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# Welfare and the Rise in Female-Headed Families[1]

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The article provides a bridge between recent marriage market research and studies of welfare incentive effects on U.S. family formation. Estimates from state and county fixed-effects models indicate significant effects of changing state Aid to Families with Dependent Children, food stamps, and Medicaid expenditure levels on county-level changes in families headed by unmarried mothers. However, neither changing welfare benefit levels nor declining economic and marital opportunities could account for recent increases in female headship. The results imply that large additional cuts in welfare payment levels would lead to only small reductions in the percentage of female-headed families with children.

## INTRODUCTION

The transformation of the family has continued apace in the United States. Between 1970 and 1993, the percentage of all families (with children) maintained by a single mother increased from 11.5% to 25.9%, while the number of children living with only their mother expanded from 7.5 to 15.6 million (Rawlings 1994; Saluter 1994). The economic and social costs have been large. A disproportionate share of children raised in

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female-headed families experience chronic poverty and various deleterious developmental and behavioral problems, including poor cognitive and emotional development, teenage pregnancy, and school dropout (Amato 1993; McLeod and Shanahan 1996; McLanahan and Sandefur 1995). Indeed, female-headed families have experienced exceptionally high rates of poverty—about 50%—over the past two decades (Casper, McLanahan, and Garfinkel 1994; Lichter 1997). They also have comprised the overwhelming share of recipients of Aid to Families with Dependent Children (AFDC), the cash assistance program that was the nation's largest prior to implementation of new welfare reform legislation in early 1997.[2]

From a public policy standpoint, it should come as no surprise that marriage is increasingly viewed as a panacea for poverty and other social problems (Blankenhorn 1995; Popenoe 1996; Waite 1995). The current retreat from marriage and the decline in the two-parent family have reinvigorated research on the etiology of family formation in the United States (e.g., Cherlin 1992; Qian and Preston 1993). Much of the interest has centered on the role played by demographic shortages of "marriageable men" and the rising employment and earnings of women (e.g., Wilson 1987; Wood 1995). At the same time, recent trends have raised new questions about whether past and current welfare policies and public assistance programs (e.g., AFDC participation and benefit levels) have undermined traditional patterns of family formation, while encouraging nonmarital fertility, divorce, and the growth in female-headed families with children (Moffitt 1994, 1995; Schultz 1994).

The main objective of our article is to provide a bridge between recent research on local marriage markets (South and Lloyd 1992; Lichter, Anderson, and Hayward 1995) and studies of welfare incentive effects on family formation (Ellwood and Bane 1985; Duncan and Hoffman 1990). Specifically, we estimate and compare the effects of *changing* state AFDC benefit levels, food stamp benefits, and Medicaid expenditures, as well as changing economic and marriage market opportunities, on recent area-level changes in female-headed families. Our goals are (1) to document the rise in female-headed families (for non-Hispanic whites, blacks, and Hispanics) during 1980–90 for all U.S. counties, (2) to evaluate, for the first time, competing explanations (e.g., welfare incentives to nonmarriage

[2] The new program is called Temporary Aid to Needy Families (TANF) in the Personal Responsibility and Work Opportunity Reconciliation Act of 1996. In the early 1990s, only about 8% of recipients received cash payments as married persons through the AFDC-Unemployed Parents program. This program provided married couples with assistance during spells of unemployment if they met certain eligibility requirements regarding previous work histories (i.e., the 100-hour rule). One purpose of this program was to reduce family stress during periods of economic hardship and prevent marital disruption and divorce.

versus mate unavailability) of increasing local-area female headship rates, and (3) to estimate various cross-state and cross-county fixed-effects models of family formation (Moffitt 1996). Research on the rise in female-headed families with children is propitious in light of current welfare debates centered on the possible implications—both good and bad—of the newly enacted welfare bill, the Personal Responsibility and Work Opportunity Reconciliation Act of 1996.

## ECONOMIC INCENTIVES AND FEMALE HEADSHIP

The retreat from marriage is inextricably linked to the growth in the number and percentages of female-headed families (Lichter 1995; Smith, Morgan, and Koropeckyj-Cox 1996). Nearly one-third of all babies today are born out of wedlock and roughly 30% of all nonmarital births are to divorced women (Ventura et al. 1995). Among blacks, increases in nonmarriage (rather than increases in nonmarital fertility rates) have accounted for the overwhelming share of the post-1960 rise in the nonmarital fertility ratio, that is, the ratio of nonmarital births to all births (Smith et al. 1996). Clearly, public policies—including welfare policies—that address the retreat from marriage may also slow or even reverse the rise in nonmarital fertility and female-headed families in the United States.

Current debates on changing patterns of family formation, especially among ethnic and racial minority groups, have centered largely on the comparative merits of theoretical perspectives that emphasize either men's or women's changing economic roles (Cherlin 1992; Oppenheimer 1997). Some argue, for example, that the rise in female headship is largely explicable in terms of growing demographic shortages of "marriageable" men. The deteriorating low-wage, low-skill labor market has reduced women's incentives to marry and has undercut the economic foundations of existing marriages. This is most apparent among black women residing in communities with large sex ratio imbalances and high male unemployment (Tucker and Mitchell-Kernan 1995; Wilson 1987). Other studies attribute changing patterns of family formation to the improving employment circumstances of American women (McLanahan and Casper 1995; Schultz 1994; Darity and Myers 1995). Employment and earnings presumably increase women's economic independence from men and reduce the incentives to marry. Improved economic status also allows women to leave unhappy marriages, unmarried mothers to live independently from other adult family members, and pregnant unmarried women to choose single motherhood over abortion, adoption, and marriage.

From a conceptual standpoint, the singular emphasis on *either* men's or women's economic roles is inappropriate. These are not mutually exclusive perspectives; men's and women's economic roles within the family

are interrelated in fundamental ways (Oppenheimer 1997). Indeed, a broader microeconomic perspective, one that emphasizes the rising economic and personal costs and the declining benefits of marriage in modern society, can subsume both (Becker 1981). The benefits from the specialization of household production along traditional gender roles—women in home production (i.e., childbearing and child rearing) and men in labor market activities—have declined with the changing economic roles of women. At the same time, the main benefit of marriage for women, traditionally one of economic support from men, has eroded as the economic position of men, especially young men with limited education or work skills, has declined relative to women's over the past two decades. The implication from Becker's rational choice model is that rising female headship results from the blurring of traditional gender and economic roles.

The problem with economic explanations that emphasize changing employment and earnings—of either men or women—is they are often inconsistent with the empirical record. For example, declining "male marriageability" implies that the economic gains from traditional marriage have declined for women. But most studies show that demographic shortages of economically attractive men do not account for the widening racial differences in family formation patterns, nor can they account for much of the recent change in marriage rates or female headship (McLanahan and Casper 1995; Lichter, LeClere, and McLaughlin 1991; Raley 1996; McLaughlin and Lichter 1997). The study by Mare and Winship (1991) is illustrative of this point. Among white men, roughly 20% of the decline in marriage rates during the 1940–85 period was due to their changing employment rates. For blacks, the role of declining "marriageable men" is even smaller. Using metropolitan-level census data, Wood (1995) showed that only about 4% of the 1970s decline in black marriage rates was due to the changing local pool of adequately employed black men.

Other research eschews the current preoccupation with the changing availability of male marriage partners but stresses instead the declining economic "costs" of remaining single and getting divorced for women. Specifically, the economic imperative for women to marry and stay married has diminished with rising female labor force participation and higher real wages. Delayed marriage and divorce impose fewer economic hardships than in the past on women. Moreover, employed single mothers can afford to live apart from their families of origin (Avery, Goldscheider, and Speare 1992). Such commonplace assertions, however, have received mixed empirical support. McLanahan and Casper (1995) reported that, among white women, changing employment and earnings accounted for 70% of the decline in marriage between 1970 and 1990. At the other extreme, Qian and Preston (1993) found that declines in marriage were most

pronounced during 1972–87 among the least educated women—a result inconsistent with arguments that emphasize the changing economic roles of women. They also suggested that change in the "force of attraction" (i.e., unobservable change in values regarding marriage) was most responsible for the retreat from marriage (see also Schoen and Kluegel 1988). Other studies show poor women are less likely to marry than nonpoor women (McLaughlin and Lichter 1997) and that employed women and high wage earners have an increased rather than decreased annual probability of marriage (Oppenheimer 1994; Lichter et al. 1992).

The mixed conclusions and apparent limitations of strictly economic explanations have revived previously discredited explanations of changing family life. This includes new research on whether the welfare state has created economic disincentives to marriage among low-income groups while setting into motion various adaptive or maladaptive cultural changes (including nonmarital fertility and divorce) felt throughout society (Moffitt 1996; Murray 1993). This resurgent interest in welfare incentives is coincident with the passage of new welfare reform legislation that includes time limits and work provisions aimed at promoting behaviors among the poor that resonate with the deeply held American values of hard work and "strong" families. During President Clinton's first term, the Department of Health and Human Services loosened the waiver process for state experimentation with welfare programs. State experimentation with welfare and social service provisions has now been institutionalized through the mechanism of federal block grants to states.

An evaluation of the incentive effects of state welfare benefit levels is clearly needed at a time when state-to-state variation in welfare provision and generosity is expected to increase substantially over the next several years. The existing literature on welfare incentives on the family is complicated and difficult to summarize neatly. Welfare effects apparently also have changed unpredictably over time (Moffitt 1996), further complicating the conventional wisdom.[3] The theory, however, is straightforward: Public assistance, especially cash assistance programs like AFDC, putatively creates economic incentives to bear children without marriage, discourages marriage, and promotes independent living among unmarried mothers. It contributes to the retreat from marriage and the rise in female-

[3] Many sociologists appear to have a much different reading of the literature than do other social scientists. At the risk of some simplification, sociologists often downplay the role of welfare incentives (based largely on their reading of pre-1980 studies) or ignore welfare altogether in empirical studies because to do so would be tantamount to "blaming the victim." On the other hand, the more nuanced analyses of economists regarding welfare incentive effects often come at the expense of ignoring or minimizing other competing explanations, including cultural or value changes, gender roles, and labor market discrimination.

headed families among low-income women by providing a "surrogate husband" in the form of a steady but modest source of income. The "antifamily" effects of welfare therefore are expected to increase over time in response to the deteriorating economic circumstances of young, less educated adults. Welfare incentive effects also should be most apparent among minority populations, a disproportionate share of whom are poor and eligible for welfare.

Others reject such arguments. First, the family has been transformed across all economic strata in America (Sweet and Bumpass 1987), a fact that militates against monocausal explanations that emphasize the role of welfare. Recent family trends suggest sweeping cultural shifts that have affected virtually all segments of American society. Second, female headship has increased over the past decade while real welfare benefit levels have become less rather than more generous (Garfinkel and McLanahan 1986). If welfare benefits were the primary determinant of single headship, theory predicts that the decline in benefit levels would have led to a decrease in this outcome. Third, "blaming" the welfare system for the problems of the poor (e.g., Murray 1993) is misplaced; rather, factors such as inadequately funded schools and family-supporting social services, few employment opportunities, racial discrimination, and neighborhood segregation and isolation are seen as root causes of various adaptive behaviors, as well as correlates of both welfare generosity and dependence. The implication is that studies of welfare incentive effects must be evaluated within a comprehensive framework that considers a variety of alternative explanations. Such is the purpose of this article.

Moffitt's (1992, 1995, 1996) recent comprehensive reviews of welfare incentives draw several useful conclusions for the purpose of our study. First, most economic studies show that welfare has significant deleterious effects on various measures of family formation, including female headship. Second, these effects have generally increased in size over the past decade (Moffitt 1994). Third, any disincentive effect of public assistance on family formation is not spurious, that is, it is not an artifact of the fact that women bearing children outside of wedlock are more likely to receive public assistance and to delay or forgo marriage (Bennett, Bloom, and Miller 1995).[4] And, fourth, welfare incentives tend to be stronger among whites than among blacks, a finding inconsistent with public perceptions and also puzzling in light of the larger share of blacks "at risk" of welfare incentive effects (i.e., being eligible for welfare by virtue of low income). Other research, however, shows that the receipt of public assistance

[4] Bennett et al. (1995, p. 57) found that "welfare recipiency accounts for a small but nontrivial portion (about one-fifth) of the negative association between nonmarital childbearing and the subsequent likelihood of marriage."

among cohabiting couples lowers marital transition rates among blacks but not whites (Manning and Smock 1995). Black women on welfare are also more likely than white women to forgo marriage rather than marry relatively low-status men (Lichter et al. 1995).

The debate today should not focus exclusively on whether welfare effects exist. The emphasis should instead be on the absolute size of welfare incentives, on the size of welfare effects compared to those of other frequently ignored "causes" (such as employment and sex ratio imbalances), and on issues of statistical design for best discerning welfare incentive effects.

## THE CURRENT STUDY

From a conceptual and analytic standpoint, our article builds most directly on the recent areal study by McLanahan and Casper (1995). They used data from the 1970, 1980, and 1990 Public Use Microdata Samples to construct marriage market indicators for the 100 largest metropolitan areas in the United States. They estimated a pooled regression model that included women's characteristics (e.g., proportion employed full-time), men's characteristics (e.g., median earnings), the local sex ratio, and state welfare benefits. They found that AFDC–food stamp benefit levels were negatively associated with metropolitan proportions currently married, but that welfare effects were small from a substantive standpoint. Substituting the low and high levels of welfare in their regression equation resulted in a difference of only about 5–7 percentage points in predicted marriage rates. Welfare incentive effects also were small in comparison to the effects of men's and women's economic circumstances.

Our analysis incorporates several important strengths of the McLanahan-Casper study, and also of other recent cross-state, cross-sectional analyses of local marriage markets (e.g., Fossett and Kiecolt 1992; Lichter et al. 1991), while addressing the weaknesses of each. First, we examine the effects of *changes* in welfare benefit levels, skill levels, and economic opportunities for women and men, and the local pool of "economically attractive" spouses on family formation. This departs from much of the existing sociological research, which downplays or ignores the role of government assistance in marriage and family formation decisions.

Second, we examine longitudinal data aggregated over small geographic areas—counties. The use of county-level economic data confers conceptual and statistical advantages in comparison to previous studies that have used gross state-level measures without regard to intrastate differences in labor market conditions or marriage opportunities (Moffitt 1994). Most marriage markets are locally circumscribed rather than defined by state boundaries.

Third, unlike McLanahan and Casper (1995), who used metropolitan area data, we need not impute welfare benefits across multistate areas, and we can consider family formation behavior in nonmetropolitan areas. Our analysis is comprehensive from a geographic standpoint, covering all counties in the 48 contiguous states.

Finally, our analysis, which uses repeated measures of both the family formation outcome variable and the other explanatory variables, can incorporate county-specific fixed-effect controls for omitted variable bias. Decisions regarding marriage and family formation are clearly more complex than indicated by the simple models specified in most empirical studies. Moreover, previous research has shown that model estimates are highly sensitive (in some instances to the point of results being eliminated or reversed) to the inclusion of alternative controls for the processes and variables that researchers can or cannot observe.[5] Omitted variable bias is addressed here with estimates of welfare incentive effects from county-level panel models of female headship.

## METHODS

### Data

The primary data for this analysis consist of cross-sectional county records drawn from the Summary Tape Files (STF) of the 1980 and 1990 decennial censuses of the United States. We matched information for each county across years to form a short panel. We excluded all observations from Alaska and Hawaii because the costs of living and ethnic composition were unrepresentative of the rest of the country. Counties that had fewer than 50 families with children under 18 in either 1980 or 1990 also were eliminated from the analysis. The result is a pooled data set containing 6,106 observations (3,053 counties matched across 1980 and 1990).[6] Short descriptions as well as means and standard deviations of the variables for the total sample of counties in 1980 and 1990 are reported in table 1.

The STF data have several features that distinguish them from the data

[5] Relevant examples include Moffitt's (1994) analysis of female headship rates and Jackson and Klerman's (1994) and Kane and Staiger's (1996) analyses of young women's fertility rates.

[6] There were a small number of cases during the 1980s where new counties split off from existing counties. For consistency, our analysis defines counties in terms of their 1980 boundaries. In some states, STF data are recorded for independent cities as well as for counties. For small independent cities whose borders rested entirely within a county, data for the city and county have been combined; for large cities, the analysis includes the city-level records as if they were counties.

TABLE 1

DESCRIPTION OF VARIABLES AND SIMPLE STATISTICS

| VARIABLE | DESCRIPTION | MEAN 1980 | MEAN 1990 |
|---|---|---|---|
| Female headship | Percentage of family households with children under 18 years old with single female heads | 16.03 (6.30) | 18.72 (6.76) |
| Sex ratio | Ratio (× 100) of males to females, 15–59 years old | 97.56 (6.96) | 98.90 (8.03) |
| Male employment | Percentage of male civilians employed | 70.20 (6.71) | 69.42 (6.75) |
| Male earnings | Median income for men working full-time, full-year (1989 dollars, in 1,000s) | 29.94 (5.01) | 29.05 (5.33) |
| Male education | Percentage of males, 25–44 years old, with a bachelor's degree or more | 24.87 (9.09) | 25.10 (10.00) |
| Female earnings | Median income for women working full-time, full-year (1989 dollars, in 1,000s) | 17.34 (2.51) | 19.18 (3.69) |
| Female education | Percentage of females, 25–44 years old, with a bachelor's degree or more | 17.51 (6.70) | 22.46 (8.78) |
| Percentage 65 and over | Percentage of the population 65 years old or older | 11.04 (3.44) | 12.41 (3.60) |
| Percentage black | Percentage of the population who are black | 11.01 (12.09) | 11.71 (12.67) |
| Percentage Hispanic | Percentage of the population who are Hispanic | 6.81 (11.01) | 8.74 (13.03) |
| Percentage rural | Percentage of the population who reside in a rural area | 26.98 (27.84) | 25.86 (27.80) |
| ln(population) | Natural logarithm of county population | 12.44 (1.62) | 12.55 (1.65) |
| Percentage Catholic | Percentage of the population who are adherents to the Catholic religion | 20.11 (15.75) | 21.14 (16.48) |
| Percentage LDS | Percentage of the population who are adherents to the Church of Jesus Christ of Latter Day Saints | 1.35 (6.55) | 1.51 (6.93) |
| Percentage conservative Protestant | Percentage of the population who are adherents to strongly antiabortion Protestant denominations | 10.93 (11.52) | 11.30 (12.27) |

TABLE 1 (*Continued*)

| Variable | Description | Mean 1980 | Mean 1990 |
|---|---|---|---|
| Maximum AFDC benefits for family of four | Maximum monthly AFDC benefits for a family of four with no other income (1989 dollars, in 100s) | 5.65 (2.07) | 4.60 (1.84) |
| Average Medicaid benefits for family of four | Average state monthly Medicaid payment for AFDC family of four (1989 dollars, in 100s) | 2.90 (.95) | 2.57 (.64) |
| Maximum combined welfare benefits | Sum of maximum AFDC and food stamps and average Medicaid benefits for a family of four (1989 dollars, in 100s) | 9.14 (1.70) | 7.48 (1.34) |

NOTE.—Statistics based on 3,053 county-level observations in each year weighted by the no. of families in each county. SDs appear in parentheses.

used in previous research. First, because the data for each county are longitudinal, our study can track changes in and determinants of family formation, including welfare payment levels both across and within counties over time. This addresses a limitation of the McLanahan-Casper study and most previous work in this area (i.e., the failure to control for other unobserved community or local-area variables). Second, the data allow us to form community-based measures of marriage opportunities, gender-specific economic opportunities, and other population characteristics. Third, these data also provide race- and ethnicity-specific measures of family formation and several other key variables. Accordingly, we conduct separate analyses for non-Hispanic whites, blacks, and Hispanics. The race-specific data sets contain 5,800 observations (data for 2,900 counties) for non-Hispanic whites; 2,366 observations for blacks; and 1,542 observations for Hispanics. Means and standard deviations for the race- and ethnicity-specific measures are reported in table 2.

## Measurement

The dependent variable for our county-level regression analysis is the percentage of family households with children under 18 headed by unmarried women.[7] Using the STF data, we measure the demographic supply of men

[7] Other research has sometimes examined the percentage of all women who are heading families with children. Welfare effects are then interpreted as reduced-form effects that have many different pathways to female headship: marriage or not, childbearing

TABLE 2

Race- and Ethnicity-Specific Variables

| | White | | Black | | Hispanic | |
|---|---|---|---|---|---|---|
| | 1980 | 1990 | 1980 | 1990 | 1980 | 1990 |
| Female headship | 12.16 | 13.98 | 41.30 | 48.10 | 19.72 | 22.31 |
| | (3.26) | (3.21) | (8.81) | (9.14) | (9.97) | (10.39) |
| Sex ratio | 98.64 | 99.45 | 88.74 | 89.24 | 100.53 | 110.91 |
| | (5.50) | (6.52) | (20.37) | (22.83) | (16.25) | (18.55) |
| Male employment | 71.57 | 71.00 | 58.82 | 57.47 | 71.73 | 70.55 |
| | (6.72) | (6.58) | (7.69) | (8.49) | (7.22) | (7.67) |
| Male earnings | 30.91 | 30.26 | 22.30 | 22.17 | 22.37 | 20.19 |
| | (5.04) | (5.79) | (5.09) | (4.49) | (3.88) | (3.31) |
| Male education | 26.39 | 26.98 | 11.01 | 12.39 | 10.49 | 10.23 |
| | (9.82) | (10.99) | (5.27) | (5.73) | (6.18) | (5.45) |
| Female earnings | 17.45 | 19.43 | 16.16 | 17.81 | 15.11 | 16.12 |
| | (2.51) | (3.92) | (3.24) | (3.90) | (1.89) | (2.41) |
| Female education | 18.26 | 23.88 | 10.89 | 13.97 | 6.98 | 9.56 |
| | (7.30) | (9.77) | (4.70) | (5.50) | (4.24) | (5.14) |
| No. of counties | 2,900 | | 1,189 | | 772 | |

Note.—Statistics in each year weighted by the no. of families of each racial/ethnic group in each county. SDs appear in parentheses.

relative to women and the "economic attractiveness" of potential male partners in the county. These are defined by the sex ratio, male employment rate, male earnings among full-time workers, and percentage of college-educated men in the county (Fossett and Kiecolt 1991; Lichter et al. 1991). The expectation is that an increasing supply of economically attractive men is negatively related to the rise in female-headed families. We construct similar measures of female economic independence, including the median earnings of full-time, full-year female workers and percentage of college-educated women. We examine the effects of both aggregate and race-specific measures of marriage market and gender-specific economic and skill-level variables on female headship. Dollar-denominated values are deflated by the Personal Consumption Deflator and expressed in constant 1989 terms.

A main objective of the analysis is to evaluate the absolute and relative

or not, independent living or not, etc. Data limitations prevent us from exploring these alternative pathways. In any event, the denominator in this alternative measure is not the population at risk of receiving AFDC; it is families with children, a fact reflected in our measure. We nevertheless have fit models with alternative measures (table A1), and the results are discussed below.

relationship between welfare generosity and female headship. For this, we have matched the county-level STF records to longitudinal state-level information on the AFDC, food stamp, and Medicaid programs (U.S. House of Representatives 1990; U.S. Office of Family Assistance 1980; and unpublished data from the U.S. Food and Nutrition Service and U.S. Health Care Financing Administration). For our measures of AFDC and food stamp generosity, we use the maximum benefit levels for a family of four with no other income. Food stamp benefits are fixed at the national rather than state level; by definition, they will not be associated with state-to-state variation in changes in female headship. Food stamps may nevertheless be important because benefits under the program are reduced by the receipt of AFDC income and thus act to narrow state-to-state variation in the overall benefits package. We use the combined benefits package as our primary measure of welfare generosity. Following Moffitt (1992), this is the sum of the maximum AFDC benefit, the adjusted food stamp benefit, and .368 times the average Medicaid benefit.[8]

Finally, the census STF data also provide other control variables. These include the age and racial/ethnic composition of each county, variables that have been highly associated with family formation in previous studies (Goldscheider and Waite 1986; Lichter et al. 1991). Moreover, unlike virtually all other studies, we include both proxy and direct measures of the cultural context of family formation. Proxy measures of urbanization (which has many cultural manifestations) include the logarithm of county population and the percentage of rural population, as defined by the U.S. Census Bureau. Previous studies have shown that nonmetropolitan women marry earlier, on average, and are less likely than urban women to bear children outside of marriage (McLaughlin, Lichter, and Johnston 1993; Heaton, Lichter, and Amoateng 1989). These proxy measures of cultural context are supplemented with direct measures, that is, longitudinal county-level information on participation in religious organizations with strong "profamily" orientations. Specifically, the Glenmary Research Center (1980–92) provides county-level estimates of the number of adherents and communicants for various religious denominations. We use these data to form estimates of the percentages of the county population that are Roman Catholic, members of the Church of Jesus Christ of Latter Day Saints (the Mormon church), or members of Protestant denomina-

[8] The adjusted food stamp amount is defined as maximum food stamp benefit − .3 max (0, maximum AFDC benefit − standard deduction), while the 0.368 in the combined benefit formula represents the fraction of AFDC recipients who actually claim services under the Medicaid program. The AFDC and food stamp amounts are deflated by the Personal Consumption Deflator, and the Medicaid amount is deflated by the Consumer Price Index for medical services.

tions with strong antiabortion positions.[9] Winkler (1994) showed that the effects of welfare payments are sensitive to the inclusion of measures of "community conservatism."

## Model Specifications

The pooled regression analysis of McLanahan and Casper (1995) provides a point of departure for our analysis. We begin with estimates from 1980–90 pooled county-level cross-section regression models of the percentage of families with children that are headed by females. Let $y_{ij}(t)$ denote the percentage of families with female heads in county $i$ of state $j$ in year $t$. Let $x_{ij}(t)$ denote the set of observed county- and year-specific economic and marriage opportunity variables. Let $s_j(t)$ denote the set of state- and year-specific welfare policy variables, and let $d(t)$ be a dummy variable indicating whether the observation came from 1990. The year dummy, $d(t)$, is used to account for "global" effects that alter the national trend in family formation (this probably includes what Qian and Preston [1994] have called changes in the "force of attraction"). With this notation, a standard regression of the relationship between family formation outcomes and local economic, marriage opportunity, and policy conditions can be written as

$$y_{ij}(t) = \beta_0 + \beta_X' x_{ij}(t) + \beta_S' s_j(t) + \beta_D d(t) + \epsilon_{ij}(t), \tag{1}$$

where $\epsilon_{ij}(t)$ represents unobserved county- and year-specific determinants of family formation. In the results that follow, this specification is estimated by applying OLS to the pooled county-level samples where each observation is weighted by the number of families in the county.

Estimates of the coefficients in model (1) are biased if the error term, $\epsilon_{ij}(t)$, includes unobserved factors that are correlated with the variables in $x_{ij}(t)$ or $s_j(t)$. For instance, states with higher welfare benefits may be more generous providers of education and social services generally. If this

[9] The specific denominations are the Assemblies of God, Associate Reformed Presbyterian Church, Christian Reformed Church, Church of God (Anderson, Indiana), Church of God (Cleveland, Tennessee), Church of the Nazarene, Evangelical Congregational Church, Free Methodist Church, Lutheran Church Missouri Synod, Mennonite Church, North American Baptist Conference, Pentecostal Holiness Church, Reformed Church in America, Southern Baptist Convention, and Wisconsin Evangelical Lutheran Synod. These denominations were identified by the National Right to Life Committee (personal communication) and Bosgra (1987) as having strong antiabortion positions. Several other denominations including the Free Will Baptists, International Church of the Foursquare Gospel, Old Order Amish, and Wesleyans were also identified as having strong antiabortion positions; however, data were not consistently available for these groups on numbers of adherents.

additional service spending is at the same time negatively related to headship and omitted from model (1), estimates of the effect of welfare benefits will be biased downward. Conversely, liberal (conservative) political and social attitudes may account for both high (low) welfare spending and high (low) rates of headship. Omitting such attitudes from regression (1) leads to upwardly biased estimates of the effects of welfare.

To control for unobserved time-invariant factors, several recent studies have estimated models that incorporate state-specific fixed effects. Using the notation from specification (1), such a model can be expressed as

$$y_{ij}(t) = \beta'_X x_{ij}(t) + \beta'_S s_j(t) + \beta_D d(t) + \mu_j + \epsilon^*_{ij}(t), \tag{2}$$

where $\mu_j$ represents a state-specific effect (with $\beta_0$ suppressed, a state-specific intercept). While the inclusion of state fixed effects likely mitigates biases associated with unobserved state spending and institutions, the specification does not control for unobserved variation in local determinants or for interactions between omitted state and local variables.[10]

The availability of repeated county-level data allows us to include a finer set of fixed-effects controls. Specifically, we estimate models of the form

$$y_{ij}(t) = \beta'_X x_{ij}(t) + \beta'_S s_j(t) + \beta_D d(t) + \tilde{\mu}_{ij} + \tilde{\epsilon}_{ij}(t), \tag{3}$$

which nest specification (2) as a special case. Because our primary interest rests with the coefficients $\beta_X$, $\beta_S$, and $\beta_D$, we can obtain consistent estimates of the relevant parameters of equation (3) by applying OLS to data that have been differenced across 1980 and 1990.[11] For purposes of comparability, we estimate variants of specifications (1), (2), and (3). However, on the basis of specification tests for the inclusion of county-specific versus state-specific or no fixed effects, the general model (3) is our preferred specification in the subsequent empirical analysis.

## RESULTS

### Changes in Female Headship

The percentage of families headed by females continued its steady rise during the 1980s. The weighted county data in table 1 indicate that the

[10] The use of the word *likely* is purposeful here as the use of fixed effects may exacerbate biases associated with other statistical problems such as measurement error in the explanatory variables or omitted local-level variables.

[11] Even after differencing, there still may be spatial correlation in the errors because of the clustering of counties within states. To account for this, we report estimates from feasible generalized least squares (FGLS) models that incorporate controls for within-state correlation.

percentage of female-headed families increased from 16.03% to 18.72% between 1980 and 1990, an absolute increase of 2.69% and a relative increase of 16.7%.[12] On both an absolute and percentage basis, the increase was largest among black women, whose headship rates rose from 41.30% to 48.10% (table 2). The increase among Hispanics was more similar to that of whites than blacks.

Over the same period, average real AFDC benefit levels declined by $105, or 19%. In terms of the combined benefits package, the decline in AFDC generosity was reinforced by a slight drop over the decade in average Medicaid benefits but partially mitigated by the income adjustment for food stamps. On net, the evidence is still one of substantial cutbacks in the real value of the welfare "safety net." The average combined monthly benefit dropped by $166, or 18%, over the decade.

Averages reweighted to reflect the different residential distributions of racial and ethnic groups (not shown) further reveal that black families were more likely to live in states with lower-than-average welfare benefits in both 1980 and 1990. The paradox is that female headship increased, especially among blacks, during a period of decline in AFDC and of divergence in the welfare safety net that increasingly favored whites over blacks. This pattern of decline provides strong evidence against the argument that welfare was responsible for the rise over the decade in female-headed families.

## Pooled Cross-Sectional Models of Female-Headed Families

Like the McLanahan-Casper study, our initial model (shown in the first column of table 3) provides estimates of welfare incentive effects from a 1980–90 pooled regression specified as in equation (1). Consistent with theory, this conventional analysis indicates that the association between welfare benefits and female headship is positive and statistically significant. At the same time, the effects are arguably small from a substantive standpoint: a $100 change in welfare payments is associated with slightly less than a one percentage point change in female headship. To give a more extreme example, if monthly AFDC payments had been *completely* eliminated and not offset by any change in food stamps or Medicaid (i.e., if the average monthly combined benefits package had decreased by $565 instead of $166 over the decade), the results indicate that headship rates would have indeed fallen but only by about half a percentage point.

A criticism of previous cross-section research is that estimates of the

[12] Since the county data are weighted by the number of families in the county, these weighted averages are algebraically equivalent to the population estimates of the percentage of female-headed families for the United States.

TABLE 3

DETERMINANTS OF FEMALE HEADSHIP REGRESSIONS WITH ALTERNATIVE AREA CONTROLS, 1980–90

| Variable | No Area Effects | State Effects[a] | County Effects[b] |
|---|---|---|---|
| Maximum combined welfare benefits | .817*** | .392*** | .838*** |
| | (.036) | (.105) | (.214) |
| Sex ratio | −.178*** | −.171*** | −.076*** |
| | (.005) | (.005) | (.008) |
| Male earnings | −.421*** | −.403*** | −.092*** |
| | (.016) | (.016) | (.023) |
| Male education | −.007 | .048*** | .045** |
| | (.018) | (.017) | (.017) |
| Male employment | −.285*** | −.270*** | −.161*** |
| | (.008) | (.009) | (.014) |
| Female earnings | .317*** | .306*** | −.432*** |
| | (.035) | (.034) | (.040) |
| Female education | .014 | −.041*** | −.159*** |
| | (.020) | (.020) | (.018) |
| Percentage 65 and over | −.185*** | −.123*** | −.135*** |
| | (.014) | (.015) | (.034) |
| Percentage black | .271*** | .316*** | .257*** |
| | (.004) | (.005) | (.017) |
| Percentage Hispanic | −.001 | .027*** | −.147*** |
| | (.004) | (.005) | (.017) |
| Percentage rural | −.087*** | −.077*** | −.019** |
| | (.003) | (.003) | (.008) |
| ln(population) | .241*** | .383*** | −2.015*** |
| | (.047) | (.049) | (.336) |
| Percentage Catholic | −.038*** | −.060*** | −.025*** |
| | (.003) | (.004) | (.005) |
| Percentage LDS | −.039*** | −.063*** | .121** |
| | (.006) | (.016) | (.055) |
| Percentage conservative Protestant | −.042*** | −.011** | −.052*** |
| | (.004) | (.005) | (.016) |
| $\text{Dummy}_{1990}$ | 3.029*** | 2.465*** | 6.053*** |
| | (.141) | (.205) | (.387) |
| $R^2$ | .834 | .862 | ... |

NOTE.—Results based on 6,106 county-level observations from 1980 and 1990 weighted by the no. of families in each county. Dependent variable is percentage of families with children under 18 years old with single female heads. SEs appear in parentheses.

[a] The 49 state effects are jointly significant at the .01 level.

[b] The 3,053 county effects are jointly significant at the .01 level.

* $P < .10$.

** $P < .05$.

*** $P < .01$.

effects of public assistance may be spurious, a result from excluding other state-level variables associated both with welfare generosity and female headship from the model. The model presented in the second column addresses this concern by including 48 state dummy variables as controls for unmeasured state-level effects (DiPrete and Forristal 1994).[13] Specification tests reveal that the state dummy variables are jointly significant at the .01 level. More important, their addition leads to a substantial (more than 50%) reduction in the estimated effect of welfare. Although the coefficient remains significantly positive, each additional $100 change in welfare benefits is now estimated to lead to only a .392 percentage point change in female-headed family households. Estimates of the coefficients for several other variables including men's and women's education and the religious affiliation variables are also very sensitive to the inclusion of state fixed effects. On balance, the concerns expressed in the existing literature about potential biases from omitted variables, even in models such as ours that are augmented by a relatively rich and detailed set of observed control variables, appear to be well founded (cf. Moffitt 1994; McLanahan and Casper 1995).

### Change Models of Female-Headed Families with County Fixed Effects

Our results so far indicate that welfare benefit levels significantly affect the percentage of female-headed families with children, even when other unobserved state effects are controlled. These models do not, however, take full advantage of the longitudinal county-level data. The possible effects of county-specific unobserved heterogeneity are addressed in the model presented in the final column of table 3, which differences each of the county-level observations over time.[14] This model provides an estimate of the effect of *changing* public assistance on *intradecade* changes in female headship at the county level, while controlling for unobserved heterogeneity between both states *and* counties (DiPrete and Forristal 1994) as well as for observed changes in local area sex ratio imbalances, local economic opportunities, and cultural factors.[15] As with the previous speci-

[13] In studies using data from a single point in time, the state dummies and welfare variables would be redundant and welfare effects could not be estimated. This is not the case using pooled data. Each state has two unique welfare variables (i.e., for 1980 and 1990), allowing us to estimate a state welfare effect independent of state effects.

[14] In this first-difference model, the dependent variable is the difference between the 1980 and 1990 percentages of families with children that were headed by women. The independent variables are similarly differenced. In terms of the coefficients for the variables of substantive interest, the specification is equivalent to having run a regression on the pooled sample with 3,052 county dummy variables.

[15] These models were estimated using the HML/2L computer program. This package does not include an estimate of $R^2$ from the model.

fication, estimation reveals that the additional fixed effect controls are jointly significant and that their inclusion leads to substantive changes in several of the coefficients of interest.

The county fixed-effects model produces estimates of the association between welfare and female headship that are much closer to pooled regression estimates from column 1 in table 3 than the state fixed effects estimates from column 2. Specifically, each $100 change in welfare benefits is associated with an .838 percentage point change in the county female headship rate. When we combine this point estimate with the actual decrease in welfare benefits over the decade, the change in welfare policy is calculated to have depressed headship rates by 1.4%. The finding of a significant positive association between welfare and headship is all the more striking because we have fit our model for female headship among all families with children, not simply for the low-income, potentially eligible population. Models estimated for such a targeted population would likely produce even stronger results than those reported here.[16]

Our difference model also is the first of its kind to examine the effects of *changing* local marriage market conditions on family formation for the 1980s. It builds directly on previous pooled cross-sectional analyses of 1980 and 1990 census data (Lichter et al. 1991; McLanahan and Casper 1995) and on Wood's (1995) study, which was restricted to the metropolitan black population using change data computed from the 1970 and 1980 censuses. Our results reveal significant effects of changing marriage market conditions on the 1980s rise in female-headed families.

Specifically, the estimates indicate that the sex ratio is negatively associated with female headship; areas with relatively fewer men to women have higher rates of female headship than do other areas. These effects, though, are small from a substantive standpoint with the coefficient implying that a decline of 13 men for every 100 women would be necessary to raise the female headship rate by a single percentage point. Men's employment and earnings are significantly negatively associated with female

[16] Our analysis has focused primarily on the effects of welfare benefits overall, without regard to the "packaging" of benefits. In some additional analyses (available upon request), we fit difference models (with county effects) that included the maximum AFDC benefits for a family of four and the average Medicaid benefits received by AFDC recipients *separately*. Our evaluation assesses claims that the rise in female headship reflects the pernicious effects of Medicaid, which provides medical coverage that is typically unavailable in the low-wage jobs held by low-income women. The results provided a straightforward conclusion: Welfare incentive effects result from changes in AFDC rather than Medicaid payment levels. The effects of changes in average Medicaid payment levels were not associated with changes in female headship for the sample of all counties, nor in the race disaggregated samples of whites, blacks, and Hispanics. There is little evidence here that Medicaid payment levels promoted the rise in female-headed families over the past decade.

headship. The magnitudes of these estimated effects, however, are again comparatively small: a 10% change in the male employment rate (an absolute change of 7 percentage points) is associated with a 1.1 percentage point change in headship while a 10% swing in male earnings is associated with only a .3 percentage point change (by way of contrast, the change in female headship associated with a 10% change in welfare payments is .7 percentage points). The results while modest nevertheless imply that the supply of marriageable men influences women's decisions to bear or raise children outside of marriage.

One piece of evidence that potentially contradicts this hypothesis is the small but signficantly positive coefficient on the percentage of men in the county who have completed at least a bachelor's degree. To the extent that this variable serves as a proxy for long-term wage opportunities, the positive coefficient runs counter to the marriage market explanation. However, to the extent that the variable reflects either a shortage of marriageble teenage and young adult men, liberal attitudes regarding sexual activity, or progressive attitudes about and a willingness to provide child support, the coefficient can be reconciled with our other results.

The results from the county fixed-effects model fail to support the female "independence hypothesis" but do support economic hypotheses regarding family formation processes. This represents strong support for Oppenheimer (1994), who claims that the recent emphasis on the deleterious effects of women's improved economic status on the family is overdrawn. Specifically, our estimates indicate that women's earnings and education are significantly negatively related to female headship. Previous cross-section research on the effects of women's changing economic roles on various family formation outcomes has been mixed with a large number of economic studies finding negative effects of female economic opportunities (see the review by Montgomery and Trussell [1986]) but with several sociological and some economic studies finding positive effects (see, e.g., McLanahan and Casper 1995; Lichter et al. 1991; Matthews, Ribar, and Wilhelm 1997). The estimated magnitudes of these associations are also much larger than those for the male earnings and education variables, in the case of the earnings variable nearly five times as large. The schooling results are consistent with higher levels of education increasing women's access to economically attractive marital partners, increasing their attractiveness to potential marital partners, and thus increasing the quality and stability of marriages generally (South 1991; Lichter et al. 1995).

As expected, high rates of female headship are significantly associated with urbanization (see also McLaughlin and Lichter 1997), high concentrations of blacks and young people, and local population declines. The percentages of Hispanics, Catholics, and antiabortion Protestants in the county, on the other hand, are negatively associated with the percentage

of families headed by women. A counterintuitive finding is that changes in the percentage of Mormons are positively associated with female headship. This last result notwithstanding, it is clear that changes in the demographic and cultural character of local areas affect the concentration of female headship, independent of changes in economic conditions.

Finally, the intercept from the differenced model can be interpreted analogously to the dummy variables in our previous specifications indicating whether the observation pertains to 1980 or 1990 (coded "1" if 1990). If the year effect is zero, the appropriate inference is that the 1980s rise in female-headed households with children was due entirely to changes in the state and county characteristics considered here. But, as shown in table 3, the net year effect of 6.053 is large and statistically significant in the model. In contrast, the percentage of family households (with children) headed by females increased by 2.68 percentage points, on average, between 1980 and 1990 (table 1). The significant time coefficient means that neither county compositional changes nor changing state welfare policy can explain the 1980s rise in female-headed families. In fact, current short-run demographic and economic trends have muted the upward rise. Other explanations clearly must be entertained, including cultural ones that emphasize changing family values or rising individualism.[17]

In some additional analyses (table A1), we evaluated the sensitivity of our results using three alternative measures of family structure. Previous research on welfare is often unclear about measurement issues (e.g., whether subfamilies are defined as families) or the unit of analysis (i.e., percentage of women heading families with children or percentage of families with children that are headed by women). Such differences may contribute to differences from study to study in reported welfare effects. Our sensitivity analysis focused on three alternative measures: (1) percentage of women heading family households with children, (2) percentage of families (including subfamilies) with children headed by women, and (3) percentage of women heading families (including subfamilies) with children. Regardless of measure used, estimates from county fixed-effect models (models specified along the lines of eq. [3]) revealed statistically significant and positive effects of changes in welfare benefit levels on the change in female headship. The welfare incentive effects are thus robust with respect to alternative measures of female headship. Our findings regarding

[17] The coefficient on the time trend is more properly interpreted as a residual that incorporates all of the trends that we have either omitted or improperly measured. So the trends in low-income wages, the availability and relative value of other forms of public assistance, child care and child support policies, the lifetime parity of married and single mothers, etc., are all potential contributors to this residual along with the cultural trends described in the text.

the negative effects of the sex ratio, men's employment, and women's economic opportunities were also robust to changes in the specification of the dependent variable, while our findings regarding men's earnings and education were somewhat more sensitive.

### Female Headship among Whites, Blacks, and Hispanics

The clear and consistent relationships between welfare and female headship for the total sample are not observed for each of the racial and ethnic groups considered here. The estimates from the 1980–90 difference or county-effects models reported in table 4 are based on counties with significant populations (more than 50 families) of whites, blacks, and Hispanics. The results, which have controlled for unobserved heterogeneity among counties, indicate significant positive effects of welfare benefit sums on female headship among blacks but not among whites and Hispanics. For blacks, the welfare effect (1.294) is over four times larger than the nonsignificant effect for whites (.302). This effect implies that each $100 change in welfare benefits is associated with a 1.294 change in the percentage of family households with children headed by females. Alternatively, this means that counties with the largest welfare cuts experienced relatively slower increases in female headship over the decade. For blacks, the observed welfare effect implies that average black female headship would decline from the observed percentage of 48.10% to 43.01%, a 5.09 percentage point drop, if welfare payments were cut in half. Such results suggest that even large cuts in welfare benefits will produce relatively small declines in female headship for blacks. They will yield no change in female headship for whites or Hispanics.

The other results in table 4 reinforce previous research that emphasizes deteriorating marriage market conditions as a causative factor in the rise in female-headed families. For each racial/ethnic group, increases over 1980–90 in the ratio of men to women and increases in men's earnings, education, and employment were associated with slower increases (or even declines) in female headship. But the effects of shortages of economically attractive men were strongest among minority populations. For example, each percentage point increase in black men's and Hispanic men's employment during the 1980s was associated with a .252 and .287 percentage point decline, respectively, in female headship. This compares with a .099 percentage decline in white female headship for each percentage point increase in white men's employment. Similarly, the effect of men's earnings on female headship was over twice as large for blacks and Hispanics. Clearly, our results indicate that changes in local economic circumstances played an especially large role in the rise in minority female-

TABLE 4

DETERMINANTS OF FEMALE HEADSHIP BY RACE/ETHNICITY

| Variable | White | Black | Hispanic |
|---|---|---|---|
| Maximum combined welfare benefits | .302 | 1.294** | −.111 |
| | (.198) | (.611) | (.803) |
| Sex ratio | −.080*** | −.041*** | −.092*** |
| | (.008) | (.009) | (.010) |
| Male earnings | −.151*** | −.346*** | −.164** |
| | (.020) | (.066) | (.066) |
| Male education | .022 | −.058 | .101** |
| | (.015) | (.045) | (.047) |
| Male employment | −.099*** | −.252*** | −.287*** |
| | (.013) | (.029) | (.030) |
| Female earnings | −.397*** | −.265** | .064 |
| | (.034) | (.098) | (.110) |
| Female education | −.132*** | −.095** | −.146*** |
| | (.015) | (.045) | (.050) |
| Percentage 65 and over | −.040 | .058 | .181 |
| | (.031) | (.121) | (.144) |
| Percentage black | .011 | −.259*** | .154* |
| | (.018) | (.044) | (.085) |
| Percentage Hispanic | −.076*** | −.163** | −.110** |
| | (.016) | (.060) | (.051) |
| Percentage rural | −.016** | .009 | −.074* |
| | (.007) | (.037) | (.039) |
| ln(population) | .582 | −5.104*** | .657 |
| | (.301) | (1.357) | (1.390) |
| Percentage Catholic | −.018*** | −.011 | −.010 |
| | (.005) | (.026) | (.013) |
| Percentage LDS | .107** | .463 | .754*** |
| | (.049) | (.428) | (.274) |
| Percentage conservative Protestant | −.021 | .085 | −.077 |
| | (.014) | (.064) | (.094) |
| $\text{Dummy}_{1990}$ | 3.820*** | 9.819*** | 3.303** |
| | (.359) | (1.075) | (1.558) |
| Observations | 5,800 | 2,366 | 1,542 |

NOTE.—Dependent variable is percentage of families with children under 18 years old with single female heads. Regressions include controls for county-specific effects and are weighted by the no. of families in each racial/ethnic group. SEs appear in parentheses.

* $P < .10$.

** $P < .05$.

*** $P < .01$.

headed families in the 1980s, a result consistent with previous research (Mare and Winship 1991).

The observed effects of women's changing economic circumstances on family formation similarly reinforces the (too often ignored) theoretical arguments and empirical evidence of Oppenheimer (1994), that is, that women's improving earnings and education promoted marriage rather than created disincentive to marriage or female independence from men. Indeed, for both white and black women, improvements in earnings were negatively associated with 1980–90 increases in female-headed families. These estimated effects, however, were generally larger and more precise among white women than among blacks and Hispanics. This is not surprising in light of existing racial and ethnic differentials in female education and earnings; fewer benefits in the form of lower female headship are likely to be observed if the education and earnings distributions are truncated at the top among minority women.

In sum, our models (with county fixed effects) indicate that welfare benefit levels are positively associated with the rise in female headship under varying model specifications and in race specific models for blacks. At the same time, our results highlight the importance of local economic opportunities, especially among men, in promoting traditional patterns of family formation. Clearly, neither monocausal explanations that focus on welfare incentives nor explanations that emphasize jobs and earnings can fully account for recent increases in female headship. From a policy standpoint, this is perhaps discouraging because it implies that cultural changes—such as value shifts—may partly underlie recent trends (Popenoe 1996). Attitudes and values are difficult to reverse by public policy, even if consensus about the need for such change exists.

## DISCUSSION AND CONCLUSION

With the passage of the Personal Responsibility and Work Opportunity Reconciliation Act of 1996, welfare reform is on the public policy and state legislative agenda. Our study revisits the subject of possible links between changing national and state welfare policy and the rise in female-headed families—a timely but contentious subject that has been given surprisingly little systematic empirical attention in the sociological literature over the past decade.[18] Our multilevel analysis of the multiple causes

[18] The voice of social science research has not often been heard in this debate. It has been drowned out by a much louder and more passionate public voice of concern about the "breakdown of the family" and the possible contributing role of welfare dependency. This is not surprising. Sociologists typically have shied away from sensitive or controversial public policy issues that might seem to "blame the victim," that might promote racial stereotypes, or that might generate empirical evidence supporting a conservative political agenda. For example, Massey's (1995) recent review essay

of the rise in female-headed families is especially appropriate at a time when state and federal welfare reform initiatives seek to restore "strong families," reduce nonmarital fertility, and promote economic independence among poor women.

Our analyses of changing female headship provided strong and consistent evidence that state welfare benefit levels were associated with changes in female headship over the 1980–90 period. Previous studies using state fixed-effects models of female headship have produced inconclusive results for the 1960s (Ellwood and Bane 1985) or have showed that state welfare benefit effects are nonsignificant (for 1968–89) when unobserved state-level variables are controlled (Moffitt 1994). Our fixed-effects analysis, which controlled for both unobserved state and county heterogeneity, indicated instead that areas with larger declines in welfare benefits had significantly slower increases in the percentage of families headed by unmarried mothers. Moreover, the welfare incentive effects observed here may in fact be underestimated if welfare has *lagged* effects on family decision-making processes (Murray 1993; Moffitt 1994). Family formation behaviors may respond slowly to changing welfare policy as perceptions of welfare benefit levels catch up with the reality of recent welfare cuts. They also may be underestimated for certain high-risk populations. For example, marriage rates were lowest in the 1980s for young, poor, and less educated women, a group that contributed disproportionately to recent increases in female-headed families and may have been influenced most by welfare availability (Qian and Preston 1993; McLaughlin and Lichter 1997).[19]

Our study perhaps built most directly on the cross-sectional study by McLanahan and Casper (1995) by further explicating the role of welfare incentives in the family formation process. First, ours is the first comprehensive study for the 1980s to pit explanations that emphasize state welfare incentives against those that stress the deteriorating economic attractiveness of male marriage partners and the rising economic independence of women. We show that the effects of *changing* welfare benefits existed under a variety of alternative and rigorous model specifications for *all* (rather than only metro) counties. The robustness of our results for the 1980s also reinforce recent speculation that welfare incentive effects may have increased since the 1970s (Moffitt 1996).

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of *The Bell Curve* argues that the attention received by this book can be traced directly to the policy vacuum left by sociologists.

[19] Some support for this argument was found in additional analyses (not reported) that revealed a significant negative interaction effect of welfare and education on female headship. The substantive interpretation is that welfare effects were largest in the counties with the least educated population, a result consistent with arguments suggesting larger welfare effects among the poor.

Second, welfare incentive effects during the 1980s were especially large for blacks (a result similarly shown but not discussed by McLanahan and Casper [1995]), while the effects for whites and Hispanics were insignificant. Such results for the 1980s contrast with previous research that typically showed that welfare effects were larger for whites than for blacks (see Moffitt 1995). The welfare incentive effects observed here for blacks are consistent with the larger share of poor and welfare-dependent persons in the black population. By virtue of their lower economic status alone, blacks should be more responsive to changing welfare benefit levels. The race-disaggregated results from our county fixed-effect models therefore help reconcile the theoretically anomalous results of previous studies.

Third, conventional economic arguments, including those that emphasize employment or welfare, cannot account for recent upswings in female-headed families with children. The "period" effects in our pooled regression models were invariably large and statistically significant. Female headship increased between 1980 and 1990, net of 1980–90 compositional changes in the sex ratio, the economic circumstances of men and women, or welfare. The fact that the size of the period effect exceeded the actual 1980–90 percentage point change in female headship indicates that recent demographic and economic changes may have contributed, on balance, to otherwise even larger increases in female headship over the past decade. In the end, welfare was not the primary or even a key factor responsible for the recent upswing in female headship. The estimated welfare incentive effects implied that even very large cuts in welfare payment levels would produce only a relatively small drop in the rate of female family headship. Little evidence exists to support the apparently widely held perception that welfare is largely responsible for the breakdown of the traditional married-couple family.

Clearly, other explanations for the rise in female-headed families are required and await additional study. For example, until recently, cultural explanations have had a bad reputation, one born of 1960s debates about a "culture of poverty" and the Moynihan report. Cultural arguments seemed to blame the poor themselves (e.g., "tangle of pathology") for their unfortunate economic circumstances (for cultural discussions, see Cherlin [1992] and Pagnini and Morgan [1996]). As in previous studies, our results also lend themselves to possible cultural interpretations.[20] Such explana-

[20] "Residual" evidence supporting the so-called minority-group hypothesis provides a similar example. Studies show that the higher fertility among minority women cannot be explained with standard social (e.g., education) and economic (e.g., income) variables. The usual inference is that this unexplained residual variation in fertility must then be due to unmeasured cultural differences between majority and minority groups. Such a residual interpretation also is used when differences in income between groups (e.g., blacks and whites, men and women) cannot be explained by objective human

tions are implied by increases in female headship even after adjusting for 1980–90 shifts in conventional social and economic variables (e.g., growing shortages of marriageable men). A cultural argument—one based on observed race differences—also is indicated by the disproportionate increase in female-headed families in counties with increasing black populations, holding marriage market and other local economic conditions constant. One interpretation, then, is that the effects of changing welfare and economic factors reported here represent only small deviations from an otherwise upward trend in female headship across many segments of American society.

Some observers have argued that the retreat from the traditional family is a result of the widespread rise in individualism at the expense of the collectivity, changing mores regarding sexuality and unmarried cohabitation, and the declining stigma associated with unmarried pregnancy and motherhood (Bumpass 1990; Thornton 1995; Popenoe 1996). Indeed, only about one-third of young people today agree that it is better to get married than to spend one's life being single, and three-fifths express moral acceptance of cohabitation before marriage (Thornton 1989). Moreover, black Americans—especially black men—are less likely than their white counterparts to indicate a desire to marry (South 1993). And, unlike white women, black women typically suffer little stigma from unmarried childbearing (Pagnini and Morgan 1996). Blacks also are less likely to marry than Mexican Americans, a group that shares the disadvantaged economic circumstances of blacks but not the same familial cultural traditions (Oropesa, Lichter, and Anderson 1994). Of course, any distinctive features of contemporary black family life may be cultural adaptations and a legacy of historical circumstances, including spatial and social isolation, chronic economic deprivation, peer group attachments, and gender role dissonance (Pagnini and Morgan 1996; Lloyd and South 1996; Schoen and Kluegel 1988). By their very nature, such sweeping cultural interpretations often defy empirical study.

Whatever its etiology, our focus on the growth of family households headed by unmarried women with children is clearly appropriate. Current welfare debates and nascent state welfare legislation remain targeted at needy unmarried women with children. Our estimates of welfare incentive effects must nevertheless be regarded as reduced-form effects; as in previous studies, we have ignored the multiple pathways through which welfare may have contributed to the rise in female-headed families with children (Ellwood and Bane 1985). Welfare incentive effects on female headship may operate indirectly through increased nonmarital childbear-

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capital or job-related characteristics. Here, the residual interpretation is usually one of unmeasured overt or subtle race or gender discrimination.

ing (Moffitt 1995), lower marriage or remarriage rates (McLanahan and Casper 1995), more independent living among unmarried women with children (Wilson 1987), a lower likelihood of resolving premarital pregnancies through abortion or marriage (Lundberg and Plotnick 1990), or more cohabitation at the expense of marriage (Manning and Smock 1995; Raley 1996). Different state welfare reform initiatives may have much different and perhaps offsetting effects on the various demographic pathways to female headship and poverty (Lichter and Gardner 1996–97). Our results should be regarded as a first step toward a fuller understanding of the etiology of family formation behaviors.

Finally, the current devolution of the federal welfare system to the states implies considerable potential for future increases in state-to-state variation in welfare policy and benefit levels. The United States currently may be undergoing a process of increased territorial differentiation—a balkanization across geographic space of cultural and economic groups (Frey 1995; Lichter 1992). The future may yield growing spatial heterogeneity in family formation processes, including the rise of female-headed families, that reinforce existing economic disparities over geographic space (Massey 1996). Indeed, the welfare incentive effects observed here may grow over time as state welfare policies further differentiate one state from another.

## APPENDIX

TABLE A1

DETERMINANTS OF FEMALE HEADSHIP: ALTERNATIVE HEADSHIP MEASURES

| Variable | Household Headship per Woman | Family Headship | Family Headship per Woman |
|---|---|---|---|
| Maximum combined welfare benefits | .378*** | 1.021*** | .538*** |
| | (.099) | (.262) | (.117) |
| Sex ratio | −.026*** | −.084*** | −.034*** |
| | (.004) | (.008) | (.004) |
| Male earnings | −.010 | −.113*** | −.008 |
| | (.011) | (.024) | (.013) |
| Male education | −.008 | .080*** | .033*** |
| | (.008) | (.018) | (.010) |
| Male employment | −.086*** | −.167*** | −.106*** |
| | (.007) | (.014) | (.008) |
| Female earnings | −.248*** | −.314*** | −.225*** |
| | (.019) | (.041) | (.022) |
| Female education | −.078*** | −.199*** | −.133*** |
| | (.009) | (.019) | (.010) |
| Percentage 65 and over | −.044*** | −.142*** | −.123*** |
| | (.016) | (.035) | (.018) |
| Percentage black | .051*** | .324*** | .129*** |
| | (.008) | (.017) | (.009) |
| Percentage Hispanic | −.076*** | −.030* | −.005 |
| | (.008) | (.017) | (.009) |
| Percentage rural | −.008** | −.019** | −.010** |
| | (.004) | (.008) | (.004) |
| ln(population) | −.244 | −3.098*** | −1.439*** |
| | (.161) | (.350) | (.186) |
| Percentage Catholic | −.014*** | −.015*** | −.009*** |
| | (.002) | (.005) | (.003) |
| Percentage LDS | .086*** | .082 | .062* |
| | (.028) | (.058) | (.033) |
| Percentage conservative Protestant | −.006 | −.067*** | −.023** |
| | (.008) | (.016) | (.009) |
| $\text{Dummy}_{1990}$ | 2.386*** | 7.566*** | 3.811*** |
| | (.179) | (.472) | (.213) |
| Mean of the dependent variable: | | | |
| 1980 | 6.919 | 18.049 | 8.018 |
| 1990 | 7.614 | 22.216 | 9.668 |

NOTE.—The weight for the first and third columns is women 15–44 years old. The weight for the second column is families with children. Regressions include controls for county-specific effects. SEs appear in parentheses.

* $P < .10$.

** $P < .05$.

*** $P < .01$.

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