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Pregnancy Wantedness and the Early Initiation of Prenatal Care

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The study examines the impact of the wantedness of a pregnancy on the demand for early prenatal care. Using a cohort of pregnant women in New York City, we estimate a prenatal care demand function in which we control for the probability of giving birth, given a woman is pregnant. We interpret this control as a measure of wantedness. The results indicate that if the black and Hispanic women who aborted had instead given birth, 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 that were actually observed for the women who gave birth. By allowing women to terminate an unwanted pregnancy, induced abortion increases the average use of prenatal care among black and Hispanic women relative to what would have been observed if the women who aborted had instead given birth.

Two recent and comprehensive reports on infant health in the United States both recommended that efforts be directed at reducing the number of unintended or unwanted pregnancies (Institute of Medicine 1985; U.S. Dept. of Health and Human Services 1986). The recommendation was based on the proposed link between pregnancy planning, improved prenatal behavior, and favorable birth outcomes. Specifically, women who choose to become pregnant are apt to be better prepared emotionally and financially for the demands of pregnancy and childbearing. Thus they may be more likely to avoid smoking, to seek out prenatal care earlier, and in general to have a more heightened concern for the impact of their behavior on the health of the fetus than are women whose pregnancies are unwanted or unintended.

As plausible as the recommendation appears, the evidence supporting it is scarce. A number of ecological analyses have documented a relationship between the use of family planning clinics and lower rates of neonatal mortality (Grossman & Jacobowitz 1981; Joyce 1987a). The availability and use of abortion services have also been associated with improvements in area-wide birth outcomes (Corman & Grossman 1985; Joyce 1987b). These authors argue that the use of contraception and abortion should be inversely related to the incidence of unwanted pregnancies and births and positively related to the increased use of prenatal care and other healthy behaviors.

At the individual level, there is some evidence that women who describe their births as wanted begin prenatal care earlier and smoke less during pregnancy (Marsiglio & Mott 1988; Weller, Eberstein, & Bailey 1987). The methodological problems of using self-assessments to measure wantedness, however, are substantial. For instance, the timing of the assessment can affect the response, since a woman's attitude toward her pregnancy and birth may be shaped by the experience itself.[1] Social circumstances can also affect a woman's response (Klerman & Jekel 1984). Teenagers and unmarried women may be more inclined to say that their pregnancies are unwanted because of the social stigmas attached to young and out-of-wedlock childbearing. Similarly, married women may feel inhibited about describing their births as unwanted, especially if their attitudes differ from those of their husbands. Perhaps the most serious drawback is that most research has focused on women whose pregnancies end in a live birth. This excludes all pregnancies that are terminated by induced abortion.

In this study we investigate the impact of pregnancy resolution on the initiation of prenatal care by treating women who give birth as a self-selected sample from the population of pregnant women. We hypothesize that pregnant women who choose to give birth differ in unobserved ways from similar women who voluntarily terminate their pregnancies. A distinguishing characteristic is the wantedness of the pregnancy.

Our approach can be described as a form of revealed preference in which only those pregnancies that are voluntarily terminated are considered to be unwanted. In statistical terms, we treat women who give birth as a censored sample from the population of pregnant women. Such a framework allows us to exploit the well-developed econometrics on selectivity bias (Heckman 1979; Maddala 1983). As a result, we estimate a demand function for prenatal care in which we control for the probability of giving birth given a woman is pregnant. We interpret this control as a measure of wantedness.

Our study is based on a cohort of pregnant women who were residents of New York City in 1984. New York is one of 14 vital registration areas in the United States that submits induced-abortion reports on individuals to the National Center for Health Statistics (Powell-Griner 1987). By combining 1984 induced-termination records with birth certificates in the same year, we were able to generate a sample of pregnant women, all of whom had conceived within a 20-month period. Another reason for choosing New York is that the proportion of pregnancies (live births plus induced abortions) that are terminated by induced abortion far exceeds national estimates.[2] Thus the use of induced abortion to distinguish between wanted and unwanted pregnancies may be most effective in an area such as New York, where the widespread availability of abortion services makes it a readily accessible option. Finally, the racial and ethnic composition of the city allows us to analyze black, white, and Hispanic women separately. The impact of pregnancy resolution on the initiation of prenatal care is likely to differ by race and ethnicity given the marked variation in abortion rates for these groups.

Analytical and Statistical Framework

Self-selection is a potential problem for researchers whenever the subjects under study have not been randomly assigned. In the economic literature, the problem of self-selection has received widespread attention following Heckman's seminal work (1976, 1979). For example, the wages of individuals who have completed a manpower training program are a biased estimate of the wages that individuals with similar observed characteristics would have earned had they gone through the program, since the individuals enrolled in the training program represent a self-selected sample from the population of all potential trainees. Put differently, those who seek out training are likely to differ in unobserved ways (more ability or more ambition) from those who do not. Thus the success of the training program at improving the earning power of graduates may be less a function of the training per se and more related to the unmeasured characteristics associated with the trainees.

The notion of self-selection can be readily applied to prenatal behavior. Contraception and abortion have provided women with the means to control the number and timing of their pregnancies and births. Consequently, women who choose to give birth represent a self-selected sample from the population of pregnant women. We hypothesize that pregnant women who choose to give birth differ in unobserved ways from similar women who voluntarily terminate their pregnancies. A characteristic that distinguishes those who give birth from those who abort is the wantedness of the pregnancy.

From a statistical standpoint, data on prenatal care can be characterized as censored with an unobserved stochastic threshold. In particular, let M_i be the month in a woman's pregnancy in which individual i began prenatal care; let X_{i1} be a vector of exogenous determinants; and let I_i^* represent a woman's desire to have a child given that she is pregnant. There exists some threshold of "wantedness" above which this woman will choose to carry to term. If the wantedness of the pregnancy is below this threshold, then the woman will terminate the pregnancy. I_i^* is unobserved, yet we do know which pregnancies were terminated by induced abortion. Thus we know the women for whom the wantedness of a pregnancy was below this threshold. The model can be written as follows:

$$M_i = X_{i1}B_1 + u_{i1} \quad \text{iff} \quad I_i^* > 0, \tag{1}$$

$$I_i^* = X_{i2}B_2 + u_{i2}, \tag{2}$$

where $I_i = 1$ if $I_{i1}^* > 0$ and $I_i = 0$ otherwise.

I_i is a dummy variable that equals 1 if a woman gives birth and 0 if she aborts. Heckman's insight was that if selection into the subsample was not random, then the expected value of the error term in equation (1) was conditional on the regressors in equation (2). This virtually ensured that the estimates of B_1 obtained by ordinary least squares would be biased given the correlation among the variables in X_1 and X_2.

Heckman's solution was to treat the problem of selectivity bias as a specification error. He showed (Heckman 1979) that unbiased estimates of equation (1) could be obtained under the assumption that the joint distribution of u_{i1} and u_{i2} was bivariate normal. His procedure was to fit equation (2) as a probit equation and to compute the inverse of Mill's ratio (λ_i) for each woman who gives birth:

$$\lambda_i = f(Q_i)/F(Q_i). \tag{3}$$

Here $Q_i = X_{i2}B_2/\sigma_2$ and f and F are, respectively, the density and distribution functions for a standard normal variable. The inverse Mill's ratio is then inserted as a regressor in equation (1), which, after adding the disturbance term (v_{i1}), becomes

$$M_{i1} = X_{i1}B_1 + (\sigma_{12}/\sigma_2)\lambda_i + v_{i1}. \tag{4}$$

The coefficient on λ_i in equation (4) estimates σ_{12} up to a positive scale factor ($1/\sigma_2$). The sign of the coefficient on lambda depends on σ_{12}, which is the covariance between the disturbance terms in equations (1) and (2). Following Trost (1981), we believe that σ_{12} has a particularly useful interpretation given the context of our model. For instance, if M_{i1} is the number of months a woman delays before initiating prenatal care, then women whose pregnancies are more wanted, all else equal, should delay the initiation of prenatal care less than women whose pregnancies are less wanted. Thus a negative coefficient on lambda ($\sigma_{12} < 0$) implies that on average, women with a greater than expected probability of not aborting (i.e., u_{i2} is higher than average) have a shorter than expected delay in the initiation of prenatal care (u_{i1} is lower than average).[3] In short, a negative and statistically significant coefficient on lambda is consistent with the interpretation that pregnancies that are more wanted delay the initiation of prenatal care for a shorter time.

Although we have argued that the wantedness of a pregnancy distinguishes women who abort from women who give birth, clearly caution is in order when giving names to unmeasured variables. For example, a model in which the cost of contraception is the underlying unobservable would generate the same predictions as a model based on wantedness. Specifically, a decrease in contraceptive costs, all else constant, diminishes the probability of an unintended pregnancy and decreases the likelihood of aborting. This lowers optimal family size and raises the investments per child (Willis 1973). On the other hand, some women will not abort a pregnancy on moral grounds, and others, adolescents in particular, choose the birthing option because the direct and indirect costs of obtaining an abortion are too high. Thus increases in the shadow price of abortion increase optimal family size and lessen the resources devoted to each child. Finally, advancements in prenatal diagnosis, such as amniocentesis and ultrasonography, yield very detailed information about the health of the fetus. Although pregnancies that are terminated because of this information are obviously unwanted, it is the health endowment of the fetus that is the underlying causal factor.[4]

The upshot is that there are numerous unobservables that have an impact on pregnancy resolution. Some yield predictions consistent with a wantedness model, and others do not. Trying to identify which is the dominant factor would require a more elaborate model as well as more data. We have emphasized the wantedness of a pregnancy because we feel it represents the most straightforward interpretation of the decision to abort or give birth. We acknowledge that some women who give birth describe their pregnancies as unwanted.[5] This means that the intensity of their subjective evaluation was insufficient to overcome the disutility, or high costs, of an abortion. Our choice of New York City should lessen the proportion of unwanted births due to the inaccessibility of abortion services.[6] Moreover, if the proportion of unwanted births is substantial, then the sign of the residual covariance will be biased toward zero or even positive.[7] The advantage of using pregnancy resolution as an indicator of wantedness is that pregnancies that are voluntarily terminated are clearly unwanted and in New York this represents a major portion of all pregnancies.

Data and Estimation

Data on births and abortions are from New York City vital statistics in 1984.[8] In that year there were approximately 105,000 singleton live births and 89,000 induced abortions among New York City residents. Our analysis is based on randomly chosen subsamples of the combined population of births and induced abortions. Specifically, we subdivided the population into three racial/ethnic groups and two age groups (less than 20 years of age and 20 years and older). The six groups and the number of observations in each group are as follows: white non-Hispanic adults (11,589), black non-Hispanic adults (11,106), Hispanic adults (10,913), white non-Hispanic teenagers (4,132), black non-Hispanic teenagers (12,437), and Hispanic teenagers (8,266).[9] Our analysis is made possible because many of the parental characteristics reported on the birth certificates are also reported on the induced-termination records. Thus by concatenating the data sets, we were able to specify an equation predicting the probability of giving birth given a woman was pregnant. A similar concatenation was done by Powell-Griner and Trent (1987) and Joyce (1988) to study pregnancy resolution. A description of the variables is provided in Table 1. The means and frequencies of the variables are presented in Table 2.

Data from the abortion and birth certificates were augmented with 1980 census data that had been aggregated up from the census tract to the health area level. The health area is the smallest geographical area identified on the birth and abortion certificates. New York City is divided into 352 health areas. The average health area contains between 15,000 and 25,000 residents. The census data enabled us to calculate the percentage of persons below the poverty level in each health area by race and ethnicity.

The vital statistics were also augmented with variables that measure the availability of various reproductive health services. By combining data from the Alan Guttmacher Institute and the New York City Department of Health, we knew the number of family planning

Table 1. Description of Variables

Variable	Definition
Prenatal care delay	The number of months from when a woman conceived until she made her first prenatal care visit
Induced abortions	The number of previous induced abortions
Spontaneous abortions	The number of previous spontaneous abortions (includes fetal deaths)
Late spontaneous abortions	The number of previous spontaneous abortions that occurred after the 19th week of gestation
Parity	Number of previous live births
Age < 18	A dichotomous variable that equals 1 if the woman is less than 18 years old
Age 35–39	A dichotomous variable that equals 1 if the woman is 35–39 years of age
Age 40 and over	A dichotomous variable that equals 1 if the woman is 40 years or older
Schooling < 9	A dichotomous variable that equals 1 if the woman completed less than 9 years of schooling
Schooling = 12	A dichotomous variable that equals 1 if the woman completed 12 years of schooling
Schooling > 12	A dichotomous variable that equals 1 if the woman completed more than 12 years of schooling
Out of wedlock	A dichotomous variable that equals 1 if a woman is not married
Medicaid	A dichotomous variable that equals 1 if the abortion or birth was financed by Medicaid
Self-pay	A dichotomous variable that equals 1 if the abortion or birth was self-financed
Family planning clinic	The number of family planning clinics per 10,000 women aged 15–44 in a health area
Abortion providers	The number of abortion providers per 10,000 women aged 15–44 in a health area
Prenatal care clinics	The number of prenatal care clinics per 10,000 women aged 15–44 in a health area
WIC center	A dichotomous variable that equals 1 if the woman resided in a health area district that contained an office for the Supplemental Nutrition Program for Women, Infants, and Children
Poverty	The race- and ethnic-specific percentage of people below the poverty level in 1980 in a health area; measure for whites includes both white Hispanics and white non-Hispanics; similar comment applies to measure for blacks
Central/South American	A dichotomous variable that equals 1 if the woman's origin or descent is from Central or South America
Mexican	A dichotomous variable that equals 1 if the woman's origin or descent is from Mexico
Cuban	A dichotomous variable that equals 1 if the woman's origin or descent is from Cuba
Other Hispanic	A dichotomous variable that equals 1 if the woman is Hispanic but the country of origin is unknown

clinics, abortion providers, and prenatal clinics by health area in 1983. These availability measures were divided by the number of women aged 15–44 in a health area in 1980. The denominators were from the 1980 census. A fourth availability measure was a dichotomous variable that equaled 1 if the woman lived in a health district that had a health center operated by the Supplemental Program for Women, Infant, and Children (WIC).[10]

Table 2. Means and Frequencies by Pregnancy Resolution, Race, Ethnicity, and Age

Variable	Whites		Blacks		Hispanics	
	Births	Abortions	Births	Abortions	Births	Abortions
			Adults			
Prenatal care delay (months)	3.26	NA	4.88	NA	4.93	NA
Schooling < 9	.08	.09	.01	.04	.15	.06
Schooling = 12	.44	.43	.45	.57	.41	.49
Schooling > 12	.49	.48	.31	.28	.18	.21
Medicaid	.08	.14	.46	.46	.53	.56
Self-pay	.10	.66	.10	.32	.10	.32
Poverty	12.18	13.60	29.38	29.76	36.46	36.99
Total spontaneous abortions	.17	.12	.22	.17	.16	.15
Total induced abortions	.22	1.04	.50	1.30	.32	1.20
Late spontaneous abortions	.01	NA	.02	NA	.01	NA
Age 35–39	.12	.10	.08	.08	.09	.07
Age > 39	.02	.02	.02	.02	.02	.01
Out of wedlock	.08	.69	.55	.76	.44	.71
Abortion providers	.63	.73	.60	.54	.60	.57
Family planning clinics	.74	.94	1.41	1.36	1.65	1.43
Parity	.86	.76	1.26	1.52	1.22	1.53
WIC centers	.30	.29	.57	.54	.41	.46
Prenatal care clinics	.67	.70	1.04	.96	.96	.87
Central/South Americans					.43	.18
Mexican					.02	.01
Cuban					.02	.03
Other Hispanics					.06	.05
Inverse Mill's ratio	.34	−.59	.74	−.60	.50	−.73
Observations	7,361	4,230	4,924	6,092	6,475	4,438
			Teenagers			
Prenatal care delay (months)	4.98	NA	5.78	NA	5.80	NA
Schooling < 9	.08	.04	.06	.06	.15	.08
Medicaid	.38	.15	.66	.49	.73	.63
Self-pay	.16	.75	.10	.32	.10	.32
Poverty	15.95	14.19	32.79	31.46	42.32	39.06
Total spontaneous abortions	.05	.02	.04	.03	.05	.04
Total induced abortions	.11	.36	.20	.44	.11	.45
Late spontaneous abortions	.002	NA	.004	NA	.003	NA
Age < 18	.29	.37	.41	.48	.38	.41
Out of wedlock	.50	.96	.93	.97	.75	.93
Abortion providers	.46	.55	.53	.51	.55	.55
Family planning clinics	.82	.84	1.54	1.42	1.79	1.69
Parity	.16	.11	.21	1.36	.27	.45
WIC centers	.31	.32	.61	.59	.52	.49
Prenatal care clinics	.55	.60	1.06	.98	1.01	.98
Central/South Americans					.17	.11
Mexican					.02	.01
Cuban					.008	.02
Other Hispanics					.04	.05
Inverse Mill's ratio	.77	−.32	.86	−.62	.62	−.73
Observations	1,225	2,907	5,201	7,236	4,482	3,784

For most of the variables, missing data were not considered a major problem. Except for previous induced and previous spontaneous abortions, less than 3% of the combined birth and abortion records lacked data on the variables of interest. If the percentage missing was evenly distributed by births and abortions, then we deleted these observations. In the case of parity and the method of finance, 1% of the observations were missing on the induced-termination records, but between 3% and 6% were missing on the birth records. To avoid altering the ratio of abortions to births, we substituted race/ethnicity- and age-specific means for the unknowns.

A similar problem existed with previous spontaneous and previous induced abortions. On the abortion certificates, approximately 1% of the records lacked data on previous induced abortions, but on the birth certificates, the percentage missing ranged from a low of 6% for white adolescents to a high of 21% for Hispanic adults. For spontaneous abortions, approximately 3% of the abortion records lacked data, but again, the birth records showed a range of 4% missing for white adolescents and 23% for Hispanic adults. To preserve the ratio of abortions to births, we substituted race/ethnicity- and age-specific means for the unknowns. The sensitivity of the results to this form of imputation is discussed later.

The birth probability equation has a dichotomous dependent variable: 1 if the woman gives birth, 0 if she aborts [eq. (2)]. Prenatal care is measured by the number of months a woman delays before seeking medical care for her pregnancy [eq. (1)]. Women who received no care are assumed to have delayed 10 months. A frequency distribution by age, race, and ethnicity for the women who gave birth is presented in Table 3. Except for women who reported no care, the distribution is relatively smooth, which suggests that women reported when care began by month rather than by trimester. The accuracy of New York City birth

Table 3. Frequency Distribution and Percentage Distribution (in parentheses) of the Month in Which Prenatal Care Began by Age, Race, and Ethnicity, New York City, 1984

	Whites		Blacks		Hispanics	
Month	Adults	Teens	Adults	Teens	Adults	Teens
1	253	38	96	99	148	99
	(3.4)	(3.1)	(1.9)	(1.9)	(2.3)	(2.2)
2	2,754	166	673	368	802	335
	(37.4)	(13.6)	(13.7)	(7.1)	(12.4)	(7.5)
3	2,355	248	1,124	738	1,505	665
	(32.0)	(20.2)	(22.8)	(14.2)	(23.2)	(14.8)
4	864	189	882	756	1,123	655
	(11.7)	(15.4)	(17.9)	(14.5)	(17.3)	(14.6)
5	414	130	585	724	817	604
	(5.6)	(10.6)	(11.9)	(13.9)	(12.6)	(13.5)
6	229	116	378	590	503	478
	(3.1)	(9.5)	(7.7)	(11.3)	(7.8)	(10.7)
7	135	92	262	505	340	353
	(1.8)	(7.5)	(5.3)	(9.7)	(5.3)	(7.9)
8	106	83	220	342	267	262
	(1.4)	(6.8)	(4.5)	(6.6)	(4.1)	(5.8)
9	69	52	137	221	160	146
	(0.9)	(4.2)	(2.8)	(4.2)	(2.5)	(3.3)
10 (no care)	182	111	567	858	810	885
	(2.5)	(9.1)	(11.5)	(16.5)	(12.5)	(19.8)
Total women	7,361	1,225	4,924	5,201	6,475	4,482

Note: Based on randomly chosen subsamples of all New York City births in 1984.

certificate data, however, is not well known. A national comparison between birth certificate data and information reported on the National Natality Survey questionnaires revealed that when a comparison was possible, the two sources agreed on the month in which care began in 42.9% of the cases and they differed by one month in 36.9% of the cases. Thus in 75.7% of the cases the discrepancy was no more than a month. In the case of no care, data from the National Natality Survey agreed with birth certificate data in 72.7% of the cases (Querec 1980).[11] If there is systematic error with respect to the month in which prenatal care began on the New York City birth certificates, it is not possible to infer in which direction it may occur. This caveat should be noted when interpreting the results.

Equations (1) and (2) are estimated simultaneously by maximum likelihood. Although the two-step estimator proposed by Heckman (1976) is consistent, maximum likelihood estimates are more efficient. To identify the model, at least one regressor from the prenatal care demand equation must not be included in the birth probability equation (Maddala 1983). Further, the model is on firmer ground if there are unique determinants of each equation. Thus we assume that the availability of family planning clinics and abortion providers has no impact on the demand for prenatal care and that the number of WIC centers and prenatal care clinics has little impact on the decision to give birth. In short, the cross-shadow price effects are restricted to be 0. Similarly, we exclude the number of previous induced abortions from the prenatal care demand equation, since experience with abortion may represent a low psychic cost of an abortion or a high cost of contraception. We excluded parity from the birth probability equation because the left side of the equation is, in essence, a measure of parity. We include parity, however, as well as late spontaneous abortions in the prenatal care equation because they act as proxies for experience with pregnancy and birth.

A unique feature of the New York City vital records is that the method of finance is included on both the birth certificates and the induced-termination records. The three categories include Medicaid, self-pay, and other third party. With respect to prenatal care, we expect women on Medicaid and women who paid for the birth themselves to delay the initiation of prenatal care longer than women who had private health insurance, including membership in a health maintenance organization. Moreover, all Medicaid recipients may face greater search costs, since not all providers accept Medicaid because of the level of reimbursement. In the birth probability equation, however, it is unclear a priori whether Medicaid recipients are more likely to abort than non-Medicaid recipients.[12] The state of New York finances abortions for Medicaid-eligible women, and thus the out-of-pocket costs are zero. Yet Medicaid status clearly measures poverty. If the opportunity costs of giving birth are lower for poor women, then Medicaid status could be positively related to the probability of giving birth (Joyce 1988). With respect to the areal measures, we expect the poverty rate to be positively related to the probability of giving birth (Cutright & Jaffe 1977) and positively related to prenatal care delay. The availability of abortion providers should increase the probability of aborting by lowering the indirect costs of accessing a provider. Given New York City's well-developed mass transit system, however, neighborhood health facilities may be a less relevant availability measure. The same may apply to prenatal care and family planning clinics.

We have also included demographic characteristics of the mother. Among adults, women more than 34 years old should be more likely to abort, since their above-average abortion ratios suggest that their pregnancies have a greater likelihood of being unplanned (Henshaw, Binkin, Blaine, & Smith 1985). For similar reasons unmarried women should be more likely to abort. With respect to prenatal care delay, unmarried women should act as a proxy for households headed by single women, especially among adults. Thus one would expect greater delay due to less income than for women in households with two potential earners. The greater the level of schooling, the greater the opportunity costs of pregnancy and childbearing and the greater the likelihood of abortion. Moreover, more-educated women obtain information regarding the availability of abortion more effectively than less-educated women (Powell-Griner & Trent 1987). Regarding prenatal care, holding income constant, more-educated women may be more aware of the epidemiological relationship between early care and birth outcomes. Thus we would expect greater schooling to be negatively related to prenatal care delay (Cooney 1985; Rosenzweig & Schultz 1983).

Among adolescents, however, a number of the demographic characteristics may be endogenous. For instance, education may determine the probability of aborting, but the years of schooling completed could be affected by the time spent pregnant. Consequently, our measure of adolescent education is a dichotomous variable that equals 1 if a teenager has completed at least 8 years of schooling and 0 otherwise. Such a low cutoff should lessen the potential problems associated with reverse causality by capturing the adolescents whose educational problems existed before they became pregnant.

A similar problem exists with marital status. Vital statistics do not indicate whether a teenager conceived inside or outside the marriage. The distinction is potentially important because the decision to give birth can be made simultaneously with the decision to get married. For example, estimates for 1980 and 1981 reveal that 28% of the first births to white adolescents and 8% of the first births to black adolescents were conceived premaritally but born inside of marriage (O'Connell & Rogers 1984). These figures are national estimates, so their applicability to New York City is unclear. For example, the proportion of out-of-wedlock adolescent births in New York City is substantially higher than the national figures.[13] This suggests that the proportion of pregnancies conceived outside of marriage but delivered inside of marriage is probably lower in New York. In particular, 93% of all black teenagers in New York City who gave birth in 1984 were unmarried (Table 2). The figures for whites and Hispanics are 50% and 75%, respectively. Thus for blacks the endogeneity of marital status in the birth probability equation appears unimportant. For Hispanics and whites we are unsure. This cautionary note should be kept in mind when reviewing the results.

Results

Maximum likelihood estimates of the birth probability equation and the prenatal care demand equation for adolescents are presented in Table 4, and those for adults are presented in Table 5.

Regardless of age, race, or ethnicity, unmarried women and women with at least one previous induced abortion are much less likely to give birth than their married or nulliparous counterparts. This is expected because unmarried women are more likely to have experienced an unintended or mistimed pregnancy (Pratt, Mosher, Bachrach, & Horn 1984). The result with respect to previous induced abortions is consistent with the interpretation that the psychic costs of abortion are lower for women who have had an abortion in the past. Moreover, these women may be more willing to use abortion as a substitute for contraception.

Women who are covered by Medicaid have a greater likelihood of giving birth than those not supported by Medicaid. Again, the finding pertains to all women regardless of age, race, or ethnicity; however, Medicaid status has a much greater impact on the probability of giving birth among adolescents than it does for adults, as measured by the magnitude of the coefficients. The results for adolescents are similar to findings from a recent study of adolescents in California (Leibowitz, Eisen, & Chow 1986). The finding is notable because for women on Medicaid, the out-of-pocket costs for an abortion are zero. One interpretation, therefore, is that economic support for adolescent childbearing as provided by Medicaid and welfare is a disincentive for pregnant teenagers contemplating an abortion. As Joyce (1988) argues, however, disentangling the economic causes of adolescent pregnancy resolution from the emotional and psychological ones may require more refined data.

Table 4. Maximum Likelihood Estimates of the Birth Probability and Prenatal Care Delay Equations (in months) for White, Black, and Hispanic Adolescents

	Whites		Blacks		Hispanics	
Variable	Birth probability	Prenatal care	Birth probability	Prenatal care	Birth probability	Prenatal care
Intercept	.916	3.706	.364	5.135	.227	5.304
	(13.24)	(15.88)	(5.76)	(22.59)	(3.32)	(18.98)
Abortion providers	−.033		.010		.012	
	(−1.35)		(.87)		(.97)	
Family planning clinic	−.016		.005		−.003	
	(−.93)		(.91)		(−.56)	
Prenatal care clinic		−.010		.012		−.021
		(−.14)		(.53)		(−.91)
WIC centers		.075		−.226		−.111
		(.47)		(−2.94)		(−1.33)
Poverty rate	.001	−.009	.002	.008	.012	.010
	(.49)	(−1.24)	(1.80)	(2.52)	(9.69)	(2.40)
Medicaid	1.028	.761	.517	.289	.377	−.372
	(18.49)	(3.05)	(21.12)	(2.45)	(10.97)	(−3.05)
Self-pay		1.335		1.243		.235
		(6.82)		(8.96)		(1.41)
Total spontaneous abortions	.270		.078		.022	
	(1.98)		(1.48)		(.34)	
Total induced abortions	−.653		−.486		−.701	
	(−18.14)		(−24.99)		(−30.97)	
Schooling	.272	.459	−.038	.824	.352	−.159
	(2.72)	(1.85)	(−.76)	(5.05)	(7.15)	(−1.27)
Age < 18	−.113	.489	−.249	.337	−.130	.289
	(−2.15)	(3.13)	(−10.14)	(4.11)	(−4.18)	(3.26)
Late spontaneous abortions		1.280		−.343		−.172
		(.51)		(−.54)		(−.26)
Out of wedlock	−1.939	1.170	−.714	.742	−.882	.676
	(−30.38)	(2.92)	(−12.53)	(4.14)	(−19.94)	(4.48)
Parity		.209		.340		.419
		(1.31)		(13.32)		(5.64)
Central/South Americans					.393	.104
					(8.43)	(.77)
Mexican					.599	.550
					(3.86)	(1.76)
Cuban					−.382	.732
					(−2.71)	(1.36)
Other Hispanics					.040	−.584
					(.58)	(−2.60)
s_1		2.388		2.672		2.738
		(35.88)		(43.47)		(51.85)
r_{12}		.047		−.321		−.187
		(.33)		(−4.15)		(−2.08)
Log likelihood	−4,499.5		−20,146		−15,727	

Note: Asymptotic t statistics are in parentheses. s_1 is an estimate of the standard deviation (σ_1) in the prenatal care delay equation, and r_{12} is an estimate of the residual correlation $[\sigma_{12}/(\sigma_1\sigma_2)]$ between the birth probability and prenatal care delay equations.

Table 5. Maximum Likelihood Estimates of the Birth Probability and Prenatal Care Delay Equations (in months) for White, Black, and Hispanic Adults

	Whites		Blacks		Hispanics	
Variable	Birth probability	Prenatal care	Birth probability	Prenatal care	Birth probability	Prenatal care
Intercept	1.523	3.306	.836	4.186	.779	4.051
	(25.40)	(39.47)	(15.84)	(26.67)	(13.47)	(25.09)
Abortion providers	−.026		.026		−.005	
	(−2.18)		(2.02)		(−.45)	
Family planning clinic	−.006		.006		.022	
	(−.59)		(.85)		(3.69)	
Prenatal care clinic		−.013		−.034		.052
		(−.83)		(−1.61)		(2.96)
WIC centers		−.006		−.355		−.119
		(−.13)		(−4.99)		(−1.80)
Poverty rate	−.004	.014	.000	.019	.001	.020
	(−2.49)	(6.25)	(.32)	(6.04)	(1.28)	(6.92)
Medicaid	.352	1.241	.163	.885	.118	.478
	(8.31)	(19.43)	(5.62)	(10.11)	(3.80)	(6.23)
Self-pay		.478		1.161		.714
		(8.47)		(9.66)		(6.37)
Total spontaneous abortions	.092		.127		.061	
	(3.56)		(5.70)		(2.25)	
Total induced abortions	−.560		−.447		−.531	
	(−46.94)		(−42.04)		(−49.13)	
Schooling < 9	−.546	.335	.461	−.188	.221	−.368
	(−5.32)	(2.86)	(5.14)	(−.97)	(4.32)	(−3.67)
Schooling = 12	−.262	−.493	−.420	−.120	−.340	−.419
	(−4.79)	(−6.84)	(−11.50)	(−2.05)	(−9.84)	(−5.18)
Schooling > 12	−.173	−.596	−.226	−.538	−.360	−.666
	(−3.03)	(−7.76)	(−5.49)	(−4.90)	(−8.29)	(−6.21)
Age 35–39	−.048	−.084	−.088	−.298	.100	−.437
	(−1.11)	(−1.24)	(−1.83)	(−2.07)	(1.93)	(−3.65)
Age 40–44	−.595	−.240	−.404	−.230	.046	−.493
	(−7.56)	(−1.39)	(−4.23)	(−.71)	(.44)	(−1.74)
Late spontaneous abortions		−.057		−.073		−.113
		(−.24)		(−.43)		(−.46)
Out of wedlock	−1.915	.885	−.654	.203	−.652	.504
	(−57.52)	(7.91)	(−22.48)	(2.08)	(−22.32)	(6.12)
Parity		.096		.180		.087
		(6.38)		(7.03)		(3.54)
Central/South Americans					.662	.306
					(20.35)	(3.60)
Mexican					.463	.550
					(3.19)	(2.17)
Cuban					−.217	−.301
					(−2.42)	(−.89)
Other Hispanics					.371	−.438
					(6.13)	(−2.59)
s_1		1.688		2.461		2.512
		(140.37)		68.99)		(75.25)
r_{12}		.014		−.101		−.172
		(.28)		(−1.83)		(−3.29)
Log likelihood	−18,836		−17,861		−20,807	

Note: Asymptotic *t* statistics are in parentheses. s_1 is an estimate of the standard deviation (σ_1) in the prenatal care delay equation, and r_{12} is an estimate of the residual correlation $[\sigma_{12}/(\sigma_1\sigma_2)]$ between the birth probability and prenatal care delay equations.

Adult women with at least 12 years of schooling are more likely to terminate a pregnancy than are women with between 9 and 11 years of schooling. A similar finding has been reported before (Powell-Griner & Trent 1987). In addition, previous pregnancy loss among adult women is positively associated with giving birth. Very young teens are more likely to abort than are older adolescents irrespective of race and ethnicity. Except for Hispanics, older adult women are more likely to abort than their younger counterparts.

Among the areal characteristics, the availability of abortion providers has the correct sign among whites but is statistically significant only for adults. Among black adults the sign is positive and significant. Hispanic adolescents living in neighborhoods of relative poverty have a greater than average propensity to give birth; white adults residing in similar neighborhoods are more likely to abort. Differences among Hispanics with respect to pregnancy resolution are pronounced. Cubans have a greater likelihood of aborting than do Puerto Ricans, but Mexicans and Central and South Americans are more likely to carry to term.

Ignoring the impact of pregnancy resolution, the results for the prenatal care demand equations are in general conformity with the literature. More-educated adult women delay less irrespective of race and ethnicity (Rosenzweig & Schultz 1983; Taffel 1980).[14] Women in areas of above average poverty initiate care later than women from less poor neighborhoods (Joyce 1987a). Residence in a health district with a WIC center is associated with less delay for black women, and neighborhood prenatal care clinics are positively related to delay among Hispanic adults. Finally, except for Hispanic teenagers, women on Medicaid and women with no health insurance begin prenatal care later than women with some other form of health insurance (Cooney 1985; General Accounting Office 1987).

Regarding other characteristics of the mother, older adult women delay less and very young adolescents delay more than women in the respective reference categories. Unmarried women begin care later than married women, and the differential is greatest for whites and smallest for blacks. Greater fertility is associated with greater delay, and women who have experienced a late spontaneous abortion respond no differently than women with no such history. Finally, as in the birth probability equation, there is substantial variation among Hispanic subgroups. Mexicans and Central and South Americans are more likely to delay prenatal care relative to Puerto Ricans; Cubans and other Hispanics delay less.

The presence of selectivity bias is measured by the residual correlation $[\sigma_{12}/(\sigma_1\sigma_2)]$ between equations (1) and (2). We find evidence of selectivity bias in the prenatal care demand equations for blacks and Hispanics irrespective of age, although for black adults, the null hypothesis of selectivity bias can be rejected at only the .07 level. There is no evidence of such bias among whites. Put differently, controlling for variations in demographic characteristics, years of completed schooling, neighborhood poverty, and method of finance, we find that black and Hispanic women who give birth represent a nonrandom draw from the population of pregnant women. The negative correlation indicates that on average, the unobserved factors that raise the probability of giving birth are positively associated with the unobserved factors that decrease delay in the initiation of prenatal care. It is our contention that by controlling for self-selection into the birth sample, we are controlling, in part, for the wantedness of the pregnancy. The negative correlation between the residuals in the pregnancy resolution equation and the prenatal care demand equation is consistent with the interpretation that pregnancies that are more wanted delay the initiation of prenatal care less.

These results imply, for white pregnant women, that the decision to abort or give birth yields no information regarding the use of prenatal care. This may be because we have no equation predicting the probability of becoming pregnant. If whites use contraception more often than blacks, and a number of sources indicate that they do, then the selection process that reduces the proportion of unplanned pregnancies may occur at an earlier point in the reproductive cycle among whites than it does for blacks (Henshaw et al. 1985; Pratt et al.

1984; Stephen, Rindfuss, & Bean 1988). Thus by analyzing only pregnant women, as opposed to sexually active women, we are unable to incorporate the impact of contraception on prenatal behavior.

As discussed earlier, there were a substantial number of missing observations on previous spontaneous and induced abortions. Race-, ethnicity-, and age-specific means were used to impute values for the missing data. The robustness of the results to this imputation was checked in several ways. First, to test whether women whose previous induced abortions were unknown differed in unobserved ways from women who reported the number of previous induced abortions, we applied the sample selection procedure previously described.[15] There was no evidence of selectivity bias among the unknowns for either spontaneous or induced abortions. Second, we deleted all of the observations on previous induced and spontaneous abortions and reestimated the equations in Tables 4 and 5. The only change of note involved black adult women. The estimated value of the residual correlation coefficient fell to $-.061$ ($t = -1.09$). We also estimated regressions predicting the number of previous induced and spontaneous abortions for each age, race, and ethnic group among women for whom the number of previous abortions were known. The regressors included age, education, marital status, the method of finance, the poverty rate, and the availability of family planning and abortion services. We used the estimates from the regressions to impute values for the missing observations. The results in Tables 4 and 5 did not differ in any appreciable manner when the imputed values were obtained by ordinary least squares as compared with the race-, ethnicity-, and age-specific means.

One means of gauging the magnitude of the selection effect on the initiation of prenatal care is to compare the observed mean number of months a woman delays before initiating care to the expected mean delay of women who aborted had they chosen to give birth. Our first comparison is based on unobserved factors only.[16] Ignoring the results for whites, we find that black teenagers who aborted would have delayed the initiation of care 1.5 months more than the mean number of months delayed that were actually observed for black adolescents. Hispanic adolescents who aborted would have delayed .8 of a month more than was observed. The relevant figures for black and Hispanic adults are .4 and .7 months, respectively. If we allow for differences in both observed and unobserved characteristics between women who gave birth and women who aborted, the results change slightly.[17] Hispanic adolescents who aborted would be expected to delay 1.0 months more than their counterparts who gave birth; black adults who aborted would delay the initiation of prenatal care, on average, .5 of a month more. There were no changes for black adolescents and Hispanic adults. As the results make clear, the expected differences in the initiation of prenatal care between aborters and those who give birth are dominated by the unobserved characteristics.

Conclusion

We have attempted to incorporate the information about the resolution of a pregnancy into the prenatal care demand equation. We accomplished this by treating women who give birth as a self-selected sample from the population of pregnant women. Instead of simply testing for selectivity bias, we have argued, however, that the sign of the residual covariance between the pregnancy resolution equation and the prenatal care demand equation may offer useful insights into the effect of unobservables on the initiation of prenatal care. In particular, women with a higher than expected probability of giving birth evidence a smaller than expected delay in the initiation of prenatal care. The result is consistent with the interpretation that women whose pregnancies are more wanted obtain prenatal care earlier than women whose pregnancies are less wanted.

We found that the black and Hispanic women who give birth differ in a statistically significant manner from their counterparts who abort. We observed no such differences for whites. We have speculated that black and Hispanic women may substitute abortion for contraception more frequently than white women.[18] As a result, the selection mechanism encouraged by abortion has a greater impact on the prenatal behavior of minorities.

Given the direct association between early prenatal care and improved birth outcomes, the results reported here suggest an important indirect effect of pregnancy resolution on infant health. Policies aimed at reducing unwanted births would increase the average use of prenatal care among women who choose to give birth. Clearly, reducing unwanted pregnancies is the preferable strategy for reducing unwanted births. The effective use of contraception among poor, young, and minority women, however, is not encouraging (Henshaw & Silverman 1988; Mosher & Bachrach 1986). Abortion remains a crucial option for many women trying to avoid an unwanted birth. If the U.S. Supreme Court overturns the decision of *Roe vs. Wade* and legalized abortion is banned, the results from this study predict that the proportion of black and Hispanic women who obtain early prenatal care will fall. The suggested impact on birth outcomes would not be favorable.

It must be noted that New York City is a unique setting. It has a well-developed market of private and public providers, a state government that continues to fund abortions for all Medicaid-eligible women, no parental notification laws with respect to minors, and a readily available system of transportation. At the same time the shadow price of contraception may be higher for minorities than for whites due to language barriers, lower levels of schooling, and less access to private gynecological care. Given these factors, the shadow price of abortion relative to contraception may be lower for black and Hispanic women, which would induce greater substitution away from contraception toward abortion.

The upshot is that the results reported here may not generalize to other areas where the shadow price of abortion is higher. Consequently, an agenda for future research would be to conduct a similar analysis in other states where abortion services are less accessible. A higher shadow price of abortion would either promote greater contraceptive use or result in more unwanted births. The latter outcome is more likely for adolescents and minorities, given the frequency with which contraception is used by these groups (Pratt et al. 1984).

Notes

[1] For example, after controlling for numerous characteristics, Marsiglio and Mott (1988) found that women who reported on their pregnancies while pregnant were more likely to have wanted the pregnancy than women who reported after giving birth. Moreover, among the subsample who reported while pregnant, wantedness was not related to early prenatal care.

[2] Forty-six percent of all pregnancies to New York City residents in 1984 were terminated by induced abortion. The average for the other 12 states that reported induced abortions to the National Center for Health Statistics was 21% (National Center for Health Statistics 1986; Powell-Griner 1987). Although New York City has a large minority population, 45% of all pregnancies to white non-Hispanic women were terminated by induced abortion.

[3] The expected probability of aborting and the expected probability of delay are based on $F(X_{i2}B_2)$ and $X_{i1}B_1$, respectively.

[4] A more complete discussion of the unobservables that determine the probability of giving birth and their implications for infant health is presented by Grossman and Joyce (1988).

[5] Data from the National Survey of Family Growth indicate that approximately 10% of cumulative births to ever-married women aged 15–44 in 1982 were described as unwanted. By race, 8% of white births and 22% of black births were unwanted at the time of conception (Pratt & Horn 1985).

[6] New York has a well-developed market of private and public abortion providers, an inexpensive system of mass transit, no parental notification laws, and one of only 14 state governments that fund abortions for Medicaid-eligible women. This explains in part why the proportion of pregnancies (live births plus induced abortions) in New York City that are terminated by induced abortion greatly exceeds national averages (see note 2).

[7] Specifically, women whose births were unwanted may have a higher than expected probability of giving birth ($u_{i2} > 0$) but a greater than expected delay in the initiation of prenatal care ($u_{i1} > 0$). This positive correlation among the residuals would lead to a rejection of a model based on wantedness.

[8] We do not include women whose pregnancies were terminated by spontaneous abortion. Early spontaneous abortions are poorly reported. The New York City Department of Health reported 4,960 spontaneous abortions in 1984. This represented less than 4.4% of all live births. Yet data from the National Survey of Family Growth indicate that the ratio of spontaneous abortions to live births is greater than .21 (Pratt, Mosher, Bachrach, & Horn 1984).

[9] For white and Hispanic adolescents, the observations represent the entire population except for records that were deleted because of missing values.

[10] There are 30 health districts in New York City. Each contains approximately 10 health areas.

[11] A similar comparison of birth certificates and hospital records in the state of New York, excluding New York City, revealed exact agreement in 53.9% of the cases and a discrepancy of 1 month or less in 82.0% of the cases (Carucci 1979).

[12] In the birth probability equation, the method of finance is reduced to two categories, Medicaid and all others. The self-pay category is difficult to interpret because third-party coverage of abortions is not common. Among women who give birth, however, third-party coverage more accurately reflects a well-insured individual. For example, among adult white women, 66% of the abortions but only 10% percent of the births were self-pay.

[13] Nationally, 89% of all black live births and 43% of all nonblack live births to women 19 years and younger were born out of wedlock. The nonblack category includes the vast majority of Hispanic births (National Center for Health Statistics 1986).

[14] The one exception is Hispanic adults, for which women with less than 9 years of schooling delay approximately one-third of a month less than women with between 9 and 11 years of schooling.

[15] Specifically, we predicted the probability that information regarding previous induced abortions was known. The dependent variable was dichotomous: 1 if the number of previous induced abortions was known, and 0 otherwise. The independent variables included age, education, marital status, the method of finance, the poverty rate, and the availability of abortion and family planning services. We then estimated the number of previous induced abortions among women whose previous induced abortions were known. We included the inverse Mill's ratio, which had been generated by an equation predicting the probability of whether data on induced abortions was known. We did this for each of the six groups. We followed the same procedure for spontaneous abortions. The coefficient of the inverse of Mill's ratio never was significant.

[16] Let M_b be the observed mean number of months a woman delays before initiating prenatal care. Let M_a be the expected delay in months that we would have observed among women who aborted had they chosen to give birth. Let $\overline{X}_b$ and $\overline{X}_a$ be the means of the determinants of M_b and M_a, respectively, and let $\hat{B}_1$ be the estimated vector of coefficients. Correcting for the selection, the difference between M_b and M_a is

$$M_b - M_a = (\overline{X}_b - \overline{X}_a)\hat{B}_1 + \hat{\sigma}_{12}/\sigma_2(\hat{\lambda}_b - \hat{\lambda}_a), \tag{5}$$

where $\hat{\lambda}_b$ is the estimated inverse Mill's ratio associated with those who give birth ($\hat{\lambda}_b > 0$) and $\hat{\lambda}_a$ is the estimated inverse Mill's ratio associated with those who abort ($\hat{\lambda}_a < 0$). Assuming $X_b = X_a$ yields the effect of unobservables alone on prenatal care delay.

[17] These calculations differ from the previous ones in that $\overline{X}_b$ and $\overline{X}_a$ in equation (5) of note 16 are allowed to differ.

[18] The speculation is based on the use of abortion and contraception by race and ethnicity. It is well documented that nonwhites have higher abortion rates and ratios than do whites (Henshaw et al. 1985). There are fewer data on Hispanics. Recent figures from a nationally representative sample indicate, however, that although Hispanics make up 8% of the population, 13% of all abortions in 1987 were to women of Hispanic origin (Henshaw & Silverman 1988). The same study reported that 57.6% of the white abortion patients were using contraception the month they became pregnant, as opposed to 41.8% of black and 42.1% of Hispanic abortion patients. Further data on contraception from the National Survey of Family Growth reveals that 46.6% of white women 15–44 years old who have had sexual intercourse used contraception at first intercourse. The figures for blacks and Hispanics were 33.6% and 26.5%, respectively (Mosher & Bachrach 1986).

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