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The $R^2 = .93$: Where Then Do They Differ? Comparing Liberal and Conservative Interest Group Ratings

Interest group ratings have long been used by social scientists to distinguish between liberal and conservative members of Congress. It is also well known that ratings by different groups are highly correlated with one another. Here, rather than focusing on the similarities between such measures, we focus on the differences between them. Although the relationship between measures is nearly linear, we find systematic robust differences between Americans for Democratic Action (ADA) and American Conservative Union (ACU) scores. Using a variety of techniques, we show that interest groups are most interested in distinguishing among their ideological friends and tend to group their ideological enemies near the bottom of the scale. Because of this, using any single interest group score to explain political phenomena (i.e., party loyalty) is likely to produce an inconsistent estimate of the impact of ideology on such phenomena. Finally, we propose and test a method that corrects for this bias.

Introduction

In order to gauge the ideology of members of Congress, scholars often rely upon ratings generated by interest groups. Two of the most commonly used ratings are those reported by the Americans for Democratic Action (ADA), a liberal interest group, and the American Conservative Union (ACU), a conservative interest group. (For

examples of those who report or compare different interest group ratings see Kritzer 1978; Ornstein, Mann, and Malbin 1996; Rohde 1994. For those who use these scores see Carson and Oppenheimer 1984; Fowler 1982; Grofman, Griffin, and Berry 1995; Grofman, Griffin, and Glazer 1990, 1991; Krehbiel 1981, 1994; Smith 1981.) Like other interest group ratings, there is a great deal of intercorrelation between these two measures.¹ The strength of this relationship is not surprising; a large portion of their utility lies in their ability to tap the liberal-conservative dimension. Thus, high ADA scores (liberalism) correspond with low ACU scores (conservatism) and vice-versa. The high correlation between these types of measures also suggests that one rating may nearly be a linear transformation of the other, and thus the two indices may be used interchangeably for social scientific purposes (Krehbiel 1981; Kritzer 1978; Rohde 1994).²

However, we must remember that groups generate these ratings for a reason: to identify their friends and expose their enemies (Fowler 1982). These rankings are based on key votes in each chamber of Congress selected by the groups themselves. Thus, different groups will select different votes for consideration. Indeed, previous research has shown only modest overlap in votes selected by different interest groups (Kritzer 1978; Poole and Daniels 1985). In 1995, for example, only one of the 20 Senate votes ranked by ADA was also chosen by the ACU. For the House in 1995, only four common votes were ranked by both groups. The groups choose different votes to rank members for the same reason they rank votes in the first place—to place like-minded members on one side of the scale and their ideological enemies on the other. As Fowler (1982, 406) states: “Thus, the net effect of the scores is to present a picture of Congress that is quite polarized.”

The high correlation between interest group rankings and each group's choice of an almost unique set of votes to rank suggests more than simply that each taps a single underlying dimension of American politics. It also implies that any variance between measures is not random; rather, it is a product of a systematic bias inherent in these types of measures. We show that each of these measures is better able to distinguish variance among those with similar ideological views than variance among those who have different ideological views. We argue here that liberal organizations like the ADA are most interested in distinguishing among members on the liberal side of the scale. While, on one hand, *they are* interested in identifying their enemies, the votes they pick are best able to distinguish among liberal members. Likewise, conservative groups such as the ACU are more interested in differences among conservative members. Thus, interest group scores tend to do

two things: (1) place their ideological enemies at or near the bottom of the scale (i.e., zero); and (2) focus on the differences among their ideological friends.

In order to demonstrate the seemingly imperceptible, yet meaningful distinctions between interest group rankings, the central hypothesis explored below is

The greater the ideological distance between the interest group and any group of congresspersons being rated, the smaller the variance in the distribution scores for that group of congresspersons.

For example, conservative interest groups (like the ACU) will tend to cluster all liberal members toward the bottom of the scale (i.e., near zero) with little variance (lower standard deviation), while their rating of conservatives will be closer to the top of the scale and more spread out.

Data Analysis

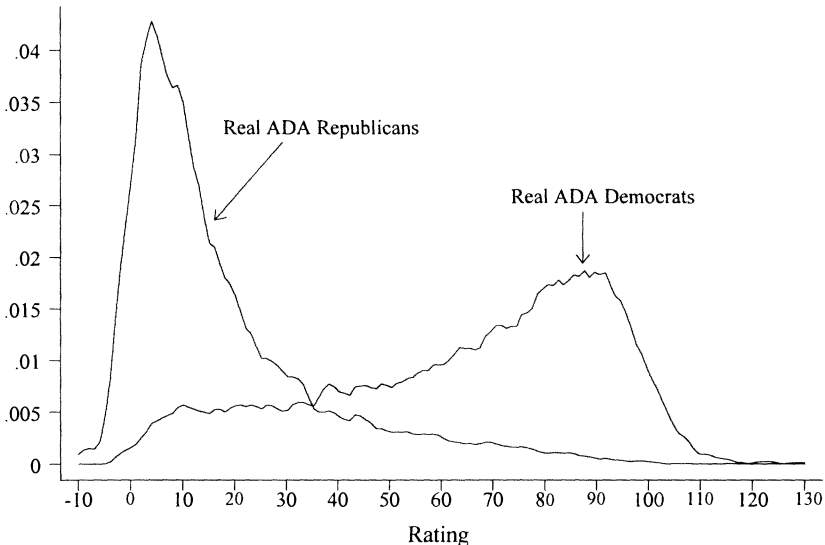
We formally test and illustrate the subtle and important differences between liberal and conservative interest group scores. Our data are ADA and ACU scores from 1971 through 1996 for both the House and the Senate.³ This sample is sufficient to ensure that our evidence is not idiosyncratic to a particular year or group of years, or to a particular chamber of Congress.

Evidence from Empirical Densities

A straightforward way to describe the ADA and ACU scores is to plot the densities of these scores by party affiliation. One limitation of such an approach is that the Republican/Democrat is a coarse division of friends/enemies. However, even such a coarse division is sufficient to establish prima facie reasonableness of our hypothesis.

In Figure 1 we have plotted the kernel density estimates (Parzen 1962) (a smoothed histogram) by party for ADA scores from 1971 to 1996 for members of the House of Representatives.⁴ Since the ADA is a liberal interest group, we expect Republican scores to be clustered around the lower bound (zero) and for the variance to be greater for Democratic members. As this figure shows, our expectations are strongly met. The Republican scores are highly peaked near zero; indeed the density of Republican scores at its mode is about 2.5 times higher than the density of Democratic scores at its mode. On the other hand, the scores of the Democrats are widely dispersed across nearly

FIGURE 1
Kernel Density Estimates by Party
for ADA Scores in the House, 1971–96



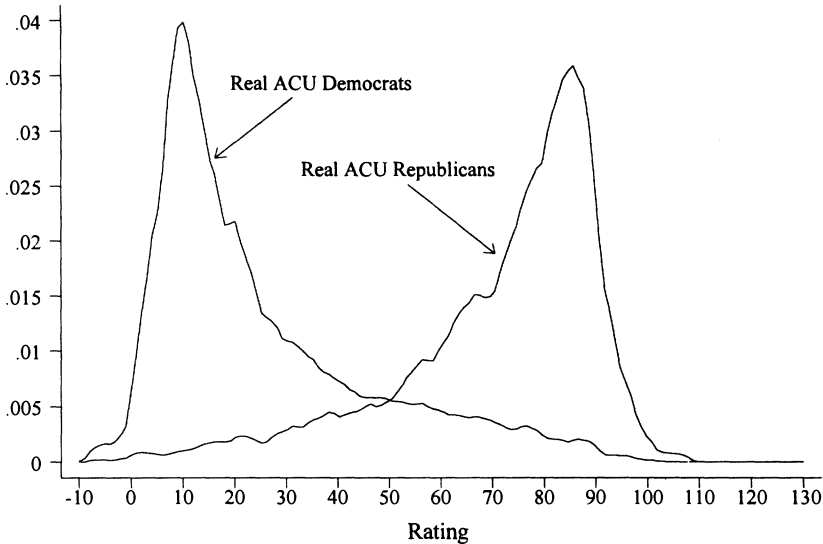
Sources: Annual Reports from the Americans for Democratic Action.

Note: For method for calculating kernel density estimates see endnote 4. Nominal scores have been converted to real scores using the method detailed in Groseclose, Levitt, and Snyder (1995).

the entire range, with a slight concentration to the right of 70.⁵ Thus, as expected the liberal interest group tends to place Republican members at the conservative end of the scale, with relatively low variance among the scores, while Democrats are spread across the continuum with a smaller concentration at the liberal end of the spectrum.⁶

In Figure 2 the same kernel densities are reported for ACU scores over the same period. Given that the ACU is a conservative group, our expectations are reversed: in this case we expect the Democrats to be tightly bunched at the liberal end near zero and the Republicans to be more dispersed across the spectrum. While the data for the ACU are not as dramatic as those for the ADA, our expectations are still generally met. Democrats are highly peaked between 0 and 10; moreover, the percentage of Democrats at the modal Democratic score is higher than the percentage of Republicans at their modal score.⁷ Again, conservative

FIGURE 2
Kernel Density Estimates by Party
for ACU Scores in the House, 1971–96



Source: American Conservative Union.

Note: For method for calculating kernel density estimates see endnote 4. Nominal scores have been converted to real scores using the method detailed in Groseclose, Levitt, and Snyder (1995).

interest groups, like liberal interest groups, tend to lump their ideological enemies at the low end of the scale and, in general, are more discriminating among their ideological friends.⁸

Another way to view the tendency of interest groups to lump their enemies at the bottom end of the scale is to ask the question if, for example, the ACU gives a member a perfect liberal score of 0, will the same member be viewed by the ADA as a “liberal hero” (i.e., have an ADA rating of 100)? For the House, there are 1,031 ACU scores of 0 for members in our sample, of these, only 18.5% also received an ADA score of 100.⁹ Thus, over 80% of the time the ADA did not similarly view these people as perfect liberals. Furthermore, only 44% of the members rated 0 by ACU had ADA scores of 90 or greater. Likewise, if the ADA saw a member as perfectly conservative (i.e., a 0 rating), only 30% of the time did the ACU rank the member a perfect 100; and

only 62% received ACU scores of 90 or higher.¹⁰ For these two interest groups, ideology is clearly in the eye of the beholder: just because a liberal group considers a member to be very conservative does not mean a conservative group will agree and vice versa. This also means that most of the differences between these scores are likely to occur at the tails of the distribution.

Our hypothesis makes further claims about the distribution of scores that can only be investigated in a bivariate context. As noted above, the relationship between ADA and ACU scores is quite strong—the R^2 from a linear regression between these two scores is .96 in the Senate and .95 in the House in 1996.¹¹ Our hypothesis suggests a restriction on the form of heteroskedasticity from a bivariate regression of ADA and ACU scores. Specifically, let y refer to an interest group score, and let x be some one-dimensional mapping of “ideology” where larger values of x imply agreement with the rater. Furthermore, assume that the rating is bounded, $\underline{x} < x < \bar{x}$ (where $\underline{x} = 0$ and $\bar{x} = 100$ in our case) and let ε be an error term that (without loss of generality) has mean zero.

$$y_{ii} = \alpha + \beta x_{ii} + \varepsilon_{ii}$$

There are several ways to characterize the bivariate relation between rating scores implied by our theory. Specifically, they involve the nature of ε . One simple representation of the type of rater bias we have outlined above is

$$\varepsilon = \gamma(\underline{x} - x)1(x < x_0) + \nu$$

where $\underline{x} < x_0 < \bar{x}$, and $0 < \gamma < 1$. Above some (possibly stochastic) threshold x_0 , distinctions are made more finely (the error is equal to ν), and below this threshold the error has an *extreme-reverting* property (this is in contrast to the more commonly encountered *mean-reverting* error). When $\gamma = 0$, this is the classical errors-in-variables case. When $\gamma = 1$, all legislators who are far enough away from the rater’s ideology are given a value 0.

In words, a rater discriminates more finely among friends than among enemies. Note that we do not require that the heteroskedasticity take *exactly* this form—indeed, our discussion suggests only that the variance of ADA scores *conditional* on a given ACU score is decreasing in the ACU score and vice versa. This simple model, however, does provide a parsimonious rendering of the type of phenomenon we expect as a consequence of our previous discussion.

Note that another view of the rating process is that there is a nonlinear mapping between ratings and ideology—the error, however,

might be homoskedastic. However, absent some a priori information on the correct class of non-linear models, this is observationally equivalent to a linear model with heteroskedasticity.¹²

We avoid modeling the bivariate relationship as a nonlinear errors-in-variable model as estimators for these models typically require strong (incredible) restrictions on the distribution of the measurement errors.¹³

Put differently, our hypothesis suggests there is a restriction on the shape of the distribution of ADA scores conditional on a given ACU score (and vice-versa). A simple way to explore this hypothesis is to examine various quantiles of the conditional density:

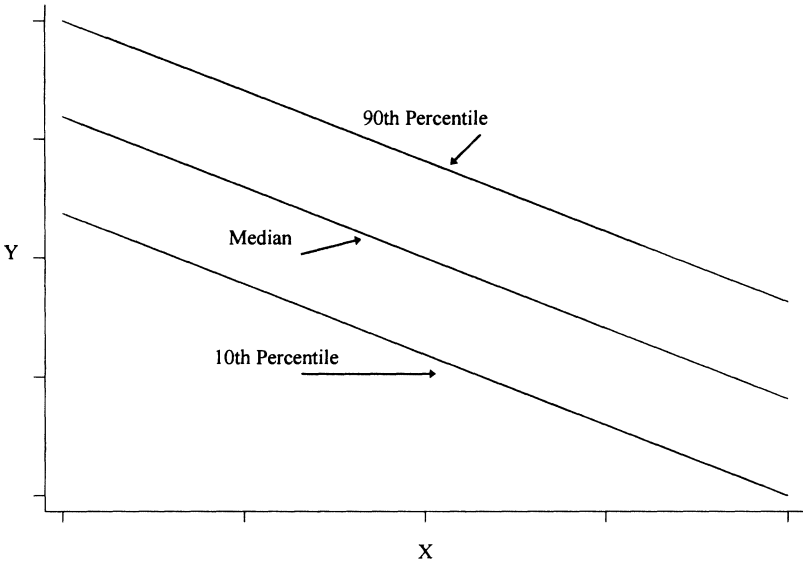
$$y_{ii}^{ADA} = \alpha + \beta y_{ii}^{ACU} + \text{error}.$$

Consider the situation when errors are homoskedastic (i.e., the variance of the errors are unrelated to the ideology of the congressperson). In this case, estimating any particular quantile of the conditional density will yield the same estimate of the slope in each equation. In Figure 3 we plot three lines from a quantile regression model with homoskedastic errors—for the median, the 10th percentile, and the 90th percentile. The key point being the 90th percentile line and the 10th percentile line are equidistant from the median line. Further, because the lines have the same slope they are parallel. Thus, if our hypothesis is *incorrect*, this is the result we should obtain.

Under our hypothesis, however, the errors will not be homoskedastic. Consider legislators with a broad range of high ACU scores: such legislators will look “broadly similar” to ADA rankers and will generally receive a narrow range of similar ADA scores. By contrast, legislators from a narrow range of low ACU scores (viewed as extremely “liberal”) will represent a broad range of more finely nuanced liberal viewpoints in the eyes of the ADA rankers.

In this case, estimates of the slope will show less steepness at lower quantiles (i.e., the 10th) and more steepness at higher quantiles (i.e., the 90th). Figure 4 plots the quantile regression¹⁴ lines for the median, 90th, and 10th percentile using ACU scores to predict ADA scores in the Senate for the period 1971–96.¹⁵ First, the three lines are clearly not parallel. Rather, the three lines form a funnel with the slope of the 90th percentile line being greater than that of median, which is greater than that for the 10th percentile. Further, the 90th and 10th percentile lines are not equidistant from the median line. Indeed the distance between the 90th and 10th percentile is greatest for low ACU scores (26.3 points apart), while it is lowest when ACU scores are

FIGURE 3
Schematic of a Homoskedastic Regression



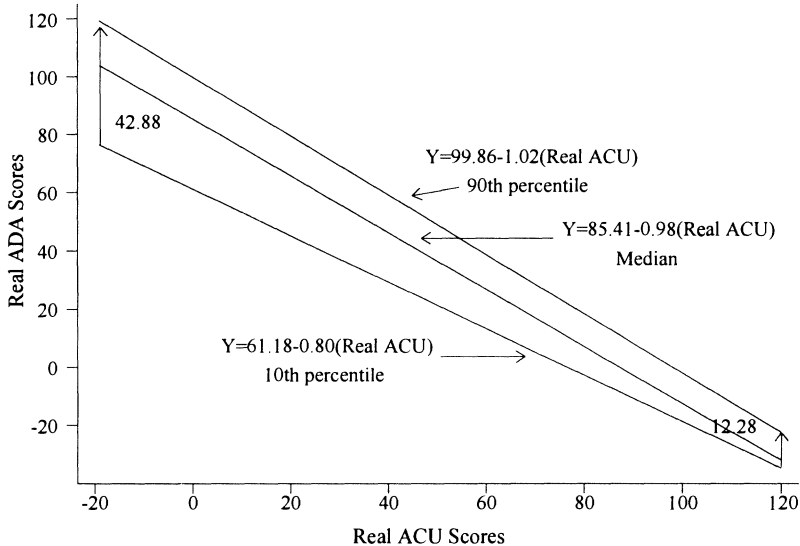
Note: The above graph depicts a hypothetical quantile regression with homoskedastic errors estimating three separate equations—one for the 90th, 50th, and 10th percentiles. See Deaton 1997.

high (20.3 points apart). Indeed it is clear from an inspection of the coefficients, that the conditional variance of the ADA scores decreases the more conservative the member. We note in passing that the mere fact that the original scores are bounded between 0 and 100 is not sufficient to produce the results we have found thus far. If one views interest group ratings as censored versions of an unbounded latent variable, one would expect heteroskedasticity, but not necessarily of the type we have found here.¹⁶

Implications for Empirical Research

Interest group scores all attempt to measure the same thing—the ideology of members of Congress. To the extent that they are all trying to measure the same thing, the problems outlined above are analogous to an intercoder reliability problem. Since these scores are tapping the

FIGURE 4
Quantile Regression Estimates for ACU
Predicting ADA Scores for the Senate, 1971–96



Sources: ADA scores are from the annual reports from Americans for Democratic Action. ACU scores are from the American Conservative Union.

Note: This figure represents the predicted values from three different quantile regressions (for the 10th, 50th, and 90th percentiles). The bootstrapped standard errors for each of the slope coefficients, derived from 1000 replications are .022 (t-statistic = -36.5) for the 10th percentile; .0082 (t-statistic = -119.67) for the median; and .0126 (t-statistic = -81.14) for the 90th percentile. For all three regressions the number of observations is 2368 from the years 1971–96. Both interest groups scores have been transformed using the method detailed in Groseclose, Levitt, and Snyder (1995).

same liberal-conservative dimension, there is a simple way to control for the bias inherent in using any one such measure to model political behavior.

Below we demonstrate one such method. Party loyalty (the proportion of times a member votes with his/her party on votes that pit a majority of one party against the majority of the other party) ought to be highly correlated with ideology. One straightforward method of modeling party loyalty as a function of ideology is:

$$\text{Loyalty} = \alpha + \beta x + \varepsilon$$

where x is ideology measured using an interest group score (ADA in this case). Assuming that ε is distributed normally with a mean of zero and a variance of σ^2 , an OLS regression will generate an unbiased estimate of β . However, suppose we observe x^* such that

$$x^* = x + v$$

where v is an error term uncorrelated with x and ε . In this case, use of the mismeasured variable x^* will yield inconsistent estimates of β in an OLS regression.¹⁷ From the above analysis we know that this is the case since ADA scores are less able to distinguish amongst the most conservative members. Specifically, we expect that the OLS model will yield slope estimates closer to zero than the true slope, which is called attenuation bias (Johnston and DiNardo 1997, 154). Thus, an OLS model will produce coefficients that are biased and inconsistent.

Under the assumption that both the ADA and ACU scores attempt to measure the same thing (albeit on different scales), and that neither is perfect, one can combine both scores to produce a better estimate of β . Under this assumption, using one measure as an instrumental variable for the other allows a consistent estimate of β to be recovered (see Fuller 1987). Table 1 reports the OLS and two-stage models for estimating party loyalty in the House for the period 1971–96 using ideology as the independent variable. Model 1 is the straightforward OLS model using ADA scores to predict party loyalty. Model 2 employs a two-stage procedure using ACU scores to instrument ADA scores in order to predict party loyalty. As expected, support for one's party in roll-call votes is linked to one's personal ideology (both models explain 76% of the variance in party loyalty across nearly 11,000 observations). Of particular interest here is the different estimates of ideology's impact on loyalty between the two models. If, as posited, there is attenuation bias in the OLS model, then the coefficient in the two-stage model ought to be larger (further from zero) than the coefficient in the OLS model. Indeed, this is the case. The value of β in the OLS model is .00741, while for the two-stage model it is .0082, an increase of nearly 11%.¹⁸ Therefore, using one measure of ideology as an instrumental variable for the other allows a consistent estimate of β to be recovered.

TABLE 1
OLS and Two-Stage Least Squares Models
of Party Loyalty in the House, 1971–96

	Model 1 (OLS)	Model 2 (2SLS)
Constant	.205*** (.002)	.1704*** (.0026)
Real ADA	.00741*** (.00004)	.0082*** (.00005)
N	10,772	10,612
R ²	.74	.73

Sources: Party loyalty 1963–95 is *Congressional Quarterly Almanac* (1962–95); 1996 is from *Congressional Quarterly Weekly Report*, December 21, 1996. ADA scores are from the annual reports from Americans for Democratic Action. ACU scores are from the American Conservative Union.

Notes: In order to model party loyalty for both parties in one equation, the party loyalty scores have been transformed for Republicans by subtracting 100 from their scores. Typically, party loyalty scores range from 0 to 100 with 100 being perfect loyalty and 0 being perfect disloyalty for both parties, thus necessitating a transformation. Both ADA and ACU scores have been adjusted using the method outlined by Groseclose, Levitt, and Snyder (1995).

*** $p < .001$.

Discussion

Interest group scores are commonly used instruments to measure the ideology of members of Congress. Nothing presented here implies that this use is inappropriate. Indeed, given the advances made by Groseclose, Levitt, and Snyder (1995), interest group scores may become all the more valuable in studying Congress. These measures are helpful in identifying conservative and liberal members of Congress, and the extremely high levels of intercorrelation between such measures suggests that they are indeed tapping the single, underlying liberal-conservative dimension that defines American politics. However, our investigation here underscores an important point: interest group ratings are better able to distinguish among their ideological friends than their ideological enemies. Thus, such measures, even the adjusted scores, are interchangeable only to a certain extent.

Other scholars have noted potential problems in using these scores similar to the ones we have examined. Snyder (1992, 340) for example, argues that because the sample of roll-call votes used to generate these ratings is biased against lopsided votes, interest group ratings “will exaggerate the degree of extremism in the distribution of legislators’ ideal points” (see also Cox and McCubbins 1993). In a similar vein, Rohde (1994, 125) argues that, notwithstanding the high correlation between these measures, different interest group scores may not be equivalent. Using hypothetical interest group ratings, he demonstrates that highly correlated scores are quite different from one another, especially with respect to the position of the median legislator. Thus, both Snyder and Rohde contend that interest group scores might not accurately reflect the actual ideology of a given member, although they employ slightly different reasoning in coming to this conclusion.

The analysis presented here generally supports these findings. However, in terms of the artificial extremism of interest group scores, our findings illustrate that the nature of extremism observed depends upon the ideology of the interest group doing the rating. For example, liberal interest groups like the ADA tend to make Republicans look more conservative by placing them all near the bottom of the scale. With regard to the equivalency of these measures, we agree that this is inherently problematic without a mechanism that corrects for the bias in these types of measures. We propose and test one such procedure above. Using a two-stage estimation process, instrumenting one interest group score with another, we are able to recover better estimates of the impact of ideology on political behavior.

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NOTES

Thanks to Tim Groseclose for his help with the transformations for “real” interest group scores. A previous version of this paper was presented at the annual meeting of the Public Choice Society in San Francisco, March 21–23, 1997.

1. They are highly *negatively* correlated since ADA rates very liberal members 100 and the ACU rates highly conservative members 100. The correlation between the two scores is $-.93$ for the Senate ($n = 2563$) and $-.92$ for the House ($n = 10,916$) for the years 1971–96.

2. Rohde (1994) makes a compelling case for not using different ratings interchangeably. Using hypothetical data, he shows that even though the correlations between scores may be very high, the distributions can differ significantly.

3. These data span the entirety of the common years for these two interest group scores.

4. The kernel density estimate of f_h of a univariate density f based on a random sample W_1, \dots, W_n is:

$$f_h(w) = \sum_{i=1}^n \frac{1}{h} K\left(\frac{w - W_i}{h}\right)$$

where h is the bandwidth and K is the kernel function. We use a Gaussian kernel and a bandwidth of 1.5. The Gaussian kernel $K(z)$ is given by:

$$K[z] = \frac{1}{\sqrt{2\pi}} \exp\left(-\frac{z^2}{2}\right)$$

(see Johnston and DiNardo 1997).

5. The mean of the Republican scores is 16.7 and the standard deviation is 18.7. The Democratic mean is 63.8 and the standard deviation is 28.1. Cox and McCubbins (1993, 67) found similar results for COPE scores (a liberal interest group).

6. Similar figures were created for the Senate; results are nearly identical and, in the interest of space, are not included.

7. The mean ACU score for Democrats is 23.4, with a standard deviation of 24.2. Republicans have a mean of 76.6, and a standard deviation of 21.8.

8. This analysis aggregates data for many years into one picture and some year-by-year variation is lost. However, for every year in our sample, both House and Senate Republican “real” ADA scores are, on average, closer to 0 than the average Democratic score is to 100. Likewise, Republicans have lower standard deviations in all but one year in the House, and for the Senate this is true better than 60% of the time. Moreover, for real ACU scores, Democrats tend to be, on average, closer to the lower bound than Republicans are to the upper bound. This is true about 92% of the time in the Senate and 85% of the years in the House. In terms of the standard deviations, we find that the Democrats have higher standard deviations 73% of the time in the Senate, while in the House only 19% of the years conform to our hypothesis.

9. These data are for the raw interest group scores and have not been transformed to real scores using the inflation index created by Groseclose, Levitt, and Snyder (1995).

10. Results for the Senate are once again nearly identical. For those members rated 0 by the ADA, only 29.6% scored 100 on the ACU, with 70.4% receiving ACU scores of 90 or better. For those senators rated 0 by the ACU, only 18.9% scored a perfect 100 by the ADA, with only 40.3% receiving scores of 90 or better.

11. The strength of this relationship varies over time. The overall R^2 (1971–96) between these two measures is .84 in both chambers.

12. The view of heteroskedasticity as arising from a difference between the conditional mean function ($y|x$) and the linear projection of y on x has been observed by others. See Chamberlain (1982) for example.

13. Hausman, Newey, and Powell (1995) note that even knowledge of the parametric form of the distribution of measurement errors is not sufficient for consistent estimation. For example, one also has to assume that true value of the regressors (i.e., ideology) are random drawings from a distribution with a known parametric form.

14. Quantile regression techniques are least absolute deviation (LAD) estimators. Here, rather than minimizing the sum of the errors squared as in OLS, the best fitting line is calculated by minimizing the absolute deviations from the median (or any specified quantile). See Deaton (1997) for further discussion of this technique.

15. The results for the same model using House data are nearly identical. The equation for the median is $-.909(\text{ACU}) + 87.3$; for the 10th percentile it is $-.734(\text{ACU}) + 60.2$; for the 90th percentile it is $-.949(\text{ACU}) + 102.3$; standard errors calculated from 1,000 bootstrapped replications are .005, .012, and .008 respectively; $n = 10,916$.

16. We did, however, estimate the model under the assumption that the scores reflected censored versions of an underlying unbounded latent variable. We used a technique suggested by Powell (1984) and an algorithm described in Buchinsky (1994) that (unlike the Tobit, for example) is consistent in the presence of heteroskedasticity, and found similar results to those we present.

17. The probability limit of $\hat{\beta}$ is given by:

$$plim \hat{\beta} = \beta \frac{\sigma^2}{\sigma_x + \sigma_v^2}$$

where the attenuating term is sometimes referred to as the “signal-to-total variance ratio.” As this ratio varies from 0 (all noise) to 1 (no measurement error), the OLS coefficient is biased towards zero. The negative of the OLS coefficients from a regression of ADA on ACU and vice versa, gives the range of estimates for this signal-to-total variance ratio. Calculation of these coefficients for our data reveals that the downward bias will be somewhere between 6 and 12% in a simple bivariate regression setting.

18. The difference between these coefficients is consistent with the estimate given by the signal-to-total variance, which suggested that this change would fall between 6 and 12%.

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