Diversity and Judicial Decision-Making:

Evidence from Affirmative Action Cases in the Federal Courts of Appeals, 1971-1999

Charles M. Cameron
Craig P. Cummings
Department of Political Science
Columbia University

Version 1.0
March 30, 2003

Draft, subject to change

We thank Lawton Cummings, Esq., for advice on doctrine and procedure, and Rajeev Dehejia, Robert Erikson, Sean Farhang, Jennifer Hill, and Greg Wawro for helpful discussions. Portions of the data were originally collected by Kevin M. Woodson for an honors thesis at Columbia College, under Cameron's direction. We gratefully acknowledge the efforts of this fine scholar. Cameron additionally acknowledges support from NSF Grant SES-0079963.

Abstract

We study the effect of racial, gender, and ideological diversity on judicial decision-making concerning affirmative action. More specifically, we examine how heterogeneity in the composition of the three-judge panels of the U.S. Courts of Appeal affected the voting of the judges sitting on those panels as they decided 179 cases concerning affirmative action, from 1971 to 1999. To do so, we create a new framework for studying interactive decision-making on collegial courts. This "social economy" approach distinguishes the effects of a judge's own characteristics from the effects of other panel members' characteristics, and both from the "peer effect" of other judges' voting choices. We illustrate practical and easily implemented statistical methods for applying this framework to voting data from the U.S. Courts of Appeals.

We find that increasing racial diversity on the panels of the Courts of Appeals substantially changed the voting behavior of other judges on the panels, an effect distinguishable from peer voting pressures. This finding suggests an apparent "deliberation effect" on racially heterogeneous panels. But we also find that peer voting pressures are potent in the Courts of Appeal, raising the possibility that highly motivated judges can "tip" the voting of colleagues on a panel. We also uncover substantial effects from judges' political ideologies, powerful effects from circuit-level precedent, and considerable responsiveness to doctrinal changes from the U.S. Supreme Court.

1. Introduction

In this paper, we study the effect of racial, gender, and ideological diversity on judicial decision-making concerning affirmative action. More specifically, we examine how heterogeneity in the composition of the three-judge panels of the U.S. Courts of Appeal affected the voting of the judges sitting on those panels as they decided 179 cases concerning affirmative action, from 1971 to 1999.

The implications of this study are obvious, particularly for the politics of judicial selection and perhaps more generally for policies aimed at enhancing the racial and gender diversity of democratic institutions. Indeed, we hope to strengthen the knowledge base relevant to current policy debates. However, in order to convincingly address the impact of judicial diversity, we must directly engage an on-going controversy concerning the foundations of judicial decision-making on collegial courts. In particular, a recent series of innovative papers have cast considerable doubt on what might be called the traditional political science approach to decision-making on collegial courts (Revesz 1997, Cross and Tiller 1998, Farhang and Wawro 2003, Gryski et al 1986, Songer and Crews-Meyer 2000, Massie et al 2002). These papers have pioneered a *contextual approach* to judicial decision-making, that stands in rather stark contrast to the traditional anomic approach largely derived from the "attitudinal model" of Supreme Court decision making. The traditional model holds that the political beliefs and attitudes of judges measurably affect their voting choices. The contextual approach explores the extent to which the characteristics of *other* judges on a panel also influence the choices of a given judge on a panel. Although the evidentiary base remains limited, in all cases in which the contextual approach has been applied to the Courts of Appeals, it has uncovered substantively significant social interactions among judges based on the diversity of the panels. But, we argue, the foundations of the contextual-based approach remain somewhat murky and interpreting contextual effects is difficult.

We create a new theoretical framework, the *social economy approach*, for studying judicial decision-making on collegial courts. The social economy approach is based on recent theoretical and empirical advances in studying social interactions in groups (Manski 2000, Becker and Murphy 2001). Applied to judicial decision making, it encompasses both the traditional approach and the contextual approach as special cases, and clarifies the differences between them. But it is broader than either, because it also allows for *peer effects* on the panels. By "peer effect," we mean that the *actions* of other judges, not merely their characteristics, may affect the voting of a given judge. Possible sources of a peer effect include a "norm of consensus" based on concerns about the perceived legitimacy of courts, strategic behavior aimed at preventing reversal in an en banc review or by the Supreme Court, and efforts to avoid the burden of writing dissents. Peer effects seem likely to be important on the U.S. Courts of Appeal, since the vast majority of votes are unanimous. Indeed, the "norm of consensus" is a standard notion in the scholarly literature on those courts. Thus, peer effects need to be addressed explicitly in studies of judicial decision making on the Courts of Appeals, and perhaps on collegial courts more generally.

Explicit attention to peer effects is particularly critical in a study of diversity, because peer effects can create a spurious diversity effect when none actually exists. In our view, a "true" diversity effect arises when the presence of Judge 1 – say, an African-American – alters the deliberative process on a panel, supplying Judges 2 and 3 – non-African-Americans – with valuable information or unfamiliar but genuinely persuasive logic. As a consequence, Judges 2 and 3 may vote differently than they otherwise would have, and diversity can be said to have improved the panel's adjudication. But suppose instead that Judge 1 simply has intense policy preferences, while Judges 2 and 3 have weak legal preferences per se but strongly wish to vote similarly to their peers on the panel. Judge 1 may tip the voting of the "go-along" judges, who will spuriously display a diversity effect absent any real change in the deliberative process. In this case, "diversity" will have altered the panel's judgment – for good or for ill, depending on one's normative preferences – but can hardly be said to have improved it.

In order to address this problem, the social economy approach distinguishes three separate effects: 1) the impact of judge *i*'s *own* characteristics (race, gender, and ideology) on judge *i*, 2) the impact of *other* judges' characteristics on judge *i* – which we take to be a diversity effect, and 3) the separate impact of other judges' *actions* on judge *i* – a peer effect. The relationship between these three effects, and the importance of untangling one from another, has not been sufficiently appreciated by advocates of either the traditional or contextual approaches to judicial decision-making.

As one would expect, untangling these three effects is difficult. Particularly tricky is the simultaneity problem inherent in interactive decision making: Judge 1's actions affect those of Judges 2 and 3, but their actions affect her's as well. Fortunately, however, untangling the effects is not impossible. Following recent work in social economics, we formulate an empirical framework allowing estimation of the three effects. We apply the framework to the affirmative action data.

The paper is organized as follows. Section 2 lays out the social economy approach to judicial decision-making, and clarifies the empirical issues in studying decision-making on collegial courts. Section 3 describes newly collected data on affirmative action cases in the U.S. Courts of Appeals. Section 4 uses simple regression models to show that the data reject the traditional model of judicial decision making, in favor of a model allowing social interactions on the panels. However, this simple approach cannot distinguish diversity from peer effects. Section 5 then investigates the social interactions on the panels, attempting to untangle the diversity effect from peer effects. We rely heavily on the fact that judges are assigned to panels through an exogenous mechanism, effectively randomizing the composition of panels and the assignment of cases. A relatively straight-forward instrumental variables technique then affords a way to differentiate the three effects. Section 6 discusses the policy implications of the findings.

2. The Social Economy of Judicial Decision-making

Background

In recent years, the burgeoning field of social economics has created new theoretical and empirical tools for studying interactions in groups (Manski 2000, Becker and Murphy 2000). Much of the new work has focused on "neighborhood" or "peer" effects. Examples include neighborhood/peer effects in crime (Glaser, Sacerdote, and Scheinkman 1996), drinking (Cremer and Levy 2001), unemployment (Topa 2001), teen pregnancy (Case and Katz 1991), teen smoking (Evans, Oates and Schwab 1992?), dropping out of school (Crane 1991), and college achievement and decisions to join social organizations (Sacerdote 2000), among others. In most cases, the basic procedure is to regress an individual's action on the average action among members of the individual's peer group, controlling for the individual's characteristics and those of the peer group. In other cases, the empirical strategy is to compare the variability of outcomes across neighborhoods with the variability of fundamentals (a theoretical implication of some social economy models is that the variability of behavior across neighborhoods may be much greater than the variability in underlying fundamentals).

These studies trace an intellectual lineage to Thomas Schelling's classic studies of externalities in actions (1971, 1978). In those studies, Schelling demonstrated how spillovers from one person's actions to another create multiple equilibria and social multipliers, leading to tipping, fads, herd behavior, band wagons, and cascades.[1] More recent theory has pursued two somewhat different approaches. The first focuses on small, local, over-lapping groups with contagion effects (Glaser and Scheinkman 2001). Conceptually, one imagines individuals located on a line or circle so that each individual interacts directly with two neighbors but may be influenced by more distant individuals as behavior diffuses through the population. The second approach imagines individuals interacting in large groups, with the average level of behavior in the group affecting the returns to each individual from his or her own action (Brock and Durlauf 2001). This approach can be extended to small groups as well (Krauth 2001).

In the following sections, we apply a social economics approach to the behavior of judges on the U.S. Courts of Appeals. A small groups version of the Brock-Durlauf "average levels" approach seems most natural in this setting, and that is the approach we take.

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[1] Sociologists have long argued for the existence of these kinds of social interactions, and have offered theoretical models emphasizing social networks (see e.g., Granovetter 1973). Coleman 1990 is particularly important in this regard.

Theoretical Framework

Judge *i*, *i* = (1, 2, 3), a member of a three judge panel, may cast a vote $v_i \in \{0,1\}$ where 1 denotes a vote in pro-affirmative action direction. The utility to judge *i* of v_i is

$$u_i(v_i) = h_i(v_i) + \theta_1 s_i(v_i, \overline{v}_i) + \varepsilon_i(v_i) \quad (1)$$

where $h_i(v_i)$ is the "private utility" to *i* from his own vote, $s_i(v_i, \overline{v}_i)$ is the "social utility" to *i* from his vote, $\overline{v}_i$ is the average value of the votes of the other judges on the panel, and $\varepsilon_i(v_i)$ is an idiosyncratic term. We discuss the private and social utility terms in turn.

We take the private utility term to reflect the value to *i* of "doing the right thing," based on his or her best available information, understanding of the law, precedents, personal convictions, and deliberation. As such, the private utility term is compatible with either a Condorcet Jury Theorem-like model of deliberation, and an attitudinal model based largely on ideological preferences. We will not supply micro-foundations for either interpretation but proceed as follows. Let the private utility of v_i be 1 if the "correct" vote in the case for *i* is v_i and 0 if the correct vote for *i* is not v_i. Let $\pi_i(v_i)$ be *i*'s assessment that v_i is the correct vote in the case. Then $h_i(v_i) = \pi_i(v_i)(1) + (1 - \pi_i(v_i))0 = \pi_i(v_i)$. In other words, judge *i*'s private utility from a vote is her assessment the vote is the correct one. In order to capture the idea that *i*'s own preferences and life experiences affect his perception of the likely correct vote, we assume π_i is a function of *i*'s personal characteristics (as in the attitudinal model). But the information supplied by other members of the panel, reflecting their knowledge and life experiences, may also affect *i*'s vote (as in a deliberative or Condorcet Jury-like model). So we allow π_i also to be a function of the mean values of the other judges' characteristics. (Use of the mean reflects that fact that neither of the other judges is distinguished or privileged, rather, it is the overall information environment they create that is critical.) More specifically, we assume

$$\pi_i(v_i) = \theta_0 + \theta_2 x_i + \theta_3 \bar{x}_i \tag{2}$$

The social utility term is the novel part of the model. We assume this term acts as follows: $s_i(v_i, \bar{v}_i) = 1 - |v_i - \bar{v}_i|$. Thus, if judges 2 and 3 both vote in the same way, judge 1 receives "1" if he also votes that way, but "0" if he votes in the opposite fashion. If judges 1 and 2 split their votes, then judge 1 receives ½ regardless of how he or she votes. The social utility term thus captures the externalities of binary choices discussed by Schelling (1971).

Although we do not provide micro-foundations for the social utility term, it can be understood in three different ways. In the first, judges value the social legitimacy of courts. As noted by Martin Shapiro, "Judges claim legitimacy by asserting they are non-political, independent, neutral servants of 'the law' " (2002). Dispensing neutral justice on the basis of ostentatiously contested votes is not conducive to maintaining this kind of legitimacy. Of course, the social legitimacy of courts is a public good for judges. An explanation of individual judicial behavior rooted in judge's concerns about the social legitimacy of courts is radically incomplete unless it shows why maintaining the public good is individually rational for judges. But one can easily imagine how maintaining the public good of legitimacy could be individually rational, due to ostracism and social pressure, in a small-population, repeat-play setting like the Courts of Appeals. Goldman supplies interview evidence on an "unwritten rule that a judge should try to achieve consensus with his colleagues and avoid public dissension" (1968: 480). In addition, the procedures employed on the Courts of Appeals seem oriented toward building consensus. For example, most panels caucus behind closed doors after hearing oral argument and attempt to achieve immediate consensus. Subsequently, a writer in the majority circulates a draft opinion. Bargaining and compromise on language then occurs.[2]

In the second, judges have preferences about legal outcomes and wish to avoid reversal by hierarchical superiors. If the Supreme Court engages in "strategic auditing" of lower court cases (as in Cameron, Segal, and Songer 2000) then a dissent can act as a signal to the Supreme Court to review and reverse a particular case. Exactly the same logic applies to a panel flouting the desires of a majority of the circuit court's membership, who can review and reverse the panel's judgment using the "en banc" procedure. An innovative empirical study by Cross and Tiller (1998) lays out the logic of "judicial whistle-blowing" in more detail. On ideologically heterogeneous panels with a potential whistle-blower, avoiding review will drive out-lying judges to conformity with Supreme Court and circuit court desires, and hence uniformity of votes. Conversely, on homogeneous but "outlier" panels, one might see votes at deviance from the desires of the Supreme Court or the district majority, but again displaying uniformity.

The third explanation, stressed particularly by Revesz 2001 and Songer (1982, 1986), hinges on the effort costs of dissents. If two members of a panel indicate an intention to vote in one direction, the third member faces an unpleasant dilemma, if he believes the correct vote is the opposite. He can dissent, thus requiring him to write a laborious dissent that will not affect the outcome of the case; or, he cast a vote with the majority despite his private beliefs – a vote that will not affect the outcome. Revesz additionally notes,

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[2] Coffin 1980 provides a vivid, behind-the-scenes account of decision making on the First Circuit. We have also discussed procedures with clerks on the D.C. and Second Circuits, who indicate similar procedures there.

> Judges on the courts of appeals are monitored with respect to the timeliness of their opinions and receive "credit" for writing majority opinions. No such credit is awarded for dissenting opinions, which do not reduce the workload that is otherwise assigned to a judge. Moreover, dissents impose costs not only on the dissenting judge, but also on the members of the majority, who must respond to the attack on the opinion. In addition, by dissenting a judge gives up whatever influence she might have on the majority opinion. (12)

In sum, it is easy to see why many judges opt for conformity.

We now define a random utility model of voting, following McFadden 1984. Substituting (2) into (1) and noting that $s_i(v_i, \bar{v}_i) = 1 - |v_i - \bar{v}_i|$ yields

$$u_i(1) = \theta_0 + \theta_1 \bar{v}_i + \theta_2 x_i + \theta_3 \bar{x}_i + \varepsilon_i(1)$$

and

$$u_i(0) = 1 - [\theta_0 + \theta_2 x_i + \theta_3 \bar{x}_i] + \theta_1(1 + \bar{v}_i) + \varepsilon_i(0)$$

Accordingly, judge *i* should cast a vote $v_i = 1$ iff

$$u_i(1) = \theta_0 + \theta_1 \bar{v}_i + \theta_2 x_i + \theta_3 \bar{x}_i + \varepsilon_i(1) \geq u_i(0) = 1 - [\theta_0 + \theta_2 x_i + \theta_3 \bar{x}_i] + \theta_1(1 + \bar{v}_i) + \varepsilon_i(0)$$

$$\Rightarrow 2\theta_0 + \theta_1(2\bar{v}_i - 1) + 2\theta_2 x_i + 2\theta_3 \bar{x}_i - 1 \geq \varepsilon_i(0) - \varepsilon_2(1)$$

We assume $\varepsilon_i(0) - \varepsilon_2(1)$ is IID and uniformly distributed on the interval $[-1,1]$. Hence,

$$\Pr(v_i = 1) = \frac{1}{2}\left(2\theta_0 + \theta_1(2\bar{v}_i - 1) + 2\theta_2 x_i + 2\theta_3 \bar{x}_i - 1 - (-1)\right)$$

$$\Pr(v_1 = 1) = \theta_0 + \theta_1(\bar{v}_i - \tfrac{1}{2}) + \theta_2 x_i + \theta_3 \bar{x}_i \qquad (3)$$

which defines a linear probability model.[3]

Note that when two members of a panel split their vote, the peer effect drops out for the third judge. But otherwise, it does not (assuming $\theta_1 \neq 0$). Indeed, if $\theta_1 \neq 0$, then the judges on a panel are in a strategic situation (a game), and the vector of votes (v_1, v_2, v_3) must yield a Nash equilibrium in the game. It is easily seen that if θ_1 is sufficiently large relative to the other terms in (3), the strategic situation has the flavor of a coordination game in which all three judges vote similarly.

Nesting of Alternative Models

We now investigate the formal relationships between the traditional and contextual models and the social economy model. We do so by deriving the reduced form for each model. This allows us to address explicitly the interpretation of the coefficients estimated in earlier work.

In matrix form, the structural equations defining the social economy model are

$$\begin{pmatrix} 1 & -\frac{1}{2}\theta_1 & -\frac{1}{2}\theta_1 \\ -\frac{1}{2}\theta_1 & 1 & -\frac{1}{2}\theta_1 \\ -\frac{1}{2}\theta_1 & -\frac{1}{2}\theta_1 & 1 \end{pmatrix} \begin{pmatrix} v_1 \\ v_2 \\ v_3 \end{pmatrix} = \begin{pmatrix} \theta_0 \\ \theta_0 \\ \theta_0 \end{pmatrix} + \begin{pmatrix} \theta_2 & \frac{1}{2}\theta_3 & \frac{1}{2}\theta_3 \\ \frac{1}{2}\theta_3 & \theta_2 & \frac{1}{2}\theta_3 \\ \frac{1}{2}\theta_3 & \frac{1}{2}\theta_3 & \theta_2 \end{pmatrix} \begin{pmatrix} x_1 \\ x_2 \\ x_3 \end{pmatrix} + \begin{pmatrix} \varepsilon_1 \\ \varepsilon_2 \\ \varepsilon_3 \end{pmatrix} \qquad (4)$$

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[3] A similar derivation but different assumptions about the error term yield logit and probit models. But the linear probability model has strong advantages here. First, we can easily derive the relationship between structural and reduced form estimates in a variety of models (as shown below). Second, we can utilize instrumental variables techniques in a straight forward way, in contrast to IV techniques for models with non-linear treatment and outcome equations. In fact, techniques for such models remain at the current frontiers of research (Angrist 2001). Hence, we focus on a structural model leading to a linear probability model.

(suppressing subscripts denoting different panels). The corresponding reduced form equations are

$$\begin{pmatrix} v_1 \\ v_2 \\ v_3 \end{pmatrix} = \begin{pmatrix} \alpha_1 \\ \alpha_2 \\ \alpha_3 \end{pmatrix} + \begin{pmatrix} \beta_1 & \gamma_1 & \delta_1 \\ \beta_2 & \gamma_2 & \delta_2 \\ \beta_3 & \gamma_3 & \delta_3 \end{pmatrix} \begin{pmatrix} x_1 \\ x_2 \\ x_3 \end{pmatrix} + \begin{pmatrix} \eta_1 \\ \eta_2 \\ \eta_3 \end{pmatrix}$$

Some algebra shows that

$$\alpha_1 = \alpha_2 = \alpha_3 \equiv a = \frac{\theta_0}{1-\theta_1}$$

$$\beta_1 = \gamma_2 = \delta_3 \equiv b = \frac{2\theta_2 + \theta_1(\theta_3 - \theta_2)}{(1-\theta_1)(2+\theta_1)}$$

$$\beta_2 = \beta_3 = \gamma_1 = \gamma_3 = \delta_1 = \delta_2 \equiv c = \frac{\theta_1\theta_2 + \theta_3}{(1-\theta_1)(2+\theta_1)}$$

$$\eta_i = \frac{2\varepsilon_i - \theta_1(\varepsilon_i - \sum_{j \neq i} \varepsilon_j)}{(1-\theta_1)(2+\theta_1)}$$

which may be estimated equation-by-equation as

$$v_i = a + bx_i + c\overline{x}_i + \eta_i \tag{5}$$

The contextual model is identical to (3), but assumes $\theta_1 = 0$ (no peer effects). In that event, the reduced form coefficients become $a = \theta_0$, $b = \theta_2$, $c = \theta_3$, and $\eta_i = \varepsilon_i$. The critical feature to note, however, is that *one cannot distinguish between the contextual and social economy models using reduced form coefficients.* In particular, a non-zero “*c*” coefficient (the coefficient on “other members’ average characteristics) may be evidence of a diversity effect as assumed in the contextual model; but it may well indicate a mix of contextual, peer, and own effects, as implied by the social economy model. Thus, interpreting a non-zero “*c*” coefficient as evidence of a diversity effect is problematic.

The traditional "own characteristics" model may also be expressed in the form of (3), specifying $\theta_1 = \theta_3 = 0$. In that event, $a = \theta_0$, $b = \theta_2$, $c = 0$, and $\eta_i = \varepsilon_i$. Critically, *estimation of the reduced form coefficients does allow one to distinguish the traditional approach from the two approaches allowing social interactions on the panels*. In other words, a non-zero "*c*" coefficient (the coefficient on "other characteristics") provides evidence of social interactions on the panels, rejecting the traditional "own effects" model. However, one cannot tell whether it is diversity or peer effects that give rise to the indicated social interaction. One must estimate the structural equations to do this.

Figure 1 displays illustrative path diagrams for the three models.

(Insert Figure 1 about here)

Empirical Issues and Strategy

Estimating the structural equations (3) rather than the reduced form equations (5) affords the obvious route to distinguishing diversity effect from peer effects. However, as Manski pointed out in a classic paper (1993), estimating systems like (3) is difficult. Two problems are important: 1) the simultaneity problem (the "reflection problem"), and 2) correlated effects, often arising from endogenous group membership.

The simultaneity problem concerns the identification of the system (3). As Manski (2000) explains

> This identification problem arises because mean behavior in the group is itself determined by the behavior of the group members. Hence, data on outcomes do not reveal whether group behavior actually affects individual behavior, or group behavior is simply the aggregation of individual behaviors. This *reflection problem* is similar to the problem of interpreting the (almost)

simultaneous movements of a person and his reflection in a mirror. Does the mirror image cause the person's movement or reflect them? (p. 128)

It is straightforward to show that (3) is not identified under conventional assumptions (that is, it fails to satisfy the rank and order conditions due to the lack of exclusions across the equations). Several scholars have advanced strategies for identifying similar systems. Two are particularly noteworthy.

The first focuses on non-linearities in individual behavior relative to aggregate behavior in the peer group. In particular, if individual behavior responds non-linearly to the group mean – for example, responds to the median rather than the mean, or displays threshold effects – this breaks the tight connection between individual behavior and group mean behavior. If so, the reflection problem is resolved and (3) could be estimated equation-by-equation. In the context of judicial decision making, this approach requires specifying more carefully how whistle-blowing, dissent avoidance, and deliberation actually work. We see this avenue as very promising, but do not pursue it further in this paper.

The second approach uses instrumental variables to break the simultaneity problem. In particular, one seeks an individual level variable or variables whose group average is *not* an element of the individual behavioral equations (Brock and Durlauf 2000). This group average variable can then be used to instrument for mean behavior in the peer group.[4] This is the approach we take in Section 5.

Correlated effects in models of social interactions arise because individuals in the same group "tend to behave similarly because they have similar individual characteristics or face similar institutional environments." (Manski 2000 p. 127). Particularly troublesome are

[4] Case and Katz 1990 affords an illustration of this approach in practice.

correlated effects that arise because of endogenous selection into groups. In the case of the panels of the federal Courts of Appeals, however, assignment of the members to panels is exogenous. Typically, circuit members are joined with one another according to a matching algorithm that assures that each member serves with every other member over some time period (e.g., two years). Cases are assigned to panels as they arrive, and litigants do not know which judges they have drawn until after they have already written and submitted briefs.[5] Thus, membership and case assignment are effectively random.[6] Random assignment of group composition removes many difficult statistical questions and greatly facilitates empirical inference.

Institutional sources of correlated effects may arise from sources other than self-selection, especially doctrinal legacies in the different circuits, changes in the Supreme Court's doctrine, and measured variables within circuits. We create specific variables (discussed below) to address doctrinal legacies within the circuits, and doctrinal changes from the Supreme Court. We also employ fixed effects for each circuit to capture otherwise unmeasured influences on member voting, specific to each circuit.

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[5] For an institutionally sensitive discussion, see Revesz 2001. Revesz notes that the D.C. Circuit, alone among the circuits, announces the composition of panels before litigants prepare briefs.

[6] To the best of our knowledge, among empirical studies of the Courts of Appeals only Ashenfelter, Eisenberg and Schwab 1995 exploits the fact of random assignment of cases.

3. The Affirmative Action Cases

Sources and Definitions

The data consist of information about 537 votes cast on 179 different panels, as judges on the U.S. Courts of Appeals decided affirmative cases from 1971-1999. Table 1 displays basic descriptive information on each of the variables used in the analysis, and some additional variables of interest.

(Insert Table 1 about here)

The cases were identified through key word searches on *Lexis-Nexis*, and reviewed individually to assure that each case involved a challenge to an existing affirmative action plan, or an employer's attempt to implement one through a consent decree or collective bargaining. In each case, the affirmative action question involved race rather than gender. The data is more extensive than cases published in *West's Federal Digest*, but does not include data on cases never published in either *Lexis-Nexus* nor *West's*.[7]

Each case, and the votes in it, were coded for a variety of information, including the case name and citation, date, court, names of judges, votes, as well as biographical data on the judges sitting on the panel as well as the type of case. Votes supportive of affirmative action were coded "1"; non-supportive votes "0". Biographical data on Courts of Appeals judges came largely from the "Auburn" data, while data on district court judges sitting by designation on circuit court panels came largely from the data base maintained by the Federal Judicial Center. On occasion, however, otherwise missing information came from the

[7] We have found no practical way to gather data on unpublished cases. Songer 1990 suggests that differences between published and unpublished cases are surprisingly few.

Almanac of the Federal Judiciary, *The American Bench*, and other biographical searches. Information coded included the identity and NOMINATE score of the appointing president, the race and gender of the judge, his or her age and birth date. Judges' racial/ethnic background was coded as African-American, Hispanic, Asian, or white, collapsed to "non-white" and "white."

Nine categories of case type were coded, indicating whether cases dealt with hiring preferences, promotion preferences, governmental contracting, set-aside plans, modifications to seniority systems in layoff plans, preferences intended to increase union membership, preferences in school admissions, preferences in training programs, preferences in recruiting programs, and challenges to the legitimacy of Executive Order (EO) 11246, which promoted the "full realization of equal employment opportunity through a positive, continuing program in each executive department and agency."

A variable to capture Circuit Court precedent was created in the following way. For each Court of Appeals, "legacy" was implemented as the average of case outcomes (pro- or anti-affirmative action) decided in the circuit prior to the extant case. Thus, this variable ranges between zero (all previous affirmative action cases in the circuit were decided in an anti-affirmative action fashion), and one.

We capture changing Supreme Court doctrine toward affirmative action by defining dummy variables for different doctrinal eras. Between 1971 and 1999, the Supreme Court decided eleven affirmative action cases. Four had a significant effect on affirmative action doctrine: 1) *Regents of the University of California v. Bakke* (1978), 2) *Firefighters Local 1784 v. Stotts* (1984), 3) *City of Richmond v. J.A. Croson* (1989), and 4) *Adarand Contractors v. Pena* (1995). The first case is clearly a landmark, with the next two moving doctrine in steadily more conservative direction. However, all three were heavily influenced by relatively moderate justices, especially Justices Powell and Stevens. The last case constituted a much more conservative departure, corresponding to the emergence of a solidly conservative majority on the high court. We discuss these cases in a brief appendix. The four cases define five eras: pre-1978, 1978-1984, 1984-1989, 1989-1995, and post-1995.

Critical variables for this study include contextual effect variables and a peer effect variable, the vote of panel colleagues. As discussed earlier, these variables indicate mean values for the panel members other than the judge in question (e.g., for judge 1, the mean race value for judges 2 and 3). Contextual variables include the mean value for panel colleagues with respect to race, gender, and ideology (NOMINATE score of appointing president), as well as age and birth date. The critical peer effect variable is, of course, the mean vote for panel colleagues. This variable, $\bar{v}_i$, takes values of 0, .5, and 1. Following the discussion in Section 2, the variable used in the statistical analysis is $(\bar{v}_i - .5)$.

Data Exploration

Figure 2 graphically displays the data, showing simple relationships between own characteristics and the probability of a pro-affirmative action vote. As shown, on average non-white judges were about 20 percentage points more likely to vote in a pro-affirmative direction than white judges. Gender appears to have had little influence on voting for affirmative action. Ideology appears to have exerted considerable influence, with the probability of a pro-affirmative action vote declining steeply for judges with ideology scores above zero (conservates).[8]

(Insert Figure 2 about here)

Figure 3 is similar to Figure 2, but focusing on the impact of contextual variables, that is, the mean values of a judge's colleagues on a panel. The left-hand figure suggests that adding a single non-white member to a panel apparently increased the probability of pro-affirmative action voting of white judges by about fifteen percentage points. There are too few observations on majority non-white panels to draw firm conclusions about such panels. The right-hand figure suggests a weak relationship, at best, between adding a female judge to a panel and the pro-affirmative action voting of the male judges. Again, there are too few cases of majority female panels to draw conclusions.

(Insert Figures 3 and 4 about here)

Figure 4 continues the examination of contextual variables, focusing on ideology. As shown in the upper left figure, context appears to have a dramatic effect on voting among judges with "liberal" scores. If a panel were composed entirely of strong liberals (so the mean value of colleagues' ideology score was less than -.34), the probability of pro-affirmative action votes was about 80 percent. But if the rest of the panel was composed of judges with conservative scores (so the mean for panel colleagues was .4 or greater) the probability of a pro-affirmative action vote fell by about 20 percentage points. One can interpret these data as supportive of strong context or peer effects.

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[8] This figure shows the fit from a locally weighted regression (a loess regression). The points have been jittered slightly to better show the distribution of observations.

The right-hand panel in Figure 4 examines similar effects among judges with conservative scores. Rather strikingly, however, the effects of ideological context were quite different for conservative judges than for liberal ones. If conservatives sat on panels with colleagues whose mean scores were moderate or conservative (above zero), the probability of a pro-affirmative action vote was about 50 percent, or slightly higher. But if a judge sat on a panel with quite liberal colleagues, the probability of a pro-affirmative action vote by the conservative decreased dramatically, to as low as 25 percent for a panel with very liberal colleagues. Because the Supreme Court's doctrine has been relatively (and increasingly) conservative on affirmative action, it is difficult not to interpret this pattern as whistle-blowing, as discussed in Cross and Tiller 1998.

Figures 5-7 display additional data on pro-affirmative action voting, by Circuit, by era of Supreme Court doctrine, and by appointing president.

(Figures 5-8 about here)

(Table 2 about here)

Table 2 presents suggestive evidence on apparent peer influences. Each column indicates a different environment created by peers' actions. Each row indicates a group that might be expected to have pro-affirmative action leanings (the top panel) or anti-affirmative action leanings (bottom panel). In the third column in each panel, the peer environment supports the group's presumptive predilections. In this case, it is no surprise to find group members voting overwhelmingly in the "natural" fashion. For example, in the 55 voting opportunities in which a conservative sat on a panel and both of the other members cast anti-affirmative action votes, the conservative also cast an anti-affirmative action vote 100% of the time.

The middle column is more intriguing, with panel members facing split votes. Thus, one might regard 50% as a reasonable baseline in this case. Instead, liberals cast pro-affirmative action votes 81% of the time, conservatives only 25% of the time, and African-Americans 100% of the time (though only 4 such opportunities existed for black judges). The effect of "predilections" seems large.

The first column is by far the most interesting, however. Here, the judges in the group presumably face cross-pressures, with personal predilections pushing one way and peer actions pushing the other. As shown, peer actions appear to be far more influential. For example, in these "peer pressure" situations, liberal judges cast pro-affirmative action votes only 15% of the time and African-American judges only 33% of the time. Conservatives, facing two pro-affirmative action votes, joined their colleagues and cast pro-affirmative action votes a remarkable 90% of the time. Of course, these data cannot address the mechanism at work, but the main point is clear: whatever the mechanism, peer influences appear profound.

4. Evidence of Social Interactions

In this section, we estimate the reduced form version of the model. Given the assumptions indicated above, equation (5) can be estimated via generalized least squares (GLS) (Achen 1986 pp. 40-42).[9]

(Insert Table 3 about here)

[9] In other words, the model is estimated with OLS, and then the standard errors are "Goldbergerized."

Table 3 displays estimates of the reduced form equations. Because of the high degree of collinearity between the year-of-birth, average-year-of-birth, age-at-time-of-decision, and average-age variables, we show two separate regressions. The first uses the birth variables, the second the age variables.

As discussed in Section 2, the reduced form coefficients in Table 3 should not be interpreted in a naïve way. For example, the reduced form coefficients on "own characteristics" such as race and ideology do not reflect just own-characteristics effects (θ_2), but a mixture of own-characteristics effects and peer effects (θ_1). (Recall from Section 2 that $b = \frac{2\theta_2 + \theta_1(\theta_3 - \theta_2)}{(1-\theta_1)(2+\theta_1)}$.) Thus, a reduced form estimate of "*b*" may indicate a non-zero effect – but this could occur when the behavioral parameter on own-effect, θ_2, is actually zero if the behavioral parameters on other-characteristic, θ_3, and peer effect, θ_1, are non-zero. That is, judge 1 might not be affected by her own characteristic, but if judges 2 and 3s are, and judge 1 is sensitive to peer effects, she will appear to be sensitive to her own characteristic. We do not suggest this effect is common, but it underscores the importance of using structural rather than reduced form estimates to make behavioral inferences.

The contextual effects in the reduced form equations are more rewarding. As discussed in Section 2, the traditional own-effects model implies that the coefficients on other-characteristics variables must be zero. As shown in Table 3, this is indeed the case for "average gender of panel colleagues," "average birth year of panel colleagues," and "average age of panel colleagues." However, the coefficient on "average ideology of panel colleagues" is clearly not zero, thus rejecting the traditional model in favor of a model with social interactions related to ideology. Of course, it is not clear whether the interaction is a diversity effect or a peer effect. The coefficient on “average race of panel colleagues” is rather large, takes the correct sign, and approaches traditional levels of significance. In our view, the reduced form equations leave open the question whether social interactions occur in connection with race.

Zero-value coefficients on contextual variables supply some leverage in interpreting the own-characteristic estimates of those variables. Recall from Section 2 that the reduced form contextual coefficient $c = \frac{\theta_1\theta_2 + \theta_3}{(1-\theta_1)(2+\theta_1)}$. If this coefficient is zero, then clearly $\theta_3 = 0$ and $\theta_1\theta_2 = 0$. But if $\theta_3 = 0$ then $b = \frac{2\theta_2 - \theta_1\theta_2}{(1-\theta_1)(2+\theta_1)}$, and if this coefficient is not zero then $\theta_2 \neq 0$, that is, the own-characteristic indeed affects own behavior. Combining the estimates of other- and own-characteristic effects in this fashion leads to the conclusion that own age affects the probability of pro-affirmative action. Gender appears not to affect own behavior. The results on birth cohort are ambiguous.

We defer discussion of case specific and circuit level variables until the next section.

5. Diversity and Peer Effects

Variables that affect own behavior, but do not affect other members’ behavior through group averages, provide instruments for disentangling own, diversity, and peer effects. That is, peer average effects in these variables can be used to instrument for mean other votes, as discussed in Section 2. Critically, age and (possibly) birth cohort provide such variables. As shown above, age affects own voting while peer average levels of age do not. Peer average birth cohort does not affect voting, but birth cohort may weakly affect own voting.

Do age and birth cohort variables make sense as instruments? Some arguments suggest the answer is "yes." Younger judges and more recent birth cohorts appear less favorably inclined toward affirmative action (controlling for ideology, precedent and many other variables) than older ones. A standard finding in social psychology is that negative attitudes toward affirmative action are associated with personal threat and competition, and perceptions that opportunity is not unequal (Kluegel and Smith 1983, Krysan 2000). It is younger judges and cohorts who are closer to career or educational environments in which affirmative action may have been perceived as personally threatening. In addition, younger judges may perceive American society as more open to equal opportunity than older judges, whose memories of Jim Crow and the battles of the civil rights era may resonate with this policy issue. Finally, positive attitudes toward affirmative action are associated with feelings of white guilt (Swim and Miller 1999). Older judges may feel more complicit in having unfairly benefited from earlier racially exclusionary policies in education and the legal profession.

Table 4 presents GLS estimates for a potential IV equation, estimating average vote of colleagues as a function of the two instruments, average birth year of panel colleagues and average age of panel colleagues. In addition, we use average gender of panel colleagues (shown in Section 4 not to affect individual behavior through group averages), although this is clearly a weak instrument because of the tenuous relationship between gender and own vote. As shown, average birth year is strongly related to average colleagues votes, average age of panel colleagues plausibly but less strongly related, and average gender of colleagues weakly related. Overall, the equation appears to perform reasonably well.

We therefore estimate structural parameters, using these variables to instrument for average panel colleagues' vote.[10] We employ Achen's generalized two-stage least squares (G2SLS) estimator (Achen 1986, pp. 44-45).[11] Achen shows that the resulting coefficient estimates are consistent, and that within the wide class of single-equation instrumental variables estimators, G2SLS is fully efficient asymptotically.

(Insert Table 5 about here)

In contrast with the reduced form estimates in Table 3, the behavioral parameters in Table 5 have straight-forward interpretations. Among own-characteristics, race, ideology, age, and birth cohort exert clear direct effects on the probability of pro-affirmative action voting. Of these effects, the largest is ideology. For example, moving from liberal score (-.35) to a conservative one (.35) reduces the probability of voting in a pro-affirmative action direction by 77 percentage points. Also large is the effect of race: non-white judges are ten percentage points more like to vote in a pro-affirmative action fashion than white judges.

From a policy perspective, the most interesting finding is the effect of average race of panel colleagues. Adding a single non-white to a panel (so the value of the variable increases from 0 to .5) directly increases the probability of a pro-affirmative action vote by his or her colleague by about five percentage points. This statistically significant effect is about half the size of the own-characteristic effect for race, and thus appears relatively substantial.

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[10] In fact, in line with the theoretical derivation in Section 2, the peer effect variable is $\overline{v}_i - .5$, average vote of panel colleague minus one-half. Thus, the variable ranges from -.5 to .5.

[11] Neither Stata nor LIMDEP provide G2SLS, but programming it is straightforward. The authors' S-Plus code for G2SLS is available on request.

Moreover, this is a direct “deliberation effect” distinct from a peer pressure effect, which magnifies the effect through feedback into the other judges’ votes.

Interestingly, the direct effect of average ideology of colleagues is apparently zero. That is to say, ideological diversity brings no deliberation effect. Rather, the impact of ideology on other judges seems to operate exclusively from through peer effects.

Peer effects are statistically significant and huge in magnitude. If the votes of colleagues are split, there is no peer effect by construction. But if both colleagues vote against affirmative action, this alone reduces the probability of a pro-affirmative action vote by the third judge by 26 percentage points, relative to a split vote. Two anti-affirmative action votes reduce the probability of a pro-vote by the third judge by 52 percentage points, relative to two pro-affirmative action votes – they tip his vote. Peer effects appear to be large indeed.

Among the remaining variables, the Supreme Court era variables and court legacy variables are particularly interesting. The “era” variables (defined relative to the most conservative, post-1995 era) indicate a continuing downward pressure on pro-affirmative action votes over time. These conform with the standard doctrinal understanding of the Supreme Court’s doctrine, and demonstrate direct responsivesness of the Courts of Appeals judges to the directions set by the high court.

The “court legacy” vote is highly statistically significant, large in magnitude, and negative in sign. The interpretation of this variable is somewhat subtle. This variable takes values between zero and one, and is the average of preceding outcomes in affirmative action cases heard in the circuit. Thus, a negative sign indicates that there is pressure to “pull in” circuit courts that established early records supporting affirmative action (like the First

Circuit), and this effect is larger for circuits for more supportive legacies. For example, if a legacy value is .5, the “reining in” effect lowers the probability of a pro-affirmative action vote by 20 percentage points. But if the legacy is .25, the effect on votes is only 10 percentage points. Over time, as the “reining in”effect gradually lowers the legacy of pro-affirmative action doctrine in the circuits, all will converge on a zero legacy (in the limit). Although this strong result surely reflects the simple linear specification of the effect, we interpret the legacy effect as reflecting continuing pressure from the Supreme Court, both to rein in affirmative action and to bring uniformity in policy across the circuits.

// In the next iteration of the paper, we will discuss the total effects of variables (direct and mediated through peer pressure), utilizing the estimated behavioral parameters and the derived reduced form equations //

6. Conclusion

In this paper, we develop a new methodology for studying decision making on collegial courts. We applied the new approach to the affirmative action cases heard in the U.S. Courts of Appeals between 1971 and 1999.

Several caveats are in order, especially in considering the policy implications of the findings. Estimating structural parameters always requires relatively strong assumptions. Our analysis is no different. Only if the instruments are plausible will the conclusions be sound and, similarly, assumptions about functional forms – e.g., the linearities we have assumed – are also critical. Thus, a healthy degree of skepticism should be directed at the results from any single study. Firm conclusions require multiple studies, data from different kinds of cases, and experience with different assumptions.

That said, we believe the picture drawn by our statistical analysis is quite plausible. Most observers of the U.S. Courts of Appeals stress the strong peer pressures at work there. Our analysis is the first to explicitly build such effects into empirical models of voting, and we indeed uncover very substantial pressures. The implications of such strong peer effects are intriguing for the governance of judicial systems, raising questions about local equilibria, tipping, and herd behavior. The findings are also interesting from the perspective of "principal-agent" approaches to the judicial hierarchy. In particular, we also find evidence of considerable responsiveness by the Courts of Appeals to doctrinal changes by the Supreme Court, and pressures to align doctrine across the circuits. Again, we believe this comports well with a variety of qualitative and historical evidence. Clearly, the ideology effects we uncover raise many additional questions. The absence of strong contextual effects, combined with the large peer effects, suggests the strategic effects discussed by Revesz and Cross and Tiller.

From a policy perspective, however, the most interesting finding concerns racial diversity. At least in the U.S. Courts of Appeals, and in affirmative action cases, racial diversity apparently works the way proponents believe it ought to: different perspectives and life experiences within a community broaden the views of the members of the community and alter their deliberations. In our view, the strongest argument in favor of diversity in the courts is social legitimacy: democratic institutions in heterogeneous societies ought to reflect the make-up of society, if only to offset memories of exclusion. This is sufficient reason to support racial and gender diversity in the federal judiciary. But racial diversity seems also to affect the decisions of federal judges, operating directly through a contextual or deliberative effect, rather than exclusively through peer pressure. In this sense, efforts to build a more diverse federal judiciary have borne fruit.

Appendix: Doctrinal Eras and Supreme Court Decisions

Each of the eleven race-based affirmative action decisions handed down by the U.S. Supreme Court between 1971 and 1999 had major policy ramifications for public or private organizations. Nevertheless, four cases appear to have had particularly significant impact on affirmative action doctrine. The first landmark decision was handed down in 1978 in *Regents of the University of California v. Bakke*. This first and most famous affirmative action case prohibited the use of racial quotas, but affirmed the use of race as one factor in the consideration of minority applications to the University of California-Davis Medical School. Because the Court split with no majority on nearly every major issue of the case, no clear ideological trend was established. However, the Court supported a race-conscious affirmative action plan for the first time, because the plan achieved a compelling government interest (which in this case was diversity in the student body). Thus, the *Bakke* decision can be interpreted as a modestly liberal outcome. The next major case, *Firefighters Local 1784 v. Stotts*, was decided in 1984. The Court ruled that preferences to racial minorities should not be afforded in layoff procedures, and that seniority is a valid criterion in reduction of force procedures, despite the disproportionate number of minorities that may consequently lose employment. The Court's decision in *Stotts* indicated the beginning of a conservative shift in affirmative action doctrine. The conservative trend was strengthened five years later in *City of Richmond v. J.A. Croson*, where the Court held that the "strict scrutiny" standard of review would be applied to any affirmative action plan enacted by a state or local government. Under the "strict scrutiny" standard, a race-based affirmative action plan will only be found constitutional if it meets both prongs of a stringent two-part test. To meet the standard of strict scrutiny, affirmative action plans must (1) be adopted to further a *compelling* governmental interest, and (2) only use racial classifications when *necessary* to achieve that compelling governmental interest.[12] The final, and arguably most conservative decision occurred in 1995 in *Adarand Contractors v. Pena*. The Court announced that strict scrutiny must be applied to race-based affirmative action plans imposed by Congress, just as it is applied to those imposed by state and local governments. The reference category for the circuit dummies is the Pre-*Bakke*, or Pre-1978, era.

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[12] Possible *compelling* interests include: the redressing of clear past discrimination (in *Croson*, Richmond was not able to provide evidence of past discrimination); diversity in the student body (*Bakke*), diversity for the police (suggested by Kennedy in *Croson*); diversity in radio station broadcasts (*Metro Broadcasting Co. v. FCC*). *Necessary* requires that a racially-neutral plan first be attempted. Quotas are not seen as necessary, because other remedies are available.

Table 1: Descriptive Statistics

Variable	Mean	Std. Dev.	Min	Max
		Own Characteristic Variables		
Vote	0.635	0.482	0	1
Race	0.073	0.260	0	1
Gender	0.076	0.266	0	1
Ideology	0.043	0.467	-0.535	0.568
Birth Year	1925	11.61	1888	1954
Age	61.6	9.963	37	95
		Contextual Effects Variables		
Average Race of Panel Colleagues	0.074	0.191	0	1
Average Gender of Panel Colleagues	0.076	0.185	0	1
Average Ideology of Panel Colleagues	0.043	0.339	-0.535	0.568
Average Birth Year of Panel Colleagues	1925	8.830	1899	1949
Average Age of Panel Colleagues	61.6	6.512	46	79
		Peer Effects Variable		
Average Vote of Panel Colleagues	0.635	0.448	0	1
		Case Specific Variables		
Outcome	0.648	0.478	0	1
Unanimous Decision	0.810	0.393	0	1
Pre-*Bakke* (Pre-1978)	0.151	0.358	0	1
Post *Bakke* (1978-1984)	0.246	0.431	0	1
Post *Stotts* (1984-1989)	0.162	0.369	0	1
Post *Croson* (1989-1995)	0.296	0.457	0	1
Post *Adarand* (Post 1995)	0.145	0.352	0	1
Hiring	0.251	0.434	0	1
Promotion	0.324	0.468	0	1
Contracting	0.223	0.417	0	1
Seniority	0.128	0.335	0	1
Union	0.067	0.250	0	1
Training	0.034	0.180	0	1
Recruitment	0.034	0.180	0	1
Schools	0.050	0.219	0	1
Executive Order 11246	0.022	0.147	0	1
		Circuit Level Variables		
Court Legacy	0.687	0.224	0	1
D.C. Circuit	0.061	0.240	0	1
1st Circuit	0.050	0.219	0	1
2d Circuit	0.111	0.315	0	1
3d Circuit	0.080	0.272	0	1
4th Circuit	0.086	0.280	0	1
5th Circuit	0.214	0.411	0	1
6th Circuit	0.158	0.365	0	1
7th Circuit	0.052	0.223	0	1
8th Circuit	0.037	0.190	0	1
9th Circuit	0.099	0.299	0	1
10th Circuit	0.019	0.135	0	1
11th Circuit	0.031	0.175	0	1

Note: N=537

Table 2. Apparent Peer Effects in Different Groups

Facing	2 Anti-AA Votes	1 Anti-AA Votes	0 Anti-AA Votes	Total
Liberal	11/71 (.15)	21/26 (.81)	134/139 (.96)	166/236 (.70)
Black	3/9 (.33)	4/4 (1)	26/26 (1)	33/39 (.85)
Woman	1/14 (.07)	3/3 (1)	23/24 (.96)	27/41 (.66)
Total	15/94 (.16)	28/33 (.85)	183/186 (.98)	226/316 (.72)

Facing	2 Pro-AA Votes	1 Pro-AA Votes	0 Pro-AA Votes	Total
Conservative	64/71 (.90)	5/20 (.25)	0/55 (0)	69/146 (.47)
White	262/281 (.93)	34/64 (.53)	12/153 (.08)	308/498 (.62)
Male	265/283 (.94)	35/65 (.54)	14/148 (.09)	314/496 (.63)
Total	496/535 (.93)	(74/149) (.54)	26/356 (.07)	691/1140 (.61)

Numbers in numerators indicate "pro-affirmative action" votes
Numbers in denominators indicate opportunities for group members (e.g., women) to cast votes under the indicated condition.
Liberal = Nominate score of appointing president less than -.39.
Conservative = Nominate score of appointing president greater than .34.
Number in parenthesis is percentage pro-affirmative action votes in cell.

Table 3: Estimates of Reduced Form Equations (GLS)

Variable	Model 1 (using Birth) Coeff.	SE	t-value	Pr(>|t|)	Model 2 (using Age) Coeff.	SE	t-value	Pr(>|t|)
Own Characteristic Variables								
Race	0.135	0.041	3.291	0.001	0.140	0.041	3.429	0.000
Gender	0.014	0.058	0.241	0.810	0.016	0.057	0.278	0.781
Ideology	-0.101	0.033	-3.063	0.002	-0.100	0.033	-3.028	0.003
Birth Year	-0.002	0.001	-1.047	0.296	-------	-------	-------	-------
Age	-------	-------	-------	-------	0.003	0.002	1.629	0.104
Contextual Effects Variables								
Avg. Race of Panel Colleagues	0.091	0.765	1.189	0.235	0.103	0.077	1.335	0.183
Avg. Gender of Panel Colleagues	0.035	0.096	0.369	0.712	0.033	0.096	0.341	0.733
Avg. Ideology of Panel Coll.	0.284	0.075	-3.336	0.000	-0.139	0.046	-3.020	0.003
Avg. Birth Year of Panel Coll.	0.000	0.002	0.185	0.854	-------	-------	-------	-------
Avg. Age of Panel Colleagues	-------	-------	-------	-------	0.002	0.003	0.601	0.548
Case Specific Variables								
Pre-*Bakke* (Pre-1978)	0.202	0.082	2.461	0.014	0.237	0.062	3.800	0.000
Post-*Bakke* (1978-1984)	0.181	0.071	2.552	0.011	0.209	0.059	3.565	0.000
Post-*Stotts* (1984-1989)	0.188	0.066	2.839	0.005	0.212	0.063	3.358	0.001
Post-*Croson* (1989-1995)	0.239	0.054	4.448	0.000	0.252	0.050	5.018	0.000
Hiring	-0.012	0.034	-0.346	0.730	0.003	0.033	0.078	0.938
Seniority	-0.115	0.056	-2.065	0.039	-0.115	0.055	2.071	0.039
Promotion	0.187	0.034	5.556	0.000	0.185	0.033	5.640	0.000
Union	0.234	0.057	4.084	0.000	0.241	0.057	4.200	0.000
Training	0.185	0.081	2.272	0.024	0.213	0.082	2.58	0.196
Recruitment	0.005	0.076	0.063	0.950	-0.028	0.076	-0.369	0.712
Schools	-0.124	0.072	-1.719	0.086	-0.114	0.069	-1.639	0.102
Executive Order 11246	0.195	0.101	1.928	0.054	0.199	0.095	2.098	0.036
Circuit Level Variables								
Court Legacy	-0.325	0.072	-4.511	0.000	-0.323	0.073	4.447	0.000
D.C. Circuit dummy	0.046	0.046	0.463	0.644	0.053	0.098	0.541	0.589
1st Circuit dummy	0.609	0.076	7.978	0.000	0.609	0.074	8.200	0.000
2d Circuit dummy	0.421	0.091	4.650	0.000	0.411	0.090	4.549	0.000
3d Circuit dummy	0.648	0.073	8.875	0.000	0.647	0.071	9.122	0.000
5th Circuit dummy	0.352	0.066	5.295	0.000	0.340	0.064	5.344	0.000
6th Circuit dummy	0.442	0.077	5.748	0.000	0.458	0.074	6.165	0.000
7th Circuit dummy	0.445	0.095	4.684	0.000	0.438	0.093	4.700	0.000
8th Circuit dummy	0.634	0.077	8.160	0.000	0.627	0.075	8.350	0.000
9th Circuit dummy	0.462	0.080	5.794	0.000	0.459	0.078	5.874	0.000
10th Circuit dummy	0.035	0.149	0.238	0.812	0.032	0.147	0.219	0.827
11th Circuit dummy	0.124	0.111	1.111	0.267	0.125	0.110	1.134	0.257
Intercept	2.299	6.204	0.371	0.711	-0.039	0.245	-0.160	0.873
F-Statistic	178.5		p < 0.000		179.6		p <0.000	

Table 4: Instrumental Variable Estimates (GLS)

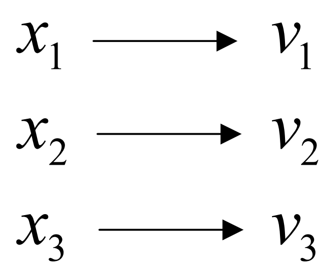
Variable	Coeff.	SE	t-value	Pr(>|t|)
Contextual Effects Variables				
Avg. Birth Year of Panel Coll.	-0.014	0.003	-5.499	0.000
Avg. Age of Panel Colleagues	-0.005	0.004	-1.411	0.159
Avg. Gender of Panel Colleagues	0.095	0.102	0.935	0.350
Intercept	26.926	4.852	5.550	0.000
F-Statistic	314.6		p < 0.000	

Table 5: Estimates of Structural Equation (G2SLS)

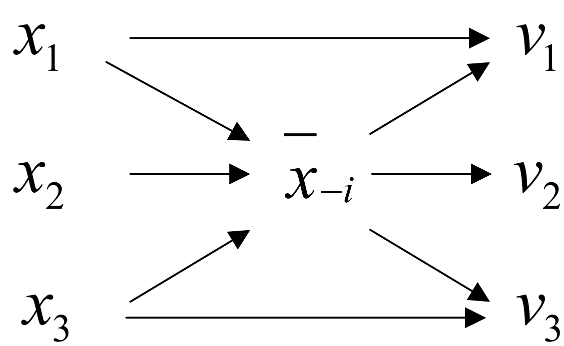
Variable	Coeff.	SE	t-value	Pr(>|t|)
Own Characteristic Variables				
Race	0.108	0.030	3.564	0.000
Gender	0.030	0.036	0.828	0.408
Ideology	-0.110	0.022	-5.085	0.000
Birth Year	0.047	0.007	6.381	0.000
Age	0.051	0.008	6.758	0.000
Contextual Effects Variables				
Avg. Race of Panel Colleagues	0.091	0.052	1.745	0.082
Avg. Ideology of Panel Coll.	0.005	0.029	0.160	0.873
Peer Effects Variable				
Average Vote of Panel Colleagues	0.519	0.250	2.072	0.039
Case Specific Variables				
Pre-*Bakke* (Pre-1978)	0.975	0.128	7.601	0.000
Post-*Bakke* (1978-1984)	0.894	0.093	9.662	0.000
Post-*Stotts* (1984-1989)	0.619	0.073	8.536	0.000
Post-*Croson* (1989-1995)	0.425	0.040	10.606	0.000
Hiring	0.000	0.023	0.011	0.991
Seniority	-0.282	0.035	-8.175	0.000
Promotion	0.110	0.022	5.060	0.000
Union	0.264	0.042	6.256	0.000
Training	0.299	0.049	6.056	0.000
Recruitment	0.005	0.043	0.104	0.917
Schools	-0.270	0.039	-7.015	0.000
Executive Order 11246	0.107	0.084	1.276	0.203
Circuit Level Variables				
Court Legacy	-0.402	0.052	-7.742	0.000
D.C. Circuit dummy	-0.019	0.058	-0.330	0.742
1st Circuit dummy	0.889	0.049	18.212	0.000
2d Circuit dummy	0.437	0.055	7.993	0.000
3d Circuit dummy	0.842	0.044	19.103	0.000
5th Circuit dummy	0.429	0.037	11.757	0.000
6th Circuit dummy	0.544	0.045	12.116	0.000
7th Circuit dummy	0.575	0.062	9.290	0.000
8th Circuit dummy	0.752	0.049	15.513	0.000
9th Circuit dummy	0.586	0.047	12.467	0.000
10th Circuit dummy	0.035	0.149	0.238	0.812
11th Circuit dummy	0.212	0.080	2.668	0.008
Intercept	-94.321	14.779	-6.382	0.000

Figure 1

Traditional Approach



Contextual Approach



Social Economic Approach

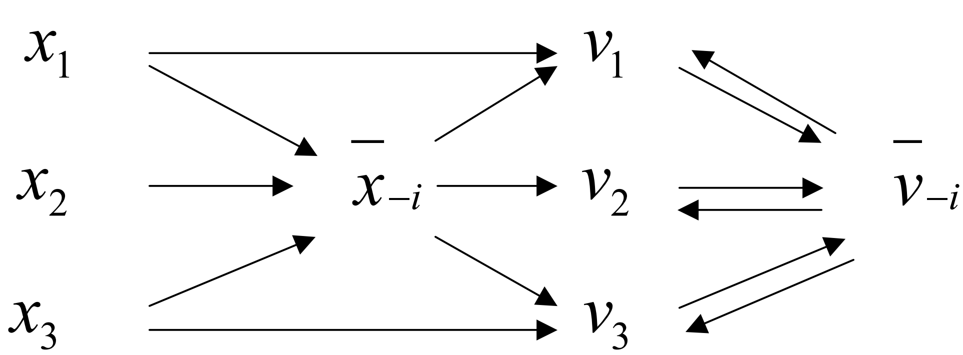


Figure 2. Voting and Own Characteristics

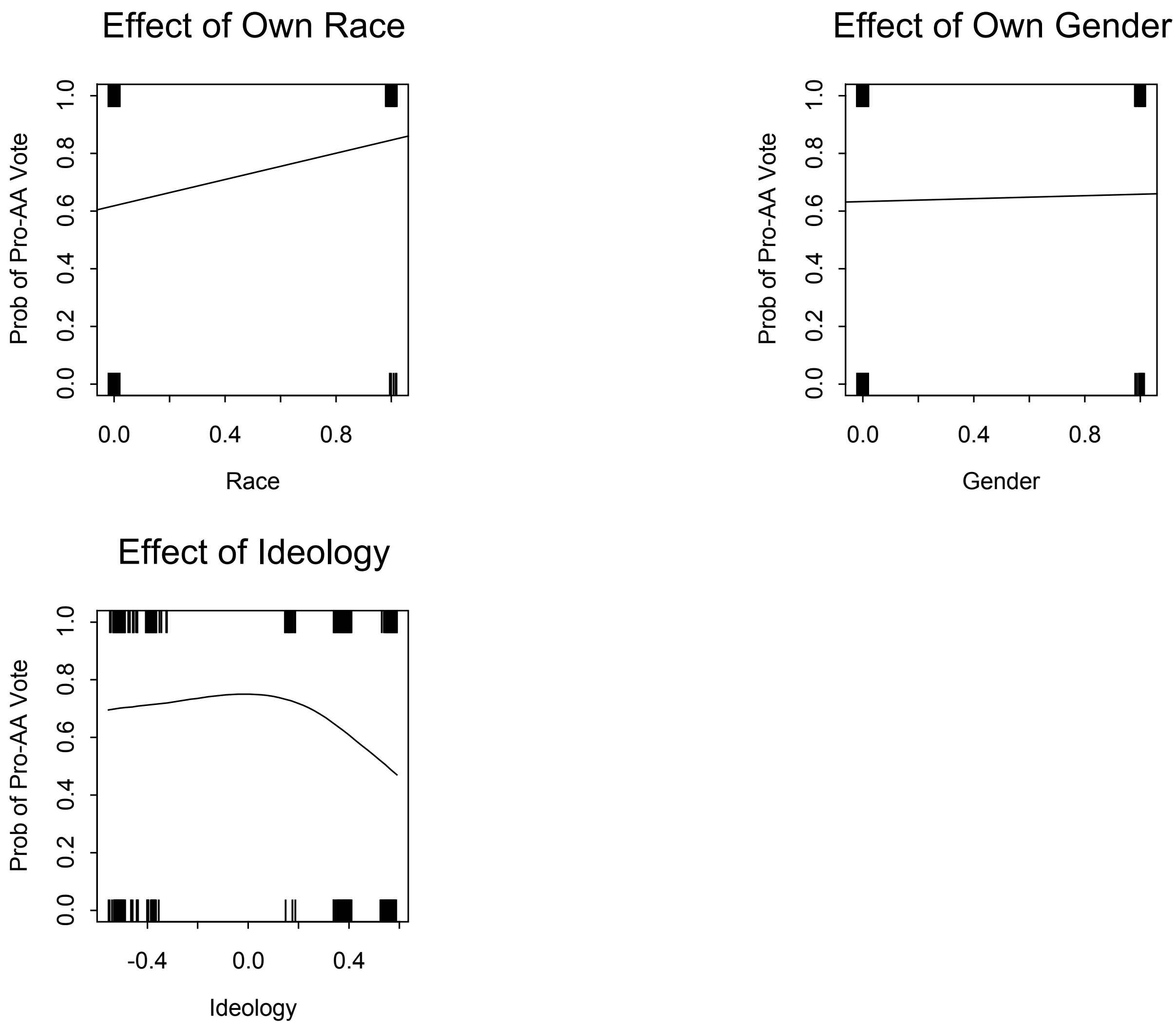


Figure 3. Voting and Contextual Variables: Race & Gender

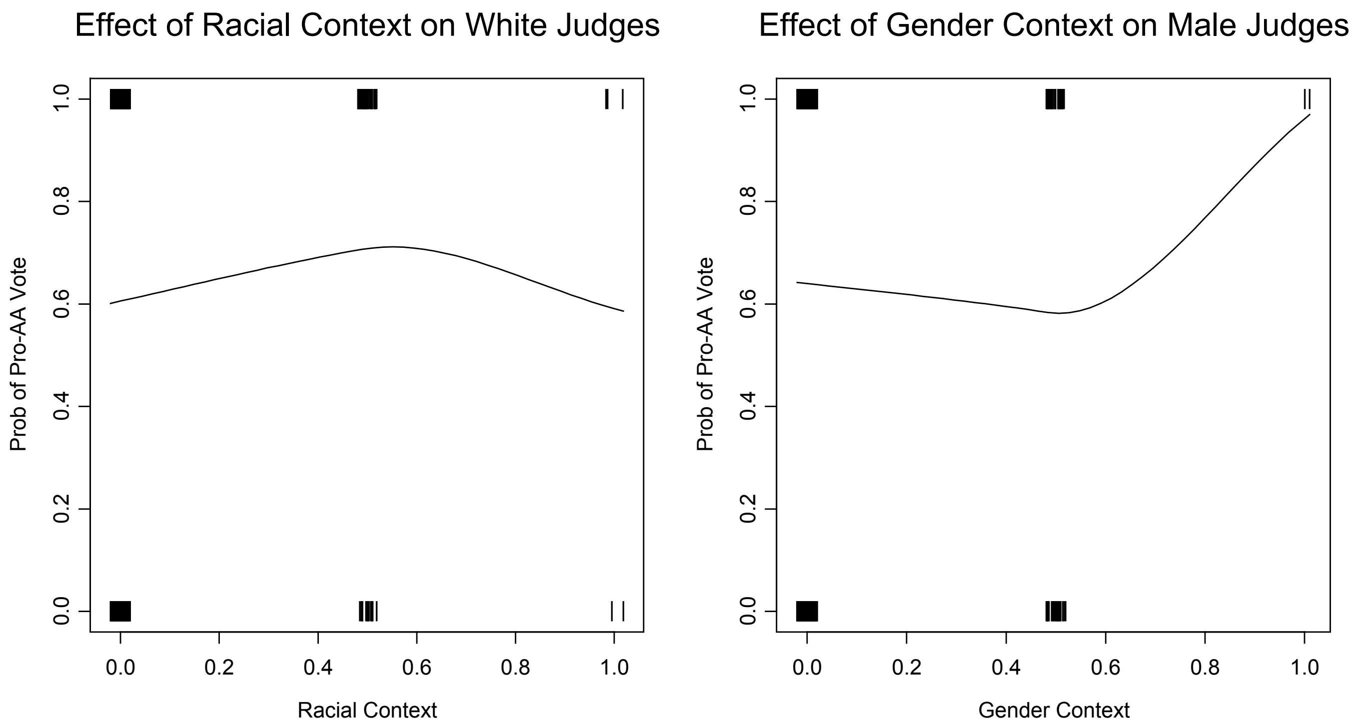


Figure 4. Voting and Contextual Variables: Ideology

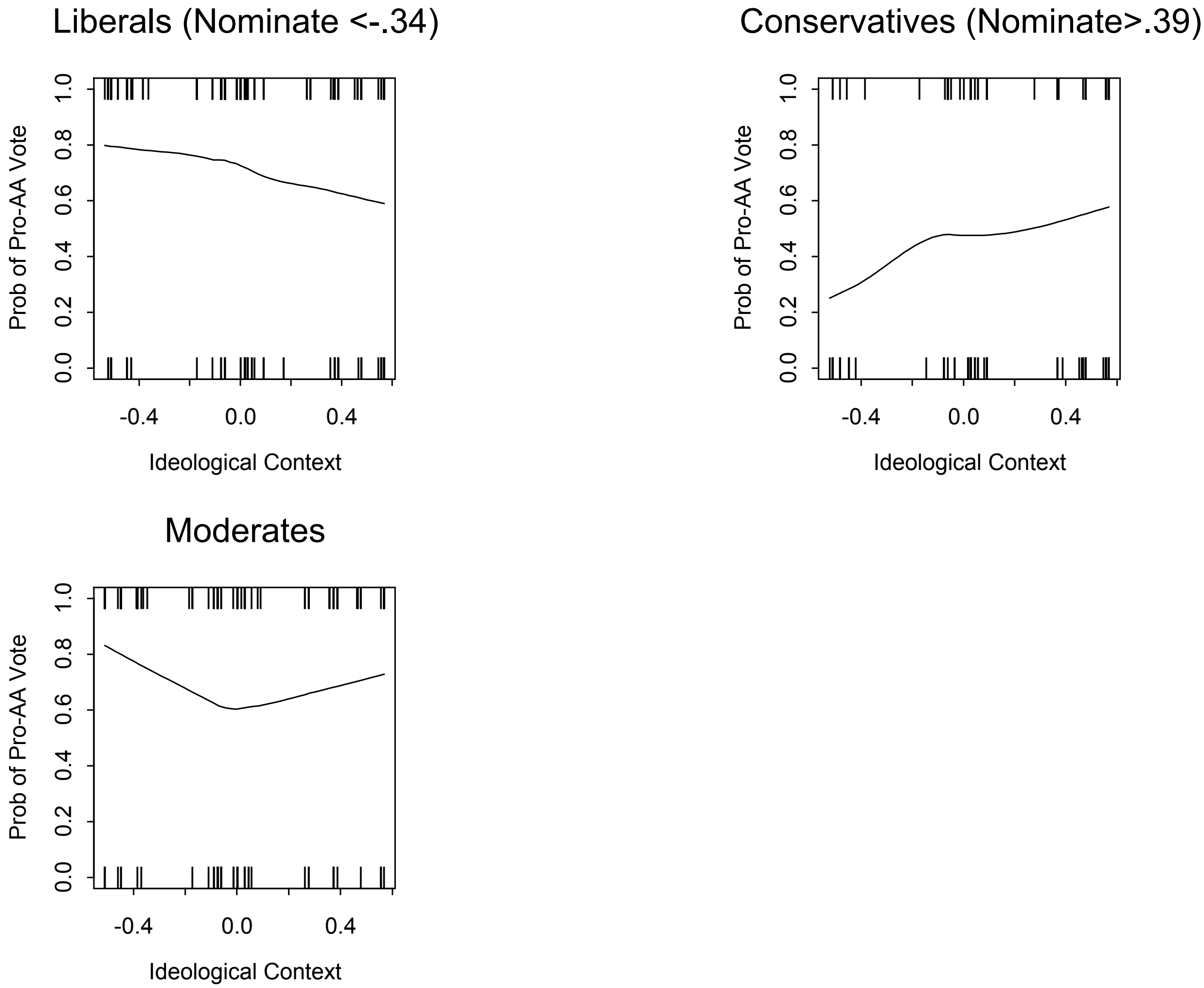


Figure 5. Pro-AA Voting by Circuit

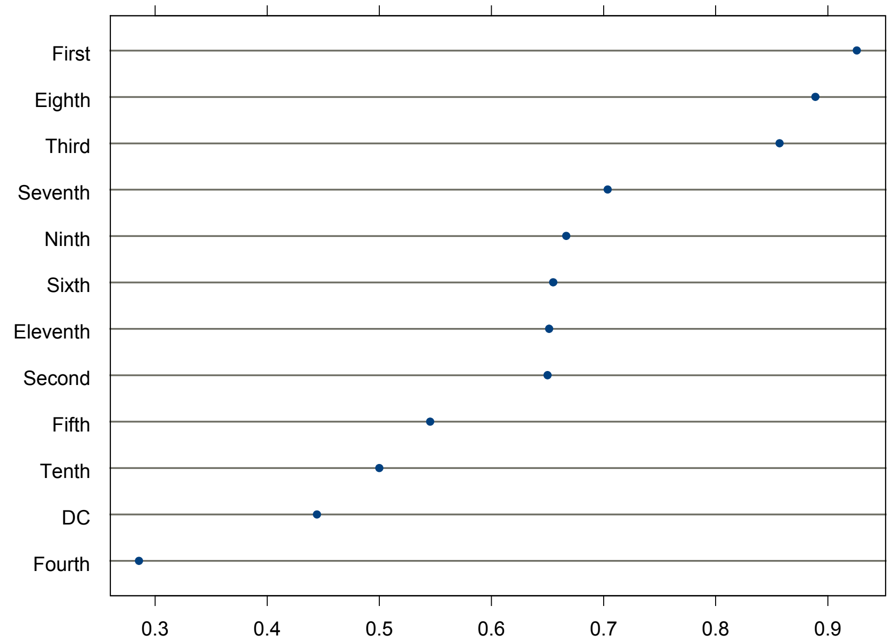


Figure 6. Pro-AA Voting by Appointing President

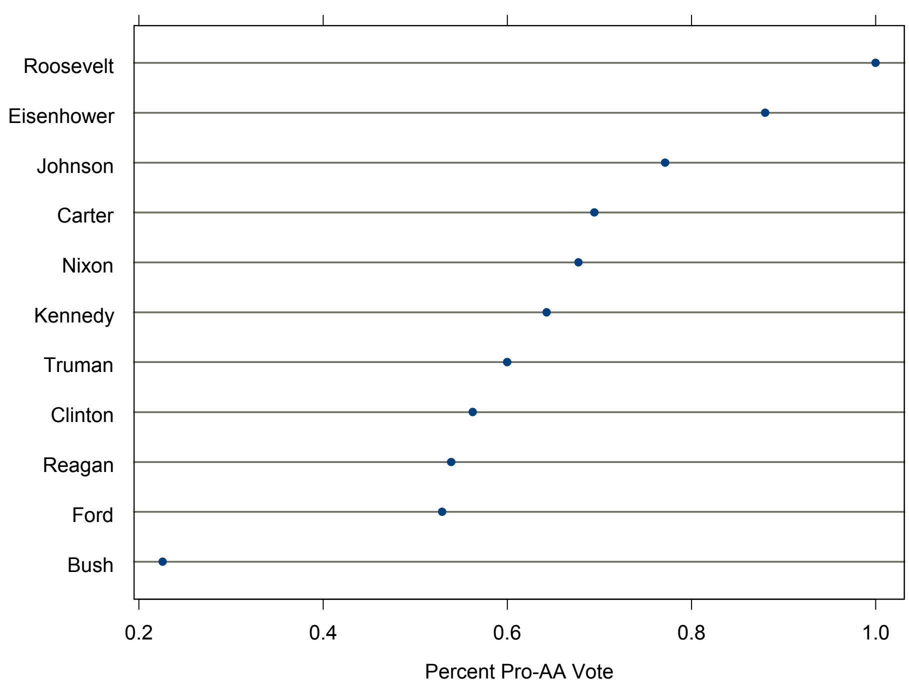
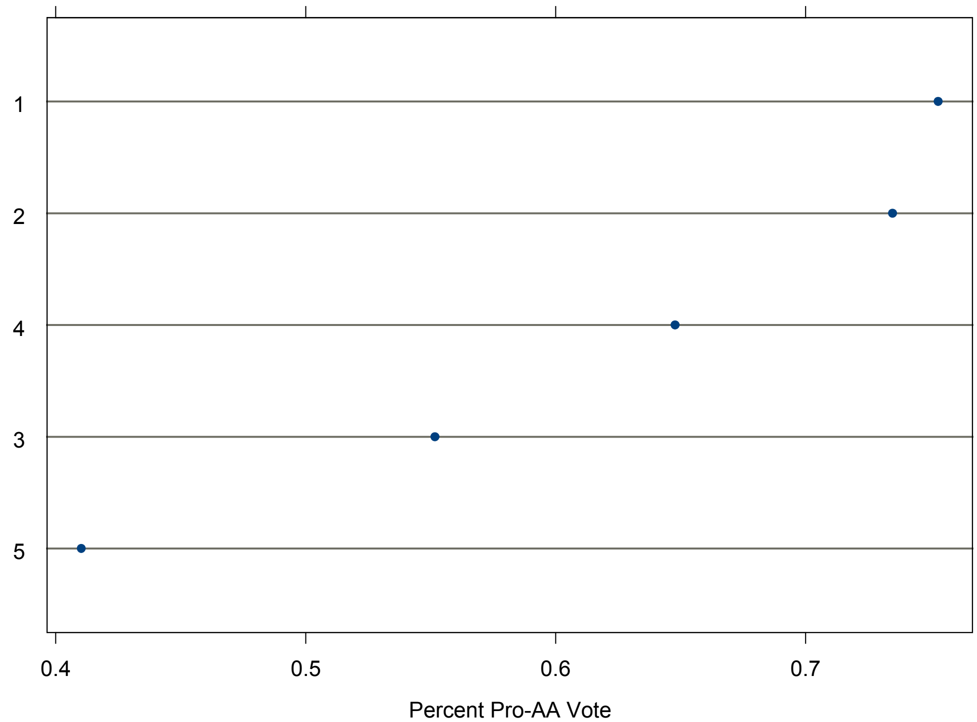


Figure 7. Pro-AA Voting by Era of SC Precedent



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