Effects of Participation in the WIC Program on Birthweight: Evidence From the National Longitudinal Survey of Youth

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Objectives. This study sought to estimate the impact on birthweight of maternal participation in the Special Supplemental Nutrition Program for Women, Infants, and Children (WIC).

Methods. WIC estimates were based on sibling models incorporating data on children born between 1990 and 1996 to women taking part in the National Longitudinal Survey of Youth.

Results. Fixed-effects estimates indicated that prenatal WIC participation was associated with a 0.075 unit difference (95% confidence interval [CI]=−0.007, 0.157) in siblings' logged birthweight. At the 88-oz (2464-g) low-birthweight cutoff, this difference translated into an estimated impact of 6.6 oz (184.8 g).

Conclusion. Earlier WIC impact estimates may have been biased by unmeasured characteristics affecting both program participation and birth outcomes. Our approach controlled for such biases and revealed a significant positive association between WIC participation and birthweight. (*Am J Public Health.* 2002;92:799–804)

Established in 1972, the Special Supplemental Nutrition Program for Women, Infants, and Children (WIC) attempts to increase the nutrition levels and general well-being of children.[1,2] Many of the evaluations of this program were conducted at least 10 years ago, and the majority have relied on data from only a few states.[3–8] Previous studies of the effects of WIC participation on children's well-being involved a risk of bias resulting from unmeasured selection factors. Not all eligible women claim benefits, and if selection into WIC involves the neediest women, studies comparing child outcomes among WIC participants and nonparticipants that fail to adjust for differences in "need" may underestimate the effects of the program. In contrast, if WIC participants are more highly motivated than nonparticipants, then studies that fail to adjust for motivation may overestimate the program's effects. Because the WIC program is locally administered, factors governing selection into the program are also likely to vary across time and place.

In the present study, we addressed these concerns by estimating WIC effects with a national sample of children and using sibling fixed-effects models to account for unmeasured heterogeneity among the mothers of sample children. Specifically, we studied a sample of children born between 1990 and 1996 to mothers participating in the National Longitudinal Survey of Youth (NLSY), and we used merged NLSY mother–child data files to estimate the effects of WIC participation on birthweight. We eliminated the biasing effects of persistent characteristics of mothers—both measurable and unmeasurable—and thus the estimates from this study constitute a methodological advance over previous studies. A disadvantage of the sample is that all of the children were born to somewhat older mothers. In addition, all of the data were reported by the mothers, and standard errors for the fixed-effects models that we estimated were considerably larger than standard errors for more conventional regression models.

THE WIC PROGRAM

WIC provides federal grants to states for supplemental foods, health care referrals, and nutrition education targeted toward low-income, pregnant, and postpartum women; infants; and children up to the age of 5 years who are at nutritional risk. In addition, participants must have incomes at or below 185% of the poverty level (in 2000, the program cutoff was $25 667 for a family of 3).

Much of the research on the effects of WIC participation on children has focused on the potential benefits of increased use of prenatal care, increased Medicaid savings, and better infant outcomes.[6,9–12] In particular, studies indicate that the WIC program is beneficial in the promotion of nutrition supplementation during pregnancy, which has been linked to more positive birth outcomes.[12–15] The WIC program has also drawn some recent criticism. Besharov and Germanis[16] argued that the benefits of the WIC program may have been overstated because many of the earlier studies suffered from issues of selection bias. In the present study, we responded to this criticism by using methodological techniques that accounted for bias arising from persistent unmeasured family characteristics that could affect both program participation and child outcomes.

METHODS

Analytic Approach

We compared sibling differences in birthweight with differences in maternal participation in WIC. A general model for the relationship between WIC participation and early child outcomes can be expressed as follows:

$$Y_i = \alpha + \beta_1 P_i + \beta_2 X_i + \varepsilon_i \; , \qquad (1)$$

where Y_i is the outcome of interest for child *i*, P_i is an indicator equal to 1 if the mother participated in WIC before the child's birth and to 0 otherwise; X_i represents a vector of persistent and time-varying maternal and family characteristics hypothesized to affect child outcome; and ε_i is an error term that includes unobserved variables and random error.

In constructing our ordinary least squares (OLS) model estimates for equation 1, we used Huber–White robust standard errors that accounted for the lack of independence of observations based on siblings born to the same mother.[17] These OLS estimations would yield biased estimates of β_1 if unobservable determinants of Y were also correlated with P.

The sibling fixed-effects estimator removes bias from time-invariant maternal and family components of ε by subtracting the average for all siblings in a given family from each child's value:

$$(2)\quad (Y_{ij}-Y^*_j)=\beta_1(P_{ij}-P^*_j)+\beta_2(X_{ij}-X^*_j)+(\varepsilon_{ij}-\varepsilon^*_j).$$

Here, subscript i references children, subscript j references families, and asterisks indicate family average values. The fixed-effects models, by subtracting the sibling mean values, remove the effects of persistent family and maternal conditions (both observed and unobserved) on β_1.

Fixed-effects models do not eliminate bias from time-varying covariates that affect both WIC participation and birth outcomes. If mothers were systematically more likely to use WIC during a second pregnancy because they had learned of the program during a difficult first pregnancy, then the results from sibling models would probably be biased by the fact that mothers were undergoing a learning process that would disproportionately favor younger siblings. However, it is likely that a considerable amount of variability in WIC participation across births is exogenous to the family and may instead be produced by state variations in program implementation.[18]

Sample

Our data were drawn from the 1996 and earlier survey waves of the NLSY, a nationally representative sample of men and women. When initially interviewed in 1979, the women in our sample were aged 14 to 21 years. The study oversampled Black, Hispanic, and economically disadvantaged White youths.[19] Beginning in 1990, the NLSY obtained maternal reports on whether WIC benefits had been received in the preceding calendar year. Our sample consisted of the 1984 children born to NLSY mothers between 1990 and 1996 for whom prenatal WIC participation status had been recorded. It is important to note that the women in our sample were between the ages of 25 and 38 years when they gave birth; thus, the sample consisted of relatively older mothers.

For our sibling-based analyses, we identified 969 children born between 1990 and 1996 who had 1 or more siblings also born between 1990 and 1996. Most of the sibling groups in this sample included just 2 children. In the majority of these groups (349 of 453), mothers did not participate in the WIC program during their pregnancies. Thirty-three of the sibling groups represented situations in which mothers' WIC participation included all of the children who were a part of this sample.

The 71 discordant-sibling groups in which siblings differed in terms of their mother's participation in the WIC program before their birth were the key source of variance in estimations of our sibling fixed-effects regression models. The number of discordant-sibling groups was consistent with previous studies on the WIC and Head Start programs[20,21] involving similar modeling techniques[22] and, as shown subsequently, provided acceptably precise estimates of β_1. The majority of the discordant-sibling groups (49, or 74%) followed the pattern of no participation in the WIC program in the first observed pregnancy and participation before a subsequent birth. However, a sizable percentage (26%) of the discordant-sibling groups followed the opposite pattern of WIC program usage.

Measurement

Prenatal WIC participation was measured with a dichotomous variable based on maternal reports of receipt of WIC benefits in the calendar year preceding the birth. Recent research indicates that a potentially important aspect of WIC participation is timing—early vs late—in terms of the mother's pregnancy.[1] The NLSY provides only annual data on receipt of WIC benefits, an analytic cost that had to be weighed against the advantages of the size, national scope, and sibling representation of our sample. Table 1 presents weighted descriptive information about the children in the total, sibling, and discordant-sibling samples. Approximately 12% of the children in both the total sample and the sibling sample had mothers who received WIC benefits during their pregnancy. Based as they are on WIC use by at least 1 sibling, mean rates of WIC use in the discordant-sibling sample were larger.

The outcome on which we focused was maternal report of child birthweight. Lower-birthweight children are at higher risk of mortality and lifelong disabilities.[23] Birthweight data from the NLSY conform quite closely to vital statistics data.[24] Moreover, research indicates that maternal reports of birthweight are accurate.[25–28] In both the total and sibling samples, the average birthweight was approximately 120 oz (3360 g). Ideally, birthweight should be adjusted for gestational age. However, gestational age data as reported by mothers in the NLSY deviate significantly from national standards derived from vital statistics data. In our empirical work, we estimated models both with and without adjustment for gestational age and found that our key results were not sensitive to this difference. Given the questionable quality of these data in the NLSY, we did not include gestational age in the analyses presented in Table 2.

One of the strengths of the NLSY data in comparison with other data sets is the rich longitudinal information that can be used to control for a relatively comprehensive set of both stable and time-varying individual and family characteristics. Our OLS models controlled for the following persistent maternal characteristics: cognitive skills (as measured by the Armed Forces Qualifying Test), self-esteem, and early deviance (e.g., stealing, fighting). We also controlled for a wide variety of intergenerational variables measured when the mother was an adolescent (e.g., number of siblings and support for literacy in the home).

Time-varying characteristics included in both OLS and fixed-effects models were mothers' reports of smoking and reports of drinking during the given pregnancy. We also accounted for maternal prenatal food stamp participation, family income (not including food stamps) in the calendar year before the birth, maternal residence in an urban area in the year before the birth, maternal education as of the year before the birth, child sex, and ethnic identity. Past research has shown significant differences in birthweight according to minority status.[23]

Finally, we controlled for birth order of the child, because first-born children may be less likely to receive WIC benefits. In our multivariate analyses, we initially controlled only

TABLE 1—Weighted Descriptive Characteristics for Samples of National Longitudinal Survey of Youth (NLSY) Children Born Between 1990 and 1996

	Total Sample (n = 1984)		Sibling Sample (n = 969)		Discordant Sibling Sample[a]			
					Group 1 (n = 19)		Group 2 (n = 49)	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Prenatal maternal variables								
Prenatal WIC participation, %	12	33	12	32	68	47	55	50
Prenatal food stamps, %	10	31	09	29	57	50	34	47
Total income (in 1000s)	47.82	26.67	50.05	26.50	20.39	19.07	18.19	18.34
Drank during pregnancy, %	52	50	52	50	40	50	35	48
Smoked during pregnancy, %	24	43	19	39	46	50	29	46
Maternal characteristics and background								
Armed Forces Qualifying Test score	74.21	19.47	76.82	18.77	63.08	20.52	54.06	20.13
Self-esteem score (1980)	1.76	0.41	1.75	0.40	1.78	0.34	1.81	0.50
Deviant behavior score (1980)	1.14	1.18	1.11	1.19	1.15	1.27	1.34	1.38
Prenatal urban residence, %	82	38	85	35	80	40	80	40
Prenatal maternal education, y	13.75	2.45	14.04	2.28	11.82	1.67	11.82	1.72
Parents' educational level, y	12.06	2.91	12.46	2.87	10.49	3.03	9.95	2.68
Worked at age of 14 years, %	51	50	47	50	21	41	40	49
Lack of literary supports, %	05	22	04	20	04	20	10	33
No. of siblings	3.33	2.20	3.26	2.22	4.04	2.80	3.95	2.44
Child demographics								
Black, %	11	31	09	29	37	49	31	47
Hispanic, %	06	24	06	24	13	34	18	38
Male, %	50	50	51	50	65	48	48	50
Gestational length, wk	38.38	1.97	38.40	1.83	38.6	1.72	38.14	1.89
Outcome variable: birthweight, oz	119.71	21.01	120.83	20.21	115.84	19.51	116.18	20.66

Note. The total sample consists of all NLSY children who were born between 1990 and 1996 and had valid information on prenatal WIC participation. The sibling sample consists of NLSY children who were born between 1990 and 1996 and had at least one sibling who was born in the same time period.
[a]Of the discordant sibling groups, 19 followed the pattern of WIC receipt in the first pregnancy but not in subsequent pregnancies (Group 1), and 49 followed the pattern of nonreceipt of WIC in the first pregnancy but not in subsequent pregnancies (Group 2).

for child demographic characteristics. Subsequently, we included a full set of prenatal variables and maternal characteristics. We did not include time-invariant prenatal and maternal characteristics in the sibling fixed-effects regression analysis, because these characteristics did not differ between siblings.

RESULTS

Table 2 presents coefficients and standard errors from OLS and fixed-effects models of the natural log of birthweight. Experimentation with raw birthweight and the square root of birthweight as alternative forms of our dependent variable produced coefficients with similar significance levels and impact estimates. The natural log transformation yielded the most consistently significant estimates of the relationship between our independent variables and birthweight. Missing data on independent variables other than WIC participation produced substantial sample size variability across our models.

Because missing dummy variables create problems in fixed-effects regression models, we adopted the strategy of substituting previous-wave values whenever possible (e.g., for family income); when this strategy was not feasible, we used mean substitution. Experimentation indicated that our results were robust across various missing data treatments. Specifically, we estimated parallel sets of coefficients in which we (1) deleted all cases involving missing data in the final set of explanatory variables, (2) used mean substitution of the missing data, and (3) set missing values to zero and included dummy variables for missing data. Table 2 notes coefficients that differed from zero at the .01, .05, and .10 levels of statistical significance; our inclusion of the .10 level was based on the relatively small sample sizes driving our fixed-effects estimates.

Estimates based on simple OLS models (Table 2) implied that participation in WIC was associated with a statistically insignificant 0.025 increase in logged birthweight. Controls for a more extensive set of maternal and child characteristics more than doubled the OLS-based estimated impact, to a statistically significant 0.055. This increase resulting from controls on observable characteristics suggested a negative selection process in

TABLE 2—Birthweight (Logged): Results of Ordinary Least Squares (OLS) and Fixed-Effects Regressions for Children in National Longitudinal Survey of Youth Born Between 1990 and 1996

	Analysis 1[a]				Analysis 2[b]			
	OLS[c]		Fixed Effects		OLS[c]		Fixed Effects	
	Coefficient	SE	Coefficient	SE	Coefficient	SE	Coefficient	SE
Prenatal WIC participation	0.025	0.016	0.034	0.034	0.055	0.020 ***	0.075	0.042*
Prenatal maternal variables								
Prenatal food stamps					−0.051	0.024**	−0.021	0.059
Total cash income					−0.002	0.012	−0.011	0.022
Drank during pregnancy					−0.030	0.013**	0.024	0.034
Smoked during pregnancy					−0.038	0.015***	−0.035	0.058
Maternal characteristics								
Armed Forces Qualifying Test score					0.001	0.001**	. . .[d]	
Self-esteem score (1980)					−0.008	0.016	. . .[d]	
Deviant behavior score (1980)					0.002	0.005	. . .[d]	
Prenatal urban residence					0.024	0.018*	. . .[d]	
Prenatal maternal education					−0.003	0.003	. . .[d]	
Maternal early background								
Parents' mean education					−0.002	0.002	. . .[d]	
Mother worked at age of 14 years					−0.021	0.013*	. . .[d]	
Lack of literary supports					0.008	0.026	. . .[d]	
No. of siblings of mother					0.002	0.003	. . .[d]	
Child demographics								
Black	−0.104	0.017***	. . .[d]		−0.092	0.018***	. . .[d]	
Hispanic	−0.028	0.016*	. . .[d]		−0.028	0.020	. . .[d]	
Male	0.024	0.011**	0.023	0.020	0.020	0.012*	0.012	0.022
First born	−0.032	0.012***	−0.037	0.023	−0.038	0.013***	−0.027	0.024
Adjusted R^2	0.04		0.02		0.08		0.03	
No. of children	1499		1229		1355		1124	

[a]Includes only prenatal WIC participation and child demographics.
[b]Includes full set of variables.
[c]OLS models were estimated with Huber–White robust standard errors accounting for the fact that observations were clustered within family units.
[d]Invariant.
$^*P<.10$; $^{**}P<.05$; $^{***}P<.01$.

which the mothers most likely to deliver low-birthweight infants were more likely to participate in WIC.

The more complete OLS model also suggested that prenatal maternal smoking and drinking were associated with significant decrements in birthweight. In addition, higher maternal cognitive skills were significantly linked to higher birthweights. We found a significant negative association between prenatal receipt of food stamps and birthweight; this result appeared to be tied to the fact that mean birthweights were considerably lower among women who received food stamps but not WIC benefits than among those who demonstrated the opposite pattern (111.83 vs 125.35 oz [3131.24 vs 3509.8 g]). Consistent with other research,[28] we observed both race and sex differences in birthweight (Table 2). Finally, firstborn children were more likely to be lower in birthweight than higher-birth-order children.

The sibling fixed-effects models presented in Table 2 also suggested a positive impact of prenatal WIC participation on birthweight. Note that the sample sizes for the fixed-effects models were based on all children but that the key WIC impact estimates were driven by the within-family WIC and birthweight differences for the 71 sibling pairs discordant in terms of maternal WIC usage.

In the case of the model that controlled only for child sex and birth order (Table 2), the estimated impact was a statistically insignificant 0.034. After the full set of prenatal maternal variables had been taken into account, prenatal WIC participation was associated with a 0.075-unit difference in siblings' logged birthweight. Note the large standard error associated with this coefficient, however. The 0.075 point estimate implied a WIC impact of 6.6 oz (184.8 g) at the 88-oz (2464-g) low-birthweight cutoff, an effect considerably larger than those revealed in previous studies. The size of the standard error argues against attaching considerable weight to this particular point estimate, but such an estimate supports the conclusion that the impact is indeed positive.

We also estimated models in which we constrained the samples in analysis 1 of Table 2 to be identical with the samples in analysis 2. The WIC coefficient did not change in the OLS model (0.023, SE=0.017), and it increased slightly in the fixed-effects model (0.052, SE=0.037). Given that the change in the WIC coefficient was slight, it appears that the change in coefficients in the models presented in Table 2 was more a function of the addition of covariates than of a difference in sample composition.

The major source of bias in our fixed-effects estimates derived from maternal conditions that influenced both pregnancy-specific WIC participation and child outcomes. To address these concerns, we considered a variety of maternal health measures in the year before the birth of the child, including the average pregnancy weight of the mother and whether she reduced calories or sodium during pregnancy, sought prenatal care, or had health conditions limiting work. These supplementary results indicated that none of the primary relationships of interest were altered substantially by inclusion of these additional variables.

We also estimated models that controlled for the effects of gestational length, testing the hypothesis that WIC may be helpful in increasing gestational length and in turn increasing birthweight. Babies born preterm had significantly lower birthweights than babies born full term, but inclusion of this control did not alter the main relationship between prenatal WIC participation and birthweight. We also assessed the interactive effects of WIC participation and birth order so as to capture the potential effects of "referral bias" (i.e., mothers may be more likely to seek WIC benefits in their second pregnancy than in their first). We did not find evidence of a significant interaction.

DISCUSSION

Previous studies on the effects of WIC participation have been based on models that may be biased by unmeasured characteristics affecting both program participation and birth and child outcomes. Our approach was to use sibling fixed-effects modeling techniques to remove potentially biasing effects of unobserved family-specific characteristics. Using this approach, we found that the positive effects of prenatal WIC participation on infant birthweight persist when the potential for bias is minimized.

There are several limitations to our approach. First, our sample consisted of children born to relatively older mothers. As a result, our findings may not generalize to children born to mothers in other age groups. Second, our sibling sample was fairly small, although it was sufficient in size to detect a 7% impact on birthweight with 90% power. Third, our sibling-based analyses may not generalize to children with no siblings. These limitations suggest the need for further research on the connection between WIC participation and birthweight. Given our finding that the impact of WIC appeared to increase with our fixed-effects controls for time-invariant family characteristics, we suggest that further research attend to the omitted-variable–bias problem.

Our results suggest a pattern in which more complete models accounting for background and prenatal factors produced the largest estimates of effects of prenatal WIC participation on birthweight. This pattern of findings implies a negative selection into WIC in which women with the greatest needs are most likely to participate. The bias-reducing methodology used here should be considered in future program evaluations. ■

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This article was accepted October 10, 2001.

Contributors

Both L. Kowaleski-Jones and G.J. Duncan analyzed the data, wrote the preliminary draft, and responded to reviewers' comments.

Acknowledgments

We gratefully acknowledge the financial support provided by the US Department of Agriculture/Institute for Research on Poverty Food Assistance Small Grants Program.

Helpful comments were provided by Andrea Beller, Douglas Besharov, Peter Germanis, Christine Ross, Sonia Salari, Ken Smith, Nick Wolfinger, and Cathleen Zick.

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