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

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I. Introduction

Infant mortality has received strong concern. The poor performance of the United Stated relative to that of many other industrialized countries is embarrassing, particularly because infant mortality is associated with poor living conditions and economic underdevelopment. Large differences in rates between blacks and whites, and between high and low income groups, have encouraged social scientists to search for causes and ways of reducing the inequalities.

Economists have used mortality rates as measures of health. Death is important and final. Accurate records are generally available. The difficulties involved in weighting different diseases according to seriousness are avoided. The study of infant mortality is particularly important. Infancy is a time of high risk. A substantial amount of resources is devoted to prenatal and obstetrical care. High infant mortality rates may indicate poor health among American children. This may be associated with lower well-being in adolescence and adulthood. Therefore, infant mortality has received the attention of many health economists.

From 1964 to 1976, the infant mortality rate in the United States declined at an annually compounded rate of 4.1 percent per year. This was an extremely rapid rate of decline compared to 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 (4.7 percent per year from 1970 to 1976 versus 3.5 percent per year from 1964 to

1970).[1]

The period from 1964 to 1976 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, Fuchs 1974b, 1978, 1979; Rogers and Blendon 1977; Eisner et al. 1978), 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. By identifying the causes of the decline in mortality in the recent past, I hope to identify policies which, if pursued more intensively in the future, will reduce the high rate of infant mortality in the United States relative to other developed countries.[2]

Before discussing the methodology for estimating the impacts of social programs and policies on infant mortality, it is useful to provide some historical information concerning the developments being studied. With the exception of abortion reform, the policy initiatives at issue are aimed primarily at blacks and other low-income persons. Therefore, they are particularly relevant to infant mortality given that black babies and babies from low-income families are much more likely to die within the first year of life than white babies and babies from high-income families (for example, MacMahon, Kovar, and Feldman 1972; Fuchs 1974a, 1974b; Gortmaker 1977; Taffel 1978).

The Medicaid program was enacted in 1965 as Title XIX of the Social Security Act. It is a joint Federal-state program designed to finance the medical care services of specified groups of needy (low-income) persons.[3] Medicaid eligibility is linked to welfare eligibility. States that elect to participate in the program (all states except Arizona have elected to do so) must cover all families covered by the aid to families with dependent children (AFDC) program. In twenty-six states AFDC is restricted to families without a father present in the home. Twenty-four states extend AFDC and Medicaid coverage to families with unemployed fathers who do not receive unemployment compensation. Seventeen states cover under Medicaid all children under the age of 21 in families with incomes below the AFDC eligibility level, regardless of the employment status of the parents or the family composition.

It is well known that AFDC income eligibility levels vary considerably among states. This factor, together with the factors mentioned above, causes a considerable percentage of low-income persons to be ineligible for Medicaid and causes this percentage to differ among states. In the case of prenatal and obstetrical care services, which obviously are very relevant in birth outcomes, variations in the treatment of first-time pregnancies among states also contribute to variations in the coverage of low-income women under Medicaid. Nineteen states do not cover any first-time pregnancies under Medicaid because the AFCD programs of those states do not cover "unborn children." In ten states first-time pregnancies are covered only if no husband is present. In eleven states they are covered only if no husband is present or if the husband is present but unemployed and not receiving unemployment compensation. In ten states all first-time pregnancies of financially eligible women are covered.[4]

In 1963 Title V of the Social Security Act of 1935 was amended to include special grants for maternal and infant care (M and I) projects. The projects are designed to provide adequate prenatal and obstetrical care to reduce the incidence of mental retardation and other conditions caused by childbearing complications as well as to lower infant and maternal mortality. By 1971, fifty-three projects were fully operative, primarily in low-income areas (Bureau of Community Health Services n.d.).

It should be noted that Medicaid and M and I projects are not the only Federal programs to finance and deliver medical care to poor persons introduced since 1963. Other examples are the neighborhood health center program and children youth projects. I focus on Medicaid and M and I projects because Medicaid is by far the largest program and because the M and I program is the only one that deals exclusively with prenatal and obstetrical care. I also assume that prenatal and obstetrical care delivered through these programs is more likely to result in favorable birth outcomes than care delivered to indigent women by hospitals in cases when there is little or no prospect for public reimbursement.

Effective birth control methods, including abortion, and Federal subsidization of family planning services for low-income women can affect the infant mortality rate by reducing the costs of birth control and therefore lowering the incidence of unwanted births, many of which are high-risk births. Federal subsidization of family planning services for low-income women originated in the 1967 amendments to the Social Security Act. Federal efforts in this area were expanded by the Family Planning Services and Population Research Act of 1970 and by the 1972 amendments to the Social Security Act (Dryfoos 1976; Davis and Schoen 1978). These subsidies go to family planning clinics organized by hospitals, state and local health departments, Planned Parenthood, and other agencies such as M and I projects and neighborhood health centers. Dryfoos (1976) reports that the number of women serviced by family planning clinics increased from 900,000 in 1968 to 3.4 million in 1974.

Prior to 1967 all states of the United States had laws which permitted abortion only when it was necessary to preserve a pregnant woman's life. Beginning in 1967 some states started to reform these laws to increase the number of circumstances under which abortions could be performed. The reformed statutes legalized abortions if there was a substantial risk that continuance of the pregnancy would seriously impair the physical or mental health of the woman, or that the child resulting from the pregnancy would be born with a serious physical or mental defect, or in cases of pregnancy resulting from rape or incest. Three states enacted such statutes in 1967; two did so in 1968; four in 1969; and three in 1970. Moreover, in 1970, four additional states (Alaska, Hawaii, New York, and Washington) enacted extremely liberal abortion laws. The statutes of these four states 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). The process of abortion reform culminated in 1973 when the Supreme Court ruled most restrictive state abortion laws unconstitutional.

It is well known that women who use oral or intrauterine contraceptive techniques have much smaller probabilities of becoming pregnant than women who use other techniques (for example, Michael 1977). The percentage of married women under age thirty-five who used the pill or the IUD rose dramatically from 1961 to 1973. In 1961 this percentage stood at approximately 3 percent. By 1973 it had increased to approximately 35 percent (Ryder 1972, Michael 1977). Of course, the diffusion of the pill and the IUD did not result entirely from actions by the Federal government. The development is important for this research however, because its impact on infant mortality must be held constant when the effects of government programs are estimated. It is also important because it meant that an extremely effective method of birth control could be offered to low-income women by Federally subsidized family planning clinics.

Following this brief description of the social policies and programs at issue, I turn to a review of the literature on infant mortality in Chapter II of the thesis. The theoretical model is discussed in Chapter III. The data and measurement of variables are described in Chapter IV, and the multivariate infant mortality functions to be estimated are specified. Empirical results are contained in Chapter V.

FOOTNOTES

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

[2]Fuchs (1978) estimates that the U. S. infant mortality rate was almost 40 percent above the median rate for 15 other developed countries in 1974.

[3]My discussion of the Medicaid program is based on Davis and Schoen (1978), Health Care Financing Administration (1978), and Ghez and Grossman (forthcoming).

[4]My information on the treatment of first-time pregnancies under Medicaid by specific states was obtained from Letty Wunglueck of the Health Care Financing Administration.

II. Review of Literature

Most of the previous literature on infant mortality by economists and social scientists has emphasized the role of socioeconomic variables and the availability of medical care. Adelman (1963) explored the factors explaining differences in age specific fertility and mortality rates among countries. Thirty-four nations which provided data for at least part of the 1947-57 period were included in the cross-sectional mortality study. The natural log of income, percentage of the labor force employed outside of agriculture and the number of physicians per 10,000 inhabitants had negative coefficients in the regression explaining the natural log of the infant mortality rate. The physicians variable was not significant at the 5% level. A positive coefficient for the rate of income growth was also insignificant. The favorable effects of income were explained by its effects on nutrition and living standards. Countries with many people employed outside of agriculture are generally more urbanized and industrialized. This may be associated with better sanitation and more up to date medical care. The positive coefficient of the growth rate was surprising, particularly because it was negative in the child and adult mortality regressions.

Fuchs (1974) regressed the natural logarithm of the neonatal mortality rate (death occurring during the first 28 days after birth divided by the number of births) and the post-neonatal mortality rate (deaths occurring during the rest of the first year divided by births) on the natural logarithm of income to obtain the income elasticity of infant mortality. He did separate runs for 48 American states and 15 developed countries. The income elasticities were negative in both the 1937 and 1965 cross-sectional runs. However, there was some evidence of a decline in the income elasticity of mortality, particularly among countries. Post-neonatal mortality was more sensitive to income than neonatal mortality. To estimate the decline in age specific mortality due to technical change, as opposed to growth of gross national product, Fuchs merged 1937 and 1965 data and regressed the mortality rates on income and a dummy variable indicating time. The latter measures the time trend due to improvements in medical knowledge. Most of the decline in infant mortality was attributed to this variable by the regressions, which assumed that the income elasticity of infant mortality remained constant over the period studied. The predicted drop in mortality rates for countries exceeded the actual fall, implying that income elasticity really decreased between 1937 and 1965.

Williams (1974) studied the influence of medical care on infant mortality. He accounted for predicted risks of births based on birthweight. The latter is negatively related to mortality and provides the best estimates of survival probability. Williams pooled cross-sectional and time series data for 1952-1968. Using a sample of 1.5 million California births, he found that a measure of availability of obstetricians and gynecologists was negatively related to infant mortality. He also used time series data for American states. Williams concluded that the reasons previous studies often did not show a significant effect for medical care is that they did not control for high risks faced by light infants. Since the contribution of medical care may be greater when a birth is light and faces high risks, it might have been useful to include an interaction variable in the regressions.

Lewit (1977) examined the demand for prenatal care and its relationship to birthweight. Data related to 67,000 New York City births, occurring during the first half of 1970, was provided by the New York City Department of Health. Regression analysis showed that both parents' schooling was positively related to use of prenatal care. Black and Puerto Rican mothers were relatively underserved. Illegitimate births received little medical care. Availability of clinic time, and obstetricians and gynecologists in a health service area had no consistent significant effects on different measures of the utilization of medical services. Surprisingly, the existence of a federally sponsored Maternal and Infant Care project in a health service area was associated with lower use of medical care. This finding may be explained by noting that M and I projects were introduced in areas with high infant mortality rates and medically underserved populations. Mothers whose past pregnancy experiences suggest high risks received more medical care than others.

Certain characteristics were associated with low birthweight: being black, Puerto Rican, illegitimate, or born to a mother of foreign origin. Surprisingly, the education levels of both father and mother have negative, insignificant coefficients in regressions explaining birthweight. Measures of past success in producing viable children are positively related to birthweight. The most important finding was the value of prenatal care. Mothers who receive adequate care generally have substantially heavier babies than those without any medical services. Thus, we can explain much of the low birthweight associated with certain racial and ethnic groups, and child characteristics by noting differences in medical care.

Gortmaker (1977) worked with data from the National Natality and the National Infant Mortality Surveys, which were conducted by the National Center for Health Statistics for 1964 and 1965. He applied methods for the analysis of multidimensional contingency tables, assuming a multiplicative relationship among independent variables in the determination of mortality. Neonatal and post-neonatal mortality were examined separately by race. Demographic variables had predictable effects; young or old mothers and women with many previous births had high risks. Previous pregnancy loss was associated with higher death rates, as was low birthweight. Income and education had most of their favorable effects independently of birthweight; while health insurance, which has its effect by increasing prenatal care, works mainly through an increase in birthweight. This casts doubts upon the effectiveness of attempts to reduce socioeconomic mortality inequalities by improving access to prenatal care, since such care does not seem to strongly influence weight specific fatality rates. The previous conclusions were based on the white sample; the number of blacks was too small to allow one to conclude that coefficients differed by race. Much of the racial mortality rate variation can be explained by demographic, socioeconomic and birthweight differences. Blacks have surprisingly low birthweights. This is responsible for the high death rate--holding weight constant blacks do not face particularly high risks. This suggests that prenatal care may reduce excess mortality among blacks.

Brooks (1978) used path analysis on 1961-65 neonatal and post-neonatal mortality data for standard metropolitan statistical areas. The study aimed at discovering the extent to which the high mortality rates associated with low income and a high percentage of blacks in the population is due to lower availability of medical care. Brooks found that little of the mortality difference was explained by physician density, obstetricians and gynecologists per birth or per capita hospital beds. In fact, the percent black variable was positively correlated with physicians per capita. These results should be evaluated cautiously, since SMSAs are large, heterogenous units. The association between the percent black variable and the presence of doctors does not necessarily imply that many physicians work in black communities--they may be located primarily in white communities and suburbs.

Economists have developed models which are useful for the analysis of infant mortality. Becker (1965) developed the notion that households produce commodities using goods purchased on the market and family members' own time. The role of education as an environmental variable, increasing efficiency in household production, was explored by Michael (1972). Grossman (1972) applied these concepts to the production of health. Using economic theories of family behavior, Becker and Lewis (1973), Ben-Porath (1973), Ben-Porath and Welch (1976) explored the demand for children. They assumed that a family's utility function includes consumption of goods, the number of births and the average quality of a birth. Quality includes any characteristics which parents value. While Ben-Porath focused on how changes in the sex ratio of births (percent male or female) affected fertility, he noted that a similar model could be applied where the desired quality characteristic is the viability or survival probability of a child. Williams (1976) emphasized the role of uncertainty and observation of past success or failure in the production of surviving infants in the decision making process. Lewit (1977) assumed that a family's utility function included consumption of goods and the number of surviving children, where the survival ratio has an exogenous component and is positively related to expenditures on prenatal and infant care. The model which follows allows us to predict the effects of programs which lower the cost of medical care and birth prevention.

III. Theoretical Model

Assume a family's utility function is described by

$$U = U(n,x) = U(\Pi b,x),$$

where n represents the number of surviving children and is equal to the product of Π (the survival probability of a birth) and b (the number of births). X is a composite of all other commodities consumed. Its unit has been set so that the price is one dollar.

The production function of Π is $\Pi = \Pi(m)$, where m is a measure of medical care per birth. The conclusions which follow remain unchanged if m is a composite of inputs, including medical care, nutrition and parents' time provided for each birth, so long as medical care represents a fixed proportion of all expenditures on m. The derivative, $\frac{d\Pi}{dm}$ or Π_m, exceeds zero. The second derivative, $\frac{d^2\Pi}{dm^2}$ or Π_{mm}, will be shown to be negative when second order conditions are met.

The income constraint is $Y = x + bmp + F + bg$. P is the price or cost of m and F is the total cost of fertility control. Assume that F is a function of births averted. Define $\hat{b}$ as $b^* - b$, where b^* is equal to the number of births which would occur in the absence of any control. Let $F - f\hat{b} = f(b^* - b)$. So f is the cost of each birth averted. G is the fixed cost of a birth and it exceeds f by assumption. It contains no medical care expenditures.

Subjecting the utility function to the income constraint I get the Lagrangian Function

$$L = U(\Pi b,x) + \lambda\left[Y - x - bmp - f(b^* - b)\right] .$$

Partial differentiation gives us the following first order conditions:

$$\frac{\partial L}{\partial x} = U_x - \lambda, \text{ where } U_x \equiv \frac{\partial U}{\partial X} \ ,$$

$$\frac{\partial L}{\partial b} = \Pi U_n - \lambda(pm - f + g) = 0, \text{ where } U_n \equiv \frac{\partial U}{\partial n} \equiv \frac{\partial U}{\partial(\Pi b)} \ ,$$

and

$$\frac{\partial L}{\partial m} = b\Pi_m U_n - \lambda bp = 0 \ .$$

From the last two equations follows that

$$\frac{pm - f + g}{p} = \frac{\Pi}{\Pi_m} \tag{1}$$

This condition describes the rule for least cost production of surviving infants. The left numerator shows the expense of a birth after subtracting the reduction in fertility control expenditures associated with a birth. The ratio between the cost of a birth and the cost of a unit of medical care should equal that of their expected contributions to the number of surviving births. The equation, describing the rule for efficiency in production, shows that Π is not affected by income or the desired number of viable infants. The last point is important. More elaborate models, such as discussed later, may lead to other predictions.

Second Order Conditions

Lewit (1977) shows that the first order conditions obtained from minimizing the costs of producing a given number of surviving infants (n^*) is identical to those obtained from utility maximization. The Lagrangian Function and the first order conditions follow:

$$L = pmb - fb + fb^* + gb + \lambda(n^* - \Pi b)$$

$$\frac{\partial L}{\partial b} = pm - f + g - \lambda\Pi = 0$$

$$\frac{\partial L}{\partial m} = pb - \lambda b\Pi_m = 0$$

The second order condition for a minimum is expressed by the bordered Hessian

$$H = \begin{vmatrix} 0 & \emptyset_b & \emptyset_m \\ \emptyset_b & L_{bb} & L_{bm} \\ \emptyset_m & L_{bm} & L_{mm} \end{vmatrix} < 0$$

where $\emptyset = n^* - \Pi b = 0$

$$\emptyset_b = -\Pi$$

$$\emptyset_m = -b\Pi_m$$

$$L_{bb} = 0$$

$$L_{bm} = p - \lambda\Pi_m$$

$$L_{mm} = -\lambda b\Pi_{mm}$$

$$\therefore$$

$$H = \begin{vmatrix} 0 & -\Pi & -b\Pi_m \\ -\Pi & 0 & p-\lambda\Pi_m \\ -b\Pi_m & p-\lambda\Pi_m & -\lambda b\Pi_{mm} \end{vmatrix}$$

$$H = -(-\Pi) \begin{vmatrix} -\Pi & p-\lambda\Pi_m \\ -b\Pi_m & -\lambda b\Pi_{mm} \end{vmatrix}$$

$$-b\Pi_m \begin{vmatrix} -\Pi & 0 \\ -b\Pi_m & p-\lambda\Pi_m \end{vmatrix}$$

$$H = \Pi^2\lambda b\Pi_{mm} < 0 \quad \therefore \quad \Pi_{mm} < 0$$

Effects of Social Programs and Socioeconomic Variables

Abortion reform or the subsidization of family planning services for low income women will reduce the savings in fertility control costs which result from having another birth. This means that f falls. One can see the effect on mortality intuitively by examining (1). Cheaper birth control increases the costs associated with a birth (of a given "quality" or viability). So families will substitute quality of births for quantity. The left side of (1) increases. Therefore, medical care must increase. This raises Π and reduces Π_m (since $\Pi_{mm} < 0$).

This can be shown formally. Rewrite (1) and differentiate with respect to f, as follows:

$$(pm - f + g)\, p^{-1} = \Pi\Pi_m^{-1} , \tag{1a}$$

$$p^{-1}(pM_f - 1) = \Pi_m^{-1}\Pi_m M_f - \Pi\Pi_m^{-2}\Pi_{mm}M_f$$

where $$M_f \equiv \frac{\partial m}{\partial f} ,$$

$$-p^{-1} = -\Pi\Pi_m^{-2}\Pi_{mm}M_f ,$$

$$\therefore \quad M_f = \frac{p^{-1}}{\Pi\Pi_m^{-2}\Pi}$$

or $$M_f = \frac{\Pi_m^2}{p\Pi\Pi_{mm}}$$

M_f is negative because $\Pi_{mm} < 0$. Since $\Pi_m > 0$, Π rises as f falls resulting in

$$\Pi_f \equiv \frac{\partial \Pi}{\partial f} = \Pi_m M_f = \frac{\Pi_m^{\;3}}{p\Pi\Pi_{mm}} \quad .$$

So the simple model predicts that liberalization of abortion laws and the subsidization of contraception will lower infant mortality.

Medicaid and Maternal and Infant Care Projects lower the costs of medical care for needy women. To examine their impact on infant mortality, differentiate (1a) with respect to p.

$$-p^{-2}(pm - f + g) + p^{-1}(m + pM_p) = \Pi_m^{\;-1}\Pi_m M_p - \Pi\Pi_m^{\;-2}\Pi_{mm}M_p$$

where $M_p \equiv \dfrac{\partial m}{\partial p}$,

$$M_p = \frac{p^{-2}(g - f)}{\Pi\Pi_m^{\;-2}\Pi_{mm}} \quad ,$$

$$M_p = \frac{\Pi_m^{\;2}(g - f)}{p^2\Pi\Pi_{mm}} < 0, \text{ since } \Pi_{mm} < 0 \text{ and } g > f \;,$$

$$\Pi_p = \frac{\Pi_m^{\;3}(g - f)}{p^2\Pi\Pi_{mm}} < 0 \qquad (2)$$

Medical care per birth and the survival probability increase as programs reduce the cost of medical care. This can be understood by noting that the cost of a birth, pm - f + g, does not fall as rapidly as pm does, since Medicaid does not lower the fixed costs of a birth. So the costs of improving infant quality decline relative to the costs of quantity of births. Since physicians per capita is generally associated with lower medical costs (holding other variables constant), one would expect a high concentration of doctors to lower infant mortality rates. The presence of many physicians could reduce the time costs of a visit, as well as reducing the monetary expense.

Some economists, including Michael (1972), Grossman (1972) and Inman (1976) have tested the hypothesis that education increases non-market productivity. Lewit (1977) developed models in which schooling affects the productivity of expenditures aimed at improving child quality. Allow education to act as a factor augmenting environmental variable. Assume that $m^* = f(e)m$, where m^* represents effective units of medical care and e is the number of years of schooling. P^*, the price of an effective unit of medical care is equal to $\frac{p}{f(e)}$. So education will lower the cost of m^* if $\frac{df(e)}{de} > 0$. We can predict the effect of a change in schooling by using a form of (2),

$$\Pi_{p^*} = \frac{\Pi_{m^*}^3 (g - f)}{p^2 \Pi_{m^*m^*}}$$

where $\Pi_{p^*} \equiv \frac{\partial \Pi}{\partial p^*}$, $\Pi_{m^*} \equiv \frac{\partial \Pi}{\partial m^*}$ and $\Pi_{m^*m^*} \equiv \frac{\partial^2 \Pi}{\partial m^{*2}}$, and multiplying by $\frac{\partial p^*}{\partial e}$ or $-pf(e)^{-2}f'(e)$ to get

$$\Pi_e = \Pi_{p^*} \frac{\partial p^*}{\partial e} = -\frac{\Pi_{m^*}{}^3 (g - f)}{p^2 \Pi_{m^*m^*}} \cdot p \frac{f'(e)}{f(e)^2},$$

where $\Pi_e \equiv \frac{\partial \Pi}{\partial e}$.

$\Pi_e > 0$ since $\Pi_{p^*} < 0$ and $\frac{dp^*}{de} < 0$. So if an increase in education multiplies the productivity of all units of m equally it will raise Π and m^*. Since $\frac{\partial m^*}{\partial e} = \frac{\partial m}{\partial e} f(e) + mf'(e)$, the effect on m is uncertain. It depends on whether the increase in education enhances the effectiveness of currently consumed m less or more than is needed to meet the new quantity demanded of m*.

Education may also exert its impact through an increase Π_o, the survival ratio when m equals zero. To simplify the analysis, first assume that Π_m is independent of e, at all levels of m. Differentiating (1a) leads to the following:

$$(pm - f + g)p^{-1} = \Pi\Pi_m^{-1}$$

$$M_e = \Pi_m^{-1}\,\Pi_e\big|^{\overline{m}} + M_e - \Pi\Pi_m^{-2}\Pi_{mm}M_e, \quad \text{where } M_e \equiv \frac{\partial m}{\partial e}$$

and $\qquad \Pi_e\big|^{\overline{m}} \equiv \dfrac{\partial \Pi}{\partial e}$ holding m constant,

$$M_e = \frac{\Pi_m \Pi_e\big|^{\overline{m}}}{\Pi\Pi_{mm}} < 0 ,$$

$$\Pi_m M_e = \frac{\Pi_m^2 \Pi_e\big|^{\overline{m}}}{\Pi\Pi_{mm}} = \text{decline in the survival ratio due}$$

to change in m resulting from change in e, $\Pi_e\big|^{\overline{m}}$ = direct increase in the survival rate from a change in e,

$$\Pi_e = \Pi_e\big|^{\overline{m}} + \frac{\Pi_m^2 \Pi_e\big|^{\overline{m}}}{\Pi\Pi_{mm}} \gtrless 0 \quad \text{as} \quad -\frac{\Pi_m^2}{\Pi\Pi_{mm}} \gtrless 1 .$$

The effect of an increase Π_o is uncertain.

If Π_o is raised by education it is likely that Π_m will fall at some or all levels of m. This is probable because Π has a maximum value of one. If Π_m falls by a constant proportion, for all units of m consumed before the increase in schooling, then the effect of this change would be similar to that of an increase in p*. It would lead to a fall in Π. If the decline in Π_m gets larger as more m is applied, then the incentive to substitute away from quality is greater. This effect should be added to those described previously.

Schooling may also influence mortality by raising the opportunity cost of parents' own time. The market value is enhanced by greater productivity. If parents' time is only involved in fixed costs of a birth, then education will increase g. The effect of this is similar to that of a reduction in f. The cost of quantity of births rises relative to the price of quality inputs, causing an increase in the survival probability. If m is a composite good which includes parents' time in fixed proportion to other inputs, then additional schooling will raise the price of m. This was earlier shown to reduce Π. Since education operates in different ways and we do not know the magnitude of all the different effects, the impact of schooling remains ambiguous.

Uncertainty

So far it was assumed that a family could predict Π perfectly. Now let Π vary stochastically around its mean value, determined by m. If parents decide the value of m before having children, then introducing uncertainty will not have a useful effect on the basic model, as Ben-Porath (1973) has shown. Allow the family to alter m after observing the outcome of previous births. If a value of Π below the anticipated level does not affect expectations about the production function of Π, then high infant mortality will not change m for future births. Real income falls with infant fatalities. The number of births will rise as parents spread the income loss to commodities other than children. However, a change in income or the number of additional surviving births desired, will not affect the optimal survival ratio for future births. So previous experience will alter m only if it changes a family's belief about the function $\Pi = \Pi(m)$. Possible ways in which a change in the expected productivity of m have an impact on mortality have been discussed in relation to education. The overall effect is ambiguous.

In the empirical section, a county's lagged infant mortality rate is used as an independent variable. To the extent that a high rate is associated with a large number of families altering their expectations, it may have the effects of a change in the production function of Π. However, the lagged rate might not involve a substantial learning effect. It may indicate that previously known factors may make a high mortality rate optimal. Barring a major change in these factors, one would expect past and current mortality to be positively correlated.

Specific Functional Form

The effects of social programs will now be analyzed using a specific functional form for the production of quality, $\Pi = 1-\alpha e^{-Bm}$. Alpha is assumed to have a positive value which does not exceed one. This equation has a number of useful properties. Π moves from a minimum of one minus alpha, when $m = 0$, towards a maximum of one, as m gets very high. Π_m is positive and Π_{mm} is negative. The derivation of first order conditions follows:

$$\Pi = 1 - \alpha e^{-Bm}$$

$$\Pi_m = \alpha B e^{-Bm}$$

$$\Pi_{mm} = -\alpha B^2 e^{-Bm}$$

$$\Pi \Pi_m^{-1} = \alpha^{-1} B^{-1} e^{Bm} - B^{-1}$$

$$(pm - f + g)p^{-1} = \Pi \Pi_m^{-1} \quad \text{from first order condition (1a)}$$

$$(pm - f + g)p^{-1} = B^{-1} e^{Bm} - \alpha^{-1} B^{-1}$$

To examine the effects of subsidized contraception, differentiate with respect to f.

$$M_f - p^{-1} = e^{Bm} M_f$$

$$p^{-1} = M_f - e^{Bm} M_f$$

$$p^{-1} = M_f (1 - e^{Bm})$$

$$M_f = \frac{p^{-1}}{1 - e^{Bm}}$$

To show that M_f is negative, it has to be proven that $e^{Bm} > 1$.

$$\Pi = 1 - \alpha e^{-Bm}$$

$$\Pi - 1 = -\alpha e^{-Bm}$$

$$1 - \Pi = \alpha e^{-Bm}$$

$$\alpha e^{Bm} = \frac{1}{1 - \Pi} > 1 \quad \text{since} \quad (1 - \Pi) < 1$$

$$e^{Bm} > 1 \quad \text{since} \quad \alpha < 1 .$$

Analyze the effect of Medicaid and Maternal and Infant Care Projects by differentiating (1a) with respect to p.

$$(pm - f + g)p^{-1} = B^{-1} e^{Bm} - \alpha B^{-1}$$

$$p^{-2}(f - g) + M_p = e^{Bm} M_p$$

$$p^{-2}(f - g) = M_p (e^{Bm} - 1)$$

$$M_p = \frac{p^{-2}(f - g)}{e^{Bm} - 1} < 0 , \text{ since } e^{Bm} > 1$$

General Utility Function

In the simple model, utility only depended on x and b. The equation $U = U(x, \Pi b, (1 - \Pi)b, b, h_1(\Pi), h_2(\Pi))$ shows a more expanded utility function. The distress brought on by fatalities depends on $(1 - \Pi)b$, the number of deaths. Births involve pain and other psychic costs. The health of a mother is affected by inputs into the quality of births and is, therefore, a function of Π. The health of surviving children is similarly affected by the use of medical care, nutrition and parents' time to increase the survival probability of a birth. The last two relationships are expressed by $h_1(\Pi)$ and $h_2(\Pi)$.

Earlier it was hypothesized that families try to produce surviving infants at the lowest possible cost (1). Π did not depend on income. In this model increasing Π could increase utility by reducing psychic costs of fatalities and childbirth and improving the health of the mother and surviving children. So parents may increase Π as income rises, even when this would not be justified for the purpose of producing viable births at a minimum cost. The model also encourages an increase in Π as p falls since the price reduction allows one to gain desired benefits at a lower expense.

Abortion reform and subsidized contraception may influence mortality in more ways than suggested by the basic model.[1] It stated that a fall in birth prevention costs raised the fixed costs of a birth by reducing the associated savings on fertility control. So parents will have fewer births and produce the desired number of viable infants by increasing the survival ratio through greater use of medical care.

The programs may also work by preventing high risk births. This is especially true for abortion reform which allows a woman to terminate a pregnancy if she suffered an illness during the period. Improved detection techniques may aid parents in determining if a birth is likely to result in an unhealthy infant.

It is important to look at how fertility control may affect the proportion of infants born to mothers of different socioeconomic groups. If relatively educated mothers, who are likely to have healthier births (according to the general utility function and previous empirical studies), are most inclined to reduce births, then the availability of fertility control may increase infant mortality rates. Educated mothers may have low costs of obtaining information about contraception and "tastes" which encourage use. However, since the variable used in this study involves subsidized family planning for the needy, one expects a negative coefficient in the mortality runs. The conventionally assumed heavy contraception use by highly schooled people suggests that abortion will be used frequently by the less educated, lowering the infant mortality rate.

It is also useful to consider how abortion and contraception will affect the age distribution of mothers. Women who want to pursue careers may postpone childbirth until the thirties or forties. The ability to terminate a pregnancy if tests detect Down's syndrome or other diseases associated with age may encourage late childbirth. These factors probably operate with a large time lag and are not important in this study.

Finally, fertility control methods may lead to increased sexual activity. However, this will affect infant mortality only to the extent that even people who do not prevent births participate more in sex. Examining all of the possible effects on our variables, it seems most likely that they will have negative effects on infant mortality in this study.

Ideally, to make predictions about the impacts of policy-manipulable variables such as the price or availability of prenatal and obstetrical care, one would estimate the production function and the input demand functions, perhaps by employing simultaneous equations estimation methods. In practice, due to data limitations and other problems, this procedure is very difficult to implement. Instead, in using almost any kind of data, the fitted survival function must incorporate elements of production and demand. That is, it is a mixture of a demand function for the probability of survival and a production function of this probability.[2] I term this hybrid equation an outcome equation from now on.

To measure the relative importance of the relevant factors in the recent U. S. infant mortality experience, I perform a cross-sectional regression analysis of variations in infant mortality rates among counties of the United States in 1971. My procedure capitalizes on variations in social programs among counties at a moment in time. Thus it provides a set if 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.

This 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.[3] 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 I 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. I incorporate this insight into the 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 of babies born to nonpoor mothers. As an identity,

$$d_j = k_j d_{pj} + (1 - k_j) d_{nj} \quad ,$$

where d_j is the observed infant mortality rate and k_j is the fraction of births to poor mothers. Specify behavioral equations for d_{pj} and d_{nj} as follows:

$$d_{pj} = \alpha_0 + \alpha_1 x_{pj} + \alpha_2 y_{pj} + \alpha_3 w_{pj} + \alpha_4 z_j \qquad (2)$$

$$d_{nj} = \beta_0 + \beta_2 y_{nj} + \beta_3 w_{nj} + \beta_4 z_j \ . \qquad (3)$$

In these equations, x_{pj} is a vector of policy variables that affects the mortality rate of poor babies alone such as Medicaid; w_{ij} ($i = p, n$) is a vector of policy variables that affects both groups such as the group specific abortion rate (legal abortions per thousand live births); y_{ij} refers to a group specific vector of basic determinants of infant mortality such as mother's schooling; and z_j is a vector of variables that has the same value for each group such as physicians per capita. Since there are no data in income specific mortality rates at the county level, substitute equations (2) and (3) into equation (1) to obtain:

$$d_j = \beta_0 + (\alpha_0 - \beta_0)\ k_j + \alpha_1 k_j x_{pj} + \alpha_2 k_j y_{pj} + \beta_2 (1 - k_j)$$
$$y_{nj} + \alpha_3 k_j w_{pj} + \beta_3 (1 - k_j)\ w_{nj} + \alpha_4 k_j z_j$$
$$+ \beta_4 (1 - k_j)\ z_j\ . \tag{4}$$

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, I assume that the income specific abortion rate (w_{ij}) is proportional to its weighted average ($w_{ij} = r_i w_j$). In addition, I 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 I fit is

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

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 I employ k_j and the product of k_j and x_{pj} as independent variables in the regression. Thus, I 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, this specification yields a direct estimate of the impact parameter (α_1).

FOOTNOTES

[1]Eugene Lewit emphasized the importance of looking at the different ways in which abortion and contraception influence mortality.

[2]For a discussion of the inherent problems in the estimation of a full production-demand model and a detailed interpretation of a health outcome function as a hybrid equation, see Edwards and Grossman (1980, forthcoming). They deal with child and adolescent health, but their treatment also pertains to infant health.

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

IV. Empirical Specification

A. Data and Measurement of Infant Mortality

The basic data set here is the Urban Institute's expanded version of the Area Resource File (ARF).[1] The ARF is a county-based data service, prepared by Applied Management Sciences, Inc., for the Bureau of Health Manpower, Health Resources Administration, U. S. Department of Health, Education and Welfare. It incorporates information from a variety of sources for 3,078 counties in the United States. These counties can also be aggregated into larger geographic areas such as county groups, Standard Metropolitan Statistical Areas, and states. Demographic and socioeconomic characteristics are taken from the 1970 Census of Population. Socioeconomic characteristics of women ages 15 to 49 come from the 1970 Census of Population, Women of Childbearing Age Tape. Deaths by age, race, and sex for the years 1969 through 1976 are obtained from the National Center for Health Statistics (NCHS) Mortality Tape. Births by race for those years are obtained from the NCHS Natality Tape. Health manpower and facilities come from the American Medical Association, the American Hospital Association, and other sources. We have added measures of the social policies and programs discussed previously to the ARF from sources indicated in Chapter IV.

There are two components of infant mortality: neonatal mortality and post-neonatal mortality. Neonatal mortality refers to deaths of infants within the first 27 days of birth. Post-neonatal mortality refers to deaths of infants between the ages of 28 and 364 days. Neonatal deaths are usually caused by congenital anomalies, prematurity, and complications of delivery; while post-neonatal deaths are usually caused by infectious diseases and accidents.

Empirical analysis is limited here to the neonatal mortality rate, defined as neonatal deaths per thousand live births. Since the causes of the two types of infant deaths are dissimilar, socioeconomic variables and social programs are likely to have different effects on each. Specifically, the social policy variables that are studied here are more relevant to neonatal mortality than to post-neonatal mortality. For instance, the former is considerably more sensitive to appropriate prenatal and obstetrical care than the latter (Lewit 1977). Moreover, since it is easier to predict neonatal risks than post-neonatal risks, high neonatal risks are likely to encourage women to have an abortion and to encourage low-income women to take advantage of subsidized family planning services. Another reason for this focus is that the neonatal mortality rate is much larger than post-neonatal 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. Ninety percent of the decline in infant mortality which occurred since 1970 involves neonatal mortality. 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 Chapter III into a single equation for the entire population.

Counties rather than states of Standard Metropolitan Statistical Areas (SMSAs) are used 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 homogenous, these problems are reduced in this 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. Moreover, the small number of births in certain counties may increase the importance of random movements or "noise" in the determination of regression coefficients.

These problems with county data can be reduced by including in the regressions only counties with a population of at least 50,000 persons in 1970. A county must also have at least 5,000 blacks for inclusion in the black regressions. There are 679 counties in the white regressions and 359 counties in the black regressions. The 679 counties include 141 million whites and the 359 counties have 18 million blacks. In addition to selecting large counties, I attenuate random elements by employing a three-year average of the race-specific neonatal mortality rate for the period 1970-72 as the dependent variable and by estimating weighted regressions, where the set of weights is the race-specific number of births in 1971.

Studied here is neonatal mortality for the period 1970-72 because measures of all independent variables are available for a year in that period or for 1969. In addition, it provides an ideal time frame to estimate the impact of abortion reform, which proceeded at a rapid pace between 1967 and the middle of 1970. Prior to 1967 all states of the United States had laws which permitted abortion only when it was necessary to preserve a pregnant woman's life. Beginning in 1967 some states started to reform these laws to increase the number of circumstances under which abortions could be performed. The reformed statutes legalized abortions if there was a substantial risk that continuance of the pregnancy would seriously impair the physical or mental health of the woman, or that the child resulting from the pregnancy would be born with a serious physical or mental defect, or in 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 were 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 (Center for Disease Control 1971, 1972, 1974).

B. Measurement of Independent Variables

Table 1 contains definitions, means, and standard deviations of the dependent and independent variables in the regressions. Wherever possible, race-specific variables are employed in the regressions. Such variables are denoted with an asterisk.

The number of active non-federal physicians per thousand population (MD) serves as a general proxy for the price and availability of medical care. 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 fully in Chapter III. Here it should be noted that there are no direct measures of births to poor women, either at the county or at the national level. Therefore, the race-specific percentage of births to such women are estimated 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 birth rate (b_j^*, the ratio of births to women ages 15 to 44) on the race-specific fraction of women in poverty (π_j^*):

$$b_j^* = \gamma_0^* + \gamma_1^* \pi_j^* \quad . \tag{6}$$

TABLE 1
Definitions, Means and Standard Deviations of Variables[a]

Variable Name	Definition
NM70-72*	Three-year average neonatal mortality rate for the period 1970-72; deaths of infants less than 28 days old per 1,000 live births (μ_w = 12.729; σ_w = 2.076; μ_b = 21.477; σ_b = 3.988)
PB*	Estimated percentage of births to mothers with family incomes less than the poverty level for the period 1969-71 (μ_w = 21.324; σ_w = 8.388; μ_b = 35.188; σ_b = 11.235)
HSP*[b]	Percentage of women aged 15 to 49 who had at least a high school education in 1970 (μ_w = 62.927; σ_w = 7.238; μ_b = 44.096; σ_b = 8.527)
MD	Active non-federal physicians per 1,000 population in 1971 (μ_w = 1.505; σ_w = 0.987; μ_b = 1.954; σ_b = 1.220)
MAXPB*	Dichotomous variable that equals one if county is in a state that covers all first-time pregnancies to financially eligible women under Medicaid (MA) multiplied by PB* (μ_w = 8.892; σ_w = 10.850; μ_b = 7.104; σ_b = 12.657)
MUXPB*	Dichotomous variable that equals one if county is in a state that covers first-time pregnancies under Medicaid only if no husband present or if husband present but unemployed and not receiving unemployment compensation (MU) multiplied by PB* (μ_w = 2.810; σ_w = 7.521; μ_b = 3.857; σ_b = 10.219)
MNXPB*	Dichotomous variable that equals one if county is in a state that covers first-time pregnancies under Medicaid only if no husband present (MN) multiplied by PB* (μ_w = 2.284; σ_w = 7.851; μ_b = 7.536; σ_b = 18.185)
MIXPB*	Dichotomous variable that equals one if the county had an M and I project that reported births in 1971 (MI) multiplied by PB* (μ_w = 5.339; σ_w = 9.390; μ_b = 16.152; σ_b = 16.577)
PMIBXPB*	Births in M and I projects in 1971 as a percentage of births to women with low income (PMIB) multiplied by PB* (μ_w = 2.174; σ_w = 5.086; μ_b = 8.470; σ_b = 12.670)

(continued on next page)

TABLE 1 (concluded)

Variable Name	Definition
UPXPB*	Percentage of women aged 15 to 44 with family income equal to or less thsn 150 percent of the poverty level who were served by organized family planning clinics in fiscal 1971 (UP) multiplied by PB* (μ_w = 639.506; σ_w = 521.843; μ_b = 1,435.559; σ_b = 741.955)
ARATE	Three-year average abortion rate for the period 1970-72 of state in which county is located; legal abortions performed on state residents per 1,000 live births to state residents (μ_w = 96.607; σ_w = 80.497; μ_b = 87.156; σ_b = 77.518)
RA	Dichotomous variable that equals one if county is in a state that reformed its abortion law by 1970 (μ_w = 0.369; σ_w = 0.483; μ_b = 0.358; σ_b = 0.480)
M66-68	Three-year average infant mortality rate for the period 1966-68, not race or age specific (μ_w = 21.517; σ_w = 3.553; μ_b = 24.380; σ_b = 3.867)

[a]Variable names ending in an asterisk(*) indicate variables that are race specific. The symbols μ_w, σ_w, μ_b, and σ_b denote the white mean, the white standard deviation, the black mean, and the black standard deviation, respectively. The white data pertain to 679 counties, while the black data pertain to 359 counties. Means and standard deviations are weighted by the race-specific number of births in 1971.

[b]Variable is available only for whites and non-whites as opposed to whites and blacks.

The regression intercept (γ_0^*) gives the birth rate of nonpoor women, and the sum of γ_0^* and γ_1^* gives the birth rate of poor women.

After fitting the regressions for white and blacks, I obtained the race-specific percentage of births to poverty women as

$$PB^* = 100 \left[(\gamma_0^* + \gamma_1^*) \quad \pi_j^* \; / \; (\gamma_0^* + \gamma_1 \pi_j^*)\right] . \tag{7}$$

It is clear that PB* is a monotonically-increasing, although non-linear, functionof the fraction of the population in poverty. Therefore, the regression coefficient of PB* summarizes the impact of poverty on infant mortality. Since poverty and family income are highly correlated, the latter is omitted from the regression.[2]

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."[3] 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 no first-time pregnancies because their AFDC programs do not cover unborn children.[4]

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 women in 1971 (PMIB). Both variables are employed because the M and I program is relatively small; there were only 53 projects in 1971. The presence of an M and I project and the number of births in the project were taken from Bureau of Community Health Services (n.d.).

The impact of variations in Federal, state, and local subsidization of family planning services is given by the percentage of women ages 15-44 with family incomes equal to or less than 150 percent of the poverty level who were served by organized family planning clinics in fiscal 1971 (UP). These clinics are organized by hospitals, state and local health departments, Planned Parenthood, and other agencies such as neighborhood health centers. This variable was taken from a survey conducted by the National Center for Health Statistics and by the technical assistance division of Planned Parenthood, then known as the Center for Family Planning Program Development and now known as the Alan Guttmacher Institute (Center for Family Planning Program Development 1974). It excludes family planning services delivered to low-income women by private physicians.

Dryfoos (1976) reports that almost all clients of family planning clinics use the pill or IUD. Therefore, the percentage of low-income women who are served by these clinics is positively related to the percentage of low-income women who select the pill or the IUD as contraceptive techniques. There is no information on the use of these techniques by other women at the county or state level, but it is known that women with at least a high school education are more likely to use them. Therefore, part of the observed effect of schooling in the regressions reflects the impact of the diffusion of the pill and the IUD on neonatal mortality.

The Medicaid, M and I, and family planning variables are interacted with the race-specific percentage of births to women in poverty. Since PB* is a percentage rather than a fraction, the regression coefficients must be multiplied by 100 to obtain the vector of impact parameters (α_I) 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 average 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.[5] 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). 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 upwards 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 states that reformed their abortion laws by 1970 and enacted generous Medicaid programs were liberal states with relatively large social welfare programs and probably lower than average infant mortality rates. In general, the use of the lagged rate as an independent variable controls for unmeasured determinants of infant mortality that are correlated with the included variables.

Ideally one would control for the initial level of mortality with the race-specific neonatal mortality rate for an interval shortly before 1964. Such a measure is not, however, available in the

Area Resource File. To the extent that the social programs at issue had an impact on mortality between 1966 and 1968, I understate their effects.[6]

FOOTNOTES

[1]I owe a special debt of gratitude to Jack Hadley for providing me with the Urban Institute's expanded version of the ARF.

[2]The regression equation for whites is

$$b_j^* = .064 + \underset{(t=15.90)}{.169\pi_j^*}, \quad \bar{R}^2 = .269, \ n = 679 .$$

The regression for blacks is

$$b_j^* = .095 + \underset{(t=6.98)}{.059\pi_j^*}, \quad \bar{R}^2 = .118, \ n = 359 .$$

In each regression, the dependent variable is a three-year average of the birth rate for the period 1969-71. The regressions are weighted by the square root of the race-specific number of women ages 15-44 in 1970. The poverty variable pertains to the fraction of families below the poverty level, rather than to the fraction of women ages 15-44. The latter variable is not available on a race-specific basis. Another reason for the use of the fraction of families in poverty is that it facilitates the trend analysis in Chapter III. The ratios of births per thousand women ages 15-44 implied by the regressions are 233 for poor whites, 64 for nonpoor whites, 154 for poor blacks, and 95 for nonpoor blacks.

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

[4]This information on the treatment of first-time pregnancies under Medicaid by specific states was obtained from Letty Wunglueck of the Health Care Financing Administration. Note that first-time pregnancies of young mothers who are themselves dependents in AFDC families would be covered under Medicaid in spite of the above provisions. States in one of the three categories, however, cover a larger percentage of first-time pregnancies than other states. Note also that the impact of Medicaid on neonatal mortality depends on the percentage of second- and higher- order births covered and on the quantity and quality of services provided. There are no data on these variables.

[5]Suppose that the neonatal mortality rate (nm_{jt}) and the legal abortion rate (a_{jt}) are measured in half-year intervals. Let the relationship between the two be

$$nm_{jt} = \beta + \delta a_{jt-1} .$$

Aggregate and average this equation over three years (six half years) to obtain

$$\overline{nm}_j = \beta + \delta\overline{a}_j \quad ,$$

where

$$a_j = \left(\sum_{t=0}^{5} a_{jt-1}\right) / 6 \quad .$$

The neotatal mortality rates pertain to the period from the first half of 1970 (70-1) to the last half of 1972 (72-2). Therefore ignore the county subscript, and write $\overline{a}$ as

$$a = a69\text{-}2 + a70\text{-}1 + a70\text{-}2 + a71\text{-}1 + a71\text{-}2 + a72\text{-}1$$

I 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. I also assume that the birth rate in the first half of 1971 equalled the birth rate in the second half of 1971, so that a71-1 + a71-2 = 2a71. Finally, I assume a72-1 = a72. Hence for these states

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

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

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

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 $\overline{a}$.

[6]Note that given lags between the enactment of the social programs at issue and their implementation and given lags between implementation and impacts on neonatal mortality, M66-68 provides an ideal control for the initial level of the mortality rate. Note also that cross-sectional studies of infant mortality sometimes control for the characteristics of high-risk births (births with relatively high probabilities of death within the first year of life). Such characteristics include the percentage of births to teenage mothers, the percentage of births to mothers over the age of 40, the percentage of illegitimate births, the percentage of fourth and higher-order births, and the percentage of low birthweight births. I exclude such variables because they are endogenous mechanisms via which poverty, schooling, medical resources, and the policy measures operate.

V. Empirical Results

Ordinary least squares regressions of white neonatal mortality rates are contained in Panel A of Table 2, and ordinary least squares regressions of black neonatal mortality rates are contained in Panel B of Table 2. For whites, the percentage of births to poor mothers has a positive and statistically significant effect on neonatal mortality, while mother's schooling has an insignificant negative effect. For blacks, the negative schooling effect is significant, but somewhat surprisingly, there is an <u>inverse</u> relationship between the percentage of births to poor black mothers and the neonatal mortality rate. For both races, the coefficient of physicians per capita is positive and not significant. Moreover, the infant mortality rate for the period 1966 to 1968 performs well as a control for the neonatal mortality rate prior to the initiation of social programs and for unmeasured determinants of mortality (see regressions A1, A3, B1, and B3).

Because the poverty viariable has the "wrong" sign for blacks, it is excluded in regressions A2, A4, B2, and B4. The main impact of this alternative specification is to increase the absolute value of the schooling effect for whites and to reduce it for blacks. Since the coefficients of the social policy variables do not change much when PB* is omitted and since the estimation of separate poverty and schooling effects "taxes" the black data, the results contained in regressions B2 and B4 are stressed here. For whites, both estimates with and without PB* are used. In part more results are

TABLE 2
Ordinary Least Squares Regressions of Neonatal Mortality Rate[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.

used for whites because trends in white neonatal mortality dominate trends in total neonatal mortality. In particular, white births account for approximately 80 percent of all births at the national level.[1]

Table 2 sheds considerable light on the roles of the social policy variables in neonatal mortality outcomes. I consider the signs, statistical significance, and magnitudes of their estimated impact parameters in turn. Nineteen of the twenty-eight social policy coefficients have the anticipated negative signs in the four white regressions. All fourteen coefficients have the anticipated negative signs in 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.[2] 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- not county-specific and must be computed subject to a number of somewhat arbitrary assumptions (see note 5, page 43). Therefore, it is probably subject to considerable measurement error, which biases its coefficient toward zero. The sizable and significant impacts of the dichotomous variable RA strengthens our confidence in the estimated coefficients of the abortion rate and confirm that the effect for blacks is larger in absolute value than that for whites.

To compare the magnitudes of the observed effects and to examine the relative contributions of schooling, poverty, and social programs to the recent U. S. neonatal mortality experience, the coefficients of regressions A1, A2, and B2 are applied to trends in the exogenous variables between 1964 and 1977.[3] National levels in 1964, 1971, and 1977 of the race-specific neonatal mortality rate and of all relevant exogenous variables except for those pertaining to Medicaid and M and I projects are shown in Table 3. The results of estimating

TABLE 3

National Levels of Selected Variables, Various Years

	1964	1971	1977
Neonatal mortality rate (NM*)			
White	16.2	13.0	8.7
Nonwhite	26.5	19.6	14.7
Physicians per thousand population (MD)	1.30	1.51	1.65
Percentage of families below poverty level			
White	11.4	7.0	6.9
Nonwhite	40.4	25.0	26.7
Percentage of births to low-income women (AB*)[b]			
White	32.1	21.3	21.3
Nonwhite	52.0	34.9	36.9
Percentage of women ages 15-49 with at least a high school education (HSP*)[c]			
White	55.8	62.9	72.5
Nonwhite	36.8	44.1	57.7
Abortion rate (ARATE)[d]	4	92	361
Percentage of low-income women served by organized family planning clinics (UP)[e]	9.0	28.5	40.0
UPXPB*[f]			
Whites	303.3	639.5	894.6
Nonwhites	673.9	1,435.6	2,010.2

TABLE 3 - SOURCES

[a]Source: Bureau of the Census (1980 and selected earlier years) unless otherwise indicated.

[b]Entries for 1964 and 1977 obtained by assuming that the ratio of the race-specific ratio of the birth rate of poor women to that of nonpoor women in 1971 also applies to 1964 and 1977.

[c]1971 entry pertains to 1970. Entries for 1964 and 1977 obtained by assuming that race-specific ratio of percentage of women ages 15 to 49 with at least a high school education to the percentage of women ages 25 and over with at least a high school education in 1970 also applies to 1964 and 1977.

[d]1964 entry pertains to 1969. 1971 entry is an average for 1970 and 1971. 1977 entry pertains to period from July 1, 1976 through June 30, 1977. Except for 1977 entry, the source is Center for Disease Control (1971, 1972).

[e]1964 entry pertains to fiscal 1965 and is taken from Cutright and Jaffe (1977). 1971 entry pertains to fiscal 1971 and is taken from Family Planning Program Development (1974). 1977 entry pertains to fiscal year 1974 and is taken from Dryfoos (1976).

[f]Entries give the mean of the product of UP and PB*, which is the relevant variable in the computations in Table 4, rather than the product of the means. The entry for 1971 is taken directly from the variables in the regressions. The entries for 1964 and 1977 assume the ratio of the mean of the product to the product of the means in 1971 also applies in 1964 and 1977. The 1977 entry is based on UP in fiscal 1974 and assumes no change in the race-specific percentage of families in poverty between 1971 and 1974.

the implied changes in neonatal mortality rates due to selected factors for the period 1964-77 and for the subperiods 1964-71 and 1971-77 are given in Table 4.

Since there is little trend in the percentage of families in poverty after 1971 and since the definition of poverty was altered beginning in 1975, the estimates in Table 4 assume no change in poverty or in PB* between 1971 and 1977. In these computations the national levels of the two M and I project measures are zero in 1964 and do not change from 1971 to 1977. The three Medicaid measures are treated in the same manner. This treatment is justified because there were few M and I projects in operation prior to 1967, and almost no trend in the number of projects or the total number of births in projects after 1971 (Bureau of Community Health Services n.d.). The Medicaid program was not enacted until July 1965, and the rules governing coverage of first-time pregnancies under Medicaid did not vary between 1971 and 1977.

Our treatment of Medicaid is somewhat controversial because the percentage of Medicaid-financed births to poor women and the real quantity of medical services per birth may have risen between 1971 and 1977. Although definitive evidence on these matters is lacking, a number of observations can be made. By 1971, the War on Poverty programs initiated by the Johnson Administration were under attack by the Nixon Administration. This is reflected in part by the failure of the percentage of families in poverty to decline after 1971 (see Table 3). Much of the observed decline over time in the relationship between income and physician visits, which Davis and

TABLE 4

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		8.4		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*.

Reynolds (1976) show was caused by Medicaid, occurred by 1971. The percentage of the poverty population that received Medicaid benefits rose by only 6 percentage points between 1970 and 1974 (Davis and Reynolds 1976; Davis and Schoen 1978). Real Medicaid benefits per recipient show no trend between 1971 and 1977 (Davis and Schoen 1978). The percentage of black mothers who started their prenatal care in the first trimester of pregnancy rose between 1969 and 1974 (Taffel 1978). Except for the last observation, this evidence justifies our treatment of Medicaid. However, the sensitivity of the results are examined with an alternative assumption described below.

As shown in Table 4, the actual decline in the white neonatal mortality rate between 1964 and 1977 was 7.5 deaths per thousand live births. Regression A1, which incorporates separate poverty and schooling effects, "explains" 2.8 of these deaths or 37 percent of the total reduction. Regression A2, which treats the schooling effect as the joint impact of schooling and poverty, accounts for 2.4 deaths or 32 percent of the total reduction. For nonwhites, the neonatal mortality rate fell by 11.8 deaths per thousand live births between 1964 and 1977.[4] Regression B2 predicts a decline of 4.8 deaths or 41 percent of the observed reduction.

A striking message in Table 4 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. For the entire period, the reduction in the white neonatal mortality rate due to abortion ranges from 1.5 to 1.7 deaths per thousand births. The comparable figure for nonwhites is a whopping 2.5 deaths per thousand births. When the two subperiods are examined separately, abortion makes the largest contribution except for nonwhites in the 1964-71 period. Here is ranks second to the impact of the rise in the use of organized family planning services by low-income women. The extremely large expansion in the abortion rate in the latter period (1971-77) provides a cogent explanation of the acceleration in the percentage rate of decline in both race-specific mortality rates and the acceleration in the absolute rate of change for whites.

The increase in the use of organized family planning services by low-income women is the second most important factor in reductions in nonwhite neonatal 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 4 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 nonwhite 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 impact of M and I projects over the entire period is dominated by the impacts of abortion reform and family planning in part because there was no change in the size of these projects between 1971 and 1977. But suppose that the absolute increase in the size of these projects had been the same in the second subperiod as it was in the first. Then their predicted impact on the nonwhite neonatal death rate would amount to 0.6 deaths per thousand births, which still is substantially smaller than the abortion and family planning effects.

Medicaid can be dismissed as a cause of the decline in white neonatal mortality; it predicts either no change or an increase in the white death rate. In the case of nonwhites, Medicaid accounts for a reduction of 0.5 deaths per thousand live births. If the somewhat controversial assumption of no change in the program between 1971 and 1977 is relaxed in the same manner as for M and I projects, a reduction of 1.0 deaths per thousand births is obtained.

There is reason to believe, however, that the preceding estimate, which is much less than the abortion impact and also less than the family planning impact, is itself too large. Based on regression B2, the nonwhite low-income neonatal mortality rate in counties in states that cover no first-time pregnancies under Medicaid is larger than the corresponding rate in states that cover all first-time pregnancies to financially eligible women by 1.4 deaths per thousand births (the coefficient of MAXPB* multiplied by 100). But this differential is smaller than the corresponding differentials for counties that cover some first-time pregnancies (see the coefficients of MUXPB* and MNXPB*). Where coverage is provided if no husband is present or if he is present but unemployed, the differential is 3.3 deaths per thousand births. Where 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, I estimate that the neonatal death rate of low-income nonwhites would fall by only 0.1 deaths per thousand births if all 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, the results presented here, 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.[5] 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 explosion of the abortion rate in this period. Medical care played an extremely limited role in this process, although there is weak evidence that M and I projects and Medicaid were of some benefit to nonwhites.

The above conclusions are subject to qualification that there are no estimates of the impact of the pill and the IUD other than those that are inferred through the use of family planning services by low-income women. They also are subject to the qualification that one cannot estimate the contribution of advances in neonatalogy. Hence, the conclusion presented here with respect to medical care pertains to the quantity of care rather than to the technology of care.

These results are relevant to current U. S. policy debates with respect to the financing of pregnancies and abortions under Medicaid and with respect to attempts by the Right-to-Life movement to enact a constitutional amendment that would outlaw abortion except when it is necessary to preserve a pregnant woman's life. As part of the Child Health Assurance Program (CHAP, a bill pending in the House of Representatives since 1978), national--as opposed to state--income standards would be established for determining the eligibility of pregnant women for Medicaid. As part of these standards, all first-time births to all financially eligible women would be financed by Medicaid. These estimates suggest that the payoffs to CHAP are extremely small. Its enactment would have no impact on the white neonatal mortality rate and would lower the nonwhite rate by only 0.1 deaths per thousand live births.

Under the Hyde Amendment, which was in effect from June 1977 until February 1980, Federal funding of abortion under Medicaid was banned except in cases where the woman's life was in danger. During this 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 (Newsweek Magazine, January 28, 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 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. 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. If the Supreme Court overturns Judge Dooling's decision, the abortion rate for poor women undoubtedly will grow more slowly than otherwise and might even fall. According to my findings, this would retard the rate of decline in the neonatal mortality rate of the poor and might even cause it to rise.[6]

The most striking implication of my 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, the nonwhite neonatal mortality rate would rise by approximately 2.8 deaths per thousand live births or by 19 percent above its 1977 level. The white neonatal mortality rate would rise by approximately 1.8 deaths per thousand live births or by 21 percent above its 1977 level. This assumes that changes in technology or other factors did not affect the relationship between abortion and mortality.

In conclusion, these results suggest that the payoffs in terms of infant deaths averted to abortion reform and family planning are larger than the payoffs to medical care. But surely the costs of the former are also smaller than the costs of the latter. However, the relative benefits of preventing an infant death through medical care or abortion are beyond the scope of this study.

Some regressions were run excluding lagged mortality to prevent biases towards zero in coefficients, which occur if social programs or favorable socioeconomic conditions had already lowered infant mortality rates by 1966-68. These biases are likely to be smaller for social programs than for the socioeconomic variables, since the former were generally introduced around 1964 or later and are likely to have had implementation lags. The new runs showed coefficients with higher absolute values for MD, PB* and HSP*. Coefficients of MAXPB* and MUXPB* are lower (or more negative), suggesting that either Medicaid or some other programs pursued by liberal states had reduced infant mortality by 1966-68. Other correlated state specific omitted variables may have had similar effects. MAXPB* generally has the most substantial negative effect of the Medicaid variables in the white regressions and is highly significant in A3. This is expected since it indicates the most liberal coverage. Only MUXPB* is significant in all of the black regressions, while MNXPB* is significant in some. The lack of significance of MAXPB* in the black runs may indicate that the correlated abortion variables are picking up some of its effect. Excluding lagged mortality raised MNXPB* in the white runs and lowered it in the black regressions.

The abortion coefficients are lower than or equal to those of the previous regressions. However, they are not much lower. This is understandable since there were only 4 abortions per thousand

TABLE 5

Ordinary Least Squares Regressions of White Neonatal Mortality Rate[a]
M66-68 Excluded

Independent Variable	Regression Number (1)	(2)	(3)	(4)
PB*	.049		.057	
	(3.57)		(4.26)	
MD	.289	.263	.274	.243
	(2.93)	(2.65)	(2.75)	(2.42)
HSP*	-.058	-.086	-.057	-.097
	(-4.05)	(-7.22)	(-3.95)	(-7.68)
MAXPB*	-.009	.007	-.024	-.009
	(-0.81)	(0.74)	(-2.56)	(-1.03)
MUXPB*	-.007	.002	-.007	.003
	(-0.66)	(0.17)	(-0.69)	(0.24)
MNXPB*	.001	.011	.006	.019
	(0.14)	(1.13)	(0.60)	(1.87)
MIXPB*	.016	.008	.010	-.0006
	(1.16)	(0.60)	(0.76)	(0.00)
PMIXPB*	-.016	-.014	-.004	.002
	(0.69)	(0.60)	(0.17)	(0.10)
UPXPB*	-.001	-.0002	-.001	-.0002
	(-2.56)	(-1.21)	(-2.56)	(-0.98)
ARATE	-.006	-.007		
	(-4.29)	(-5.12)		
RA			-.607	-.669
			(-3.41)	(-3.72)
CONSTANT	15.819	18.459	15.386	
$\bar{R}^2$	.160	.145	.152	

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

TABLE 5

Ordinary Least Squares Regressions of Black Neonatal Mortality Rate[a]
M66-68 Excluded

Independent Variable	Regression Number (1)	(2)	(3)	(4)
PB*	-.137		-.123	
	(-3.78)		(-3.51)	
MD	.424	.619	.349	.542
	(1.92)	(2.84)	(1.62)	(2.57)
HSP*	-.179	-.076	-.187	-.088
	(-4.41)	(-2.47)	(-4.72)	(-3.11)
MAXPB*	-.009	-.021	-.012	-.015
	(-0.35)	(-0.81)	(-0.57)	(-0.69)
MUXPB*	-.047	-.042	-.049	-.042
	(-2.16)	(-1.89)	(-2.30)	(-1.95)
MNXPB*	-.013	-.033	-.012	-.030
	(-0.90)	(-2.48)	(-0.85)	(-2.33)
MIXPB*	.008	.010	.006	.008
	(0.37)	(0.45)	(0.28)	(0.35)
PMIXPB*	-.039	-.038	-.043	-.042
	(-1.39)	(-1.32)	(-1.55)	(-1.49)
UPXPB*	-.0002	-.0007	-.0001	-.0005
	(-0.68)	(-2.08)	(-0.14)	(-1.47)
ARATE	-.009	-.007		
	(-2.20)	(-1.59)		
RA			-1.909	-1.918
			(-4.18)	(-4.13)
CONSTANT	35.073	25.928	34.784	26.431
$\bar{R}^2$	.087	.053	.119	.090

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

births in 1969. So abortion reform probably had little effect in 1966-68 and the inclusion of the lagged mortality rate does not seriously bias the abortion coefficients towards zero.

The coefficients of Maternal and Infant Care Projects variables were usually higher in the new runs. This is understandable since the projects were located in areas with high mortality rates. PMIXPB* had lower coefficients in the new black regressions. Coefficients of the family planning variable were sometimes higher in the new results. This may also be explained by noting that runs without the lagged mortality rate do not control for reverse causality. Counties with high values of UPXPB* may contain many young, old, or unhealthy women. High risks of unfavorable birth outcomes may encourage fertility control. Heavy use of family planning in a county might also indicate a low demand for children and be associated with inferior prenatal care. Finally, family planning clinics were primarily set up in poverty areas, which generally have high infant mortality rates. My attempts to hold poverty constant are imperfect since I do not have a poverty measure specifically for mothers of childbearing age.

The new runs allow for extrapolations which explain a greater proportion of the reduction in infant mortality rates over the 1964-1977 period. Socioeconomic variables, especially education, make greater contributions than in the old extrapolations. The effect of the abortion variable is increased -- it still exceeds that of the education and poverty variables combined, for the 1964-1977 interval. The importance of family planning fell, especially for blacks. The

TABLE 6
Contribution of Selected Factors to Reductions in
Neonatal Mortality Rates, 1964-1977
Based on Regressions Excluding M66-68

	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		8.4		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	-.1	a	a	a	a	a	-.2	-.1	a
PB*	.5	b	.5	a	c	b	b	b	b
HSP*	1.0	1.4	.4	.6	.6	.8	.6	.6	1.0
ARATE	2.1	2.5	.5	.6	1.6	1.9	2.5	.6	1.9
UPXPB*	.6	.1	.3	a	.3	a	.9	.5	.4
M and I Projects[d]	a	a	a	a	c	c	.2	.2	c
Medicaid[e]	a	a	a	a	c	c	.6	.6	c
Total explained reduction	4.1	3.8	1.7	1.2	2.5	2.7	5.6	2.4	3.3
Percentage explained	54.6	50.7	53.1	37.5	58.1	62.8	47.5	34.8	67.3

[a] Less than .1 in absolute value.
[b] Variable omitted from regression.
[c] No change in variable.
[d] Combined contribution of MIXPB* and PHIBXPB*.
[e] Combined contribution of MAXPB*, MUXPB*, and MNXPB*.

small contributions of the maternal and infant care projects decreased.

Now I will examine the probable impact of proposed government policies, using the new coefficients. The CHAP amendment, which would extend Medicaid coverage to all financially eligible pregnant mothers would lower the white neonatal mortality rate by about .5 deaths per thousand births and the nonwhite rate by about .14. A ban on abortion could raise the white rate by about 2.6 and the nonwhite rate by 2.8. This confirms the previous conclusions presented about the importance of abortion reform.

The regressions which include lagged mortality provide the most reliable estimates of the impact of social programs and policies. The other runs are most useful for obtaining coefficients of poverty and education variables. Using these results, one can explain a substantial proportion of the recent decline in neonatal mortality.

FOOTNOTES

[1]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 is somewhat more complicated because the simple correlation between the death rate and PB* is negative.

(b) There are few studies of the race-specific impact of poverty on infant mortality. Using a special sample of births and subsequent infant deaths taken by the National Center for Health Statistics, Gortmaker (1977) reports results similar to mine. White babies are more likely to die in poverty families than in nonpoverty families, but this relationship does not hold for black babies.

(c) The unimportance of physicians per capita in these regressions mirrors findings reported by Brooks (1978) in a study of variations in infant mortality rates among SMSAs. The positive and significant coefficients of MD in black regressions B2 and B4 are not troublesome because other coefficients are not sensitive to the exclusion of this variable. The MD variable is retained here because there is almost no trend in it between 1964 and 1977. Hence its retention does not cloud the forecasts and backcasts that follow.

(d) Low income black births are more likely to occur outside of a hospital than high income births. Deaths occurring in a hospital are more likely to be reported.

[2]For Medicaid, I always accept the hypothesis that no member of the set given by MAXPB*, MUXPB*, and MNXPB* has a non-zero coefficient at the 5 percent level. For M and I projects, I accept the hypothesis that no member of the set given by MIXPB* and PMIBXPB* has a non-zero coefficient in five of six cases. The exception pertains to regression A4.

[3]In the discussion that follows, I use the terms "contribution to reduction in mortality" and "magnitude" as synonyms in a loose sense. If α_i is the regression coefficient of variable x_i and Δ_i is its change over time, then its contribution to the change in mortality is $\Delta_i\alpha_i$. Clearly $\Delta_i\alpha_i$ can be large even if α_i is small. I use $\Delta_i\alpha_i$ as a measure of magnitude for two reasons. First, the independent variables are measured in different units. Second and more importantly, in a fundamental sense we view an exogenous variable as having a small impact if it makes a minor contribution to the reduction in infant mortality during a period of rapid decline.

Note that the regressions with the continuous abortion rate are employed in the computations that follow to facilitate the discussion of the impacts of future trends in this variable. These conclusions are not altered when the dichotomous abortion variable is employed. Since abortion makes a large contribution to the decline in neonatal mortality (see Table 4), the possibility examined is the possibility that its effect is nonlinear by entering the square of ARATE as well as ARATE in the regressions. Since the square term was always insignificant, there is no evidence of a nonlinear effect.

[4]Data for nonwhites are shown because separate time series for blacks are not available. Note that changes in the lagged mortality rate are not relevant in the forecasts and backcasts in Table 4 because the underlying model is not a dynamic one. Rather, the lagged rate serves as a proxy for the initial level (before 1964), which does not change by definition. In econometric terminology our model is one with "fixed effects" rather than one with "state dependence."

[5]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.

[6]The above conclusion is strengthened by the results of a study by Leibowitz, Eisen, and Chow (1979). They examine the decision by unwed pregnant teenagers to have an abortion or to deliver their babies in a sample of California residents in the period 1972-74. Teenagers who were eligible for Medicaid financing of an abortion or a delivery were more likely to choose to keep their babies. I believe that this effect would be even stronger if Medicaid financed deliveries but not abortions. Ultimately such a financing arrangement might retard the reduction in family size in poverty families or cause it to increase. Since the health and cognitive development of children are negatively related to family size (for example, Lewit 1977; Edwards and Grossman 1979, forthcoming), many aspects of the welfare of poor children would be harmed if Medicaid financing of abortions were banned.

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