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DAUGHTERS AND LEFT-WING VOTING

Andrew J. Oswald and Nattavudh Powdthavee*

Abstract—What determines human beings' political preferences? Using nationally representative longitudinal data, we show that having daughters makes people more likely to vote for left-wing political parties. Having sons leads people to favor right-wing parties. The paper checks that our result is not an artifact of family stopping rules, discusses the predictions from a simple economic model, and tests for possible reverse causality.

1. Introduction

IN remarkable research, the sociologist Rebecca Warner and the economist Ebonya Washington have shown that the gender of a person's children seems to influence the attitudes and actions of the parent.

Warner (1991) and Warner and Steel (1999) study American and Canadian mothers and fathers.[1] The authors' key finding is that support for policies designed to address gender equity is greater among parents with daughters. This result emerges particularly strongly for fathers. Because parents invest a significant amount of themselves in their children, the authors argue, the anticipated and actual struggles that offspring face, and the public policies that tackle those, matter to those parents. In the words of Warner and Steel (1999), "Child rearing might provide a mechanism for social change whereby fathers' connection with their daughters undermines . . . patriarchy." The authors demonstrate that people who parent only daughters are more likely to hold feminist views (for example, to favor affirmative action). By collecting data on the voting records of U.S. congresspersons, Washington (2004) is able to go beyond this. She provides persuasive evidence that congresspersons with female children tend to vote liberally on reproductive rights issues such as teen access to contraceptives. In a revision, Washington (2008) argues for a wider result: that the congresspersons vote more liberally on a range of issues, such as flexibility for working families and tax-free education. Her data, compiled partly but not wholly from voting record scores compiled by the three interest groups of the National Organization for Women, the American Association of University Women, and the National Right to Life Coalition, cover a cross-section of 828 members of four congresses of the U.S. House of Representatives for the years 1997 to 2004.[2] As her final sentence puts it: "Not only should we consider the influence that parents have on children's behavior, but we should acknowledge that influence may flow from child to parent" (Washington, 2008).

In this paper, we use nationally representative random samples of men and women to generalize these results to voting for entire political parties.[3] We document evidence that having daughters leads people to be more sympathetic to left-wing parties. Giving birth to sons, by contrast, seems to make people more likely to vote for a right-wing party. Our data, which are primarily from Great Britain, are longitudinal. We also report corroborative results for a German panel. Access to longitudinal information gives us the opportunity, one denied to previous researchers, to observe people both before and after they have a new child of either gender. We can thereby test for political switching.

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[1] Kamo and Warner (1997) explore the same phenomenon in Japanese data.

[2] The interest groups choose a subset of votes in each session and base the score on the percentage of time that the congressperson votes in agreement with that group's position. They generally choose only twenty of the thousand votes that are held each congress. Ebonya Washington uses all votes that split along party lines during the time period—over 2,000 votes in total. Hence, Washington uses many votes that are not covered by the interest groups.

[3] Our work was stimulated by hearing, in 2005, Ebonya Washington present the results of Washington (2004).

Although panel data cannot resolve every difficulty of establishing cause-and-effect relationships, they allow sharper testing than can simple cross-section data.

Following the earlier literature, we think of the gender of a child arriving in the household as a kind of exogenous event. Hence, we have the character of an experiment, where "nature randomly assigns the child gender" (Washington, 2008). It is then possible to study what happens after a new child enters a household, and in particular to see whether girl babies and boy babies have different observable consequences. Consistent with the idea of causality flowing from the gender of children on to later parental attitudes, we find that when compared to the year before the birth, men and women alter their political opinions. Daughters tilt their parents to the left; sons tilt them to the right.

A difficulty for all analysis of this sort is the possibility of endogenous family stopping rules. The problem is that certain kinds of voters may choose to cease having offspring after they achieve some desired gender mix within their children, thereby spuriously creating a form of reverse causality where attitudes determine the gender pattern in the children. For example, imagine that people with right-leaning attitudes tend to stop having children after a baby girl is born, while left-leaning people stop after a baby boy is born. Then there will emerge a positive association between right-wingness among parents and statistically disproportionate percentages of sons. The reason is that the only families with long strings of daughters are the left-wing parents, and the only families with long strings of sons are the right-wing parents. Nevertheless, we can solve this by looking at, say, the gender of the first child who is born.[4] Whatever one's stopping rule as a parent, one starts with some initial baby, and unlike the later composition of the family, the gender of that first baby is uncontrollable (given no selective abortion). Hence we report, later in the paper, results for firstborns alone.

A lucid overview of much of the research in this field is provided in Lundberg (2005). The research literature finds, for example, that the gender of children appears to affect both labor supply decisions and parents' attitudes to their own roles in the family.[5] Moreover, female politicians raise different questions in political debates than men. The interesting recent work by Campbell (2004) documents systematic gender differences in modern British political attitudes. The author tabulates answers given in the 2001 British Election Survey. She shows that the single most important concern to males is low taxes. For females, by contrast, it is the quality of the National Health Service. Norris (2002) studies the gradual shift to the left of women in Britain's politics since World War II. More broadly, our paper is relevant to the ideas of Benabou and Tirole (2003) on parental-child interactions, and it fits within work on the nature of endogenous preferences (see, for instance, Bowles 1998).

Political institutions vary from one nation to another, so we are not sure how far these results will hold across countries. However, because of their statistical robustness and the generality of the issues, we would conjecture that a version of the same results will be found more widely in international panel data on voting (such data sets are currently rare).[6]

The next section of the paper sets out a (stylized) model in which it is rational for male and female parents gradually to alter their voting preferences. Our framework has an economic flavor. What happens behind the formal analytics is that because, by assumption, there is pay discrimination against women and females derive greater utility from public goods like community safety, it transpires that unmarried women are intrinsically more left wing than unmarried men.[7] When compared to males, women prefer a larger supply of the public good and a greater tax rate on income: the reason is that their marginal utility from the first is relatively high and the tax penalty they face from the latter relatively low. As men acquire female children, however, those men gradually shift their political stance and become more sympathetic to the "female" desire for a steeper income tax schedule and a larger amount of the public good, so they become more left wing. Similarly, a mother with many sons becomes sympathetic to the "male" case for lower taxes and a smaller supply of public goods and becomes more right wing. In practice, these forces may operate at a subconscious level. Our paper assumes, following the tradition of economic modeling, that people optimize as if they are conscious of deeper motives.

II. Analytical Framework

Assume a world in which people earn real income y and there is an amount of public good denoted P. The public good—it might be thought of as the safety of the community or the quality of the environment—is funded out of tax revenue. There is a single tax rate, t, levied on personal income. Assume the political shade of government in this world can be captured by a single variable, r, the shade of "red" of this society.[8]

[4] This argument generalizes to the nth child; it does not merely hold good for the first child.

[5] Work on gender in a variety of such settings has been done by Angrist and Evans (1998), Ben-Porath and Welch (1976), Bird (2005), Butcher and Case (1994), Chattopadhyay and Duflo (2004), Edlund (1999), Kohler, Behrman, and Skytthe (2005), Lundberg and Rose (2002), Norris (2004), Morgan, Lye, and Condran (1988), Oswald and Powdthavee (2008a), and Peresie (2005).

[6] In 2007, Andrew Leigh from the Australian National University wrote to us to say that he had managed to replicate a version of our finding on Australian microdata: see Leigh (2008).

[7] Another case might be that of state pensions and medical care, which, because females live longer than males, are of natural particular concern to women.

[8] We use red in the historical sense that goes back at least to the era of Karl Marx, not in the sense used in recent U.S. Democrat and Republican conventions.

Assume the existence of a monotonic relationship $P(t)$ between the supply of the public good and the tax rate. This is increasing and differentiable; greater income taxes lead to a larger supply of the public good. Define a left-wing society, with a high value of r, as one that provides a relatively large amount of the public good and funds this with a relatively high tax rate on income. Right-wing societies, by contrast, have low P and low t. Let the tax rate be $t = t(r)$, and assume $t(r)$ is increasing, monotonic, and differentiable. Write the amount of the public good

$$P = P(t(r)) = p(r), \tag{1}$$

namely, as a reduced-form function of the political shade of the society.

Consider an unmarried male who has no children. Assume he has separable utility function

$$V = v(P) + y(1 - t) \tag{2}$$

where the function $v(P)$ captures the utility from the public good, and $v(.)$ is differentiable, increasing, and strictly concave.

In choosing his society's optimal political color, r, this male voter balances a desire for low taxes with a desire for the public good. An unmarried male's utility maximization decision is the choice of the level of r that maximizes

$$V = v(p(r)) + y(1 - t(r)) \tag{3}$$

so that

$$\frac{\partial V}{\partial r} = v'(p(r))p'(r) - yt'(r) = 0, \tag{4}$$

after assuming, as will be done throughout this paper, that the citizen's maximand $V(r)$ is well behaved.

Now consider an unmarried female voter. In this world, a childless woman's utility function is assumed to take the form

$$U = (1 + \alpha)v(p(r)) + y(1 - \delta)(1 - t), \tag{5}$$

where a nonnegative parameter, α, captures any extra relative weight that females put on the public good P relative to the males, and another nonnegative parameter, δ, is the degree of pay discrimination, if any, within the society. These seem the relevant characteristics to explore. We later examine the effects of variations in these. A woman's optimal shade of political red is not identical to a man's. Hers is given by

$$\frac{\partial U}{\partial r} = (1 + \alpha)v'(p(r))p'(r) - y(1 - \delta)t'(r) = 0, \tag{6}$$

which is usefully written as

$$v'(p(r))p'(r) - yt'(r) = -\delta yt'(r) - \alpha v'(p(r))p'(r) \tag{7}$$

and contrasted with the condition in the male equation in equation (4). This leads to:

Proposition 1. *Unmarried women's voting preferences lie strictly to the left of unmarried men.*

The function U is increasing and concave; the right-hand-side term of equation (7) is negative; hence the optimal political shade of red, r^*, is higher among females than males. Other results follow:

Proposition 2.

(i) The greater is their income, y, the less left wing are unmarried individuals (of either sex).
(ii) The greater is discrimination, the more left wing are unmarried females.
(iii) The greater is females' weight on P, the more left-wing are unmarried females.

Consider income, y. For men, the sign of the cross-partial of the maximand with respect to r and y is given by the term

$$-t'(r) < 0 \tag{8}$$

and for women by

$$-(1 - \delta)t'(r) < 0, \tag{9}$$

which establishes the early part of the proposition. The others also follow from the cross-partials of equation (5).[9]

For married people, assume that, where h is a weight less than unity, when in a couple the maximand of a person is instead the convex combination

$$W = h[v(p(r)) + y(1 - t(r)] + (1 - h) \times [(1 + \alpha)v(p(r)) + y(1 - \delta)(1 - t)] \tag{10}$$

in which the values of W will later be denoted for the case of married men by $W = V^m$ and for married women by $W = U^w$. At an optimum, the equivalent to equations (4) and (6) is

$$v'(p(r))p'(r) - yt'(r) = (h - 1)[\delta yt'(r) + \alpha v'(p(r))p'(r)], \tag{11}$$

where this parameterization imposes the same value of weight h on men and women. An alternative assumption is that people put their own utility above that of their spouse.

[9] The first-order condition for maximizing $J(x, a)$ is $J_x = 0$. Around that turning point, $J_{xx}dx + J_{xa}da = 0$, which can be written simply as $dx/da = -J_{xa}/J_{xx}$. But J_{xx} is negative by the second-order condition for a maximum. Hence the sign of the comparative static result dx/da is determined solely by the sign of the cross-partial J_{xa}.

In such a case, define that weight for males as h^m and the weight for females as h^f, with for men $h^m \geq h^f$. As h lies in the unit interval, the right-hand side of this equation is greater than that in equation (7) for unmarried women and smaller than that in equation (4) for unmarried men. As $W(.)$ is concave:

Proposition 3.

(i) Unmarried women are more left wing than married people, who in turn are more left wing than unmarried men.
(ii) In the case where $h^m > h^f$ holds as a strict inequality, the married women are strictly more left wing than the married men.

How then might parents be affected by having male and female offspring? Take a man with f female children and m male children. One assumption is that he might put some weight on his own preferences and some weight on the preferences of his offspring. A strict Darwinian might even argue that he would be put complete weight on his children's utilities, but that is an extreme. Hence, define an equivalent to the earlier V function—this time for a married man with children. Let the preferences of a father be represented by the new utility function

$$V^c = \gamma V^m + (1 - \gamma)[fU + mV] \tag{12}$$

in which V^m is utility of a married man without children (which takes a value determined from equation 11), the assigned weight on own utility is γ and that on the children's utility is an assigned weight $1 - \gamma$. Here the individual acts somewhat like a welfare planner (and if all weights are 0.5, it is exactly family utilitarianism). For simplicity, equation (12) imposes a steady state in utilities and ignores discounting. Male children are assigned within their male parent's maximand the same utility function as that of childless males, V, and female children are assigned the utility function of childless females, U. This might seem myopic, because parents may bear in mind that their own children will reproduce, but such extra terms eventually disappear algebraically.

Put more intuitively, a father takes on some of the preferences of his female offspring. For their sake, if only subconsciously, he begins to vote accordingly. The optimal political shade of the father is given by

$$\frac{\partial V^c}{\partial r} = \gamma \frac{\partial V^m}{\partial r} + (1 - \gamma)m \frac{\partial V}{\partial r} + (1 - \gamma) f \frac{\partial U}{\partial r} = 0, \tag{13}$$

where, as before, we are concentrating on the case of interior optima. Under these assumptions:

Proposition 4. *The more daughters a person has, the more he or she votes to the left. The more sons the person has, the more he or she votes to the right.*

As the number of daughters, f, rises, the optimal political shade of red of this individual, r^*, also increases. The sign of dr^*/df is given by the sign of the partial derivative of equation (13) with respect to female children, f. That cross-partial's sign is determined solely by

$$(1 - \gamma) \frac{\partial U}{\partial r}. \tag{14}$$

Unmarried women are the most left wing of the four groups. Hence around the r^*, that is optimal for married men, the derivative $\partial U/\partial r$ is strictly positive. Similar results apply for females; the algebra is omitted.

This framework is a deliberately simple one. It is not designed to explain details of the political world. Our aim instead is to try to contribute to thinking about possible sources of gender differences—to allow us to say something about averages within a population.

III. Empirical Testing

The paper proposes an empirical exploration of these ideas. The main source used in the analysis is the British Household Panel Survey (BHPS). This is a nationally representative random sample of British households, containing over 10,000 adult individuals, conducted between September and Christmas of each year from 1991 (see Taylor et al., 2002). Respondents are interviewed in successive waves; households that move to a new residence are interviewed at their new location; if an individual splits off from the original household, the adult members of the new household are also interviewed. Children are interviewed once they reach 11 years old. The sample has remained representative of the British population since the early 1990s. Numbers of adult children are not recorded fully in the data set, so this paper focuses on offspring who live at home. Relatively little research appears to have been done on political preferences in BHPS data. Some exceptions are Sanders and Brynin (1999) and the work of Johnston, Jones, et al. (2005) and Johnston, Sarker, et al. (2005), but these do not explore the influence of children on their parents' politics.

A chief focus here is on which political party an individual supports. The exact question used (AV8 in the survey) is as follows, with, for illustration, British people's mean answers given for the year 1991: "Which party do you regard yourself as being closer to than the others?"

Conservative (3,110 individuals, 46.3%)
Labour (2,707 individuals, 40.3%)
Liberal Democrats (698 individuals, 10.4%)
Scottish National Party (91 individuals, 1.4%)
Plaid Cymru (7 individuals, 0.1%)

Green Party (76 individuals, 1.1%)
Other parties (22 individuals, 0.3%)
Other answer (7 individuals, 0.1%)
Don't know/no answer (3,546 individuals)

In the later analysis, we measure "left wing" by using individuals' expressed support for the Labour Party or Liberal Democrat Party. We measure "right wing" by using expressed support for the Conservative Party. Because they are hard to classify on a political left-right scale and numbers are small, individual voters for other political parties are eventually eliminated from the data set. Clearly it is not possible in this way, or any simple way, to do justice to the complexities of human beings' political preferences. A trade-off exists between tractability and generality. Nevertheless, there is agreement that Labour is to the left (it has traditionally promoted socialist ideas) and the Conservatives are to the right (it has promoted the free market). The Liberal Democrats are more centrist, and thus in between the two larger parties, but have often been seen as closer to the left than the right. The Labour and Liberal Democrats are combined only for simplicity; the results of our paper do not rest on such an aggregation. Later analysis will not distinguish between whether the individual survey respondent is literally happier when his or her political party is in power, though it is natural to assume so (and Di Tella & MacCulloch, 2005, find evidence for that in Western Europe). It is clear from these data, moreover, that many voters say they are undecided. We assume in the paper that this is inevitable in empirical work on political preference and, for simplicity, later generally leave aside these observations.

Before we move to a formal analysis of the data set, it is natural to mention the political complexion of current female members of Parliament in Great Britain. At the time of writing, there are 127 women in the House of Commons, the main legislative body. Of those, 17 are Conservative. More than 100 of the women are Labour or Liberal Democrat. Such a highly unequal division between right wing and left wing among female politicians contrasts with an approximately equal split among male politicians. This fact suggests some kind of connection between gender and political beliefs.

While the theoretical model may apply generally, this paper will be silent empirically on a large range of nations. Women in the United States, for instance, are known to be more pro-Democrat in general than men, and this tendency has grown over the past few decades (Edlund & Pande, 2002; Box-Steffensmeier, De Boef, & Lin 2004). Greenberg (1998) concludes: "There is no question that, in general, women are more likely than men to favor activist government, the sort of agenda traditionally associated with the Democratic Party." Nevertheless, it is not clear how, for example, the principles of Britain's Labour Party should be viewed relative to those of the U.S. Democratic Party. In modern data, Inglehart and Norris (1999) find some evidence of a more widespread female tendency to vote left in other countries (although in older data, this was less common). Further research will be needed to compare the paper's patterns with non-British ones. Moreover, the paper is unable to say how long-standing the patterns in the data have been; it is known that in the 1950s, both British and American females were more right wing than they are today, and it is not easy to speculate on any role for child gender during that era. Here our analysis's contribution is inevitably weak.

In this British data set, which spans the years 1991 to 2005, we examine the voting intentions of adults. There are approximately 80,000 observations on political party preferences. These are longitudinal data (this is an unbalanced panel), and there is much stability, year-on-year, in a person's political views. Approximately two-thirds of people in this sample express a preference for the left, in our terminology, which we take as synonymous with either Labour or Liberal Democrat. In the raw data, the split between men and women is similar (approximately two-thirds of the population being left leaning), although this makes no allowance for different ages or cohort effects. As we shall see later, unmarried men vote to the right and unmarried women to the left.

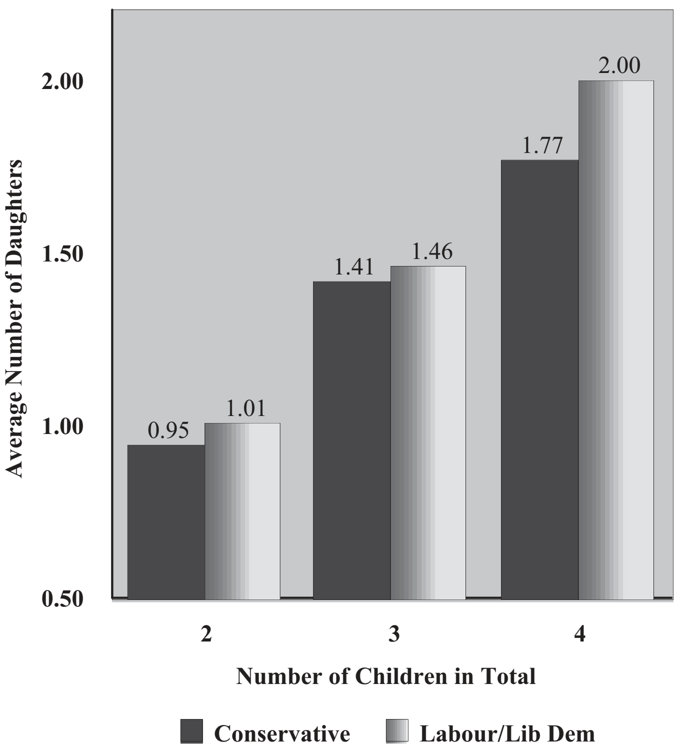
Means and standard deviations for the raw BHPS data are provided in the appendix. The mean number of children in a household is 0.70, with a standard deviation of 1.02. Approximately 3% of the sample (the denominator here includes people out of the labor force) say they are unemployed; 7% are self-employed; 7% look after the home; 25% are retired; 47% are males; 59% are married; 9% are widowed; 10% have as their highest qualification a university bachelor's degree, while 2% have a master's or doctorate. Mean age is 48 years old. These personal characteristics are viewed here as additional influences beyond the gender effect studied in the earlier section's formal model.

As suggested by the theoretical framework, we ask whether the gender of a person's children makes a difference to that individual's political preferences. Because the sex of babies is random, the gender mix of the family might potentially be viewed as exogenous. But such an argument is not quite complete. Family size is chosen. Some families will for personal and cultural reasons have different "stopping rules" (perhaps go-on-until-a-boy-is-born-and-then-stop, and so on). Nevertheless, the individual gender of a child is approximately out of a parent's control. One feasible exception is that in principle, some babies might be aborted because of their sex, as measured by a scan in the womb. However, abortion is legal in Britain only where the mother's physical or mental health is at stake. A referee has pointed out that there is some chance women pregnant with a child of the "wrong" sex would calculatingly declare that they cannot handle another baby and seek an abortion; so if right-wing parents aborted daughters disproportionately, then we would get signs of our later pattern. However, there are two objections to this interpretation. First, it is perhaps not easy to believe that such an activity (if discovered, the doctors would be struck off—banned for life—for taking part in it) could go on at the large level required to generate the strong daughters–left-voting correlation that we observe in the data. Second, and more important, it is hard to see how it could be an explanation longitudinally for the key correlation. Because we see our effect in panel data, not just in cross-sections, the selective-abortion thesis would require that for some reason, a switch to the left among parents preceded the selected birth of a female.

The paper's emphasis is on the correlation between the gender composition of offspring and the voting preferences of parents. In the formal analysis, we generally combine natural children and any stepchildren of the head of the household (that is, other stepchildren are omitted). But we check what happens when the types are separated.

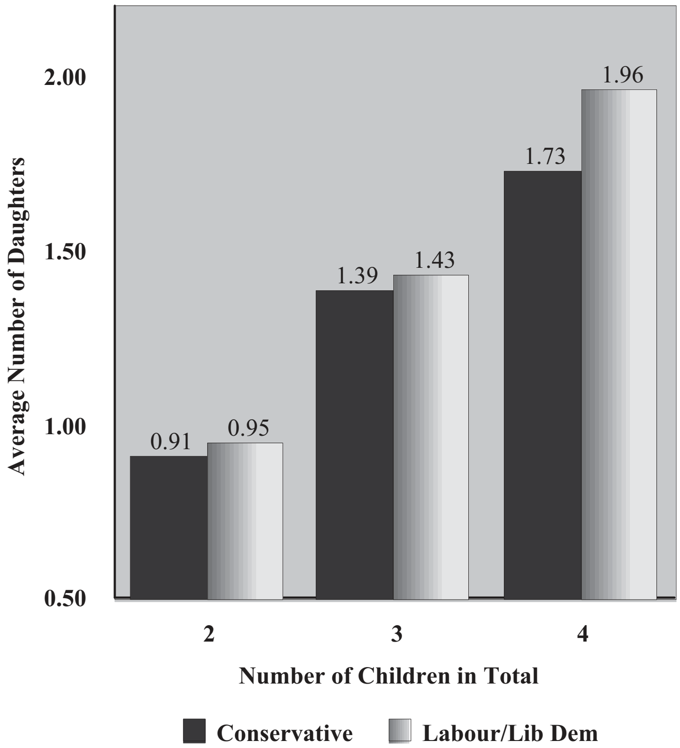
Figure 1 gives a first flavor of the result. Its columns show, for randomly selected British voters, that left-wing voters have systematically higher proportions of female children. Among families with two children, the mean number of daughters among left voters exceeds the mean number of sons; the same is true for people with three children; and the same holds among those with four children. Figure 1 includes children who are on the household roster (so those children who are dependents aged 0 to 15 *and* children who are over 15 but remain at home). It does not count children who have left the household. When, as a check, the sample is restricted solely to those aged under 16, in figure 2, the same pattern emerges. Because size of family is endogenous and is likely to be correlated with people's characteristics and innate preferences, the comparisons here are deliberately across groups with equal numbers of offspring. This result should nonetheless be treated cautiously. Once the standard errors are adjusted for clustering, it is not possible to reject the null hypothesis that, for any number of children c, the number of daughters equals the number of sons for supporters of each political wing. Even so, such a test throws away statistical information because it does not pool the findings from all six columns in, for example, figure 2. We return later to other tests of statistical significance. Figure 3 switches to a graph in which political preference is on the y-axis. The comparison in this case is between people with only three sons and those with only three daughters. Of those with sons, 66% vote for the Labour Party and the Liberal Democrat Party. Among

Figure 1.—Proportion of Daughters and Voting Preferences in Great Britain, 1991–2004



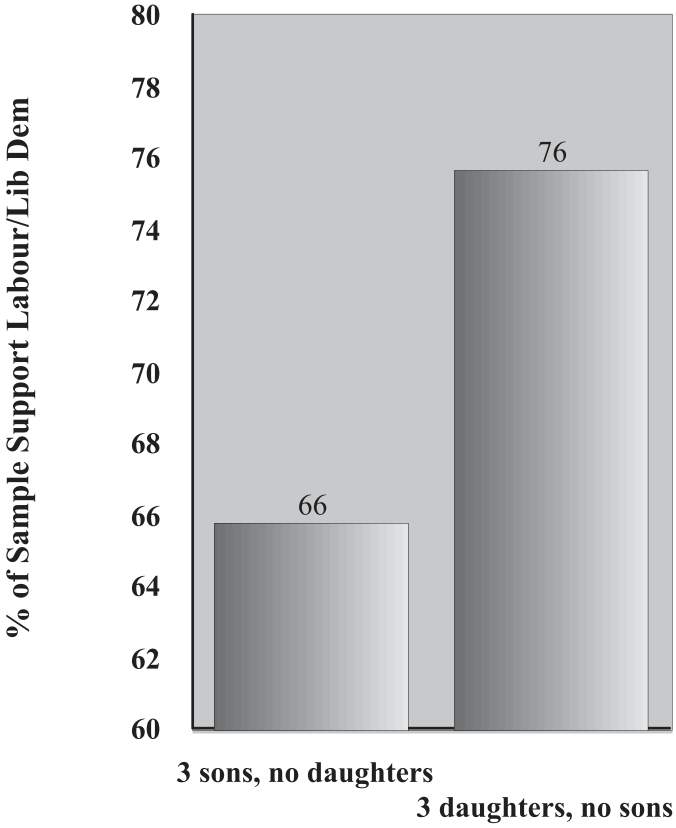
Note: There were 3,859 (7,453) observations preferring Conservative (Labour/Liberal Democrats [Lib/Dems]) over other parties with two children; 1,171 (2,534) observations preferring Conservatives (Labour/Lib Dems) with three children; and 217 (601) observations preferring Conservatives (Labour/Lib Dems) with four children. The t-test statistics (p-value) of whether the mean number of daughters between the two groups is equal are −2.535 [0.000] (N of children = 2), −3.999 [0.000] (N of children = 3), and −2.577 [0.000] (N of children = 4). The adjusted t-test statistics (p-value) for clustering by personal identification of whether the mean number of daughters between the two groups is equal are −0.822 [0.411] (N of children = 2), −1.354 [0.176] (N of children = 3), and −0.844 [0.377] (N of children = 4).

Figure 2.—Proportion of Daughters (Aged under 16) and Voting Preferences in Great Britain, 1991–2004



Note: There were 2,581 (5,233) observations preferring Conservative (Labour/Lib Dems) over other parties with two children aged under 16; 778 (1,682) observations preferring Conservatives (Labour/Lib Dems) with three children aged under 16; and 115 (376) observations preferring Conservatives (Labour/Lib Dems) with four children aged under 16. The t-test statistics (p-value) of whether the mean number of daughters aged under 16 between the two groups is equal are −2.199 [0.000] (N of children = 2), −1.914 [0.056] (N of children = 3), and −3.293 [0.000] (N of children = 4). The adjusted t-test statistics [p-value] for clustering by personal identification of whether the mean number of daughters between the two groups is equal are −0.980 [0.164] (N of children = 2), −0.924 [0.356] (N of children = 3), and −1.687 [0.097] (N of children = 4).

FIGURE 3.—PROPORTION OF PEOPLE SUPPORTING PARTIES BY THE GENDER OF THEIR CHILDREN



Note: There were 503 observations with three sons and no daughters and 473 observations with three daughters and no sons. The t-test statistics (p-value) of whether the proportion of people supporting either Labour or Liberal Democrats between the two groups is equal is −3.035 [0.002]. The adjusted t-test statistics (p-value) for clustering by personal identification of whether the proportion of people supporting either Labour or Liberal Democrats between the two groups is equal is −1.531 [0.127].

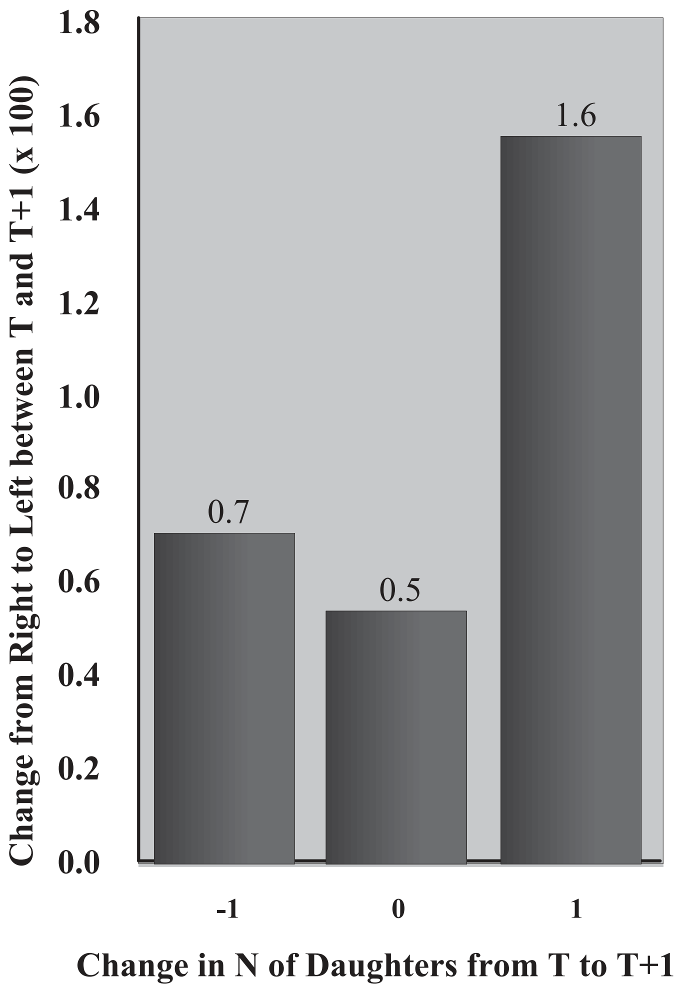
those with daughters, 76% vote Labour or Liberal Democrat.

The advantage of longitudinal data is that we can examine the interesting case of political switchers. The remaining figures (from figure 4 onward) measure on the y-axis the size of the leftward move. As people have their daughters and sons, we can observe what happens. In the third column of figure 4, those who have an additional daughter—there are approximately 1,000 such households—shift during that year disproportionately to the left. The y-axis of figure 4 gives the proportion of changes in the voting preference from $t - 1$ to t. Define a value of $0 =$ no change in the voting preference. If it takes a value of -1, then it means the person voted for Labour/Liberal Democrat at $t - 1$ and then switched to Conservative at t, and vice versa for the value of 1 (switching from Conservative at $t - 1$ to Labour/Liberal Democrat at t). Because most people did not alter their vote, the means of the number of daughters here are quite small. Hence in the figures we multiply the numbers by 100.

The effect captured in figure 4 is suggestive but not statistically different from 0 at the 5% level. Sharper evidence is provided in figure 5. Plotted on the horizontal axis is the net growth, within the year, of daughters relative to sons. There are 1,883 observations on households with a negative net change in daughters and 1,924 with a positive net change. As is clear visually from figure 5, there is a strong association between having daughters and moving leftward politically. In this case, the effect is significant at the 1% level.

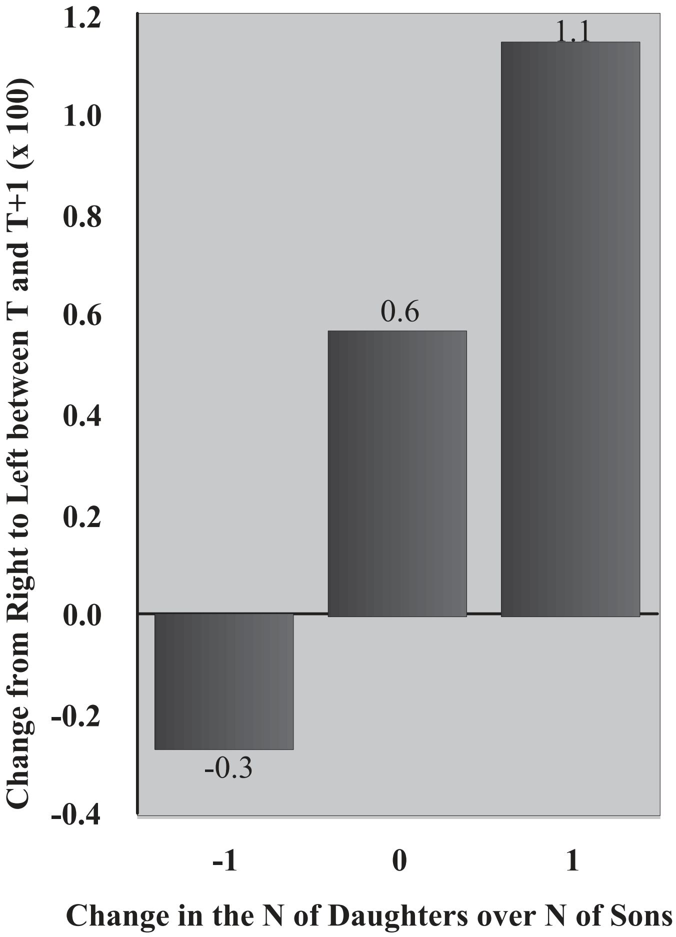
Family structure is not exogenous. Arguably, therefore, a good experiment stems from the impact of the gender of, say, a firstborn child; this test is not subject to bias from family stopping rules. Thus firstborn children are studied in figure 6. Once again, acquiring a daughter is associated with people turning toward the Labour and Liberal Democratic parties, and having a son with parents tilting instead to the Conservative party. The size of the effect is approximately the same as earlier.

FIGURE 4.—PROPORTION OF PEOPLE SWITCHING POLITICAL PARTY AFFILIATION AND CHANGE IN THE NUMBER OF DAUGHTERS FROM T TO $T + 1$



Note: There were 993 observations with at least one daughter leaving the household roster between T and $T + 1$ (net change $= -1$), 45,214 observations with no change in the number of daughters, and 967 observations with at least one additional daughter in the household roster (net change $= +1$). The adjusted t-test statistics (p-value) for clustering by personal identification of whether the average change in the voting preference between the two groups (-1 and $+1$) is the same is −1.078 [0.281].

FIGURE 5.—POLITICAL PARTY AFFILIATION SWITCHING AND CHANGE IN THE NUMBER OF DAUGHTERS OVER THE NUMBER OF SONS FROM T TO $T + 1$



Note: There were 1,883 observations with a negative net change in the number of daughters over the number of sons between T and $T + 1$, 43,259 observations with no change in the number of daughters relative to the number of sons, and 1,924 observations with a positive net change in the number of daughters over the number of sons between T and $T + 1$. The adjusted t-test statistics (p-value) for clustering by personal identification of whether the average change in the voting preference between the two groups (-1 and $+1$) is the same is -2.649 [0.008].

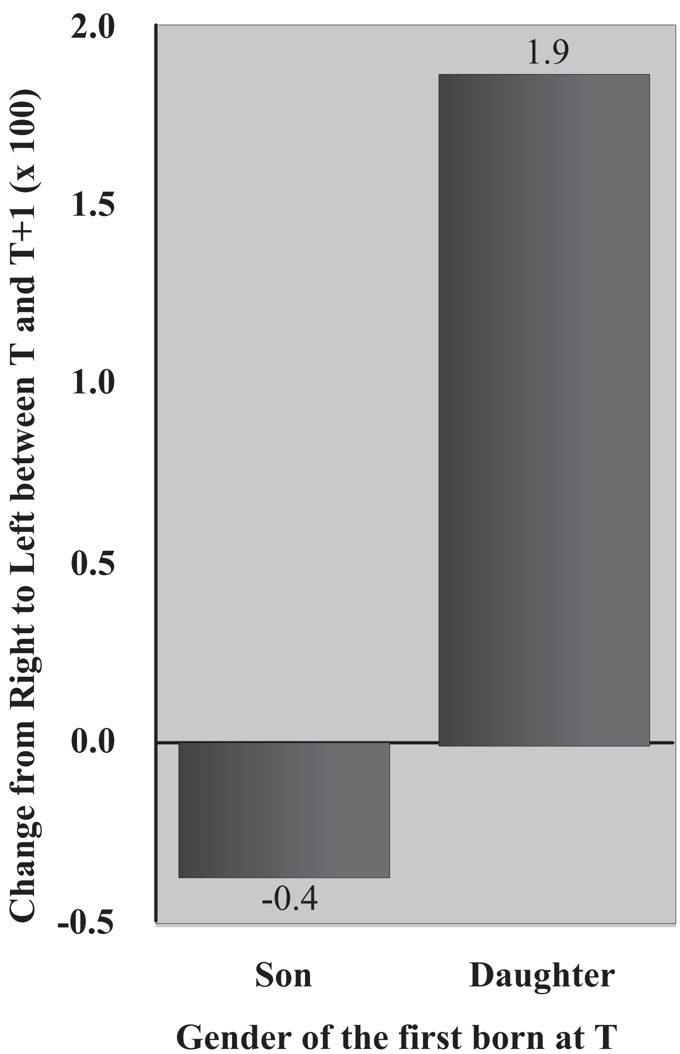
Figure 7 sets out an equivalent finding for Germany. Here the data source is the German Socioeconomic Panel, a larger panel than the British BHPS data set. Switching toward the left (detailed definitions are given later) is once again disproportionately preceded by having a daughter, and the reverse for the arrival of a son.

To control for confounding influences, BHPS regression equation evidence is set out in table 1. This begins, in its first three columns, with elementary logit equations in which the dependent variable is a binary variable to capture voting left. The key independent variable is the number of daughters. Here we follow the empirical strategy in the work of Washington (2004): the specification allows the effect of pure family size to be held constant. When the number of children is controlled for, the coefficient on the number of daughters tells us about the proportional influence of the gender composition of offspring. Column 1 of table 1, in which only basic demographic variables are held constant, estimates the coefficient on the number of daughters at 0.100 with a standard error of 0.044. There is also an effect on left-wing voting from the age variable: older people vote to the right. Regional dummies also have strong effects (but are not reported explicitly). The poorer north of Great Britain is known to be more supportive of left-wing parties. "Wave dummies" here are year dummies for each wave of the BHPS surveys.

Column 2 of table 1 incorporates a list of extra variables: controls for marital status, income, education, employment type, and other personal characteristics. As before, there remains a positive link, with a coefficient of almost the same size, between having daughters and voting for the Labour and Liberal Democrats.

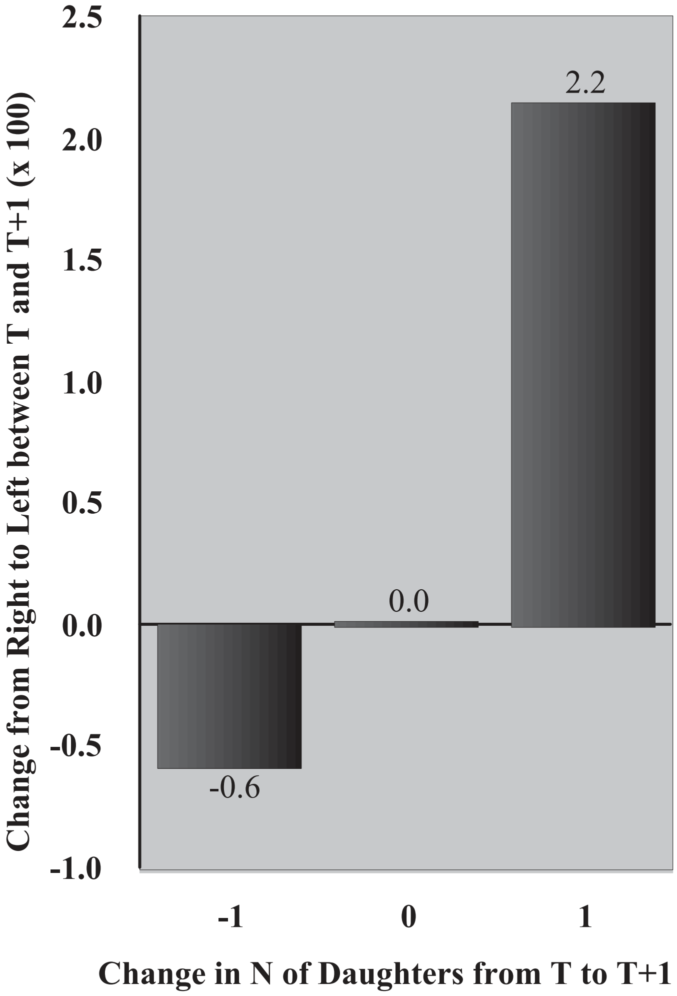
To check the theoretical framework's ideas, column 3 of table 1 introduces separate dummy variables for married male, married female, and single female. The omitted cat-

FIGURE 6.—POLITICAL PARTY AFFILIATION SWITCHING AND THE GENDER OF THE FIRSTBORN



Note: The sample is restricted to those with a firstborn at T. There are 825 sons and 804 daughters born during the BHPS sample. The adjusted t-test statistics (p-value) for clustering by personal identification of whether the average change in the voting preference between the two groups (-1 and $+1$) is the same is -2.473 [0.014].

FIGURE 7.—PROPORTION OF PEOPLE SWITCHING POLITICAL PARTY AFFILIATION AND CHANGE IN THE NUMBER OF DAUGHTERS FROM T TO $T+1$: GERMAN SOCIOECONOMIC PANEL DATA, 1985–2002



Note: There were 871 observations with at least one daughter leaving the household roster between T and $T+1$ (net change $= -1$), 46,928 observations with no change in the number of daughters, and 651 observations with at least one additional daughter in the household roster (net change $= +1$). The adjusted t-test statistics (p-value) for clustering by personal identification of whether the average change in the voting preference between the two groups (-1 and $+1$) is the same is -2.713 [0.007].

egory here is for unmarried males. Consistent with the analysis predictions, the coefficients on the three categories rise monotonically: the numbers are 0.030, 0.146, and 0.275**. This makes it possible to conclude, at the 5% level, that single women are further to the left than single men. The difference is fairly large, at approximately half the ceteris paribus cross-sectional effect of having a university degree.

Consistent with the theoretical model in the earlier part of the paper, the results of columns 2 and 3 of table 1 find that high-income people lean rightward. Highly educated people tend to be left wing; self-employed individuals tend to be (strongly) right wing. Interestingly, mental strain, as captured by the commonly used GHQ score, enters positively in this left-wing voting probability equation, as do widowed, divorced, and disabled. The construction of GHQ scores, a measure that amalgamates answers to twelve psychiatric strain questions, is described in Oswald and Powdthavee (2008b). As would be expected, other independent variables enter the political-preference equations (Alesina & La Ferrara, 2005, discuss the micro-determinants of taste for redistribution), but the paper does not explore these in detail. Many of the variables in our equations are likely to be endogenously determined, but it seems useful to observe that the daughters' effect survives their inclusion.

The last three columns of table 1 turn to fixed-effect logit estimates (denoted L-FE in the table). This has methodological advantages over the related work on cross-section data by Rebecca Warner and Ebonya Washington. For well-understood reasons, there may be omitted variables that are correlated with both voting preferences and the nature of people's families. Hence there is a case for using an estimator that can difference out the unobservable personal characteristics. Although the usual criticisms of non–fixed effects estimation are possibly less powerful in this setting (because the gender mix of the children is somewhat difficult for parents to control), it is natural to explore the structure of a fixed-effects voting equation.

Now, in column 4 of table 1, the coefficient on the daughters' variable is 0.363, with a standard error of 0.141. The coefficient is similar in columns 5 and 6, which separate into subsamples for mothers' and fathers' voting. The standard errors now weaken a little. It is not possible to reject the null hypothesis that 0.383 is equal to 0.343. Hence, the influence of child gender appears to be similar for male parents and female parents.

At the suggestion of a referee, table 2 turns to specifications in which natural daughters (those born biologically to the parent) and step-, foster, or adopted daughters are separated into two groups. The bulk of the daughters effect appears to come through the coefficient on natural daughters.

Table 3 now breaks the longitudinal data into daughters and sons "entering" and "leaving" the parental household. Although, perhaps inevitably, standard errors are not always small, the clearest effects come from children when they enter. In the third column of table 3, the coefficient on more daughters is 0.333, and that on more sons is -0.378; in each case, the null of 0 cannot quite be rejected at the 5% level (though equality of coefficients can). Some of the other coefficients move around and are poorly defined.

Table 4 checks that the paper's correlation is not being produced by reverse causality. We examine people's voting preferences before they have a child. Being left wing in time $t-1$ is not predictive within this equation of having a daughter in period t. Table 5 also checks that the main result is not produced by some unusual interaction with income. A test of the case of firstborns is also done. Table 6 demonstrates, though the size of the effective sample is inevitably reduced and the standard errors worsened, that the same tenor of results is found among parents of firstborn children.

TABLE 1.—EQUATIONS FOR THE PROBABILITY OF VOTING FOR A LEFT-WING PARTY AT TIME *T*: LOGITS AND FIXED EFFECT LOGITS, BHPS, 1991–2005

Vote Left Wing at *t*	Logit All	Logit All	Logit All	L-FE All	L-FE Female	L-FE Male
Number of all daughters	0.100**	0.093**	0.093**	0.363**	0.383*	0.343*
	(0.044)	(0.046)	(0.045)	(0.141)	(0.210)	(0.199)
Number of all children						
1	0.051	−0.047	−0.053	−0.369**	−0.308	−0.393*
	(0.048)	(0.057)	(0.057)	(0.153)	(0.215)	(0.224)
2	−0.048	−0.230***	−0.237***	−0.432*	−0.169	0.697**
	(0.048)	(0.082)	(0.082)	(0.233)	(0.331)	(0.344)
3	−0.041	0.228*	0.237*	−0.109	0.297	−0.555
	(0.103)	(0.125)	(0.125)	(0.347)	(0.517)	(0.485)
4	0.313*	−0.025	−0.038	0.318	0.900	−0.108
	(0.187)	(0.219)	(0.219)	(0.613)	(0.954)	(0.857)
5	0.230	−0.106	−0.122	−0.858	−1.630	−0.625
	(0.318)	(0.368)	(0.368)	(0.868)	(1.341)	(1.297)
6	0.145	−0.174	−0.192	−0.699	11.671	−2.239
	(0.553)	(0.646)	(0.646)	(1.290)	(366.953)	(1.658)
Married		−0.044		0.212	0.278	0.350
		(0.069)		(0.210)	(0.317)	(0.295)
Married male			0.030			
			(0.086)			
Married female			0.146			
			(0.111)			
Single female			0.275**			
			(0.111)			
Cohabited		0.086	0.215**	0.134	0.230	0.187
		(0.070)	(0.090)	(0.182)	(0.287)	(0.247)
Widowed		0.174*	0.344***	0.409	0.187	1.102**
		(0.102)	(0.123)	(0.322)	(0.438)	(0.549)
Divorced		0.258**	0.405***	0.338	0.504	0.245
		(0.106)	(0.122)	(0.286)	(0.425)	(0.406)
Separated		0.068	0.212	0.020	0.175	−0.049
		(0.137)	(0.148)	(0.327)	(0.483)	(0.461)
Male	−0.0217	−0.010	0.106			
	(0.040)	(0.043)	(0.077)			
Age	−0.015**	0.014*	0.014*			
	(0.005)	(0.007)	(0.007)			
$Age^2/100$	0.004	−0.023***	−0.023***	−0.030	−0.026	−0.029
	(0.005)	(0.007)	(0.007)	(0.025)	(0.034)	(0.038)
Roman Catholic		0.157**	0.158**	0.036	−0.024	0.129
		(0.077)	(0.077)	(0.221)	(0.299)	(0.335)
Church of England		−0.575***	−0.574***	−0.032	0.009	−0.125
		(0.047)	(0.047)	(0.119)	(0.169)	(0.172)
Other religion		−0.332***	−0.332***	−0.103	−0.071	−0.164
		(0.062)	(0.062)	(0.172)	(0.239)	(0.255)
Income (in £1,000)		−0.039***	−0.039***	0.007	0.004	0.008
		(0.003)	(0.003)	(0.005)	(0.006)	(0.008)
Household size		0.033	0.035*	−0.052	−0.065	−0.029
		(0.021)	(0.021)	(0.057)	(0.087)	(0.078)
First degree		0.466***	0.467***	−0.161	−0.238	−0.188
		(0.073)	(0.073)	(0.258)	(0.355)	(0.387)
Higher degree		0.794***	0.794***	0.279	0.086	0.435
		(0.154)	(0.154)	(0.486)	(0.763)	(0.659)
Self-employed		−0.725***	−0.728***	0.034	0.158	−0.089
		(0.069)	(0.069)	(0.161)	(0.283)	(0.199)
Unemployed		0.367***	0.367***	−0.262	−0.115	−0.446*
		(0.074)	(0.074)	(0.200)	(0.317)	(0.267)
Retired		−0.041	−0.041	−0.157	0.138	0.574**
		(0.065)	(0.065)	(0.148)	(0.202)	(0.227)
Maternity leave		0.266*	0.274*	0.085	0.174	
		(0.146)	(0.146)	(0.406)	(0.415)	
Family care		−0.014	−0.009	0.118	0.170	1.543*
		(0.064)	(0.064)	(0.150)	(0.165)	(0.903)
Full-time student		−0.049	−0.056	0.318	0.517*	0.156
		(0.073)	(0.073)	(0.206)	(0.305)	(0.297)
Disabled		0.475***	0.478***	−0.613**	−0.647*	−0.849**
		(0.103)	(0.103)	(0.252)	(0.380)	(0.349)
Government training scheme		0.487*	0.495*	−0.498	1.339	−1.482**
		(0.250)	(0.250)	(0.560)	(0.861)	(0.726)
Other		−0.050	−0.050	0.304	0.618	−0.447
		(0.184)	(0.184)	(0.453)	(0.560)	(0.835)

TABLE 1.—(CONTINUED)

Vote Left Wing at t	Logit All	Logit All	Logit All	L-FE All	L-FE Female	L-FE Male
Mental distress (GHQ-12)		0.031***	0.031***	0.032***	0.031**	0.043**
		(0.005)	(0.005)	(0.011)	(0.015)	(0.019)
Constant		0.872***	0.940***			
		(0.236)	(0.249)			
Regional dummies	Yes	Yes	Yes	Yes	Yes	Yes
Wave dummies	Yes	Yes	Yes	Yes	Yes	Yes
Log likelihood	−56,863.63	−49,436.37	−49,422.74	−3,443.891	−1,784.184	−1,621.478
N	93,044	84,440	84,440	10,826	5,798	5,028

Note: The first three columns are pooled cross-section logits; they allow for clustering within person. The second set of three columns, headed L-FE, is fixed-effect logits. Reference groups: no stated religion, lower than first-degree education, employed full time, never married (and single male in the second column). Standard errors are in parentheses.
* < 10%; ** < 5%, *** < 1%.
Source: British Household Panel Study.

The emphasis so far in the paper has been on whether the null hypothesis of 0 can be rejected. How large, and therefore how significant for social science, are the effects from child gender on parental voting? The most persuasive estimates are arguably likely those in logits with fixed effects. Results for other estimators (equations available upon request) are given below as a contrast. For each daughter, holding family size constant, a parent is approximately 2 percentage points more likely to vote left as follows:

Calculated Size of the Effect from Each Extra Daughter (Percentage Increase in the Likelihood of Voting Left)

Logit with random effects: 1.8 percentage probability points
OLS with fixed effects: 1.4 percentage probability points
Logit with fixed effects: 2.7 percentage probability points

TABLE 2.—EQUATIONS FOR THE PROBABILITY OF VOTING FOR A LEFT-WING PARTY AT TIME t: VARIABLES FOR TYPE OF CHILDREN, BHPS 1991–2005

Vote Left Wing at t	L-FE All	L-FE Female	L-FE Male
Number of natural daughters	0.409***	0.387*	0.451*
	(0.147)	(0.213)	(0.216
Number of step/foster/ adopted daughters	−0.181	−0.799	−0.076
	(0.506)	(1.589)	(0.552)
Number of natural children			
1	−0.355**	−0.283	−0.390*
	(0.155)	(0.217)	(0.229)
2	−0.550**	−0.191	−0.990***
	(0.239)	(0.333)	(0.362)
3	−0.014	0.313	−0.457
	(0.365)	(0.522)	(0.533)
4	0.462	0.906	−0.001
	(0.653)	(0.955)	(0.970)
5	−0.661	−2.268	−0.645
	(0.943)	(1.674)	(1.355)
6	−0.564	12.715	−2.249
	(1.325)	(642.753)	(1.699)
Number of step/foster/ adopted children			
1	0.071	0.789	−0.16
	(0.408)	(1.281)	(0.463)
2	0.58	−10.226	0.324
	(0.677)	(2370.390)	(0.723)
3	0.805	—	0.629
	(1.278)		(1.320)
Log likelihood	−3,440.254	−1,783.632	−1,617.632
N	10,827	5,799	5,028

Note: Same controls as in table 1. Standard errors are in parentheses. In the first row, it is not possible to reject at 5% the null hypothesis that coefficient 0.387 is equal to 0.451.
* < 10%; ** < 5%, *** < 1%.
Source: BHPS.

TABLE 3.—EQUATIONS FOR THE PROBABILITY OF VOTING FOR A LEFT-WING PARTY AT TIME t: VARIABLES FOR CHILDREN LEAVING AND ENTERING THE HOUSEHOLD, BHPS 1991–2005

Vote Left Wing at t	L-FE All	L-FE All	L-FE All
Fewer daughters from $t-1$	0.172		0.233
	(0.213)		(0.216)
More daughters from $t-1$	0.392*		0.333*
	(0.186)		(0.190)
Fewer sons from $t-1$		0.110	0.076
		(0.207)	(0.210)
More sons from $t-1$		−0.398*	−0.378*
		(0.190)	(0.194)
Number of all children at $t-1$	0.174*	0.086	0.104
	(0.078)	(0.077)	(0.089)
Log likelihood	−3,276.670	−3,276.975	−3,274.717
N	10,294	10,294	10,294

Note: Same controls as in table 1. Reference groups: no change in the number of natural children from $t-1$, no change in the number of natural daughter from $t-1$. Variables such as fewer daughters include grown-up daughters leaving the home and daughter deaths and any other reason (we do not know the exact reasons). Standard errors are in parentheses.
* < 10%; ** < 5%, *** < 1%.
Source: BHPS.

TABLE 4.—CHECKING FOR THE EFFECTS OF LEFT-WING VOTING AT $T-1$ ON THE PROBABILITY OF HAVING A FIRST DAUGHTER (LOGIT)

First Daughter	All
Vote Left Wing at $t-1$	−0.099
	(0.164)
Personal controls	Yes
Regional dummies	Yes
Wave dummies	Yes
Log likelihood	−539.5585
N	830

Note: Dependent variable: First daughter = 1, first son = 0. Standard errors are in parentheses. Sample contains only individuals with their first daughter (no record of other children—sons or daughters—in the household).
* < 10%; ** < 5%, *** < 1%.
Source: BHPS.

TABLE 5.—TESTING FOR INTERACTION EFFECTS BETWEEN DAUGHTERS AND INCOME

Vote Left Wing at *t*	Logit All	L-FE All
Number of all daughters	0.220***	0.494***
	(0.063)	(0.164)
Income (in £1,000)	−0.036***	0.009*
	(0.003)	(0.005)
Number of daughters × Income	−0.017***	−0.016
	(0.006)	(0.010)
Number of all children		
1	−0.029	−0.350**
	(0.057)	(0.153)
2	−0.222***	−0.429*
	(0.083)	(0.233)
3	−0.259**	−0.153
	(0.126)	(0.348)
4	−0.124	0.18
	(0.223)	(0.620)
5	−0.203	−1.051
	(0.379)	(0.875)
6	−0.414	−0.953
	(0.658)	(1.304)
Log likelihood	−49,415.17	−3,442.684
N	84,440	10,826

Note: Same controls as in table 1. Standard errors are in parentheses.
* < 10%; ** < 5%, *** < 1%.
Source: BPHS.

TABLE 6.—LEFT-WING VOTING EQUATIONS AND THE FIRSTBORN

Vote Left Wing at *t*	L-FE All	L-FE All
Firstborn daughter	0.191	—
	(0.341)	
	—	−0.292
Firstborn son		(0.319)
Log likelihood	−2,553.213	−2,552.951
N	7,904	7,904

Note: Same controls as in table 1. Standard errors are in parentheses.
* < 10%; ** < 5%, *** < 1%.
Source: BHPS.

The numbers in the case of firstborn children, as in table 6, are similar in size, at slightly more than 2 percentage points per daughter.

It seems interesting to go a little further. In the spirit of the research literature described earlier, and especially Washington (2004), we can ask empirically whether other attitudes are altered by having daughters rather than sons. Table 7 is an attempt to shed light on this. It uses answers to various attitudinal questions from the panel; these are coded on a 5-point scale so that, for simplicity here, cardinality is assumed. Each of the four columns in table 7 is a GLS regression equation, with a different dependent variable each time. The number of daughters enters negatively in a "Cohabitation is not all right" equation; positively in a "Homosexuality is not wrong" equation, although in this instance the standard error is not well determined; positively in a "It is not true that a husband should earn while the wife stays at home" equation; and positively in a "It is not true that children need a father as much as mother" equation. Following the questions discussed in Johnston and Pattie (2000), it would be possible to look more deeply into attitudinal issues, but we have not done so in this paper. There are no questions in the British Household Panel survey on the area of life covered particularly by the work of Washington (2004), namely, that of people's attitudes toward women's issues such as abortion, but, like her, we find here that the gender mix of children is correlated with parents' social attitudes toward family matters.

A number of robustness checks, some suggested by seminar participants in presentations of the paper, were undertaken. By using a set of dummy variables, we have found that the influence of the number of daughters seems to be monotonic up to around five children (where, because of the rarity in modern data of large families, the size of sample

TABLE 7.—ATTITUDES REGRESSIONS (RANDOM EFFECTS)

	Cohabiting Is Not All Right			Homosexuality Is Not Wrong			Not True That: Husband Should Earn, Wife Should Stay at Home			Not True That: Children Need Father as Much as Mother		
	All	Female	Male	All	Female	Male	All	Female	Male	All	Female	Male
Number of all daughters	−0.020*	−0.014	−0.026	0.016	0.007	0.025	0.027***	0.013	0.040***	0.020**	0.019	0.016
	(0.011)	(0.014)	(0.016)	(.013)	(0.017)	(0.021)	(0.010)	(0.013)	(0.015)	(0.008)	(0.012)	(0.012)
Number of all children												
1	−0.033**	−0.035*	−0.025	0.013	−0.023	0.031	−0.031**	−0.045***	−0.036*	0.063***	0.114***	−0.042***
	(0.013)	(0.018)	(0.020)	(0.016)	(0.021)	(0.026)	(0.012)	(0.016)	(0.019)	(0.011)	(0.015)	(0.015)
2	−0.051***	−0.063**	−0.026	0.030	−0.017	0.051	−0.063***	−0.073***	−0.081***	0.063***	0.132***	−0.070***
	(0.020)	(0.026)	(0.030)	(0.024)	(0.030)	(0.038)	(0.018)	(0.024)	(0.027)	(0.015)	(0.022)	(0.021)
3	−0.045	−0.057	−0.015	0.056	0.019	0.050	−0.091***	−0.090**	−0.127***	0.064***	0.123***	−0.056*
	(0.029)	(0.039)	(0.044)	(0.035)	(0.045)	(0.056)	(0.027)	(0.035)	(0.041)	(0.023)	(0.032)	(0.031)
4	0.020	0.013	0.045	0.090	0.000	0.157*	−0.231***	−0.230***	−0.280***	0.055	0.132***	−0.093*
	(0.046)	(0.061)	(0.070)	(0.056)	(0.070)	(0.090)	(0.042)	(0.056)	(0.065)	(0.036)	(0.051)	(0.050)
5	−0.075	−0.093	−0.038	0.126	0.021	0.200	−0.312***	−0.316***	−0.374***	−0.013	0.045	−0.142*
	(0.076)	(0.100)	(0.116)	(0.092)	(0.115)	(0.149)	(0.072)	(0.095)	(0.111)	(0.062)	(0.088)	(0.086)
6	0.235**	0.357**	0.066	−0.389***	−0.270	−0.676***	−0.560***	−0.596***	−0.656***	−0.096	−0.061	−0.211*
	(0.119)	(0.155)	(0.187)	(0.145)	(01.79)	(0.239)	(0.104)	(0.141)	(0.154)	(0.091)	(0.132)	(0.122)
R^2 (within)	0.002	0.002	0.003	0.006	0.006	0.008	0.003	0.006	0.001	0.004	0.008	0.003
N	53,599	29,212	24,390	53,419	29,098	24,324	74,694	40,588	34,108	74,715	40,607	34,110

Note: Standard errors are in parentheses. Responses are coded: 1 = strongly agree to 5 = strongly disagree. Standard errors are in parentheses.
* < 10%; ** < 5%, *** < 1%.
Source: BHPS.

becomes small). This issue seems important but demands a larger data set if it is to be examined truly persuasively. Splitting the number-of-daughters variable into two age classes does not alter the main conclusion. Once again, it seems likely that a larger data set would be needed if the aim is to find out whether it is young children, rather than older children, who are disproportionately responsible for the shaping of political attitudes.

As a further check on reverse causality, we tested extensively for signs of the Trivers-Willard hypothesis (1973). This is the idea that causality might flow from parental characteristics or the environment on to the gender of babies being born: "In species with a long period of parental investment after birth of young, one might expect biases in parental behavior toward offspring of different sex, according to the parental condition; parents in better condition would be expected to show a bias toward male offspring" (p. 90). This is related to Bateman's principle (1948) that females invest more in offspring and therefore become the scarce resource that are competed over by males. In interesting work, Kanazawa and Vandermassen (2005) have recently proposed a generalized version of the Trivers-Willard hypothesis, which they call gTWH. Nevertheless, whatever we do, in these data a person's voting color in time t does not seem to be predictive of a new child's gender in $t + 1$. Insofar as we can tell, causality is running from the gender of the child, not toward it.

Table 8, as a further check, changes to the German Socioeconomic Panel. It suggests a similar pattern. The key coefficient, in the first column of table 8, is 0.383, with a standard error of 0.105. However, breaking the sample into female and male parents shows in the second column of table 8 that the coefficient on daughters is very poorly determined in the sample of women voters. We are not sure how to interpret this. Nevertheless, using longitudinal data from 1984 to 2003, which here provides a sample of approximately 16,000 Germans' recorded political preferences, in the full sample, we find quite strong corroborative evidence for the earlier result on the British data. We measure left-wing political preferences here as expressed support for a combination of Social Democratic Party and Free Democratic Party. The alternatives, the right wing in this classification, are the combined Christian Union and Christian Democrat parties. Other, and alternative, specifications are provided in Oswald and Powdthavee (2005b). Further discussion of the German case, and associated regression equations, is available on request. Earlier online versions of our work, with a wider range of specifications and other checks, are reported in Oswald and Powdthavee (2005a, 2006).

Table 8.—Checking the Voting Result on German Data, 1984–2003: Fixed Effect Logits

Vote Left Wing at t	All	Female	Male
Number of daughters	0.383***	−0.141	0.813***
	(0.105)	(0.160)	(0.143)
Number of all children			
1	−0.107	0.049	−0.262*
	(0.103)	(0.152)	(0.143)
2	0.025	0.552**	−0.384*
	(0.151)	(0.236)	(0.203)
3	0.105	1.102***	−0.712**
	(0.225)	(0.353)	(0.307)
4	−0.036	1.384**	−1.056**
	(0.347)	(0.549)	(0.461)
5	−1.496**	−0.159	−2.481***
	(0.644)	(0.911)	(0.930)
6	−1.314	0.310	−1.689
	(1.056)	(1.709)	(1.630)
7	10.395		8.657
	(618.331)		(590.717)
$Age^2/100$	−0.094***	−0.144***	−0.051*
	(0.019)	(0.027)	(0.027)
Log of household income	−0.073	−0.087	−0.009
	(0.063)	(0.083)	(0.099)
Number of years spent in school	0.047	−0.005	0.090*
	(0.040)	(0.067)	(0.050)
Employed full time	−0.013	−0.035	0.056
	(0.076)	(0.104)	(0.114)
Disabled	0.322***	0.074	0.534***
	(0.108)	(0.173)	(0.143)
Single	0.205	0.425	0.129
	(0.189)	(0.299)	(0.251)
Widowed	0.126	0.257	−0.158
	(0.173)	(0.229)	(0.277)
Divorced	0.052	0.110	0.037
	(0.181)	(0.266)	(0.259)
Separated	−0.053	−0.049	−0.090
	(0.208)	(0.299)	(0.302)
East Germany	1.048	0.573	−10.752
	(1.323)	(1.947)	(651.378)
Regional dummies	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes
Log likelihood	−5,451.296	−2,490.613	−2,915.163
N	16,099	7,515	8,584

Note: Reference groups: married, not disabled, and West Germany. Left wing = 1 if Social Democrats and Free Democratic Party, 0 = Christian Union and Christian Democrats.
* < 10%; ** < 5%, *** < 1%.
Source: German Socioeconomic Panel (GSOEP).

IV. Conclusion

This paper explores the roots of political preference. Our work builds on, and attempts to generalize, innovative research by Rebecca Warner on children's influence on parents' views on feminist issues and affirmative action and by Ebonya Washington on children's influence on congresspersons' views on issues such as reproductive rights.

The paper finds evidence that having daughters makes people more sympathetic to left-wing parties. Acquiring sons, by contrast, makes individuals more right wing. Ceteris paribus, in our panel data, every extra daughter (or son) leads a person to be approximately 2 percentage points more likely to vote left (or right). Our data come principally from Great Britain, but we show that the basic result can be replicated for German microdata. The checks described in the paper suggest that the result seems not to be an artifact of family stopping rules, or of some unusual biological causal chain from politics to the later sex of offspring, or of selective abortion.[10]

A long-standing idea in Western society is that parents influence the behavior and psychology of their children. Following previous research, the analysis suggests the reverse idea: that children shape their parents.[11] This paper, which could be seen as a study of endogenous preferences, also sets out a formal framework with an economic flavor. The model describes a world in which, because of wage discrimination and different female preferences over public goods, parents rationally tilt to the left if they have daughters and to the right if they have sons. Our analytical framework has this prediction. Whether the model's ideas are truly the right explanation for the pattern we witness in the data seems an important topic for further research.

[10] Nor could we find longitudinal evidence in the data set that couples stay together more when a son is born (Dahl & Moretti, 2005), so it is apparently not a by-product of that.

[11] In passing, a (tentative) conjecture can be made. It is that left-wing individuals may be disproportionately people who come from extended families where, over recent past generations, many females have been born. Having many daughters pushes parents to the left; by the time the children are old enough to acquire a political sense, their parents have passed on some of those left-wing opinions to their sons and daughters; if those children then go on to have daughters themselves, those left-wing views, inherited from their parents, become strengthened among the sons and daughters of the next generation. In this way, strings of daughters through the generations might lead to left-wing families today. Strings of sons would have the opposite effect. Whether there is empirical support for this conjecture is an open question.

REFERENCES

Alesina, A., and E. La Ferrara, "Preferences for Redistribution in the Land of Opportunities," *Journal of Public Economics* 89 (2005), 897–931.

Angrist, J. D., and W. N. Evans, "Children and Their Parents' Labor Supply: Evidence from Exogenous Variation in Family Size," *American Economic Review* 88 (1998), 450–477.

Bateman, A. J. "Intra-Sexual Selection in Drosophila," *Heredity* 2 (1948), 349–368.

Benabou, R., and J. Tirole, "Intrinsic and Extrinsic Motivation," *Review of Economic Studies* 70 (2003), 489–520.

Ben-Porath, Y., and F. Welch, "Do Sex Preferences Really Matter?" *Quarterly Journal of Economics* 90 (1976), 285–307.

Bird, K., "Does It Really Matter If Men Dominate Politics?" University of Essex working paper (2005).

Bowles, S., "Endogenous Preferences: The Cultural Consequences of Markets and Other Economic Institutions," *Journal of Economic Literature* 36 (1998), 75–111.

Box-Steffensmeier, J. M., S. De Boef, and T. M. Lin, "The Dynamics of the Partisan Gender Gap," *American Political Science Review* 98 (2004), 515–528.

Butcher, K. F., and A. Case, "The Effect of Sibling Sex Composition on Women's Education and Earnings," *Quarterly Journal of Economics* 109 (1994), 531–563.

Campbell, R., "Gender, Ideology and Issue Preference: Is There Such a Thing as a Political Women's Interest in Britain?" *British Journal of Politics and International Relations* 6 (2004), 20–46.

Chattopadhyay, R., and E. Duflo, "Women as Policy Makers: Evidence from a Randomized Policy Experiment in India," *Econometrica* 72 (2004), 1409–1443.

Dahl, G. B., and B. Moretti, "The Demand for Sons," unpublished paper, University of Rochester (2005).

Di Tella, R., and R. MacCulloch, "Partisan Social Happiness," *Review of Economic Studies* 72 (2005), 367–393.

Edlund, L., "Son Preference, Sex Ratios, and Marriage Patterns," *Journal of Political Economy* 107 (1999), 1275–1304.

Edlund, L., and R. Pande, "Why Have Women Become More Left-Wing? The Political Gender Gap and the Decline in Marriage," *Quarterly Journal of Economics* 117 (2002), 917–961.

Greenberg, A., "Deconstructing the Gender Gap," paper presented at the Midwest Political Science Association Annual Meeting (April 1998).

Inglehart, R., and P. Norris, "The Developmental Theory of the Gender Gap: Women and Men's Voting Behavior in Global Perspective," JFK School of Government, Harvard University working paper (1999).

Johnston, R., K. Jones, C. Propper, R. Sarker, S. Burgess, and A. Bolster, "A Missing Level in the Analyses of British Voting Behavior: The Household as Context as Shown by Analyses of a 1992–1997 Longitudinal Survey," *Electoral Studies* 24 (2005), 201–225.

Johnston, R., and C. Pattie, "Inconsistent Individual Attitudes within Consistent Attitudinal Structures: Comments on an Important Issue Raised by John Bartle's Paper on Causal Modelling of Voting in Britain," *British Journal of Political Science* 30 (2000), 361–374.

Johnston, R., R. Sarker, K. Jones, A. Bolster, C. Propper, and S. Burgess, "Egocentric Economic Voting and Changes in Party Choice: Great Britain 1992–2001," *Journal of Elections, Public Opinion and Parties* 15 (2005), 129–144.

Kamo, Y., and R. L. Warner, "The Effect of Children's Sex on Parents' Gender-Role Attitudes: An Extension Using Japanese Data," *Journal of Comparative Family Studies* 28 (1997), 204–210.

Kanazawa, S., and G. Vandermassen, "Engineers Have More Sons, Nurses Have More Daughters: An Evolutionary Psychological Extension of Baron-Cohen's Extreme Male Brain Theory of Autism and Its Empirical Implications," *Journal of Theoretical Biology* 233 (2005), 589–599.

Kohler, H. P. J. R. Behrman, and Skytthe, A. "Partner Plus Children Equals Happiness? The Effects of Partnerships and Fertility on Well-Being," *Population and Development Review* 31 (2005), 407–422.

Leigh, A., "Do Daughters Make You Left Wing? Evidence from Australia," Australian National University mimeograph (2008).

Lundberg, S., "Sons, Daughters, and Parental Behavior," *Oxford Review of Economic Policy* 21 (2005), 340–356.

Lundberg, S., and E. Rose, "The Effects of Sons and Daughters on Men's Labor Supply and Wages," this REVIEW 84 (2002), 251–268.

Morgan, S. P., D. N. Lye, and G. A. Condran, "Sons, Daughters, and the Risk of Marital Disruption," *American Journal of Sociology* 94 (1998), 110–129.

Norris, P., "Gender and Contemporary British Politics" (pp. 38–59), in C. Hay (Ed.), in *British Politics Today* (Cambridge: Polity Press, 2002).

——— *Electoral Engineering: Voting Rules and Political Behavior* (Cambridge: Cambridge University Press, 2004).

Oswald, A. J., and N. Powdthavee, "Daughters and Left-Wing Voting," University of Warwick, working paper (2005a).

——— "Further Results on Daughters and Left-Wing Voting: Germany 1985–2002," University of Warwick, working paper (2005b).

——— "Daughters and Left-Wing Voting," IZA discussion paper 2103, (2006).

——— "Death, Happiness, and the Calculation of Compensatory Damages," *Journal of Legal Studies* 37(S2) (2008a), S217–S252.

——— "Does Happiness Adapt? Longitudinal Evidence on Disability with Implications for Economists and Judges," *Journal of Public Economics* 92 (2008b), 1061–1077.

Peresie, J. L., "Female Judges Matter: Gender and Collegial Decision-Making in the Federal Appellate Courts," *Yale Law Journal* 114 (2005), 1759–1790.

Sanders, D., and M. Brynin, "The Dynamics of Party Preference Change in Britain, 1991–1996," *Political Studies* 47 (1999), 219–239.

Taylor, M. F., J. Brice, N. Buck, and E. Prentice-Lane, *British Household Panel Survey User Manual* (Colchester: University of Essex, 2002).

Trivers, R. L., and D. E. Willard, "Natural Selection of Parental Ability to Vary the Sex Ratio of Offspring," *Science* 179 (1973), 90–92.

Warner, R. L., "Does the Sex of Your Children Matter? Support for Feminism among Women and Men in the United States and Canada," *Journal of the Marriage and Family* 53 (1991), 1051–1056.

Warner, R. L., and B. Steel, "Child Rearing as a Mechanism for Social Change: The Relationship of Child Gender to Parents' Commitment to Gender Equity," *Gender and Society* 13 (1999), 503–517.

Washington, E., "Family Matters: How Daughters Affect Their Legislator Fathers' Voting on Women's Issues," Yale University working paper (2004).

——— "Family Matters: How Daughters Affect Their Legislator Fathers' Voting on Women's Issues," Yale University working paper (2005).

——— "Family Socialization: How Daughters Affect Their Legislator Fathers' Voting on Women's Issues," *American Economic Review* 98:1 (2008), 311–332.

APPENDIX

BHPS Data Description and Summary

Variables	Description	All Mean	All SD	Males Mean	Males SD	Females Mean	Females SD
Vote left-wing parties	Political party affiliation; 0 = Conservatives (British right-wing party) 1 = Labour/Liberal Democrats (British left-wing parties)	0.651	(0.476)	0.648	(0.477)	0.653	(0.476)
Number of daughters	Number of daughters	0.330	(0.640)	0.317	(0.630)	0.341	(0.648)
Number of children	Number of children	0.704	(1.024)	0.681	(1.023)	0.725	(1.025)
Men	Gender (male = 1)	0.474	(0.499)				
Age	Age	48.460	(18.403)	47.634	(18.011)	49.205	(18.719)
$Age^2/100$	$Age^2/100$	26.870	(18.862)	25.933	(18.183)	27.716	(19.417)
Roman Catholic	Religion: Roman Catholic	0.085	(0.279)	0.077	(0.266)	0.093	(0.290)
Church of England	Religion: Church of England	0.263	(0.440)	0.221	(0.415)	0.300	(0.458)
Other religion	Religion: Other religion	0.127	(0.333)	0.109	(0.312)	0.142	(0.350)
Self-employed	Employment status, self-employed = 1	0.071	(0.256)	0.108	(0.311)	0.037	(0.188)
Unemployed	Employment status, unemployed = 1	0.032	(0.176)	0.045	(0.208)	0.020	(0.140)
Retired	Employment status, retired = 1	0.250	(0.433)	0.232	(0.422)	0.266	(0.442)
Maternity leave	Employment status, maternity leave = 1	0.011	(0.103)	0.000	(0.016)	0.020	(0.141)
Housewife/looking after home	Employment status, housewife/looking after home = 1	0.071	(0.256)	0.007	(0.082)	0.129	(0.335)
Student	Employment status, student = 1	0.043	(0.202)	0.043	(0.204)	0.042	(0.200)
Disabled	Employment status, disabled = 1	0.035	(0.184)	0.041	(0.199)	0.030	(0.170)
Government training scheme	Employment status, government training scheme = 1	0.002	(0.041)	0.002	(0.048)	0.001	(0.033)
Other employment	Employment status, other employment = 1	0.003	(0.057)	0.003	(0.052)	0.004	(0.061)
Income (*1,000)	Annual household income per capita, adjusted to CPI index (in £1,000)	9.798	(8.076)	10.236	(8.098)	9.404	(8.036)
Married	Marital status, married = 1	0.588	(0.492)	0.626	(0.484)	0.554	(0.497)
Living as a couple	Marital status, living with a partner = 1	0.090	(0.286)	0.096	(0.294)	0.085	(0.279)
Widowed	Marital status, widowed = 1	0.094	(0.291)	0.044	(0.205)	0.138	(0.345)
Divorced	Marital status, divorced = 1	0.052	(0.222)	0.041	(0.199)	0.062	(0.241)
Separated	Marital status, separated = 1	0.014	(0.118)	0.012	(0.108)	0.016	(0.126)
Education: First degree	First degree education, i.e., undergraduate levels	0.099	(0.299)	0.108	(0.311)	0.091	(0.288)
Education: Higher degree	Higher degree education, i.e., postgraduate levels	0.025	(0.155)	0.029	(0.169)	0.020	(0.141)
Mental distress (GHQ-12)	Measure of mental distress (GHQ-12)	1.883	(2.901)	1.599	(2.671)	2.141	(3.071)
Attitude questions							
Cohabitation is all right	Cohabitation is all right; 1 = strongly agree, 5 = strongly disagree	2.241	(1.000)	2.217	(0.990)	2.262	(1.008)
Homosexuality is wrong	Homosexuality is wrong; 1 = strongly agree, 5 = strongly disagree	3.193	(1.206)	2.973	(1.238)	3.388	(1.142)
Husband should earn, wife should stay at home	Husband should earn, wife should stay at home; 1 = strongly agree, 5 = strongly disagree	3.415	(1.125)	3.299	(1.117)	3.520	(1.122)
Children need father as much as mother	Children need father as much as mother; 1 = strongly disagree	1.823	(0.762)	1.766	(0.698)	1.874	(0.812)