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State Abortion Policy, Geographic Access to Abortion Providers and Changing Family Formation

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Context: One of the goals in cutting welfare payments and setting time limits on welfare receipt is the reduction of out-of-wedlock childbearing among poor women. Yet such changes may increase the demand for abortion at the same time that access to abortion has decreased, throwing into doubt the potential effect of these changes on the proportion of women who are heading families.

Methods: State and county fixed-effects models were used to estimate the effects of factors influencing abortion availability—geographic access, parental notification requirements and Medicaid funding restrictions—on the county-level proportion of women heading households.

Results: The decline in geographic access to abortion providers during the 1980s accounted for a small but significant portion of the rise in the percentage of women heading families (about 2%). Restrictions on Medicaid funding for abortion accounted for about half of the increase in female headship among blacks, while new state parental notification requirements contributed modestly to the rise in the proportion of white women heading single-parent families.

Conclusions: Welfare reform legislation and attempts to reduce the availability of abortion services in the United States appear to be working at cross-purposes. Cutbacks in access to abortion may have contributed modestly to the increase in the proportion of women heading households.

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Current government policies seemingly reflect mixed if not contradictory goals for the American family. The newly implemented Personal Responsibility and Work Opportunity Reconciliation Act of 1996 is aimed at strengthening the traditional two-parent family while discouraging out-of-wedlock childbearing. New time limits on welfare receipt and mandated work requirements have imposed additional "costs" on unmarried childbearing.

An explicit aim of the legislation is to promote economic self-sufficiency among welfare-dependent single mothers, while also reducing the share of children in poverty. Some policymakers believe that the knowledge that welfare is less generous than in the past may motivate sexually active unmarried women to become better contraceptive users and encourage pregnant single women to marry their partners.

Time-limited welfare, mandated work requirements and the imposition of family caps on benefits (in some states) may also have the unintended effect of increasing the demand for abortion services among low-income women with unplanned pregnancies.[1] At the same time, many states are passing laws aimed at restricting geographic and legal access to reproductive health and abortion services. Recent Supreme Court decisions now allow states to require abortion providers to notify parents of abortions performed on minors, to impose restrictions on Medicaid funding for abortion and to create 24-hour waiting periods. An unintended effect of such restrictions may be an accelerated growth in nonmarital births that, in turn, increases the proportion of unmarried women heading families.

Our study addresses a straightforward question with important implications for public policy: Are new barriers to abortion access likely to contribute to increases in the proportion of women who head households in the United States? In this article, we estimate state and county fixed-effects models of the impact of geographic access to abortion providers, parental consent and notification requirements, and of Medicaid funding restrictions on recent changes in family headship rates among women.

BACKGROUND

Few observers disagree that shortages of local-area abortion providers and the imposition of new legal restrictions on abortion mean that fewer women can resolve an unintended pregnancy through abortion. One recent study found that observed declines in the proportion of women living in counties with abortion providers reduced the abortion rate by 1.2% between 1988 and 1992.[2] Another study reported that abortions are reduced by nearly 25% among low-income women when states restrict Medicaid funding.[3]

Abortion access is especially relevant among unmarried pregnant women. Indeed, data from the 1995 National Survey of Family Growth suggest that only 44% of all pregnancies among never-married women are intended, compared with 81% among married women.[4] Not surprisingly, abortions also occur disproportionately among unmarried women.

Whether restrictions on abortion lead to increases or decreases in female headship, however, is ambiguous. The conventional view is that restrictive abortion policies or other barriers to abortions will lead to higher fertility, especially among unmarried women, and to an increased share of unmarried women heading families with children.

Another, less common view is that new restrictions on abortion will instead reduce unmarried childbearing and female family headship because legal restrictions and the lack of geographic access to abortion may increase women's motivation to avoid unwanted pregnancy.[5] Thus, restrictions may lower both demand for abortion services and nonmarital fertility rates.

Proponents of the conventional view argue that the lack of geographic access to reproductive health care or abortion providers increases the economic costs (e.g., out-of-pocket travel expenses) and information costs of obtaining an abortion.[6] As these costs increase, abortion rates decline and alternative pregnancy resolutions become more likely to be adopted, including childbearing (both in-wedlock and out-of-wedlock).

Moreover, parental notification or consent requirements seem to discourage abortion among pregnant teenagers, especially if parental involvement increases the likelihood of other resolutions to unwanted pregnancy.[7] Pregnant teenagers may choose unmarried childbearing if parents are opposed to abortion for moral or other reasons. Restrictions on Medicaid funding for abortion also mean that unmarried pregnant women, especially those who are economically disadvantaged, will be less able to pay the costs of ending unwanted pregnancies.[8]

Such arguments are commonly used to buttress claims that recent restrictions on abortion have led inexorably to more unintended births and to higher rates of headship among unmarried mothers, many of whom are poor and welfare-dependent. The consequences of new restrictions on abortion access seem anathema to the stated policy goals of recent welfare reform legislation. Newly imposed time limits, family caps on welfare benefits, and restrictions on independent living among minor mothers are aimed at raising the monetary and nonmonetary costs of out-of-wedlock childbearing and female headship. Yet geographic and legal restrictions on abortion access may also have effectively increased the costs of aborting a pregnancy. It is unclear, however, whether the costs of abortion increased in the 1980s compared with the cost of bearing a child outside marriage or of other pregnancy resolutions (e.g., adoption or fosterage).

An alternative view recognizes that decisions regarding how to resolve unintended, nonmarital pregnancies are preceded by the behaviors that led to pregnancy, including decisions regarding whether and how frequently to engage in sexual activity and whether and how to practice contraception. In this view, geographic or legal barriers to abortion not only increase the costs of abortion but also those of unwanted pregnancy, leading unmarried women (and their partners) to take greater precautions to avoid conception. If these effects are strong enough, restrictions on abortion access may reduce nonmarital births and female headship rates.

Indeed, one recent study found that declines in abortion access, including greater distance to an abortion provider and restrictions on Medicaid funding, actually led to small but statistically significant declines in teenage birthrates.[9] Each 100-mile increase in distance to a provider was associated with a decline of two births per 1,000 white teenage females. The closing of abortion clinics between 1977 and 1988 was similarly associated with declines in teenage birthrates, but by slightly less than 0.1%. Such results, however, were less robust among blacks, and geographic and legal restrictions on abortions had smaller effects on birthrates among unmarried teenagers than among married teenagers.

Despite a voluminous literature on the rise in female-headed families, few if any studies have evaluated the role of changing state abortion policies and declining geographic access to reproductive health care, including abortion providers.[10] In this article, we evaluate the effects of county-level changes in abortion access between 1980 and 1990 on the local-area percentage of women heading households with children, using repeated measures of family formation, as well as measures of changes between 1980 and 1990 in state abortion policies and geographic access to abortion providers. Unlike most previous researchers, we adjust our estimates of state abortion policy effects both for the effects of observed county-level social and economic indicators known to be associated with the rise in female headship and for unobserved state and county fixed effects.

DATA AND VARIABLES

Our analysis draws on cross-sectional county records from the summary tape files of the 1980 and 1990 decennial censuses of the United States. We match information for each county across years to form a two-period comparison. Counties from Alaska and Hawaii are excluded because their ethnic composition and cost of living are unrepresentative of the rest of the country. We also eliminated counties with fewer than 100 women of reproductive age (15-44 years) in either 1980 or 1990. The resulting pooled data set contains 6,132 observations (3,066 counties or county equivalents matched across 1980 and 1990).*

The census data, which are aggregated to the county level, have several advantages that are useful for this analysis. First, they show family formation outcomes for all counties in the United States—including rural areas, where reproductive health services are least likely to be available. Second, because the census data are matched longitudinally, our study tracks changes in family formation and its determinants within counties over time and can control for unobserved time-invariant, county-specific factors. Third, they can be linked with other county-level data on abortion and physician availability and can thus be related to appropriate local measures of the availability of reproductive health services. Finally, the census data include race- and ethnicity-specific variables and permit separate analyses for black, Hispanic and non-Hispanic white women.

FEMALE-HEADED HOUSEHOLDS

The dependent variable in our analysis is the percentage of women of reproductive age who are single heads of households with children younger than 18. As an alternative, we conduct some sensitivity analyses using another measure that more closely reflects the total incidence of single parenthood—the percentage of women in each census who are single heads of either households or subfamilies within households. Although the measure that combines family and subfamily headship is more complete than the household headship measure, it still misses some single mothers living with their children (e.g., boarders within rooming houses and some cohabiting mothers). In addition, it is not available by racial and ethnic group.[†]

Previous studies have focused on the determinants of early childbearing, nonmarital childbearing or both. Our approach is more comprehensive. While our measures of female headship reflect nonmarital births that were not followed by a marriage, they also include marital births that were followed by separation and divorce. The focus on headship thus accounts for a direct effect of abortion access on out-of-wedlock childbearing, and also for the possibility that births resulting from unwanted marital pregnancies strain (or strengthen) weak marriages.

Besides being more general, our decision to consider headship, rather than the particular routes to that outcome, is appropriate from the perspective of welfare policy, where the principal concern is the size of the welfare-eligible population. The main drawback to this approach is that the direct effect of abortion access on nonmarital childbearing is not identified. We might be suspicious, for instance, if our results reflected mostly a relationship between abortion availability and divorce (or unmeasured determinants of these two variables). To account for this possibility, we have conducted sensitivity analyses that explicitly control for local divorce rates.

ACCESS TO SERVICES

Our key independent variables, which capture geographic and legal access to reproductive health services, have been gathered from other sources and merged with the census data. Our primary variable for geographic access to abortion services is the number of abortion providers per 1,000 women aged 15-44 years in each county. The local data on number of providers come from surveys conducted in 1979 and 1988 by The Alan Guttmacher Institute (AGI). The AGI data have also been used to form a dummy variable indicating whether any abortion providers were present in the county and to form measures of the distance from the population-weighted geographic center of the county to the similarly defined centers of the nearest in-state and out-of-state counties with providers. These alternative measures are used in some sensitivity analyses.

Legal restrictions on abortion services are measured by the number of years out of the five preceding each census (either 1975-1979 or 1985-1989) that parental notification or consent requirements and Medicaid-funding restrictions were in effect in each state. Annual data on state legal restrictions come from a study by Matthews and colleagues.[11] Five-year histories are used because injunctions by some state courts in the enforcement of these provisions introduce too much variability into simpler single-year measures.

Data from the Bureau of Health Professionals Area Resource File are used to construct measures of the number of active obstetrician-gynecologists involved in patient care per 1,000 women aged 15-44 years in each county. Obstetrician-gynecologists provide a variety of medical services that reduce the incidence of fertility, such as prescribing contraceptives, referring patients for abortions or performing abortions themselves. However, these physicians also monitor pregnancies and perform deliveries. Thus, it is unclear whether they have a net positive or negative effect on fertility and headship rates.

To control for changes in the generosity of public assistance programs, we also use state-level data on welfare benefits in our analysis. Our measure of welfare generosity is the maximum monthly combined benefit (adjusted for inflation) for a family of four with no other income under the Aid to Families with Dependent Children, Food Stamp and Medicaid programs.[12]

Our empirical analysis also includes numerous other independent longitudinal county-level explanatory variables, assembled from the census files and other sources, that control for local marriage opportunities, gender-specific economic opportunities, and other population and institutional characteristics that previous research has shown to be associated with female headship. Specific variables include the sex ratio; men's and women's inflation-adjusted median full-time incomes; men's and women's education; men's employment; the log of the population; and the percentages of the population in each county that are older than 65, black, Hispanic, rural, Catholic, divorced, adherents of the Church of Jesus Christ of Latter Day Saints or antiabortion Protestant. Because these variables are secondary to our present analysis and have already been explicitly considered in a previous published study, we do not discuss them further here.[13]

ANALYTIC APPROACH

We fit regression models of county female-headship rates that include the measures of access to abortion and reproductive health services and other variables noted above as explanatory variables. Each of the regressions also includes either state, county, or county and state-by-year fixed effects to control for unobserved variables.

The use of fixed-effect controls is a significant feature of our study. Estimates for ordinary regression analyses are biased if key determinants of female headship, such as community values, state policies, urbanization or the provision of social services, are correlated with the availability of reproductive health services but omitted from the model.[14] For example, if a progressive social policy climate is positively associated both with the availability of abortion providers and with female headship within counties, the exclusion of this variable from the regression would lead to upward bias in the estimated effect of abortion access on headship.

The state fixed-effects model is equivalent to a regression specified to include a dummy variable for each state in the sample. The county fixed-effects model is equivalent to including a dummy variable for each county, and the state-by-year fixed effects are equivalent to interacting the state dummy variables with a dummy for the year of observation.

Each of the fixed-effects specifications controls for a different type of omitted variable. The state fixed-effects model controls for state-specific factors that do not vary over time. The county fixed-effects specification is more general; it accounts for both state- and county-specific factors that are time-invariant (i.e., state fixed-effects would be redundant in this model).[†] Finally, adding the state-by-year effects to these models absorbs all of the state-specific variation (e.g., the variation in measured and unmeasured state-level policy measures, such as changing Medicaid eligibility), but also makes it impossible to estimate independent effects of these variables in this model.

RESULTS

Table 1 shows population-weighted descriptive statistics for the dependent and key independent variables. These statistics are calculated for all women and for black, Hispanic and non-Hispanic white women. (As with the data for all women, the race- and ethnicity-specific data come from counties with at least 100 such women in both 1980 and 1990.) The figures indicate that between 1980 and 1990, the percentage of all women who were single heads of households rose by about 10%, from 6.9% to 7.6%. Black women had substantially higher rates of single headship, but each racial and ethnic group experienced similar percentage-point increases during the intercensal period.

Table 1. Mean values (and standard deviations) for variables used in regression analyses of factors influencing county-level rates of female household headship, by race and ethnicity, 1980 and 1990

Variable	All		White		Black		Hispanic	
	1980 (N=3,066)	1990 (N=3,066)	1980 (N=3,004)	1990 (N=3,004)	1980 (N=1,422)	1990 (N=1,422)	1980 (N=1,053)	1990 (N=1,053)
% of women who are single heads of households	6.93 (2.32)	7.62 (2.46)	5.19 (1.14)	5.62 (1.33)	18.30 (4.38)	19.57 (4.23)	9.90 (5.21)	10.25 (4.88)
No. of abortion providers per 1,000 women aged 15-44	0.05 (0.05)	0.04 (0.05)	0.05 (0.06)	0.04 (0.05)	0.05 (0.04)	0.05 (0.04)	0.07 (0.05)	0.06 (0.04)
No. of ob-gyns per 1,000 women aged 15-44	0.44 (0.28)	0.52 (0.32)	0.42 (0.27)	0.50 (0.30)	0.57 (0.33)	0.66 (0.39)	0.51 (0.22)	0.57 (0.26)
Parental notification laws (years in effect)	0.70 (1.33)	0.66 (1.55)	0.70 (1.33)	0.71 (1.60)	0.79 (1.39)	0.59 (1.44)	0.50 (1.08)	0.20 (0.88)
Medicaid abortion-funding restrictions (years in effect)	1.25 (1.13)	3.06 (2.39)	1.28 (1.13)	3.16 (2.37)	1.30 (1.05)	3.12 (2.37)	1.09 (1.27)	2.28 (2.48)
Maximum monthly public assistance (in $00s)	9.20 (1.70)	7.52 (1.35)	9.20 (1.66)	7.51 (1.30)	8.81 (1.84)	7.20 (1.36)	9.52 (1.91)	7.92 (1.65)

Note: Statistics based on county-level observations in each year are weighted by the number of women aged 15-44 in each county.

The data from Table 1 indicate modest declines across the decade in the number of abortion providers per 1,000 women; the declines range from 13% to 19%, depending on whether all women or specific racial or ethnic groups are considered. These declines coincide with dramatic decreases in the availability of Medicaid funding, but small increases in geographic access to obstetrician-gynecologists. The time pattern of parental consent and notification requirements is mixed across groups, with the incidence of such requirements increasing slightly across the decade for white women but decreasing for minority women (because white and minority women are distributed differently across states). Combined public assistance benefits declined for all groups.

We report results from several models of the determinants of female headship. We examined the effects of using different outcome measures of female headship, adjusting for either state or county fixed effects, controlling for within-state clustering, incorporating interactions between state dummies and time (i.e., fixed-effects controls for unmeasured changes in state policy and other variables), including alternative measures of abortion access and estimating models for different racial and ethnic groups. These alternative approaches are helpful in establishing how well our estimates take into account the biases resulting from different types of omitted variables.

ABORTION ACCESS AND FEMALE HEADSHIP

The first column in Table 2 shows coefficients produced by a regression of household headship that controls only for state fixed effects. The results are presented mostly for purposes of comparison with previous studies that have used only state-level data or have incorporated limited fixed-effects controls.[15] The estimates indicate that household headship rates among women are positively and significantly related to the number of abortion providers and obstetrician-gynecologists within a county and to parental consent and notification restrictions, Medicaid abortion-funding restrictions and welfare benefits within a state.

Table 2. Regression coefficients (and standard errors) showing the estimated effects of selected variables on county-level rates of female headship, by definition of headship and type of effect

Variable	Heads of households			Heads of families or subfamilies		
	State effects	County effects	County and state/time effects	State effects	County effects	County and state/time effects
No. of abortion providers per 1,000 women aged 15-44	0.90 (0.33)**	-1.22 (0.35)**	-1.18(0.35)**	0.89 (0.35)*	-1.25 (0.41)**	-1.28(0.40)**
No. of ob-gyns per 1,000 women aged 15-44	0.84 (0.07)**	0.81 (0.12)**	0.83(0.13)**	1.06 (0.07)**	0.68 (0.14)**	0.69(0.14)**
State requires parental consent or notification	0.05 (0.01)**	0.06 (0.03)*	na	0.02 (0.01)	0.01 (0.03)	na
State restricts Medicaid funding for abortion	0.15 (0.02)**	-0.01 (0.04)	na	0.12 (0.02)**	-0.03 (0.04)	na
Maximum public assistance benefits	0.18 (0.04)**	0.37 (0.10)**	na	0.27 (0.05)**	0.54 (0.12)**	na
State fixed effects	yes	no	no	yes	no	no
County fixed effects	no	yes	yes	no	yes	yes
Controls for within-state clustering	no	yes	no	no	yes	no
State/time interactions	no	no	yes	no	no	yes
R^2	*0.83*	*na*	*0.96*	*0.89*	*na*	*0.97*

*p<.05 **p<.01 *Notes:* Estimates are based on 6,132 county-level observations from 1980 and 1990, weighted by the number of women aged 15-44 in each county. The data in Tables 2-4 are adjusted for the effects of sex ratio; men's and women's median full-time income; men's and women's education; men's employment; the logarithm of the population; and the percentages of the population in each county that are older than 65, black, Hispanic, rural, Catholic, divorced, adherents of the Church of Jesus Christ of Latter Day Saints or antiabortion Protestants. na=not applicable

The positive coefficient for the availability of obstetrician-gynecologists suggests that this variable acts more as a proxy for lower childbirth costs than for contraceptive costs. This interpretation contrasts with that offered by Matthews and colleagues, who found that the availability of obstetrician-gynecologists was positively associated with overall birthrates.[16] The positive coefficients for legal restrictions suggest that such restrictions lead to fewer abortions and, consequently, to more births among women who are unmarried or in unstable marriages. However, the conclusion that reduced access to abortion services leads to higher headship rates appears to be undermined by the positive correlation between headship rates and the number of abortion providers within the county.

In this case, appearances are deceiving. When the regression model is respecified to incorporate county-specific effects (column 2), the contradictory results for geographic access and legal access to abortion services disappear. Specifically, the positive coefficient for the number of abortion providers becomes significantly negative. The effect on female headship of the change in abortion availability between 1980 and 1990 is calculated by multiplying the observed change in abortion availability (-0.01) by the coefficient in column two (-1.22). The change in female headship attributable to reduced access to abortion is 0.012 (roughly 2% of the 0.69 percentage-point increase in headship). In contrast, the coefficients for the number of obstetrician-gynecologists and the presence of parental consent and notification remain significantly positive, while the coefficient for Medicaid restrictions becomes small and nonsignificant.

The difference in the estimated impact of number of abortion providers between models with state-specific effects and those with county-specific effects mirrors the findings of Kane and Staiger.[17] Specification tests reveal that the county controls not only are significant overall but also significantly improve the fit of the model over that of the first regression. The improvement in fit indicates that there are unobserved county-specific determinants of headship rates, and that they are correlated with access to abortion providers.

It is possible, however, that the estimated effects of geographic access to abortion providers and the number of obstetrician-gynecologists are sensitive to other unmeasured state policies or characteristics (such as other laws, the provision of sex education, changing Medicaid eligibility or the general outlook of the state legislatures and courts). To examine this possibility, we replace the two state-level abortion policy variables and the state-level welfare variable with a general set of state-by-year dummy variables. The results (shown in the third column of Table 2) are essentially the same as those in the previous column, indicating that our estimates are not sensitive to the inclusion or exclusion of additional state controls.

The next three columns of Table 2 show results from regressions that are comparable to those in the first three columns, but use family or subfamily headship rather than household headship as the dependent variable. The results of these models are similar to the results from the first set of regressions. The only substantive difference is that the coefficients for parental consent and notification requirements become much smaller and lose their significance. We conclude from this exercise that our estimates of the effects of geographic access to reproductive health services are not affected by minor changes in the definition of headship.

RACIAL AND ETHNIC DIFFERENCES

We also estimate separate county fixed-effects models (with and without state-time interactions) of the percentage of women heading households with children for whites, blacks and Hispanics. Race-disaggregated results are reported in Table 3 and, for purposes of comparison, are estimated from models similar in functional form to the models specified in the second and third columns of Table 2.

Table 3. Regression coefficients (and standard errors) showing the estimated effects of selected variables on county-level rates of female headship, by race and ethnicity and type of effect

Variable	White (N=6,008)		Black (N=2,844)		Hispanic (N=2,106)	
	County effects	County and state/time effects	County effects	County and state/time effects	County effects	County and state/time effects
No. of abortion providers per 1,000 women aged 15-44	-0.78 (0.30)*	-0.76(0.30)*	-3.30 (2.38)	-2.89(2.40)	-3.43 (1.84)	-3.13(1.85)
No. of ob-gyns per 1,000 women aged 15-44	0.33 (0.11)**	0.35(0.11)**	0.43 (0.63)	0.35(0.64)	0.27 (0.63)	0.25(0.64)
State requires parental consent or notification	0.05 (0.02)*	na	0.06 (0.11)	na	0.09 (0.10)	na
State restricts Medicaid funding for abortion	-0.03 (0.03)	na	0.37 (0.15)*	na	0.16 (0.13)	na
Maximum public assistance benefits	0.03 (0.09)	na	1.21 (0.41)**	na	0.16 (0.38)	na
Controls for within-state clustering	yes	no	yes	no	yes	no
State/time interactions	no	yes	no	yes	no	yes
R^2	*na*	*0.87*	*na*	*0.84*	*na*	*0.94*

*p<.05 **p<.01 *Notes*: Estimates are based on county-level observations from 1980 and 1990, weighted by the number of women aged 15-44 of each racial or ethnic group in each county. Regressions also include county fixed effects. na=not applicable.

The results generally confirm our earlier findings regarding the effects of geographic access to reproductive health services. Access to abortion providers had a negative effect on household headship rates for women in all three racial and ethnic groups. The coefficient estimates are statistically significant only for white women; because of the large standard errors, the estimates for black women and Hispanic women are not significant. The estimated effects of the availability of obstetrician-gynecologists are positive for all three groups, but are significantly positive only for whites.

The estimated effects of parental consent and notification requirements follow a similar pattern—positive for all three groups, but significant only for whites. The coefficients for Medicaid restrictions show more variability, however. The estimated effects are significantly positive for blacks, positive but nonsignificant for Hispanics, and negative and nonsignificant for whites. The differences in the Medicaid results across groups might be explained by the greater salience of this program for blacks (who are disproportionately likely to be poor and, therefore, more likely to rely on public assistance) than for whites. The change in Medicaid restrictions accounted for 52% of the 1.27 percentage-point increase in headship among black women. The estimates for public assistance reinforce this interpretation, as they too are significantly positive for black women but not for white or Hispanic women.

SUBSTITUTION OF ALTERNATIVE MEASURES

To test the sensitivity of our results, we reestimate our aggregate models using alternative measures for several variables. Table 4 reports results for three models that are based on all women and incorporate county fixed effects and the general set of state-by-year controls (as in the model in the third column of Table 2), but use different measures for geographic access to reproductive health services and the demand for such services.

Table 4. Regression coefficients (and standard errors) showing the estimated effects of selected variables on county-level rates of female headship, adjusted for provider presence, proximity and demand			
Variable	Adjusted for presence of provider	Adjusted for proximity of provider	Adjusted for demand for providers
No. of abortion providers per 1,000 women aged 15-44	-1.04(0.42)**	-1.04(0.42)**	-1.11(0.33)**
Abortion clinic in county	-0.04(0.07)	-0.04(0.20)	na
Distance to nearest in-state abortion clinic (log)	na	0.00(0.05)	na
Distance to nearest out-of-stateabortion clinic (log)	na	-0.09(0.06)	na
No. of ob-gyns per 1,000 women aged 15-44	0.82(0.13)**	0.87(0.14)**	0.71(0.12)**
Distance to nearest ob-gyn (log)	na	0.01(0.02)	na
No. of births per 1,000 women aged 15-44	na	na	-0.01(0.00)*
No. of women aged 15-44 (log)	na	na	3.41(0.60)**
% of adult women who are divorced	na	na	0.44(0.02)**
R^2	*0.96*	*0.96*	*0.96*

*p<.05 **p<.01. *Notes*: Estimates are based on 6,132 county-level observations from 1980 and 1990, weighted by the number of women aged 15-44 in each county. Regressions also include county fixed effects and state/time interactions. na=not applicable.

The variable for the number of abortion providers per 1,000 women reflects two different aspects of availability—geographic proximity and congestion. In the first two respecified models, we use alternative measures that relate more closely to proximity. In the first model (column 1), we consider whether headship is affected by the simple presence of a provider rather than by the number of providers in a county. The estimates provide little support for this view. Indeed, controlling for the presence of an abortion provider does not alter the coefficient estimate for the number of providers, and the coefficient for this dummy variable is statistically nonsignificant.

In the next model (column 2), we add controls for proximity to providers (the logarithms of distance to the nearest in-state and out-of-state abortion clinics and to the nearest obstetrician-gynecologist). None of these additional distance controls is significantly associated with the level of family headship among women. The inclusion of these additional controls also does little to attenuate the effect of either number of abortion providers per 1,000 women or number of obstetrician-gynecologists per 1,000 women on county-level growth in female headship. The results suggest that congestion (i.e., lines and waiting for appointments) at abortion clinics and physicians' offices is an important element of availability.

A final sensitivity analysis addresses the question of whether local growth in the number of abortion providers or obstetrician-gynecologists is market driven and therefore endogenous to changing patterns of family formation in a county. Simply, local changes in reproductive health services may be a market response to increasing fertility in the county or to growth in the number of women of childbearing age. Higher fertility and population growth would, in turn, place increasing proportions of women at risk of female headship.

To control for this and for the possibility that our results primarily reflect an effect through divorce, the model in the last column includes measures for changes in the number of births per 1,000 women aged 15-44, the local divorce rate and the logarithm of the number of women aged 15-44. The three variables are individually and jointly significant. The coefficient estimates for the divorce rate and the number of women are positive, while the coefficient for the birthrate is small and negative. Despite their significance, the inclusion of these controls does not substantially alter our underlying estimates of the effects of access to abortion providers and obstetrician-gynecologists.

CONCLUSIONS

Recent welfare reform legislation and new legal restrictions on abortion are seemingly working at cross-purposes. The cutbacks in welfare have increased the costs of unmarried childbearing, while arguably increasing the demand for abortions to end unintended pregnancies. At the same time, reductions in the local availability of abortion providers, new parental notification requirements and cutbacks in federal funding for abortions have made abortions more difficult to obtain.

Our results suggest that the public policy goal of reducing unmarried childbearing and the proportion of women heading families may be undermined—at least in part—by increasing geographic and legal barriers to abortion. Our estimates from fixed-effects models, which control for unobserved heterogeneity across counties, indicate that female headship decreases with increasing availability of abortion providers and increases with the availability of obstetrician-gynecologists.

The substantive impact of abortion availability, however, is not large: Our estimates indicate that the nearly 20% decrease in providers accounted for less than 2% of the overall growth in headship across the decade. The increase in the availability of obstetrician-gynecologists was a more important factor, explaining nearly 10% of the increase in headship.

The effects of new legal restrictions on abortion—Medicaid funding restrictions and parental notification requirements—were less conclusive overall. State restrictions on Medicaid funding for abortions were significantly associated with increases in female headship among blacks but not among other racial or ethnic groups. State parental consent and notification requirements, however, were significantly associated with the rise in family formation among unmarried white women but not in other groups. The estimates of these effects might have been larger if our data had permitted analyses more narrowly restricted to younger or poorer women, those most likely to have been affected by the policies.

Our findings must be kept in proper perspective. Any additional costs (financial, social or legal) associated with decreasing abortion access arguably have not been prohibitive, if judged by the roughly 1.4 million abortions performed each year.[18] Moreover, roughly one-quarter of all pregnancies end in abortion; a disproportionate share of abortions continue to be obtained by unmarried teenagers with unintended pregnancies.† Our results nevertheless cast some doubt on recent claims that the knowledge that abortions are now more difficult to obtain may have resulted in changes in sexual behavior or contraceptive practice to avoid unintended pregnancy.[19] A cautious interpretation of our results is that cutbacks in abortion access have instead contributed modestly to the upswing in female headship, and that this has occurred primarily because an increasing share of unmarried women are choosing nonmarital fertility over abortion, especially as access to abortion providers has diminished over the past decade or so.

Finally, our findings linking declining abortion access to rising female headship do not, in themselves, constitute sufficient basis for rescinding existing restrictive abortion legislation or initiating efforts to make abortion services more geographically accessible. Policies aimed at reducing the number of abortions are motivated by many considerations—moral, legal and social. At a minimum, our results highlight the potential unintended consequences of newly imposed abortion restrictions and declining abortion access in the context of welfare reform legislation aimed at strengthening the traditional two-parent family while discouraging growth of single-parent families. They also add new information to ongoing debates about welfare and abortion legislation in the context of often conflicting or competing public policy goals.

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*Our analysis defines counties in terms of their 1980 boundaries. For small, independent cities (e.g., in Virginia) whose borders rest entirely within a county, we combine data for the city and the county. The analysis treats data for large cities as if they were county records.

†Another limitation is that the subfamily measure for 1980 is underestimated because of a coding error by the Census Bureau. The use of county-specific fixed effects should control for this systematic measurement error. Because of its various limitations, we restrict our use of the subfamily measure to a few sensitivity analyses.