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Self-selection in decision-making of pregnancy resolutions, demand for prenatal care and birth outcome production function

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SELF-SELECTION IN DECISION-MAKING OF PREGNANCY RESOLUTIONS DEMAND FOR PRENATAL CARE AND BIRTH OUTCOME PRODUCTION FUNCTION

by

Guoen Liu

A Dissertation submitted to the Graduate Faculty in Economics in partial fulfillment of the requirements for the degree of Doctor of Philosophy, The City University of New York.

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Abstract

SELF-SELECTION IN DECISION-MAKING OF PREGNANCY RESOLUTIONS DEMAND FOR PRENATAL CARE AND BIRTH OUTCOME PRODUCTION FUNCTION

by

Guoen Liu

Adviser: Professor Michael Grossman

This dissertation investigates women's self-selection behavior in decision-making for pregnancy resolutions, the self-selection effects on demand for prenatal care and production of birth outcomes. Three major hypotheses are tested: A. Women make sequential decisions concerning their pregnancies: the first decision is to select between giving birth or obtaining abortion; the second decision is to determine an optimal demand for health inputs, such as prenatal care, which in turn determine the infant health through a health production function; B. The nature of the sequential decision-making suggests that the second decision is a sub-outcome of the first decision. Thus, women who select to give birth, as a censored sample, may not be a random sample drawn from the entire population of pregnant women due to the self-selection effect in the first stage; C. The self-selection effect may also exist with the heterogeneity of health endowment and the endogeneity of prenatal care in a birth outcome production function. Since prenatal care demand and birth outcome production only pertain to those who selected into the sample of giving birth, failure to take into account these two self-selection effects would considerably bias the estimates of behavioral demand for health inputs and the infant health technology.

First, to rationalize the framework of modeling for the decision-making in pregnancy resolutions, a random utility differential function is introduced to govern the mechanism of self-selection: giving birth or obtaining an abortion. Probit maximum likelihood estimates are presented for a probability function of giving birth. Second, the information of sample selection

in pregnancy resolutions is then incorporated into a model of demand for prenatal care and birth outcome production in which self-selection effects are controlled by the methods of Heckman (1979), Lee et al. (1980), and Rosenzweig and Schultz (1983). Third, an empirical investigation is conducted for both the model controlling for self-selection effects as well as those without the correction of self-selection effects. The empirical study is based on the 1984 vital statistics for all induced abortions and live births in the Commonwealth of Virginia. Adolescents are excluded from the data base to minimize the problem of endogeneity of health inputs. The estimation is made respectively for four cohorts: urban whites (13,863), rural whites (6,196), urban blacks (13,528), and rural blacks (7,291).

The major findings of this dissertation are as follows. A. Women's behavioral self-selection in pregnancy resolutions introduces an upward bias to the true marginal product of prenatal care in the birth outcome production function; a significant self-selection bias is found in the demand for prenatal care among urban whites and rural blacks. B. Women's adverse self-selection behavior in input use appears to dominate the favorable selection effect. Therefore, it is observed that a downward bias would have occurred with the marginal product of prenatal care if no action was taken to control the endogeneity of prenatal care and heterogeneity of health endowment. C. Unobservables in the decision model of pregnancy resolutions are inversely correlated with those in the demand for prenatal care and birth outcome production functions. This evidence suggests that the shadow price of abortion tends to be dominant over the other unobserved factors involved in the model system.

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CHAPTER ONE

INTRODUCTION

Following the theoretical framework of a household production model pioneered by Becker (1965), there has appeared extensive theoretical and empirical investigations concerning the economics of fertility, derived demand for health or health related inputs, and health production function (Grossman 1972; Becker 1973; Willis 1973; Michael 1973; Gortmaker 1979a, 1979b; Harris 1982; Rosenzweig and Schultz 1983; Easterlin and Crimmins 1985; Montgomery 1987; Leibowitz, Eison, and Chow 1986; Powell-Griner and Trent (1987); Corman, Joyce, and Grossman 1987; Serrato 1989; Joyce and Grossman 1990; Heckman and Wlaker 1990; Frank, Strobino, Salkever, and Jackson 1991; Grossman and Joyce 1991).

In the context of the economics of fertility, researchers focus primarily on the households' long run decisions such as optimal total number of children, timing and spacing of children, dynamic changes in fertility or fertility control (Willis 1973; Easterlin and Crimmins 1985; Heckman and Walker 1990). However, a household's short run decision-making, concerning a marginal fertility, arises when an unexpected outcome occurs. A woman, for example, due to failure of contraception, may have an unplanned pregnancy which is obviously out of her long term optimal fertility plan. That is, given to be pregnant, a woman is facing a decision in resolving her pregnancy -- carrying it to term or obtaining an abortion. How do women differ in the probability of having an unexpected pregnancies? What is the decision rule for the resolution of such unwanted pregnancies? What are the basic arguments involved in the rule?

Easterlin and Crimmins (1985) demonstrate that household fertility is jointly determined by three forces: the supply of and the demand for children, and the cost of fertility control. Moreover, modernization also changes these three forces in such a way that supply tends to rise, demand declines, and cost of fertility control falls. Availability of new contraceptive techniques or more efficient induced abortion, for example, provides households with more choices regarding deliberate fertility control. This argument strengthens the notion that, for a given pregnancy, choosing to give birth becomes highly optional or self-selective. Moore and Caldwell (1976), Leibowitz et al. (1986), and Serrato (1989) present an economic model of teenager's premarital pregnancy resolutions: abortion, birth and marriage, and out-of-wedlock birth. To test the influences of public welfare such as Aid to Families with Dependent Children (AFDC) payments, Medicaid, and sex education on the decision-making, Serrato (1989) sets up a framework with three utility functions for three alternatives faced by decision maker. He adopts a nested logit model (NL) for the estimation of his model. Grossman and Joyce (1989, 1990) employ a concept of optimal number of children, derived from Becker's quantity-quality model for children, and compare it with the actual number of children at present to govern the pregnant women's selection decision in pregnancy resolutions.

Since only those women who decide to give birth would have to make the second decision, i.e., their demand for health inputs, which in turn would determine the health of the fetus through a health production function. A more important issue then arises -- whether the women's self-selection behavior in pregnancy resolutions is independent from the behavior with decision-making for the demand for health inputs and consequently the health production function.

Literally, a health production function describes a technological or biological relationship between health inputs and its outcome. Yet the level of health inputs that are employed to produce an expected or desired health outcome depends on the household's preference orderings about child health in the parental utility function. Following Becker (1965, 1973) and Grossman (1972), households maximize their utility functions subject to constraints of income, time, and health technology. Thus, a household's demand for health related goods is derived from the demand for health which directly augments the household's utility. Apparently, the derived demand for health inputs are behavioral variables which, in turn, make the health production also be characterized by the household's tastes. If an unobserved factor affects household behavior and therefore involves determining both demand for inputs and health production, then the estimates of the biological effects of the inputs on health, based on the observed information for variables controlled in the model, would be biased from their true counterparts. Health endowment, as an example of unobservables, appears to affect women's behavior adversely in their demand for prenatal care, but favorably in health production, assuming that the information is known to the individuals. That is, women with poor health endowments would be more likely to seek more or earlier prenatal care but end up a lighter birth weight than women with better health endowments. Rosenzweig and Schultz (1983) show that the household's adverse selection behavior in health inputs due to the heterogeneity of health endowment may bring misleading observed information for the researchers estimating the health technology. They present a health production function controlling for unobserved factors such as health endowment by using two stage least square estimation (2SLS). Their findings provide a good explanation of why epidemiological correlation studies have not always found a beneficial effect of the timing of prenatal care on birth weight (Eisner et al. 1979). Similar arguments are made by Harris (1982). Recent works by Corman, Joyce, and Grossman (1987), Joyce (1987), and Frank et al. (1991) have emphasized the structural estimates of health technology controlling for the effect of adverse selection in health inputs. Possibly, on the other hand, a favorable selection may also characterize women's attitude toward the use of health inputs (Gortmaker 1979; Institute of Medicine 1985). The efficacy of medical care would be overestimated if women who initiate more medical care may also simultaneously consume more of other healthy goods or have better living environment for the growth of the fetus both financially and psychologically. For example, a woman seeking more prenatal care may also eat more nutritious goods, suffer less stress, smoke less, engage in more appropriate exercises, and receive less potentially harmful substances than women who begin initiating prenatal care later. The rationale behind the favorable selection is that the initiation of prenatal care may in part act as an indicator of "wantedness" of the pregnancy (Klerman and Jekel 1984; Cooney 1985; Brown 1989).

All of these previous studies attempt to address the issue of self-selection in the demand for health inputs and health technology, based on pregnant women who selected to carry their fetus to term. This is because the demand for health inputs and the health production function can only be observed for the women giving birth while the others who obtain an abortion yield no information for the health inputs (e.g. prenatal care) and health outcome(e.g. birth weight). Now the crucial question is whether the women who obtained abortion, had they chosen to carry their fetus to term, would have behaved in the same manner as those who actually gave birth. Statistically, the issue is whether the women who give birth are a sub-sample selected from the entire population of pregnant women in a random manner. Put differently, the observed data set is a censored sample in that actual health inputs and health outcome are used for women giving birth, while a zero value is set to the same variables for women obtaining abortion. This topic of sample selection has been the subject of an enormous recent literature, both theoretical and applied. Research in statistics and applied econometrics show that if a sample is drawn randomly, the estimates given by the usual statistical techniques would be consistent except for the loss of efficiency. However, if the sample selection is not random, the results would lose not only the efficiency but also the consistency which is one of the most important properties for a statistical estimator (Heckman 1976, 1979: Griliches, Hall, and Hausman 1978; Maddala 1985).

Following Heckman's work (1976), the problem of sample selection has received widespread attention (Willis and Rosen 1979; Wales and Woodland 1980; Nakosteen and Zimmer

1980; Nelson 1984; Maddala 1983,1984; Killingsworth 1983; Dhrymes 1984; Amemiya 1984). Yet in the field of demand for prenatal care and birth weight production, although the notion of sample selection appears particularly applicable, little has been done until Grossman and Joyce (1988, 1990) formally considered the impacts of women's self-selection of pregnancy resolutions on their behavioral demand for prenatal care and the health productions for New York City in 1984. They argue that the resolution of pregnancy itself is characterized by self-selection, which in turn shapes the behavior of women who give birth and the women who obtain abortion. Consequently the observed health technology and birth outcomes would be seriously biased if the effect of sample selection is disregarded. They apply Heckman's two-step procedure and present the first health production functions that simultaneously control for self-selection in the resolution of pregnancy as live births or induced abortions and in the use of prenatal care services. Their infant health technology and demand for prenatal care functions are refined by incorporating the information conveyed in the women's behavior in pregnancy resolutions.

As an elaboration of the Grossman-Joyce approach, this study conducts a further theoretical rationalization of the modeling and empirical investigation of women's self-selection behavior in the context of decision-making in pregnancy resolution, self-selection of prenatal care demand and infant health production functions. First, it refines the framework of modeling for decision-making in pregnancy resolutions by adopting the random utility function approach. As a decision maker, a pregnant woman faces two choices: giving birth or obtaining abortion. The well-defined utility maximization approach of consumer theory is not applicable because the outcomes are not in a continuous space of alternatives but are "corner" solutions, the points where the usual first-order conditions for an optimum do not hold. To formulate a decision mechanism dealing with corner solutions, a random utility model, originally developed by McFadden (1973, 1974), Manski (1977), Nakosteen and Zimmer (1980), Ben-Akiva and Lerman

(1989), is introduced in the context of pregnancy resolutions. Manski (1973) proposes four distinct sources of randomness: unobserved attributes, unobserved taste variables, measurement errors and imperfect information in random utility functions. The random utility model can fully capture the unobservables involved in this study such as the cost of contraception, the cost of abortion, the health endowment, and the household's preferences. A utility differential function of giving birth vs. abortion is proposed. This enables a derivation of a probit probability function for women making a selection of giving birth or obtaining abortion. Second, the information of self-selection generated by the decision model of giving birth is then incorporated into the model of prenatal care demand and the health production function. Moreover, women's self-selection in input use, due to the endogeneity of prenatal care and heterogeneity of health endowment in health production, is also considered (Rosenzweig and Schultz 1983, 1988). Third, the husband's educational level is introduced into the demand function for prenatal care. It is hypothesized that the demand for prenatal care is an outcome of joint decisions made by both wife and husband. Moreover, since some women who give birth are not married or without husbands, their behavior of demand for prenatal care may differ from those having husbands, the difference can be, in part, captured by including a dummy variable of illegitimacy in the model. Finally, an empirical investigation of the model controlling for self-selection effects and also those without the correction of self-selection is conducted respectively for four adults cohorts: Urban Whites, Rural Whites, Urban Blacks, and Rural Blacks, using the 1984 micro vital records on all induced abortions and live births in the Commonwealth of Virginia. Adolescents are excluded to minimize the problem of endogeneity. Study on the adolescent cohort will be conducted in a separate study. The empirical performance of the model corrected for self-selection bias is also compared with the outcomes yielded by other approaches (OLS, 2SLS). In addition, a comparison of the findings for the present study and those achieved by others is made. The estimation methodology is based on a probit two-stage approach (Heckman 1979; Lee, Maddala, and Trost 1980; Griliches, Hall, and Hausman 1978; Greene 1990). Wu-Hausman (Wu 1973 and Hausman 1978) test is conducted to identify the endogeneity of prenatal care in health production functions.

This dissertation is organized in seven chapters. Chapter one presents the introduction of this study. Chapter two reviews the extensive body of the literature on economics of fertility, decision model for pregnancy resolutions, demand function for prenatal care, and health production function. Comments on previous researcher's methodologies and conclusions are included. Chapter three introduces a random utility function for the birth-abortion binary choices and proposes a utility differential function of giving birth vs. abortion as a self-selection mechanism through which a probit probability function of giving birth is derived. Chapter four first explores the issue of self-selection in derived demand for prenatal care and infant health production functions. Then, a model corrected for the effects of self-selection is developed. Chapter five describes the data base, model specification and computational framework. Chapter six presents the empirical report and discussion of the findings in different models, emphasizing the one controlling for self-selection. Chapter seven ends the dissertation with a brief summary and concluding remarks.

CHAPTER TWO

REVIEW OF THE LITERATURE

2.1 Economics of Fertility and Decision-Making in Pregnancy Resolutions

Households' optimal number of children, spacing and timing of children in life-time horizon, and the dynamics in fertility as well as fertility control have been fruitfully studied in the context of fertility economics (Willis 1973; Michael 1973; Goldman and Grossman 1986; Leibowitz et al. 1986; Easterlin and Crimmins 1985; Montgomery 1987; Powell-Griner and Trent 1987; Serrato 1989; Heckman and Walker 1990; and Grossman and Joyce 1991). The household production model, introduced by Becker (1965), almost always serves as the fundamental framework underlying most of these analyses.

In an attempt to examine the most widely and frequently observed influence from education to fertility, Michael (1973) formulates an economic model of human fertility. This model focuses on the mechanisms through which education may affect a couple's fertility decision and particularly documents the effects of education on the contraceptive use with respect to the timing and the spacing of children.

Two of the commodities augmenting a household's utility function are made to be associated with fertility: "family life," a function of the couple's offspring, and "sexual gratification," analytically related to the number of children. In order to address the economic forces of fertility plan, a probability of conception is introduced, as a function of the couple's sexual gratification, fecundity, contraception use, and time input. Households are assumed to alter the probability of conception by varying the resources of time and money in the use of contraceptive techniques, given their socioeconomic and demographic characteristics. Michael argues that a household's decision regarding fertility control may be made independently in each period on the basis of the static marginal benefit and marginal cost from an additional child. A household would seek to raise the probability of contraception by not using the contraception, if the net benefits, being as the difference between the marginal benefits and marginal costs of an additional child, turn out to be positive. If, on the other hand, the net benefits are negative, then the couple may engage in contraceptive use. Furthermore, the net benefit variable is shown to be inversely correlated with the discrepancy between the completed fertility demand (N^*) and the actual fertility (N).

Michael outlines four channels through which the households' educational level may influence the determination of fertility: education and subjective preferences concerning the good of child; education and opportunity cost of child; education and one's productivity; and education and the life expectancy (Grossman 1972). Based on the 1965 National Fertility Study with 5,600 U.S. non-Catholic women aged 25 and 40, he observes a negative influence of education, through the use of contraceptives, on households' birth probabilities consistently across all cohorts by age, color, and parity. Michael attributes the negative role of education on fertility to two distinct forces: the variations in effectiveness of different contraceptive techniques or the differential effectiveness of a given technique at different educational levels.

The households' behavior with respect to fertility and fertility control also presents a dynamic pattern over time. One of the most notable dynamic feature of fertility for population throughout the would is a declining trend in fertility and a spread of deliberate fertility control over time. The causes and determinants of such changes in the people's reproductive behavior are well addressed in an enormously influential work, <u>The Fertility Revolution - A Supply and Demand Analysis</u> (Easterlin and Crimmins 1984).

Easterlin and Crimmins introduce a household utility model characterized by the forces of supply of and demand for children, and the cost of fertility regulation via which modernization plays its role in influencing people's behavior of fertility and fertility control. Following Becker (1960, 1965) and Willis (1974), Easterlin and Crimmins also treat children as a normal good in a household utility function. In a derived household demand function for children, the number of children demanded is expected to vary directly with income, prices of other goods relative to children, and inversely with the strength of tastes for other goods relative to children and the economic as well as non-economic costs of children. Moreover, child quality is also controlled in their model as an additional good to the quantity of children and other goods consumed by parents. This allows them to test how and why households with different status would differ in the demand for number and quality of children.

The supply of children is defined as the product of natural fertility without deliberate regulation and probability of child survival. Easterlin and Crimmins also make a distinction between the biological reproductive potential and the natural fertility. The first refers to the maximum number of children in the most favorable conditions whereas the latter shows the number of children in actual living conditions with no conscious effort made to regulate fertility. The natural fertility depends on many factors such as period of exposure to intercourse, fecundity, duration of postpartum infecundity, spontaneous intrauterine mortality, and sterility. These factors vary with physiological and cultural conditions that may inadvertently reduce childbearing and consequently the observed natural fertility is likely to be below the reproductive potential.

Motivation for fertility regulation, one of the most interesting notion made in this study, is defined as the difference between the potential supply of and demand for children. An excess supply, that is, supply exceeds demand, would induce parents to seek a deliberate fertility control. On the other hand, if the potential supply falls short of demand, then there should be no such desire for parents to limit fertility consciously.

Motivation, however, only acts as a necessary condition for deliberate fertility control. Whether parents actually take an action to regulate their fertility also depends upon the cost of regulation as well. Fertility regulation imposes both psychic costs as well as market costs which include the direct monetary costs and also the costs of time involved. Subjective drawbacks associated with dis-utility for use of control methods such as abstinence or withdrawal, for example, are psychic costs. Regulation techniques such as induced abortion, sterilization, or contraceptive pills apparently are priced in time and money as well. Fertility regulation will actually be adopted if the motivation outweighs the cost of regulation. Conversely, if the cost of regulation is considered to be heavier than the motivation, no adoption of regulation will be made even if there may be potential unwanted children.

Based on the concepts stated above, fertility is jointly determined by the three key variables: supply, demand, and costs of regulation. Thus, fertility pattern changes as any of these variables changes. This framework allows Easterlin and Crimmins to explain the fertility revolution in a very good manner through the links from modernization to the three variables. Several aspects of modernization are particularly examined: innovations in public health and medical care, innovation in education, urbanization, development of new goods or techniques relevant to fertility control, and the establishment of family planning programs.

To be specific, better public health and medical care tend to increase the potential supply of children. The effects can either go through an increase in potential reproductivity within marriage or through an income effect on probability of survival because a healthier one is more likely to be more productive which, in turn, raises per capita income. Improvement in education may decrease the demand for the number of children and lower the costs of regulation for more relevant information or knowledge. At the same time it may also increase the potential supply of children as well. Urbanization is likely to reduce the demand for children by reducing tastes or lowering the prices of other goods relative to children, and it may also lower the costs of regulation costs due to, say, more availability of regulation services or information. New goods specifically related to fertility control such as modern condoms, oral contraceptive pills and IUD would reduce the costs of regulation. Finally the family planning programs also tend to lower the costs of regulation via offering services well below market prices. A reduction in costs, of course, should raise the adoption of regulation and as a consequence the probability of receiving an unwanted pregnancy should be lowered.

The empirical work is provided at both individual and aggregate levels. The individual data set is drawn from Sri Lanka and Columbia for the mid-1970s. The data for the Indian state of Karnataka and Taiwan from 1950s to 1970s is used for aggregate time series analysis. A cross sectional investigation is also made at aggregate level on the basis of ten states in India in 1970. All data sets are truncated to sub-data sets with women aged 35-44 who have been continually married and have had at least two children. Among the most striking findings, the motivation is found to be positively correlated with the adoption of regulation. Furthermore, in terms of contribution to the motivation, the supply force is demonstrated to be at least as important as demand and possibly more important. The regulation costs are measured by the number of contraceptive methods known to the respondents. The rationale for the approach is that those having more information of regulation methods would be more likely to choose the less costly method than their counterparts. This variable and other proxies for the cost of regulation show a consistent negative association with use of control. The fertility revolution - declining trend in fertility and increasing deliberate use of fertility control is well explained by the modernization process through its effect on the supply of, demand for children, and costs of regulation. Fertility declines because the demand for children falls and the widespread use of control techniques rises over time. The increase in adoption of regulation is a consequence of both upward changes in motivation and availability of control methods as well as a downward change in the costs of regulation. The authors, however, end up with limited evidence of an independent effect of regulation costs on the use of control, once motivation is taken into account. This indicates a dominant importance of motivation in the adoption and use of fertility control.

To refine Easterlin and Crimmins' framework, Montgomery (1987) presents an application of a switching regression model to World Fertility Survey (WFS) data and provides the estimates for fertility determination controlled for a sample-selection effect. Montgomery argues that the framework employed by Easterlin and Crimmins can be refashioned as a demand function for children with rationing constraints. Let b^* be the optimal number of children and B be the fecundity endowment. The b is assumed to be inversely associated with the surviving probability, shadow price of child, and positively related to the market as well as psychic costs of contraception. For the case of $b^* < B$, the optimal fertility demand for fertility is observed $b = b^*$. If $b \geq B$, on the other hand, the observed number of children is actually $b = B$. In other words, the observed number of children b is the minimum of b^* and B: $b = \min(b^*, B)$. Clearly, the observed b should not be employed directly to estimate the optimal demand function for children b^*. To identify the demand function, the observations with $b = B$, assuming B is known, are excluded from the sample. This exclusion leaves a truncated sample composed of observations with $b = b^* < B$. Apparently, OLS approach is not applicable to the truncated data set due to the sample selection bias (Heckman 1976, 1979). Alternatively, Montgomery employs the Tobit maximum likelihood procedure for the estimation which yields consistent estimates.

The data set for the empirical tests is from the 1975 WFS for women aged 40 and over in Sri Lanka and the 1976 WFS in Columbia. The truncation mechanism for $b \geq B$ is governed by women's behavior of "ever-use of contraceptives." Marriage duration is treated as a key determinant of the fertility constraint equation. As expected, the marriage duration is found to be positively associated with the fertility supply; parental education exclusively shows a force in reducing fertility demand; and survival probability presents a negative role in fertility demand. To test the selection bias with the sample truncation nature, a comparison is made between the estimates of OLS and of a Switching model. Very interestingly, the demand function by OLS appears flatter (more elastic) than the one arrived by the Switching model. Moreover, both education and marriage duration are found to increase the births averted by the use of contraceptives. It ranges from 2.2 number of births averted by contraception for the lowest of education and marriage duration to nearly 5.5 for the highest. Finally, the study reports a fertility demand function b^* with the fecundity supply constraint B. The results suggest, as expected, that the fecundity supply constraint B acts as a positive shift factor in the fertility demand function.

The optimal number of children, spacing and timing of giving birth are virtually the long run decision-making concerning fertility in life time horizon. However, a household's short run decision-making, concerning a marginal fertility for a given pregnancy, is also critically important. A woman, for example, due to a random failure of contraception, may have an unplanned pregnancy according to her long term optimal fertility plan. She then must make a decision in resolving her pregnancy - carrying it to term or terminating it. How do women differ in the probability of having an unexpected pregnancy? How do they make such short run decision in the resolution of a given pregnancy?

Leibowitz, Eison, and Chow (1986) apply an economic model to the short run decision-making behavior in resolving a given pregnancy for never-married, never-before-pregnant teenagers. The data are collected on 386 pregnant girls aged 13 to 19 who received either abortion or prenatal services from a variety of health providers in Ventura County, California between 1972 and 1974. A conditional logistic regression model (McFadden 1974) is employed to estimate the decision-making process. The decision variable for the pregnancy resolutions has three outcomes: to bear the baby without marrying; to bear the baby with marrying; or to abort. According to the consumer theory in economics, a teenager would rationally choose the alternative that maximizes her perceived utility as a function of the corresponding perceived costs and benefits. The perceived benefits of the second choice, for example, range from the emotional rewards of parenthood to the economic role of AFDC payment. The perceived opportunity costs would include both psychological and economic costs of having a baby. The determinants of such decision-making include schooling, as a proxy for a teenager's value of time; self-reported grade point average as an additional cost component of having a baby; two age-dummy variables (13-15 and 16-17) to monitor the variation in costs of having a baby as age changes; self-support and public aid AFDC payments to gauge the financing effect, and an ethnicity dummy variable (Mexican-American versus others) to test a potential difference between the two cohorts.

It is found that girls who are currently enrolled in school would be less likely to keep the baby than others. Self-report grade point also yields a negative role in keeping the baby. Turning to the age variable, they discover that girls aged 16 to 17 are more likely to carry the fetus to term than those aged 18 to 19. The girls of Mexican-American origin are more likely to marry and continue the pregnancy than other cohorts. The study also indicates that the public aids such as AFDC or Medicaid appear to have a positive contribution to the likelihood of not aborting the pregnancy. The self-support variable, however, tends to make girls more likely to abort than to become a single mother, but not significantly more likely to seek abortion than to marry.

In a similar study of teenagers, Serrato (1989) adopts a random utility approach to the decision-making mechanism in the resolution of pregnancy. Assuming the disturbances to be distributed in accordance with the generalized extreme value distribution, he derives a nested logit model (NL) to estimate the pregnant teenagers' probabilities of choosing an abortion or a birth with marrying as a sub-choice. Serrato emphasizes on the role that public policies would play in young women's decision concerning the resolution of a premarital pregnancy.

The data is drawn for 1053 single women aged 14 to 21 over the years of 1979 to 1986, from the National Longitudinal Study of Youth Work Experience (NLSY). This sample is also augmented with the state-year specific variables of AFDC programs, Medicaid programs, abortion providers, prenatal involvement levels, and unemployment rates from various sources. The results are interesting. For example, for Hispanic and white women the AFDC appears to reduce the likelihood of giving birth. Medicaid funded abortions tend to increase the probability of aborting the first premarital pregnancy. For the black women, however, an opposite role of these two variables is arrived. In agreement with the similar studies by Grossman and Joyce (1989, 1990), the provider of abortion shows a significant effect in favor of seeking the option of abortion. It is also found that the likelihood -of resolving a premarital pregnancy by abortion would increase with mother's educational level. Moreover, the women's wage income, sex education (for Hispanic and Whites only), and drug use all yield a positive role to the selection of obtaining abortion.

The evidence obtained shows that single women's probability of marrying before giving birth appears to increase with the variables: education, AFDC, Medication needy program, Ribicoff/CHAP program, Catholic (dummy), and south (dummy). The factors which would make women more likely to give birth while keeping a single status are: real wage rate, college education, sex education, expected AFDC payment, and mother-headed households.

With the multivariate analysis, Powell-Griner and Trent (1987) examine the socioeconomic determinants of women's induced abortions, based on the vital statistics files from National Center for Health Statistics (NCHS). The statistics provides the records for the three pregnancy outcomes: induced abortions (192,798), spontaneous abortions or fetal deaths (5,130), and live births (475,940) covering the nine reporting areas[1] in 1980. A log-linear model is employed to test the impacts of several explanatory variables (race, marital status, parity, education, area of residence, age) on households' behavior in resolving a given pregnancy by an induced abortion against giving birth or spontaneous abortion. A net distribution of induced abortions against one factor conditional on controlling the others simultaneously is provided. The simple bivariate relationship between the pregnancy outcome and its determinants is also presented for comparison.

It is hypothesized that the true distribution of pregnancy resolution against each factor would be distorted by others uncontrolled. For example, with the simple bivariate analysis, women with no previous births tend to seek abortion more readily than those with previous births. Yet, a reversed result for the parity would be arrived if all the other factors are controlled at the same time. Induced abortion odds appear to decline with ages up to the age of 29 years old, while women aged 35 and over are the second highest. However, the oldest age cohort shows the greatest likelihood of selecting an abortion if other factors are controlled simultaneously in the net distribution model. The rationale behind the finding yielded by the net distribution model is that the older women would find an unwanted pregnancy more costly than younger women. Metropolitan women are found to be more likely to seek induced abortion than non-metropolitan women. This finding is consistent with previous studies (Sullivan et al. 1977; Henshaw et al. 1981, 1985; and Henshaw and O'Reily 1983). Education, single status, and race (blacks vs. whites, as expected, tend to increase one's likelihood of seeking abortion option. Greater variation in abortion odds between married and single is also found for whites than for blacks. This finding implies that married non-whites may more widely use induced abortion as

[1] Missouri, Montana, Oregon, South Carolina, Tennessee, Utah, Vermont, Virginia, and New York city.

a contraception than do whites to resolve a pregnancy (Tietz 1977). Yet, unmarried blacks seem more likely to give birth as a means of achieving adult status (Hogan and Kitagawa 1985).

2.2 Prenatal Care Demand and Birth Outcome Production Functions

Prenatal medical care has been widely considered as one of the important determinants of birth outcome in the medical as well as economics of health literature. The behavioral nature of seeking the prenatal care by pregnant women is also well accepted by investigators. Yet, the efficacy of prenatal care on fetal health, neonatal and postneonatal mortality, and the feedback effect of fetal health status and endowment on the demand for prenatal care, i.e. self-selection effect, have been the subject of serious dispute for nearly four decades (Shwartz and Vinyard 1965; Kessner et al. 1973; Gortmaker 1979a; Harris 1982; Rosenzweig and Schultz 1983; Joyce 1987; Corman et al. 1987; Weller et al. 1987; Brown 1989; Joyce and Grossman 1990; Grossman and Joyce 1990).

In an attempt to correct some past deficiencies, the data upon all births and infant deaths for the native whites and blacks in New York City in 1968 are carefully examined by Gortmaker (1979a) using methods for the analysis of multidimensional contingency tables. The sample consists of 90,339 total births with 9,512 low birth weight infants (less than 2,500 grams at birth), 1,447 neonatal deaths (deaths to infants during the first 27 days of life), and 482 postneonatal deaths (deaths to infants occurring 28 days through 11 months). Gortmaker's study particularly emphasis the responsibility of the lack of control for a wide variety of selective factors for the observed but distorted or invalid role of prenatal care on the health outcomes of the newborn. The odds of a low birth weight infant, neonatal mortality, and postneonatal mortality are predicted as three outcomes of birth. Prenatal care is classified into three levels: inadequate, intermediate, or adequate. In addition to the prenatal medical care, the other explanatory factors of birth outcomes, considered to be controlled for, are parents' education, mother's age, parity, wedlock status, medical conditions observable during pregnancy, type of hospital service (private vs. general), and birth weight (for mortality equations only).

Several interesting findings are arrived in his study. First of all, the relatively high risk of low birth weight significantly tends to be associated with those with inadequate care, as apposed to adequate care, for both native whites and blacks. Yet, no credit could be granted to the role of prenatal care in reducing either neonatal mortality or postneonatal mortality if other factors are under control. Interestingly, a significant effect of care on the two mortalities would be observed if no correction is made for the other determinants. This finding suggests that prenatal care exerts its positive effect upon infant mortality virtually via the channel of reducing the risk of low birth weight. Second, variables of education, wedlock status, and the type of hospital services (private vs. general) are all found to contribute to the differences in the birth outcomes. Therefore, it is indicative of the fact that prenatal medical care can only partially capture the variations in infant mortality. Moreover, the racial differences in outcomes can not be entirely attributed to the differences in seeking prenatal care (Frank et al. 1991). Third, after controlling for those stated selective factors, blacks and some whites with inadequate care appear to bear substantially high risk of low birth weight relative to those with adequate care if the care is delivered on general services. In contrast, no statistically significant importance of receiving prenatal care to birth weight would be reached if the care is provided via private services for majority of the whites. This finding may reflect the nature of women's self-selection into a program of prenatal care due to the roles of some relevant but uncontrolled factors such as diet, smoking, drinking, and drug abuse during pregnancy. It is because of the lack of these factors controlled in the model, a caution is called for by the author about drawing a causal inferences for the efficacy of prenatal medical care to the birth outcomes according to the estimates.

Many investigators have noted that the timing or quantity of prenatal care and the duration of pregnancy are endogenously correlated (Well et al. 1958; Kane 1964; Russell and Burke 1975; Lewit 1977; Harris 1982). The role of prenatal care in the production of infant health would apparently be confounded by the nature of the endogeneity between prenatal care and infant gestational age. Harris (1982) conducts an extensive study inquiring how prenatal care is correlated with the duration of pregnancy and, more important, how the prenatal medical care favorably influences the outcome of pregnancy, controlling the infant gestational age.

Based on a data set of 138,943 live births, among which 1,229 died within 28 days, in Massachusetts during 1975-1976, Harris finds that the quantity of prenatal care yields a substantial positive effect on infant neonatal mortality rate, without controlling the age of gestation. In particular, beyond three prenatal care visits, the neonatal mortality rate rapidly declines. After about ten visits, however, the marginal efficacy of prenatal care seems to be not significant at all. Such behavior of prenatal care is attributed to women's adverse selection effect in seeking the care and the causal relation between prenatal care and the duration of pregnancy. In a time stochastic model treating both prenatal care and premature termination as two competing risks, the maximum likelihood estimates suggest that prenatal care reduces the risk of early termination. In addition, education (negative), illegitimacy (negative), and prior fetal loss (positive) are all revealed a statistically significant effects on the rate of termination of pregnancy. Moreover, the results also display that the initiation of prenatal care increases with gestational age.

In a model incorporating unobservables associated with infant health characteristics, Harris points out that fetal selection may introduce another type of bias in measuring the efficacy of prenatal care. The initial fetus population is likely to have heterogeneous health endowments. If the risk of early termination of pregnancy vary with the health endowments, the pregnancy outcome and the timing of prenatal care would be selective. For example, late prenatal care may act as an indictor of better health endowment for those infants with relatively low risk of early termination. This inverse relation between prenatal care delay and the duration of pregnancy may not be captured by the observed variables included in the model. As the consequence of a failure of controlling the gestational age, the efficacy of prenatal care may be overestimated. On the other hand, the efficacy of prenatal care may be underestimated if mothers who perceive their babies to be less fit initiate care earlier, while those with healthier babies may delay care.

In order to sort out the direct marginal product of prenatal care for fetal growth, a model controlling for the gestational age and other potentially confounding factors is specified. In particular, the effect of trimester of initiation of care on rate of intrauterine growth is estimated by maximum likelihood approach for 6,736 blacks. An evidence is found for statistically significant effects on fetal growth of gestational age, legitimacy status, prior perinatal loss, and parity. Prenatal care appears to increase the fetal growth compared to those without care at all. The estimated marginal impact of care, however, is statistically insignificant at 5% level. This finding may suggest that the contributing effect of prenatal care on the fetal growth primarily go through the effect of care on gestational age. The estimates shows that the direct effect of care is about 60 grams while the indirect effect via gestational age would be about 110 grams.

Harris finally tests how prenatal care would affect infant mortality in a model controlling both gestational age and birth weight as well as some other observable. Interestingly, he finds that prenatal care yields its favorable impact on the probability of perinatal survival almost indirectly via its effects on birth weight and gestational age. The effects of care on perinatal mortality via birth weight also appears to be dominant. The direct effect on prenatal mortality of prenatal care is weakly positive in one model and is even negative in another. This finding is consistent with the hypothesis of fetal selection and previous study by Gortmaker (1979a).

In a study dealing with infant health production technology, Rosenzweig and Schultz (1982) present a health production model controlling for adverse selection effect introduced by heterogeneity of health endowment. Their study extensively reveals how the exogenous variations in health endowment would affect women's behavior in selecting health related inputs and as a consequence, a significant positive impact on child health of early prenatal care could be masked by the adverse selection bias.

Applying the notion of household production model introduced by Becker (1965), Rosenzweig and Schultz set a household utility function with the direct arguments of regular goods X_i, health related goods Y_j, and child health H_i. The child health is governed by a health production function with the inputs of Y_j, health inputs I_k, and a health endowments factor u which is assumed to be unobservable to researchers. Rosenzweig and Schultz point out that the observed input-constant relationship between Y_m and H would differ from the true marginal product of Y_m, if the dependency of Y_m and other health inputs I_k on health endowment u is disregarded. The bias arises as the heterogeneity of health endowment affect women's selection of prenatal care and other health inputs for remedial purpose. For example, a pregnant woman with some information on her poor health endowment, from prior histories of pregnancy complications or prior birth outcomes reflecting child health, may seek earlier or more prenatal care than others. The dependency on health endowment of these health inputs is shown by the household's reduced-form demand functions for these inputs, where the health endowment as well as other income and price factors act as direct arguments. In addition to the demand for prenatal care, smoking, the order of the current live birth, and the age of the mother at the birth are all treated as endogenous or behavioral variables having a direct technical relationship with birth weight, the index of infant health employed in this study.

Statistically, Rosenzweig and Schultz indicate that OLS estimation approach would yield a biased health technology, based on that observed hybrid relationship between prenatal care and infant health. To isolate the true technology from the hybrid relationship, 2SLS approach is employed in this study. With respect to the modeling, Rosenzweig and Schultz adopt the Translog as well as Cobb-Douglas functions to model the health technology treating the four inputs, delay, smoking, parity, and age as endogenous variables. All of the four behavioral variables are regressed, in the first stage by OLS, on a set of identified exogenous variables at individual level: race, and parents' educational levels; and at state or SMSA level: husband's life cycle income, metropolitan residence, health expenditures, health department family planning, cigarette price index, milk price index, hospital family planning, population per doctor, OB-GYN per capita, manufacturing jobs, service jobs, unemployment rates, hospital beds per capita, sales tax on cigarettes, and the size of the SMSA for inhabitants of SMSAs. The race dummy variable is employed to test for a hypothesized difference in the health inputs as well as in infant weight at birth between whites and blacks. Moreover, in order to identify the effect of gestation on fetal growth and thus infant weight at birth, a birth weight function standardized for gestation is also estimated.[2] The empirical work is based on the 1967, 1968, and 1969 U.S. National Natality Followback Surveys. It is a national sample data set of about 10,000 legitimate live births.

Rosenzweig and Schultz find the birth weight technology to be characterized by Cobb-Douglas function. Prenatal care delay during women's pregnancy shows no appreciable effect on birth weight according to the OLS model, whereas the 2SLS estimates suggest a significant contribution of prenatal care to infant weight at birth. This finding confirms the existence of adverse selection bias in seeking prenatal care within the observed hybrid health technology, due

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[2] The fetal growth in weight is projected by a cubic function from the sample data as follows:

$$\text{Birth weight} = \underset{(7.7)}{10{,}107} - \underset{(10.0)}{1{,}042}\ \text{weeks} + \underset{(10.4)}{37.8}\ \text{weeks}^2 - \underset{(10.9)}{0.398}\ \text{weeks}^3$$

where the absolute values of t-statistics are reported in parentheses.

to the heterogeneity of health endowment. Interestingly, it is demonstrated that the beneficial effect of prenatal care are no longer evident, no matter whether or not the health heterogeneity is controlled for, in the standardized birth weight model. This result suggests that prenatal care yields its effect on fetal growth primarily by extending gestation (Harris 1982). The results also indicate that the heterogeneity of endowment in OLS model tends to mask the significant negative effects of maternal smoking on the rate of fetal growth. Based on the Cobb-Douglas 2SLS estimates, the effect on birth weight standardized for gestation of smoking by mother would increase by nearly 75% over its effect on actual birth weight. Moreover, parity is beneficially found to affect birth weight as twice large by 2SLS approach as indicated by OLS. As expected, the race dummy variable yields a significant difference in birth weight in favor of the whites. Evidence is also obtained for women with favorable health endowments to bear children at significantly later ages. The age, however, appears to exert little effect on birth weight.

Two of the noticeable points for this study merit attention. First, as also indicated by the authors, the exogenous variables employed in the study to estimate the four behavioral health inputs may be not really determined by external forces independently from the model system. For example, the governmental health and hospital expenditures per capital may be more clustered at the areas where the people's poor health endowments or health conditions are identified and thus call for more health programs establishment by local governments. Second, the data set used in the study only pertains to the married women who are currently with husbands and also give a live birth. These women, from a statistical point of view, are apparently a sub-sample selected from the entire population of pregnant women. If the sample selection is not made in a random manner, a self-selection bias may be introduced by the nature of incidental sample censoring into the demand for behavioral inputs as well as health technology under consideration.

Neonatal mortality rates in the United Sates have sharply declined from 17.9 deaths in 1964 to 7.7 deaths in 1982 per one thousand live births. What factors or health inputs would account for such trend? Corman, Joyce, and Grossman (1987) present an infant health production function that simultaneously consider the effects of a variety of health inputs on race-specific neonatal mortality rates in the United States. Since health endowment and prenatal care may be two confounding variables in influencing mortality rates, a particular emphasis is placed on the direct effect of prenatal care on neonatal mortality by controlling for low birth weight, as a proxy for the health endowment, in the model. In addition, perinatal and neonatal care, the use of abortion services, the use of contraception services, and maternal cigarette smoking are also treated as endogenous determinants of the mortality rates. To deal with the endogeneity, 2SLS estimation is emphasized compared to the estimation by OLS. To reduce the endogeneities of these health inputs, a reduced form regression is conducted for each of the endogenous variables upon a set of exogenous factors such as price and availability measures, relevant socioeconomic characteristics, risk measure, and the biological endowments.

Based on a modified sample of 677 counties for whites and 357 counties for blacks which account for about 80% of population of the U.S. in 1977, both OLS and 2SLS estimates are provided for the two model specifications with/without controlling for the risk of low birth weight. The Wu-test shows an evidence of non-endogeneity of health inputs for both whites and blacks in OLS models controlling for the incidence of low birth weight. Abortion services and prenatal care are found to be statistically significant in reducing the neonatal mortality rates. In the case of prenatal care, the results show a consistency with previous studies (Gortmaker 1979a; Harris 1982). For example, a considerably large negative effect of prenatal care would be observed in the model without controlling for low birth weight, whereas the magnitude as well as significance of the effect would be reduced if the incidence of low birth weight is also simultaneously controlled for. For blacks, prenatal care even yields no significant effect on mortality rates after holding the risk of low birth weight constant. Again, this finding suggests that prenatal care makes its favorable contribution in reducing mortality mainly through the reduction of risk of low birth weight. The risk of low birth weight itself, as expected, presents the most important contribution in determination of neonatal mortality rates. An evidence is also found for the significant impacts of smoking (whites only), maternal nutrition program, high risk women aged 15-19 or 40-44 (whites only). In sum, the estimates basically underscore the qualitative and quantitative contributions to the recent downward trend in neonatal mortality rates in U.S. of health programs such as abortion services, nutrition programs (WIC), and prenatal medical care.

Self-selection problem has been fruitfully investigated in the literature of economics and applied econometrics. Yet, little had been systematically done in the context of prenatal care and birth outcome production functions although researcher have noted it (Harris 1982; Maddala 1985; Rosenzweig and Schultz 1983) until a study by Grossman and Joyce (1988). In a series of studies by Grossman and Joyce (1988, 1990) and Joyce and Grossman (1990), the issue of self-selection effects associated with both behavioral endogeneity of demand for prenatal care and sample selection in the resolution of pregnancy is introduced and also well addressed.

Grossman and Joyce (1988, 1990) argue that pregnant women face a selection problem concerning the resolution of pregnancies as live births or induced abortions. Moreover, the outcome of the selection would likely further influence women's prenatal behavior such as prenatal care demand and, of course, the birth outcomes. Thus, as an important contribution, they incorporate the decision of a pregnant woman to give birth or to obtain an abortion into the demand function for prenatal care as well as the infant health production functions. With regard to the outcome of pregnancy resolutions, the selection is favorable if women with better infant health endowment or women who desire to make relatively large investment in infants are more likely to give birth. The selection would be adverse if women who would make relatively small investments are more likely to give birth. The investments in infant health range from prenatal care services, nutritious food, appropriate exercise, to use of fewer drugs, smoking or other potential harmful substances. The other type of self-selection in use of health inputs pertaining to women who give births are also concerned in this study. Due to the heterogeneity of health endowments, adverse selection arises as women who anticipate a problematic birth seek out more remedial care while women with better infant health status seek out less (Rosenzweig and Schultz 1983, 1988; Harris 1982). The self-selection in input use could be favorable if those women with better health endowment are also seeking more prenatal care services and other healthy inputs. Statistically, the first type of self-selection effect is caused by the censoring nature of sample selection whereas the second type of self-selection effect is associated with the heterogeneity of health endowments and the endogeneity of prenatal care in the production function of birth outcome. To control for the two types of selection effects simultaneously, the estimation of their model is conducted based on the TSLS probit method introduced originally by Heckman (1976, 1979) and later developed by Lee, Maddala, and Trost (1980).

Based on a combined population of 105,000 births and 89,000 induced abortions taken from New York City vital statistics in 1984, they particularly examined a randomly chosen sub-samples of 11,591 pregnancies to whites aged 20 years and older and 11,016 pregnancies to blacks of the same age. For the demand functions for prenatal care, evidence is found for a self-selection effect due to the sample censoring in pregnancy resolution in black cohort but not in white cohort. In the birth weight production function, the results also suggest a race-specific selection bias for blacks only. In particular, compared to the black women who actually gave births, those who aborted would have more delayed the initiation of prenatal care and would have given birth to lighter infants if they had selected the option of giving birth. According to the specification of unobservables in their model, this finding suggests that the cost of contraception, one of the unobservables, appears to dominant the others. Grossman and Joyce argue that the racial differences with respect to the selectivity bias may be due to a greater shadow price of contraception for blacks than for whites. Further, the shadow price is apt to vary more among blacks than it does among whites (Pratt et al. 1984; Henshaw et al. 1985; Stephen et al. 1988). In the case of marginal product of prenatal care in birth weight production function, prenatal care shows a weaker positive efficacy in increasing birth weight for whites than for blacks. Compared to the estimates given by OLS model, the magnitude of efficacy of prenatal care would increase for both whites and blacks as the endogeneity of prenatal care is corrected in TSLS model. However, the Wu-test yields no evidence to reject the null hypothesis of exogeneity of prenatal care in the birth weight production function. This finding suggests that the correction of self-selection effect due to sample censoring may also capture, to some extent, the effects associated with the selection of endogenous health inputs. As a consequence, as the authors claimed, in this study more credit and emphasis should be granted to the OLS model controlling for the sample selection bias than to the TSLS model controlling for both sample selection as well as endogeneity.

In a recent study, Frank et al. (1991) present their updated estimates of race-specific impact of prenatal care on birth weight outcomes. Following the work by Corman et al. (1987) and Joyce (1987), a modified version of quasi-structural production function is tested. The analysis is made for a pooled aggregate data set at county level: 2,317 counties with population of at least 10,000 whites and 660 counties with at least 5,000 blacks in the U.S. during 1975 through 1984. Since this is an aggregate data set, Frank et al. place a particular emphasis on a potential primary source of bias in the estimation of the birth weight production function using aggregate county-specific data. This is because the women's health endowments may not be well distributed in a random manner among the population in the counties selected into the sample. Rather, it is likely to have a distribution of women with similar health endowments geographically clustering at some particular counties. To control for the potential unmeasured heterogeneity of health endowments across counties, a fixed effects model is introduced. In the model, the birth outcome is measured by county-specific low birth weight rate (LBW) by race, i.e. the percentage of infants weighing less than 25,00 grams at birth in a county. In order to gauge the direct role of prenatal care in first trimester in birth outcome production, use of abortion services, cigarette smoking, birth order (percentage of first births), and income are simultaneously controlled in the model. In addition, four time dummy variables are included to correct for time trends in the pooled data.

According to the weighted least square estimates for the fixed effects model, they find no empirical evidence to reject the null hypothesis of exogeneity of health inputs for blacks. For whites, Wu-test still fail to reject the exogeneity at 0.05 level but would reject at the 0.10 level. Their finding shows an agreement with that of Maeshiro (1979) and Grossman and Joyce (1988, 1990). For this reason, the single stage estimates are particularly focused on. Based on the estimates, initiation of prenatal care in the first trimester and per capital income are found to be the two significant determinants of low birth weight rates for whites. For blacks, prenatal care effect remains significant whereas income effect is no longer significant statistically. Moreover, the marginal product of prenatal care also shows a larger magnitude for blacks (-0.018) than for whites (-0.007). Frank et al. attribute this racial difference in efficacy of prenatal care partially to the improvements in the types of services available to minority women during 1974-1976.

The descriptive statistics indicates a substantial racial gap of LBW between whites (5.3%) and blacks (12.0%) in 1984. To analyze how prenatal care contributed to this difference, an interesting projection is made by controlling prenatal care initiation at different levels. Based on the mean of first trimester initiation at 66.5% for blacks and at 82.50% for whites, it is found that the LBW of blacks would have decreased from 12.80% to 11.79% if 82.50% of blacks had initiated the care in first trimester. This reduction is expected to lead to a drop in the ratio of black to white LBW rates from 2.28 to 2.22. Further, the LBW, with 100% of blacks initiating the care in first trimester, would go down slightly to 11.4%, leading to a corresponding drop in the black/white LBW ratio to 2.16. The policy implication by the projection suggests that the racial difference in county-specific LBW between blacks and whites should not be primarily attributed to the racing difference in prenatal care services. Therefore, rather than relying on expanding early initiation of prenatal care to blacks, something else (e.g. welfare programs, health education) may be the better candidates of a strategy to reduce the high LBW of blacks relative to whites.

CHAPTER THREE

DECISION-MAKING IN THE RESOLUTION OF PREGNANCY

3.1 General Remarks

Consumer theory in microeconomics provides a well-defined utility framework dealing with consumer choice problems in a continuous space of alternatives. In the context of women's decision-making concerning pregnancy resolutions, however, the decision maker faces two exclusively discrete options: giving birth or obtaining abortion. This makes the utility maximization approach not applicable since the outcome ends up at a "corner" solution, a point where the usual first-order conditions for an optimum do not hold. The widespread use of induced abortion, since its legalization by the Supreme Court in 1973, has permitted much greater choices in the number and timing of births. According to Easterlin and Crimmins (1985), the changes in fertility, throughout the world, have been characterized by a declining in demand for, a rise in supply of children, and a fall in the costs of fertility regulation (induced abortion and contraceptive techniques) accompanied with economic and social modernization. This fertility revolution makes the modern women's resolution of pregnancy highly deliberate or self-selective. If a women's pregnancy is wanted or expected on plan, she would normally carry it to term unless it is not permissible medically for health problem, or fetal health endowment is well below a critical level to survive. If a pregnancy is unwanted or unexpected, the woman has to make a self-selection decision between giving birth or obtaining abortion. Given demand for fertility and costs of fertility control, the higher the supply of children or reproductivity, the more likely a woman is to end up an unwanted pregnancy. The supply of children reflects both probability of survival and natural fertility which may be determined, for example, by the period of exposure to intercourse or fecundity or sterility. In contrast, holding supply and costs of fertility control constant, higher demand would more likely lead to unexpected pregnancy. The demand for children reflects parents' tastes or preference for children since children directly augment parents' utility (Becker 1973). The fertility regulation refers to both contraceptive methods and induced abortion. Therefore, the costs should reflect the expenses of the two methods. Costs of contraception include the market costs (money, time, and information costs) and also a psychic component, that is, the disutility associated with the idea or practice of contraception. A joint production with adoption of contraception (Pollack and Wachter 1975), for example, at some extent, may reduce the gratification yielded by sexual intercourse. Households with this problem, therefore, may have a moral objection to adoption of a contraceptive instrument for the high psychic displeasure associated with it. Efficacy of contraception also differs across individuals due to the knowledge of using it or individual fecundities (Michael 1973). The costs of induced abortion, as an alternative to contraception, depends on individual attitudes toward the notion of this method. Evidently the probability of choosing to give birth is inversely influenced by the costs of induced abortion. In sum, since not all pregnancies are wanted or planned and would be necessarily carried to term, therefore, a rational self-selection decision with regard to pregnancy resolutions is deliberately made during women's pregnancy period. Now the issue becomes what the decision mechanism for such self-selection is.

The economics of fertility conventionally emphasizes the overall picture of fertility and fertility control such as the life time optimal number of children, spacing or timing of children, and the determinants of these variables or the way they change with modernization. Can this framework be employed to formulate the decision structure concerning women's self-selection for a given pregnancy? For example, a long run optimal fertility plan can be derived from Becker's children quality-quantity model (1973). Households are assumed to have their children over time according to their long term fertility plan (Willis 1973; Becker and Lewis 1973). In particular, Grossman and Joyce (1989,1990) define a dynamic version of children for the ith woman in year t (N_{it}^*) as a threshold, and the N_{it} as the actual number of children for the ith woman in year t. They then argue that a pregnant woman will give birth provided $N_{it}^* \geq N_{it}$ and will abort provided $N_{it}^* < N_{it}$. The approach of applying a long-run optimal fertility plan to present birth decision, however, has made two implicit assumptions. First, it simply assumes that the households make a long term fertility plan and also as expected behave according to such plan. This assumption may not be plausible at some point. According to Michael (1974), women may suffer from a particular type of myopia: although they consider long-run repercussions of their fertility-control decision in terms of future costs and benefits of children, they may not determine an optimal fertility-control strategy in such a way. In other words, in each period, an independent decision may be made with respect to fertility control. Second, and more important, the long run fertility approach may not be fully able to capture the impacts of random disturbances (health endowment, costs of induced abortion or contraception) on the present decision-making in resolution of pregnancy. For example, a religious couple may end up with an unwanted pregnancy due to a failure of contraception. If the couple has a moral or religious objection to induced abortion, the fetus would then still be carried to term even if the actual number of children N_{it} is greater than the N_{it}^* from the long term optimal plan at present ($N_{it}^* < N_{it}$).

One cannot deny that the stock of children (not including the current fetus) affects the decision concerning pregnancy resolutions, but the argument is that the theory of the optimal total number of children may not be able to fully capture the effects of random disturbances on the selection decision for a given pregnancy. Therefore, one of the hypotheses in this study is that women's decision with regard to the present pregnancy resolution is deliberately made on the basis of comparing the utilities of having an additional baby less its corresponding cost and that of obtaining abortion alternatively at present.

3.2 Approach of Random Utility Functions

In econometric literature dealing with multi-choice problems under uncertainty, a random utility model has been well applied by researchers (McFadden 1973, 1974; Manski 1977; Nakosteen and Zimmer 1980; Ben-Akiva and Lerman 1989; and Greene 1990). The basic idea is as follows. If all alternative choices are uncertain, random utilities can be assigned to each choice. The underlying theory shows that a rational individual selects an alternative yielding the highest utilities. Since the utilities are random, the probability of choosing one alternative against others is considered. Manski (1973) identifies four distinct sources of randomness: unobserved attributes, unobserved taste variables, measurement errors and imperfect information. Therefore, the notion of random utility model can be readily applied in the context of decision-making for pregnancy resolutions where some important but hard-to-measure factors are inevitably involved.[3]

A Utility Model of Fertility

A pregnant woman's fertility behavior can be described by a household utility function as follows:

$$\begin{gathered} U_i = U(N_i + D_i,\ h_i,\ I_i - C_{Di};\ S_i) \\ C_{Di} = p_{bi}(N_i + D_i)h_i + (1 - D_i)p_{ai} \\ \frac{\partial U}{\partial N_i} > 0\ ;\ \frac{\partial U}{\partial h_i} > 0\ ;\ \frac{\partial U}{\partial (I_i - C_{Di})} > 0; \end{gathered} \qquad (1)$$

where the i stands for the individual index, N_i shows the stock of children, D_i is a dichotomous decision variable: it equals 1 if the option of giving birth is selected and 0 if abortion is ended,

[3] An application of random utility function in the decision-making of pregnancy resolution can be found in the work by Serrato (1989). In testing the influence of public assistance, sex education, and labor market opportunities on the decision-making of premarital teenager's pregnancy resolution, Serrato has applied the concept of random utility function and used a NL model for estimation.

p_{bi} is the unit health cost of a child, p_{ai} is the cost of obtaining an induced abortion, I_i is a vector of constraints over one's living resources such as money income, time, health endowments, psychic endowments or religions, I_i-C_{ai} then shows the rest of resources available for maintaining other desired living conditions (as a plausible assumption I_i-C_{ai} is held to be positive), S_i denotes the relevant household's characteristics such as education, age, marital status, which would affect the marginal impact of those utility arguments specified, and h_i shows an average child health level demanded by parents through a health production function which is specified as:

$$h_i=h(m_i,\ x_{hi}\ ;\ e_i) \tag{2}$$

where the m_i, x_{hi}, e_i stand respectively for the inputs of medical care, other health related inputs, and the health endowment for individual household i.

Birth Utility Function

As one decides to carry the fetus to term a birth, the utility function is generated by equation 1 as:

$$U^{bi}=U(N_i+1,\ h_i,\ I_i-C_{bi};\ S_i) \tag{3}$$
$$C_{bi}=p_{bi}(N_i+1)h_i$$

The household utility function of fertility indicates that a child, as a normal good, directly augments parental utility. The marginal benefits of having additional baby depend upon the family's socioeconomic variables such as income or parental education, and demographic variables such as parental race, age, costs of contraception, history of births as well as health endowments. Costs of contraception, for example, may most likely make a negative contribution to the wantedness of a pregnancy which may in turn reduce the marginal utilities of additional baby. Put differently, the costs of contraception are positively associated with the likelihood of receiving an unwanted pregnancy which in principle shifts the parental marginal utility of a child downward. Following Becker and Lewis (1973), the costs of having a baby P_{bi}, also termed as shadow price of a child, can be roughly considered as two parts: variable cost π_{hi}, which directly depends on child health level h_i, and a "fixed" cost π_f for having each baby.

To be specific, suppose a child health production function is:

$$h_i = h(m_i,\ x_{hi};\ e_i) \tag{1}$$

where m_i denotes the medical care input, x_{hi} is a vector of the other health related inputs such as time, baby food, clothing and woman's healthy behavior such as proper diet, appropriate exercise, smoking, drinking, and avoidance of stress, and e_i shows the health endowment. All these health inputs have positive prices. Fixed costs measure the direct monetary fees for birth, mother's health depreciation of giving birth, and psychic costs as well. For example, an infant may interpret or reduce a couple's social activities or leisure time; a woman's good figure may be affected unfavorably by carrying a fetus to term; an unmarried woman may suffer from the social stigma of being assumed to be sexually active or promiscuous if a baby is born.

Abortion Utility Function

The utility function associated with selecting abortion option is reduced from equation 1 as a woman decides to abort the fetus, that is, $D_i = 0$:

$$U^{ai} = U(N_i,\ h_i,\ I_i - C_{ai};\ S_i) \tag{5}$$
$$C_{ai} = p_{bi} N_i h_i + p_{ai}$$

Clearly, by selecting the abortion option, the potential costs of having a baby ($p_{bi}h_i$) can be averted. Put differently, the marginal benefits of obtaining abortion are yielded through a subsidy by not paying the potential shadow price of $p_{bi}h_i$ for additional baby. Torres and Forrest (1988), in a survey of abortion patients in 1987, found four major factors contributing to abortion women's decision. Of 1,900 respondents, 76% said that having a baby would change her life, 68% said that they cannot afford to have the baby, 51% said that they wanted to avoid single parenthood, and 31% said that obtaining an abortion would prevent others from knowing that they had sex or were pregnant. The abortion option, on the other hand, also costs at a positive price. The abortion costs include a direct abortion cost P_a and the foregone marginal utilities associated being with a child otherwise. The direct abortion cost can also be seen as a sum of the monetary costs and health cost component. Monetary costs include direct abortion fees, time, and searching costs for information. The health costs reflect both physical health cost and moral cost. Studies in medical literature show that the abortion option may reduce the women's future reproductivity, which is an example of physical health cost. Moral cost can be found for parents with a moral objection to induced abortion due to, say, religions. Very religious parents, for example, are much less likely to abort an unwanted pregnancy than less religious parents. Thus, less religious women would be more likely to select into the abortion sample while the more religious women would be more likely to select into the birth sample, holding everything also constant. It is plausible to assume that the health costs dominant the monetary costs.

Rational parents are assumed to make their decisions regarding pregnancy resolutions on the basis of comparing the above birth and abortion utilities. Thus, a mechanism governing women's binary self-selection behavior in pregnancy resolutions is defined as:

Self-Selection Mechanism of Pregnancy Resolution

Select giving Birth if $U^{bi} \geq U^{ai}$

Select obtaining abortion if $U^{bi} < U^{ai}$

3.3 Probability Function of Giving Birth - A Probit Model

The utility functions of birth and abortion are random because of some unobservable random disturbances evidently involved. First, the event of becoming pregnant is virtually considered to be a random outcome because it depends upon a variety of hard-to-control factors such as motivation of fertility control, efficacy of contraception, and couple's fecundity, which in turn would partially determine the wantedness of a pregnancy.

Second, given to be pregnant, women's utility of giving a birth or obtaining an abortion depends on such factors as health endowment, and abortion costs which are not practically controllable either. Finally, the measurement errors and imperfect information over the variables explicitly considered in the utility functions may introduce the third type of randomness in modeling the decision-making for pregnancy resolutions. Utility functions become random as these underlying factors cannot be explicitly controlled but, instead, treated as stochastic error terms in the model.

To make the binary selection decisions under the random utility framework operational, define a random utility differential function of giving birth vs. abortion as follows:

Random Utility Differential Function of Giving Birth Against Abortion

$$\Delta U_i = U^{b_i} - U^{a_i} = \Gamma(X_{ui}) + \varepsilon_{ui} \tag{6}$$

where the utility differential is decomposed into a deterministic or systematic components, $\Gamma(X_{ui})$, and a stochastic component ϵ_{ui} for simplicity. To be specific, assume that $\epsilon_{ui} \sim N(\mu_{ui}, \sigma_{ui})$. The probabilities associated with each choice are formulated as:

Probability of giving birth $= \Pr(\Delta U_i \geq 0)$

Probability of obtaining abortion $= \Pr(\Delta U_i < 0)$

In particular, for estimation purpose, the probability function of giving birth vs. abortion can be further demonstrated as:

Probability Function of Giving Birth Against Abortion

$$Pr(\ Birth \mid Birth,\ Abort) = Pr(\Delta U_i \geq 0) = Pr(\varepsilon_{ui} \leq \Gamma(X_{ui})) \tag{7}$$

$$= \int_{-\infty}^{\Gamma(X_{ui})} \frac{1}{\sqrt{2\pi\sigma_{ui}^2}} \exp\left[-\frac{1}{2}\left(\frac{\varepsilon_{ui}-\mu_{ui}}{\sigma_{ui}}\right)^2\right] d\varepsilon_{ui} \tag{8}$$

$$= \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{\frac{\Gamma(X_{ui})-\mu_{ui}}{\sigma_{ui}}} \exp\left[-\frac{1}{2}\mu_{ui}^2\right] d\varepsilon_{ui} \tag{9}$$

$$= F\left(\frac{\Gamma(X_{ui})-\mu_{ui}}{\sigma_{ui}}\right) \tag{10}$$

where the F denotes the standardized normal cumulative distribution function (CDF).

For any individual pregnant woman, once a proper estimate of the standardized utility differential index $\Delta U_i^* = [\Gamma(X_{ui}) - \mu_{ui}] / \sigma_{ui}$ (simply a standardized normal variable) is obtained, is straightforward to predict how likely she would select to carry the fetus to term.[4] Since the CDF is a monotonically increasing function of the normal index variable ΔU_i^*, for individual i, the greater the index, the higher the probability to select the option of giving birth.

Analytically, based on the model setting, the contributions of some important unobservables to the birth differential utility function can be well expected. Practice of contraception, for example, differs across families since it has a positive price in terms of direct monetary cost, time costs, efficiency (Westoff 1988) and psychic costs due to the joint production

[4] A probit likelihood approach will be employed to estimate the probability function of giving birth. Chapters five and six present the estimation procedure and empirical works.

(Grossman 1972, Pollack and Wachter 1975). Easterlin and Crimmins (1985) found a significantly negative association between costs of fertility control and use of the control for married women aged 35-44 in both Colombia and Sri Lanka during 1970s. Studies by Henshaw and Silverman (1988) also indicate that, due to the costs of contraception, a number of abortion patients did not use any contraceptive method around the time of conception. Jones at al. (1988) point out that experiences from other countries has shown that lower abortion rates can be achieved through improved contraceptive use. In short, with other factors equal, the lower the cost of contraception, the higher the adoption of contraceptive method for fertility control, the higher the probability of having a wanted pregnancy, and consequently the higher the marginal utility of giving the present birth. That is, the price of contraception makes a negative contribution to the birth differential utilities.

Second, health endowment, which is also hard-to-measure, may significantly characterizes women's selection behavior with regard to the pregnancy resolutions, as previously indicated. The selection is favorable towards to giving birth if women whose fetuses have better health endowments are more likely to carry fetuses to term or the converse. Rosenzweig and Schultz (1983, 1985) show that pregnant women with a poor health endowment may have to use more medical care for the problematic fetus. Put differently, the poorer the endowment, the higher the cost of having a baby or the converse. In addition, for medical reasons, a very poor health endowment (well below some critical point) may make a selection of giving birth not permissible. Therefore, a positive contribution of health endowment to the differential utilities of giving birth ($\partial \Delta U_i / \partial e_i$) is expected.

Third, as the abortion utility function shows, the abortion cost (p_{ai}) is negatively associated with the utility of selecting abortion ($\partial U^{ai} / \partial p_{ai} < 0$). Obviously, a woman with higher cost of abortion (p_{ai}) would be less likely to abort than others with lower abortion costs

($\partial \Delta U_i / \partial p_{ai}$), with everything else equal. That is, the costs of abortion act as a positive argument in the utility differential function. It is worthwhile to notice how the costs of contraception and costs of abortion, though both as costs of birth control, differ (opposite) with respect to their effects in the utility differential function of giving birth. Finally, a negative relationship between the shadow price of having a baby and the birth differential utility function is anticipated. It was showed earlier that the shadow price of having a baby is defined a fixed cost π_f plus a variable cost (π_{hi}) which directly depends on children's health level maintained or desired by parents. According to Grossman's health investment model (1972), people with higher education demand higher health than that of people with lower education. That is, the health or quality of children acts as a substitute good in parental utility function. For a given income and other resources, as education increases, an increase in demand for health would reduce the demand for quantity of children. Alternatively, children's health embodies a given set of inputs of time and goods, a rise in demand for health means a rise in the shadow price of children. Therefore, with a increase in the price of children relative to other goods with fixed prices, the demand for number of children would fall. In short, if a more educated household demands a higher quality of children, the it would be more likely to select into the abortion group than other households with a lower educational level. Evidence for the rationale of the negative association between probability and cost of giving birth can also be found through opportunity costs of taking care of baby. People making any decision are always subject to the scarce resource-time constraint (Becker 1965). The more time input required to take care of a baby means less time left for working, leisure, or obtaining something else. Put differently, given the same time inputs for given level of health, the higher one's price of time (wage rate), the more costly it will be to have a baby. A recent studies by Heckman and Walker (1990), based on a Swedish longitudinal data set, finds that rising female wages delay times for all conceptions and reduce total conceptions.

The result is robust across a variety of empirical specifications. As usual, if mother's education is used as a proxy for her wage rate (unit cost of time input), one can then logically expect a negative correlation between education and birth rates (Nerlove 1973; Butz and Ward 1979). Teenagers, for another example, have much higher likelihood of selecting the abortion option than that of adults. A study by Torres and Forrest (1988) shows that 92% of teenagers under age 19 indicated that having baby would interrupt their current life. One of the explanation might be that, due to the discontinuation or interruption of their education or their relatively independent single life style, many teenagers would bear a higher opportunity cost of having a baby than adults.

CHAPTER FOUR

SELF-SELECTION IN PRENATAL CARE DEMAND AND HEALTH PRODUCTION

4.1 General Remarks

Pregnant women's prenatal care demand and their infant health production have received widespread attention (Rosenzweig and Schultz 1983, 1988; Kleinman et al. 1984, 1989; Harris 1982; Poland and Giblin 1987; Chao et al. 1984; Knoll 1986; Corman, Joyce, and Grossman 1987; Joyce 1987; Joyce and Grossman 1990; Grossman and Joyce 1990; and Frank et al. 1991). Rosenzweig and Schultz (1983, 1988); Corman et al. (1987); and Joyce (1987) present the birth outcome production function and demand function for health inputs in the context of a static economic model of household production. In Grossman and Joyce's study (1990, 1991), women's decision for pregnancy resolutions is, for the first time, incorporated into the estimation of prenatal care demand and birth weight production functions.

From a statistical standpoint, the individuals obtaining prenatal care and giving birth are a sub-sample self-selected from the total population of pregnant women. If the sample selection mechanism is random, then the application of usual statistical analysis to this issue presents no problems except for the efficiency of the resulting parameter estimates (Maddala 1985). A major problem in econometrics arises if the selection is in a non-random way. Such non-random selection may result in both losses of efficiency and of biasedness or consistency of the estimated coefficients for a model that fails to take into account the nature of non-random selection (Griliches, Hall, and Hausman 1978). In the context of pregnant women's prenatal care demand and infant health production function, the selection of birth sample from total pregnant population is characterized by women's behavioral self-selection governed by the birth utility differential model presented in Section II. The birth sample is systematically generated by the birth utility differential function. Put differently, the pregnant population is incidently censored, according to the birth utility differential function, by reporting a single value of zero to those women who obtained abortion. It is, therefore, one of an important hypotheses in this study that the self-selected birth sample is unlikely a random sample drawn from the entire population of pregnant women. Therefore, the estimates of prenatal care demand and infant health production function, without taking into account the nature of sample censoring, would be considerably biased and inconsistent. In addition, self-selection bias may also arise with heterogeneity of health endowment and endogeneity of prenatal care in health production function within the birth sample only (Rosenzweig and Schultz 1982, 1983, 1988; Harris 1982;).

4.2 Self-Selection in Derived Demand for Prenatal Care

Rational demand for prenatal care of the current fetus should only pertain to those women who decide to carry the fetus to term. For women in birth sample, an individual household i's preference is described by a birth utility function defined earlier:

$$U^{bi}=U(N_i+1,\ h_i,\ I_i-C_{bi};\ S_i) \tag{11}$$

The household maximizes the total utility subject to a given health production function:

$$h_i=h(m_i,\ x_{hi}\ ;\ e_i) \tag{12}$$

and a vector of budget constraints:

$$(N_i+1)\left[\pi_{mi}m_i+\pi_{hi}x_{hi}+\pi_{fi}\right]\ \leq\ I_i \tag{13}$$

where the π_j's denote the unit prices for health inputs ($j=m_i,x_{hi}$), π_{fi} is an average fixed cost of a child, and I_i is a boundary vector of resources available to individual i.

As usual, the first-order conditions for maximizing the utility function subject to the two constraints simultaneously generate a demand function for infant health (h_i), conditional on the stock of children (N_i) plus 1, and a vector of demand functions for the resources (I_i-C_{bi}) available for other utility arguments.

The demand for infant health, through the health technology, leads household to have a derived demand for health inputs, m_i and x_{hi}, in a reduced form. As stated above, the demand for prenatal care only pertains to the women giving birth with a total number of children resulting at N_i+1. According to the self-selection mechanism for pregnancy resolutions, a woman selecting to give birth is assumed to have had higher utilities of giving a birth than that of obtaining an abortion (that is, $\Delta U_i = U^{bi} - U^{ai} \geq o$). Clearly, this selection mechanism is characterized by women's self preferences and characteristics. Thus, if some important unobservable determinants of the selection mechanism exist but are explicitly omitted, and the estimation of prenatal care demand is made on the non-randomly self-selected birth sample only, then the resulting estimates would be statistically biased. Based on the biased estimates, any inference about the relationship between women's prenatal care demand and its determinants would be obviously unreliable, simply because the self-selected birth sample does not fairly represent the natures of the entire population.

One way to test how the demand for prenatal care on the birth sample omitting abortion women's behavior would be biased is to apply the concept of "wantedness" (Joyce and Grossman 1990). Given the proposition that the more wanted the pregnancy, the more prenatal care is demanded, then failing to consider those women who selected into the abortion sample would bias the estimates of prenatal care demanded in upward direction if giving birth is assumed to be an outcome of wanted pregnancies and obtaining an abortion is considered as an consequence of unwanted pregnancies. Of course, if the last critical assumption is dropped, one cannot arrive at this conclusion.

In particular, consider how the health endowment, cost of contraception, and cost of abortion would distort the estimates of a model disregarding the sample selection problem with the three unobservables. It was pointed out previously that these three variables in principle are crucially involved with women's first decision in pregnancy resolutions but it is fairly hard to observe or control them in practice. Presumably in the timing and demand for quantity and quality of prenatal care, however, women who actually give birth would differ from those with similar "observed characteristics" who obtained abortion had they given birth otherwise. The decisions of pregnancy resolutions, and of how much prenatal care demanded are rationally assumed to be made in a sequential manner. That is, the first decision is whether or not to carry the fetus to term; the second decision is then the initiation of prenatal care for fetus if giving birth is made (Maddala 1985; Brown 1989). Some factors which control women's decision in first stage -- the resolution of pregnancy, would doubtlessly shape the second decision, i.e. the demand for prenatal care.

Abortion cost, for example, affects women's behavior in the first decision of pregnancy resolutions. But, it may also influence the demand for prenatal care in the second stage. Abortion cost was shown as the sum of direct market cost and health cost (both physically and mentally or morally). The health cost component, assumed to be dominant, differs across individuals. Recall the positive association between abortion cost and probability of selecting into birth sample. It is then fairly straightforward to infer that the women with a higher abortion cost would be more likely to end up a higher stock of children (N_i) than women with a lower abortion cost, having everything else equal. In addition, on the basis of the first-order condition of the birth utility function with respect to medical care demand m_t, one can easily show that N_i acts as a negative argument for the derived demand function for health care m_i. The force at work here is the higher the optimal number of children, and the lower the optimal amount of resources allocated to each child (Becker and Lewis 1973; Willis 1973; Easterlin and Crimmins 1985). Therefore, abortion cost would be expected to be negatively associated with prenatal care demand. A biased estimate of demand for prenatal care, based only on women giving birth and no consideration of the self-selection effect from first to second stages, would be obtained if abortion cost differs between women giving birth and those obtaining abortion.

The cost of contraception, p_c, is assumed to be related to the monetary price, time, availability, contraceptive efficiency or knowledge, and more important, the psychic preferences. Empirical evidence indicates that a number of couples consider it as a disutility instrument and therefore refuse to use it (Pollack and Wachter 1975; Henshaw and Silverman 1988). It follows that an increase in the cost of contraception raises the probability of having an unwanted pregnancy. Therefore it is likely to delay initiation of prenatal care or cause less demand for prenatal care for an unwanted fetus, given that a decision to give birth is made. This negative relationship between the price of contraception and the prenatal care demand can also be tested through the impact of contraception on stock of children and allocation of resources to children. A reduction in p_c lowers the stock of children (N_t is smaller) and hence increases the health care as well as other resources allocated to each birth. Since it has been shown that the cost of contraception inversely affects the likelihood of women's selection of giving birth, there is no reason to reject the hypothesis that, on average, the women in abortion cohort would bear higher a contraception cost than the women in birth cohort. If so, a demand for prenatal care estimated with no consideration of women obtaining abortion and self-selection effect would be biased in an upward direction.

The Self-selection bias of prenatal care demand due to the influence of health endowment may be adverse. Rosenzweig and Schultz (1983) argue that it is likely that women who have information on poor health endowments from prior histories of pregnancy complications or of prior birth outcomes reflecting low child health may seek more or earlier prenatal medical care than women with a good health endowment. More evidence of such adverse selection in the use of health inputs can be found in studies by Inman (1976), Eisner et al. (1979) and Harris (1983). Harris (1983) points out that women with previously established high risks or who perceive their babies to be less fit, may initiate care earlier or make more frequent prenatal visits, while those with uneventful past pregnancies may delay care. Recall the role of health endowment in the decision-making of pregnancy resolutions: pregnancies with poor endowments would be less likely carried to term. That is, health endowments of women obtaining an abortion appear to be poorer than those of women actually giving birth. In this case, due to the adverse selection effect associated with health endowment, the prenatal care demand would be under-estimated in this censored birth sample.

Parental reproductive health endowment, however, may create a bias to estimates of demand for prenatal care in a an upward direction. A couple with a better endowment in fecundity, for example, may be exposed to a high probability of conception (Michael 1973). Holding everything else equal, they may be more likely to end up an unwanted pregnancy or an higher stock of births. For a given pregnancy, the probability of choosing an abortion for a fecund couple would be higher than sub-fecund couples. Consequently, had a fecund couple obtaining abortion chosen to carry the fetus to term, they would not demand prenatal care as promptly as sub-fecund couples. This is because the wantedness of a given pregnancy for a fecund couple is likely to be lower than sub-fecund couples. In addition, the behavior of fecund couples' demanding less prenatal care can also be explained by their allocation of resources among their relatively higher stock of children. Overall, the ultimate impact of the above two opposite self-selection behaviors on prenatal care demand would depend upon which one is dominant.

A recent study by Joyce and Grossman (1990) empirically presents a self-selection bias in the demand for early prenatal care. By assuming that only those pregnancies voluntarily terminated are unwanted, they find that if the Black and Hispanic women who aborted had instead given birth, then they would have delayed the initiation of prenatal care, on average, more than three quarters of a month longer than the mean number of months of delay actually observed for those women who had given birth.

In sum, the above discussion reveals that a demand function for prenatal care, estimated with a incidentally censored birth sample, would not be able to capture the true relationship between demand for prenatal care and its determinants, if no consideration of self-selection nature is made. The major reason is that some of the hard-to-measure variables, such as health endowments, the shadow price of abortion, and the cost of contraception play an important role in women's selection decision in first stage, i.e. resolution of pregnancy. The same variables, in turn, may also interfere with the women's second decision -- the demand for prenatal care.

4.3 Self-Selection in Health Production Functions

To analyze self-selection effects in health production, it is very important to be aware of the difference between the effect from the censored nature of birth sample self-selected from the entire population and the effect due to the endogeneity of prenatal care in the health production function within the context of the censored birth sample.

Self-Selection Bias Due to the Censored Nature of Birth Sample

For health production technology, the self-selection bias due to censored birth sample refers to a systematic difference between the estimated marginal products of health inputs, based upon a censored birth sample, and the true counterparts if the relevant information on abortion sample is also taken into account in estimation. Again, health endowment may cause such bias.

Given the marginal impact of health endowment on probability of giving birth to be positive, it is plausible to suggest then that women in the censored birth cohort would be more likely to have better health endowment than women in the abortion cohort. Assuming that a better child health endowment makes a positive contribution to fetal health, given other health related inputs constant, then the true technical product of health care would be doubtlessly over-estimated if only the censored birth sample is employed. Moreover, women's other behavior may also bias the estimation of health technology if the behavior also plays a role in the selection decision of pregnancy resolutions. Suppose some pregnant women decide to abort their fetus mainly because they could not give up smoking, given the knowledge that smoking has a noticeable negative impact on fetal health.[5] Thus, had those smoking women given birth, the technical marginal product of prenatal care in health production function would be potentially lower than that based on censored birth sample only. The outcome, however, might be ambiguous if there exist some other favorably behavioral selections in health inputs which would make considerably high positive contribution to fetal health.

Self-Selection Bias Due to the Endogeneity of Parental Care in Health Production

In addition to the self-selection bias caused by the nature of sample censoring, within in the birth sample, a self-selection problem may also arise if there exist possible endogeneity of prenatal care and heterogeneity of health endowment in health production function. One source of the bias to the true technical relationship between health and its inputs appears to be associated with women's adverse selection behavior in inputs use. Health endowments differ across individuals. If women know, to some extent, their health endowments and then take them into

[5] According to the report by Grossman and Joyce (1990), everything equal, the birth weight of a smoking mother will be 42.4 grams less than that of a non smoking mother.

account in selecting their optimal demand for health inputs, a bias in the estimation of health technology would occur if the endowment variable is not controlled for in a model.

According to Rosenzweig and Schultz (1983), the demand for health care is a behavioral variable depending on health endowment in such manner that women with a poor endowment would seek more or early prenatal care. Consequently the true technical contribution of prenatal care to fetal health (h_m) may be underestimated by an ordinary least square approach (OLS) disregarding the negative impact of poor health endowment on fetal health. This kind of women's adverse selection behavior in input use suggests why epidemiological correlational studies have not always found a beneficial effect of the timing of prenatal care on birth weight (Eisner et al. 1979).[6] Harris (1983) also indicates that birth weight by gestation age has a very strong positive relationship with the initiation of prenatal care but at a diminishing rate. In this study, he also shows that women with high risks or fetal health problems would likely initiate early prenatal care while their infant birth weight, however, tends to be lower due to poor health endowment or possibly shorter gestation. On the other hand, self-selection in health inputs might be favorable to health production. The basic forces that generate the favorable effects are due to a positive correlation between the demand for prenatal care and for other health related inputs or behaviors which would increase the efficacy of prenatal care. A pregnant woman who initiates early care, for example, may simultaneously have a better living environment for the growth of fetus such as eating more nutritious goods, suffering less stress, stopping smoking, narcotics, or other potentially harmful substances, and engaging in appropriate exercises than women who begin prenatal care late. How the overall marginal product of prenatal care would be affected

[6] Probably because of the adverse self-selection effect in health input, in public health and epidemiology fields, the issue of effectiveness of prenatal care is still open at some extent to debate. Some, who extremely question about its positive contribution to the growth of fetus' health, even call for a reduction or termination by government in financing or subsiding the programs for prenatal care and instead argue for more resources devoted to some other health welfare programs such as WIC.

by the two opposite selections, of course, depends upon which one would be dominant.

In sum, health endowments appear to be heterogeneous across individuals. The heterogeneity of health endowments, in turn, would interfere with both the selection of prenatal care demand and the health production as well. Based on a model failing to consider such selection behavior, the true marginal effect of prenatal care on health production would be overestimated in the case of a dominant favorable selection or be under-estimated if adverse selection is dominant.

4.4 An Econometric Model Corrected For Self-Selection Bias

In the applied econometric field, the theory and computational techniques used to deal with the problem of self-selection in an incidentally censored sample have been fruitfully developed and applied to many areas (Tobin 1958; Amemiya 1973, 1981; Hanoch 1976; Maddala 1974, 1985; Hausman and Wise 1977; Heckman 1974, 1976, 1979; and Griliches, Hall, and Hausman 1978; Willis and Rosen 1979; Lee, Maddala, and Trost 1980; Wales and Woodland 1980; Nakosteen and Zimmer 1980; Killingsworth 1983; Nelson 1984; Greene 1983, 1990). The earliest formal application of the self-selection model to demand for prenatal care and health production is made by Grossman and Joyce (1989). Previous discussions have revealed the potential self-selection biases in various aspects of demand for prenatal care and health production functions based on an incidentally censored birth sample. A model controlling for the self-selection biases is presented as follows.

Recall the hypothesis that pregnant women make sequential decisions regarding their pregnancy: the first decision is to make a choice of binary alternatives -- giving birth or having abortion, and the selection mechanism is governed by a utility differential function of giving birth vs. abortion (ΔU_i); the second decision is to select an optimal demand for health inputs, in particular, the derived demand for prenatal care (m_i). The demand for health inputs, in turn, determine infant health outcome through an health production function (h_i).

The model system is written as follows:

$$\Delta U_i = B^u X_{u_i} + \varepsilon_{u_i}(p_{ai}\ ,\ p_{ci}\ ,\ e_i) \geq 0\ ; \tag{14}$$

$$m_i = B^m X_{m_i} + \varepsilon_{mi}(p_{ai}\ ,\ p_{ci}\ ,\ e_i)\ ; \tag{15}$$

$$h_i = B^h X_{h_i} + \beta^h_m m_i + \varepsilon_{hi}(s_i\ ,\ e_i)\ ; \tag{16}$$

where B^u, B^m, B^h are three vectors of coefficients for the three equations, the corresponding variable vectors X_{ui}, X_{mi}, X_{hi} represent systematic components respectively for the three functions; ϵ_{ui}, ϵ_{mi}, ϵ_{hi} denote stochastic terms consisting of unobserved factors, of which three significant unobservables p_{ai}, p_{ci}, e_i are particularly specified (defined as before), and the s_i stands for the hard-to-measure health inputs in the health production function which are assumed to have similar arguments as the demand function for prenatal care. For convenience, the intercepts are suppressed.

The birth utility differential function (equation 14) acts as women's self-selection mechanism in pregnancy resolutions; parental care demand function (equation 15) and health production function (equation 16) are sub-outcomes of the birth utility differentials being greater than zero. That is, equation 14 pertains to full sample $I = I_a + I_b$, while equations 15 and 16 are observed only for women who give birth (I_b).

The conditional expectations of prenatal care and health production functions on birth sample I_b are:

$$\begin{aligned} E(m_i \mid \Delta U \geq 0) &= E(B^m X_{m_i}) + E(\varepsilon_{m_i} \mid \varepsilon_{u_i} < B^u X_{u_i}) \\ &= B^m X_{m_i} + E(\varepsilon_{m_i} \mid \varepsilon_{u_i} < B^u X_{u_i}) \\ &\neq E(m_i), \quad \text{if } cov(\varepsilon_{m_i}, \varepsilon_{u_i}) \neq 0; \end{aligned} \tag{17}$$

$$\begin{aligned} E(h_i \mid \Delta U > 0) &= E(B^h X_{h_i} + \beta_m^h m) + E(\varepsilon_{h_i} \mid \varepsilon_{u_i} < B^u X_{u_i}) \\ &= B^h X_{h_i} + \beta_m^h E(m_i \mid \Delta U \geq 0) + E(\varepsilon_{h_i} \mid \varepsilon_{u_i} < B^u X_{u_i}) \\ &\neq E(h_i), \quad \text{if } cov(\varepsilon_{h_i}, \varepsilon_{u_i}) \neq 0 \end{aligned} \tag{18}$$

From the previous discussion, the three stochastic disturbances ϵ_{ui}, ϵ_{mi}, and ϵ_{hi} are assumed to be correlated with each other. Thus, the expectations conditional on the censored birth sample (sample I_b) for prenatal care demand (m_i) and health production (h_i) would differ from the corresponding unconditional population mean (sample $I = I_a + I_b$). If so, the estimates of model made by OLS approach would be considerably biased. This biasedness with sample selection is reflected by the non-zero truncated conditional means of random disturbances.

In addition, the health production technology also involves another type of biasedness, due to the possible endogeneity of prenatal care m_i and heterogeneity of health endowment e_i in health production. To be specific, assuming the covariance matrix of vector X_h and m_i, $cov(X_h, m_i)$, is zero and m_i is independent from s_i, the partial derivative of equation 16 with respect to variable m_i yields:

$$\frac{\partial h_i}{\partial m_i} = \beta_m^h + \left(\frac{\partial \varepsilon_{hi}}{\partial e_i}\right)\left(\frac{\partial m_i}{\partial e_i}\right)^{-1} \tag{19}$$

As long as the second term is not zero, the estimated marginal product of prenatal care ($\partial h_i / \partial m_i$) would be biased from the corresponding true technical counterpart (β_m^h). Based on the formulation of equation 15, one can verify that:

$$\frac{\partial m_i}{\partial e_i}=\frac{\partial \varepsilon_{mi}}{\partial e_i}=\beta_e^m \neq 0 \tag{20}$$

Here the bias of selection in input arises from the correlation between prenatal care m_i and the unobserved health endowment e_i.

Following Johson and Kotz (1972), Griliches, Hall, and Hausman (1978), assuming a normal distribution for each disturbance term and a joint normal distribution for each pair of them,[7] the incidently truncated conditional mean of disturbance in equation 17 can be shown as:

$$\begin{aligned} &E(\varepsilon_{m_i} \mid \varepsilon_{u_i} < B^u X_u) \\ &= \frac{\sigma_{um}}{\sigma_{ui}^2} E(\varepsilon_{u_i} \mid \varepsilon_{u_i} < B^u X_u) \end{aligned} \tag{21}$$

where σ_{um} is the covariance of ϵ_{ui} and ϵ_{mi}.

For the purpose of computation, standardize the truncated normal variable ϵ_{ui} at $B^u X_u$ by setting a standard normal variable $\omega_{ui} = \epsilon_{ui}/\sigma_{um}$ with a zero mean, thus:

$$\begin{aligned} E(\varepsilon_{ui} \mid \varepsilon_{ui} < B^u X_{ui}) &= \sigma_{ui} E(\omega_{ui} \mid \omega_{ui} < B^{ui} X_{ui}/\sigma_{ui}) \\ &= \frac{\sigma_{ui}}{F(B^u X_{ui}/\sigma_{ui})} \int_{-\infty}^{B^u X_{ui}/\sigma_{ui}} \omega_{ui} f(\omega_{ui}) d\omega_{ui} \\ &= \frac{\sigma_{ui} f(B^u X_{ui}/\sigma_{ui})}{F(B^u X_{ui}/\sigma_{ui})} = \sigma_{ui}\lambda_{ui} \end{aligned} \tag{22}$$

where use is made of the fact that the normal density function is symmetric and $f(-\infty)=0$, $df(\omega)/d\omega=-\omega f(\omega)$, and the λ_{ui} is the inverse of Mills ratio. Substitution of the truncated mean of ϵ_u into equation 19 yields:

---

[7] Following Newey, Powell, and Walker (1990), estimates of the self-selection model with normality are not very sensitive to different parametric assumptions made on error term. Greene (1990) points out, however, that if the underlying disturbances are not normally distributed, the estimators may be inconsistent. Non-normality is an essentially difficult problem in this setting. Research on alternative distributions and estimation approaches is still ongoing. See Duncan (1983, 1986), Kalbfleish and Prentice (1980), Goldberger (1983), Amemiya (1984), Maddala (1986), and Mroz (1987) for more information.

$$E(\varepsilon_{mi} \mid \varepsilon_{ui} < B^{u}X_{ui}) = \frac{\sigma_{um}}{\sigma_{u}} \lambda_{ui} \tag{23}$$

By the same token, the truncated conditional mean of ϵ_{hi} in equation 18 can be written as:

$$E(\varepsilon_{hi} \mid \varepsilon_{ui} < B^{u}X_{ui}) = \frac{\sigma_{uh}}{\sigma_{u}} \lambda_{ui} \tag{24}$$

Note the interesting features of inverse of Mills ratio λ_{ui}.[8] First, the denominator, CDF at B_uX_{ui}/σ_u, simply represents the probability that a pregnant woman selects the option of giving birth:

$$\begin{aligned} F(B^{u}X_{u}/\sigma_{u}) &= P(\epsilon_{u} < B^{u}X_{u}/\sigma_{u}) = P(\varepsilon_{u} < B^{u}X_{u}) \\ &= P(\Delta U > 0) = P(\mathit{Birth} \mid \mathit{Birth}, \mathit{Abortion}) \end{aligned} \tag{25}$$

Second, the λ_{ui} is a monotonically decreasing function of the standard normal variable $\omega_{ui}=B^{u}X_{ui}/\sigma_{u}$ such that:

$$\frac{\partial \lambda_{ui}}{\partial \omega_{ui}} < 0 \; ; \quad \lim_{\omega \to -\infty} \lambda_{ui} = \infty \; ; \quad \lim_{\omega \to \infty} \lambda_{ui} = 0 \; ; \quad \text{where} \quad \omega_{ui} = B^{u}X_{ui}/\sigma_{ui}. \tag{26}$$

Following Heckman (1974, 1976), the self-selection bias of demand function for prenatal care , due to the sample truncation incidently, can be corrected by inserting the λ_{ui} as an additional regressor into the original demand function for prenatal care. To control both sample selection bias as well as endogeneity bias, a two-stage probit methodology is suggested by Lee, Maddala, and Trost (1980). That is, the health production function simultaneously employs the inverse Mills ratio to correct for sample selection and an instrumental variable for prenatal care

[8] For the women who obtained abortions, the inverse of Mill's ratio will be a negative value: $P(\Delta U<0)=P(\varepsilon_{u_i}<-B^{u}_{i}X_{u})$, $E(\varepsilon_{ui} \mid \varepsilon_{ui}<-B^{u}X_{ui}) = -\sigma_{ui}\frac{f(B^{ui}X_{ui}/\sigma_{ui})}{1-F(B^{ui}X_{ui}/\sigma_{ui})} = \sigma_{ui}(-\lambda_{i})$.

to deal with the bias with endogeneity of prenatal care demand. It can be proven that the two-stage probit approach yields consistent estimators.

In practice, the estimates of λu_i is first generated by a probit birth probability model derived from the birth utility differential function. Then, the demand function for prenatal care and the health production function corrected for both two self-selections are formulated as follows:

$$m_i=B^m X_{mi}+\frac{\sigma_{um}}{\sigma_u}\hat{\lambda}_{ui}+v_{mi}\ , \qquad E(v_{mi})=0$$
$$where \quad \frac{\sigma_{um}}{\sigma_u}\lambda_{ui}=\frac{\sigma_{um}}{\sigma_u}\hat{\lambda}_{ui}+v_{mi} \quad cov(v_{mi},\ \varepsilon_{ui})=0 \tag{27}$$

$$h_i=B^h X_{hi}+\beta_m^h m_i+(\frac{\sigma_{uh}}{\sigma_u})\hat{\lambda}_{ui}+v_{hi}\ , \qquad E(v_{hi})=0$$
$$where \quad (\frac{\sigma_{uh}}{\sigma_u})\lambda_{ui}=(\frac{\sigma_{uh}}{\sigma_u})\hat{\lambda}_{ui}+v_{hi}\ , \quad cov(\varepsilon_{ui},\ v_{hi})=0 \tag{28}$$

One can verify that the estimated vectors of B^m, B^h, and β_m^h yielded by the two models are consistent. The coefficients of λ_{ui} are self-defined as covariances σ_{um} and σ_{uh} respectively for prenatal care and health technology, up to a positive scale factor $(1/\sigma_u)$. Technically, however, they would measure the marginal contributions of the inverse Mills ratio to prenatal care demand and health production respectively. The latter leads to a particularly interesting interpretation of signs. Recall the fact that λ_{ui} is a decreasing monotonic function of $B^u X_{ui}$ and that CDF $[F(B^u X_{ui}/\sigma_u)]$ is a monotonically increasing function of $B^u X_{ui}$. The higher the marginal utility of giving birth (more wantedness), the more prenatal care would be demanded if giving birth is selected. For the prenatal care demand equation, thus, a negative coefficient of λ_{ui} ($\sigma_{um} < 0$) implies that holding other factors constant, women with higher probability of giving birth would initiate more or earlier prenatal care than average (i.e., larger $F(B^u X_{ui})$, smaller λ_{ui}, and then more of m_i). If the prenatal care variable is indexed by the number of months a woman delays before initiating parental care, as this study will do later, then a positive coefficient of λ_{ui} (σ_{um} > 0) is expected to be consistent with such interpretation. For the health production function, a negative coefficient of λ_{ui} (σ_{uh} < 0) may be appropriately expected. This expectation is consistent with the hypothesis of favorable selection in input: women with higher utility of giving birth seek early or more prenatal care and simultaneously select more other health related inputs and consequently would be more likely to end up with better infant health than their counterparts with lower probability of selecting into the birth sample (that is, higher $F(B^u X_{ui})$, lower λ_{ui}, and thus greater h_i).

The forces that determine the signs of covariance σ_{um} and σ_{uh}, however, are yielded by a number of unobservables from the disturbances. To explore the nature of these covariances in a simple manner, additive functions for each of the stochastic disturbances ϵ_{ui}, ϵ_{mi}, ϵ_{hi} in the model system are assumed as follows:

$$
\begin{aligned}
\varepsilon_{ui} &= \beta_a^u\, p_{ai} + \beta_c^u\, p_{ci} + \beta_e^u\, e_i \\
\varepsilon_{mi} &= \beta_a^m\, p_{ai} + \beta_c^m\, p_{ci} + \beta_e^m\, e_i \\
\varepsilon_{hi} &= \beta_s^h\, s_i + \beta_e^h\, e_i
\end{aligned}
\tag{29}
$$

Taking a deviation for each equation and multiplying by each other, three pairwise covariances between the disturbance terms result as:

$$
\begin{aligned}
\sigma_{um} &= \beta_a^u \beta_a^m \sigma_{p_a}^2 + \beta_c^u \beta_c^m \sigma_{p_c}^2 + \beta_e^u \beta_e^m \sigma_e^2 \\
\sigma_{uh} &= \beta_a^u \beta_s^h \sigma_{sp_a} + \beta_c^u \beta_s^h \sigma_{sp_c} + \beta_e^u \beta_e^h \sigma_e^2 \\
\sigma_{mh} &= \beta_a^m \beta_s^h \sigma_{sp_a} + \beta_c^m \beta_s^h \sigma_{sp_c} + \beta_e^m \beta_e^h \sigma_e^2
\end{aligned}
\tag{30}
$$

where use is made of the assumptions that p_{ci}, p_{ai}, and e_i are mutually independent and that s_i and e_i are uncorrelated. To be consistent with the data set and analysis that are adopted in this study, from now on, the prenatal care variable is specifically measured by the number of months delayed before initiating prenatal care. Recall the expectations about the signs of the marginal products of p_{ci}, p_{ai}, e_i, and s_i with respect to prenatal care demand and health productions:

$$\beta_a^u>0,\ \beta_c^u<0,\ \beta_e^u>0;\ \beta_a^m>0,\ \beta_c^m>0,\ \beta_e^m>0;\ \beta_s^h>0,\ \beta_e^h>0;\ \sigma_{sp_c}<0,\ \sigma_{sp_a}<0\ .$$

All of these coefficients are involved in determining the signs of σ_{um}, σ_{uh}, and σ_{mh}:

$$\begin{aligned}
\sigma_{um} &= \underset{(+)}{\beta_a^u\beta_a^m\sigma_{p_a}^2} + \underset{(-)}{\beta_c^u\beta_c^m\sigma_{p_c}^2} + \underset{(+)}{\beta_e^u\beta_e^m\sigma_e^2} \\
\sigma_{uh} &= \underset{(-)}{\beta_a^u\beta_s^h\sigma_{sp_a}} + \underset{(+)}{\beta_c^u\beta_s^h\sigma_{sp_c}} + \underset{(+)}{\beta_e^u\beta_e^h\sigma_e^2} \\
\sigma_{mh} &= \underset{(-)}{\beta_a^m\beta_s^h\sigma_{sp_a}} + \underset{(-)}{\beta_c^m\beta_s^h\sigma_{sp_c}} + \underset{(+)}{\beta_e^m\beta_e^h\sigma_e^2}
\end{aligned} \tag{31}$$

Clearly, since in principle the p_{ci}, p_{ai}, e_i, and s_i all vary, the signs of the three covariances would be virtually ambiguous. Nevertheless, it is still possible to shed light, to some extent, on their roles in the determination of the demand for prenatal care and health technology. That is, based on the observed estimates of the three pairwise covariances (up to a positive scale factor), one can in turn identify or infer which of the unobservables would play a dominant role over the others in determining the signs of the covariances.

To be specific, suppose the health endowment e_i and cost of contraception p_{ci} were the same for every body. Then one would end up with $\sigma_{um} > 0$ and $\sigma_{uh} < 0$, as an outcome of considering abortion cost alone. If the actual estimates turn out to support such a sign pattern, then it at least implies that the cost of abortion would be likely dominant over the other variables in shaping a woman's self-selection behavior. Moreover, it also tells that the disturbance in the utility function of giving birth is negatively correlated with that of health production functions while it is positively correlated with that of the prenatal care delay function. This is consistent with the prior interpretation about the nature and the coefficients of the inverse of Mill's ratio in both the demand for prenatal care and in health production functions. That is, as the price of abortion gets higher, it makes selection of birth more likely, but delays initiating prenatal care longer, and consequently tends to decrease health outcome. This may be termed as the **cost of abortion model** (Grossman and Joyce 1989, 1990).

With the same approach, one can formulate a **health endowment model** by assuming that the cost of contraception and cost of abortion are homogeneous across all individuals, emphasized by Rosenzweig and Schultz (1982,1983), Harris (1982), Corman et al. (1987), and Frank et al. (1991). This model, where both $\sigma_{um} > 0$ and $\sigma_{uh} > 0$, is in agreement with a woman's adverse selection behavior in inputs. For instance, a woman with a poor health endowment is more likely to select the abortion option, to seek early or more prenatal care (delay less), and to end up with a lighter birth weight.

Moreover, if there were no variations in the health endowment e_i and cost of abortion p_{ai}, then a **contraception cost model** arises, where $\sigma_{um} < 0$ and $\sigma_{uh} > 0$. This model indicates that a reduction in the contraception price raises the probability of giving birth, lessens the delay of prenatal care, and makes a positive contribution to health production.

In sum, two types of self-selection biases are explored: one is caused by the censored nature of birth sample and the other is due to the endogeneity of prenatal care and heterogeneity of health endowment in the health production function. Theoretically, they are hypothesized to be different, both of them should be controlled simultaneously. Yet a further testing on their difference or causation needs to be done. The Wu test is conducted for this task. Estimation of the model is based on probit two stage method developed by Lee, Maddala, and Trost (1980) and LIMDEP by Greene (1989, 1990).

CHAPTER FIVE

DATA AND ESTIMATION METHODOLOGY

5.1 Data Description

This study uses data on births and abortions collected from the Commonwealth of Virginia in 1984. The original data base contains 84,747 births and 32,606 induced abortions for all residents of Virginia as well as all the births and induced abortions that occurred in Virginia. The records for non-Virginia residents and the observations that are missing for the variables employed in the study are excluded. This results in 78,365 births and 27,650 induced abortions in the data set. Of this modified data set, there are 53,842 births (80.0%) and 13,441 abortions (20.0%) for white adults; 14,582 births (70.0%) and 6,237 abortions (30.0%) for black adults; 5,925 births (51.2%) and 5,637 abortions (48.8%) for white teenagers; 4,016 births (63.2%) and 2,335 abortions (36.8%) for black teenagers. Following the work by Hofferth and Hayes (1987), Leibowitz, Eisen, and Chow (1986), and Serrato (1989), the adolescent status such as education, age, and marriage may be simultaneously determined with the decision-making of giving birth and initiation of prenatal care, while the determination of these variables are believed to be less relevant to the decision-making in pregnancy resolutions for adults. Therefore, to minimize the endogeneity problem with these variables, teenagers are excluded in the present study.[9] The estimation for blacks is made on full records while for white adults it is based on a sub-sample randomly drawn from the corresponding combined population (67,283) of births and abortions. The random sub-sample ends up with 16,091 births and 3,968 abortions for white adults. The adults distributions of birth and abortion by race and residence are presented in the following Table 5-1.

[9] A study for the teenagers will be conducted in another project, a continuation of the present.

TABLE 5-1
Birth/Abortion Distributions
By Race and Residence

Options	Whites		Blacks		Total
	Urban	Rural	Urban	Rural	
Birth	10,878 (78.5)*	5,213 (84.1)	9,059 (67)	5,523 (75.8)	30,673 (75)
Abort	2,985 (21.5)	983 (15.9)	4,469 (33)	1,768 (24.2)	10,205 (25)
Total	13,863	6,196	13,528	7,291	40,878

* The figures in parenthesis are percentages by columns.

The data set is augmented with some county-specific socioeconomic variables from the Area Resource File (ARF) 1984. This augmentation makes it possible to test if income variables such as race-specific poverty line or medium family income influence women's decision-making in fertility planning and demand for prenatal care. Other variables which are incorporated with the major data set include the Supplemental Food Program for Women, Infants, and Children (WIC) from WIC Local Agency Directory 1986, published by the United States Department of Agriculture; the number of abortion providers in 1984 (categorized into two types: 30-390 and over 400), number of family planning clinics in 1983, and total females aged 15 to 44 in 1984, provided by Alan Guttmacher Institute (AGI). To be comparable, the absolute number of abortion providers and family planning clinics as well as WIC centers are divided by county-specific female populations aged 15-44 times 10,000. Thus, the study eventually employs the county-specific number of WIC centers, abortion providers, and family planning clinics per 10,000 female aged 15-44 for estimation. All the augmentations are made by county.[10] Moreover, all counties are sub-grouped as urban or rural areas, according to the classification

[10] Since all the independent cities in Virginia are treated as equivalent to counties, this study, through consistently coding among different data sets, finally ends up with 142 modified counties which might not be the same as official total number of counties in Virginia.

Table 5-2
Definition of Variables

Name of Variables	Descriptions
Birth Weight	Infant birth weight in grams
Prenatal Care Delay	Number of months delayed before seeking prenatal care
Terminations < 20W	Number of previous abortions before the 20th week of
Terminations > 20W	Number of previous abortions after the 20th week of gestation
Parity Alive	Number of previous living births who are still alive
Parity Dead	Number of previous living births who have died
WIC Centers	WIC centers per 10,000 females aged 15-44 by county
Abortion Providers	Abortion providers per 10,000 females aged 15-44 by county
Family Planning Clinics	Family planning clinics per 10,000 females aged 15-44 by
White Below Poverty (%)	Percentage of the whites below the poverty line in 1979
Black Below Poverty (%)	Percentage of the blacks below the poverty line in 1979
Male	Dummy variable that equals 1 if the infant is male
Birthdummy	Dummy variable that equals 1 if the women selects giving birth
Illegitimacy	Dummy variable that equals 1 if the woman is not married
Urban	Dummy variable that equals 1 if the woman lives in an urban
Mother Age 35-39	Dummy variable that equals 1 if the woman is 35-39 years of
Mother Age > 40	Dummy variable that equals 1 if the woman is 40 years or
Mother Education 10-12	Dummy variable that equals 1 if the woman has 10-12 years of
Mother College	Dummy variable that equals 1 if the women has college
Father Education 10-12	Dummy variable that equals 1 if the father has 10-12 years of
Father College	Dummy variable that equals 1 if the father has college
λ	The inverse of Mill's ratio

of metropolitan counties from State and Metropolitan Area Data Book 1986 by the U.S. Department of Commerce, Bureau of the Census. Race and area of residence are thought likely to interact with a number of explanatory variables. They may also be endogenously determined in the model. In order to minimize these possible problems of multicollinearity or endogeneity and also to test how the model would differ for different races and residences, this study conducts race-area specific estimations of the model respectively for four cohorts: urban whites, rural whites, urban blacks, and rural blacks.

5.2 Probability Function of Giving Birth

Through probit maximizing likelihood approach, a probability function of giving birth is derived from the random utility differential function, acting as women's self-selection mechanism in pregnancy resolutions. The random birth utility differential function is treated as a reduced form consisting of a deterministic component ($B^u X_{ui}$) and a random component ϵ_{ui}. Based on the previous discussion, the basic arguments in the deterministic component are specified as the number of abortion centers per 10,000 females aged 15-44, the number of family planning clinics per 10,000 females aged 15-44, race-specific percentage below poverty, previous early abortions within 20 weeks, mother's education (less than high school, high school, and college), mother's age (20-34, 35-39, and over 40), and illegitimacy (marital status).

Recall how the mechanism of self-selection in pregnancy resolutions works: a pregnant woman selects the option of giving birth provided the utility differential is greater than or equal to zero and selects the option of abortion otherwise. Thus, women who are in the birth sample would have had a positive birth utility differential when the decision was made while women who are in abortion sample should have had a negative birth utility differential. Denoting D_i as a random dummy variable with the value of one for a woman giving birth and zero for a woman obtaining abortion, a probit model of giving birth is derived as follows:

Selection Mechanism

$$D_i = 1 \quad if \quad \Delta U = B^u X_{ui} + \varepsilon_{ui} \geq 0$$
$$D_i = 0 \quad if \quad \Delta U = B^u X_{ui} + \varepsilon_{ui} < 0 \qquad (32)$$

$$Prob\ (D_i = 1) = F(\phi_i)$$
$$Prob\ (D_i = 0) = 1 - F(\phi_i) \qquad (33)$$

Probit Likelihood Function

$$L = Prob(D_1 = d_1,\ D_2 = d_2,\ ....D_I = d_I)$$
$$= \prod_{i=1}^{I} [F(\phi_i)]^{d_i}\ [1 - F(\phi_i)]^{1-di} \qquad (34)$$

where the $\phi_i = B^{ui}X_{ui}/\sigma_{ui}$, $F(\phi_i)$ represents the probability of selecting birth for a pregnant woman with the characteristic X_i. Maximization of this probit likelihood function yields the estimates of coefficient vectors β^u, μ_{ui}, and σ_{ui}. The inverse of Mills ratio, $\lambda_{ui} = f(B^u X_{ui}/\sigma_u)/F(B^u X_{ui}/\sigma_u)$, is also generated by estimating the model as well.

5.3 Prenatal Care Delay and Birth Weight Production Functions

Demand for prenatal care is measured by the number of months a woman delays seeking the prenatal care for her pregnancy. Women who received no prenatal care are treated as having ten months delay. Prenatal care delay is simply a negative correlate of demand for prenatal care. The function for prenatal care delay is specified as a linear reduced form with the explanatory variables of mother's age, mother's education, and illegitimacy. It also takes into account the number of WIC centers per 10,000 females aged 15-44, race-specific percentage of people below the poverty line, previous births still living (parity), and previous late abortions (after 20 weeks). In addition, since the prenatal care delay pertains to women who give birth and most of them are married, it is interesting to test how husbands would play a role in the initiation of prenatal care.

The hypothesis is that initiation of prenatal care would be a likely outcome of the decision jointly made by both wife and husband. To test this hypothesis, educational dummy variables for fathers are included and are also categorized by the same three levels for mother's education. The difference between unmarried women and married women with husbands, on the other hand, is in part captured by a dummy variable of illegitimacy (that equals 0 for married women and 1 for single women). The inverse of Mill's ratio (λ_{ui}) is inserted into the model to control for the self-selection effect with birth sample censoring. The coefficient of λ_{ui} represents the covariance of the two random disturbances in birth probability function and prenatal care delay function, up to a positive scale factor, $1/\sigma_{ui}$.

Birth weight is employed as an indicator of infant health, h_i. Birth weight has been repeatedly found to be a critical determinant of perinatal survival or infant mortality, and of subsequent child growth and development in an extensive medical literature as well as in the economics of health and demography (Beck and van den Berg 1975; Chernichovsky and Coate 1979; Eisner et al. 1979; Niswander, Gordon et al. 1972; Harris 1982; Rosenzweig and Schultz 1983; Corman et al. 1987; Grossman and Joyce 1990; Frank et al. 1991). According to Rosenzweig and Schultz (1984), during 1964-1965 in the U.S.A., 18% of infants weighing less than 2,500 grams did not reach their first birthday, while only 0.97 % of the infants weighing more than 2,500 grams died in the first year. They also point out that birth weight is not strongly determined by geneticity but is significantly associated with socioeconomic or demographic characteristics such as mother's education, income, parities, prenatal care, and age. Taking the data from Massachusetts in 1969-1970, Harris (1984) also shows a dramatic decline in the trend of neonatal mortality rate per 1,000 infants as birth weight increases up to 2,500 grams.

The birth weight production is specified as a structural function of endogenous prenatal care delay and other socioeconomic variables as well as biological or demographic factors. To be specific, it includes previous living births who have died, previous living births still living, total previous late abortions (after 20 weeks), child's sex, mother's educational level, mother's age, as well as the inverse of Mills Ratio, λ_{ui}, yielded by the birth selection model. Treating prenatal care delay as only one endogenous variable in the birth weight function does not imply, however, that other arguments are not possibly determined by the model. For example, variables such as mother's age, education, and parity could be simultaneously determined with the decision-making in pregnancy resolutions and demand for prenatal care and other health inputs, which would in turn endogenously influence birth weight production (Rosenzweig and Schultz 1983). Treating prenatal care demand as only one endogenous variable in the birth weight function provides an abstraction advantage to focus on the test of self-selection effect of demand for prenatal care on health production. Moreover, both prenatal care demand and birth weight production functions are estimated to ascertain whether the cost of contraception, the cost of abortion, or the health endowment is a dominant unobserved determinant of reproductive outcomes. The estimation of other endogenous variables is not essential in accomplishing this goal.

5.4 Identification and Marginal Effects of Determinants

As shown earlier, the functions for birth utility differential, demand for prenatal care, and the birth weight production are formulated as a recursive model. The birth utility differential function serves as a sample selection mechanism incidently censoring pregnant population into either birth sample or abortion sample. The last two functions pertain to the birth sample only and prenatal care, as an endogenous variable, enters into birth weight production function. Identification problems arise as random disturbances in these functions are interactively correlated. According to econometric theory (Gujarati 1988), in a model system with three equations and three endogenous variables, order condition for identification needs less than two (at most two) exogenous variable(s) to be excluded from each equation. This condition is obviously met by all the three equations (there are 6, 4, and 9 exogenous variables respectively excluded from equations of birth probability, demand for prenatal care, and birth weight production). With regard to the rank condition, it requires one non-zero determinant of order 2×2 constructed from coefficients of the variables excluded from each equation. One can verify that all the three equations do meet the rank condition as well.

Based on the functions of prenatal care demand and birth weight production controlled for self-selection in sample censoring by including the inverse of Mills ratio, λ_{ui}, the marginal product of a particular argument in vector of X_{hi} or X_{mi} is not just the corresponding coefficient β_i^h ($=\partial h_i/\partial x_{hi}$) or β^m_i ($=\partial m_i/\partial x_{mi}$) if the particular independent variable x_i also appears in birth probability function.

Recall the prenatal care demand equation:

$$m_i = B^m X_m + \frac{\sigma_{um}}{\sigma_u}\hat{\lambda}_i + v_{m_i}\,, \qquad E(v_{mi})=0 \tag{35}$$

Following Greene (1990), the partial derivative of m_i with respect to a particular independent variable x_{mi} is:

$$\frac{\partial E(m_i \mid \varepsilon_{ui} \geq -B^u X_m)}{\partial x_{mi}} = \beta_i^m - \beta_i^u(\sigma_{um}/\sigma_u^2)(\lambda_i^2 + \lambda_i \phi) \tag{36}$$

where the ϕ_i is as defined as earlier. β_i^m measures direct effect of x_i on the mean of prenatal care $E(m_i)$ for a typical pregnant woman in the population while the second part of the equation stands for indirect influence of x_i on the mean of m through its presence in the λ_i. For example, if $\sigma_{um} > 0$, the second part of the partial derivative serves to reduce the marginal product of x_i provided

x_i makes positive contributions in selection of giving birth ($\beta_i^u > 0$) or to increase the marginal product of x_i if $\beta_i^u < 0$. $\sigma_{um} > 0$ also implies that the mean of prenatal care in birth sample appears to be higher than the population mean of m.[11]

For birth weight production, the marginal effect of a regressor x_{hi} consists of three components if the regressor, at the same time, appears in both prenatal care demand and selection function of giving birth as well:

$$\frac{\partial E(h_i \mid \varepsilon_{ui} \geq -B^u X_h)}{\partial x_{hi}} = \beta_i^h - \beta_i^u(\sigma_{um}/\sigma_u^2)(\lambda_i^2+\lambda_i\phi_i) - \beta_i^u(\sigma_{uh}/\sigma_h^2)(\lambda_i^2+\lambda_i\phi_i) \tag{37}$$
$$= \beta_i^h - \beta_i^u(\lambda_i^2+\lambda_i\phi_i)[(\sigma_{um}+\sigma_{uh})/\sigma_u^2]$$

Here β_i^u stands for the direct technical contribution of x_{hi} to infant health for every one in the population. The second component measures additional impact of x_{hi} specific to those women in birth sample alone. In particular, it indicates that the total marginal effect of the argument x_{hi} on birth weight production interacts with direct marginal product of x_{hi} (β_i^h), the behavior of self-selection in giving birth (β_i^u), the correlation between its own disturbance and the one in the birth probability function (σ_{uh}), and the association between disturbances in prenatal care and in birth probability function (σ_{um}). The last term, σ_{um}, reflects the self-selection effect with the endogeneity of prenatal care in health production.

5.5 Expected Self-Selection Effect

On the basis of equations for prenatal care and health production corrected for selection biases, self-selection effect on both equations can be gauged by comparing the observed means of care and birth weight to the corresponding means for those who aborted had they chosen to give birth or to the means of entire population:

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[11] Similar interpretations of the marginal effects for labor can be found in the work by Killingsworth (1984).

$$\bar{M}_b = B^m \bar{X}_{bm} + \frac{\sigma_{um}}{\sigma_u}\bar{\lambda}_b \tag{38}$$

$$\bar{M}_a = B^m \bar{X}_{am} + \frac{\sigma_{um}}{\sigma_u}\bar{\lambda}_a \tag{39}$$

where the sub index b stands for women in birth sample and the sub index a for women in abortion sample. The expected mean of prenatal care delay for a typical pregnant woman with characteristics $\bar{X}_i$ (i=b, a) would be $E(M_i)=B^m\bar{X}_{mi}$. Self-selection effect, due to sample censoring, makes the two cohorts to have a biased mean of prenatal care compared to the expected mean of it by $(\sigma_{um}/\sigma_u)\bar{\lambda}_i$. Between the two cohorts, a potential difference in prenatal care delay controlling for other characteristics ($X_{bm}=X_{am}$) is measured by the term of $\sigma_{um}/\sigma_u(\bar{\lambda}_b-\bar{\lambda}_a)$.

By the same token, one can also gauge the self-selection effects on infant health production. The sample specific means of birth weight can be illustrated as follows:

$$\bar{H}_b = B^h \bar{X}_{hb} + \beta^h_m \bar{M}_b + (\frac{\sigma_{uh}}{\sigma_u})\bar{\lambda}_{ub} \tag{40}$$

$$\bar{H}_a = B^h \bar{X}_{ha} + \beta^h_m \bar{M}_a + (\frac{\sigma_{uh}}{\sigma_u})\bar{\lambda}_{ua} \tag{41}$$

where the H_i (i=b, a) represents the means of birth weight respectively for women giving birth or those who aborted had they selected to give a birth. The last term in both equations simply shows the self-selection bias, due to the sample censoring, on the mean of birth weight for a typical pregnant woman characterized by X_{hi} (i=b, a). If women's behavioral selection (adverse selection or favorable selection), due to endogeneity of M_i, is taken into account, then the gross self-selection biases would be sum of the two effects: $\bar{\lambda}_i[\beta_m^h(\sigma_{um}/\sigma_u)+(\sigma_{uh}/\sigma_u)]$. Subtracting equation 41 from equation 40 describes how women giving birth would differ, with respect to birth weight outcomes, from women selecting abortion. Of course, one can accomplish the goal by either allowing M_{mi} and X_{hi} to vary or controlling them at the same level.

CHAPTER SIX

EMPIRICAL PERFORMANCE OF THE MODEL CORRECTED FOR SELF-SELECTIONS

The maximum likelihood estimates for the probit model predicting the probabilities that pregnant women select into birth sample are given in Table 6-8 for white adults and Table 6-9 for black adults. Tables 6-10 and 6-11 present estimates for race specific prenatal care delay functions. To test the significance of self-selection bias due to sample censoring, each function has been estimated by both OLS and Heckman's two-stage procedure. Birth weight production functions are presented in Tables 6-12, 6-13, 6-14, and 6-15 for urban whites, rural whites, urban blacks, and rural blacks respectively. For each cohort, four different estimates yielded by OLS, OLS-Select, 2SLS, and 2SLS-Select are provided in appropriately marked columns.

6.1 Findings for Probability Functions of Giving Birth

The probability function of giving birth is estimated using a probit likelihood model. The estimation is made on 20,059 white adults and 20,819 black adults. Among the pregnant whites adults, 21% of those residing in urban areas obtain a abortion while only 16% of others living in rural areas abort. For black pregnant adults, the abortion ratio of urban women is 33% and 24% for rural cohort. The relatively low abortion ratio for women residing in rural areas for both whites and blacks, compared to those in urban areas, may reflect the fact of less access to abortion services (Shelton 1976; Sullivan et al. 1977; Griner and Trent 1987).

The results from estimation quite strongly support prior expectations about the role of most variables influencing women's selection decision in pregnancy resolutions. In particular, women's educational levels are measured by three dummy variables: less than high school, high school, and college. It is found that the educational level serves to reduce the probability of giving birth with an increasing rate across all cohorts. A examination of the sample means of mother's college education by race also gives evidence for such educational impact on pregnancy resolutions. For women giving birth, 49% of whites and 30% of blacks have college education whereas, for women obtaining an abortion, 51% of whites and 42% blacks are educated at college level. This finding is consistent with many previous studies dealing with the similar issue (Michael 1973; Easterlin and Crimmins 1985; Leibowitz et al. 1986; Grossman and Joyce 1990; Serrato 1989; and Heckman 1990). For example, Easterlin and Crimmins (1985) show that education would lower the costs of contraception by providing more or better information on various means of fertility control which would, in turn, increase one's adoption of deliberate fertility control (including induced abortion) and its efficacy as well. Moreover, education tends to reduce the demand for children by shifting tastes of children in an unfavorable manner and decreasing the prices of other goods relative to children. They argue that tastes of children and life style with them are essentially old goods, while education presents images of new life styles competitive to children. More educated households may also put greater emphasis on the quality of children at the expense of quantity of children. As a consequence, the more educated would then substitute quality for quantity of children. Finally, since education is often interpreted as proxy for a woman's wage, the negative association between education and probability of giving birth may indicate the notion of the opportunity cost of having a baby with respect to education.

Mother's age is another significant determinant of decision making in giving birth. When compared to women aged 20 to 34, those aged between 35 to 39 and over 40 are more likely to select the option of abortion. This finding may basically reflect the influence of risk from higher ages on pregnancy, fetal health growth, and delivery. It also indicates the notion of diminishing marginal utility of having an additional baby, holding everything else constant, since age is positively correlated with the stock of children. Moreover, in magnitude, the age coefficient for

40 and over is two or three times larger than the coefficient for age between 35-39. Descriptive statistics in Table 6-1 also shows a fairly low distribution of women giving birth in high age groups: 7.3 % pregnant whites and 4.4 % pregnant blacks aged 35 to 39; for women aged over 40, the ratio of birth is even lower to 0.8 % for whites and 0.7 % for blacks.

The probability of giving birth shows a strong inverse dependence on previous early abortions (less than 20 weeks). Such finding is highly significant across all cohorts. In general, an induced abortion would likely occur before 20 weeks. The variable, termination < 20 weeks, can be treated as a good proxy for induced abortion. Thus, the result reveals that pregnant women with a higher number of previous early abortions would have a higher likelihood of selecting an abortion than those with lower number of previous early abortions. Simple descriptive statistics present a consistency with this finding. All pregnant white women, for example, have a mean of 0.4 terminations within 20 weeks whereas the sub mean for white women obtaining abortion is 0.61. For blacks, all pregnant women have a mean of 0.50 for the same variable whereas, for those who abort, it is 0.69. Following Tietz (1978) and Cates (1984), several reasons can be made to explain this role of previous induced abortion in the selecting the option of abortion: 1) women who have already had abortions are generally in their prime reproductive years; 2) they are demonstrably sexually active; 3) they have proven fecundity; 4) they are willing to accept an induced abortion as a way of preventing an unwanted pregnancy; and 5) they have demonstrated their ability to overcome barriers that may deter other women from obtaining abortion services. Put differently, a woman with a large number of previous induced abortions may bear a high cost of contraception or low abortion cost, and hence she may be more likely to select abortion option.

As expected, illegitimacy has an extremely negative influence on selection of giving birth, in both magnitude as well as in significance level. For whites, 80% of pregnant women are married while 93 % of those giving birth are married. Overall, the rate of legitimacy for blacks is low relative to whites, but there still exits a considerable difference between all black pregnant women (42%) and those giving birth (52%). Grossman and Joyce (1990) also arrive at a similar

TABLE 6-1
Means and Standard Deviations of Key Variables for Race-Specific Pregnant Women

Variables	Whites	Blacks
Abortion Providers	0.2779 (0.3144)*	0.3784 (0.3822)
Family Planning Clinics	1.1123 (1.3117)	1.3107 (1.4692)
Below Poverty (%)	7.8071 (3.7835)	24.7371 (6.5232)
Terminations > 20W	0.0398 (0.2515)	0.0617 (0.2882)
Terminations < 20W	0.4019 (0.7388)	0.5019 (0.8219)
Mother Education 10-12	0.4685 (0.4990)	0.6023 (0.4894)
Mother College (> 12)	0.4893 (0.4999)	0.3359 (0.4723)
Father Education 10-12	0.3181 (0.4657)	0.2250 (0.4176)
Father College (> 12)	0.3956 (0.4890)	0.1204 (0.3255)
Mother Age 35 - 39	0.0706 (0.2561)	0.0466 (0.2108)
Mother Age > 40	0.0105 (0.1020)	0.0088 (0.0941)
legitimacy	0.7966 (0.4025)	0.4217 (0.4938)
Observations	20059	20819

*Numbers in parenthesis are deviations.

conclusion. Referring to the components of birth cost defined earlier, this finding should not be surprising since it simply confirms a fact that an unmarried woman would bear higher social as well as financial costs of giving a birth than a married woman (Torres and Forrest 1988).

TABLE 6-2
Means and Standard Deviations of Key Variables for Pregnant Women By Race and Residence

Variables	Whites		Blacks	
	Urban	Rural	Urban	Rural
Abortion Providers	0.3674 (0.3174)*	0.0778 (0.1913)	0.5128 (0.3785)	0.1287 (0.2356)
Family Planning Clinics	0.7424 (0.6370)	1.9401 (1.9160)	0.9007 (0.7249)	2.0715 (2.0732)
Below Poverty (%)	6.1258 (2.6389)	11.5463 (3.2453)	21.8452 (5.0330)	30.1028 (5.4954)
Terminations > 20W	0.0400 (0.2586)	0.0394 (0.2351)	0.0636 (0.2867)	0.0580 (0.2908)
Terminations < 20W	0.4349 (0.7585)	0.3281 (0.6870)	0.5571 (0.8570)	0.3995 (0.7420)
Mother Education 10-12	0.4279 (0.4978)	0.5592 (0.4965)	0.5835 (0.4930)	0.6372 (0.4808)
Mother College (> 12)	0.5439 (0.4981)	0.3670 (0.4820)	0.3693 (0.4826)	0.2740 (0.4461)
Father Education 10-12	0.2685 (0.4432)	0.4290 (0.4950)	0.2096 (0.4070)	0.2536 (0.4351)
Father College (> 12)	0.4488 (0.4974)	0.2765 (0.4473)	0.1389 (0.3459)	0.0861 (0.2804)
Mother Age 35 - 39	0.0791 (0.2698)	0.0516 (0.2213)	0.0495 (0.2170)	0.0411 (0.1986)
Mother Age > 40	0.0109 (0.1038)	0.0097 (0.0979)	0.0088 (0.0934)	0.0092 (0.0954)
legitimacy	0.7860 (0.4102)	0.8205 (0.3838)	0.4223 (0.4939)	0.4207 (0.4937)
Observations	13863	6196	13528	7291

*Numbers in parenthesis are deviations.

Of the county-specific socioeconomic factors incorporated in the model, most of these aggregate variables function as expected in resolving a pregnancy. The number of abortion providers by county, for example, is assumed to reflect the availability of obtaining abortion acting as a negative correlate of abortion cost. The estimates support a prior hypothesis that a higher number of abortion providers would lower the probability for a woman to select giving

TABLE 6-3
Means and Standard Deviations of Key Variables for Race-Specific Women Giving Birth

Variables	Whites	Blacks
Birth Weight	3436.028 (577.850)	3159.027 (640.700)
Prenatal Care Delay	2.3717 (1.2679)	2.8944 (1.5549)
Male	0.5186 (0.4997)	0.5093 (0.4999)
WIC Center	0.2589 (0.5081)	0.3101 (0.6220)
Parity Alive	0.8404 (0.9848)	1.1755 (1.2198)
Parity Dead	0.0212 (0.1556)	0.0328 (0.2039)
Abortion Providers	0.2637 (0.3088)	0.3541 (0.3769)
Family Planning Clinics	1.1515 (1.3531)	1.3742 (1.5397)
Below Poverty (%)	7.8664 (3.8332)	24.8773 (6.6780)
Terminations > 20W	0.0186 (0.1546)	0.0316 (0.2069)
Terminations < 20W	0.3554 (0.7020)	0.4197 (0.7593)
Mother Education 10-12	0.4695 (0.4991)	0.6235 (0.4845)
Mother College (> 12)	0.4852 (0.4998)	0.2994 (0.4580)
Father Education 10-12	0.3965 (0.4892)	0.3212 (0.4670)
Father College (> 12)	0.4931 (0.5000)	0.1719 (0.3773)
Mother Age 35 - 39	0.0726 (0.2595)	0.0443 (0.2058)
Mother Age > 40	0.0080 (0.0888)	0.0072 (0.0846)
legitimacy	0.9328 (0.2504)	0.5232 (0.4995)
Observations	16091	14582

*Numbers in parenthesis are deviations.

TABLE 6-4
Means and Standard Deviations of Key Variables for Race-Specific Women Obtaining Abortions

Variables	Whites	Blacks
Birth Weight	-	-
	-	-
Prenatal Care Delay	-	-
	-	-
Male	-	-
	-	-
WIC Center	0.2796	0.3067
	(0.4760)	(0.5139)
Parity Alive	0.7112	1.1405
	(1.0030)	(1.1471
Parity Dead	0.0136	0.0152
	(0.2138)	(0.1373)
Abortion Providers	0.3355	0.4349
	(0.3299)	(0.3887)
Family Planning Clinics	0.9534	1.1623
	(1.1140)	(1.2771)
Below Poverty (%)	7.5660	24.4093
	(3.5650)	(6.1342)
Terminations > 20W	0.1258	0.1320
	(0.4624)	(0.4124)
Terminations < 20W	0.5905	0.6942
	(0.8466)	(0.9243)
Mother Education 10-12	0.4645	0.5527
	(0.4988)	(0.4973)
Mother College (> 12)	0.5058	0.4214
	(0.5000)	(0.4938)
Father Education 10-12	-	-
	-	-
Father College (> 12)	-	-
	-	-
Mother Age 35 - 39	0.0625	0.0519
	(0.2421)	(0.2219)
Mother Age > 40	0.0209	0.0130
	(0.1431)	(0.1132)
legitimacy	0.2447	0.1844
	(0.4300)	(0.3878)
Observations	3968	6237

*Numbers in parenthesis are deviations.

TABLE 6-5
Means and Standard Deviations of Key Variables for White Women Giving Birth by Residence

Variables	W/Urban	W/Rural
Birth Weight	3451.151 (573.773)	3404.471 (585.062)
Prenatal Care Delay	2.2880 (1.2316)	2.5463 (1.3237)
Male	0.5174 (0.4997)	0.5210 (0.4996)
WIC Center	0.2493 (0.3100)	0.2789 (0.7719)
Parity Alive	0.8171 (0.9691)	0.8891 (1.0150)
Parity Dead	0.0215 (0.1596)	0.0205 (0.1471)
Abortion Providers	0.3534 (0.3130)	0.0766 (0.1950)
Family Planning Clinics	0.7488 (0.6490)	1.9918 (1.9310)
Below Poverty (%)	6.1118 (2.6834)	11.5280 (3.2402)
Terminations > 20W	0.0162 (0.1400)	0.0238 (0.1812)
Terminations < 20W	0.3862 (0.7269)	0.2912 (0.6423)
Mother Education 10-12	0.4186 (0.4934)	0.5755 (0.4943)
Mother College (> 12)	0.5527 (0.4972)	0.3443 (0.4752)
Father Education 10-12	0.3422 (0.4745)	0.5099 (0.5000)
Father College (> 12)	0.5720 (0.4948)	0.3286 (0.4697)
Mother Age 35 - 39	0.0817 (0.2740)	0.0535 (0.2251)
Mother Age > 40	0.0087 (0.0930)	0.0063 (0.0793)
legitimacy	0.9388 (0.2398)	0.9202 (0.2710)
Observations	10878	5213

*Numbers in parenthesis are deviations.

TABLE 6-6
Means and Standard Deviations of Key Variables for Black Women Giving Birth by Residence

Variables	B/Urban	B/Rural
Birth Weight	3163.87 (637.926)	3151.072 (645.203)
Prenatal Care Delay	2.7940 (1.5212)	3.0590 (1.5951)
Male	0.5091 (0.4000)	0.5095 (0.5000)
WIC Center	0.2776 (0.2746)	0.3633 (0.9451)
Parity Alive	1.1173 (1.1883)	1.2709 (1.2894)
Parity Dead	0.0315 (0.1966)	0.0349 (0.2154)
Abortion Providers	0.4929 (0.3788)	0.1263 (0.2369)
Family Planning Clinics	0.9220 (0.7656)	2.1159 (2.1007)
Below Poverty (%)	21.7172 (5.1472)	30.0606 (5.5716)
Terminations > 20W	0.0306 (0.1962)	0.0333 90.2235)
Terminations < 20W	0.4750 (0.7996)	0.3290 (0.6786)
Mother Education 10-12	0.6048 (0.4889)	0.6542 (0.4757)
Mother College (> 12)	0.3358 (0.4723)	0.2397 (0.4270)
Father Education 10-12	0.3129 (0.4637)	0.3348 (0.4720)
Father College (> 12)	0.2074 (0.4055)	0.1137 (0.3175)
Mother Age 35 - 39	0.0471 (0.2119)	0.0397 (0.1952)
Mother Age > 40	0.0066 (0.0811)	0.0081 (0.0899)
legitimacy	0.5411 (0.4983)	0.4939 (0.5000)
Observations	9059	5523

*Numbers in parenthesis are deviations.

birth. The mean value of abortion providers across different cohorts also yield evidence for such finding (0.26 for whites and 0.35 for blacks in the birth sample vs. 0.36 for whites and 0.43 for blacks in the abortion sample). The number of abortion providers, however, is statistically significant only for urban whites although it has a correct sign for all other cohorts. This may be because the variable is of several shortcomings. First, it is augmented with the main individual data sets from another aggregate data source (county specific). Second, Independent cities in Virginia are treated as counties and the coding for such treatment might not be perfect. Third, the border crossing between Virginia and Washington D.C. may influence the availability and therefore the selection of abortion services for residents in Virginia.

Evidence is found for family planning clinics' positive role in selecting the option of giving birth. That is, women living in an area with more family planning clinics available are more likely to carry their fetus to term than those living in counties with less family planning clinics. This is defensive on the grounds that the availability of family clinics may serve as a negative correlate of costs of contraception or giving a birth. Following Easterlin and Crimmins (1985), family planning programs chiefly affect fertility regulation costs -- the time and money required to obtain information about methods and contraception supplies. Moreover, it may also reduce the subjective drawbacks associated with the adoption of contraception. These influences on access and attitudes toward contraception, in turn, would increase the wantedness of a given pregnancy which is critically associated with selection of giving birth (Klerman and Jekel 1984; Joyce and Grossman 1990). The means of family planning clinics also give favorable evidence that women giving birth are more likely those residing in the counties with more family planning clinics (1.15 for whites and 1.37 for blacks) than those obtaining abortions (0.95 for whites and 1.16 for blacks). No significant evidence is found for the impact of the county-specific poverty variable on its residents' pregnancy resolutions. This may be because the explanatory power of poverty, as an income variable, may be reduced due to its possible multicollinearity with other variables such as education, in addition to the reasons given to abortion providers above. Nevertheless, it is still kept in the model on a belief that the dangers of omitting it may be greater than those stemming from either measurement error or multicollinearity.

6.2 Findings For Prenatal Care Demand Functions

Prenatal care demand functions are estimated by the two-step procedure suggested by Heckman (1976, 1979). Tables 6-10 and 6-11 respectively report the estimates for white and black cohorts. For comparison, OLS estimates are also provided in the two tables. 10,878 urban whites, 5,213 rural whites, 9,059 urban blacks, and 5,523 rural blacks are examined. Evidence for a significant self-selection bias is found in urban whites and rural blacks (at 5% significance level). Three positive coefficients of the inverse of Mills ratio are obtained for the two white cohorts and urban blacks and a negative coefficient for rural blacks. As stated earlier, a positive coefficient of λ_i implies that unobserved factors affecting decision-making in pregnancy resolutions are positively correlated with unobservables involved in delay of seeking prenatal care. Except for rural blacks, this finding may be interpreted as an outcome yielded by a model where costs of abortion dominates other unobservables. Since a woman with a higher cost of abortion may be more likely to give an unwanted birth, she may then delay, more than the average, initiation of prenatal care. The result also indicates that the observed mean of prenatal care delay in the birth sample appears to be higher than the expected mean for the population with the same characteristics X_m had those obtaining an abortion chosen to give birth. This is because, in principle, the first part of prenatal care demand equation, i.e. the deterministic B^mX_m, represents the mean of m_i for the population observably characterized by X_m. The second part with λ_i captures the self-selection effect on m_i, attributed to the unobservables. Apparently a positive coefficient of λ_i serves to increase the delay of obtaining prenatal care. The magnitude of the selection effect on care, however, appears not to be considerable. Urban whites giving birth, for example, are observed to have a delay of 2.29 month on average before to initiating prenatal care. Yet, only about 4 days of this delay could be attributed to the selection effect.

In contrast, for rural blacks, results show a consistency with a model emphasizing the cost of contraception. The finding seems to be in agreement with the reports from Grossman and Joyce (1990, 1991) and Henshaw et. al. (1985). A possible explanation for the outcome is that the shadow price of contraception might be higher and varies more for rural blacks than for all other cohorts.

Attention has been called to the husband's role in initiating prenatal care for women. A hypothesis is made that the demand for prenatal care would be an outcome of the joint decision by both wife and husband. It is further anticipated that husband's role should be positive in the demand for prenatal care and also increases with educational level. Interestingly, the result provides solid evidence for such a hypothesis across all cohorts (Rosenzweig and Schultz 1983). In terms of the coefficients' magnitude, husbands with less than 9 year's schooling make smallest but still positive contribution while husbands with a college education make the largest contribution to initiation of prenatal care. In other words, the marginal productivity of education in the initiation of prenatal care accelerates with educational level. Illegitimacy, a dummy variable, is used to capture in part the difference in demand for prenatal care between married women and single women. Estimates for the coefficient of illegitimacy turn out to be positive as well as significant for both rural cohorts but not significant for women residing in urban areas. This may imply that urban people are more liberal than rural people with respect to seeking early prenatal care for an illegitimate pregnancy. A conservative unmarried woman, for example, may take longer time to decide to have an out-of-wedlock birth before initiating prenatal care than a liberal unmarried woman. For urban women, therefore, the status of illegitimacy may not interrupt the timing of initiating care as significantly as for rural people, if urban women are assumed to be less conservative relative to rural women.

Other significant factors include mother's education (high school and college levels) and mother's age (35-39). Both as expected increase the demand for prenatal care. According to Grossman (1973), Becker and Lewis (1973), Rosenzweig and Schultz (1983), and Easterlin and Crimmins (1985), parental education serves to substitute quality of children for quantity of children. Following the health production function specification earlier, prenatal care acts as one of critical factors in fetal health production. The more educated women should, therefore, initiate an earlier prenatal care to raise the quality of children than the less educated women. Similar outcomes can be also found in the research by Rosenzweig and Schultz (1983), Fingerhood, Makuc and Kleinman (1989), and Joyce and Grossman (1990). Mother's age > 40 shows no effect on prenatal care delay.

6.3 Findings For Birth Weight Production Functions

The birth weight production function is hypothesized to be characterized by self-selections due to both sample censoring and endogeneity of prenatal care. The model controlling for both selection effects is estimated by probit two-stage methods, developed by Lee, Maddala and Trost (1980) and Greene (1990). For comparison, the estimates given by OLS, OLS-Selection, 2SLS, and 2SLS-Selection are presented. Tables 6-12 through 6-15 report all these estimates by race and area of residence.

Self-selection effects are found to be negative in health production functions. All coefficients of λ_i by race and area of residence are consistently negative and statistically significant in OLS-Select model controlled for selection with sample censoring. This finding, along with that for prenatal care demand, presents an agreement with a model emphasizing the cost of abortion. 2SLS-Select yields significant self-selectivity effects for both urban and rural blacks, yet it is not significant for whites although the direction of the selection effect still remains in consistency.[12] The sign for the coefficient of the inverse Mills ratio reflects how unobservables between pregnancy resolutions and health production are interacted. A negative coefficient, i.e. $cov(\epsilon_{ui},\epsilon_{hi}) < 0$, demonstrates that disturbances raising the likelihood of giving birth would be functioning to lower birth weight. It would also imply that the actual mean of birth weight for women giving birth would be lower than the potential mean of birth weight for the entire pregnant population had those who aborted chosen to give birth. The following Table 5-7 presents such expected reductions in birth weight due to sample selection effect only, controlling all explanatory variables including prenatal care delay.

TABLE 6-7
The Average Reductions in Birth Weight for Birth Group
(grams) Attributed to Self-Selectivity By Race and Residence

	Whites		Blacks	
	Urban	Rural	Urban	Rural
OLS-SELECTION	-21.4 (3451.2)*	-32.9 (3404.5)	-105.3 (3163.9)	-84.8 (3151.1)
2SLS-SELECTION	-12.4 (3451.2)	-11.8 (3404.5)	-95.8 (3163.9)	-58.8 (3151.1)

*figures in parenthesis are means

The finding, dominance of the abortion cost, may imply a relatively high abortion cost or great variations in abortion cost for Virginia residents. Reviewing the Federal or State policies

[12] In contrast, a strong evidence of selectivity bias is also found only among black women but not among whites in New York City by Joyce and Grossman (1990). In this work, however, the effect of self-selection is revealed to be positive in birth weight production which supports a model that emphasizes the cost of contraception. The considerable low cost of abortion cost in New York City relative to other states such as Virginia can be in part seen from its high abortion rates. In this study, of the total 194,000 pregnant women from New York City in 1984, 89,000 women (46 %) end up with induced abortions. Based on their randomly chosen sub-samples, the race specific abortion rates respectively are 36 % for whites, 55 % for blacks and 41 % for Hispanics.

regarding funding for abortion services may shed light on this implication.[13] Until August 1977, the Federal government, in cooperation with state governments, funded abortions for eligible women through Medicaid (Title, XIX of Social Security Act). In the 12 months ending September 30, 1977, 4,000 abortions were paid for by Medicaid in Virginia. In August, 1977, the Hyde Amendment took effect and prohibited the Federal funds for abortion services except in cases where the continuation of the pregnancy threatened the pregnant women's life. Although some states still voluntarily continued to pay for most abortions for Medicaid-eligible women or under court order to do so,[14] the number of publicly-funded abortions fell significantly from about 240,900 to about 191,911 in the U.S. as a consequence. The Virginia Commonwealth experienced the sharpest drop in Medicaid abortions among all states from 4,000 in 1977 to 10 in 1978. In contrast, almost 90 percent of the publicly-funded abortions in 1978 took place in California, New York, Pennsylvania, and Michigan. In New York, for example, the drop was only about 16.7 percent, from 50,000 to 41,560. This might be attributed, in part, to New York's varieties of availability as well as relatively low price of abortion services as a consequence of market competition. In 1979, the number of funded abortions was 51 for Virginia while for New York it was 50,310. Apparently the Hyde Amendment may have raised uneven financial burdens on induced abortion for residents in different states. Everything else held constant, an increase in the cost of abortion would increase the likelihood for women, particularly for the low-income women, to carry an unwanted fetus to term or to obtain illegal abortion. The unwantedness of fetus, in turn, would doubtlessly lower birth outcome.

The relatively high cost abortion in Virginia also can be seen from its relatively low

[13] The sources are from Abortion 1977-1979 Need & Services In the United States, Each State & Metropolitan Area, edited by Stanley K. Henshaw et al., The Alan Guttmacher Institute.

[14] The nine states, Alaska, Colorado, Hawaii, Maryland, Michigan, New York, North Carolina, Oregon, Washington, and D.C., keep state-funding Medicaid or medically necessary abortions. Seven other states including California, Illinois, Massachusetts, Pennsylvania, Virginia, West Virginia, and Wisconsin are under court order to do so.

availability of abortion services compared to other states. In Virginia in 1984, the average number of abortion providers per 10,000 women aged 15-44 was 0.28 for all pregnant whites and 0.38 for all pregnant blacks. For those giving birth it was even lower, 0.26 for whites and 0.35 for blacks. In contrast, in New York City in 1984, the availability of abortion providers was much higher at 0.63 for whites and 0.60 for both blacks and Hispanics (Joyce and Grossman 1990a). The relative low overall abortion ratio for Virginia, 25%, compared to 29% for U.S. (Henshaw, Forrest and Blaine 1984), may also shed light into the nature of high abortion cost.

As theoretically expected, the estimates of prenatal care delay illustrates a negative role in influencing birth weight across all cohorts and methods. That is, the longer a woman delays the initiation of prenatal care, the lighter the infant will be at birth. With OLS results, each month of prenatal delay would reduce birth weight by about 13.1 grams for urban whites, 6.3 grams for rural whites, 14.8 grams for urban blacks, and 22 grams for rural blacks. If the self-selection effect with pregnancy resolutions is controlled for by OLS-Select model, then the marginal products of prenatal care for all cohorts are reduced to 9 grams for urban whites, 0.04 grams for rural whites, 9 grams for urban blacks, and 18 grams for rural blacks. The downward shifts may indicate that the self-selection due to sample censoring is favorable to the direct marginal effect of prenatal care in health production.

More interestingly, if the self-selection effect introduced by endogeneity of prenatal care, in additional to the effect of incidental sample selection, is also controlled by 2SLS-Select, then the direct marginal contribution of prenatal care to health production is dramatically improved by 8.3 times to 75.1 grams for urban whites, 2440 times to 107.4 grams for rural whites, 3.8 times to 35.8 grams for urban blacks, and 4.8 times to 85.6 grams for rural blacks. The upward shifts imply a very strong adverse selection in demand for prenatal care due to unobserved endogenous interactions among the demand for prenatal care, health endowment, and health outcomes.

The upward bias due to sample censoring and downward bias due to endogeneity of care can be further confirmed by comparing the result yielded by 2SLS-Select with the regular 2SLS estimates. On the basis of the estimates from 2SLS-Select, 2SLS pushes up the direct marginal effect of prenatal care for all cohorts (139.3 vs. 75.1 grams for urban whites; 142.6 vs. 107.4 grams for rural whites; 197.1 vs. 35.8 grams for urban blacks; and 160.3 vs. 85.6 grams for rural blacks). Similar findings are also obtained by Rosenzweig and Schultz (1983), and Grossman and Joyce (1991).[15]

A Wu-test is conducted to test the endogeneity of prenatal care. For a null hypothesis of no endogeneity of care at 5 percent significance level, it would be highly rejected across all cohorts in the OLS model where no any action was taken to control self-selection effect at all. This test provides very favorable evidence to the self-selection biases which would otherwise distort the observed marginal product of prenatal care in a downward direction. Furthermore, if the self-selection effect due to sample censoring is controlled in the OLS-Select model, then no more evidence can be obtained to reject the null hypothesis of exogeneity of prenatal care for urban whites, urban blacks, and rural blacks. Yet, for rural whites, the null hypothesis would still remain not rejected. That means, in addition to the control of self-selection effects with sample censoring, self-selection with endogeneity of prenatal care would still need to be corrected to get an unbiased estimate. These results of testing are highly consistent with the estimated coefficients of prenatal care delay and their corresponding t-ratios. For rural whites, for example, prenatal care delay has no significant role in the OLS-Select model (marginal product

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[15] For the women giving birth in New York City 1984, Grossman and Joyce demonstrate the self-selection biases with the marginal product of prenatal care in such a pattern: -4.5 from OLS, -4.4 from OLS-Select, and -123.145 from 2SLS-Select for whites; -12.6 from OLS, -12.4 from OLS-Select, and -37.428 from 2SLS-Select for blacks. Rosenzweig and Schultz, using national survey data of about 10,000 legitimate live births for 1967, 1968, and 1969, also arrive at the similar results about the adverse selection bias of endogeneity. Based on a Cobb-Douglas health production function, measured in logs, they present the marginal impacts of prenatal care delay on birth weight as -0.00178 for OLS and -0.0682 for 2SLS.

is as small as -0.04 grams with a t-ratio of -0.007), a signal calling for further correction of selection bias with endogeneity. The 2SLS-Select model, therefore, ends up with a marginal product of care at -107.4 grams with a t-ratio of -2.8, a dramatic improvement.

Other striking findings include the roles of mother's education, infant's sex (male), and parity alive. These all make a significantly positive contribution to infant birth weight across all cohorts. The positive effect of education, once again, enhances the notion emphasized earlier that parental education is favorably correlated with quality of children at the expense of quantity of children. Infant sex apparently captures the biological difference in birth weight between males and females. Males are found, for all cohorts, to gain more weight, at a range of 128 to 148 grams, than females at birth. Parity alive probably reflects, at some extent, women's reproductivity, health endowment, or other established healthy environment (Rosenzweig and Schultz 1983).

The same possible reasons for the positive marginal product of parity alive, in contrast, also shed light into a negative marginal impact of parity dead in the past on present infant health. The results significantly indicate that parity dead in the past would be negatively associated with the birth weight at present. In particular, additional parity dead would, holding everything else constant, lower birth weight by a range of 96 grams to 214 grams, depending upon estimation approaches and cohorts. The inverse correlation between infant mortality rate and birth weight has been remarkably pointed out by researchers as well as policy makers (Institute of Medicine 1985; Corman et al. 1987; Kleiman et al. 1987; Frank et al. 1991). Another negative determinant of birth weight is the number of previous late abortions. Since a late abortion is defined to occur after 20 weeks, it is most likely a spontaneous abortion. If so, one should not be surprised with the result, i.e. the higher the previous late abortions, the lighter the birth weight would be, assuming spontaneous abortion to be an indicator of women's poor health endowment and health productivity. Of course, one would not be able to observe such role of late abortions if women do not significantly behave in such a way in obtaining an induced abortion more likely before 20 weeks. This might be a good explanation of why a significantly negative role of late abortion is observed for both urban whites and urban black as well, but not significant for rural people.

The health production model does not explicitly include illegitimacy because it has been employed in the prenatal care function. It is found that illegitimacy is highly correlated with the predicted prenatal care which enters, as an endogenous variable, into the birth weight function in the second stage. Estimation of an alternative model, including illegitimacy, shows that the marginal effects of some important variables such as prenatal care delay and illegitimacy per se would be seriously distorted against expectations due to the high multicollinearity between predicated prenatal care and illegitimacy in the birth weight production function (the correlation coefficients of illegitimacy and predicated prenatal care delay are -0.64 for urban whites, -0.65 for rural whites, -0.68 for urban blacks, and -0.70 for rural blacks). Age, however, appears to exert little or no effect on birth weight (Rosenzweig and Schultz 1983).

TABLE 6-8

PROBIT PROBABILITY FUNCTION OF GIVING BIRTH WHITE ADULTS

VARIABLES	WHITE/URBAN	WHITE/RURAL
Intercept	2.20555 (21.396)	2.23111 (15.922)
Abortion Providers	-0.123804 (-2.396)	-0.109012 (-0.886)
Family Planning Clinics	0.0639296 (2.124)	0.0310955 (2.333)
White Below Poverty (%)	-0.0088008 (-1.164)	-0.00546157 (0.731)
Terminations < 20W	-0.118059 (-6.353)	-0.193510 (-6.438)
Mother Education 10-12	-0.528817 (-5.777)	-0.566079 (-5.427)
Mother College	-0.552431 (-6.020)	-0.867055 (-8.187)
Mother Age 35-39	-0.213348 (-3.882)	-0.0640202 (-0.607)
Mother Age > 40	-0.781227 (-6.484)	-1.17170 (-6.543)
Illegitimacy	-2.30977 (70.300)	-1.93961 (-38.938)
Likelihood Ratio (χ^2)*	6321.3	1917.0
Prediction Ratio	90.3 %	89.6 %
Observations	13863	6196

*The critical $\chi(9)$ at the 0.5 percent level is 23.6.

TABLE 6-9

**PROBIT PROBABILITY FUNCTION OF GIVING BIRTH
BLACK ADULTS**

VARIABLES	BLACK/URBAN	BLACK/RURAL
Intercept	2.03166 (23.729)	2.12666 (16.682)
Abortion providers	-0.0390194 (-1.090)	-0.109012 (-1.541)
Family Planning Clinics	0.0747891 (4.056)	0.00707861 (0.798)
Black Below Poverty (%)	-0.00331139 (-1.164)	0.00214097 (0.669)
Terminations < 20W	-0.171607 (-12.695)	-0.278791 (-12.949)
Mother Education 10-12	-0.633374 (-9.999)	-0.679486 (-9.119)
Mother College	-1.04003 (-16.026)	-1.11339 (-14.151)
Mother Age 35-39	-0.220447 (-4.027)	-0.264524 (-3.132)
Mother Age > 40	-0.688744 (-5.636)	-0.657863 (-3.896)
Illegitimacy	-1.07931 (40.849)	-0.902193 (-23.910)
Likelihood Ratio (χ^2)*	2374.4	1011.6
Prediction Ratio	73.3 %	77.5 %
Observations	13528	7291

* The critical $\chi^2(9)$ at the 0.5 percent level is 23.6.

TABLE 6-10

DEMAND FUNCTIONS FOR PRENATAL CARE (DELAY)
WHITE ADULTS

VARIABLES	WHITE URBAN		WHITE RURAL	
	OLS	OLS-SELECT	OLS	OLS-SELECT
Intercept	2.77511 (28.935)	2.78753 (28.833)	2.48338 (22.999)	2.48304 (23.019)
WIC Center	0.307302 (8.128)	0.301904 (7.982)	-0.083691 (-3.744)	-0.0835667 (-3.742)
White Below Poverty (%)	-0.0159879 (-3.617)	-0.0161561 (-3.660)	0.0409211 (7.622)	0.0410593 (7.611)
Parity Alive	0.143892 (11.732)	0.14336 (11.713)	0.181127 (9.964)	0.180920 (9.953)
Terminations > 20W	0.0733703 (0.900)	0.0673596 (0.827)	-0.0122726 (-0.129)	-0.0129621 (-0.136)
Father Education 10-12	-0.118972 (-1.593)	-0.109813 (-1.462)	-0.342249 (-5.073)	-0.341357 (-5.057)
Father College	-0.327162 (-4.287)	-0.317102 -(4.133)	-0.55188 (-7.291)	-0.550447 (-7.257)
Mother Education 10-12	-0.307588 (-4.245)	-0.404471 (-4.789)	-0.196866 (-2.838)	-0.205589 (-2.617)
Mother College	-0.466102 (-6.209)	-0.567205 (-6.472)	-0.335514 (-4.266)	-0.349475 (-3.3553)
Mother Age 35-39	-0.101329 (-2.354)	-0.138382 (-3.031)	-0.170831 (-2.163)	-0.172175 (-2.176)
Mother age > 40	0.0518854 (0.419)	-0.121047 (-0.859)	-0.347288 (-1.582)	-0.378603 (-1.477)
Illegitimacy	0.608975 (7.269)	-0.143807 (-0.437)	0.624213 (7.313)	0.566772 (2.197)
Inverse of Mill's Ratio	-	0.638172 (2.364)	-	0.0695729 (0.236)
F-statistic*	76.9778	71.0126	64.7416	59.3403
R^2	0.0722935	0.0727269	0.120436	0.120445
Sample Mean	2.28801	2.28801	2.54633	2.54633
Observations	10878	10878	5213	5213

*The critical $F(12,\infty)$ at the 1 percent level is 2.18.

TABLE 6-11

**DEMAND FUNCTIONS FOR PRENATAL CARE (DELAY)
BLACK ADULTS**

VARIABLES	BLACK URBAN		BLACK RURAL	
	OLS	OLS-SELECT	OLS	OLS-SELECT
Intercept	3.22727 (22.865)	3.23282 (22.745)	2.19387 (13.708)	2.16933 (13.494)
WIC Center	0.272869 (4.799)	0.270799 (4.735)	-0.0607771 (-2.761)	-0.0621943 (-2.827)
Black Below Poverty (%)	-0.0231579 (-7.506)	-0.0231836 (-7.517)	0.0217454 (5.803)	0.0216047 (5.771)
Parity Alive	0.146151 (10.410)	0.146012 (10.402)	0.162098 (9.191)	0.16367 (9.278)
Terminations > 20W	-0.0803432 (-1.018)	-0.0806752 (-1.023)	0.096243 (1.034)	0.0999422 (1.076)
Father Education 10-12	-0.182687 (-1.710)	-0.181215 (-1.696)	-0.204869 (-2.021)	-0.210403 (-2.065)
Father College	-0.435323 (-3.912)	-0.432683 (-3.880)	-0.303124 (-2.590)	-0.322598 (-2.740)
Mother Education 10-12	-0.031654 (-0.460)	-0.0476002 (-0.560)	0.0252441 (0.352)	0.127223 (1.464)
Mother College	-0.262830 (-3.482)	-0.290005 (-2.544)	-0.205207 (-2.407)	-0.0179038 (-0.145)
Mother Age 35-39	-0.174103 (-2.323)	-0.18044 (-2.329)	-0.299562 (-2.734)	-0.24574 (-2.188)
Mother age > 40	0.0776452 (0.402)	0.0574893 (0.283)	-0.155263 (-0.659)	-0.042145 (-0.175)
Illegitimacy	0.208402 (1.978)	0.180415 (1.315)	0.342289 (3.461)	0.491644 (4.019)
Inverse of Mill's Ratio	-	0.0580491 (0.318)	-	-0.442765 (-2.094)
F-statistic*	56.8149	52.0836	35.7005	33.1049
R^2	0.0646161	0.0646265	0.0665184	0.0672492
Sample Mean	2.79402	2.79402	3.05903	3.05903
Observations	9059	9059	5523	5523

*The critical $F(12,\infty)$ at the 1 percent is 2.18.

TABLE 6-12

**BIRTH WEIGHT PRODUCTION FUNCTIONS
WHITE URBAN ADULTS**

VARIABLES	OLS	OLS-SELECT	2SLS	2SLS-SELECT
Intercept	3244.67 (90.001)	3263.48 (90.274)	3604.03 (45.983)	3438.23 (27.965)
Parity Dead	-115.538 (-3.381)	-113.829 (-3.338)	-127.199 (-3.589)	-113.445 (-3.295)
Parity Alive	50.6640 (8.631)	48.9343 (8.341)	67.0736 (9.801)	57.9815 (6.830)
Terminations > 20W	-127.042 (-3.261)	-126.797 (-3.261)	-116.175 (-2.877)	-121.667 (-3.087)
Male	128.213 (11.789)	128.958 (11.878)	128.692 (11.431)	129.191 (11.785)
Mother Education 10-12	98.9777 (2.959)	96.8469 (2.901)	36.3566 (0.992)	65.9262 (1.665)
Mother College	163.316 (4.882)	153.788 (4.598)	60.8052 (1.527)	105.754 (2.264)
Mother Age 35-39	11.9548 (0.584)	17.2668 (0.844)	-3.83884 (-0.179)	6.92286 (0.318)
Mother Age > 40	55.5228 (0.939)	84.1571 (1.425)	62.6164 (1.023)	75.0304 (1.249)
Prenatal Care Delay	-13.0554 (-2.903)	-9.09565 (-2.002)	-139.292 (-5.650)	-75.0714 (-1.685)
Inverse of Mill's Ratio	-	-106.187 (-5.519)	-	-61.4301 (-1.713)
F-statistic	32.9351	32.7136	-	11.4507
Wu-test*	29.2	2.25	-	-
R^2	0.02655	0.0292239	-	0.0104273
Sample Mean	3451.15	3451.15	3451.15	3451.15
Observations	10878	10878	10878	10878

*The critical $F(1,\infty)$ at the 5 percent level is 3.84.

TABLE 6-13

BIRTH WEIGHT PRODUCTION FUNCTIONS
WHITE RURAL ADULTS

VARIABLES	OLS	OLS-SELECT	2SLS	2SLS-SELECT
Intercept	3153.11 (84.073)	3160.49 (84.368)	3559.40 (42.674)	3458.81 (30.793)
Parity Dead	-153.750 (-2.818)	-155.501 (-2.857)	-175.780 (-3.078)	-159.563 (-2.855)
Parity Alive	38.2463 (4.512)	37.4215 (4.433)	64.0020 (6.396)	57.3761 (5.128)
Terminations > 20W	21.7782 (0.493)	35.0496 (0.796)	30.0227 (0.650)	32.2206 (0.711)
Male	147.455 (9.213)	147.993 (9.279)	136.239 (8.094)	147.986 (9.036)
Mother Education 10-12	144.058 (4.649)	151.374 (4.886)	81.1119 (2.365)	99.4397 (2.709)
Mother College	224.732 (6.854)	238.905 (7.274)	114.533 (2.890)	146.784 (3.130)
Mother Age 35-39	-11.4445 (-0.312)	-10.9916 (-0.301)	-42.4651 (-1.099)	-34.3021 (-0.894)
Mother Age > 40	-88.0249 (-0.865)	-5.2755 (-0.052)	-123.351 (-1.159)	85.1113 (-0.780)
Prenatal Care Delay	-6.26492 (-1.006)	-0.0444125 (-0.007)	-142.552 (-5.580)	-107.413 (-2.784)
Inverse of Mill's Ratio	-	-180.614 (-5.249)	-	-64.9554 (-1.192)
F-Test	17.8373	18.774	-	-
Wu-test*	33.37	8.44		
R^2	0.0299309	0.0348328	-	-
Sample Mean	3404.47	3404.47	3404.47	3404.47
Observations	5213	5213	5231	5231

*The critical $F(1,\infty)$ at the 5 percent level is 3.84.

TABLE 6-14

**BIRTH WEIGHT PRODUCTION FUNCTIONS
BLACK URBAN ADULTS**

VARIABLES	OLS	OLS-SELECT	2SLS	2SLS-SELECT
Intercept	2992.23 (89.119)	3045.02 (89.285)	3523.25 (40.034)	3115.37 (27.394)
Parity Dead	-100.604 (-2.970)	-96.4098 (-2.861)	-96.0629 (-2.604)	-96.2213 (-2.850)
Parity Alive	25.4634 (4.224)	22.8743 (3.807)	49.8757 (6.629)	26.5703 (3.207)
Terminations > 20W	-166.472 (-4.917)	-168.198 (-4.990)	-182.331 (-4.936)	-170.283 (-5.020)
Male	134.805 (10.164)	134.713 (10.207)	133.333 (9.232)	134.783 (10.192)
Mother Education 10-12	100.144 (3.433)	137.989 (4.673)	78.6269 (2.463)	131.528 (4.218)
Mother College	184.969 (6.003)	234.484 (7.483)	87.0905 (2.376)	216.152 (5.121)
Mother Age 35-39	23.7912 (0.743)	36.1599 (1.135)	-13.8908 (-0.393)	29.7017 (0.888)
Mother Age > 40	44.8372 (0.542)	106.296 (1.297)	44.9739 (0.499)	100.801 (1.219)
Prenatal Care Delay	-14.7734 (-3.330)	-9.4625 (-2.126)	-197.146 (-7.062)	-35.7833 (-0.877)
Inverse of Mill's Ratio	-	-227.879 (-8.872)	-	-207.348 (-5.087)
F-statistic	24.0970	29.6763	-	26.0908
Wu-test*	52.46	0.42	-	-
R^2	0.0234056	0.0317572	-	0.0280278
Sample Mean	3163.88	3163.88	3163.88	3163.88
Observations	9059	9059	9059	9059

*The critical $F(1,\infty)$ at the 5 percent level is 3.84.

TABLE 6-15

BIRTH WEIGHT PRODUCTION FUNCTIONS
BLACK RURAL ADULTS

VARIABLES	OLS	OLS-SELECT	2SLS	2SLS-SELECT
Intercept	3058.12 (84.255)	3091.66 (83.803)	3478.67 (35.758)	3283.31 (28.007)
Parity Dead	-181.877 (-4.513)	-184.954 (-4.592)	-214.743 (-4.983)	-185.227 (-4.541)
Parity Alive	29.6399 (4.055)	26.9881 (3.692)	50.8919 (5.693)	37.8365 (3.8.91)
Terminations > 20W	-36.9682 (-0.957)	-38.4154 (-0.999)	-21.9596 (-0.537)	-31.7298 (-0.809)
Male	127.643 (7.416)	127.896 (7.457)	131.909 (7.253)	127.811 (7.352)
Mother Education 10-12	55.7031 (1.896)	94.9977 (3.125)	42.6879 (1.372)	77.4166 (2.393)
Mother College	126.409 (3.781)	191.645 (5.382)	63.0518 (1.670)	141.905 (3.074)
Mother Age 35-39	-40.7967 (-0.902)	-21.7045 (-0.481)	-93.8730 (-1.914)	-53.3496 (-1.082)
Mother Age > 40	-1.64551 (-0.017)	45.7606 (0.473)	-38.7641 (-0.376)	13.2317 (0.132)
Prenatal Care Delay	-21.9939 (-4.015)	-18.0016 (-3.270)	-160.262 (-5.348)	-85.5715 (-2.163)
Inverse of Mill's Ratio	-	-236.203 (-5.423)	-	-163.711 (-2.689)
F-statistic	13.1870	14.8325	-	-
Wu-test*	24.73	3.06	-	-
R^2	0.0210742	0.0262044	-	-
Sample Mean	3151.07	3151.07	3151.07	3151.07
Observations	5523	5523	5523	5523

*The critical $F(1,\infty)$ at the 5 percent level is 3.84.

CHAPTER SEVEN

CONCLUDING REMARKS

This dissertation applies an econometric model dealing with sample selection (Heckman 1976, 1979; Lee et al. 1980) to the context of women's pregnancy resolutions, prenatal care demand, and the infant health production function. In particular, it incorporates the self-selection effects of women's pregnancy resolutions and of the endogeneity of prenatal care into the context of demand for prenatal care and health production function, following the original work by Grossman and Joyce (1988). A random birth utility differential function (McFadden 1973, 1974; Manski 1977; Ben-Akiva and Lerman 1989) is introduced to govern the self-selection mechanism of pregnancy resolutions. Estimation of the decision model for pregnancy resolutions is conducted by Maximum Likelihood Probit method. The decision function for pregnancy resolutions and the demand function for prenatal care are in a reduced form. The infant health production function, measured by birth weight, is treated as a structural form with prenatal care as an endogenous variable. Heckman's two stage procedure is employed to estimate the prenatal care demand function controlled for self-selection bias introduced by sample censoring. Estimation of the infant health production function is conducted by several methods with an emphasis on both OLS-Selection and 2SLS-Select, introduced by Lee et al. (1980).

Results of the estimation lead to the major findings of this study. A. Women's self-selection behavior in pregnancy resolutions causes an upward bias to the direct marginal product of prenatal care in health production; a significant self-selection bias in demand for prenatal care is found among urban whites and rural blacks. B. Women's adverse self-selection behavior in input use appears dominant to favorable selection effects and, therefore, introduces a downward bias on the marginal product of prenatal care due to its endogenous association with heterogeneous health endowment. C. Unobserved factors in the decision model of pregnancy resolutions tend to be negatively correlated with those in the health production function and positively correlated with those in the delay of seeking prenatal care (except for rural blacks). This finding is consistent with the outcomes yielded by a model where the shadow price of abortion is dominant over other unobservables. D. Presence of a husband or his education makes a positive contribution in the initiation of prenatal care, compared to single women or women with husbands educated less than 10 years. Women's educational level likewise makes a positive contribution here.

Much more work could be done on the issue of women's decision-making in pregnancy resolutions and self-selection effects on prenatal care demand and birth outcomes technology. One of the important issues may remain with the test of self-selection effects on the health production function controlling for the endogeneity of prenatal care as well as other endogenous choice variables such as women's age, gestational age, education, smoking, income, and marital status (Kessner et al. 1973; Lewit 1977; Gortmaker 1979; Harris 1982; Rosenzweig and Schultz 1983; Corman et al. 1987; Frank et al. 1991). Especially for adolescents, since they have not completed their important choices regarding their future life yet, the model concerning the selection of fertility, initiation of prenatal care, and health technology should particularly consider the endogeneities of those behavioral choice variables. Furthermore, it might be interesting to keep track of the self-selection impacts on the demand for neonatal intensive care and infant neonatal mortality rates.

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