

Decision-Making Under a Norm of Consensus: A Structural Analysis of Three-Judge Panels

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July 3, 2007

Abstract

This paper estimates a structural model of decision-making in judicial panels under a norm of consensus. Using data from asylum and sex discrimination cases in the courts of appeals, the model estimates ideology parameters for individual judges as well as a “cost” of dissent. I show that a positive cost of dissent for both the majority and the minority is necessary to reconcile the high rate of unanimity with the variation in individual judges’ voting records. The parameter estimates of the structural model show that the dissent rate substantially understates the actual level of disagreement within panels and that consensus voting obscures the impact of ideology on case outcomes. A significantly positive cost of dissent for the majority also implies that judges will sometimes compromise to avoid a dissent by another judge, and hence, that case outcomes are not determined purely by majority rule. The methodology developed in this paper can also be used to derive more accurate estimates of judicial ideology that control for consensus voting.

Appellate courts in the United States, like many deliberative bodies, operate under an informal norm of consensus. Judges value unanimity, and will often compromise in order to reach agreement with their colleagues. Thus, each judge’s vote in a case will be determined not only by that judge’s own preferences, but also by the preferences of the other judges on the court. This interaction poses a significant challenge for the empirical analysis of decision-making in multimember courts: when only final votes are observable, the determinants of judicial behavior may be obscured by the unobservable influence of group deliberation (Howard 1968). This difficulty is compounded by the fact

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that the primary influences on judicial behavior – ideology and the merits of a case – cannot be easily quantified.

To disentangle the effects of these influences on case outcomes, this paper develops a model of group voting under a norm of consensus. The model assumes that judges have one-dimensional ideological preferences, allows for unobserved heterogeneity among cases, and imposes a “cost” of dissent. Exploiting the random assignment of judges to cases, the paper estimates the structural parameters of the model using data sets of asylum and sex discrimination cases in the U.S. Courts of Appeals.

The results reveal a strong pattern of consensus voting: in the asylum data, the “cost” of writing a dissenting opinion is about twice the ideological “distance” between the median Democratic appointee and the median Republican appointee. The issuance of a dissent imposes a cost on the majority judges as well, creating a dynamic in which a majority may choose to compromise in the presence of a credible threat to dissent. The estimates derived in this paper thus provide strong evidence that the intensity of judges’ preferences influences case outcomes, and therefore, that the decision rule in three-judge panels departs from the median voter theorem.

The results also show that the courts are more ideologically polarized than judges’ voting records indicate. The intuition for this is simple: since the norm of consensus has a moderating influence on judges’ votes, controlling for it will yield ideology estimates for judges that are more extreme than their voting records.

The structural approach taken in this paper contrasts with the previous empirical literature on voting in judicial panels.¹ Because judicial ideology is not directly observable, these papers proxy for ideology using variables such as the party of the appointing president, “common space” ideology scores (Giles *et al.* 2001), race, and gender. This literature controls for the effects of group deliberation by regressing judges’ votes on the judge’s own characteristics and on the characteristics of the other judges on a panel. However, since the proxy variables capture ideology with measurement error, these estimates may have substantial bias.

Because the methodology developed in this paper directly estimates ideology for individual judges in a way that controls for group voting behavior and case heterogeneity, it can facilitate empirical research on judicial decision-making in situations where subtle influences on judicial decision-making might otherwise be obscured by the imprecision of proxy variables. This is especially important in situations where proxy variables are poor predictors² or where there is

¹Some recent examples include Cox and Miles (2007), Miles and Sunstein (2006), Peresie (2005), Sunstein, Schkade, and Ellman (2004), Farhang and Wawro (2004), Tiller and Cross (1998), Revesz (1997).

²For example, Sisk and Heise (2005) find that party affiliation has no predictive value for judges’ votes in religious freedom cases, although they find that other demographic variables are significant. Similarly, Sunstein *et al.* (2004) find that party affiliation is uncorrelated with votes in criminal appeals. In state courts, foreign courts, and earlier periods in history, judicial appointments may be less partisan, and these variables may be poor indicators.

insufficient variation in these variables.³

For example, the methods developed here could be useful to estimate how judicial behavior is influenced by career concerns⁴ or appellate review.⁵ The ideology estimates derived in this paper can be used to construct a measure of “judicial bias” that can be used to test its impact on contracting decisions,⁶ the evolution of law⁷, and the incidence of litigation.⁸ These estimates may be used to construct instrumental variables for case outcomes whenever judges are randomly assigned to cases.⁹ Finally, the analysis provided here may be applicable in other settings in which consensus voting is observed, such as in regulatory agencies, academic committees, and the Federal Open Market Committee.¹⁰

The results here are also relevant to the economic literature on committee decision-making. Much of this literature addresses the aggregation of information, particularly when acquisition or transmission are costly; the results here suggest that the process of preference aggregation in collegial committee settings warrants closer examination. The results here, derived from a natural experiment involving expert decision-makers and real stakes, complement the experimental literature on committee decision-making.¹¹

The paper is organized as follows. Section 1 constructs a model of decision-making in appellate panels, in which judges balance their ideological preferences with their desire for unanimity. Section 2 discusses the technique for estimating the parameters of the model. Section 3 describes the data. Section 4 discusses the estimation results for the judges and for the parameters governing the dynamics of panel decision-making. Section 5 uses the model and the estimated parameters to make out-of-sample predictions. The first part of Section 5 estimates how often a unanimous decision results from actual agreement among the judges, as opposed to judicial compromise, and how often the majority position dominates when compromise occurs. The second part compares the distribution of outcomes when cases are decided by single judges and three-judge panels, to provide a sense of how panel decision-making moderates rulings. Section 6 concludes. All proofs are in the appendix.

1 Conceptual Framework

We model the norm of consensus using a “cost” of dissent. This may be interpreted most easily as a disutility from conflict. Judges themselves have often

³This is especially true in some state courts, where one party has dominated appointments. For example, in most southern states for the century after Reconstruction, there would be little or no variation among judges in race, gender, or party of appointment.

⁴See e.g., Levy (2005).

⁵See e.g., Shavell (2006), McCubbins *et al.* (1995).

⁶See e.g., Gennaioli (2005).

⁷See e.g., Gennaioli and Shleifer (2007a, b), Ponzetto and Fernandez (2007).

⁸See e.g., Gennaioli and Shleifer (2006).

⁹Some recent examples of this research design include Kling (2006) and Schoar and Chang (2006).

¹⁰See e.g., Belden (1989), Havrilesky and Gildea (1991), Chappell *et al.*, (2004).

¹¹See e.g., Kahneman *et al.* (1998), Plott & Fiorina (1978).

emphasized the important of “collegiality” in appellate courts (e.g., Edwards 2003, Coffin 1994).

The cost of dissent can also be motivated by the effort required to write, or respond to, a dissenting opinion. (Posner 1993). A dissenting opinion weakens the legitimacy of the panel’s ruling, and frequent dissents can diminish the authority of the court. Dissenting opinions also impose a cost to the majority by increasing the likelihood that the ruling will be overturned. (Kastellec 2007, Cross & Tiller 1998) For the dissenting judge, the “cost” of dissent also reflects the fact that issuing too many dissents may diminish their signaling value; Ginsburg (1990) refers to this as the “danger of crying wolf too often.”

In the spatial model we develop here, judges’ utility is based on one-dimensional ideological preferences and the cost of dissent. Let each case be represented as a cutoff point η on the real line, where each judge’s preferred outcome in a case depends on her position relative to the cutoff point. Similarly, we represent judge i ’s propensity to favor plaintiffs as a point a_i in ideology space. Each judge may choose a ruling $v_i \in \{P, D\}$, representing a vote in favor of the plaintiff or defendant, respectively.¹² When $a_i > \eta$, judge i will prefer to vote P , and when $a_i < \eta$, judge i will prefer to vote D . In the context of asylum cases, we can think of judges with greater a_i as being more sympathetic to asylum claimants, and judges with lower a_i as being more sympathetic to the government. Cases with lower η are stronger cases for the asylum claimant.

We represent the ideological component of the judge’s utility as follows: the judge gets utility 0 from voting in her preferred direction, and utility $-|a_i - \eta|$ from voting against her preferred position. We model the preference for consensus by imposing a cost c_d on the dissenting judge and a cost c_m on each of the majority judges when there is a dissent, where we restrict $c_m \leq c_d$.¹³ Each judge’s total utility is therefore

$$U_i = \begin{cases} \max\{0, a_i - \eta\} - I_m c_m - I_d c_d, & \text{if } v_i = P \\ \max\{0, \eta - a_i\} - I_m c_m - I_d c_d, & \text{if } v_i = D \end{cases}$$

where I_m and I_d are indicator variables for judge i being in the majority and the dissent, respectively. When cases are unanimous, $I_m = I_d = 0$.

Deliberation is modeled as a multi-stage game. At stage zero, nature randomly chooses a voting order for the judges. Each permutation is equally likely, and the voting order is common knowledge. At stages one through three, the judges cast their votes, which are irrevocable.¹⁴

¹²In cases at the appellate level, the term “plaintiff” will always denote the original plaintiff in the case.

¹³We impose this restriction to ensure that, in the case of a split, a judge always prefers to be in the majority. Otherwise, a situation could arise in which two opposing judges each try to push the pivotal judge toward the other, a scenario we find implausible.

Note that the likelihood function will take the same form in the case where $c_m > c_d$, so the model can still be estimated.

¹⁴We model deliberation as a multi-stage game with irrevocable votes in order to guarantee a unique subgame perfect equilibrium. Although it would be natural to model deliberation as a cooperative game, the possibility of multiple equilibria makes estimation impractical. Note that when a cooperative model of deliberation has a unique equilibrium, it will coincide with

When $\eta < \min\{a_i\}$ or $\max\{a_i\} < \eta$, the decision will be unanimous: all of the judges will have the same preference, and none of them will have an incentive to switch sides. The following results focus on the case where there is *ex ante* disagreement. Let d denote the judge in the minority position, m_1 denote the median judge, and m_2 denote the more extreme majority judge, so that $a_d < \eta < a_{m_1} < a_{m_2}$ or $a_d > \eta > a_{m_1} > a_{m_2}$.

The following cases will describe the unique subgame perfect equilibrium. All proofs are provided in the appendix.

Case 1 *The minority judge will always switch sides if $|a_d - \eta| < c_d$ and either $|a_{m_1} - \eta| > c_m$ or $|a_{m_2} - \eta| > c_d$.*

This case captures the notion of the “collegial concurrence”: from a judge’s perspective, not every case of disagreement merits a dissenting opinion. When the “ideological distance” between the judge’s preferred decision rule and the case cutoff is small, the minority judge will forgo the dissent. Although the content of judicial opinions is not incorporated into the model, it is plausible that a judge may withhold a dissent in order to negotiate for a more limited holding. The latter two conditions ensure that at least one of the majority judges is unwilling to compromise; otherwise, case 4 applies.

Case 2 *The judges in the majority will always vote against their preferences, and the minority judge will prevail, if $|a_d - \eta| > c_d$ and $|a_{m_2} - \eta| < c_m$.*

This case captures the possibility of “minority rule.” If the minority judge feels strongly enough to dissent, and the majority’s preferences are weak, they would choose to give in to the minority rather than face a dissenting opinion.

Case 3 *All judges will vote their true preferences, and the minority judge will dissent, if $|a_d - \eta| > c_d$ and either $|a_{m_1} - \eta| > c_m$ or $|a_{m_2} - \eta| > c_d$.*

This case corresponds to the situation in which compromise is impossible. The condition $|a_{m_1} - \eta| > c_m$ guarantees that neither majority judge is sufficiently averse to a dissent to switch sides. The condition $|a_{m_2} - \eta| > c_d$ means that the more extreme majority judge cannot be reconciled with the majority judge, so that consensus is impossible. Although the case cutoff is unobserved in the data, this condition requires that there be sufficient ideological separation between the sides in order for a dissent to occur.

Case 4 *If $|a_d - \eta| < c_d$ and $|a_{m_2} - \eta| < c_m$, there will always be a unanimous outcome. The position of the minority judge will prevail only if judge d votes first; otherwise, the panel will vote in favor of the majority position.*

Here, all of the judges have weak ideological preferences in the case at hand, and no one is willing to dissent; the only question is which side prevails. The

the above game; when a cooperative model has multiple equilibria, the random voting order “assigns” probabilities to each of them.

case outcome is determined in the model by the sequential voting format: the judge who gets to vote first determines the direction of the panel decision, and the other judges will follow.

Case 5 *If $|a_d - \eta| > c_d$, $|a_{m_1} - \eta| < c_m$ and $c_m < |a_{m_2} - \eta| < c_d$, then there will be a unanimous vote in favor of the minority position if judge m_1 votes before m_2 ; otherwise, all judges will vote their true preferences, and the minority judge will dissent.*

Here, the minority judge will not switch sides; the only question is what the majority judges do. The judges in the majority will always vote together, but judge m_2 prefers to vote sincerely and let the minority judge dissent, while judge m_1 prefers to switch sides. Thus, the voting order of these two judges determines the equilibrium outcome.

Case 6 *If $|a_d - \eta| < c_d$, $|a_{m_1} - \eta| < c_m$ and $c_m < |a_{m_2} - \eta| < c_d$, there will always be a unanimous outcome. The position of the minority judge will prevail only if the voting order is (d, m_1, m_2) ; otherwise, the panel will vote in favor of the majority position.*

This is similar to the last case, except that the minority judge is no longer willing to dissent. The only way that the minority position can prevail is if judge d votes first and judge m_1 votes before m_2 .

The probability of equilibrium outcomes is illustrated in Figures 1–2. Figure 1 shows the case where the minority judge is within a distance of c_d of the case cutoff. In this case, the minority judge is always willing to compromise, and deliberation will always result in a unanimous opinion. The only times that the minority position can prevail are when the majority judges also strictly prefer consensus. The lower-left region in Figure 1 illustrates case 4; there will be a $\frac{1}{3}$ probability of a unanimous pro-defendant vote. The adjacent regions illustrate case 6, and therefore have a $\frac{1}{6}$ probability of a unanimous pro-defendant vote.

Figure 2 shows the case in which the minority judge is further than c_d from the case cutoff. In this case, the minority judge will never switch sides, and the only question is whether the majority will capitulate in order to avoid a dissent. In Figure 2, the lower-left region corresponds to case 2; since the threat of dissent is credible, the majority judges will always yield to the minority. The adjacent regions correspond to case 5; the probability of each outcome will be $\frac{1}{2}$, depending on which of the majority judges votes first.

The following corollaries provide the conditions for independent voting and majority rule in panels.

Corollary 7 *When $c_d = c_m = 0$, judges will vote in favor of their preferred outcomes, and will not be influenced by the other judges on the panel.*

Proof. When $c_d = c_m = 0$, disagreement is not costly, and hence all judges will vote in favor of their preferred outcomes. ■

Corollary 8 *When $c_m = 0$, all case outcomes will be determined by majority rule.*

Proof. When $c_m = 0$, the conditions in cases 2, 4, 5, and 6 can never hold. Since these cases are the only ones that allow for the panel to vote in favor of the minority position, the panel decision must coincide with the majority’s preferred outcome. ■

Corollary 7 shows that when $c_d = c_m = 0$, there will be desire to reach consensus, and we will not observe judges influencing each others’ votes in the data. All cases will be decided by majority vote, and the proportion of unanimous rulings observed in the data will reflect the actual degree of agreement among the judges. Because the hypothesis $c_d = c_m = 0$ can be tested on the data, the model provides a general test for a consensus voting in judicial panels that does not rely on proxy variables for ideology, as in previous studies.

When $c_m = 0$ and $c_d > 0$, corollary 8 shows that the minority judge may still switch sides, and judges will still influence each others’ votes, but case outcomes will be the same as under independent voting. A test of $c_m = 0$ on the data therefore shows if panel decisions adhere to majority rule, or if preference intensity has an effect on case outcomes.

2 Estimation

Several additional assumptions will be necessary to estimate the above model. To allow for the possibility that judges’ positions may vary over cases, we assume $a_{it} = \alpha_i + \varepsilon_{it}$, where $\varepsilon_{it} \sim N(0, 1)$.¹⁵ The parameter α_i represents the ideology of judge i , where a greater α_i corresponds to a more liberal judge.

A second complication arises because the case cutoff η_t is not observable. We treat it as a random variable, with $\eta_t \sim N(0, \sigma^2)$, where σ is a parameter to be estimated. This accounts for the fact that the judges’ positions in a particular case may be highly correlated; in some cases, liberals and conservatives will agree on the merits. The parameter σ represents the magnitude of the correlated component of the judges’ preferences: a larger σ (keeping all other parameters constant) means that there will be more “easy cases” and consequently more *ex ante* agreement.

The assumption that the distribution of case cutoffs has mean zero is without loss of generality, since no restrictions are imposed on the judge ideology parameters. When characteristics of individual cases are observed in the data, however, it is possible to allow the mean to vary with these characteristics. In this case, we let $\eta_t \sim N(x\beta, \sigma^2)$, where x is a vector of characteristics, and β is a vector of coefficients to be estimated.

Given that the model predicts the probability of each outcome conditional on the judge positions $\{a_i\}$ and the case cutoff η_t , we can construct a likelihood function that predicts the probability of each outcome given the judge ideologies

¹⁵Because the model is invariant to scale transformations, we normalize the variance of the error term to be 1 to ensure that the model is identified.

$\{\alpha_i\}$ and σ . The technical details for deriving the likelihood function are in the appendix.

An additional complication arises when particular judges appear only a small number of times in the data. Estimation of the ideology parameter α_i for such a judge would lead to imprecise ideology estimates for that judge, as well as bias in the estimates of the other parameters due to the “incidental parameter problem.” (Neyman and Scott 1948) When there are fewer than a threshold number of observations for a particular judge, we can assign this judge to a cluster based on observable characteristics. For this data, judges who appear infrequently are grouped into clusters of Democratic and Republican appointees. For these judges, we incorporate an additional error term into the likelihood function to account for within-cluster heterogeneity.

The parameters to be estimated are the judge ideology parameters α_i , the costs of disagreement c_d, c_m , and the standard deviation of the random effect, σ . The log-likelihood function is maximized using Newton’s method with multiple starting points to ensure a global optimum. Standard errors are derived from the inverse Hessian matrix, evaluated at the maximum likelihood estimate.

3 Data

The main results in this paper come from a data set of 1892 asylum cases decided by the Ninth Circuit Court of Appeals between 1992 and 2000.¹⁶ The legal standard for asylum cases – that the petitioner must demonstrate a “well-founded fear of persecution on account of race, religion, nationality, membership in a particular social group, or political opinion” – is relatively malleable and allows judges substantial discretion. At the same time, these cases are highly fact-specific, and have relatively low salience, meaning that the cost of dissent in a particular case may be high relative to the ideological “benefit.” The asylum data includes unpublished as well as published opinions.¹⁷

In asylum cases, there are typically no intermediate outcomes – the claimant either wins asylum or is deported – and the settlement rate is low.¹⁸ This is important because studies have shown in other contexts that judges may influence both case outcomes and settlement decisions (e.g., Waldfogel 1995, de Figueiredo 2005). Moreover, the panel of judges is announced less than two weeks before oral argument in the Ninth Circuit (Revesz 2000), after most of the litigation costs have been sunk. Thus, selection bias due to settlement should be minimal.

The overall rate of asylum grants during the period of study is 18%. We derive individual ideology estimates for the 65 judges who appear at least 10 times in the asylum data. Figure 4 provides a histogram of asylum grant rates

¹⁶This data was previously examined in Law (2005, 2006).

¹⁷In fact, 92% of the decisions are unpublished. Note that “unpublished” decisions are still accompanied by written opinions, but these opinions may not be cited as precedent in future cases.

¹⁸The settlement rate for asylum cases in the Ninth Circuit ranged between 6% and 8% from 1994 to 1999, and rose to 14% in 2000 (Palmer et al. 2005).

by judge. Most judges have grant rates in the range of 5-30%, and 95% of the asylum cases are decided unanimously.

The second data set consists of 977 sex discrimination and sexual harassment cases decided by three-judge panels in the U.S. Courts of Appeals between 1995 and 2002. The data set was taken from a larger data set collected and made available by the Chicago Judges Project of the University of Chicago Law School.¹⁹ This data set includes only published opinions, and may therefore not be fully representative of sex discrimination cases.²⁰ Nevertheless, the estimates for the sex discrimination data provide a useful comparison to the asylum data.

There are 110 judges who participated in at least 10 cases in the sex discrimination data set; the rest are grouped together by party of the appointing president. Plaintiffs won 40% of the sex discrimination cases, and 93% of these were decided unanimously. Figure 5 provides a histogram with pro-plaintiff voting rates by judge.

For each case, the data provides the identities and votes of the three judges. A case is coded as “*P*” if the court provides *any* relief to the plaintiff; otherwise, it is coded as a “*D*.”²¹ The asylum data also includes country of origin, which is coded using the Freedom House Civil Liberties Index to measure political conditions.

3.1 Random Assignment

A key identifying assumption in this paper is that judges are randomly assigned to panels and to cases. Without random assignment, the model would be unidentified due to the “reflection problem” (Manski 1993); it would be impossible to distinguish between selection of judges into cases and the effect of ideology and consensus voting.

According to Ninth Circuit Rules, the random assignment of judges to cases is done by computer. The algorithm randomly assigns three judges to each panel, subject to the constraints that workload should be equalized among the judges, and each pair of judges should work together an approximately equal number of times during a two-year period. Each panel may include senior judges

¹⁹The cases were gathered using a Lexis search of the terms “sex! discrimination” and “sex! harassment.” (Sunstein et al. 2004) The 9th Circuit and the D.C. Circuit removed because there are too few observations per judge in these circuits to be able to derive meaningful estimates. In addition, one case was removed from the data because it was decided by only two judges after the death of the third judge.

²⁰An examination of cases from a subsample of the period showed that about 83% of all sex discrimination decisions were published. Several studies have shown that published opinions are not be a relevant sample of all decisions. (Law 2005, Siegelman and Donohue 1990) Also, since the decision to publish is made by the panel deciding the case, it is possible that unobserved case characteristics could be correlated with the ideologies of the judges on the panel.

²¹Note that this method of coding overstates the degree of unanimity in the data. Some cases may be coded as (*P, P, P*) even though a judge may have dissented on some portion of the ruling. This seems to be rare in the asylum data but does occur occasionally in the sex discrimination data.

and at most one judge sitting by designation, but every panel must have at least one active Ninth Circuit judge.²² The only exception to random assignment occurs when a case returns to the appeals court after a previous appeal; in this case, it will typically be assigned to the panel that heard the first appeal.

Cases are grouped together on the court calendar by issue and complexity, in order to ensure an even workload. Cases presenting similar issues may be grouped together. Thus, a single panel may hear multiple asylum appeals from claimants of the same nationality. Although the grouping of cases together is nonrandom, the random assignment of these cases to panels ensures that case characteristics will be uncorrelated with the ideologies of the judges.

As an additional test of random assignment, we examine the number of Democratic judges assigned to claimants for each country of origin. Because of the way cases are clustered – a single panel may hear up to 50 asylum cases together, many of which may be from the same country – we test this hypothesis using Monte Carlo simulation.²³

For each country of origin, we examine all judges assigned to claimants from that country, and calculate the percentage of those judges who are Democratic appointees. Under the assumption of random assignment, that percentage should be close to the percentage of Democratic appointees in the entire pool. Then, we simulate the assignment of cases by randomly matching clusters of cases with three-judge panels that decided cases in the same year,²⁴ and calculate the percentage of Democratic appointees associated with claimants from each country in each simulated matching. By repeating this random assignment simulation many times, we obtain a simulated distribution of the number of Democratic judges assigned to claimants from each country.

For each country of origin, we compare the actual percentage of Democratic appointees for each country with the simulated distribution. For country k , let x_k be the actual percentage of Democratic appointees, and $F_k(\cdot)$ be the empirical cumulative distribution derived from the simulations. Then under the assumption of random assignment, we should expect the test statistic $F_k(x_k)$ to be uniformly distributed.

For example, there were 318 claimants from Nicaragua in the data, and 40% of the judges assigned to those cases were Democrats. In the simulated matchings, Nicaraguan claimants were assigned a lower percentage of Democrats 29.5% of the time, and a higher percentage 70.5% of the time. Thus, the test statistic for Nicaragua would be 0.295.

We now compute these test statistics for the 26 most prominent countries of origin,²⁵ and an additional test statistic for all remaining countries grouped

²²Judges sitting by designation are either district judges or judges from other circuits.

²³The simulation method here closely follows the method developed in Abrams, et al (2006).

²⁴The random matching of panels and cases within the same year is necessary because the Ninth Circuit trended Democratic during much of the period of study and migration from various countries peaked at different times.

Note that the year of assignment is not available in the data. We therefore designate a year for a cluster using the mean year of filing for cases within the cluster, rounded off to the closest year.

²⁵These are the countries with at least ten cases.

together. The minimum test statistic is 0.005 and the maximum is 0.954. Figure 3 shows the empirical cumulative distribution of the test statistics, which closely resembles the uniform distribution. A chi-square test and a Kolmogorov-Smirnov test both fail to reject the assumption that these test statistics are distributed uniformly. Thus, these simulation results are consistent with the assumption that cases are randomly assigned to judges.

4 Results

4.1 Judge Ideology

For each judge estimated individually, we report α_i , the location in ideology space, and p_i , the probability that the judge would vote in favor of the plaintiff in a case selected at random, where

$$p_i = \int \Phi(\alpha_i + \sigma\eta)\phi(\eta) d\eta \quad (1)$$

The ideology estimates and probabilities for all judges in the asylum cases can be found in Table 1.²⁶ Although the standard errors are large for many of the individual judges, the heterogeneity in grant preferences is quite evident. Figure 6 shows the distribution of preferences in a histogram. There is a large spike in the range of 0-5%, corresponding to about one-fourth of the judges who very rarely vote in favor of asylum. The rest of the judges are distributed somewhat evenly over the rest of the distribution, with a small dropoff above 45%.

The disparities in voting rates among the judges is dramatic. On the pro-asylum extreme, Judge Reinhardt voted in favor of asylum 62% of the time, compared to 4% of the time for Judge Wallace. However, these voting rates include those cases in which these judges may have compromised with their colleagues on the panel. The estimated probabilities of these judges favoring asylum – how they would behave if their vote were pivotal – are even more extreme: 92% for Judge Reinhardt and 1% for Judge Wallace. The disparities at the panel level are also extreme: the most liberal panel in the data has a predicted pro-asylum probability of 85%, while there are over 100 panels with a predicted pro-asylum rate of less than 1%.

The histogram in Figure 7 shows the dispersion of probabilities for the individual judges in the sex discrimination data. Note that the dispersion of predicted probabilities is much wider than the dispersion of actual votes in Figure 5. Since panel deliberation has a moderating effect on judges' votes, controlling for these effects reveals judges to be more extreme in both data sets than their voting records would suggest.

Figure 8 provides a histogram comparing Democratic and Republican appointees in asylum cases. Although Democratic appointees are more favorable

²⁶For simplicity of exposition, the data on Freedom House Civil Liberties scores is ignored in the asylum analysis, except in section 4.3.

toward granting asylum, there is significant heterogeneity within both parties. Most notably, both parties are heavily represented in the 0-5% range. Figure 9 provides the same party breakdown for sex discrimination. While there is still significant within-party heterogeneity, the differences between Democrats and Republicans is instantly evident in this graph.

The individual estimates derived here may also be used to compare aggregate differences among groups of judges. For example, the mean probability of a pro-asylum vote is 36.6% for Democratic appointees and 18.7% for Republican appointees, a statistically significant difference.²⁷ In sex discrimination cases, the mean voting rates for Democratic appointees is 63.5% compared to 29.4% for Republican appointees, also a significant difference.²⁸ The difference between male and female judges is significant in sex discrimination cases (39.9% for males compared to 58.4% for females), but not in asylum cases (28.1% for males compared to 33.9% for females).²⁹

These results have significant implications for empirical research that relies on political party as a proxy for ideology. Differences between Democratic and Republican appointees are indeed significant, and party is therefore a valid predictor of voting behavior in sex discrimination cases. However, there is substantial within-party heterogeneity, which means that using party as a proxy variable will lead to substantial measurement error.

4.2 Structural Parameters

Table 2 also provides estimates of the distribution of ideology parameters, as well as c_d and c_m , the costs of dissent, and σ , the standard deviation of case cutoffs. In the asylum cases, an outvoted judge in the minority is willing to “travel” an ideological distance of $c_d = 1.71$ in order to avoid issuing a dissent. Judges in the majority will travel a distance of $c_m = 1.36$ to avoid a dissenting opinion. These distances are both quite substantial compared to the distance of 0.85 between the median Democrat and the median Republican.

The fact that these costs of dissent are large compared to the degree of ideological heterogeneity suggests that many unanimous opinions may be the result of compromise rather than sincere agreement. Thus, although ideology still has a large impact on case outcomes, it has a much more muted effect on individual votes. Any method of estimating ideology that does not control for consensus voting will thus provide biased estimates.

In the sex discrimination data, the cost of dissenting ($c_d = 3.21$) is statistically significant and roughly comparable to the distance between the median Democrat and the median Republican (a distance of 2.98). However, the cost of dissent to the majority is not statistically significant from zero.³⁰ One striking

²⁷The standard error of the difference is 3.1%.

²⁸The standard error of the difference is 13.0%.

²⁹The standard error on the difference is 3.8% in the asylum cases and 8.9% in the sex discrimination cases. For an alternative analysis of voting by gender in the Sunstein sex discrimination data, with similar conclusions, see Boyd et al. (2007).

³⁰In fact, the likelihood function is maximized at a point where the restriction $c_m \geq 0$ is

difference between the asylum data and the sex discrimination data is the estimate of the case-specific random effect σ . In the sex discrimination estimates, σ is significant and of the same order of magnitude as the ideology estimates; in the asylum cases, σ is much smaller than the ideology parameters. Moreover, a likelihood ratio test on $\sigma = 0$ fails to reject in the asylum data but overwhelmingly rejects in the sex discrimination data.³¹

The variability in case cutoffs in the sex discrimination data shows that there are unobserved case characteristics that shift judges' preferences in a correlated fashion. Although we do observe these in the data, this suggests a role for institutional constraints on judicial decision-making; judges may be bound in a common way by text, precedent, or appellate oversight. On the other hand, this behavior seems to be absent in the asylum data, suggesting that case characteristics play a much weaker role in asylum cases. Although unpublished opinions are commonly believed to involve straightforward applications of the law, the opposite seems to be the case here: the low salience and low probability of further review may give judges greater latitude to indulge their own preferences.

In the asylum data, but not the sex discrimination, the cost of dissent for the majority is positive and significant: a likelihood ratio test clearly rejects the restriction $c_m = 0$ ³². This means that some majorities will be willing to compromise in the presence of a credible threat to dissent, and therefore, that consensus voting changes case outcomes as well as individual votes. These empirical results stand in contrast to the formal decision rule: that cases are decided by majority vote. In collegial settings, when agents are averse to conflict, committee decisions may thus derive from a process that departs from the formal decision rule. The greater willingness to compromise in asylum cases suggests that judges are more willing to defer to their colleagues in non-precedential, low-salience cases.

Note that we can easily reject the restriction $c_d = c_m = 0$, corresponding to the assumption that judges vote independently.³³ This is consistent with the conclusion in Sunstein et al. (2004), although the analysis here does not rely on the use of proxy variables for ideology.

4.3 Case Characteristics

For simplicity, the model used to derive the estimates in the previous section did not incorporate case characteristics. Since cases are randomly assigned to judges, and hence each judge should have a similar distribution of cases, these

binding.

³¹The χ^2_1 test statistic is 1.6 for the asylum data and 22.6 for the sex discrimination data, compared to the 99% critical value of 6.6.

³²The test is significant at the level $p = 0.05$ in the sex discrimination data. For the asylum data, it is highly significant: the chi-square test statistic is 33.2, far above the 99% critical value of 6.3.

³³The χ^2_2 test statistic is 416.4 for the asylum data and 164.1 for the sex discrimination data, compared to the 99% critical value of 9.2.

characteristics are not necessary to estimate ideology, but they offer the possibility of increased precision. Furthermore, incorporating case characteristics directly into the model enables us to estimate the impact of case factors on outcomes.

Although we do not have information on individual cases in the sex discrimination data, we do have country of origin for the asylum cases, which is coded using the Freedom House Civil Liberties Index (henceforth “CL”). This index codes countries on a scale from 1 to 7, where a rating of 1 corresponds to the “most free” countries and 7 corresponds to the “least free” countries. We estimate the data with dummy variables for each CL score (where the dummy for the CL = 7 is omitted).

The results are shown in table 3.³⁴ These estimates show only a weak relationship between the Civil Liberties Index and asylum outcomes. For cases from the second-highest category countries (CL = 2), there was a significantly negative effect on grant rates. The only other estimate which is statistically significant is the coefficient for the second worst category (CL = 6), which is also negative. Note that all of these coefficients are small compared to the magnitude of the ideology parameters. The difference in impact between a country with a CL score of 2 and a country with the CL score of 7 is less than half the difference between the median Democratic judge and the median Republican judge.

The small magnitude of the coefficients on the CL scores should be interpreted in light of the fact that the cases presented for appeal constitute a non-random subsample of all asylum cases. The data only includes claimants who were initially denied asylum and subsequently appealed. Thus, the effect of unobservable case characteristics may be negatively correlated with the effect of the CL scores.

5 Simulations

One benefit of a structural model is that it lends itself to simulations and out-of-sample predictions. In this section, we will explore a few applications. First, we will estimate the frequency of compromise. We will use Monte Carlo simulation to estimate the proportion of unanimous decisions that were the result of a compromise by one or more of the judges. Second, we will estimate how appellate rulings would be different if they were decided by a single judge rather than a three-judge panel. By simulating these results for the actual panels in the data, and also for individual judges, we can get a sense of how panel decision-making increases predictability in rulings.

³⁴The category corresponding to CL = 1 is dropped because there are only two claims from the “most free” category, both of which were denied asylum.

5.1 Dynamics of Compromise

To understand the degree of fluidity in panel votes we estimate the probability that one or two judges would have switched votes in order to achieve unanimity. We estimate these probabilities using Monte Carlo simulation, generating random values for η_t and $\varepsilon_{it}, \varepsilon_{jt}, \varepsilon_{kt}$ for the actual cases that occur in the data and including only those values that result in the observed vote. For these simulated results, we use Bayes' rule to estimate the probability of each combination of initial preferences. The results are presented in Table 4.

Several observations can be made from these results. First, although more than 90% of cases were decided unanimously, the judges initially agreed in only about 49% of the asylum cases and 57% of the cases. In the cases in which the judges did not initially agree, they were able to reach agreement more than 80% of the time. These results reveal a large gap between observed rates of unanimity and the predicted rate of agreement.

Second, when the judges did compromise, they settled on the minority position roughly one-third of the time in both data sets. Although cases are formally decided by majority rule, this suggests that panel decision-making in practice can lead to substantially different results. This result is consistent with the example of "leisure-seeking" judges given in Posner (1993), in which two relatively "indifferent" judges go along with an "opinionated" judge in order to avoid having to respond to a dissenting opinion.

Third, we can estimate that about 20% of all votes changed during deliberations in asylum cases. Note that if votes are interpreted "sincerely," without any correction for collegial voting behavior, this would be equivalent to a "garbling" of 15-20% of votes.

5.2 Individual Judge vs. Three-Judge Panels

In the federal courts, appeals are usually decided by three-judge panels. An obvious justification for using multimember courts is that they should provide more consistency than individual judges. In this section, we can compare simulated outcomes from single judges and three-judge panels.

We estimate the probability the plaintiff has of winning based on the strength of her claim, using actual panel compositions that were observed in the data but varying the case cutoff. Figure 10 shows the probability of a ruling in favor of asylum, using the parameters estimated in Section 4, conditional on a given case cutoff η . We can also estimate the probability of each outcome if cases were decided by a single judge. For simulated single-judge rulings, we weight the likelihood of a judge being assigned by the number of times that judge occurred in the data. The solid line in figure 10 shows the simulated probability of a ruling in favor of the plaintiff for three-judge panels; the dashed line shows the probability for a single judge.

In figure 10, the horizontal axis measures the quantile of η , the case cutoff, which represents the strength of the plaintiff's claim. The point 0 on the x -axis represents the weakest claim, and the point 1 represents the strongest claim.

The probability of winning is less than 10% for the weakest 40% of cases. For cases in the middle, however, there is substantial uncertainty, and the outcome will hinge on which judges are chosen to hear the case. Even for the very strongest claims, there is a substantial chance that asylum will be denied if the claimant draws an unsympathetic panel.

The dashed line in figure 10 shows the probability of outcomes if the cases were decided by a single judge chosen at random. Note that the outcomes are somewhat less predictable under this assumption. These results confirm the intuition that a panel of three judges will rule more moderately and more predictably than a single judge.

Figure 11 provides the corresponding graph for sex discrimination cases. Here, there is a bit more predictability at both extremes, but the graph still suggests that the outcome of most cases will depend on the panel of judges who hear the case.

A central feature of the rule of law is that judges must apply the law consistently. If the legal standards employed by judges were in fact perfectly uniform, then the graphs in figures 10 and 11 would resemble step functions: below a certain threshold, all claims would be denied, and above that threshold, all claims would be granted. While perfectly uniform enforcement is never attainable, both of the graphs reveal a gap between the application of the law in practice and the ideal under the rule of law.

6 Conclusion

This paper uses a structural model to analyze deliberation in three-judge panels, estimating a “cost” of dissent as well as ideology parameters for individual judges. The results show that ideology, case characteristics, and collegial influences all have a significant impact on judicial decision-making.

The model here could potentially be extended to incorporate interaction between different levels of the judicial hierarchy. Another possible extension would be to allow for some heterogeneity in the cost of dissent across judges, cases, or judicial circuits. With richer data on individual cases, it may also be possible to more precisely estimate the influence of ideology and case factors on outcomes.

The estimates derived here may also be used directly as ideology “scores” in other settings. When ideology is correlated across areas of law, these estimates may be used as alternatives to traditional proxy variables for predicting voting behavior. This will be especially relevant in areas of law in which data is sparse, and the methods developed here are not applicable.

A Appendix

For simplicity in the following proofs, we assume that $a_d < \eta < a_{m_1} < a_{m_2}$. The case of $a_d > \eta > a_{m_1} > a_{m_2}$ is dealt with symmetrically.

PROOF OF CASE 1: The minority judge will always switch sides if $|a_d - \eta| < c_d$ and either $|a_{m_1} - \eta| > c_m$ or $|a_{m_2} - \eta| > c_d$.

The minority judge's utility from switching will be $-|a_d - \eta|$, and the utility from not switching will be $|a_d - \eta| - c_d$. Hence the minority judge will switch sides if $|a_d - \eta| < c_d$ and if she believes that the other judges will not change their votes. Similarly, when $|a_{m_2} - \eta| > c_d$, judge m_2 will always vote P , irrespective of the other judges' votes. Since $a_{m_1} > \eta$, judge m_1 will always vote P if judge m_2 votes P . When $|a_{m_1} - \eta| > c_m$, judges m_1 and m_2 would both vote P even if judge d dissented. Thus, either of the conditions in the proposition are sufficient to induce judge d to switch sides.

PROOF OF CASE 2: The judges in the majority will always vote against their preferences, and the minority judge will prevail, if $|a_d - \eta| > c_d$ and $|a_{m_2} - \eta| < c_m$.

In this case, judge d will always have higher utility from not switching. Given judge d 's position, judges m_1 and m_2 will have utility $-|a_{m_i} - \eta|$ from switching and utility $|a_{m_i} - \eta| - c_m$ from voting their preferred positions, provided that the other majority judge plays the same action. The conditions in the proposition guarantee that both will strictly prefer switching; hence, this will be the unique subgame perfect equilibrium.

PROOF OF CASE 3: All judges will vote their true preferences, and the minority judge will dissent, if $|a_d - \eta| > c_d$ and either $|a_{m_1} - \eta| > c_m$ or $|a_{m_2} - \eta| > c_d$.

If $|a_d - \eta| > c_d$ and $|a_{m_2} - \eta| > c_d$, then these two judges have an irreconcilable disagreement, and one of them will dissent in either case. Since the median judge will incur the cost c_m in either case, she will maximize her utility by voting in favor of her preferred outcome. If $|a_d - \eta| > c_d$ and $|a_{m_1} - \eta| > c_m$, then judges m_1 and m_2 have utility $|a_{m_i} - \eta| - c_m$ if they both vote in favor of their preferred outcome, and a maximum utility of $-|a_{m_1} - \eta|$ if they join the minority judge. Thus voting their true preference will always be the dominant strategy.

PROOF OF CASE 4: If $|a_d - \eta| < c_d$ and $|a_{m_2} - \eta| < c_m$, there will always be a unanimous outcome. The position of the minority judge will prevail only if judge d votes first; otherwise, the panel will vote in favor of the majority position.

Note that under these conditions, each judge strictly prefers every unanimous outcome over every nonunanimous outcome. In this case, the judge who votes first may vote in favor of his preferred outcome, knowing that the other judges will choose to vote in the same direction. Thus, the judge who votes first determines the outcome.

PROOF OF CASE 5: If $|a_d - \eta| < c_d$, $|a_{m_1} - \eta| < c_m$ and $c_m < |a_{m_2} - \eta| < c_d$, there will always be a unanimous outcome. The position of the minority judge will prevail only if the voting order is (d, m_1, m_2) ; otherwise, the panel will vote in favor of the majority position.

In this case, judge d prefers switching sides to dissenting, and judge m_1 prefers any unanimous outcome over any nonunanimous outcome. If either m_1 or m_2 vote first, they can therefore vote their true preference, knowing that the other judges will agree. If the vote order is (d, m_1, m_2) , then judge m_2 will vote in the same direction as judge m_1 . Knowing this, judge d can vote in favor of her preferred outcome. Judge m_1 will join judge d , since he prefers unanimity over disagreement, and judge m_2 will follow. If the order is (d, m_2, m_1) and if judge d voted her preferred position, then judge m_2 would vote in the opposite direction. Once there was already disagreement, judge m_1 would join m_2 , and thus d would be in the dissent. Since this is a suboptimal outcome for judge d , this cannot be a subgame perfect equilibrium. Hence in this case, judge d will switch sides in the first stage, and position of judges m_1 and m_2 will prevail.

PROOF OF CASE 6: If $|a_d - \eta| > c_d$, $|a_{m_1} - \eta| < c_m$ and $c_m < |a_{m_2} - \eta| < c_d$, then there will be a unanimous vote in favor of the minority position if judge m_1 votes before m_2 ; otherwise, all judges will vote their true preferences, and the minority judge will dissent.

In this case, judge d strictly prefers voting her preferred position over compromise. Both judges m_1 or m_2 would prefer compromising over being a lone dissenter, so the last one of the majority judges to vote will follow the first. Since judge m_1 prefers switching sides over having judge d dissent, while judge m_2 prefers voting sincerely and having judge d dissent, the outcome depends on which judge votes first.

DERIVATION OF THE LIKELIHOOD FUNCTION:

Let y_{it} be judge i 's preference in case t . Then for any case cutoff η_t ,

$$\begin{aligned} \Pr(y_{it} = P) &= \Pr(\alpha_i + \varepsilon_{it} > \eta_t) \\ &= \Phi(\alpha_i - \eta_t) \end{aligned}$$

To derive the probability of a unanimous ruling, recall that such a ruling can occur three ways: from *ex ante* agreement, the minority switching sides, or the majority switching sides. The probability of a unanimous “ P ” vote following from *ex ante* unanimity is

$$\Phi(\alpha_1 - \eta_t)\Phi(\alpha_2 - \eta_t)\Phi(\alpha_3 - \eta_t)$$

Applying the results in section 1, the probability of arriving at a unanimous “ P ” ruling after the minority judge switches sides is

$$\sum_{\substack{i=1 \\ j,k \in \mathcal{S}_t - \{i\}}}^3 [\Phi(\alpha_i - \eta_t + c_d) - \Phi(\alpha_i - \eta_t)] \left[\Phi(\alpha_j - \eta_t)\Phi(\alpha_k - \eta_t) - \frac{1}{6}\Psi^-(\alpha_j, \alpha_k) \right]$$

where

$$\begin{aligned}\Psi^-(\alpha_j, \alpha_k) &= [\Phi(\alpha_j - \eta_t) - \Phi(\alpha_j - \eta_t - c_m)] [\Phi(\alpha_k - \eta_t) - \Phi(\alpha_k - \eta_t - c_d)] \\ &\quad + [\Phi(\alpha_j - \eta_t) - \Phi(\alpha_j - \eta_t - c_d)] [\Phi(\alpha_k - \eta_t) - \Phi(\alpha_k - \eta_t - c_m)]\end{aligned}$$

We can similarly derive the probability of the majority switching sides to reach a unanimous “ P ” vote:

$$\sum_{\substack{i=1 \\ j,k \in S_t - \{i\}}}^3 \left[\frac{1}{6} \Phi(\alpha_i - \eta_t) + \frac{1}{3} \Phi(\alpha_i - \eta_t - c_d) \right] \Psi^+(\alpha_j, \alpha_k)$$

where

$$\begin{aligned}\Psi^+(\alpha_j, \alpha_k) &= [\Phi(\alpha_j - \eta_t + c_m) - \Phi(\alpha_j - \eta_t)] [\Phi(\alpha_k - \eta_t + c_d) - \Phi(\alpha_k - \eta_t)] \\ &\quad + [\Phi(\alpha_j - \eta_t + c_d) - \Phi(\alpha_j - \eta_t)] [\Phi(\alpha_k - \eta_t + c_m) - \Phi(\alpha_k - \eta_t)]\end{aligned}$$

Hence

$$\begin{aligned}\Pr((P, P, P) \mid \alpha_1, \alpha_2, \alpha_3, c_d, c_m, \eta_t) &= \\ &\sum_{\substack{i=1 \\ j,k \in S_t - \{i\}}}^3 \Phi(\alpha_i - \eta_t + c_d) \Phi(\alpha_j - \eta_t) \Phi(\alpha_k - \eta_t) \\ &- \frac{1}{6} \sum_{\substack{i=1 \\ j,k \in S_t - \{i\}}}^3 [\Phi(\alpha_i - \eta_t + c_d) - \Phi(\alpha_i - \eta_t)] \Psi^-(\alpha_j, \alpha_k) \\ &+ \frac{1}{6} \sum_{\substack{i=1 \\ j,k \in S_t - \{i\}}}^3 [\Phi(\alpha_i - \eta_t) + 2\Phi(\alpha_i - \eta_t - c_d)] \Psi^+(\alpha_j, \alpha_k) \\ &- 2\Phi(\alpha_1 - \eta_t) \Phi(\alpha_2 - \eta_t) \Phi(\alpha_3 - \eta_t)\end{aligned}$$

It follows that

$$\Pr((D, D, D) \mid \alpha_1, \alpha_2, \alpha_3, c_d, c_m, \eta_t) = \Pr((P, P, P) \mid -\alpha_1, -\alpha_2, -\alpha_3, c_d, c_m, \eta_t)$$

To derive the probability that one judge dissents, we use the conditions given in cases 3 and 5:

$$\Pr((P, P, D) \mid \alpha_1, \alpha_2, \alpha_3, c_d, c_m, \eta_t) = \Phi(-\alpha_3 + \eta_t - c_d) \left[\Phi(\alpha_1 - \eta_t) \Phi(\alpha_2 - \eta_t) - \frac{1}{2} \Psi^-(\alpha_1, \alpha_2) \right]$$

Given that η_t is in fact unobserved, we must integrate over the distribution of η_t to get the unconditional probability:

$$\Pr((v_1, v_2, v_3) \mid \alpha_1, \alpha_2, \alpha_3, c_d, c_m, \sigma) = \int \Pr((v_1, v_2, v_3) \mid \alpha_1, \alpha_2, \alpha_3, c_d, c_m, \sigma \eta) \phi(\eta) d\eta$$

which can be estimated using Gauss-Hermite quadrature. The log-likelihood function is therefore

$$\log L(\alpha, c_d, c_m, \sigma | V) = \sum_{i=1}^n \log \Pr((v_{i1}, v_{i2}, v_{i3}) | \alpha, c_d, c_m, \sigma)$$

When there are observed case characteristics, we let the case cutoff be $x\beta + \eta$, where x is a vector of characteristics and β is to be estimated. The above expression can then be substituted into the likelihood function.

An additional complication arises when particular judges appear only a small number of times in the data. When the number of observations for a judge is below a certain threshold, we assign this judge to a “cluster” based on observable characteristics. For this data, these judges are grouped into clusters of Democratic and Republican appointees. We assume that a clustered Democratic judge has ideology parameter $\alpha_i \sim N(\alpha_{dem}, \tau^2)$ and a clustered Republican judge has ideology parameter $\alpha_i \sim N(\alpha_{rep}, \tau^2)$, where α_{dem} , α_{rep} , and τ are additional parameters to be estimated. In the case of a clustered Democratic judge,

$$\Pr(y_{it} = P) = \Phi\left(\frac{\alpha - \eta_t}{1 + \tau^2}\right)$$

and the likelihood function can be constructed in a similar manner as above.

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Figure 1

Probability of Equilibrium Outcomes when Minority Judge is Willing to Compromise

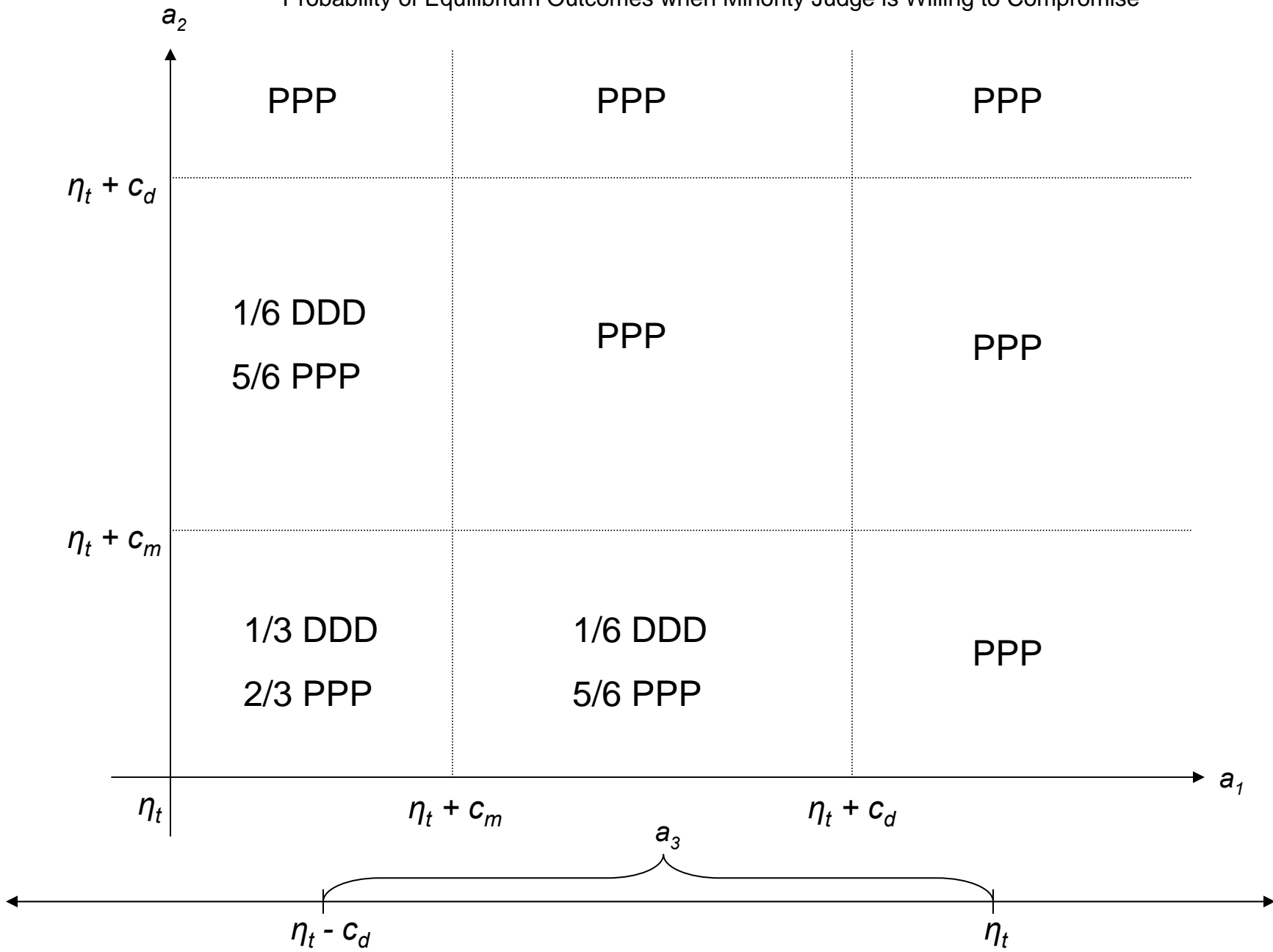


Figure 2

Probability of Equilibrium Outcomes when Minority Judge is Unwilling to Compromise

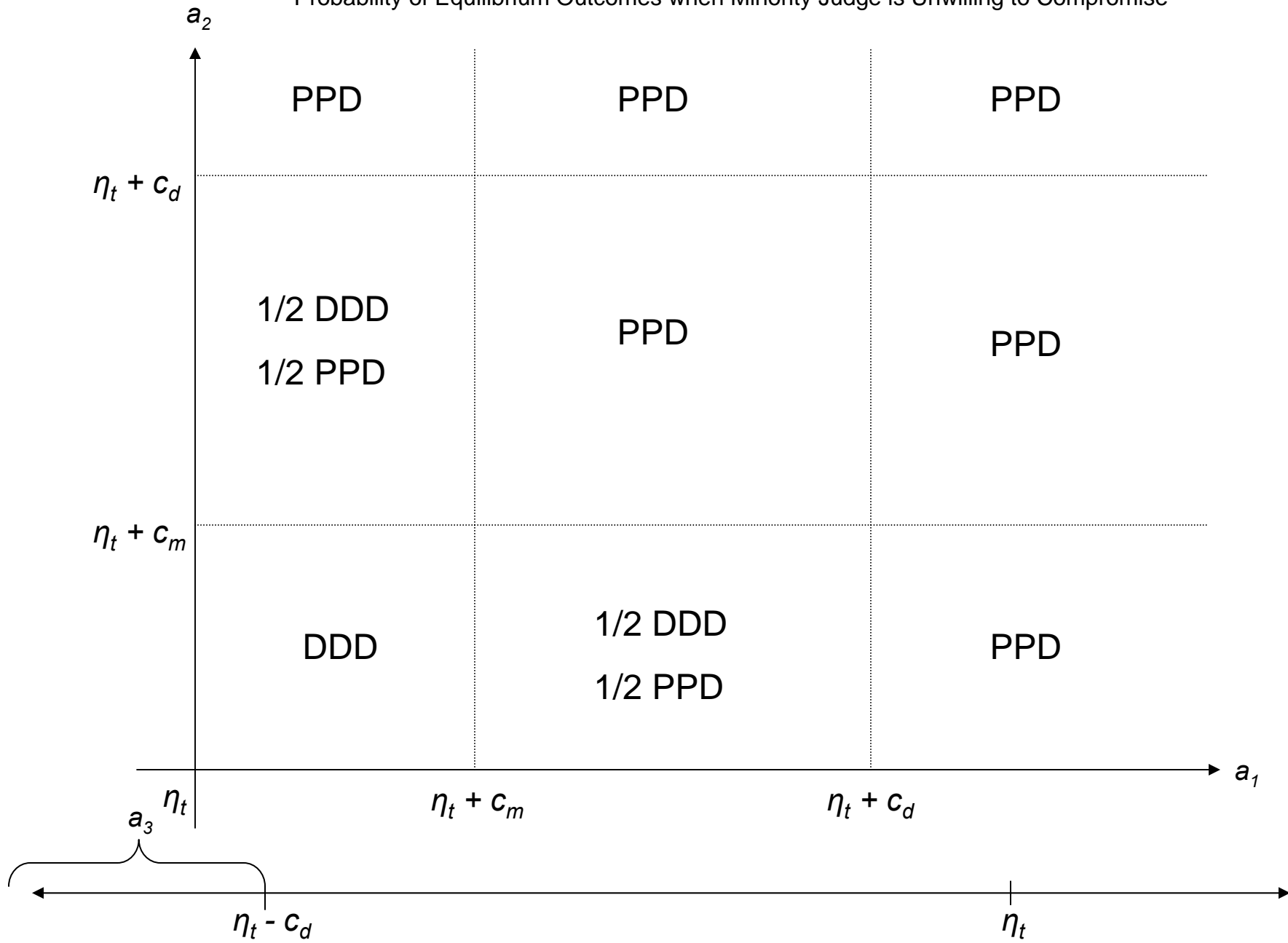


Figure 3

Empirical Distribution of Test Statistics for Random Assignment

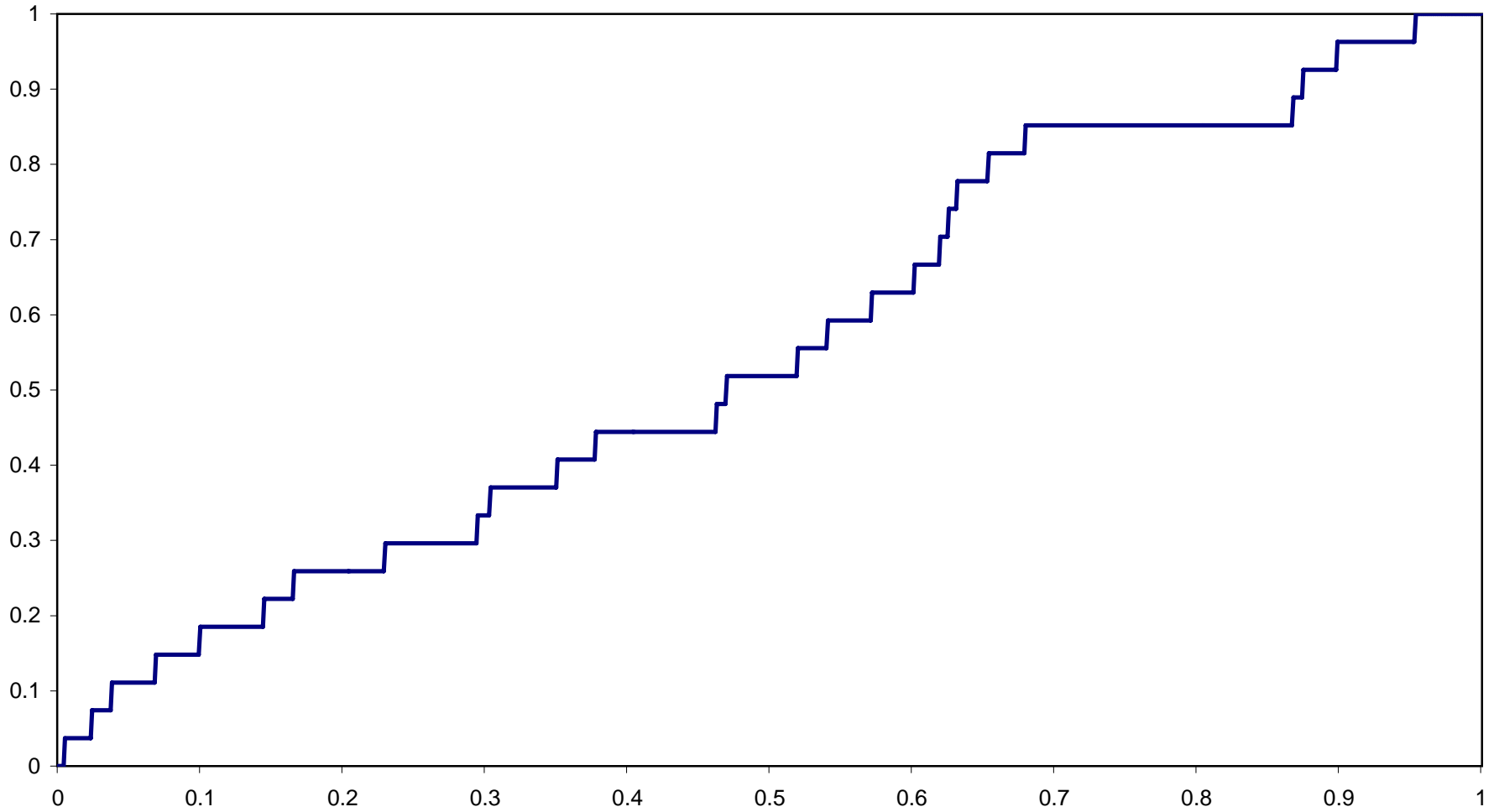


Figure 4

Grant Rate in Asylum Cases, by Individual Judge

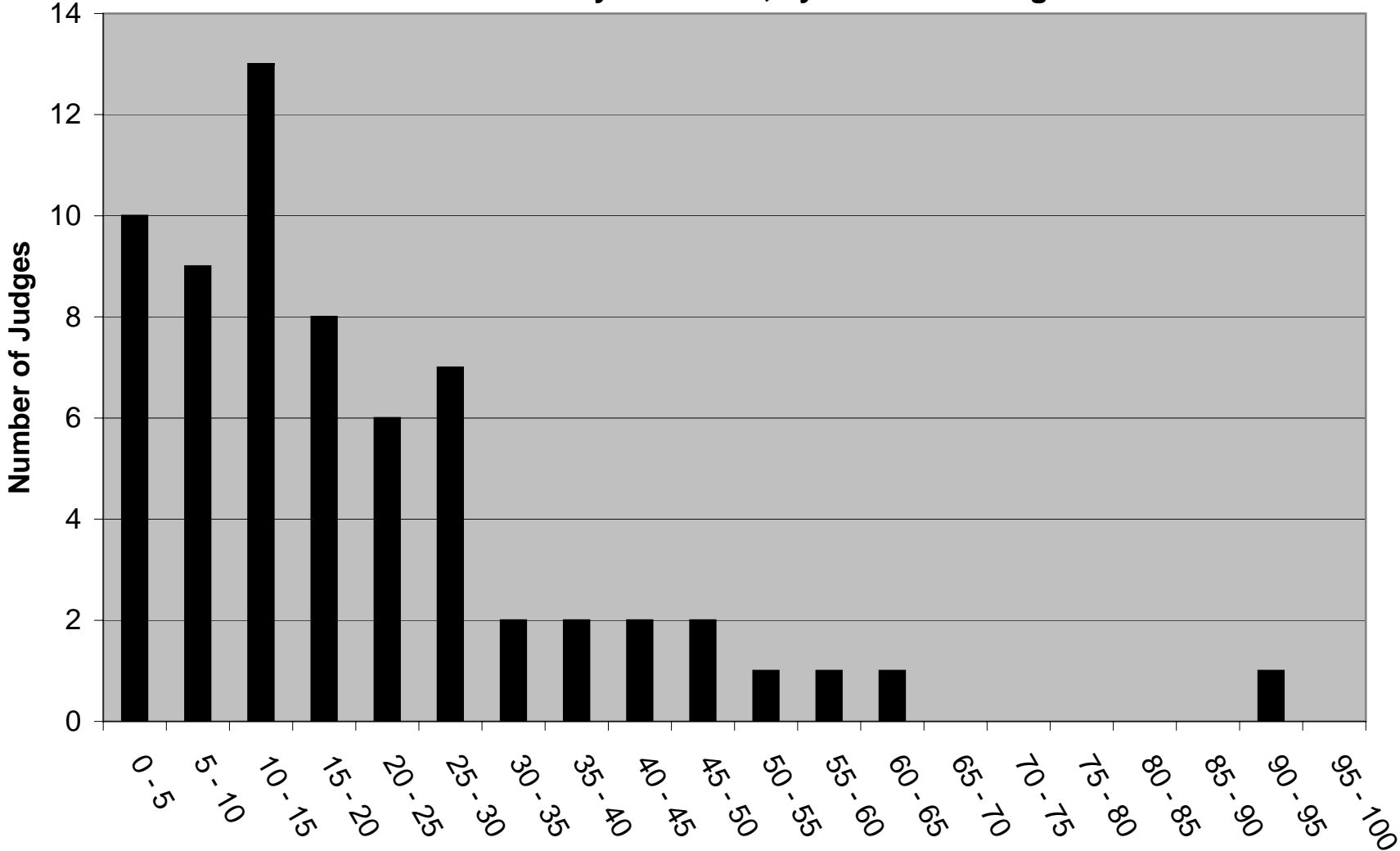


Figure 5

Pro-Plaintiff Voting Rate in Sex Discrimination Cases, by Individual Judge

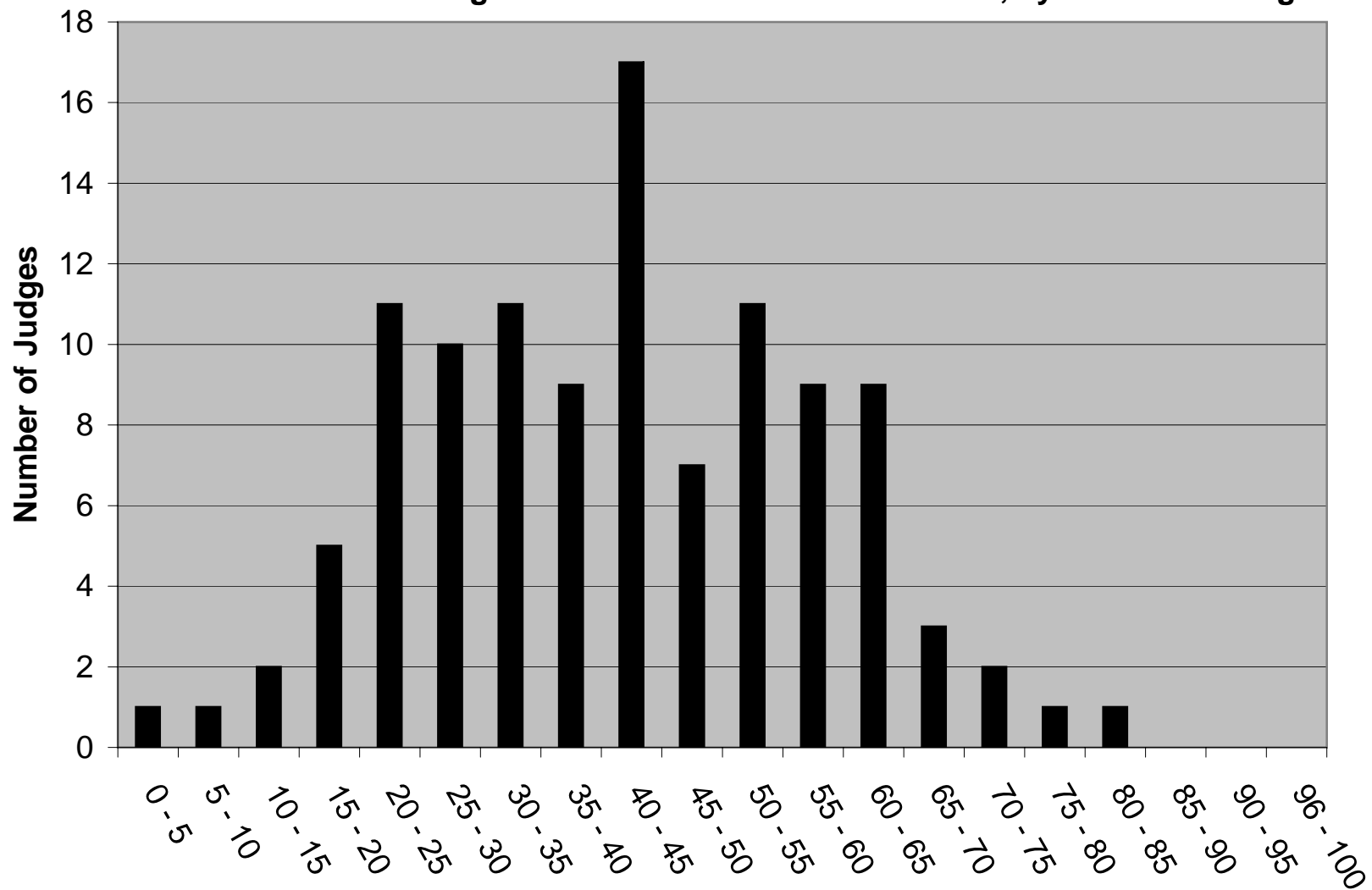


Figure 6

Estimated Probability of Pro-Asylum Preference for Individual Judges

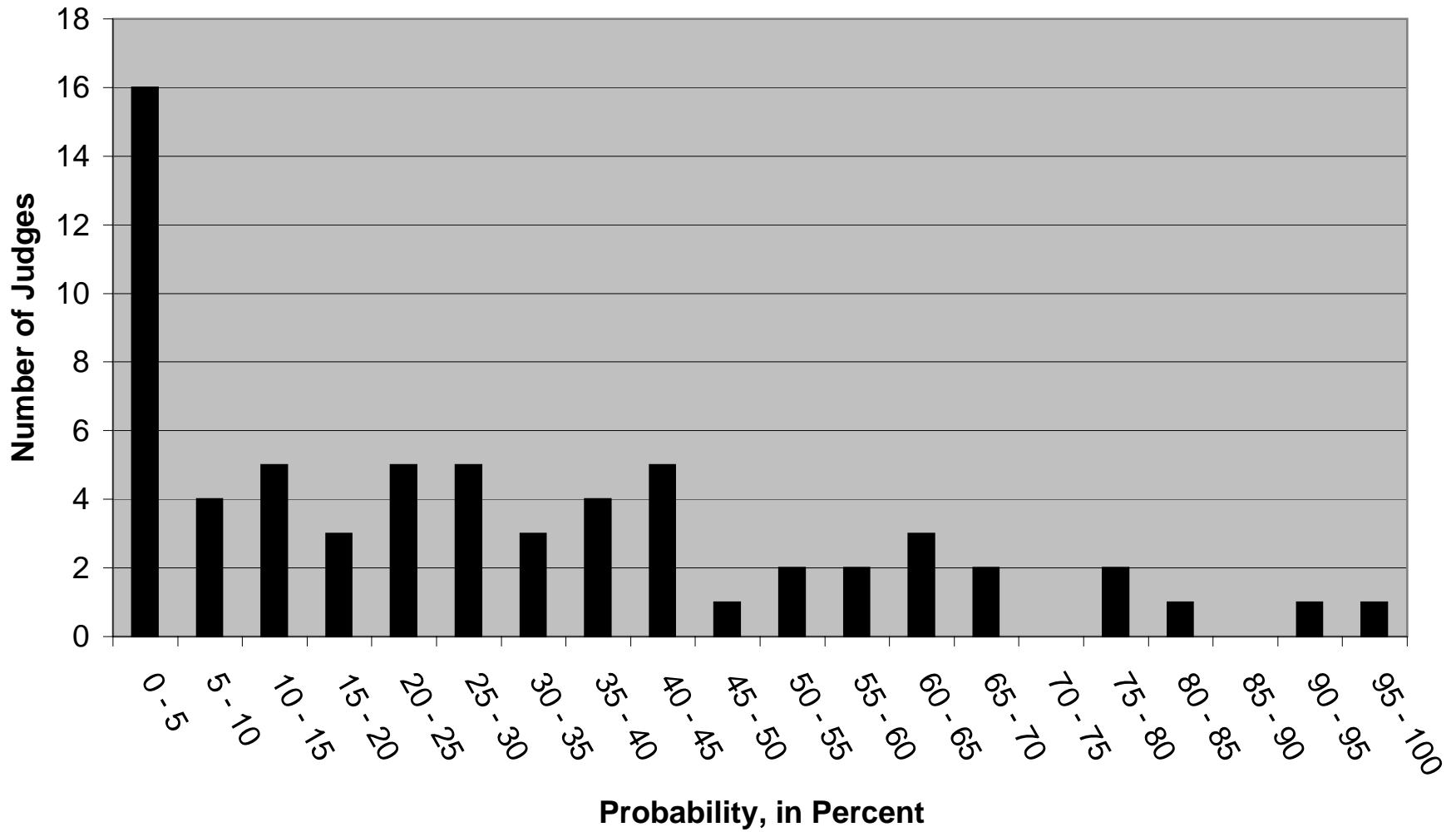


Figure 7

Estimated Probability of Pro-Plaintiff Preference for Individual Judges
in Sex Discrimination Cases

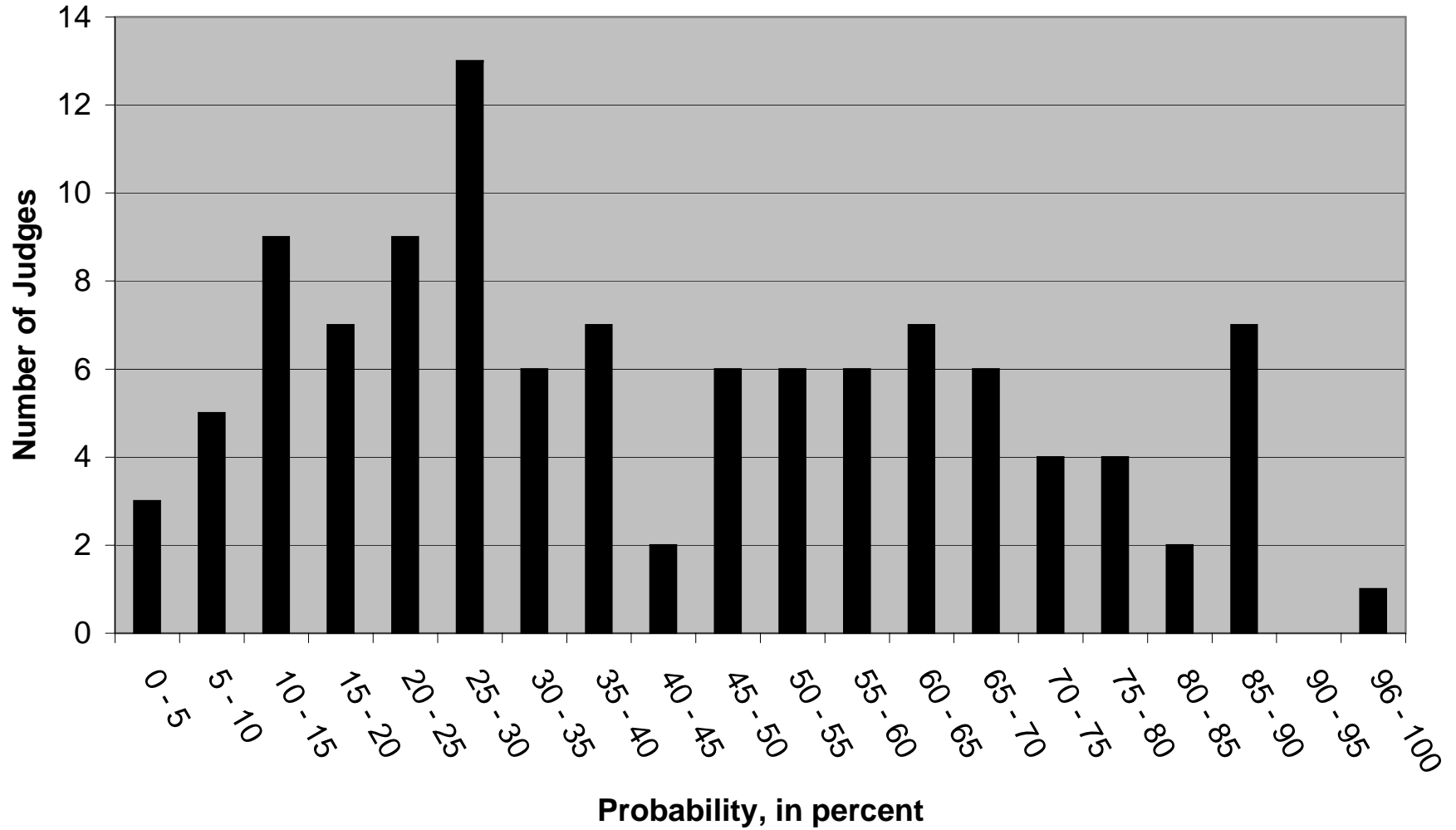


Figure 8

Estimated Probability of Pro-Asylum Preference for Individual Judges,
by Party of Appointment

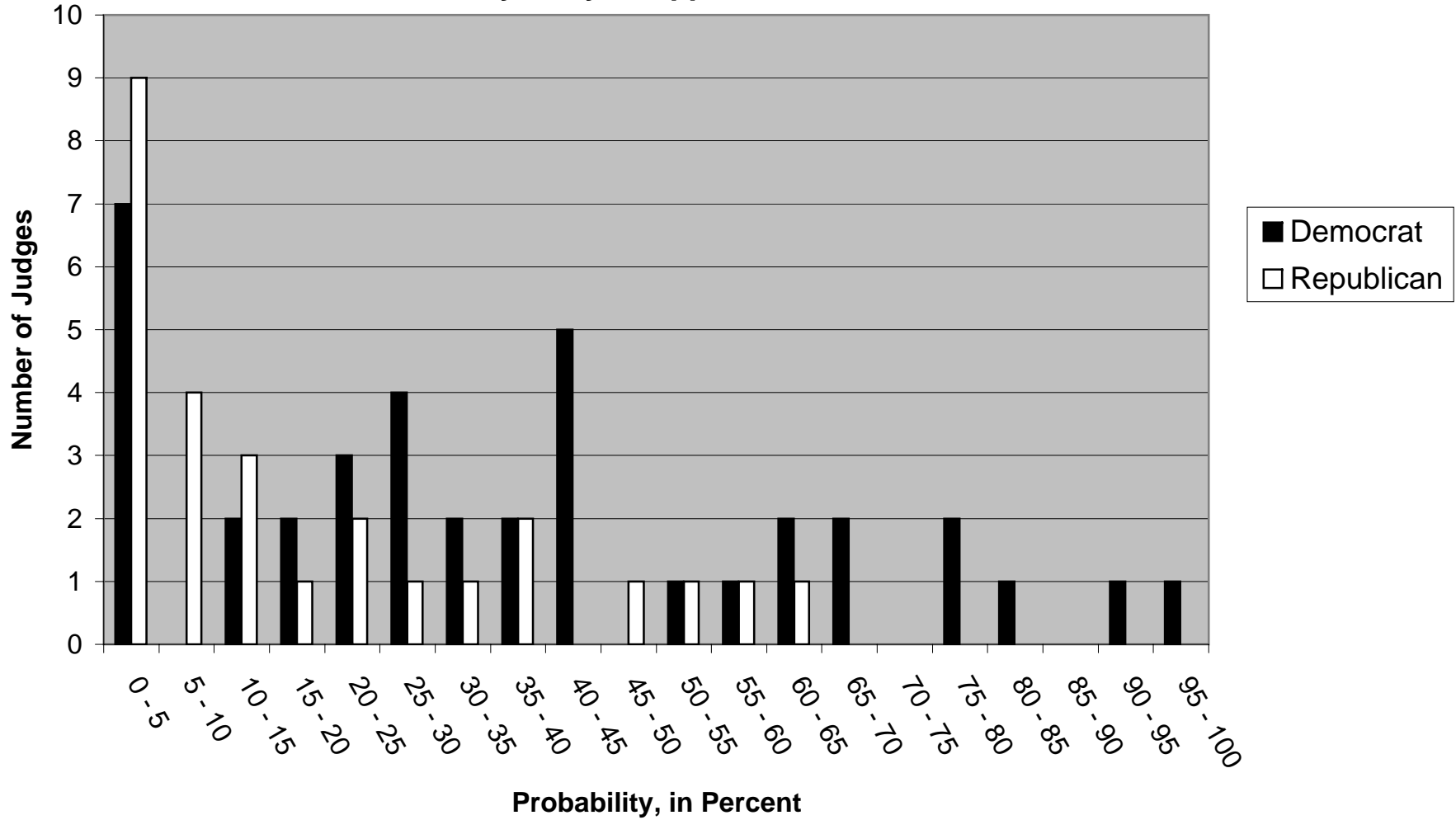


Figure 9

Estimated Probability of Pro-Plaintiff Preference for Individual Judges in Sex Discrimination Cases, by Party of Appointment

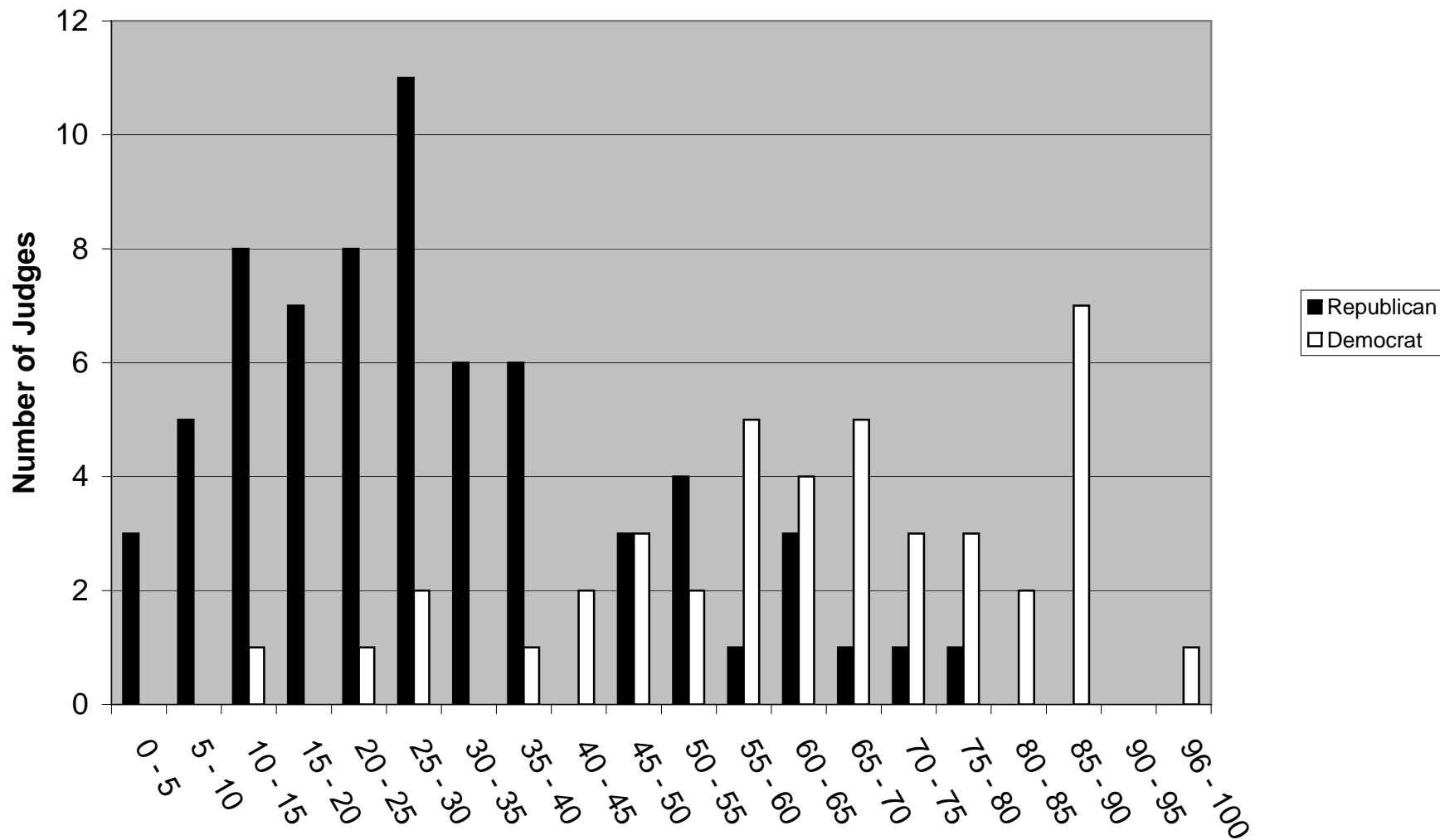


Figure 10

Probability of Asylum Grant, Conditional on Strength of Claim

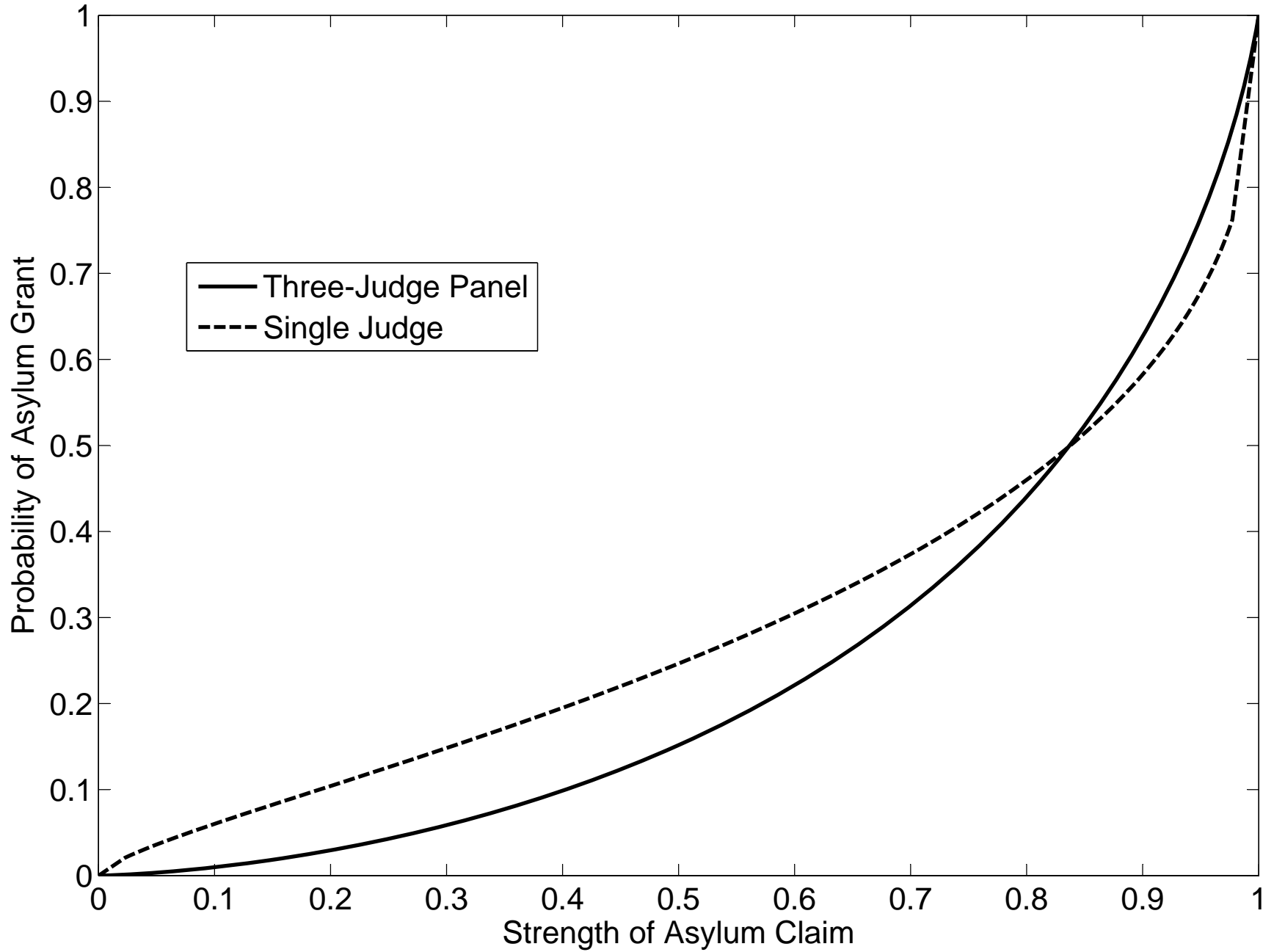


Figure 11

Probability of Ruling for Plaintiff,
Conditional on Strength of Sex Discrimination Claim

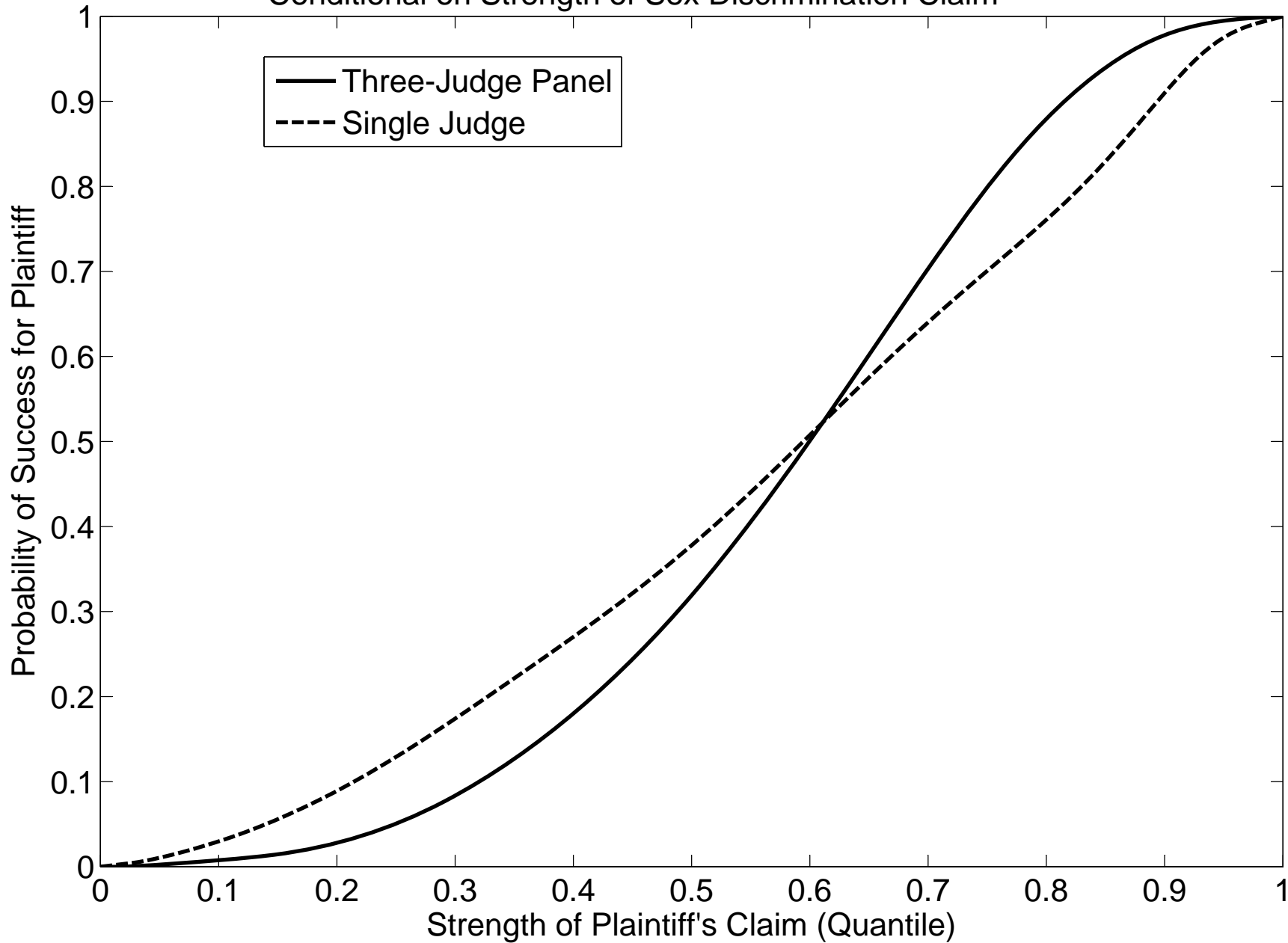


Table 1: Ideology Estimates of Individual Judges in Asylum Cases

Judge	Spatial Estimate	Pro-Asylum Probability	Judge	Spatial Estimate	Pro-Asylum Probability
Alarcon	-1.02 (0.59)	18% (14%)	Nelson, T.G.	-1.06 (0.26)	17% (6%)
Aldisert*	0.44 (0.43)	66% (15%)	Noonan	0.32 (0.24)	61% (9%)
Beezer	-0.90 (0.28)	20% (7%)	Norris	-0.53 (0.57)	31% (18%)
Berzon	0.07 (0.67)	53% (25%)	O'Scannlain	-2.63 (0.36)	1% (1%)
Boochever	-0.30 (0.46)	39% (16%)	Paez	-0.26 (0.66)	41% (23%)
Browning	-0.21 (0.23)	42% (8%)	Poole	-2.34 (0.91)	2% (3%)
Brunetti	-1.21 (0.30)	13% (6%)	Pregerson	0.92 (0.16)	80% (4%)
Canby	-1.03 (0.34)	17% (8%)	Rawlinson	-2.70 (0.74)	1% (1%)
Choy	-0.43 (0.39)	35% (13%)	Reinhardt	1.52 (0.19)	92% (2%)
Farris	-2.27 (0.46)	2% (2%)	Rymer	-2.27 (0.39)	2% (1%)
Ferguson	0.54 (0.26)	69% (8%)	Schroeder	-0.59 (0.21)	29% (6%)
Fernandez	-1.79 (0.35)	5% (3%)	Schwarzer*	-0.39 (0.58)	36% (20%)
Fitzgerald*	-0.40 (0.71)	36% (24%)	Shadur*	0.76 (0.66)	76% (19%)
Fletcher, B.	0.79 (0.21)	77% (6%)	Shea*	-0.20 (0.83)	43% (30%)
Fletcher, W.	-0.42 (0.33)	35% (11%)	Silverman	-1.29 (0.37)	12% (6%)
Goodwin	-0.61 (0.24)	29% (7%)	Skopil	-0.89 (0.56)	21% (15%)
Gould	-1.90 (1.02)	4% (8%)	Sneed	-1.35 (0.43)	11% (7%)
Graber	-1.88 (0.55)	4% (4%)	Tallman	-0.74 (0.86)	25% (25%)
Hall	-1.20 (0.35)	14% (7%)	Tang	0.30 (0.61)	61% (21%)
Hawkins	-0.27 (0.19)	40% (7%)	Tashima	-1.25 (0.32)	13% (6%)
Hug	-0.47 (0.24)	33% (8%)	Thomas	0.31 (0.20)	61% (7%)
King*	0.19 (0.46)	57% (17%)	Thompson	-1.51 (0.34)	8% (4%)
Kleinfeld	-0.78 (0.28)	24% (8%)	Trott	-1.56 (0.32)	8% (4%)
Kozinski	-1.97 (0.38)	4% (2%)	Wallace	-2.72 (0.44)	1% (1%)
Kravitch*	1.82 (1.02)	95% (9%)	Wardlaw	-0.72 (0.29)	25% (8%)
Lay*	0.14 (0.49)	55% (18%)	Weiner*	-0.57 (0.79)	30% (25%)
Leavy	-2.13 (0.42)	3% (2%)	Wiggins	-1.93 (0.45)	4% (3%)
Magill*	0.13 (0.59)	55% (22%)	Wood, H.*	-1.41 (1.09)	10% (17%)
McKeown	-0.60 (0.38)	29% (12%)	Wright	-0.03 (0.56)	49% (21%)
Moskowitz*	-0.74 (0.95)	25% (28%)	Zilly*	-1.94 (1.05)	4% (8%)
Nelson, D.W.	-0.15 (0.29)	44% (10%)			

Notes: Table provides ideology parameter estimates for all judges who appear at least 10 times in the data. Standard errors are shown in parentheses. Standard errors are estimated using the inverse Hessian matrix, evaluated at the maximum likelihood estimate. Estimated probabilities of pro-plaintiff preference are derived from equation (1); standard errors for probabilities are computed using the delta method. Judges who only have votes in a single direction are excluded. District judges and judges from other circuits are marked with an asterisk.

Table 2: Estimates of Structural Parameters

Asylum			Sex Discrimination	
Distribution of Judicial Ideology	Spatial Estimate	Pro-Plaintiff Probability	Spatial Estimate	Pro-Plaintiff Probability
25th Percentile	-1.79	5.1%	-2.19	22.7%
Median	-0.74	25.0%	-0.98	36.9%
75th Percentile	-0.20	42.6%	1.02	63.7%
Median Republican	-1.35	10.8%	-1.81	26.8%
Median Democrat	-0.50	32.3%	1.17	65.5%
Model Parameters				
c_d (Cost of dissent for minority judge)	1.71	(0.10)	3.21	(0.47)
c_m (Cost of dissent for majority judge)	1.36	(0.28)	0.00	
σ (Standard deviation of case cutoff)	0.44	(0.21)	2.75	(0.81)

Notes: Top table provides estimates of costs of disagreement and the standard deviation of the case cutoff. The bottom table shows quantiles of the distribution of the judges in ideology space. Standard errors are shown in parentheses. Standard errors are derived from the inverse Hessian matrix, evaluated at the maximum likelihood estimate. Predicted probabilities of pro-plaintiff preference are derived from equation (1).

Table 3: Impact of Nationality in Asylum Cases

Freedom House

Civil Liberties Index

CL = 2	-0.63	(0.23)
CL = 3	-0.09	(0.13)
CL = 4	-0.14	(0.14)
CL = 5	-0.06	(0.17)
CL = 6	-0.42	(0.21)
CL = unidentified	-0.56	(0.24)

Notes: Table provides estimates of dummy variables for Freedom House Civil Liberties Index, coded by country of origin. The case CL = 7 is excluded. The case CL = 1 is dropped because there were only two such cases, both of which lost. Standard errors are shown in parentheses.

Table 4: Simulation of Vote Changes in Asylum Cases

Vote Changes	Percent	
initial agreement	48.9%	(5.8%)
one judge switched	32.5%	(6.0%)
two judges switched	13.3%	(3.1%)
dissent	5.3%	

Notes: Table provides simulated estimates of votes changes based on parameter estimates, using votes observed in the data. Simulated standard errors are shown in parentheses. Dissent rate is observed exactly.