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ABSTRACT

This paper specifies and estimates an econometric model of low and very low birthweight rates for counties in the U.S. for the years 1975-1984. We focus on the impact of several specific public policy actions on use of prenatal care and the subsequent effect on birthweight outcomes. Our results point to strong racial differences in the impact of prenatal care on low birthweight rates. We also find that for the white population changes in income eligibility standards and expanded availability of publicly financed maternal and infant clinics have the strongest impacts on low birthweight rates.

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I. Introduction

Following a decade (1969-1979) when the percentage of women receiving prenatal care in their first trimester of pregnancy rose steadily, the years since 1979 have showed no gains in the timely initiation of prenatal care (NCHS 1979, 1983). The percentage of women initiating prenatal care in the first trimester was the same in 1983 as it was in 1979 (Swartz 1984). Although there is some uncertainty concerning the exact contribution of adequate prenatal care to past reductions in low birthweight (LBW) rates (prior to 1979), there is a growing consensus that expanding access to prenatal care is a key component in any strategy to shift the birthweight distribution upward (IOM 1985).

National trends in health care funding for the poor may be contributing to problems with access to care. According to recent estimates the number of Americans without insurance is growing. In 1979 roughly 14.6% of the non-aged population was uninsured compared to 17.5% in 1986 (CRS 1988a). The growth rate in the number of uninsured is even higher for those with incomes below 200% of the federal poverty line. Moreover, increasingly restrictive income eligibility standards for Medicaid arising from failure to adjust the standard for increases in inflation seriously decreased the number of eligible poor women. The 1985 Institute of Medicine Report on Preventing Low Birthweight (IOM 1985) emphasized problems of financial and geographic access to high quality prenatal care as an impediment to improving the birth weight distribution.

The analysis reported here is concerned with developing an econometric model of the effects of specific resources and public policy actions for poor and near poor women on the rate of low birthweight and very low birthweight (VLBW) births. Specifically, our model focuses on the direct effects of the use of prenatal care and abortion services on rates of LBW and on the indirect affects of the availability of prenatal care providers, Medicaid program structure and the availability of abortion providers on birthweight.

Previous research has examined the impacts of individual policy changes at the state or local level (Strobino, Chase, Kim, et al. 1986, Peoples, et al. 1984). Few studies have examined the effect of a variety of public programs on birthweight for nationally representative data on births. In this study we make use of a quasi-structural birthweight production function to establish the empirical relations. The data consist of pooled time series and cross-sections of counties for the years 1975 to 1984. We study all counties in the United States with populations of 10,000 or more whites or 5,000 or more blacks (based on the 1980 census).

The paper is organized into 6 sections. The second section presents some general theoretical remarks regarding specification of the empirical model; the third section describes the data and measurement of key variables; the fourth section discusses estimation issues. Results are presented in the fifth section. The paper concludes with a discussion of findings.

II. Theoretical Remarks

The analytical point of departure for our study is Grossman's (1972) thesis that households both produce and demand health, and the application of this thesis to the study of birth outcomes (Lewit 1983, Rosenzweig and Schultz 1982, Corman, Grossman and Joyce 1987). Birth outcomes can be viewed as determined by a "production" process. Through various clinical, lifestyle and environmental mechanisms, a number of "inputs" (physician services, diet, and exercise) act in combination to produce a particular birth outcome.

Our conceptual approach differs from previous research in this area in several respects. The most important difference, described below, is in the degree to which we believe the birthweight and gestational age production functions can be identified. We also estimate age as well as race specific models of birth weight outcomes. A third difference is our reliance on data from a panel of counties over a ten year period rather than a cross section of counties for a single time period.

The system of equations described below is designed to allow us to trace the mechanisms by which several factors influence birthweight at the county level. The basic model we specify consists of six equations:

(1) $b = b(v, a, s, f, g, x, e)$

(2) $g = g(v, a, s, f, x, e)$

(3 - 6) $v, a, s, f = h(m, I, x, z, e)$

The first two equations are structural production functions, while the last four equations are input demand functions. Equation (1) states that the rate of low (or very low) birthweight births (b) depends on use of prenatal care (v), use of abortion services (a), use of family planning services (f), the rate of preterm births (g), a vector of exogenous risk factors (x), and the mother's biological endowment that is presumed to be at least partially observable to the mother but not to the researchers (e). Equation (2) treats the rate of preterm births or gestational age as an endogenous variable. This assumption is controversial and sensible arguments exist for treating prematurity as either endogenous or exogenous (IOM 1985, Herron, Katz and Creasy 1981). As we will see shortly this issue is not one of major practical concern in this analysis. Equations (3-6) are input demand equations which are posited to depend on the structure of the Medicaid program (m), the level of income (I), a vector of exogenous risk factors (x), the availability of various maternity services (z) and the unobserved health endowment of the mother (e).

There are several important features of this model which influence the choice of an estimation strategy, the final structure of the model and the interpretation of results. The first is that the form of the model is a recursive one, where the disturbance terms across equations are correlated. Rosenzweig and Schultz (1982) point out that the implication of correlated disturbances is that direct estimation of the production functions (1) and (2) will lead to biased coefficient estimates. The reason for the bias is that a pregnant woman may have knowledge of factors regarding her health endowment (e.g. genetical make-up) that may influence both her choice of inputs (say use of prenatal care to monitor a possible problem) and the birth outcome. The implication of such potential bias is that ordinary least squares may not be the most appropriate estimation technique for this model (see discussion of estimation below).

A second important feature of the model relates to the specification of the prematurity production function (equation (2)). Corman, Grossman and Joyce (1987) and Joyce (1987) omit smoking from the preterm birth production function and thereby assume that smoking has no impact on gestational age, we are reluctant to make this assumption. To our knowledge only one study supports this proposition, (the work by Rosenzweig and Schultz (1982)) which did not include a representative sample of black child bearing women. If this assumption is not made then the rank and order conditions for identification of equations (1) and (2) are not met. (Because all the right hand side variables in (2) also appear in (1), inclusion of a predicted value for g in (1) based on an estimate of (2) results in singularity of the cross-product matrix.) Practically, this is of small consequence since a substantial portion of counties are missing data on gestational age.

A third feature of the model is that for variables such as family planning services we do not directly observe the volume of services consumed in a county. Rather, we substitute determinants of utilization, the availability of family planning clinic services (f') and income (I), into the production function.

With these points in mind we rewrite the model described in equations (1-6) in the following manner:

(7) $b = b(v, a, s, f', I, x, e)$

(8-10) $v, a, s = h(m, I, x, z, e)$

We refer to this model as a quasi structural model of birthweight outcomes. It is quasi structural in that we have substituted exogenous variables for both the endogenous family planning and preterm birth variables. The result is that the impact of, for example, prenatal care (V) on birthweight is made up of two components: the direct impact of care on birthweight and the indirect impact which works through the effect of prenatal care on gestational age ($db/dv = \partial b/\partial v + \partial b/\partial g \; \partial g/\partial v$). We are unable to seperate the two effects in our model. Thus, we lose some ability to understand the mechanism by which factors such as use of prenatal care influence low birthweight rates.

III. Data Used in the Analysis

The empirical analysis of the quasi-structural birthweight production functions uses the county as the unit of observation. The study design is a pooled time series cross-sectional study for the years 1975 to 1984. The study population consists of all counties in the United States with populations of 10,000 or more whites or 5,000 or more blacks (based on the 1980 census).

The empirical strategy consists of estimating a series of regression models which trace the structural relationships outlined in the theoretical model given above. Separate regressions are estimated for black and white birthweight outcomes and also by three age groups (less than 20 years, 20-34 years and 35 and older). The reasons for separate race and age specific regressions is that race and age are thought to interact with a variety of other explanatory variables. Moreover, the LBW rate for blacks is roughly twice as large as the rate for whites. Race and age are also correlated with use of prenatal care, abortion services and other explanatory variables.

We combine data from birth records for two years for each county data point to reduce problems related to random fluctuations in rates of LBW births. We estimate our models for counties for even numbered years in 1976-1984. Data from the birth records are combined for the following two year pairs: 1975-76, 1977-78, 1979-80, 1981-82, and 1983-84. The result is a panel of 2,192 counties which meet the white population criterion and 686 counties which meet the black criterion.

Additional criteria were applied for inclusion of counties in the study sample. Counties with more than 30% of observations missing from items on birth records were eliminated from the analysis. Arizona was eliminated from the analysis file since it did not have a Medicaid program during the study period. These sample selection criteria led to reductions in the sample of counties to 2,137 counties in the white sample and 660 counties in the black sample.

1. Data and Measurement of Birthweight Outcomes

The source of data for the measurement of birth outcomes as well as several of the explanatory variables discussed below is the national natality files for the years 1975-1984 produced by the National Center for Health Statistics (NCHS). These files contain information reported on standard birth certificates. During the 10 year study period, some states reported data for 100% of births while others reported a 50% random sample of births. In order to obtain data reflecting the full population of births we multiplied data for states reporting 50% by two.

Two race and age specific measures of (LBW) are used in the analysis. The first is the standard measure of LBW, defined by the percentage of infants weighing less than 2500 grams at birth. The second measure is very low birthweight (VLBW), the percentage of infants weighing less than 1500 grams at birth. The VLBW measure is included because of the exceedingly high mortality rate among these infants relative to those weighing 1500 to 2500 grams (Strobino, Kim, Crawley, et al. 1985, Lee, et al 1980). One disadvantage of this measure is the very low rate of such births.

The age of the mother and the race of the infant are also obtained from the birth records. The reliability of the reporting of maternal age on birth records is excellent as evaluated by making comparisons with census data (NCHS 1985). The infant's race is determined by the races of the parents. When both parents are of the same race the child is assigned that race. If neither parent is white the child is assigned the father's race. If the race of only one parent is known the child is assigned that race. A 1981 study showed that 99.4% of white birth records and 98.6% of nonwhite birth records were complete (NCHS 1981).

2. Measurement and Definition of Endogenous Inputs

The variable used to define the use of prenatal care in our analysis is the percentage of women initiating care in the first trimester of pregnancy.

It has the advantage of being unrelated with the length of pregnancy for live born infants. Corman, Grossman and Joyce (1987) also argue that it does not reflect frequent use by women who develop complications during their pregnancy or who begin pregnancy with underlying medical problems.[1]

The use of abortion services is measured as a two year average of the predicted rate of abortions among residents of a county per fertile woman (15-44 years). The source of the data on the volume of abortions are surveys conducted by the Alan Guttmacher Institute (AGI). The AGI collects data on the volume of abortions and the providers of abortions by county of occurrence. The reporting of data by occurrence created difficulties for the analysis because observations on birth outcomes are reported by county of residence. We therefore used data on county of occurrence to create synthetic estimates of abortion rates by county of residence.[2]

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[1] Standards for prenatal care include not only the early initiation of care but also scheduling of prenatal visits, the frequency of visits, of course, increases with the duration of pregnancy (Kotelchuck 1987).

[2] The specific algorithm involved the following steps: We aggregated abortion data to the regional level using the Health Systems Agency (HSA) regions as a measure of a distinct medical region. The number of abortions at the HSA level we refer to as A_j. We then estimated a regression of the form A_j=A(POP, PROV, I, MDPOP) where POP is the female population of the jth HSA between the ages of 15 and 44, PROV is the number of abortion providers in the HSA, I is the per capita income, and MDPOP is the physician to population ratio. Based on the regression results we used the estimated coefficients to predict the number of abortions in all counties that were in HSAs with positive numbers of abortions. This predicted value is given the symbol A_{ij} denoting the number of abortions in the ith county within the jth HSA. We then created the ratio $R_{ij} = A_{ij}/ \text{sum } A_{ij}$. R is therefore the share of the predicted total abortions within the jth HSA accounted for by the ith county. In order to obtain the final estimate of the number of abortions in a given county we added a constraint that the sum of abortions across counties within an HSA has to add up to the actual total number of abortions. Thus the estimated volume of abortions for the ith county is R_{ij} A_j.

We estimated the predicted levels of abortion for each sample county for all the even numbered years. In order to determine the validity of this approach to estimation we obtained data from vital statistics in Tennessee which records

The third endogenous input in the production function model is the level of smoking by women in a county. We were only able to obtain data on the volume of cigarette sales by state and year for the study period. These data are obtained from the Tobacco Institute and were made available to us by Michael Grossman of the National Bureau of Economic Research. The smoking data have several important shortcomings. First they are not race or age specific and are collected at the state rather than the county level. Second, cigarette sales data can be misleading indicators of consumption because of border crossing for the purpose of purchasing cigarettes in low tax states and bootlegging of cigarettes. Nevertheless, we believe the dangers of omitted variables bias are greater than those stemming from measurement error.

3. Measurement of Exogenous Risk Factors

Five exogenous variables measuring socio-demographic risk, are included in the quasi-structural birthweight production function: parity, education, percentage of female headed households in the community, rate of urbanization and the poverty rate. There is a well established relationship between birth order and LBW (Chase 1977, Selvin and Garfinkel 1972, and Strobino 1982). Birth order is measured by the percentage of first births in an age and race specific group in a given county and obtained from the NCHS natality files.

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all abortions in the state. We used the vital statistics data on the volume of abortions by county for 1982. We compared the vital statistics data with our predicted estimates of the volume of abortions via calculation of a rank correlation statistic. The estimated correlation coefficient for the predicted and vital statistics volume of abortions was 0.875 which was significantly different from zero at the 0.01 level. The results suggest that we were able to develop rather accurate predictions of the volume of abortions by county of residence. It should be pointed out that this variable is neither age nor race specific. The final step involved dividing the predict volume of abortions in a county by the female population age 15 to 44 years.

Education is measured as the median education of persons 25 years and older in the county. It was obtained from the 1980 census compiled in the Area Resource File (ARF) which is administered by the Health Resources and Services Agency in the U.S Department of Health and Human Services (USDHHS). Schooling has been found to be strongly related to rates of LBW (Strobino 1982, Hardy and Mellits 1977 and Rosenwaike 1971). Ideally we had wanted to measure the schooling of the women giving birth from the birth records. However, a number of states did not report maternal education on their birth records for a number of years and California, Texas and Washington did not report education data for any study years. Estimation of the models using education data from the birth certificate would have resulted in a loss of a number of important states.

Out of wedlock births have also been shown to be related to rates of LBW and of preterm births. This variable is generally thought to be a proxy for a variety of factors related to poverty and low social and economic status (USDHHS Secretary's Task Force on Black and Minority Health 1986). Because of frequent missing data on birth records and the lack of information on marital status in 9 states until 1980, we measured the percent of female headed households in the county for 1980. This variable was obtained from census data in the ARF.

Two other factors are included to measure community risk factors: the 1980 poverty rate and the percentage of a county that is urban. The poverty rate, obtained from the ARF, is defined on a race specific basis and measures the percent of persons living in a county in 1980 who have incomes below the federal poverty line. It substitutes for unobservable inputs in the production function such as use of family planning or nutrition intake. The percent of a county's population living in urbanized areas was obtained from the 1980 census.

A final exogenous variable is the availability of family planning services per fertile female as measured by the number of family planning clinics per capita in a county. As discussed above we included this variable because a measure of the use of family planning services was not available. The data to measure this variable were obtained from surveys for the years 1975, 1981 and 1983. For the two year pairs where the data were not available (1977-78 and 1979-80) we interpolated values for counties.[3]

4. Exogenous Variables in the Input Demand Equations

Variables measuring the availability of medical care have been included in numerous studies of health care utilization on the grounds that the supply of medical resources influences important dimensions of the accessibility (travel time, waiting time, willingness of providers to take on new patients etc.). Counties where availability of services is greatest are presumed to have enhanced access to care and in turn greater utilization of care than counties with fewer available services. While some of these variables have not been strong predictors of the use of prenatal care in previous studies, the changes in the supply of care in the study period were substantial. In addition, previous studies have often used non-specific indicators of availability of care. This research sought to improve the measurement of availability of health care resources by focusing specifically on a number of sources of prenatal care.

The first measure of availability is the number of obstetricians and

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[3] We are grateful to Stanley Henshaw of the Alan Guttmacher Institute for his assistance in obtaining and using these data.

gynecologists (OB/GYNs) in patient care per fertile female in a county. These data were obtained from the ARF for the even numbered study years. We also measured the non-OB/GYN physician to population ratio. The third measure, a proxy variable for the availability of hospital based resources such as outpatient clinics, is the ratio of hospital beds to population. These data were taken from the American Hospital Association's Annual Survey for the study years.

Another set of measures of availability pertain specifically to public provision of prenatal care services. Since no data on these services are readily available we constructed a data set for all U.S. Counties which combined information from public documents, a survey of State Maternal and Child Health (MCH) programs and telephone interviews with state MCH directors and/or their staffs. A detailed report of the data collection method and some general description of public MCH programs in the United States for the years 1975-1984 is available from the authors.

We use two key variables from the public prenatal care data set in the input demand equations. They are: the availability of routine and comprehensive prenatal care. Routine prenatal care is defined as prenatal care involving only medical testing and monitoring, including blood tests, monitoring fetal growth, maternal blood pressure measurement and other such procedures. Comprehensive prenatal care includes routine prenatal care plus ancillary services such as maternal counseling, educational activities and nutritional supervision. These variables are both measured as dichotomous indicators that take a value of one if the program was present, and zero otherwise.

The AGI provided data on family planning and abortion providers. The family planning variable was described above. The availability of abortion providers was measured by two variables. The first is the number of providers who performed a relatively small number of abortions (30 to 390 annually) per fertile female. The second measure of abortion availability measured the number of providers performing a large number (over 400 annually) of abortions per fertile female. These data were available for all study years.

In addition to variables representing the availability of health care services, we included variables that measured the structure of the Medicaid and AFDC programs. Four variables were used to characterize Medicaid program structure: the AFDC income eligibility standard, whether or not the state allows first time pregnant women to enroll in Medicaid, whether the state Medicaid program limits the number of visits to physicians or outpatient departments of hospitals, and whether or not a state Medicaid program pays for abortions. Data on AFDC came from abstracting of AFDC state plan summaries. Medicaid information came from four sources: the Health Care Financing Administration (Medicare-Medicaid Data Book, Various years), the Intergovernmental Health Policy Project (IHPP) (Recent and Proposed Changes in State Medicaid Programs, various years), special studies commissioned by HCFA (LaJolla Associates 1983), and the House Ways and Means Committee (Reports by the Institute for Medicaid Management 1976-1980).

The AFDC income eligibility standard is measured in nominal dollars. The Medicaid program characteristics are all measured as dichotomous indicators. The variable measuring limits is somewhat heterogeneous since it covers all types of limits to both physicians and outpatient departments. We experimented with disaggregation of this variable and found severe collinearity. The problem of heterogeneity of Medicaid limits is somewhat mitigated by the fact that the most common form of limit is to constrain visits to one per month. Table 1 provides brief definitions of all independent variables used in the analysis.

IV. Estimation

The model described above by equations (7-10) suggests a system of equations that is recursive with correlated errors. The presence of the variable e in all structural equations causes the error terms between equations to be correlated. Rosenzweig and Schultz (1982) show that the consequences of direct estimation of the structural production function is bias in the relationships of interest. They propose the use of instrumental variables and a two stage least squares estimator as a means of ensuring the inputs in the structural production function are orthogonal to the error term of the production function. Work by Joyce (1987) supports this view by showing that the hypothesis of zero correlation between health care inputs and neonatal mortality is always rejected.

An alternative estimation strategy, with panel data, is to directly estimate the structural production function and to allow each county to have a separate intercept. This fixed effects approach would take account of all fixed unobserved county influences on birthweight. The key assumption underlying this approach is that the unobserved variables are, in fact, fixed. Since one unobserved group of factors is the health endowment of women and the women in the data set change in each year (depending on child bearing) it is unreasonable to assume these factors are stable overtime. We, therefore, adopt an instrumental variables approach and rely on a two stage least squares estimator.

A second estimation issue relates to the form of our dependent variables. The dependent variables are percentages of LBW and VLBW births. One can view the measure as the probability that a birth is either LBW or VLBW, for which a linear probability model is used with micro-data. However, with grouped data the use of percentages suffers less from various difficulties than does the linear probability model (Maddala 1983). In this case, Maddala (1983) proposes the use of weighted least squares in order to obtain appropriate variance estimates. The weight for the LBW equations is given by:

$$(11)\ W = N^{.5}\ (p(1-p))^{-.5}$$

where N is the number of births in a county and p is the percentage that are LBW births. This weighting scheme serves to reduce the impact of random fluctuations by weighing counties with more births more heavily in the regressions.

A third estimation issue concerns the use of panel data. The combining of time series and cross sectional data for use with ordinary least squares (or two stage least squares) results in a regression disturbance term that may contain time series, cross section and random disturbance elements. The possibility of auto-correlation, and possible correlation of error terms with right hand side variables, suggest that some alternative to simple two-stage least squares or ordinary least squares is necessary.[4]

We investigated three alternative approaches to simple two stage least squares. These were: 1) variance components, 2) fixed effects, and 3) a hybrid of instrumental variables and fixed effects. The variance component estimates

[4] Only the second state of estimation with two stage least squares requires some adjustment, because it is only necessary to obtain consistent estimates of the reduced form parameters with two stage least squares. Efficient estimates are not necessary.

were rejected because they produced estimates that were quite sensitive to changes in model specification. The fixed effects model was problematic because of multicollinearity. The collinearity between the county dummy variables and the right hand side variables was sufficiently severe that it produced large standard errors and unstable coefficient estimates.

We chose to use a modified version of the instrumental variables estimator proposed by Hausman and Taylor (1981). For simplicity let us write the second stage estimator as:

$$(12)\ b_{it} = X_{it}B + a_i + u_t + e_{it}$$

where X is a vector of all the right hand side variables in equation (7) above, a is a time invariant unobserved error component, u is a time varying cross sectionally fixed error component and e is a random error. Our approach consists of two elements. First, we treat u as a fixed effect and allow for a separate intercept for each two year time period. Second we develop a set of instruments, Z, which are time invariant. Thus we rewrite (12) as :

$$(12')\ b_{it} = X_{it}B + Z_i d + u_t + e_{it}.$$

We choose elements of Z so that they measure factors we believe to be unobserved and related to birth outcomes and possibly correlated with the Xs. They include variables such as percentage of female headed households and the rate of urbanization. These variables serve to remove the systematic portions of the cross sectional error that covaries with the Xs. The remainder of the error is assumed to be random. This method produced rather stable estimates that were not sensitive to small changes in model specification.

V. Results

In this section we present the estimation results. We first focus on the quasi-structural production function results for the LBW and VLBW outcomes. We then present key estimates from the input demand functions for prenatal care, since the relationship between public policy and birthweight are posited to work through the use of prenatal care. In the interest of brevity we do not present the full set of input demand equation estimates.

The age and race specific quasi-structural production functions are presented on Tables 2 and 3. Each table presents separate estimates for LBW and VLBW outcomes. We begin our discussion by summarizing patterns in the results by race and outcome. We focus on race specific results because the differences across races are the most pronounced.

A. Results for LBW and VLBW Rates Among Blacks

Table 2 presents the age-specific quasi-structural LBW and VLBW production function estimates for black women. We discuss the set of results as a group. They represent a complicated set of relationships between the inputs in the production function and the two measures of birthweight.

The coefficient estimates for smoking indicate that for all models the volume of cigarettes consumed in a state is positively related to the rates of LBW and VLBW. The parameter estimates are significant for each model for the less than 20 year old and 20 to 34 year old age groups. The coefficient estimates for the 35 to 44 year age group are considerably less precise and generally do not attain conventional levels of significance.

The results for the coefficient estimates of the abortion variable are consistently positive. The significance levels vary with the specific outcome measure being considered. The coefficient estimates for the abortion variable are positive and significant (at the 0.05 level) in the VLBW production function for black teenage women and women 20 to 34 years old. They are significant in the LBW production function only for teenagers. The estimates for the older women (35 years and over) are small for the LBW and VLBW outcomes and are never significant at conventional levels.

The abortion results differ notably from those reported by Joyce (1987) and Corman, Joyce and Grossman (1987) and from our expectations. While our results are not directly comparable to those of Joyce (1987) and Corman, Joyce and Grossman (1987), both studies report results that show: (1) the county abortion rate is associated with a reduction in neonatal mortality among black women and (2) the abortion rate appears to reduce the rate of LBW births. We explored several alternate specifications in order to probe the stability of the findings. By and large the coefficient estimates were quite stable with respect to substantial changes in model specification.

In general the coefficient estimates suggest a weak negative relationship between our measure of prenatal care use and rates of LBW. The estimates are conistently negative across age groups. The prenatal care coefficient estimates for VLBW are all positive, although only the coefficient for women aged 20-34 years is significantly different from zero.

The family planning variable was specified as an interaction with poverty based on the hypothesis that availability of family planning clinics would have a larger impact on birth outcomes in areas of greater poverty.[5] Overall, the results for the teenage population and the over 35 years group indicate that the availability of family planning services is associated with significant reductions in the LBW rate. No significant impact was found for the 20 to 34 year old age group. Contrary to expectation

[5] The significance of the full effect is evaluated as:

$$t = (\beta FP + \beta FPXPOV)/\sqrt{VAR(\beta FP) + VAR(\beta POV) + 2\ COV(\beta FPA)}$$

the interaction effect tends to attenuate the impact of family planning availability.

The pattern of results by age for the impact of family planning availability on the VLBW rate are the reverse of those found for LBW. The impact of greater family planning availability is associated with a significant reduction in the VLBW rate only among 20 to 34 year old black women. The estimated impact for the two other age groups was positive and was not significantly different from zero. The findings are consistent with results for neonatal mortality reported by Corman, Joyce and Grossman (1987). We expected relatively larger impacts for teenage women, but the differential findings by age for the VLBW and LBW rates were unexpected.

The measure of parity, the percentage of first births, yielded parameter estimates that were negative for all age groups among black women. The estimates were significantly different from zero for teenagers and 20 to 34 year old women for the LBW outcome and for 20 to 34 year olds for the VLBW outcome. The estimates for the 35 to 44 year old group were very imprecise.

The estimates for the variable measuring the percentage of female headed households were directly and significantly related to both LBW and VLBW rates for all age groups. The magnitude of the estimated impact of this variable differed notably by age. For example the estimated coefficient in the LBW equation was four times larger for the 35 to 44 year old group than for the teenage group. It was four times larger for the oldest mothers compared to the 20 to 34 year olds. The percentage of female headed households was included in the model primarily as an proxy for the unobserved cross sectional error. Thus, while there is a substantive interpretation relating to the impact of female headed households and birth outcomes this set of coefficient estimates should also be interpreted broadly as an indicator of differences in county social and economic environments.

Higher rates of urbanization were also associated with higher rates of LBW and VLBW births. They are statistically significant in all specifications of the quasi-structural production function. The coefficient estimates are considerably larger for the two younger age groups than for the 35 to 44 year old group. Again, this variable was included as a proxy for the unobserved cross-sectional error.

The estimates of the impact of poverty on birthweight outcomes differed across age groups and by outcome for the black women. The race specific poverty rate was positively and significantly related to the LBW and VLBW rate among black teenage women. In contrast, the estimates for the 20 to 34 year old age group were negative and had a significant impact only for the VLBW rate. The results for the older mothers (35-44) show positive coefficient estimates for the LBW outcome measure and negative estimates for the VLBW outcome measure. The estimate in the LBW model was significant at conventional levels. The most obvious explanation for these results is that black teenage mothers suffer from the consequences of poverty to a much larger extent than do other age groups. Other explanations relate to the crudeness of the poverty rate as a measure of economic stress of women of child bearing age in a county.

B. Results for White LBW and VLBW Rates

The age-specific quasi-structural LBW and VLBW production functions are shown in Table 3. There are some important differences in these results as compared to the estimated coefficients for black women. Of particular note are the estimates for the effect of first trimester initiation of prenatal care.

The estimates for the smoking variable in the quasi-structural production functions for white women were quite similar to those estimated for blacks with one exception. For the 35 to 44 year old white women the estimates were generally positive although not statistically significant.

The results for the use of abortion services display a more complicated pattern of estimates than in the case of black women. The results for the two younger age groups are particularly sensitive to model specification (based on other model estimates not reported here). The coefficient estimates for the abortion variable in the VLBW models are consistently positive and significant for all age groups. The coefficient for the LBW model is positive for teenagers and negative for women aged 20 to 34 years but neither are significntly different from zero at conventional levels. The coefficient estimate for the women 35 years old and over was positive and significant for both outcomes. Once again these results run counter to both our expectations and previous research (Corman, Joyce and Grossman 1987). One implication of both sets of results is that abortions may be undertaken, largely for reasons other than concern over maternal or infant health. Indeed, women who seek abortions differ markedly on social and economic factors from women who continue their pregnancies (Hoffreth 1987).

The estimates for the impact of prenatal care in the birthweight production functions for white women display some similarities to the findings for black women. However, overall the estimated impact of prenatal care use among white women is much larger and is significantly negative for LBW. The findings also suggest that prenatal care has the strongest impact for teenage mothers. The coefficients for the prenatal care measure in the LBW model is significantly larger than the coefficient for the other two age groups.[6] The estimates in the VLBW models for teenagers and 20-34 year olds are both positive and significant. These findings are the reverse of the hypothesized effects.

The effects of family planning availability on the LBW and VLBW rates are negative and significant for all age groups except for the LBW rate among women aged 20 to 34 years. Thus counties where more family planning services are available have lower rates of LBW and VLBW birth, other things equal. These results are generally consistent with those reported in the analyses of black births.

The results for the parity measure among white women differ substantially across age groups and outcome measure. The estimates for women under twenty years of age suggest that counties with larger portions of first births among teenagers have lower rates of LBW births. The estimate exhibited a high degree of precision (0.01 level). The estimate for VLBW was essentially zero. In contrast the estimates for the 20 to 34 year old age group suggest a negative effect though not significant, of first births on the rate of LBW and a positive effect for VLBW. For the 35 to 44 year group the parity variable is positive and statistically significant for both measures of outcome. These findings are consistent with previous studies that show reduced rates of LBW for first births among teenagers and increased rates for older primiparous women.

6

$t = \beta_1 - \beta_2 / \sqrt{VAR_1 + VAR_2}$ assuming independence which may not be entirely reasonable.

The impact of the female headed household indicator is estimated to be positive for all specifications. The estimates are all very significant ($p<0.01$). The rate of urbanization is also estimated to be positively and significantly (at the 0.01 level) associated with both the LBW and VLBW rates. Moreover, the estimates are robust to changes in specification (based on results not reported here).

The poverty rate is generally estimated to have a significant negative impact on the LBW and VLBW rates. These results are generally significant. These are results that contrast with all our prior expectations. One explanation of this result may be related to greater availability of services to women in areas of high white poverty, such as WIC services, which were not measured in our study..

C. Results for the Prenatal Care Input Demand Model

Tables 4 and 5 report estimates for the input demand functions for prenatal care. Each table presents a set of race and age specific regression equations. We focus on the prenatal care equations since the initial goal of this research was to trace the impact of public financing policies on the use of prenatal care and the subsequent impact on rates of LBW births. The tables show a strong impact of several public financing mechanisms on the use of prenatal care.

The first major finding is that the AFDC Need (income eligibility) standard is positively related to measures of prenatal care use and is statistically significant in the prenatal care input demand function for all age and race groups except white women aged 35 or older. The elasticity estimates for the Need standard was largest for black teenagers with a value of 0.12. This figure means that a 10% increase in the Need standard would lead to a 1.2% increase in the percentage of women receiving prenatal care in the first trimester. The coefficient estimates are quite stable across age and race groups.

A second finding is that counties in states where Medicaid will pay for abortions tend to have higher rates of prenatal care use among black women than do otherwise similar counties. However, The estimates are generally not significant at conventional levels. Among white women aged less than 20 and 20 to 34 years, the coefficients are negative and significant suggesting prenatal care use is lower in states where Medicaid will pay for abortions.

A third set of results are negative results. The estimates for other dimensions of Medicaid benefit structure that we measure are generally either not significant or run counter to our expectations. Limits on reimburseable days and eligibility of first time pregnant women were negatively related to use of prenatal care for older black women while limits were positively related to use for white teenagers. Otherwise they had no impact on prenatal care.

Two of the most important policy determinants of prenatal care use were the indicators of the availability of local publicly funded prenatal care programs. Both the availability of routine prenatal care and the availability of comprehensive prenatal care programs were positively related to prenatal care use (the one exception is routine care for black women 35 to 44 years). The coefficient estimates for comprehensive care are significant at least the 0.05 level for all ages of black women and for white women under 35 years of age. The magnitude of these effects is often quite large. For example, the coefficient estimate for comprehensive prenatal care for black women 35 years or older implies that 3.6 percentage points more women initiate care in the first trimester in counties with such services than in otherwise similar counties.

The impact of the availability of routine care on use of prenatal care is somewhat smaller and not significant for black women. Its impact is quite large for white women under 35 years. There is no significant impact of the OB-GYN density on the early initiation of care among black women. For teenage white women the availability of OB-GYNs is positive and significant ($p<0.10$) in the model of prenatal care use. The elasticities implied by the coefficient estimates are generally quite small (less than 0.01). The availability of physicians that are not OB-GYNs is generally estimated to lead to greater use of prenatal care services for black women under 35 years. The estimate for the 35 to 44 year age group is essentially zero. The coefficient estimate for white women is negative in all three age categories but it is significant only for white teenagers. This finding runs counter to our expectations.

The estimated impact of hospital availability on use of prenatal care is dramatically different across racial groups. The number of beds per capita is positively and significantly related to use of prenatal care among blacks (except older women) and negative for whites. This finding is likely to reflect racial differences in where prenatal care is received. The availability of abortion providers and family planning programs are estimated to have a different impact on use of prenatal care across age groups. There is a strong positive effect of family planning clinics on use of prenatal care for both races and all age groups except white women over 35 years. Negative estimates were obtained for the impact of abortion availability on prenatal care use for the two younger age groups of both races.

The level of educational attainment in a county had a positive and significant association with early initiation of prenatal care for all race and age groups except black teenagers. Most impact estimates were rather small especially for white women. Poverty levels were generally weakly related to use of prenatal care. Our estimates indicate that higher levels of poverty are most often negatively associated with use of prenatal care, particularly for whites. The relationship between poverty and prenatal care use for blacks is far less clear; among black teenagers, the poverty rate is positively related to early use of care.

VI. Discussion of Findings

There are several important findings of our investigation of the impact of resources and financing policies on use of prenatal care and LBW and VLBW rates. In this section we direct our attention towards findings that have implications for policy regarding the financing of prenatal care and other maternal and child health services. We will weave together implications of the research with policy implications and comparisons to other studies.

Our empirical models show that prenatal care has a significant effect on LBW for whites and little impact for blacks. However, for the 20 to 34 year old age women, the coefficient for prenatal care is larger for black women than for white women. A second important pattern to note is that the largest impact of prenatal care on rates of LBW births is for white teenagers.

The racial differences in the impact of prenatal care on LBW are somewhat surprising. We would have expected the LBW rate to be more sensitive to increases in prenatal care for black women. While we do not have a rigorous explanation for the racial differences, data from the 1982 National Survey of Family Growth provide a clue. Roughly 50% of black women in the survey received prenatal care from a clinic, while 40% obtained care from a private physician or group practice. In contrast, 30% of white women received care in a clinic and 63% from a private physician or group. These racial differences in the context in which prenatal care is obtained suggest that it may be important to learn more about the process of care within different organizational settings. Little appears to be known about the process nor the content of prenatal care (OTA 1988). The development of such knowledge will expand our ability to interpret results such as those reported above.

The differences in prenatal care impact by age conform with our expectations. Because teenage women who become pregnant often face difficult circumstances such as low income, single parenthood and family disruption, receipt of medical attention might be particularly beneficial. In general our findings suggest that increasing access to prenatal care for both races will lead to improvements in the LBW rate. The magnitude of the reductions varies by race. For example, 10 percentage point increase in early initiation among black teenager would lead to a 0.1 percentage point decline in the LBW rate (from 14% to 13.9%). Among white teenage women the same change in first trimester by 10 percentage points (from 60% to 70%) is estimated to reduce the LBW rate from 7.5% to 5.5%. An increase of 10 percentage points in the first trimester care would result in a 1.7 percentage point decline in the LBW rate (from 12.08% to 10.38%) among black women aged 20 to 34 years and a 0.8 percentage point decline for white women of the same age.

One unexpected finding which warrants comment is the positive and significant impacts of early initiation of care on VLBW rates for white teenagers and for black and white women aged 20 to 34 years. These findings may be related to the mechanism that has been hypothesized to generate the effects of prenatal care on LBW. It is often presumed that prenatal care will effect LBW rates by increasing the weight of newborns or by prolonging gestation. It is entirely possible that it will have this effect at the lower end of the birthweight continuum, as well. (That is, for infants weighing less that 1500 grams.) Prenatal care may also have an effect on the probability of survival of light newborns. Thus counties with higher rates of early prenatal care initiation may have high VLBW rates because the rate of survival of very small infants may be higher. Since change in a relatively small number of births can produce large effects on the VBLW rate the positive impact of prenatal care seems plausible.

Given the results for prenatal care a critical policy question to answer concerns what types of policy initiatives might lead to increases in use of prenatal care, particular among poor women. The relevant findings for addressing this question are the estimates from the input demand functions. Since these findings differ by age and race group we cannot offer a single set of answers. We focus our attention on teenagers and black women since they are the group with the lowest rates of use of prenatal care. Our discussion addresses two main policy issues: (1) expansion of the Medicaid income eligibility standard, and (2) expansion of publicly funded MCH programs.

One policy option, the expansion of Medicaid eligibility for pregnant women to 185% of the poverty line, is now available to all states. Adopting this option would entail increases in the effective need standard by $800 to $1300 per month in most states. According to our estimates increasing the AFDC need standard by $1250 in counties meeting our black population criterion, will result in a 0.3 percentage point decline in LBW rates for black teenagers (Table 6). Adoption of the OBRA-87 income eligibility standard would have a substantial impact on the LBW rate among white teenagers. The AFDC changes would increase early initiation by 6 percentage points, which in turn would lead to a reduction of 3.4 percentage points. Thus the LBW rate would fall from 8.68% to 5.28%. The effects for the older white women are much more modest.

The effect of expanding availability of local government funded prenatal care programs has a significant but modest impact on the LBW rate. Table 6 shows that the impact of the MCH programs are greatest for white teenagers and black women aged 20 to 34 years. Expansion of MCH program would decrease the LBW rate among white teenagers by 0.47 percentage points, a 5.5% reduction in the rate. The LBW rate among 20 to 34 year old black women would fall by about 0.23 percentage points in response to expanded MCH programs, or a 3.1% reduction in the LBW rate.

There are several limitations of our research that we wish to comment on here. Perhaps the most important constraint on our research was our inability to incorporate measures of preterm births into the models for birth weight. We faced two important problems which require further investigation. The first was that we were reluctant to impose the identifying restrictions on our model that have often been used in prior work. The foundation for these restrictions, we believe, is not strong. A more solid basis for identification of the structural parameters of a model which includes both measures of LBW and preterm births as endogenous variables is required. The second problem was one of data availability. More complete reporting and verification of gestational age information from birth records is necessary.

A second important constraint relates to our inability to measure the prices paid by various buyers of prenatal care. Economic factors within the Medicaid program may constrain access to prenatal care by setting prices that make providers reluctant to provide prenatal care to Medicaid patients. Some evidence to support this notion has been reported in the literature (IOM 1988). These issues require systematic treatment in the context of a complete model of birth outcomes.

Our results suggest that the structure of state MCH programs are key variables related to access to prenatal care. The responses to our survey of these programs also indicate considerable variation in the manner in which services are organized and financed. For example some states and counties provide services in public health clinics, others contract with non-profit agencies while still other contract with for-profit organizations. There are also some indications in the survey data that contracting arrangements might vary. Developing an understanding on how the organization of care and the nature of contracting effects the delivery of services we believe is an important topic for further research.

Our study relies on aggregate county data. While there is much policy relevant information to be gleaned from these data, analysis of micro data on individuals would clearly advance our understanding of behavioral responses to economic factors. Studies to date using mico- data have either been limited with respect to the variation in economic conditions (Harris 1982) or have relied on non-representative samples (Rosenzweig and Schultz 1982). National data on individual women who give birth that includes detailed information on their health related behavior and economic situations would contribute greatly to improving our knowledge in this area.

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TABLE 1

Variable Descriptions

Variable	Definition
Smoking state	Number of cigarettes smoked in a per capita
Abortion Rate	Estimated number of abortions by residents of a county divided by female population age 15-44
1st Trimester	Percentage of women in an age and race specific category initiating prenatal care in the first trimester
Parity	The percentage of births in an age and race specific category that were first births
Fem HH80	Percentage of households in a county in 1980 that were female headed
% Urban	The percent of county's population in 1980 that resides in an urban area
POVERTY	The rare specific percentage of households falling below the federal poverty standard in 1980.
Mcaid Abort	State Medicaid Plan Covers Abortion Services = 1, 0 Otherwise
Need	The four person family AFDC Need Standard Measured in dollars per month
Limits	The State Medicaid program limits. Ambulatory Vistis = 1, 0 Otherwise
1st Preg	Firt time pregnant women are eligible for Medicaid = 1, 0 Otherwise

TABLE 1 (Cont'd.)

Variable Descriptions

Variable	Definition
Compreh	County has an MCH program that offers comprehensive prenatal care = 1, 0 Otherwise
Routine	County has an MCH program that offers routine prenatal care = 1, 0 Otherwise
OBS/POP	The number of obstetricians/gynecologists in patient care divided by female population age 15-44.
MD/POP	Number of Physicians that are not Ob-Gyn per capita
BED/POP	Number of acute care general hospital beds per capita
FP Clinic	Number of Family Planning Clinics divided by female population age 15-44.
Lg Abort	Number of abortion providers performing over 400 abortions per year divided by female population age 15-44.
Sm Abort	Number of abortion providers performing fewer than 400 abortions per year divided by female population age 15-44.
Schooling	Median educational attainment of adults in county.

TABLE 2

Quasi-Structural Production Function Estimates

Black Women by Age*

Variable	LBW <20	LBW 20-34	LBW 35+	VLBW <20	VLBW 20-34	VLBW 35+
Smoking[a]	$0.13e^{-4}$	$0.11e^{-4}$	$0.51e^{-7}$	$0.10e^{-4}$	$0.12e^{-4}$	$0.27e^{-6}$
	(2.45)	(3.16)	(0.05)	(3.89)	(6.91)	(0.16)
Abortion Rate[a]	39.61	5.90	5.07	5.73	4.07	0.47
	(6.11)	(1.29)	(0.69)	(4.29)	(1.98)	(0.26)
1st Trimester[a]	-0.01	-0.17	-0.08	0.008	0.02	0.009
	(1.35)	(0.88)	(0.87)	(0.73)	(3.33)	(1.21)
Parity	-0.03	-0.05	-0.02	-0.005	-0.01	-0.002
	(1.71)	(4.35)	(0.97)	(0.91)	(2.14)	(0.53)
Fem HH 80[b]	0.12	0.10	0.50	0.08	0.05	0.03
	(4.12)	(5.13)	(8.58)	(6.11)	(5.35)	(3.48)
% Urban[b]	0.01	0.014	0.07	0.02	0.01	0.01
	(2.99)	(4.35)	(9.76)	(13.43)	(7.59)	(3.30)
FP Clinics	-2.79	-0.014	-1.36	0.18	-1.19	0.05
	(2.76)	(0.03)	(3.61)	(0.45)	(4.41)	(0.63)
POVxFP	0.05	0.007	0.001	-0.008	-0.02	0.00
	(2.25)	(0.51)	(0.21)	(0.96)	(3.56)	(0.00)
POVERTY[b]	0.03	0.01	0.034	0.02	-0.01	-0.001
	(2.18)	(0.96)	(1.73)	(2.80)	(2.83)	(0.47)
Intercept[c]	11.68	8.59	10.78	-2.62	-2.17	-0.92
	(6.36)	(6.32)	(5.23)	(3.89)	(3.51)	(1.90)
F	10.60	10.07	26.35	29.93	47.48	4.67

*t statistic in parentheses

[a] endogenous variable

[b] time invariant

[c] time dummies included

TABLE 3

Quasi-Structural Production Function Estimates

White Women by Age*

Variable	LBW <20	LBW 20-34	LBW 35+	VLBW <20	VLBW 20-34	VLBW 35+
Smoking[a]	$0.24e^{-4}$	$0.16e^{-4}$	$0.63e^{-4}$	$0.13e^{-5}$	$0.27e^{-6}$	$0.12e^{-6}$
	(6.67)	(12.67)	(1.58)	(1.84)	(0.66)	(0.20)
Abortion Rate[a]	-5.92	1.08	28.5	3.03	1.15	2.07
	(1.16)	(0.77)	(7.01)	(2.51)	(2.29)	(3.18)
1st Trimester[a]	-0.20	-0.08	-0.08	0.02	0.008	-0.01
	(10.18)	(7.30)	(2.99)	(3.56)	(2.51)	(1.48)
Parity	-0.09	-0.002	0.03	0.002	0.004	0.004
	(8.06)	(0.50)	(4.22)	(0.82)	(2.76)	(2.83)
Fem HH 80[b]	0.14	0.09	0.14	0.02	0.02	0.01
	(6.78)	(12.83)	(5.20)	(5.80)	(8.27)	(3.29)
% Urban[b]	0.02	0.006	0.04	0.006	0.003	0.003
	(7.82)	(7.52)	(16.48)	(12.41)	(10.62)	(7.49)
FP Clinics	-1.61	0.07	-3.53	-0.62	-0.41	-0.25
	(3.71)	(0.47)	(9.82)	(7.33)	(9.13)	(4.03)
POVxFP	0.12	0.02	0.17	0.03	-0.01	-0.01
	(4.46)	(2.45)	(7.27)	(5.84)	(5.00)	(3.25)
POVERTY[b]	-0.12	-0.04	-0.12	-0.02	-0.02	-0.01
	(7.03)	(4.69)	(5.24)	(5.79)	(5.41)	(2.30)
Intercept[c]	21.62	9.01	4.68	1.28	0.97	0.41
	(15.67)	(10.41)	(2.90)	(4.55)	(3.75)	(1.48)
F	97.27	64.56	71.77	61.20	100.24	16.75

*t statistic in parentheses

[a] endogenous variable

[b] time invariant

[c] time dummies included

TABLE 4

Selected Estimates From Prenatal Care Demand Equations

Black Women by Age[1]

Variable	<20	20-34	35+
Mcaid Abort*	0.62	0.004	1.76
	(1.52)	(0.20)	(1.15)
Need*	0.027	0.02	0.009
	(2.06)	(2.89)	(1.80)
Limits*	0.36	-0.35	-7.29
	(0.34)	(0.34)	(1.99)
1st Preg*	-0.22	-0.72	-9.09
	(0.22)	(0.45)	(2.65)
Compreh.	1.88	1.65	3.57
	(3.10)	(3.27)	(3.36)
Routine	1.18	0.43	-1.49
	(1.64)	(0.73)	(1.23)
OBS/POP	-0.42	2.64	0.26
	(0.16)	(1.21)	(0.06)
MD/POP	1.11	0.99	-0.06
	(2.23)	(2.49)	(0.08)
BED/POP	0.47	0.28	0.04
	(3.41)	(2.43)	(0.15)
FP Clinic	3.20	3.20	5.69
	(3.50)	(4.14)	(5.92)
Lg Abort	-5.49	-1.79	22.88
	(0.43)	(0.46)	(1.01)
Sm Abort	-2.72	-8.71	6.57
	(0.58)	(0.86)	(0.94)
Schooling	-0.04	1.19	3.26
	(0.09)	(2.72)	(3.82)
POVERTY	0.19	-0.01	-0.073
	(1.72)	(0.12)	(0.42)
R^2	0.11	0.19	0.18
F	9.07	18.60	16.01

* Includes poverty interaction effect

[1] t statistic in parentheses

TABLE 5

Selected Estimates From Prenatal Care Demand Equations

White Women by Age[1]

Variable	<20	20-34	35+
Mcaid Abort*	-2.37	-1.874	0.03
	(3.33)	(2.22)	(0.03)
Need*	0.02	0.007	0.003
	(4.10)	(2.33)	(0.69)
Limits*	1.03	0.52	0.51
	(1.88)	(0.62)	(0.53)
1st Preg*	-0.81	-0.75	0.40
	(1.15)	(1.33)	(0.44)
Compreh.	0.50	1.69	0.72
	(1.56)	(5.01)	(1.71)
Routine	2.36	2.30	0.61
	(6.15)	(5.81)	(1.24)
OBS/POP	1.80	-0.95	1.86
	(1.34)	(0.71)	(1.14)
MD/POP	-0.75	-0.82	-0.25
	(3.11)	(0.99)	(0.78)
BED/POP	-0.01	-0.10	-0.20
	(3.07)	(1.85)	(2.19)
FP Clinic	4.62	5.28	0.31
	(3.77)	(4.72)	(0.23)
Lg Abort	-16.36	-14.42	3.79
	(3.30)	(3.19)	(0.75)
Sm Abort	-6.48	-2.94	3.61
	(4.43)	(1.83)	(1.99)
Schooling	0.60	1.07	0.89
	(2.53)	(4.94)	(3.20)
POVERTY	-0.07	0.17	-0.83
	(1.58)	(1.40)	(5.56)
R^2	0.11	0.09	0.20
F	32.54	25.76	67.18

* Includes poverty interaction effect

[1] t statistic in parentheses

TABLE 6

Impacts of AFDC Changes and MCH Programs

on LBW Rates [*, 1, 2]

Age Group	Prenatal Care Measure	
	AFDC	MCH
<20		
B	-0.27	-0.02
W	-3.4	-0.47
20-34		
B	-3.4[3]	-0.23
W	-0.48	-0.18
35-44		
B	0	0
	-0.20	-0.07

[*]Changes in AFDC Need => 185% of poverty.

[1]These changes are relevant to (a) states that have a need standard of 62% of poverty or less and (b) counties without either comprehensive or routine PNC progrms. It is worth pointing out that 67% of states have need standards below 50% of poverty and 95% of states have need standards below 75% of poverty.

[2]These analyses used 1980 Poverty levels.

[3]Not significantly different from zero.