimates. We also find evidence that augmenting birthweight among lower-birthweight babies, but not among higher-birthweight babies, has significant labor market payoffs. However, shifting the distribution of birthweights in developing countries to that in the United States would reduce world earnings inequality by less than 1%, far less than indicated by the cross-country correlation between per-worker GDP and birthweight.

I. Introduction

A major health-related policy objective in most developed countries is to increase the birthweights of children in low-birthweight populations. Many low-birthweight infants who survive infancy, for example, are claimed to suffer cognitive and neurological impairment that limits the returns to human capital investments in them, their productivities, and their earnings as adults, and that, for females, increases the probability of having low-birthweight babies.¹ The effects of birth outcomes and other physical attributes on human capital and wages have also been the concern of a recent economics literature and of much bigger literatures in other disciplines.²

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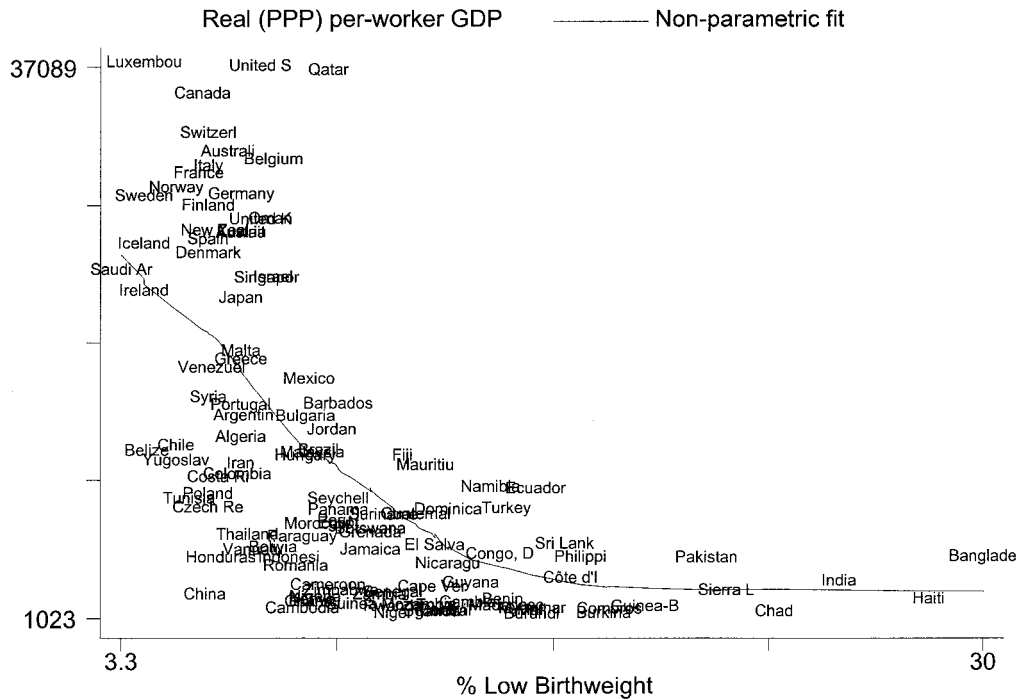
¹ In the epidemiological literature, for instance, Barker and others, whose research is summarized in Barker (1992, 1998), increasingly have emphasized the associations between low birthweight and stunting or wasting at birth on the one hand and adult health on the other, with possible effects through adult health on adult productivity. Based on animal and epidemiological studies, they find that malnourishment at birth is associated with coronary heart disease, hypertension, non-insulin-dependent diabetes, raised serum cholesterol, and abnormal blood clotting in later life.

² Pojda and Kelley (2000), for example, cite 198 references in their survey on the determinants of and the effects of low birth weight.

A number of studies have identified the determinants of birthweight within the context of behavioral models that incorporate endowment heterogeneity and evaluated major policy changes in terms of their effect on weight at birth (Rosenzweig & Schultz, 1983; Grossman & Joyce, 1990; Rosenzweig & Wolpin, 1995; Currie & Gruber, 1996; Currie & Hyson, 1999). However, the literatures concerned with the consequences of weight gain at birth generally do not distinguish between the policy-relevant effects of increasing the nutrients received by a fetus and possible genetic and family background influences on fetal development. It is possible that infants with genetically determined low weights at birth also have genetic endowments that make them less healthy as adults, and that increasing their weight at birth would have little lasting effect on adult life achievements. It also is possible that the genetic endowments of such infants are correlated with those of their parents and thereby associated with the household resources available to support child development. What is of policy interest is the effect of increasing the birthweight of a child with given genetic endowments and family background. Conventional cross-sectional and longitudinal data sets cannot answer that question unambiguously.

Obtaining improved estimates of the effect of intrauterine nutrient intake on adult health and earnings is also potentially useful for shedding light on the proposition that health conditions play a major role in determining the world distribution of income (Bloom & Sachs, 1998; Gallup, Sachs, & Mellinger, 1999). There is considerable variation in the incidence of low birthweight across countries. We have assembled a data set for 112 countries by merging country-specific information on rates of low birthweight (birthweight < 2500 g) in the decade of the 1990s, put together by UNICEF (2001) from official governmental sources and surveys, with information on country-specific, per-worker PPP-adjusted GDP in 1989. Figure 1, which plots these two variables, indicates that there is a strong negative association between the log of PPP per-worker GDP and the percentage of low-birthweight babies. Although it is doubtful that this association largely reflects cross-country genetic differences, it confounds the effects of the health of babies on their adult productivity with the effects of income on child health and the influence of common factors, such as institutions, that both improve health outcomes and augment incomes. Recent work has used micro evidence on the productivity effects of health to separate out the contribution of health in determining cross-country differences in incomes and to estimate the returns to increasing birthweights in low-income countries (Acemoglu, Johnson, & Robinson, 2001, 2003; Alderman & Behrman, 2003; Shastry & Weil, 2003). Such explorations using micro data are obviously also dependent on the appropriate

FIGURE 1.—LOW BIRTHWEIGHT AND LOG PER-WORKER GDP AROUND THE WORLD



identification of health effects, including those of better fetal nutrition.

Despite the importance of understanding the impact of improved nutrition in the womb on micro and aggregate outcomes, researchers generally have ignored the estimation problems that arise from endowment heterogeneity in assessing the effects of birthweight. Two recent exceptions are Conley and Bennet (2000) and Almond, Chay, and Lee (2002). Conley and Bennett, using sibling data from the PSID, find that conventional estimates substantially understate the negative effect of low birthweight on the probability of timely high school graduation. A problem with their approach, however, is that it assumes that prenatal input differences between siblings are orthogonal to their endowment differences, contrary to the evidence in Rosenzweig and Wolpin (1995). Almond et al. match the birth records of twins with death records to examine how differences in the birthweight of twins affect the probability of infant mortality. However, they cannot identify those twin pairs that share the same genetic endowments, so that their within-twin estimates confound in part the effects of nutrient intake and genetic endowments.

In this paper we obtain estimates of the impacts of in utero nutrient intake on adult outcomes within a context in which individuals are born with contemporaneously and intergenerationally correlated health and earnings endowments and resource allocations may be attentive to endowments. We show in particular that with information on the birthweights of genetically identical twins it is possible, with fewer assumptions than are needed to identify schooling returns using twins (Ashenfelter & Krueger, 1994;

Behrman, Rosenzweig, & Taubman, 1994), to identify the causal effect of increasing birthweight on adult anthropometrics, schooling, and earnings and to draw inferences about the role of endowments in affecting resource allocation behavior.

In obtaining our estimates we use new survey data on monozygotic (MZ) female twins collected by the authors from a sample from the Minnesota Twins Registry, the largest birth-certificate-based twins registry in the United States. We focus on women in this paper because of the suggestion in most previous studies that the earnings of women are more sensitive to anthropometrics and the claim that lack of sufficient weight for women at birth has important effects across generations (Averett & Korenman, 1996; Conley & Bennett, 2000; Gortmaker et al., 1993; Haskins & Ransford, 1999; Sarlio-Lahteenkorva & Lahelma, 1999). In particular we look at the effects of increasing birthweight on schooling attainment and on adult body mass, height, and earnings, comparing estimates that control for endowment effects with those that do not, to assess the confounding role of genetic influences and other family endowments. We also estimate the extent to which the intergenerational birthweight relation is due to the heritability of body mass and preferences. Finally, we use our estimation methodology to estimate how reducing the birthweight gap between low- and high-income countries would reduce world earnings disparities.

Our estimates provide a number of clear results. First, they indicate that increasing fetal growth has a significant positive effect on schooling attainment that is underestimated by 50% if there is no control for genetic and family

background endowments as in cross-sectional estimates. Second, our estimates indicate that intrauterine nutrient consumption does not have any persistent effects on adult BMI—increasing birthweight is not a cause of adult obesity. Third, our results indicate that the heritable component of birthweight plays the dominant role in the intergenerational correlation of birthweights. Fourth, the estimates indicate that intrauterine nutrient intake significantly affects adult height, consistent with the literature that makes use of height statistics to gauge childhood nutritional investments over time and across countries (Fogel, 1994; Schultz, 1999; Steckel, 1995; Strauss & Thomas, 1998). Fifth, although we find evidence that augmenting birthweight, particularly among lower-birthweight babies, would have significant labor-market payoffs, we show that the strong cross-country correlation between incomes and birthweight substantially overstates the reduction in world earnings inequality that would arise from reducing cross-country disparities in birthweights.

II. Identification of Prebirth Nutrient Intake Effects on Adult Outcomes

We discuss how information on the birthweights of MZ twins can be used to identify the effects of an exogenous increase in intrauterine nutrient consumption on schooling, adult physical characteristics, adult wages, and birthweights for one's children when there is endowment heterogeneity among individuals that is correlated across generations.³ The key identifying assumption is that although the average of the nutrient intakes for any twin pair may be correlated with their common endowment, which may reflect their genetic heritage and resource allocation decisions of their parents, the birthweight difference within an MZ twin pair reflects purely random differences in nutrient intakes that may arise, for example, from differences in womb position. Twin-specific intakes that are expressed in birthweight differences are thus orthogonal to their identical endowments.

Consider a linear representation of the birthweight production function expanded to distinguish between common and specific nutrient intakes for each twin in a pair:

$$B_{jk} = C_{jk} + C_k + B_{0k}, \quad (1)$$

where B_{jk} = birthweight for twin j in twin pair k , C_{jk} = deviation in nutrient intake for twin j from average (common) nutrient intake of pair k , C_k = nutrient intake common to both twins in pair k , and B_{0k} = common genetic endowment to the pair. In any model in which parents care about the consequences of the health of their children, the covariance between C_k and B_{0k} will be nonzero due to

³ We broadly define nutrient consumption to include all factors contributing to fetal development, including, for example, oxygen. We have not been able to find in the biomedical literature any evidence that birthweight differences between members of MZ twinships reflect other differences, such as mutations just after cells divide or random deformities, that are unrelated to nutrient consumption differences.

optimizing behavior as long as either (i) parents, who can (only) choose the average or common level of inputs, know B_{0k} ,⁴ or (ii) the child's health endowment is correlated with the parents' endowments and that affects their resource constraint. However, differences in the twin-specific intakes C_{jk} cannot be functions of B_{0k} or of C_k .

The reduced-form linear relation between an adult twin's log wage (or any other adult outcome) and his or her intrauterine nutrient intake for twin j in family k is given by

$$\ln W_{jk} = \alpha_1(C_{jk} + C_k) + \alpha_2 B_{0k} + \mathbf{Z}_k \boldsymbol{\alpha}_3 + \mu_k + v_{jk}, \quad (2)$$

where the α 's are coefficients, \mathbf{Z}_k is a vector of exogenous family and community characteristics determining postbirth human capital investments in the twin (such as income and prices of inputs like schooling), μ_k is the common earnings endowments of the twins, and v_{jk} is a random, twin-specific error. Neither the nutrient inputs nor the endowments are observed in equations (1) and (2), only birthweights. However, the effect of in utero nutrient intake α_1 can be identified if we substitute for the unobserved inputs in equation (2) the observed birthweight using equation (1) and the difference across the twins. We then get the simple difference regression in terms of the observable birth outcome:

$$\Delta \ln W_{jk} = \alpha_1 \Delta B_{jk} + \Delta v_{jk}, \quad (3)$$

where Δ is the twin difference operator.

To highlight the identification problem using conventional cross-sectional data, it is useful to relate the observed variable moments to the unmeasured variables and the coefficients in equation (2). The regression coefficient in equation (3) exploiting twin differences is $\text{cov}(\ln \Delta W_{jk}, \Delta B_{jk}) / \sigma^2(\Delta B_{jk})$, where $\text{cov}(\Delta \ln W_{jk}, \Delta B_{jk}) = \alpha_1 \sigma^2(\Delta C_{jk})$ and $\sigma^2(\Delta B_{jk}) = 2\sigma^2(\Delta C_{jk})$, which can be solved for α_1 . In contrast, the coefficient obtained from a cross-sectional regression of the log wage on birthweight for nonidentical individuals or siblings neither identifies the effect of variation in nutrient intake nor provides insights into the roles of endowments in input choices. To see this, note that the regression coefficient in this case is $\text{cov}(\ln W_{jk}, B_{jk}) / \sigma^2(B_{jk})$, which is given by⁵

$$\begin{aligned} & \{\alpha_1 [\sigma^2(C_{jk}) + \sigma^2(C_k) + \text{cov}(C_k, B_{0k})] + \alpha_2 [\sigma^2(B_{0k}) \\ & + \text{cov}(C_k, B_{0k})] + \alpha_3 [\text{cov}(B_{0k}, \mu_k) + \text{cov}(C_k, \mu_k)]\} / \\ & [\sigma^2(C_{jk}) + \sigma^2(C_k) + \text{cov}(C_k, B_{0k}) + \sigma^2(B_{0k})]. \quad (4) \end{aligned}$$

It is clear from equation (4) that a cross-sectional regression of wages on birthweight does not identify the effect α_1 of increasing nutrient intake on wages, and indeed, depending on the covariances between endowments and the response

⁴ Or parents are uncertain about B_{0k} and parental expectations of B_{0k} are positively correlated with the actual values of B_{0k} .

⁵ For simplicity, we have dropped from this expression the price and income terms (which are presumably observables).

of inputs to endowments, the sign of the regression coefficient could be opposite that for α_1 .

The contrast between the cross-sectional relationship between birthweight and the wage in equation (4) and the effect obtained using the within-MZ estimator, which provides α_1 , can provide some information on the existence of endowment effects. First, if there were no endowment influences—no effect of variation in B_0 or in μ_k —the cross-sectional and within-MZ estimates would be identical. If they are different, two comparisons of the cross-sectional and within-MZ estimators are of particular interest. One may find from the within-MZ estimator that α_1 is positive while the cross-sectional birthweight estimated effect is zero. In that case it must be true that, assuming that both the birthweight and earnings endowments positively contribute to earnings ($\alpha_2, \mu_k > 0$), either the birthweight and earnings endowments are negatively correlated or nutrient intakes are provided less to those with higher endowments, or both. A second polar case is when α_1 is zero. In that case the cross-sectional birthweight relationship arises solely from endowment-allocation effects and covariances, which are thus identified.

Note that an alternative estimation scheme using parental characteristics as instruments to predict intrauterine nutrient consumption and thus the effect of birthweight on the child's adult wage is ruled out for two reasons: First, parental characteristics affect postbirth human capital investments and thus belong in the reduced-form wage equation. Second, parental characteristics may be correlated with the child's endowments, as is consistent with the empirical estimates for the determination of child schooling in Behrman and Rosenzweig (2002). We show empirically below that there is an intergenerational correlation in birthweight endowments.

Finally, there is a relatively large literature exploiting differences between MZ twins in schooling attainment to identify schooling effects on earnings in the presence of unmeasured earnings endowments (for example, Ashenfelter & Krueger, 1994; Behrman et al., 1994). The problem with this application of twinning is that postbirth differences in resources allocated across genetically identical twins such as schooling may reflect postconception, earnings-relevant differences in twin characteristics unobservable to the researcher but apparent to the parents. The evidence on birthweight differences across MZ twins suggests that such twins are no longer identical after conception, and these differences may influence postbirth parental resource allocations. Evidence of significant effects of birthweight on adult outcomes thus would call into question the estimates from twin-based schooling studies, which assume that schooling differences across genetically identical twins are orthogonal to other unmeasured human capital attributes of the twins. In contrast, differences in intrauterine growth across MZ twins cannot be due to parental choice—they are random with respect to parental and child endowments.

III. Data

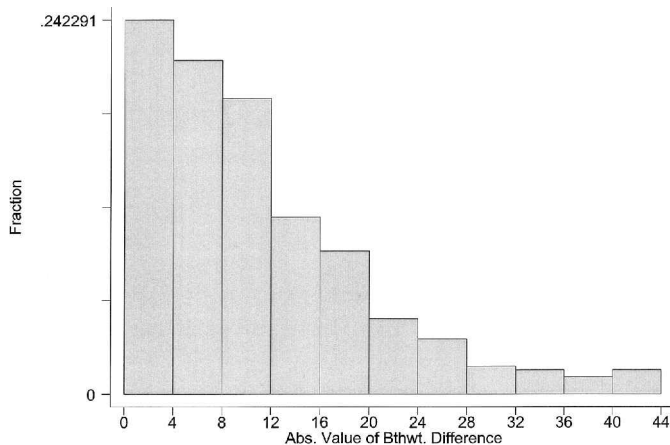
To implement the twins-based methodology to identify the long-term consequences of intrauterine nutrition, we use information from a recent survey of twins that we undertook. The twins data were obtained by resurveying a subset of the twins from the Minnesota Twin Registry (MTR) based on a survey instrument designed by us in collaboration with the Temple University Institute of Survey Research. The MTR is the largest birth-record-based twins registry in the United States, assembled between 1983 and 1990 starting with birth records on all twins [both MZ and dizygotic (DZ)] born in Minnesota in 1936–1955. Details of the sample and its characteristics are in Lykken et al. (1990). The MTR staff obtained from the Minnesota State Health Department all birth certificates reporting multiple births. These birth certificates provide information on both the birthweights and the gestation of the twins. The MTR staff located over 80% of the twins and sent them a four-page biographical questionnaire (BQ) in the mid 1980s, with an introductory newsletter describing the project and a letter signed by Minnesota's governor urging participation.⁶ 80% of the individuals contacted, 71% of concordant pairs, supplied information for the BQ.

Our survey instrument was mailed out in May 1994 to the 5,862 members of same-sex twin pairs who had filled out the BQ and for whom the MTR had current addresses. An additional 776 members of same-sex pairs for whom updated addresses had been located between May and September 1994 were sent questionnaires in November 1994. Altogether 3,682 twins returned completed questionnaires, for a response rate of surviving twins of over 60%. Information obtained included the height and weight of the twins at the time of the survey, their schooling attainment, their work experience and wages on the last job, their parents' characteristics, and the birthweights for their first four children.⁷ The estimates in this paper are based on the returned questionnaires from (i) all women ($N = 1418$), (ii) women in MZ twin pairs ($N = 804$), (iii) twin mothers ($N = 1207$), and (iv) MZ-twin-mother pairs ($N = 608$) for whom the key birthweight, gestation, and self-reported height, weight, and earnings variables are available.

⁶ Among the first six birth cohorts studied (birth years 1936, 1937, 1938, 1949, 1954, and 1955), 27% of the same-sex pairs identified through birth certificates are known to have been broken by the death of at least one of the members prior to our survey. This raises the issue of whether the sample of intact adult twin pairs is selective. Almond et al. (2002) show, using matched birth and death records, that differences in the birthweights of twins do not importantly affect differences in their likelihood of dying within the first year of life. This suggests that death may not be selective, although in Almond et al.'s study identical and nonidentical twins could not be distinguished, so that the pure effect of in utero intake on infant mortality is not identified.

⁷ The item response on returned questionnaires is very high, exceeding that on recent Current Population Surveys and the 1990 Census. For example, only 9% of ever employed workers in our sample did not answer the questions on earnings or self-employment income; on the CPS more than 20% do not.

FIGURE 2.—DISTRIBUTION OF THE ABSOLUTE VALUE OF BIRTHWEIGHT DIFFERENTIALS (OZ.) AMONG IDENTICAL TWINS



There are two features of the data that are particularly relevant to the analysis of birthweight effects on adult outcomes using within-twin differences across twins. First, the birthweight information is based on measures from birth certificates, and thus the birthweights for the twins are not subject to recall error. It is well known that estimates based on within-twin or sibling-pair differences (Bishop, 1976; Griliches, 1979) are particularly prone to bias from measurement error, so that accurate measures of birthweight are critical. Moreover, because the birthweights of twins are assessed at the same time and by the same measurer, most of the measurement error will be common to the twins and thus will be eliminated using within-twin estimators. The second feature of the sample, and of twins in general, relevant to the identification of birthweight effects is that there is substantial variation in birthweight between genetically identical twins, variation that is not correlated with their common environment or genes. In our sample, the average absolute value of the difference in birthweights within pairs of MZ twins is 10.5 ounces, with substantial variation across pairs (figure 2).⁸ The corresponding figure for same-sex nonidentical twin pairs is 11.2 ounces. Thus there is ample and real within-twin variation to obtain precise estimates of birthweight input effects using twin difference methods.

For each twin we have constructed a measure of fetal growth based on the birth certificate information—birthweight divided by gestation. We use this measure rather than birthweight in obtaining our initial estimates for two reasons. First, the literature has suggested that this measure better reflects the healthiness of a child, as the leading proximate cause of low birthweight in the United States and other developed countries is preterm births (Pojda & Kelley, 2000; Strauss, 2000). In our sample, 20.5% of the twins were, by the standard definition (less than 37 weeks of

⁸ Having a greater birthweight in a twin pair is not correlated with being first-born in our data. The average difference between first- and second-born twins in our sample is a statistically insignificant 0.5 ounces.

gestation), preterm births. Second, although within-twin differences in birthweight are by definition for the same gestation, variation in birthweight across nontwin births also reflects variation in gestation. Normalizing birthweight by gestation makes the two estimates comparable.⁹

Another important feature of the survey is that, to maximize the size of the sample of female twins that could be used for analysis of wage determination and to avoid sample selectivity in recognition of the intermittent labor force participation of women, the earnings on the last job were elicited rather than only earnings in the year prior to the survey. A well-known problem in analyzing the wages of women is that at any given survey date many women may not be in the labor force, resulting in a selective sample of female earners (Gronau, 1974; Heckman, 1974). Only 82% of the women in the sample, for example, worked in 1993, the survey reference year. But 97% of the sample members worked at some point in their lives, 91% in the five years preceding the survey. Finally, to take into account the variability in work time during the year in which earnings are reported, the wage, earnings, and time-worked information was used to construct full-time annual earnings based on either earnings in 1993 or on the last job, inflated in the latter case by the relevant CPI.

The existence of information on standard family background characteristics—parental schooling and occupation—permits us to explore to what extent the differences between the cross-sectional and within-MZ estimates of the effect of fetal growth is changed with control for these characteristics. For this purpose we converted the occupation information for the twins' fathers into a singulate measure of lifetime earnings by using information from the 6% sample from the 1990 Census on all men aged 40 through 59 who worked in 1989 and resided in the states of Minnesota and Wisconsin, reflecting the principal residence states of the twins' parents.¹⁰ The occupational earnings equation we estimated ($n = 19,183$) is of the following form:

$$\ln y_{ik} = \eta_{0k} + \sum_k \eta_{1k} O_k + \sum_k \eta_{2k} O_k age_i + \eta_{3k} age_i + \sum_k \eta_{4k} O_k S_i + \eta_{5k} S_i + \eta_{5k} S_i age_i + u_{ik}, \quad (5)$$

where $\ln y_{ik}$ = log wage and salary plus self-employment earnings for individual i in occupation k , O_k = indicator for occupation k , S_i = schooling attainment, and u_{ik} is a

⁹ We could make the cross-sectional and within-twin estimates of birthweight comparable by also including gestation in the cross-sectional specifications, but the birthweight coefficient would then in any case reflect fetal growth. Estimates of the effects of variation in birthweight in which gestation is included as a separate variable do not differ importantly from the fetal growth estimates reported here.

¹⁰ 74.9% of the surviving parents of the respondents resided in Minnesota and Wisconsin at the time of the survey. The next most represented state (California) was the residence of only 3.2% of parents. We do not make similar estimates for the twins' mothers, because high proportions were housewives or worked in family enterprises (including farms).

TABLE 1.—MEANS AND STANDARD DEVIATIONS: MINNESOTA TWINS SAMPLE

	Mean	Standard Deviation
All female twins (<i>N</i> = 1418):		
Fetal growth (oz. per week of pregnancy)	2.34	.420
Birthweight (oz.)	90.2	17.9
BMI (lb./in. ²)	0.0361	0.00738
BMI (kg/m ²)	25.4	5.20
Height (in.)	64.7	2.67
Weight (lb.)	150.9	33.0
Schooling (yr.)	13.8	2.20
Hourly wage (1993 \$)	11.45	7.20
Age	45.6	5.45
Mother's schooling (yr.)	11.4	2.66
Father's schooling (yr.)	10.7	3.26
Father's occupational earnings (\$)	28,637	14,298
Female twins with at least one own (biological) child (<i>N</i> = 1207):		
Birthweight of twin's first child (oz.)	118.1	19.7
Fetal growth of mother	2.34	.426
Birthweight of mother	90.1	18.0
Age	45.9	5.51

random disturbance term. Based on these estimates and the information on the principal occupation and schooling attainment of each father of the twins available in the data, we computed the occupational earnings of each father at age 50.¹¹

IV. Generalizability and Validity of Twins-Based Birth Outcome Estimates

It is important to consider whether the analysis of twins can be generalized to larger populations of interest. For example, twins are significantly smaller at birth than are nontwins. Table 1 reports descriptive statistics for the two samples of female twins that we use. As seen in table 1, the average birthweight of twins is 5 lb. 10 oz. The average birthweight in the general population is 7 lb. 6 oz. Moreover, by the standard definition of low birthweight—below 5 lb. 8 oz. (2.5 kg)—almost half of the twins were low-birthweight.

There are two issues with respect to the size differences between twin and nontwin births. The first is whether these differences at birth reflect genetic differences between twins and nontwins. The evidence suggests, however, that the smallness of twins at birth is the result of twinning rather than the selectivity of parents who produce twins. Figure 3 displays the average birthweights of the twins in the sample who are mothers, the birthweights of their first-born children (almost none of whom are twins themselves), and the average birthweights of the first-borns of white mothers in the 1998 round of the NLSY. As can be seen, the average birthweight of the (nontwin) children of the twins, 7 lb.

6 oz., is the same as that of the nontwin children of nontwin mothers in the general population.¹²

A second issue with respect to the size differences between twins and nontwins at birth relates to whether or not the relationship between birthweight and/or fetal growth and adult outcomes is nonlinear. If, for example, the effects of increasing fetal growth on adult outcomes are stronger at low levels of fetal growth and perhaps even nonmonotonic, then estimates based on the twin population may overstate the average treatment effect for the general population of singletons in developed countries such as the United States. Figure 4 provides the distribution of fetal growth rates for the twins sample and for the singleton births of white mothers occurring between 1979 and 1983 in the NLSY. As can be seen, more than three-fourths of twin births are in the lower half of the singleton distribution. To obtain estimates that are more relevant to the general United States population and to assess whether the effects of birthweight on outcomes are nonlinear, we will estimate the birthweight-outcome relationships both using the unweighted twins sample and using sample weights based on the distribution of singleton births as given in figure 4. Given that the supports of the two distributions overlap except for the very highest-fetal-growth births, the weighted within-twin estimates will be more generalizable to the general population in the presence of nonlinearities.

The possibility of heterogeneous treatment effects by birth outcome levels may also create problems for inferences about the source of bias between OLS and within-MZ-twins estimates of birth outcome effects obtained from the same population. Because differences in birth outcomes within a twin pair are relatively small, the within-twin estimator provides a local average treatment effect (LATE) over the support of the distribution of (within-twin) average birth outcomes. In particular, if the true relationship between nutrient intake in the womb and adult outcomes differs by the level of the within-twin average nutrient intake, then the full-sample within-twin estimate of α_1 is

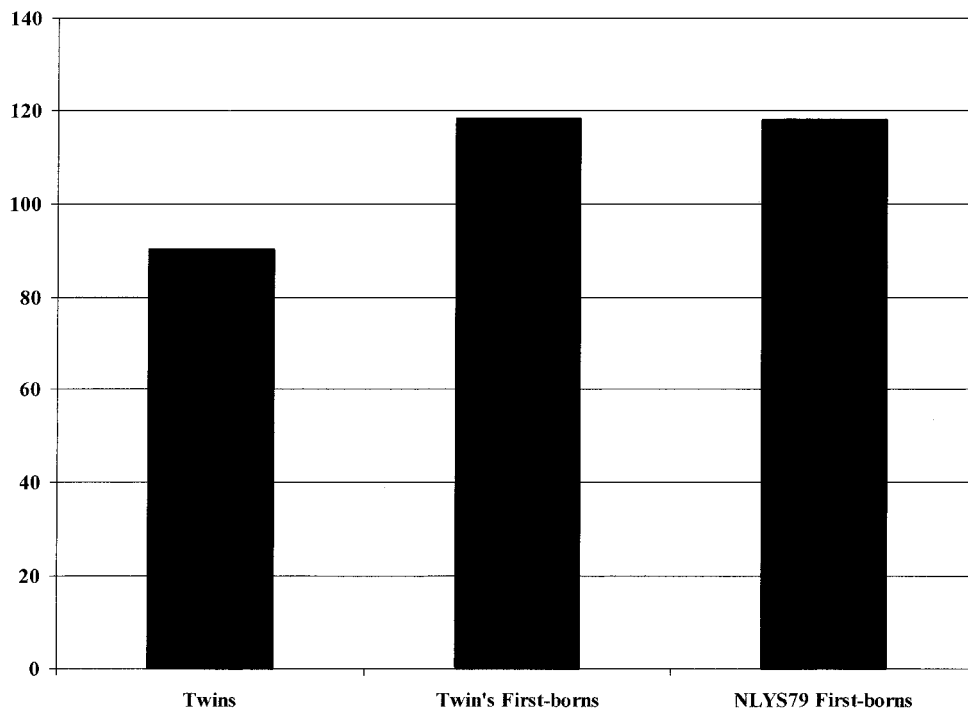
$$\alpha_1 = \sum_i f_i \alpha_{1i}, \tag{6}$$

where f_i is the proportion of twin pairs whose average fetal growth is B_i . The OLS cross-sectional estimate, however, imposes a linear functional form over the support of the birth-outcome distribution. We will thus estimate nonparametric regressions to assess whether or not nonlinearities can account for the difference between the OLS and within-MZ-twins estimates. An advantage of the within estimator is that we can obtain unbiased within-twin LATE estimates for any interval of the *average* birth outcomes of twin pairs (for example, in the low-birthweight range) because all

¹² This is not because the twins who are mothers had higher birthweights than the twins who did not have any children. Indeed, the mean birthweight of the twin mothers was 3 oz. less than that of the total sample of female twins.

¹¹ The estimates are available on request from the authors.

FIGURE 3.—AVERAGE BIRTHWEIGHTS (OZ.): TWINS, THEIR FIRST-BORNS, AND WHITE NLSY79 MOTHERS' FIRST-BORNS



endogenous sources of selectivity for the pair are common to the twins and thus are swept out.

Another aspect of the issue of whether or not estimates based on differences between twins are generalizable to the general population concerns the effect of twinning on parental resource allocation behavior. Life cycle birth outcome effects reflect in part decisions by parents to allocate resources in response to birth outcome differences. In any family with more than one birth, given a common resource constraint, the allocation of resources to children reflects a competition between siblings, with at the margin one child obtaining more resources at the expense of another. The issue is thus whether the competition among twins for parental resources is significantly different than among siblings whose births are separated over time.¹³ In an environment with no credit constraints and forward-looking parents, decisions even about a first child are based on expectations about the characteristics of subsequent children, and the resource allocations for twins and nontwin siblings may not be very different. Given credit constraints, however, the possibility arises that the competition between twins may be more acute than among spaced siblings and thus within-twin estimates overstate the effects of birth outcome differences across the general population.

Empirical evidence suggests that the spacing between nontwin siblings does not affect within-sibling estimates.

¹³ Because of intrafamily responses to sib differences, no estimator based on siblings may be relevant to households with only one child. One-child (completed) families are rare. In the 1990 U.S. Census, among ever-married white women aged 45–60 who had any children, only 10% had only one child.

Olneck (1977) compared the correlations between 375 siblings for earnings, test scores, schooling, and occupation at two life cycle points between siblings with age differences of 3 years or less and more than 3 years, found no statistically significant difference in any but current occupation, and concluded that exception to be due to “error.” We examined the relationship between birthweight and height for 494 sibling pairs aged 15–21 in the 1998 round of the NLSY. We used a within-sibling estimator, interacting the interval in years between each siblings’ birth and his or her birthweight. The estimates we obtained are

$$\text{height (in.)} = 0.0319 \times \text{birthweight (oz.)} - 0.0069 \times \text{birthweight} \times \text{spacing}, \quad (7)$$

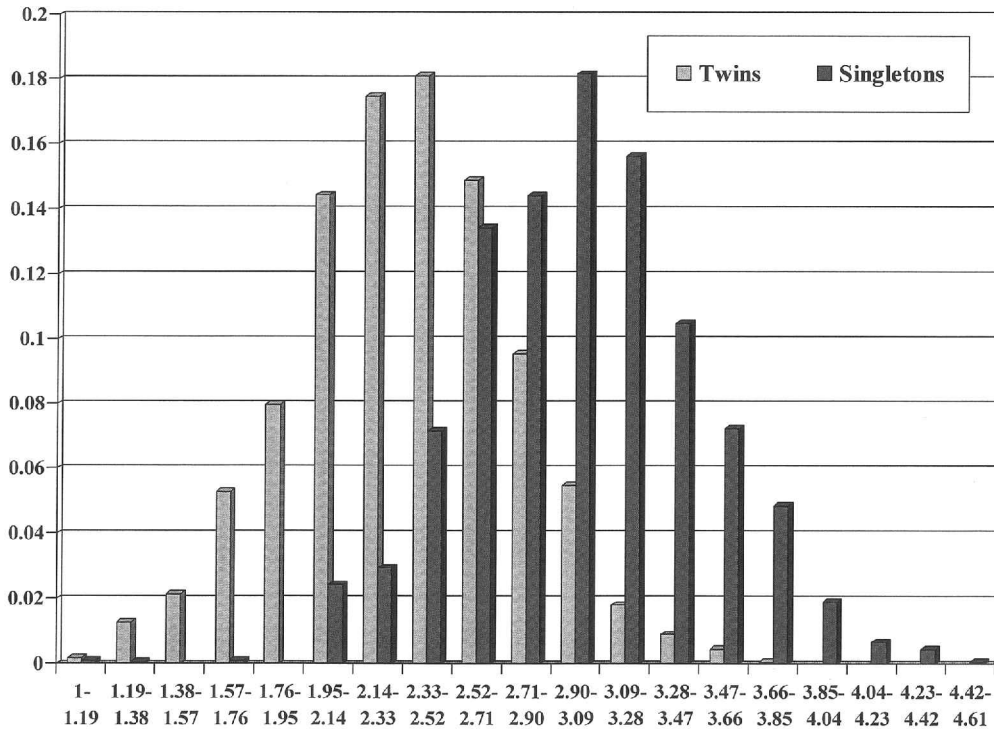
(3.75) (0.49)

where the absolute values of the *t*-ratios are in the parentheses. Estimates of the birthweight relationship for closely spaced siblings and those spaced up to 6 years evidently do not differ.¹⁴

The final issue we consider about the generalizability of twins-based estimates of birth outcome effects is whether birth outcome differences between MZ twins reflect the same in utero growth factors, other than genetic factors, that affect the development of single fetuses. There is evidence that the increased fetal stress associated with sharing the womb results in acceleration of the maturation of the lung and other organs, notably the brain (Amiel-Tison & Gluck,

¹⁴ Over 28% of the sibling pairs were born approximately 1.5 years apart; 16% were born 4 or more years apart.

FIGURE 4.—DISTRIBUTION OF FETAL GROWTH RATES (LB./IN.²): FEMALE TWINS AND SINGLETONS



1995). This might mean that a shorter gestation for twins may have less deleterious effects than for singleton births and twins-based estimates will understate the effect of intrauterine development. There is also some evidence that the larger of two twins is more susceptible to diabetes. If this does not hold for singletons, it again means that *ceteris paribus* the effect of twins' birthweight may be an underestimate of the effect of singletons' birthweight. Differences in the intake of nutrients and oxygen within the womb will most likely reflect differences in fetal positions, which may be more unusual than those for single births and which may affect oxygenation more than nutrients. In that case the within-twin estimates may overstate the efficacy of programs that emphasize increased nutrient consumption by the mother compared with initiatives that discourage maternal smoking. Clearly more research on the fetal development of twins and nontwins is needed for a more precise interpretation of birth outcome effects, however they are estimated.

V. Estimates

A. Fetal Nutrient Intake Effects on Human Capital, Physical Attributes, and Earnings

Table 2 reports OLS and within-MZ estimates of the relationships between fetal growth and, in the order of the table, subsequent adult schooling attainment, adult BMI,¹⁵

adult height, and the log wage, using the unweighted twins sample. Two OLS estimates are given. The first includes only fetal growth and age (which controls for secular trends, for example, in schooling). The second in addition includes mother's and father's schooling and father's occupational earnings.

The estimates for schooling in the first three columns indicate that, whatever the estimation procedure, a child's fetal growth and his or her subsequent schooling attainment are significantly and positively related. However, the OLS estimate understates α_1 for schooling attainment by more than 50%. This is the case whether or not the observed family background controls are included in the OLS estimates. In fact, though the observed family background estimates have significant positive coefficient estimates, their inclusion reduces the gap between the OLS and the within-MZ estimate of the effect of fetal growth only by approximately a fifth. Most of the difference between the simple OLS estimates and the within-MZ estimates is evidently not due to the failure to control for these standard family background variables. The difference between the OLS and the within-MZ estimates suggest that in the population either human capital inputs tend to be negatively correlated with health endowments or health and ability endowments are negatively correlated. The within-MZ point

¹⁵ The relation between BMI and health is nonlinear. For very undernourished populations, increases in BMI indicate better health. But for increases above the normal range, increases in BMI indicate being

overweight or obese. For the population that we study, however, undernourishment is not an issue. The mean BMI is in the overweight range (>25 kg/m²; see table 1), so increases in BMI for this sample basically imply moving from normal to overweight to obese status.

TABLE 2.—OLS AND WITHIN-MZ TWIN ESTIMATES OF FETAL GROWTH ON OWN SCHOOLING ATTAINMENT, BMI, HEIGHT, AND \ln WAGE

Variable	Schooling			BMI ($\times 10^{-3}$)			Height			In Wage		
	OLS	OLS	Within-MZ	OLS	OLS	Within-MZ	OLS	OLS	Within-MZ	OLS	OLS	Within-MZ
Fetal growth	0.313 (2.06) ^a	0.385 (2.76) ^a	0.657 (3.11) ^b	1.06 (2.04) ^a	0.903 (1.76) ^a	0.292 (0.39) ^b	1.52 (8.06) ^a	1.50 (8.07) ^a	1.48 (8.72) ^b	0.0269 (0.81) ^a	0.0329 (1.01) ^a	0.190 (2.47) ^b
Age	-0.0429 (3.53)	-0.0054 (0.47)	—	0.130 (3.20)	0.0904 (2.15)	—	0.0261 (1.77)	0.0282 (1.80)	—	-0.00657 (2.58)	-0.0031 (1.19)	—
Mother's schooling	—	0.166 (5.35)	—	—	-0.0267 (0.25)	—	—	0.0388 (1.02)	—	—	0.0176 (2.71)	—
Father's schooling	—	0.112 (3.70)	—	—	-0.216 (2.03)	—	—	-0.0257 (0.66)	—	—	0.00598 (0.89)	—
Father's earnings ($\times 10^4$)	—	0.160 (3.38)	—	—	-0.317 (1.51)	—	—	0.0072 (0.89)	—	—	0.00345 (1.98)	—
Constant	15.0 (22.4)	9.60 (12.8)	—	27.7 (12.2)	33.3 (11.8)	—	59.9 (72.1)	59.5 (58.1)	—	2.55 (17.7)	2.01 (11.9)	—
N	1418	1418	804	1418	1418	804	1418	1418	804	1418	1418	804

^aAbsolute value of robust t -statistic with clustering by twin pairs.

^bAbsolute value of robust t -statistic.

estimate of α_1 indicates that efforts taken by a mother to increase fetal growth such that birthweight is increased by 1 lb. at birth (an increase in fetal growth of 0.4 oz./week) would result in almost a third of a year more of schooling for her child.

The results for BMI suggest clearly that genetics play a role in BMI, though there is some inverse association with family background. The within-MZ estimate of α_1 for body mass is essentially zero—fetal inputs that increase birthweight (and schooling attainment) do not result in additional body fat later in life. Thus the primary reason for the positive association between fetal growth and adult BMI is evidently the influence of genetic endowments, as indicated in the expression (4).¹⁶ This suggests that increasing birthweight, for given genetic endowments, does not result in obesity of the child later in life. However, increasing birthweight does result in increased adult height. The within-MZ point estimate, which is not different from the OLS estimate, suggests that a 1 lb. increase in birthweight brought about by increased womb nutrients would increase adult height by 0.6 in.

Finally, the within-MZ point estimate of the reduced-form effect of fetal growth on wages indicates that augmenting a child's birthweight by a 1 lb. increases her adult earnings by over 7%. That increasing fetal growth and birthweight increases earnings is not surprising given the estimates suggesting that increasing birthweight increases schooling. However, the OLS estimate of the fetal-growth-earnings relationship indicates the absence of any correla-

tion between fetal growth and the wage.¹⁷ This is also consistent with the results for schooling, suggesting a general negative correlation between the genetic component of birthweight and ability endowments or inputs. These relationships and variation in endowments in the population evidently obscure the fact that increasing fetal growth augments human capital and earnings.

B. Are There Important Nonlinearities in Birth Outcome Effects?

As discussed, the comparisons between the OLS and within-MZ estimates that shed light on the role of endowments may be misleading if there are important nonlinearities between fetal growth and the adult outcomes in the cross section. We estimated nonparametric regressions, using a sensitive bandwidth of 0.4, for each of the four dependent variables in table 2. The cross-sectional nonparametric and OLS relationships between fetal growth and schooling, BMI, height, and the log wage are depicted in figures 5–8, respectively. The graphs suggest that for schooling, height, and BMI the relationship is monotonic and, over most of the range of fetal growth values, linear, although the slopes for schooling and height appear to be slightly larger in the bottom quarter of the distribution of fetal growth rates. Spline estimates indicate, however, that these slope differences are not statistically different for either outcome. The differences between the OLS and within-MZ estimates for schooling, BMI, and height thus

¹⁶ This result is consistent with the findings in the literature on the heritability of adult BMI based on adoptions, MZ twins reared apart, and comparisons of MZ and DZ twins (Vogler et al., 1995; Allison et al., 1996; Herskind et al., 1996).

¹⁷ This is the case for both the simple and extended OLS estimates. However the comparison between the point estimates of the two provides an illustration of the possibility that the estimate of the effect of fetal growth in the extended OLS estimate need not be between the estimates in the simple OLS and the within-MZ cases.

FIGURE 5.—ADULT SCHOOLING ATTAINMENT (YR.) AND FETAL GROWTH (LB./IN.²): OLS AND NONPARAMETRIC ESTIMATES

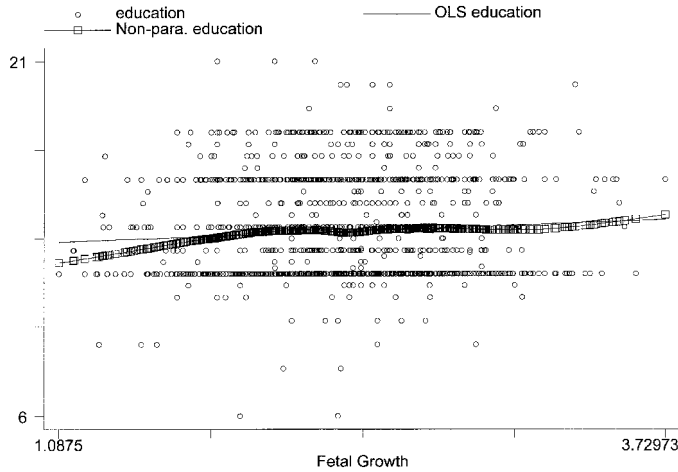


FIGURE 7.—ADULT HEIGHT (IN.) AND FETAL GROWTH (LB./IN.²): OLS AND NONPARAMETRIC ESTIMATES

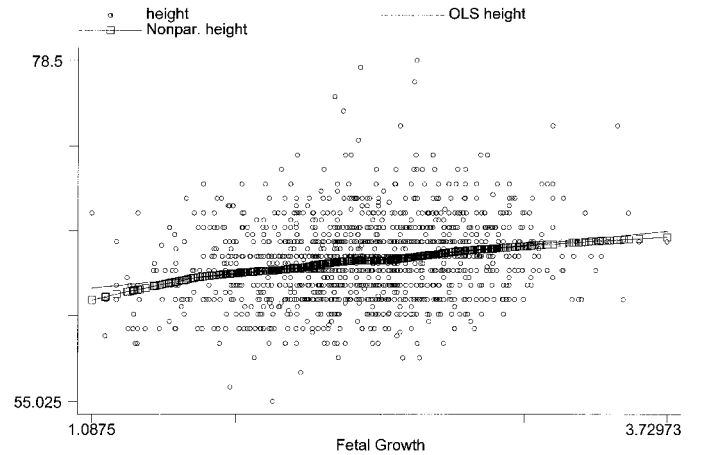


FIGURE 6.—ADULT BMI (LB./IN.²) AND FETAL GROWTH (LB./IN.²): OLS AND NONPARAMETRIC ESTIMATES

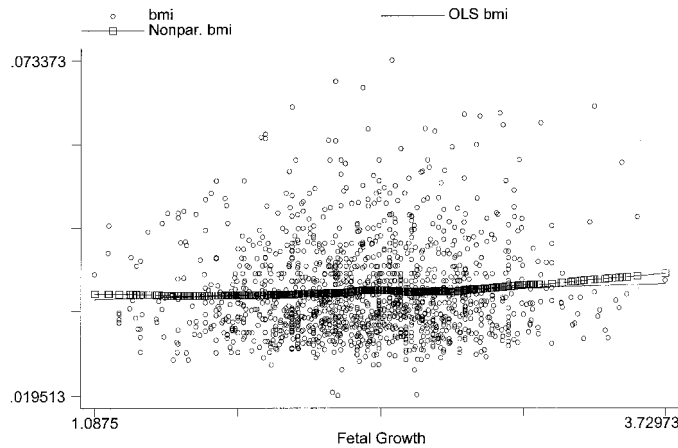
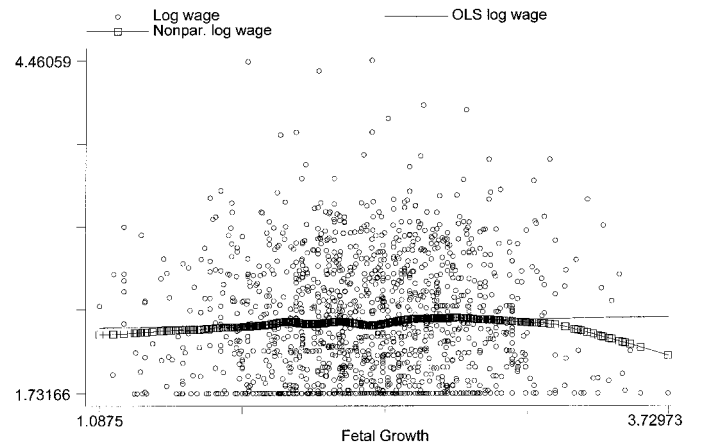


FIGURE 8.—LOG WAGE (\$/YR.) AND FETAL GROWTH (LB./IN.²): OLS AND NONPARAMETRIC ESTIMATES



appear solely to reflect biases due to endowment heterogeneity.

The cross-sectional fetal growth effect on the log wage, however, is nonmonotonic—positive through the bottom quarter of the fetal growth distribution of the sample, and negative in the top quarter. If the within-MZ estimator just provided an estimate of the average of local OLS slope coefficients seen in figure 8, the within-MZ estimate would be smaller than the OLS estimate. The fact that it is larger for the same sample when the effects of endowments are eliminated reinforces our interpretation as endowment heterogeneity bias.

Figures 5–8 do not say anything about the shape of the relationships between fetal growth rates and the adult outcomes based on within-twin differences. Indeed, if the within-MZ relationship between fetal growth and the log wage is similar to that depicted in figure 8, the estimate in table 2 may substantially overstate the fetal growth effect on earnings for singleton births, given that most of the fetal growth rates for twins are clustered in the bottom half of the

distribution for singletons. To explore the shape of the relationship between fetal growth and adult outcomes net of endowment heterogeneity we first test whether it is only crossing the low-birthweight threshold that matters for adult outcomes, as assumed in many if not most studies of the effects of birthweight effects. Appendix table A1 reports within-MZ estimates of the effects of low birthweight, as conventionally defined, on the same set of dependent variables with and without the fetal growth measure included in the specification. These estimates suggest that the continuous birthweight specification dominates—in no case is the coefficient on the indicator of low birthweight statistically significant when fetal growth was also included in the specification.¹⁸

To further assess the importance of nonlinearities and the generalizability of the twin-based results to the singleton population of births, we reestimated the fetal growth

¹⁸ Richards et al. (2001) also find that birthweight is positively correlated with adult outcomes above the low-birthweight threshold, although their results are based on cross-sectional estimates.

TABLE 3.—OLS AND WITHIN-MZ TWIN ESTIMATES OF FETAL GROWTH ON OWN SCHOOLING ATTAINMENT, BMI, HEIGHT, AND \ln WAGE: SAMPLE WEIGHTED USING U.S. SINGLETON DISTRIBUTION OF FETAL GROWTH RATES

Variable	Schooling			BMI ($\times 10^{-3}$)			Height			ln Wage		
	OLS	OLS	Within-MZ	OLS	OLS	Within-MZ	OLS	OLS	Within-MZ	OLS	OLS	Within-MZ
Fetal growth	0.152 (0.67) ^a	0.278 (1.36) ^a	0.625 (2.72) ^b	1.34 (1.61) ^a	1.02 (1.27) ^a	0.351 (0.04) ^b	1.62 (6.18) ^a	1.62 (6.26) ^a	1.14 (7.13) ^b	0.00585 (0.12) ^a	0.0231 (0.49) ^a	0.143 (1.78) ^b
Age	-0.0435 (2.71)	0.0051 (0.32)	—	0.147 (3.01)	0.0867 (1.68)	—	0.0185 (0.96)	0.0256 (1.25)	—	-0.00328 (1.40)	-0.00269 (0.82)	—
Mother's schooling	—	0.214 (5.83)	—	—	-0.0357 (0.26)	—	—	0.0529 (1.12)	—	—	0.0199 (2.55)	—
Father's schooling	—	0.0890 (2.32)	—	—	-0.363 (2.49)	—	—	-0.0161 (0.33)	—	—	0.0141 (1.83)	—
Father's earnings ($\times 10^4$)	—	0.138 (2.31)	—	—	-0.119 (0.50)	—	—	0.0601 (0.57)	—	—	0.00165 (0.81)	—
Constant	15.4 (17.2)	9.15 (8.89)	—	26.1 (8.7)	34.2 (9.31)	—	60.0 (58.2)	59.1 (46.0)	—	2.36 (17.5)	1.97 (9.62)	—
<i>N</i>	1,418	1,418	804	1,418	1,418	804	1,418	1,418	804	1,418	1,418	804

^aAbsolute value of robust *t*-statistic with clustering by twin pairs.

^bAbsolute value of robust *t*-statistic.

equations, using both OLS and the within-MZ estimator, weighting the sample using the U.S. singleton distribution of fetal growth rates shown in figure 4. Table 3 reports the weighted estimates. Given that the weighting gives substantially more prominence to higher fetal growth rates and given the decline in the cross-sectional slopes seen for schooling and the log wage as fetal growth increases observed in figures 5 and 8, it is not surprising to find that the weighted OLS estimates of fetal growth rates for these dependent variables are substantially lower than their unweighted counterparts, by 28% for schooling and by 30% for the log wage in the specifications including the parental controls, and by 49% and 78% in the simple specifications. The within-MZ estimates of α_1 are, however, more robust to the reweighting—the coefficient for schooling drops by less than 5%, and the coefficient for BMI remains statistically insignificant. The coefficients for height and the log wage drop by less than 25%. Inferences about biases due to endowment heterogeneity are thus robust to the reweighting of the sample.

The within-MZ coefficient on the log wage loses statistical significance through reweighting. Inspection of the within-MZ log wage coefficients across the range of the weighted distribution of twin-pair average fetal growth rates indicates that the relationship is positive and monotonic but the slope declines monotonically as average fetal growth increases. The within-MZ estimate of the birthweight effect on log earnings for the bottom third of the U.S. singleton distribution of fetal growth rates (fetal growth rate = 2.45) is statistically significant— $\alpha_1 = 0.2$ ($t = 2.17$); but $\alpha_1 = 0.08$ for the top third ($t = 0.33$). These results thus imply that focusing policy on lower-birthweight populations is warranted if the criterion is to increase overall productivity, but not if the goal is to increase overall levels of schooling.

C. How Much Will Reducing World Birthweight Inequality Reduce World Earnings Inequality?

We can use our data on twins and our within-twin estimator to assess how reducing differences in birthweight across low- and high-income countries would contribute to the reduction in world income inequality. In particular, we provide illustrative estimates of how much eliminating the birthweight gap between a low-income country and the United States reduces the disparity in their incomes, estimate how much the world variation in country-specific birthweights affects worldwide income inequality, and assess the extent to which the cross-country relationship between birthweight and income observed in figure 1 overstates the productivity effect of birthweight enhancement due to the existence of confounding unmeasured factors affecting birthweight and incomes.

The formula for computing the percentage earnings gain for a country j from closing the birthweight gap between it and some target country is

$$\%earnings\ gain_j = \sum_i f_{ij} \alpha_{1i} \cdot birthweight\ gap_j. \quad (8)$$

Given the nonlinearities we have identified in the effects of intrauterine consumption on adult earnings, the estimated birthweight gap effect will depend on the country's birthweight distribution, the set of country-specific f_i 's. To illustrate these estimated effects, we picked two countries with available information on birthweights—Malaysia and India. We chose Malaysia because it is in the middle range of all countries with respect to income and birthweight and because there are good data on birthweights (but less good data on gestation), from the Malaysia Family Life Survey (MFLS) for 1976–1977, that have been analyzed in other

TABLE 4.—PERCENT LOW-BIRTHWEIGHT BIRTHS AND MEAN BIRTHWEIGHT FOR SELECTED COUNTRIES

Country	% Low Birthweight	Mean Birthweight	Number of Births	Source
India	22.6	98.3 (24.5) ^a	8650	1998/99 DHS
Philippines	16.2	111.0 (26.3)	4337	1998 DHS
Pakistan	16.0	111.9 (28.8)	607	1991 DHS
Malaysia	14.3	108.3 (19.2)	3941	1976/77 MFLS
Senegal	11.2	110.9 (26.5)	2239	1997 DHS
Uganda	11.2	113.5 (28.2)	1889	1995 DHS
Brazil	9.1	114.3 (21.4)	4427	1996 DHS
Peru	9.0	114.6 (23.4)	10654	1996 DHS
United States—whites	8.9	118.2 (20.9)	1711	NLSY79, 1988 round

^a Standard errors in parentheses.

work studying the determinants of birthweight (Rosenzweig & Wolpin, 1988). We chose India because India appears to have the world’s highest incidence of low birthweight among countries for which there is available nationally representative information on birthweights [Demographic and Health Surveys (DHS), 1998–1999].¹⁹

Because of the absence of accurate worldwide information on gestation, we use estimates based on birthweight data. Table 4 provides the percentage of low-birthweight babies and mean birthweight that we have computed for a number of countries with available survey data on birthweight. The table indicates that the United States–Malaysia mean birthweight gap is about 10 oz. (9.2%)—comparable to between the United States and a number of African and Asian countries in the 1990s—and that for India in 1999 is 20 oz. (20.3%). Neither the U.S. distribution of birthweights for singleton births nor the distribution for twin births measures well the birthweight distribution for Malaysia or India. Indeed, the proportion of low-birthweight babies among the U.S. twins (49.7%) is more than twice as high as that for India (22.6%). The 1976–1977 birthweight distribution for Malaysia is between that for the U.S. twins and singletons, as depicted in figure 9. Given the evident nonlinearities in birthweight effects, we use the Malaysia birthweight distribution to estimate the birthweight effect for Malaysia, and the India birthweight distribution to obtain the India birthweight effect based on the within-MZ estimator.

Table 5 reports the within-MZ estimates of the effect of birthweight on log earnings based on the twins sample using no weights, the Malaysia weights, and the India weights. The estimates are also reported for the low-birthweight

population within each distribution and, for Malaysia and India, for the bottom half of their birthweight distributions. Similar to the estimates for fetal growth on log earnings, the estimates for birthweight by weighting scheme and by location in the distribution within a country’s birthweight distribution indicate that birthweight effects on log earnings are stronger at lower birthweights. Interestingly, the strongest birthweight effect is within the low-birthweight part of the distribution for India, with an average 20-oz. increase in birthweight resulting in a 14.8% increase in earnings.

The estimates reported in table 5 based on the relation (8) indicate that closing the average U.S.–Malaysia and U.S.–India birthweight gaps by a uniform increase in birthweight across the birthweight distribution for each country would increase earnings per worker by 4.1% and 8.6%, respectively. More realistically, confining the increase in birthweight to the bottom half of the birthweight distribution to make up the average gap would increase earnings by 4.8% and 9.2% in the two countries. These are not trivial effects. However, the U.S. per-worker GDP gap in PPP dollars for both countries is substantial—287% for Malaysia (1977) and more than an order of magnitude for India using 1989 GDP figures. Thus the birthweight gap plays a very small role in explaining the earnings gap for either Malaysia or India.

To estimate how much the worldwide variation in birthweight can account for world inequality in GDP per worker, we need to convert the available low-birthweight percentages by country to mean birthweights. The accurate, birth-record-based information on birthweights from our twins survey indicates that the distribution of birthweights is normal. Figure 10 shows the quantile plot of the birthweight distribution of twins against the normal, which exhibits a very close fit. Given the assumption of normality, we then need only assume a value for the standard deviation σ . To obtain this we used the empirical country-specific low-birthweight proportions and mean birthweights for the nine countries reported in table 4 to estimate σ , using the relationship

$$B_i = 88.18 - \Phi^{-1}(P_i)\sigma, \tag{9}$$

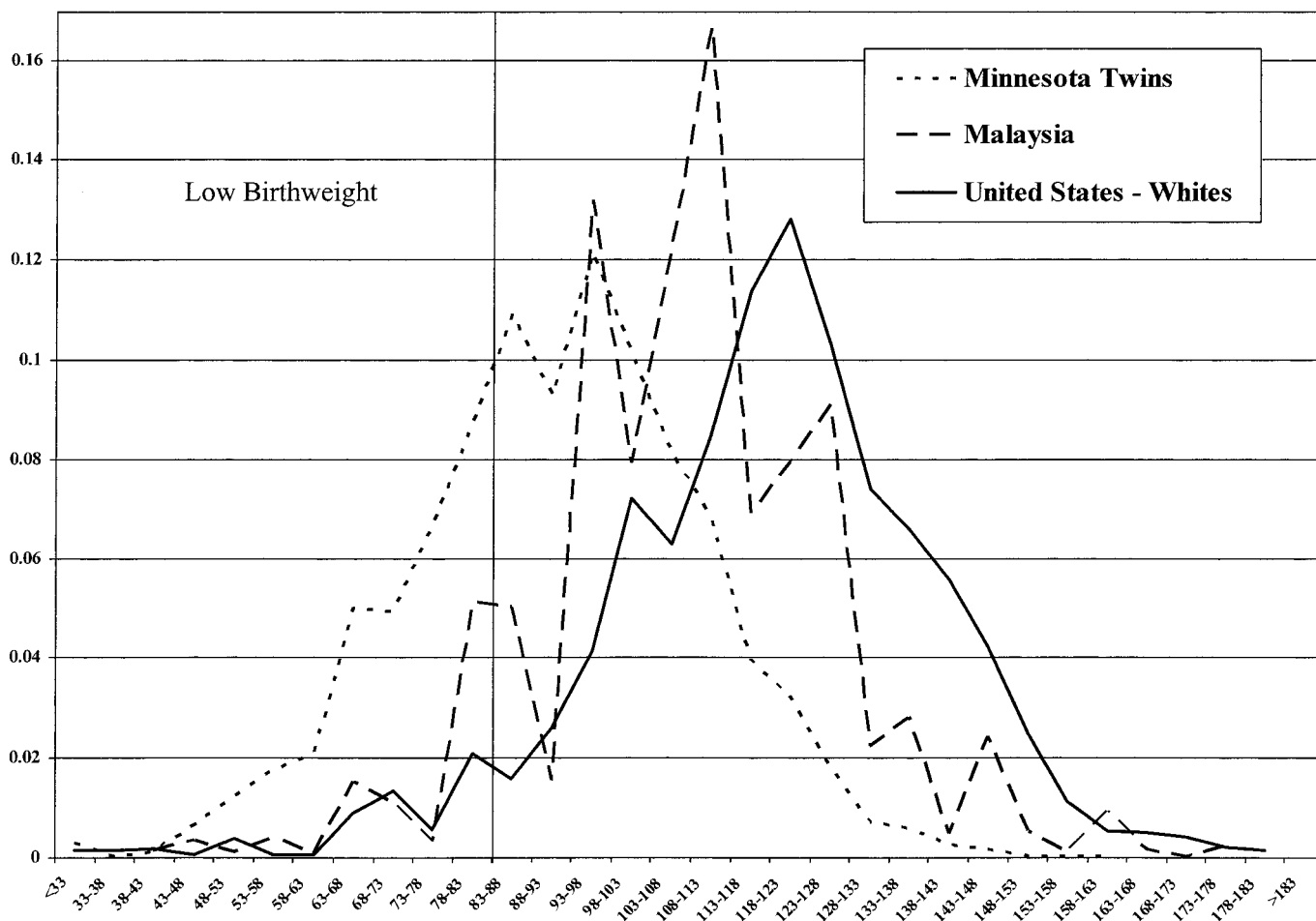
where Φ^{-1} is the inverse normal distribution, P_i is the proportion of low-birthweight births in each country, B_i is the country’s mean birthweight, and 88.18 is the 2500 g low-birthweight cutoff in ounces. The regression estimate of σ is 20.3 ($t = 25.3$), with an R^2 of 0.988.²⁰ Using equation (9) and the estimated σ to estimate the mean birthweight for the 112 countries for which we have both low-birthweight proportions and PPP GDP per worker, we obtain an estimate of the cross-country world variance in birthweight of 32.6 and a mean birthweight for the world of 115.0 oz.²¹

²⁰ This is close to the standard deviation of the U.S. white birthweight distribution of 20.9 in table 4.

²¹ If the countries are weighted by population, the average world birthweight is 112.9 oz., almost two-thirds of a pound less than the average birthweight in Sweden.

¹⁹ In figure 1 the incidence of low birthweight in India is exceeded in Haiti and Bangladesh, but there are no available data on birthweights for Bangladesh or for Haiti for nationally representative samples of mothers.

FIGURE 9.—DISTRIBUTION OF BIRTHWEIGHTS (OZ.) BY COUNTRY AND BIRTH TYPE



There are two methods of estimating the contribution of the inequality in birthweight to world earnings inequality. The first is to estimate a regression of log PPP per-worker GDP on mean birthweight across the 112 countries. Such a regression yields an R^2 of 0.443.²² This method assumes that all of the observed correlation between mean birthweights and earnings across the world corresponds to the causal effect of birthweight on earnings. If we use instead the within-MZ estimated effect of birthweight on log earnings based on the Malaysia birthweight distribution in table 5, which eliminates any role for reverse causation or corre-

lated endowments, we find that the world inequality in birthweight could account for less than 1% of world earnings inequality, measured by the log variance in GDP per worker $[(0.00413)^2 \times 32.6/1.16]$.

D. Does Increasing Birthweight Affect the Birthweight of the Next Generation of Children?

The literature also suggests that improving birth outcomes will produce healthier children in the next generation. We can assess this proposition using the information on the birthweights of the biological children of the twins. The reduced-form correlation between maternal and child birthweight reflects the heritability of endowments, the

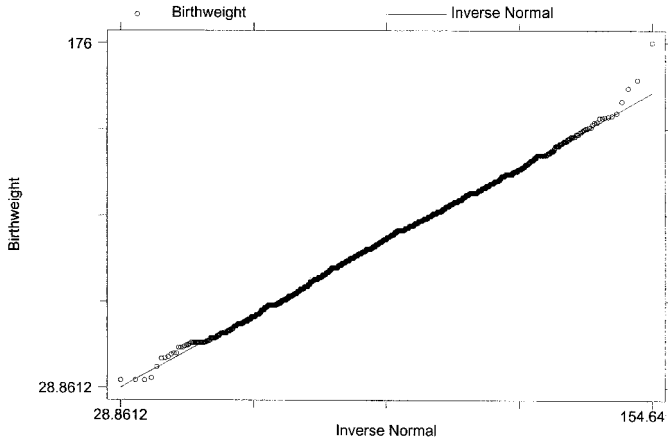
²² The regression estimate is $\log \text{ real PPP GDP/worker} = -6.69(t = 4.06) + 0.136(t = 9.45) \times \text{birthweight}$.

TABLE 5.—WITH-MZ ESTIMATES OF BIRTHWEIGHT EFFECTS ON THE LN WAGE, BY COUNTRY-SPECIFIC SAMPLE WEIGHTS

	No Weights		Malaysia Weights			India Weights		
	Full Sample	L.B.	Full Sample	L.B.	<50%	Full Sample	L.B.	<50%
Birthweight	0.00478 (2.41) ^a	0.00643 (2.35)	0.00413 (2.04)	0.00612 (2.23)	0.00502 (2.33)	0.00430 (2.24)	0.00736 (2.45)	0.00452 (2.10)
N	812	404	812	404	632	812	404	616

^a Absolute value of t-ratio in parentheses.

FIGURE 10.—Q-Q PLOT OF THE BIRTHWEIGHT DISTRIBUTION AGAINST THE NORMAL: ALL TWINS



intergenerational transmission of nutritional habits and preferences, and the fact that the mother’s birthweight may affect her choices of inputs that affect her child’s birthweight. The last is the policy-relevant effect. For example, we have seen that higher birthweight is associated with higher schooling levels. If more-educated mothers make more investments in fetal development (in part because they marry more-educated husbands), then this will be reflected in the intergenerational birthweight correlation. By estimating the intergenerational effect of birthweight across mothers who are identical twins, both the mother’s genetic contribution and any common preferences for child investments “inherited” from the twin’s parents are eliminated. What remains in the maternal birthweight difference is only that part of birthweight that reflects different inputs, which are neither heritable nor reflective of maternal preferences. Thus, the differences between the mothers’ own fetal growth rates within a twin pair can only affect the difference in their children’s birth outcomes by either influencing input choices made by the mothers while pregnant or influencing the heritable endowment of the fathers via the marriage market (for example, through the schooling effect). To the extent that assortative mating by endowments is strong, this latter effect will be negligible.

The first two columns of table 6 report the OLS and within-MZ estimates of the effects of the mother’s fetal growth on her first own biological child’s birthweight for the unweighted sample of MZ twin mothers. The last two columns report the estimates using the U.S. singleton frequency weights. Both OLS estimates indicate that the mother’s fetal growth and her child’s birthweight are significantly and positively associated. The point estimates suggest that a 1 lb. increase in the mother’s birthweight is associated with a 2.3 oz. (unweighted) to 2.8 oz. (weighted) increase in her child’s birthweight. Differencing out the common input components of birthweight for the two twin mothers (reflecting both genes and the preferences of their

family for prenatal investments), however, reduces the relationship between the mother’s birthweight and her child’s birthweight by 75%, to insignificance, in the unweighted sample and changes the coefficient sign to negative in the weighted sample. The within-MZ estimates of the intergenerational birthweight effect thus suggest the lack of importance of birthweight inputs correlated with differences in the twin mother’s weight at birth. Evidently, improving fetal growth in today’s generation will not significantly impact the birth outcomes of the next generation.

VI. Conclusion

In this paper we have used data on female twins to estimate the relationships between fetal growth and schooling, adult physical characteristics, and wages. We demonstrated that differences in birthweights between identical twins can be used to identify the life cycle consequences for physical attributes, schooling, and earnings from increasing fetal nutrient intake and show that these intrauterine nutritional effects are confounded in cross-sectional estimates with variations in genetic and in family background endowments. Our empirical results suggest that there are real payoffs to increasing body weight at birth. Increasing birthweight increases adult schooling attainment and adult height for babies at most levels of birthweight, but has no effect on adult body mass. The effect of increasing birthweight on schooling, moreover, is underestimated by 50% if there is no control for genetic and family background endowments as in cross-sectional estimates, and the use of standard family background variables—parental schooling and father’s earnings—to reduce endowment heterogeneity does not reduce these biases significantly.

We also find evidence that augmenting birthweight among lower-birthweight babies, but not among higher birthweight babies, has significant labor market payoffs. For example, our estimates indicate that increasing the average birthweight of U.S. babies in the bottom half of the birthweight distribution to the U.S. mean (about 17 oz.) would increase their lifetime earnings by 6%. Increasing the

TABLE 6.—OLS AND WITHIN-MZ TWIN MOTHER ESTIMATES: THE INTERGENERATIONAL EFFECT OF MOTHER’S OWN FETAL GROWTH ON HER CHILD’S BIRTHWEIGHT

Variable	Unweighted Sample		Weighted Sample	
	OLS	Within-MZ	OLS	Within-MZ
Mother’s fetal growth	7.48 (5.00) ^a	1.87 (0.51) ^b	8.84 (3.93)	-1.56 (0.40)
Current age of mother	-0.198 (1.88)	—	-0.259 (1.82)	—
Constant	109.6 (19.1)	—	109.7 (14.9)	—
N	1,207	608	1,207	608

^a Absolute value of robust *t*-statistic with clustering by twin pairs.
^b Absolute value of *t*-statistic.

average birthweight of low-birthweight babies by the same amount would increase their earnings by almost 10%, and further increasing their average birthweight to the U.S. mean would increase earnings by 26%. However, our estimates also suggest that the cross-country correlation between incomes and birthweight substantially overstates the reduction in world earnings inequality that would arise from reducing cross-country disparities in birthweight.

Our findings have some important implications for policy. The significantly positive effects of fetal growth on schooling, adult height, and earnings support programs that serve to increase weight at birth, such as Medicaid in the United States (Currie and Gruber, 1996), and the use of public resources to do so from an efficiency perspective if schooling has positive externalities, as usually is assumed. Moreover, our results suggest that targeting populations with the potential for low birthweight will be particularly efficacious in enhancing the earnings prospects of the next generation. Our results indicate, moreover, that such efforts will not increase the incidence of obesity among adults. However, our findings do not indicate any support for policies increasing fetal growth on the grounds of intergenerational nutritional links between mothers and their children.

Our results also have implications for countries in which couples increasingly rely on medical procedures that augment fertility in part as a consequence of fertility postponement. Our findings suggest that such procedures, to the extent that they increase multiple births, impose nontrivial costs on children's subsequent development. On average, twins have 28 oz. less birthweight than singletons, which our estimates indicate translates into a reduction in lifetime earnings for such children of over 12%.

Finally, while these policy implications are important for countries such as the United States, they are potentially more important for many developing countries, which have substantially lower distributions of birthweights. Our estimates suggest that the gains to per-worker GDP from reducing the developed–developing-country birthweight gap can be as high as 9% in low-birthweight countries. However, shifting the distribution of birthweights in developing countries to that in the United States would reduce world earnings inequality by less than 1%, far less than indicated by the cross-country correlation between per-worker GDP and birthweight.

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APPENDIX

TABLE A1.—WITHIN-MZ TWIN ESTIMATES OF LOW BIRTHWEIGHT AND FETAL GROWTH ON OWN SCHOOLING ATTAINMENT, BMI, HEIGHT, AND ln WAGE, WITHOUT AND WITH FETAL GROWTH INCLUDED IN SPECIFICATION

	Schooling		BMI ($\times 10^{-3}$)		Height		ln Wage	
	Without	With	Without	With	Without	With	Without	With
Low birthweight	-0.270 (1.80)	0.0326 (0.17)	0.0279 (0.05)	0.258 (0.38)	-0.608 (5.11)	-0.0106 (0.07)	-0.0384 (0.70)	0.0744 (1.07)
Fetal growth	—	0.686 (2.53)	—	0.522 (0.54)	—	1.40 (6.83)	—	0.256 (2.60)
N	804	804	804	804	804	804	804	804

^aAbsolute value of robust *t*-statistic with clustering by individuals.

^bAbsolute value of robust *t*-statistic.