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THE IMPACT OF INDUCED ABORTION ON BIRTH OUTCOMES IN THE U.S.

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ABSTRACT

This paper examines the impact of induced abortion on birth outcomes by treating abortion as an endogenous input into the production of infant health. To gauge the direct and indirect effect of abortion, three measures of infant health are considered simultaneously: the neonatal mortality rate, the percentage of low-birth weight births, and the percentage of preterm births. All three are race-specific and all pertain to large counties in the U.S. in 1977. Because the utilization of health inputs may be conditioned on the expected birth outcome, estimates obtained by two-stage least squares are emphasized. The results make clear that abortion is an important determinant of infant health. This suggests that by reducing the number of unwanted births, abortion enhances the healthiness of newborns of a given weight and gestational age, as well as improving the distribution of births among high-risk groups. Moreover, these direct and indirect effects differ by race.

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The Impact of Induced Abortion on Birth Outcomes in the U.S.

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Induced abortion has been frequently mentioned as a potential causal factor in the accelerated decline in neonatal mortality over the past decade (Eisner et al 1978; Kleinman et al 1979; David and Seigal 1983). Initial research generally supported the link between increased use of abortion and improved birth outcomes (Lanman, Kohl, and Bedell 1974; Glass et al 1974; Quick 1978). However, not until Grossman and Jacobowitz (1981) and Corman and Grossman (1985) was the impact of abortion services quantified in a multivariate context. These two studies , however, employed a reduced form model and therefore could only speculate as to the mechanisms through which abortion operated on infant health.

The purpose of this paper is to outline the pathways through which abortion[1] impacts on birth outcomes. In particular, to what extent has the process of fetal selection, made possible by abortion, resulted in healthier infants of a given birth weight; and to what extent has abor-

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[1] Throughout this paper, all references to abortion refer to induced abortion.

tion improved the distribution of of births among high-risk groups and thereby lowered the incidence of prematurity?

In order to address this question, abortion is treated as an endogenous input in a structural model of infant health. Production functions are estimated for three birth outcomes: the race- and county-specific neonatal mortaltiy rate, the percentage of low-birth weight births, and the percentage of preterm births. Attempts are made to control for other determinants of infant health such as prenatal and neonatal care. The latter has been repeatedly cited as a primary factor in the increased survivability of premature babies (Paneth et al. 1982; McCormick 1985). Moreover, recent insights by Rosenzweig and Schultz (1982,1983a,1983b) with respect to health heterogeneity will also be incorporated.

II. Analytical Framework

Economic models of the family developed by Becker and Lewis (1973) and Willis (1973) have been used by numerous authors to study theoretically and empirically the determinants of birth outcomes. (Grossman and Jacobowitz 1981; Rosenzweig and Schultz 1982,1983a,1983b; Lewit 1983; Corman and Grossman 1985) Accordingly, I assume that the parents' utility function depends on their own consumption, the number of births and the health of each child at birth. Moveover, the infant health production function depends on the quality and quantity of medical care, the own time of the mother, nutritional intake, and the mother's reproductive efficiency. The latter factor includes an unobserved genetic component which affects an infant's health and survivability at birth as well as other aspects of a mother's efficiency in household production.

Maximization of the utility function subject to the production and resource constraints generates a demand function for infant health. The interaction between the demand and production of infant health spawns the derived demands for medical care and the other inputs.

To clarify the preceeding ideas and to lay the foundation for the empirical implementation, the basic model can be written as follows:

$$d=f_1(n,m,a,c,b,z,e) \qquad (1)$$

$$b=f_2(m,a,c,s,g,z,e) \qquad (2)$$

$$g=f_3(m,a,c,s,r,z,e) \qquad (3)$$

$$r=f_4(a,c,x,z,e) \qquad (4)$$

$$n,m,a,c,s=f_j(p,y,x,z,e) \qquad (5)-(9)$$

The probability that an infant's health deteriorates to the point that he/she does not survive the first month of life is represented by equation (1). Equations (2) and (3) give the likelihood that a newborn is of low weight (less than 2,500 grams or 5.5 pounds) or preterm (less than 37 weeks gestation) respectively. The basic set of inputs used in the production of infant health are neonatal intensive care (n), prenatal care (m), abortion (a), organized family planning use (c), and smoking (s). Z is a measure of environmental conditions and e is an unobserved genetic component specific to each family and which is henceforth refered to as an infant's endowed health. Equation (4) shows the determinants of the endogenous risk factor which in this model is the probability that a baby is born to a woman less than 20 or older than 39

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[2] Other endogenous risk factors include the spacing between births, marital status, parity, and education attainment. The rationale for choosing age as opposed to the others is given in Section III.

years of age.[2] X is a demographic measure that determines the age of a mother at the time of birth.

Equations (1)-(4) are the production functions and they constitute the structural equations of the model. Equations (5)-(9) are the input demand functions for they relate the use of an input to its price and availability (p), income and resource constraints (y), and the other exogenous determinants mentioned previously.

The empirical focus is on equations (1)-(3) which give the direct or risk-specific effect of an input on a particular birth outcome. To assess the indirect effect of an input on a specific aspect of infant health the right-hand-side birth outcome in each structural equation is replaced by its endogenous determinants.[3] Specifically, substituting equation (4) into (3), equations (3) and (4) into equation (2), and equations (2), (3), and (4) into equation (1) yields what will be refered to in this paper as the quasi-structural equations:

$$d=f_{10}(n,a,c,m,s,x,z,e) \qquad (10)$$

$$b=f_{11}(a,c,m,s,x,z,e) \qquad (11)$$

$$g=f_{12}(a,c,m,s,x,z,e) \qquad (12)$$

Although the model has meaningful interpretations at the family

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[3] To clarify the notion of a direct and indirect effect consider the following model:

$$b = a_0 + a_1X + a_2g$$

$$g = c_0 + c_1X$$

where b is birth weight, g is gestational age and X is an input such as prenatal care. The direct effect of prenatal care on birth weight is a_1. The indirect effect, c_1a_2, is obtained by substituting for g in the birth weight equation.

level, the empirical analysis pertains to county-level data. The only variables whose roles are expected to be altered by aggregation are abortion and family planning. The difference arises because the impact of abortion and family planning cannot be directly incorporated into a production function of infant health based on a sample of individual births.

Conceptually, the expanded use of abortion and family planning have enabled individuals to more effectively plan the number and timing of their offspring. The choice of whether to give birth or not has created a potentially significant self-selection problem. The issue is analogous to the more commonly cited examples of self-selection found in the literature (Maddala, 1977). Specifically, the neonatal mortality rate based on a sample of women who choose to give birth may understate the neonatal mortality rate that would prevail if all pregnant women tried to carry to term.[4]

To adjust for selectivity bias at the individual level,[5] Heckman's two-stage procedure could be applied to a chort of women observed from

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4 The fact that in 1977 the abortion rate for whites and non-whites was 20 and 59 respectively, and the abortion ratio was 333 and 679 respectively suggest that the magnitude of the self-selectivity bias may be substantial. The abortion rate is the number of abortions per women ages 15-44 while the abortion ratio is the number of abortions per thousand live births (Statistical Abstract of the United States, 1984).

5 The selectivity effect described in this example refers to the bias caused by abortion. To correct for the potential self-selection due to contraception, it would be necessary to have data on a cohort of sexually active women observed over their reproductive cycle--a far more demanding data requirement than is needed to measure the self-selection generated by aboriton.

6 This geographical restriction insures that the price and availability of abortion is relatively equal for most women.

conception and living in the same county or metropolitan area.[6] A sample selection criterion could be established predicting the probability of not obtaining an abortion. Given all the necessary assumptions, the appropriate adjustment factor (the inverse of the Mill's ratio) could be estimated and entered into the structural equation as a right-hand-side variable.[7]

At the aggregate level, the selectivity effect of abortion can be controlled for directly in the production function. Since the abortion rate measures the probability that a woman between the ages of 15 and 44 will voluntarily terminate a pregnancy, it can be viewed as an indicator of the extent to which unwanted births have been averted. Counties in states which offer relatively greater financing and accessibility to abortion should have a higher abortion rate and thus, fewer unwanted births than counties in which abortion is less readily available. All else equal, this should improve the distribution of births among high-risk groups (Sklar and Berkov 1973; Shelton 1977; Brann 1979) and lower the incidence of prematurity and neonatal mortality.

In addition, the use of abortion and contraception may be associated with the frequently cited improvement in birth weight-specific mortality (Kleiman et al 1979; Lee et al 1980; Williams and Chen 1983). As noted above, advances in perinatal and neonatal care have undoubtedly played a major role in the delcine of birth weight-specific neonatal

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[7] Besides computational ease, the advantage of Hechman's procedure over direct maximization of the appropriate liklihood function is the useful interpretation of the adjustment factor. Given that the inverse of the Mill's ratio is a "monotone decreasing function of the probability that an observation is selected into the sample" (Heckman p. 156), it could be interpreted as the probability of not obtaining an abortion, and as such, could be considered a measure of the "wantedness" of a birth (Joyce 1985).

mortality. However, newborns of a given weight may also be healthier because less healthy fetuses are aborted and because births that are better planned may be more wanted (Tietze 1984; NCHS 1985).

III. Empirical Implementation

A. Data and Measurement of Variables

Race-specific data on births by weight, gestational age and mother's age from 1976 through 1978 are from the National Center for Health Statistics (NCHS) Natality Tape. Neonatal deaths over the same period are from the NCHS Mortality Tape. Socioeconomic characteristics are taken from the Census of Population, and estimates on smoking are from Eugene Lewit of the National Bureau of Economic Research. Data on neonatal intensive care are from the American Hospital Association and from Ross Planning Association of Ross Labratories (1982). Information relating to family planning and abortion are from the Alan Guttmacher Institute. Finally, measures pertaining to the services and programs used in the reduced form input demand equations are from Corman and Grossman (1985).

Counties are used as the unit of observation instead of states or SMSA's because they are more homogenous with respect to socioeconomic characteristics and medical resources. However, small counties present two potential problems: first, people may travel outside the county for medical services and second, sparsely populated counties with few births may show large fluctuations in birth outcomes due to random movements. To minimize these difficulties, only counties with a 1970 population of 50,000 or more in states that reported gestational age on birth certificates are included in the sample. For the black sample there is an additional criterion of at least 5,000 blacks in the county. There are 632 counties in the white regressions and 327 counties in the black regressions. Both samples account for approximately 72 percent of the total U.S. white and black populations.

Three measures of infant health are analyzed: the neonatal mortality rate, the percentage of low-birth weight births, and the percentage of births for which gestational age is less than 37 weeks. All are race-specific and all are three-year averages centered on 1977.[8] The interrelationships among the three birth outcomes are well-documented.[9] For instance, light births are 40 times more likely to die in the first month of life than normal weight infants and approximately 50 percent of low-birth weight births are preterm.

Table 1 contains a description of the dependent and independent varibles used in the estimation of the three structural and three quasi-structural birth outcome production functions. Table 2 displays their means and standard deviations. The availablity and resource measures used in the reduced form input demand equations are described in the Appendix. Except for the abortion rate, neonatal intensive care, and per capita smoking, all variables in the structural model are race- and county-specific.

The five basic inputs used in the production of infant health are neonatal intensive use, prenatal care use, the abortion rate, the use of

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8 Three-year averages help to attenuate random elements arising in counties with few infant deaths.

9 For a recent summary see Part I of Perinatal Epidemiology, edited by Michael Bracken.

Table 1

Definitions of Varibles[a]

Variable	Definitions
Neonatal mortality rate*	Three-year average neonatal mortality rate centered on 1977; deaths of infants less than 28 days old per 1,000 live births
Low birth weight*	Three-year average percentage of low-birth weight (2,500 grams or less) live births centered on 1977
Preterm*	Three-year average percentage of preterm (gestational age less than 37 weeks) live births centered on 1977
Teenage family planning users*	Percentage of women aged 15-19 who used organized family planning clinics in 1975
Abortion rate	Three-year average state-specific resident abortion rate centered on 1976; abortions performed on state residents per 1,000 women aged 15-44
Prenatal care*[b]	Three-year average percentage of live births for which prenatal care began in the first trimester (first three months) of pregnancy centered on 1977
Neonatal intensive care	Sum of state-specific hospital inpatient days in Level II, Level III, or Levels II and III neonatal intensive care units in 1979 per state-specific three-year average number of births centered on 1977
Cigarettes	State-specific daily number of cigarettes smoked per adult 18 years and older in 1976
Births to young and old women*	Three-year average percentage of live births to women aged 15-19 centered on 1977
Young and old women*[c]	Ratio of women ages 15 to 19 and 40 to 44 over the total number of women 15 to 44 in 1975
Ln population density	The natural logarithm of the ratio of the population in 1975 to the area per square mile

Notes to Table 1

[a] An asterisk (*) next to a variable means that it is race-specific. All variables are calculated for counties unless otherwise specified.

[b] Four states, Kansas, Pennsylvania, Ohio, and Virginia have data on the percentage of women who initiated prenatal care in one of the first three months for 1978 only. Because this variable has trended upwards nationally, data for 1976 and 1977 by county in these particular states were estimated by deflating the 1978 figures by the national rates of growth between 1976 and 1978. The three-year average for the counties in these states incorporated these estimates for 1976 and 1977.

[c] The variable is available for whites and nonwhites as opposed to whites and blacks.

Table 2

Means and Standard Deviations of Variables[a]

	Whites		Blacks	
	mean	std	mean	std
Neonatal mortality*	8.817	1.643	16.342	3.323
Low-birth weight*	5.952	.734	12.988	1.248
Preterm*	7.384	.884	15.648	2.138
Teen family planning users*	9.182	6.512	24.035	9.712
Abortion rate	25.336	9.050	25.035	8.866
Neonatal intensive care	.650	.401	1.519	1.045
Prenatal care*	78.710	8.302	59.868	11.057
Cigarettes	7.404	.533	7.480	.362
Births to young/old women*	14.346	3.912	29.826	4.817
Young and old women*	.336	.022	.350	.026
Ln population density	6.369	1.557	7.315	1.691
Sample size	632		327	

[a]An asterisk (*) next to a variable means that it is race-specific. Means and standard deviations are weighted by the three year sum of race-specific births in 1976-1978.

family planning clinics by teenagers, and adult per capita smoking. Except for smoking, all the inputs should be negatively related to the measures of infant health descirbed above.

This study focuses on teenage family planning use as opposed to the utilization of family planning clinics by all women of childbearing age for several reasons.[10] First, only the former measure is race-specific. Second, teenagers are disproportionate users of family planning clinics (Torres and Forrest 1983). Third, adolescent births and births to unmarried teenagers have well-documented, above average risks of prematurity and neonatal mortality (Taffel 1980; Elster 1984). Moreover, Forrest (1981) reports that for every ten adolescents enrolled in family planning clinics in 1975, three pregnancies were averted the following year. Consequently, measures of infant health may be more sensitive to orgnaized teenage family planning use than to variations in use by all women.[11]

The smoking variable is taken from Lewit (1982) who estimated it from a micro-level study of the demand for cigarettes (Lewit and Coate 1982). Lewit applied the coefficients of the fitted demand functions to state means of the independent variables to arrive at the figures used

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[10] The Alan Guttmacher Institute provided the data on teenage family planning use. Their estimates were adjusted to compensate for cross-county utilization. However, in 26 counties measuring nonwhite use and in 2 counties measuring white use the figures appeared abnormally high (above 50 percent or greater than 2.5 standard deviations from the mean adjusted for the outliers). In these cases the state mean was substituted.

[11] Work by Corman, Joyce and Grossman (1985) found that when organized family planning use by teenagers, and orgnaized family planning use by all women ages 15 to 44 were entered as separate variables in a regression predicting neonatal mortality, the measure of adolescents use always had the correct sign and consistently dominated the broader indicator of organized family planning use.

here. The advantage of Lewit's variable over the readily-available tax-paid sales per state is that his measure adjusts for the substantial "bootlegging" of cigarettes at both the individual and group level. Because of this smuggling, data from tax-paid sales underestimate consumption in high-tax states and overestimate it in low-tax states.

The endogenous risk factor (r) present in the gestational age production function (equation 3) is represented by the three-year average of births to women less than 20 or greater than 39 years of age. The risk of prematurity among women at both ends of the age spectrum are well-documented (Institute of Medicine, 1985). Furthermore, births to women in these age groups are correlated with other risk factors such as out-of-wedlock births, births to women with low levels of education and births of high parity. For instance, 28 percent of white teenagers and 83 percent of black teenagers who gave birth in 1977 were unmarried (Ventura, 1984).

Two points concerning the use of this risk factor deserve mention. First, using age alone, instead of including the other risk factors cited above, is preferable for several reasons. There is a substantial overlap between age and the other socio-demographic characteristics and therefore, the multicollinearity generated by endogenizing such highly correlated risk factors is likely to be substantial. Furthermore, by 1978, eight states including New York, California and Texas did not report marital status on birth certificates. Consequently, inclusion of that variable would have reduced the sample size substantially.

Second, the model assumes that the percentage of births to women in these age groups impacts on birth weight and neonatal mortality only indirectly. This is necessary in order to identify the system of structural equations (1)-(3). This restriction does not affect the estimated total effect of an input on infant health. However, to the extent that this and other risk factors impact on mortality and low birth weight independent of the relevant intermediate birth outcome, the direct effect of an input may be overstated.

The natural logarithm of the population density and the ratio of women ages 15 to 19 and 40 to 44 over the total number of women 15 to 44 constitute the environmental (z) and the exogenous risk (x) components respectively. The former may pick-up the possible impact of pollution or stress from overcrowding on infant health (Lave and Seskin 1973; Bakketeig et al 1984) . The ratio of young and old women to total women of childbearing ages should vary directly with the percentage of births in these high-risk age groups and hence, vary inversely with infant health.

B. Estimation

Following Rosenzweig and Schultz (1982,1983a,1983b), it is anticipated that the residuals in the structural equations will be correlated with the health inputs.[12] This expectation is based on the assumption that individuals have some information concerning their genetic health endowment which is unknown to the researcher but which causes the parents to alter their behavior with respect to their use of inputs. For example, ultrasound and amniocentesis are very direct methods of obtaining information concerning the health of the fetus. Responses to such information range from abortion to more intensified prenatal and perina-

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[12] The notion that an unobserved input may cause the coefficients of an estimated production function to be biased and inconsistent has long been a part of the literature on the estimation of production functions (Mundlak and Hoch 1965). In his seminal work on health production functions, Grossman (1972) also noted the danger of assuming the inputs to be uncorrelated with error term.

tal care. Such remedial behavior may create a correlation between the use of the health inputs and the unobserved endowment term imbedded in the residuals. In short, the use of health inputs may not only affect the birth outcome, but the anticipated birth outcome may also affect utilization. Because of the potential reverse causality, ordinary least squares may yield biased and inconsistent estimates.

To test whether a significant correlation between the production function residuals and the health inputs does exist, Wu's T_2 statistic (Wu 1973) as described by Nakamura and Nakamura (1981) will be applied. If the null hypothesis of zero correlation between the error term and the regressors is rejected, then two-stage least squares (TSLS) will be used to estimate the birth outcome production functions. In the first stage, family planning, abortion, neonatal intensive care and prenatal care will be regressed on the set of reduced form determinants described in the Appendix.[13] They include measures of female schooling, female poverty levels, the proportion of high-risk women, Medicaid and AFDC benefits, population density, as well as availability measures for neonatal and prenatal care, abortion and organized family planning. The underlying assumption of this two-stage procedure is that the resource and availability measures vary independently of the health endowment.

It is anticipated, however, that the correlation between the residuals and the health inputs will be reduced by holding constant the

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[13] Cigarette smoking is also endogenous but it excluded from the first stage becuase the measure used is already a predicted value generated from a micro reduced form demand function for cigarettes. Technically, this procedure amounts to assuming that the quantity of cigarettes consumed depends only on the predictors employed by Lewit (income, price, education, age, sex and race) and not on the availability measures of the other inputs.

percentage of low-birth weight births, preterm births, and births to young and old women in the estimation of equations (1)-(3). This should occur because each of these variables may be an effective proxy for the health endowment. Examination of the Wu statistic should offer evidence to this effect. For instance, if the percentage of low-birth weight births adequately controls for the health endowment in the neonatal mortality equation, then barring other forms of misspecification, the Wu statistic should indicate no significant correlation between the error term and the righ-hand-side health inputs.

Finally, the infant health production functions are assumed to be linear. All regressions are weighted by the race-specific number of births between 1976 and 1978. Work by Joyce (1985) with similar data found that a Cobb-Douglas and logistic functional form generated results that differed little qualitatively.

IV. Results

A. Estimates of the Infant Health Production Functions

Estimates of the structural and quasi-structural production functions are presented in Tables 3 and 4 respectively. The first-stage results are given in the Appendix. This paper focuses on infant health production functions, and therefore, the estimated input demand equations are discussed only briefly. A more complete discussion can be found in Joyce (1985).

The relationship between input use and input availibility is well-supported by the data (Tables A-3 and A-4). Increases in family planning

Table 3

Structural Equations of Infant Health

Independent Variables	Whites: Neonatal Mortality OLS (3-1)	Neonatal Mortality TSLS (3-2)	Low Birth Weight OLS (3-3)	Low Birth Weight TSLS (3-4)	Preterm OLS (3-5)	Preterm TSLS (3-6)	Blacks: Neonatal Mortality OLS (3-7)	Neonatal Mortality TSLS (3-8)	Low Birth Weight OLS (3-9)	Low Birth Weight TSLS (3-10)	Preterm OLS (3-11)	Preterm TSLS (3-12)
Teen family planning users*[b]	-.030	-.047	.007	.006	-.007	.010	-.031	-.047	.001	-.013	.001	.030
	(-3.23)	(-2.41)	(1.91)	(.90)	(-1.71)	(1.27)	(-1.70)	(-3.18)	(.18)	(-.99)	(.12)	(1.30)
Abortion rate[b]	-.034	-.023	-.010	-.014	.003	.008	-.043	-.081	-.005	-.0004	-.061	-.076
	(-4.78)	(-1.76)	(-3.73)	(-3.30)	(.96)	(1.43)	(-1.84)	(-1.77)	(-.62)	(-.02)	(-3.63)	(-3.09)
Prenatal Care*[b]	-.022	-.007	-.008	-.019	-.017	-.0003	.007	.071	.005	-.019	-.034	-.079
	(-2.82)	(-.32)	(-2.45)	(-2.36)	(-3.96)	(-.03)	(.46)	(1.48)	(.99)	(-.84)	(-3.63)	(-3.42)
Neonatal intensive care*[b]	-.195	-1.291	---	---	---	---	-.460	-1.828	---	---	---	---
	(-1.31	(-2.37)	---	---	---	---	(-2.86)	(-3.08)	---	---	---	---
Cigarettes[b]	---	---	.114	.142	.052	.084	---	---	.126	.264	.476	.619
	---	---	(2.57)	(3.06)	(.98)	(1.50)	---	---	(.76)	(1.54)	(1.56)	(2.01)
Low birth weight*[b]	.752	1.239	---	---	---	---	1.056	1.340	---	---	---	---
	(8.56)	(4.25)	---	---	---	---	(7.25)	(2.69)	---	---	---	---
Preterm*[b]	---	---	.416	.490	---	---	---	---	.318	.433	---	---
	---	---	(13.92)	(5.97)	---	---	---	---	(10.89)	(4.12)	---	---
Births to young/old women*[b]	---	---	---	---	.130	.172	---	---	---	---	.154	.098
	---	---	---	---	(13.68)	(8.73)	---	---	---	---	(6.32)	(2.11)
Ln population density	.131	.042	.094	.093	.182	.222	.130	.225	.346	.393	.010	-.062
	(3.18)	(.62)	(5.94)	(5.18)	(9.02)	(8.62)	(1.05)	(1.18)	(9.18)	(7.45)	(.15)	(-.68)
Constant	6.520	3.605	2.283	2.489	5.667	2.604	3.746	1.584	4.335	2.763	10.930	14.418
	(6.72)	(1.12)	(4.40)	(2.10	(10.94)	(2.04)	(1.62)	(.19)	(2.82)	(.96)	(3.76)	(3.78)
Adjust R^2	.198		.372		.389		.183		.378		.308	
F	26.90	19.23	63.24	47.17	81.21	53.66	13.14	5.06	33.95	22.66	25.16	22.51
N	632	632	632	632	632	632	327	327	327	327	327	327
Wu test, F=	4.30		10.89		5.05		2.87[c]		5.27		3.41	

[a] Asymtotic t-ratios in parentheses. The critical asymtotic t-ratios at the 5 percent level are 1.64 for one-tailed and 1.96 for a two-tailed test. In this table and others that contain regression results, the F-ratio associated with each regression is significant at the 1 percent level unless otherwise indicated.

[b] Endogenous

[c] Significant at the 5 percent level.

Table 4

Quasi-Structural Equations of Infant Health

Independent Variables	Whites: Neonatal Mortality OLS (4-1)	Whites: Neonatal Mortality TSLS (4-2)	Whites: Low Birth Weight OLS (4-3)	Whites: Low Birth Weight TSLS (4-4)	Whites: Preterm OLS (4-5)	Whites: Preterm TSLS (4-6)	Blacks: Neonatal Mortality OLS (4-7)	Blacks: Neonatal Mortality TSLS (4-8)	Blacks: Low Birth Weight OLS (4-9)	Blacks: Low Birth Weight TSLS (4-10)	Blacks: Preterm OLS (4-11)	Blacks: Preterm TSLS (4-12)
Teen family planning users*[b]	-.015	-.023	.006	.012	-.006	-.007	-.026	-.150	.003	.003	.005	.032
	(-1.51)	(-1.05)	(1.41)	(1.40)	(-1.09)	(.79)	(-1.34)	(-2.94)	(.43)	(.17)	(.48)	(1.27)
Abortion rate[b]	-.046	-.063	-.015	-.026	-.010	-.020	-.062	-.149	-.020	-.042	-.080	-.113
	(-6.24)	(-5.83)	(-4.60)	(-5.46)	(-2.54)	(-3.58)	(-2.10)	(-3.13)	(-1.95)	(-2.57)	(-4.74)	(-4.19)
Prenatal Care*[b]	-.049	-.097	-.029	-.063	-.049	-.082	.001	-.019	-.004	-.053	-.039	-.107
	(-6.26)	(-7.39)	(-8.47)	(-10.80)	(-12.11)	(-12.27)	(.05)	(-.48)	(-.75)	(-4.42)	(-3.88)	(-5.43)
Neonatal intensive care*[b]	-.037	-.501	---	---	---	---	-.387	-1.572	---	---	---	---
	(-.24	(-.99)	---	---	---	---	(-2.23)	(-2.65)	---	---	---	---
Cigarettes[b]	.543	.552	.203	.213	.220	.222	.661	.932	.253	.354	.294	.624
	(4.77)	(4.68)	(4.05)	(3.93)	(3.75)	(3.60)	(1.20)	(1.56)	(1.31)	(1.72)	(.92)	(1.84)
Young/old women*	10.875	14.266	1.104	5.074	-1.140	1.928	7.188	-4.911	10.338	2.816	9.330	-3.702
	(3.52)	(3.86)	(.75)	(2.99)	(-.72)	(1.00)	(.78)	(-.42)	(3.21)	(.70)	(1.75)	(-.55)
Ln population density	.260	280	.126	.152	.068	.087	.497	.589	.369	.298	-.051	-.145
	(6.21)	(5.86)	(6.89)	(7.18)	(3.15)	(3.62)	(3.74)	(3.49)	(7.92)	(5.38)	(-.66)	(-1.58)
Constant	4.671	7.919	5.922	7.244	9.853	11.564	7.960	17.662	5.477	11.331	14.762	21.812
	(3.16)	(4.35)	(9.13)	(8.68)	(12.98)	(12.15)	(1.23)	(2.16)	(2.42)	(4.05)	(3.94)	(4.73)
Adjusted R^2	.152		.178		.222		.051		.173		.229	
F	17.09	19.04	23.73	29.91	31.10	31.19	3.50	4.33	12.39	13.78	17.18	18.17
N	632	632	632	632	632	632	327	327	327	327	327	327
Wu test, F=	11.19		25.85		16.23		3.43		10.51		11.10	

[a] Asymtotic t-ratios in parentheses. The critical asymtotic t-ratios at the 5 percent level are 1.64 for one-tailed and 1.96 for a two-tailed test. In this table and others that contain regression results, the F-ratio associated with each regression is significant at the 1 percent level unless otherwise indicated.

[b] Endogenous

clinics, abortion providers and hospitals with neonatal intensive care units are strongly associated with increased family planning use by teenagers, the abortion rate and the number of days in a neonatal intensive care unit per birth respectively. In short, the own shadow price effects are as hypothesized. This is a noteworthy result for it reinforces the notion that accessibility to medical services, the fees for which are often covered by insurance, plays a major role in the utilization of these resources.

Turning to the estimated production functions, the total effect of abortion on race-specific birth outcomes in the quasi-structural regressions is negative and significant in every case, regardless of whether the estimates were obtained by OLS or TSLS. However, since the null hypothesis of zero correlation between the inputs and the error term is never accepted, this discussion will concentrate on the TSLS results. The differences between the OLS and TSLS estimates are considered below.

In general, the coefficient of abortion is approximately two times greater for blacks than it is for whites. This underscores the importance of estimating race-specific infant health production functions whenever possible. Moreover, these results support the conclusions reached by Corman and Grossman (1985) from their reduced form estimates as to the significance of abortion for neonatal mortality.

A comparison of the estimated structural and quasi-structural equations (Tables 3 and 4) sets forths the direct and indirect effect of abortion on infant health. For example, comparing regressions 3-2 and 3-8 with 4-2 and 4-8 reveals that the abortion coefficient falls by 63 percent in the case of whites and 48 in case of blacks when the percentage of low-birth weight births is held constant in the neonatal mortality equation. Nevertheless, for both races the risk-specific effect of abortion remains significant at the .05 level for a one-tailed test. This suggests that the process of fetal selection encouraged by abortion may serve to improve the survivability of risk-specific births as well as reducing the incidence of prematurity.

A notable result is that the direct effect of abortion on black neonatal mortality (-.081) is greater than its indirect effect through prematurity (-.068).[14] Furthermore, the direct effect is obtained from a specification that controls for the percentage of low-birth weight births, the most important predictor of newborn survivability (Institute of Medicine 1985). One possible explanation is that perinatal and neonatal care are incompletely measured. As Corman and Grossman (1985) point out, the increased availability of abortion lowers the cost of fertility control, increases the relative cost of a birth and enhances the resources devoted to each birth. Given the link between increased use and availability (see Appendix), the utilization of abortion may be correlated not only to the quantity, but also the quality of perinatal and neonatal care. Measures for the latter are lacking. A more speculative explanation for the sizeable direct effect of abortion is that better planned births may be "more wanted." Data showing that black births have experienced the greatest increase in "wantedness" is coincident with this result (NCHS 1985).

The impact of abortion on low-weight and preterm births also differs by race. In the case of whites, the abortion coefficient in the low-birth weight equation falls but remains significant at the .01 level

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[14] The direct effect is measured by the coefficient of abortion in regression 3-2; the total effect is from the abortion coefficient in regression 4-2; the indirect effect is obtained from subtracting the former from the latter.

when the percentage of preterm births is held constant (regression 3-4 versus 4-4). However, controlling for the percentage of births to young and old women completely eliminates the effect of abortion on preterm births (regressions 3-6 versus 4-6). Clearly, then, one means by which abortions improves infant health is by shifting the age distribution of births.

The results are reversed with respect to blacks. It is evident from comparing regressions 3-10 and 4-10 that abortion has no effect on low birth weight holding the percentage of preterm births constant; yet, unlike in the white preterm regressions, the effect of abortion falls by only twenty-six percent and remains significant at the .01 level when the percentage of births to women in high-risk age categories is included (regressions 3-12 versus 4-12). Put differently, abortion use by whites lowers the percentage of full-term light births, while abortion use by blacks lowers the percentage of premature light births. Full-term light infants, it should be noted, are likely to have congenital defects and life-long health problems (Beck and van der Berg 1975). Thus, these results imply that white women have a greater propensity to identify and abort defective fetuses than black women. Blacks on the other hand, may use abortion primarily as a substitute for other forms of contraception thus lessening the link between abortion and fetal selection. Further evidence for this explanation is the fact that the black abortion rate that is almost triple the white rate.

Organized teenage family planning use operates inconsistently. For both blacks and whites its effect on neonatal mortality increases in absolute value and becomes significant when the percentage of low-birth weight births is held constant. Its coefficient is insignificant and often positive in the other birth outcome equations. One explanation is that a high percentage of teenage users may be indicative of a relatively high proportion of sexually active adolescents. Even if family planning use is effective in preventing pregnancies among users, the proportion of teens who give birth may still be greater than in counties where both adolescent sexual activity and orgnaized family planning use are relatively less. A better measure might be the ratio of teenage users to the number of sexually active teens. Unfortunately, such data is unavailable.

The negative association between neonatal mortality and organized teenage family planning use may reflect the fact that high-use counties may be indicative of users that have been integrated into a network of prenatal and perinatal care. For example, Chaime et al (1982) note that clinics in counties that serve a relatively large proportion of teens at risk of becoming pregnant tend to provide additional services such as prenatal care. In short, although births to these young women are more likely to be problematic, they may also receive better care, and thus their offspring may have a higher probability of survival.

Estimates from the quasi-structural regressions leave little doubt as to the positive correlation between early initiation of prenatal care and infant health. Whether this reflects the importance of medical intervention or is due to the unobservable set of characteristics associated with a self-selected group of women for whom early prenatal care is but one expression of healthy behavior cannot be determined from the data. The results do indicate that for blacks, the direct effect of prenatal care on the prevention of low birth weight is statistically zero (regression 3-10) and that the improvement in infant health associated with early care is derived from prolonged gestation (regression 3-12). This is consistent with Harris'(1982) findings based on a sample of black births in Massacusetts between 1975 and 1976. For whites, the coefficient of prenatal care in the estimated structural equation of low birth weight is significant at the .01 level (regression 3-4). A significant direct effect of prenatal care on birth weight is frequently cited in the literature (Rosenzweig and Schultz 1983b; Lewit 1983; Showstack et al 1984).

The relationship between neonatal intensive care and an increased likelihood of survival is confirmed by the data (regressions 3-2, and 3-8). Since neonatal intensive care use is determined by birth weight, the results from the structural equations are emphasized. Stated differently, the indirect effect of neonatal intensive care is zero. However, replacing low birth weight with its endogenous determinants, and using instrumental variable techniques to correct for the correlation between neonatal intensive use and the disturbance term should yield a coefficient that is equivalent to the one obtained by the specification including birth weight. As a means of gauging the statistical difference between the neonatal intensive coefficients from the two specifications, it was assumed the coefficients from the regressions including birth weight (regressions 3-2 and 3-8) were the true population parameters. Thus, it was tested whether the neonatal intensive care coefficients from the equations excluding birth weight (regressions 4-2 and 4-8) differed significantly from these "true" estimates. The null hypothesis of no difference was easily accepted. The t-statistics were 1.06 and .30 for whites and blacks respectively.

As aniticipated, smoking is inversely related to infant health. The results for whites are congruent with the frequently reported finding (Meyer et al 1976; Institute of Medicine 1985) of a significant deleterious effect of maternal smoking on birth weight (regressions 4-

4). Even when the percentage of preterm birth is held constant in the equation predicting low birth weight, the coefficient of smoking remains significant at the .01 level (regression 3-4). In the case of blacks, the results are less convincing. Only once does the smoking coefficient in a birth outcome equation have a t-ratio greater than two (regression 3-12). The lack of a consistently significant effect between smoking and black infant health may result from the fact that the smoking measure is neither race- nor sex-specific.

As mentioned in the introduction, the decline in the U.S. neonatal mortality rate over the past decade has been largely attributed to advances in the management of newborn care (Kleiman et al 1979; Lee et al 1980; David and Siegal 1983) . However, in all these studies the relative contribution of abortion to the increased healthiness of newborns has not been calculated. Therefore, to compare the importance of abortion and neonatal intensive care use on early infant deaths, the percentage change in the neonatal mortality rate given a one percent increase in each of these inputs was computed.[15]

Because the decline in neonatal mortality has been primarily an improvement in birth-weight specific survivability, coefficients from the structural equations (regressions 3-2 and 3-8) were used to calculate the elasticities. For both races, reductions in the neonatal mortality rate are more responsive to changes in neonatal intensive care

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15 A potential problem with such an exercise is that an increase in abortion use resulting from a change in the availability of abortion may, for example, cause the use of family planning clinics to fall. In short, these computations do not provide reduced form effects that are required to evaluate the potential impact of alternative policies to lower neonatal mortality rates. Despite these shortcomings, they do reflect the benefits of expanding the use of one input while holding the others constant, and thus, offer insights as to relative effects.

use. The elasticity is -.095 for whites and -.170 for blacks. The corresponding magnitudes for a one percent change in the abortion rate are -.066 and -.012 respectively. The finding that abortion has two-thirds the impact of neonatal intensive care on infant deaths is qualified by the effect of abortion on prematurity. Specifically, including the reduction in low-weight and preterm births brought about by a one percent increase in the abortion rate raises the abortion elasticity to -.181 and -.228 for whites and blacks respectively. Both are larger than the corresponding neonatal intensive care elasticities.

These results should be interpreted with some caution. First, the number of inpatient days in a Level II or Level III neonatal intensive care hospital is one measure of the management of high-risk births. The increase in cesarean sections, for example, has been associated with significant gains in perinatal survivability among infants weighing 2000 grams of less (Williams and Chen 1983). Second, the indirect effect of abortion on mortality via prematurity may be overstated due to the restriction that such risk factors as age, marital status, and parity impact on birth weight only through gestation.

Nevertheless, the relative magnitude of the risk-specific response of neonatal mortality to changes in the abortion rate is an important finding. What cannot be overemphasized is that this result is obtained while holding constant the percentage of low-birth weight births. As mentioned above, there is overwhelming evidence that low birth weight is the most important predictor of neonatal mortality. Moreover, these results are very similar to the rankings obtained from Corman and Grossman's (1985) reduced form estimates.

B. Health Endowment Effects on Input Utilization

As discussed previously OLS may produce biased and inconsistent estimates if the correlation between the health inputs and the residuals is not corrected. Comparison of the quasi-structural estimates (Table 4) obtained by OLS as opposed to TSLS supports this hypothesis. For instance, the TSLS prenatal care coefficient in Table 4 is at least two times greater than it OLS counterpart for both races in all three birth outcomes. The impact of neonatal intensive care on neonatal mortality increased by a factor of four when estimated by TSLS. Although this difference is not as great with respect to the abortion coefficient, OLS consistently yields a smaller effect of abortion on infant health.

Making the same comparisons with the structural equations (Table 3) shows a more inconsistent pattern. This is in line with the contention of Corman, Joyce and Grossman (1985) that controlling for the relevant birth outcomes and endogenous risk factors may be an effective proxy for the endowment term. Although the Wu test indicates that a correlation between the health inputs and the residuals persists in the structural equations, this could be caused by any form of misspecification such as an incorrect functional form or some other ommitted variable. Moreover, the Wu statistic in the structural equations is half the magnitude of its quasi-structural counterpart in five of six instances. This suggest that when controlling for the appropriate intermediate birth outcome, the tradeoff between efficiency and consistency becomes more difficult to evaluate.

As further evidence that the use of health inputs is conditioned in part, on the unobserved health endowment imbedded in the error term, Rosenzweig and Schultz (1983b) regressed each input on the residuals from the equation predicting birth weight. Their results were consistent with the interpretation that women demand more prenatal care, for example, based on the expectation of a potentially adverse birth outcome.

A similar procedure was followed here. However, instead of using the residuals from one equation, each health input was regressed on the first principal component extracted from the three sets of residuals obtained from the structural equations of mortality, birth weight, and gestation. The same regressions were then reestimated, but the first principal component extracted from the residuals of the the quasi-structural equations was used instead.

The objective was two-fold: first, to see whether the results obtained by Rosenzweig and Schultz could be duplicated at the county-level; and second, to compare the regression results using the principal component from the structural equation residuals as the explanatory variable with the results from using the principal component obtained from the quasi-structural residuals. If the percentage of low-birth weight births in the mortality equation, the percentage of preterm births in the birth weight equation, and the percentage of births to high-risk women in the gestation equation are effecitve proxies for the health endowment, then one should see little evidence of remedial behavior when the health inputs are regressed on the principal component extracted from the residuals of the structural equations.

Principal components seemed an improvement over using the residuals from one equation. If the three equations are reasonably well-specified, then the element that explains the greatest portion of the variance in the all three residual series should be the health endowment.[16] Furthermore, since only the sign and statistical significance of the relationship between the principal component and the health inputs is relevant, the fact that the residuals are in different units is not a problem.

The results are presented in Table 5. Use of the endowment measure estimated from the quasi-structural residuals supports the conclusions reach by Rosenzweig and Schultz (1983b). In particular, counties in which the probability of an adverse outcome is higher than predicted evidence greater utilization of prenatal and neonatal care. This holds for whites and blacks. The same is true for abortion but only in the case of blacks is the effect significant. The result that organized teenage family planning use is lower when the probability of an adverse outcome is higher than expected may be due to the same factors outlined previously for its inconsistent impact on infant health.

There is little evidence of remedial behavior on the part of pregnant women when the same inputs are regressed on the endowment measure constructed from the residuals of the structural equations. This is interpreted as support for the use of the relevant intermediate birth outcome as a proxy for the aggregate health endowment. Moreover, the percentage of variance accounted for by the first principal component falls substantially when the residuals from the structural equations are used. This is also consistent with the interpretation that the endowment is effectively controled for in the structural equations.

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[16] If the residual from only one series were used, like those from the mortality equation, there is a greater liklihood that an omitted variable, such as the quality of care at delivery, may dominate the effect of the endowment. However, since the quality of care at delivery is not a determinant of birth weight, or gestational age, there will be less of a likelihood that a potential omitted variable common to all three birth outcomes will dominate the variation in the health endowment.

Table 5

Comparison of T-statistics Associated with Regressing each Health Input on the Estimated Health Endowment Obtained from the Residuals of the Structural and Quasi-structural Equations.[a]

	Endowment measure from quasi-structural equations		Endowment measure from structural equations	
	Whites	Blacks	Whites	Blacks
Abortion rate	.96	1.81	-.05	-.39
Prenatal care*	9.80	7.01	-4.08	-.89
Teen family planning use*	-.86	-1.05	-1.10	3.23
Neonatal intensive care	2.38	2.40	-.17	1.92
Proprotion of variance accounted by first principal component	.60	.56	.40	.41

[a] The critical t-ratios at the 5 percent level are 1.64 for a one-tailed test and 1.96 for a two-tailed test.

V. Conclusion

The results from this study underscore the need to incorporate the effect of induced abortion in examining the determinants of birth outcomes. Although the reported association between abortion and infant health needs to be examined at the individual level, this study becomes part of a growing body of literature confirming this association in both a reduced form (Grossman and Jacobowitz 1981; Corman and Grossman 1985) as well structural model of infant health. Furthermore, a blueprint now exist for pursuing this research with individual births.

There is growing speculation that the benefits of neonatal technology may be reaching the point of diminishing returns (McCormick 1985). Consequently, if the U.S. is to sustain the rate of decline in early infant deaths that it has enjoyed over the past twenty years, greater emphasis will have to placed on lowering the incidence of prematurity. The findings from this study support the contention that by preventing unwanted births, abortion reduces the percentage of preterm and low-weight infants. Few would argue that averting unintended pregnancies is the most preferable strategy for reducing unwanted births. However, the dramatic rise in deliveries to unmarried teenagers over the past decade must be addressed if substantial gains are to be made in lowering the incidence of prematurity. Unless dramatic changes occur in teenage sexual activity, or contraceptive use among adolescents increases significantly, abortion will remain an important option for many pregnant women. According to the estimates from this study, legislative attempts to ban abortion would have a negative impact of birth outcomes.

Finally, the results from the quasi-structural regressions also support the proposition of Rosenzweig and Schultz (1982, 1983b) that the correlation between the health inputs and the unobserved endowment term tends to mask the effect of the behavioral inputs on infant health. This source of bias may become increasing more important with respect to abortion as rapidly advancing techniques of prenatal diagnosis, such as amniocentesis, enjoy more widespread utilization. Such procedures dramatically enhance the information known to parents concerning the health of the fetus. Decisions to abort, based on this knowledge, augment the process of fetal selection and increase the correlation between the use of abortion and the endowment term.

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Appendix

Table A-1

Definitions of Reduced Form Variables

Variable	Definitions
Percent poor*[b]	Percentage of women aged 15-44 with family income less than 200 percent of the poverty level in 1980
High school education*[c]	Percentage of women aged 15-49 who had at least a high school education in 1970
Medicaid elegibility-1	Dichotomous variable that equals one if county is in state that covered first-time pregnancies under Medicaid to financially eligible women in the period 1976-1978
Medicaid elegibility-2	Dichotomous variable that equals one if county is in state that covered first-time pregnancies under Medicaid only if no husband was present or if the husband was present but unemployed and not receiving unemployment compensation in the period 1976-1978
Medicaid elegibility-3	Dichotomous variable that equals one if county is in state that covered first-time pregnancies under Medicaid only if no husband was present in the period 1976-1978
Medicaid coverage	Dichotomous variable that equals one if county is in state in which Medicaid paid for newborn care under the mother's number but allowed pregnant women to register their "unborn children" with Medicaid in 1981
Medicaid payments	State specific average annual Medicaid payment per adult recipient in AFDC families in fiscal 1976
Family planning clinics	Number of organized family planning clinics in 1975 per 1,000 women aged 15-44 with family income less than 200 percent of the poverty level in 1975
Community health centers[d]	Sum of maternal and infant care (M and I) projects and community health centers (CHCs) in 1976 per 1,000 women aged 15-44 with family income less than 200 percent of the poverty level in 1975; numerator termed Bureau of Community Health Services (BCHS) projects

Abortion providers	Three-year average of abortion providers centered on 1976 per 1,000 women aged 15-44 in 1975
Neonatal intensive hospitals	Sum of state-specific number of hospitals with Level II, Level III, or Levels II and III neonatal intensive care units in 1979 per 1,000 women aged 15-44 in state in 1975
AFDC payments	Three-year average AFDC payment per recipient centered on 1977
Young and old women*[c]	Ratio of women ages 15 to 19 and 40 to 44 over the total number of women 15 to 44 in 1975
Ln population density (POPDEN)	The natural logarithm of the ratio of the population in 1975 to the area per square mile

Notes to Table A-1

[a]An asterisk (*) next to a variable means that it is race-specific. All variables are calculated for counties unless otherwise specified.

[b]Variable is available for nonblacks and blacks as opposed to whites and blacks

[c]Variable is available for whites and nonwhites as opposed to whites and blacks

[d]Since the numerator of this variable is not race-specific, the denominator also is not race-specific. The denominator is obtained by applying the race-specific percentage of women aged 15-44 with family income less than 200 percent of the poverty level in 1980 to the race-specific number of all women aged 15-44 in 1975.

Table A-2

Means and Standard Deviations of Reduced Form Variables[a]

	Whites		Blacks	
	mean	std	mean	std
Percent poor*	26.276	8.073	55.192	9.463
High school education	63.446	6.752	44.095	9.243
Medicaid elegibility-1	.427	.495	.286	.452
Meidicaid elegibility-2	.151	.358	.179	.383
Medicaid elegibility-3	.091	.287	.938	.240
Medicaid coverage	.920	.272	.938	.240
Medicaid payments	442.526	144.018	439.709	138.249
Family planning clinics	.271	.194	.272	.215
Community health centers	.019	.036	.026	.033
Abortion providers	.056	.044	.057	.037
Neonatal intensive hospitals	.010	.044	.010	.003
AFDC payments	78.944	22.341	71.739	27.662
Young and old women*	.336	.022	.350	.026
Ln population density	6.369	1.557	7.315	1.691
Sample size	632.		327.	

[a]An asterisk (*) next to a variable means that it is race-specific. Means and standard deviations are weighted by the three year sum of race-specific births in 1976-1978.

Table A-3

Ordinary Least Squares Input Demand Equations--Whites[a]

Independent Variables	Teenage Family Planning Users	Abortion Rate	Prenatal Care	Neonatal Intensive Care
Abortion providers	36.991	60.097	-9.777	1.756
	(6.80)	(9.79)	(-1.49)	(4.38)
Family planning clincs	9.267	1.350	1.404	-.189
	(7.90)	(1.02)	(.99)	(-2.18)
Community health centers	33.404	-14.372	-10.650	.636
	(5.54)	(-2.11)	(-1.47)	(1.43)
Neonatal intensive hospitals	241.713	-891.570	88.830	9.338
	(3.76)	(-12.31)	(1.15)	(1.97)
Medicaid elegibility-1	-.027	5.206	1.536	.158
	(-.38)	(6.44)	(1.78)	(3.00)
Medicaid elegibility-2	-1.762	-4.250	5.068	.187
	(-2.11)	(-4.51)	(5.04)	(3.04)
Medicaid elegibility-3	.490	.338	1.037	.229
	(.75)	(.46)	(1.32)	(4.75)
Medicaid payment/recipient	.004	.016	-.006	.001
	(2.46)	(8.18)	(-2.74	(5.87)
Medicaid coverage	-3.857	4.390	1.388	-.128
	(-4.82)	(4.86)	(1.44)	(-2.18)
High school education*	.011	-.087	.341	.0001
	(.30)	(-1.97)	(7.21)	(.04)
Percent poor*	.087	-.001	-.490	-.003
	(2.66)	(-.04)	(-12.45)	(-1.31)
AFDC payments	.015	.012	-.028	-.003
	(.81)	(.60)	(-1.28)	(-2.29)
Young/old women*	-59.377	-26.296	79.301	.255
	(-5.22)	(-2.05)	(5.79)	(.30)
Ln population density	.019	.481	47.498	-.020
	(.12)	(-2.64)	(6.65	(1.71)
Constant	28.924	28.241	47.498	.520
	(3.19)	(4.22)	(6.65)	(1.19)
F	29.44	67.50	38.32	7.27
Adjsuted R^2	.387	.596	.453	.122
Sample size	632	632	632	632

[a]The t-ratios are in parentheses. The critical t-ratios at the 5 percent level are 1.64 for a one-tailed test and 1.96 for a two-tailed test. The F-ratio associated with each regression is significant at th e 1 percent level.

Table A-4

Ordinary Least Squares Input Demand Equations--Blacks[a]

Independent Variables	Teenage Family Planning Users	Abortion Rate	Prenatal Care	Neonatal Intensive Care
Abortion providers	-8.520	26.801	9.292	4.834
	(-.57)	(3.08)	(.56)	(1.41)
Family planning clincs	10.373	3.005	3.550	-.112
	(4.30)	(2.16)	(1.35)	(-.40)
Community health centers	44.606	-14.189	26.023	.247
	(2.85)	(-1.57)	(1.52)	(.14)
Neonatal intensive hospitals	-278.433	-1014.390	440.760	107.04
	(-1.47)	(-9.24)	(2.12)	(4.84)
Medicaid elegibility-1	-.348	7.015	-3.015	-.030
	(-.17)	(6.09)	(-1.38)	(-.13)
Medicaid elegibility-2	2.009	-6.648	3.620	.199
	(1.19)	(-6.80)	(1.96)	(1.01)
Medicaid elegibility-3	-4.705	-.622	5.175	.789
	(-2.95)	(-.67)	(2.96)	(4.24)
Medicaid payment/recipient	-.019	.017	-.016	.001
	(-3.84)	(6.07)	(-3.03)	(1.31)
Medicaid coverage	.253	-.196	-7.856	-.643
	(.12)	(-.17)	(-3.53)	(-2.71)
High school education*	.014	.064	.355	-.003
	(.15)	(1.15)	(3.36)	(-.31)
Percent poor*	.113	.026	-.319	-.002
	(1.37)	(.55)	(-3.53)	(-.16)
AFDC payments	-.066	-.030	.052	.012
	(-1.32)	(-1.06)	(.96)	(2.07)
Young/old women*	-4.343	-64.309	-.192	-.286
	(-.18)	(-4.54)	(-.01)	(-.10)
Ln population density	1.243	-.029	-1.806	-.058
	(3.23)	(-.42)	(-4.29)	(-1.30)
Constant	21.876	45.911	79.014	.286
	(1.66)	(6.04)	(5.49)	(.19)
F	10.38	59.26	13.12	44.00
Adjsuted R^2	.287	.714	.342	.166
Sample size	632	632	632	632

[a]The t-ratios are in parentheses. The critical t-ratios at the 5 percent level are 1.64 for a one-tailed test and 1.96 for a two-tailed test. The F-ratio associated with each regression is significant at th e 1 percent level.