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Abstract

Women may want greater taxes and redistributional spending than men if male incomes exceed female incomes and men and women are not always married. Cross-section evidence from opinion polls provides indirect support for this view. Women express greater favorability toward redistributional spending than do men; the difference is explained largely but not entirely by income differences. The empirical analysis provides partial support for Meltzer and Richard (1981), and suggests that increases in singleness after World War II may have acted to increase differences by sex in desired redistributional spending.

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We study differences by sex in desired spending on programs that redistribute from higher-income to lower-income households. Examples of such programs include welfare, food stamps, government-provided medical care, and government-provided child care. These programs tend to redistribute income because the taxes used to finance the programs increase with income but benefits fall with or are independent of income.

Differences in desired redistributional spending can arise for several reasons. A first plausible possibility is that women may be more "altruistic" than men in the sense that they have greater pure tastes for redistributional policies.[1] A second is that women earn less on average than men, so women gain net income on average and men lose net income on average from redistributional policies. For instance, average earnings of single women are roughly 80 percent of average per-person earnings of married couples, while average earnings of single men are roughly 115 percent of average per-person earnings of married couples.[2] Such earnings differences may reflect differing labor-market engagements that arise because women bear children.[3]

The idea that relative income is a determinant of desired redistributional

[1] Evidence from experimental studies of altruism is mixed. Dictator games provide perhaps the most information because individuals in such games chose how much to redistribute voluntarily to others. In such games, Bolton and Katok (1995) find no male-female differences, Eckel and Grossman (1998) find that women are more altruistic than men, and Andreoni and Vesterlund (1998) find that women are sometimes more altruistic and sometimes less altruistic than men. In public-good games, Brown-Kruse and Hummels (1993) find that men are more cooperative than women, but Nowell and Tinkler (1994) find the opposite. Mixed results on altruism also characterize the psychological literature—see Eagly (1995) for a survey.

[2] Calculated from data in Census Historical Income Tables, Table H-9 (online).

[3] See e.g. Mincer and Polachek (1974), Gronau (1988), Korenman and Neumark (1992).

spending also underlies Meltzer and Richard's (1978, 1981) median-voter model of redistributional spending.[4] The Meltzer-Richard model abstracts from the possibility that individuals marry and share income. This abstraction matters because income sharing in marriage redistributes between the sexes, which may reduce or eliminate the effect of public redistribution policies on relative male/female consumption. Expected marital status is therefore a further potential determinant of differences by sex in attitudes toward government. A male/female difference in average earnings may cause little or no average difference by sex in attitude toward redistributional spending in a "traditional" world in which all men and women are married and share income, for instance, but may have a substantial effect in a world in which many men and women expect to live at least part of their adult lives as singles.

We study cross-section data from the National Election Studies (NES) of opinions of roughly 32,000 adults over 1952-1992 to test for sex-based differences in desired redistributional spending, and to examine how much of these differences are explained by a person's income and how much are explained by the person's sex. To make clear the main theoretical issues and specify the econometrics, we begin in section II by modeling how an individual's desired redistributional policy depends on expected future time spent in different marital statuses and on expected incomes in the different statuses.

Empirical analysis is in section III. Broadly, we find two things. First, women on average express a greater desire than men for redistributional spending. Second, only a small part of the difference appears to be due to a greater female pure taste for redistribution ("altruism"); most appears to be explained by income differences in different marital statuses. Specifically, the analysis provides support for the idea in Meltzer and Richard that desired redistributional policy depends inversely on

[4] See also Romer (1975), Roberts (1977).

income.[5]

The dependence of desired redistribution on income differences by marital status may have further implications. The ratio of economically active adults (persons aged 18-64) living in single-adult households to the total number of economically active adults (the single rate) rose from 0.25 in 1950 to 0.40 in 1998.[6] Given earnings differences by sex, this upward trend in singleness may have acted to increase differences by sex in attitudes toward redistributional spending, which in turn may help understand trends in spending and political allegiances. For instance, Federal spending on redistributional programs has also increased since 1950. Because female voters outnumber male voters,[7] a natural hypothesis is that increases in singleness have been one force behind the increased redistributional spending.[8;9]

Causality under this hypothesis is reversed from that under the common argument that welfare causes some single women to have children.[10] The hypothesis here is that singleness causes welfare programs rather than that welfare programs cause singleness.[11] Because empirical studies based on time series or on cross sec-

[5] Tests of the Meltzer-Richard hypothesis are mixed. Peltzman (1980), Inman (1978), Murrell (1985), and Brosio and Marchese (1988) find little support, while Mueller and Murrell (1986) and Husted and Kenny (1997) find support.

[6] The 1950 figure is calculated from Census Bureau, (1975); the 1998 figure is from Summary Table B, March 1998 CPS (online).

[7] In the 1996 Presidential elections, 115 votes were cast by females for every 100 votes cast by males. There were also 113 females registered to vote for every 100 registered males (CPS/Census, P20-504, Table 4, online).

[8] A number of redistributional spending programs were introduced in the 1930s, but the single rate started to increase notably only after about 1950. This suggests that redistributional spending is also driven by factors other than singleness.

[9] Lott and Kenny (1998) provide econometric support for the view that the female vote helps explain rising public spending and other political outcomes.

[10] See Murray (1984).

[11] Akerlof et al. (1996) argue that the increased prevalence of female heads of household has in

tions across states generally find that increases in welfare only poorly explain the increased prevalence of female heads of household, it may be that the empirically stronger causality runs from singleness to redistributional spending.[12] We study the causality issue briefly in section IV.

II. THEORETICAL ISSUES

Expected Marital Status

The main issues can be laid out briefly by thinking of policy as a linear tax under which government imposes a levy on all labor income at rate t and uses the revenue to make transfer payments in amount r per economically active person.[13] Under the government budget constraint, the redistributional payment equals average tax revenue or $r = ty$, where y is average labor income, that is, total labor income of all economically active individuals divided by the number of such individuals. Under the constraint, a value of the tax rate t determines a value of r in general equilibrium. In this way, t provides an index of redistributional policy, with greater r requiring greater t as long as the Laffer curve slopes upward.

To understand how an individual's expected marital status affects his or her desired redistributional spending, consider an individual indexed by sex $s = m$ (male) or $s = f$ (female) who spends a portion π of his or her future lifetime single (marital

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turn been caused by improvements in birth control due to freer abortion and better contraception.

[12] Moffitt (1992) surveys studies of the effects of welfare on female headship, concluding that the "failure to find strong benefit effects is the most notable characteristic of this literature."

[13] The linear tax assumption plays a more crucial role in Romer (1975), Roberts (1977), and Meltzer and Richard (1981): it ensures that the median-voter theorem applies. The assumption's main inaccuracy for our purposes is that actual redistributions in the form of welfare payments, food stamps, and medicaid benefits go proximately to women, so the assumption may understate differences by sex in desired redistributional spending.

status $n = 1$) and spends the remaining portion $1 - \pi$ married ($n = 2$). (In the empirical section, we consider three marital statuses: never married, married, and widowed or divorced.)

A single individual of sex s finances consumption (c_{1s}) from the sum of labor and transfer income according to the budget

$$c_{1s} = (1 - t) w_s l_{1s} + r, \tag{1}$$

where w_s is the individual's wage and l_{1s} is hours of labor chosen by the individual when single. A married couple has pooled consumption equal to pooled income:

$$c_{2m} + c_{2f} = (1 - t)(w_m l_{2m} + w_f l_{2f}) + 2r, \tag{2}$$

where w_m and w_f are wages and l_{2m} and l_{2f} are hours worked by the married man and married woman. For an individual of sex s, let s' denote the sex of the spouse.

Utility depends on consumption, own leisure, and spousal leisure if married. For a single individual, spousal leisure is zero, which we represent compactly by writing utility with $l_{1s'}$ equal to available hours h. In this section, we suppress dependence of utility on r due to "altruistic" tastes for redistribution.

A single individual of sex s chooses (c_{1s}, l_{1s}) to maximize utility $U_s(c_{1s}, l_{1s}, h)$ subject to (1) and nonnegativity constraints. Married individuals choose c_{2m}, c_{2f}, l_{2m}, and l_{2f} to maximize household utility $U_m(c_{2m}, l_{2m}, l_{2f}) + U_f(c_{2f}, l_{2f}, l_{2m})$ subject to (2) and nonnegativity constraints.[14]

[14] Under this household-utility-function approach, the use of family income is independent of the partner who receives the income. A growing body of empirical work suggests to the contrary that the use of household income depends on the partner who receives the income (Schultz 1990, Thomas 1990, Browning et al. 1994, Hoddinott and Haddad 1995, Lundberg et al. 1997, Pezzin and Schone 1997, Grey 1998). Such a dependence can arise in a model in which a couple maximizes a function of differences between each partner's utility and the partner's "threatpoint" utility, if the

An individual's expected future utility is the weighted average of the utilities the individual would achieve if single and if married, with weights equal to portions of future lifetime expected to be spent in each marital status:

$$EU_s = \pi U_s(c^*_{1s}, l^*_{1s}; h) + (1 - \pi) U_s(c^*_{2s}, l^*_{2s}, l^*_{2s'}), \qquad (3)$$

where stars (*) denote the labor and consumption levels that maximize $U_s(c_{1s}, l_{1s}; h)$ subject to (1) when $n = 1$, and maximize $U_m(c_{2m}, l_{2m}, l_{2f}) + U_f(c_{2f}, l_{2f}, l_{2m})$ subject to (2) when $n = 2$.

Under the expected-utility formulation (3), government policy can have a social-insurance property. If an individual would be poorer as a single and hence have a greater marginal utility of income as a single, for instance, then public redistribution that transfers resources from higher-income to lower-income households simultaneously transfers from the lower-marginal-utility married state to the higher-marginal-utility single state. Such transfers from lower-marginal-utility to higher-marginal-utility states also characterize insurance. Public redistribution differs from voluntarily purchased insurance, however, because the nature of redistribution is that some individuals win and some lose.

For a rational, forward-looking individual who views policy as fixed over time, desired redistribution may be thought of as the value of the index t that maximizes expected future utility (3). The first-order condition for a maximum is

$$\frac{dEU_s}{dt} = \pi \frac{dU_s(c^*_{1s}, l^*_{1s}; h)}{dt} + (1 - \pi) \frac{dU_s(c^*_{2s}, l^*_{2s}, l^*_{2s'})}{dt} = 0, \qquad (4)$$

latter depends on the income the partner receives (Manser and Brown 1980, McElroy and Horney 1981, Chiappori 1988, 1992, Apps and Rees 1988, Lundberg and Pollak 1993). A threatpoint specification would introduce an additional channel by which income redistribution would increase mean or median female control over resources, which would strengthen sex-based differences in desired redistributional spending.

which sets to zero the weighted average of the effects of marginal t on utilities conditional on being single and married, with weights equal to portions of time spent in each marital status. If the individual expects to spend no time as a single, for instance, then $\pi = 0$, and (4) reduces to the condition for choosing t to maximize utility conditional on being married.

To evaluate $dU_s(c_{1s}^*, l_{1s}^*, h)/dt$, differentiate U_s totally and "envelope out" terms by applying the derivative of the government budget constraint $r = ty$ and the first-order conditions for maximizing $U_s(c_{1s}, l_{1s}, h)$ subject to (1) to get

$$\frac{dU_s(c_{1s}^*, l_{1s}^*, h)}{dt} = \frac{\partial U_s(c_{1s}^*, l_{1s}^*, h)}{\partial c_{1s}}[-(w_s l_{1s} - y) + t\,dy/dt]. \qquad (5)$$

Thus the rise in a single person's utility from a marginal change in t equals the marginal utility of income times the sum of two terms. The first is the negative of the individual's relative income; the second is the change in tax revenue that arises because taxation distorts labor supply decisions. We exploit the former to specify econometrics in the section III: an individual's desired redistribution should depend on relative income when single. From (4), the importance of relative income when single should be greater the greater the portion of time the individual expects to be single.

For the married state, the analogue of (5) is

$$\frac{dU_m(c_{2m}^*, l_{2m}^*, l_{2f}^*)}{dt} + \frac{dU_f(c_{2f}^*, l_{2f}^*, l_{2m}^*)}{dt} = \frac{\partial U_m(c_{2m}^*, l_{2m}^*, l_{2f}^*)}{\partial c_{2m}}[-(w_m l_{2m} + w_f l_{2f} - 2y) + 2t\,dy/dt].$$

Thus the rise in household utility ($U_m + U_f$) given an increase in redistribution (t) is greater the lower the household's total earnings, and also is greater the less that taxes reduce aggregate labor. For econometric purposes, a married person's desired redistribution should depend on relative household income conditional on being married $(w_m l_{2m} + w_f l_{2f} - 2y)$. As above, the importance of this relative income should be greater the greater the portion of time the individual expects to be married.

Changes in Singleness

If an increase in the single rate coincides with an increase in expected singleness (π) and men on average earn more than women, then divergence in fiscal desire between the sexes may also occur. To see this, consider the situation of a woman who would have relatively low income as a single so $dU_f(c^*_{1f}; l^*_{1f}; h)/dt > 0$; and hence around the optimum $dU_f(c^*_{2f}; l^*_{2f}; l^*_{2m})/dt < 0$; also consider the situation of a man who would have relatively high income as a single so $dU_m(c^*_{1m}; l^*_{1m}; h)/dt < 0$ and $dU_m(c^*_{2m}; l^*_{2m}; l^*_{2f})/dt > 0$. Given that female voters outnumber male voters, it is useful to study the woman's situation first and in more detail.

Differentiate (4) with respect to π to obtain

$$\frac{d^2 EU_f}{dt\,d\pi} = \frac{dU_f(c^*_{1f}; l^*_{1f}; h)}{dt} - \frac{dU_f(c^*_{2f}; l^*_{2f}; l^*_{2m})}{dt} + \pi \frac{d^2 U_f(c^*_{1s}; l^*_{1s}; h)}{dt\,d\pi} + (1-\pi)\frac{d^2 U_f(c^*_{2f}; l^*_{2f}; l_m)}{dt\,d\pi}. \tag{6}$$

If $d^2 EU_f/dt\,d\pi$ is positive and the second-order condition for maximization of (3) holds, then an increase in π leads to an increase in desired redistribution (t).

The first two terms on the right-hand side of (6) represent direct effects and are both positive: an increase in singleness increases the weight on utility when single, when a woman is relatively poor and desires much redistribution, and decreases the weight on utility when married, when the woman is relatively well-off and desires little redistribution. The second two terms reflect general-equilibrium feedbacks on consumption and labor that arise because an increase in singleness alters aggregate labor, wages, and government revenue. It is not possible to sign these feedbacks.

As long as the feedbacks do not produce large declines in dU_f/dt, then $d^2 EU_f/dt\,d\pi$

is positive and an increase in singleness leads to an increase in the woman's desired redistribution. For the man, on the other hand, a rise in singleness leads to a decrease in desired redistribution as long as feedbacks are not too great.

III. EVIDENCE OF DIFFERENCES IN FISCAL DESIRE BY SEX

To examine whether differences in relative incomes explain desired redistribution and also explain sex-based differences in desired redistribution, we study National Election Studies (NES) surveys of political opinions of individuals aged 25-65 taken in election years from 1952 to 1992. The NES provides a non-longitudinal cross section of respondent traits as well as attitudes toward a variety of issues.[15] We focus on attitudes toward redistributional spending on food stamps, social security, federal job guarantees, government-supplied health care, and government spending generally. These have been sampled in recent years from five to 11 times. We also study party affiliation because this may be a broad proxy for redistributional policy and because party affiliation has been sampled in each of the 20 NES surveys from 1952-92. (We omit 1954 from the sample because the age variable was not measured continuously in the 1954 NES but rather took one of seven discrete values.)

As a control, we study attitudes toward defense spending. Because the redistributional component of defense spending may be minor, variables that explain desired redistributional spending may perform less well in explaining sex-based differences in desired defense spending.

[15] The total number of individuals sampled is 37,456. Each survey samples opinions of approximately 2000 individuals. Not every question was asked in every survey, so the number of responses on each question varies.

Method

The theory provides three predictions. First, greater income should generally reduce an individual's desired redistributional spending. Second, because of this and because women have lower earnings than men on average, women should desire greater redistributional spending than men on average. Third, the expected-utility structure of the theory implies that the importance of expected future income in a marital status as a determinant of desired redistributional spending should be proportional to expected time in the status. This suggests that expected future income in a marital status should enter a regression of desired redistributional spending as the interaction between expected per-adult household income in the status and expected future time in the status. Predicted signs of interaction coefficients are negative.

Accordingly, the empirical strategy involves three stages. In the first stage, we examine whether favorability toward a given type of redistributional spending and party affiliation differ by sex. This is done by testing whether mean male and mean female favorabilities differ, and also by regressing opinion responses on an indicator for sex (female = 1, male = 0) and testing whether the coefficient on the female indicator (the female coefficient) differs from zero.

In the second stage, we add current income to the regressions to test the prediction that coefficients on incomes are negative. If women on average desire greater redistributional spending because female incomes on average are lower, then inclusion of income should also reduce magnitudes of female coefficients. We use Wald tests to judge whether female coefficients change when the income is added as a regressor. To further evaluate the relative importances of sex and income in explaining differing responses by sex, we use analysis of variance to estimate the portion of the variance of opinion responses explained by sex and income in the stage-II regressions.

In the third stage, we construct year-by-year estimates of future incomes conditional on marital status and expected future times spent in each marital status, add present value of interactions between the two to the stage-I regressions, and test whether the interactions have negative coefficients. Because this specification may be theoretically preferred to the stage-II specification as a way to include income, the stage-III regressions may capture the effects of income better than the stage-II regressions. This has two consequences. First, stage-III estimates of female coefficients may be less contaminated by effects of income. Second, if lower female incomes explain part of differences by sex in desired redistribution, then better measurement of income should further reduce magnitudes of female coefficients. Again we use Wald tests to judge whether female coefficients change when income (interaction) terms are added, and we also use anovas to estimate the portion of the variance of opinion responses that is explained by sex and interaction terms.

NES data alone do not contain sufficient information to calculate (income-in-status) $\times$ (time-in-status) interactions. Namely, the NES contains a measure of current annual family income censored into five intervals, and contains no direct information about the expected portion of an individual's future lifetime spent in each marital status. Of course, current income may proxy for expected future income but the proxy is imperfect for at least four reasons. First, censoring is coarse. Second, year-to-year income fluctuations may introduce noise if the goal is to measure expected future income. Third, current income plus other variables may predict expected future incomes better than does current income alone. Fourth, theory suggests that opinions about redistributive spending may depend on incomes in each of the marital statuses in which the individual may find himself or herself in the future.

We use a synthetic cohort from the 1993 Survey of Income and Program Participation (SIPP) to calculate the interactions. The SIPP is an on-going survey conducted by the Census Bureau that follows a panel of individuals for 2.5 to four years. We use a sample of 10,179 men and 12,176 women aged 25 and older who participated in the entire three years of the 1993-95 SIPP panel. A desirable property of the SIPP is that it measures income continuously and not as a censored variable.

First, we estimate marital-status-dependent age-income profiles separately for men and women using random-effects regression over all three years (1993-95) of annualized SIPP data with age, age squared, race, and education in addition to current marital status as controls. Next, we estimate marital status transition probabilities using a probit model estimated over the second and third years (1994-95) of the SIPP data with sex, race, education, and previous-year income as controls. The use of previous-year income and the choice not to base estimation also on the first year of the SIPP data is dictated by concern that marital-status transitions may affect incomes. Because widowed and divorced individuals may differ substantially from never-married individuals, we consider three marital statuses: never married (n = 1); married (n = 2); and widowed or divorced (n = 3).

We then insert data from the NES into the estimated income-profile and transition-probability functions to compute status-dependant age-income profiles and transition probabilities for each individual in the NES. Beginning from an individual's current marital status in the NES, we recursively apply the demographic relationships between transition probabilities and the probability of being in a given marital status in each future year of life, computing the latter probabilities. In this way, we exploit the longitudinal structure of the SIPP to calculate transition probabilities and capture how individual differences in current marital status create cross-section variation in expected probabilities of being in each marital status in each future year of life. For each future year of a person's expected lifetime, we multiply income in a status by the probability of being in the status, and finally compute the present value sum of these interactions with actuarial discounting rates equal to mortality rates.

A more detailed description of how interaction variables are computed is in an appendix.

NES responses are coded on either a seven-point approval scale or a four-point increase/decrease scale.[16] We estimate such ordered, discrete choices as ordered logits. For a particular issue, the latent fiscal desire t_i^* of individual $i = 1; \ldots; N$ is a function of individual characteristics and other controls X_i with coefficient vector β' and disturbance ε_i:

$$t_i^* = \beta' X_i + \varepsilon_i; \qquad i = 1; \ldots N; \tag{7}$$

where we do not observe t_i^* directly but rather observe one of the $J = 4$ or $J = 7$ discrete response choices. (Subscripts to mark issues are suppressed; separate versions of (7) are estimated for each issue.) For ordered logits on the four-level issues (social security and food stamps), "cut out spending" was favored by a small number of respondents so this category is added to "reduce spending." For the anovas below, we retain all four levels. Neither treatment has much effect on results.

An econometric concern is that ambiguity or uncertainty in opinions may differ across individuals so the variance of the disturbance in the choice function (ε_i) may plausibly differ across individuals. To allow for such heteroskedasticity, we assume that the variance of individual i's choice depends on a vector of explanatory variables Z_i according to:

$$Var(\varepsilon_i) = \exp(Z_i \gamma)^2 :$$

Inclusion of any variable in Z_i that also appears in X_i would make it difficult to interpret the size and significance of the variable's coefficient in X_i, so we do not

[16] Choices on the latter are "increase spending," "keep spending the same," "decrease spending," or "cut out spending entirely."

include the female indicator, incomes, or expected time in a marital status in Z_i. The main candidates for inclusion in Z_i are therefore age, race, education, current marital status, current income, and year of survey. Because we do not have strong priors, we adopt the somewhat conservative strategy of including only age, education, and year of survey.

Results

Stage I.|

First, mean male and mean female favorabilities differ. For each issue, the mean female favorability toward redistributional spending exceeds the mean male favorability, and the difference is statistically significant at the 99.9 percent level.

This is confirmed in stage-I regressions summarized in table 1. In these regressions, only the female indicator and year of survey are included in X_i: Year of survey accounts for year-specific and issue-specific shifts in mean opinions such as occurred in 1992 when both men and women were unusually positive toward government health care, and as occurred in 1980 when both men and women were unusually positive toward defense spending. We do not include individual controls for age, race, education, current marital status, current income, or other variables such as the presence of children in stage-I regressions because our strategy is to compare female coefficients when income is not included (stage I) with coefficients when income is included (stages II and III) in order to evaluate the extent to which sex-based differences are due to income differences. Inclusion in stage I of individual controls that predict individual income would partially defeat this strategy. A small number of background controls also helps avoid simultaneity bias from controls that might be endogenous, such as the presence of children and work status.

For each type of spending, we see in the table 1 regressions that women differ from men at the 99.9 percent confidence level in favorability toward redistributional spending. (Standard errors are reported in parentheses below estimated coefficients.) In all cases, women desire greater redistributional spending. Women tend to identify with the Democratic Party. Women also desire less defense spending.

To test whether the vector Z_i provides information about the variance of ε_i, we compute the likelihood ratio $LR = 2(L_{he} - L_{ho})$, where L_{he} is the log-likelihood of the heteroskedastic model based on Z_i and L_{ho} is the log-likelihood of the standard homoskedastic model. The ratio LR has a χ^2 distribution with degrees of freedom equal to the number of variables in Z_i, and equals zero under the null of homoskedasticity. We see in the table that the homoskedastic model is rejected in favor of the heteroskedastic model at the 99.9 percent level for all issues. Accordingly, we use the heteroskedastic model in all stage-II and stage-III regressions below. In an appendix, we provide results for the homoskedastic model; these are broadly similar to those from the heteroskedastic model.

Stage II.|

In stage-II regressions, we add current income as measured in the NES to X_i; the female indicator and year of survey are also included as in stage I. Income may have different effects depending on the number of adults per household because returns to scale or specialization may make income more effective at generating utility in two-adult households. To allow the coefficient on income to differ by number of adults in the household, we interact income with a dummy for single or married.

The second set of rows in table 1 summarize stage-II results. The line Wald test reports the test variable and significance [in brackets] for a two-tailed test of the hypothesis that the female coefficient has the same value as in the stage-I model. For each redistributional issue, inclusion of current income causes the female coefficient to move closer to zero, as expected. For the two spending areas that most clearly target low-income women (government health care and food stamps) the relative decline in the female coefficient is the greatest. For instance, ratios of the estimated female coefficient in the stage-II regression to that in the stage-I regression were 0.60 and 0.56, respectively, for these two issues, but 0.86, 0.82, and 0.77 for the other three issues. For three of the five issues, the change in the female coefficient is significant at the 10 percent level.

Estimates for party identification show the same pattern as those for specific issues. Women tend to identify more with the Democratic Party, but the female coefficient is roughly half as large when current income is included, which would indicate that about half of the female Democratic preference may be explained by lower female incomes. The Wald test for this difference is highly significant. For the non-redistributional issue (defense spending), by contrast, the female coefficient exhibits essentially no change when current income is included.

All income coefficients but one are negative for issues other than defence spending; most are also significant. This provides support for the theory here and more generally for the view in Meltzer and Richard (1978, 1981) that desired redistribution depends negatively on income.

To judge further the relative importance of income in explaining sex-based differences, table 2 presents anova decompositions of variations in categorical variables that take value one if the individual gave a particular response (out of the four or seven possible responses) into variations attributable to sex and income. Reported partial sums of squares measure explanatory powers of the sex and income variables; associated p-values are in square brackets. In most instances, the explanatory power of income exceeds that of sex, and when it does not, the difference is small. The effect of income exceeds that of sex in 37 of 43 instances in which partial sums of squares for both the female indicator and an income variable differ significantly from zero at the ten percent level. Of these 43, the explanatory power of income is more than ten times that of sex in nine instances. In 14 other instances, the partial sum of squares for the female indicator differs insignificantly from zero, and the partial sum of squares for an income variable is significant and greater than the partial sum of squares for the female indicator. In 13 of the 14 instances, the difference is greater than a factor ten.

Stage III.|

Stage-III regressions are like stage-II regressions except that the effects of income are measured by discounted sums of interactions of expected future income in a marital status (never married, married, or widowed or divorced) times expected future time spent in the status, instead of by interactions of current income times a dummy for the current number of adults per household (one for unmarrieds or two for marrieds).

As with the stage-II results, female coefficients are positive for all five specific redistributional issues, but each stage-III female coefficient is smaller than the corresponding stage-II coefficient. This provides additional support for the model; better measurement of income should reduce the magnitudes of female coefficients if lower female incomes partly explain of differences by sex in desired redistribution.

In detail, ratios of female coefficients in stage III regressions to stage I regressions are 0.12, 0.06, 0.43, 0.18, and 0.42, respectively, for the five types of redistributional spending; these ratios are substantially lower than corresponding ratios reported above from the stage-II regressions. Wald tests indicate that changes in female coefficients from stage-I to stage-III regressions are highly significant for all redistributional issues, again suggesting that some of the greater female desire for redistribution is due to lower female incomes. Stage-III female coefficients for attitudes toward government spending, health care, and food stamps are indistinguishable from zero at the five percent level, which suggests that nearly all of the greater relative female desire for these forms of redistributional spending may be explained by lower female incomes.

For party identification, inclusion of income-status interactions causes the sign of the female coefficient to reverse from that in the stage-I regression.[17] The interpretation would be that females have a natural proclivity for the Republican Party, but that lower female incomes tend to make women identify with Democrats. This is possible if a party represents more than just a position on redistributional issues. For defense spending, inclusion of income (interactions) again leads to essentially no decline in the coefficient on the female indicator from that in the stage-I regression.

Interaction coefficients for nondefense issues were mainly negative. For instance, 13 nondefense interaction coefficients were negative and significant and three were positive and significant. Negative interaction coefficients suggest that the greater the time an individual can expect to be in a given marital status, the greater is the tendency for desired redistributional spending to decrease with income in the status.

Table 3 presents anovas similar to those in table 2 except that the stage-III interactions between expected future income and time in each status replace current NES income interacted with the single/married dummy. The effect of income exceeds that of sex in 20 of 30 instances in which partial sums of squares for both the female indicator and an income variable differ significantly from zero. Of those, the explanatory power of income is more than ten times that of sex in seven instances. More notably, partial sums of squares for the female indicator differ insignificantly from zero in only nine of 36 cases in table 2 but in 22 of 36 cases in table 3. As a

[17] Kaufmann and Petrocik (1997) find that women tended more to vote and identify with the Republican party than did men from 1948–1960. They also find that, since 1964, male allegiances have shifted while female allegiances have remained fairly constant, so women have tended more to vote and identify with the Democratic party than did men from 1964–1996.

result, there are 30 instances in table 3 in which the partial sum of squares for the female indicator differs insignificantly from zero, and the partial sums of squares for an income variable is significant and greater than the partial sum of squares for the female indicator. In 24 of these instances, the difference is greater than a factor ten. Thus the explanatory power of income relative to that of sex is greater when incomes are measured using interaction variables suggested by the theory in section II than when incomes are measured as current incomes interacted with the single/married dummy.

IV. TENTATIVE ANALYSIS OF TIME-SERIES RELATIONSHIPS BETWEEN SINGLE RATES AND REDISTRIBUTIONAL SPENDING

Given male/female earnings differences, the findings above suggest that increasing singleness since about 1950 may have acted to increase differences by sex in attitudes toward redistributional spending.[18] Because women are a majority, this may in turn help explain why some empirical studies have failed to find much of an effect of welfare on single parenthood (Moffitt, 1992). As noted at the outset, the issue is the extents to which singleness causes welfare and to which welfare causes singleness. To shed brief light on the issue, we ask whether single rates Granger cause various forms of Federal redistributional spending and whether Federal redistributional spending Granger causes single rates. These tests are atheoretical and simplistic, of course, but provide exact statements about how well lagged single rates predict welfare spending and how well lagged welfare spending predicts single rates.

[18] A force that operates in the opposite direction is that female incomes have been increasing over time relative to male incomes. For instance, the female/male ratio of hours-adjusted median weekly wage and salary income for 25-34 year-olds increased from around 0.68 in the early 1970s to 0.76 in 1987 (Goldin, 1990).

(In a regression of the current value of one variable on lagged values of itself and another variable, the null is that coefficients on lagged values of the other variable are zero.)

Results of tests for spending on AFDC/TANF, food and nutrition, housing assistance, Medicaid, Medicare, Social Security and other retirement, and Unemployment Insurance are in table 4. Spending is measured both as annual amounts and as annual amounts divided by gross domestic product.[19] As measures of the single rate, we consider both an overall single household rate defined as the number of unmarried households divided by the total number of households, and a female single household rate defined as the number of unmarried-female households divided by the total number of households.[20] The tests employ five years of lags.

Of the spending categories considered, the most redistribution from men to women may occur with AFDC/TANF, food and nutrition, housing assistance, and Medicaid. In no instance can we conclude at a ten percent confidence level that spending on one of these four categories Granger causes a single rate, and in all instances but two (AFDC/TANF not divided by GDP), we can conclude that single rates Granger cause spending. Results are similar but not as strong for the other three spending categories (Medicare, Social Security, and Unemployment Insurance). Thus the stronger statistical causation is that changing singleness partly explains changing welfare levels, and not vice verse.

[19] Source: Office of Management and Budget, Budget for Fiscal Year 2000, Historical Tables 10.1 and 11.3 (online).

[20] Source: Census table HH-1 (online)

SUMMARY

Empirical analysis of polling data supports the idea that women desire greater redistributional spending than men. Most of this difference is explained by income differences. Interestingly, however, a greater relative female desire for redistributional spending persists and is sometimes significant even when controls for income are added. This might occur if women are more altruistic than men. The timing of welfare spending and singleness also suggests that increased singleness over time may explain rising welfare spending better than rising welfare spending explains increased singleness.

APPENDIX: CONSTRUCTION OF INTERACTION VARIABLES

Income Profiles

Denote the age-income profile estimated from SIPP data by f. Consider an individual sampled in the NES in year μ. Let age denote the individual's age in NES data in year μ and let $\tau = 0, \ldots, 84 - age$ index the current and future years of the individual's lifetime so the individual's age in year $\mu + \tau$ is $age + \tau$. (We truncate at age 84 because the mortality tables we apply only go to age 84; this means that we treat people as dying on their 85th birthday.) Similarly let sex, $race$, and $education$ denote the individual's sex, race, and education in the year-μ NES data. Race is reported as white, black, and other (which are only 1.5 percent of total NES observations). Education is one of five classes. We suppress notationally the dependence of age, sex, $race$, and $education$ on μ; the dependence should be understood.

If an individual were sampled by the NES in 1994 (the average year over which f is estimated), the individual's predicted income in year $1994 + \tau$ for $\tau = 0, \ldots, 84 - age$ would be $I_{n_\tau, \tau} = f(age + \tau, sex, race, education, n_\tau)$, where n_τ is the individual's marital status in $\mu + \tau$ $(= 1994 + \tau)$. Because we consider individuals sampled by the NES in years before 1994, we use census/CPS data to adjust for the generally lower incomes that prevailed before 1994. Adjustments are made separately for men and women and by race to account for differing relative incomes by sex and by race. Because available census data do not provide incomes for all three racial categories, we calculate adjustments only for whites and blacks, and apply the white adjustment factors to individuals in the racial group "other." The adjustment factor for year μ, sex s, and race ρ = white or black is denoted $\gamma_{\mu, s, \rho}$, which is calculated as the ratio of the group's median income in μ divided by its median income in 1994. For an individual sampled in year $\mu < 1994$, the individual's predicted income in year

$1 + \tau$ for $\tau = 0; \ldots; 84 - age$ is then $I_{n_\tau;\tau} = \gamma_{1;sex;race} f(age + \tau; sex; race; education; n_\tau)$:

Marital Status Transition Probabilities

From each marital status, it is possible to transit into only one other status; specifically, the only possible transitions are from never married to married, from married to widowed or divorced, and from widowed or divorced to married. This means that it suffices notationally to consider only the probability of remaining in the same marital status. Let g denote this probability function as estimated from SIPP data over 1994-1995.

If an individual were sampled by the NES in 1994, we would take the probability the individual remains in marital status n during year $1994 + \tau$ for $\tau = 0; \ldots; 84 - age$ as depending on demographic variables and on the individual's predicted income τ years after 1994 according to $p_{n;\tau} = g(age + \tau; sex; race; education; I_{n_\tau;\tau}; n_\tau)$. Because we consider individuals sampled by the NES in years before 1994, we adjust transition probabilities to reflect the generally greater marriage rates that prevailed before 1994. Adjustments are made separately by sex and race. As above, we calculate adjustments only for whites and blacks, and apply the white adjustment factors to individuals in the racial group "other."

Adjustments are by linear interpolation. For each sex-race combination with $s = m; f$ and $\rho =$ black, white, we estimate the marriage rate, defined as the number of married individuals divided by the total number of individuals, as a linear function of time and a constant. From the estimated function, we find a "base" year $B_{s;\rho}$ defined as the year in which the predicted marriage rate for the sex-race combination equals one. We then calculate adjustment factors $\varphi_{n;s;\rho}$ under the assumption that in year $B_{s;\rho}$, the probability that a never-married or widowed or divorced individual of sex s and race ρ would remain unmarried was zero, and the probability that a married individual of sex s and race ρ would remain married was one. Specifically, for an individual sampled in year $\mu < 1994$; the predicted probability of remaining in marital status n during year $\mu + \tau$ for $\tau = 0; ...; 84 - age$ is $p_{n;\tau} = \varphi_{n;sex;race}g(age + \tau; sex; race; education; I_{n_\tau;\tau}; n_\tau)$, where $\varphi_{1;sex;race} = \varphi_{3;sex;race} = [(\mu - B_{sex;race})/(1994 - B_{sex;race})]$ and $\varphi_{2;sex;race} = [(1994 - \mu)/(1994 - B_{sex;race})]$.

The estimated probabilities $p_{n;\tau}$ of remaining in a marital status recursively determine the probabilities $\pi_{n_\tau;\tau}$ that the individual is in marital status n_τ in year $\mu + \tau$ for $\tau = 0; ...; 84 - age$: An individual in marital status n in the NES data has $\pi_{n;0} = 1$ and $\pi_{n';0} = 0$ for each $n' \neq n$: The recursive probabilities that the individual is in each marital status in future year $\tau = 1; ...; 84$ are then

$$\begin{aligned}
\pi_{1;\tau} &= \pi_{1;\tau-1}p_{1;\tau-1};\\
\pi_{2;\tau} &= \pi_{2;\tau-1}p_{2;\tau-1} + \pi_{1;\tau-1}(1 - p_{1;\tau-1}) + \pi_{3;\tau-1}(1 - p_{3;\tau-1});\\
\pi_{3;\tau} &= \pi_{3;\tau-1}p_{3;\tau-1} + \pi_{2;\tau-1}(1 - p_{2;\tau-1}):
\end{aligned}$$

Interactions

Interaction terms $\pi_n I_n$ for $n = 1; 2; 3$ are the discounted sums of present values over each year of an individual's remaining lifetime of the product of the expected proportion of the year the individual will be in marital status n; and the individual's expected future income in the year conditional on being in marital status n:

$$\pi_n I_n = \sum_{\tau=0}^{84-age} \pi_{n;\tau} I_{n;\tau}(1 + d_{age+\tau})^{\tau} / \sum_{\tau=0}^{84-age}(1 + d_{age+\tau})^{\tau}; \qquad n = 1; 2; 3;$$

where d_a is a discount rate equal to the mortality rate at age a; and the denominator is expected future years of life. The mortality tables we use are differentiated by sex and by race (black, or white plus other), so we differentiate the d_a by sex and race.[21]

The interaction variables we include in the stage-III regressions are $\pi_1 I_1$, $\pi_2 I_2$, and $\pi_3 I_3$, which are computed from available NES data on the individual's age, race, education, current marital status, and current income plus the functions f and g estimated from the SIPP.

Marital status portions $\pi_n = \sum_{\tau=0}^{84-age} \pi_{n,\tau}(1+d_{age+\tau})^{\tau} / \sum_{\tau=0}^{84-age}(1+d_{age+\tau})^{\tau}$ for $n = 1, 2, 3$ might also be included in the regressions. This would seem to allow individuals with low expected future income in a status to want greater redistributional spending the greater the probability of being in the status. With negative coefficients on interaction terms and without inclusion of π_1, π_2, and π_3, a greater probability of being in a status with low expected future income would appear to reduce estimated favorability toward redistributional spending, contrary to theoretical prediction.

Inclusion of the π_n as regressors is impractical as an econometric matter because it leads to a regressor set with substantial multicollinearity as evaluated by variance inflation factor tests. On the other hand, the three marital status portions π_1, π_2, and π_3 sum to one, so including them in a regression under a constraint that the three have the same coefficient would simply add a constant to the regression. Under this constraint, omission of the π_n does not rule out individuals with low expected future income in a status wanting greater redistributional spending the greater the probability of being in the status.

[21] Mortality data are from the National Center for Health Statistics (1996) and apply to mortality in 1996. In preliminary calculations, we scaled these rates crudely to fit conditions in earlier years, but the adjustment factors were so small that there seemed little point in making the adjustments.

REFERENCES

[1] Akerlof, George, Janet Yellen, and Michael Katz, "An Analysis of Out-of-Wedlock Childbearing in the United States," Quarterly Journal of Economics 111, May 1996, 277-317.

[2] Andreoni, James and Lise Vesterlund, "Which is the Fair Sex? Gender Differences in Altruism," working paper, 1998.

[3] Apps, Patrica and Ray Rees, "Taxation and the Household," Journal of Public Economics, 35, April 1988, 335-69.

[4] Bolton, Gary and Elena Katok, "An Experimental Test for Gender Differences in Beneficent Behavior," Economic Letters, 48, 1995, 287-92.

[5] Brosio, Giorgio and Carla Marchese, "The Growth of Government under Different Redistributive Regimes: A Long Term Study of the Italian Case," Public Finance Quarterly, 16, Oct. 1988, 439-63.

[6] Brown-Kruse, Jamie and David Hummels, "Gender Effects in Laboratory Public Goods Contribution: Do Individuals Put Their Money Where Their Mouth is?," Journal of Economic Behavior and Organization, 22, 1993, 255-67.

[7] Browning, Martin, François Bourguignon, Pierre-André Chiappori, and Valérie Lechene, "Incomes and Outcomes: A Structural Model of Intrahousehold Allocation," Journal of Political Economy, 102, Dec. 1994, 1067-96.

[8] Census Bureau, Historical Statistics of the US, 1790-1970, Washington: USGPO, 1975.

[9] Chiappori, Pierre-André, "Rational Household Labor Supply," Econometrica, 56, Jan. 1988, 63-90.

[10] ————, "Collective Labor Supply and Welfare," Journal of Political Economy, 100, June 1992, 437-67.

[11] Eagly, A.H., "The Science and Politics of Comparing Women and Men," American Psychologist, 50, 1995, 145-58.

[12] Eckel, Catherine and Philip Grossman, "Are Women Less Selfish than Men?: Evidence from Dictator Experiments," Economic Journal, 108, May 1988, 726-735.

[13] Council of Economic Advisers, Economic Report of the President. Washington, US-GPO, 1986, 1997.

[14] Goldin, Claudia, Understanding the Gender Gap, New York: Oxford Univ. Press, 1990.

[15] Gray, Jeffrey, "Divorce-Law Changes, Household Bargaining, and Married Women's Labor Supply," American Economic Review, 88, June 1998, 628-42.

[16] Gronau, Reuben, "Sex-related Wage Differentials and Women's Interrupted Careers—The Chicken or the Egg," Journal of Labor Economics, 6, July 1988, 277-301.

[17] Hoddinott, John and Lawrence Haddad, "Does Female Income Share Influence Household Expenditures? Evidence from Côte D'Ivoire," Oxford Bulletin of Economics and Statistics, 57, Feb. 1995, 77-96.

[18] Husted, Thomas and Lawrence Kenny, "The Effect of the Expansion of the Voting Franchise on the Size of Government," Journal of Political Economy, 105, Feb. 1997, 54-82.

[19] Inman, Robert, "Testing Political Economy's 'as if' Proposition: Is the Median Voter Really Decisive?" Public Choice, 33, 1978, 45-65.

[20] Korenman, Sanders and David Neumark, “Marriage, Motherhood, and Wages" Journal of Human Resources, 27, Spring 1992, 233-55.

[21] Kaufmann, Karen and John Petrocik, “The Revenge of the Soccer Moms? Gender as a Cleavage in American Politics," working paper, 1997.

[22] Lott, John and Lawrence Kenny, “How Dramatically Did Women's Suffrage Change the Size and Scope of Government?," working paper, 1998.

[23] Lundberg, Shelly and Robert Pollak, “Separate Spheres Bargaining and the Marriage Market," Journal of Political Economy, 101, Dec. 1993, 988-1010.

[24] ————, Robert Pollak, and Terence Wales, “Do Husbands and Wives Pool Their Resources? Evidence from the United Kingdom Child Benefit," Journal of Human Resources, 32, Summer 1997, 463-80.

[25] Manser, Marilyn and Murray Brown, “Marriage and Household Decision Making: A Bargaining Analysis," International Economic Review, 21, Feb. 1980, 31-44.

[26] McElroy, Marjorie and Mary Horney, “Nash-Bargained Household Decisions: Toward a Generalization of the Theory of Demand," International Economic Review, 22, June 1981, 333-49.

[27] Meltzer, Allan and Scott Richard, “Why Government Grows (and Grows) in a Democracy," Public Interest, 52, Summer 1978, 111-18.

[28] ————, “A Rational Theory of the Size of Government," Journal of Political Economy 89, Oct. 1981, 914-27.

[29] Mincer, Jacob and Solomon Polachek, “Family Investment in Human Capital: Earnings of Women," Journal of Political Economy, 82, Mar. 1974, S76-108

[30] Moffitt, Robert, "Incentive Effects of the U.S. Welfare System: A Review," Journal of Economic Literature, 30, March 1992, 1-61.

[31] Mueller, Dennis and Peter Murrell, "Interest Groups and the Size of Government," Public Choice, 48 1986, 125-45.

[32] Murray, Charles, Losing Ground, New York: Basic Books, 1984.

[33] Murrell, Peter, "The Size of Public Employment: An Empirical Study," Journal of Comparative Economics, 9, Dec. 1985, 424-45.

[34] National Center for Health Statistics, Mortality Tables, 1996.

[35] Nowell, Clifford and Sarah Tinkler, "The Influence of Gender in the Provision of a Public Good," Journal of Economic Behavior and Organization, 25, 1994, 25-36.

[36] Peltzman, Sam, "The Growth of Government," Journal of Law and Economics, 23, Oct. 1980, 209-87.

[37] Pezzin, Liliana and Barbara Schone, "The Allocation of Resources in Intergenerational Households: Adult Children and Their Elderly Parents," American Economic Review, 87, May 1997, 460-64.

[38] Roberts, Kevin, "Voting over Income Tax Schedules," Journal of Public Economics, 8, Dec. 1977, 329-40.

[39] Romer, Thomas, "Individual Welfare, Majority Voting, and the Properties of a Linear Income Tax," Journal of Public Economics, 4, Feb. 1975, 163-85.