Peer Effects on the United States Supreme Court

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Abstract

Using data on essentially every US Supreme Court decision since 1946, we estimate a model of peer effects on the Court. We consider both the impact of justice ideology and justice votes on the votes of their peers. To identify these peer effects we use two instruments. The first is based on the composition of the Court, determined by which justices sit on which cases due to recusals or health reasons for not sitting. The second utilizes the fact that many justices previously sat on Federal Circuit Courts and are empirically much more likely to affirm decisions from their "home" court. We find large peer effects. Replacing a single justice with one who votes in a conservative direction 10 percentage points more frequently increases the probability that *each* other justice votes conservative by 1.63 percentage points. In terms of votes, a 10 percentage point increase in the probability that a single justice votes conservative leads to a 1.1 percentage increase in the probability that *each* other justice votes conservative. Finally, a single justice becoming 10% more likely to vote conservative increases the share of cases with a conservative outcome by 3.6 percentage points–excluding the direct effect of that justice–and reduces the share with a liberal outcome by 3.2 percentage points. In general, the indirect effect of a justice's vote on the outcome through the votes of their peers is typically several times larger than the direct mechanical effect of the justice's own vote.

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1 Introduction

Economists have long been interested in the impact of one's social, educational, and workplace environment–and the characteristics of other agents in that environment–on one's own behavior and outcomes.[1] The presence of positive spillovers, or *peer effects* in such settings would suggest a range of interesting policy interventions that could improve educational and labor-market outcomes.

Notwithstanding this, there are two formidable obstacles to identifying peer effects. The first is that the externalities created by peer effects should presumably be internalized by the market's price mechanism or, failing that, by firms, or even governments. Only when none of these three institutions internalize the externality can one hope to observe it in equilibrium outcomes.

The second obstacle is an econometric one. There is typically a mechanical link between the characteristics of individuals and those of their peer group. It is natural, then, to look at settings where there is random variation in the peer group.[2]

In this paper we sidestep these two obstacles by studying a unique laboratory for estimating peer effects: the Supreme Court of the United States. As we will discuss in detail below, both the structure of which justices sit on which cases, as well as the fact that many justices previously sat on Federal Circuit Courts of Appeals, [3] provides us with compelling instruments to identify peer effects.

In addition to this, the composition of the Supreme Court and the rulings it makes are of intrinsic interest, given their impact on legal outcomes. Furthermore, understanding the extent to which justices with a particular ideological standpoint can influence the votes of other justices is important for understanding the optimal strategy for an administration in nominating justices. This, in turn, speaks to the characteristics and design of legal institutions.

Setting aside the issue of nominations being successful, it easy to see how the ideological po-

[1]In the context of education, the concept of peer effects dates to at least the so-called "Coleman Report" (Coleman et al. (1966)).

[2]See also Manski (1993) and Manski (2000) regarding identification issues and, in particular, the *Reflection Problem.*

[3]Epstein et al. (2009) find strong evidence that federal judges are highly inclined to rule in favor of their respective home circuit court. We find the same effect and utilize it as an instrument.

sition of the optimal nominee will depend on the existence and magnitude of justice peer effects. First, note that justices will optimally employ a simple decision rule to maximize their utility; given two possible voting options they will vote for that which is closer to their ideal point (which is a function of their ideological preferences, the case characteristics,[4] and potentially the ideologies of other justices if peer effects exist). Since decisions depend on majority voting then, if there are no justice peer effects, the median justice will be pivotal, and a case outcome will reflect her position. It is thus tempting to think that the ideal appointment is one that shifts the median justice closest to the view of the President.

However if peer effects exist then voting decisions of justice j can be affected by the ideological position of justice i, and thus the Court's disposition will not merely be a function of the median justice's ideal point. This leads to a disjuncture between a justice's *ideological ideal point* and their *effective ideal point*, with the latter including the impact of peer effects. Where peer effects are a function of ideological positions, this means that the effective ideal point of justice j depends on the ideological positions of the other justices. This suggests that the President, in choosing a nominee, should consider her ability to affect the Court's rulings through her impact on other justices, as well as her own ideological position.

The approach we take to estimating peer effects on the Supreme Court is as follows. We first consider ideology as the channel through which peer effects operate. To do so we measure justice ideology by estimating a linear probability model of justice votes as a function of case characteristics and justice dummy variables in our model of voting behavior. We then add these peer ideologies as additional explanatory variables. Since, unlike some other courts, Supreme Court cases do not involve random assignment of justices, and because there is relatively slow turnover of justices, identifying peer effects is challenging. We tackle this challenge by observing that recusals and absences provide a plausibly exogenous source of peer variation on a given case. Using this approach, we find clear evidence of ideology-based peer effects. In particular, we find that replacing a single justice with one who votes conservative 10 percentage points more frequently on average increases the probability that *each* other justice votes in the conservative direction by 1.63 percentage points.

[4]Note that the presence of case characteristics means we are not taking a strictly legal realist position, but allowing for a mixture of judicial motives.

An alternative possible channel is for peer effects to operate through the votes of the justices, not ideology *per se*. Here, identifying a true peer effect requires exogenous variation in voting propensity across justices–i.e. a variable which directly affects how a given justice votes in a given case, but not the votes of other justices, except through the vote of the directly-affected justice. We utilize the fact that justices who have previously served on a Circuit Court of Appeals vote differently when a case comes from their "home" court, rather than another Circuit Court. This provides us with an instrument with the above mentioned properties. We find that a percentage point increase in the proportion of peers casting conservative votes in a case makes a justice 0.9 percentage points more likely to vote conservative. In the typical full bench (9 justices) case this implies a ten percentage point increase in the probability that a single justice votes in the conservative direction leads to a 1.1 percentage point increase in the probability that *each* other justice casts a conservative vote.

Finally, we examine whether the peer effects that we find actually change pivotal votes, and hence case outcomes, or if they merely affect the size of the majority. If peer effects merely push a decision from 6-3 to 5-4, or *vice versa*, then they are of limited practical interest.[5] We again utilize the home court instrument, except that variables are now aggregated at the case level, and we consider how a single justice's vote affects the collective voting behavior of their peers. We find strong evidence that peer effects can be pivotal. A single justice becoming 10% more likely to vote conservative increases the share of cases with a conservative outcome by 3.6 percentage points–excluding the direct or mechanical effect of that justice–and reduces the share with a liberal outcome by 3.2 percentage points.

To highlight the magnitude and importance of the effects we estimate, one can consider the impact of replacing the late justice Antonin Scalia with President Obama's nominee, Judge Merrick Garland (chief judge of the United States Court of Appeals for the District of Columbia Circuit.) Using "Judicial Common Space" (Epstein et al. (2007)) measures of ideology we find that the Supreme Court justice whose average score is closest to that of Judge Garland is Justice John Paul Stevens. Using our peer effect estimates we find that replacing Justice Scalia with Judge Garland

[5]Of course, the credibility of the Court, and how political it looks, is an important issue, and is plausibly affected by the size of the majority in a case. 5-4 decisions breaking along the lines of the party of the appointing President, for instance, may be seen as particularly political and this could be damaging to the image of the Court.

would make *each* other justice 5.1% more likely to vote liberal on a given case. On the other hand, the Supreme Court justice with estimated ideology closest to that of President Trump's nominee, Judge Neil Gorsuch (of the 10th Circuit), is Justice Scalia, so the analogous effect of appointing Judge Gorsuch would be trivially zero – a difference of 5.1%.

We are certainly not the first authors to consider the issues of judicial ideology and peer effects. There is a significant literature estimating the ideological position of judges and justices. For instance, Martin and Quinn (2002) develop a dynamic item response model and estimate justice ideal points that can be time-varying via Markov Chain Monte Carlo methods, and Martin et al. (2005) use the Martin-Quinn method to estimate the median Supreme Court justice on Courts dating from 1937. If one thinks that peer effects operate through the characteristics of judges, then understanding judicial ideology is a necessary first step to study them, and it is arguably of interest in its own right.

Perhaps closer to our paper is the literature on panel effects on lower courts. A large literature considers peer effects (often referred to as "panel effects") on U.S. Circuit Courts of Appeals.[6] Different authors emphasize different channels, such as: deliberation, group polarization, or aversion to dissent. Fischman (2015) argues that peer effects are best understood by reference to peers' votes rather than characteristics, and reanalyzes 11 earlier papers on Circuit Court "panel" voting, as well as new data. He finds that–across the board–each judge's vote increases the probability that a given judge votes in the same direction by approximately 40 percentage points. He replaces the characteristics of panel colleagues with their votes, so the votes are endogenous, but colleague characteristics can be used as an instrument for colleague votes, assuming that they have no direct causal effect. Boyd et al. (2010) considers the impact of female judges and, using Rubin (1974)'s "potential outcomes" approach, only finds strong effects for sex discrimination cases, suggesting an information channel is operative rather than alternative theories of influence.[7]

Finally, Epstein and Jacobi (2008) suggest that the power of the median justice is due to bargaining power, not personality. They claim that ideological remoteness of the median justice gives them a greater range of the ideological spectrum over which they are pivotal.

[6] For three notable examples, in addition to those mentioned below, see Revesz (1997), Miles and Sunstein (2006) and Posner (2008).

[7] See also Peresie (2005).

Relative to this large literature, we see our contribution as threefold. One, we focus on the United States Supreme Court rather than Federal Circuit Courts of Appeals. Two, we analyze a simple and intuitive voting model using a novel identification strategy for both the ideological channel and the vote channel. And three, we focus on both peer effects and their impact in altering case outcomes.

Once one is convinced that peer effects exist, the real question, of course, is what is driving them. As we mentioned above, in the context of lower courts, several possibilities have been raised, including: deliberation, group polarization, and aversion to dissent. We return to the question of what is driving the effects we find in this paper in our concluding remarks, where we also offer estimates of our effects by issue area.

The remainder of the paper is organized as follows. Section 2 outlines our estimation approach, and discusses the data we use. Section 3 contains our analysis of the ideological channel for peer effects, while Section 4 analyzes the voting channel. Section 5 focuses on case outcomes, rather than just the peer effects themselves, and Section 6 contains some concluding remarks.

2 Model and Data

2.1 Framework

A natural approach to modeling voting decisions is to estimate a random utility model. Let j denote justice, c denote case and t denote year. The ideological direction of the vote by each justice present in each case, d_{jct}, is either conservative (1) or liberal (0).[8] Justices choose the option that maximizes their utility. Define u_{jct} as the net utility that a justice derives from voting conservative

[8] Note that cases can occur where the context of the case is distant from the ideological middle ground, such that justices may face a choice between a highly conservative (liberal) position and a mildly conservative (liberal) position. The theoretical framework provided by the random utility model merely requires that the median justice can be determined as being closer to one of the voting options; their ideological ideal point need not be situated between the two alternatives.

rather than liberal. Then,

$$d_{jct} = \begin{cases} 1 & \text{if } u_{jct} \geq 0 \\ 0 & \text{otherwise} \end{cases} \tag{1}$$

We consider two different mechanisms through which peer effects may exist. In the first model presented below, peer effects work directly through ideological positions, with the preferences of justice j directly influenced by the ideological positions of the other justices $\{i\}\backslash j$. In the terminology of Manski, this is a contextual peer effect since justice ideology is predetermined with respect to their interactions with other justices. Under this mechanism, the voting decisions of justice i gravitate to (or are repeled from) positions consistent with the ideology held by other justices, without considering how those other justices actually vote in the same case. This peer mechanism, if it exists, implies that justices affect the underlying ideological disposition of each other and hence affect votes by this means.

An alternative mechanism is that, rather than fundamentally shifting ideology for all cases, the effects of peers on the their colleagues operate through their own votes, jointly affecting their respective votes on a case-by-case basis. Since outcomes of justices and their peers are jointly determined, this fits within the framework of Manski's endogenous peer effects (Manski (1993)). If peer effects operate via an effect of the vote of each justice on the votes of their colleagues, this does not preclude there from being an effect of peers on ideology. However it does imply that peers affect ideological preferences of other justices only when they vote in a manner consistent with their established ideology.

Peer effects could operate through either or both mechanisms. Indeed, the first mechanism, where peers effect ideological positions, may merely be a reduced form for the second, where peer effects operate through the voting decisions of a justice's peers, and the probability of those vote decisions is in large part driven by peer ideological positions. Alternately, these channels need not be identical, as it is possible that the ideology of peers continues to have an effect on voting decisions independently from how a justice's peers vote in a given case.

2.2 Data

We use data from the Supreme Court Database.[9] This database contains a wide range of information for almost the entire universe of cases decided by the Court between 1946 and 2013.[10] The data provides a rich array of information for each case, including the case participants, the legal issue area the case pertains to, the court term in which the case was heard and opinions were issued, and further identifies the winning party and overall vote margin. Particularly relevant for the analysis in this paper, the data includes the identity and voting decision of each justice, for each case in which they were involved, such that decisions of individual justices, and their relationship with the identity and voting decisions of the peer justices, can be analyzed. For almost all cases, votes are identified according to their ideological disposition, categorized as either liberal or conservative, with codification following an explicit set of rules, with the exceptions being for cases without any clear ideological underpinning, or occasions where a justice recuses themselves from voting. Finally, it also contains identifying data including case and vote identification numbers, and citation numbers used in official reports.

These data are augmented with additional information on each justice from the U.S. Supreme Court Justices Database developed by Epstein et al.[11] In particular, this provides information on which, if any, Circuit Court of Appeals a justice previously served on, and the length of their tenure on that court. This turns out to be useful as justices sometimes hear cases that come from a court they previously worked on, and thus this data allows any home bias towards their affiliated court to be accounted for.

2.3 Descriptive Statistics

In its entirety, the data provide information about 116 362 votes (including recusals) from 12 981 cases. Restricting attention to the relevant subset of votes used in this paper (excluding recusals and votes issued in cases without any discerned ideological direction), the data contains 110 729

[9] `http://supremecourtdatabase.org/documentation.php?`

[10] For example, non-orally argued cases with *per curiam* decisions are not included unless the Court provided a summary, or one of the justices wrote an opinion.

[11] `http://epstein.wustl.edu/research/justicesdata.html`

votes with identified ideological direction[12] from 12 779 cases, three quarters of which involve a vote by all nine serving justices. Considering directional votes, the distribution of votes by ideological direction is closely balanced, with 48% being issued in the conservative direction. In contrast, the majority (55%) of lower court decisions in cases reviewed by the Supreme Court are in the conservative direction.[13] This reversal is symptomatic of a strong tendency towards overturning lower court decisions; in the dataset 58% of votes made by justices and 60% of Supreme Court opinions are in the reverse direction to the source court's decision. This tendency towards overturning is a natural consequence of the Supreme Court's operations; since it reviews only a small fraction of cases and chooses which cases to hear, there is a natural tendency towards selecting to hear cases in which a preponderance of justices believe (it is likely that) an incorrect decision had been made by the relevant lower court.

Table 1 breaks down these aggregate proportions across several stratifications of the data. Of the 11 high-level legal-issue-area categories in the database with a nontrivial number of votes in our sample,[14] the distributions of vote ideology over the entire 1946-2013 range of court terms vary from 29% conservative for Federal Taxation cases to 60% conservative in Privacy cases. Separating instead by the Circuit Court of Appeals that previously heard the case (for the $\sim$60% of cases that source from such a court) the conservative share of votes ranges from 43% for cases from the Seventh Circuit to 54% for Ninth Circuit cases.[15] There is a larger degree of variation in vote ideology proportions across justices, with conservative vote share ranging from 22% for William O. Douglas to 72% for Clarence Thomas (see Table 11 in Appendix A for details), while Figure 1 further illustrates how the conservative vote share has varied over time.

[12] A small number of cases result in tied votes, following which the votes of individual justices are typically not made public. Provided that the case had a lower court decision with stated ideological direction, so that the case is known to have ideological relevance, the vote direction for each justice is coded as 0.5 by convention.

[13] There are a small number of cases with directional Supreme Court votes but unspecified lower court vote direction. This accounts for 1% of directional Supreme Court votes.

[14] There are another 4 issue area categories which collectively make up less than 0.1% of the sample, for 15 issue area categories in the entire database.

[15] The Ninth Circuit is often considered as being strongly liberal, which recalling the Supreme Court's endogenous case selection and its overall tendency towards overturning the decisions it reviews, is consistent with this high conservative vote share.

Table 1 – Descriptive Statistics for Directional Votes

	Votes	Cases	Vote Direction *(Cons. %)*	Lower Court *(Cons. %)*	Overturn *(%)*
Total	110,729	12,779	47.58	55.03	58.18
Legal Issue Area					
Criminal Procedure	22,549	2,585	52.12	63.07	60.23
Civil Rights	18,435	2,112	44.87	53.47	58.71
First Amendment	9,895	1,140	45.92	56.66	56.25
Due Process	4,975	577	42.57	53.65	59.84
Privacy	1,483	169	60.35	30.21	57.38
Attorneys	1,122	130	43.23	52.05	60.34
Unions	4,387	506	45.25	57.53	55.87
Economic Activity	21,447	2,500	42.28	48.82	57.20
Judicial Power	17,041	1,976	58.32	54.18	58.33
Federalism	5,805	670	43.65	56.66	58.23
Federal Taxation	3,415	394	29.49	56.78	52.71
Circuit Court					
Federal	937	107	46.21	43.00	62.82
First	2,125	246	47.01	40.82	51.40
Second	8,107	934	48.35	50.70	54.85
Third	5,008	575	51.54	49.84	54.21
Fourth	4,471	512	45.96	60.88	55.40
Fifth	7,907	914	43.49	65.12	60.88
Sixth	5,558	644	47.59	50.55	60.17
Seventh	5,523	645	42.97	59.07	58.63
Eighth	4,046	465	45.30	48.60	57.94
Ninth	11,835	1,359	54.30	38.27	62.80
Tenth	3,153	367	51.03	51.22	60.01
Eleventh	2,203	247	44.80	67.68	57.10
D.C.	6,961	818	52.15	51.13	59.46

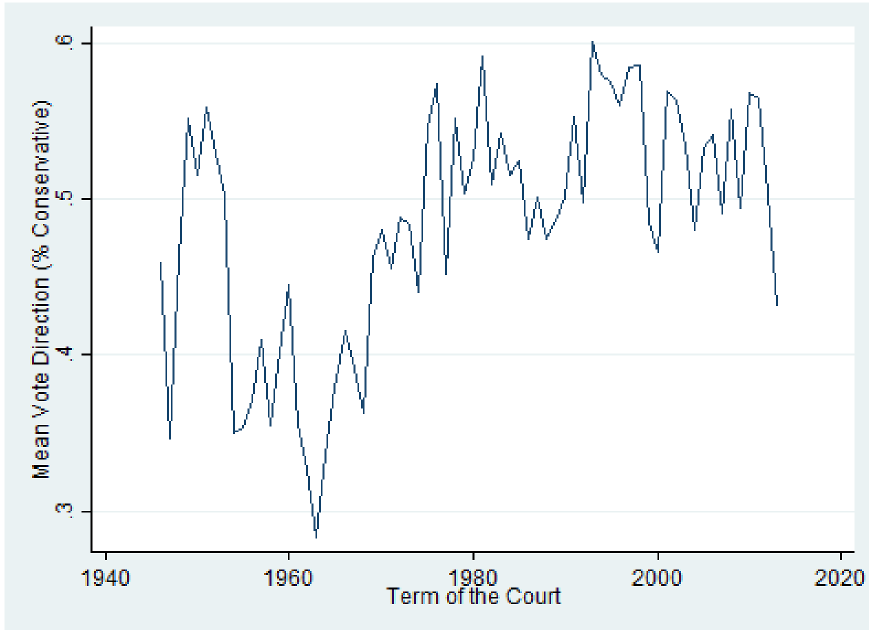


Figure 1 – Evolution of Conservative Vote Share by Term

3 Peer Ideology Effects

In order to estimate the effect of peer ideology on the voting decisions of a justice, a two-step procedure is utilized. This is motivated by the need to first generate estimates of justice ideology. These individual ideology measures are then combined in order to construct measures of peer ideology. Finally, the peer effect estimation can be undertaken.

More specifically, the first step involves estimation of a linear probability model[16] of justice votes as a function of a set of case characteristics along with dummy variables for justices. The dummy coefficient for each justice provides an estimate of the respective justice's ideal point in the ideological spectrum. By virtue of the linear probability model framework, the estimated justice coefficients are strictly interpretable as the fraction of cases (in the appropriate excluded dummy categories) in which the respective justice will make a conservative (rather than liberal) vote.[17] These justice coefficients can then be extracted and used to create proxies for peer ideology, including but not limited to the mean ideological position of contemporaneous peers.

[16]The panel data structure with a predominance of dummy variables in the estimated model favors OLS estimation.

[17]To abstract away from the potentially unclear 'excluded categories' note that differences between justice coefficients reflect the difference in the proportion of cases in which justices issued conservative votes.

In the second step, these peer measures are added as an additional explanatory variable to the first-stage regressions. Nonzero coefficients on peer ideology indicate the presence of peer effects (rejecting the null hypothesis of the absence of peer effects). We estimate several specifications with different sets of controls for case and justice characteristics. In order to prevent peer variables from containing information about the case not present[18] in the covariates, for each specification the peer variables utilized in the second stage are those constructed from the analogous first-stage regression (that is, with the same set of covariates in both stages). Concluding that this two-step procedure yields unbiased estimates of peer ideology effects presents several econometric challenges, which are discussed in detail below.

3.1 Empirical Specification

In the baseline model, the hypothesized utility function (also interpretable as the probability that a justice will issue a conservative vote) is of the form

$$\begin{aligned} u_{jct} =& \alpha_j + \gamma_c + l_c + lc\_dec_c \times \beta_1 + I\left[j \in app_c\right] \times \left[\beta_2 + \beta_3 \times app\_tenure_j\right] \\ &+ lc\_dec_c \times I\left[j \in app_c\right] \times \left[\beta_4 + \beta_5 \times app\_tenure_j\right] + \varepsilon_{jct} \end{aligned} \tag{2}$$

where α_j is a justice fixed effect, γ_c is a fixed effect for the Circuit Court of Appeals (if any) that previously heard the case, l_c is a fixed effect for the legal issue area the case pertains to, lc_dec_c is the ideological direction of the decision made by the lower court, which the Supreme Court is reviewing. Further, $I\left[j \in app_c\right]$ is an indicator for whether the case sourced from a Circuit Court of Appeals for which the justice previously served, and app_tenure_j is the number of years that the justice previously served on a Circuit Court of Appeals (if any). These latter two variables are interacted with the decision of the lower court.

Subsequent specifications add further precision to the model. The second specification adds fixed effects for court term δ_t to control for systematic drift in ideology of the Court over time.[19] Since justices may conceivably have differing ideological preferences across different issue areas

[18]This is problematic in any particular case, which is why we subsequently use an instrumental variables approach.

[19]Since there is no anchor on, or exact measure of, the ideology of cases heard over time, term dummies account for systematic changes in justice ideology net of changes in the ideological composition of cases heard.

(that is, a single ideological dimension may not fully characterize justice ideological preferences) a third specification incorporates justice by issue area fixed effects α_j^l (replacing α_j and l_c). A fourth specification further adds issue area by term fixed effects δ_t^l to account for any differential systematic (across justices) ideological drift by issue area (replacing δ_t). The precise rationale for these specifications, in terms of the exogenous variation in peer ideology that they capture to identify peer effects, is discussed in detail in Section 3.3 below.

3.2 First Stage Results

The four specifications of the linear probability model outlined in Section 3.1 are estimated by OLS. Standard errors are clustered by case to account for unobserved case characteristics providing a common within-case shock to the votes of all justices. Given the purpose of extracting proxies for ideology, it is desirable that the specifications yield stable ideology measures. Table 2 shows the correlations between different measures, weighting equally by directional votes. The correlations vary from 0.86 to 0.99, and are particularly high when considered separately by issue area (Models 3 and 4). Further, the potential empirical relevance of any peer ideology influences is inherently restricted by the influence of own ideology on voting decisions. If votes are not substantially driven by ideology, peer effects based on the transmission of ideology are unlikely to have meaningful effects. However, the model estimates shown in Table 3 demonstrate that justice ideology is an extremely important determinant of votes; in each specification the justice dummy variables have substantial explanatory power over vote direction after controlling for all other covariates, with marginal contributions to model R^2 of between 0.0805 and 0.1111.

Table 2 – Ideology Measures Correlation Matrix

	Model 1	Model 2	Model 3	Model 4
Model 1	1.0000			
Model 2	0.9859	1.0000		
Model 3	0.8611	0.8709	1.0000	
Model 4	0.8614	0.8716	0.9854	1.0000

For Models 3 and 4 where justice ideology differs by issue area, ideology scores are normalised by issue area to remove level differences between models.

While most of the model coefficients are not of particular interest, several interesting results are worth a brief discussion. First, the coefficients for a conservative (liberal) lower court opinion (compared to the omitted category of an indeterminate lower court ideological direction) being negative (positive) reflect the tendency of the Supreme Court to overturn many decisions that it reviews. Second, a consistent pattern of *home court bias* is evident. Previous service on a Circuit Court of Appeals (a justice's *home court*) affects how a justice votes when hearing a case sourced from that court (i.e., when they are *at home*). Justices who had previously served on a Circuit Court of Appeals are less likely to overturn the opinion of a *home court case* (the interaction coefficient for a vote of a home justice hearing a case with a conservative (liberal) lower court decision is positive (negative)). However this bias diminishes with home court tenure. The interaction coefficients of the lower court opinion in home court cases with length of a justice's tenure on the home court operate in the reverse direction. Justices with long Circuit Court tenures are instead more likely to overturn lower court decisions when hearing a case sourced from their home court.

Table 3 – First Stage Results - Justice Vote Direction (Conservative %)

	(1) Vote Direction	(2) Vote Direction	(3) Vote Direction	(4) Vote Direction
Conservative LC	-0.030	-0.061	-0.074*	-0.089
	(0.029)	(0.041)	(0.043)	(0.059)
Liberal LC	0.116***	0.085**	0.070	0.054
	(0.029)	(0.041)	(0.043)	(0.059)
Justice Home Court	0.110***	0.108***	0.121***	0.123***
× Conservative LC	(0.032)	(0.031)	(0.029)	(0.028)
Justice Home Court	-0.155***	-0.157***	-0.138***	-0.139***
× Liberal LC	(0.032)	(0.032)	(0.031)	(0.031)
Justice Home Court Tenure	-0.012***	-0.012***	-0.012***	-0.012***
× Conservative LC	(0.004)	(0.004)	(0.003)	(0.003)
Justice Home Court Tenure	0.015***	0.016***	0.014***	0.014***
× Liberal LC	(0.003)	(0.003)	(0.003)	(0.003)
Circuit Court FE	Yes	Yes	Yes	Yes
Justice FE	Yes	Yes	No	No
Issue Area FE	Yes	Yes	No	No
Term FE	No	Yes	Yes	No
Justice x Issue Area FE	No	No	Yes	Yes
Term x Issue Area FE	No	No	No	Yes
R-squared	0.1370	0.1446	0.1753	0.2101
Δ R-squared	0.0894	0.0805	0.1111	0.1071
Observations	110729	110729	110729	110729

Δ R-squared is the marginal explanatory power of justice ideology on vote direction, measured as the increase in model R-squared collectively due to the justice fixed effects (or justice by issue area fixed effects). * p<0.10, ** p<0.05, *** p<0.01

3.3 Second Stage Results

Ideally, estimating the effect of the average ideology of a justice's peers would involve adding a variable $\overline{a^l_{\text{-}j}}$ measuring the average peer ideology to the specification in Equation 2, yielding

$$\begin{aligned} u_{jct} =& \alpha_j + \gamma_c + l_c + \beta_p \times \overline{a^l_{\text{-}j}} + lc\_dec_c \times \beta_1 + I\left[j \in app_c\right] \times \left[\beta_2 + \beta_3 \times app\_tenure_j\right] \\ &+ lc\_dec_c \times I\left[j \in app_c\right] \times \left[\beta_4 + \times \beta_5 \times app\_tenure_j\right] + \varepsilon_{jct} \end{aligned} \tag{3}$$

However since justice ideology is unobservable, the peer variable that we actually utilize is the proxy $\widehat{\overline{a^l_{\text{-}j}}}$ constructed as the average fixed effect (i.e. ideological position) of the concurrently serving justices, using the extracted first stage coefficients. This enables the model to be estimated, with the estimate of β_p, which measures strength of peer effects, being of particular interest. A positive coefficient indicates that judges are pulled towards the ideological position of their peers.

One difficulty in identifying peer effects in a context such as the Supreme Court is that there is very little panel rotation. For example, unlike other courts, cases do not involve random selection of a subset of justices, and further, the cohort of justices evolves only slowly over time. Intuitively, these features complicate the task of separating peer effects from joint ideological drift of justices over time.

However, while cases before the Supreme Court are generally heard by the full panel of justices, justice recusals provide a natural source of exogenous variation in the peers voting on a given case. In fact, as noted in Section 2.3, at least one justice is absent due to a recusal (or other factor such as illness) in roughly 1/4 of all cases. This variation in Court composition is particularly useful in that it allows the effect of peers to be considered both when they are active (voting on a case) and absent (recused). Intuitively, any peer effect that a justice may have should be attenuated or eliminated entirely when a justice does not vote on or otherwise participate in a case (i.e. if recused, it would be considered improper for them to discuss the case with the other justices).[20]

[20]Note that in addition to the mechanism considered, where a justice's ideology affects their peers while they are present on the court, justice ideology may also have permanent effects on peers by influencing the peers' viewpoint or manner of thinking in an enduring manner. This will not be identified by these tests, as it will largely be soaked up by the justice and time FE. Thus our method at best captures only some of the channels through which peer effects may operate.

To take advantage of this, for each of the four first-stage model specifications, three peer variables are created as the average ideology of (1) all other peers, (2) other justices active in a case, and (3) the justices absent from a case (set to zero if no justices are absent). Equation (3) is estimated using each of these peer measures in turn, with a further specification jointly testing the effect of active and absent peer ideology. Since most cases involve no absent justices, the specifications containing this variable also include a dummy indicating whether any justices are absent.

To properly identify peer effects, these regressions require the implicit assumption that the residual variation in peer ideology induced by recusals is exogenous with respect to unobserved case characteristics. These estimates would be biased if the fact that a justice with particular ideology was recused provided information about the ideological tendency of the case. For example, if justices are more likely to recuse themselves when they would counterfactually either be in the minority or vote opposite to their general disposition, the court will contemporaneously issue disproportionately conservative (liberal) votes when endogenous recusals make the composition of peers more conservative (liberal). Such a phenomena would create the appearance of peer effects even if they do not exist. The reverse, and equally problematic bias, would occur if recusals are more frequent when in the majority. Given these threats to identification, the absent peer regressions operate as placebo tests to detect the presence of endogenous recusal bias. If the ideology of recused justices provides information about unobserved case characteristics, then the regressions using the ideology of absent justices as the relative peer measure should find this variable to have strong explanatory power.[21] Furthermore, if peer effects do not truly exist, the ideology of active peers should have no effect once controlling for ideology of absent justices. Hence by comparing the coefficients on the different peer measures, the appropriateness of using recusal based variation in peer ideology to isolate peer effects can be established.

With this strategy in mind, the results of these estimations for each of the four first stage models are shown in Table 4. The results for the first model, where the peer measure is based on justice fixed effects, and there are no term dummies, are shown in the first panel of Table 4. The first column reports results using the mean ideology of all peers to measure the peer effects.

[21] This magnification reflects the fact that the Court exhibits a strong degree of agreement in decisions, for example 37% of cases in the sample involve a unanimous decision. Accordingly if recusals on average provide even a small amount of information about a justice's counterfactual vote, substantial information is conveyed about the overall vote of remaining justices.

Table 4 – Peer Ideology Second Stage Results - Justice Vote Direction (Conservative %)

	Model 1: Justice FE			
	(1)	(2)	(3)	(4)
	Vote Direction	Vote Direction	Vote Direction	Vote Direction
Mean All Peer Justices	-0.012			
	(0.106)			
Mean Active Peer Justices		0.143		0.105
		(0.099)		(0.102)
Mean Absent Peer Justices			-0.199**	-0.191**
			(0.083)	(0.084)
R-squared	0.5487	0.5488	0.5491	0.5491
Observations	110729	110729	110729	110729

	Model 2: Justice, Term FE			
	(1)	(2)	(3)	(4)
	Vote Direction	Vote Direction	Vote Direction	Vote Direction
Mean All Peer Justices	-0.789			
	(0.968)			
Mean Active Peer Justices		1.311***		1.468***
		(0.371)		(0.511)
Mean Absent Peer Justices			-0.162*	0.038
			(0.085)	(0.120)
R-squared	0.5527	0.5531	0.5529	0.5531
Observations	110729	110729	110729	110729

	Model 3: Justice by Issue Area, Term FE			
	(1)	(2)	(3)	(4)
	Vote Direction	Vote Direction	Vote Direction	Vote Direction
Mean All Peer Justices	0.390**			
	(0.154)			
Mean Active Peer Justices		0.562***		0.583***
		(0.129)		(0.138)
Mean Absent Peer Justices			-0.029	0.027
			(0.068)	(0.070)
R-squared	0.5689	0.5691	0.5687	0.5691
Observations	110729	110729	110729	110729

	Model 4: Justice by Issue Area, Term by Issue Area FE			
	(1)	(2)	(3)	(4)
	Vote Direction	Vote Direction	Vote Direction	Vote Direction
Mean All Peer Justices	0.030			
	(0.800)			
Mean Active Peer Justices		1.245***		1.838***
		(0.275)		(0.305)
Mean Absent Peer Justices			0.015	0.157***
			(0.046)	(0.051)
R-squared	0.5869	0.5875	0.5870	0.5877
Observations	110729	110729	110729	110729

Models estimated with associated set of covariates used in analogous first stage regression, see Table 3. Peer variables are constructed using the analogous first stage justice coefficients estimates. * $p<0.10$, ** $p<0.05$, *** $p<0.01$

Here the primary identification comes from variation in a justice's cohort over time as their peer justices retire (or die) and are replaced on the court by new appointees. This yields a small and insignificant peer effect estimate. The second column reports results using our preferred active peers measure, which also utilizes variation in justices voting on each case. This yields a small but insignificant peer effect (of a justice's mean peer) of 0.143. In the third column, the placebo measure of absent peers yields a negative estimated coefficient of -0.199, which is likely due to the Court's average ideology being relatively stable across time, such that the ideologies of absent and active justices tend to be negatively correlated.[22] Including both absent and active peer measures jointly yields similar coefficients, although only the absent justice measure is significant.[23]

Since this model lacks controls for time (such as term fixed effects), changes in the ideological composition of the Supreme Court as justices are replaced are not well distinguished from joint ideological drift of justices over time. In the presence of exogenous ideological drift (due to changing norms, beliefs, and preferences of society), new justice appointments will have voting records (and thus estimated ideology) that tend to reflect this drift.[24] Thus the average peer ideology measures will tend to comove with ideological drift and voting propensities,[25] and thus produce positive estimates of peer effects even when they do not exist.

The results for the second model, which attempts to alleviate this concern through the addition of term fixed effects, are shown in the second panel of Table 4. Since the all peer measure is based off justice fixed effects, for a given justice it is constant for all cases in a year, except due to infrequent cohort changes arising from mid-year appointments. While changes in the cohort of justices produces variation in a justice's ideology relative to their peers over time, it does so in a common way for all continuing justices.[26] Accordingly the all peer measure is close to collinear

[22]This argument is particularly pertinent in models with term fixed effects; there is little variation in Court ideology within term, so this negative correlation is stronger.

[23]Nor is it surprising, here, that the absent measure stays significantly negative when controlling for the active measure. Since the active measure captures variation in peer ideology both within term through recusals and across term through appointment of new justices, the absent justice measure which captures only the variation based on recusals is thus a tighter proxy for this variation in ideology. This should largely be corrected once controlling for term.

[24]Even if ideological drift does not coexist with political circumstances that cause a justice with ideology particularly in the given direction to be appointed, it remains that a new appointee of average ideology (of that period) will on average have a voting record favoring that ideological direction, as they are similarly impacted by the ideological drift of the era.

[25]This reveals an important nuance; if ideological drift adjusts the ideological composition of cases the Supreme Court considers by an equal amount, the net effect on voting propensities is zero. Hence the relevant consideration is ideological drift of justices net of changes in the ideological composition of cases the Supreme Court hears.

[26]Since in constructing a mean ideology of other justices, each involves replacing the retiring justice's ideology esti-

with the combination of term and justice fixed effects, which yields the very noisy coefficient estimate shown in Column 1.

In contrast, identification of peer effects using the active peers measure comes from within-term variation, due to recusals, in the panel of justices hearing a particular case. This yields a substantial and tightly estimated peer effect coefficient of 1.311. This implies, for example, that replacing a justice with another who votes in the conservative direction 10 percentage points more frequently on average would increase the conservative vote probability of all other justices by 1.64 percentage points, generating a cumulative 0.13 extra conservative votes by the peer justices per case (i.e., $0.0164 \times 8 = 0.13$). As with the first model, the absent peers measure yields a small negative estimate, which disappears when jointly including the active and absent peer measures (again reflecting the negative correlation between these variables), while the active peers measure increases slightly.

The third panel of Table 4 show results from the third model which utilizes a richer specification where justice ideology is allowed to vary by legal issue area. Since the term fixed effects are common across issue areas, this allows the peer variables to gain identification through differential variation in the ideology of peers by issue areas over time when justices are replaced by new appointees (since the common component of issue-area specific changes is differenced out by the term dummies). An alternate framing is that changes in the cohort of justices produces variation in the ideology of peers, and while this is common amongst continuing peers, it nonetheless differs by issue area. Using this richer model of ideology, the all peers measure yields an estimated peer effect coefficient of 0.390, while the active peers measure which gains additional identification from recusal-driven variation in peers gives an estimate of 0.562. For the thought experiment of replacing a single justice with another who votes in the conservative direction 10 percentage points more frequently, the latter estimate implies an increase of 0.7 percentage points in conservative vote probability (and thus $0.007 \times 8 = 0.056$ additional conservative votes per case). Further, the placebo absent peers specification yields a tightly estimated insignificant coefficient, and the results vary little when the absent and active peer coefficients are jointly estimated.

Since Model 3 incorporates justice ideology (and thus peer measures) that differ by issue area,

mate with the new justice's score.

but only a single set of controls for term, it is vulnerable to the criticism that peer effects identified off changes in Court composition are not well distinguished from issue-area-specific ideological drift over time. Similar to the argument above, given exogenous ideological drift specific to an issue area, new justice appointments will on average have voting records and thus estimated ideology that captures this drift. Thus for issue areas where idiosyncratic (i.e. issue specific) ideological drift is pertinent, average peer ideology measures for cases of that issue area will tend to co-move with ideological drift and voting propensities, upwardly biasing the peer effects estimates.

The results for the fourth model, which controls for this differential ideological drift through incorporation of term fixed effects by issue area, are displayed in the final panel of 4. Analogously to the second model, the term by issue area dummies soak up almost all variation in the all peers measure, such that the associated coefficient is imprecisely estimated. However the active peer measure, which is identified through within-year-and-issue-area variation in ideology of a justice's voting peers across cases due to recusals, yields a positive and significant peer effects coefficient of 1.245. By contrast, the placebo measure of absent justices yields a precisely estimated statistically zero coefficient. These results change slightly under joint estimation of the effect of active and absent peers; the estimated effect of active peers is nontrivially higher at 1.838 while the coefficient on absent peers is rendered significant albeit relatively small. It is unclear whether this final result is indicative of a statistical artifact or captures a real but relatively small peer effect of justices even when not voting on a case. However, recall the absent peer specification is only partially a placebo test, and may still capture some true peer effects.

3.4 Endogenous Justice Ideology

While these results are collectively strongly indicative of substantial positive peer ideology effects, there are several notable issues with the estimation procedure. Most notable is that the justice fixed effects from the first stage are used to construct the peer ideology measure utilised in the second stage. However, if peer effects are present, then the first stage is misspecified. As a result, each justice's own ideology measure will be contaminated by her peers' ideology, which in turn means that the peer ideology measures that we construct will be contaminated by a justice's own ideology (see Appendix C for a detailed derivation). However, as shown in Appendix C, when we do fixed-effects estimation in the second stage, the justice j-specific effect that potentially "contaminates" the peer measure washes out. This is because such contamination is invariant across observations for a given justice. Nevertheless, the measurement error in the peer ideology measure is shown to generate an attenuation bias. This implies that our findings regarding the magnitude of peer effects are conservative.

A second issue is that the ideology estimates constructed from the first stage estimates are based on each justice's full voting record, rather than being limited to their previous votes. This is a practical approach, as the larger a voting history the ideology variables are based upon, the less noisy a proxy it should be, reducing attenuation bias caused by measurement error. This means the ideology estimates are not predetermined in a temporal sense. However, to the extent that future votes reflect a predetermined ideological propensity, this is not an issue, but a failure of strict exogeneity will arise if there is ideological drift that is in part due to past cases and decisions.[27]

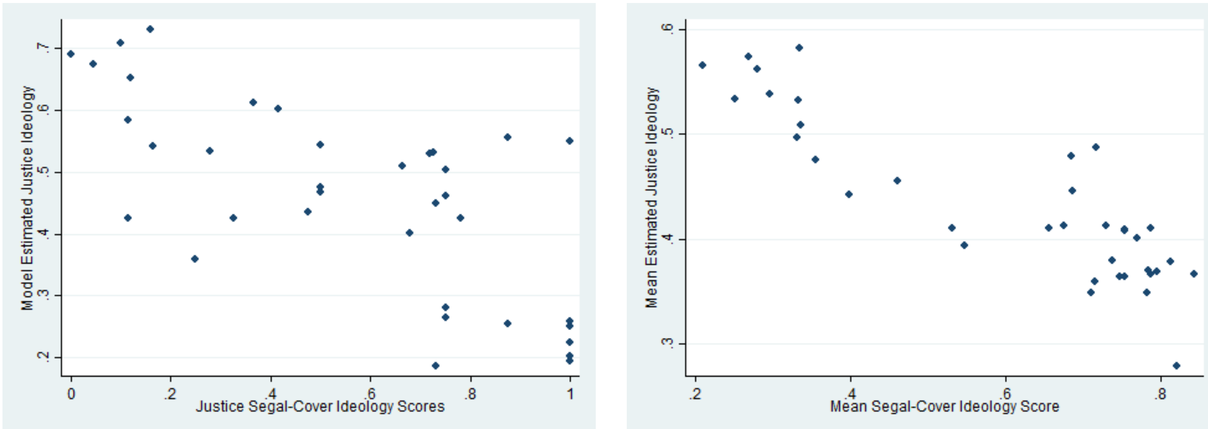
Given these potential problems, the obvious approach is to instrument for the peer effect variable using a predetermined (to Supreme Court tenure, and thus voting behavior) measure of justice ideological preferences. Segal-Cover scores (Segal and Cover (1989)), calculate estimates of justice ideology based on textual analysis of newspaper editorials between nomination by the President and the Senate confirmation vote, thus predating any of the justice's Supreme Court votes.[28,29] While Segal-Cover scores are again at best a noisy proxy of true justice ideology, since they are based on pre-Court tenure observables the error they contain should be substantially independent of the mismeasurement error in the constructed ideology estimates.

Accordingly the peer effect regressions in Table 4 are re-estimated, using the mean Segal-

[27]Note also that, in finite samples, individual votes have a non-vanishing effect on the justice ideology estimates. Unobserved characteristics of the contemporaneous case thus affect the justice coefficients in the first stage, causing the peer measures to be positively correlated with unobserved case characteristics in the second stage. While this effect is very slight if a justice is observed to vote on many cases, it nonetheless produces upwardly biased coefficients for the all and active peer ideology measures.

[28]Formally, the coding from editorial text to ideology score was undertaken much later when Segal and Cover developed these scores, and the coding process involves some subjectivity (it does not, for example, follow a simple decision rule). However the scores remain plausibly exogenous to subsequent voting behavior of justices.

[29]Three of the justices in the sample sat on the court for several months as recess appointments before being nominated and confirmed by the US Senate through normal procedures, so their Segal-Cover scores, which stem from this later nomination, are not truly predetermined to all their votes. However the scores still predate the vast majority of their votes (98-99%), and the results are robust to adjusting the recess votes.



(a) Ideology by justice

(b) Ideology by mean peer

Figure 2 – Relationship between Segal-Cover ideology estimates and Model (2) ideology estimates

Cover score of justice peers (all others, active peers, and absent peers in turn as appropriate) as an instrument for their true ideology. Usefully, as demonstrated in Figure 2, Segal-Cover scores are strongly correlated with the model estimated justice ideology scores, with an even tighter relationship between the mean Segal-Cover score and ideology estimate of peers (since averaging over multiple justices reduces noise).[30] Using Segal-Cover scores as an instrument involves the identifying assumption that the pre-Court tenure perceived ideology of justices only affects how their peers vote through their own true ideology (note that this is much more credible in specifications with time-based controls for ideological drift). For the first two models where justice ideology is common across all issue areas, it is sufficient to use a single Segal-Cover score variable as the instrument. However, in the specifications with justice ideology differing by legal issue area, Segal-Cover scores interacted with issue area dummies are used as instruments to capture variation in the slope (and intercept) of the relationship between overall perceived ideology and observed voting propensity by issue area (without this, the first stage fitted peer ideology measures would not capture any differences between issue areas).

Results for these estimations are shown in Table 5. These results are generally consistent with the OLS estimates shown before. Peer effects are consistently found to be positive and of meaningful magnitude, in particular for the active peer measures where identification comes from changes in Court ideology due to recusals. The results are generally consistent with what we

[30]Note these correlations are negative, because Segal-Cover scores are coded on a spectrum of 0 (conservative) to 1 (liberal), the reverse orientation to the voting propensity measure used in this paper.

found above–peer effects are positive and substantial– albeit the point estimates are slightly lower and less precise. This suggests that on net, the bias introduced by simultaneous determination of the ideology variable with peer effects, and more general measurement error in the ideology variable (potential attenuation bias) are relatively small. The placebo specifications testing peer effects of absent justices again find effects relatively close to zero, largely statistically insignificant, and of unstable sign. While some specifications find negative peer effects, these primarily involve the all peers measure, where peer effects are less convincingly identified, the exclusion restriction is less plausible, and point estimates are very noisy.

3.5 Case Selection Bias

Since many characteristics of individual cases are not observed, an implicit assumption underlying the analysis is that these unobserved characteristics do not systematically vary with the ideology of the Court as the cohort of justices changes over time. As noted above, this is particularly pertinent because case characteristics have an overwhelming influence on individual votes; in fact, in the full dataset of directional votes 37% of cases yield unanimous opinions.

Since justices select which cases the Supreme Court will hear, one important potential source of bias is that the characteristics of cases chosen will depend on justice ideology, due to an underlying strategic objective. For example, a natural strategic aim of a majority coalition of justices with similar ideology is to enshrine their own preferences in precedent (or move precedent in their preferred direction). Winning cases thus becomes an instrumental goal. The appointment of a new justice that shifts the majority balance to some coalition may make them more willing to take on cases that are more ideological (in their favored direction) and thus offer a greater prospect of setting important precedent. By definition, these more ideological cases are harder than usual for such a grouping to win compared to the null set of cases they could instead hear (otherwise an earlier less-powerful coalition would have already caused the case to be heard). This occurs because the more ideological (in the favored direction) the case is, the greater likelihood that given justices will vote in the opposite direction.[31] If this endogenous case selection does exist, the re-

[31] Implicit in this idea is that if a majority wins all cases by too large a margin, they could have chosen harder targets and still been successful.

Table 5 – Peer Ideology IV (Segal-Cover) Results - Justice Vote Direction (Conservative %)

	Model 1: Justice FE			
	(1)	(2)	(3)	(4)
	Vote Direction	Vote Direction	Vote Direction	Vote Direction
Mean All Peer Justices	-0.156			
	(0.117)			
Mean Active Peer Justices		-0.083		-0.067
		(0.114)		(0.113)
Mean Absent Peer Justices			-0.120	-0.119
			(0.121)	(0.121)
First Stage F-Statistic	54075	43013	664	
Observations	110729	110729	110729	110729

	Model 2: Justice, Term FE			
	(1)	(2)	(3)	(4)
	Vote Direction	Vote Direction	Vote Direction	Vote Direction
Mean All Peer Justices	-1.704			
	(1.375)			
Mean Active Peer Justices		1.304***		1.239*
		(0.497)		(0.692)
Mean Absent Peer Justices			-0.160	-0.012
			(0.117)	(0.160)

	Model 3: Justice by Issue Area, Term FE			
	(1)	(2)	(3)	(4)
	Vote Direction	Vote Direction	Vote Direction	Vote Direction
Mean All Peer Justices	0.222			
	(0.245)			
Mean Active Peer Justices		0.411*		0.518**
		(0.220)		(0.227)
Mean Absent Peer Justices			-0.027	0.017
			(0.108)	(0.110)
First Stage F-Statistic	865	735	85	
Observations	110554	110554	110554	110554

	Model 4: Justice by Issue Area, Term by Issue Area FE			
	(1)	(2)	(3)	(4)
	Vote Direction	Vote Direction	Vote Direction	Vote Direction
Mean All Peer Justices	-1.921*			
	(1.162)			
Mean Active Peer Justices		0.811**		1.351***
		(0.396)		(0.422)
Mean Absent Peer Justices			0.028	0.143**
			(0.064)	(0.064)
First Stage F-Statistic	67	103	214	
Observations	110554	110554	110554	110554

Models estimated with associated set of covariates used in analogous OLS regression, see Tables 3 & 4. Peer variables are constructed using the analogous first stage justice coefficients estimates. Segal-Cover peer measures used as instruments are constructed from individual justice Segal-Cover scores.
* p<0.10, ** p<0.05, *** p<0.01

sulting *case selection bias* will appear to manifest itself as a negative peer effect (and thus bias this estimate downwards), since movements of the Court's ideological composition in one direction will change the distribution of cases heard, moving the average vote of continuing justices in the opposite direction.

The existence of such a mechanism cannot be tested merely by looking at the relationship between observed votes and justice ideology, because this does not separate the effects of peer effects and case selection upon votes, and hence little can be said about unobserved case characteristics. Assessing justice ideology based on voting propensity when this is possibly affected by case selection complicates matters further.

A more fruitful approach is to consider the relationship between Segal-Cover scores (as a predetermined measure of justice ideology, identified separately from votes) and case characteristics that are known to be viewed as particularly conservative or liberal. If observable case characteristics are impacted in one direction, it seems most plausible that this will be true of unobservable case characteristics also. Given this, recall that a substantial majority of Supreme Court decisions are to overturn the lower court ruling. Accordingly, reviewing a larger number of conservative (vis a vis liberal) lower court decisions is behavior that would intuitively be consistent with a comparatively liberal Court, if case selection effects exist. Figure 3 documents the share of lower court directional opinions in the liberal direction by natural court (a period during which no personnel change occurs), and its relationship with justice ideology. This reveals a strong relationship as hypothesized, with more liberal Supreme Court cohorts (high average Segal-Cover scores) mostly reviewing conservative lower court opinions, and vice versa.

This reveals an additional rationale for controlling for term in the models considered above. To the extent that case selection is governed by the justices jointly, irrespective of whether a justice will be recused or not,[32] this means that case selection effects will be common (at least by issue area) within a natural court. Term dummies thus capture this effect, so the peer effect coefficients are not biased. However, in specifications without term dummies, downward bias will result, which may partially explain the estimates for Model 1 in Tables 4 and 5.

[32]This does not require that a justice who will ultimately recuse themselves from the case still participate in selecting the case to be heard, but rather that their recusal does not change the probability that the case is selected to be heard.

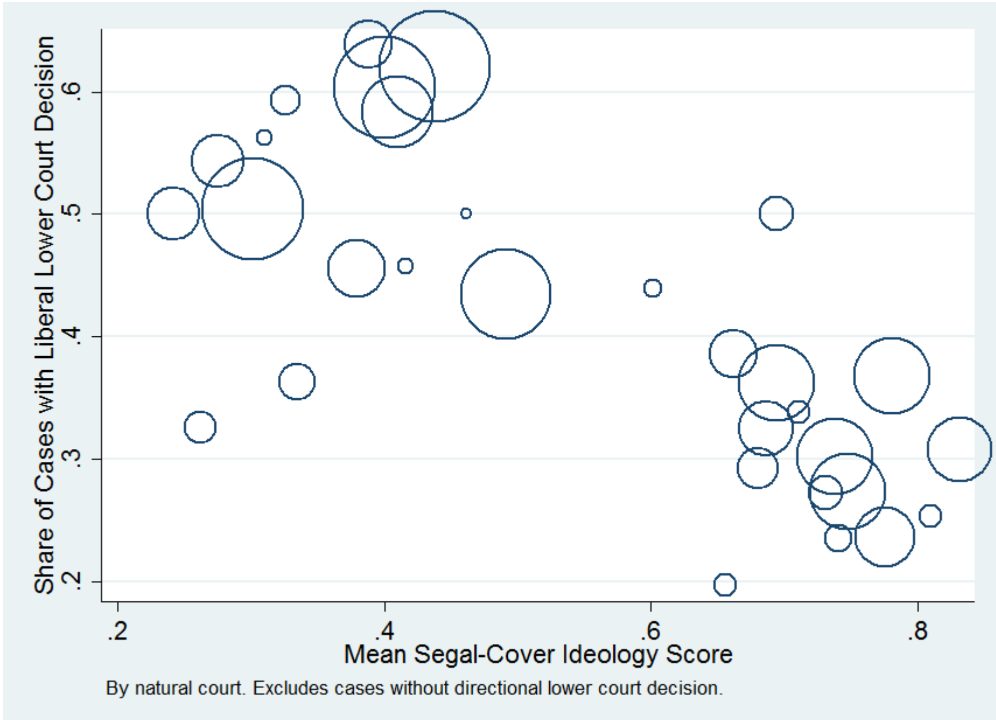


Figure 3 – Endogenous Case Ideology Selection

4 Peer Vote Effects

An alternate possible peer effect mechanism is that justices influence their colleagues through their own votes, so that justices' respective votes are jointly determined on a case-by-case basis. The intuition behind such a mechanism is simple: any attempt a justice makes to influence how their peers vote on a given case will reflect their own voting disposition. Accordingly, the first mechanism where peers affect ideological positions discussed in Section 3 may merely be a reduced-form representation of this structural relationship through votes, since vote probabilities are in large part driven by justice ideology.

4.1 Empirical Specification and Vote Endogeneity

To estimate the effect of the votes of peers on a justice's vote, a similar specification to Equation (3) is used, except that peer effects are captured through a variable reflecting the mean vote of other justices $\overline{d_{\text{-}j,ct}}$ in the same case (rather than their ideology).

$$
\begin{aligned}
u_{jct} = &\alpha_j + \gamma_c + \delta_t + l_c + \beta_p \times \overline{d_{\text{-}j,ct}} + \beta_1 \times lc\_dec_c + I\left[j \in app_c\right] \times \left[\beta_2 + \beta_3 \times app\_tenure_j\right] \\
&+ lc\_dec_c \times I\left[j \in app_c\right] \times \left[\beta_4 + \beta_5 \times app\_tenure_j\right] + \varepsilon_{jct}
\end{aligned}
\tag{4}
$$

This again includes justice and term fixed effects to control for systematic variation in vote ideology propensities across justices and time. However, unlike previously, focus is given to the simpler specification without justice by issue area and term by issue area fixed effects, since identifying the peer vote effect mechanism does not require generating precise estimates of justice ideology or exogenous variation in the ideology of peers.[33]

As before, β_p captures the relationship with peers, with a positive coefficient indicating that justices are inclined to vote in accordance with their peers. However, β_p cannot be interpreted as a consistent estimate of peer effects since votes are jointly determined. Unobserved case characteristics which affect the ideological position of a case drive the votes of both a specific justice and their peers, yielding an omitted variable bias in the OLS estimates. Since these unobserved case characteristics include almost everything material to the case,[34] the vote of peers provides substantial information about the nature of the case. Recalling that in the full sample 37% of cases involve a unanimous vote, even the vote of a single justice has very substantial predictive power over how other justices vote.

Accordingly, very strong correlations can exist between votes, irrespective of the existence of peer effects. Table 6 documents these strong correlations, showing the OLS estimates from regressions of vote direction on three different measures of the votes of other justices as the endogenous variable. Column 1 uses the mean vote direction (proportion conservative) of other justices in the case. Columns 2 and 3 explore the predictive power provided by the votes of *home* justices in *home court* cases; defined as those which are sourced from the Circuit Court of Appeals on which the justice previously served. Column 2 shows the estimated relationship with the mean vote of other

[33]Furthermore, the instrument used for votes (see below) is by definition unrelated to issue area or term, and empirically the correlation appears small. i.e. all of the results are robust to adding these controls.

[34]The observed case characteristics include only the legal issue area, the lower court decision, the Circuit Court of Appeals (if any) that the case stems from, and the term in which the case is heard by the Supreme Court. These jointly explain relatively little of the variation in case vote outcomes.

home justices in that same case.[35] Since the relationship between the votes of home justices should be more predictive when they are more numerous, Column 3 considers the relationship with the net vote direction of other home justices, constructed as the number of other home justices issuing conservative votes less liberal votes, divided by the total number of other justices present in the case.[36] This thus captures both both the frequency of home justices and their level of agreement in a particular case. As expected, each of these regressions reveals a strong relationship between the votes of justices, but due to endogeneity bias this provides no insight into the existence of peer effects.

Table 6 – Peers Vote Effects OLS (Endogenous) - Justice Vote Direction (Conservative %)

	(1) Vote Direction	(2) Vote Direction	(3) Vote Direction
Peer Vote Mean	0.860***		
	(0.003)		
Home Peer Vote Mean		0.444***	
		(0.014)	
Net Home Peer Vote Mean			1.467***
			(0.050)
Circuit Court FE	Yes	Yes	Yes
Justice FE	Yes	Yes	Yes
Issue Area FE	Yes	Yes	Yes
Term FE	Yes	Yes	Yes
R-squared	0.7252	0.5699	0.5692
Observations	110729	110729	110729

* $p<0.10$, ** $p<0.05$, *** $p<0.01$

4.2 Instrumental Variables Estimation Results

To identify any true peer vote effects it is necessary to isolate exogenous variation in voting propensity across justices. This requires a variable which directly affects how a justice votes in a given case, but has no plausible rationale for affecting the votes of others except through the vote of the directly affected justice. While typical observed case characteristics produce variation in votes across cases, they do so simultaneously for all justices, so direct and peer effects cannot

[35]Since this is by convention set to zero in cases where no home justices are present, such as any case not from a Circuit Court of Appeals, a dummy variable is added to indicate the presence of another home justice.

[36]For example, if there is a single home peer justice, and they vote liberal, this variable is $-1/8$. If there are three home peers, of which two vote liberal and the other conservative, the variable is also $-1/8$. If there are two home peers, and both vote liberal, it is $-1/4$.

be separated. More fruitfully, as mentioned in Section 3.2, justices who have previous service on a Circuit Court of Appeals vote differently when hearing cases that are sourced from their home court. In particular, justices who had short tenures on a Circuit Court of Appeals are on average less likely to overturn a lower court opinion, while the reverse is true for justices with long home court tenures. Figure 4 documents this tendency by plotting the differential in the rate at which justices overturn decisions in cases from their home court compared to all other cases, against the duration of home court tenure, for each of the 19 justices who previously served on a Circuit Court of Appeals.

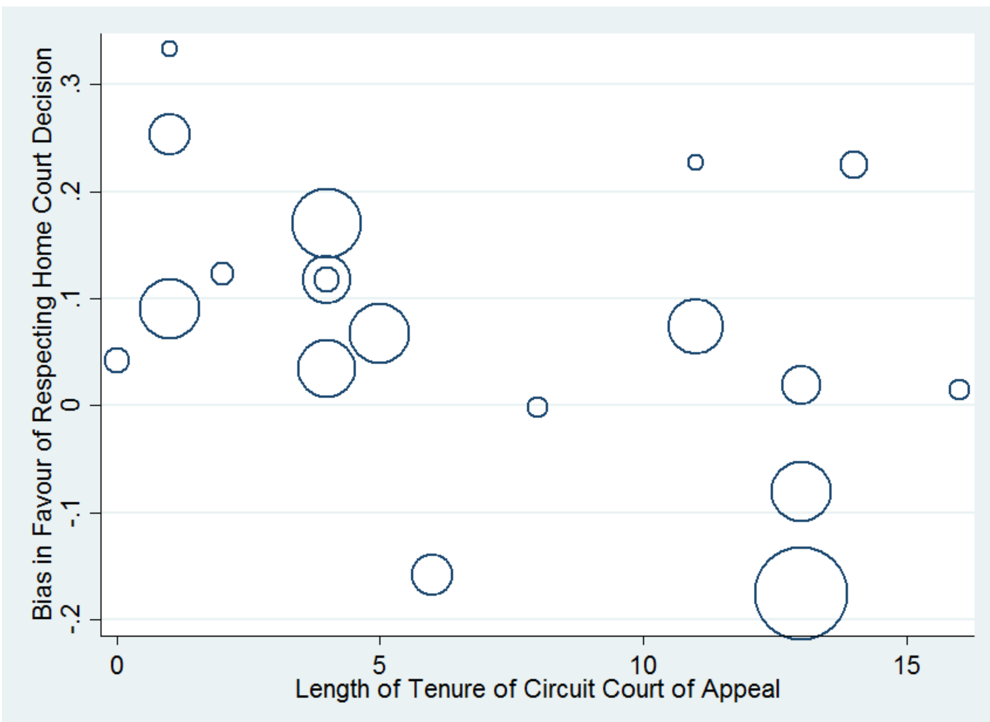


Figure 4 – Home Court Bias in Overturn Rate of Lower Court Decisions

It is thus possible to consistently estimate Equation (4) by Two-Stage Least Squares, using the share of other justices *at home* $\frac{1}{N-1}\sum_{j\neq i} I[j \in app_c]$ and the average length of home court tenure per justice $\frac{1}{N-1}\sum_{j\neq i}(I[j \in app_c] \times app\_tenure_j)$ in a case (where the denominator counts both home and *away* justices) as instruments for the votes of peer justices. Since the home court relationships affect overturn rates, to capture the effects on vote ideological direction (the dependent variable) these two variables are interacted with the ideological direction of the lower court opinion.[37] This method relies on the exclusion restriction that a justice's vote is affected by the presence of home justices and the length of their home tenure only through the votes of the home justices (directly) and other away justices (indirectly, through the potential peer mechanism).

To mitigate any possibility that the instruments are contaminated by some selection effect regarding which justices are present and vote in respective cases, two different specifications of the instruments are considered. These both utilize the share of other justices at home and the average length of home court tenure per justice, but in one specification the instruments are defined using all justices on the Supreme Court while the other only uses the justices active in each respective case.[38]

Consistent with Figure 4, Table 7 shows there is a strong relationship between these home justice variables and voting propensities. The pattern of justices with short (long) home tenure being respectively less (more) likely to overturn lower court decisions (indicated by the +, -, - + pattern of the four coefficients) is evident irrespective of whether all, or only active justices, are considered. Estimates in the latter case (see the bottom of Table 7) are generally slightly larger, which is consistent with the inclusion of absent justices adding noise to the instruments.

The second stage estimates exploit this natural variation in justice votes driven by home court affiliation to estimate the extent to which a justice's vote is causally affected by the votes of their peers. These estimates, documented in Table 8, show that the strong correlation between justice votes is not solely due to unobserved case characteristics. Indeed, the IV estimates in Table 8 are fairly similar to the OLS results in Table 6.

The indicated magnitude of peer effects is sizeable and of practical significance for each of the peer measures mentioned above. Columns 1 and 2 show that, holding all else equal, a percentage point increase in the proportion of peers issuing a conservative vote in a case makes a justice 0.9 percentage points more likely to vote conservatively. In the typical full panel case (with 8 peer justices), this means that a single peer experiencing a 10 percentage point increase in conservative vote probability yields a direct effect of 1.1 percentage points on each other justice.

[37] This is only for liberal and conservative lower court decisions. In cases where the lower court opinion is not of specifiable direction, overturning the lower court is not well defined.

[38] If there are no selection effects to be concerned about, the latter specification is more intuitive since the endogenous variable can only utilize the votes of active justices.

Table 7 – Peers Vote Effects IV First Stage- Peer Vote Measures

	Peer Vote Mean		Home Peer Vote Mean		Net Home Peer Vote Mean	
	(1)	(2)	(3)	(4)	(5)	(6)
Share of Peers at Home						
× Conservative	0.199		0.384*		0.177**	
	(0.170)		(0.211)		(0.090)	
× Liberal	-0.567***		-1.403***		-0.280***	
	(0.175)		(0.229)		(0.094)	
Peer Mean Years at Home						
× Conservative	-0.038		-0.085***		-0.025***	
	(0.026)		(0.030)		(0.009)	
× Liberal	0.078***		0.217***		0.050***	
	(0.020)		(0.025)		(0.008)	
Share of Active Peers at Home						
× Conservative		0.303*		0.470**		0.198**
		(0.173)		(0.237)		(0.093)
× Liberal		-0.573***		-1.420***		-0.304***
		(0.177)		(0.256)		(0.095)
Active Peer Mean Years at Home						
× Conservative		-0.056**		-0.085***		-0.027***
		(0.026)		(0.031)		(0.009)
× Liberal		0.070***		0.231***		0.053***
		(0.020)		(0.027)		(0.009)
R-squared	0.6871	0.6872	0.5853	0.5870	0.0841	0.0885
Observations	110729	110729	110729	110729	110729	110729
First Stage F-Statistic	5.263	5.426	25.037	26.177	23.055	24.510
First Stage P-Value	0.000	0.000	0.000	0.000	0.000	0.000

* p<0.10, ** p<0.05, *** p<0.01

Columns 3 to 6 focus explicitly on the effect that the votes of home justices have on their peers. A percentage point increase in the proportion of home peers who issue a conservative vote in a case makes the votes of their peers on average 0.3 percentage points more conservative. Accordingly, in cases with a single home justice, switching their vote has a 30 percentage point effect on peer votes. The final two columns allow the peer effect of an additional home justice being in a case to be calculated; such a change produces a one-eighth change in the net home peer vote mean variable, and thus has a 14 percentage point effect on the conservative vote probability of peers.[39]

Table 8 – Peer Vote Effects IV Second Stage - Justice Vote Direction (Conservative %)

	(1)	(2)	(3)	(4)	(5)	(6)
	Vote Direction		Vote Direction		Vote Direction	
Peer Vote Mean	0.902***	0.874***				
	(0.037)	(0.041)				
Home Peer Vote Mean			0.336***	0.302***		
			(0.067)	(0.063)		
Net Home Peer Vote Mean					1.280***	1.131***
					(0.271)	(0.248)
Observations	110729	110729	110729	110729	110729	110729
First Stage F-Statistic	5.263	5.426	25.037	26.177	23.055	24.510

* $p<0.10$, ** $p<0.05$, *** $p<0.01$

When considering endogenous effects, it is possible that initial shocks to voting propensities are propagated from justice to justice. In fact, different propagation mechanisms, which amount to differing peer effect mechanisms, can yield a common average peer effect coefficient. For insight, consider the following stylized examples, with a single home justice experiencing a shock to her vote propensity. Now let λ be the direct effect of one justice's vote on the vote of the other justices, scaled down by the number of peers. We shall refer to this as the *direct effect*. We consider three natural possibilities of how the direct effect translates into the total effect on the vote of a justice.

First, it may be that the vote of a justice affects each other justice only directly, with no propagation through the votes of other justices. This occurs when justices provide information to each

[39]By virtue of the specification, the effect of a home justice switching the ideological direction of their vote is assumed to be twice as large.

other; the vote probability of the individual justice is a sufficient statistic for her signal. This signal can affect the vote probability of each peer justice, but have no subsequent spillovers, because any vote changes by the peer justices are understood to be in response to the initial justice's signal and thus provide no additional information. In such a context, an initial shock of magnitude k to the home justice's vote probability shifts the vote probability of each peer by $\frac{\lambda k}{N-1}$, with no multiplier effect occurring. The lack of multiplier effects means that the home peer vote variable changes by a large amount relative to the mean peer vote measure, limiting the coefficient on home peer votes. In expectation the peer vote mean variable for away justices shifts by $\frac{(N-2)\frac{\lambda k}{N-1}+k}{N-1}$, so the average peer coefficient is

$$\beta_p^1 = \frac{\lambda}{\frac{N-2}{N-1}\lambda + 1}.$$

Given our estimate of $\beta_p = 0.874$ and that $N = 9$, this implies a λ of 3.7. This implies a direct effect of a given justice's vote on the vote probability of any other justice of 3.7/8=0.46, under this (perhaps implausible) hypothetical.

Second, suppose that indirect propagation does occur. For example, in addition to the direct peer effect arising due to the shock experienced by the home justice, suppose justices further respond equally strongly to the induced changes in the votes of their other peers. However, suppose that the home justice experiences no indirect peer effects reflecting back on themselves; as above their initial change in vote probability is a sufficient statistic for the information content they provide. Then an initial shock of magnitude k to the home justice's vote probability produces a direct effect of $\frac{\lambda k}{N-1}$ on the vote probability of each peer, which is then multiplied by $(1-\frac{N-2}{N-1}\lambda)^{-1}$ through the indirect propagation mechanism. Compared to the first propagation mechanism, the mean peer vote variable changes by a large amount relative to the home peer vote measure, with the multiplier effects amplifying the coefficient on home peer votes. In expectation the peer mean vote variable for away justices shifts by

$$\left((N-2) \times \frac{1}{1-\frac{N-2}{N-1}\lambda} \times \frac{\lambda k}{N-1} + k\right) \Big/ (N-1)$$
$$= \frac{k}{(N-1)-(N-2)\lambda},$$

so the average peer coefficient is

$$\begin{aligned}\beta_p^2 &= \frac{\lambda}{1-\frac{N-2}{N-1}\lambda} \Bigg/ \left(\frac{N-2}{1-\frac{N-2}{N-1}\lambda} \times \frac{\lambda}{N-1} + 1 \right) \\ &= \lambda.\end{aligned}$$

Third, suppose that indirect propagation does occur for all justices, including the justice initially experiencing the shock. Then an initial shock of magnitude k to the home justice's vote probability produces a direct effect of $\lambda k/(N-1)$ on the vote probability of each peer, with an immediate reflection on the home justice of $\lambda \times \dfrac{\lambda k}{N-1}$. These effects are then amplified by a factor of $\left(1-\lambda \times \frac{\lambda+N-2}{N-1}\right)^{-1}$. In expectation the total effect on the peer mean vote variable is

$$\begin{aligned}&\left((N-2) \times \frac{\lambda k}{N-1} \times \frac{1}{1-\lambda \times \frac{\lambda+N-2}{N-1}} + \frac{\lambda^2 k}{N-1} \times \frac{1}{1-\lambda \times \frac{\lambda+N-2}{N-1}} + k \right) \Bigg/ (N-1) \\ &= \frac{k}{(N-1) \times \left(1-\lambda \times \frac{\lambda+N-2}{N-1}\right)}\end{aligned}$$

for away justices and

$$\frac{\lambda k}{(N-1) \times \left(1-\lambda \times \frac{\lambda+N-2}{N-1}\right)}$$

for the home justice who experiences the initial shock. Where the average peer coefficient β is identified off variation in the peer vote mean variable for away justices, it is given by

$$\begin{aligned}\beta_p^3 &= \frac{\lambda}{1-\lambda \times \frac{\lambda+N-2}{N-1}} \Bigg/ \left((N-2) \times \frac{\lambda}{N-1} \times \frac{1}{1-\lambda \times \frac{\lambda+N-2}{N-1}} + \frac{\lambda^2}{N-1} \times \frac{1}{1-\lambda \times \frac{\lambda+N-2}{N-1}} + 1 \right) \\ &= \lambda.\end{aligned}$$

Thus, in both case 2 (which we might call "partial reflection"), and case 3 (which we might call "full reflection") we find that $\beta_p = \lambda$. Technically, adding reflection back to the home justice scales up the effect of each justice on each other justice proportionally, leaving the solution to the fixed point problem unchanged.

4.3 Exogeneity of Home Court Occurrences

A natural concern with using the home court status of justices as an instrument for justice voting propensity is that the cases which the court hears are chosen by justices. Hence a justice's previous tenure on a Circuit Court of Appeals may affect the nature of cases that are chosen to be heard from their prior court, relative to other courts. For example, it seems plausible that the same bias that leads justices to have an increased (decreased) propensity to overturn decisions from their home court could also lead them to advocate disproportionately for (against) the Supreme Court reviewing decisions from their home court to begin with.

Crucially, were a case selection bias of this form to exist, it is far from clear that this would bias the IV estimates upwards. First, consider a justice biased towards the home court, who may try to prevent home cases from being reviewed by the Supreme Court. Intuitively, their lobbying to prevent cert being granted is most likely to be successful for cases with below average ex ante overturn probability (based on case characteristics and facts).[40] Selecting out these cases would thus increase the average overturn propensity observed for home cases that reach the Supreme Court, and falsely look like a negative peer effect. Conversely, suppose a justice biased against their home court desires to have additional cases from their home court reviewed by the Supreme Court. Since the Supreme Court has a disproportionate tendency to overturn lower court decisions, it is plausible that the marginal home case that the justice may persuade the Supreme Court to hear has lower than average overturn probability, by virtue of it not otherwise being reviewed.[41]

Moreover, the data regarding the frequency of cases from each Circuit Court of Appeals fails to show any clear link to the presence of home justices. Considering each Circuit Court in turn, Figure 5 separates cases into three groups; cases (irrespective of where they are sourced from) where no justice with previous tenure on the considered Circuit Court is on the Supreme Court, and then those with short and long home tenure justices from the considered court respectively

[40]This may be tempered by the home justice having greater incentive, and thus investing greater effort, to prevent cases with high overturn probability from being reviewed.

[41]This effect may be weak since the Supreme Court chooses to hear only a small proportion of cases over which it has jurisdiction, even when it would counter-factually view the lower court as having made an incorrect decision. The qualifier that justices may focus their lobbying on cases with higher perceived overturn probability also applies.

(note that the latter two can be present simultaneously). For each group, we report the share of cases from the respective Circuit Court. In general, the relative frequency of cases from each Circuit Court is similar regardless of the presence of a justice with tenure from that same court, or the length of that tenure. The most notable exception is an artifact of a consistent increase in the share of cases from the relatively liberal 9th Circuit over time, combined with Justice Kennedy, who had previously served on the 9th circuit, being on the Supreme Court since 1988.

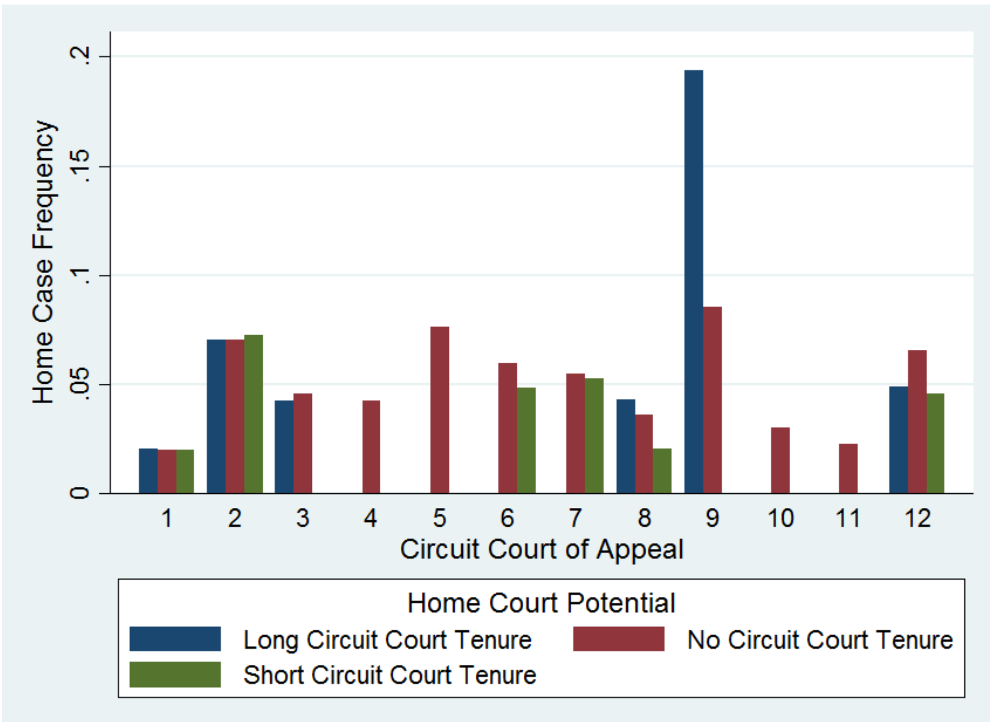


Figure 5 – Home Court Effect in Selection of Circuit Court of Appeals Cases

It is also worth restating that the possible concern that the first stage relationship could be an artifact of recusal behavior does not appear to be merited. As shown in Table 7, the relationship between length of home court tenure and propensity to overturn the Circuit Court's decision holds irrespective of whether all home justices are considered, or only those actively participating in each respective case. This suggests the IV estimates are not being driven by justice decisions about whether to participate or be recused from a case providing information about unobserved (to the econometrician) case facts.

5 Case Outcomes

The two sections above provide strong evidence that peers affect votes. However, if these peer effects do not change pivotal votes and thus alter the direction of case outcomes by switching majority decisions, they are of diminished practical interest. Accordingly it is necessary to establish whether the peer effects documented above are a general phenomena, or only affect votes in cases that are not tightly decided, such that case majority outcomes are not altered.

A natural first pass is to consider the distribution of the number of votes by Supreme Court justices to overturn the lower court decision of different cases. Since short (long) tenure home justices have reduced (increased) propensities to overturn lower court decisions, we would expect cases with short tenure home justices to have fewer justices on average overturning the lower court decision than those with long tenure home justices (with cases with no home justices falling somewhere in-between). Figure 6 shows the cumulative distribution of the number of overturn votes for these three groups of cases, defining long home tenure as more than 8 years, and restricting the sample to cases voted on by a full panel (9 justices) and including at most one home justice. The number of overturn votes in long home tenure cases first-order stochastically dominates that in short tenure cases, while the distribution for cases with no home justice mostly falls between. The magnitude of the difference is substantial and stable across the distribution, with lower court decisions overturned 9 percentage points more often in the long home tenure cases.

Notwithstanding the lack of controls, Figure 6 does not in itself tell us anything about peer effects, since it does not disentangle the change in the home justice's own vote from the votes of the other (*away*) justices. Accordingly, Figure 7 isolates the effect on the away justices by plotting the distribution of overturn votes in these cases once the home justice is excluded.[42] Again, the number of overturn votes for cases with a long tenure home justice first-order stochastically dominate those in short tenure home justice cases. However, as expected the exclusion of the home justice reduces the distance between the distributions, with a 4 percentage point difference in the proportion of cases with at least half of the away justices voting to overturn the lower court

[42]To make cases where there are no home justices comparable, the distribution is calculated by applying equal (1/9) weight to dropping each justice in turn. By comparison, 8-justice cases make a poor placebo group since there is a clear aversion in the data to producing tied votes, which distorts the shape of the cumulative distribution.

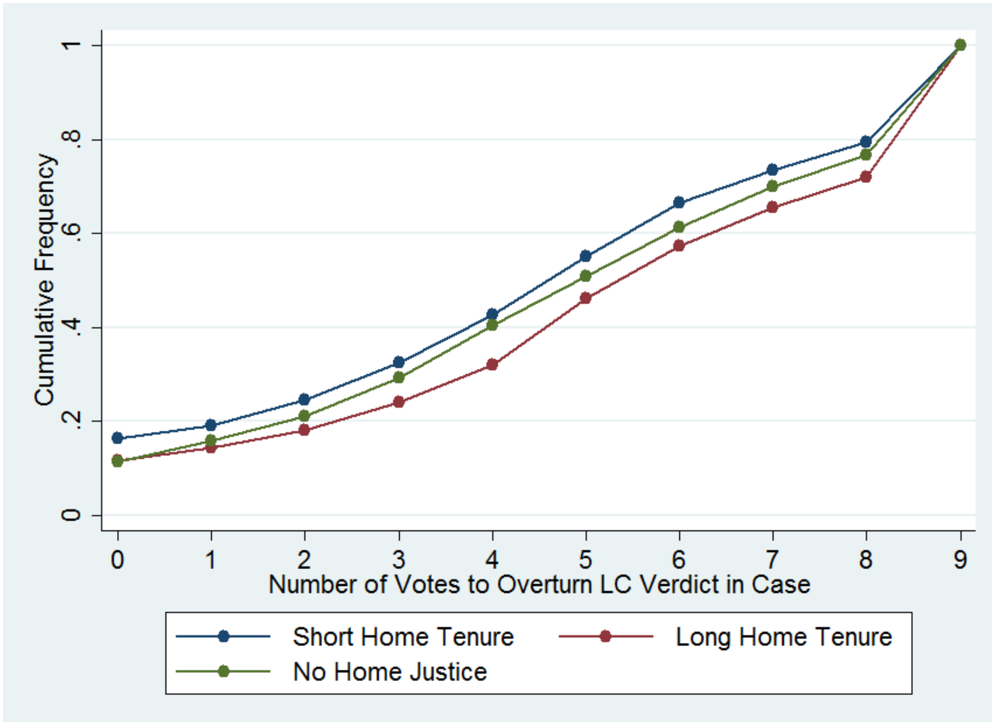


Figure 6 – Distribution of Overturn Votes by Presence of Home Justice

opinion. These results are indicative that the difference in Figure 6 is due to effects upon both the home justice's own vote propensity (consistent with the first stage IV results) and the distribution of overturn votes by peers. In particular, both of these figures suggest that peer effects operate at all levels of case closeness, rather than occurring only in one-sided cases.

5.1 Instrumental Variables Estimation Results

To determine whether the peer effects in case outcomes are statistically significant after controlling for covariates it is possible to use a similar procedure to that discussed in Section 4.2, except that variables are aggregated at the case level. In particular, all the regression analysis in Sections 3 and 4 considered the effect of some characteristic of her peers (namely ideology and votes respectively) on the votes of a single justice, but identifying whether peer effects can change pivotal votes instead requires that we consider a single justice and analyze how their vote affects the collective voting behavior of their peers. Disentangling peer effects from the mechanical effect of a justice's vote on the majority outcome requires excluding the vote of the justice whose perspective

Figure 7 – Distribution of Overturn Votes of Other (Away) Justices by Presence of Home Justice

is taken. As in Section 4, a justice's own vote is jointly determined with the votes of peers, so the home circuit court variables are again used as instruments which identify exogenous variation in votes. However, due to the change in perspective, such that the individual justice's vote is the regressor of interest, the pertinent instruments are the home court status of this individual justice. Correspondingly, the home court status of peers can be used as an additional control variable to explain variation in the measures of peer voting behavior.

Using these instruments, the effects of a justice's own vote on several different measures of the collective votes of peers are considered. First, the number of conservative votes by other justices in a case serves to verify the existence of peer vote effects when analyzed at the case level, and also is potentially informative about the extent of propagation. However it sheds no light on the case circumstances (i.e., types of cases) under which peer effects operate. Second, in order to measure whether peer effects change pivotal votes, we define a case's opinion as *potentially conservative* or *potentially liberal* if, not counting a justice's own vote, enough others (for example, at least four out of eight) vote in the pertinent direction that such a opinion outcome is mathematically possible.[43] Considering the potential outcome variables, rather than the overall majority outcome, enables us to estimate the effect of a justice's own vote on the decision's direction, excluding the mechanical effect of her own vote.

Estimating these case outcome effects in a rigorous manner requires restricting the samples and altering the covariates to properly isolate peer effects. In the first set of analyses for each of the dependent variables, the sample is restricted to cases with a full panel of justices, and either zero or one justice at home. In cases with a home justice, full weight is given to the home justice's observation (such that the dependent variable considers the votes of only the away justices), while in cases where all justices are away each is given equal (1/9) weight. This means that in total, each case is weighted equally, and there is no arbitrariness in which justices are included in the dependent or independent variables.

Letting $d_{\text{-}j,ct}$ represents the pertinent collective vote measure of the peer justices and d_{jct} the vote of the justice whose perspective is taken, the regressions estimated have the following functional form.

$$\overline{d_{\text{-}j,ct}} = \alpha_j + \gamma_c + \delta_t + l_c + \beta_1 \times lc\_dec_c + \beta_p \times d_{jct} + \varepsilon_{\text{-}j,ct} \tag{5}$$

A few details of the estimation are worth noting. Since outcomes are considered at the case level (except for the exclusion of the single justice whose perspective is taken) and only full panel cases are considered, the natural control for justice cohorts would be to include fixed effects for the natural court, but these are nearly perfectly collinear with term and are hence omitted.[44] Since the justice whose perspective is taken (and whose vote is thus not considered in the dependent variable) varies across (and within) cases for a given cohort, the dependent variable is mechanically affected by the excluded justice's ideology. This is addressed by the inclusion of justice fixed effects. Finally, the home circuit court peer variables are dropped since by construction only the votes of away justices are considered in the dependent variables.

[43]Thus it is possible that a case can have both conservative and liberal potential from the perspective of some justice when they are voting, with their own vote deciding the actual decision direction.

[44]Recall most of the within-year variation in justice cohort is due to recusals, which are excluded since only full panel cases are considered here.

Results for these analyses are shown in the first three columns of Table 9. The coefficient for the home justice's vote direction in Column 1 is consistent with previous individual-level analysis; a single (implicitly home) justice switching their vote changes the net votes of their (away) peers by approximately 2.8 votes collectively. This highlights the empirical importance of peer effects; the indirect effect of a justice's vote on the total vote outcome through the votes of their peers is several times stronger than the direct mechanical effect of the justice's own vote. Columns 2 and 3 verify that these peer effects alter case outcomes. The home justice switching their vote from liberal to conservative has a substantial peer effect in the same direction, increasing the share of cases with conservative outcome potential[45] by 36 percentage points and reducing the share with liberal potential by 32 percentage points.

A second set of regressions expand the sample of cases considered to all cases with a full panel of justices, such that cases with more than one home justice are included, and each justice observation is weighted equally. For each observation the dependent variables are again constructed using the votes of all the other peer justices in the respective case. The specifications considered are as above, except that since the dependent variables can incorporate the votes of home justices, the home circuit court peer variables for the peer justices are added as controls, yielding a functional form of

$$\begin{aligned}\overline{d_{\text{-}j,ct}} =& \alpha_j + \gamma_c + \delta_t + l_c + \beta_1 \times lc\_dec_c + [\beta_2 + \beta_3 \times lc\_dec_c] \times \frac{1}{N-1} \sum_{i \neq j} I[i \in app_c] \\ &+ [\beta_4 + \beta_5 \times lc\_dec_c] \times \frac{1}{N-1} \sum_{i \neq j} (I[i \in app_c] \times app\_tenure_j) + \beta_p \times d_{jct} + \varepsilon_{\text{-}j,ct}\end{aligned} \tag{6}$$

The estimates from these specifications are shown in Columns 4 through 6 of Table 9. In each case the results are similar to those in the first three columns, and again provide strong evidence that peer effects shift pivotal votes. According to the point estimates, an individual justice switching their vote from liberal to conservative increases the probability that the other justices collectively vote in a manner that produces conservative outcome potential by 32 percentage points, and decreases liberal outcome potential by 40 percentage points.

While both sets of estimates find large effects on the potential ideological direction of case

[45]For example, shiting the vote of other justicess from 3-5 or less to 4-4 or more.

Table 9 – Peer Vote Effects on Verdict Direction Outcomes IV

	(1) Conservative Peer Votes	(2) Conservative Potential	(3) Liberal Potential	(4) Conservative Peer Votes	(5) Conservative Potential	(6) Liberal Potential
Vote Direction	2.781***	0.361***	-0.323**	3.037***	0.321**	-0.397***
	(0.683)	(0.128)	(0.133)	(0.750)	(0.152)	(0.146)
Observations	67576	67576	67576	84267	84267	84267
First Stage F-Statistic	11.792	11.792	11.792	9.755	9.755	9.755

* p<0.10, ** p<0.05, *** p<0.01

opinions, a plausible argument can be made that these estimates are biased towards zero. This possibility arises because, due to the need to exclude a justice's own vote, cases can be categorized as having both conservative and liberal potential. Suppose justices do not care purely about the margin of victory, but attach additional importance to winning any majority. A justice who plans to issue a conservative (liberal) vote has great incentive to attempt to influence the votes of peers to achieve at least a 4-4 split, and hence *cause* a *potentially conservative (liberal)* outcome. Once this is achieved, however, they would have a lesser incentive to go further and *prevent* a *potentially liberal (conservative)* outcome (requiring a split of at least 5-3), because their desired majority outcome is achieved regardless due to their own vote. Such a mechanism would systematically reduce the between-group difference in the outcome potential measures and hence bias the estimated effect of the considered justice's vote downwards. Note that this relates to reduced effort (once a majority is achieved), and could arise even if justices exert full effort to try to win a majority.

6 Discussion and Conclusion

We have presented a variety of estimates suggesting that the ideology and voting behavior of a justice affects the voting behavior of other justices. Moreover, our estimates imply that these peer effects can be pivotal and thus affect case outcomes–and the magnitudes are meaningful.

This raises the question of why these effects exist and what drives them. As we mentioned in the introduction, a variety of explanations have been offered in the context of lower courts, including: deliberation, group polarization, aversion to dissent, or deference to expertise. It is challenging to provide compelling evidence distinguishing between these different channels. We do know, by virtue of the fact that peer effects can cause a change in outcome, that dissent aversion (not wanting be an outlier justice on a case) cannot be the whole story.

Having said that, it is not easy to distinguish between justices persuading each other, being deferential to each other on areas of expertise, or even some form of *horse trading*. We can, however, get some sense of whether Posner (2008)'s *deference effect* is at work. Under that hypothesis, roughly put, justices defer to other justices who have expertise in a certain area of law, other than for highly politicized issues. As Posner puts it: "The hotter the issue (such as abortion, which nowadays is much hotter than, say, criminal sentencing), the greater the explanatory power of the political variable."

In the following table, we estimate our peer effect coefficient (including justice by issue area and term by issue area fixed effects) separately for the 11 issue areas. To facilitate precise estimates, these are peer vote regressions, with identification for each issue area coming from using the mean active justice ideology estimates in that issue area as instruments for the mean vote of peers.[46] A first thing to note is that these are fairly coarse categories that typically include some "hot" issues and some uncontroversial ones. Second, some of the first stage F-statistics indicate weak instrument problems, and some of the standard errors are large (the Privacy and Unions coefficients, for instance, are almost completely uninformative).

Notwithstanding that, it is noteworthy that, relative to the average coefficient of about 0.6, the issue areas with stronger peer effects include: Attorneys, Economic Activity, Judicial Power and Federal Taxation, all of which are arguably on the "cooler" end of the political spectrum. Conversely, First Amendment, Civil Rights and Due Process have lower-than-average coefficients and are arguably on the "hotter" end of the political spectrum.[47]

There may be less "deference to expertise" on the Supreme Court compared to Circuit Courts of Appeals because, given the types of cases the Supreme Court hears, it is less likely that a justice has particular expertise in the area. Justice Breyer and administrative law is an example of a justice with particularly relevant expertise, but it is not easy to think of a large number of such examples.

[46]See Appendix B for further discussion of peer vote effect regressions that utilise estimated peer ideology as an instrument for peer votes.

[47]We again emphasize caution in interpreting the Civil Rights and Due Process coefficients given the large standard errors.

Table 10 – Peer Effects by Issue Area - Justice by Issue Area, Term by Issue Area FE

	Coefficient	Standard Error	First Stage F-Statistic	Observations
Criminal Procedure	0.601***	0.077	29.143	22549
Civil Rights	0.346	0.350	3.370	18435
First Amendment	0.547***	0.162	8.674	9895
Due Process	0.442	0.535	1.427	4975
Privacy	2.096	2.376	0.226	1483
Attorneys	0.724***	0.196	3.239	1122
Unions	-0.109	2.594	0.179	4387
Economic Activity	0.674***	0.105	11.508	21447
Judicial Power	0.806***	0.067	10.020	17041
Federalism	0.853***	0.092	3.685	5805
Federal Taxation	0.760***	0.145	4.123	3415

Coefficients of mean peer vote on justice vote. Coefficients calcualted from separate regressions for each issue area, using ideology of active peer justices for that issue area as instrument. * $p<0.10$, ** $p<0.05$, *** $p<0.01$

Nonetheless, the "hot button" issue effect that Posner conjectures, and that we provide some evidence for, could still operate in the absence of deference to expertise. It could simply be that on "hot button" cases justices decide ideologically, and on other cases they are more persuadable by their colleagues.

It would be highly desirable to precisely estimate our coefficient for each "issue"–which is more granular than "issue area"–but given our identification strategy and the number of cases, it is not possible to draw meaningful conclusions from that exercise.

Our estimates of peer effects also speak to the broader issue of the optimal strategy for a president nominating a justice. This requires balancing the proximity of the justice's ideology to that of the president, as well as the effect they will have on their peers. An immediate implication of this is that optimal nominations are "court specific" in the sense that they depend on the existing justices, as well as presidential preferences.

Finally, the magnitude of the peer effects that we estimate implies that the indirect effect of a justice's vote on the outcome through the votes of their peers is several times larger than the direct mechanical effect of the justice's own vote. Thus, the replacement for a particularly liberal or conservative justice (such as Justice Scalia) is particularly consequential in that it has the potential to have a material impact on case outcomes.

Appendices

A Justice Ideology Point Estimates

Table 11 orders the justices from the 1946-2013 period according to their estimated ideology, from most liberal to most conservative. The estimates from the second model in Section 3 are used to provide a single point estimate per justice. This model controls for term fixed effects and thus the justice estimates abstract from joint ideological drift in the views of justices and secular changes in the ideological composition of cases heard by the Supreme Court. Rather than ranking justice ideology in an absolute sense, this attempts to measure the ideology of justices relative to their social milieu. Alternatively, by accounting for time effects that would affect any justice serving in an equivalent context, these ideology scores are interpretable as estimating the relative ideologies of any set of justices had they counterfactually been on the Supreme Court at the same time.

B Peer Vote Effects Robustness Tests

Section 4 estimates the effect of peer votes on a justice's vote using home court affiliation of peers as an instrument for their votes. This provides evidence on the effect of a given justice changing their vote, holding everything else about them and the case under consideration constant.

The peer ideology measures constructed in Section 3 provide an alternative set of instruments for the peer vote regressions, with identification again coming through recusal driven variation in peer ideology. Results for these alternative regressions are shown in Table 12, using the peer ideology estimates from the fourth specification. Since this specification estimates each justice's ideology by issue area and also allows for differential ideological drift by issue area over time, the peer vote regressions here include justice by issue area and issue area by term fixed effects to prevent mechanical bias. When using the ideology of active justices as an instrument, the estimated peer vote effects are substantial and statistically significant, providing further credence to the results in Section 4.

Table 11 – Justice Ideology Estimates

Justice	Ideology Estimate	Segal-Cover Score	Conservative Vote Proportion	Party of Appointing President
W. O. Douglas	0.1784	0.730	0.2154	Democratic
W. B. Rutledge	0.1864	1.000	0.2336	Democratic
F. Murphy	0.1940	1.000	0.2424	Democratic
T. Marshall	0.2176	1.000	0.2802	Democratic
W. J. Brennan	0.2438	1.000	0.2930	Republican
H. L. Black	0.2483	0.875	0.2820	Democratic
A. Fortas	0.2517	1.000	0.3082	Democratic
E. Warren	0.2583	0.750	0.2703	Republican
A. J. Goldberg	0.2735	0.750	0.2404	Democratic
J. P. Stevens	0.3508	0.250	0.3889	Republican
R. B. Ginsburg	0.3937	0.680	0.3863	Democratic
H. A. Blackmun	0.4174	0.115	0.4790	Republican
D. H. Souter	0.4174	0.325	0.4183	Republican
S. Sotomayor	0.4174	0.780	0.3712	Democratic
S. G. Breyer	0.4263	0.475	0.4160	Democratic
E. Kagan	0.4403	0.730	0.3963	Democratic
P. Stewart	0.4533	0.750	0.5046	Republican
T. C. Clark	0.4591	0.500	0.4764	Democratic
B. R. White	0.4680	0.500	0.5201	Democratic
F. M. Vinson	0.4967	0.750	0.5635	Democratic
F. Frankfurter	0.5022	0.665	0.5394	Democratic
S. Minton	0.5217	0.720	0.5688	Democratic
S. F. Reed	0.5242	0.725	0.5708	Democratic
H. H. Burton	0.5260	0.280	0.5669	Democratic
L. F. Powell	0.5327	0.165	0.6084	Republican
C. E. Whittaker	0.5353	0.500	0.5516	Republican
R. H. Jackson	0.5420	1.000	0.6157	Democratic
J. Harlan II	0.5471	0.875	0.5729	Republican
W. E. Burger	0.5755	0.115	0.6574	Republican
S. D. O'Connor	0.5927	0.415	0.6245	Republican
A. M. Kennedy	0.6042	0.365	0.6042	Republican
J. G. Roberts	0.6430	0.120	0.6126	Republican
W. H. Rehnquist	0.6659	0.045	0.7134	Republican
A. Scalia	0.6804	0.000	0.6793	Republican
S. A. Alito	0.6992	0.100	0.6653	Republican
C. Thomas	0.7221	0.160	0.7157	Republican

Notwithstanding previous remarks dealing with endogeneity concerns regarding the use of the constructed ideology measures, Table 13 provides further estimates using the predetermined Segal-Cover scores of peers as instruments for the votes of peers in a case. To avoid weak instrument problems, unlike in Section 3.4 the Segal-Cover scores are not interacted with issue area.[48] For the results where the first-stage F-statistic for the instruments suggest there is sufficient variation to gain identification, this yields peer vote effects of a very similar magnitude to those in Table 12. One point worth making is that these ideology based measures yield somewhat smaller coefficients than the home court affiliation measures in Table 8. This seems likely to stem from the fact that these ideology measures gain identification through recusals, and thus exogenous variation in the votes of peers is driven by changes in the set of peers itself. It seems plausible that a change in the mean vote of a justice's peers would be more convincing and thus have a larger peer effect when produced by a given justice changing their mind (and thus holding everything else about that justice constant), compared to when produced by a change in the set of peers present (i.e. if only Nixon can go to China, only Scalia can convince you to vote liberal).

Table 12 – Peer Vote Effects IV Using Ideology Estimate Instruments

	(1) Vote Direction	(2) Vote Direction	(3) Vote Direction	(4) Vote Direction
Peer Vote Mean	0.064	0.599***	-0.247	0.671***
	(1.600)	(0.055)	(0.923)	(0.041)
Observations	110729	110729	110729	110729
Instruments	All Peers	Active Peers	Absent Peers	Active & Absent
First Stage F-Statistic	0.347	57.926	1.778	39.246

* p<0.10, ** p<0.05, *** p<0.01

Table 13 – Peer Vote Effects IV Using Segal-Cover Score Instruments

	(1) Vote Direction	(2) Vote Direction	(3) Vote Direction	(4) Vote Direction
Peer Vote Mean	1.274***	0.643***	0.695***	0.652***
	(0.206)	(0.078)	(0.097)	(0.077)
Observations	110554	110554	110554	110554
Instruments	All Peers	Active Peers	Absent Peers	Active & Absent
First Stage F-Statistic	2.021	20.371	10.114	10.196

* p<0.10, ** p<0.05, *** p<0.01

[48] In that analysis, the endogenous variable was the constructed peer ideology measure, with which the Segal-Cover ideology score is more closely related.

C Peer Effect Measure and Justice Ideology

We remarked in the text that the justice ideology measures obtained in the first stage are contaminated by any peer effects of other justices, and that this in turn causes our peer ideology measures to be contaminated by a justice's own ideology. However, this contamination is washed out when we use fixed effects in the second-stage regressions. We now demonstrate this formally.

Let α_j^p and $\bar{\alpha}_{-j}^p$ denote our proxies for justice ideology and the peer variable, respectively. Now let votes, which are influenced by the true measures α_j and $\bar{\alpha}_{\text{-}j}$, follow

$$d_{jc} = \alpha_j + \beta\bar{\alpha}_{\text{-}j} + \varepsilon_{jc}.$$

If court composition is unchanged during the tenure of justice j then $\bar{\alpha}_{\text{-}j}$ is constant during her tenure. So if we estimate

$$d_{jc} = \alpha_j^p + \xi_{jc},$$

we will obtain (in large samples, by the Khintchine Law of Large Numbers (hereafter "KLLN")) the following proxy

$$\alpha_j^p = \alpha_j + \beta\bar{\alpha}_{\text{-}j}.$$

We now construct the following:

$$\begin{aligned}
\bar{\alpha}_{\text{-}j}^p =& \frac{1}{N-1}\sum_{k\neq j}\alpha_k^p \\
=& \frac{1}{N-1}\sum_{k\neq j}\left(\alpha_k + \beta\bar{\alpha}_{\text{-}k}\right) \\
=& \left(\frac{1}{N-1}\sum_{k\neq j}\alpha_k\right) + \beta\left(\frac{1}{N-1}\sum_{k\neq j}\bar{\alpha}_{\text{-}k}\right) \\
=& \bar{\alpha}_{\text{-}j} + \frac{\beta}{N-1}\left(\frac{1}{N-1}\{(\alpha_2 + ... + \alpha_j + ... + \alpha_N) + (\alpha_1 + \alpha_3 + ... + \alpha_j + ... + \alpha_N) + ...\right. \\
& \left. +(\alpha_1 + ... + \alpha_j + ...\alpha_{N-2})\}\right) \\
=& (1+\beta)\bar{\alpha}_{\text{-}j} + \frac{\beta}{N-1}\left(\alpha_j - \bar{\alpha}_{\text{-}j}\right) = \left(1 + \frac{N-2}{N-1}\beta\right)\bar{\alpha}_{\text{-}j} + \frac{\beta}{N-1}\alpha_j.
\end{aligned}$$

This expression makes clear what is meant by saying that the peer effect measure is contaminated by a term due to the justice's own ideology (the $\frac{\beta}{N-1}\alpha_j$ term).

Now suppose that justice j is observed sitting on a number of different courts $g = 1, ..., G$ with relative frequency π_g, each with a different group of $N-1$ other justices. The true model is now

$$d_{jc} = \alpha_j + \beta\bar{\alpha}_{\text{-}j,g} + \varepsilon_{jc}.$$

If one estimates

$$d_{jc} = \alpha_j^p + \xi_{jc},$$

with a large number of cases per court, the KLLN implies that one obtains

$$\alpha_j^p = \alpha_j + \beta\left\{\pi_1\bar{\alpha}_{\text{-}j,1} + ... + \pi_G\bar{\alpha}_{\text{-}j,G}\right\}.$$

Let us now construct

$$\bar{\alpha}^p_{\text{-}j,g} = \frac{1}{N-1}\sum_{k\neq j}\alpha_k^p.$$

As we showed above, this is of the form

$$\bar{\alpha}^p_{\text{-}j,g} = \left(1 + \beta\frac{N-2}{N-1}\right)\bar{\alpha}_{\text{-}j,g} + \frac{\beta}{N-1}\alpha_j. \tag{7}$$

Now suppose we run the regression

$$d_{jc} = \gamma_j + \theta\bar{\alpha}^p_{\text{-}j,g} + \omega_{jc}.$$

where γ_j are justice fixed effects in this second-stage estimation and θ is the key estimated parameter that captures peer effects.

As this estimation incorporates justice fixed effects, let us de-mean (7) over g.

$$\begin{aligned}\overline{\bar{\alpha}^{p}_{\text{-}j,g}} &= \sum_{g=1}^{G} \pi_g \left(1+\beta\frac{N-2}{N-1}\right)\bar{\alpha}_{\text{-}j,g} + \sum_{g=1}^{G}\pi_g\left(\frac{\beta}{N-1}\alpha_j\right) \\ &= \left(1+\beta\frac{N-2}{N-1}\right)\sum_{g=1}^{G}\pi_g\bar{\alpha}_{\text{-}j,g} + \frac{\beta}{N-1}\alpha_j.\end{aligned}$$

It follows that

$$\bar{\alpha}^{p}_{\text{-}j,g} - \overline{\bar{\alpha}^{p}_{\text{-}j,g}} = \left(1+\beta\frac{N-2}{N-1}\right)\left(\bar{\alpha}_{\text{-}j,g} - \sum_{i=1}^{G}\pi_i\bar{\alpha}_{\text{-}j,i}\right),$$

and observe that the α_j drops out as claimed.

This leaves us with the fixed-effects regression:

$$d_{jc} - \overline{d}_{jc} = \theta\left(1+\beta\frac{N-2}{N-1}\right)\left(\bar{\alpha}_{\text{-}j,g} - \sum_{i=1}^{G}\pi_i\bar{\alpha}_{\text{-}j,i}\right) + \left(\omega_{jc} - \bar{\omega}_{jc}\right), \tag{8}$$

where the first parenthetical term on the right-hand side is the *attenuation factor* and the second parenthetical terms is the “correct” regressor.

In large samples we obtain

$$\beta = \theta\left(1+\beta\frac{N-2}{N-1}\right),$$

and thus

$$\theta = \beta/\left(1+\beta\frac{N-2}{N-1}\right). \tag{9}$$

Therefore θ is consistent for β if $\beta = 0$, it is *attenuated* if $\beta > 0$ and it is *exaggerated* if $\beta < 0$. Note that tests for the existence of peer effects will still be consistent, as $\beta = 0$ under the null (see Wooldridge (2010, pp.158-160), where in his notation, $G = 0$ so 2SLS standard errors and test statistics are valid).

Note that in our case $N = 9$. So for example, if the true parameter β were equal to 0.5 then, in large samples, the $plim_{n\to\infty}$ of the fixed effects estimator of θ would be 0.348 (as the attenuation factor is 1.44). This illustrates the sense in which our fixed effects estimates of peer effects are conservative.

In the above example, it is assumed that justices are matched with different groups of peers over time, but we assume nothing about the degree of similarity (or overlap) between those groups. In the actual data, there is of course substantial continuity of the court (and hence in the composition of peers) over time. We have conducted a Monte Carlo analysis where we find that including such continuity in peer group composition reduces the attenuation bias below the level implied by (9).

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