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Can targeted transfers improve birth outcomes? Evidence from the introduction of the WIC program

Hilary Hoynes *, Marianne Page, Ann Huff Stevens

University of California, Davis and NBER, United States

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ABSTRACT

The goal of federal food and nutrition programs in the United States is to improve the nutritional well-being and health of low income families. A large body of literature evaluates the extent to which the Supplemental Program for Women Infants and Children (WIC) has accomplished this goal, but most studies have been based on research designs that compare program participants to non-participants. If selection into these programs is non-random then such comparisons will lead to biased estimates of the program's true effects. In this study we use the rollout of the WIC program across counties to estimate the impact of the program on infant health. We find that the implementation of WIC led to an increase in average birth weight and a decrease in the fraction of births that are classified as low birth weight. We find no evidence that these estimates are driven by changes in fertility or selection into live births. Our preferred estimates suggest that WIC initiation raised average birth weight by 2 g, or by 7 g among infants born to mothers with low education levels. These translate into estimated birth weight increases among participating mothers of approximately 18 to 29 g. Estimated treatments on the treated impacts among infants born to participating mothers with low education are of similar magnitude.

1. Introduction

The goal of the Supplemental Nutrition Program for Women, Infants, and Children (WIC) is to improve the nutritional well-being of low income pregnant and postpartum women, infants, and children under the age of five. Many studies show that pregnant women who participate in WIC give birth to healthier infants than those who do not, and this has contributed to broad support for the program: since its inception in the mid 1970s, the number of WIC participants has grown to 8.7 million, at an annual cost of 6.2 billion dollars.[1]

Recently, however, the validity of existing studies – most of which use comparisons between participants and non-participants to estimate WIC's effects – has come under question. Several researchers (Besharov and Germanis, 2001; Bitler and Currie, 2005; Brien and Swann, 2001; Chatterji et al., 2002; and Kowaleski-Jones and Duncan, 2002) have drawn attention to the fact that selection into the WIC program is non-random. If pregnant women who participate in WIC are healthier, more motivated, or have better access to health care than other eligible women, then comparisons between the children of participants and non-participants could produce positive program estimates even if the true effect is zero. Conversely, if WIC participants are more disadvantaged than other mothers, such comparisons may understate the program's impact.

Recent studies have used several different approaches to address this problem. Bitler and Currie (2005),Joyce et al. (2005, 2008) and Figlio et al. (2009) compare outcomes among more narrowly defined treatment and control groups; Brien and Swann (2001), Chatterji et al. (2002) and Kowaleski-Jones and Duncan (2002) include maternal fixed effects in their regression analyses; and Brien and Swann (2001) and Chatterji et al. (2002) utilize limited state variation in WIC program parameters. These approaches yield smaller WIC estimates, yet they, too, suffer from identification problems. For example, even estimates based on comparisons of observationally similar participants and non-participants may suffer from omitted variables bias. Likewise, within family estimates may be driven by changes in family circumstances between births. And while evaluations of other programs aimed at helping disadvantaged families – including AFDC/TANF and Medicaid – commonly leverage significant variation in eligibility and benefit rules across states, the parameters of the WIC program, like other food and nutrition programs in the United States, exhibit little geographic variation.[2] Ultimately, as noted in a recent review of WIC studies by Ludwig and Miller (2005), WIC analyses are challenged by the absence of a "clearly exogenous source of identifying variation (that is, a randomized or natural experiment that drives variation across low-income women in WIC enrollment)."[3]

This study addresses this problem by exploiting variation in WIC program introduction across geographic areas and over time. WIC was first established as a pilot program in 1972, and WIC sites were

* Corresponding author. Tel.: +1 530 564 0505; fax: +1 530 752 9382.
E-mail addresses: hwhoynes@ucdavis.edu (H. Hoynes), mepage@ucdavis.edu (M. Page), annstevens@ucdavis.edu (A.H. Stevens).

[1] For a recent summary of the extensive WIC literature see Currie (2003). WIC program information available on USDA website, see http://www.fns.usda.gov/pd/wisummary.htm.

[2] Another difference between WIC and AFDC/Food Stamps is that WIC is administered by local non-profits.

[3] Ludwig and Miller (2005).

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established in different counties and in different years between 1972 and 1979. This feature of the program's introduction allows us to perform a difference-in-differences analysis, in which we compare changes in infant health within a set of "treatment" counties (those adopting WIC in a given year) to changes within a set of "control" counties (those who have not yet adopted WIC).

We find that when WIC is made available by the third trimester, average birth weight in the county increases by approximately 2 g. This estimated effect is driven by women with low levels of education and women living in high poverty counties — precisely the women who are most likely to be eligible for program benefits. Among women with low levels of education, WIC increases average birth weight by 7 g and reduces the fraction of births that are classified as low birth weight by 1.4%. Using estimates of WIC participation rates, these results for low educated women suggest a 10 (11) percent increase (decrease) in average birth weight (fraction low birth weight) of children born to WIC participants. Since we find no evidence that WIC affects fertility or characteristics of mothers giving birth, our estimates are unlikely to be generated by indirect effects on selection into birth.

In the next section we provide a brief description and history of the WIC program and in Section 3 we review the prior WIC literature. In Section 4 we present our research design and in Section 5 we describe our data. We present our results in Section 6 and we conclude in Section 7.

2. Background

The goal of the WIC program is to increase the nutritional well-being among low-income pregnant/post-partum women, infants and young children by providing food packages and nutritional counseling. Five types of individuals are eligible for WIC: pregnant women, post-partum women with a child under six months, breastfeeding women with a child under 12 months, infants, and children under age five. Participants must live in households with family incomes below 185% of the poverty line or become eligible through participation in another welfare program such as Medicaid, Temporary Assistance to Needy Families, or Food Stamps. They must also be certified to be at nutritional risk, but virtually all financially eligible persons appear to satisfy this requirement (Ver Ploeg and Betson, 2003). Food packages are typically provided in the form of vouchers that can be used to purchase specific items from participating supermarkets. WIC maintains a list of approved foods, which must contain protein, calcium, iron, and Vitamins A and C.[4] Post-partum women have access to free infant formula and (in later years of the program) breastfeeding services.

The WIC program was first established as a pilot program in 1972, and became permanent in 1975. The program was developed in direct response to policy recommendations highlighting health deficits among low-income individuals that might be reduced by improving their access to food. It was further recognized that, by providing food at "critical times" to pregnant and lactating women and young children, it might be possible to prevent a variety of health problems (Oliveira et al., 2002).

WIC sites were established in different counties between 1972 and 1979, with legislation requiring that the program be implemented first in "areas most in need of special supplemental food" (Oliveira et al., 2002). The first WIC program office was established in January 1974 in Kentucky, and had expanded to include counties in 45 states by the end of that year. WIC was intended to supplement food stamp benefits and the authorizing legislation specifically did not preclude a person from WIC participation if they were already receiving food stamps.[5]

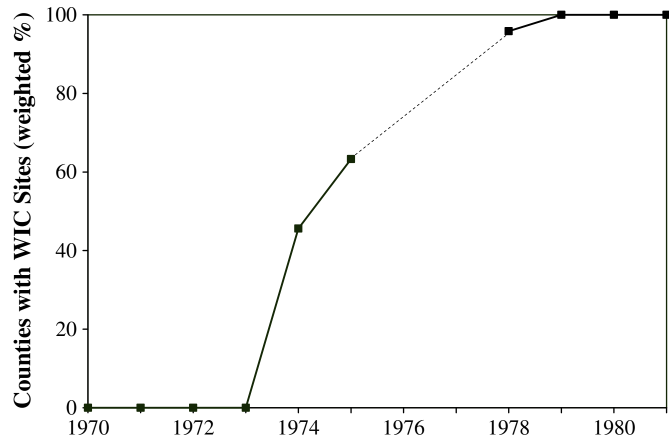

Fig. 1. Weighted percent of counties with WIC programs, 1970–1981. Notes: As noted in the text, we do not have any data on WIC coverage for 1976 and 1977. In 1978, we observe data for some but not all states. The 1978 value is calculated for this subsample of counties. See text and Appendix A for details.

We have obtained data on the year the first WIC programs were implemented in a county. These data were assembled from five documents listing all active WIC programs as of 1974, 1975, 1978, 1979, and 1989 and are more fully described in Section 5 and Appendix A. Fig. 1 documents the aggregate rollout of WIC by presenting the fraction of counties (weighted by 1970 population) that had WIC programs in place in each year. The figure clearly shows that there was a dramatic increase in exposure to the program between 1974 and 1979.[6]

Fig. 2 presents a map of the U.S. counties in 1974, 1975, 1978 and 1979. In each panel, black counties identify those counties that had a WIC program in place. White counties did not have a WIC program. Gray counties are counties for which we have no information (the reason for this missing data is explained below). As is clear from these figures, there is a considerable amount of geographic variation in the timing of WIC implementation, both within and across regions. Our identification strategy hinges on this county level variation in WIC "treatment."

3. Existing literature

There are many studies that examine the impact of WIC on birth outcomes, breastfeeding, and nutritional intake (see Currie, 2003 for a review of the literature), and most find that women who participate in WIC give birth to healthier infants than non-participants (Currie, 2003; Devaney et al., 1990). At issue, however, is the extent to which such studies provide information about the program's true causal effect. If WIC recipients differ from non-recipients in other ways, then some of the differences in children's outcomes may reflect differences in the mothers' characteristics. For example, if pregnant women who participate in WIC are healthier, more motivated, or have better access to health care than other eligible women, comparisons between the children of participants and non-participants could produce positive program estimates even if the true effect is zero (Besharov and Germanis, 2001; Brien and Swann, 2001; Chatterji et al., 2002; and Kowaleski-Jones and Duncan, 2002). Conversely, if WIC participants are more disadvantaged than other mothers, such comparisons may understate the program's impact. A detailed examination of participating vs. non-participating mothers' characteristics suggests that WIC mothers are negatively selected from the pool of eligibles (Bitler and Currie, 2005).

[4] WIC approved foods include juice, fortified cereal, eggs, cheese, milk, dried beans, tuna, carrots, and iron-fortified infant formula.

[5] Participation in the commodity distribution program, however, disqualified individuals from WIC participation (Oliveira et al., 2002). But the CDP was being phased out during the 1970s as the FSP expanded to a national program.

[6] Note that 1976 and 1977 are omitted because we have no information for those years.

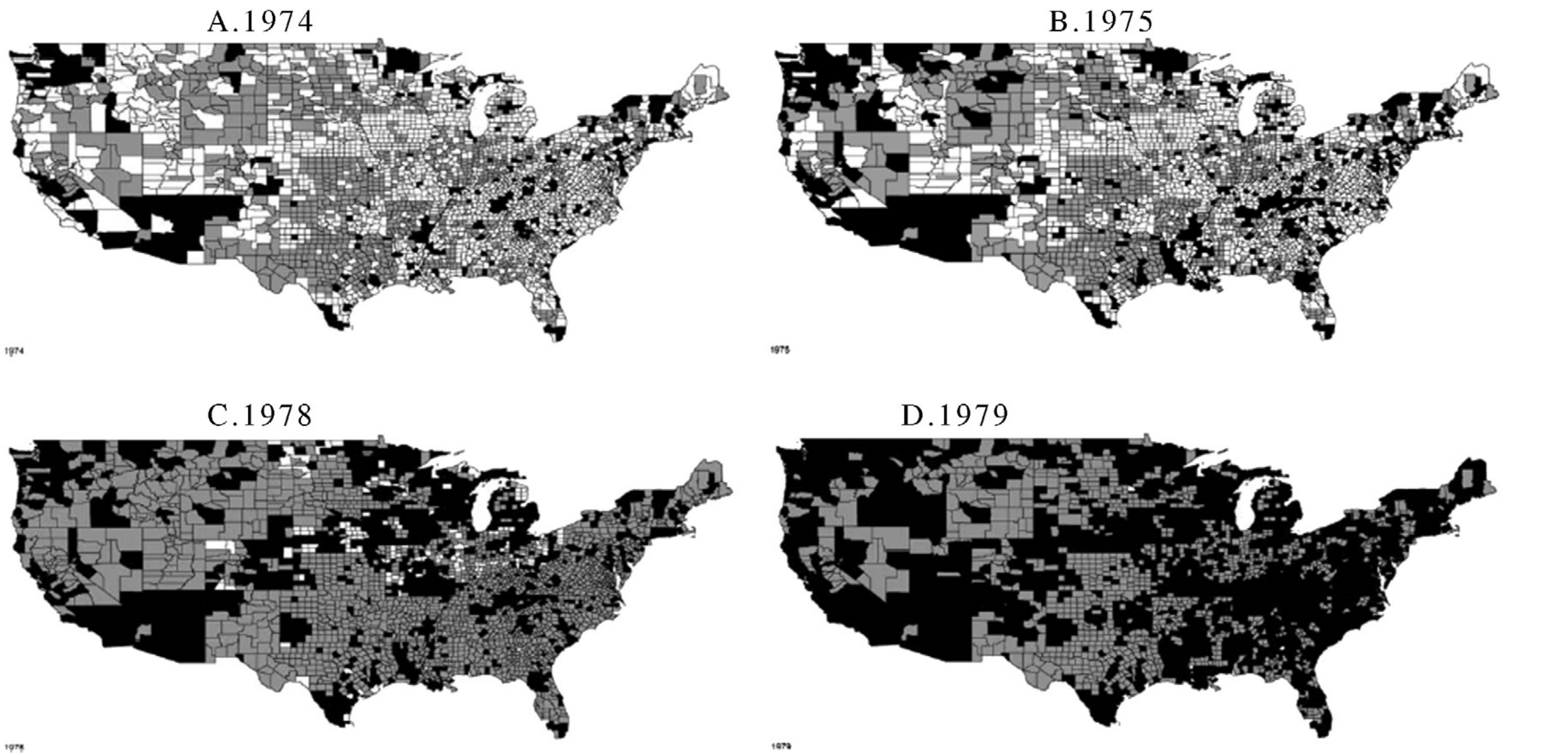

Fig. 2. WIC implementation by county in 1974, 1975, 1978, and 1979. Notes: Black denotes a county with a WIC program, white denotes a county without a WIC program in place, and grey denotes a county with missing information. Counties with missing information are counties that we do not observe with a WIC program by 1979 or, in 1978, counties in states not covered by the National WIC Evaluation. See text and Appendix A for details.

In the wake of these concerns, a new wave of WIC studies has recently emerged that employs alternative estimation strategies. One approach taken is to compare outcomes among more narrowly defined treatment and control groups (e.g., Bitler and Currie, 2005; Joyce et al., 2005, 2008, and Figlio et al., 2009). For example, Bitler and Currie (2005) create a control group based on Medicaid funded births, and employ selection correction models. Figlio et al. (2009) also create more narrowly defined treatment and control groups by focusing on those whose income puts them on just either side of the eligibility cutoff. These approaches make some progress on the problem, but because differences between participants and non-participants remain, they do not fully eliminate selection concerns.

Another set of studies controls for unobserved family background characteristics by comparing outcomes among siblings who participated in WIC to outcomes among those who did not (Brien and Swann, 2001; Chatterji et al., 2002; and Kowaleski-Jones and Duncan, 2002). There are several drawbacks to this approach: first, it is well known that within-family comparisons are likely to exacerbate measurement-error problems that bias estimates towards zero (Griliches, 1979). Second, there may be spill-over effects from the participating sibling to the non-participating sibling, which will lead to underestimates of the program's true effect. Finally, between-birth changes in economic or health conditions of other family members may be correlated with between sibling differences in program participation. In such cases, selection biases will not be eliminated.

In the larger program evaluation literature, these selection problems can often be avoided by comparing individuals living in states with different program parameters. Many studies of the AFDC program, for example, are based on this type of identification strategy (Moffitt, 1992; Blank, 2002). Unfortunately, WIC is a federal program for which there is very little geographic variation in either eligibility criteria or benefit levels. Brien and Swann (2001) and Chatterji et al. (2002) compare infant health outcomes across states with different program rules,[7] but the variation in program rules turns out to have limited power in predicting WIC participation.

The estimation strategy employed by Rush et al. (1988) comes closest to our own. Like our study, Rush et al. use county-level variation during the early years of WIC, but instead of focusing on program implementation, the authors identify the effects of WIC using variation in WIC *penetration* — roughly the fraction of eligible women participating in each county between 1972 and 1980. Since this variable captures both the presence of a program in the county and county-level participation rates, non-random selection remains a potential concern.

In summary, the literature to date is dominated by studies that are subject to selection bias. The magnitude of WIC's effects is thus largely unknown, even though this knowledge is crucial to determining whether the program is a success. We are able to make significant inroads on this problem by comparing within county changes in infant health across treatment and control counties. The gains from our methodological improvements outweigh the disadvantages of studying an earlier period. To be sure, the United States has experienced many demographic, social, and economic changes since the inception of WIC, but our results will still be informative to today's policy debates because to date, no one has been able to separately estimate the causal effects of the program from the effects of individual characteristics that might independently affect their health.

4. Research design

Cross-county variation in WIC initiation forms the basis of our estimation strategy. Similar strategies have been used in recent studies of social programs such as the food stamp program (Almond et al., forthcoming; Currie and Moretti, 2008; and Hoynes and Schanzenbach, 2009), Head Start (Ludwig and Miller, 2007), Medicare (Finkelstein and McKnight, 2008), family planning programs (Bailey, 2009), Title I (Cascio et al., 2010), and a larger literature examining impacts of the Great Society and Civil Rights era (for example see Almond et al., 2006). Our basic regression model is:

$$y_{ct} = \alpha + \delta WIC_{ct} + \gamma_1 Z_c t + \gamma_2 GT_{ct} + \eta_c + \lambda_t + \theta_{st} + \varepsilon_{ct}, \qquad 1$$

where y_{ct} is an outcome variable measured for county c in year t, and η_c and λ_t are county and year fixed effects. We also show estimates produced by models with and without state by year fixed effects (θ_{st}). WIC_{ct} is the WIC treatment variable, which is equal to one if county c has a WIC program in place in year t.[8] We match births occurring in the first quarter of the calendar year with program and other county-level variables for the previous year, based on evidence that the third trimester is the most important in determining birth weight.[9]

Eq. (1) includes both county and year fixed effects, so that identification comes from cross-county variation in the timing of program introduction. An unbiased estimate of the program impact requires that there are no contemporaneous county level *trends* that are correlated with WIC introduction and infant health. The introduction of WIC took place during a period of tremendous expansion in cash and noncash transfer programs, as the War on Poverty and Great Society programs expanded. County-level variation in WIC roll-out is central to disentangling WIC from these other programs. We control for possible confounders by including two sets of county level control variables. First, to account for possible changes in other social programs, we include three measures of per capita government transfers GT_{ct} (cash public assistance, medical care, and retirement and disability programs), which are measured annually at the county level (U.S. Bureau of Economic Analysis, 2007).[10] We also include an indicator for food stamp program (FSP_{ct}) availability in the county-year. These data have been previously collected by Hoynes and Schanzenbach (2009).

Following Hoynes and Schanzenbach's work on the implementation of the food stamp program, we also include a set of county level variables Z_c that might be correlated with program introduction. Hoynes and Schanzenbach (2009) find that counties that adopted the food stamp program early were also counties with a relatively high fraction of elderly, young, black, and low income residents, whereas counties that adopted later were more likely to be rural. These county characteristics explain very little of the variation in food stamp adoption dates, but because we wish to adopt a conservative approach we include them in our analyses. Specifically, we include 1970 (pre-treatment) measures of the above variables and interact them with a linear time trend.[11] We also directly estimate the relationship between these variables and the probability that WIC is adopted. The results of this exercise are discussed in Section 6. Because most means tested programs are administered at the state level, we also include a full set of state-year dummies θ_{st}. This helps avoid possible

[7] States have had some scope to set WIC eligibility and policies including: income allowances, processes for declaring income, linkages with other programs, and specifics of eligible food items.

[8] Ideally we would match birth outcomes to WIC availability according to the quarter and year in which WIC was implemented. Unfortunately, we do not have this level of detail on the timing of WIC adoption.

[9] See the literature review of Rush et al. (1980). For example, the cohort exposed to the Dutch Famine in the third trimester had lower average birth weight than cohorts exposed earlier in pregnancy (Painter et al., 2005). Almond et al. (forthcoming) also find that birth weight gains associated with maternal participation in food stamps was concentrated in third trimester exposure.

[10] Most of these programs (such as Aid to Families with Dependent Children, Medicaid, Medicare, and Social Security) are federal or state-administered programs and, therefore, we expect their impacts will be absorbed by the state-year fixed effects.

[11] The vector Z_c includes county variables measured prior to WIC, and is non time varying. Any fixed characteristics of the county will be captured by the county fixed effects. But by interacting the county characteristics with a linear time trend we hope to control for possible differences in *trends* across counties.

contamination from coincident timing in the implementation or expansion of other means-tested programs.

All estimates are weighted using the number of births in the county-year cell and standard error estimates are clustered at the county level. To further protect against estimation problems that might be associated with thinness in the data, we drop all county-year cells where there are fewer than 25 births. The results are not sensitive to this sample selection.

We focus on two dependent variables: mean birth weight (in grams) and the fraction of newborns classified as "low birth weight" (less than 2500 g). In some analyses we also look at how the introduction of WIC affected the birthrate among women between the ages of 15 and 44. Our choice of dependent variables is largely dictated by the availability of county-level time-varying measures of infant health, and the lack of comparable data on other health or nutritional outcomes for women, older infants, or children. Nevertheless, pregnant women have always been a key part of the WIC target population, and towards the end of our sample period (1978) approximately 20% of WIC participants were pregnant or postpartum women (Oliveira et al., 2002).[12] The program goal of improving nutrition among this group suggests that infant health should be a key outcome for evaluating WIC efficacy. Birth weight outcomes are both important in their own right and predictive of later health and socioeconomic success (Currie and Hyson, 1999; Black et al., 2007 and Oreopoulos et al., 2006).

Our "program implementation" research design identifies the impact of WIC on the population, where the relevant population is the particular sample used in the regression. This is different from (but related to) typical estimators in the WIC literature, where researchers have estimated the impact of individual participation in WIC on individual outcomes. To make our estimates comparable, we scale our estimates up by the fraction of the population that participates in the program. This is the usual manipulation from the "intent-to-treat" effect to the average effect of the "treatment on the treated."

5. Data

5.1. WIC program data

The policy variable of interest is the year that each county first implemented WIC. Since the date that each county first began WIC services is not available from a unified source, we compile information from several directories and congressional filings that provide lists of local agencies that directly provided WIC services. More details, including the full names of these directories, are included in Appendix A. In short, we have county-level information on WIC local agencies for the years 1974, 1975, 1978, 1979, and 1989. These years span the period over which the WIC program expanded. The first WIC office (post-pilot program period) was opened in Kentucky in 1974 and by the end of that year WIC served approximately 88,000 women and infants. Five years later, in 1979, the program served 1.5 million women, infants and children.[13]

Our main analyses are based on data from 1971 through 1975, and 1978 through 1982. We do not include 1976 and 1977 because we have no information on which counties began offering WIC services during those years. In addition, our information for 1978 is incomplete — we have county-level information on WIC for only 13 states that participated in a WIC migrant study. If, in 1978, a county is not in one of the 13 states, but is known to have offered WIC services in a previous year, we retain that county in the sample. The WIC implementation variable is set to missing for the other counties that are in the 37 states not covered in the 1978 directory, and they are not included in our regressions.[14] Our results are robust to the exclusion of 1978.

Table 1
Mean characteristics of counties included and excluded in sample.

	Included counties	Excluded counties
1970 census variables		
Population	242,707	79,753
Percent black	11.0	10.0
	(10.8)	(15.6)
Percent urban	78.2	59.5
	(25.4)	(28.3)
Percent on farm	2.6	6.0
	(5.4)	(7.6)
Percent under 5	8.6	8.4
	(1.0)	(1.1)
Percent over 65	9.6	10.4
	(2.9)	(3.7)
Percent of persons poor	9.9	11.3
	(6.9)	(8.2)
1970 county per capita transfers (2000$)		
Family assistance	$116	$63
	(100)	(55)
Medical payments	$300	$230
	(170)	(111)
Retirement and disability	$742	$757
	(203)	(218)
Per capita income	$18,563	$16,865
	(4150)	(4013)
1970 birth outcomes		
Average birthweight	3275	3294
	(61)	(76)
Fraction low birthweight	0.079	0.076
	(0.017)	(0.020)
Number of counties	2059	1041
Fraction of the population (1970)	0.86	0.14
Fraction of births (1970)	0.85	0.15

We focus on the set of counties that had established a local WIC agency by 1979 because our next available data source is ten years later (1989). Our sample includes about two thirds of all counties in the United States (2059 out of 3100) and covers 86% of the 1970 population (85% of births). Fig. 2 shows that the counties that had not adopted WIC by 1979 were dispersed throughout the U.S. Table 1 presents characteristics of our sample counties compared to those excluded from the analysis. Sample counties are more urban and populous, with a smaller share of farm residents. Average income is also somewhat higher in our sample, and a smaller fraction of the population is below the poverty line. On the other hand, there is little difference in 1970 infant health outcomes between the included and excluded counties.

5.2. Vital statistics natality data

Our birth outcomes are taken from vital statistics records, which are coded from birth certificates and available beginning in 1968. Depending on the state and year, the data represent either a 100% or 50% sample of births, and include about 2 million observations per year. Reported birth outcomes include birth weight, gender, and (in some state-years) gestational length. The data also include (limited) information on the age and race of the mother, maternal education (in some state-years), and each infant's county and year of birth. This information allows us to link individual natality outcomes to our WIC indicator, and to collapse the data into county-year cells covering 1971–1975 and 1978–1982. Our dependent variables are the mean birth weight in each county-year cell, and the fraction of infants in each county-year who are classified as low birth weight. Our observation period ends in 1982; three years after all counties in our sample have adopted the program.

[12] Recently, the fraction of children participating in WIC has grown. In 2007 pregnant women made up just 11% of WIC recipients (Oliveira and Frazao, 2009).

[13] For program data see http://www.fns.usda.gov/pd/wisummary.htm.

[14] This explains the greater number of counties in 1978 that are shaded grey in Fig. 2.

5.3. Other controls

5.3.1. County population data

The CANCER-SEER population data provide estimates of the population of women ages 15–44 by county-year.[15] These population estimates are considered to be an improvement over population estimates interpolated from the Census because they are based on a sophisticated algorithm that incorporates information from Vital statistics, IRS migration files and the Social Security database. These population data are used together with our natality data to construct fertility rates — defined as births per 1000 women ages 15–44. In some analyses, we use these fertility rates as an additional dependent variable.

5.3.2. County control variables

We have obtained a number of pre-treatment economic and demographic control variables from the 1970 IPUMS. These variables include the percent of the 1970 county population that: lives in an urban area, on a farm, is black, is less than 5, is 65 or over, or is poor. The IPUMS also provides a measure of the county population in 1970, which we include in log form. We also include measures of annual, county real per capita income and government transfers including cash public assistance benefits (Aid to Families with Dependent Children AFDC, Supplemental Security Income SSI, and General Assistance), medical spending (Medicare and Military health care), and cash retirement and disability payments using data from the Bureau of Economic Analysis, Regional Economic Information System (REIS).

6. Results

6.1. County adoption of WIC

The validity of our research design hinges on the exogeniety of county WIC start dates. Thus, we begin by empirically investigating the determinants of county implementation. The written history of WIC provides little intuition as to which types of counties were likely to be early adopters, so, as described in Section 5, we have simply gathered data on as many pre-treatment county characteristics as we could, and estimate the relationship between these characteristics and WIC initiation. We estimate three separate cross-county regressions. In the first regression, the dependent variable is equal to 1 if the county had a WIC program in place by 1974, and 0 if otherwise. In the second regression, the dependent variable is an indicator for whether the county had adopted WIC by 1975. In the third regression, the dependent variable is an indicator for whether WIC was in place by 1978. The regressors of interest are the county level variables described above. Each regression includes state fixed effects, and is weighted by the 1970 county population.

Table 2 shows that few of these variables have any predictive power. Notably, there is no association between 1970 infant health outcomes and the timing of WIC start-up. We do find that counties with a larger population and a higher poverty rate are more likely to implement WIC programs earlier, but taken as a whole, the quantitative importance of these predictors is small. For example, the negative and significant coefficients on the fraction poor in a county suggest that counties in the highest quartile of poverty rates are between one-half and one percentage point more likely to adopt WIC in a given year than those in the lowest quartile of poverty rates. This is a very small effect relative to the 60% of counties that had implemented WIC by 1975.

Table 2 makes clear that most of the variation in the timing of WIC implementation is unexplained, and this may be related to some early difficulties in getting the program launched. Historical documents suggest that, during the program's initial years, there was excess demand for WIC services (U.S. Congress, 1976), and that several lawsuits were filed by individual communities attempting to get the program started. This is consistent with relatively idiosyncratic factors determining both which counties started WIC programs first, and which individuals received initial program services. Nonetheless, in order to control for possible differences in county-level trends that might be spuriously correlated with the county treatment effect, our main regressions include all of our 1970 county variables interacted with linear time trends.

Table 2
Determinants of county WIC start date.

	Probability that county has adopted WIC in		
	1974	1975	1978
1970 census variables			
Percent black	0.001	−0.002	0.001
	(0.002)	(0.003)	(0.002)
Percent urban	0.000	0.004***	0.000
	(0.002)	(0.001)	(0.001)
Percent on farm	0.008	0.011**	0.002
	(0.005)	(0.004)	(0.004)
Percent under 5	0.003	−0.012	−0.021
	(0.015)	(0.013)	(0.018)
Percent over 65	−0.016	−0.024**	−0.010
	(0.010)	(0.010)	(0.009)
Percent of persons poor	0.014**	0.008*	0.006***
	(0.007)	(0.004)	(0.002)
Log (population)	0.160***	0.098***	0.038**
	(0.035)	(0.023)	(0.014)
1970 county per capita transfers (2000$)			
Family assistance	−0.126	0.955	0.009
	(0.654)	(0.503)	(0.145)
Medical payments	−0.064	−0.231	−0.116
	(0.343)	(0.261)	(0.073)
Retirement and disability	0.135	0.245**	0.091
	(0.087)	(0.100)	(0.062)
Per capita income	0.004	0.006	0.003
	(0.008)	(0.006)	(0.002)
1970 birth outcomes			
Average birthweight	0.000	−0.001	0.000
	(0.001)	(0.001)	(0.000)
Fraction low birthweight	2.583	−0.250	0.871
	(2.153)	(1.357)	(0.752)
State fixed effects	X	X	X
Number of observations	1911	1950	1123
R squared	0.42	0.47	0.20

These are county level regressions where the dependent variable equals 1 if the county has adopted WIC by the given year. See text for information on sample. Standard errors are in parentheses and ***, **, and * indicate that the estimates are significant at the 1%, 5%, and 10% levels.

6.2. Effects of WIC on infant health

Table 3 provides our main estimates of the effect of WIC on infant health. The first two columns of Table 3 show the estimated relationship between WIC and average birth weight, and the second two columns show the estimated effect of WIC on the fraction of births that are classified as low birth weight. All of the estimates are based on regressions that control for county fixed effects, 1970 county-level variables interacted with linear time, and the time-varying county-level variables described in Section 5. The second column in each panel adds state by year fixed effects, which helps control for the possibility that changes in state-level social programs might be correlated with provision of WIC services. We have also estimated these models including controls for the cell-level distribution of race, father's presence, and mother's age and education; including these controls reduces the point estimate somewhat but does not substantively change our findings.

The first column of Table 3 shows that WIC availability increased average birth weight by a statistically significant 2.7 g. Adding state-year fixed effects to the regression changes the estimate only slightly,

[15] See National Cancer Institute http://seer.cancer.gov/popdata/download.html.

Table 3
Impacts of WIC introduction on birth outcomes, full sample.

	Average birth weight (in grams)		Share low birth weight (<2500 g)	
WIC implementation	2.7**	2.3*	0.0000	0.0001
	(1.22)	(1.21)	(0.0004)	(0.0004)
Observations	18,517	18,517	18,517	18,517
R-squared	0.91	0.92	0.77	0.78
Mean of dependent variable	3316	3316	0.072	0.072
Coeff/mean	0.1%	0.1%	−0.1%	0.0%
Controls				
1970 county char * time	X	X	X	X
REIS transfers, per capita income	X	X	X	X
Year fixed effects	X	X	X	X
County fixed effects	X	X	X	X
State * year fixed effects		X		X

Each parameter is from a separate regression of the outcome variable on the WIC implementation dummy. The sample includes county-year means for years 1972–1975 and 1978–1982 where cells with less than 25 births are dropped. Data for 1976 and 1977, and for many states in 1978, are missing due to incomplete WIC data. In addition to the fixed effects, controls include 1970 county variables (log of population, percent of population black, poor, urban, on farm, age<5, age>65) each interacted with a linear time trend, per capita county transfer income (cash assistance, medical care, and retirement and disability benefits), and county real per capita income. Estimates are weighted using the number of births in the cell and are clustered on county. Standard errors are in parentheses and ***, **, and * indicate that the estimates are significant at the 1%, 5%, and 10% levels.

Table 4
Impacts of WIC introduction on birth outcomes, by maternal education.

	Average birth weight (in grams)	Share low birth weight (<2500 g)
Less than high school		
WIC implementation	7.0***	−0.0014
	(2.50)	(0.0010)
Coeff/mean	0.2%	−1.4%
Observations	14,831	14,831
Mean of dependent variable	3205	0.102
High school		
WIC implementation	1.6	−0.0003
	(1.75)	(0.0007)
Coeff/mean	0.0%	−0.4%
Observations	16,022	16,022
Mean of dependent variable	3329	0.069
More than high school		
WIC implementation	0.0	0.0005
	(2.00)	(0.0008)
Coeff/mean	0.0%	1.0%
Observations	13,773	13,773
Mean of dependent variable	3388	0.054
Controls		
1970 county char * time	X	X
REIS transfers, per capita income	X	X
Year fixed effects	X	X
County fixed effects	X	X
State * year fixed effects	X	X

Each parameter is from a separate regression of the outcome variable on the WIC implementation dummy. The sample includes county-year means for years 1972–1975 and 1978–1982 where cells with less than 25 births are dropped. Data for 1976 and 1977, and for many states in 1978, are missing due to incomplete WIC data. In addition to the fixed effects, controls include 1970 county variables (log of population, percent of population black, poor, urban, on farm, age<5, age>65) each interacted with a linear time trend, per capita county transfer income (cash assistance, medical care, and retirement and disability benefits), and county real per capita income. Estimates are weighted using the number of births in the cell and are clustered on county. Standard errors are in parentheses and ***, **, and * indicate that the estimates are significant at the 1%, 5%, and 10% levels.

to 2.3 g. Both estimates are statistically significant. These are small effects on the overall population, indicating an increase in birth weight of one-tenth of one percent of the mean (labeled as "Coeff/Mean" on this and future tables), but recall that most pregnant women are not eligible for WIC and would not have been affected by the program's implementation. The remaining columns in the table suggest that the small increase in birth weight does not appear to occur around the 2500 g margin.

An earlier debate in the WIC literature (Joyce et al., 2008) has focused on the mechanisms by which WIC might produce increases in birth weight. In particular, Joyce et al. raise concerns about previous findings which suggest that WIC participation increases birth weights partially through increases in gestational age. They argue that clinical evidence suggests little or no link between dietary supplementation and reductions in pre-term births. As a result, any effects of the WIC program on gestational age are likely to be spurious. We agree that gestation is not the component of birth weight variability that should be moved by the introduction of WIC. To verify that the reported effects on birth weight are not coming about as the result of changes in gestational age, we have repeated our main analysis replacing birth weight with an indicator for whether the birth was pre-term (less than 37 weeks gestational age). We find no evidence of WIC effects on gestational age; all coefficients are small and far from statistical significance.[16]

What we really care about, however, is WIC's impact on the population that actually receives assistance. In order to appropriately scale our estimates, we need an estimate of the fraction of births that were to women receiving WIC during the 1970s. Unfortunately, information on early WIC participation is not sufficiently detailed to calculate this statistic. In particular, we do not have detailed information on how many participants were women, infants, and children in the earliest years of the program. Our TOT estimates are, therefore, based on a range of statistics that we have gathered from a variety of sources over several different years. For example, the U.S. House of Representatives Green Book (1991) indicates that in 1980, 411,000 women participated in WIC. Using detailed participation categories for 1988 provided by the USDA's National Survey of WIC Participants (2001), we estimate that approximately 281,000 of these women participated as pregnant women. We also know that there were 3.6 million births in that year. When we combine these two statistics, we obtain an estimated WIC participation rate of 8%.[17] We then use this participation rate to convert our overall estimates into estimates of the effect of treatment on the treated by dividing our overall estimate of 2.3 g by 0.08. This yields an estimated TOT effect of approximately 29 g.[18]

Alternatively, Rush et al. (1988) state that 13% of all births in 1980 were to mothers on WIC.[19] If we divide 2.3 g by 0.13 then our estimated TOT effect is 18 g. Importantly, both of these TOT estimates are smaller than most estimates in the literature. Devaney et al. (1990), for example, produce estimates that range from 51 to 113 g across several states. Bitler and Currie (2005) estimate that WIC increases average birth weight among WIC participants by 62 g, even after including observable controls for selection. Interestingly, Rush et al. (1988) which also focus on the early years of the WIC program (but we argue is still subject to selection bias) find that state WIC availability increases average birth weight by 23 g, which is close to our estimate. The average effects of the program today may be smaller

[16] Results are available from the authors upon request. We have also estimated models where the dependent variable is the fraction of births below a gestation-varying threshold (known as small-for-gestational-age models). These regressions yielded results very similar to those produced when the dependent variable is an indicator for low birth weight.

[17] The National Survey of WIC Participants reports that of the total caseload in 1988, 13% were pregnant women, 3% were breastfeeding women, and 6% were postpartum women (USDA, 2001). In 1980, breastfeeding women were not eligible for WIC, so we use this information to estimate that roughly two-thirds (13/(13 + 6)) of all women on WIC were pregnant. Multiplying 411,000 participating women by 2/3 gives us an estimated 289,000 pregnant women participants, compared with 3.6 million total births, for an estimated participation rate among pregnant women of 8%.

[18] We acknowledge that our estimated participation rates are imperfect and that mismeasurement could lead to either higher or lower TOT estimates. Given that the program faced funding constraints in its early phases it is likely that participation rates were lower than our estimates suggest. Thus, the true TOT effects may be higher. Our estimates reflect our best guess given the information that is available.

[19] Rush et al. do not elaborate on how this statistic was calculated.

than our estimated effects for two reasons. First, the program was initially implemented in areas most in need, where one might expect the resulting birth weight increase to be highest. Second, participation rates have grown substantially over time. If the increases in participation rates have been driven by relatively more advantaged women, who are likely to be healthier, then the average benefits of WIC today will be smaller than in the early years of the program.[20]

Since WIC is a means tested program we expect its impact to be concentrated among families with low incomes. Although the vital statistics birth certificate data does not include data on family income, it does include information on maternal education, which we use to identify groups that are more or less likely to be affected by WIC. We present results by maternal education in Table 4, where we collapse our data by county, year, and maternal education based on three categories: less than high school, high school, and greater than high school.[21] We find that among women with less than a high school education, the availability of WIC in the county of birth increases average birth weight by a statistically significant 7.0 g. We also estimate a 1.4% (not statistically significant) decrease in the probability of being born with low birth weight. Moving down the rows of Table 4 we see that as the education of the mother increases, the magnitude of the estimated treatment effects moves towards zero. As expected, we see that the estimated impact of WIC on mothers with more than a high school degree is very small and statistically insignificant. Finally, because some states do not have mother's education available on the birth certificate in all years, we have re-estimated the analysis in Table 4 using a balanced sample, composed only of states for which we observe maternal education in all years. This does not substantially change our results.

Our estimates for women with less than a high school education can be translated into a treatment on the treated effect by using a WIC participation rate specific to this population. Rush et al. (1986) indicate that in 1983, 55% of pregnant WIC participants had less than a high school education, 34% had graduated from high school, and 11% had more than a high school education. We combine this information with the number of births in 1980 and the two alternative participation rates described above (8% and 13%) to obtain two sets of treatment on the treated estimates by mothers' education level. These range from 23–38 g for births to mothers who have less than a high school education, to 16–27 g among mothers with exactly 12 years of education.[22] As would be expected, these estimates among lower educated women are larger than our overall estimate, but still on the low end of estimates that are based on comparisons between participants and non-participants. Our estimated effect of WIC on the fraction of infants with low birth weight (−0.0014) implies an effect on participants of less than 1 percentage point, or roughly an 8% increase in the probability of being below the low birth weight threshold. Currie's (2003) review summarizes earlier studies as finding effects on the probability of low birth weight of 10 to 43%. Bitler and Currie (2005) estimate that WIC reduces the probability that a baby is classified as low birth weight by 29%. Among low educated women, our results for WIC's impact on the probability of a low birth weight are again below many previously published estimates.

Table 5
Impacts of WIC introduction on birth outcomes, by quartiles of 1970 county poverty.

	Average birth weight (in grams)	Share low birth weight (<2500 g)
Highest poverty quartile		
WIC Implementation	7.1***	−0.0005
	(2.35)	(0.0010)
Coeff/mean	0.2%	−0.6%
Observations	10,464	10,464
Mean of dependent variable	3286	0.080
Lowest poverty quartile		
WIC implementation	−0.2	0.0011
	(2.87)	(0.00089)
Coeff/mean	0.0%	1.7%
Observations	1,632	1,632
Mean of dependent variable	3348	0.064
Controls		
1970 county char * time	X	X
REIS transfers, per capita income	X	X
Year fixed effects	X	X
County fixed effects	X	X
State * year fixed effects	X	X

Each parameter is from a separate regression of the outcome variable on the WIC implementation dummy. The sample includes county-year means for years 1972–1975 and 1978–1982 where cells with less than 25 births are dropped. Data for 1976 and 1977, and for many states in 1978, are missing due to incomplete WIC data. In addition to the fixed effects, controls include 1970 county variables (log of population, percent of population black, poor, urban, on farm, age<5, age>65) each interacted with a linear time trend, per capita county transfer income (cash assistance, medical care, and retirement and disability benefits), and county real per capita income. Estimates are weighted using the number of births in the cell and are clustered on county. Standard errors are in parentheses and ***, **, and * indicate that the estimates are significant at the 1%, 5%, and 10% levels. Quartiles are assigned using 1970 county poverty rates (weighted using county population).

We show the results of a similar exercise in Table 5, where we stratify our sample according to the fraction poor in 1970. We present results for counties whose 1970 poverty rates were in the top quartile (where the average poverty rate is 20%) and the bottom quartile (where the average poverty rate is 4%).[23] We expect that, given that WIC is means tested, counties with higher poverty rates should be more affected by the initiation of the program. Table 5 shows that among counties with the highest poverty rates, WIC adoption raised average birth weight by a statistically significant 7 g and reduced the fraction born with low birth weight by an insignificant 0.6% of the mean. The estimated impacts in low poverty counties (where we would expect very low WIC participation rates) have signs opposite of the expected effects for WIC participants and are not significantly different from zero.

Taken together, Tables 4 and 5 provide further evidence that we are truly identifying the impact of WIC implementation. Our estimated impacts appear to be concentrated among groups where WIC participation should be highest. We find no impacts within groups where WIC participation is expected to be low.

[20] A related point is that participation rates in the first few years of the program are likely to have been even smaller than the 8% estimate for 1980. The total caseload for 1975, for example, is less than half the caseload for 1980, although 60% of (population-weighted) counties had a WIC program at that point. Because our results are based on data from 1972 through 1982, the relevant participation rate is a mixture between these lower rates from the earlier years and the rates we use for 1980. Thus, a treatment on the treated estimate for the first few years of the program would probably be larger than the estimates for the full period reported above.

[21] State reporting of maternal education varies during the 1970s. In 1977, for example, about 80% of births were in states that provided maternal education. For this analysis, we drop state-years missing data on education.

[22] We estimate these treatment-on-the-treated effects using two approaches. In the first, we multiply the total number of births in 1980 (3.6 million) by 13% to get an estimated number of WIC births in 1980 (468 thousand). We then multiply this number by the fraction of WIC births in each education category (0.55, 0 .34 or 0.11) in 1983, to get an estimate of the number of WIC births in 1980 by mothers' education category. This number is then divided by the total number of births by education category in 1980, provided by the Vital Statistics data. This produces an estimate of WIC participation rates by education group. We then divide our estimated birth weight effects for each education group by the estimated participation rate for each group. In the second approach we multiply the number of women on WIC in 1980 (411 thousand) that is provided by the 1991 Green Book, by our estimate of the fraction who were pregnant (68%), which is based on statistics provided by the National Survey of WIC Participants, Final Report (2001). This gives us an estimate of the total number of WIC births (281 thousand) in 1980. We multiply this number by the fraction of births in each education category (0.55, 0.34, 0.11) to get an estimate of the number of WIC births by education. These estimates are then divided by the total births to each education group to get an estimated WIC participation rate for each education group. The two methods give us WIC participation rates for women with less than a high school degree of 0.18 and 0.30. Finally, we divide the estimated birth weight effects for each education group by the estimated participation rates for each education group.

[23] Poverty quartiles are assigned using the 1970 population as weights. In Table 5, the larger number of observations in the high poverty quartile regressions reflects the fact that smaller counties (in 1970) had higher poverty rates.

Although the estimates are not shown,[24] we have also looked at the effects of WIC introduction by race of the mother, age of the mother, presence of father's information on the birth certificate, and gender of the child. Estimation stratified by gender of the child and father's presence show no statistically distinguishable patterns of effects across groups, although the point estimates are generally larger for those cases with no father on the birth certificate, as would be expected if these are births to more disadvantaged, single, women.[25] Stratifying by race, we also do not find estimates that are statistically different from one another. The lack of a clear pattern of results by race largely reflects a lack of precision. For blacks, in particular, the number of observations falls because there are more likely to be county-year cells with birth counts that are too small to be useful (as blacks are not only a small share of the total population but they are also more geographically concentrated). It is, however, somewhat surprising that we do not find even marginally significant effects of WIC introduction for births to blacks, given that they are likely to be less advantaged than whites. One possibility is that, in the early years of the program introduction, blacks were far less likely to participate in WIC than disadvantaged whites. In its early years, WIC was frequently administered through local health centers. If blacks were less likely to be served by, or connected to these centers than whites, their early participation in WIC may also have been low.[26] At this point, we have little direct evidence on participation rates by race in WIC's early years, and so cannot state with certainty whether the lack of strong program estimates is due to low participation rates, lack of statistical power, or truly small effects during the program's initial years.

Finally, the effects of WIC may depend upon the other resources available to low-income, pregnant women. During most of the period of our analysis, women who were eligible for WIC would also have been eligible for AFDC. Given this, WIC might be expected to have a smaller impact on the birth outcomes of low income women residing in states with relatively generous cash welfare programs. To test this hypothesis we calculate the state's 1970 AFDC maximum monthly benefit level for a family of 3, divided by the 1970 poverty threshold, and then use this measure to rank each state's level of welfare generosity. We then re-estimate the impact of WIC separately among counties in the top and bottom quartiles of the AFDC benefit distribution. The results of this exercise are shown in Table 6. While the estimates are not statistically significant, the pattern is consistent with our priors: the positive effects of WIC are largest in states with low AFDC benefits.

WIC's nutritional benefits are targeted on pregnant women with the aim of improving fetal development and reducing the incidence of low birth weight. Our point estimates suggest that WIC reduces the percent of births below 2500 g, but the estimates are not statistically significant. To further investigate the impact of WIC on the distribution of birth weight, we estimate a series of models relating the introduction of WIC to the probability that birth weight is below a given gram threshold, specifically: 1500, 2000, 2500, 3000, 3250, 3500, 3750, and 4000. Our regressions include state-year fixed effects (as in the 4th column of Table 3) and are estimated for two groups: births to mothers with less than a high school education and births to mothers living in the highest poverty quartile counties. The results are presented in Fig. 3 where we plot percent impacts — the estimated coefficient divided by the mean of the dependent variable. We find that the largest percent reduction in the probability of birth weight below a certain threshold occurs at the bottom of the birth weight distribution. The impacts become gradually smaller as the birth weight threshold is increased, and reaches near zero for births below 3750 g.

Table 6
Impacts of WIC introduction on birth outcomes, by quartiles of state AFDC/Medicaid income eligibility.

	Average birth weight (in grams)	Share low birth weight (<2500 g)
Highest AFDC benefit quartile		
WIC implementation	−1.4	0.0013
	(2.95)	(0.0010)
Coeff/mean	0.0%	1.8%
Observations	1760	1760
Mean of dependent variable	3303	0.073
Lowest AFDC benefit quartile		
WIC implementation	2.7	−0.0004
	(2.17)	(0.0008)
Coeff/mean	0.1%	−0.5%
Observations	7966	7966
Mean of dependent variable	3298	0.079
Controls		
1970 county char * time	X	X
REIS transfers, per capita income	X	X
Year fixed effects	X	X
County fixed effects	X	X
State * year fixed effects	X	X

Each parameter is from a separate regression of the outcome variable on the WIC implementation dummy. The sample includes county-year means for years 1972–1975 and 1978–1982 where cells with less than 25 births are dropped. Data for 1976 and 1977, and for many states in 1978, are missing due to incomplete WIC data. In addition to the fixed effects, controls include 1970 county variables (log of population, percent of population black, poor, urban, on farm, age<5, age>65) each interacted with a linear time trend, per capita county transfer income (cash assistance, medical care, and retirement and disability benefits), and county real per capita income. Estimates are weighted using the number of births in the cell and are clustered on county. Standard errors are in parentheses and ***, **, and * indicate that the estimates are significant at the 1%, 5%, and 10% levels. Quartiles are assigned using 1970 maximum state AFDC benefits for a family of three (weighted using county population).

7. Further robustness checks

The results above suggest that average county birth weight increased following WIC adoption. An important question is whether

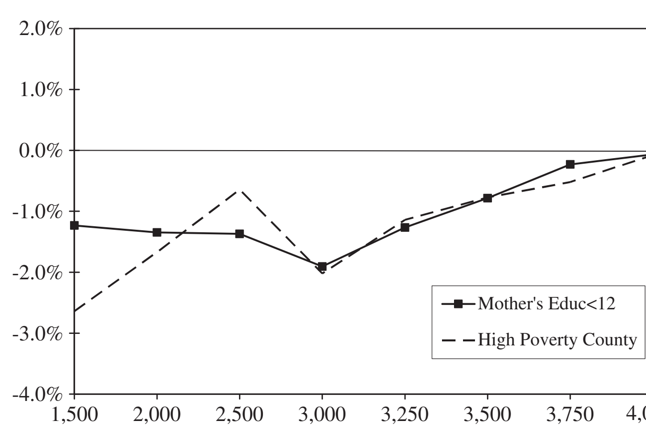

Fig. 3. Effects of WIC implementation on distribution of birth weight, percent impacts (coeff/mean). Notes: The graph plots the coefficient on the WIC implementation dummy in a model where the dependent variable is the fraction of births in the county-year that is below each specified number of grams. The specification is given by column (2) in Tables 4 and 5.

[24] Results are available from the authors on request.

[25] In some states in some years a variable is provided for whether the birth is “illegitimate” or legitimate. In all years, there is a variable for father's age. We code a father as “absent” if either the birth is coded as illegitimate or the father's age is missing.

[26] In the South, there is anecdotal evidence that some clinics serving blacks may have had difficulty in getting their WIC proposals through the state bureaucracy to the federally funded program. A report on the status of WIC from 1975 recounts that “In Mississippi, an application from a black community-controlled clinic, The Voice of Calvary Health Center, sat for months without action from the state, in spite of complaint letters to Congress, the press, USDA and the U.S. Commission on Civil Rights.” (Fleming, 1975).

the counties that implemented WIC early were otherwise similar to counties that initiated WIC at the end of the period. Our focus on within county changes and inclusion of extensive controls for differential county trends help make the case that our research design is truly capturing the program's causal effect. We also employ an event study analysis to help focus more directly on how birth outcomes among early adopters compare to birth outcomes in other counties in the years just prior to WIC implementation. Specifically, we estimate the following equation

$$y_{ct} = \alpha + \sum_{i=-3}^{3} \delta_i 1(\omega_{ct} = i) + \gamma_1 Z_c t + \gamma_2 GT_{ct} + \eta_c + \lambda_t + \theta_{st} + \varepsilon_{ct}. \quad (2)$$

In this analysis, the treatment variables are contained in the ω terms, a series of dummy variables that capture the number of years before and after WIC is implemented in county *i*. For example, ω_0 is set equal to one in the year a county first implements WIC, and ω_{-2} is set equal to one if two years prior to implementation. We omit year -1 from the set of event time variables, so that all of the estimated effects are measured relative to the year just prior to implementation. As in our main specification, the event study model includes county- and time-varying information on other government transfer programs (GT_{ct}), 1970 county characteristics interacted with a linear trend ($Z_c t$), and county (η_c) and year (λ_t) fixed effects. We estimate models with and without state by year fixed effects (θ_{st}). If our treatment effects are being driven by the WIC program itself, and are not capturing other underlying differences in counties around the time of adoption, we should expect the event time variables prior to the date of implementation to be equal to zero (suggesting no pre trend bias), and those at and after implementation to be positive (birth weight analyses) or negative (low birth weight analyses).

The coarseness of the WIC adoption data forces us to modify the sample slightly when we estimate this model. Specifically, recall that we have no data on implementation during 1976 or 1977 (and for this reason those years are not included in our main model). Further, we only have data for county WIC programs for a subset of states in 1978. In order to conduct a meaningful event study, we include only those counties for which we can measure the time before and after implementation. This means we include all counties that are first observed to implement WIC in 1974, 1975, 1978 or 1979. Among counties that first adopted WIC in 1978 our WIC variable faces potential measurement error since adoption could have occurred at any time during 1976, 1977, or 1978. This will bias our estimated effects towards zero. We also include counties observed in 1978 (in the subset of states observed in 1978) that have not yet implemented WIC. Among this sample of counties implementing in 1974, 1975, 1978 or 1979, we construct event time variables for all years from 3 years prior to implementation to 3 years after (so that the panel is balanced). This gives us a sample of approximately 13,600 county-year cells, or roughly three-quarters the size of our main sample. The sample includes fewer counties, since we must drop all counties that are not observed in 1978 but have implemented by 1979, but it includes more years per county since we now include 1976 and 1977 in the analysis.

Our event study results are shown in Fig. 4. The figure includes four panels, the top two are for less educated mothers only, and the bottom two are for the full sample. The results confirm our earlier findings: following WIC implementation, there is an increase in birth weight that is strongest among less-educated mothers. Similarly, there is a slight decline in the fraction of infants born with low birth weight, although the low birth weight result is only visible for less educated mothers. Among these women, we see no systematic differences in birth weight in the years prior to implementation. The estimates are robust to adding state by year fixed effects. This provides strong evidence that our research design is identifying the causal impact of WIC.

The next set of analyses address the concern that the same forces leading to an association between WIC and better birth outcomes, also lead to a change in the composition of births. In particular, if improvements in fetal health lead to fewer fetal deaths, there could be a *negative* compositional effect on birth weight from higher likelihood that “marginal” fetuses survive. In addition, WIC may lead to increased fertility among disadvantaged women if children are a normal good.[27] Both these factors might lead to endogenous sample selection and contribute towards downward biased estimates. To evaluate this possibility, we estimate the impact of WIC on total births and the characteristics of mothers giving birth. The results of this exercise are shown in Table 7. The first two columns present estimates of the impact of WIC implementation on the fertility rate, which is defined as total births per 1000 women ages 15–44. The estimates are very small and statistically insignificant for both the full sample and the counties in the highest poverty quartile.[28] Therefore, the total level of fertility does not appear to be affected by WIC.[29] The remaining columns of the table relate county WIC implementation to observable maternal characteristics. WIC access appears to have no impact on the percent of births that are to mothers with less than a high school education, the fraction of births to minority mothers, or the fraction of births to mothers with no father present. However, the results show very small but statistically significant effects on births by age of mother — showing a decrease to mothers 15–24 and an increase to mothers 25–34. If this small change in the distribution of births biases our estimates then the bias is likely upward: mothers between the ages of 25 and 34 generally give birth to healthier babies than those who are younger.[30] In other words, the presence of these selection effects may mean that WIC's true impact is even smaller than our estimated effects.

We next perform some checks on our WIC policy variables. Rush et al. (1986), in their national WIC evaluation, include a tabulation of the number of counties served, by year and state, from 1973 to 1981 for the 25 states included in the study (see Appendix A for list of covered states). We can use this, in a state-year research design, to provide indirect evidence on our WIC policy variable. To implement this robustness check, we collapse our data to the state-year level. We estimate models similar to those presented above and control for state and year fixed effects, state specific linear time trends, state real per capita income and state real per capita government transfers. We present estimates using three alternative WIC policy variables: (1) Rush et al.'s fraction of counties in the state that have WIC programs in place, (2) the fraction of counties in the state with WIC programs in place from our data, and (3) the weighted fraction (using 1970 population) of counties in the state with WIC in place from our data. Because neither our data nor the Rush et al. (1986) data covers all states, we present estimates for a common sample followed by all available observations. In all models we weight by the state population. The results, provided in Table 8, show that when we limit the analysis to the common sample, our unweighted state WIC implementation variable and the comparable variable from Rush et al. (1986) provide very similar results for both the overall sample and among births to less educated mothers. Note that our precision is reduced significantly when we move from the county to the state identification strategy. The state-year approach is useful in comparing our data on WIC implementation with the data

[27] The existing literature suggests that the elasticity of fertility with respect to additional transfers from income support programs is very small (Moffitt, 1998).

[28] Note that we cannot estimate fertility models by education group because the county population does not break out education groups.

[29] We also estimated fertility models by age of woman (15–24, 25–34, and 35–44) and found small statistically insignificant effects, as in the full sample.

[30] The youngest mothers are more likely to have low birth weight infants than are mothers aged 24 to 35. Friede et al. (1987) report that 5.5% of births to white mothers aged 15 to 24 are low birth weight, compared with 4.2% of births to those aged 25 to 29.

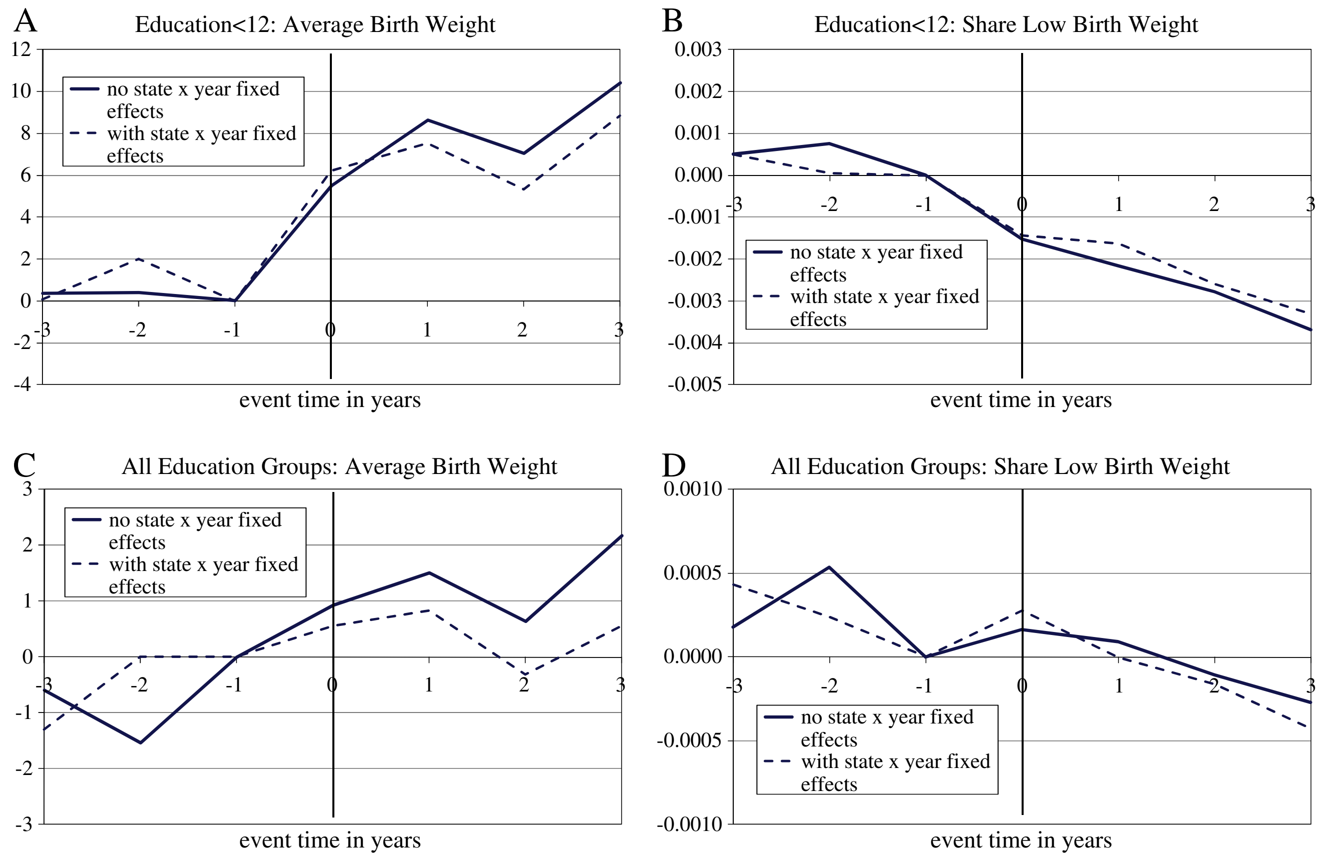

Fig. 4. Event study estimation of effects of WIC implementation. Notes: Each figure plots coefficients from an event-study analysis. Coefficients are defined as years relative to the year WIC is implemented in the county. The sample is a balanced county sample, where a county is included only if there are at least 3 pre-years and 3 post-years of data. The specification includes controls for 1970 county controls interacted with linear time, county per capita transfers, and fixed effects for county and year; with and without state by year fixed effects.

used by Rush et al., and the results are broadly comparable, but only by using our main county-level approach can we identify treatment effects with the necessary degree of precision to draw strong conclusions.

Our results are also robust to several additional specification checks. The results do not hinge on the inclusion of data from 1978, where WIC status can only be directly observed for a small sample of counties. They are also robust to the exclusion of our county-level control variables. Further, as shown in Table 3, our estimates are similar whether or not we include state-year fixed effects. As might be expected, controlling for county and time fixed effects is important. Further, we provided placebo tests in Tables 4 and 5, showing small

Table 7
Impact of WIC introduction on fertility and maternal characteristics.

	Fertility births per 1000 women 15–44		Characteristics of mother in natality sample								
	All	All highest quartile poverty	Educ<12	Educ=12	Educ>12	White	Black	Mom 15–24	Mom 25–34	Mom 35–44	No father
WIC implementation	−0.3	0.5	−0.001	0.002	−0.001	−0.002	0.001	−0.003**	0.003**	0.001	−0.002
	(0.21)	(0.44)	(0.007)	(0.009)	(0.007)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
Observations	18,517	10,464	17,972	17,972	17,972	18,517	18,517	18,517	18,517	18,517	18,517
R-squared	0.91	0.90	0.69	0.48	0.66	0.99	0.99	0.95	0.95	0.8	0.96
Mean of dep var	71.41	80.87	0.26	0.43	0.31	0.81	0.16	0.51	0.44	0.05	0.15
Controls											
1970 county char * Time	X	X	X	X	X	X	X	X	X	X	X
REIS transfers, per capita income	X	X	X	X	X	X	X	X	X	X	X
Year fixed effects	X	X	X	X	X	X	X	X	X	X	X
County fixed effects	X	X	X	X	X	X	X	X	X	X	X
State * year fixed effects	X	X	X	X	X	X	X	X	X	X	X

Each parameter is from a separate regression of the outcome variable on the WIC implementation dummy. The sample includes county-year means for years 1972–1975 and 1978–1982 where cells with less than 25 births are dropped. Data for 1976 and 1977, and for many states in 1978, are missing due to incomplete WIC data. In addition to the fixed effects, controls include 1970 county variables (log of population, percent of population black, poor, urban, on farm, age<5, age>65) each interacted with a linear time trend, per capita county transfer income (cash assistance, medical care, and retirement and disability benefits), and county real per capita income. Estimates for fertility (mother's characteristics) are weighted using population (births) in the cell and are clustered on county. Standard errors are in parentheses and ***, **, and * indicate that the estimates are significant at the 1%, 5%, and 10% levels. Quartiles are assigned using 1970 county poverty rates (weighted using county population). The shares by education group and race are means among the sample of nonmissing observations.

Table 8
Impact of WIC introduction on birth outcomes using state-year policy variable.

	All available observations		Common sample			
	Average birth weight (in grams)	Share low birth weight (<2500 g)	Average birth weight (in grams)	Average birth weight (in grams)	Share low birth weight (<2500 g)	Share low birth weight (<2500 g)
				Less than HS		Less than HS
Unweighted fraction of counties						
WIC	4.5	−0.0003	2.6	4.7	−0.0007	−0.0016
	(4.5)	(0.0009)	(4.3)	(7.3)	(0.0014)	(0.0033)
Observations	783	783	167	167	167	167
Dep Mean	3334	0.070	3328	3224	0.070	0.100
Weighted fraction of counties						
WIC	1.8	−0.0002	1.0	1.8	0.0000	−0.0015
	(4.5)	(0.0008)	(4.1)	(10.5)	(0.0014)	(0.0043)
Observations	783	783	167	167	167	167
Dep mean	3334	0.070	3328	3224	0.070	0.100
Unweighted fraction of counties from Rush et al. (1986)						
WIC	1.7	0.0007	3.3	3.9	−0.0007	−0.0029
	(3.5)	(0.0009)	(3.7)	(10.0)	(0.0009)	(0.0037)
Observations	225	225	167	167	167	167
Dep Mean	3320	0.070	3328	3224	0.070	0.100
Controls						
REIS transfers, per cap inc	X	X	X	X	X	X
Year fixed effects	X	X	X	X	X	X
County fixed effects	X	X	X	X	X	X
State * linear year	X	X	X	X	X	X

Each parameter is from a separate regression of the outcome variable on the WIC implementation dummy in a state-year research design. The treatment variable is the fraction of counties in a state-year observation that have WIC programs in place. The first panel, using policy variable defined from Rush et al. (1986), includes data for 25 states over the years 1973–1981. The second and third panels include our sample (described above 1972–1975 and 1978–1982) collapsed to the state-year level. We construct the state treatment both unweighted (to match Rush et al.) and weighted by the 1970 county population. The "common sample" provides estimates for the sample of state-years that are common to both samples. Estimates are weighted using the number of births in the cell and are clustered on state. Standard errors are in parentheses and ***, **, and * indicate that the estimates are significant at the 1%, 5%, and 10% levels.

(but imprecisely estimated) results for groups with lower probabilities of being impacted by WIC.

8. Conclusion

This paper presents new evidence on the effect of the WIC program on infant health. Because the parameters of the WIC program have not varied tremendously over time and exhibit relatively little geographic variation, it has proven difficult to identify the program's effects. Until recently, most WIC studies have compared participants' outcomes to the outcomes of non-participants, leading to biased estimates. Attempts to control for non-random selection into WIC have confirmed that participation is unlikely to be random. We overcome this problem by exploiting county-level variation in the original adoption of the program between 1974 and 1979. We use the gradual roll-out feature of WIC implementation at the county level to identify the effect of WIC separately from the effect of other policy changes that occurred during the late 1970s and early 1980s. This research design provides an important contribution to the WIC evaluation literature, which is free of bias due to non-random participation decisions.

We find that the introduction of this transfer program improved birth outcomes, and that the effects were largest among populations most likely to be eligible for the program. Among women with less than a high school education, for example, the availability of WIC in the county of birth increased average birth weight by approximately 7 g, implying a treatment on the treated effect of 23 to 38 g. Similar birth weight increases are observed in high poverty counties. In contrast, estimated effects for more highly educated mothers and low poverty counties are small and statistically insignificant. The shift in birth weights is larger at the bottom of the birth weight distribution, where birth weight is more closely linked to other long-run outcomes. These effects on birth weight may lead to substantial and persistent effects on earnings and health. Black et al. (2007), for example, find that a 10% increase in birth weight is associated with a 1% increase in adult earnings. Our upper bound estimates suggest that WIC increases average birth weight by about 1%, which would translate into a 0.1% increase in annual earnings, or about $375.[31]

Given simple estimates of the cost of providing WIC to pregnant women,[32] it is tempting to use these earnings estimates to calculate a cost-benefit ratio for the program. At least three major considerations, however, make this exercise difficult. First, the population of WIC participants today looks very different (and is less disadvantaged) than the WIC population in the early years of the program. Besharov and Call (2009) argue that the current WIC program is poorly targeted to needy families as the result of program expansions that include families quite far up in the income distribution. Our estimates probably apply most directly to the more disadvantaged among current WIC participants. Second, as noted in Section 6.2, our TOT estimates rely on indirect estimates of the numbers of women who received WIC during its early years. Third, a benefit-cost calculation based only on earnings benefits from increased birth weight would ignore a number of other potential benefits, including improvements in long-run health and education. Despite these limitations our analysis provides the most credible U.S. evidence to date that a targeted nutritional assistance program aimed at pregnant women can produce economically (and statistically) significant improvements in infant health. The birth weight increases that we estimate among early program participants suggest that the value of benefits to needy participants (when appropriately summed over many years of improved earnings and health) can be substantial. In contrast, the cost of providing WIC to pregnant women is relatively low. Further research, using equally credible methods for isolating the causal effect of WIC on children's health, is required to ascertain the full extent of the WIC program's benefits today.

[31] Assuming median earnings in 2008 of $37,544 (Bureau of Labor Statistics).

[32] For example, WIC program information (see http://www.fns.usda.gov/pd/wisummary.htm) estimates the total cost of benefits per year at just over $500 for FY2009.

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Appendix A. WIC county implementation

This paper uses data on the geographic rollout of WIC to identify the program's impacts on birth outcomes. In particular, our regressions use a binary variable equaling one if the county has a WIC program in place in year *t*. The date of the initial implementation of WIC services in each county is not available from a single unified source. As such, we make use of several directories and congressional filings that each provide a list of local agencies that directly provided WIC services (hereafter "local agencies") operating at a particular moment in time. We compile this information into a single database which indicates when WIC services were provided in each county.[33] The following is a list of our primary data sources:

Source	Contents
"Hearings Before the Select Committee on Nutrition and Human Needs – Part 3 – Oversight: WIC program", December 28, 1974.	List of local agencies that had been approved for operation by the USDA as of December 1974.
"Women and Children First or Last — Report on the Special Supplemental Food Program for Women Infants and Children," Virginia Fleming, Children's Foundation, 1975.	List of local agencies approved prior to April 1975, as well as a list of the additional agencies that were approved on April 1, 1975.
"Special Supplemental Food Program for Women, Infant, and Children: WIC Program Directory of Local Agencies," USDA, Food and Nutrition Service, Supplemental Food Programs Division, Washington. 1978.	List of counties served by the WIC program as of March 1978 in the 13 states (Colorado, Illinois, Indiana, Iowa, Kansas, Michigan, Minnesota, Missouri, Nebraska, North Dakota, Ohio, Texas, and Wisconsin) that participated in the Migrant WIC program.
"Directory of Special Supplemental Food Programs for Women, Infants and Children," Children's Foundation, Washington DC, 1979.	List of local agencies operating in the United States in 1979.
"WIC Directory of Special Supplemental Food Programs for Women, Infants, and Children," Center for Budget and Policy Priorities, Washington D.C, 1989	List of local agencies operating in the United States in 1989.
"Evaluation of the Special Supplemental Food Program for Women, Infants and Children, 5-Volumes.," Report to USDA by Rush et al., 1986.	Count of counties served by state and year from 1973 to 1981 for 25 states in the National WIC Evaluation. (Unfortunately, it does not list these counties.)

We assign county-level WIC implementation primarily using the first four sources in the table above. Our first list of WIC agencies comes from a document generated as part of congressional hearings on food and nutrition programs (U.S. Congress, 1974). This provides a list of 216 local agencies; we assign 1974 as the implementation year for these programs.[34] Our second source is a report by the Children's Foundation which provided a list of local agencies established prior to April 1975 as well as a list of 45 additional agencies that were approved by the USDA on April 1, 1975 (Fleming, 1975). We assign 1975 as the implementation year for the (new) agencies that appear in this document. Our third source documents WIC agencies in 1978 in the thirteen states (Colorado, Illinois, Indiana, Iowa, Kansas, Michigan, Minnesota, Missouri, Nebraska, North Dakota, Ohio, Texas, and Wisconsin) that were included in the Migrant WIC Study (USDA, 1978). We assign these programs as "on" in 1978 and "off" in 1975 (but we do not know about their status in 1976 or 1977). Finally, we use the WIC directory published by the Children's Foundation in 1979 to identify new WIC agencies as of 1979 (Children's Foundation, 1979).[35]

Matching WIC agencies to counties

According to national regulations, "local WIC programs may be administered by any health, welfare, or private non-profit agency meeting the statutory requirements" (U.S. Congress, 1981). In practice the majority of early local agencies were health departments or clinics, but a number of welfare and community action programs administered WIC programs and subcontracted out the provision of health services (U.S. Congress, 1981).

Local WIC agencies can be further subdivided by their geographical scope or mandate. During the 1974 to 1979 period, the majority of local agencies served a designated county (e.g. The Jefferson County Health Department) or group of counties (e.g. Tri-County District Health Service: Morgan, Limestone and Lawrence Counties). Occasionally district or regional agencies were listed without an explicit list of member counties. Under these circumstances the county membership is assumed to be the same as the membership for said regional agency in the closest year for which we have membership information. (For example, the North Health District in Georgia was listed without member counties in 1974. Member counties were listed for this agency in 1975, so we use this 1975 list in our 1974 data.) In a handful of occasions, regional membership was not given in any of the directories. In this case, we attempt to establish historical membership using online sources.

Local agencies not specifically affiliated with a particular county or group of counties, were often affiliated with a particular city (e.g. The Oakland Children's Hospital) or with an independent welfare agency operating in some regional area (e.g. The Lewis & Clark Children & Youth Project, Helena). In these cases, state and city information was used to determine in which county the local agency was located.[36] Such agencies were assumed to cover the county in which they were located. Local agencies affiliated with Native American tribes or military bases were assumed to cover the county which contained the appropriate reservation or base.

Main sample

In our empirical work, we limit our analysis to the set of counties that have a local WIC agency in place by 1979 and include 2059 of the 3100 counties. This sample accounts for more than 86% of the U.S. population in 1970 (and 85% of births in 1970). We construct the sample in this way because our next source for WIC agencies is ten years later, in 1989.

Other details

We assume that once a WIC agency is open, it stays open for the rest of our sample period. Our sources sometimes indicate that an

[33] We have had numerous conversations with WIC program administrators at the USDA. It appears that this information was not tracked by USDA. Thus, we relied on the congressional reports and WIC program directories described in the text to construct our measure.

[34] The vast majority (at least 207) of these agencies were scheduled to open by or before April 1974, the date of the Committee report (U.S. Congress, 1974, page 345).

[35] There were many references to Children's Foundation Directories for other years. Unfortunately, we were not able to locate these documents.

[36] This was generally done by typing the city and state names and the word "county" into Google and scanning the resultant government, Wikipedia, or city-data.com listings. At times, this process was automated using a web scraping tool and the website http://www.townsusa.org.

agency covers only part of the county; possibilities include serving a subset of the population (military, Native Americans, migrants) or some other unspecified partial coverage. Given that the number of counties indicating partial coverage is relatively small, we choose here to assign these counties to have a program in place.

We do not have any information on WIC agencies in 1976 or 1977. Hence we drop births in the natality data in 1976 and 1977. In 1978, our source is limited to the 13 states participating in the WIC migrant study. If a county is in one of the 37 other states, and we have already observed a WIC agency from our earlier sources, we keep that county

Number of counties by WIC implementation status[a]

Year	No WIC program in place	WIC program in place	Missing[b]	Never observed by 1979	Total
1974	1690	349	0	1039	3078
1975	1491	578	0	1038	3077
1978	154	1038	847	1039	3078
1979	0	2059	0	1041	3100

[a] The number of counties varies slightly from one year to the next because of missing data for either the Census or REIS data that is merged with our Natality sample. This occurs only for very small counties accounting for a very small number of births. The results are unaffected by their exclusion. Balancing the panel of counties also has no impact on the results.

[b] As noted in the text, a county is missing in 1978 if it is not included in the 1978 data source and there is no WIC agency that appears in that county in either of the earlier years (1974, 1975).

in the sample. However, for the remaining counties in the 37 states not included in the 1978 data, we set the WIC implementation variable to missing.

The table above summarizes our WIC implementation variable. In particular, we tabulate the number of counties by WIC implementation status for the four years (1974, 1975, 1978, and 1979) when we observe WIC agencies.

For all counties in all years prior to 1974, we set the WIC policy variable to 0. For all counties in all years 1979 and later, we set the WIC policy variable to 1.

All coverage data was compared against secondary sources including the National WIC Evaluation (Rush et al., 1986) and, where available, individual state filings. The National WIC Evaluation includes a tabulation of the number of counties served, by year and state, from 1973 to 1981 for the 25 states includes in the study.[37] We compare county counts from these tables with those from our dataset and find them to be generally consistent. Further, in extensions to our main models, we use a state-year identification strategy and compare results from our WIC policy variable to those constructed using the National WIC Evaluation study. We find very similar results for the two sets of policy variables, which give us additional confidence in our archival work.

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[37] The states in the Rush et al. (1986) table include Arizona, Colorado, Connecticut, DC, Delaware, Florida, Georgia, Idaho, Illinois, Iowa, Kansas, Kentucky, Louisiana, Maine, Michigan, Mississippi, Missouri, New York, North Carolina, Oregon, Pennsylvania, Rhode Island, South Dakota, Utah and Washington.

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