NBER WORKING PAPER SERIES

VARIATIONS IN INFANT MORTALITY RATES AMONG COUNTIES OF THE UNITED STATES: THE ROLES OF SOCIAL POLICIES AND PROGRAMS

Michael Grossman

Steven Jacobwitz

Working Paper No. 615

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge MA 02138

January 1981

The research reported here is part of the NBER's research program in Health Economics. Any opinions expressed are those of the authors and not those of the National Bureau of Economic Research.

NBER Working Paper #615
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ABSTRACT

The purpose of this paper is to shed light on the causes of the rapid decline in the infant mortality rate in the United States in the period after 1963. The roles of four social policies are considered: Medicaid, subsidized family planning services for low-income women, maternal and infant care projects, and the legalization of abortion. The most striking finding is that the increase in the legal abortion rate is the single most important factor in reductions in both white and nonwhite neonatal mortality rates. Not only does the growth in abortion dominate the other social policies, but it also dominates schooling and poverty.

Michael Grossman
Department of Economics
City University of New York
Graduate School
33 West 42nd Street
New York, New York 10036

(212) 790-4426

Steven Jacobowitz
Department of Economics
Montclair State College
Upper Montclair, New Jersey 07043

(201) 893-5255

National Bureau of Economic Research
15-19 West 4th Street, 8th Floor
Washington Square
New York, New York 10012

(212) 598-3321

VARIATIONS IN INFANT MORTALITY RATES AMONG COUNTIES OF THE UNITED STATES: THE ROLES OF SOCIAL POLICIES AND PROGRAMS

Michael Grossman and Steven Jacobowitz*

From 1964 to 1977, the infant mortality rate in the United States declined at an annually compounded rate of 4.4 percent per year. This was an extremely rapid rate of decline compared to the figure of 0.6 percent per year from 1955 to 1964. The reduction in mortality proceeded at an even faster pace in the 1970s than in the late 1960s (5.2 percent per year from 1971 to 1977 versus 3.8 percent per year from 1964 to 1971).[1]

The period from 1964 to 1977 witnessed the introduction of Medicaid, maternal and infant care projects, Federally subsidized family planning services for low-income women, the legalization of abortion, and the widespread adoption of oral and intrauterine contraceptive techniques. These developments have been pointed to in discussions of the cause of the acceleration in the downward trend in infant mortality (for example, Eisner et al., 1978; Lee et al., 1980), but the question has not been studied in a multivariate context. Moreover, the relative contribution of each factor has not been quantified. The purpose of this paper is to estimate the impacts of social policies and programs on infant mortality.

I. Analytical Framework

Economic models of the family and household production developed by Becker and Lewis (1973) and Willis (1973) provide a fruitful theoretical framework to generate multivariate health outcome functions and to assess the roles of social programs and policies in these functions. Ben-Porath (1973), Ben-Porath and context of an economic model of the family, these developments raise the "optimal" survival probability and lower the "optimal" number of births. In addition, they will lower the observed infant mortality rate if less healthy fetuses are less likely to be conceived or more likely to be aborted.[4]

To measure the relative importance of the above factors in the recent U.S. infant mortality experience, we perform a cross-sectional regression analysis of variations in infant mortality rates among counties of the United States in 1971. Our procedure capitalizes on variations in social programs among counties at a moment in time. Thus it provides a set of impact coefficients to identify the contribution of each program net of basic determinants of infant mortality such as poverty, schooling levels, and the availability of physicians. After estimating the regression, we apply its coefficients to national trends in the exogenous variables between 1964 and 1977 to "explain" the trend in infant mortality.

Our methodology has a number of desirable properties. It mitigates the multicollinearity problems that almost certainly would arise in a time-series regression analysis for the U.S. as a whole.[5] Moreover, the state-of-the-art in neonatalogy, which has changed over time and is difficult to quantify, is constant in the cross section. Finally, with the exception of abortion reform, the social programs that we study are aimed at poor persons. Therefore, the appropriate way to measure their impacts is to interact the policy variables with the fraction of births to poor women. We incorporate this insight into our basic regression specification.

The last point is worth spelling out in more detail. Let d_{pj} be the infant mortality rate of babies born to poor mothers (infant deaths divided by live births) in the j^{th} county, and let d_{nj} be the infant mortality rate

Equation (4) gives a multiple regression of d_j on eight variables (vectors): k_j, $k_j x_{pj}$, $k_j y_{pj}$, $(1 - k_j)\, y_{nj}$, $k_j w_{pj}$, $(1 - k_j)\, w_{nj}$, $k_j z_j$, and $(1 - k_j)\, z_j$. Attempts to estimate this equation would be plagued by severe problems of multicollinearity and by the absence of income-specific measures of certain variables such as the legal abortion rate. Therefore, we assume that the income-specific abortion rate (w_{ij}) is proportional to its weighted average ($w_{ij} = r_i w_j$). In addition, we assume that schooling of poor mothers in a given county is proportional to schooling of nonpoor mothers ($y_{pj} = s y_{nj}$). The actual equation that we fit is

$$d_j = \beta_0 + (\alpha_0 - \beta_0)\, k_j + \alpha_1 k_j x_{pj} + \delta_2 y_{nj} + \delta_3 w_j + \delta_4 z_j \quad , \qquad (5)$$

where δ_2 estimates $\alpha_2 k_j s_p + \beta_2 (1 - k_j)$, δ_3 estimates $\alpha_3 k_j r_p + \beta_3 (1 - k_j)\, r_n$, and δ_4 estimates $\alpha_4 k_j + \beta_4 (1 - k_j)$. The important point to note is that we employ k_j and the product of k_j and x_{pj} as independent variables in the regression. Thus, we employ a specification that explicitly recognizes that the impact on the observed infant mortality rate of policies aimed at the poor is larger the larger is the fraction of births to poor mothers ($\partial d_j / \partial x_{pj} = k_j \alpha_1$). Moreover, our specification yields a direct estimate of the impact parameter (α_1).[6]

II. Empirical Specification

A. Data and Measurement of Infant Mortality

Our basic data set is the Urban Institute's expanded version of the Area Resource File (ARF). The ARF is a county-based data service, prepared by Applied Management Sciences, Inc., for the Bureau of Health Manpower, Health postneonatal mortality. For instance, the former is considerably more sensitive to appropriate prenatal and obstetrical care than the latter (Lewit 1977). Another reason for our focus is that the neonatal mortality rate is much larger than postneonatal mortality rate; it was three times as large in 1971. Consequently, trends in the infant mortality rate are dominated by trends in the neonatal mortality rate. Obviously, one cannot hope to explain trends in the infant mortality rate without being able to explain trends in the neonatal mortality rate.

Separate regressions are fitted for white neonatal mortality and for black neonatal mortality. Black neonatal mortality rates are much higher than white rates. In a non-race-specific regression, one would enter the percentage of black births to control for race differences. But this variable would be highly correlated with the percentage of births to low-income women, schooling, and other independent variables. By fitting race-specific regressions, we reduce multicollinearity and allow the coefficients of the independent variables to vary between races. Linear regressions are estimated because a linear specification facilitates the aggregation of the two income-specific mortality rate functions given in Section I into a single equation for the entire population.

We use counties rather than states or Standard Metropolitan Statistical Areas (SMSAs) as the units of observation. SMSAs and states are very large and sometimes heterogenous. Income, schooling levels, medical resources and other variables may vary greatly within an SMSA or a state. Since counties are much more homogeneous, these problems are reduced in our research. A weakness with the use of counties is that the small size of some of these areas may mean that people may receive medical care outside the county.

cases of pregnancy resulting from rape or incest. By 1970, twelve states had enacted such statutes. Moreover, in 1970 four additional states enacted extremely liberal abortion laws which placed no legal restriction on the reasons for which an abortion may be obtained prior to the viability of the fetus (Center for Disease Control 1971). After the middle of 1970, there was no significant changes in abortion laws until 1973 when the Supreme Court ruled most restrictive state abortion laws unconstitutional. Concurrent with these reforms, the U.S. ratio of legal abortions per thousand live births rose from 4 in 1969 to 180 in 1972 and to 361 in 1977 (Center for Disease Control, 1971; 1972; 1974; Bureau of the Census, 1980).

B. Measurement of Independent Variables

Wherever possible, race-specific variables are employed in the regressions. Such variables are denoted with an asterisk. Except for the Medicaid and abortion measures, all variables are county-specific.

The number of active non-federal physicians per thousand population (MD) serves as a general proxy for the price and availability of medical care.[7] The roles of the percentage of births to poverty mothers (PB*) and the percentage of women of childbearing ages who had at least a high school education (HSP*) were discussed in Section I. Here we note that there are no direct measures of births to poor women, either at the county or at the national level. Therefore, we estimate the race-specific percentage of births to such women by assuming that the race-specific birth rate of poor women does not vary among counties and that the race-specific birth rate of nonpoor women does not vary among counties. Under these conditions, one can compute race-specific birth rates of poor and nonpoor women by regressing the race-specific low-income women is a superior variable to the actual percentage of such births, even if the latter were available.[10]

The social policy and program measures contain variables pertaining to Medicaid, maternal and infant care (M and I) projects, the use of organized family planning clinics by low-income women of childbearing ages, and abortion reform. In the case of prenatal and obstetrical care services, variations among states in the treatment of first-time pregnancies under Medicaid contribute to substantial variations in the percentage of pregnant low-income women whose medical care is financed by Medicaid. In particular, nineteen states cover <u>no</u> first-time pregnancies because their aid to families with dependent children (AFDC) programs do not cover "unborn children."[11] The treatment of first-time pregnancies of low-income women under Medicaid by the state in which the county is located is described by three dichotomous variables (MN, MU, MA). MN equals one for counties in states that cover first-time pregnancies only if no husband is present. MU equals one for counties in states that provide coverage if no husband is present or if the husband is present but unemployed and not receiving unemployment insurance. MA equals one for counties in states that provide coverage to all financially eligible women, regardless of the presence or employment status of the husband. The omitted category pertains to counties in states that cover <u>no</u> first-time pregnancies because their AFDC programs do not cover unborn children.[12]

The presence of an M and I project that reported positive births in 1971 is denoted by the dichotomous variable MI. A second measure of the impact of M and I projects is given by the number of births in an M and I project in 1971 as a percentage of our estimated births to low-income percentage rather than a fraction, the regression coefficients must be multiplied by 100 to obtain the vector of impact parameters (α_1) associated with policies aimed at low-income women [see equations (2), (4), or (5)].

The role of abortion reform is measured by a three-year average of the legal abortion rate for the period 1970-72 in the state in which the county is located (ARATE). The measure is an average of legal abortions performed on state residents per 1,000 live births to state residents and is derived from information reported by the Center for Disease Control (1971, 1972, 1974). It is assumed that abortions performed in the first half of a given year affect the neonatal mortality rate in the second half of that year. The computation also takes account of the extremely low legal abortion rates before the second half of 1970 in states that reformed their abortion laws in 1970. The assumptions required to estimate the abortion rate are somewhat arbitrary.[13] Therefore, in some regressions the rate is replaced by a dichotomous variable that identifies counties in states that reformed their abortion laws by the middle of 1970 (RA).

The final variable in the regressions is a three-year average of the infant mortality rate for the years 1966-68 (M66-68). Theoretically, this is an important variable to include in the analysis because programs such as M and I projects and subsidized family planning clinics for low-income women were designed to service target populations with poor health indicators. Consequently, estimates of their impacts are biased toward zero if the initial level of the mortality rate is omitted from the regression. In the case of abortion reform and liberal treatment of first-time pregnancies under Medicaid, the exclusion of the lagged mortality rate might overstate their contributions to reductions in neonatal mortality. This is because most of the

TABLE 1
Ordinary Least Squares Regressions of Neonatal Mortality Rates[a]

Independent Variable	Panel A: White Regressions (A1)	(A2)	(A3)	(A4)	Panel B: Black Regressions (B1)	(B2)	(B3)	(B4)
PB*	.037 (3.00)		.042 (3.45)		-.147 (-4.14)		-.133 (-3.83)	
MD	.144 (1.60)	.122 (1.35)	.124 (1.37)	.097 (1.07)	.227 (1.03)	.450 (2.05)	.172 (0.79)	.393 (1.84)
HSP*	-.015 (-1.14)	-.036 (-3.13)	-.013 (-0.96)	-.037 (3.22)	-.124 (-2.93)	-.017 (-0.49)	-.137 (-3.31)	-.035 (-1.08)
MAXPB*	.004 (0.39)	.016 (1.83)	-.003 (-0.39)	.008 (1.00)	.0004 (0.00)	-.014 (-0.53)	-.007 (-0.31)	-.010 (-0.46)
MUXPB*	.003 (0.29)	.010 (1.03)	.004 (0.44)	.012 (1.24)	-.038 (-1.78)	-.033 (-1.51)	-.041 (-1.97)	-.034 (-1.61)
MNXPB*	-.006 (-0.67)	.001 (0.13)	-.002 (-0.21)	.007 (0.77)	-.010 (-0.75)	-.032 (-2.47)	-.010 (-0.73)	(-.030) (-2.32)
MIXPB*	-.005 (-0.36)	-.011 (-0.87)	-.008 (-0.67)	-.017 (-1.37)	-.007 (-0.30)	-.003 (-0.15)	-.007 (-0.34)	-.005 (-0.20)
PMIBXPB*	-.022 (-1.06)	-.020 (-0.98)	-.015 (-0.76)	-.011 (-0.56)	-.033 (-1.19)	-.032 (-1.13)	-.037 (-1.35)	-.036 (-1.31)
UPXPB*	-.001 (-2.99)	-.0003 (-1.94)	-.001 (-2.80)	-.0003 (-1.58)	-.0003 (-0.86)	-.001 (-2.37)	-.0001 (-0.34)	-.001 (-1.76)
ARATE	-.004 (-3.25)	-.005 (-3.91)			-.009 (-2.25)	-.007 (-1.58)		
RA			-.549 (-3.43)	-.592 (-3.69)			-1.751 (-3.89)	-1.773 (-3.86)
M66-68	.274 (12.34)	.280 (12.53)	.281 (12.73)	.288 (13.04)	.260 (3.98)	.240 (3.61)	.235 (3.65)	.217 (3.31)
CONSTANT	7.554	9.400	7.045	9.094	27.184	17.998	27.618	19.238
$\bar{R}^2$	.315	.307	.317	.305	.125	.084	.149	.116
F	29.38	31.05	29.54	30.80	5.64	4.30	6.70	5.68

[a]t-ratios in parentheses. The critical t-ratio at the 5 percent level of significance is 1.64 for a one-tailed test. The eight F-ratios are significant at the 1 percent level.

the two relevant black regressions (B2 and B4). The exceptions in the white regressions pertain to the coefficients of the variables that identify liberal coverage of first-time pregnancies under Medicaid (MAXPB*, MUXPB*, MNXPB*). Given the high degree of intercorrelation among the variables in the regression and the imprecise measures used, the preponderance of negative effects is an important and impressive finding.

In terms of statistical significance, the hypothesis that no member of the set of social policy variables has a non-zero effect on neonatal mortality always is rejected at the 1 percent level. With respect to the four specific policies, in general abortion and the use of subsidized family planning services by low-income women have significant impacts, while Medicaid and M and I projects do not.[16] Specifically, for whites the abortion rate (ARATE) achieves significance at all conventional levels in regressions A1 and A2. A similar comment applies to the dichotomous variable that denotes abortion reform by the middle of 1970 (RA) in regressions A3 and A4. For blacks, RA is significant at all levels in regression B4, while ARATE is significant at the 6 percent level, but not at the 5 percent level, in regression B2. For whites, the interaction between the percentage of low-income women who use organized family planning clinics and the percentage of births to low-income women (UPXPB*) is significant at the 5 percent level in the first three regressions and at the 6 percent level in the fourth. For blacks, UPXPB* is significant at the 5 percent level in both regressions.

The significance of the abortion rate is notable because this variable is neither race- nor county-specific and must be computed subject to a number of somewhat arbitrary assumptions (see note 13). Therefore, it is probably subject to considerable measurement error, which biases its coefficient toward

TABLE 2

Contribution of Selected Factors to Reductions in Neonatal Mortality Rates, 1964-1977

	Panel A: Whites						Panel B: Nonwhites		
	1964 - 1977		1964 - 1971		1971 - 1977		1964-1977	1964-1971	1971-1977
Observed reduction in neonatal mortality rate (deaths per thousand live births)	7.5		3.2		4.3		11.8	6.9	4.9
Annually compounded percentage rate of decline in neonatal mortality rate	4.9		3.2		6.9		4.6	4.4	4.9
Contribution of selected factors to observed reduction in neonatal mortality rate	Reg.A1	Reg.A2	Reg.A1	Reg.A2	Reg.A1	Reg.A2	Reg.B2	Reg.B2	Reg.B2
MD	a	a	a	a	a	a	-0.2	-0.1	-0.1
PB*	0.4	b	0.4	b	a	c	b	b	b
HSP*	0.2	0.6	0.1	0.3	0.1	0.3	0.3	0.1	0.2
ARATE	1.5	1.7	0.4	0.4	1.1	1.3	2.5	0.6	1.9
UPXPB*	0.6	0.2	0.3	0.1	0.3	0.1	1.4	0.8	0.6
M and I Projects[d]	0.1	0.1	0.1	0.1	a	c	0.3	0.3	c
Medicaid[e]	a	-0.2	a	-0.2	a	c	0.5	0.5	c
Total explained reduction	2.8	2.4	1.3	0.7	1.5	1.7	4.8	2.2	2.6
Percentage explained	37.3	32.0	40.6	24.7	34.9	39.5	40.7	31.9	53.1

[a]Less than .1 in absolute value.

[b]Variable omitted from regression.

[c]No change in variable.

[d]Combined contribution of MIXPB* and PMIBXPB*.

[e]Combined contribution of MAXPB*, MUXPB*, and MNXPB*.

The increase in the use of organized family planning services by low-income women is the second-most important factor in reductions in nonwhite neonatality mortality for the entire period (1.4 deaths per thousand live births) and the most important factor in 1964-71 (0.8 deaths per thousand live births). For whites, the estimate of the contribution of family planning is sensitive to the inclusion in or exclusion from the regression of the percentage of births to poor women. When PB* is included, it dominates all the other factors except for abortion in the entire period and in the two subperiods. Its effect is weaker when PB* is omitted and is no larger than the impact of M and I projects in the earlier subperiod.

There is reason to believe that we understate the impact of the use of all family planning services as opposed to organized services by low-income women. This is because our measure excludes services delivered by private physicians. National trends in the percentage of low-income women serviced by private physicians contained in Family Planning Program Development (1974), Dryfoos (1976), and Cutright and Jaffe (1977) suggest that the estimates in Table 2 should be multiplied by a factor of 1.6. This adjustment makes family planning a more important contributor to neonatal death rate reductions than M and I projects in the computations based on regression A2. It suggests that the predicted reductions of 1.4 nonwhites deaths per thousand births and between 0.2 and 0.6 white deaths per thousand births due to family planning are conservative lower-bound estimates of the true impact.

M and I projects have small impacts on white neonatal mortality regardless of the regression specification employed. For nonwhites the effect is somewhat more substantial; it amounts to a decline of 0.3 deaths per thousand births for the years during which the projects were expanding. Of course the deaths per thousand births. If coverage is provided only if no husband is present, the differential is 3.2 deaths per thousand births. Given these differentials and the percentage of counties in each category, we estimate that the neonatal death rate of low-income nonwhites would fall by only 0.1 deaths per thousand births if <u>all</u> states covered all first-time pregnancies. This computation implies that any increase in the percentage of Medicaid-financed births between 1971 and 1977 had a minor impact, at best, on nonwhite neonatal mortality.

To summarize, our results, when combined with information on the use of the pill and the IUD by women of all income classes, provide a coherent explanation of the U.S. neonatal experience from 1964 to 1977. After a period of relative stability, the neonatal mortality rate began to decline following 1964 as a lagged response to the extremely rapid increase in the percentage of women who used the pill and the IUD between 1961 and 1964.[19] The decline was further fueled by the increase in the percentage of low-income women who used subsidized family planning services between 1965 and 1971 and by the dramatic rise in the legal abortion rate between 1969 and 1971. The acceleration in the rate of decline in the mortality rate between 1971 and 1977 was due primarily to the literal explosion of the abortion rate in that period. Medical care played an extremely limited role in this process, although we have weak evidence that M and I projects and Medicaid were of some benefit to nonwhites.

The above conclusions are subject to the qualification that we have no estimates of the impact of the pill and the IUD other than those that we infer through the use of family planning services by low-income women. They also are subject to the qualification that we cannot estimate the contribution except in cases where the woman's life was in danger. During that period, 28 states refused to pay for "medically necessary" abortions. The other 22 states continued to finance most abortions for Medicaid-eligible women by paying the Federal share as well as the state share. As a result the number of Federally financed abortions declined from approximately 250,000 per year before 1976 to less than 3,000 in 1978 (Trussell et al., 1980). Federal funding of abortions resumed temporarily in February 1980, pending a review by the U.S. Supreme Court of a ruling by Federal District Judge John F. Dooling, Jr. that declared the Hyde Amendment unconstitutional. In June 1980 the Supreme Court reversed Judge Dooling's decision and upheld the constitutionality of the Hyde Amendment.

In spite of the Hyde Amendment, the abortion rate continued to rise between 1977 and 1978. In part, this trend reflects the continued diffusion of a relatively new method of birth control. In part, it reflects a substitution of private for Federal funds by roughly 80 percent of women who would have been eligible for Federal financing in the absence of the amendment (Trussell et al., 1980). One can speculate, however, that the abortion rate would have risen at a more rapid rate between 1977 and 1978 in the absence of the Hyde Amendment. Given the recent Supreme Court ruling, the abortion rate for poor women probably will grow slower than otherwise and might even fall. According to our findings, this will retard the rate of decline in the neonatal mortality rate of the poor.

Taken at face value, the most striking implication of our study pertains to a constitutional ban on abortions. The current U.S. abortion rate is 400 abortions per thousand live births, while the rate in 1969 was 4 abortions per thousand live births. If a ban reduced the rate to its 1969 level, our

FOOTNOTES

[*]Research for this paper was supported by a grant from the Robert Wood Johnson Foundation to the National Bureau of Economic Research. We are indebted to Gary Becker, Willard Cates, Ann Colle, Joy Dryfoos, Linda Edwards, Victor Fuchs, Louis Garrison, Eugene Lewit, Robert Michael, Charlotte Muller, and Cathy Schoen for their comments on an earlier draft. In addition, we would like to thank Joy Dryfoos, Edward Duffy, Jack Hadley, and Letty Wunglueck for providing us with the data to conduct our research. This is a revision of a paper presented at the World Congress on Health Economics, Leiden University, the Netherlands, September 8-11, 1980. A preliminary version of the paper also was presented at a session sponsored by the American Economic Association and the Health Economics Research Organization at the annual meeting of the Allied Social Science Associations, Atlanta, Georgia, December 28-30, 1979. An expanded version of the paper is available on request. This version contains means and standard deviations of the variables in the regression analysis, sources for the extrapolations in Table 2, and a detailed description of the assumptions that underlie these extrapolations.

[1]The above computations are based on data contained in Bureau of the Census (1980).

[2]Descriptive and historical information concerning the social programs is available in the expanded version of this paper (available on request), and details on abortion reform are provided in Section II. Briefly, Medicaid, enacted in 1965 as Title XIX of the Social Security Act of 1935, is the

increasing the level of sexual activity in general. In spite of these factors, we feel that the hypothesis that abortion reform lower the infant mortality rates is very plausible. In part this is because we control for the use of family planning services in the regression analysis in Section III.

[5]We rejected the strategy of fitting a pooled cross-sectional time-series mortality function because county- or area-specific time series for a number of key independent variables are not available.

[6]A more general formulation of the above model can be developed by decomposing the observed infant mortality rate in the j^{th} county into rates associated with a variety of birth characteristics such as mother's age, mother's income, parity, birth weight, and legitimacy status of the birth:

$$d_j = \sum_{i=1}^{m} k_{ij} d_{ij} \quad .$$

In this equation k_{ij} is the fraction of births in the i^{th} category and d_{ij} is the infant mortality rate associated with that category. An example of one such category is an illegitimate, low-birth weight birth to a low-income, teenage mother with no previous live births. The social policies that we study can lower the observed infant mortality rate by lowering the fraction of births in high-risk categories (categories where d_{ij} is higher than on average) and by lowering the mortality rate in a given risk category (d_{ij}). Our regression estimates incorporate both effects because we do not control for characteristics such as the percentage of births to teenage mothers, the percentage of births to mothers over the age of 40, the percentage of illegitimate

233 for poor whites, 64 for nonpoor whites, 154 for poor blacks, and 95 for nonpoor blacks.

[9] In regressions not shown in Section III, median family income was included as an independent variable. Its coefficient was not significant.

[10] From equation (5), the reduced form effect of x_{pj} on d_j is

$$\frac{\partial d_j}{\partial x_{pj}} = (\alpha_0 - \beta_0) \frac{\partial k_j}{\partial x_{pj}} + \alpha_1 k_j + \alpha_1 x_{pj} \frac{\partial k_j}{\partial x_{pj}} \quad .$$

Note that

$$k^*_{jp} = \pi^*_j \, b^*_{jp}/b^*_j \quad ,$$

where b^*_{jp} is the race-specific birth rate of poor women in the j^{th} county. Clearly, this variable is not held constant in our regressions. Note that reduced form effects also could be estimated by expressing k_j as a function of a set of variables, including the social policies, in equation (5). This results in an extremely complicated functional form. Specifically, it includes the level of each social policy measure, the square of that measure, and its product with each of the other measures. Such an equation is not tractable from the standpoint of estimation.

[11] This list of states includes Arizona which has no Medicaid program.

[12] Our information on the treatment of first-time pregnancies under Medicaid by specific states was obtained from Letty Wunglueck of the Health Care

We have data for a70-2, a71 (the abortion rate during the entire year of 1971), and a72. For states that reformed their abortion laws before 1970, we assume that a69-2 + a70-1 = a70-2 due to the rapid upward trend in the abortion rate during this period. We also assume that the birth rate in the first half of 1971 equaled the birth rate in the second half of 1971, so that a71-1 + a71-2 = 2a71. Finally, we assume a72-1 = a72. Hence for these states

$$\bar{a} = (1/3) \quad (a70\text{-}2) + (1/3) \quad (a71) + (1/6) \quad (a72) \quad .$$

For states that reformed their laws in the middle of 1970, we assume a69-2 = a70-1 = 0. Hence, given the other two conditions used above,

$$\bar{a} = (1/6) \quad (a70\text{-}2) + (1/3) \quad (a71) + (1/6) \quad (a72) \quad .$$

Since the law for New York State had no residency requirements, states near New York are treated in the same manner as New York in the computation of $\bar{a}$.

[14]Age- and race-specific infant mortality rates for years prior to 1969 are not available on the Area Resource File.

[15]Space limitations prevent us from discussing the effects of poverty, schooling, and physicians in detail and from presenting additional specifications of the basic regressions. Note the following:

(a) The variables PB* and HSP* are highly correlated for whites (r = -.6) and for blacks (r = -.8). The insignificant regression coefficients of HSP* in regressions A1 and A3 are due in part to multicollinearity. This phenomenon may also contribute to the black results, although the explanation

[18]One might argue that we understate the impacts of schooling and poverty by holding constant an average infant mortality rate centered on the year 1967. Although it is reasonable to suppose that changes in social policies had no impacts until after 1967, this assumption may not be reasonable in the cases of schooling and poverty. This is because the trends in these variables were continuous from 1960 to 1970. To examine the roubustness of our conclusion that abortion dominates schooling and poverty, we re-estimated the contributions of these variables from regressions that exclude the lagged mortality rate. Although the contribution of schooling rises relative to the contribution in Table 2, it is still smaller than that of abortion. Note that if county-level fixed effects that lower mortality are positively correlated with schooling, we overstate the schooling coefficient by excluding the lagged mortality rate.

[19]Ryder (1972) reports that in 1961 the percentage of married women under age 35 who used the pill stood at approximately 3 percent. By 1964, it had increased to approximately 16 percent.

Center for Family Planning Program Development. Need for Subsidized Family Planning Services: United States, Each State and County, 1971. (New York: Planned Parenthood - World Population, 1974).

Cutright, Phillips, and Jaffe, Frederick S. Impact of Family Planning Programs on Fertility: The U.S. Experience. New York: Praeger Publishers, 1977.

Dryfoos, Joy G. "The United States National Family Planning Program, 1968-74." Studies in Family Planning, 7, No. 3 (March 1976).

Eisner, Victor et al. "Improvement in Infant and Perinatal Mortality in the United States, 1965-1973." American Journal of Public Health, 68, No. 4 (April 1978).

Fuchs, Victor R. "Some Economic Aspects of Mortality in Developed Countries." In The Economics of Health and Medical Care, edited by Mark Perlman. London: MacMillan, 1974.

Gortmaker, Steven L. "Poverty, Race, and Infant Mortality in the United States." Institute for Research on Poverty Discussion Paper No. 404-77, University of Wisconsin. (March 1977).

Lee, Kwang-Sun, et al. "Neonatal Mortality: An Analysis of the Recent Improvement in the United States." American Journal of Public Health, 70, No. 1 (January 1980).

Lewit, Eugene M. "Experience with Pregnancy, the Demand for Prenatal Care, and the Production of Surviving Infants." Ph.D. dissertation, City University of New York Graduate School, 1977.

Ryder, Norman B. "Time Series of Pill and IUD Use: United States, 1961-1970." Studies in Family Planning 3, No. 10 (October 1972).

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