Appendix 2. Additional Results

2A. Medicaid Eligibility and Utilization

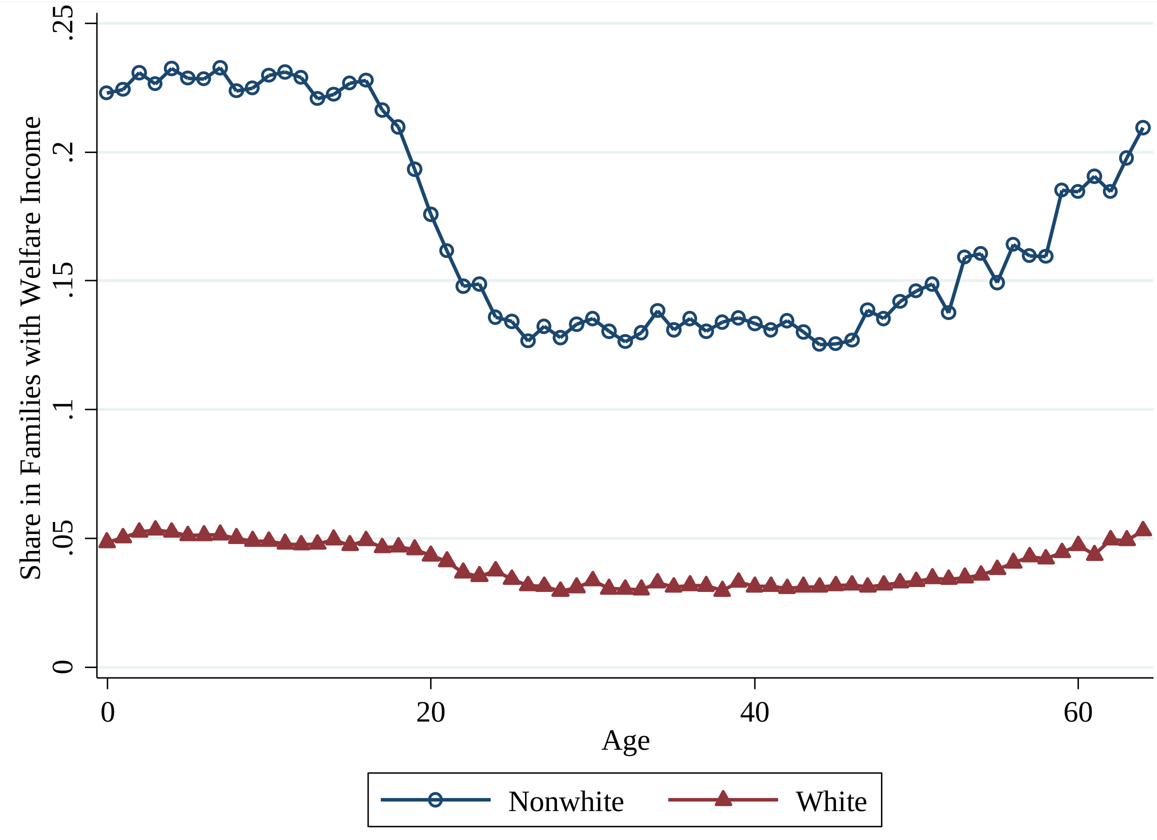
Appendix Table 2.A1. State Welfare Rates in the Year of Medicaid Implementation

	AFDC Rate in the Year of Medicaid Implementation		Medicaid Implementation	
State	Nonwhite Women	White Women	Month	Year
Alabama	12.4	1.2	1	1970
Arkansas	11.9	1.0	1	1970
California	11.9	2.3	3	1966
Colorado	15.4	2.4	1	1969
Connecticut	14.2	1.0	7	1966
Delaware	11.5	0.7	10	1966
District of Columbia	3.8	0.1	7	1968
Florida	17.6	0.9	1	1970
Georgia	6.0	0.8	10	1967
Idaho	7.0	1.5	7	1966
Illinois	11.1	0.5	1	1966
Indiana	13.2	0.9	1	1970
Iowa	15.8	1.5	7	1967
Kansas	12.3	1.0	6	1967
Kentucky	9.8	2.1	7	1966
Louisiana	7.5	0.9	7	1966
Maine	6.1	2.0	7	1966
Maryland	9.9	0.7	7	1966
Massachusetts	12.9	1.5	9	1966
Michigan	8.8	0.9	10	1966
Minnesota	16.3	1.4	1	1966
Mississippi	15.8	0.9	1	1970
Missouri	11.9	1.1	10	1967
Montana	15.5	0.9	7	1967
Nebraska	14.5	0.9	7	1966
Nevada	13.0	0.7	7	1967
New Hampshire	3.6	0.7	7	1967
New Jersey	22.6	2.1	1	1970
New Mexico	7.4	2.8	12	1966
New York	11.5	1.6	5	1966
North Carolina	9.0	0.9	1	1970
North Dakota	18.5	1.0	1	1966
Ohio	10.1	0.8	7	1966
Oklahoma	16.2	1.7	1	1966
Oregon	11.0	1.7	7	1967
Pennsylvania	10.1	1.0	1	1966
Rhode Island	19.2	2.1	7	1966

South Carolina	3.5	0.4	7	1968
South Dakota	25.5	1.2	7	1967
Tennessee	10.9	1.4	10	1969
Texas	2.9	0.6	9	1967
Utah	13.8	2.1	7	1966
Vermont	0.4	1.4	7	1966
Virginia	5.9	0.6	7	1969
Washington	7.5	1.6	7	1966
West Virginia	11.0	4.4	7	1966
Wisconsin	11.1	0.8	7	1966
Wyoming	6.8	1.3	7	1967

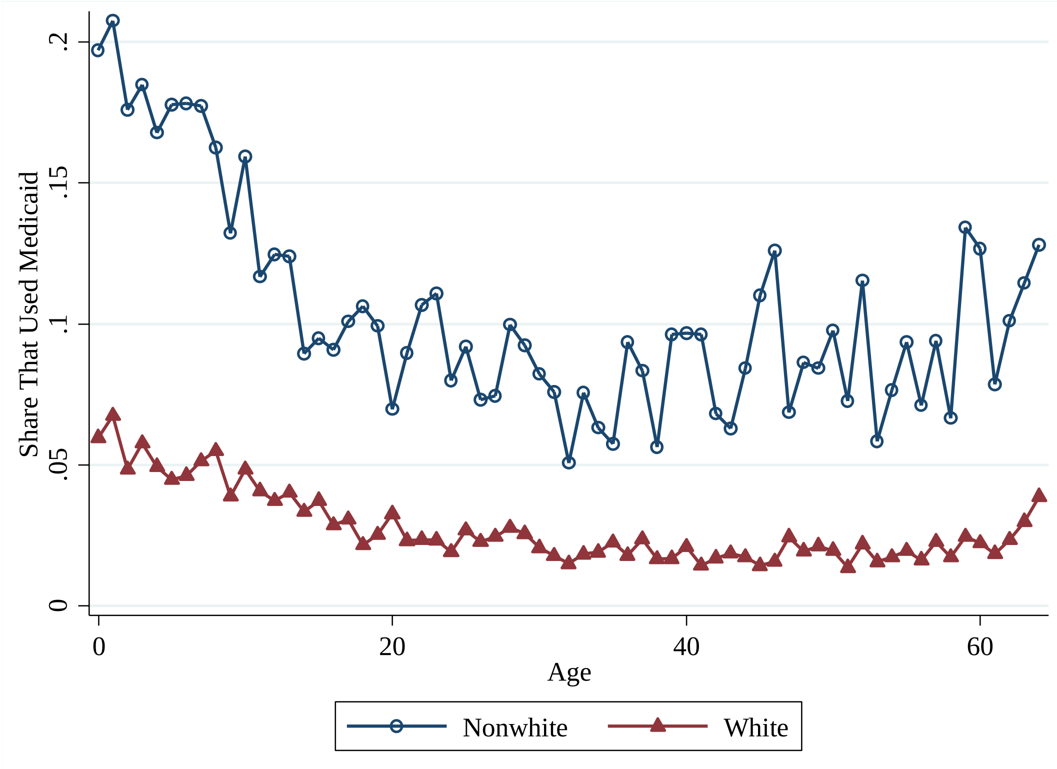
Notes: Race-specific AFDC rates are calculated as described in text and in appendix 1. AFDC rates for women are per woman ages 15-54. Medicaid implementation dates are widely available, including in DHEW (1970).

Appendix Figure 2.A1. Age-Specific Rates of Welfare Receipt in the 1970 Census



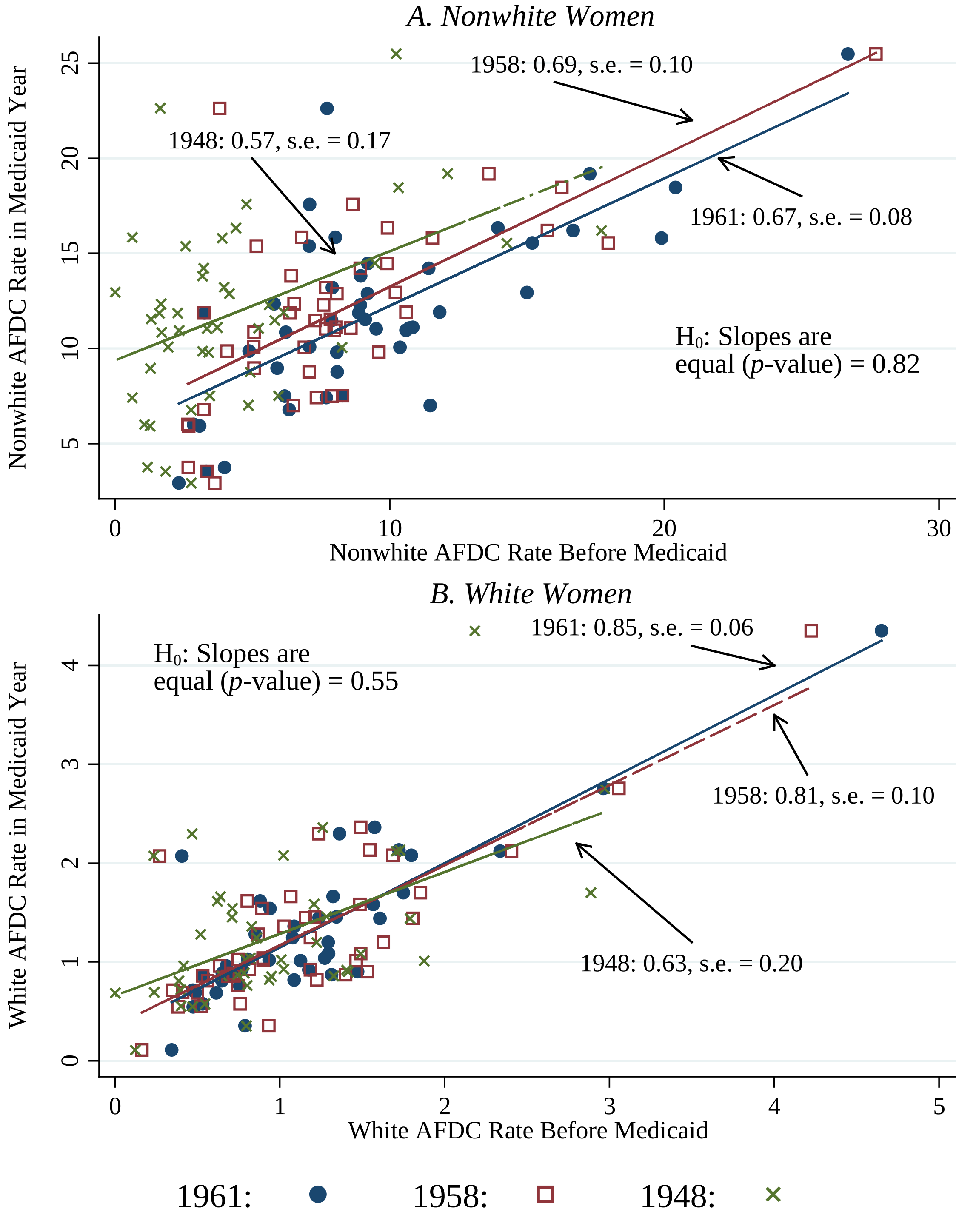
Notes: Data from the 1970 Census of Population State Sample Forms 1 and 2. The figures plots the share of respondents who lived in a household where at least one person reported positive welfare income. The average welfare receipt is higher than in figure 2 because the Census question is not restricted to AFDC. This increases the adult welfare rate the most, but it does not necessarily mean that their Medicaid eligibility rates were higher because this includes General Assistance, a state program not included in the definition of categorical (Medicaid) eligibility. Source: Ruggles et al. (2010)

Appendix Figure 2.A2. Age-Specific Rates of Medicaid Receipt in the 1976 Survey of Income and Education



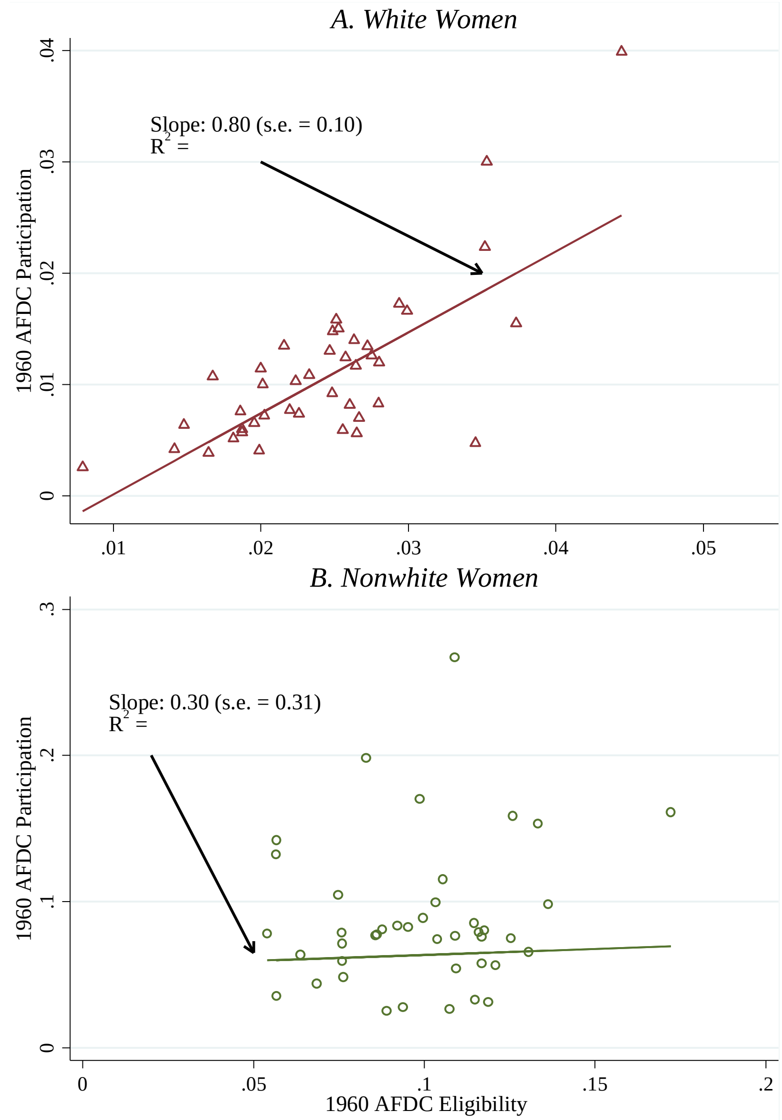
Notes: The figures plots the share of respondents who report *using* Medicaid in the previous year. 3,819 observations (out of 440,815; 0.87%) are missing and are dropped from the calculation. Source: 1976 Survey of Income and Education (US Department of Commerce 2006)

Appendix Figure 2.A3. Stability in Cross-State AFDC Variation, 1948 through Medicaid Implementation



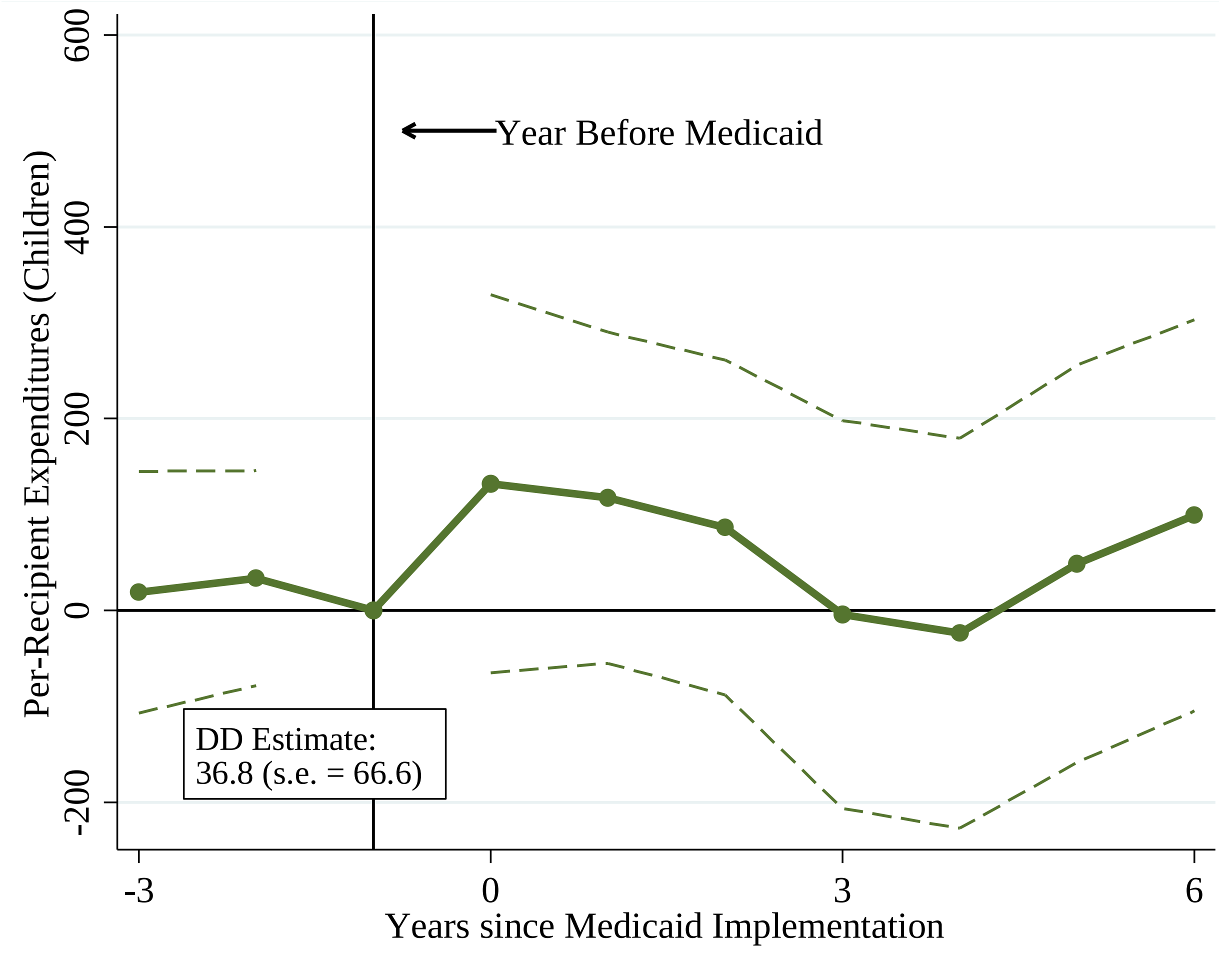
Notes: The figure presents scatter plots and fitted values of the relationship between the paper's primary measure of categorical eligibility—the AFDC rate in the year of Medicaid implementation (y-axis)—and three measures of AFDC rates in years *prior* to each state's Medicaid year. The results show that the cross-state variation in AFDC was very stable over time. For both white and nonwhite women, pre-Medicaid AFDC rates strongly predict AFDC rates in the year of Medicaid and the relationship itself does not change over time. *p*-values from a test that the slopes are equal using a robust regression to minimize the influence of outliers (Berk 1990) are 0.34 and 0.32.

Appendix Figure 2.A4. The Relationship between AFDC Eligibility and AFDC Participation in 1960



Notes: The figure shows the relationship between state-level measures of AFDC eligibility and observed AFDC participation. I use the 1960 Census and a table of AFDC "needs standards" (one of several income eligibility thresholds) from 1961 (DHEW 1963; table 40) to calculate the share of women between 20 and 64 who are unmarried family heads, with at least one qualifying child (under 16 or under 18 and attending school), and monthly "countable income" (earnings minus "other" income and income of qualifying children) below the average family-size specific needs threshold in her state. This calculation ignores eligibility criteria such as coverage of unborn children, asset tests, the "payment test" (which compares adjusted income to a lower payment threshold). More importantly, I cannot account for more subjective eligibility criteria, such as requirements that heads accept work, "man-in-the-house" or "suitable home" provisions, or caseworker practices such as underbudgeting (requiring additional paperwork to increase recipients' grants after the birth of a child, deducting child support amounts regardless of whether or not the support order was paid; Piven and Cloward 1971). The importance of these criteria for actual categorical Medicaid eligibility strengthens the identification strategy based on observed AFDC rates because, as this figure shows, they lead to nonwhite AFDC rates that are orthogonal to the factors that determine eligibility. Source: 1960 Census (Ruggles et al. 2010).

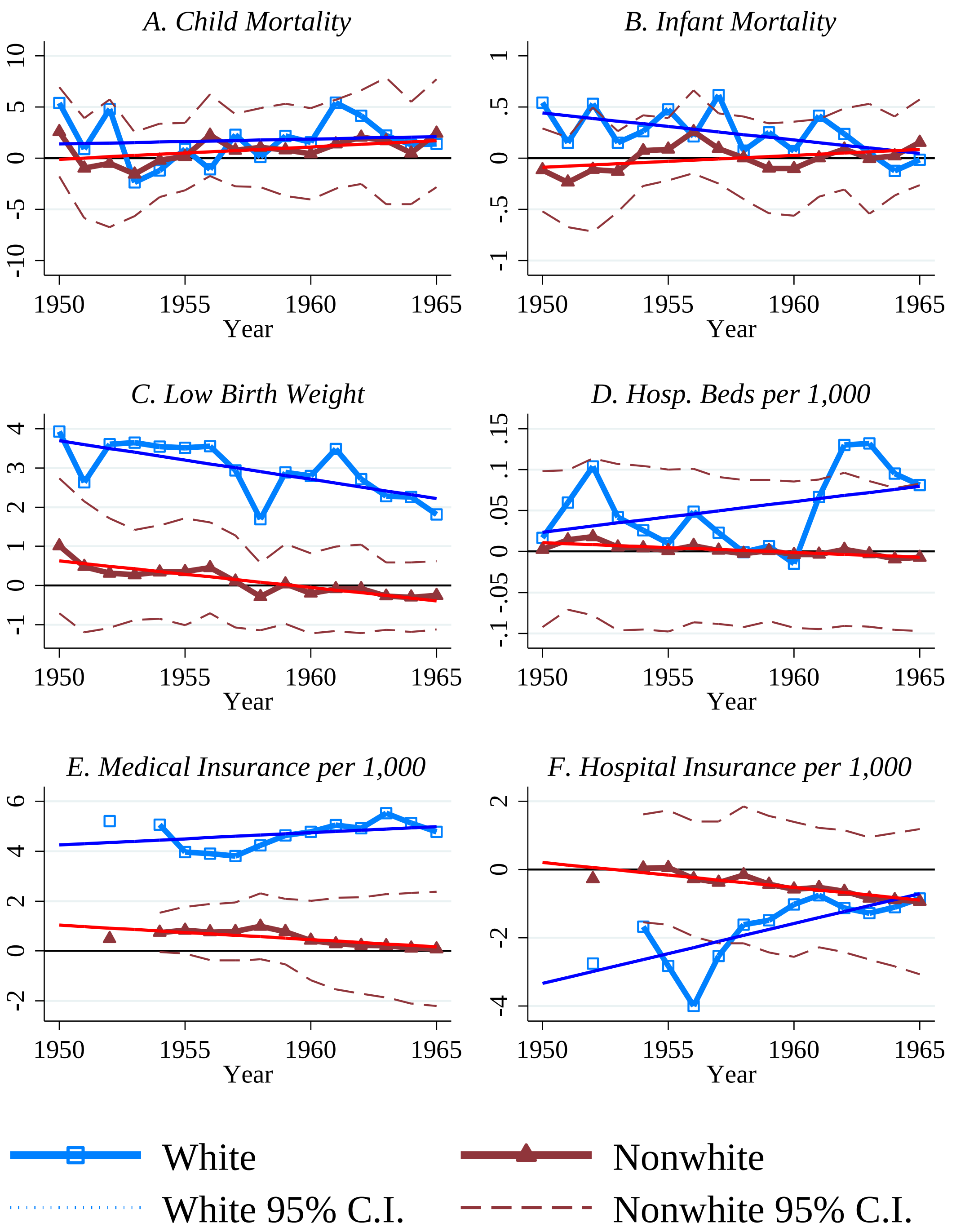
Appendix Figure 2.A5. Medicaid's Effect on Public Insurance Expenditures Per Child Recipient



Notes: The figure plots event-study coefficients on the interaction between $AFDC_s^*$ and Medicaid event-time dummies. The dependent variable is the ratio of total public insurance spending on children divided by the number of child recipients, expressed in 2012 dollars. The variable is set to zero in cells with zero child recipients. The figure shows that the generosity (or intensity of utilization per recipient) did not vary with Medicaid eligibility, either before or after Medicaid implementation.

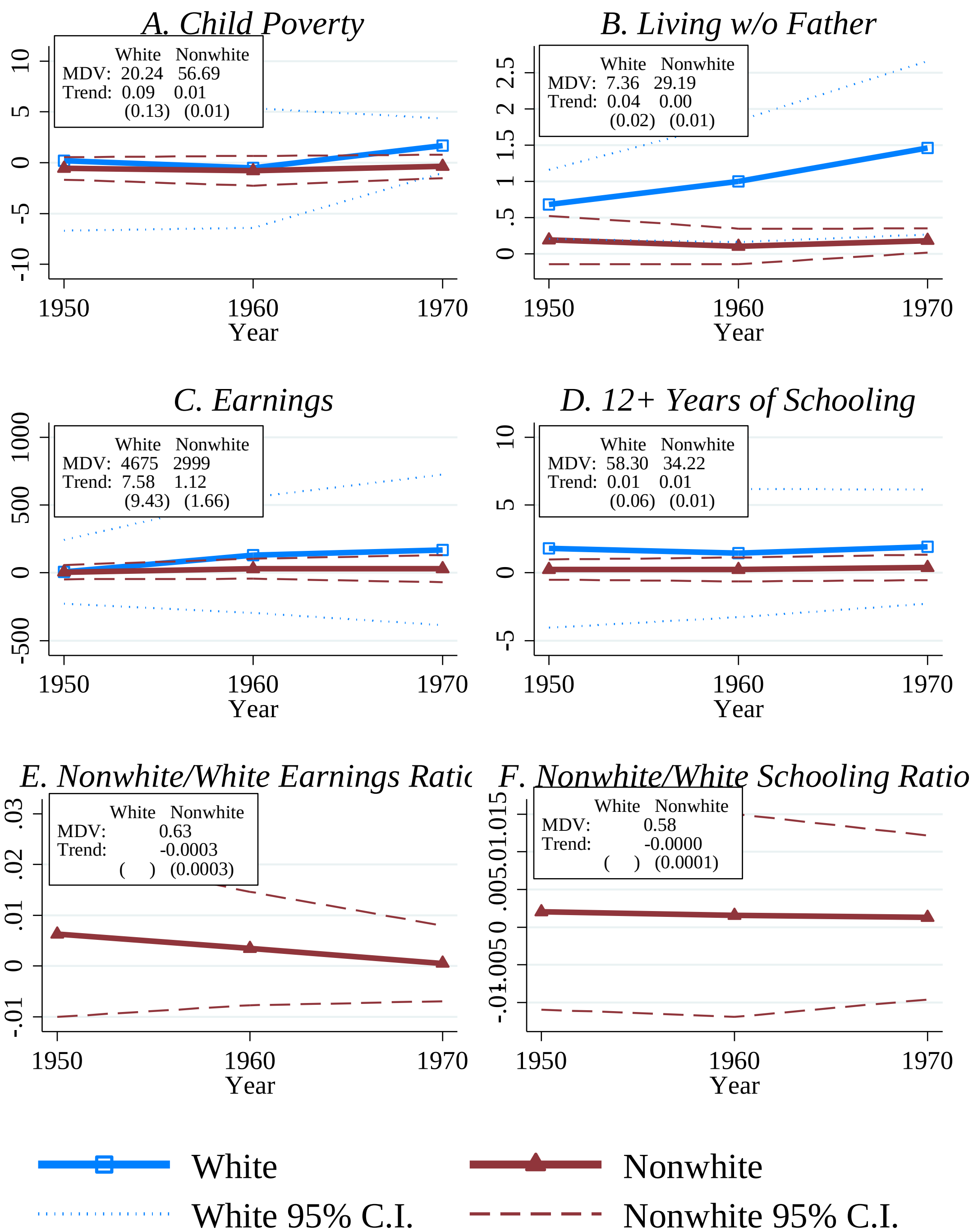
2B. Additional Support for the Research Design

Appendix Figure 2.B1. Annual Pre-Medicaid Relationship between $AFDC_s^*$ and State Characteristics



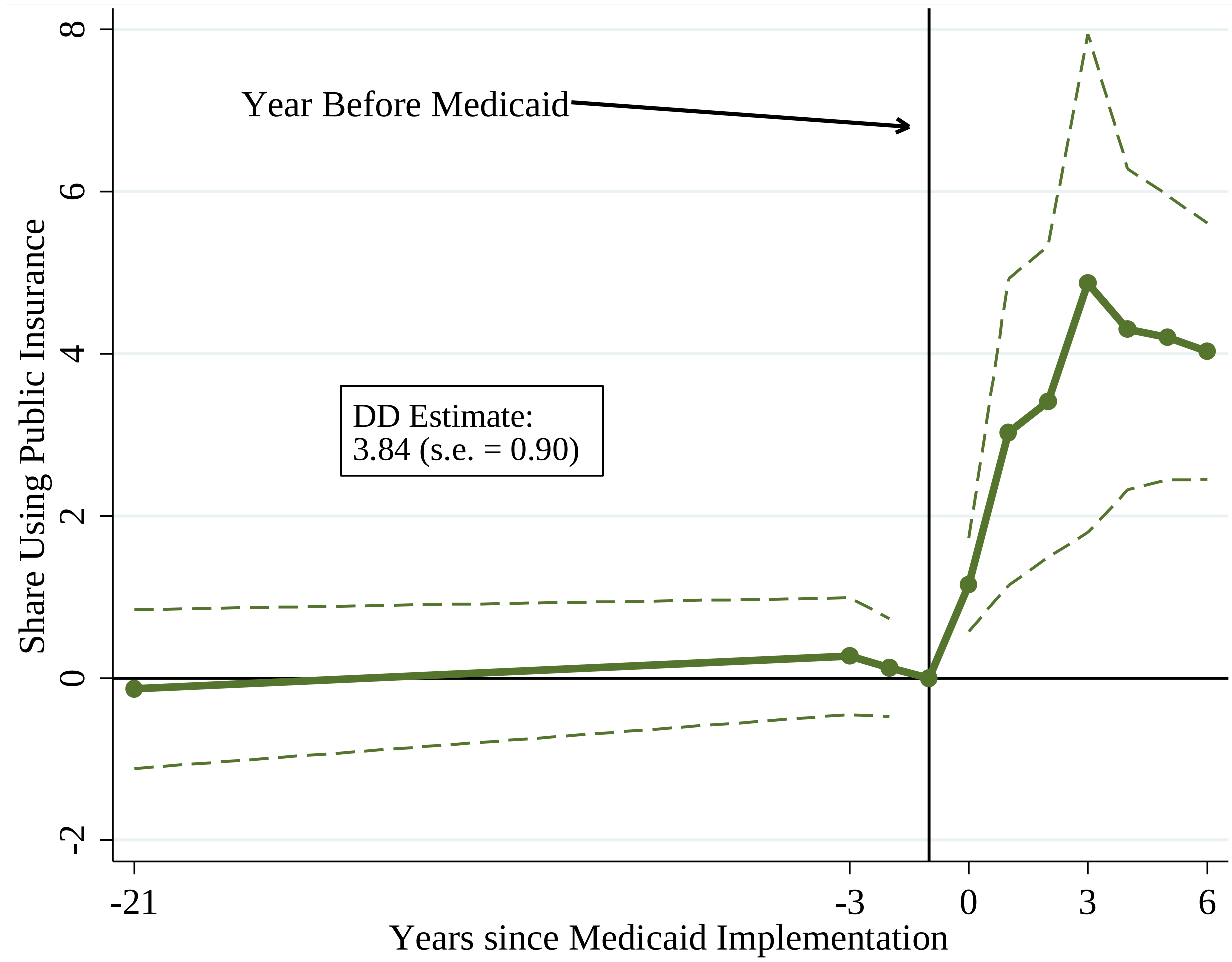
Notes: The figure plots estimated coefficients on interactions between year dummies and $AFDC_s^*$ from regressions that use the characteristic in each panel as an outcome variable. There is no omitted interaction so each coefficient is the annual regression slope between the outcome and $AFDC_s^*$. The linear trend from table 1 is superimposed in each figure. White confidence intervals are very wide and omitted for readability. The results support the conclusion of table 1 that characteristics did not change differently before Medicaid in a way that was correlated with $AFDC_s^*$. They also show that the linearity restriction in table 1 is generally reasonable.

Appendix Figure 2.B2. Relationship between $AFDC_s^*$ and Civil Rights Outcomes: Earnings and Education Outcomes by Race, 1950-1970



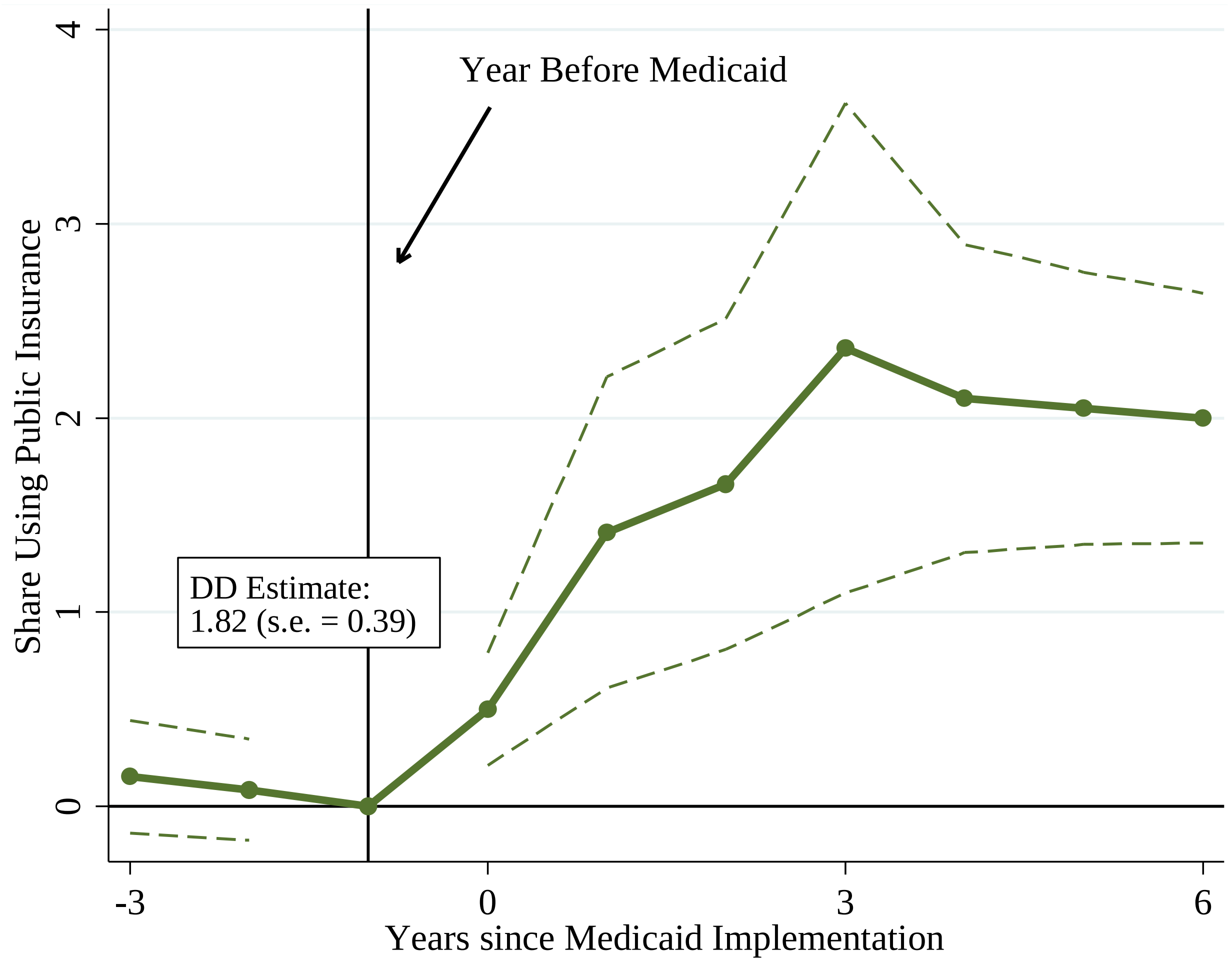
Notes: For details on the specification see notes to figure 2.B1. The figure tests whether there were differential changes in socioeconomic outcomes before Medicaid (1950-1960) or during the overlapping roll out of Medicaid and the Civil Rights era (1960-1970). The results support the conclusion of panel B of table 1, that trends in socioeconomic characteristics before Medicaid were uncorrelated with $AFDC_s^*$, and they show that the rapid changes that occurred *during* the 1960s were also uncorrelated with $AFDC_s^*$.

Appendix Figure 2.B3. First-Stage Estimates Adding Zero Participation before 1950



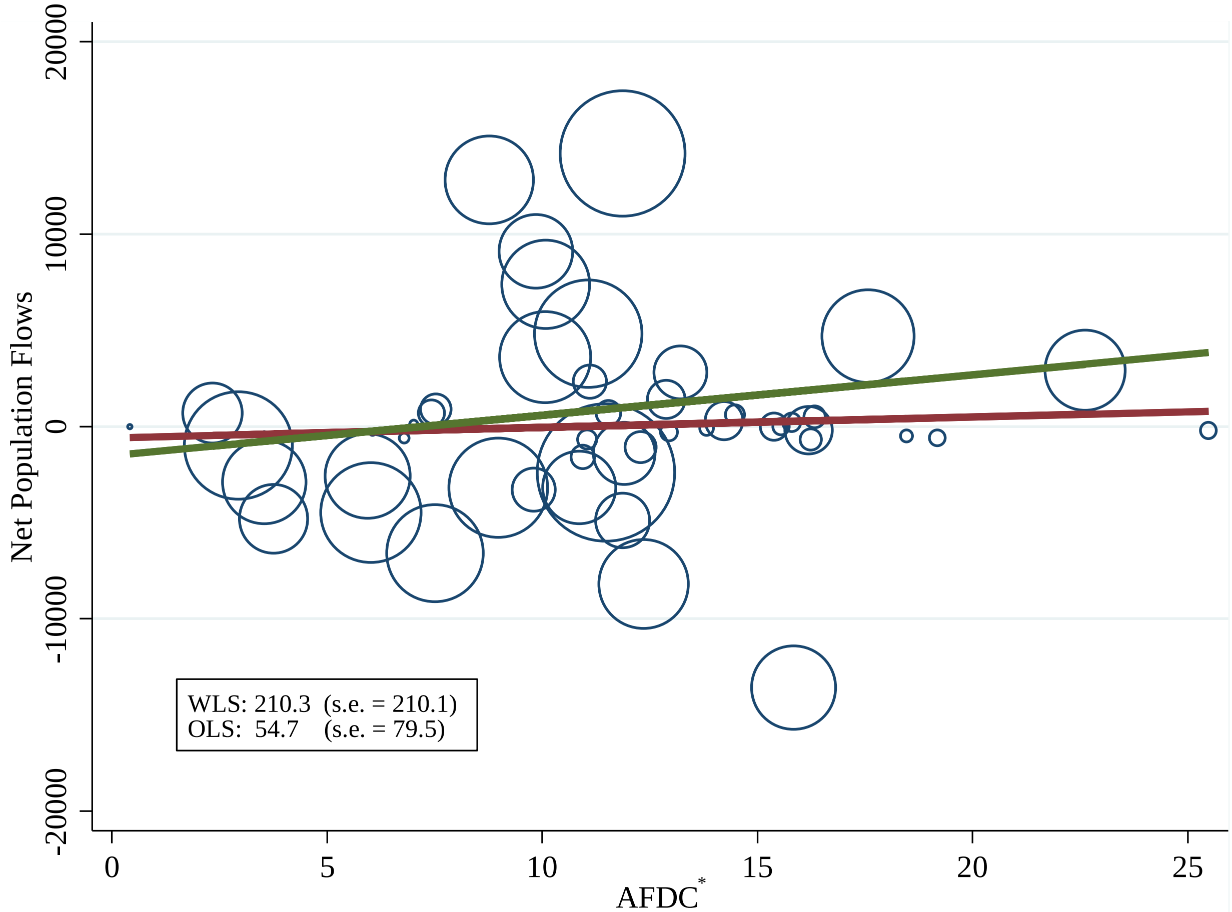
Notes: This figure plots event-study estimates comparable to figure 5 but imposing zero public medical care receipt in event-time -21. This reflects the fact that specific federal reimbursement for public assistance medical care was not available until 1950 (event-time -21 for 1970 Medicaid states), and so it can be assumed to equal zero before then. These first-stage estimates are nearly identical to those presented in figure 5 (3.96 vs. 3.98).

Appendix Figure 2.B4. First-Stage Estimates Using Child AFDC Rates



Notes: This figure plots estimates comparable to figure 5 that use the *child* AFDC rate to calculate $AFDC_s^*$ instead of the rate among women 15-54. Child AFDC rates are about twice as big as rates among women, so the point estimate is about half of that in figure 5 (1.92/3.98 = 0.48). Reduced-form mortality effects that use the child AFDC rate are in figure 2.C5.

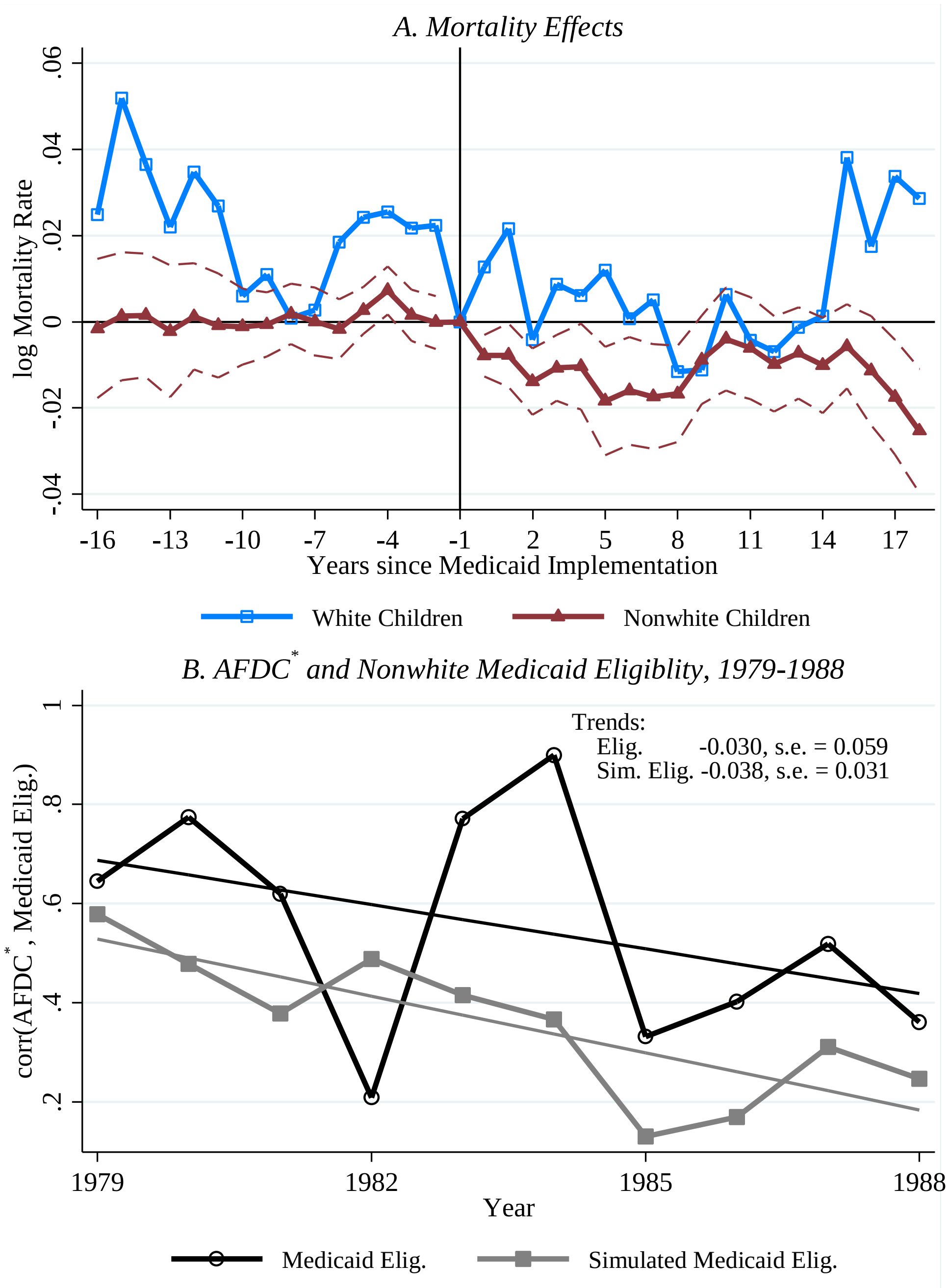
Appendix Figure 2.B5. The Relationship between Net Population Flows (1965 and 1970) and Initial AFDC Rates, Nonwhite Families with Children



Notes: Data include nonwhite families with children from the 1970 Census, form 2 state sample (Ruggles et al. 2010). The net population flow for a state is the weighted count of respondents who moved to that state between 1965 and 1970 minus the count of respondents who left that state between 1965 and 1970. The figure plots each state's net population flow against its value of $AFDC_s^*$ and includes univariate regression slopes with and without weighting by the total (weighted) count of respondents living in each state in 1965. The figure shows that population movements during Medicaid implementation were uncorrelated with AFDC-based Medicaid eligibility.

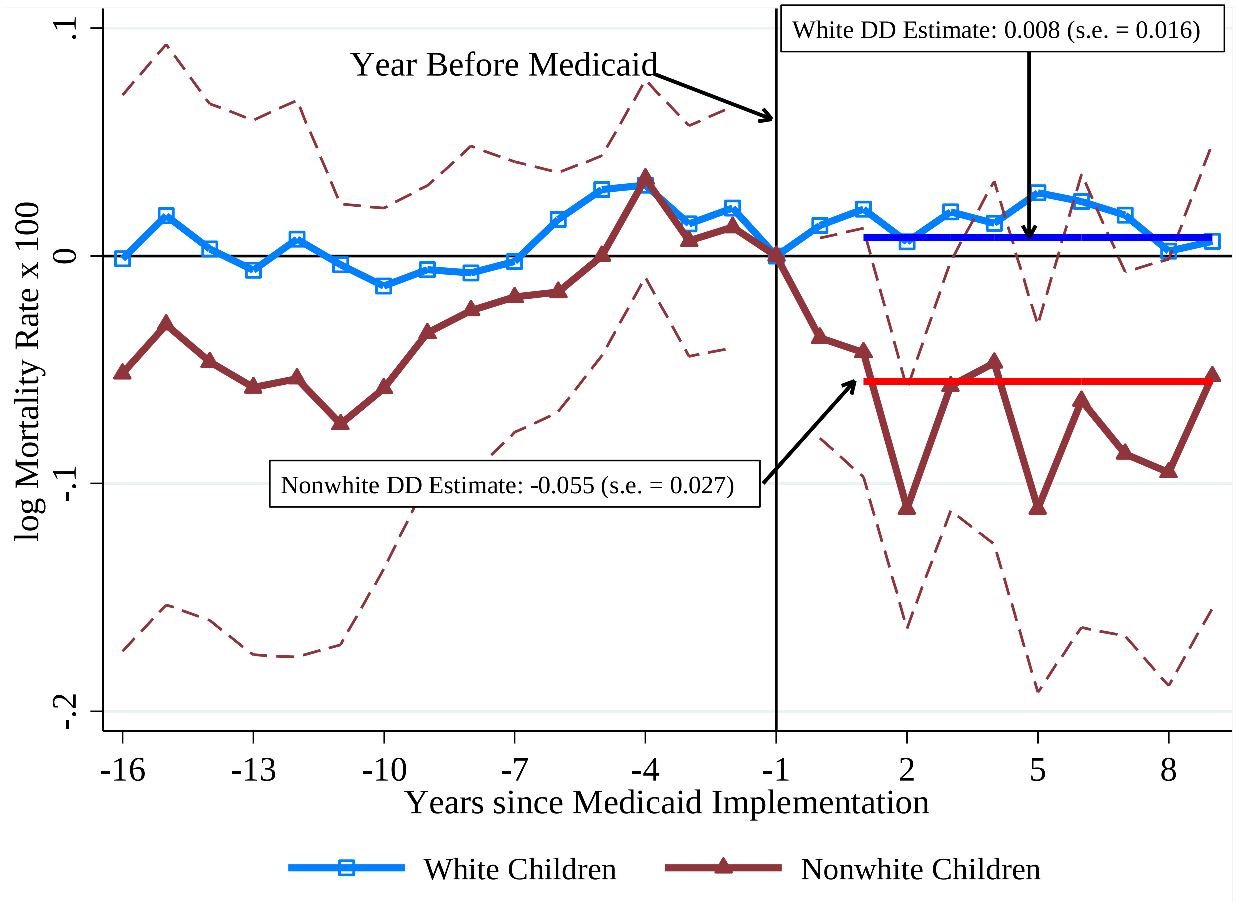
2C. Additional Mortality Event-Study Results

Appendix Figure 2.C1. Regression-Adjusted Estimates of Medicaid's Intention-to-Treat Effect on Child Mortality by Race (1950-1988) and the Relationship between $AFDC_s^*$ and Medicaid Eligibility Outside the Main Sample Period (1979-1988)



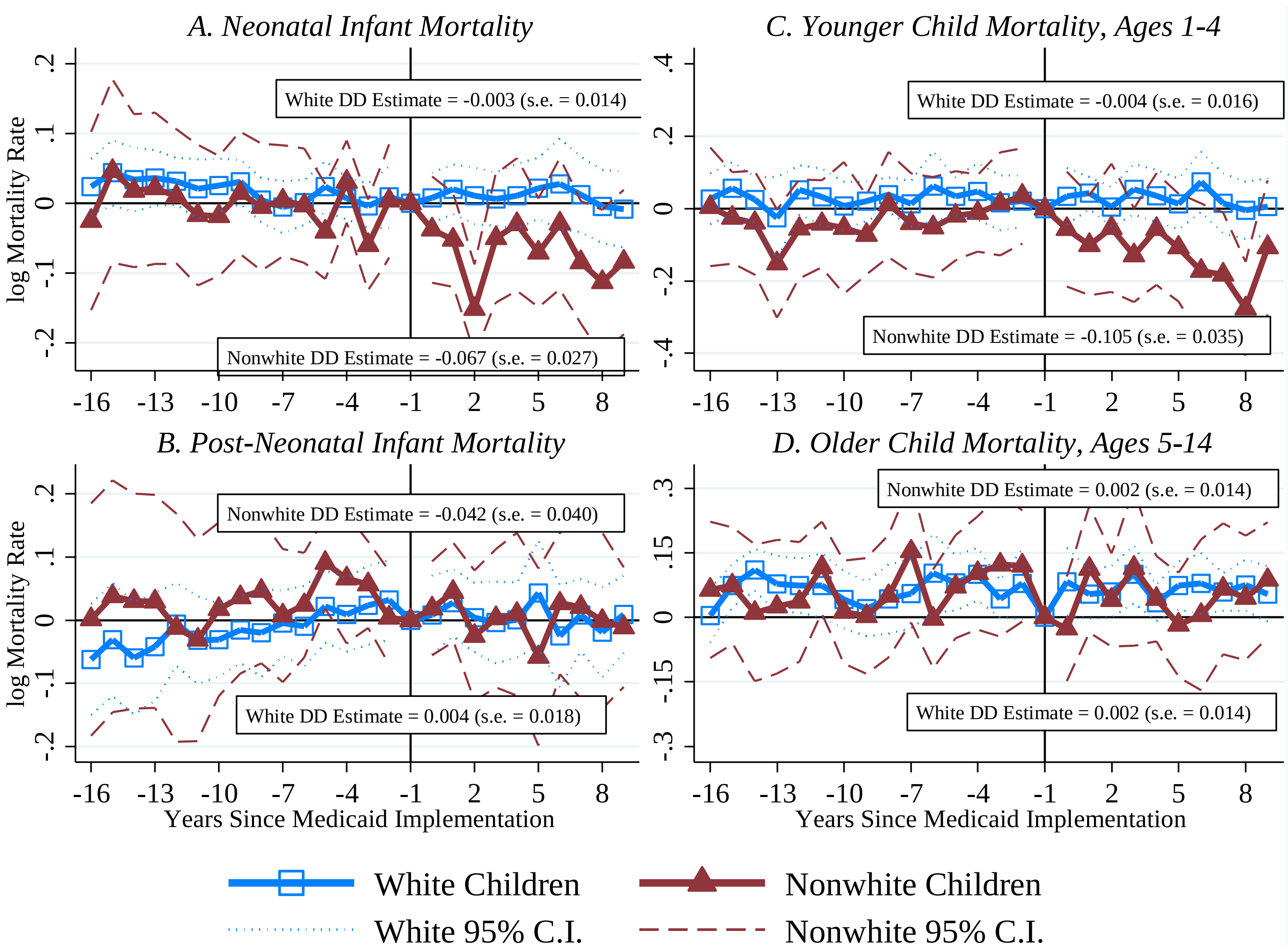
Notes: See notes to figure 6. The sample extends through 1988 and so the estimated treatment effects cover event years -16 through 18. The second panel uses data on Medicaid eligibility and simulated Medicaid eligibility for infants, children 1-4 and children 5-14 from 1979-1988 to demonstrate that the 1980's expansions worked to eliminate the cross-state variation that came from initial AFDC rates. The figure plots annual correlations between $AFDC_s^*$ and the eligibility variables as well as linear trends. I thank Laura Wherry for sharing the eligibility data. The figure shows that the reduced-form mortality effects based on $AFDC_s^*$ shrink at the same time that the first-stage relationship between $AFDC_s^*$ and eligibility measures falls.

Appendix Figure 2.C2. Regression-Adjusted Estimates of Medicaid's Effect on Age-Adjusted Child Mortality Rates by Race, Binary Specification



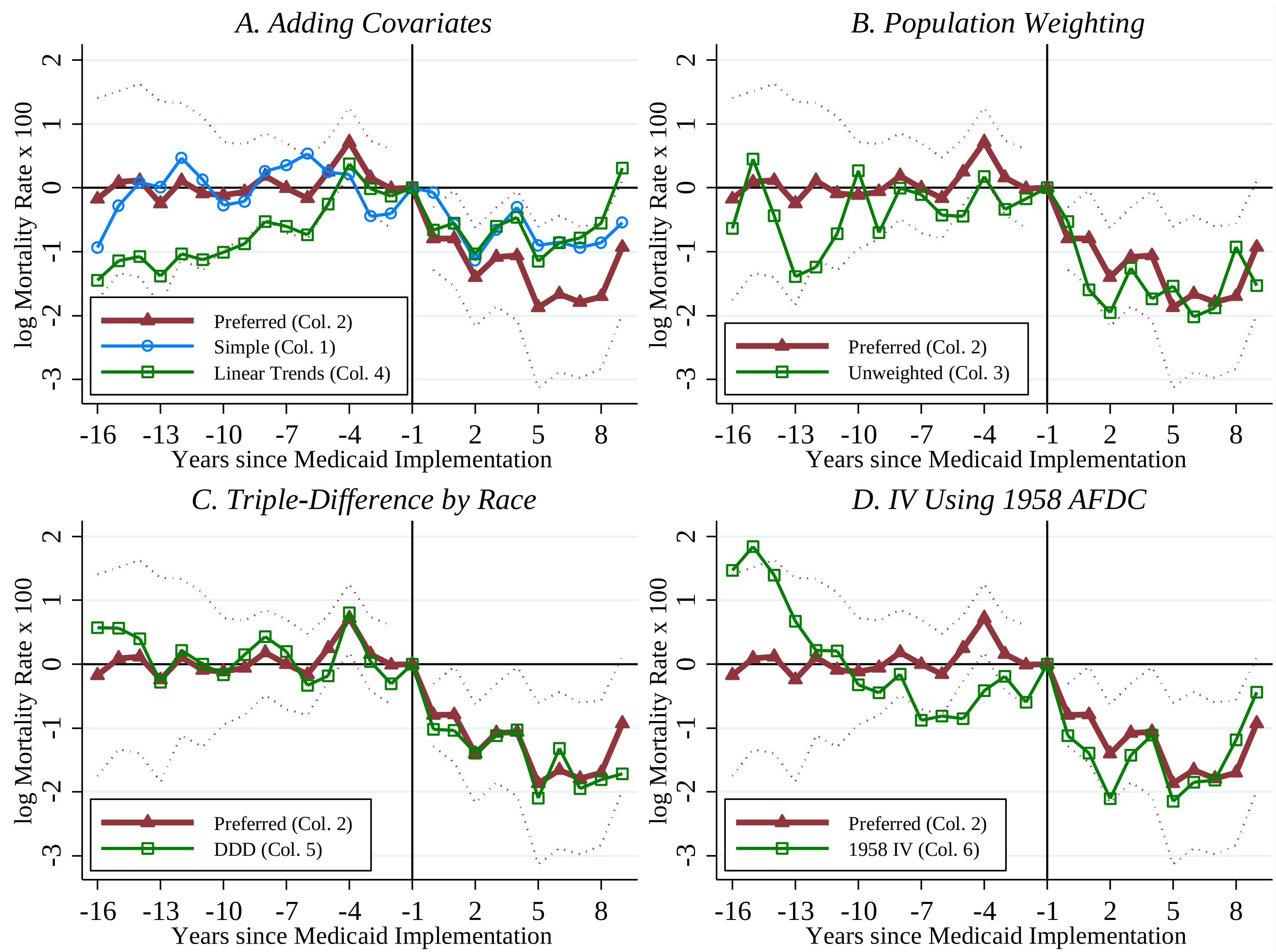
Notes: The figure plots reduced-form treatment effects equivalent to figure 6 that replace $AFDC_s^*$ with a dummy equal to one for states with $AFDC_s^*$ greater than the (race-specific) median value.

Appendix Figure 2.C3. Regression-Adjusted Estimates of Medicaid's Effect on Mortality Rates by Race and Age, Binary Specification



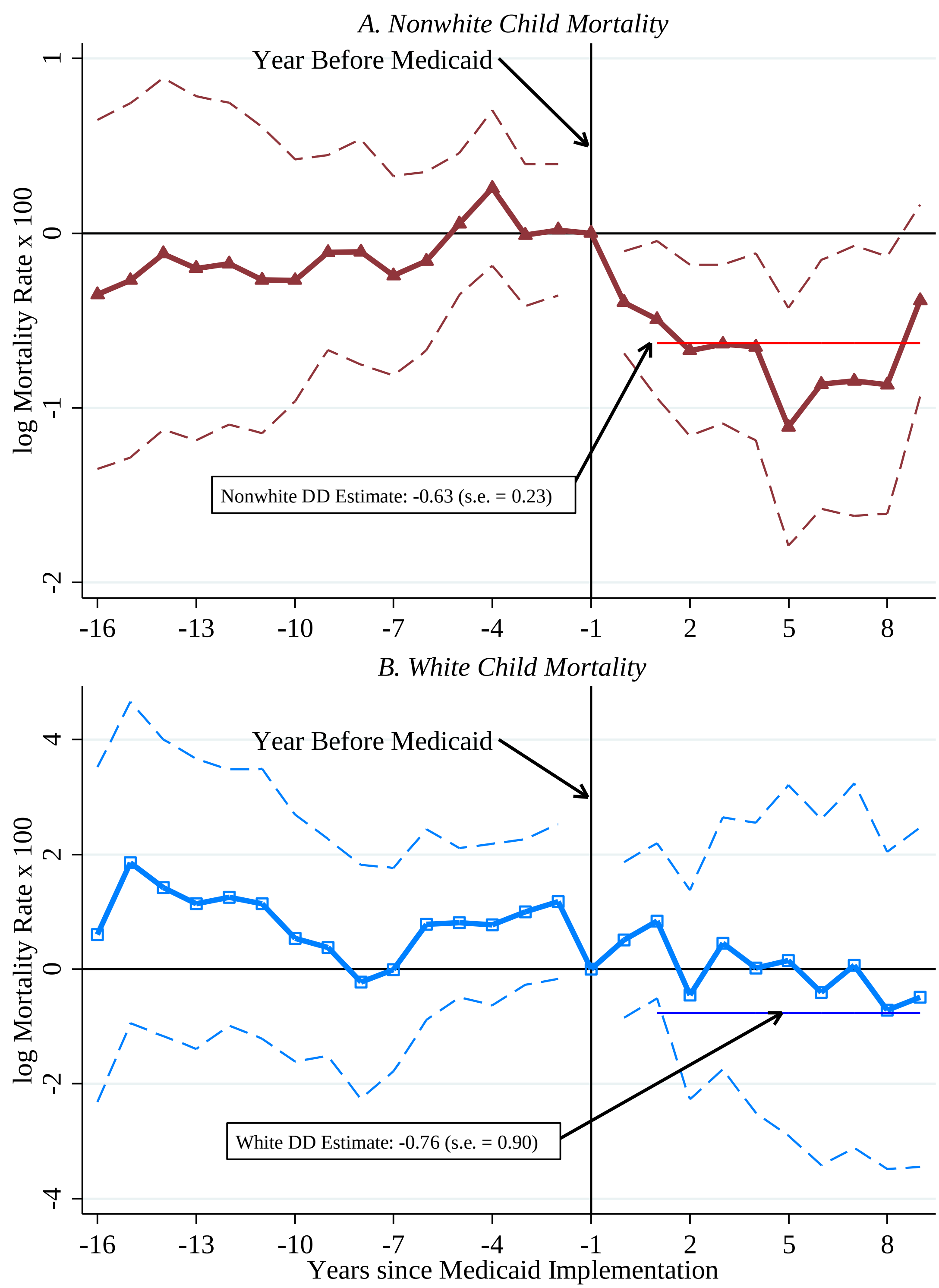
Notes: See notes to figure 6 and appendix figure 2.C2. Each panel plots event-study estimates for age-group-specific mortality by race.

Appendix Figure 2.C4. Estimates of Medicaid's Effect on Age-Adjusted Nonwhite Mortality Rates by Specification



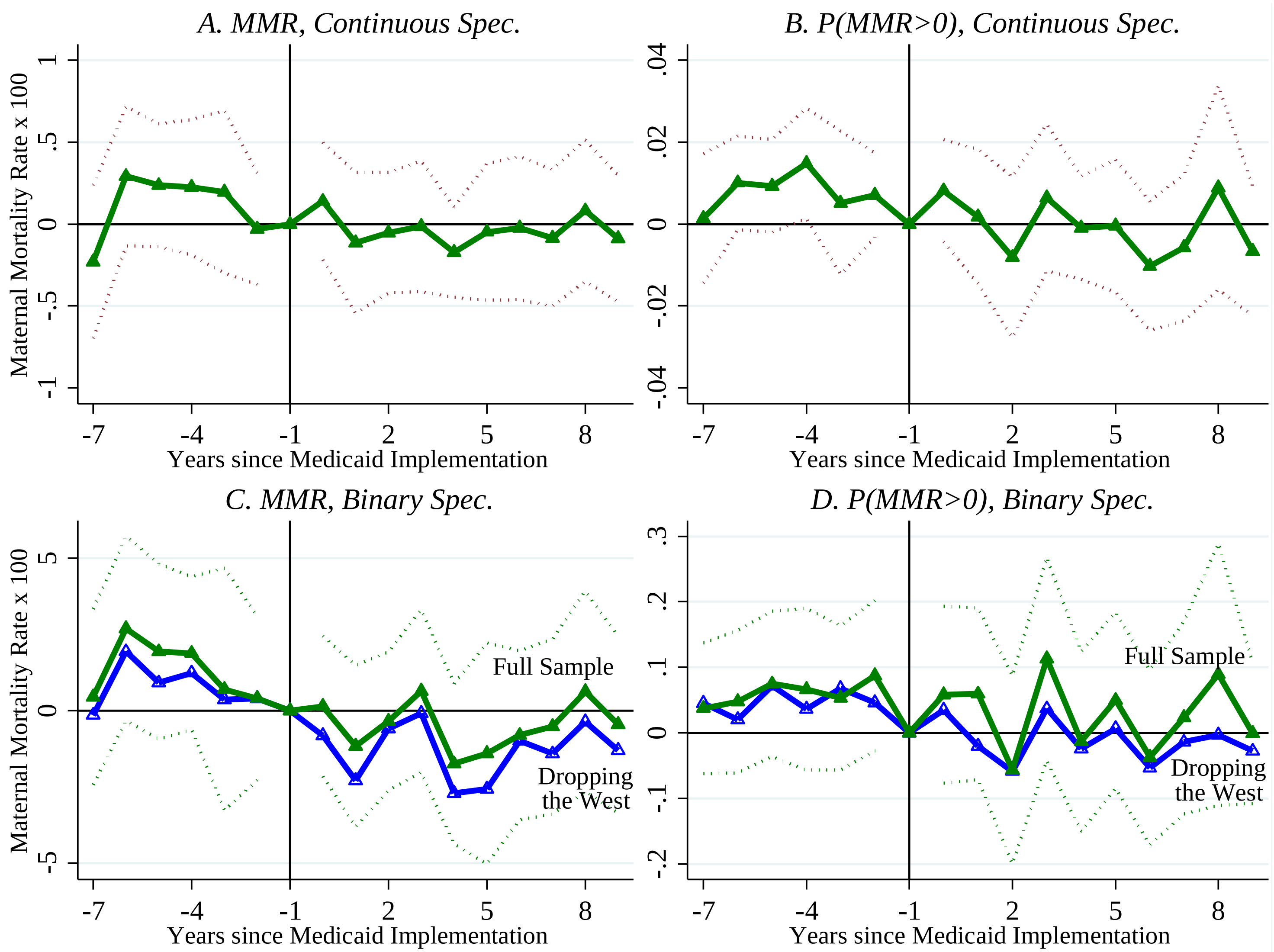
Notes: The bold line in each panel is reproduced from panel A of figure 6. The other specifications correspond to the columns of table 3.

Appendix Figure 2.C5. Estimates of Medicaid's Effect on Age-Adjusted Nonwhite Child Mortality Rates using Child AFDC Rates



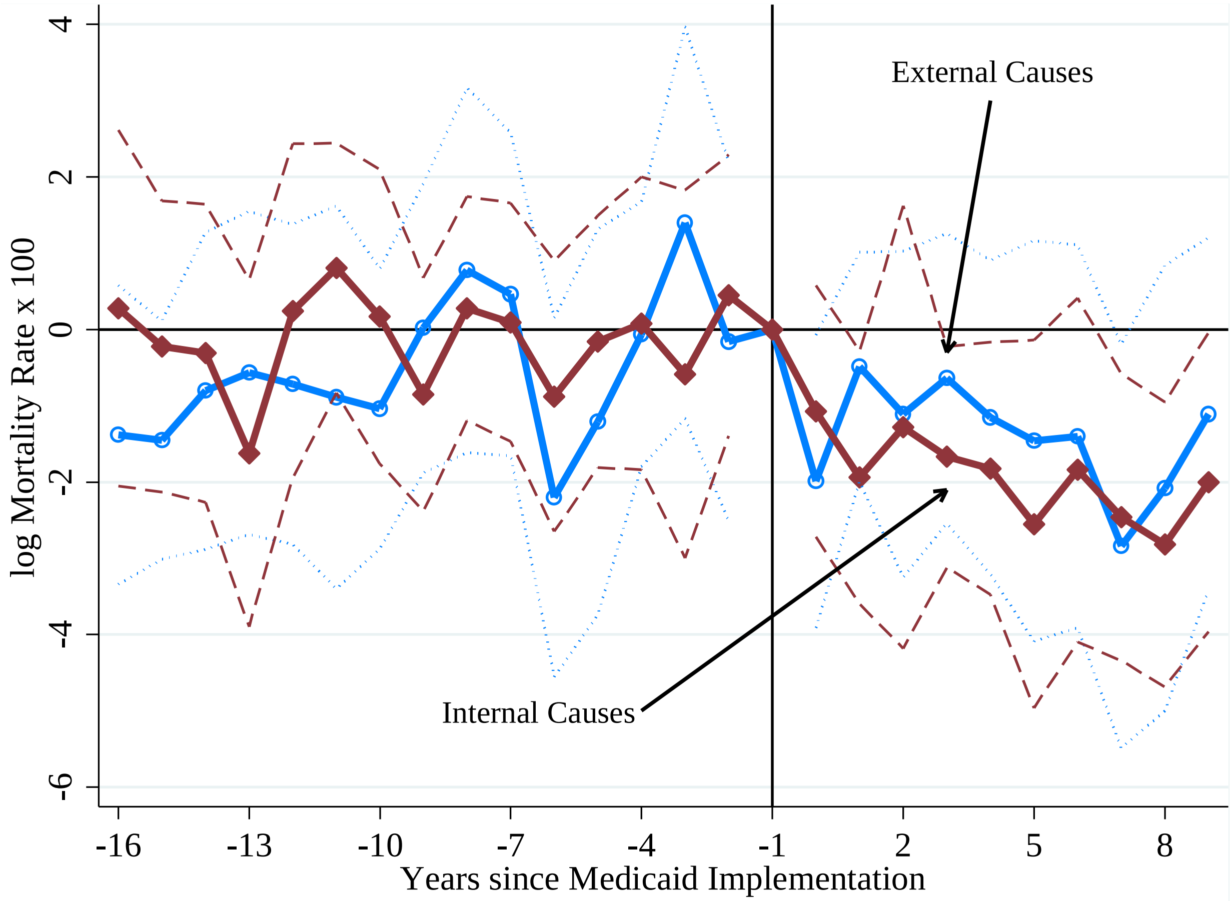
Notes: The figure plots event-study estimates comparable to figure 6, but using the child AFDC rate to calculate $AFDC_s^*$ (see appendix figure 2.B4). These effects imply a nearly identical ATET to those presented in the paper: -18.2 percent (-0.63/1.92/1.8, see appendix 4) versus -19.7 percent (see figure 10).

Appendix Figure 2.C6. Estimates of Medicaid's Effect on Age-Adjusted Nonwhite Maternal Mortality Rates by Specification



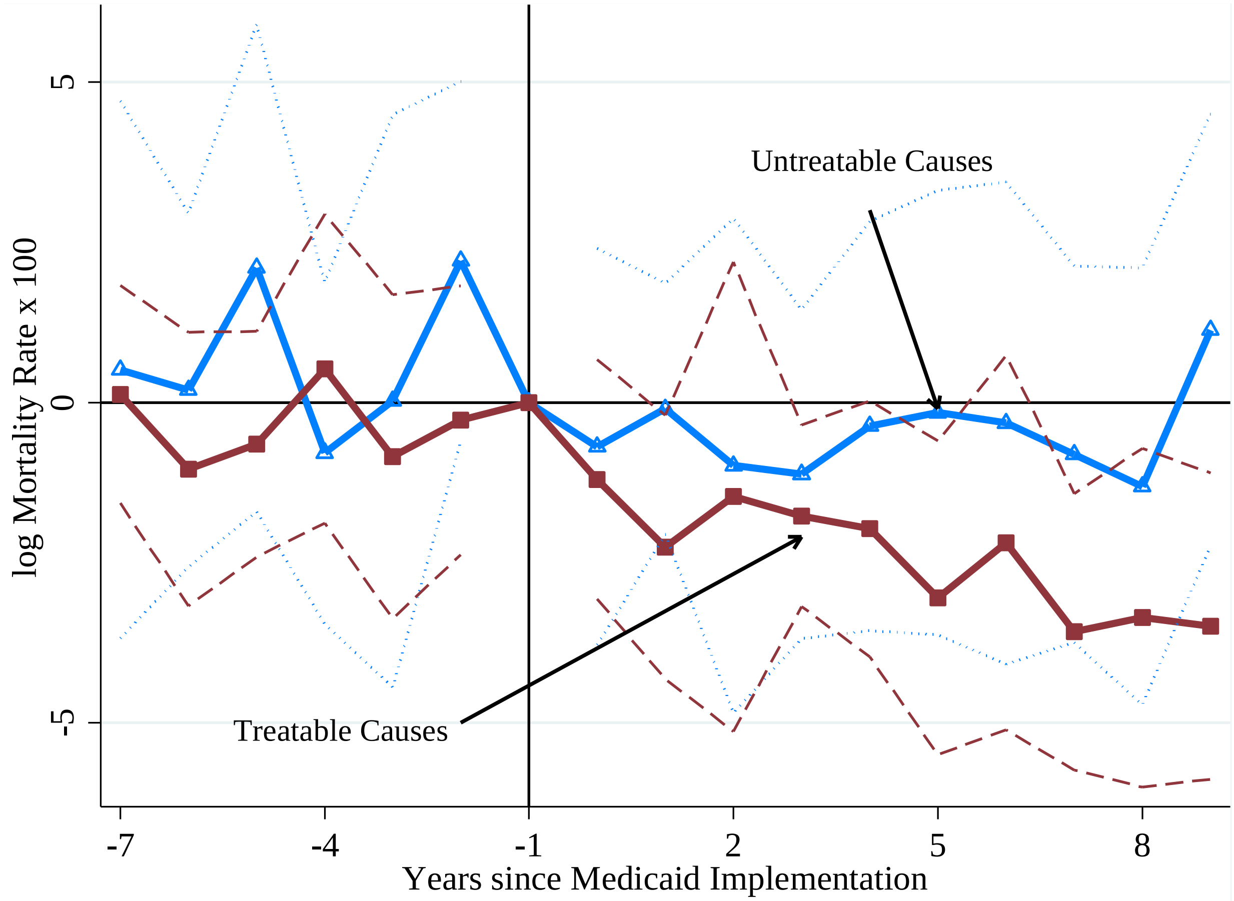
Notes: The figure plots event-study estimates for maternal mortality for continuous and binary measures of mortality and AFDC. These correspond to column 4 of table 5 and the discussion in section IV.D.

Appendix Figure 2.C7. Estimates of Medicaid's Effect on Age-Adjusted Nonwhite Internal- and External-Cause Mortality Rates



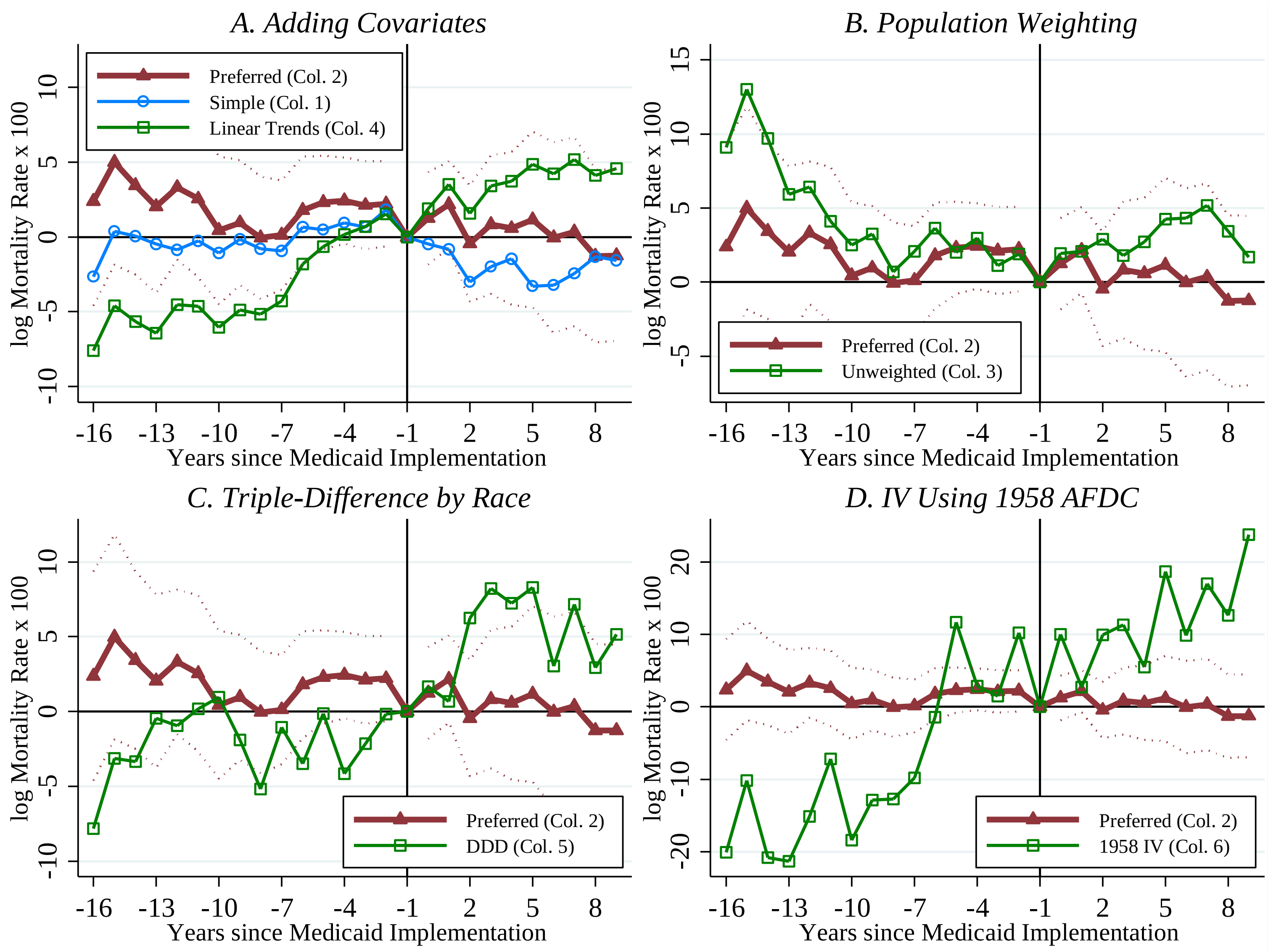
Notes: the figure plots event-study estimates that correspond to columns (4) and (5) of table 7.

Appendix Figure 2.C8. Estimates of Medicaid's Effect on Age-Adjusted Nonwhite Treatable- and Untreatable-Cause Mortality Rates



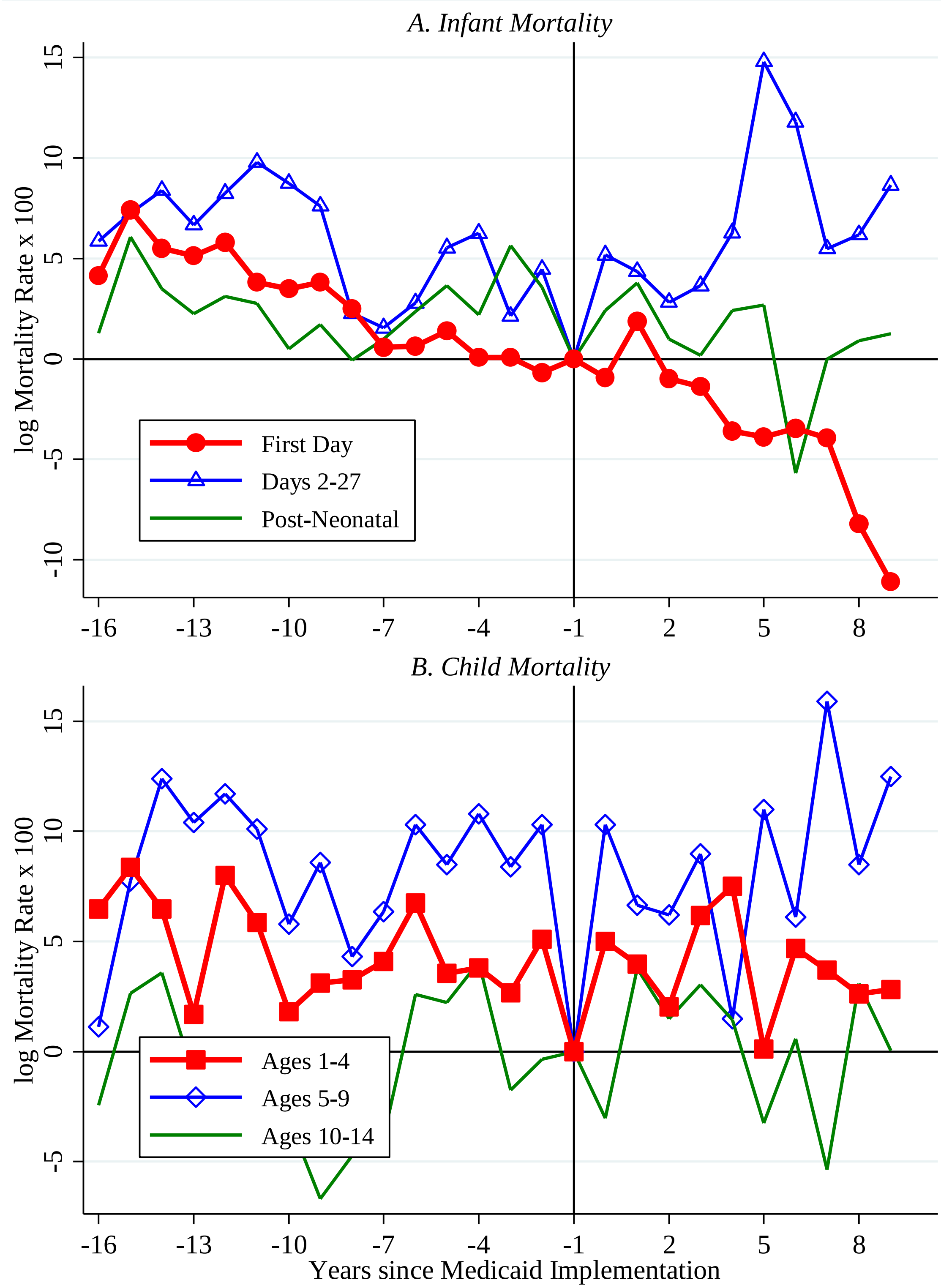
Notes: the figure plots event-study estimates that correspond to columns (6) and (7) of table 7.

Appendix Figure 2.C9. Estimates of Medicaid's Effect on Age-Adjusted White Mortality Rates by Specification



Notes: The bold line in each panel is reproduced from panel B of figure 6. The other specifications correspond to the columns of table 3.

Appendix Figure 2.C10. Estimates of Medicaid's Effect on Age-Specific White Mortality Rates



Notes: This figure plots age-specific morality results for whites, comparable to figure 7. See note to figure 6.

Appendix 2D. Additional Mortality Difference-in-Difference Results

Appendix Table 2.D1. Robustness to Alternative Samples

Dependent Variable:	log Nonwhite First Day Infant Mortality	log Nonwhite Neonatal Infant Mortality	log Nonwhite Mortality, Ages 1-4	log Nonwhite Child Mortality, 0-14
Sample:				
Baseline	-1.50	-1.47	-2.25	-1.41
	[0.47]	[0.4]	[0.54]	[0.34]
Add Arizona as a Control	-1.33	-1.27	-1.90	-1.26
	[0.46]	[0.4]	[0.58]	[0.33]
Drop Deep South	-1.77	-1.87	-2.23	-1.72
	[0.44]	[0.41]	[0.65]	[0.43]
Drop South	-1.61	-1.95	-2.57	-1.84
	[0.37]	[0.34]	[0.83]	[0.41]
Drop South and Border	-1.66	-1.65	-2.87	-2.06
	[0.65]	[0.55]	[0.99]	[0.65]
Drop Low-Black-Share States	-1.56	-1.41	-2.45	-1.39
	[0.51]	[0.45]	[0.55]	[0.36]
Drop Early Abortion States (CA, NY, WA)	-1.55	-1.39	-2.46	-1.43
	[0.53]	[0.46]	[0.63]	[0.38]

Note: All samples exclude Alaska and Hawaii because they are not observed in the Vital Statistics data before 1958. The second row adds Arizona as a control (setting all Medicaid dummies equal zero). The third row drops the Deep South, which includes Alabama, Georgia, Louisiana, Mississippi, and South Carolina. The fourth row drops states where, in the 1960 Census (Ruggles et al. 2010), fewer than 50 percent or nonwhite children were black : Alaska, Arizona, Idaho, Montana, New Mexico, North Dakota, South Dakota, Utah, Washington, and Wyoming. The final row drops early abortion states to demonstrate that the results are not driven by changes in the composition of births. (Alaska and Hawaii also legalized abortion in 1970, but are already omitted from the full sample.) This is a relevant sample restriction for child deaths as well, not because of a potential abortion effect, but because California and New York were by far the largest Medicaid states in terms of enrollment and expenditures. The final row shows that none of the results are sensitive to their inclusion.

Appendix Table 2.D2. Medicaid's Effect on Log White age-Adjusted Child Mortality Across Specifications

	(1)	(2)	(3)	(4)	(5)	(6)
	A. Grouped Event-Study Estimates					
Pre-Medicaid						
(Years -16 to -12)×AFDC*	-0.68	3.22	8.74	0.26	-2.61	4.98
	[3.66]	[2.83]	[3.59]	[2.58]	[8.19]	[5.64]
(Years -11 to -8)×AFDC*	-0.57	0.99	2.63	-1.16	-1.31	0.32
	[2.24]	[2.13]	[2.41]	[2.09]	[5.74]	[3.49]
(Years -7 to -2)×AFDC*	0.61	1.85	2.29	0.93	-1.44	2.62
	[1.44]	[1.37]	[1.6]	[1.3]	[4.34]	[2.21]
Post-Medicaid						
(Year 0)×AFDC*	-0.49	1.18	2.02	1.34	1.90	-1.87
	[1.33]	[1.55]	[1.98]	[1.6]	[4.86]	[3.4]
(Years 1 to 4)×AFDC*	-1.82	0.71	2.39	1.36	5.84	3.63
	[1.25]	[1.86]	[2.09]	[2.16]	[3.75]	[2.55]
(Years 5 to 9)×AFDC*	-2.39	-0.15	3.85	1.02	5.50	3.60
	[1.92]	[2.83]	[3.2]	[3.6]	[5.38]	[3.68]
R^2	0.89	0.98	0.96	0.99	1.00	0.97
DD Test (*p*-value)	0.56	0.12	0.01	0.44	0.75	0.01
	B. Difference-in-Differences Estimates					
Post-Medicaid×AFDC*	-2.06	-1.50	-0.53	0.51	7.14	1.17
	[1.55]	[1.9]	[2.74]	[2.26]	[3.95]	[2.44]
Bootstrap *p*-value	(0.21)	(0.427)	(0.873)	(0.83)	(0.043)	(0.657)
R^2	0.89	0.98	0.96	0.99	1.00	0.97
Observations	1,410	1,410	1,410	1,410	2,828	1,380
Covariates	High-AFDC FE, Time-to-Medicaid Dummies	(1) + State FE, Medicaid-timing-by-year FE, region-by-year FE, X_{st}	(2), unweighted	(2) + state-specific linear trends	Pooled Races, (2)*White + state-by-year FE	(2), IV using 1958 AFDC Rates
Mortality Rate in *t*-1*	198.2 deaths per 100,000					

Notes: This table is comparable to table 3, and shows age-adjusted child mortality across specifications for whites.

Appendix Table 2.D3. Difference-in-Differences Estimates of Medicaid's Effect on Log White Mortality by Age

	(1)	(2)
Infant Mortality		
First Day	-6.36	-5.79
	[2.88]	[2.60]
	(0.11)	(0.08)
Neonatal	-4.22	-3.57
	[2.27]	[1.97]
	(0.13)	(0.12)
Post-Neonatal	2.33	-1.62
	[4.20]	[2.95]
	(0.67)	(0.57)
Infant	-2.52	-3.2
	[2.34]	[1.95]
	(0.28)	(0.14)
Child Mortality		
Ages 1-4	2.22	-0.396
	[2.42]	[1.80]
	(0.59)	(0.85)
Ages 5-9	-0.703	0.765
	[1.16]	[2.29]
	(0.56)	(0.82)
Ages 10-14	-0.0872	1.34
	[1.29]	[1.34]
	(0.95)	(0.28)
Covariates	High-AFDC FE, Time-to-Medicaid Dummies	(1) + State FE, Medicaid-timing-by-year FE, region-by-year FE, X_{st}

Notes: The table plots DD estimates for age-specific white mortality for the simplest specification (column 1 of appendix table 2.D2) and the preferred specification (column 2 of appendix table 2.D2).

Appendix Table 2.D4. Medicaid's Effect on Log Nonwhite Neonatal Infant Mortality Across Specifications

	(1)	(2)	(3)	(4)	(5)	(6)
			A. Grouped Event-Study Estimates			
Pre-Medicaid						
(Years -16 to -12)×AFDC*	-0.22	0.31	-0.08	0.13	-0.09	0.16
	[0.5]	[0.73]	[0.85]	[0.79]	[0.76]	[1.2]
(Years -11 to -8)×AFDC*	-0.17	0.371	0.33	0.21	0.18	-1.04
	[0.33]	[0.55]	[0.77]	[0.69]	[0.67]	[1.1]
(Years -7 to -2)×AFDC*	0.01	0.08	-0.63	-0.01	-0.07	-1.81
	[0.34]	[0.45]	[0.64]	[0.5]	[0.54]	[0.98]
Post-Medicaid						
(Year 0)×AFDC*	-0.24	-0.71	-1.19	-0.69	-0.74	-1.67
	[0.23]	[0.52]	[1.1]	[0.52]	[0.65]	[0.65]
(Years 1 to 4)×AFDC*	-1.06	-1.15	-2.22	-1.05	-1.14	-2.39
	[0.23]	[0.47]	[0.95]	[0.49]	[0.55]	[0.74]
(Years 5 to 9)×AFDC*	-1.71	-1.37	-1.05	-1.09	-1.46	-2.58
	[0.41]	[0.51]	[0.85]	[0.66]	[0.61]	[0.71]
R^2	0.70	0.93	0.78	0.94	0.99	0.92
DD Test (*p*-value)	0.10	0.53	0.37	0.97	0.68	0.12
			B. Difference-in-Differences Estimates			
Post-Medicaid×AFDC*	-1.32	-1.47	-1.40	-1.00	-1.29	-1.50
	[0.2]	[0.4]	[0.53]	[0.46]	[0.57]	[0.67]
Bootstrap *p*-value	(0.002)	(0.007)	(0.005)	(0.015)	(0.033)	(0.129)
R^2	0.70	0.93	0.78	0.94	0.99	0.93
Observations	1,405	1,405	1,350	1,405	2,815	1,397
Covariates	High-AFDC FE, Time-to-Medicaid Dummies	(1) + State FE, Medicaid-timing-by-year FE, region-by-year FE, X_{st}	(2), unweighted	(2) + state-specific linear trends	Pooled Races, (2)*Nonwhite + state-by-year FE	(2), IV using 1958 AFDC Rates
Mortality Rate in *t*-1*			24.7 deaths per 1,000 live births			

Notes: For details on dependent variables see notes to table 3, for details on specification and sources see notes to table 3. The *p*-value from a Hausman of the equality of the weighted and unweighted estimates in columns 2 and 3 is 0.128 for the DD model and 0.065 for the grouped event-study model. (Deaton 1997; Solon et al. 2015).

Appendix Table 2.D5. Medicaid's Effect on Log Nonwhite Child Mortality Ages 1-4 Across Specifications

	(1)	(2)	(3)	(4)	(5)	(6)
			A. Grouped Event-Study Estimates			
Pre-Medicaid						
(Years -16 to -12)×AFDC*	0.61	-0.754	-0.54	-0.13	-0.05	-0.16
	[0.58]	[0.92]	[2.08]	[1.26]	[1.15]	[1.19]
(Years -11 to -8)×AFDC*	0.88	-0.411	0.50	-0.16	-0.27	-0.31
	[0.39]	[0.89]	[1.9]	[1.1]	[1.08]	[1.1]
(Years -7 to -2)×AFDC*	0.42	-0.77	-0.86	-0.69	-0.65	-0.40
	[0.43]	[0.72]	[1.83]	[0.74]	[1.01]	[1.05]
Post-Medicaid						
(Year 0)×AFDC*	-0.26	-1.90	-1.74	-1.84	-2.03	-3.23
	[0.49]	[1.06]	[2.07]	[1.04]	[1.7]	[1.48]
(Years 1 to 4)×AFDC*	-0.26	-2.27	-2.39	-2.24	-2.71	-2.64
	[0.54]	[0.72]	[1.94]	[0.76]	[1.14]	[0.98]
(Years 5 to 9)×AFDC*	0.35	-3.38	-3.14	-3.35	-3.16	-2.91
	[0.67]	[0.93]	[1.54]	[1.3]	[1.2]	[1.32]
R^2	0.63	0.89	0.71	0.90	0.98	0.89
DD Test (*p*-value)	0.15	0.25	0.16	0.42	0.84	0.97
			B. Difference-in-Differences Estimates			
Post-Medicaid×AFDC*	-0.49	-2.23	-2.43	-1.66	-2.59	-2.51
	[0.59]	[0.55]	[1.01]	[0.85]	[0.6]	[0.73]
Bootstrap *p*-value	(0.44)	(0.001)	(0.003)	(0.11)	(0.002)	(0.001)
R^2	0.63	0.89	0.71	0.90	0.98	0.89
Observations	1,362	1,362	1,340	1,362	2,772	1,359
Covariates	High-AFDC FE, Time-to-Medicaid Dummies	(1) + State FE, Medicaid-timing-by-year FE, region-by-year FE, X_{st}	(2), unweighted	(2) + state-specific linear trends	Pooled Races, (2)*Nonwhite + state-by-year FE	(2), IV using 1958 AFDC Rates
Mortality Rate in *t*-1*			161.7 deaths per 100,000 children			

Notes: For details on dependent variables and sample see notes to table 6, for details on specification and sources see notes to table 3. A Hausman test rejects the null hypothesis that the weighted and unweighted estimates in columns 2 and 3 are equal for the grouped event-study model (*p*-value = <0.01) but not the DD model (*p*-value = 0.747) (Deaton 1997; Solon et al. 2015).

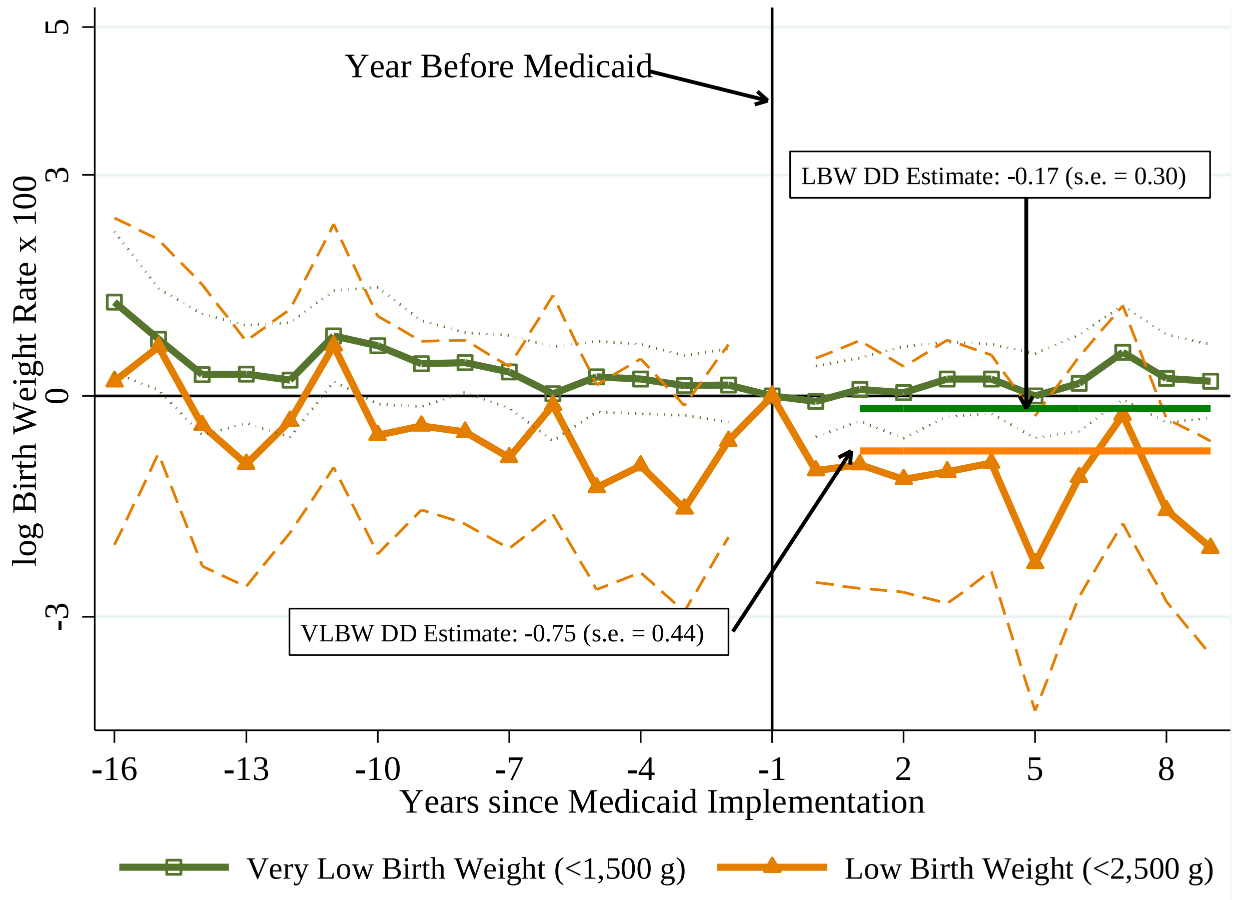
Appendix Table 2.D6. Medicaid's Effect on Log Nonwhite First-Day and Neonatal Infant Mortality Controlling for Post-Neonatal Infant Mortality

	(1)	(2)	(3)
Additional controls:	log PNMR	Birth-weight-by-year interactions	Birth-weight-by-region interactions
First 24 hours	-1.42	-1.03	-1.32
	[0.46]	[0.42]	[0.41]
Neonatal Period (before 28 days)	-1.25	-1.08	-1.34
	[0.35]	[0.35]	[0.38]

Notes: The table presents estimated effects on first-day and neonatal infant mortality that control for the log of post-neonatal infant mortality (column 1), interactions of year fixed effects with low and very low birth weight (column 2), and interactions of region fixed effects with low and very low birth weight (column 3). The results show that Medicaid's effect on the earliest infant mortality rates are not affected by controlling for subsequent mortality rates that are likely to reflect omitted factors such as hospital desegregation or changes in socioeconomic conditions, or more flexible specifications of the birth weight variables.

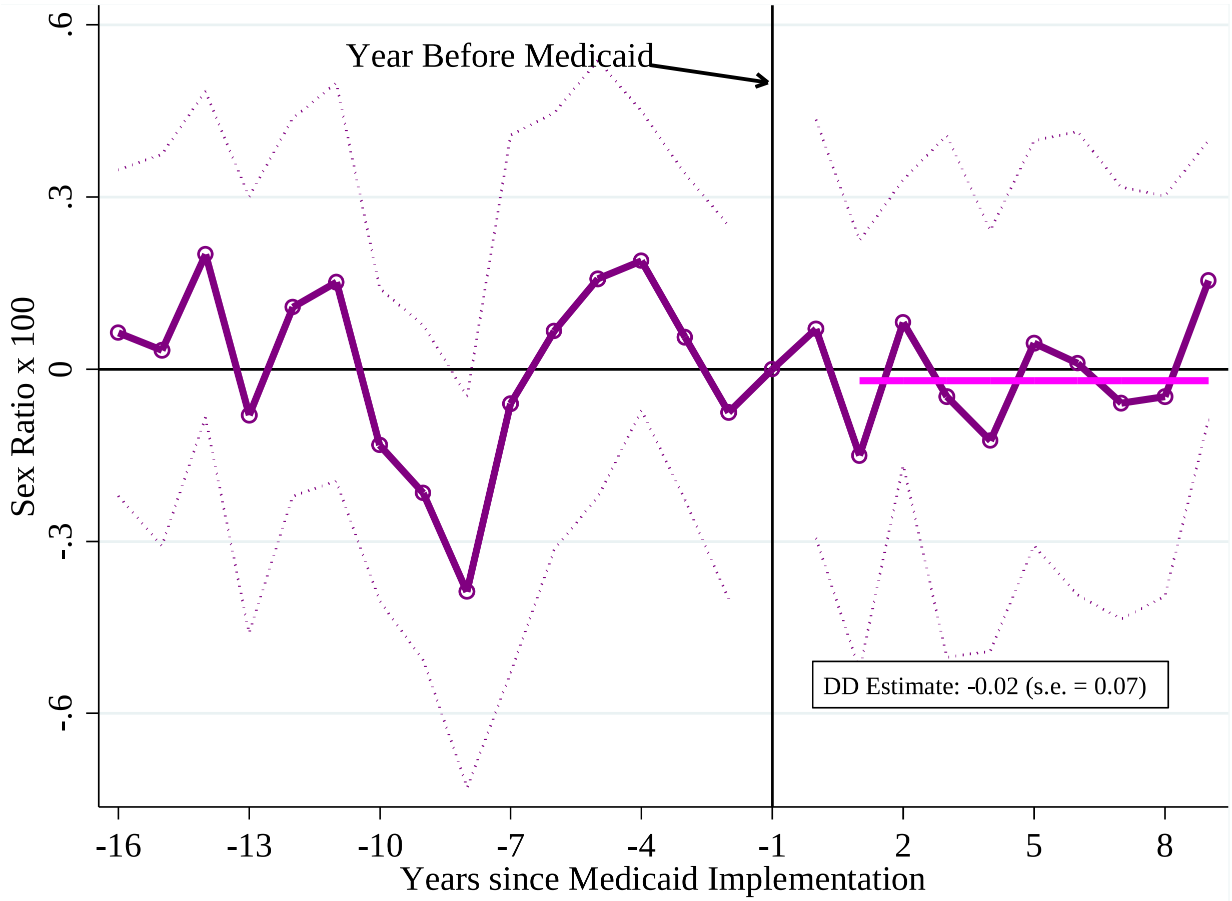
Appendix 2E. Additional Non-Mortality Event-Study Results

Appendix Figure 2.E1. Estimates of Medicaid's Effect on log Nonwhite Low and Very Low Birth Weight Rates



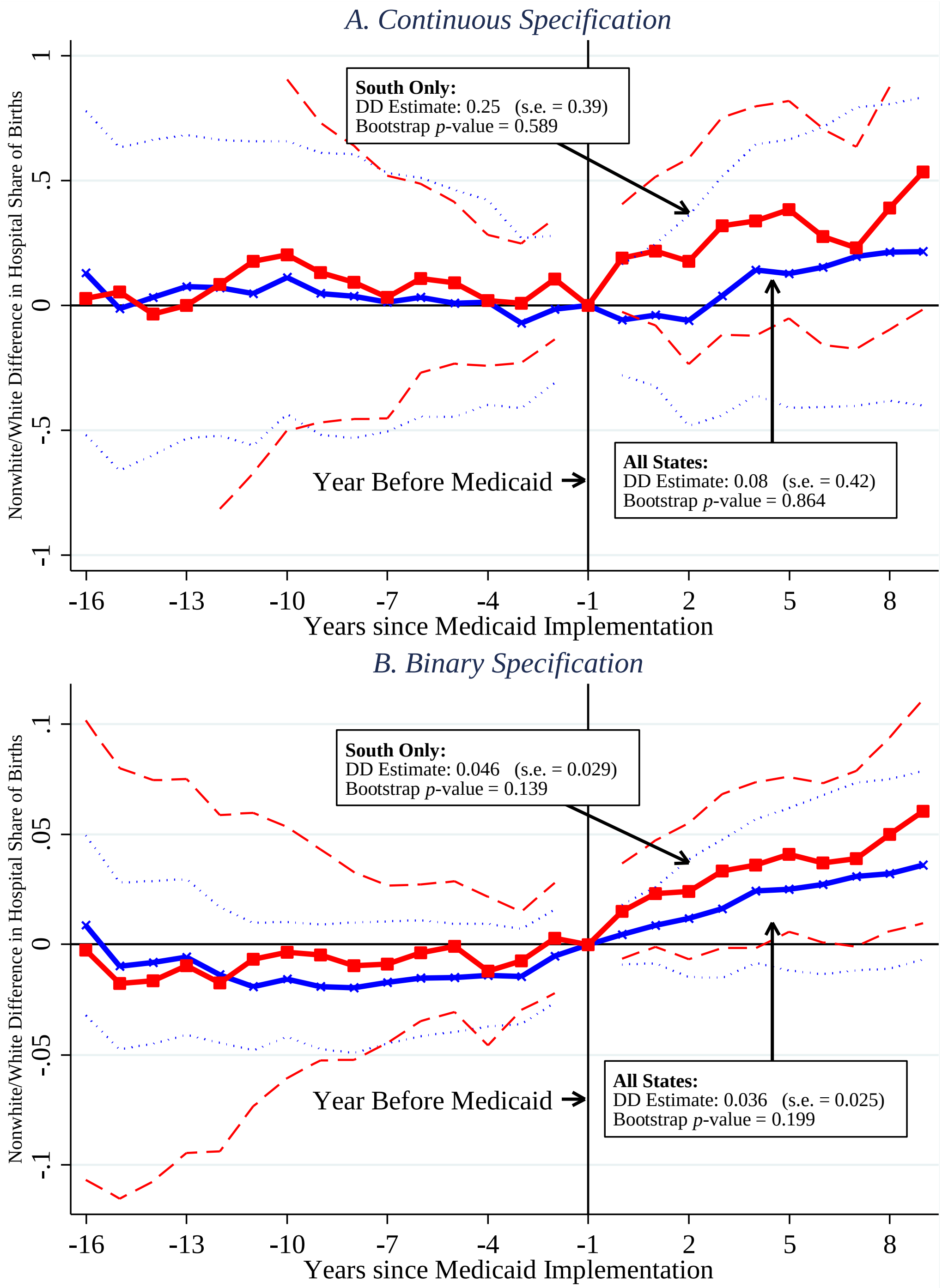
Notes: The figure plots event-study estimates that correspond to columns (1) and (2) of table 5.

Appendix Figure 2.E2. Estimates of Medicaid's Effect on the Nonwhite Sex Ratio at Birth



Notes: The figure plots event-study estimates that correspond to column (3) of table 5.

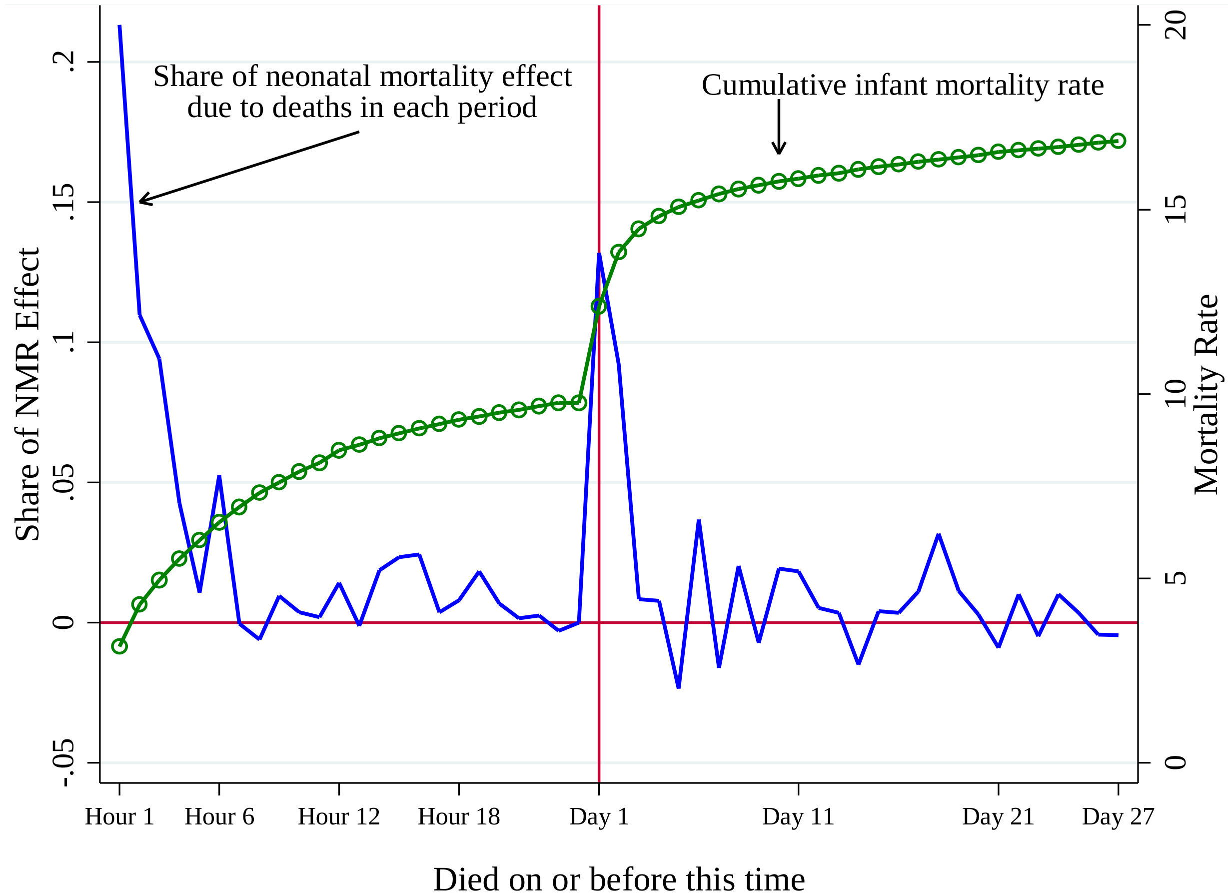
Appendix Figure 2.E2. Estimates of Medicaid's Effect on the Racial Gap in Hospital Births



Notes: The figure plots event-study estimates of Medicaid's effect on the nonwhite-white difference in hospital birth shares using $AFDC_S^*$ (panel A) and a binary measure that equals one if $AFDC_S^*$ is greater than the median. Using the racial gap addresses some measurement error in the hospital shares and controls implicitly for state-by-year effects. Specifications that use only Southern states omit continuous covariates, replace region-by-year effects with interactions between year fixed effects and a Deep South dummy, include baseline event-time dummies (rather than Medicaid-by-year fixed effects), and do not use population weights. Standard errors are clustered by states and the bootstrap p-values come from 1,000 draws of a wild cluster bootstrap percentile-t procedure (Cameron et al. 2008). The data were collected by Amy Finkelstein and Heidi Williams with support from NIA grant P30- AG012810 and publicly are available through NBER.

Appendix 2F. Results Related to the Interpretation of the Effects

Appendix Figure 2.F1. Decomposition of Neonatal Effects by Hour and Day at Death



Notes: The figure plots the share of infant mortality effects in figure 8 that come from each hour and day. Because figure 8 uses cumulative mortality rates, the effect of each point in levels comes from subtracting adjacent coefficients and multiplying by period-specific mortality rates. Dividing by the neonatal mortality estimate times the baseline NMR yields the solid blue line above. The open circles show baseline mortality rates at each hour and day of death in 1965.

Appendix Table 2.F1. Medicaid's Effect on levels and Poverty Gaps in Age-Adjusted Nonwhite Child Mortality, and Number Needed to Treat, 1966-1979

(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
Year	Share on Medicaid	Pop. (millions)	Mort. Rate	Poor/ Overall Mort.	Mortality Among Treated (4)*(5)	Mortality Among the Untreated: [(4) - (2)*(6)]/ [1-(2)]	Counterfactual Mortality Among Treated: (6)/[1-0.2]	Counterfactual Mortality Rate: (4) + 0.2*(8)*(2)	Lives Saved: [(9)-(4)]*(3)*10	Proportional Effect on Aggregate Mortality: [(4)-(9)]/(9)	Proportional Effect on Poverty Gap Mortality: - 0.2*(8)/ [(8)-(7)]	Number Needed to Treat: 1/ [0.3*(8)]
1965			425.5					425.5				
1966	0.08	8.18	405.2	*1.80*	729.4	377.8	911.7	419.4	1,164	-3.4%	-34%	548
1967	0.18	8.30	373.2	1.76	657.4	308.9	821.7	403.5	2,515	-7.5%	-32%	608
1968	0.18	8.41	364.1	1.72	627.4	307.9	784.3	391.7	2,324	-7.0%	-33%	638
1969	0.24	8.53	359.5	1.68	605.6	281.1	757.0	396.0	3,121	-9.2%	-32%	661
1970	0.24	8.59	348.1	1.65	573.0	276.9	716.3	382.6	2,960	-9.0%	-33%	698
1971	0.27	8.65	309.3	1.61	497.3	238.2	621.6	343.4	2,951	-9.9%	-32%	804
1972	0.25	8.66	295.9	1.57	464.3	238.5	580.3	325.4	2,554	-9.1%	-34%	862
1973	0.27	8.65	285.7	1.53	437.3	229.3	546.6	315.3	2,563	-9.4%	-34%	915
1974	0.30	8.62	270.5	1.49	403.7	213.6	504.6	300.7	2,605	-10.0%	-35%	991
1975	0.33	8.63	260.2	1.45	378.3	201.7	472.9	291.5	2,706	-10.8%	-35%	1,057
1976	0.35	8.61	252.7	1.42	357.7	196.5	447.2	283.9	2,686	-11.0%	-36%	1,118
1977	0.35	8.60	236.5	1.38	325.6	188.7	407.0	264.9	2,443	-10.7%	-37%	1,229
1978	0.35	8.61	233.6	1.34	312.7	191.3	390.9	260.9	2,348	-10.5%	-39%	1,279
1979	0.35	8.65	219.2	*1.30*	285.0	184.0	356.3	244.1	2,149	-10.2%	-41%	1,403

Notes: Nonwhite child Medicaid receipt comes from multiplying the share of all children on Medicaid (see figure 1) by the ratio of Medicaid participation rates for nonwhite children 0-14 and all children 0-19 in the 1976 Survey of Income and Education, 2.7. The treatment effect on the level of mortality uses the ATET estimate in figure 4: -0.20. Column

Appendix Table 2.F2. Cost Effectiveness Calculations for Age-Adjusted Nonwhite Child Mortality

	(1)	(2)	(3)	(4)	(5)	(6)
Year	Real Child Medicaid Expenditures (billions)	Real Child Medicaid Expenditures, Ages 0-14 (billions): (1)*0.78	Lives Saved	Life Years Gained: (3)*65.5	Cost per life saved (millions): 1,000*(2)/(3)	Cost per Discounted Life Year Gained: 10^6*(5)/[(1-0.97^{65})/(1-0.97)]
1966	$1.30	$1.02	1,164	76,225	$0.87	$30,416
1967	$2.58	$2.01	2,515	164,704	$0.80	$27,820
1968	$3.01	$2.35	2,324	152,197	$1.01	$35,129
1969	$3.85	$3.00	3,121	204,439	$0.96	$33,504
1970	$4.47	$3.48	2,960	193,863	$1.18	$40,960
1971	$6.06	$4.73	2,951	193,261	$1.60	$55,753
1972	$5.73	$4.47	2,554	167,256	$1.75	$60,959
1973	$5.13	$4.00	2,563	167,885	$1.56	$54,379
1974	$6.08	$4.74	2,605	170,598	$1.82	$63,357
1975	$6.66	$5.19	2,706	177,216	$1.92	$66,794
1976	$6.94	$5.41	2,686	175,949	$2.02	$70,152
1977	*$9.25*	$7.22	2,443	160,030	$2.95	$102,833
1978	*$9.91*	$7.73	2,348	153,796	$3.29	$114,634
1979	*$10.57*	$8.25	2,149	140,788	$3.84	$133,563
Average per Year	**$5.83**	**$4.54**	**2,506**	**164,158**	**$1.83**	**$63,590**
Total	**$81.55**	**$63.61**	**35,087**	**2,298,208**		

Notes: Child Medicaid expenditures for 1966-1976 are taken from published tables (DHEW 1967; 1968; 1969; 1971b; a; 1972b; a; 1974a; b; 1975b; a; 1976a; b). To obtain estimated spending for 1977-1979 I use state-specific linear fitted values in calendar year.

Appendix Table 2.F3. Medicaid's Effect on Levels and Poverty Gaps in Nonwhite Neonatal Mortality, and Number Needed to Treat, 1966-1979

(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
Year	Share on Medicaid	Population (millions)	Mortality Rate	Poor/ Overall Mortality	Mortality Among Treated (4)*(5)	Mortality Among the Untreated : [(4) - (2)*(6)]/[1-(2)]	Counterfactual Mortality Among Treated: (6)/[1-0.31]	Counterfactual Mortality Rate: (4) + 0.2*(8)*(2)	Lives Saved: [(9)-(4)]*(3)*10	Proportional Effect on Aggregate Mortality: [(4)-(9)]/(9)	Proportional Effect on Poverty Gap Mortality: - 0.3*(8)/[(8)-(7)]	Number Needed to Treat: 1/[0.3*(8)]
1965			25.7					25.7				
1966	0.09	0.61	25.1	1.24	31.1	24.5	45.0	26.3	759	-4.8%	-66%	74
1967	0.23	0.59	24.0	1.25	30.1	22.2	43.6	27.1	1,844	-11.5%	-61%	77
1968	0.22	0.57	23.2	1.26	29.3	21.5	42.5	26.1	1,660	-11.1%	-61%	78
1969	0.28	0.59	22.8	1.27	28.9	20.4	41.9	26.4	2,130	-13.7%	-58%	79
1970	0.30	0.62	21.6	1.28	27.8	19.0	40.3	25.4	2,332	-14.8%	-57%	83
1971	0.32	0.62	19.9	1.29	25.7	17.1	37.3	23.6	2,291	-15.8%	-55%	89
1972	0.33	0.58	19.4	1.30	25.3	16.4	36.7	23.2	2,210	-16.4%	-54%	91
1973	0.34	0.56	18.1	1.32	23.8	15.1	34.5	21.7	2,067	-16.9%	-53%	97
1974	0.38	0.56	17.4	1.33	23.0	13.9	33.4	21.3	2,204	-18.5%	-51%	100
1975	0.42	0.57	17.0	1.34	22.7	12.9	32.9	21.3	2,424	-20.1%	-49%	101
1976	0.44	0.57	16.5	1.35	22.3	12.0	32.3	21.0	2,546	-21.2%	-48%	103
1977	0.44	0.61	14.9	1.36	20.2	10.6	29.3	18.9	2,441	-21.3%	-47%	114
1978	0.44	0.62	14.2	1.37	19.5	10.0	28.2	18.1	2,417	-21.5%	-47%	118
1979	0.44	0.66	13.0	1.38	18.0	9.1	26.1	16.6	2,379	-21.6%	-46%	128

Notes: Nonwhite infant Medicaid receipt comes from multiplying the share of all children on Medicaid (see figure 1) by the ratio of Medicaid participation rates for nonwhite infants and all children 0-19 in the 1976 Survey of Income and Education, 3.2. The treatment effect on the level of mortality uses the ATET estimate in figure 4: -0.30.

Appendix Table 2.F4. Cost Effectiveness Calculations for Nonwhite Neonatal Mortality

	(1)	(2)	(3)	(4)	(5)	(6)
Year	Real Child Medicaid Expenditures (billions)	Real Child Medicaid Expenditures, Ages 0-14 (billions): (1)*0.06	Lives Saved	Life Years Gained: (3)*65.5	Cost per life saved (millions): 1,000*(2)/(3)	Cost per Discounted Life Year Gained: 10^6*(5)/[(1-0.97^{65})/(1-0.97)]
1966	$1.30	$0.08	759	49,731	$0.10	$3,586
1967	$2.58	$0.15	1,844	120,753	$0.08	$2,919
1968	$3.01	$0.18	1,660	108,751	$0.11	$3,782
1969	$3.85	$0.23	2,130	139,506	$0.11	$3,777
1970	$4.47	$0.27	2,332	152,757	$0.11	$3,999
1971	$6.06	$0.36	2,291	150,075	$0.16	$5,523
1972	$5.73	$0.34	2,210	144,728	$0.16	$5,419
1973	$5.13	$0.31	2,067	135,368	$0.15	$5,188
1974	$6.08	$0.36	2,204	144,350	$0.17	$5,760
1975	$6.66	$0.40	2,424	158,778	$0.16	$5,735
1976	$6.94	$0.42	2,546	166,792	$0.16	$5,693
1977	*$9.25*	$0.56	2,441	159,916	$0.23	$7,916
1978	*$9.91*	$0.59	2,417	158,314	$0.25	$8,566
1979	*$10.57*	$0.63	2,379	155,823	$0.27	$9,283
Average per Year	**$5.83**	**$0.35**	**2,122**	**138,975**	**$0.16**	**$5,510**
Total	**$81.55**	**$4.89**	**29,704**	**1,945,644**		

Notes: Child Medicaid expenditures for 1966-1976 are taken from published tables (DHEW 1967; 1968; 1969; 1971b; a; 1972b; a; 1974a; b; 1975b; a; 1976a; b). To obtain estimated spending for 1977-1979 I use state-specific linear fitted values in calendar year.

Appendix Table 2.F5. Medicaid's Effect on Aggregate and Poverty Gaps in Nonwhite Child Mortality (Ages 1-4), and Number Needed to Treat, 1966-1979

(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
Year	Share on Medicaid	Population (millions)	Mortality Rate	Poor/ Overall Mortality	Mortality Among Treated (4)*(5)	Mortality Among the Untreated: [(4) - (2)*(6)]/[1 -(2)]	Counterfactual Mortality Among Treated: (6)/[1-0.31]	Counterfactual Mortality Rate: (4) + 0.2*(8)*(2)	Lives Saved: [(9)-(4)]*(3) *10	Proportional Effect on Aggregate Mortality: [(4)-(9)]/(9)	Proportional Effect on Poverty Gap Mortality: -0.3*(8)/[(8)-(7)]	Number Needed to Treat: 1/[0.3*(8)]
1965			181.2					181.2				
1966	0.08	2.15	179.5	1.85	332.1	165.6	481.3	192.0	269	-6.5%	-46%	693
1967	0.22	2.13	159.1	1.82	289.5	123.3	419.5	187.1	596	-15.0%	-42%	795
1968	0.21	2.11	157.9	1.79	282.7	125.7	409.7	184.0	550	-14.2%	-43%	814
1969	0.26	2.09	148.5	1.76	261.4	109.0	378.8	179.0	636	-17.0%	-42%	880
1970	0.28	2.11	134.4	1.73	232.5	96.0	337.0	163.8	619	-17.9%	-42%	989
1971	0.30	2.10	126.9	1.70	215.7	88.6	312.7	156.1	614	-18.7%	-42%	1,066
1972	0.31	2.15	119.3	1.67	199.2	82.9	288.6	147.2	601	-19.0%	-42%	1,155
1973	0.32	2.17	121.4	1.64	199.1	84.7	288.6	150.1	624	-19.1%	-42%	1,155
1974	0.36	2.17	107.4	1.61	173.0	71.3	250.7	135.0	600	-20.4%	-42%	1,330
1975	0.39	2.16	101.1	1.58	159.7	63.2	231.5	129.3	608	-21.8%	-41%	1,440
1976	0.42	2.11	97.2	1.55	150.7	59.2	218.3	125.3	592	-22.4%	-41%	1,527
1977	0.42	2.08	97.8	1.52	148.6	61.7	215.3	125.5	578	-22.1%	-42%	1,548
1978	0.42	2.11	99.6	1.49	148.4	64.9	215.0	127.2	583	-21.8%	-43%	1,550
1979	0.42	2.15	94.3	1.46	137.7	63.5	199.5	120.0	552	-21.4%	-44%	1,670

Notes: Nonwhite young child (1-4) Medicaid receipt comes from multiplying the share of all children on Medicaid (see figure 1) by the ratio of Medicaid participation rates for nonwhite children 1-4 and all children 0-19 in the 1976 Survey of Income and Education, 3.2. The treatment effect on the level of mortality uses the ATET estimate in figure 4: -0.30.

Appendix Table 2.F6. Cost Effectiveness Calculations for Nonwhite Child Mortality (Ages 1-4)

	(1)	(2)	(3)	(4)	(5)	(6)
Year	Real Child Medicaid Expenditures (billions)	Real Child Medicaid Expenditures, Ages 0-14 (billions): (1)*0.21	Lives Saved	Life Years Gained: (3)*65.5	Cost per life saved (millions): 1,000*(2)/(3)	Cost per Discounted Life Year Gained: 10^6*(5)/[(1-0.97^{65})/(1-0.97)]
1966	$1.30	$0.27	269	17,595	$1.02	$35,476
1967	$2.58	$0.54	596	39,047	$0.91	$31,593
1968	$3.01	$0.63	550	36,010	$1.15	$39,974
1969	$3.85	$0.81	636	41,635	$1.27	$44,293
1970	$4.47	$0.94	619	40,557	$1.51	$52,713
1971	$6.06	$1.27	614	40,207	$2.07	$72,151
1972	$5.73	$1.20	601	39,374	$2.00	$69,717
1973	$5.13	$1.08	624	40,896	$1.73	$60,101
1974	$6.08	$1.28	600	39,328	$2.13	$73,993
1975	$6.66	$1.40	608	39,800	$2.30	$80,073
1976	$6.94	$1.46	592	38,807	$2.46	$85,633
1977	*$9.25*	$1.94	578	37,833	$3.36	$117,107
1978	*$9.91*	$2.08	583	38,179	$3.57	$124,325
1979	*$10.57*	$2.22	552	36,179	$4.02	$139,931
Average per Year	**$5.83**	**$1.22**	**573**	**37,532**	**$2.11**	**$73,363**
Total	**$81.55**	**$17.13**	**8,022**	**525,448**		

Notes: Child Medicaid expenditures for 1966-1976 are taken from published tables (DHEW 1967; 1968; 1969; 1971b; a; 1972b; a; 1974a; b; 1975b; a; 1976a; b). To obtain estimated spending for 1977-1979 I use state-specific linear fitted values in calendar year.

Appendix 3. ESTIMATES USING THE STAGGERED TIMING OF MEDICAID IMPLEMENTATION

One strategy to identify the effect of Medicaid implementation is to estimate difference-in-difference models that use variation in *when* states implemented Medicaid (Decker and Gruber 1993; Strumpf 2011). I do not use this source of variation because there is strong evidence that earlier and later Medicaid states are not comparable. Finkelstein (2007, fn. 4) concludes that "the timing of state implementation of Medicaid was not random with respect to hospital outcomes" and I argue that the same holds with respect to mortality rates.[1]

26 states implemented Medicaid in 1966, 16 more from 1967 to 1969 and 7 states established programs in 1970 at the latest date stipulated in the original legislation.[2] Because Medicaid increased federal reimbursement for public assistance costs, "the order in which states moved in establishing Medicaid programs was dictated by concerns about maximizing the federal share of vendor programs" (Stevens and Stevens 1974, pp. 80). This incentive led "more affluent industrial states" to adopt Medicaid earlier than poorer states with smaller welfare programs (Fein 1986, pp. 115). Strumpf (2011, table 2) shows that local government expenditures on public welfare and health programs are half as large in later Medicaid states than earlier ones. Relative to earlier states, later Medicaid states had significantly higher 1960 child poverty rates.[3] Figure 3.1 shows log mortality rates in for each Medicaid timing group with linear trend estimates for the calendar years in which no states had Medicaid (1959-1965). These estimates

[1] Strumpf (2011) also estimates triple-difference models that use the presence of children, states' decisions to implement a "medically needy" program, and whether or not women are black and live in the South. Decker and Gruber (1993) also present the coefficient on a post-Medicaid/AFDC interaction, but without controlling for the baseline Medicaid timing dummies, which means that the estimate is identified both by AFDC variation and Medicaid timing. Their results use the 1964-1967 NNFBS.

[2] Alaska (1972) and Arizona (1982) missed this deadline, although the threat to withhold reimbursements were "not only not made but never considered seriously" (Stevens and Stevens 1974, pp 137).

[3] For nonwhite children between 1 and 4 the 1960 poverty rate was 52 percent in the 1966 Medicaid states and 83 percent in the 1970 Medicaid states (p-value of the difference <0.0001); for white children the corresponding child poverty rates were 0.17 and .027 (p-value of the difference = 0.017).

show that earlier and later Medicaid states were on different mortality paths already in the early 1960s, and F-tests that the slopes are equal across the groups reject the null hypothesis of equality near or below the 5 percent level for nonwhite children (p-value =0.059) and white children (p-value = 0.02).

There are also limitations inherent in difference-in-differences estimates based only on variation in treatment timing. Bitler et al. (2003) show that in a model in which all units are treated but at different times, the difference-in-difference estimate (with year fixed effects but not unit fixed effects) only uses variation from the periods in which some units are treated and others are not. Meer and West (2013) consider a specification with unit and time fixed effects and identify potential problems with using "a standard difference-in-differences model to identify treatment effects if there is staggered treatment intensity and the treatment affects the growth of the outcome variable."

The results below show event-study and difference-in-difference estimates from a version of this timing-only estimator. It includes state fixed effects, region-by-year fixed effects, continuous covariates and the Medicaid event-time dummies:

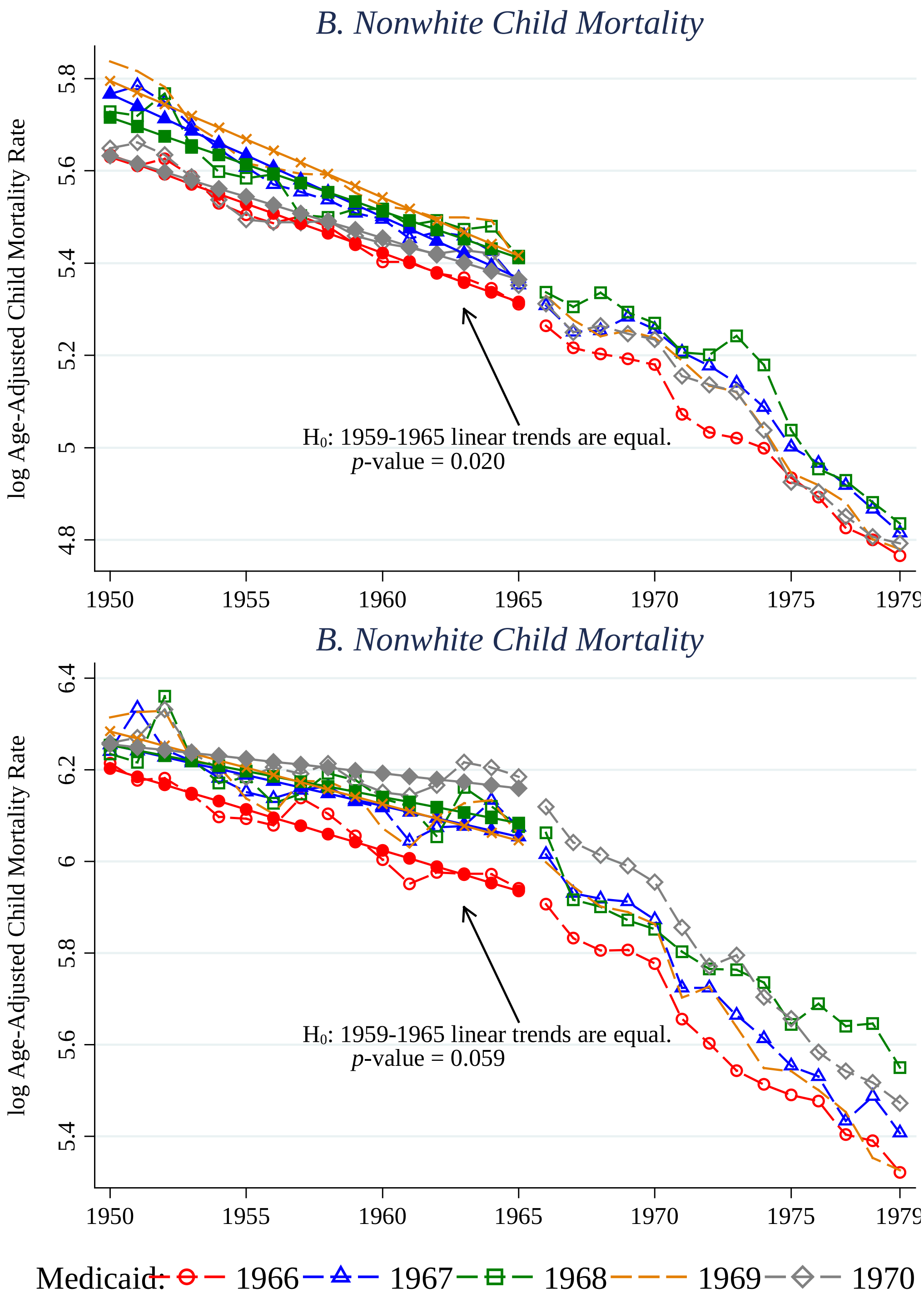
$$\ln\left(ASMR_{st}^{k}\right) = \boldsymbol{x}_{\boldsymbol{st}}'\widetilde{\boldsymbol{\beta}}_{\boldsymbol{k}} + \sum_{y=-17}^{-2} \tilde{\pi}_{y}^{k} 1\{t - t_{s}^{*} = y\} + \sum_{y=0}^{10} \tilde{\gamma}_{y}^{k} 1\{t - t_{s}^{*} = y\} + u_{st}^{k} \qquad (A1)$$

The tildes are meant to distinguish the coefficients from those in equation 1. The dependent variable is the log of age-adjusted child mortality.

The event-study results (figure 3.2) are clearly driven by strong negative trends in mortality that are correlated with Medicaid timing. Perhaps surprisingly, the associated DD results (table 3.1) are very close to zero. Figure 3.4 plots the year fixed effects (for the Northeast region) from the event-study and DD models. The year fixed effects for the restricted DD specification capture a large part of the strong negative trend that is apparent in the event-study results. This follows from the small amount of variation in Medicaid timing, and it is why the event-study results appear strongly negative (they absorb part of the time trend) but the DD results are small (the year effects account for most of the time trend). The timing-only estimator appears to be confounded by differential trends across the timing groups that (a) are distinguishable in the raw data, (b) clearly drive the event-study estimates and (c) (because they are confounded with the year effects) lead to DD estimates that suggest that Medicaid had no effect on mortality. Furthermore, figure 3.3 shows that the first-stage effects of Medicaid implementation on public insurance use are relatively small and quite imprecise.

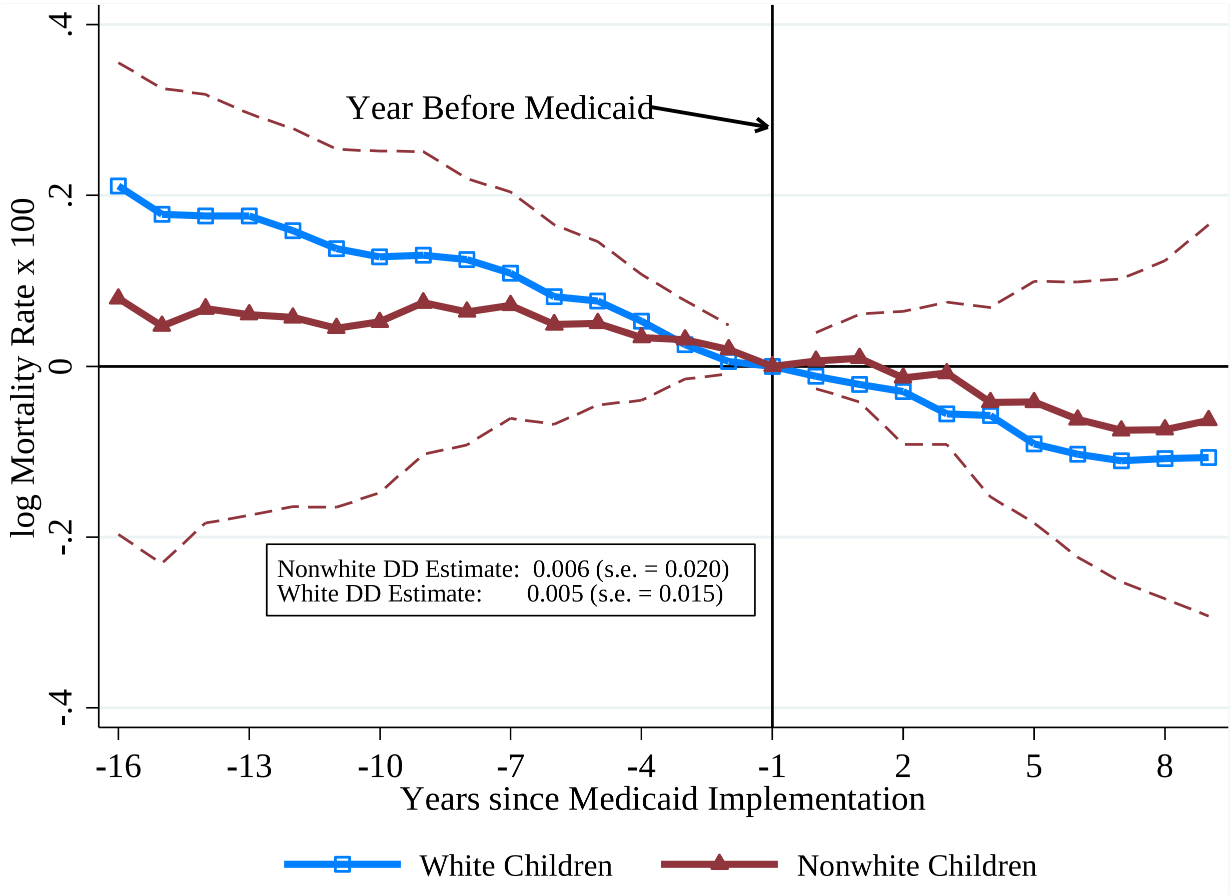
None of these concerns is present for the estimator in the main text based on categorical eligibility. The first stage is strong (figure 5). $AFDC_S^*$ is uncorrelated with pre-trends in state characteristics (table 1) and mortality rates (figures 6 and 7). DD estimates correspond closely to the event-study results (table 3). And the estimated year effects are invariant to the specification of the treatment variable (panel B of appendix figure 3.4).

Figure 3.1. Pre-Medicaid Mortality Trends by Medicaid Timing Group



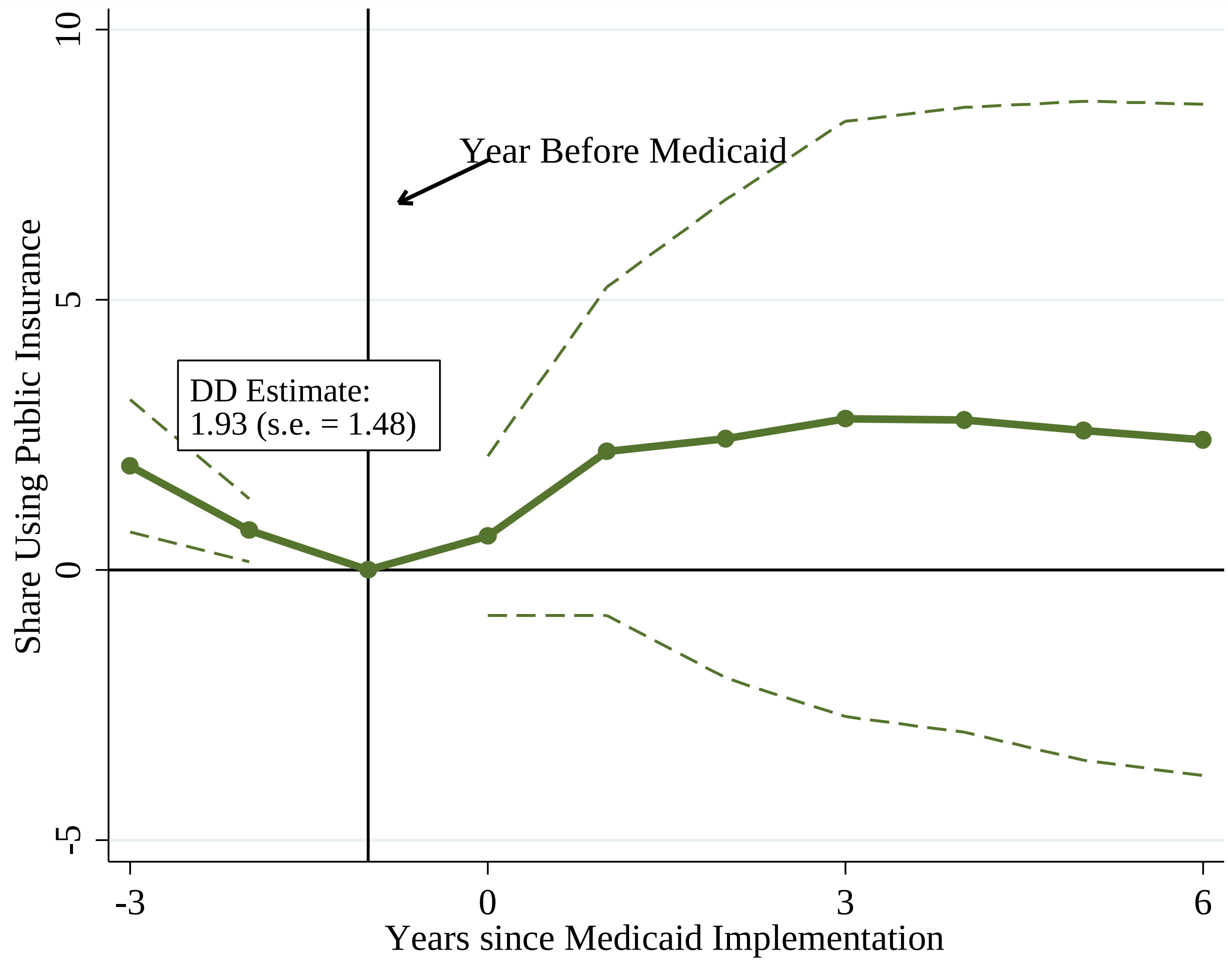
Notes: The figure plots mean log mortality rates (dashed lines and open symbols) for each group of states that implemented Medicaid in 1966, 1967, 1968, 1969 and 1970. For calendar years that precede Medicaid entirely (1950-1965), linear trends are laid over the mortality rates. A test that these slopes are equal rejects the null with a p-value of 0.02 (white) and 0.059 (nonwhite). Using 1959-1965 only leads to a very strong rejection of the null of common trends for nonwhites: p-value<0.0001, $F(4,47)=10.15$. The earliest Medicaid states had strong reductions in nonwhite child mortality in the early 1960s, while the latest Medicaid states actually had slight increases.

Figure 3.2. Event-Study Estimates for log Child Mortality Using Medicaid Timing



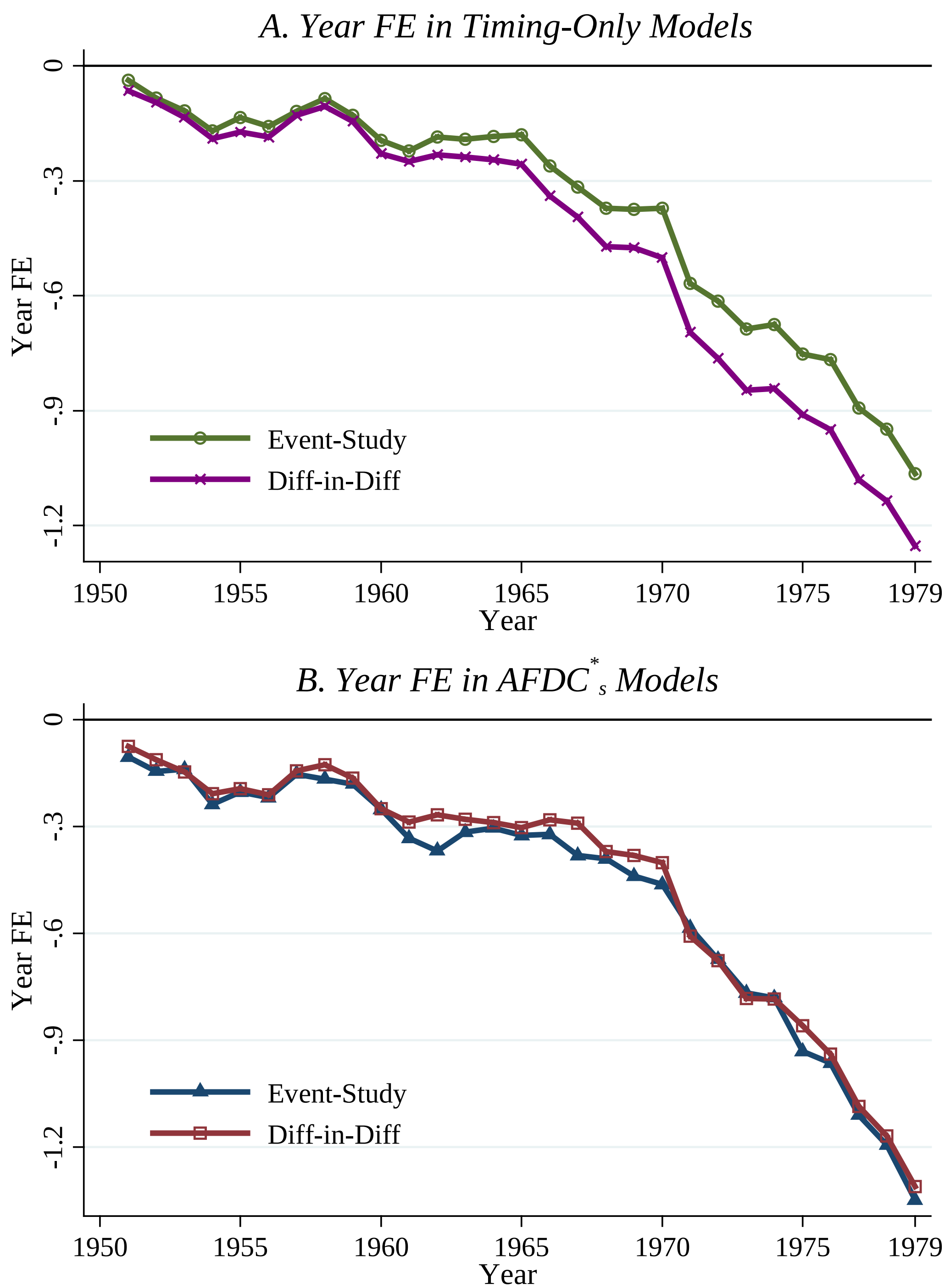
Notes: The figure plots (weighted) event-study coefficients for white and non-white age-adjusted child mortality from the preferred specification (see above) but for an estimator that only uses the staggered timing of Medicaid implementation to identify the effects. Unlike the estimates in figure 6, but consistent with the evidence in appendix figure 3.1, both series display strong trends before Medicaid and no apparent trend break afterwards. This is evidence against the validity of the timing-only estimator.

Figure 3.3. First-Stage Estimates Using Medicaid Timing



Notes: The figure plots first-stage event-study coefficients from the timing-only model for the share of children receiving public insurance.

Figure 3.4. Year Fixed Effects in Event-Study and Difference-in-Difference Models That Use Medicaid Timing versus $AFDC_s^*$



Notes: Panel A plots the estimated year dummies (the northeast region is omitted from the region-by-year fixed effects) from a timing-only model of log nonwhite child mortality with region-by-year fixed effects (the omitted year for each region is 1950). The year dummies fall much more in the DD specification than in the event-study specification, which suggests that the event-study coefficients plotted in figure 3.2 represent underlying time trends. Panel B plots the same fixed effects from the model in figure 6 that uses $AFDC_s^*$. The specification of the treatment variable does not affect the year fixed effects in this model (which fully controls for Medicaid timing through Medicaid-timing-by-year fixed effects).

Table 3.1. Difference-in-Difference Estimates for log Child Mortality Using Medicaid Timing

	(1)	(2)	(3)	(4)	(5)
			A. Log White Child Mortality (0-14)		
Post-Medicaid	0.004	0.005	0.0002	0.006	
	[0.012]	[0.015]	[0.017]	[0.015]	
R^2	0.97	0.98	0.95	0.98	
			B. Log Nonwhite Child Mortality (0-14)		
Post-Medicaid	-0.024	0.006	0.007	0.011	0.008
	[0.017]	[0.02]	[0.036]	[0.017]	[0.025]
R^2	0.93	0.95	0.85	0.96	1.00
Covariates	State FE, Year FE, Time-to-Medicaid Dummies	State FE, Region-by-year FE + X_{st}	(3), unweighted	(3) + state-specific linear trends	Pooled Races, (2)*Nonwhite + state-by-year FE

Notes: The *p*-value from a Hausman test of the difference between the weighted and unweighted estimates (columns 3 and 4) for nonwhite mortality rejects the null hypothesis that they are equal with a *p*-value of 0.079. For white mortality the *p*-value is 0.689

Appendix 4. RE-SCALING QUASI-EXPERIMENTAL ESTIMATES

Quasi-experimental studies of Medicaid's effect on mortality discussed in section I estimate the reduced-form intention-to-treat effect (ITT) of a given policy "instrument" on an aggregate mortality rate. A general version of the estimating equation relates a mortality measure, M_{ast}, for age group a, in state s at time t,to covariates, $\boldsymbol{X}'_{\boldsymbol{ast}}$, and the policy variable, z_{ast}:

$$M_{ast} = \boldsymbol{X}'_{\boldsymbol{ast}}\boldsymbol{\beta} + \gamma z_{ast} + \nu_{ast} \qquad (A1)$$

The policy instrument in Currie and Gruber (1996a; 1996b) is the share of a national sample of children or women in the March CPS who are eligible for Medicaid in each state, year, age group cell. In Wherry and Meyer (2013) the instrument is the discontinuous jump in eligibility that occurs for children born just after September 30, 1983 (and $\boldsymbol{X}'_{\boldsymbol{ast}}$ contains polynomials in birthdate). In Sommers et al. (2012a) the instrument is a dummy for being in a treatment state (New York, Maine, or Arizona) after an expansion of Medicaid eligibility (it has no a subscript). In the OHIE (2012) the instrument is a dummy for winning the eligibility lottery (ie. an individual-level equation and instrument). In this paper the instrument is the interaction of $AFDC_s^*$ with a post-Medicaid dummy (it also has no age subscript).

A. *Rescaling Intention-to-Treat Effects into Average Treatment Effects on the Treated*

The translation of ITT effects into average treatment effects on the treated (ATET) follows from a standard two-equation model for the share of children on public insurance and the effect of insurance coverage on mortality. The first-stage equation for children's insurance use (p_{ast}; not just *public* insurance) is:

$$p_{ast} = \boldsymbol{X}'_{\boldsymbol{ast}}\boldsymbol{\Gamma} + \tau_a z_{ast} + e_{ast} \qquad (A2)$$

And the structural equation for mortality is:

$$M_{ast} = \boldsymbol{X'_{ast}\alpha} + \delta_a \, p_{ast} + u_{ast} \qquad (A3)$$

A1 is obtained by substituting A3 into A2. If z_{ast} is a valid instrument for p_{ast} in A3 (ie. $z_{ast} \perp u_{ast}$), then the reduced-form coefficient (ITT) equals the product of the first-stage and the structural coefficient:

$$\gamma = \tau_a \delta_a$$

Estimates of γ from existing Medicaid papers are shown in panel B of table 4.1 and the corresponding first-stage estimates of τ_a are shown in panel B of table 4.2. The ratio of these coefficients ($\frac{\gamma}{\tau_a}$) is an estimate of the effect of Medicaid coverage on the *level* of mortality among new Medicaid recipients *who gained coverage*—the average treatment effect on the treated (ATET). This exercise is valid if the policy *only* affects the mortality of those who are induced to move from uninsured to insured (ie. there are no spillovers to children whose coverage status did not change).

Since the estimates considered here span 50 years over which time aggregate mortality rates fell, treatment effects in levels may differ simply because baseline mortality rates differ. To facilitate comparisons across studies I express ATET estimates as a proportion of baseline mortality among newly-covered Medicaid recipients. For existing studies (whose effects are in levels) this requires an estimate of the mortality rate in a particular time period (pre-treatment) for a particular sub-group (new Medicaid recipients). Unfortunately, the variables that determine Medicaid eligibility (eg. income, family structure) are not available in the Vital Statistics data, and surveys that collect this information for samples of decedents have only been conducted rarely. Table 4.3 lists the sources and calculations I use to approximate the correct denominator of the proportional ATET for existing Medicaid papers.

I assume that the mortality rate for the poor in the relevant demographic group (first column of tables 4.1 and 4.2) is an accurate guess for the mortality rate among new Medicaid recipients. On one hand, recent eligibility expansions covered many families with incomes *above* the poverty line, so the mortality rate among the poor may be too large and, therefore, understate the proportional ATET. On the other hand, adverse selection into Medicaid, especially for those who enroll at the point of care (Sommers et al. 2012b), implies that the average mortality rate among the poor may be lower than the true rate for new Medicaid recipients. To the extent that new enrollees under each policy change differ from the average poor person, proportional ATET estimates constructed using mortality rates among the poor may be too large or too small. For an alternative approach to rescaling these effects based on assuming a value for the proportional ATET see table S4 in Sommers et al. (2012a)

Because only one dataset contains information on the income and demographics of living and (eventually) dead respondents (the National Longitudinal Mortality Study, NLMS), I calculated the mortality rate among the poor as the aggregate mortality rate ($\frac{D}{N}$) times the ratio of poverty rates among decedents and among the population ($\frac{D_{poor}}{D}\left[\frac{N_{poor}}{N}\right]^{-1}$):

$$\frac{D_{poor}}{N_{poor}} = \frac{D_{poor}}{N_{poor}} \cdot \frac{D}{N} \cdot \frac{N}{D} = \frac{\frac{D_{poor}}{D}}{\frac{N_{poor}}{N}} \cdot \frac{D}{N} \qquad (A4)$$

The poverty rate among a given year’s decedents can be calculated from retrospective surveys of decedents’ family members (HHS 1969; HHS and NCHS 1986; HHS and NCHS 2005) or from survey data with linked information on deaths (NLMS). The poverty rate among the living can be calculated from the Current Population Survey or for infants, from retrospective surveys of mothers (HHS and NCHS 1999, 2001). The aggregate mortality rate is calculated from Vital

Statistics data and is given in each paper. These statistics are listed in table 4.3. The final column shows the implied mortality rate among the poor based on (A4), which is higher than average in all cases.

Because the estimates in this paper refer to *log* mortality rates, they are already expressed as proportional changes in the overall mortality rate. Think of δ_a as the treatment effect on mortality levels divided by $\frac{D}{N}$, the aggregate mortality rate. This means that the adjustment factor to rescale them to the proportional effect among treated children is slightly different. Using equation (A4) I adjust my effects not by the estimated level of mortality among new enrollees, but by the ratio of poor-to-aggregate mortality rates. This is shown in table 4.4. Data on poverty for infant deaths and live births are available in the 1960s as are data on poverty for adult deaths (and the adult population), but similar data for children are not. For the child results I use data from 1983 on child *and* adult deaths to adjust the adult scaling factor calculated from the 1960s data. See notes to table 4.4

The proportional ATET estimates are shown in table 4.5. They equal the reduced-form ITT effect on mortality ($\hat{\gamma}$), divided by the corresponding first-stage effect on *any* insurance coverage ($\hat{\tau}$), divided by the scaling factor ($\hat{f}$) that adjusts for the higher baseline mortality of new Medicaid recipients. (I further adjust first-stage estimates based on survey data because of well-documented under-reporting.)

$$\widehat{ATET} = \frac{\hat{\gamma}}{\hat{\tau}}\hat{f} \qquad (A5)$$

These are the estimates plotted in figure IX.

B. Calculating Standard Errors for the Average Treatment Effect on the Treated

The proportional ATET is a non-linear function of several estimated parameters, and calculating its standard error is not straightforward. Ordinarily, a resampling procedure could be used to capture any distribution of the components or covariances between them that exist in the analysis sample. Unfortunately, the pieces that make up $\widehat{ATET}$ come from different datasets with different years of coverage and so resampling is not feasible. Instead I use a parametric bootstrap procedure which leverages the asymptotic normality of the regression estimates (already assumed in the inference procedures used throughout the paper), and the fact that the scaling factor is a function only of proportions to draw a series of bootstrap samples from normal random variables (for the ITT and first-stage coefficients) and binomial random variables (for the proportions that make up the scaling factor). A simplified version of this approach is described in Johnston and DiNardo (1997, pg. 365-366). The algorithm is as follows:

1. ITT: Store 10,000 draws from a normal distribution with mean equal to the point estimates and standard deviation equal to the standard errors reported in table 4.1.
2. First-Stage: Store 10,000 draws from a normal distribution with means equal to the point estimates and standard deviation equal to the standard errors reported in table 4.2.
3. Scaling factor: For each component of the scaling factor create 10,000 samples from a uniform distribution with the number of observations listed in tables 4.3 and 4.4 and store the share of those draws that are less than or equal to the proportions (ie. the poverty rates) listed in tables 4.3 and 4.4.
4. Bootstrap ATET: for each of the 10,000 replications calculate the ATET according to A5. Make an adjustment for underreporting of 0.85 (Davern et al. 2007), which is *not* bootstrapped. The aggregate mortality rates reported in each paper are not bootstrapped since they are based on the universe of deaths reported in Vital Statistics data.

The output of this procedure is a dataset of 10,000 replications of the proportional ATET. I use a modified percentile method to create confidence intervals from this bootstrap sample. The lower end of the confidence interval is the 5th percentile of the draws that are below the mean and the upper end of the confidence interval is the 95th percentile of the draws that are above the mean. This yields a confidence interval that covers 95 percent of the bootstrap estimates. Applying the standard percentile method (which chooses the smallest interval that covers 95 percent of the bootstrap draws) tends to create much tighter confidence intervals by choosing an upper bound very close to (but never greater than) zero. This follows from the skewness in the empirical distribution of the ATETs. The modified percentile method is more conservative in that it always yields wider confidence intervals.

C. Covariance between Reduced-Form and First-Stage Effects

The procedure outlined above assumes that all the parameters that make up the ATET are independent (this assumption would not be necessary in a nonparametric bootstrap). To check the importance of this assumption I repeat the procedure for a range of values of the correlation coefficient between the reduced-form and the first-stage.[4] The results are plotted in figure 4.1. The solid line is the ATET estimate and the dashed lines are the confidence intervals under the assumption that the correlation between the ITT and the first-stage, ρ = -1, -0.9,-0.8,…0.8, 0.9, 1. Figure 4.1 yields two important conclusions. First, no value of this correlation leads the confidence intervals to cross zero. This is not surprising, since a test that the reduced form is significantly different from zero is equivalent to a test that the structural coefficient is different from zero (Angrist and Krueger 2001; Chernozhukov and Hansen 2008). Second, for overall child mortality and for neonatal infant mortality, this exercise never yields confidence intervals that cross -100 percent (the logical lower bound). The confidence intervals clearly widen as the

[4] I thank Alejandro Molnar for this suggestion.

two parameters are more positively correlated, but the main points from figure IX are not affected.

Table 4.1. Intention-to-Treat Effects of Medicaid Policy Changes on Mortality

Population	Paper	Notes	Source	ITT Effect on Mortality Rates
		A. Estimates from this Paper (logs)		
Nonwhite Children (0-14)			Table 3, Column 2, Panel B	-1.41 [0.34]
Nonwhite Neonates			Table 4, Column 3, Panel B	-1.47 [0.4]
Younger Nonwhite Children (1-4)			Table 7, Column 1, Panel B	-2.23 [0.55]
		B. Existing Estimates of Medicaid's Effect on Mortality (levels)		
Infants	Currie and Gruber (1996a)	Smallest overall mortality estimate	Table 3, Column 6, Row 1	-2.82 [0.69]
Children (1-14)	Currie and Gruber (1996b)	Outcome is per 10,000 children.	Table VI, Column 1, Row 1	-1.277 [0.48]
Black Teens (15-18)	Meyer and Wherry (2012)	Smallest mortality estimate, internal causes. Eligibility gain is 0.8 years.	Table 7, Column 8, Row 6	-0.34 [0.15]
Adults (20-64)	Sommers, Baicker and Epstein (2012)	SE calculated from upper CI: (-11.9+19.6)/1.96	Table 2, Column 2, Row 1	-19.6 [3.92]
White Adults (20-64)	Sommers, Baicker and Epstein (2012)	SE calculated from upper CI: (-8.2+14)/1.96	Table 2, Column 2, Row 2	-14 [2.96]
Nonwhite Adults (20-64)	Sommers, Baicker and Epstein (2012)	SE calculated from upper CI: (-17.3+41)/1.96	Table 2, Column 2, Row 3	-41 [12.1]

Notes: standard errors in square brackets.

Table 4.2. First-Stage Effects of Medicaid Policy Changes on Any Insurance Coverage

Population	Paper	Notes	Source	First-Stage Effect on Insurance Coverage
A. Estimates from this Paper				
All children (0-19)			Table 2, Column 1, Panel B	3.83 [0.94]
B. Existing Estimates of Medicaid's Effect on Mortality				
Infants	Dave, Decker, Kaestner and Simon (2008)	Administrative data, 1986-1991. Model without state trends.	Table 3, Column 5, Row 1	0.163 [0.05]
Children (1-14)	Cutler and Gruber (1996)	Actual estimate is for *uninsurance*, and is negative.	Table IV, Row 1, Column 3	0.119 [0.02]
~7 year old children, family income between 60% and 140% of FPL	Card and Shore-Sheppard (2004)	Contemporaneous effect, not cumulative by ages 15-18.	Table 3, Last Row, Column 6	0.1 [0.05]
Adults (20-64)	Sommers, Baicker and Epstein (2012)	SE calculated from upper CI: (-0.024+0.032)/1.96	Table 3, Column 2, Row 1	0.032 [0.004]
White Adults (20-64)	Sommers, Baicker and Epstein (2012)	SE calculated from upper CI: (-0.018+0.033)/1.96	Table 3, Column 2, Row 2	0.033 [0.01]
Nonwhite Adults (20-64)	Sommers, Baicker and Epstein (2012)	SE calculated from upper CI: (-0+0.028)/1.96	Table 3, Column 2, Row 3	0.028 [0.01]

Notes: standard errors in square brackets.

Table 4.3. Baseline Mortality Among New Medicaid Recipients for Existing Studies: Estimated Mortality Rates for the Poor

Population	Dataset	Notes	(1) Poverty Rate Among Decedents	(2) Poverty Rate	(3) Reported Baseline Mortality Rate	(4) Adjusted Mortality Rate for Poor: (3)*((1)/(2))
Infants	1980 National Natality Followback Survey	Not all mothers sampled at one year post-birth. No income or infant death data for out-of-wedlock births.	0.21 (178)	0.15 (7,936)	12.6 per 1,000 live births	17.49 per 1,000 live births
Children (1-14)	National Longitudinal Mortality Study	The follow-up period is 11 years rather than 1 year. This is the poverty rate in the survey year of children who died in the follow-up period.	0.26 (871)	0.2 (210,430)	3.81 per 10,000 children	4.85 per 10,000 children
Black Teens (15-18)	1993 National Mortality Followback Survey, 1993 March CPS		0.66 (136)	0.36 (1,069)	2.35 per 10,000 teens	4.3 per 10,000 teens
Adults (20-64)			0.27 (11,213)	0.11 (89,373)	320 per 100,000 adults	758.03 per 100,000 adults
White Adults (20-64)			0.24 (7,272)	0.09 (76,669)	309 per 100,000 adults	788.75 per 100,000 adults
Nonwhite Adults (20-64)			0.38 (3,941)	0.22 (12,704)	361 per 100,000 adults	621.61 per 100,000 adults

Notes: sample sizes in parentheses.

Table 4.4. Baseline Mortality Among New Medicaid Recipients for This Paper: Estimated Mortality Rates for the Poor

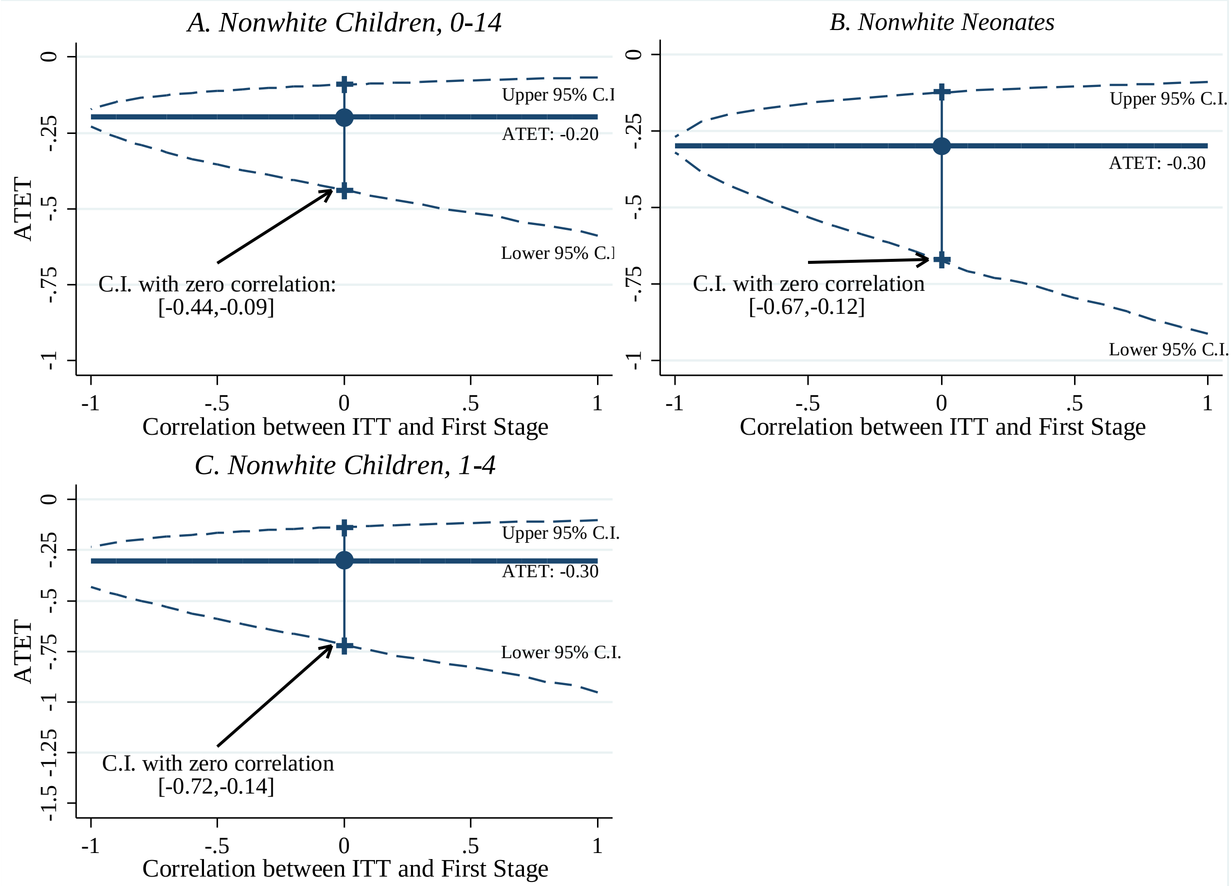
		1964-1966 Poverty Rate		1968 Poverty Rate		1983 Poverty Rate		1983 Poverty Rate		Decedent/ Population Poverty Rate
Population	**Dataset**	**Infant Decedents**	**Infant Population**	**Adult Decedents**	**Adult Population**	**Adult Decedents**	**Adult Population**	**Child Decedents**	**Child Population**	
Nonwhite Children (0-14)	1966-1968 National Mortality Followback Study; 1993 National Longitudinal Mortality Study; 1968, 1983 March CPS			0.63	0.31	0.38	0.27	0.54	0.43	*1.80*
				(2,572)	(5,981)	(11,046)	(59,442)	(202)	(38,693)	
Nonwhite Neonates	National Infant Mortality Study 1964-1966, National Natality Followback Study 1964-1966	0.60	0.48							*1.24*
		(2,315)	(1,091)							
Younger Nonwhite Children (1-4)				0.63	0.31	0.38	0.27	0.57	0.44	*1.85*
				(2,572)	(5,981)	(11,046)	(59,442)	(34)	(10,373)	

Notes: Notes: sample sizes in parentheses. No data exist on decedents under age 18 from the 1960s. The scaling factor for nonwhite children in the 1960s is calculated in two steps. First, I use data on the scaling factor (decedent poverty rate/overall poverty rate) for adults and children in the National Longitudinal Mortality Study (representative of the 1983 population) to get a child/adult ratio. Second, I assume that this *ratio* was constant over time and multiply it by the scaling factor for nonwhite adults in the 1968 National Mortality Followback Survey.

Table 4.5. Proportional Average Treatment Effects of Medicaid on Treated Recipients

Population	Paper(s)	Adjustment for Medicaid underreporting:	Additional Notes	Proportional Effect of Medicaid on Mortality Rates of New Recipients
		A. Estimates from this Paper		
Nonwhite Children (0-14)				*-20%*
Nonwhite Neonates				*-31%*
Younger Nonwhite Children (1-4)				*-31%*
		B. Estimates from Existing Papers		
Infants	Currie and Gruber (1996a), Dave, Decker, Kaestner and Simon (2008)	None, administrative coverage data.		*-99%*
Children (1-14)	Currie and Gruber (1996b), Cutler and Gruber (1996)	0.85 (Davern, Klerman and Ziegenfusi 2007)		*-188%*
Black Teens (15-18)	Meyer and Wherry (2012), Card and Shore-Sheppard (2004)	0.85 (Davern, Klerman and Ziegenfusi 2007)	Divided by 0.8 to reflect less than a full year of eligibility gain.	*-84%*
Adults (20-64)	Sommers, Baicker and Epstein (2012)	0.85 (Davern, Klerman and Ziegenfusi 2007)		*-69%*
White Adults (20-64)	Sommers, Baicker and Epstein (2012)	0.85 (Davern, Klerman and Ziegenfusi 2007)		*-46%*
Nonwhite Adults (20-64)	Sommers, Baicker and Epstein (2012)	0.85 (Davern, Klerman and Ziegenfusi 2007)		*-200%*

Figure 4.1. 95% Modified Percentile Method Bootstrap Confidence Intervals Under Different Values of the Correlation between the Reduced Form and First Stage Estimates



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