The Effects of Economic Conditions and Access to Reproductive Health Services on State Abortion Rates and Birthrates

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Abstract:
The effects that such factors as wages, welfare policies and access to physicians, family planning clinics and abortion providers have on abortion rates and birthrates are examined in analyses based on 1978-1988 state-level data and longitudinal regression techniques. The incidence of abortion is found to be lower in states where access to providers is reduced and state policies are restrictive. Calculations indicate that decreased access may have accounted for about one-quarter of the 5% decline in abortion rates between 1988 and 1992. In addition, birthrates are elevated where the costs of contraception are higher because access to obstetrician-gynecologists and family planning services is reduced. Economic resources such as higher wages for men and women and generous welfare benefits are significantly and consistently related to increased birthrates; however, even a 10% cut in public assistance benefits would result in only one birth fewer for every 212 women on welfare. Economic factors showed no consistent relationship with abortion rates. (Family Planning Perspectives, 29:52-60, 1997)

Article:
Policy discussions concerning factors that influence reproductive behavior reach a level of intensity seldom matched by other matters of public discourse. Especially controversial have been discussions regarding the effects of Medicaid funding restrictions, 24-hour waiting periods and parental consent requirements for minors on the incidence of abortion; the effects of family planning programs on rates of births and abortions; and incentives for out-of-wedlock childbearing attributed to the Aid to Families with Dependent Children (AFDC) program.

These questions reflect more general but perhaps less widely publicized issues that are well suited to economic analysis, such as the cost and accessibility of reproductive health services and changes in women's and men's economic opportunities. Economic models are particularly useful for sorting out the determinants of reproductive behavior in the United States because of the substantial heterogeneity that characterizes not only the population's values and preferences, but also its economic resources and access to different types of health care.[*]

Accordingly, economists have conducted numerous empirical studies of the socioeconomic and political deteiminants of fertility in the United States. This literature, which is summarized elsewhere,[1] has recently expanded to include research on the determinants of abortion.[2] While researchers have generally found support for basic economic hypotheses, a number of specific results, including most of the policy questions raised above, remain in dispute.

However, there remain unexplored issues and weaknesses in the literature. An important shortcoming of many studies is the omission of relevant economic and policy variables. For instance, analyses of women's abortion decision-making have failed to include the gender-specific measures of economic resources and labor-market opportunities that have commonly appeared in studies of fertility. Both types of studies have tended to include

only limited measures reflecting access to reproductive health services. A related problem, which recent analyses have begun to address, involves potentially confounding effects from unobserved and imperfectly measured variables. Finally, only a few studies have examined abortion and fertility behavior in tandem.[3]
In this article, we investigate the determinants of annual abortion rates and birthrates using state-level data from the years 1978-1988. The analysis is based on an economic model in which the behavior leading to pregnancy and to pregnant women's decisions on whether to carry the pregnancy to term depends on the women's resources, direct costs, opportunity costs, attitudes and preferences for children.

Our empirical analysis complements and extends previous research in several respects. First, we consider a comprehensive set of explanatory variables. To describe the direct costs of contraception, abortion and births, we include longitudinal variables for the number and geographic distribution of family planning clinics, abortion providers and obstetrician-gynecologists within states. State policy indicators are used to measure access to reproductive health services. To describe resources and opportunity costs, we include state- and year-specific measures of women's and men's property incomes (i.e., nonwage) and available wage rates. To gauge the political and social climate, we examine party affiliations of state executives, legislators' voting records and attitude measures drawn from opinion surveys. For many of our primary variables, we have identified supplemental measures that permit us to check the robustness of our results (i.e., we check whether the estimated effect of each variable maintains its sign and magnitude and remains significant when assumptions underlying the analytic model change).

Second, although the use of such elaborate controls accounts for a considerable portion of the cross-state differences in abortion rates and birthrates, standard regression estimates may still be biased by the omission of relevant variables. Hence, in our empirical analysis, we employ fixed-effects regression methods (i.e., we account for state-specific dummy variables) to control for unobserved heterogeneity in attitudes and institutions across states.

Third, the study examines abortion rates and birthrates together, allowing us to consider the logical consistency of the results, check for "cross-policy" effects (e.g., the effect of abortion policy on fertility) and examine whether policies affect reproductive outcomes by altering the number of pregnancies that occur. Furthermore, since the presence of measurement error in an independent variable should bias the estimates for both abortion rates and birthrates toward zero, this parallel analysis may help us distinguish a small effect of an independent variable from an artifact of measurement error.

Theoretical Framework

We begin by sketching a simple, stylized model of the economic determinants of behavior that may lead to pregnancy and the decision of how to resolve a pregnancy. The model is offered to generate broad predictions and motivate the empirical analysis that follows. Because economic models of fertility have been extensively discussed elsewhere,[4] and since data limitations preclude us from undertaking a detailed structural analysis, we keep the theoretical discussion brief.

A live birth results from a number of behavioral and biological factors. For simplicity, our theoretical analysis groups the behavioral determinants of fertility into two decisions. The first decision is the level of effort used to avoid or achieve a pregnancy (contraceptive effort). This decision reflects such behavioral factors as entry into a sexual relationship, frequency of intercourse and choice of contraceptive method, and such biological factors as postpartum infecundability and onset of sterility. The second decision, faced by women who conceive, is how to resolve the pregnancy.

These decisions can be examined through a straightforward economic analysis of fertility.[5] We assume that women have preferences regarding contraceptive use, abortion, the presence of children and the consumption of commodities. These preferences embody women's personal, cultural and religious values. We also assume that women's preferences are constrained by their available time and economic resources, the availability and costs of alternative reproductive health services, and the time and financial requirements of raising children and producing household commodities. This approach reveals that contraceptive effort and pregnancy resolution are related and that each involves direct expenditures of time and money, as well as opportunity costs in terms of forgone earnings and consumption possibilities.

To examine the implications of the model, we first consider conditional predictions regarding pregnancy resolution. The model predicts that if the direct costs of abortion are increased by such factors as higher service fees, reduced availability or tighter legal restrictions, pregnant women will be deterred from obtaining abortions. On the other hand, if the direct costs of having children are increased by such factors as higher delivery costs and less generous AFDC and Medicaid payments, women will be more likely to terminate unintended pregnancies.

Higher levels of wealth from sources other than earnings increase the affordability of both abortions and children; however, we assume that the presence of children is an economic "good," and we therefore expect childbearing to dominate. (The opposite relationship is possible if income increases the demand for child "quality" — e.g., improved health and education — and, in turn, raises the price of having more children.[6]) An increase in women's wages raises both wealth and the opportunity costs of childbearing; consequently, its effect on pregnancy resolution is ambiguous.

Returning to contraceptive effort, the model suggests that the above factors affect the unconditional rates of abortions and births to the extent that they affect the likelihood of pregnancy. For instance, higher costs associated with obtaining an abortion or having a child reduce the expected indirect utility associated with pregnancy (i.e., the woman's valuation of alternatives) and thereby encourage contraceptive effort. Thus, policies that restrict access to abortion not only deter pregnant women from obtaining abortions but also deter women from becoming pregnant in the first place. Because these effects are reinforcing, such policies would lower the incidence of abortions. For births, however, the separate effects of these policies run counter to one another, leading to ambiguous net effects.

A similar analysis can be used to examine other policy changes. For example, we would expect more generous AFDC benefits to be associated with increased rates of pregnancies and births, but the overall effect on abortion rates would be less clear.

Previous Research

Nearly every social science discipline has contributed to examinations of various aspects of this theoretical model. In some cases, the model's predictions have been confirmed (e.g., the effect of policy restrictions on abortion levels),[7] but other results have been less conclusive (e.g., the effect of AFDC benefits on birthrates).[8] A number of careful reviews of the literature are available.[9] For brevity, we focus on two studies (by Blank and colleagues[10] and by Jackson and Klerman[11]) that have addressed methodological issues from previous research, plus a few others that address economic hypotheses.

A central methodological issue is that ordinary regression estimates are biased if important determinants of reproductive behavior, such as the values underlying women's preferences, are correlated with the observed independent variables but are omitted or only partially accounted for in the regression equation. Indeed, Blank and collaborators found that the estimated effects of parental involvement laws and Medicaid funding restrictions on abortions were sensitive to the use of state-specific dummy variable controls for unobserved heterogeneity (fixed effects). They also showed that both parental involvement and funding restrictions had negative but statistically nonsignificant effects on abortion rates classified by the woman's state of residence; Medicaid restrictions had significant effects on rates by the state in which the abortion occurred, a finding confirmed by at least one other study.[12] However, other studies found that both types of restrictions significantly reduced the incidence of abortion among minors.[13]

These fixed-effects studies have had more uniform results in other areas: They have found that the incidence of abortion is positively and significantly associated with both the availability of providers and women's per capita income.

Unfortunately, several issues have been ignored or only lightly researched. For example, only one fixed-effects abortion study has examined access to reproductive health services, other than abortions,[14] and no abortion study — fixedeffects or otherwise has separately examined the economic resources of women and men. Although women's and men's economic resources and opportunities have not been considered in the abortion studies, measures of these characteristics have figured prominently in empirical analyses of fertility. For example, a 1994 cross-sectional analysis of welfare and fertility included women's wages and properly incomes and potential spouses' wages, arguing that omitting these conditions would likely bias the estimated effects of AFDC generosity on births.[15] As in most fertility analyses, the study found a significant negative relationship between women's wages and fertility; unexpectedly, however, property income also had negative effects. Men's wages had a positive association with childbearing among younger women, while AFDC generosity typically had nonsignificant effects.

These results, however, also appear to be sensitive to the inclusion of fixed effects. Preliminary results from Jackson and Klerman based on fixed-effects regressions indicated that AFDC benefits had large positive effects on the birthrate among white women but no association among blacks. In addition, men's earnings had positive effects on births among whites and mixed effects among blacks; however, the earnings results were very sensitive to the inclusion of other economic variables.

With respect to other policy variables, preliminary evidence suggests that parental involvement laws are associated with reduced levels of childbearing.[16] However, evidence regarding the fertility effects of Medicaid funding restrictions has been mixed.[17]

Methodology Data and Variables

We employ an extensive collection of state-level data covering the period 1978-1988. Detailed descriptions of the data and their sources appear in the appendix (page 59). The use of state-level information confers some advantages over using data that are less aggregated. Primarily, it permits us to examine a wide array of measures covering a moderately long period. The data, which describe behavior related to highly personal issues, also are less susceptible to reporting problems than are some microlevel data sets and can be made nationally representative. On the other hand, when using aggregate data, we lack individual-level controls and cannot isolate effects among particular groups of women (e.g., the effects of Medicaid restrictions only among poor women). In addition, with these data, we can make only limited inferences about individual behavior.

The study focuses on two reproductive outcomes — births and abortions. Annual data on total births are available from several sources; we use information from the Area Resource File.[18] Obtaining accurate data on abortions is more problematic. The best available information comes from The Alan Guttmacher Institute (AGI), which regularly surveyed abortion providers through the 1970s and 1980s.[19] (However, provider information was not collected for 1983, 1986, 1989 or 1990.) On the basis of these provider reports, AGI estimated the numbers of abortions by state of occurrence and by state of residence.

Unfortunately, the rates by state of occurrence are difficult to analyze, since unknown numbers of women cross state borders to obtain an abortion. According to the Centers for Disease Control and Prevention, which also collects state-level abortion data, the proportion of abortions obtained out of state declined from 44% in 1972 to less than 10% after 1980.[20] However, the low proportion in later years masks considerable variation across states. For example, AGI estimates that in 1985, nonresidents accounted for roughly half of the abortions performed in North Dakota and the District of Columbia but fewer than 1% of those performed in California.[21]

The AGI data by patients' state of residence may contain more error than the figures based on state of occurrence; furthermore, since the procedure used to estimate residence changed in 1978, we exeluded state-of-residence data from earlier years. Nevertheless, we use the state-of-residence estimates because they allow us to match information on abortion incidence with data on the characteristics of women who may have an abortion.

The explanatory variables for our analysis (shown with their population-weighted mean values for the United States in Table 1) can be grouped into a few broad categories. A key set of measures describes the accessibility of reproductive and general health services. These are the numbers of abortion providers, family planning clinics and obstetrician-gynecologists per 1,000 women aged 15-44; the proportion of women living in counties with each service; the average distance to the nearest in-state and out-of-state abortion provider; and the proportion of the population enrolled in a health maintenance organization (HMO).

We interpreted accessibility as a proxy for the direct cost of a service. For the family planning and abortion availability measures, the implications are straightforward: Increased access reduces the effective costs and should increase reliance on contraception and abortion, respectively. For obstetrician-gynecologist availability and HMO membership, the implications are less clear, because access to these services reduces the costs of all types of reproductive health care.

Another set of variables describes women's and men's economic resources and opportunities. The primary data are calculated from the 1979-1989 March supplements of the Current Population Survey (CPS).[22] Pooling individual-level data from these CPS files, we have averaged property incomes by sex, state and year to form gender-specific, longitudinal estimates of wealth. The pooled data are also used to construct ordinary averages and selectivity-adjusted imputations of women's and men's hourly wages for each state and year (i.e., imputations that account for the fact that not all women and men work).

To assess whether the estimation results are sensitive to the specification of the CPS variables and for general purposes of comparability, we supplemented the CPS measures with longitudinal state-level data on gender-specific unemployment rates and per capita total personal income and on average annual manufacturing and retail earnings.[23] An advantage of these measures is that they are representative for all states. (Although the CPS is nationally representative, weighted observations from it may not be representative at the state level.) The disadvantages are that they only indirectly measure the variables of interest, they may be more endogenous than the measures of wage or property income, and the variables for total income and sector-specific earnings may not accurately reflect gender-specific opportunities.

A third group of explanatory variables describes state policies that may influence reproductive decisions. Some of these restrictions on Medicaid funding for abortion, and parental consent and notification laws for abortion — are intended to do so. Others — AFDC benefits for a family of four and average Medicaid benefits distributed to AFDC recipients — may alter incentives for childbearing even if that is not their purpose.

To control for attitudes toward abortion and other institutional and population characteristics, we include in the analysis several political and demographic variables. These generally have standard interpretations, and many have appeared in previous studies.

Analytic Approach

In this article, we use regression analysis to examine the determinants of state abortion rates and birthrates. For each outcome, we report estimates from two ordinary least-squares regressions. The first incorporates the explanatory variables and dummy variables for each year and each state (i.e., time and state fixed effects). The second adds interactions of each state dummy variable with a linear time trend. To make the results nationally representative, we have weighted each state's observations by the number of women aged 15-44 in 1980. (We also ran models that omitted the state dummy variables and treated the state-specific controls as random effects. Tests indicated that we could reject these models.)

These two models, which account for potential confounding influences of omitted state-specific factors, generate estimates that are robust under a broad set of circumstances. The models may be especially useful in reducing biases among the potentially related health service accessibility and policy variables. Specifically, changes in the availability of health care providers and shifts in policy probably both influence and are influenced by changes in reproductive behavior. However, endogeneity in these variables seems more likely to stem from shifts in long-term rather than short-term conditions. To the extent that the models control for such long-term movements, biases in the service provision and policy variables should be mitigated.

There are disadvantages to this approach. First, these procedures can exacerbate biases associated with measurement error and simultaneity in the explanatory variables. Second, influences of time-invariant observed variables cannot be identified separately from the state dummies. A related issue is that because the analysis involves data for a relatively short time period, it may be difficult to identify the effects of variables that change only gradually within states, such as the population measures. For the regressions including both state dummies and state-time interactions, this difficulty becomes even greater because the model does not easily distinguish the effects of variables that, within states, follow an essentially linear trend, such as the abortion policy measures.

Results Baseline Analyses

In the first regression for determinants of state abortion rates (Table 2, page 56), the significant positive coefficient for abortion provider access and significant negative coefficients for Medicaid funding restrictions and parental consent or notification laws indicate that abortion demand is negatively associated with the procedure's direct costs. The significant positive coefficient for HMO membership suggests that this variable may also be a proxy for the immediate costs of obtaining an abortion or an abortion referral. Of the other health service measures, access to family planning clinics has a small and marginally significant positive association with abortions ($p<.10$), while access to an obstetrician-gynecologist has no significant effect.

There is little evidence of effects of income or opportunity cost. Neither the gender-specific economic variables, AFDC benefits nor Medicaid coverage has a significant effect on the abortion rate. Among the remaining control variables, the proportions of blacks and women aged 35-44 have significant effects in the anticipated direction. Unexpectedly, however, the presence of a Republican governor is associated with a significant increase in the abortion rate.

The results for the second model indicate that the added interactions of state dummies and linear time trends are jointly significant. Nevertheless, several results are consistent with those of the first model. The coefficient for accessibility to abortion providers remains positive and significant, confirming that the direct cost of abortion services, as measured by proximity to providers, affects the incidence of abortions. The models are likewise consistent in generating nonsignificant results for the economic variables. Although this partly reflects imprecision in the estimates, it also reflects some genuinely small effects. A final consistent result, and a puzzling one, is the apparent positive effect of Republican governors on abortion rates. We can offer no explanation for this result beyond suggesting that it may reflect that other conditions in states with Republican governors indirectly influence abortion, such as stricter eligibility and work requirements for welfare or increased incarceration rates for young men.[*]

Some differences in the two models' results indicate that a number of variables are sensitive to the inclusion of state-time controls for unobserved heterogeneity. The effect of the proportion of women aged 15-19 becomes significant when state-time interactions are included. The effects of access to obstetrician-gynecologists, congressional voting records and Medicaid generosity become marginally significant, although for Medicaid generosity, the effect is in the opposite direction of what would be expected. As anticipated, the effect of parental consent or notification laws loses significance.

In the analysis showing effects on birthrates, three consistent results appear. First, the effect of obstetrician-gynecologist availability is significant (although only marginally so in the second regression) and negative, which suggests that the effects of contraceptive cost captured by this variable dominate those of delivery costs. Second, several economic and policy-related variables are positively associated with birthrates. Women's and men's wages and AFDC benefits have significant positive coefficients in both models; women's property income has a significant positive effect in the first model. Men's property income also has positive coefficients, hut the results fall slightly short of attaining statistical significance. The positive coefficients for AFDC benefits and the men's variables are anticipated. However, the positive associations of women's wages are somewhat surprising and suggest that the income effects of wages dominate the effects of opportunity costs. Third, the proportion of teenagers and women aged 35-44 are each estimated to have significant negative effects on birthrates in both models. As in the abortion regressions, the state-time interactions are jointly significant (not shown).

For the remaining variables, either the sign of the coefficient changes between models or the measure attains or loses statistical significance. These inconsistencies prevent us from drawing firm conclusions.

Alternative Access Measures

In the following analyses, we modify the access measures, but retain the other explanatory variables. We present only the coefficients for the alternative measures, since those for the other explanatory variables do not change appreciably across the respecifications.

Table 3 shows the results of analyses using alternative measures of access to health services. For the first set of regressions, all measures of service access except the proportion of women living in counties with an abortion provider are excluded. The estimated effects of provider access on both abortion rates and birthrates are consistent with those shown in Table 2,

In the next set of regressions, the variables describing the proportions of women living in counties with abortion providers, family planning clinics and obstetrician-gynecologists are replaced with logarithms of the statewide numbers of these providers per 1,000 women. The statewide counts may capture problems of lines and waiting at facilities more accurately than the original measures do. However, given the increasing geographic isolation and urban concentration over the 1980s of some types of facilities, especially abortion clinics, the statewide count variables may not adequately reflect physical proximity.

As in Table 2, the results in the second panel of Table 3 show that abortion provider access has a significant positive effect on abortion rates. In contrast to our earlier results, however, access to family planning clinics and to obstetrician-gynecologists now have no significant effect in either model, and the magnitude of the coefficients for HMO membership changes somewhat. Two substantial changes are evident in the birthrate regressions. Abortion provider access is now estimated to have a small and nonsignificant effect in each model, while access to family planning clinics is estimated to have a significant negative association with birthrates.

The next regressions examine the effects of the distance to the nearest in-state and out-of-state providers. (Regressions that included either the in-state measure alone or both measures plus the other health service variables were also run; the reported results are not sensitive to either respecification.) These variables represent refined measures of access for women who live in counties without an abortion provider. Consistent with our previous findings, these results show that as distance to an abortion provider (in-state or out-of-state) increases, the abortion rate declines significantly. These measures have no statistically significant effects on birthrates.

In the final set of regressions in Table 3, we replace the variable for the proportion of women living in counties with any abortion provider with variables representing the proportions in counties with small, medium and large facilities (defined as those with annual caseloads of fewer than 25 abortions, 25-400 abortions and more than 400 abortions, respectively). By doing this, we can examine the extent of endogeneity bias in the abortion access results. (Specifically, smaller facilities, which presumably have lower fixed operating costs, are going to be more sensitive to short-term shifts in abortion demand than their larger counterparts.) The estimates, which indicate that the access results in the longitudinal models are driven primarily by large facilities, provide evidence that our findings reflect the effects of exogenous variation. These measures generally have no statistically significant effects on birthrates.

Alternative Economic Measures

As with the access measures, we conducted analyses based on alternative sets of economic variables; Table 4 (page 58) shows the results. For the first set of regressions, gender-specific unemployment rates replace the CPS wage measures. Because the unemployment rates are based on representative state-level surveys, they provide a check on whether the wage results come about because of possible problems with representativeness in the CPS data. The unemployment variables also measure an alternative dimension of labormarket opportunities.
In the abortion regressions, the results confirm our earlier findings that women's wages have no significant effects. However, whereas Table 2 showed positive but nonsignificant effects of men's wages, the coefficient in the first model in Table 4 is negative and significant, implying a positive effect of men's wages on abortions. (While this result runs counter to the predictions of most economic models, it conforms with some recent game theoretic results.[24]) As for birthrates, assuming that wages are higher when employment is high and vice versa, the estimates support our earlier findings of a positive association between both women's and men's wages and birthrates.

For the next set of regressions, we substituted measures of annual earnings in the retail and manufacturing sectors for the wage variables for women and men, respectively. Like the unemployment measures, the sector-specific earnings variables are representative at the state level and may better capture both wage and employment opportunities. However, they may only roughly approximate gender-specific opportunities. Both models indicate that retail earnings have a strong positive effect on abortion rates, while manufacturing earnings have no significant effect. Assuming that retail and manufacturing earnings are suitable proxies for women's and men's salaries, the results suggest that the abortion rate is influenced more by effects associated with either the immediate affordability of abortions or women's opportunity costs than by the standard income effects (long-term resources available for childrearing).

Large effects of opportunity costs, however, seem to be ruled out by estimates indicating strong positive effects of retail earnings on birthrates; these results support our earlier findings regarding the positive effects of women's wages. The findings for manufacturing earnings support our earlier results for men's wages, in that the coefficients are both positive, but they differ in that only the coefficient for the second model attains statistical significance.

In the last set of regressions in Table 4, we have replaced the gender-specific property income variables with a single measure of total per capita personal income. Because we feel that the total income variable is problematic, this exercise is performed mostly for comparability with previous research. As in the results reported by Blank and colleagues,[25] per capita income has a strong, positive association with abortion rates. Total per capita income also has a large and significantly positive effect on birthrates in both models. The pattern of results is consistent with the income variable, capturing the effects of immediate affordability for abortions and either immediate affordability of obstetric services or long-term affordability of births. Interestingly, the respecification leads to a negative and significant estimate for women's wages in the abortion model with only state dummies, a result that accords with the positive effects found for births.

Alternative Dependent Variables
As a final sensitivity analysis, we examined alternative regressions in which the dependent variable was the natural logarithm of the sum of the abortion rate and the birthrate (the approximate pregnancy rate); the number of abortions by state of occurrence; and the proportion of pregnancies that end in abortion. For brevity, detailed results from these regressions are not reported here, but are available upon request from the authors. Although some changes occurred across specifications, the results for the pregnancy models are mostly similar to those reported here for births; the few noticeable differences can be traced to differences in results for abortion rates and birthrates. The results for the other two outcomes generally conform with our estimates for abortions by state of residence.

Discussion
For the period 1988-1992, AGI documented substantial declines in both the incidence and the availability of abortions in the United States.[26] Nationwide, the abortion rate fell by 5%, from 27.3 abortions per 1,000

women of reproductive age to 25.9 per 1,000. At the same time, the proportion of women living in counties with an abortion provider fell from 71% to 69%, and the number of providers per 100,000 women fell from 4.4 to 4.0.[*] Our results, which consistently indicate that decreased access to providers leads to lower rates of abortion, conform with the broad directions of these trends. They also can be used to determine how much of the decline in the incidence of abortion was directly attributable to reductions in access

The dependent variables in our models are expressed as logarithms. Accordingly, coefficients for the variables expressed as levels indicate the estimated percentage change of the abortion rate or birthrate due to a unit change in the independent variable. Coefficients for variables expressed as logarithms indicate the percentage change in the dependent variable due to a percentage change in the independent variable.

Thus, given the coefficient in the second model in Table 2 for the proportion of women living in counties with an abortion provider (.587), the two-point drop in this measure between 1988 and 1992 reduced the abortion rate by an estimated 1.2%, to 27.0 abortions per 1,000. Similarly, the 9% decline in the number of providers per 100,000 women, applied to the coefficient for this variable in the second column of Table 3(.170), suggests that the decline in accessibility reduced the abortion rate by 1.5%, to 26.9 per 1,000. Hence, our results indicate that decreased access accounted for roughly 24-30% of the 5% decline in the abortion rate. Our data also indicate that the drop in the proportion of women in their late teens accounts for most of the remaining decline.[†]
In addition, the regression results indicate that the availability of reproductive health services affects fertility. Decreased access to obstetrician-gynecologists is associated with higher birthrates, and more restrictive abortion policies and greater access to obstetric and gynecologic services are consistently estimated to have negative effects on abortion rates, although the results are not significant in all specifications.

Our findings suggest that higher wages for women and men and more generous AFDC benefits are associated with increased fertility rates. There is also weak evidence that higher property incomes for women and men are linked to elevated birthrates. The estimated effects of women's wages are a little surprising because they imply that income effects dominate opportunity-cost effects for women's fertility decisions; however, our sensitivity tests indicated that the results for women's wages are robust.

For abortion outcomes, the economic results are much more sensitive to the particular measures employed, and no consistent pattern appears. The AFDC results suggest that the effects of welfare on reproductive behavior have been greatly exaggerated by some in the current policy debate. According to these findings, a 10% cut in AFDC benefits would have reduced the birthrate by 0.45%, implying a reduction of roughly one birth for every 212 women receiving AFDC.[‡]

Because we have presented estimates of abortion and birth models in parallel, the results can be used to infer whether changes in abortions and births reflect changes in contraceptive effort. For instance, the finding that higher wages increase birthrates but do not significantly affect abortion rates suggests that when women's economic conditions improve, their contraceptive effort declines. Similarly, in the models including interactions between state and time, a change in contraceptive effort resulting from access to abortion can be inferred from the results indicating that access increases abortion rates but does not significantly lower birthrates.
Several implications for public policy emerge from our results. The consistent findings regarding the negative effects of obstetrician-gynecologist availability on birthrates suggest that convenient access to prescription contraceptives is an important determinant of reproductive behavior. This interpretation receives modest support from weaker evidence indicating that obstetrician-gynecologist availability may also reduce the incidence of abortions.

Our results also indicate that policies that either expressly or indirectly reduce women's access to abortion services decrease their use of the procedure. The Supreme Court has generally held regulations to be invalid if they place substantial obstacles in the path of women seeking abortions prior to fetal viability. The contentious issue of whether policies go too far in restricting access is being resolved by the Court under its standard of "undue burden." While the Court has applied its test one restriction at a time, our findings of independent effects from several aspects of availability suggest another approach — namely, that the standard be broadened to consider the entire constellation of restrictions and factors affecting abortion access within states.

*Geographic heterogeneity in the effective costs of reproductive health services seems likely to increase in view of a series of Supreme Court decisions increasing states' authority to regulate abortion and the government's general authority to control procedures within publicly funded family planning clinics. See: Webster v. Reproductive Health Services, 442 U.S. 490 (1989); Hodgson v. Minnesota, 497 U.S. 417(1990); Ohio v. Center for Reproductive Health, 497 U.S. 502 (1990); Planned Parenthood of Southeastern Pennsylvania v. Casey, 112 S. Ct. 1759 (1991); and Rust v. Sullivan, 59 U.S. L. W. 4451 (May 21, 1991).
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*This factor remained significant in models including alternative state policy, health service access and economic controls. In regressions disaggregated overtime, the effect was significantly positive in all periods and became stronger in later periods, results that run counter to the evolution of the Republican abortion platform. Blank and colleagues also obtained counterintuitive results for some party affiliation variables (see: reference 2).

*AGI reported that 70% of women lived in counties with an abortion provider in 1992; we believe that this figure, which is based on the number of women recorded in the 1990 census rather than the 1992 intercensal estimate, is slightly high.

† AGI attributed a smaller role to age composition; the difference in the proportion of teenagers between the 1990 census and the 1992 intercensal estimates accounts for much of the difference between AGI's conclusion and ours.

‡In 1988, a 0.45% drop in the birthrate would have meant 0.3 fewer births per 1,000 women. That year, 58.1 million women were of reproductive age and 3.7 million families received AFDC. Thus, approximately 6.37% (or 63.7 per 1,000) women of reproductive age were receiving AFDC. Assuming that the reduction in births occurred only among these women, there would have been 0.3 fewer births per 63.7 AFDC recipients of reproductive age, or one birth fewer per 212 AFDC recipients.

Corrections

In "The Effectiveness of the Yuzpe Regimen of Emergency Contraception" [28:58-64 & 87], by James Trussell, Charlotte Ellertson and Felicia Stewart, the number of women listed in Table 3 (p. 61) as having been treated in the Percival-Smith et al. 1987 study should be 648, not 622, and the total number of women treated in all studies included in the table should be 3,539, not 3,534.
In "Determinants of Early Implant Discontinuation Among Low-Income Women" [28:256-260], by Debra Kalmuss et al., Table 4 (p. 259) shows logistic regression coefficients, not odds. The heading should read: Logistic regression coefficients indicating impact of characteristics on implant discontinuation within six months, by whether regression includes implant side effects.

Table 1. Variables used in regression analyses of factors influencing abortion rates and birthrates, and nationally representative means (and standard deviations)
Legend for Chart:

A - Variable Mean
B - Outcome
A: Abortion rate (by stale of residence)
B: 28.12 (10.54)
A: Birthrate
B: 66.58 (7.93)
Service accessibility
A: Proportion of women in counties with abortion providers
B: 0.72 (0.22)
A: Proportion of women in counties with small providers[†]
B: 0.53 (0.27)
A: Proportion of women in counties with medium providers[‡]
B: 0.62 (0.25)
A: Proportion of women in counties with large providers[§]
B: 0.57 (0.22)
A: No. of abortion providers per 1,000 women
B: 0.05 (0.03)
A: Avg. miles to nearest in-state provider (hundreds)[‡]
B: 0.14 (0.15)
A: Avg. miles to nearest out-of-state provider (hundreds)[‡]
B: 0.94 (0.68)
A: Proportion of women in counties with family planning clinics
B: 0.92 (0.10)
A: No. of family planning clinics per 1,000 women
B: 0.10 (0.05)
A: Proportion of women in counties with obstetrician-gynecologists
B: 0.91 (0.10)
A: No. of obstetrician-gynecologists per 1,000 women
B: 0.48 (0.11)
A: Per capita HMO membership
B: 0.07 (0.07)
Economic
A: Avg female property income (000s of S)
B: 0.29 (0.12)
A: Avg. male property income (000s of $)
B: 0.36 (0.11)
A: Avg. selectivity-adjusted imputed female wage ($)
B: 4.44 (1.06)
A: Avg. unadjusted female wage ($)
B: 7.34 (0.76)
A: Avg. selectivity-adjusted imputed male wage ($)
B: 9.38 (1.14)
A: Avg. unadjusted male wage ($)
B: 10.41 (0.91)
A: Avg. per capita total personal income
B:(000s of $) 14.53 (2.05)
A: Avg. annual retail earnings (000s of $)
B: 12.91 (1.36)
A: Avg. annual manufacturing earnings (000s of $)
B: 28.30 (4.01)
A: Female unemployment rate
B: 0.07 (0.02)

A: Male unemployment rate
B: 0.07 (0.02)
State policy
A: Wedicaid abortion funding restrictions[‡‡]
B: 0.52 (0.48)
A: Parental consent or notification law[‡‡]
B: 0.11 (0.31)
A: Maximum monthly AFDC benefits for family of four ($)
B: 474.08 (182.87)
A: Avg. monthly Medicaid benefits per AFDC recipient ($)
B: 63.69 (18.07)
Political and attitudinal
A: Republican governor[‡‡]
B: 0.44 (0.50)
A: Antiabortion voting index for state's congressional delegation[§§]
B: 0.47 (0.23)
A: Antiabortion attitude index‡‡]
B: 4.21 (0.84)
Demographic and other
A: Proportion of population black
B: 0.12 (0.08)
A: Women 15-19 as proportion of women of reproductive age
B: 0.18 (0.02)
A: Women 35-44 as proportion of women of reproductive age
B: 0.27 (0.02)
A: No. of rapes per 100,000 individuals
B: 35.44 (12.33)
†Providers that perform fewer than 25 abortions pet year. ‡Providers that perform 25-400 abortions per year. §Providers that perform more than 400 abortions per year ‡‡Excluding Alaska and Hawaii. ‡‡When states with this variable in effect are coded as 1.0. §§When pro-choice votes are coded as 1.0. ‡‡On a scale of 0-12. with 0 indicating approval of each of six circumstances for abortion and 12 indicating approval for none. Notes: In this and subsequent tables, unless noted otherwise, statistics are based on 1976-1982, 1984-1985 and 1987-1988 data from the 50 states and the District of Columbia; data are weighted by the number of women aged 15-44 in 1980 in each slate; and percentages and rates are based on women aged 15-44. For a detailed description of the data and sources, see the appendix (page 59).

Table 2. Ordinary least-squares regression results (and standard errors) indicating effects of selected variables on state abortion rates and birthrates
Legend for Chart:
A - Variable
B - Abortion rate: Fixed state-time effects
C - Abortion rate: Adj. for state-time interaction

D - Birthrate: Fixed statetime effects
E - Birthrate: Adj. for statetime interaction
Service accessibility
A: % of women in counties with abortion providers
B: 0.329[**] (0.129)
C: 0.587[***] (0.150)
D: -0.089[**] (0.044)
E: 0.021 (0.035)
A: % of women in counties with family planning clinics
B: 0.120[*] (0.066)
C: 0.120 (0.091)
D: 0.004 (0.022)
E: 0.008 (0.021)
A: % of women in counties with obstetrician-gynecologists
B: -0.372 (0.255)
C: -0.523[*] (0.302)
D: -0.814[***] (0.087)
E: -0.136[*] (0.071)
A: Per capita HMO membership
B: 0.698[***] (0.218)
C: -0.180 (0.418)
D: 0.182[**] (0.074)
E: 0.100 (0.099)
Economic
A: Log of female property income
B: -0.007 (0.013)
C: -0.019 (0.012)
D: 0.009[**] (0.004)
E: 0.001 (0.003)
A: Log of male property income
B: -0.009 (0.016)
C: -0.011 (0.015)
D: 0.009 (0.006)
E: 0.005 (0.003)
A: Log of female wage
B: -0.032 (0.078)
C: 0.123 (0.080)
D: 0.152[**] (0.027)
E: 0.080[**] (0.019)
A: Log of male wage
B: 0.166 (0.103)
C: 0.083 (0.105)
D: 0.178[***] (0.035)
E: 0.053[**] (0.025)
State policy
A: Medicaid abortion funding restrictions
B: -0.056[***] (0.020)
C: -0.029 (0.022)
D: -0.019[***] (0.007)
E: -0.005 (0.005)
A: Parental consent or notification law
B: -0.032[**] (0.016)

C: -0.012 (0.021)
D: -0.021[***] (0.006)
E: 0.008[*] (0.005)
A: Log of AFDC benefits
B: -0.037 (0.051)
C: 0.025 (0.069)
D: 0.124[***] (0.017)
E: 0.045[***] (0.016)
A: Log of Medicaid benefits
B: -0.016 (0.024)
C: 0.049[*] (0.029)
D: -0.022[***] (0.008)
E: 0.025[***] (0.007)
Political and attitudinal
A: Republican governor
B: 0.051[***] (0.009)
C: 0.029[***] (0.010)
D: 0.005 (0.003)
E: -0.001 (0.002)
A: Antiabortion congressional voting index
B: -0.029 (0.034)
C: -0.064[*] (0.035)
D: -0.014 (0.012)
E: -0.009 (0.008)
A: Antiabortion attitude index
B: 0.008 (0.010)
C: 0.011 (0.010)
D: 0.001 (0.003)
E: -0.002 (0.002)
Demographic and other
A: % of population black
B: 2.552[**] (1.161)
C: -9.734 (6.149)
D: 1.163[***] (0.397)
E: -0.157 (1.450)
A: % of women 15-19
B: 1.998 (1.334)
C: 3.928[**] (1.858)
D: -2.667[***] (0.456)
E: -2.329[***] (0.438)
A: % of women 35-44
B: -3.853[***] (0.951)
C: -0.925 (2.903)
D: -1.593[***] (0.325)
E: -3.140[***] (0.684)
A: Log of rape rate
B: -0.008 (0.033)
C: 0.025 (0.048)
D: -0.075[***] (0.011)
E: 0.037[***] (0.011)
A: R2
B: 0.974

C: 0.984
D: 0.966
E: 0.990
*p<-10. **p<-05. ***p<.01. Notes In this and subsequent tables, the dependent variables are the natural logarithms of the state-level abortion rate and birthrate. All regressions include year-specific dummy variables.

Table 3. Ordinary least-squares regression results (and standard errors) indicating effects of various measures of access to reproductive health services on abortion rates and birthrates
Legend for Chart:
A - Variable
B - Abortion rate: Fixed statetime effects
C - Abortion rate: Adj. for statetime interaction
D - Birthrate: Fixed statetime effects
E - Birthrate: Adj. for statetime interaction
Excluding other access measures
A: % of women in counties with abortion providers
B: 0.396[***] (0.130)
C: 0.573[***] (0.150)
D: -0.098[**] (0.049)
E: 0.023 (0.035)
"Per woman" access
A: Log of abortion providers per 1,000 women
B: 0.131[***] (0.035)
C: 0.170[***] (0.037)
D: 0.010 (0.013)
E: 0.0003 (0.009)
A: Log of family planning clinics per 1,000 women
B: -0.007 (0.020)
C: -0.015 (0.044)
D: -0.042[***] (0.007)
E: -0.037[***] (0.010)
A: Log of obstetrician-gynecologists per 1,000 women
B: -0.194 (0.120)
C: -0.207 (0.148)
D: -0.189[***] (0.043)
E: -0.041 (0.034)
A: Per capita HMO membership
B: 0.534[**] (0.225)
C: -0.541 (0.427)
D: 0.202[**] (0.080)
E: 0.052 (0.099)
Abortion provider distance[†]
A: Avg. miles to nearest in-state provider
B: -0.217[***] (0.081)
C: -0.136[*] (0.079)
D: -0.045 (0.031)
E: -0.010 (0.019)
A: Avg. miles to nearest out-of-state provider
B: -0.155[**] (0.070)
C: -0.270[***] (0.083)

D: 0.015 (0.027)
E: -0.008 (0.019)
Size of abortion provider
A: % of women in counties with small providers
B: 0.012 (0.041)
C: 0.045 (0.046)
D: 0.0002 (0.014)
E: -0.020[*] (0.011)
A: % of women in counties with medium providers
B: 0.032 (0.066)
C: 0.010 (0.069)
D: -0.016 (0.022)
E: 0.017 (0.016)
A: % of women in counties with large providers
B: 0.190[**] (0.087)
C: 0.157[*] (0.090)
D: 0.035 (0.030)
E: 0.017 (0.021)
A: % of women in counties with family planning clinics
B: 0.142[**] (0.067)
C: 0.124 (0.093)
D: 0.003 (0.023)
E: 0.008 (0.021)
A: % of women in counties with obstetrician-gynecologists
B: -0.361 (0.256)
C: -0.549[*] (0.310)
D: -0.839[***] (0.088)
E: -0.148[**] (0.071)
A: Per capita HMO membership
B: 0.744[***] (0.218)
C: -0.050 (0.428)
D: 0.159[**] (0.075)
E: 0.092 (0.099)
*p<.10. **p<.05. ***p<.01. †Excluding Alaska and Hawaii. Note: All regressions also include coefficients for economic, state policy, political and attitudinal, demographic and other, and time variables, as in Table 2.

Table 4. Ordinary least-squares regression results (and standard errors) indicating effects of various economic measures on abortion rates and birthrates
Legend for Chart:
A - Variable
B - Abortion rate: Fixed statetime effects
C - Abortion rate: Adj. for statetime interaction
D - Birthrate: Fixed statetime effects
E - Birthrate: Adj for state-time interaction
Unemployment
A: Log of female property income
B: -0.012 (0.013)
C: -0.018 (0.012)
D: 0.009[**] (0-004)

E: 0.001 (0.003)
A: Log of male property income
B: -0.011 (0.016)
C: -0.009 (0.015)
D: 0.012[**] (0-005)
E: 0.006[*] (0.003)
A: Female unemployment rate
B: -0.271 (0.575)
C: -0.129 (0.541)
D: -0.819[***] (0.196)
E: -0.633[***] (0.121)
A: Male unemployment rate
B: -1.039[**] (0.476)
C: -0.538 (0.472)
D: -0.379[**] (0.162)
E: -0.093 (0.106)
Wages
A: Log of female property income
B: -0.009 (0.013)
C: -0.019 (0.012)
D: 0.012[***] (0.004)
E: 0.001 (0.003)
A: Log of male property income
B: -0.012 (0.016)
C: -0.012 (0.014)
D: 0.008 (0.005)
E: 0.004 (0.003)
A: Log of average annual retail earnings
B: 0.467[**] (0.185)
C: 1.002[***] (0.287)
D: 0.496[***] (0.059)
E: 0.391[***] (0.063)
A: Log of average annual manufacturing earnings
B: -0.150 (0.220)
C: -0.065 (0.300)
D: 0.068 (0.069)
E: 0.270[***] (0.066)
Income
A: Log of per capita total personal income
B: 0.650[***] (0.131)
C: 0.975[***] (0.197)
D: 0.330[***] (0.044)
E: 0.412[***] (0.042)
A: Log of female wages
B: -0.170[**] (0.079)
C: -0.009 (0.081)
D: 0.100[***] (0.026)
E: 0.332[*] (0.017)
A: Log of male wages
B: -0.014 (0.105)
C: 0.059 (0.102)
D: 0.101[***] (0.035)

E: 0.050[**] (0.022)
*p<.10. **p<.05. **p<.01 Note: All regressions also include coefficients for provider availability, state policy, political and attitudinal, demographic and other, and time variables, as in Table 2.

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Appendix: Data Sources

Reproductive Outcomes

Abortion rates. Abortion rates per 1,000 women of reproductive age by state of occurrence and state of residence, 1978-1982, 1984-1985 and 1987-1988, were obtained from S. K. Henshaw and J. Van Vort, 1994 (see: reference 26); and AGI, unpublished data.

Birthrates. Annual county-level data on number of births were obtained from the Area Resource File (see: reference 18), aggregated to state levels and divided by the number of women of reproductive age.

Health Services

Abortion providers. Numbers of providers of various sizes, by county, 1978-1982, 1984-1985 and 1987-1988, were obtained from AGI. These data were combined with data on women of reproductive age to calculate the number of women in counties with each type of provider. The number of providers and number of women in counties with providers were aggregated to state levels and divided by the total number of women of reproductive age to obtain the variables used in the analysis.

Distance to abortion providers. AGI county-level abortion provider data were combined with distances calculated using the latitude and longitude coordinates of the population-weighted centroids of each U.S. county at the time of the 1980 census. Distances between each pair of centroids were calculated along "great circle arcs." (See: A. Robinson et al., Elements of Cartography, fifth ed., John Wiley and Sons, New York, 1994, pp. 74-75.) For each county, distances to the nearest in-state and out-of-state counties with abortion providers were calculated (if the county had a provider, the distance to the nearest in-state provider was set at one mile). State averages of these distances, weighted by the number of women of reproductive age, are used in the analysis.

Family planning clinics and obstetrician-gynecologists. County-level numbers of clinics in 1975, 1981 and 1983 were obtained from AGI; county-level numbers of nonfederal obstetrician-gynecologists in patient care in 1977-1981, 1983, 1985-1986 and 1988-1989 were obtained from Quality Resource Systems, Fairfax, Va. Quadratic interpolation (rounded to integers and constrained to nonnegative values) was used to impute provider counts for intermediate years and, for clinics, to extrapolate data after 1983. Actual and imputed data were combined with data on women of reproductive age to calculate the number of women in counties with providers. The number of providers and number of women in counties with providers were aggregated to state levels and divided by the total number of women of reproductive age to obtain the variables used in the analysis.
Per capita HMO membership. Annual county-level numbers of HMO members were obtained from the Area Resource File (see: reference 18), aggregated to state levels and divided by the total state population.

Economic Variables
Properly income. Individual-level interest, dividend, property, and estates and trusts income for women aged 15-44 and men aged 15-54 were obtained from the 1979-1989 March Current Population Survey (CPS) files (see: reference 22). Consistent top-coding was applied across all years, and the top-coded observations were eliminated. Data were aggregated to state and annual levels using the March CPS weights and then deflated using the personal consumption deflator (PCD).
Wages. Individual-level hours and wage- or salary-income data for women aged 15-44 and men aged 15-54 were obtained from the March CPS files. Self-employed workers, farmers, unpaid workers and individuals with missing or inconsistent information were excluded. Consistent top-coding was applied, and the top-coded observations were eliminated. Hourly wages were computed by dividing annual earnings by annual hours and deflating by the PCD. Individuals whose deflated hourly wages were less than 25 cents or more than $250 were excluded. Wage regressions were estimated using a standard two-stage selectivity-correction procedure. In the first stage, individual-level employment probit models were estimated for women and men separately by state. In the second stage, the logarithms of hourly wages were regressed on the predicted inverse Mill's ratio; on annual dummy variables; on dummy variables for black and other origins; on potential work experience; on experience squared; on years of elementary, secondary and postsecondary schooling; and on interactions of experience, secondary schooling and postsecondary schooling with quadratic time trends. Like the probit models, the wage regressions were run separately by state. Results from the regressions were used to predict wages for the entire CPS sample (workers and nonworkers). Predicted wages were aggregated to state levels using the March weights.
Per capita income and average annual earnings. Annual state-level data on total personal income and on total earnings and employment in the manufacturing and retail sectors were obtained from the U.S. Bureau of Economic Analysis, 1992 (see: reference 23). Personal income data were deflated by the PCD; total earnings were divided by total employment and deflated by the PCD to obtain average earnings.

Policy Variables
Medicaid funding restrictions. States with enforced Medicaid funding restrictions similar to those of the Hyde Amendment were coded 1, and other states were coded 0. States with restrictions for part of the year were assigned a fraction representing the amount of time the restriction was in effect. The data come primarily from J. Merz, "A Review of Abortion Policy: Legality, Medicaid Funding, and Parental Involvement, 1967-1994," Rand, Santa Monica, Calif., 1994. Where possible, the data have been reconciled with: D. Bush, "Fertility-Related State Laws Enacted in 1982," Family Planning Perspectives, 15:111-116, 1983; J. Currie et al., 1996 (see: reference 17); R. B. Gold, "Publicly Funded Abortions in FY 1980 and 1981," Family Planning Perspectives, 14:204-207, 1982; R. B. Gold and S. Guardado, "Public Funding of Family Planning, Sterilization and Abortion Services, 1987," Family Planning Perspectives, 20:228-233, 1988; R. B. Gold and J. Macias, "Public Funding of Contraceptive, Sterilization and Abortion Services, 1985," Family Planning Perspectives, 18:259-264, 1986; R. B. Gold and B. Nestor, "Public Funding of Contraceptive, Sterilization and Abortion Services, 1983," Family Planning Perspectives, 17:25-30, 1985; and B. Nestor and R. B. Gold, "Public Funding of Contraceptive, Sterilization and Abortion Services, 1982," Family Planning Perspectives, 16:128-133, 1984.
Parental notification and consent laws. States with enforced parental notification or consent laws were coded 1, and other states were coded 0. States with restrictions for part of the year were assigned a fraction representing

the amount of time the restriction was in effect. The data come primarily from J. Merz, "A Review of Abortion Policy: Legality, Medicaid Funding, and Parental Involvement, 1967-1994," Rand, Santa Monica, Calif., 1994. Where possible, the data have been reconciled with: D. Bush and P. Donovan, "Fertility-Related State Laws Enacted in 1981," Family Planning Perspectives, 14:63-67, 1982; M. D. Greenberger and K. Connor, "Parental Notice and Consent for Abortion: Out of Step with Family Law Principles and Policies," Family Planning Perspectives, 23:31-35, 1991; and T. Sollom and P. Donovan, "State Laws and the Provision of Family Planning and Abortion Services in 1985," Family Planning Perspectives, 17:262-266, 1985.

AFDC benefits. Data were obtained from the following sources and deflated by the PCD: U.S. House of Representatives, Committee on Ways and Means, Green Book: Background Material and Data on Programs Within the Jurisdiction of the Committee on Ways and Means, U.S. Government Printing Office (GPO), Washington, D.C., 1981-1990; and U.S. Social Security Administration, Characteristics of State Plans for Aid to Families with Dependent Children, GPO, Washington, D.C., various years.

Medicaid benefits. Annual unpublished state-level data from the U.S. Health Care Financing Administration on medical vendor payments made on behalf of AFDC recipients were divided by the number of AFDC recipients in each state, converted into monthly amounts and deflated by the PCD. The resulting figures approximate the insurance value of Medicaid to the average AFDC recipient.

Political Variables

Republican governor. States with a Republican governor were coded 1, and other states were code d 0. (The District of Columbia was coded according to its mayor's party affiliation.) Data were obtained from R. Glashan, American Governors and Gubernatorial Elections, 1775-1978, Meckler Books, Westport, Conn., 1979; and M. Mullaney, American Governors and Gubernatorial Elections, 1979-1987, Meckler Books, Westport, Conn., 1988.

Congressional delegation voting index. Data on key abortion votes were obtained from National Committee for a Human Life Amendment, Key Votes on Abortion, U.S. Senate and U.S. House of Representatives, Washington, D.C., 1978-1989. Votes were coded 0 if prochoice, ½ if a nonvote (e.g., abstentions, votes of "present") and 1 if antiabortion, as reported in Inter-University Consortium for Political and Social Research, United Slates Congressional Roll Call Voting Records, University of Michigan, Ann Arbor, 1989. Annual averages of these indexed votes were constructed for each state's House and Senate delegation. A final index was formed using a simple average of the chamber-specific figures.

Antiabortion attitudes index. In 1978, 1980 and each year from 1982 to 1988, the General Social Survey asked whether individuals approved of abortions under the following circumstances: the child would have a serious birth defect, the mother was married but wanted no more children, the mother's health was in danger, the mother could not afford more children, the pregnancy resulted from rape, and the mother was single and the father did not agree to marry her (see: National Opinion Research Center, General Social Survey, 1972-1997: Cumulative Codebook, University of Chicago Press, Chicago, 1992). For each condition, individuals who approved were coded 0, individuals with no opinion were coded 1 and individuals who disapproved were coded 2. The scores for each individual were summed to form an index of opposition to abortion (Cronbach's alpha=.85), and this index was averaged across regions. Quadratic interpolation was used to impute data for 1979 and 1981.

Demographic and Other Characteristics

Age and race. Annual county-level number of women in each age-group and number of black women were obtained from the following source and aggregated to state levels: U.S. Bureau of the Census, Intercensal Estimates of the Population of Counties by Age, Sex, and Race: 1970-1980 and 1980-1989, Washington, D.C., 1984 and 1992.

Rape rates. Annual state-level data were obtained from U.S. Federal Bureau of Investigation, Crime in the United States, GPO, Washington, D.C., 1978-1988.