WHY HAVE WOMEN BECOME LEFT-WING? THE POLITICAL GENDER GAP AND THE DECLINE IN MARRIAGE*

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The last three decades have witnessed the rise of a political gender gap in the United States wherein more women than men favor the Democratic party. We trace this development to the decline in marriage, which we posit has made men richer and women poorer. Data for the United States support this argument. First, there is a strong positive correlation between state divorce prevalence and the political gender gap—higher divorce prevalence reduces support for the Democrats among men but not women. Second, longitudinal data show that following marriage (divorce), women are less (more) likely to support the Democratic party.

I. INTRODUCTION

If only women had voted in the 2000 U. S. Presidential election, the Democratic candidate Al Gore would have won a landslide victory: 54 percent of female voters cast their vote for him. However, 53 percent of men voted for Bush [Voter News Service exit poll, reported in *The New York Times,* November 12, 2000]. This striking difference in political preferences between men and women is a significant feature of the present political landscape [Becker February 1997; Inglehart and Norris 2000; Norris forthcoming]. However, it is a recent development.

Until the mid-1960s, women were consistently more conservative than men [Duverger 1955; Harvey 1998]. In the 1980s a significant number of men, so-called Reagan Democrats, switched party allegiance to the Republicans, leading to a political hegemony of the right. The 1990s saw previously conservative voting women, so-called Soccer Moms, moving to the left, resulting in the Clinton years [Stark 1996]. The consequence is that over the past

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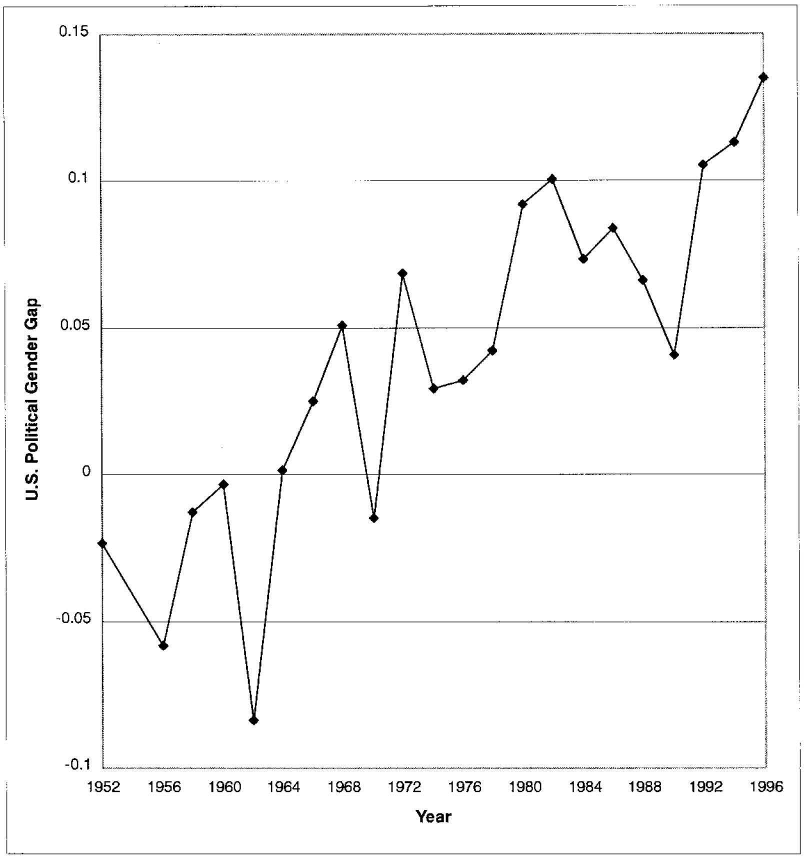

FIGURE I

U. S. Political Gender Gap

Notes. The U. S. Political Gender Gap is defined as the difference between the proportion of female and male respondents who are Democrat. The gap is constructed using respondent-level information from the National Election Studies data 1952–1996, where the sample is restricted to respondents aged 18–64 years. A respondent is defined as a Democrat if he/she states self to be a Strong-, Weak-, or Independent-leaning Democrat. Appendix 1 provides a full description of the National Election Studies sample.

twenty years the gap between men's and women's political preferences has reversed direction, and it has become significant to the extent that in the last two elections men and women would have chosen different presidents.

Figure I illustrates the evolution of this political gender gap in the United States between 1952 and 1996. The period saw the

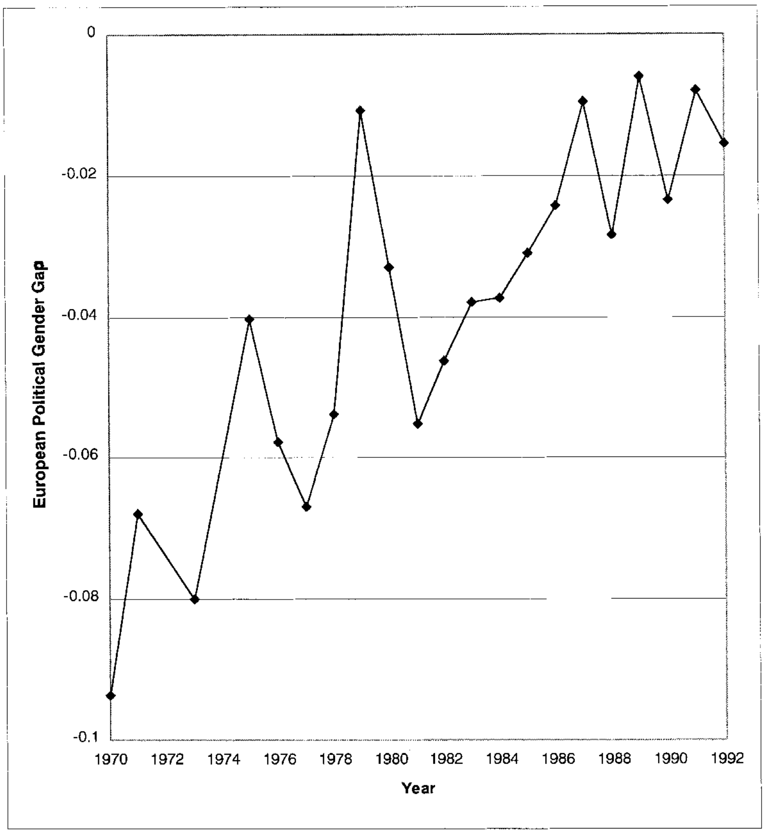

FIGURE II
European Political Gender Gap

Notes. The European Political Gender Gap is the population weighted average Gender Gap for ten European countries. For each country the European Political Gender Gap is defined as the difference between the proportion of female and male respondents who favor the Left. For each country the gap is constructed using respondent level information from the annual Eurobarometer Surveys, where we include information on all respondents aged fifteen and over. A respondent favors the Left if his/her stated party preference is for a Left-wing party in the country. The Eurobarometer Survey provides complete identification of all parties in a country as Left/non-Left. The countries are Germany, Italy, France, the Netherlands (1970–1992), Denmark, Ireland, Luxemborg (1973–1992), United Kingdom (1970, 1973–1992), and Greece (1980–1992).

gap between the proportion of women and men who identify themselves as Democrats increase from −2 to 12 percent. A nearly identical trend is evident in Europe (Figure II).

The United States also witnessed a fall of over a quarter in the proportion of currently married adults, and a threefold rise in the proportion of currently divorced individuals in the last three decades.[1] We argue that men transfer resources to women in marriage. We further argue that this decline in marriage made women poorer relative to men and thereby contributed to the political gender gap. Underlying the latter argument is the assumption that individual political party affiliation is determined by (per capita) income through its effect on preferences with respect to redistribution. This hypothesis provides the following testable predictions.

First, it implies that a decline in marriage has affected political preferences principally among middle-income voters. While a poor man is richer if unmarried, he is still sufficiently poor to favor redistribution; similarly, rich women, while poorer if unmarried, remain rich enough to oppose redistribution. However, among the middle-income group, marital status impacts income sufficiently to affect political preferences. Second, the political impact of increased nonmarriage will depend on its incidence across middle-income groups.[2] For instance, if a relatively poor, i.e., left-leaning, couple divorces, support for the left will fall if the man becomes rich enough to favor the right. Conversely, if a relatively rich, i.e., right-leaning, couple divorces, support for the left will rise if the woman's income falls sufficiently. Third, if nonmarriage first affects the poor and thereafter extends upward in the income distribution, then we would expect men to shift right before women shift left.

Our empirical analysis focuses on testing the first prediction, and we find robust evidence. We note, however, that the two other predictions are consistent with stylized facts [Stark 1996].

First, we analyze survey data from the biennial National Election Studies (1964–1996) to examine whether changes in aggregate divorce risk affected male and female political preferences differently. We use two proxies for divorce risk: the extent of state-level divorce computed from the Current Population Survey, and the passage of unilateral divorce laws. We find a strong positive correlation between increased divorce risk and the political gender gap. We only find this correlation among middle-

1. Between 1964 and 1996 the proportion of adults aged 18–64 currently married fell from 84 to 58 percent, and those divorced rose from 3 to 10 percent (Current Population Survey, authors' calculations).

2. We use the term nonmarriage to emphasize that this category covers all individuals, including cohabitants, who are currently not married.

income respondents, irrespective of whether we measure political preferences by an individual's party affiliation or redistributive preferences.

Second, we directly examine how changes in marital status affect an individual's party affiliation. To this end, we analyze three waves of the Youth Parent Socialization Survey, a longitudinal study that interviewed a nationally representative sample of 1965 high school graduates in 1965, 1973, and 1982. We find that marriage and divorce affect a woman's party affiliation significantly more than they do a man's. Marriage tends to make a woman more Republican, whereas divorce tends to make her more Democratic. We find no evidence of a shift in political preferences presaging divorce for either sex. That is, changes in political affiliation between 1965 and 1973 do not predict changes in marital status between 1973 and 1982.

A number of alternative explanations for the evolution of the gender gap have been proposed. Our analysis investigates their relevance.

It has been suggested that the rise in female labor force participation makes women more likely to favor the left by increasing their awareness of labor market discrimination or raising demand for state-subsidized child care, or both. We find, however, that the correlation between divorce risk and the gender gap is robust to the inclusion of controls for both individual and aggregate labor force participation. We also find that working makes middle-income women, but not poor or rich women, more likely to favor the Democrats. An interpretation consistent with our hypotheses is that, for this group, women's decisions to work have been predicated on a fall in income from deteriorating marriage market conditions [Johnson and Skinner 1986]. We also show that increases in aggregate female labor force participation had no impact on political preferences other than for the richest 5 percent of households, where men became more Democratic.

An alternative explanation invokes the recent adoption of conservative stances on issues such as abortion rights or a woman's role in the family by the political right. The suggestion is that women will oppose these policies more than men. However, our empirical analysis shows that the issue of abortion rights did not affect men's and women's political preferences differently. This is in line with other surveys that consistently show no significant gender differences in either opinions or intensity of preferences on these issues [Mansbridge 1980; Cook and Wilcox 1991].[3] We find that the correlation between divorce risk and the gender gap for middle-income respondents is robust to the inclusion of controls for an individual's attitudes on social and religious issues.

The remainder of this paper is organized as follows. Section II situates our paper within the existing literature, and discusses the rationale underpinning our view of marriage. We provide a theoretical example to illustrate our proposed link between marriage, the gender gap, and overall demand for redistribution. Sections III and IV present our empirical findings. Section V concludes.

II. Background

Evidence of a growing political gender gap, in both redistributive and party preferences, has been documented in many surveys: for the United States, the National Election Studies [Chaney, Alvarez, and Nagler 1996; Montgomery and Stuart 1999], CBS News and New York Times quarterly surveys [Box-Steffensmeier, Boef, and Lin 2000], the General Social Surveys [Shapiro and Mahajan 1986; Alesina and Ferrara 2000]; and for Western European countries, the World Values Survey [Inglehart and Norris 2000]. In a similar vein, Lott and Kenny [1999] argued that female suffrage is behind the growth of government.

The papers most closely related to our study are Montgomery and Stuart [1999] and Box-Steffensmeier, Boef, and Lin [2000]. These papers note that changing demographics, especially the rise of nonmarriage, are correlated with the emergence of the political gender gap. Our innovation lies in providing an explanation for the likely effects of marriage on male to female income inequality, and in identifying several refutable predictions concerning the relationship between nonmarriage, the gender gap, and the overall demand for redistribution.

II.A. Marriage

We argue that marriage affects male to female income inequality because within marriage men transfer resources to

3. For instance, the General Social Surveys show that 41 percent of men and 39 percent of women supported abortion on request by the woman (question was asked in 1977–2000), and that 72 percent of men and 75 percent of women favored the Equal Rights Amendment (ERA) (question asked in 1982), authors' calculations.

women in exchange for sex and for access to children. This is because women are more discriminating than men in partner selection [Trivers 1972] and are vested with default property rights to the children they bear (e.g., Glendon [1996]).[4] Family law only recognizes one default parent, the mother. However, both parents may find it mutually beneficial to assign parental rights to the father as well. The outright sale of children is almost universally condemned. However, all known societies have devised contracts that link fathers to their children, and these contracts, however varied, are known as marriage (e.g., Morgan [1877], Mair [1953], and Posner [1992]). Hence, one way to understand marriage is to view it as a contract under which women provide men with parental rights [Edlund 1998], and in the majority of cases, also sex.[5] If women are compensated for this transfer, a decline in marriage may represent a shortfall in income for women.

This view of family formation is consistent with several stylized facts: women, on average, earn less than men; spouses' potential earnings are positively correlated [Becker 1991; Mare 1991; Qian and Preston 1993; Juhn and Murphy 1997]; high male relative to female earnings are conducive to marriage [Blackwell and Lichter 2000; Blau, Kahn, and Waldfogel 2000]; on divorce, female income falls substantially, with remarriage the main route to economic recovery [Weitzman 1985; Duncan and Hoffman 1985, 1988; Page and Stevens 2001].

Moreover, this view of marriage, unlike that proposed by Becker [1973], can account for the absence of negative sorting in the marriage market despite an increasing number of career-oriented women—women in high powered professions rarely marry men who specialize in so-called household production.

Another shortcoming of the framework proposed by Becker is that it is ill-suited to explain out-of-wedlock fertility, a family form that involves children, possibly cohabitation, but not marriage; or polygamy, a family form that involves marriage and children, but not necessarily cohabitation. By contrast, this paper's proposed view of marriage is consistent with all known

4. Both of these reasons may stem from the fact that already at conception, the female has made a greater parental investment than the male [Trivers 1972].

5. Rape in marriage is only recently recognized, and in some U. S. states is treated more leniently than rape outside marriage. Moreover, consistent with the view that women sell sex to men, rape may be considered theft, and rape of a woman a more serious offense than rape of a man [Posner 1992].

forms of marriage, including polyandry, polygyny, time-limited marriages [Posner 1992], marriages that continue after the death of the husband [Mair 1953], and many aspects of prostitution [Edlund and Korn 2002]. It is also consistent with the observation that marriage can be a source of income for women (e.g., Ellis [1936] and Goldin [1997]) and women only;[6] the empirical rejection of the unitary household model [Udry 1996; Lundberg, Pollak, and Wales 1997]; and many aspects of nonmarriage.

II.B. The Rise of Nonmarriage

The last three decades have witnessed a rapid decline in marriage, driven by delayed age of first marriage, increased out-of-wedlock childbearing, and divorce. Marriage has always been a more tenuous affair among the poor (e.g., Myrdal [1944], Göransson [1993], Smith [1996], and Edin and Lein [1997]), and the recent decline started earlier, and has been more dramatic, among low income groups. For instance, between 1972 and 1987 the marriage rate fell by 58 percent, 42 percent, and 24 percent for men with less than high school education, high school education, and some college education, respectively [Qian and Preston 1993]. We outline possible explanations for this development, and their implications for male-female inequality.

Contraceptives. If marriage is a contract in which women provide sex, then a possible reason for the fall in marriage may be lower male willingness to pay for this. The oral contraceptive is a female-controlled, low cost contraceptive that was approved by the Food and Drug Administration in 1960. It is a prescription drug that initially was only available to married women, but became available to unmarried women in the late 1960s [Goldin and Katz 2002]. Abortion is another female-controlled contraceptive. Abortion was legalized in 1970 in five U. S. states including California and New York, and nationally in 1973 with *Roe v Wade*. While abortion was medically feasible long before that, legalization lowered its cost.

Female-controlled contraceptives lowered women's marginal cost of supplying sex. One consequence may have been a reduc-

6. The Napoleonic Code states that "The husband owes protection to his wife, the wife obedience to her husband. The wife is obliged to live with her husband, and to follow him to every place where he may judge it convenient to reside: the husband is obliged to receive her, and to furnish her with every thing necessary for the wants of life, according to his means and station." Book 1, title V, chapter VI.

tion in the transfers women receive in marriage, since male willingness to pay for marriage partially derives from sexual access. Moreover, those interested in sex, but not children, no longer needed to marry [Akerlof, Yellen, and Katz 1996]. Hence, these contraceptives are likely to have reduced male to female income transfers, directly through lower marriage rates, for instance by raising the age at marriage [Goldin and Katz 2002], and higher divorce rates, and indirectly in marriage through an improved male bargaining position.

A potentially linked development was the passage of unilateral divorce laws in the 1970s, often considered as a proximate cause of increased nonmarriage [Friedberg 1998]. While the reasons for the timing of the divorce law reforms are not well established, these reforms were preceded by a buildup in popular demand for mutual consent divorce, which may have made their passage, if not inevitable, the next logical step [Phillips 1988; Glendon 1996]. One should note that divorce alone does not predict lower transfers to women since if coupled with remarriage it allows for serial polygyny and thus effectively raises demand for wives (cf. Becker [1991]). This points to the role of contraceptives in lowering demand for wives and divorce as a conduit for the subsequent cheapening of marriage.

Female labor force participation. The last three decades have seen a sharp rise in female labor force participation [Goldin 1990; Costa 2000]. If marriage is based on comparative advantages, as proposed by Becker [1973], then the narrowing of the gender wage gap seemingly suggests an explanation for the fall in marriage: lower gains from trade. However, given the rise in the number of high wage women, and the worsening labor market for low skilled men, it is unclear whether gains from trade have actually diminished.

Alternatively, if a man's role in marriage is to be the provider, then women's greater earnings ability may imply a decline in marriage (e.g., Edlund [1998]). However, this cannot be the only reason nonmarriage rose. If so, we would not expect nonmarriage to be associated with a feminization of poverty [Fuchs 1989; Smith and Ward 1989].

Welfare. Another explanation is that policies which target poor single-parent families, Aid to Families with Dependent Children (AFDC) in particular, have encouraged nonmarital fertility (e.g., Murray [1984]; for recent contributions see Rosenzweig

[1999] and Nechyba [2001]). AFDC afforded low income women the possibility of having children independently of a male provider (marriage). However, its level was too low to affect marital decisions of individuals other than the very poor.[7] The growing prevalence of nonmarriage increasingly involves groups not directly affected by welfare policies.

Marriage squeeze. Husbands tend to be older than their wives. This can give rise to a marriage squeeze if cohorts are of different sizes. Grossbard-Shechtman [1993] proposed that the baby boom that followed World War II created a marriage squeeze for women in the mid-1960s to early 1970s and men in the early 1980s, and that this prompted the observed changes in marriage patterns. According to this theory, the marriage market for females should have improved in the early 1980s. However, marriage has declined steadily since the mid-1960s. Moreover, it is unclear whether the magnitude of the effect was sufficient to cause a substantial reduction in male transfers to women. Other than a marriage squeeze, variations in cohort sizes can be absorbed through an adjustment of the spousal age gap. Finally, sex ratios have varied before, without the posited effect.[8]

II.C. Example

This subsection provides a simple example to illustrate how increasing nonmarriage generates a gender gap in political preferences and affects the aggregate demand for redistribution.

Consider a large population of an equal number of men and women. Let i be a continuous within gender income rank index, $i \in [0,1]$. Both men and women supply one unit of labor. Earnings y are distributed according to the density function $f(y)$ for women and $m(y)$ for men, with the corresponding cumulative distribution functions $F(y)$, $M(y)$. Moreover, $f(\cdot)$ and $m(\cdot)$ have compact supports, share a common lower support, $y \geq 0$, and $F(0) = M(0)$. $\bar{y}$ is the unconditional mean of y. We assume the male income distribution first order stochastically dominates the

7. For instance, in 1993, the maximum AFDC for a family of three was $367 a month in Illinois, the median state in this respect [Edin and Lein 1997, p. 35].

8. For instance, the United States suffered roughly 290,000 military casualties in World War II [Britannica Online], the vast majority of whom were young and male. This should have tilted the balance against marriage for women in the 1950s—a decennium in which the breadwinner-housewife model was at its apogee.

TABLE I
PER CAPITA INCOME RELATIVE TO MEAN INCOME, BY INCOME GROUP

Group	Income Rank	Nonmarried Woman	Nonmarried Man	Married
poor	0–i_M	$<$	$<$	$<$
lower middle income	i_M–i_P	$<$	$>$	$<$
upper middle income	i_P–i_F	$<$	$>$	$>$
rich	i_F–1	$>$	$>$	$>$

female, with the dominance strict (at least) at the mean income $\bar{y}$; i.e., $F(y) \geq M(y)$ and $F(\bar{y}) > M(\bar{y})$.

We assume that sorting is positive on income y, i.e., if woman i marries, she marries man i, and vice versa. Within marriage, men and women obtain a fixed share of household income, for simplicity, 50/50.[9]

We refer to the proportion of nonmarried individuals as the nonmarriage rate, ν. For simplicity, and in keeping with stylized facts, we assume that nonmarriage declines with income in the following way:[10]

$$\nu(i) = \begin{cases} 1 & \text{if } i \leq \nu, \\ 0 & \text{otherwise.} \end{cases} \tag{1}$$

The rank of the man and woman earning the mean income are $i_M \equiv M(\bar{y})$ and $i_F \equiv F(\bar{y})$, respectively, unless the highest ranked woman earns less than the mean income, in which case $i_F = 1$. Let i_P denote the rank of the individuals who (would) form the couple earning twice the mean income.[11]

Table I summarizes individual income, relative to mean income by income group and gender. The first column gives an income group label for each rank interval, and the second the intervals. The third and fourth columns give nonmarried female and male income, respectively, and the fifth gives married couple per capita income, relative to the mean income.

Two parties, left and right, compete in elections. These parties favor different redistributive policies. If elected, the left party

9. For simplicity, we assume a fixed income share. However, a sufficient assumption is that men transfer income to women in marriage.

10. Edlund and Pande [2001] also consider less restrictive assumptions on the nonmarriage pattern.

11. Formally, $i_P \equiv P(\bar{y})$, where $P^{-1}(\bar{y}) = (F^{-1}(\bar{y}) + M^{-1}(\bar{y}))/2$.

implements full redistribution, and the right party none. Taxation is on a household per capita basis, i.e., household income divided by the number of members (one or two).[12] In this environment sincere voting is optimal. Individual utility increases in income. Hence, those with income below mean income, $\bar{y}$ favor the left, and those with income above $\bar{y}$ the right.[13] From Table I it is clear that since per capita income determines party preference, only the political preferences of the middle-income groups change with marital status. Nonmarriage causes lower middle-income men to favor the right, and upper middle-income women to favor the left.[14]

Gender gap. Let l^f be the share of women and l^m the share of men who favor the left. We define the gender gap as $\gamma = l^f - l^m$. Clearly, $l^f = l^m$ corresponds to no gender gap, and $l^f > l^m$ to a leftist gender gap.

Table I affords the following observations.

1. If everyone is married, there is no gender gap, $\gamma = 0$.
2. Positive nonmarriage corresponds to a nonnegative gender gap, and the gap is strictly positive if there is nonmarriage among the middle-income groups.
3. The gender gap increases in nonmarriage if and only if nonmarriage increases among the middle-income groups.

Demand for redistribution. While the gender gap increases weakly in nonmarriage, support for the left may or may not. The reason is that for every woman who becomes poorer from nonmarriage, a man becomes richer. Support for the left is $l = (l^f + l^m)/2$. When nonmarriage is restricted to the poor, $\nu < i_M$, nonmarried men may be richer than if they were married, but remain poor enough to favor the left. Hence, an increase in nonmarriage among this group does not change the support for the left. By contrast, among the lower middle-income group, nonmarriage implies that the left loses the support of men. Hence, increasing nonmarriage in this group entails a decline in the overall support for the left. Among the upper middle-income group, the effect of nonmarriage on support for the left is the

12. Qualitatively similar results obtain as long as the higher income spouse (i.e., the man) pays higher taxes and receives fewer transfers when single than married, and the converse is true of the lower income spouse.

13. In this framework the median voter is decisive [Meltzer and Richard 1981]. For a critique of the median voter model, see Mulligan [2001].

14. Note that, if women are sufficiently poor, relative to men, then the rich group need not exist.

reverse—an unmarried, but not a married, woman votes for the left. Consequently, nonmarriage among this group favors the left. Last, among the rich, nonmarriage has no effect on political preferences, since albeit poorer if unmarried, women are rich enough to side with the right irrespective of marital status. It follows that l is a nonmonotone function of nonmarriage. Formally,

$$l'(\nu) = \begin{cases} 0 & \text{if } \nu < i_M, \\ < 0 & \text{if } i_M < \nu < i_P, \\ > 0 & \text{if } i_P < \nu < i_F, \\ 0 & \text{if } i > i_F. \end{cases} \tag{2}$$

Clearly, if in the absence of nonmarriage among the middle-income groups the left and the right enjoy equal support, then the men who switch to favoring the right when $\nu \in [i_M, i_P]$ will be pivotal for the right. Whether the women who switch left for $\nu \in [i_P, i_F]$ can tip the balance in favor of redistribution depends on whether their group size exceeds that of lower middle-income men. This is the case if nonmarriage is sufficiently high ($\nu > 2i_P - i_M$) and the upper middle-income group is larger than the lower middle-income group.[15] In any circumstance, nonmarriage alters the political preferences of lower middle income men and upper middle income women.

Another prediction offered by this example is that if nonmarriage first increases among the poor and subsequently spreads to higher income groups, then we would first observe that lower middle-income men switch party allegiance to the right, followed by upper middle-income women switching left. Figure III uses National Election Survey data to depict gender differences in Republican party identification by education level. We observe a clear sequencing: a pronounced jump in Republican support among men with no more than high-school education in the 1984 election (Reagan Democrats) followed by a fall in support among women with some college education starting in the 1992 election (Soccer Moms) (also see Stark [1996]). That support for the left increases in nonmarriage only when nonmarriage affects the upper middle-income group provides one explanation for the recent adoption of conservative social policies that purportedly encourage marriage by the right (it was not until the 1992 election

15. For instance, this is the case if $f(y)$ and $m(y)$ are symmetric, single peaked, and share a common lower support.

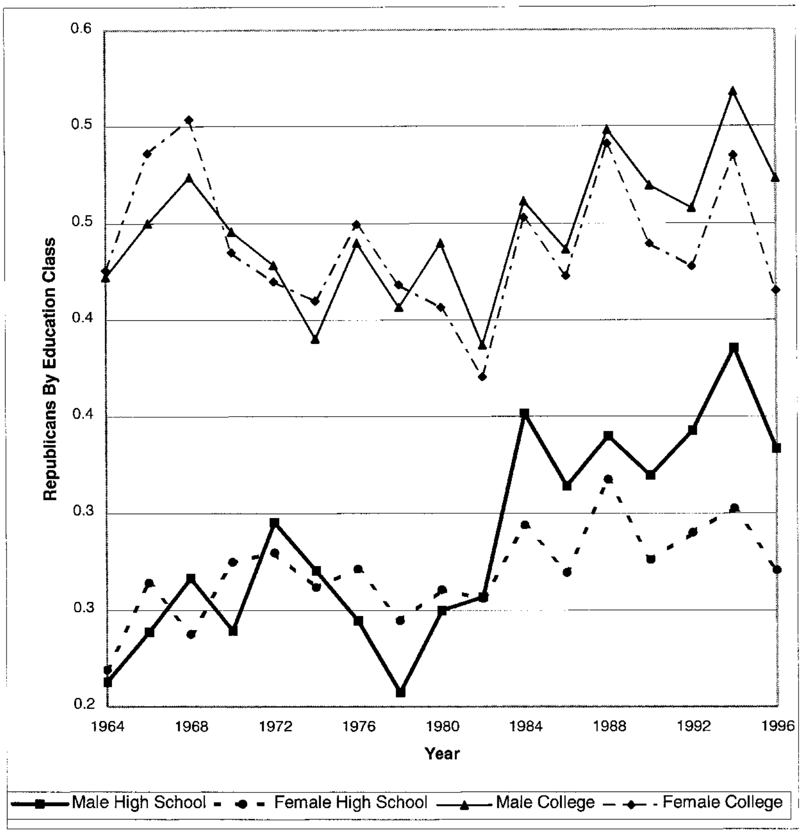

FIGURE III
Reagan Democrats and Soccer Moms

Notes. This figure graphs by education the proportion of male and female respondents who are Republicans. We use information on 18–64 year old respondents from the National Election Studies surveys. Male College is the proportion of men with at least some college education who are Republican, Female College is the proportion of women with at least some college education who are Republican, Male High School is the proportion of men with no college education who are Republican, Female High School is the proportion of women with no college education who are Republican. A respondent is a Republican if his/her own stated party identification is Strong-, Weak-, or Independent-leaning Republican. Appendix 1 provides a full description of the National Election Studies sample.

that family values became a prominent feature of Republican political campaigns [Whitehead 1993]).

Children. An important consequence of nonmarriage is the increasing prevalence of families in which children live with only one biological parent, usually the mother. To examine how our results change if voting decisions are based on preferences over public provision of goods and services benefiting children, rather than redistribution to adults, we consider a simple example. Assume that income and marriage patterns in the society are as in the above example. In addition, every woman has one child, and the elected party can redistribute resources via a fixed transfer to each child. The left, but not the right, favors such redistribution. We assume that this transfer is a public good for married parents (e.g., they share custody), while for nonmarried parents it is the mother (the custodian) who benefits from the transfer.

If everyone is married, then there is no gender gap, and aggregate support for the left will depend on the male and female income distributions. The main difference from our earlier example arises when there is nonmarriage among the poor. With child transfers such nonmarriage engenders a gender gap since, irrespective of own income, only married men benefit from redistribution. Hence the popular support for child transfers declines with nonmarriage for a larger interval. For any level of nonmarriage child targeted transfers attract lower popular support than general transfers.[16]

Therefore, our earlier results are qualitatively similar to a situation where women are more likely to be single parents than men and preferences over child-transfers also affect voting. In the presence of children the gender gap is driven, not by women's having greater child responsibilities than men have, but by the interaction of such gender differences with rising nonmarriage. This is similar to the mechanism identified by our earlier example where it was not women's earning less than men but their not being married that drove the gender gap.

III. Divorce Risk and the Gender Gap

This section presents evidence on how increases in aggregate divorce risk have impacted on the political gender gap. Our data are drawn from the biennial National Election Studies (NES), for individual level information, and the March Current Population

16. Obviously, childless women may align their preferences with unmarried men and thus attenuate the gender gap. Equally, if noncustodian fathers benefit from child-related transfers, then poorer nonmarried fathers may favor child-targeted redistributions.

Surveys, for state-level aggregates, and span the period 1964 to 1996.

III.A. Data and Descriptive Statistics

We restrict the sample to the period 1964 to 1996, and respondents aged 18–64. This leaves us with seventeen survey rounds and approximately 1400 respondents per survey. The average respondent was 39 years old, 54 percent were female, and 65 percent married. Between 1964 and 1996 the proportion of married respondents fell from 80 to 57 percent (Table II).

Roughly 90 percent of the respondents had at least completed grade school, and 80 percent were in the labor force at the survey date. The NES only identifies a respondent's annual family income percentile. We distinguish between three income groups: (i) 0–33 percentile (poor); (ii) 34–95 percentile (middle income); and (iii) 96–100 percentile (rich). Since, relative to the per capita income distribution, such a classification places unmarried respondents "too low" and married respondents "too high," our regressions allow income coefficients to vary by marital status.

To avoid sample selection issues related to actual voting, we measure a respondent's political preferences as his/her stated partisan identification. The survey question asks respondents to indicate party preference on a seven-point scale ranging from Strong Democrat to Strong Republican. We collapse responses to this question to a dummy measure **idemocrat** which equals 1 if respondent stated self to be a Strong-, Weak-, or Independent-leaning Democrat.[17] Fifty-four percent of female, and 48 percent of male, respondents identified themselves as **idemocrat.**

To ascertain that an individual's party and redistributive preferences are aligned, we use a direct measure of individual redistributive preferences as an alternative dependent variable. The dummy **govspend** equals 1 if the respondent states that the government should provide many services (and implicitly increase spending and taxes). This variable is only available since 1982.

To examine whether male to female differences on social issues, rather than income differences, lie behind the emergence of the political gender gap, we make use of attitudinal questions

17. We find qualitatively similar results using a stronger measure of political affiliation: **democrat,** a dummy variable that equals 1 only if the respondent stated self to be a Strong or Weak Democrat.

TABLE II
DESCRIPTIVE STATISTICS 1964–1996 NES, CPS

Variable		Percentage (Standard deviation)		
A. NES data		All	Men	Women
Demographics				
married		65.7	69.2	62.1
age [years]		39.2	39.4	39.1
		(12.6)	(12.5)	(12.7)
Black		11.3	9.3	12.9
cohort	–1910	3.2	3.2	3.19
	1911–1942	46.1	45.9	46.2
	1943–1958	35.5	35.7	35.3
	1959–	15.0	14.9	15.1
Economic characteristics				
education	Less than 9 years	9.1	10.0	8.3
	9–12 years	50.0	44.3	54.8
	some college	21.7	22.5	21.0
	college +	19.0	23.0	15.7
In labor force	labor	81.5	97.5	68.1
family income				
percentile	0–33	26.2	20.4	31.1
	34–95	67.8	72.9	63.6
	96–100	5.83	6.6	5.1
Preferences				
	idemocrat	51.8	48.4	54.6
	govspend	67.1	60.1	73.6
	pro-choice	54.9	55.6	54.3
	equal roles	65.6	66.1	65.2
religion	Protestant	63.6	60.4	66.3
	Catholic	24.2	24.4	24.0
	Jewish	2.14	2.33	1.97
	church	47.8	40.8	53.6
salient issue	social	12.3	11.3	13.1
	economics	33.3	37.2	29.9
	welfare	22.1	18.6	25.1
number of observations		24,140	11,007	13,133
B. CPS-state				
pdivorced		6.60		
		(2.81)		
plabor		59.60		
		(9.36)		
number of observations		336		

All NES descriptives refer to the sample of 18–64 year old respondents in the survey years 1964–1996 for whom demographic and economic characteristics are available (N = 24,140); with the following exceptions: labor (N = 23,106) spans 1968–1996; equal roles (N = 15,812) and pro-choice (N = 17,470) 1972–1996; **govspend** (N = 9,947) 1982–1996; and church (N = 23,986) 1970–1996; social, economics, and welfare are available for the entire period, but missing values reduce the sample size to N = 19,903. See Appendix 1 for precise variable definitions.

on women's issues (abortion and equal roles), the relative political salience of social, welfare, and economic issues for the respondent, and religiosity. There were no significant gender differences on women's issues and the salience of social issues. However, more women emphasized welfare issues. Religiosity exhibited significant gender differences; 53 percent of female, but only 40 percent of male, respondents attended church regularly.

We proxy for the divorce risk facing an individual by two different aggregate measures. Our first measure, **pdivorced,** is the divorce incidence in a state, as captured by the proportion of adult population that is currently divorced. This variable is constructed from March Current Population Survey data. To ensure representativeness, our unit of aggregation is the CPS-state which often includes multiple U. S. states. Overall, there are 21 CPS-states (for details, see Appendix 1).

Our second measure, **unilat,** is the passage of unilateral divorce laws by U. S. states. This captures changes in divorce risk arising from alterations in the legal framework governing marriage dissolution. Following Gruber [2000], we define unilateral divorce to be available when divorce can be filed on a no-fault ground, and there is no separation requirement. Thus, the **unilat** variable equals 0 until the year these laws were introduced, and then 1. Appendix 2 provides information by state on the year unilateral divorce laws were passed and the party affiliation of the then state's governor (source: Gruber [2000]). Over our sample period, Democrat and Republican governors were equally likely to pass such laws, suggesting bipartisan support.

III.B. Basic Results

In order to provide a baseline against which we can compare subsequent findings, we examine how the political gender gap varied across years. We estimate an OLS linear probability regression of the form,

$$d_{ikt} = c_k + \tau_t + \phi_1 f_{ikt} + \phi_2 (f_{ikt} \times \tau_t) + \epsilon_{ikt}, \tag{3}$$

where d_{ikt} is the **idemocrat** dummy for individual i, c_k are CPS-state dummies, τ_t are year dummies, f_{ikt} is a female dummy (female in text). The coefficient ϕ_2 provides a measure of the trend in, and $\phi_1 + \phi_2$ the level of, the gender gap unexplained by our controls.

Table III, column (1), reports the results. While the regression includes the full set of female × year interaction terms, to

TABLE III
INDIVIDUAL DETERMINANTS OF DEMOCRATIC PARTY IDENTIFICATION
DEPENDENT VARIABLE: **IDEMOCRAT**

	(1)	(2)	(3)	(4)	(5)
female	−0.005	−0.017	−0.024	−0.034	−0.084***
	(0.021)	(0.020)	(0.021)	(0.021)	(0.033)
female × 1968	0.058	0.044	0.043	0.042	0.036
	(0.036)	(0.032)	(0.036)	(0.035)	(0.036)
female × 1972	0.075***	0.072***	0.073***	0.075***	0.058***
	(0.020)	(0.018)	(0.020)	(0.020)	(0.022)
female × 1976	0.039	0.046*	0.054**	0.054**	0.011
	(0.029)	(0.026)	(0.026)	(0.027)	(0.033)
female × 1980	0.100***	0.107***	0.105***	0.107***	0.038
	(0.034)	(0.031)	(0.034)	(0.034)	(0.052)
female × 1984	0.080***	0.079***	0.079**	0.080***	−0.007
	(0.030)	(0.029)	(0.031)	(0.030)	(0.052)
female × 1988	0.070**	0.077***	0.087***	0.088***	−0.009
	(0.029)	(0.028)	(0.025)	(0.024)	(0.050)
female × 1992	0.107***	0.115***	0.115***	0.117***	−0.000
	(0.029)	(0.028)	(0.029)	(0.028)	(0.073)
female × 1996	0.139***	0.150***	0.148***	0.151***	0.022
	(0.032)	(0.035)	(0.033)	(0.031)	(0.074)
Married	—	−0.051***	−0.066***	−0.067***	−0.067***
		(0.008)	(0.023)	(0.024)	(0.024)
Black	—	0.357***	0.338***	0.340***	0.339***
		(0.028)	(0.028)	(0.028)	(0.028)
age	—	0.006***	0.009***	0.009***	0.009***
		(0.002)	(0.002)	(0.002)	(0.002)
age^2 ($\times 10^{-3}$)	—	−0.051**	−0.082***	−0.082***	−0.082***
		(0.024)	(0.023)	(0.023)	(0.023)
cohort:					
1911–1942	—	0.039*	0.038*	0.038*	0.038*
		(0.023)	(0.022)	(0.022)	(0.022)
1942–1952	—	0.052*	0.049*	0.050*	0.050*
		(0.027)	(0.027)	(0.027)	(0.027)
1959–	—	0.024	0.016	0.017	0.017
		(0.030)	(0.031)	(0.031)	(0.031)
religion:					
Catholic	—	0.077***	0.075***	0.075***	0.076***
		(0.015)	(0.016)	(0.016)	(0.016)
Protestant	—	−0.098***	−0.099***	−0.099***	−0.099***
		(0.015)	(0.016)	(0.017)	(0.017)
Jewish	—	0.238***	0.291***	0.291***	0.293***
		(0.039)	(0.037)	(0.038)	(0.038)
education:					
<9 years	—	—	0.067***	0.066***	0.066***
			(0.021)	(0.021)	(0.021)
9–12 years	—	—	0.049***	0.049***	0.050***
			(0.013)	(0.013)	(0.013)
some college	—	—	0.010	0.009	0.010
			(0.011)	(0.011)	(0.011)

TABLE III
(CONTINUED)

		(1)	(2)	(3)	(4)	(5)
family income percentile:						
	0–33	—	—	0.140***	0.140***	0.142***
				(0.028)	(0.028)	(0.027)
	34–95	—	—	0.153***	0.152***	0.153***
				(0.028)	(0.028)	(0.028)
	married × 0–33	—	—	0.051**	0.052**	0.051**
				(0.026)	(0.026)	(0.026)
	married × 34–95	—	—	0.006	0.007	0.007
				(0.024)	(0.024)	(0.024)
nonmarriage:						
	pdivorced	—	—	—	—	−2.116**
						(0.937)
	female × **pdivorced**	—	—	—	—	1.802**
						(0.921)
Constant		0.831***	0.250***	0.059	0.022	0.150*
		(0.008)	(0.052)	(0.056)	(0.055)	(0.091)
other dummies:						
	year	yes	yes	yes	yes	yes
	CPS-state	yes	yes	yes	yes	yes
	female × CPS-state	no	no	no	yes	yes
Adj. R^2		0.020	0.091	0.097	0.098	0.098
N		26,215	25,848	24,140	24,140	24,140

OLS regression results, with robust standard errors adjusted for CPS-state-clustering, are reported in parentheses. The excluded categories are female × year—1964; education—college educated; cohort group—pre-1911 cohort; income—96–100 percentile. Coefficients for female × year interactions are only reported for the years of presidential elections, however, all regressions include the full set of interaction terms. *indicates significance at 10 percent, ** at 5 percent, and *** at 1 percent.

avoid clutter, Table III reports the coefficients only for Presidential election years. Relative to 1964 (the omitted year), apart from 1972, no significant gender gap exists until 1980. However, with the exception of 1990, all years since 1980 show a significant Democratic gender gap. Comparing point estimates, the gender gap rose sharply in the early 1980s, then stabilized and fell, before rising again in the 1990s. To use popular parlance, the first phase corresponded to the Reagan Democrat years and the last to the Soccer Mom years.

To investigate the relative roles of individual characteristics and divorce risk in explaining this trend, we reestimate the above regression and sequentially include these two sets of covariates. Our final regression is of the form,

$$(4) \quad d_{ikt} = c_k + \tau_t + \phi_1 f_{ikt} + \phi_2(f_{ikt} \times \tau_t) + \phi_3 X_{ikt} + \phi_4(f_{ikt} \times c_k) + \phi_5 \nu_{kt} + \phi_6(f_{ikt} \times \nu_{kt}) + \epsilon_{ikt},$$

where X_{ikt} is the vector of individual demographic and economic controls. ν_{kt} is our primary measure of divorce risk, **pdivorced,** that varies by year and CPS-state. In all regressions we cluster standard errors by CPS-state. This is to correct for two potential problems. First, grouped error terms which arise from the fact that our unit of observation, the individual, vary at a more disaggregate level than **pdivorced.** Second, **pdivorced** is serially correlated. Bertrand, Duflo, and Mullainathan [2001] show that such clustering can help reduce the bias in standard errors that this causes.

Column (2) of Table III reports results for the regression which includes individual demographic controls. Consistent with existing research, we find that Black, Catholic, Jewish, and older respondents are significantly more likely to be **idemocrat.** Column (3) includes information on economic attributes. Democratic support falls monotonically with education. Poor and middle-income individuals are more favorable toward the Democratic party than the rich. However, the relationship is nonmonotone, with the poor less likely to be Democratic than middle-income individuals. A potential explanation is that the poor include individuals with high lifetime income, for instance, college students. Comparing across columns (1)–(3), we see that the inclusion of individual controls improves our regression fit, but does not explain the trend in the gender gap.

As a precursor to analyzing the role of **pdivorced** in explaining this gender gap, column (4) reports regressions that include a set of interaction terms female × CPS-state. The latter accounts for omitted CPS-state variables which affect men and women differentially. These interaction terms are jointly significant in explaining Democratic party affiliation, but not in explaining the trend in the political gender gap.

Finally, column (5) includes our measure of divorce risk—**pdivorced** and female × **pdivorced**—as explanatory variables. The coefficients on the controls for individual characteristics remain unaffected. However, both the economic magnitude and the statistical significance of the female × year set of interaction terms are dramatically lowered. No significant unexplained trend in the gender gap remains after 1980. Figure IV illustrates how the inclusion of **pdivorced** improves our ability to predict the

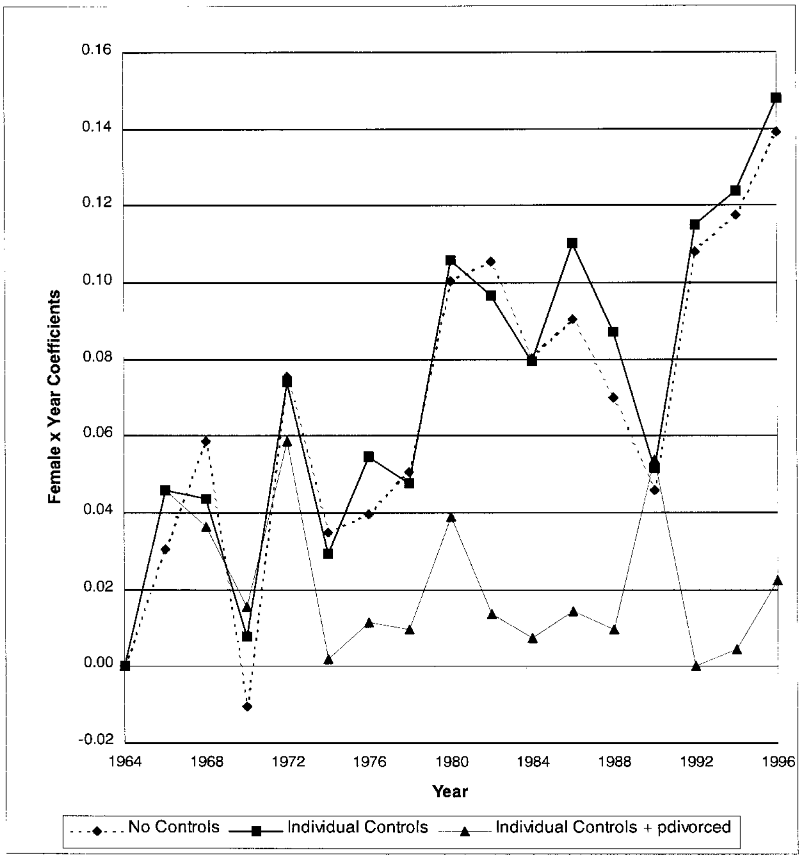

FIGURE IV
Time Trend in the Gender Gap

Note. This figure graphs the coefficients for the set of female × year interaction terms which are reported in Table III. No Controls refers to column (1), Individual Controls to column (3), and Individual Controls + pdivorced to column (5).

trend in the gender gap, it graphs the sets of coefficients on the female × year terms reported in Table III, columns (1), (3), and (5), respectively.

Between 1964 and 1996 the gender gap increased by 13.4 percentage points, and **pdivorced** from 3 to 10 percent. A back-of-the-envelope calculation using the point estimate for female × **pdivorced** in column (5) suggests that the rise in **pdivorced** can explain a gender gap of 12.6 percentage points, or 94 percent of the observed gap.

TABLE IV
NONMARRIAGE AND DEMOCRATIC PARTY IDENTIFICATION
DEPENDENT VARIABLE: **IDEMOCRAT**

	Family income percentile							
	All incomes		0–33		34–95		96–100	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
female	−0.084***	−0.080*	0.049	0.116**	−0.149**	−0.195**	0.179	0.002
	(0.033)	(0.048)	(0.045)	(0.041)	(0.067)	(0.084)	(0.213)	(0.273)
pdivorced	−2.116**	−1.816*	0.165	0.581	−2.680***	−2.681***	−2.137	−2.877
	(0.937)	(0.999)	(1.514)	(1.586)	(0.986)	(0.993)	(2.823)	(3.061)
female × **pdivorced**	1.802**	1.837*	0.516	−1.253	2.656***	3.611***	0.349	1.904
	(0.921)	(1.036)	(1.715)	(1.741)	(0.921)	(1.120)	(3.672)	(4.355)
female × married	—	0.002	—	−0.139**	—	0.047	—	0.212
		(0.043)		(0.054)		(0.061)		(0.234)
pdivorced × married	—	0.470	—	−1.301	—	0.001	—	1.019
		(0.485)		(0.809)		(0.577)		(1.702)
female × **pdivorced** × married	—	0.066	—	2.190	—	−1.175	—	−2.065
		(0.557)		(0.780)		(0.778)		(3.224)
Adj. R^2	0.097	0.097	0.080	0.081	0.095	0.096	0.139	0.138
N	24,140	24,140	6,343	6,343	16,388	16,388	1,409	1,409

OLS regression results, with robust standard errors adjusted for CPS-state clustering, are reported in parentheses. Controls are included for year dummies, CPS-state dummies, female × CPS-state interactions, and all the other covariates in column (5) of Table III except that the income covariates are not included in specifications that divide the sample by income groups. *indicates significance at 10 percent, ** at 5 percent, and *** at 1 percent.

Table IV investigates how the impact of **pdivorced** on political preferences varies with income group and marital status. The endogenous nature of individual income and marital status raises the concern that selectivity bias may underlie apparent income group or marital status effects. We, therefore, first report results for the entire sample, and for each income group provide two specifications: one that does not distinguish between individuals by marital status, and one that does. All regressions include the individual controls in column (5) of Table III, except the income covariates in specifications that divide the sample by income groups (columns (3)–(8)).

Columns (1) and (2) of Table IV report results for the entire sample. Comparing across the two, we find that **pdivorced** does not affect the political preferences of married and unmarried respondents differentially. As this is the case for all specifications we consider, in subsequent tables we do not report specifications

TABLE V
NONMARRIAGE AND PREFERENCE FOR REDISTRIBUTION
DEPENDENT VARIABLE: **GOVSPEND**

	Family income percentile			
	All incomes (1)	0–33 (2)	34–95 (3)	96–100 (4)
female	−0.280**	0.132	−0.055	−0.763
	(0.123)	(0.176)	(0.142)	(0.512)
pdivorced	−1.917**	−2.049	−1.923*	−0.222
	(0.912)	(3.084)	(1.115)	(5.439)
female × **pdivorced**	4.714***	3.701	5.059***	3.385
	(1.469)	(3.252)	(1.860)	(5.805)
Adj. R^2	0.089	0.038	0.084	0.101
N	9,969	2,505	6,880	584

OLS regression results, with robust standard errors adjusted for CPS-state clustering, are reported in parentheses. Controls are included for year dummies, CPS-state dummies, female × CPS-state interactions, and all the other covariates in column (5) of Table III except that the income covariates are not included in specifications that divide the sample by income groups. * indicates significance at 10 percent, ** at 5 percent, and *** at 1 percent.

that control for marital status. Columns (3)–(8) report results by income group. An increase in **pdivorced** is associated with a statistically significant Democratic gender gap only for the middle-income group (percentiles 34–95). Moreover, the magnitude of the effect is largest for this group. Among the middle-income group, increased divorce risk turns men away from the left. A one percentage point increase in divorce risk lowers the likelihood that a male respondent is an **idemocrat** by 2.7 percentage points, but leaves that of women unchanged (column (5)). Within this group we find that, relative to nonmarried women, married women are significantly less likely to be **idemocrat.** However, the impact of divorce risk on women's political preferences does not differ by marital status.

III.C. Robustness

How well does an individual's party affiliation, as captured by **idemocrat,** correlate with his/her redistributive preferences? To examine this, Table V reports results for regressions that use a measure of individual redistributive preferences, **govspend,** as the dependent variable. Column (1) reports results for the entire sample. Increases in **pdivorced** have a significant and differential effect on male and female redistributive preferences. Col-

TABLE VI
DIVORCE LAW LIBERALIZATION
DEPENDENT VARIABLE: **IDEMOCRAT**

	Family income percentile			
	All incomes (1)	0–33 (2)	34–95 (3)	96–100 (4)
female	0.718***	−0.068	0.207***	0.02
	(0.026)	(0.056)	(0.040)	(0.135)
unilat	−0.065***	−0.051	−0.064**	−0.085
	(0.022)	(0.044)	(0.023)	(0.067)
female × **unilat**	0.069***	0.091**	0.067**	0.087
	(0.025)	(0.042)	(0.033)	(0.081)
Adj. R^2	0.102	0.089	0.100	0.170
N	24,140	6,343	16,388	1,409

OLS regression results, with robust standard errors adjusted for clustering at the state level, are reported in parentheses. Controls are included for year dummies, state dummies, female × state interactions, and all the other covariates in column (5) of Table III except that the income covariates are not included in specifications that divide the sample by income groups. *indicates significance at 10 percent, ** at 5 percent, and *** at 1 percent. There were no respondents from the following states: Alaska, Hawaii, Idaho, Montana, North Dakota, Rhode Island, and Vermont.

umns (2)–(4) estimate this regression by income group. As with party affiliation, the differential effect of divorce risk on male and female political preferences is limited to the middle-income group.

The results in Tables IV and V paint a consistent picture of how increased divorce risk affects the political preferences of the middle-income group. However, there are differences in how divorce risk affects men's and women's party affiliations and redistributive preferences. First, at 32 percentage points, the redistributive preference gender gap is more than double the Democratic gender gap. Second, increased divorce risk alters men's party affiliation but women's desire for redistribution. Taken together, these findings are suggestive of a shift in party platforms.

The other measure of divorce risk we explore is the passage of unilateral divorce laws, **unilat.** Table VI presents the results for this measure. Column (1) tells us that the liberalization of divorce laws was associated with the emergence of a political gender gap. Moreover, this effect varied by income group. The passage of unilateral divorce laws left the political preferences of the rich unaffected (column (4)), but had a gender differential effect on the political preferences of the middle-income group (column (3)). For this group easier divorce made men, but not women, abandon the Democratic party. These results are consistent with our findings for **pdivorced.** The only difference is that, unlike **pdivorced,** the passage of unilateral divorce laws also affected the political preferences of the poor. Easier divorce made women more likely to identify with the Democratic party. This last effect is sensitive to the introduction of controls for marital status—introduction of marital status controls suggests that this effect is primarily driven by married women.

In Edlund and Pande [2001] we provide additional robustness checks. Arguably, the impact of **pdivorced** on expected income, and therefore political preferences, should be more muted for the young or old. Moreover, if aggregate divorce risk is primarily driven by divorce among the young to middle-aged, we would expect movements in aggregates to concern older individuals less. In line with these arguments we found that increases in **pdivorced** were associated with a political gender gap only among the 25–40 age group. We also considered alternative specifications. First, to check that racial differences in marriage patterns and political behavior do not drive our results, we reestimated our regressions for the sample of White respondents. Second, to ensure that **pdivorced** does not simply pick up state-specific trends in political preferences, we estimated the regressions including a CPS-state-specific linear trend. Third, to check that the results are not sensitive to the choice of a linear specification, we also used a Probit specification. These modifications did not qualitatively alter our main results.

III.D. Competing Hypotheses

This subsection provides evidence on three alternative explanations for the emergence of the political gender gap: female labor force participation, women's issues, and religious and social values.

Female labor force participation. The increase in female labor force participation over the last three decades has been accompanied by changes in female educational profile, own-earned income, and social and political attitudes. An alternative hypothesis is that the political gender gap was engendered by the social and economic changes wrought by women's mobilization into the labor force.

TABLE VII
LABOR FORCE PARTICIPATION AND DEMOCRATIC PARTY IDENTIFICATION
DEPENDENT VARIABLE: **IDEMOCRAT**

	Family income percentile					
	0–33		34–95		96–100	
	(1)	(2)	(3)	(4)	(5)	(6)
female	−0.089	−0.311	−0.086	−0.119	0.150	1.460**
	(0.107)	(0.333)	(0.083)	(0.133)	(0.191)	(0.715)
pdivorced	0.287	0.781	−2.609**	−2.668***	−2.193	−5.496**
	(1.483)	(1.439)	(1.048)	(1.006)	(2.886)	(2.565)
female × **pdivorced**	−0.507	−0.720	2.410**	2.312**	2.058	5.647**
	(1.776)	(1.754)	(0.986)	(1.075)	(3.788)	(3.004)
labor	−0.001	−0.000	−0.050	−0.049	−0.048	−0.058
	(0.035)	(0.035)	(0.052)	(0.052)	(0.144)	(0.139)
female × labor	0.000	−0.000	0.111**	0.110**	0.096	0.105
	(0.035)	(0.035)	(0.053)	(0.053)	(0.163)	(0.159)
plabor	—	−0.737	—	0.054	—	2.986**
		(0.677)		(0.294)		(0.957)
female × **plabor**	—	0.341	—	0.089	—	−3.206**
		(0.526)		(0.346)		(1.642)
Adj. R^2	0.081	0.081	0.097	0.097	0.141	0.146
N	6,124	6,124	15,643	15,643	1,339	1,339

OLS regression results, with robust standard errors adjusted for CPS-state clustering, are reported in parentheses. Controls are included for year dummies, CPS-state dummies, female × CPS-state interactions, and all the other covariates in column (5) of Table III except that the income covariates are not included in specifications that divide the sample by income groups. * indicates significance at 10 percent, ** at 5 percent, and *** at 1 percent.

We test this hypothesis in two ways. First, we examine whether being in the labor force affects male and female political preferences differentially (Table VII). The relationship between **pdivorced** and the political gender gap is robust to including this information. Relative to a man, labor force participation only affects the political preferences of middle-income women. Being in the labor force makes a middle-income woman (relative to a man) 11 percentage points more likely to be an **idemocrat** (column (3)). The response to own labor force participation among middle-income women is consistent with an interpretation of women's working (for this group) being associated with a more precarious economic situation.

Second, we examine whether changes in the proportion of women in the labor force in a CPS-state (denoted as **plabor**) affect political preferences.[18] Between 1964 and 1996 **plabor** rose from 44 to 71 percent. It is possible that increases in this aggregate were correlated with changing attitudes which, in turn, altered men's and women's political preferences. Alternatively, if increases in **plabor** are associated with increased nonmarriage, then the effects we attribute to **pdivorced** may simply proxy for labor market effects. Table VII reports the results for regressions that include **plabor.** Among the poor and middle-income group, we find no effect (columns (2) and (4)). Instead, among the rich, increases in **plabor** increase male sympathy for the Democratic party, while women are largely unmoved (column (6)). This suggests that among the rich, increases in aggregate female labor force participation muted rather than contributed to the political gender gap. Throughout, our main results for divorce risk remain robust to the inclusion of labor force participation variables.

Social and religious values. In Table VIII we provide evidence on how changing social and religious values have impacted on male and female political preferences. We first consider changing attitudes on women's issues. The past three decades have seen women's issues become politically divisive. In particular, the Democratic party has come to champion abortion rights (vested with the woman) and the Republican party the pro-life position. Republicans have also become associated with so-called family values that prescribe a traditional homemaking role for women. It is commonly believed that these policy differences have divided the electorate along gender lines. Moreover, some believe that the onset of feminism and increasing male-to-female differences on women's issues lie behind the rise in nonmarriage. If correct, we may have misattributed the impact of women's issues on the political gender gap to increased divorce risk.

Rows (1) and (2) of Table VIII explore this possibility by examining how respondent's attitudes on these issues condition his/her political preferences. In row (1) we include a dummy for whether the respondent supports a woman's right to choose abortion (pro-choice). Respondents who are pro-choice are 6 percentage points more likely to identify themselves as **idemocrat** (a slightly higher percentage of men than women are pro-choice). Moreover, relative to men, women who are pro-choice are 3 percentage points more likely to favor the left. The latter effect,

18. **plabor** is constructed from March Current Population Surveys.

TABLE VIII

SOCIAL AND RELIGIOUS VALUES, INCOME PERCENTILES 34–95 ONLY DEPENDENT VARIABLE: **IDEMOCRAT**

				Abortion and equal roles			
	female	**pdivorced**	female × **pdivorced**	pro-choice	female × pro-choice	equal roles	female × equal roles
(1)	−0.070	−3.125***	2.613**	0.061***	0.031*		
	(0.097)	(1.145)	(1.203)	(0.015)	(0.018)		
(2)	−0.134	−3.292***	2.849**	0.058***	0.018	0.032*	0.041*
	(0.117)	(1.210)	(1.419)	(0.015)	(0.019)	(0.017)	(0.024)

				Political salience					
	female	**pdivorced**	female × **pdivorced**	social	female × social	economics	female × economics	welfare	female × welfare
(3)	−0.156	−2.923***	2.991***	−0.070***	−0.015	−0.011	−0.012	0.084**	−0.052**
	(0.058)	(1.024)	(1.118)	(0.019)	(0.027)	(0.013)	(0.015)	(0.015)	(0.022)

				Religion and church attendance							
	female	**pdivorced**	female × **pdivorced**	Catholic	female × Catholic	Protestant	female × Protestant	Jewish	female × Jewish	church	female × church
(4)	−0.094	−2.603***	2.692***	0.088***	−0.008	−0.071***	−0.059	0.276***	0.052	−0.015	−0.042**
	(0.080)	(0.937)	(0.911)	(0.028)	(0.040)	(0.027)	(0.042)	(0.041)	(0.047)	(0.016)	(0.020)

OLS regression results, with robust standard errors adjusted for CPS-state clustering, are reported in parentheses. Controls are included for year dummies, CPS-state dummies, female × CPS-state interactions, and all the other covariates in column (5) of Table III except that the income covariates are not included. Adjusted R^2 and number of observations for the regressions are (1) 0.103 and 11,785; (2) 0.111 and 9,795; (3) 0.106 and 13,637; and (4) 0.098 and 16,284, respectively. * indicates significance at 10 percent, ** at 5 percent, and *** at 1 percent.

however, is statistically insignificant once we control for marital status.

To examine how feminist sympathies affect political preferences, row (2) includes information on whether the respondent believes men and women should have an equal role in society. Respondents who believe in equal roles are 3 percentage points more likely to be **idemocrat.** Moreover, relative to men, women who believe in equal roles are 4 percentage points more likely to favor the left.

The estimated relationship between divorce risk and male-to-female political preferences remains robust to the inclusion of these attitude variables. While clearly shaping political preferences, the relatively weak gender differential effects associated with women's issues suggests that the parties' diverging stance on these issues has not been an important determinant of the gender gap.

The second possibility we consider is whether gender differences in the political salience attached to social and economic issues drove the gender gap. We construct three dummies: social which equals 1 if the respondent believed that the most important problem facing the nation related to public order issues including crime, civil rights and social, religious or moral decay; economics which equals 1 if the respondent believed that the most important problem facing the nation related to economic, business, and consumer issues; and welfare which equals 1 if the respondent believed that the most important problem related to welfare issues such as child care, education, the elderly, and health care.

Slightly more women than men consider social issues to be the most important issue. While respondents who believe social issues to be the most important are 7 percentage points less likely to be **idemocrat,** this effect does not vary by gender. More men than women consider economics to be the most salient issue. However, this view does not significantly impact on party affiliation for either sex. By contrast, those who consider welfare to be the most important issue are 8 percentage points more likely to favor the left, and within this group it is men who are the most left-leaning (row (3)). A possible explanation is sample selection: markedly more women than men held this view.

Finally, we consider the role of religion. The last three decades have seen a marked decline in both religiosity, and moral values. At the same time, politically active religious movements such as the Moral Majority and the Christian Coalition emerged, movements that are mainly associated with the Republican party. While women are traditionally portrayed as the bedrock of religiosity and public morality, one may wonder whether the decline in religiosity affected women to a greater extent and thereby led to a political gender gap.

Row (4) explores this possibility. Our main result remains robust: higher divorce risk turns middle-income men, but not women, away from the Democratic party. While religious denomination is a significant predictor of political behavior, there are no significant gender differences in the extent to which religious belief conditions political behavior. In contrast, the intensity of religious belief, as captured by frequency of church attendance, affects male and female political behavior differentially. The dummy variable church equals 1 if the respondent attended church at least twice a month. Controlling for religious denomination, we find that church attendance makes women, relative to men, four percentage points less likely to be an **idemocrat.** While suggesting that the decline in church attendance has made women less right-leaning, this finding raises the question of why less religious women favor the left.

In Edlund and Pande [2001] we also examined whether the political gender and racial gaps were linked. For if ideologically feminism shared common ground with the civil rights movement, we might expect the gender gap and the Black-White gap to exhibit similar trends. However, this was not the case. Black support for the Democratic party increased dramatically in the first half of the 1960s, peaked at over 90 percent in 1968, and has since been falling off. Moreover, Black men rather than Black women led this early shift to the Democratic party.

IV. Marital Status and Political Preferences: Evidence from Longitudinal Data

The previous section identified a strong positive correlation between aggregate divorce risk and the political gender gap. This section complements the analysis with longitudinal data that allow us to examine how actual changes in an individual's marital status impact on his/her political preferences. Our analysis exploits the observation that changes in own marital status are not fully anticipated. Hence the realization of such a change is a valid instrument for changing individual expectation regarding marital status.

We use the three publicly available waves of the Youth Parent Socialization Survey (YPSS).[19] This survey started in the spring of 1965 with a national survey of high school seniors. Subsequent surveys were conducted in 1973 and 1982. A total of 1135 respondents (567 men and 568 women) completed all three waves, providing an unadjusted retention rate of 68 percent.[20] Respondents were 18 years old in the first wave and 35 in the last.

IV.A. Characteristics of YPSS Respondents

Descriptive statistics for the YPSS sample are presented in Table IX. The sample design implies that all respondents had at least completed high school. The earnings distribution reflects the fact that the average educational attainment in the sample exceeded the national average. Only 10 percent of the respondents in 1973, and 14 percent in 1982, were in the bottom thirty-third percentile of the national income distribution. For this reason (and because of the relatively small sample size), we do not report results separately by income group. Between 1965 and 1973, 63 percent of the men and 73 percent of the women married. By 1982, 10 percent of female, and 6 percent of male, respondents had divorced. The survey years also saw most respondents have children. In 1973, 50 percent of the women and 40 percent of the men had at least one child. By 1982, this figure had risen to over 70 percent for both sexes.

Changes in a respondent's marital status between 1973 and 1982 affected his/her income. Irrespective of gender, divorce between 1973 and 1982 lowered a respondent's family income. The decline in family income was, however, much sharper for a woman who divorced. Conversely, marriage between 1973 and 1982 raised a man's, but lowered a woman's, earnings. These effects were mainly driven by changes in labor supply, especially for women. For this reason, we choose not to use income variables as covariates in the analysis.

The class of 1965 lay at the heart of the protest generation. In their early adulthood they were witnesses to sweeping political and social changes such as the rise of the civil rights and women's liberation movement. The impact of some of these events on

19. The survey was specifically designed to study political socialization and was conducted by the Survey Research Center and Center of Political Studies of the University of Michigan [Jennings and Markus 1984]; also see Appendix 1.

20. Jennings and Markus [1984] showed that the attrition caused no apparent bias.

TABLE IX
DESCRIPTIVE STATISTICS, YPSS

Variable		percentage 1965	1973	1982
female		50.0	50.0	50.0
age [year]	Men	18.2	26.2	35.2
	Women	18.0	26.0	35.0
Family formation				
married	Men	0.0	63.3	74.6
	Women	0.0	73.0	71.3
divorced	Men	0.0	2.4	5.9
	Women	0.0	3.6	10.2
child	Men	0.0	39.8	74.9
	Women	0.0	51.9	79.2
Political preferences				
democrat	Men	29.6	30.0	25.6
	Women	35.1	39.0	37.1
idemocrat	Men	51.3	47.1	41.4
	Women	61.2	53.7	53.0
Other				
equal roles	Men	n.a.	31.9	44.2
	Women	n.a.	31.5	52.1
church	Men	74.6	21.5	28.9
	Women	87.1	32.7	43.6
union	Men	n.a.	20.8	28.0
	Women	n.a.	6.1	9.0

n.a. = not available. The union variable in 1982 is available for 471 men and 487 women. All values reported are means for the 1135 YPSS respondents.

respondents' social and political outlooks can be gauged from the YPSS survey. In 1973 one-third of both male and female respondents favored equal roles for men and women. By 1982, gender differences had emerged with 52 percent of the women, but only 44 percent of the men, favoring equal roles. Another indicator of changing social mores is church attendance. Between 1965 and 1973 church attendance fell from over 70 to under 35 percent for both sexes. Between 1973 and 1982 church attendance recovered, but remained well below 50 percent. Throughout, women were more likely to attend church. Finally, unionization increased over the period. Although more men than women were unionized, the increase was marginally greater among women (between 1973 and 1982, unionization increased from 21 to 28 percent among men and 6 to 9 percent among women).

The YPSS and NES survey questions on party affiliation are identical. In addition to **idemocrat,** we also use a stronger measure of Democrat identification **democrat** as the dependent variable. The latter dummy equals 1 only if the respondent identifies self as a Strong- or Weak-Democrat. Between 1965 and 1982 the proportion of respondents who identified themselves as **idemocrat** fell, with **democrat** affiliation exhibiting a similar, though nonmonotone, trend. Moreover, relative to nondivorced women, divorced women were more likely to identify themselves as **democrat.** The converse was true of divorced men. Edlund and Pande [2001] present transition matrices for how changes in a respondent's marital status affected his/her **democrat** affiliation between 1973 and 1982. These showed that every woman who identified as **democrat** in 1973 and divorced between 1973 and 1982 remained **democrat** in 1982; while only half of the men who divorced between the last two survey waves remained **democrat** in the latter wave. Moreover, while the category non-**democrat** (Republicans and Independents) gained male support, the gain was greater among men who divorced. The **idemocrat** measure produced qualitatively identical, but more muted, results.

IV.B. Estimation and Results

We use an OLS linear probability regression model to estimate how changes in individual i's marital status at time t impact on his/her Democratic affiliation:

$$d_{it} = \tau_t + \chi_i + \phi_1 m_{it} + \phi_2 \delta_{it} + \phi_3 \theta_{it} + \phi_4(f_i \times m_{it}) + \phi_5(f_i \times \delta_{it}) + \phi_6(f_i \times \theta_{it}) + \epsilon_{it}, \tag{5}$$

where m_{it} is a marriage dummy (married) and δ_{it} a divorce dummy (divorced). τ_t denotes the year dummies, and χ_i a time-invariant individual fixed effect. Thus, unlike our NES-based analysis which exploited CPS-state-year variation in divorce rates for identification, this analysis identifies the impact of marital status on political preferences from changes in individual marital status between successive waves of the YPSS survey. ϕ_4 and ϕ_5 capture the gender differential effect of marriage and divorce, respectively. Finally, to examine how other time-varying individual characteristics mediate the relationship between mari-

TABLE X
MARITAL STATUS AND DEMOCRATIC PARTY IDENTIFICATION

	Dependent variable:					
	democrat			**idemocrat**		
	(1)	(2)	(3)	(4)	(5)	(6)
married	−0.034	0.036	0.023	−0.031	0.029	0.077
	(0.031)	(0.039)	(0.050)	(0.029)	(0.037)	(0.050)
female × married	−0.005	−0.095*	−0.082	−0.033	−0.107**	−0.130*
	(0.031)	(0.050)	(0.073)	(0.029)	(0.047)	(0.074)
divorced	−0.270***	−0.276***	−0.274***	−0.108	−0.071	−0.106
	(0.077)	(0.079)	(0.087)	(0.073)	(0.075)	(0.081)
female × divorced	0.377***	0.290***	0.294**	0.218**	0.160*	0.228*
	(0.093)	(0.100)	(0.110)	(0.088)	(0.095)	(0.103)
child	—	−0.105***	−0.080*	—	−0.090**	−0.111***
		(0.037)	(0.044)		(0.035)	(0.044)
female × child	—	0.096*	0.068	—	0.083*	0.111*
		(0.049)	(0.058)		(0.046)	(0.059)
church	—	0.018	−0.033	—	−0.004	0.006
		(0.029)	(0.050)		(0.028)	(0.050)
female × church	—	−0.051	0.097	—	−0.035	−0.014
		(0.040)	(0.067)		(0.038)	(0.068)
union	—	—	0.086*	—	—	0.127***
			(0.040)			(0.049)
female × union	—	—	0.010	—	—	−0.016
			(0.088)			(0.089)
equal roles	—	—	0.020	—	—	−0.003
			(0.039)			(0.040)
female × equal roles	—	—	−0.037	—	—	0.009
			(0.054)			(0.054)
R^2 within	0.010	0.014	0.007	0.020	0.024	0.006
N	3,385	3,385	2,090	3,385	3,385	2,090

OLS regression results are reported, with standard errors in parentheses. The regressions in columns (1), (2), (4), and (5) consist of observations of YPSS respondents for the 1965, 1972, and 1983 waves, while regressions in columns (3) and (6) are based on the 1973 and 1982 waves only. All regressions include individual and year fixed effects. * indicates significance at 10 percent, and ** at 5 percent.

tal status and political preferences, we sequentially include elements of a vector of time-varying individual characteristics denoted θ_{it} in our regression.

Table X, column (1), tells us that marriage lowers the likelihood that a woman, relative to a man, is a **democrat.** This effect, however, is statistically indistinguishable from 0. In contrast, divorce has a strong and significant gender differential effect on political preferences—it makes a man 27 percentage points less likely to be a **democrat.** Divorce implies a political gender gap of 38 percentage points. Since roughly 8 percent of the sample were divorced by 1982, a back-of-the-envelope calculation suggests that divorce can account for 3 percentage points (0.08×0.38) of the gender gap.

Column (2) includes information on whether the respondent has a child, and on the respondent's degree of religiosity (as measured by church attendance). Having a child makes a respondent 10 percentage points less likely to be a **democrat.** The effect differs across men and women. It is much more muted for women, and we cannot reject the hypothesis that the negative relationship between having a child and **democrat** affiliation is restricted to men. We speculate that gender differences in preferences for tax-financed support of single parents may lie behind this. Since single parents tend to be mothers, such support favors mothers over fathers. In contrast, church attendance does not affect political preferences significantly.

Column (3) includes information on union membership, and the respondent's views on gender equality. As information on these two variables is only available since 1973, the sample size is reduced accordingly. Unionization makes respondents 8 percentage points more likely to be **democrat,** and there is no evidence of gender differences. However, we find no evidence that respondent views on gender equality impact political preferences.

Columns (4)–(6) report reestimates of these regressions, using **idemocrat** as the dependent variable. Our findings are qualitatively identical. However, comparing the effect of divorce on the two measures of political affiliation reveals interesting differences. Divorce loosens the extent of male Democratic affiliation. In particular, it significantly lowers the likelihood that a man is a **democrat** but not the likelihood that he is a **idemocrat.** By contrast, divorce makes erstwhile non-**idemocrat** women roughly 20 percentage points more likely to favor the Democratic party. Finally, a broader definition of Democratic affiliation strengthens the positive relationship between unionization and Democratic affiliation.

The early adulthood years for the class of 1965 coincided with the rise of the women's liberation movement. This raises the concern of omitted variable bias. While we cannot rule out the possibility that, for instance, feminism caused respondents to simultaneously change both their political behavior and their marital status, we can test for reverse causality, i.e., whether changes in political preferences presaged divorce. To do so, we ran fixed effect regressions where the dependent variable was a dummy that equaled one if the respondent changed marital status between 1973 and 1982 and the explanatory variable of interest was a dummy for whether the respondent changed political affiliation between 1965 and 1973. We found that neither leftward nor rightward switches in political affiliation between 1965 and 1973 predicted divorce between 1973 and 1982 [Edlund and Pande 2001].

V. Summary and Discussion

If marriage transfers resources from men to women, then the dramatic decline in marriage over the last 30 years made men richer and women poorer. This, we hypothesize, would impact on the political preferences of middle-income groups but not those of the poor or the rich. We present empirical evidence consistent with this hypothesis. Increased societal incidence of divorce, or the actual experience of divorce, both affect men's and women's political preferences in such a way as to increase the gender gap, and the effect is largely confined to the middle-income group.

Concurrent with the rise in nonmarriage, women improved their ability to earn their own income, by obtaining better qualifications, and greater acceptance at all levels in the workforce. While the changes in the marriage and labor markets are clearly linked, it is unclear which drove which. The introduction of the Pill may have reduced transfers from men to women, suggesting that greater female labor market presence is largely a response to this shortfall. However, this is not to deny the possibility of either a direct labor market effect on political preferences or that labor market gains outweighed the marriage market losses for a substantial subset of women. In fact, we find that working makes middle-income women favor the left. Throughout, the gender differential effect of divorce risk on support for the Democratic party among the middle-income group remains robust.

While the discussion centered on how increasing nonmarriage affected the political gender gap, the empirical testing focused on divorce. Divorce is not the only reason for nonmarriage. The age of first marriage has risen, as has the level of out-of-wedlock fertility. An alternative measure of the rise in nonmarriage is the fall in the proportion of adults who are currently married. In Edlund and Pande [2001] we show that this decline is uncorrelated with the gender gap. This is consistent with the view that later age of marriage often reflects greater human capital investments, especially on the part of women (possibly in response to increased risk of divorce) and with the fact that in the United States, out-of-wedlock fertility is so far not common among the middle-income groups.

Over the past thirty years, the principal political parties have adopted sharply diverging stances on social issues [Adams 1997]. It is not immediately clear how these stances relate to their long-standing ideologies or historical constituencies. One could argue that the fiscal libertarianism espoused by the Republican party would be a good fit with an equally libertarian position on issues of personal choice such as abortion. It is equally surprising that the Democrats should have been willing to alienate the Catholics and evangelical Christians, groups who have historically formed part of their constituency, by adopting a pro-life stance [Erikson and Tedin 1994]. One possible explanation afforded by this paper is that parties adopt social policies that promote family formation patterns conducive to their preferred redistributive policies.

Finally, the paper suggests a way of measuring the overall changes in the relative economic fortunes of men and women. Analyzing changes in political proclivities allows us to examine both the effects of improved labor market opportunities for women and the income effects associated with shorter marriages.

APPENDIX

The data sources are abbreviated as NES for National Election Studies cumulative file 1948–1998; CPS for Annual Current Population Survey March Supplement 1964–1996; YPSS for Youth Parent Socialization panel survey; youth section 1965, 1973 and 1982 waves. In all data sets no answer, do not know, and not applicable are coded as missing values. The NES and CPS samples are restricted to respondents aged 18–64 years.

NES and YPSS variables

Demographics:

female (NES and YPSS) Dummy equals 1 if respondent is female.

married (NES and YPSS) Dummy equals 1 if respondent is married and living with spouse; for YPSS dummy also equals 1 if spouse is in military service.

divorced (YPSS) Dummy equals 1 if respondent is divorced.

Black (NES) Dummy equals 1 if respondent is African-American.

age (NES and YPSS) Respondent age in years.

cohort (NES) Four cohort dummies were created: Cohort born (i) prior to 1910; (ii) 1911–1942; (iii) 1943–1958; and (iv) after 1959.

Economic characteristics:

education (NES) Original question: 1964–1972 How many grades of school did you finish? 1974–1996 What is highest grade of school or year of college you have completed? Four education dummies were created (i) **0–8 grade** Grade school or less; (ii) **9–12 grade** Completed grade school but no more than high school; (iii) **some college** completed high school, some college education but no college degree; (iv) **college** Completed college or higher degree.

labor (NES) Dummy equals 1 if respondent is in the labor force at the time of the survey.

income (NES) Three family income dummies were created: annual family income in (i) 0–33 percentile (poor); (ii) 34–95 percentile (middle income); and (iii) 96–100 percentile (rich).

union (YPSS) Dummy equals 1 if respondent is a union member.

Preferences:

Democrat (NES and YPSS) Original question: Generally speaking, do you think of yourself as a Republican, a Democrat, an Independent or what? Prompted answers coded as 1 = Strong Democrat; 2 = Weak Democrat; 3 = Independent-Democrat; 4 = Independent-Independent; 5 = Independent-Republican; 6 = Weak Republican; 7 = Strong Republican. **idemocrat** dummy equals 1 if respondent answered 1–3 from above classification; and **democrat** dummy equals 1 if respondent answered 1–2 from the above classification. In the 1965 wave of the YPSS, the categories were slightly different: 11 = Strong Democrat; 12 = Not very strong Democrat; 13 = yes, Democrat; 14 = No, neither; 15 = Yes, Republican; 16 = Not very strong Republican; 17 = Strong Republican. **idemocrat** dummy equals 1 if respondent answered 11–13 from the above classification; and **democrat** dummy equals 1 if respondent answered 11–12 from above classification.

govspend (NES) Dummy equals 1 if respondent answered 4 through 7, on a 7-point scale, where 1 was Government should provide many fewer services: reduce spending a lot; and 7 was Government should provide many more services: increase spending a lot.

pro-choice (NES) Dummy equals 1 if respondent stated that abortion should be permitted if, due to personal reasons, the woman would have difficulty in caring for the child, or that abortion should never be forbidden, since one should not require a woman to have a child she does not want.

equal roles (NES and YPSS) Original question: Recently there has been a lot of talk about women's rights. Some people feel that women should have an equal role with men in running business, industry and government. Others feel that women's place is in the home. And other people have opinions somewhere in between. Where do you stand? Dummy equals 1 if respondent states men and women should have equal roles.

religion (NES) Based on respondent's religious identity, three dummies: Catholic, Protestant, and Jewish.

church (NES) Dummy equals 1 if respondent attends church two or more times a month.

social (NES) Dummy equals 1 if respondent stated that the most important problem government should try to take care of was social (includes crime, drugs, civil liberties and nonracial civil rights, women's rights, abortion rights, gun control, family/social/religious/moral decay, church and state, etc.)

economics (NES) Dummy equals 1 if respondent stated that the most important problem government should try to take care of was economics, business, and consumer issues (includes foreign investment, tariffs/protection of U. S. industries, international trade deficit/balance of payments, immigration, interstate commerce/transportation)

welfare (NES) Dummy equals 1 if respondent stated that the most important problem government should try to take care of was social welfare (includes population, child care, aid to education, the elderly, health care, housing, poverty, unemployment, welfare etc.)

CPS variables

CPS household weights used to create population shares. Sample restricted to respondents aged 18–64. **pdivorced** created using information on CPS respondent marital status, while **plabor** used information on all adult individuals in household.

pdivorced Proportion of individuals in CPS-state aged 18–64 currently divorced.

plabor Proportion women in CPS-state aged 18–64 currently in the labor force.

CPS-state The correspondence between CPS-state and individual U. S. states is as follows: New England—Maine, New Hampshire, Vermont, Massachusetts, and Rhode Island; East North Central—Michigan and Wisconsin; West North Central—Minnesota, Iowa, Missouri, North Dakota, South Dakota, Nebraska, and Kansas; Middle Atlantic—Delaware, Virginia, Maryland, and West Virginia; South 1—North Carolina, South Carolina, and Georgia; South 2—Alabama and Mississippi; South 3—Arkansas, Oklahoma, and Louisiana; Border—Kentucky and Tennessee; Mountain—Montana, Idaho, Wyoming, Utah, Nevada, Colorado, New Mexico, and Arizona; Pacific—Washington, Alaska, Hawaii, and Oregon. For all other states the correspondence is one-to-one.

Unilateral divorce series

unilat Dummy equals 1 for all years from when a state introduces a no-fault ground for divorce and has no separation requirement and follows classification by Gruber [2000].

Description of YPSS survey

In 1965 the students interviewed were chosen from a national probability sample of 97 secondary schools selected with a probability proportionate to school size. At each school, 15–21 randomly designated seniors were interviewed, for a total of 1669 respondents (dropouts were eliminated from the sample). In 1973, 1119 of these were reinterviewed, and an additional 229 completed mailback questionnaires. In 1982, 1135 were reinterviewed (of which 177 completed the mailback questionnaire). This reflected a retention rate of 68 percent between 1965 and 1982, and a rate of 84 percent between 1973 and 1982.

APPENDIX 2: YEAR OF INTRODUCTION OF NO-FAULT GROUND AND MAX 3 YEARS SEPARATION REQUIREMENT (**UNILAT**)

State	Year of **unilat**	Governor year of **unilat**	State	Year of **unilat**	Governor year of **unilat**
Alabama	1971	D	Nebraska	1972	D
Alaska	1935	D	Nevada	1967	D
Arkansas	s.c.	—	New Hampshire	1971	R
Arizona	1973	R	New Jersey	s.c.	—
California	1970	R	New Mexico	1933	D
Colorado	1972	R	New York	s.c.	—
Connecticut	1973	R	North Carolina	s.c.	—
Delaware	1968	D	North Dakota	1971	D
Florida	1971	R/D	Ohio	s.c.	—
Georgia	1973	R	Oklahoma	1953	D
Hawaii	1972	D	Oregon	1971	R
Idaho	1971	R/D	Pennsylvania	s.c.	—
Illinois	1984	R	Rhode Island	1975	D
Indiana	s.c.	—	South Carolina	s.c.	—
Iowa	1970	R	South Dakota	1985	R
Kansas	1969	D	Tennessee	s.c.	—
Kentucky	1972	D	Texas	1970	D
Louisiana	s.c.	—	Utah	1987	R
Maine	1973	R	Virginia	s.c.	—
Maryland	s.c.	—	Vermont	s.c.	—
Massachusetts	1975	D/R	Washington	1973	R
Michigan	1972	R	Washington, DC	s.c.	—
Minnesota	1974	D	West Virginia	s.c.	—
Mississippi	s.c.	—	Wisconsin	1972	D
Missouri	s.c.	—	Wyoming	1977	D
Montana	1973	D			

Source: Year of **unilat** from Gruber [2000].
R = Republican; D = Democrat; s.c. = still consent. If there was a shift of power the year preceding **unilat,** the party affiliations are given as preceding year/year.

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