Political Analysis, 8:4

Cutpoint-Adjusted Interest Group Ratings

Michael C. Herron
Department of Political Science, Scott Hall, 601 University Place, Evanston, IL 60208-1006
e-mail: m-herron@northwestern.edu

While it is very common for Congressional researchers to use interest group ratings as measures of legislator policy preferences, this paper argues that the manner in which such ratings are calculated implies that they may poorly approximate the underlying legislator preferences on which they are based. In light of this, the paper develops a technique designed to adjust interest group ratings so that they more closely correlate with legislator preferences. It argues based on Monte Carlo simulations that the technique produces adjusted ratings that improve on unadjusted ratings, and it applies the adjustment technique to historical ratings published by the Americans for Democratic Action.

1 Introduction

STUDIES OF CONGRESS that test theories of party control (Moscardelli et al. 1998), analyze committee composition (Krehbiel 1992; Groseclose 1994), and attempt to determine if political party membership has an independent effect on legislator voting decisions (Krehbiel 1995; Binder et al. 1999) require measures of legislator policy preferences. These preferences are not directly observable, however, and this has led many Congressional researchers to use interest group ratings as proxies for them.

This paper offers an analysis of interest group ratings and the extent to which they are suitable measures of legislator preferences. Its results are based on the theory of spatial voting, and, accordingly, on any given political issue Congressional legislators are presumed to be endowed with real–valued ideal points that reflect true policy preferences. Furthermore, it is assumed that each Congressional vote can be summarized by a roll call cutpoint that divides a chamber of Congress into "yea" and "nay" coalitions. When an interest group selects a set of roll calls on which to base its ratings, it is therefore implicitly selecting a distribution of roll call cutpoints. It is the contention of this paper that the relation between underlying legislator preferences and a set of interest group ratings will almost always be distorted by features in the associated cutpoint distribution.

This paper focuses on two types of what are called cutpoint distortions. A cutpoint distortion of the first type is said to occur when interest group ratings mischaracterize the difference between the median Democrat and the median Republican in a single Congressional

Author's note: Computer code and ADA cutpoints are available from the *Political Analysis* website. The author thanks Tim Groseclose for providing Americans for Democratic Action ratings from 1947 to 1992. The calculation of optimal cutpoint parameters as described in the body of this paper was carried out using CFSQP Version 2.5d.

chamber. For example, a given set of interest group ratings can make the median Democrat in a chamber appear to have very different preferences than the corresponding median Republican even when these two individuals actually have similar preferences. Using the terminology of Snyder (1992), such interest group ratings would be considered artificially extreme. Despite the fact that Snyder's article on artificial extremism was published in 1992, and despite Poole and Rosenthal's (1998, Chap. 8) essential confirmation of Snyder's arguments, it is fair to say that the vast majority of Congressional studies which rely on interest group ratings disregards the potential measurement problems associated with artificial extremism or what is known here as cutpoint distortions of the first type.

A cutpoint distortion of the second type is said to occur when interest group ratings mischaracterize the extent of Democratic and Republican intraparty preference heterogeneity. For example, interest group ratings may make Democrats appear to be more homogeneous than Republicans even if, based on actual preferences, they are actually more heterogeneous. In a similar manner, an interest group can create a set of ratings based on a symmetric distribution of legislator ideal points, but the resulting ratings may be asymmetrically distributed.

The second type of cutpoint distortion is problematic in light of empirical research that seeks to determine the degree to which political party membership exerts an independent influence on legislator voting behavior in Congress (e.g., Krehbiel 1995; Brady et al. 1998; Krehbiel 1998; Binder et al. 1999; Snyder and Groseclose 2000). If interest group ratings distort the extent of intraparty preference heterogeneity across the Democratic and Republican parties, then Congressional researchers will find it difficult if not impossible to use such ratings in attempts to disentangle in an accurate way the independent impact of party membership, as opposed to policy preferences, on legislator voting behavior.

This paper argues that the properties of interest group ratings, and in particular, the potential for cutpoint distortions, can be studied with an eye toward improving the fit between such ratings and actual legislator policy preferences. And it asserts that, under a set of weak conditions, adjustments to interest group ratings can be made in light of the two types of distortions described above. The adjustment method proposed in this paper leads to what are called *cutpoint-adjusted interest group ratings*, and it is the claim of this paper that cutpoint-adjusted ratings, while not perfect legislator preference proxies, will be more accurate compared to regular or unadjusted interest group ratings.

Beyond Snyder's discussion of artificial extremism, other researchers have leveled criticisms at interest group ratings (e.g., Poole and Daniels 1985; Hall and Grofman 1990; Jackson and Kingdon 1992; Cox and McCubbins 1993; Brunell et al. 1999). For example, Hall and Grofman (1990) offer a behavioral critique of interest group ratings: they argue that members of Congress cast roll call votes in a manner that is deferential to key Congressional committee figures. It follows from this position that, on account of legislator behavior, distillations of voting records and hence interest group ratings may not reflect underlying legislator policy preferences.

Despite the many critiques, however, interest group ratings continue to be practically ubiquitous in Congressional studies. And it appears that the primary reason Congressional scholars have relied on interest group ratings as legislator preference measures is that such ratings are plausibly jurisdiction-specific. It is logical, for instance, to employ ratings produced by the Leadership Conference on Civil Rights in a study of civil rights voting (e.g., Hutchings 1998). General ideology ratings of members of Congress, for example, NOMINATE scores (Poole and Rosenthal 1997), are not jurisdiction-specific. Such ratings, therefore, may not be applicable to Congressional research that depends heavily on comparisons of, say, committee composition across legislative jurisdictions (Maltzman 1997). Thus, it is to be expected that interest group ratings will maintain their widespread position in Congressional research, and it follows that efforts to improve the extent to which they accurately proxy for legislator policy preferences are warranted.

The interest group rating adjustment method developed here can be contrasted with two recent proposals for alternative correction methods. First, Groseclose, Levitt, and Snyder (1999) (hereafter GLS) present a technique that, they assert, allows researchers to compare ADA ratings across sessions and chambers of Congress. The key difference between the GLS approach and the technique proposed here is that GLS assume that ADA ratings are accurate within individual Congressional chamber sessions. In fact, problems such as artificial extremism indicate that such an assumption is dubious. In contrast, the interest group rating adjustment method described in this paper begins with the conjecture that interest group ratings for individual Congressional chamber sessions are almost certainly inaccurate, and it seeks to adjust for inaccuracies.[1]

Second, Brunell et al. (1999) propose that researchers employing ADA or ACU ratings as legislator ideology measures in regression studies use a two-stage least-squares procedure in which ratings of the ACU are used as instruments for those of the ADA, or vice versa. This method does not produce stand-alone adjusted interest group ratings, and therefore a researcher carrying out a study that is not regression-based cannot use it. Furthermore, because the Brunell et al. method does not produce stand-alone ratings, it has the feature that different implementations of it to the same set of ADA and ACU ratings can produce different adjustments for the said ratings. This is not helpful since the adjustments that ostensibly need to be applied to ADA ratings should be application-independent. In contrast to the Brunell et al. method, this paper describes a technique that produces stand-alone, adjusted interest group ratings that can be applied to all research contexts.

The next section describes the spatial voting framework used here to analyze interest group ratings, and it illustrates the impact of cutpoint distortions. The paper then proposes a method to adjust interest group ratings for cutpoint distortions, explains the method's assumptions, and shows how cutpoint-adjusted interest group ratings are produced. After presenting a sequence of numerical examples that illustrate the value of cutpoint-adjusted interest group ratings, the paper considers cutpoint-adjusted ADA ratings.

2 Legislator Policy Preferences, Roll Call Cutpoints, and Interest Group Ratings

Given a political issue area, for example, defense policy, let p_i, $i = 1, \ldots, N$, denote legislator i's *p*reference or ideal point. Henceforth, the terms "legislator preference" and "legislator ideal point" are used interchangeably. Let $r_i \in [0, 1]$ denote an interest group's *r*ating of legislator i, where, in the case of defense, r_i could denote the American Security Council's rating. Large values of r_i are associated with legislators favored by the interest group, and small values with those who are disfavored.[2]

Since legislator ideal points have neither an absolute scale nor a location, an affine transformation of a collection of ideal points will yield an equivalent collection of ideal points. Therefore, at issue in an analysis of whether interest group ratings are good preference proxy variables is not whether $r_i = p_i$. Rather, what is important is whether there is a linear relationship between these two. If indeed there is a linear relation between a set of interest group ratings and associated legislator preferences, then the ratings perfectly proxy for the

[1] For related work on preferences and scaling, see Aldrich and McKelvey (1977) and Poole (1998).

[2] Interest group ratings are traditionally reported as numbers from 0 to 100, but they are scaled to lie between 0 and 1 for the purposes of this paper.

underlying preferences. If such a linear relation does not exist, however, then the ratings are not perfect preference proxies. Therefore, a maintained hypothesis in all analyses that use interest group ratings is that the ratings are linear functions of actual legislator ideal points.

Suppose that an interest group uses a collection of J roll call votes to form r_i, $i = 1, \ldots, N$. Let c_j be the cutpoint for roll call vote j, $j = 1, \ldots, J$. Spatial voting theory implies that all legislators i with $p_i > c_j$ vote in the same direction on bill j. For the examples in this section, assume, without loss of generality, that $p_i > c_j$ implies that i votes in favor on roll call j. Furthermore, again without loss of generality, assume that $p_i \in [0, 1]$.

It is assumed throughout this paper that roll call cutpoints c_j, $j = 1, \ldots, J$, have a beta distribution with shape parameters α_c and β_c. Therefore, it follows that

$$r_i = F(p_i; \alpha_c, \beta_c) \tag{1}$$

where $F(\cdot; \alpha_c, \beta_c)$ is the beta distribution function with parameters $\alpha_c > 0$ and $\beta_c > 0$. When $\alpha_c = \beta_c = 1$, the distribution of roll call cutpoints is uniform on [0, 1], $F(t; 1, 1) = t$ for $t \in [0, 1]$, and in such a case, $r_i = p_i$, and, trivially, r_i is a linear function of p_i. For all other values of α_c and β_c, however, r_i will not be a linear function of p_i. The family of beta distributions is quite flexible, and using it to model roll call cutpoints enables consideration of a wide variety of cutpoint distributions.

Equation (1) can be used to illustrate potential problems associated with the assumption that r_i is linearly related to p_i. In particular, Fig. 1 displays the relation between the distribution of legislator ideal points p_i and the distribution of roll call cutpoints when the latter is characterized by $\alpha_c = 10$ and $\beta_c = 5$. Of the 12 vertical lines in the upper half of Fig. 1, suppose that the shorter 7 reflect Democratic ideal points and the longer 5 denote Republicans.[3] The unimodal curve in the figure describes the beta density of the roll call cutpoint distribution.

The lower panel in Fig. 1 displays the resulting interest group ratings r_i when Eq. (1) is applied to the 12 vertical lines in the top panel. There is not a linear relation between r_i and p_i in the figure: in particular, the Democratic values of r_i are compressed, the corresponding Republican values are expanded, and the correlation between the actual preference values and the resulting ratings is 0.93, a number less than 1. Furthermore, Fig. 1 contains evidence of both types of cutpoint distortions.

First, let med_D^p denote the median *D*emocratic policy *p*reference or ideal point, and let med_D^r denote the median Democratic interest group *r*ating. Define med_R^p and med_R^r similarly. While the values depicted in the top panel in Fig. 1 are based on $|med_R^p - med_D^p| = 0.30$, the bottom panel leads to $|med_R^r - med_D^r| = 0.57$. Thus, the Democratic–Republican rating difference is greater than it actually is based on true ideal points.

Similarly, the variance of the legislator ideal points in Fig. 1 is 0.033, whereas the variance of the interest group ratings is 0.098. These two numbers indicate that the interest group ratings exaggerate the extent of preference heterogeneity. In conjunction with the median expansion noted above, this illustrates a cutpoint distortion of the first type, in particular, artificial extremism.

Second, let var_D^p denote the variance of the *D*emocratic *p*references or ideal points and let var_D^r denote the variance of the resulting Democratic interest group *r*ating. As before, define var_R^p and var_R^r similarly. For the values depicted in Fig. 1, $var_D^p = 0.012 >$

[3]The party assumption is made entirely for convenience. Furthermore, the import of Fig. 1 would not be diminished if there were overlap between Democratic and Republican ideal points.

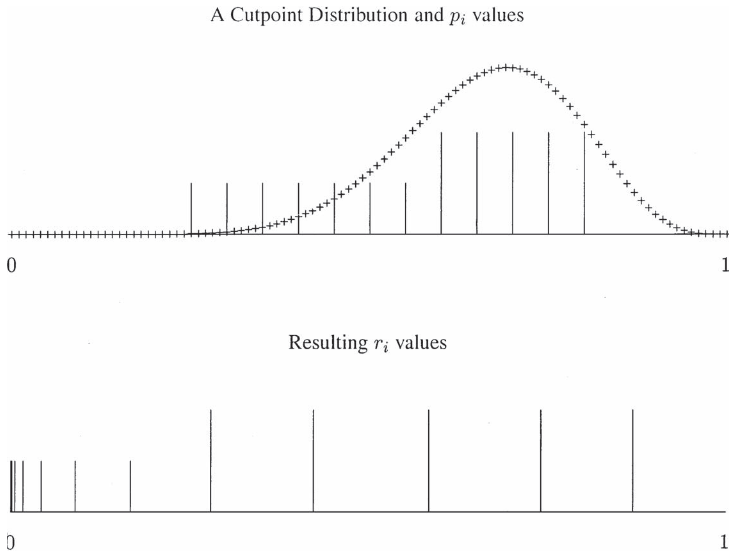


Fig. 1 Relationship between p_i, r_i, and a roll call cutpoint distribution ($\alpha_c = 10$ and $\beta_c = 5$).

$0.0063 = var_R^p$; this leads to a Democratic–Republican variance ratio of approximately 1.9. However, this variance ratio is distorted by the roll call cutpoint distribution depicted in the upper plot in Fig. 1. Indeed, it is clear from the lower plot that Republican interest group rating heterogeneity is much greater than Democratic interest group heterogeneity. Namely, $var_R^r = 0.056 > 0.0038 = var_D^r$; the corresponding Democratic–Republican ratio is 0.068, a number less than 1, not to mention less than 1.9. In other words, although Democratic ideal points are more dispersed than Republican ideal points, Democratic interest group ratings are less dispersed. This illustrates a cutpoint distortion of the second type.

As there are seven Democrats and five Republicans in the Fig. 1 legislature, one would characterize the legislature as being controlled by Democrats. This is reflected in a cutpoint distribution that implies that most winning coalitions consist of almost all Democrats combined with left-leaning Republicans. Because of the supermajoritarian cloture rules in the Senate and the supermajoritarian veto override provision in both chambers of Congress, cutpoint distributions that are not centered over the distribution of legislator ideal points should be the norm (Krehbiel 1998).[4]

Now assume that the two extreme Democrats in Fig. 1 are replaced by two correspondingly extreme Republicans. This leads to Fig. 2, which provides a second example of cutpoint distortions. Again, in this figure there is not a linear relationship between p_i values in the upper plot and resulting r_i values in the lower: the variance of Democratic interest group ratings is augmented, while the spread of Republican ratings is depressed. As in Fig. 1, the correlation of ideal points and resulting ratings is 0.93.

From the upper plot in Fig. 2, the Democratic–Republican difference in medians is $|med_R^p - med_D^p| = 0.30$. However, based on the resulting interest group ratings,

[4]Figure 1 is based on the notion that an interest group samples randomly from a collection of roll calls taken over the course of a Congressional session. This is, in fact, a best-case scenario. If an interest group were deliberately to pick key roll calls to, say, compress the ratings variance of its legislative enemies, then the resulting ratings could be highly skewed and off-center compared to the distribution of legislator ideal points.

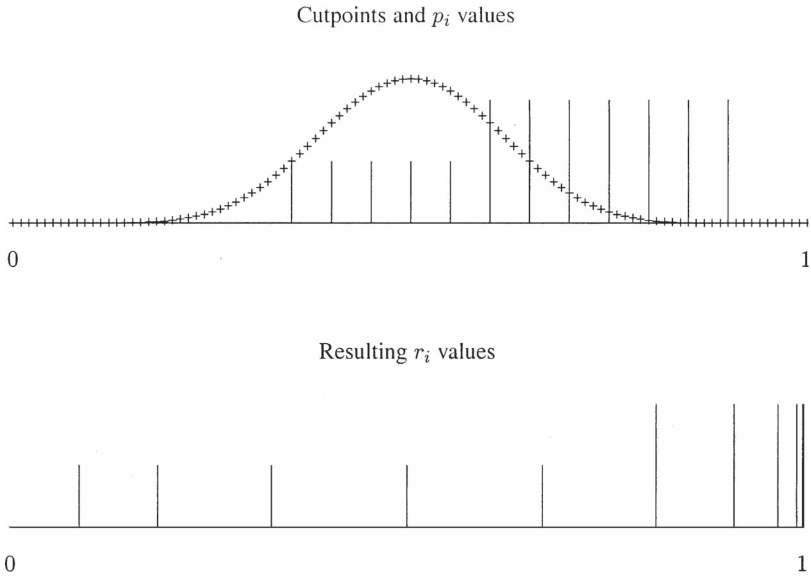


Fig. 2 Relationship between p_i, r_i, and a roll call cutpoint distribution ($\alpha_c = 10$ and $\beta_c = 10$).

$|med_R^r - med_D^r| = 0.66$. And, the variance of legislator ideal points is 0.033 while the variance of the interest group ratings is 0.12. Thus, the roll call cutpoint distribution shown in Fig. 2 leads to artificial extremism as measured by overall ideal point and interest group rating variance.

Based on an analysis of the interest group ratings in Fig. 2, it would be logical to conclude that the Republicans were highly unified but that the Democrats were not. According to actual legislator preferences, however, this conclusion does not follow. It is an artifact of the distribution of roll call cutpoints—in particular, of a cutpoint distortion of the second type.

Figure 2 was constructed based on a simple change from Fig. 1. Namely, in the former two, extreme leftist legislators were replaced by extreme rightists; this gave control of the hypothetical legislature to the Republicans, and the cutpoint distribution moved accordingly. Yet while the legislative composition in Figs. 1 and 2 is similar, the distribution of interest group ratings is starkly different. Thus, the figures have two implications. First, considered independently of one another, they draw attention to the consequences of cutpoint distributions and illustrate that the relation between legislator ideal points and interest group ratings can be complicated and nonlinear. Second, the figures show that comparing interest group ratings across different chambers (or time periods, for that matter) is something that should never be done. Relatively minor changes in legislature composition, if they affect the mean of the roll cutpoint distribution, can have major implications for interest group ratings.

It is important to point out that the number of bills an interest group uses to construct its ratings is not necessarily related to the accuracy of the resulting ratings. If, for instance, one group samples 10 times from the cutpoint distribution in Fig. 2, another group considers 15 bills, and still another group 100, the three sets of ratings will be almost identical. In other words, the accuracy of group ratings does not necessarily increase with the number of bills considered.

On account of the presence of the beta distribution function in Eq. (1), for arbitrary cutpoint parameters α_c and β_c it is, in general, not possible to derive analytic results as to the extent of cutpoint distortions. However, beyond the examples of Figs. 1 and 2, the impact of cutpoint distortions can be seen via the following example.

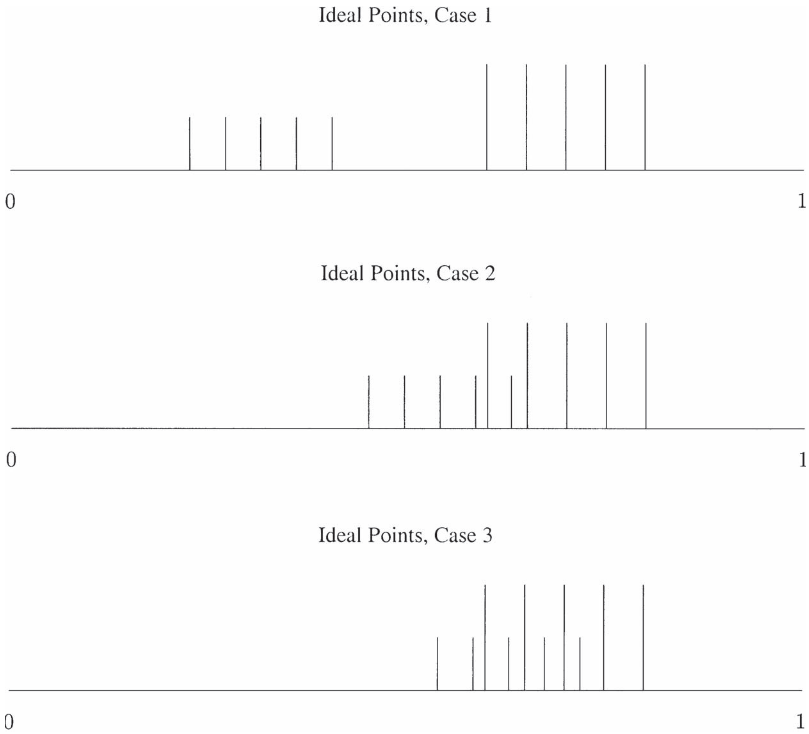


Fig. 3 Three cases of Democratic and Republican ideal points.

Consider three distributions of Democratic and Republican ideal points as shown in Fig. 3. In the three cases, Republican ideal points are identical and the three sets of Democratic ideal points have identical variances but different locations. In Case 1 there is a greater Democratic–Republican preference difference than in Cases 2 and 3, and so forth. Let α_c and β_c range from 1 to 40 in the integers. This defines $40^2 = 1600$ different cutpoint distributions, and the extent of the cutpoint distortions caused by these 1600 different distributions is captured in Table 1.

The column in Table 1 labeled "Artificially Extreme (Median)" lists the percentage of cases of α_c and β_c for which the absolute Democratic–Republican median difference as calculated by interest group ratings is greater than the absolute difference as calculated by actual ideal points. The three percentages in the column are all fairly large, and, furthermore, Table 1 suggests that the closer the median Democrat and Republican, the smaller the distortionary effect of interest group ratings as measured by absolute median differences.

The third column in Table 1, labeled "Artificially Extreme (Variance)," lists the percentages of cases for which the variance of interest group ratings is greater than the variance of

Table 1 Cutpoint distortions for the ideal point distributions in Fig. 3

Case	*Artificially extreme (median)*	*Artificially extreme (variance)*	*Variance ranking preserved*
1	58.9%	67.2%	49.1%
2	41.1%	52.6%	31.8%
3	32.3%	41.3%	26.9%

the underlying ideal points. According to this measure of artificial extremism, the distortionary effects of interest group ratings are diminished as the variance of actual legislator preferences gets lower. This means that relatively homogeneous legislatures are less likely to be characterized by artificial extremism as measured by overall variance.

Finally, in all three cases in Fig. 3, as measured by ideal point variance the Democrats are slightly more homogeneous than the Republicans. And, the column in Table 1 labeled "Variance Preserved" lists the percentage of the 1600 cutpoint distribution for which the Democratic interest group ratings have a lower variance than the Republican ratings. The relatively low values in the column indicate the extent of cutpoint distortions and the fact that interest group ratings will often fail to identify correctly the more homogeneous party.

The two detailed examples and the results in Table 1 illustrate the following points. First, the difference in median Democratic and Republican ideal points may be very different from the corresponding difference as measured by interest group ratings. Second, the variance of interest group ratings may not reflect the variance of underlying legislator ideal points. Third, even if the variance of Democratic ideal points is greater (lesser) than the variance of Republican ideal points, the variance of Democratic interest group ratings may be less (greater) than the variance of Republican ratings. Fourth, if an interest group forms a set of ratings using roll call votes with nonminimal winning coalitions, then a symmetric distribution of legislator ideal points will be transformed into an asymmetric distribution of interest group ratings. These four points constitute serious problems that impinge on the accuracy of interest group ratings insofar as they proxy for legislator policy preferences.

3 Adjusting Interest Group Ratings for Cutpoint Distortions

This section proposes a method to adjust interest group ratings for cutpoint distortions and explains the procedure for generating what are called cutpoint-adjusted interest group ratings. The first step in the cutpoint adjustment procedure involves another rescaling of interest group ratings. Recall that actual ratings r_i lie between 0 and 100, and in the previous section r_i values were scaled to lie between 0 and 1. The cutpoint adjustment technique rescales interest group ratings so that they lie in $[0.05, 0.95]$. That is, define $r_i' = 0.9r_i + 0.05$. Arbitrary rescalings preserve all information available in interest group ratings.

Key to the cutpoint adjustment method is the following observation: were a researcher to have a collection of interest group ratings r_i' and also know the corresponding values of the cutpoint parameters α_c and β_c, then by inverting Eq. (1), underlying legislator ideal points could be inferred in a straightforward manner:

$$p_i = F^{-1}(r_i'; \alpha_c, \beta_c) \tag{2}$$

Inverting interest group ratings as in Eq. (2) would uncover the true p_i values on which a set of interest group ratings r_i' was based.[5] The difficulty inherent in applying Eq. (2), however, is that α_c and β_c are not observable and cannot be calculated from interest group ratings alone.

Suppose for the moment, however, that a researcher with a set of interest group ratings actually did know the distribution of Democratic ideal points. In such a situation, the researcher could combine her knowledge of the Democratic ideal point distribution with

[5]Technically, on account of the rescaling of interest group ratings, the F in Eq. (2) is different from the F in Eq. (1). However, the paper could have started with r_i' in Eq. (1) and the same logic as explained previously would still be valid.

observed Democratic interest group ratings to solve via Eq. (1) for the appropriate cutpoint parameters α_c and β_c. Then Republican ideal points could be inferred using Eq. (2). In other words, a sufficient condition for employing Eq. (2) is knowledge of the distribution of Democratic ideal points along with exact values of Democratic interest group ratings.

This is the basis for the cutpoint adjustment method. Even with a set of interest group ratings, researchers of course never know the distribution of associated Democratic ideal points any more than they know the values of the pertinent cutpoint parameters. However, the cutpoint adjustment method is based on assuming a specific distribution for Democratic ideal points and then applying the logic specified above. Note that the method does *not* assume that Democratic ideal points are themselves known. Rather, it makes a weaker assumption about the shape of the histogram that Democratic ideal points would inhabit if they were known.

Consider a normalization such that Democratic ideal points p_i are assumed to have a specified distribution G. The cutpoint adjustment procedure asks, What values of the cutpoint parameters α_c and β_c best transform via Eq. (2) the observed Democratic interest group ratings r_i' so that inferred Democratic ideal points are distributed G? These values of α_c and β_c can then be used to infer Republican ideal points *without* the assumption that the latter ideal points fit any particular distribution.[6]

The specific procedure for calculating cutpoint-adjusted interest group ratings is as follows. Given a normalizing distribution function G, let μ_k^G denote the kth moment of G and let $\hat{\mu}_k^G$ denote the kth moment of the inferred Democratic ideal points $\hat{p}_i$, generated from Eq. (2), when $\alpha_c = \hat{\alpha}_c$ and $\beta_c = \hat{\beta}_c$. Note that $\hat{\mu}_k^G$ is a function of $\hat{\alpha}_c$ and $\hat{\beta}_c$. Let $i \in D$ imply that legislator i is a Democrat and let m be a positive integer. If there are N_D Democratic legislators, $\hat{\alpha}_c$ and $\hat{\beta}_c$ are chosen to minimize the following objective function:

$$\begin{aligned}\sum_{k=1}^{m} \left(\mu_k^G - \hat{\mu}_k^G(\hat{\alpha}_c, \hat{\beta}_c)\right)^2 &= \sum_{k=1}^{m} \left(\mu_k^G - \frac{1}{N_D} \sum_{i \in D} \hat{p}_i^{\,k}\right)^2 \\ &= \sum_{k=1}^{m} \left(\mu_k^G - \frac{1}{N_D} \sum_{i \in D} (F^{-1}(r_i'; \hat{\alpha}_c, \hat{\beta}_c))^k\right)^2 \qquad (3)\end{aligned}$$

The optimization problem implicit in Eq. (3) selects $\hat{\alpha}_c$ and $\hat{\beta}_c$ so that the first m moments of the resulting inferred distribution of Democratic ideal points are as close as possible to the corresponding m moments of G, the normalizing distribution.[7] After employing Eq. (2) with $\alpha_c = \hat{\alpha}_c$ and $\beta_c = \hat{\beta}_c$, the resulting values of $\hat{p}_i$ for all legislators i are the cutpoint-adjusted interest group ratings.

A useful feature of this paper's rating adjustment technique is that it allows cutpoint distributions to be skewed compared to legislator ideal points, and so forth. Indeed, based on the logic of Fowler (1982) and the results of Brunell et al. (1999), one might conjecture that the cutpoint distribution used by a politically tendentious interest group may be highly skewed. Because beta distributions are not restricted to being symmetric, this conjecture can be internalized in the cutpoint-adjustment procedure. In other words, the critique that

[6]The use of party membership to determine those legislators that have ideal point distributed G is not a function of the fact that interest groups tend to have an affinity for one party or another. Rather, what is needed is a normalizing group, and it is logical to use as such a group a collection of legislators who have ideal points that are presumably unimodal. It is quite plausible to think that the Democratic members of Congress fit this category.

[7]The choice of m will influence the extent to which the inferred Democratic ideal points will be approximately distributed G.

interest groups choose cutpoint distributions to facilitate political goals—i.e., highlighting legislative enemies—does not nullify the validity of the cutpoint-adjustment method.[8]

4 Monte Carlo Analysis of the Cutpoint Adjustment Method

The cutpoint adjustment procedure attempts to infer legislator preferences subject to normalizing Democratic ideal points so that they are distributed G, and throughout this paper G is assumed to be a beta distribution with both shape parameters set to 50. If Democratic preferences were truly distributed G, then the normalization would not be binding, and in this case the procedure would uncover true Democratic and Republican preferences. Clearly, if true Democratic ideal points are distributed G, then a researcher who needed a measure of legislator preferences would be best served by using cutpoint-adjusted interest group ratings as opposed to unadjusted ratings.

However, it is reasonable to inquire, What if the Democratic ideal points associated with a collection of interest group ratings are not distributed G? In such a situation, the cutpoint adjustment procedure operated on a set of interest group ratings will be relying on a normalizing distribution that is incorrect. Will the cutpoint adjustment procedure produce meaningful inferred legislator preferences or will unadjusted interest group ratings, which do not depend on the choice of G, be better preference proxies than the cutpoint-adjusted ratings? This section addresses this question through Monte Carlo experiments, and it presents a series of simulations in which a distribution for underlying legislator preferences p_i is specified, interest group ratings r_i are calculated based on a set of cutpoint parameters α_c and β_c, and cutpoint-adjusted $\hat{p}_i$ values are derived using the cutpoint adjustment procedure. Finally, each simulation considers whether unadjusted ratings or adjusted ratings are better proxies for underlying legislator preferences p_i.

All simulations are based on a chamber of 435 legislators with 200 Democrats and 235 Republicans. Democratic ideal points are assumed to have a beta distribution with shape parameters seven and five, and Republican ideal points are distributed beta with parameters three and five. These two ideal point distributions are shown in Fig. 4. Note that the actual distribution of Democratic ideal points is not G, where, as noted previously, G is beta with both shape parameters set to 50. Since the distribution of Democratic ideal points violates the normalizing assumption that is part of the cutpoint adjustment technique, it follows that the resulting Monte Carlo results are conservative.[9]

The Monte Carlo experiments proceed as follows. For given values of α_c and β_c and 435 randomly drawn ideal points from the two densities shown in Fig. 4, Eq. (1) is used to generate interest group ratings r_i and rescaled ratings r_i'. Then the cutpoint adjustment procedure via Eq. (3) calculates $\hat{\alpha}_c$ and $\hat{\beta}_c$ and it subsequently uses Eq. (2) to generate $\hat{p}_i$ for all legislators i. Based on 25 simulations—that is, repetitions of the above draws for each pair of α_c and β_c—Table 2 displays the first set of results.

In Table 2, the columns labeled α_c and β_c list the actual cutpoint parameters used to derive interest group ratings. The four results columns show that the correlation between p_i and $\hat{p}_i$, denoted $\rho(p_i, \hat{p}_i)$, is for all cases of α_c and β_c greater than the correlation $\rho(p_i, r_i)$ between p_i and r_i. In other words, for the legislator preferences and roll call cutpoint

[8]The reason that the cutpoint-adjustment procedure uses $r_i' \in [0.05, 0.95]$, as opposed to $r_i \in [0, 1]$, is because many sets of actual interest group rating contain a plethora of 0's and 100's. If Eq. (2) were to use r_i, many inferred $\hat{p}_i$ values would accordingly be 0 and 1 for the simple reason that the beta distribution has positive support on [0, 1] for all possible values of α_c and β_c.

[9]Results akin to those that follow were also obtained using underlying ideal point distributions that are truncated normal.

Table 2 Simulation results, Part 1

α_c	β_c	$\rho(p_i, r_i)$	$\rho(p_i, \hat{p}_i)$
1	20	0.40	0.44
2	20	0.49	0.55
3	20	0.59	0.65
4	20	0.66	0.72
5	20	0.72	0.77
6	20	0.77	0.82
7	20	0.81	0.86
8	20	0.84	0.88
9	20	0.86	0.90
10	20	0.89	0.91
11	20	0.90	0.92
12	20	0.92	0.94
13	20	0.93	0.95
14	20	0.93	0.95
15	20	0.94	0.96
16	20	0.94	0.96
17	20	0.94	0.96
18	20	0.94	0.96
19	20	0.94	0.96
20	20	0.93	0.95

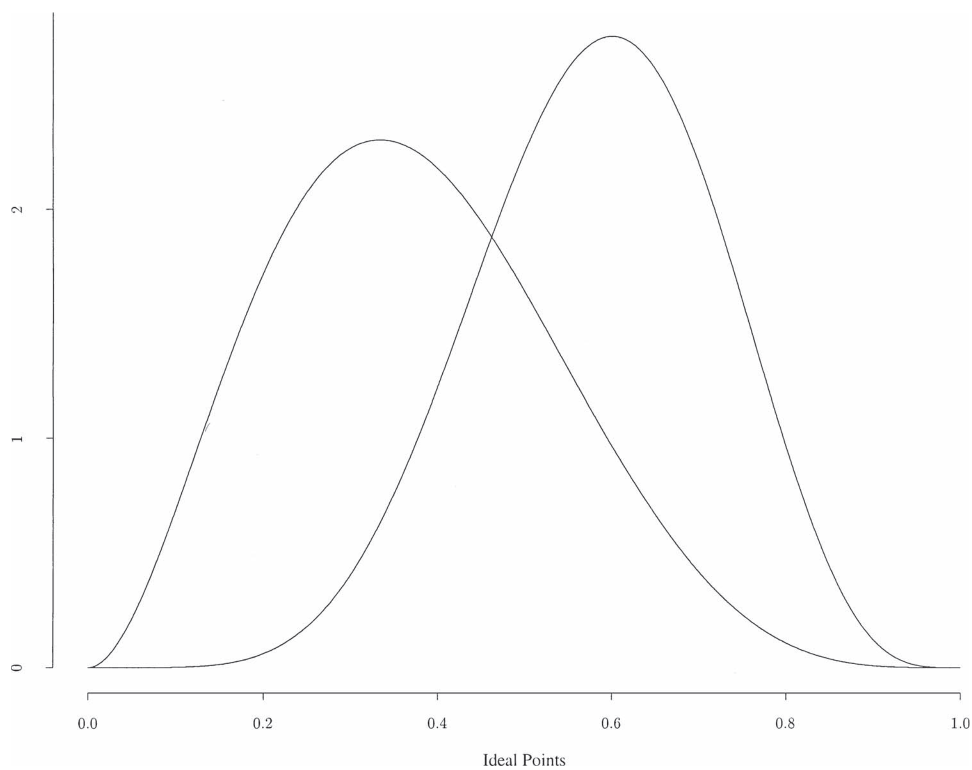


Fig. 4 Assumed distribution of Democratic and Republican preferences.

Table 3 Simulation results, Part 2

Case 1			*Case 2*		
γ	$\rho(p_i, r_i)$	$\rho(p_i, \hat{p}_i)$	γ	$\rho(p_i, r_i)$	$\rho(p_i, \hat{p}_i)$
0.0	0.93	0.95	0.0	0.92	0.96
0.1	0.94	0.96	0.1	0.93	0.96
0.2	0.95	0.96	0.2	0.93	0.97
0.3	0.96	0.96	0.3	0.94	0.97
0.4	0.96	0.97	0.4	0.95	0.97
0.5	0.97	0.97	0.5	0.96	0.98
0.6	0.98	0.98	0.6	0.97	0.98
0.7	0.98	0.98	0.7	0.97	0.99
0.8	0.97	0.99	0.8	0.98	0.99
0.9	0.94	0.98	0.9	0.99	0.99
1.0	0.89	0.96	1.0	1.00	1.00

distributions considered in Table 2, cutpoint-adjusted interest group ratings $\hat{p}_i$ are better preference proxies than unadjusted interest group ratings r_i. Importantly, this conclusion is valid despite the fact that it is based on Democratic ideal points that are not distributed G.

Because all correlations in Table 2 are less than 1, it follows that in none of the cases are the cutpoint-adjusted ratings $\hat{p}_i$ perfect linear transformations of legislator ideal points p_i. This implies that the cutpoint adjustment technique will not perfectly uncover legislator policy preferences when the actual Democratic preference distribution is different than G. This is, in fact, the reason that $\hat{p}_i$ values are called *cutpoint-adjusted* ratings rather than *cutpoint-corrected* ratings. Nonetheless, Table 2 indicates that, even though not perfect, cutpoint-adjusted interest group ratings do a better job of proxying for legislator preferences job compared to unadjusted ratings.

One potential weakness of the results in Table 2 is that they are based on cutpoint distributions that are beta. From the perspective of the cutpoint adjustment procedure, this is fortuitous: recall that the procedure assumes that the roll call cutpoint distribution is indeed a member of the beta family. Hence, the next set of simulation results considers two cases in which interest group ratings are calculated with nonbeta cutpoint distributions.

Suppose that $\gamma \in [0, 1]$. In Case 1, let the roll call cutpoint distribution function be defined as $(1-\gamma)F(\cdot; 20, 20)+\gamma F(\cdot; 1, 4)$, where, as before, $F(\cdot; \alpha, \beta)$ is the beta distribution function with shape parameters α and β. In Case 2, suppose that the roll call cutpoint distribution function is $(1-\gamma)F(\cdot; 20, 20)+\gamma F(\cdot; 0.5, 0.75)$. In both Case 1 and Case 2 the cutpoint distribution function is a mixture of two beta distributions, and simulation results based on these two cases are displayed in Table 3.[10] In all rows of Table 3 feature $\rho(p_i, \hat{p}_i) \geq \rho(p_i, r_i)$, and it follows that the cutpoint-adjustment procedure produced improved preference measures compared to unadjusted interest group ratings even when the roll call cutpoint distribution used to create the ratings was a mixture of two beta distributions.

In light of the simulation evidence, it is useful to consider a potential critique of the cutpoint adjustment procedure. One might ask, If the objective of the cutpoint adjustment

[10] Some numbers in Table 3 are written as "1.00" due to rounding. The distributions used in Cases 1 and 2 are arbitrary, but additional simulation results suggest that the results in Table 3 are robust to changes in the two cases.

procedure is to infer legislator policy preferences, then by assuming that Democratic preferences have a specified distribution, is not the procedure actually assuming what it intends to estimate? There are two responses to this critique. First, the assumption that Democratic ideal points are distributed G when G is a beta distribution with both shape parameters set to 50 says little other than the fact that such ideal points are unimodal and symmetric. This is an intuitive assumption and one that is probably valid. As Snyder points out, many constituency characteristics such as income are unimodally distributed, and, if the preferences of members of Congress reflect constituency features, then one would expect Congressional ideal points to be unimodal.

Second, the Monte Carlo evidence indicates that cutpoint-adjusted interest group ratings are better preference proxies even in a specific case when the normalizing assumption regarding G is wrong. Thus, it is not sufficient to argue that the cutpoint adjustment procedure is flawed simply because Democratic ideal points are not distributed G.

5 Application: Cutpoint-Adjusted ADA Ratings

The paper now applies the interest group ratings adjustment technique to ratings produced by the Americans for Democratic Action (ADA), a politically liberal interest group. While the cutpoint-adjustment technique can be applied to any set of interest group ratings, the focus on those of the ADA is a consequence of the fact that ADA ratings are perhaps the most commonly used set of legislator preference measures within the literature on Congress.

5.1 *1980, 1992, and 1998 ADA Ratings for the House of Representatives*

Plots 1 and 2 in Fig. 5 display 1980 ADA ratings (scaled to lie within [0.05, 0.95]) for the House of Representatives.[11] The median Democratic ADA rating is 0.64 and the median Republican rating is 0.16. This ordering is intuitive on account of the ADA's liberal orientation and its general preference for Democrats over Republicans, and the unadjusted ADA ratings reflect a preference difference of $0.64 - 0.16 = 0.48$, or 300%. Furthermore, from these ratings it appears that the Republicans are more homogeneous than the Democrats: the Republican–Democratic variance ratio is $0.052/0.093 = 0.56$. Overall, the variance of the ADA House ratings is 0.091.

The upper two histograms in Fig. 5 imply that the unadjusted 1980 ADA House ratings were bimodal, with spikes near the endpoints of the unit interval. Since a bimodal distribution for House member preferences is odd in light of the fact that most constituency characteristics are unimodally distributed, the data in Fig. 5 are suspicious and the possibility for artificial extremism and consequently ADA rating inaccuracy should be taken seriously. Assuming, as usual, that G is a beta distribution with both shape parameters set to 50 and setting $m = 5$, optimal $\hat{\alpha}_c$ and $\hat{\beta}_c$ values from Eq. (3) are $\hat{\alpha}_c = 37.53$ and $\hat{\beta}_c = 40.46$. This leads to a roll call cutpoint distribution with a mean of 0.48, and cutpoint-adjusted ADA ratings for the 1980 House are depicted in plots 3 and 4 in Fig. 5.

Plot 3 in Fig. 5 depicts three densities, all of which relate to cutpoint-adjusted ADA ratings for the 1980 House. The histogram in the plot displays the cutpoint-adjusted ADA ratings for House Democrats, the dotted line is the density of the normalizing distribution G, and the solid curve is the smoothed density of the cutpoint-adjusted ratings. The fit among the three densities in plot 3 is quite pleasing—particularly when comparing the smoothed

[11]The smoothing curve in Fig. 5, along with all smoothing curves in this paper, was generated using a normal kernel with an appropriate bandwidth.

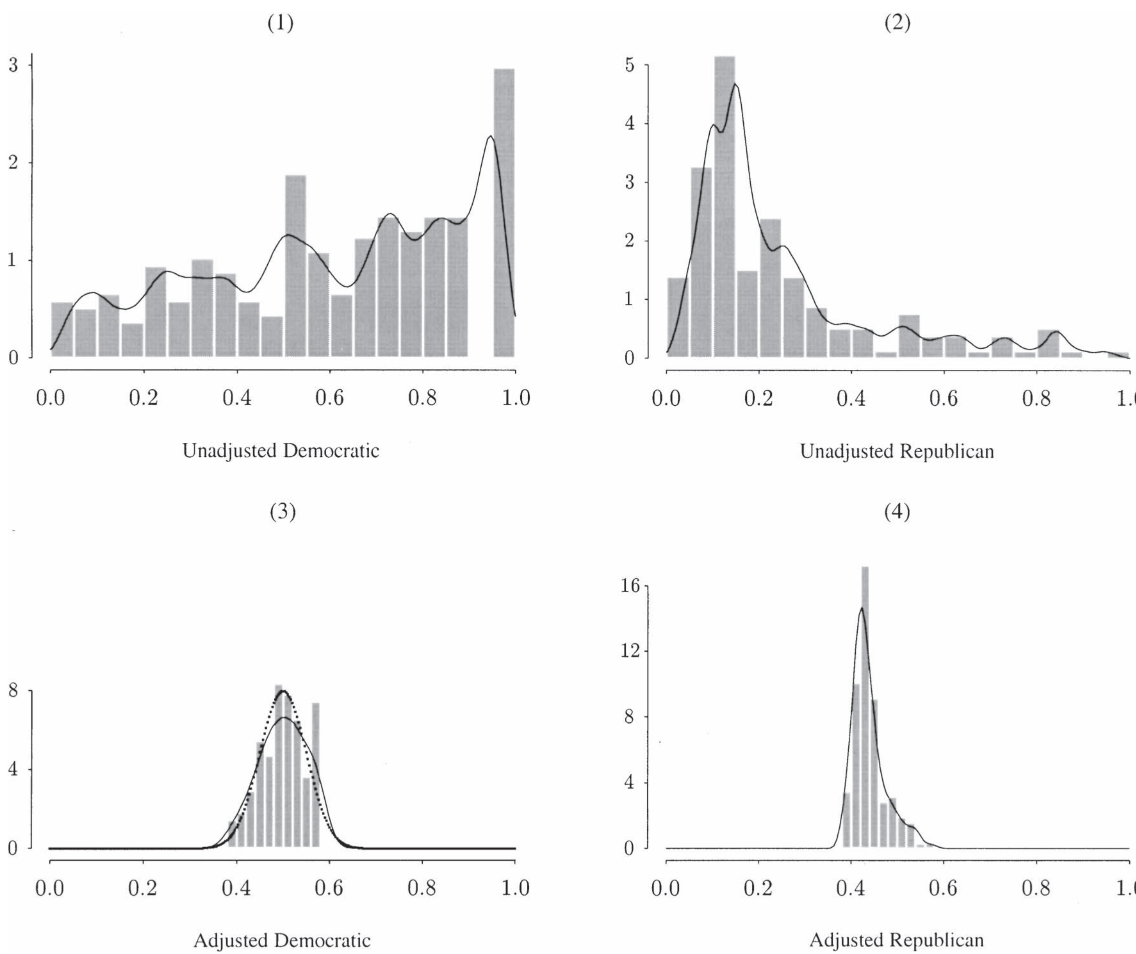


Fig. 5 Unadjusted and cutpoint-adjusted 1980 ADA House ratings.

density (solid line) and the density (dotted line) of G—and it suggests that the selection of G and the use of a beta distribution to model roll call cutpoints were reasonable.[12]

Plot 4 in Fig. 5 depicts cutpoint-adjusted ADA ratings for House Republicans. Compared to the cutpoint-adjusted Democratic ratings, the new Republican ratings are skewed right. This implies that the vast majority of Republican Representatives in 1980 were relatively conservative but that a small group was more liberal. One can also see this in plot 2, but plot 4 indicates that the right skewness in the unadjusted ADA ratings is, not surprisingly, exaggerated. In particular, the skewness of the Republican ADA ratings is 0.030, but the skewness of the cutpoint-adjusted ADA ratings is 0.0015. Thus, there was greater symmetry across Republican legislators than the 1980 ADA ratings implied.

The adjusted Democratic median is, by construction, 0.50 (this is the median of G). The adjusted Republican median is 0.42, and this leads to a median difference of $0.50 - 0.42 = 0.08$, or approximately 19%. Recall that the original ADA ratings indicated that the Democratic–Republican difference was 300%. This reflects, according to the cutpoint adjustment technique, an exaggeration of the difference between the median House Democrat and the median House Republican. Finally, the adjusted Republican-to-Democrat variance ratio is 0.54, slightly less than 0.56, the original ADA variance ratio. It follows that the ADA ratings did not do a particularly poor job of assessing the extent of across-party preference heterogeneity.

Similar features can be seen from 1992 ADA House ratings as displayed in Fig. 6. Plots 1 and 2 in the figure display unadjusted ratings, and these ratings appear artificially extreme as did the 1980 ratings. The cutpoint-adjusted ADA ratings for 1992—shown in plots 3 and 4—indicate a good fit in the Democratic plot, and they reduce the extremism of the median House Democratic versus the median House Republican. In particular, the median difference according to the unadjusted 1992 ADA ratings is $0.77 - 0.19 = 0.58$, or 305%. The corresponding difference in terms of cutpoint-adjusted ratings is $0.50 - 0.36 = 0.14$, or approximately 39%. Moreover, the Republican–Democratic variance ratio as measured by the unadjusted House ratings is $0.020/0.030 = 0.67$, while the corresponding ratio measured by adjusted ratings is $0.0016/0.0026 = 0.62$. While the unadjusted ADA ratings imply that Democratic preferences in 1992 were more heterogeneous than Republican preferences, according to the cutpoint-adjusted ratings they understate this fact.

The overall impact of the cutpoint adjustment as applied to 1980 and 1992 ADA ratings can be easily observed in Fig. 7. This figure contains four histograms, each with a corresponding smoothing curve. The histograms on the left in the figure pertain to 1980 ADA ratings for the entire House, and those on the right are associated with 1992 House ratings.

The top two histograms in Fig. 7 have the types of appearances frequently associated with interest group ratings. Namely, the two histograms are bimodal, their political centers clearly have relatively low densities, and so forth. In contrast, the cutpoint-adjusted ratings in the figure reflect more of a convincing picture of Congress in 1980 and 1992: there is strong evidence of unimodality in the cutpoint-adjusted ADA ratings, and, to the extent that one could argue in favor of bimodality in the adjusted ratings, the modes are not as extreme relative to the unadjusted ADA ratings.

Furthermore, the tails of the cutpoint-adjusted ratings are less heavy than the tails of the original ADA ratings: the unadjusted ADA kurtosis in 1980 is -1.40 and the corresponding

[12]The only potentially troubling feature in plot 3 (Fig. 5) is the right tail of the histogram of adjusted ADA ratings. However, judging by the smoothed density, the tail is not as sharp as it appears. It is important to consider smoothed densities of cutpoint-adjusted ratings along with histograms of adjusted ratings because interest groups often use only a small number of bills, 10 to 15 in some cases, to form their ratings.

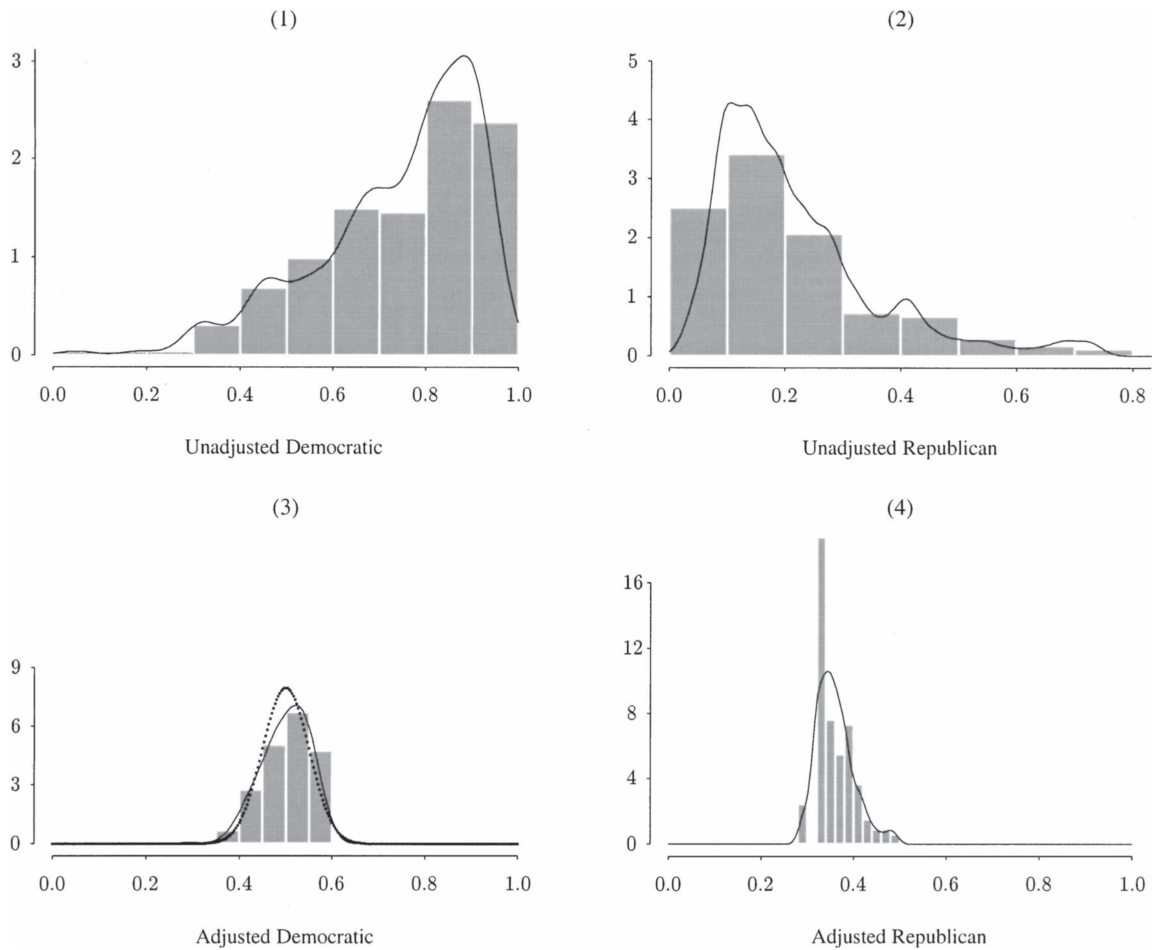
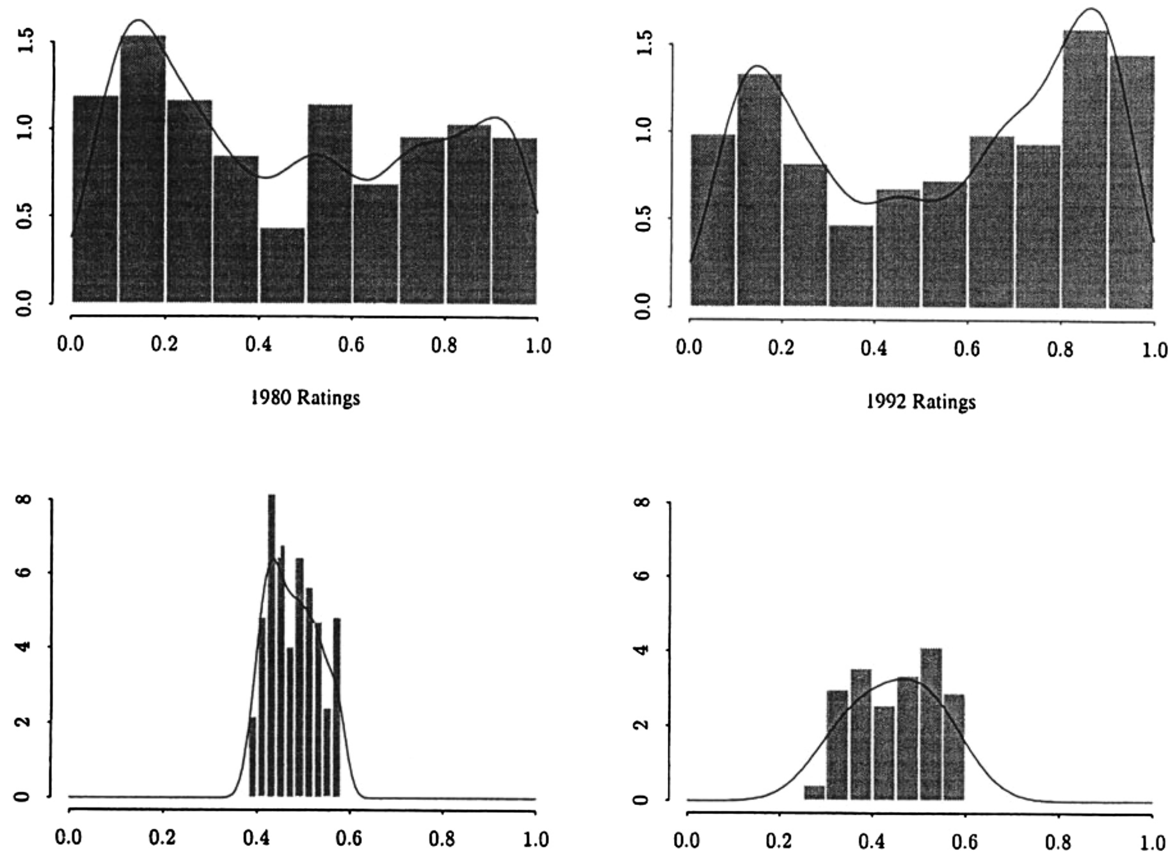


Fig. 6 Unadjusted and cutpoint-adjusted 1992 ADA House ratings.



1980 Cutpoint–Adjusted Ratings

1992 Cutpoint–Adjusted Ratings

Fig. 7 Comparison of 1980 and 1992 unadjusted and cutpoint-adjusted ADA ratings.

adjusted kurtosis is −1.05. Similarly, in 1992 the unadjusted ADA kurtosis is −1.50 and the adjusted kurtosis is −1.3. Since decreases in negative values of kurtosis indicate distributions with less mass in their tails, it follows that the adjusted ADA ratings imply the existence of a greater political center compared to that under the unadjusted ratings. In light of concerns over artificial extremism, this finding is intuitive and compelling, and it suggests that the cutpoint adjustment procedure is uncovering plausible distributions of House member preferences.

Although not pictured here, a similar analysis was conducted on 1998 House ADA ratings with the intent to determine whether cutpoint-adjusted ratings correlated better with constituency characteristics compared to unadjusted ratings. One would expect that a House member's 1998 ADA rating is highly correlated with the faction of the population in his or her district that voted for Bill Clinton in the 1996 presidential race. If cutpoint-adjusted ADA ratings represent an improvement over unadjusted ratings, then there should be a greater correlation between the former and the districtwide Clinton vote share as opposed to the correlation between unadjusted ratings and the Clinton share.

Indeed, for 1998, the correlation between unadjusted ADA House ratings and the Clinton vote share is 0.689, and the corresponding cutpoint-adjusted correlation is 0.691. This difference is extremely small, but it is positive in favor of the cutpoint-adjusted ratings. The fact that the improvement is even positive suggests that the mathematical details of the adjustment procedure is most likely addressing something systematic in how interest group ratings relate to underlying legislator preferences. Overall, and in line with the Monte Carlo evidence presented in Section 4, the single correlation difference suggests that the cutpoint adjustment procedure leads to small but measurable improvements compared to unadjusted ratings.

5.2 *A Comparison of ADA Ratings, 1974–1992*

As noted in the Introduction (Section 1), GLS offer a method for adjusting the scale and location of interest group ratings, ADA ratings, in particular, and they argue that their method fosters across-time and across-chamber comparisons of interest group ratings. However, the GLS method and its associated assumptions are fundamentally different from the cutpoint adjustment technique presented in this paper. The key to the GLS approach is the assumption that interest group ratings are valid preference measures within individual chamber-years. It is precisely this type of assumption that, according to Snyder (1992) and the results presented earlier in this paper, should not be made. In particular, since GLS only adjust the scale and location of ADA ratings across different chambers of Congress, they do not address the potentially nonlinear relation between ratings and underlying legislator preferences that can result from cutpoint distortions.

Based on values of $\hat{\alpha}_c$ and $\hat{\beta}_c$ from 1947–1992 for the House and Senate, Fig. 8 plots the Democratic–Republican difference in mean cutpoint-adjusted ADA rating for each chamber-year. For all years, the Democratic mean cutpoint-adjusted ADA rating is, by construction, one-half. However, as described earlier, the Republican mean is left unconstrained by the cutpoint adjustment procedure.

Both panels in Fig. 8 illustrate that the difference between the Democratic party and the Republican party has been growing in the past several decades, and this is consonant with results of Poole and Rosenthal (1997). Such a conclusion follows from the fact that the mean Democratic–Republican difference as measured by cutpoint-adjusted ADA ratings has increased, particularly from 1970 onward. Evidence from the House indicates that the Democratic and Republican parties were more polarized in the late 1940s and very early 1950s than they were in the 1960s and 1970s.

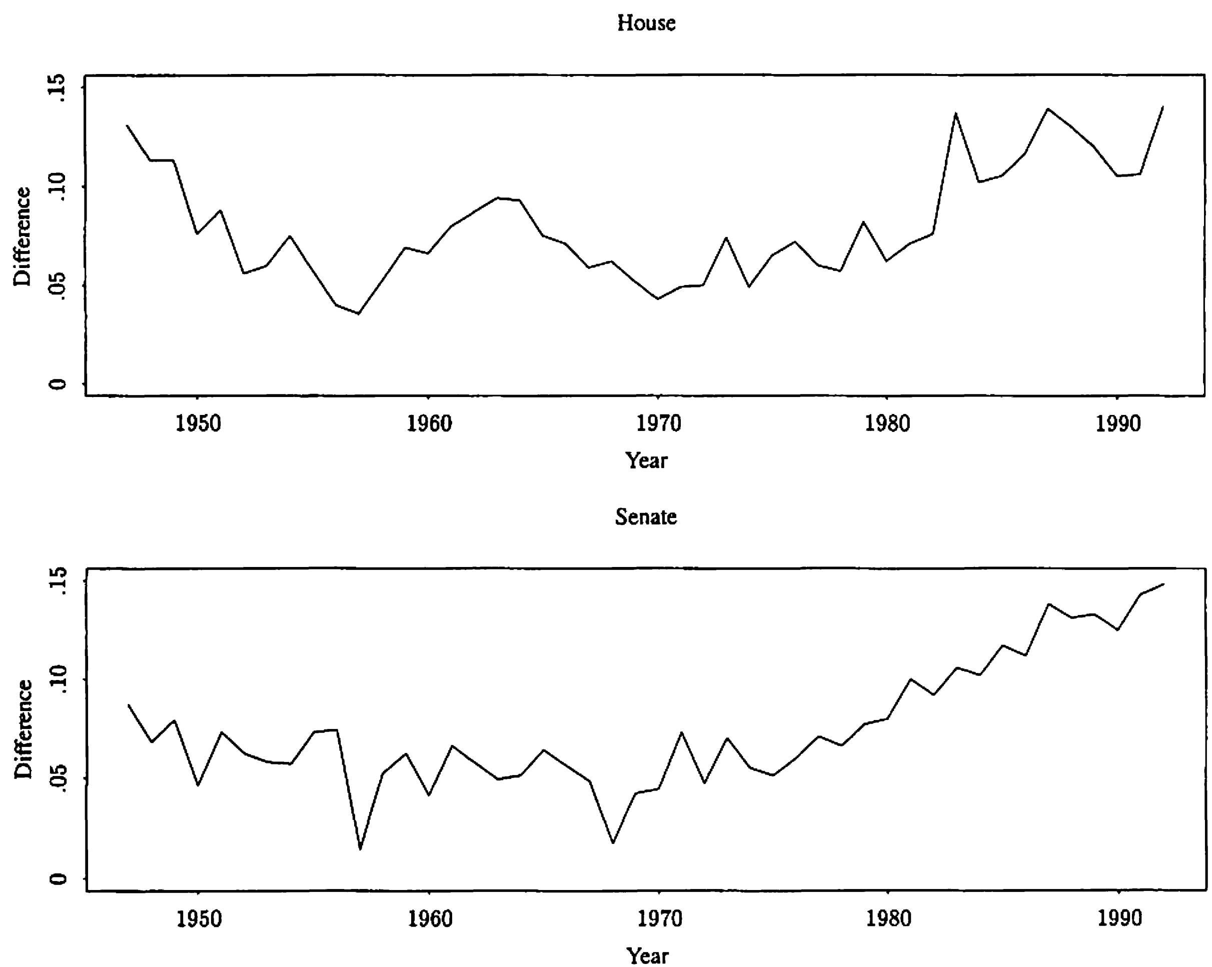


Fig. 8 Democratic–Republican difference in cutpoint-adjusted ADA ratings.

As pointed out by GLS, comparing unadjusted ADA ratings across time is not possible. However, the cutpoint adjustment technique offers a minimal form of across-time comparison, namely, comparison of the difference between the mean Democrat in a chamber and the mean Republican. Comparing cutpoint-adjusted ratings for individual legislators is not possible unless one wants to make the GLS assumption that legislators have preferences that are constant subject only to mean zero shocks each year. In theory, then, one could create a set of inflation and cutpoint adjusted ADA ratings based on GLS and the technique presented here.

6 Conclusion

Congressional researchers frequently use interest group ratings as measures of legislator policy preferences, and this paper has considered whether this is a reasonable thing to do. Broadly speaking, it is not. As explained in detail, calculating interest group ratings as percentages of "correct" roll call votes—"correct" as specified by the political predilections of a particular group—leads to what have been labeled cutpoint distortions. Such distortions affect the mapping from legislator preferences to interest group ratings, and they generate situations in which interest group ratings are not representative of legislator policy preferences.

The paper has described two types of cutpoint distortions. First, as noted by Snyder (1992), interest group ratings will often mischaracterize the difference between the median Democrat in a Congressional chamber and the corresponding median Republican. Second, interest group ratings will often mischaracterize the extent of intraparty preference heterogeneity and they will often fail to identify correctly which of the two parties, Democratic or Republican, is more homogeneous. Both types of cutpoint distortions are serious and they can lead to inaccuracies insofar as ratings are used to proxy for legislator policy preferences.

In light of cutpoint distortions, the paper proposes an adjustment technique that seeks to improve the fit between a set of interest group ratings and the corresponding legislator policy preferences that generated them. Under weak conditions, the technique produces adjusted interest group ratings that are more reflective of legislator preferences than unadjusted ratings, and application of the cutpoint adjustment to historical ADA ratings shows that the technique can have significant consequences for rating interpretation. Although for purposes of exposition this paper has focused on the popular ADA ratings, the measurement improvements elicited by the cutpoint adjustment procedure apply to all forms of interest group ratings.

Because it is not possible to verify the accuracy of interest group ratings, to some extent researchers will always face an arbitrary choice when selecting a preference measure for a study of Congress. Nonetheless, for the following three reasons, it is the contention of this paper that cutpoint-adjusted interest group ratings should be preferred to unadjusted ratings. First, as illustrated with ADA data from 1980 and 1992, the former portray a more intuitive picture of preference distributions within Congress: adjusted ratings are less polarized and less skewed compared to unadjusted ratings. Second, cutpoint-adjusted ratings are based on a model of interest group ratings whereas unadjusted ratings are essentially model- and theory-free. Third, there are very good reasons to believe that, within individual Congressional chamber sessions, interest group ratings will be inaccurate measures of legislator ideology. The cutpoint adjustment procedure seeks to correct for such inaccuracies.

Krehbiel (2000) argues that roll call-based legislator preference measures such as interest group ratings are limited in their ability to help researchers determine whether political party membership has an independent impact on roll call voting. His argument seems compelling, but it does not invalidate this paper's purpose of seeking to uncover the preferences that drive roll call voting in Congress. Whether or not these preferences are party-induced is an important issue in its own right, but it is not directly relevant to the objective here. Namely, it is argued in this paper that interest group ratings will do a poor job of approximating legislator policy preferences regardless of whether they are party-induced, constituency-based, or some combination of both. Under weak conditions, cutpoint-adjusted ratings will do a better job.

References

Aldrich, John H., and Richard D. McKelvey. 1977. "A Method of Scaling with Applications to the 1968 and 1972 Presidential Elections." *American Political Science Review* 71(1):111–130.

Binder, Sarah A., Eric D. Lawrence, and Forrest Maltzman. 1999. "Uncovering the Hidden Effect of Party." *Journal of Politics* 61(3):815–831.

Brady, David, Judith Goldstein, and Daniel Kessler. 1998. "Does Party Matter in Senators Voting Behavior: An Historical Test Using Tariff Votes over Three Institutional Periods." Paper presented at the annual meeting of the American Political Science Association, Boston.

Brunell, Thomas L., William Koetzle, John DiNardo, Bernard Grofman, and Scott L. Feld. 1999. "The $R^2 = .93$: Where Then Do They Differ? Comparing Liberal and Conservative Interest Group Ratings." *Legislative Studies Quarterly* 24(1):87–101.

Cox, Gary W., and Mathew D. McCubbins. 1993. *Legislative Leviathan: Party Government in the House*. Berkeley: University of California Press.

Fowler, Linda. 1982. "How Interest Groups Select Issues for Rating Voting Records of Members of the U.S. Congress." *Legislative Studies Quarterly* 7:401–413.

Groseclose, Tim. 1994. "Testing Committee Composition Hypotheses for the U.S. Congress." *Journal of Politics* 56(2):440–458.

Groseclose, Tim, Steven D. Levitt, and James M. Snyder, Jr. 1999. "Comparing Interest Group Scores across Time and Chambers: Adjusted ADA Scores for the U.S. Congress." *American Political Science Review* 93(1):33–50.

Hall, Richard L., and Bernard Grofman. 1990. "The Committee Assignment Process and the Conditional Nature of Committee Bias." *American Political Science Review* 84(4):1149–1166.

Hutchings, Vincent L. 1998. "Issue Salience and Support for Civil Rights Legislation Among Southern Democrats." *Legislative Studies Quarterly* 23(4):521–544.

Jackson, John E., and John W. Kingdon. 1992. "Ideology, Interest Group Scores, and Legislative Votes." *American Journal of Political Science* 36(3):805–823.

Krehbiel, Keith. 1992. *Information and Legislative Organization*. Ann Arbor: University of Michigan Press.

Krehbiel, Keith. 1995. "Cosponsors and Wafflers from A to Z." *American Journal of Political Science* 39(4):906–923.

Krehbiel, Keith. 1998. *Pivotal Politics: A Theory of U.S. Lawmaking*. Chicago: University of Chicago Press.

Krehbiel, Keith. 2000. "Party Discipline and Measures of Partisanship." *AJPS* 44(2):212–227.

Maltzman, Forrest. 1997. *Competing Principals: Committees, Parties, and the Organzation of Congress*. Ann Arbor: University of Michigan Press.

Moscardelli, Vincent G., Moshe Haspel, and Richard S. Wilke. 1998. "Party Building Through Campaign Finance Reform: Conditional Party Government in the 104th Congress." *Journal of Politics* 60(3):691–704.

Poole, Keith T. 1998. "Recovering a Basic Space from a Set of Issue Scales." *American Journal of Political Science* 42(3):954–993.

Poole, Keith T., and R. Steven Daniels. 1985. "Ideology, Party, and Voting in the U.S. Congress, 1959–1980." *American Political Science Review* 79(2):373–399.

Poole, Keith T., and Howard Rosenthal. 1997. *Congress: A Political-Economic History of Roll Call Voting*. New York: Oxford University Press.

Snyder, James M., Jr., and James M. 1992. "Artificial Extremism in Interest Group Ratings." *Legislative Studies Quarterly* 17(3):319–345.

Snyder, James M., Jr., and Tim Groseclose. 2000. "Estimating Party Pressure in Congressional Roll-Call Voting." *AJPS* 44(2):193–211.