The Separation of Powers, Court Curbing, and Judicial Legitimacy

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A major focus of judicial politics research has been the extent to which ideological divergence between the Court and Congress can explain variation in Supreme Court decision making. However, conflicting theoretical and empirical findings have given rise to a significant discrepancy in the scholarship. Building on evidence from interviews with Supreme Court justices and former law clerks, I develop a formal model of judicial-congressional relations that incorporates judicial preferences for institutional legitimacy and the role of public opinion in congressional hostility towards the Supreme Court. An original dataset identifying all Court-curbing legislation proposed between 1877 and 2006 is then used to assess the influence of congressional hostility on the Court's use of judicial review. The evidence indicates that public discontent with the Court, as mediated through congressional hostility, creates an incentive for the Court to exercise self-restraint. When Congress is hostile, the Court uses judicial review to invalidate Acts of Congress less frequently than when Congress is not hostile towards the Court.

During each of the two recent Supreme Court confirmation hearings in the United States Senate, the Judiciary Committee noted its concern with what it perceived to be an overly aggressive Supreme Court and an imbalance in the separation of powers. Chairman Arlen Specter opened the confirmation hearings of John Roberts by commenting, "I'm very much concerned about what I conceive to be an imbalance in the separation-of-powers between the Congress and the court. I am concerned about what I bluntly say is the denigration by the court of congressional authority."[1] Senator Specter's comments highlight a claim about American democracy which has been at the center of a lively academic debate for decades: the Supreme Court is an insulated legal body, free to make decisions away from the political pressures of Congress and the Executive.

In this article, I challenge that claim. First, relying on evidence from interviews with Supreme Court justices and former law clerks, I establish a set of assumptions about judicial preferences that departs from previous separation-of-powers models. Second, I formalize an interaction between the Court and Congress that rests on those assumptions. The formalization yields empirically testable comparative statics. Third, I test those predictions with an original dataset identifying every Court-curbing bill introduced in Congress since Reconstruction. The analysis here suggests a new interpretation of the separation-of-powers mechanism that drives congressional constraints on judicial decision making.

The article proceeds as follows: the first section briefly reviews the existing debate concerning the separation-of-powers model and the literature on Court-curbing; the second section establishes assumptions about judicial and legislative preferences; the third section then uses that theory of preferences to develop a formal model of congressional-judicial relations which yields empirically testable predictions; the fourth section presents an empirical analysis of an original dataset to test the hypothesis; and the last section discusses the results and offers some concluding remarks.

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[1]United States Senate Confirmation Hearings for John Roberts, Nominee as Chief Justice of the United States, Senate Judiciary Committee, September 12, 2005, Arlen Specter, Opening Statement.

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Separation-of-Powers Theory and Court-Congress Relations

In the study of judicial-congressional relations, two related bodies of literature have asked whether and to what degree Supreme Court decision making is constrained by congressional preferences. One literature has examined the separation-of-powers (SOP) model, which asserts that Congress's power to reverse statutory decisions creates incentives for sophisticated judicial decision making (see, e.g., Epstein and Knight 1998; Ferejohn and Shipan 1990; Marks 1989; Segal and Spaeth 2002, chap. 8). A generation of models has posited that when the Court's ideal policy lies outside of the set of Pareto-optima—the policies immune from congressional reversal—then the Court will have an incentive to choose strategically the best policy from among those Congress cannot reverse. In the context of statutory construction the empirical support for this model has been mixed at best (Bergara, Richman, and Spiller 2003; Eskridge 1991; Gely and Spiller 1990; Martin 2005; Segal 1997; Spriggs and Hansford 2001). On the other hand, there is some evidence that the SOP model may be applicable to the Court's constitutional decisions (Dahl 1957; Meernik and Ignagni 1997), and some case study–based analyses stand out as the primary examples (Clinton 1994; Gely and Spiller 1992; Knight and Epstein 1996). Recently, scholars have begun to examine more closely the extent to which the SOP model extends to constitutional interpretation and have found mixed evidence of the SOP mechanism at work in constitutional decisions (Epstein, Knight, and Martin 2001; Harvey and Friedman 2006; Sala and Spriggs 2004; Segal and Westerland 2005; Segal, Westerland, and Lindquist 2007; Spriggs and Hansford 2001).

A second body of literature has asked whether legislative attempts to limit judicial power have brought about substantive changes in judicial decision making. This literature distinguishes Court curbing from statutory reversals. Studies of Court curbing by Congress have attempted to identify specific periods of Court-curbing activity and to demonstrate that these actions by Congress lead to reversals in judicial policy (Handberg and Hill 1980; Hansford and Damore 2000; Nagel 1965; Rosenberg 1992; Stumpf 1965). These studies have generally found that periods of Court curbing are followed by marked periods of judicial deference to legislative preferences and posit that the risk of actual changes to the Court's institutional power—through jurisdiction stripping, Court packing, or other legislative means—will create an incentive for sophisticated decision making by the Court. However, no systematic study of this mechanism has yet appeared, perhaps because the incidence of significant risk of fundamental changes to the Court's power has been historically rare.[2]

I unite these two bodies of research to propose an alternative interpretation of separation-of-powers theory. In particular, while previous research has recognized the link between Congress's power to restrict judicial power and the separation-of-powers model (Baum 2006, 72–91), I argue that Court curbing in Congress may affect judicial decision making independent of any threat of enactment. I argue that Court curbing can affect judicial independence because it can be a credible signal about waning judicial legitimacy, which is essential for the efficacy of an independent judiciary. Because Congress is more directly connected to the public than the Court, observing institutional signals such as Court curbing can help solve an informational problem confronting a Court concerned about its standing with the public.

Assumptions about Institutional Preferences

The theory of judicial-congressional relations that I develop here rests on a set of assumptions about institutional preferences that departs in several important ways from previous scholarship. In this theory, the Court has preferences for both policy and institutional legitimacy, while Congress has both policy and position-taking preferences. I establish these assumptions with evidence from existing literature as well as personal interviews with three Supreme Court justices and 10 former law clerks, conducted between November 2006 and March 2007.[3]

Judicial Preferences

That the Court has preferences over policy outcomes is an uncontroversial claim resting on a large body of empirical scholarship, the paradigmatic example of which is Segal and Spaeth (2002). Moreover, that courts have preferences for institutional legitimacy is similarly well demonstrated in the literature (Baum 2006; Caldeira 1987; Caldeira and Gibson 1992, 1995; Carrubba 2009; Hausseger and Baum 1999; Lasser 1988; Rogers 2001; Staton 2006;

[2]This dearth of research may also be due to the difficulty of acquiring data on Court curbing (Segal, Westerland, and Lindquist 2007).

[3]As a condition of the interviews, I guaranteed these individuals anonymity. Details available upon request.

Stephenson 2004; Vanberg 2005). This study departs from the literature on institutional legitimacy in an important way—I argue that congressional hostility towards, and attacks on, the judiciary indicate a lack of judicial legitimacy and public prestige. In particular, the justices believe that legislative attacks on the Court are signals about a lack of public support for the Court. Thus, while the justices have their own information about public opinion and the Court, they can, and do, update their beliefs by observing political activity concerning the Court. In an interview, one Supreme Court justice commented, "The Court is pretty good about knowing how far it can go. . . . Congress is better than we are, especially the House. They really have their finger on the pulse of the public." Similarly, another justice commented,"We read the newspapers and see what is being said—probably more than most people do. . . . We know if there is a lot of public interest; we have to be careful not to reach too far," a sentiment echoed by numerous other Court insiders.

Further, considerable evidence demonstrates that the Court is concerned about political criticism. Research on the importance of institutional legitimacy for judicial power provides evidence that the Court is sensitive to how it is perceived by the public and members of the bar (Baum 2006; see also Epstein and Knight 1998, chap. 5; Klein and Morrisroe 1999; Staton 2006), while other scholarship demonstrates that the Court has an incentive to protect its institutional legitimacy by avoiding institutional confrontations and acts on that incentive (Caldeira 1987; Carrubba 2009; Hausseger and Baum 1999; Lasser 1988; Stephenson 2004; Vanberg 2005; see also Marshall 1989, 2004).

The scholarly literature distinguishes between *diffuse* support and *specific* support for the Court. Whereas diffuse support refers to broad institutional support for the Court as an institution (possibly despite unpopular rulings), specific support refers to public support for a particular decision. Court curbing may simply be a signal of a loss of specific support, but that is important information for the Court, because continued losses of specific support may have a deleterious effect on the Court's diffuse support in the aggregate. One Supreme Court justice commented, "Once the public ceases to believe that the Court is not a political institution, they will no longer support the Court." Another justice observed that the Court's being "perceived as acting legitimately. . .[is] predicated on whether the public understands that we are a court and act [in] a legitimate way." Indeed, the scholarly literature similarly shows that when public support for the Court declines, the public will increasingly support efforts to politically sanction the Court and restrict judicial power (Caldeira and Gibson 1995; Gibson and Caldeira 1995, 1998, 2003; Gibson, Caldeira, and Baird 1998). For example, Caldeira and Gibson claim that individuals who have no diffuse support, or institutional loyalty, for the Court will be willing to "accept, make, or countenance major changes in the fundamental attributes of how the high bench functions or fits into the U.S. constitutional system" (1992, 638). Using the rubric of "rational anticipation," McGuire and Stimson suggest "a Court that strays too far from the broad boundaries imposed by public mood risks having its decisions rejected" (2004, 1019). Mondak and Smithey summarize the point nicely: "A disgruntled public may not only refuse to cooperate with a Supreme Court decision but may also pressure elected officials to resist implementation of judicial orders" (1997, 1114). That is, the judiciary is given no positive powers and depends heavily upon political will to give effect to its decisions. The Court is therefore faced with an implementation problem. Scholars of the courts cite diffuse support as a resource necessary for overcoming this implementation problem (Caldeira 1986; Murphy and Tanenhaus 1990; Stephenson 2004; Weingast 1997; see also Carrubba 2009).

As such, despite the Supreme Court's nominal insulation from the American people, the justices have strong incentives to be concerned with their public standing. They recognize that erosion of public support and institutional legitimacy has negative consequences for the Court's power and institutional integrity. The justices themselves corroborate the claim that a loss of public support leads to an erosion of institutional legitimacy that negatively affects the Court's efficacy as a governing institution. Speaking at a conference on judicial independence in 2003, former Chief Justice William Rehnquist (2003) noted that past preservation of the independence and integrity of the Court has been "dependent upon the public's respect for the judiciary" and that "[t]he degree to which that independence will be preserved will depend again in some measure on the public's respect for the judiciary."

Indeed, historical examples suggest the Court does at times recognize the limits of its independence and exercises self-restraint for fear of acting without public support and inflicting irreparable institutional damage. Notable examples include the Supreme Court's reluctance to consider the constitutionality of antimiscegenation laws in the wake of *Brown v. Board of Education* (Klarman 2004, 321) and its continued reluctance to address widespread prayer in public schools, despite the Court's declaration that such practices violate the constitution. This point is made generally by Lasser (1988), who argues that the historical pattern has in fact been one of judicial self-restraint precisely at those times when it is aware that the political situation is too perilous. For my purposes here, though, it is important not that the Court *has* at any point lost public support but rather that the justices behave *in anticipation* of a lack of public support. The historical record suggests the Court has at times been reluctant to forge ahead with its policy agenda for fear of acting outside of the broad contours of public support.

Thus, while justices have preferences over policy outcomes, they also have a preference for institutional legitimacy. Importantly, members of the Court believe that political attacks on the judiciary evince an erosion of public support and a decline in the Court's institutional efficacy. For this reason, *political attacks on the Court serve as signals of a lack of specific support for the Court, which in turn indicates that further judicial recalcitrance will not be tolerated and that the Court will not be able to effectively set policy*. The mechanisms by which this may take place are several, but, primarily, continued judicial recalcitrance could lead to the impeachment of justices, the reluctance of lower court judges to heed Supreme Court precedent, or the refusal of elected officials to implement judicial decisions. For example, for fear of electoral reprisal, an elected official would find it in her interest to disregard a judicial decision perceived as illegitimate by the public.

Legislative Preferences

Congress has preferences for both policy outcomes and position taking. A policy outcome refers to the Court's decision in a case. That policy concerns are a primary motivation for legislators is a claim deeply rooted in legislative scholarship (Fenno 1973). Position taking, by contrast, refers to official, observable activities by members of Congress that reflect their constituents' opinions. Because the public can hold legislators accountable for misrepresenting their preferences (Arnold 1990, 56–57; Canes-Wrone, Brady, and Cogan 2002), members of Congress will generally have an interest in correctly taking position in line with public opinion, which is a central activity in the pursuit of reelection (Mayhew 1974). I assume members of Congress have an interest in position taking in order to create a public record to which they can point to demonstrate they have taken some action to secure the goals of their constituents. It does not matter whether those actions lead directly to policy changes, because constituents may not hold their members of Congress accountable for policy outcomes but rather blame Congress as an institution for failed policy initiatives (Parker and Davidson 1979). Thus, an important consideration for members of Congress is to take public positions that are visible and popular with their constituents (Arnold 1990, chap. 4).

Engaging in Court curbing and other political attacks on the Court can be reasonably considered a position-taking endeavor, because it is an effective way to help build support from an issue constituency. Interest groups concerned with the judiciary and its role in American politics closely monitor legislative activity concerning the Courts and draw their supporters' attention to legislators' actions and positions. Indeed, previous scholarship demonstrates that constituent preferences are a primary determinant of legislative responses to, and attacks on, the judiciary (Clark and McGuire 1996). Major Supreme Court decisions have even occupied central positions in presidential campaigns (Stephenson 1999). What is more, the sheer number of Court-curbing proposals that are introduced in Congress with great regularity, but never earn so much as a committee hearing, suggests that Court curbing is driven at least in part by interested contingents or groups from a member's constituency. Finally, I note a positive correlation between the number of Court-curbing bills introduced and negative public opinion about the Court.[4]

I assume that there is a potential trade-off between legislative electoral interests in position taking and legislative policy preferences. That is, when the public supports the Court, Congress may have to balance its interest in accurately representing public preferences and its own interests in seeing its preferred policy realized.

The Model

In this section, I develop a formal model of legislative-judicial interactions that rests on the assumptions described above. Specifically, the model relies on two assumptions. First, making a decision that is politically and publicly rejected due to a lack of public support for the judiciary is costly for the Court; second, failure to represent constituency preferences has negative political consequences for the legislature.

[4]I have assessed this relationship in several ways. First, I compare the number of bills introduced in Congress with the few public opinion polls that do exist. For each, I find a positive correlation. Second, I performed a factor analysis on a series of variables thought to be related to public disapproval of the Court—ideological divergence between the Court and the public, reversal of state court decisions, precedents overturned, closely divided decisions—and found a positive correlation between the number of Court-curbing bills and the two primary factors derived from that analysis.

Elements of the Model

Players and Sequence of Play. In the model, there is a Supreme Court, J, and a Congress, C. At the beginning of the game, Nature selects a state of the world, $\Omega \in \{H, L\}$, which is observed only by Congress. A "high" state, $\Omega = H$, indicates a state of the world in which an unconstrained decision by the Court will be rejected by political and public actors; a "low" state, $\Omega = L$, indicates a state where an unconstrained decision by the Court will be accepted by political and public actors. Next, Congress must choose a signal about the state of the world, $\omega \in \{h, l\}$. A signal h represents congressional attacks on the Court. The Court must then make a decision, $d \in \{u, c\}$, where u represents an "unconstrained" decision and c represents a "constrained" decision. After the Court makes its decision, the state of the world is revealed and payoffs accrue.

Beliefs. The Court's uncertainty about the state of the world is characterized by a belief, $\Pr(\Omega = H) = p$. Upon observing Congress's signal about the state of the world, the Court updates its beliefs according to Bayes' Rule.

Payoffs. Ceteris paribus, the Court prefers to make an unconstrained decision rather than a constrained decision. However, if the state of the world is $\Omega = H$, then the Court prefers not to make an unconstrained decision. In particular, if it plays $d = u$, the Court receives b_j if $\Omega = L$ and $-b_j$ if $\Omega = H$. If the Court plays $d = c$, then it receives 0. The fact that these payoffs are symmetric does not affect the analysis that follows.

Congress, on the other hand, has preferences over both policy outcomes and position taking. In particular, Congress receives $-b_c$ if the Court plays $d = u$ and $+b_c$ if the Court plays $d = c$; the parameter b_c represents the degree of ideological divergence between Congress and the Court and, therefore, the policy utility associated with receiving one or the other's preferred policy. In addition, Congress receives $-\varepsilon$ if $\omega \neq \Omega$ and $+\varepsilon$ if $\omega = \Omega$. This dimension of Congress's payoffs represents congressional preferences for representing constituent interests. Attacking a popular Court or refusing to do something about an unpopular Court may have significant electoral consequences for a legislator.[5]

[5]In analyzing the model below, I assume $b_c < 2\varepsilon$. Relaxing this assumption does give rise to an additional pooling equilibrium, but this additional equilibrium does not affect any of the implications derived from the model.

Strategies. A strategy for Congress is a mapping from the state space into a signal, $\omega : \Omega \rightarrow \{h, l\}$. A strategy for the Court is a mapping from its prior belief about the state of the world and the signal it receives into a decision, $d : [0, 1] \times \{h, l\} \rightarrow \{u, c\}$.

Analysis

To analyze the model, I seek perfect Bayesian equilibria. In particular, I consider three types of equilibria that characterize equilibrium behavior for all values of the model's parameters. From these equilibrium types, I derive predictions about the relationship between judicial and congressional behavior. Proofs of formal results are gathered in the appendix.

First, consider the case where the electoral benefit is greater than the policy benefit from a constrained decision. Under this condition, an equilibrium can be supported where Congress perfectly represents the state of the world. In particular, when $\varepsilon \geq b_c$, there exists a separating equilibrium in which Congress sends a high signal whenever the public has lost confidence in the Court and sends a low signal whenever the public has confidence in the Court. The Court makes a constrained decision if it observes a high signal and an unconstrained decision if it observes a low signal.

Proposition 1. *When $b_c \leq \varepsilon$, there exists a unique perfect Bayesian equilibrium in which Congress plays $\omega = h$ if $\Omega = H$ and $\omega = l$ if $\Omega = L$, and the Court makes an unconstrained decision if and only if it observes a low signal from Congress and a constrained decision otherwise.*

The separating equilibrium indicates that for sufficiently high electoral incentives, Congress will always accurately reveal the level of public support for the Court, even though it would be able to receive a better policy outcome if it were to deviate, and the Court will be able to perfectly update its belief about the state of the world. When Congress signals that the Court has lost public confidence, the Court will make a constrained decision, and when Congress signals that the Court has not lost public confidence, the Court will make an unconstrained decision.

On the other hand, when Congress's benefit from a constrained decision by the Court is greater than the electoral benefit associated with representing constituents' preferences, then the type of equilibrium that can be supported will depend on the Court's prior belief. First, suppose the Court has a sufficiently high prior belief—that it believes it is more likely than not that it has lost public support ($p > \frac{1}{2}$). When this is the case, then there will exist a pooling equilibrium in which Congress always sends a high signal and the Court always makes a constrained decision.

Proposition 2. *When $b_c > \varepsilon$ and $p > \frac{1}{2}$, there exists a unique perfect Bayesian equilibrium in which Congress always sends a high signal, $\omega = h$, and the Court always makes a constrained decision.*

This equilibrium captures an interesting and striking dynamic. When the Court believes *ex ante* that it is sufficiently likely that the public will reject an unconstrained decision, then there will be an incentive for the Court to make a constrained decision upon observing a high signal, even though it knows Congress may be bluffing and falsely representing a high state. However, the risk that the Congress is misrepresenting public opinion is not great enough to justify the Court making an unconstrained decision. Because it knows this, Congress will always prefer to misrepresent a low state and signal a lack of public support for the Court. The key to this equilibrium is that the Court's prior is sufficiently high—when the prior belief drops below the critical threshold ($p = \frac{1}{2}$), then this equilibrium will not be sustainable, and there will exist a hybrid equilibrium.

In this semiseparating equilibrium, Congress sends a high signal whenever the public does not support the Court but also sends a high signal with some probability $q(p)^*$ when the Court has public confidence. The specific probability with which Congress bluffs, $q(p)^*$, is increasing in p—that is, the more likely the Court thinks it has lost public confidence, the higher the probability with which Congress can bluff in equilibrium. The Court makes a constrained decision probabilistically whenever it observes a high signal and makes an unconstrained decision whenever it observes a low signal.

Proposition 3. *When $b_c > \varepsilon$ and $p \leq \frac{1}{2}$, there exists a unique perfect Bayesian equilibrium in which Congress sends a high signal whenever the Court has lost public confidence and sends a high signal with probability $q(p)^* = \frac{p}{1-p}$ when the Court has public confidence and sends a low signal otherwise. Upon observing a high signal, the Court makes a constrained decision with probability $m = \frac{2\varepsilon - b_c}{b_c}$ and makes an unconstrained decision otherwise.*

This equilibrium reveals a dynamic that compliments the pooling equilibrium characterized by Proposition 2. In particular, in the pooling equilibrium, the Court was willing to constrain itself upon observing a high signal, even though it believes that there is a very high chance that Congress is bluffing and misrepresenting the true state of the world. That equilibrium, it was shown, can only be supported when the Court has a sufficiently pessimistic belief about its public support.

On the other hand, when the Court's prior belief falls below that critical threshold, the pooling equilibrium cannot be supported. Congress cannot bluff with probability 1 when the public supports the Court; instead, Congress must be honest about the state of the world with some positive probability. This behavior characterizes a hybrid, semi-separating equilibrium. Under this condition, the Court will be willing to probabilistically constrain itself whenever it observes a high signal. However, in order to maintain that behavior, it must be the case that Congress is willing to be sufficiently honest about the state of the world. That is, Congress must bluff with a low enough probability that the Court is willing to believe Congress or at least expect that the probability it is being misled is small enough, relative to the risk of not believing Congress and losing public support.

Interpretation and Comparative Statics

The three equilibria characterized above provide a framework for thinking about the conditions that give rise to constrained decision making by the Court. Moreover, the analysis demonstrates that, under a wide range of conditions, Congress will be (at least somewhat) honest about the nature of public support for the Court—there can be (at least semi-) separation. This result is not necessarily obvious; indeed it is notable that Congress sometimes will be willing to accept divergent policy decisions in order to establish an expectation by the Court that a hostile congressional signal is a credible indicator about the state of public opinion. The hybrid equilibrium demonstrates that this choice can be supported in equilibrium even if the policy benefit associated with constrained judicial decision making is greater than the electoral benefit from position taking.

The empirical interpretation I give to congressional signals here is the introduction of Court-curbing bills in Congress. In the context of the theoretical model, the signal must have two features. First, the signal must be publicly observable, because this is the way in which legislators derive electoral benefit from "position taking." If it were not publicly observable, then the informative value of the signal in the model would disappear, because the justices could not place any credibility in Court curbing as an indicator of public opinion. Court curbing is a public activity, and sponsoring legislation is a key way in which legislators publicly position-take. Second, the Court must think of the activity as a signal about public support. The evidence presented above about judicial interpretation of

Court curbing, the extent to which the Court pays attention to these bills, and the correlation between them and public opinion all suggest the Court does in fact consider these bills indicators of its standing with the public. To be sure, while I operationalize congressional signals with Court curbing, I leave open the possibility that other possible signals may exist.

From the model I derive a set of empirical predictions. First, the model yields a prediction about the Court's responsiveness to Court curbing. In each equilibrium, the Court will be weakly more likely to make a constrained decision when it observes a high signal—i.e., Court curbing—than when it observes a low signal—i.e., no Court curbing.

H1: The Court should be (weakly) more likely to make a constrained decision in the presence of Court curbing than in the absence of Court curbing.

Second, the model also predicts that the Court should be weakly more likely to make a constrained decision as its prior expectation that it lacks public support increases.

H2: The Court should be (weakly) more likely to make a constrained decision as its prior belief that it has lost public support increases.

Third, the model yields predictions concerning movement across equilibria. Specifically, the movement across the equilibria suggests two interactive relationships among the model's parameters. First, as ideological divergence between the Court and Congress increases, the effect of Court-curbing bills on the Court should (weakly) decrease. That is because the Court will be (weakly) more likely to be in the hybrid equilibrium, in which it sometimes disregards Court-curbing signals.[6]

H3: The constraining effect of Court curbing on the Court should (weakly) decrease as the degree of policy divergence between the Court and Congress increases.

It is noteworthy that this interactive prediction runs against the intuition that follows from an alternative theory in which Court curbing is constraining because of the threat of enactment. If the Court is worried about the enactment of Court curbing, then one might expect that the constraining effect of the legislation would actually *increase* when there is greater ideological divergence between the institutions.

[6]As noted above, I assume that $b_c \geq 2\varepsilon$. If this assumption is relaxed, there exists a pooling equilibrium when $b_c > 2\varepsilon$. In that equilibrium, Congress always introduces Court curbing and the Court always makes an unconstrained decision. The existence of that equilibrium is fully compatible with the comparative statics and this hypothesis.

Finally, the effect of Court curbing on the Court should (weakly) increase as the Court becomes increasingly pessimistic about its level of public support. This is because the Court will be (weakly) more likely to find itself in the pooling equilibrium as its prior belief that it has lost public support increases. For any level of policy divergence between the Court and Congress, increases in that prior belief (weakly) increase the probability that the Court will make a constrained decision upon observing Court curbing.

H4: The constraining effect of Court curbing on the Court should (weakly) increase as the Court's prior belief about its public support becomes more pessimistic.

Empirical Analysis

The empirical analysis proceeds in two stages. To test the theoretical model's predictions, one would ideally employ an operationalization of the model's exogenous parameters—the Court's prior belief and the policy benefit associated with judicial review—to identify which observations occur in which equilibria. However, such measures are not available. I therefore present, first, an empirical analysis of the model's predictions that hold across the equilibria. Second, I present an analysis that considers the interactive relationships concerning movement across the equilibria.

Data

Dependent Variable. The theoretical model is one of "constrained" versus "unconstrained" judicial decision making. One might interpret this concept in any one of a number of contexts, such as constitutional or statutory decison making. I present an analysis of constitutional decision making because it is in this area that the Court's concern for institutional legitimacy might be most significant. Following Segal and Westerland (2005, 1339–41), I assert that a more constrained Court will invalidate fewer federal laws.[7] This operationalization implicitly assumes that constitutional invalidations displease Congress. While there may be instances when Congress supports the use of judicial review, there are several reasons why this assumption is justified. First, constitutional

[7]The latent independent variable of interest here is the level of judicial independence. My empirical measure assumes only that the use of judicial review to invalidate Acts of Congress will be correlated with judicial independence.

invalidations come disproportionately from recently enacted laws by ideologically divergent majorities (Clark and Whittington 2007). Second, the use of judicial review to invalidate an Act of Congress underscores judicial power vis-à-vis Congress, the institutional implications of which have significant theoretical ramifications. Third, a popular majority—the source of the Court's concern for its legitimacy—could most likely undermine any statute with which it disagreed. That is, a constitutional challenge to a law requires a case, which requires support for the law by the political majority (Congress must fund and the president must enforce the law). One might expect, then, that these cases generally involve laws that have at least some public support. Therefore, while Congress may not oppose the invalidation of a statute, given a majority could undermine or reverse the policy on its own, it would most likely prefer to use its own powers rather than concede the power to the Court to do so and accept whatever policy the Court may choose instead. After all, a constitutional decision by the Court cannot be easily overturned.

One might be concerned that a more appropriate dependent variable would be the *proportion* of laws struck down rather than the raw number (or, alternatively, the decision in a specific case). There are, however, complications with such a measure. First, the decision to decide a case on constitutional grounds may itself be a strategic one. That is, the observation of laws upheld might be affected by Court curbing. This effect would make interpretation of such a measure unclear. Indeed, the model predicts that the Court will exercise self-restraint in the face of institutional signals about waning judicial legitimacy. This may not necessarily mean the Court will uphold laws that it would otherwise strike down; rather it may simply be that the Court declines to reach constitutional questions. If this is the case, then the appropriate dependent variable is indeed the number of laws struck down rather than the proportion of constitutional cases in which the Court strikes a law down. Second, there is no dataset of laws upheld by the Court, and what exactly constitutes a constitutional "uphold" is not self-evident (Clark and Whittington 2007; Whittington 2005). Despite these limitations, though, the primary results reported here hold when one uses all laws decided on constitutional grounds (as coded by Spaeth 2008) to calculate a proportion of laws struck down.

To identify the number of laws held unconstitutional each year, I reference a dataset recently compiled by Keith Whittington (2005; Clark and Whittington 2007). Whittington's count of the number of laws held unconstitutional each year is positively correlated with the list maintained by the Congressional Research Service ("CRS list") ($r = 0.77$), though there are important discrepancies between the two lists. Specifically, over 75 years, the CRS list identifies at least one fewer case annually than the Whittington list. The number of laws invalidated each year is shown below in Figure 1.

Independent Variables. To test the empirical hypotheses derived above, I model the number of federal laws held unconstitutional each year as a function of the level of Court curbing in Congress, the degree of ideological divergence between the Court and Congress, and a set of controls.

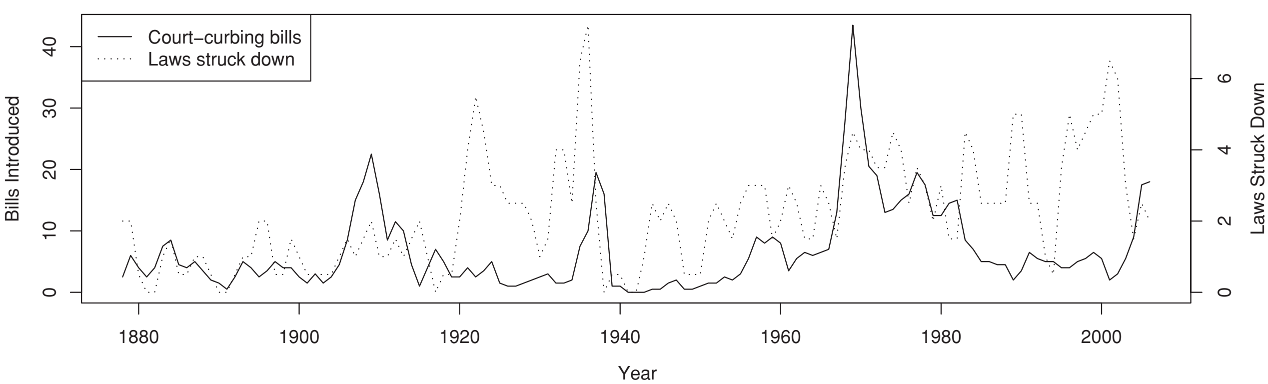
Court curbing. Scholars interested in hostility between Congress and the Court have suffered from the unavailability of data on Court curbing. Segal, Westerland, and Lindquist note that "data on court curbing legislation proposed. . .is difficult to acquire" (2007, 16), which is perhaps part of the reason. Indeed, the one widely cited dataset on these proposals (Nagel 1965; Rosenberg 1992) is no longer available. Moreover, the Nagel/Rosenberg data do not include constitutional amendments or resolutions; these types of legislation are equally important for my purposes, because they represent public discontent with the Court just as much as other types of Court-curbing proposals. Therefore, I assemble a new dataset, identifying all Court-curbing bills introduced during the 130-year period from 1877 through 2006. A Court-curbing bill is defined as a legislative proposal to restrict, remove, or otherwise limit judicial power.

Because of the differing availability of search methodologies across time, three sources were used to identify the relevant legislation. For the period 1877 through 1937, I read the House and Senate journal indices, searching for all bills introduced under a set of exhaustive index terms.[8] For the period 1937 through 1989, I read the *Digest of Public General Bills and Resolutions*, which contains synopses of every bill introduced in Congress. From this source, I identified all bills referred to the Judiciary Committee of either chamber. Finally, for the period 1989 through 2006, I used the online THOMAS search engine to search for all bills indexed under the same exhaustive set of terms used to search the House and Senate journals. I then read each bill identified in each search to determine which bills constitute Court-curbing legislation.[9]

[8]Specifically, I read all bills indexed under "Courts," "Supreme Court," "Judges," "Justices," "Judiciary," "Judicial Power," "Constitution," and "Constitutional Amendments."

[9]A variety of validity checks ensure that there is no systematic bias across search methodologies. In particular, I have performed all three searches for the years 1981 through 1989 and found a 100% agreement rate across the different search techniques.

FIGURE 1 Two-year Moving Average of Court-curbing Bills Introduced in Congress and Federal Laws Invalidated, 1877–2006



The typical Court-curbing bill is what might be characterized as an institutional assault on the Court rather than a case-specific effort to reverse a Court decision. That is, Court-curbing bills are generally wholesale-level responses to (potentially) a series of retail-level problems. Indeed, seen as position-taking endeavors, this empirical finding makes sense. Legislators seeking to garner political support and earn position-taking points with their constituents might find it more "profitable" to introduce a broad bill than a narrowly focused bill directed at a single decision. Indeed, in an interview for this research, one congressman observed that Court curbing could be done with a "sledgehammer" or a "scalpel." The congressman noted that while the "scalpel" might be easier to enact and have more direct policy consequences, the "sledgehammer" gets a lot of public attention, so it is generally used.

Figure 1 shows the two-year moving average of the number of Court-curbing bills introduced over time, from 1877 through 2006. These data suggest there have been several distinct periods of particular hostility towards the Court as well as several distinct periods during which Congress has not been hostile towards the Court. Most recently, there was an increase in the level of congressional hostility towards the Court during the 109th Congress. While a full treatment of the substantive issues that motivated each of the periods of hostility revealed in Figure 1 is beyond the scope of this article, there are several highly intuitive interpretations of each of the periods of Court curbing. For example, Court curbing in the late 1800s and early 1900s was generally focused on labor-related issues; attacks during the late 1930s were largely concerned with economic regulation; and congressional hostility towards the Court during the late 1970s and early 1980s was driven primarily by salient issues related to the welfare state and the "culture war."

The variable $CourtCurbing_t$ is a logistic transformation of the number of bills introduced in Congress each year.[10] This operationalization allows for a nonlinear relationship between the number of Court-curbing bills introduced and the number of laws struck down. However, all of the results reported here are robust to any one of a variety of alternative operationalizations, including a log transformation or no transformation at all. I use the logistic transformation, though, because the theoretical model predicts that the Court should respond differently to "high" signals than it does to "low" signals. The logistic transformation allows for a relationship where the largest effect of the covariate is a change from low levels of Court curbing—i.e., fewer than 10 bills introduced—to high levels of Court curbing—i.e., more than 10 bills introduced.

The Court's outlook. To measure the Court's prior belief about its public support, one would ideally reference a public opinion poll about the Court. Unfortunately, public opinion data about the Court are notoriously sparse. While some data on public support for the Court have been compiled (Caldeira 1987; Durr, Martin, and Wohlbrecht 2000), they are available for a relatively small number of years. In particular, the General Social Survey (GSS) asks a national sample of respondents how much "confidence" they have in "the people running the Supreme Court." This question has been asked regularly since 1973. To measure the Court's level of public support, I use the percent of respondents saying "hardly any"

[10]Specifically, $CourtCurbing_t = 1/(1 + \exp(-\,bills_t/2)) - .5$, where $bills_t$ indicates the number of Court-curbing bills introduced in year t.

each year the question is asked. In the years the question is not asked, I use the average of the preceding and following years, though none of the results reported below depend on this. Unfortunately, because the year is the unit of analysis here, using these data significantly limits the scope of the analysis.

To overcome this limitation, I consider an alternative proxy measure for the Court's outlook about its public support. Durr, Martin, and Wohlbrecht (2000) show that when the Court's decisions become increasingly liberal (conservative) while the public becomes increasingly conservative (liberal), public support for the Court decreases. I adopt the Durr et al. variable, *Divergence*, which is an index of the divergence between the public mood and the ideological distribution of the Court's decisions. I extend their data backward and forward to cover the years 1953–2003. I then use the two-year prior moving average of the variable *Divergence* as a proxy measure for the Court's expectation. This operationalization is helpful because it is a factor that has been demonstrated to be a strong predictor of public support for the Court, is exogenous to the dependent variable of interest here (constitutional invalidations), and allows us to extend the analysis over a longer period of time than the GSS measure. In the analysis, I estimate first a model on the full sample of data excluding a measure of public opinion, second a model using the proxy (divergence) measure on a limited sample for which it is available, and third a model using the direct (GSS) measure, though using the most limited sample—only those years for which the GSS is available.

Ideological divergence. To measure the degree of ideological divergence between the Court and Congress, I identify the partisan affiliation of each justice on the bench during each year (Epstein et al. 2007) as well as the partisan control of each chamber of Congress. I then code *Opposite Party Congress*$_t$ as 1 whenever both chambers of Congress are controlled by one party in year t and a majority of the justices have the opposite party affiliation (otherwise, 0). Of course, other SOP analyses have made use of more fine-grained measures of ideology. However, while such measures exist for Congress over the course of American history, sophisticated estimates of judicial ideology are only available for the modern Supreme Court and therefore cannot be used to study the entire period under investigation here. It is noteworthy, though, that the primary results reported here hold when the empirical model is restricted to the modern era for which these data are available. In particular, replicating the Segal and Westerland (2005) analysis, which uses the Bailey (2007) estimates of congressional and judicial ideology, yields strong, substantively similar results. I present these results in the "Robustness" section below.

Natural court fixed effects. Finally, I include fixed effects for the natural court—the period of time during which the Court's composition does not change. These fixed effects allow me to account for heterogeneity in the number of laws struck down that may be correlated with natural court periods (or other temporal periods within which natural courts may be nested, such as the identity of the chief justice or the median of the Court). They also control for the ideological composition of the bench, in the absence of sophisticated measures of judicial ideology across the period of study and any idiosyncrasies that may be associated with specific eras in the Court's history. Importantly, the inclusion of these fixed effects allows for the nonparametric estimation of a time trend that may exist in the frequency of judicial review over time and therefore help account for changing patterns in the use of judicial review over time. The results reported below are robust to alternative controls, such as fixed effects for the Chief Justice or the identity of the median justice (or exclusion of the fixed effects).

Estimation

The dependent variable in these analyses is the number of laws struck down in a given year and therefore constitute event-count data (i.e., data that can only take on nonnegative integer values), but they are also time-series data. While the primary empirical model used to model event-count data is the Poisson regression model, there is an ongoing debate in the literature about the best way to treat count data that are also time-series data. The primary concern is that the data may exhibit time-series dynamics such as persistence, short-term autocorrelation, or cyclical dynamics. When the data exhibit such time-series dynamics, standard event-count models will not be efficient; when the data, however, do satisfy the assumptions of independence and equidispersion, then standard event-count methods, such as the Poisson regression model, are appropriate (Brandt et al. 2000).

To assess how best to model the data here, I have performed several diagnostic tests. First, data that suffer from serial correlation generally exhibit overdispersion. However, a likelihood-ratio test indicates the data do satisfy the Poisson assumption of equidispersion (results are reported in Table 1). Next, I have evaluated the data to determine whether there is serial correlation. The Portmaneau (Q) statistic indicates that the data are not serially correlated (results are reported in Table 1), and

TABLE 1 Effect of Court Curbing on Frequency of Constitutional Invalidations of Acts of Congress

	GSS Measure	Proxy Measure	No Prior Measure	GSS Measure	Proxy Measure
$CourtCurbing_{t-1}$	−1.53	−1.79	−1.50	−5.00	−1.54
	(0.61)	(0.44)	(0.46)	(2.72)	(0.33)
$CourtOutlook_t$	−0.06	−0.76	–	−0.17	−0.73
	(0.03)	(0.35)		(0.11)	(0.72)
$OppositePartyCongress_t$	0.10	0.26	−0.10	0.35	0.55
	(0.69)	(0.70)	(0.40)	(1.04)	(0.79)
$CourtCurbing_{t-1} \times$	–	–	–	0.28	−0.03
$CourtOutlook_t$				(0.22)	(1.80)
$CourtCurbing_{t-1} \times$	–	–	–	−0.82	−0.70
$OppositePartyCongress_t$				(1.43)	(1.01)
Constant	2.84	2.04	1.94	4.16	1.97
	(0.49)	(0.15)	(0.15)	(1.35)	(0.13)
N	32	50	127	32	50
Overdispersion	$\overline{\chi}^2_{[1]} = 0.00$	$\overline{\chi}^2_{[1]} = 0.00$	$\overline{\chi}^2_{[1]} = 0.00$	$\overline{\chi}^2_{[1]} = 0.00$	$\overline{\chi}^2_{[1]} = 0.00$
Serial Correlation	15.95	32.40	49.53	13.85	32.71
	$p \leq .32$	$p \leq .09$	$p \leq .14$	$p \leq .46$	$p \leq .09$

Notes: Dependent variable is number of Acts of Congress invalidated in year t; Poisson coefficients shown (standard errors clustered on natural court in parentheses); natural court fixed effects not shown; overdispersion test is likelihood-ratio test of null hypothesis that $\alpha = 0$ and the data satisfy equidispersion assumption; serial correlation test is Portmanteau Q-statistic testing null hypothesis that there is no autocorrelation in the data.

further diagnostics indicate that any potential weak serial correlation is fully accounted for by the natural court indicators.[11] Finally, to test the possibility that the dependent variable is a unit root, I performed the Augmented Dickey-Fuller GLS test on the logged number of laws struck down. The test strongly rejects the null hypothesis of a unit root at the 1% level ($Z(t) = 7.67$); therefore a Poisson model is appropriate for the data. Thus, I model the data using an empirical specification as follows:

$$\begin{aligned} LawsStruckDown_t(x_t) = \exp(&\beta_0 + \beta_1 CourtCurbing_{t-1} \\ &+ \beta_2 CourtOutlook_t \\ &+ \beta_3 OppositePartyCongress_t \\ &+ \gamma_t \mathbf{Court}_t) \end{aligned} \tag{1}$$

where $LawsStruckDown_t(x_t)$ is the expected number of laws struck down in year t, conditional on the covariates at year t, and $\mathbf{Court}_t$ is a matrix of indicator variables controlling for the current natural court. To help account for any heterogeneity that may be correlated with the composition of the Court, I estimate this model with standard errors clustered on the natural court.

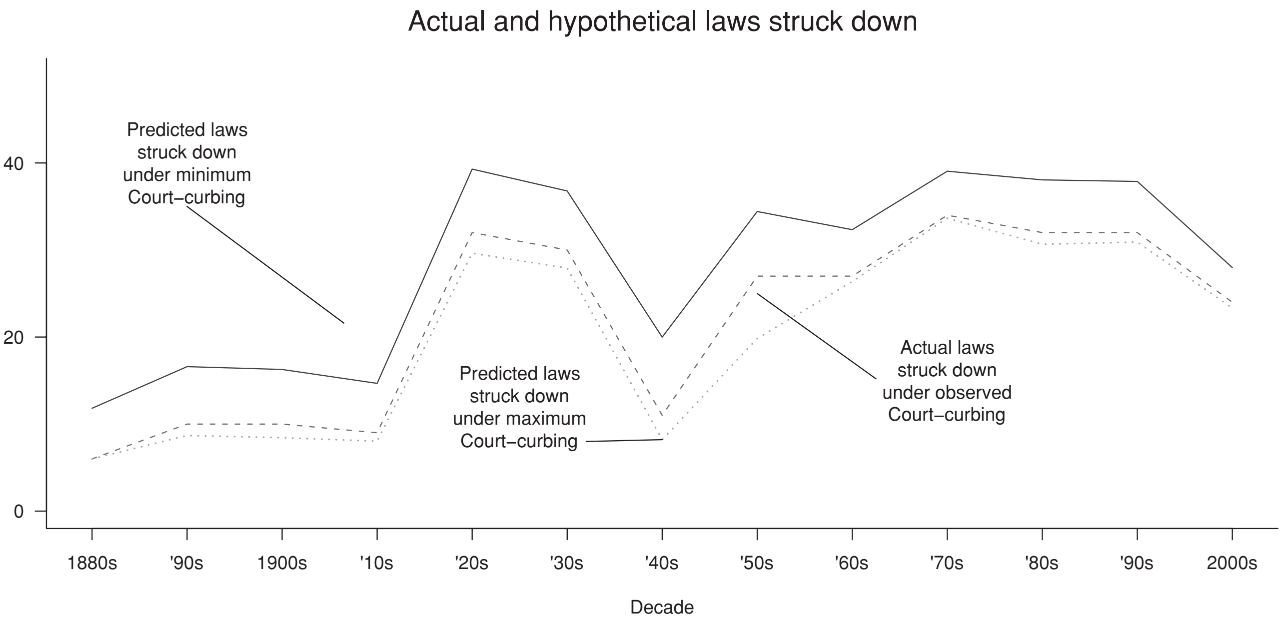
Results

The results of this analysis are reported in the first three columns of coefficients in Table 1. As these results make clear, the number of laws struck down each year is inversely related to Court curbing the previous year. The negative and statistically significant coefficient associated with $CourtCurbing_{t-1}$ indicates that an increase in the level of Court curbing in one year is associated with a decrease in the number of laws held unconstitutional the following year. In addition, the negative and statistically significant coefficient associated with $CourtOutlook_t$ in each model indicates that as the level of public support for the Court decreases, the Court strikes down fewer laws.[12] These two findings provide direct support for Hypotheses 1 and 2 outlined above.

[11]The Q-statistic does suggest there may be some weak serial correlation in the models using the *Divergence* measure of the Court's outlook. However, reestimating the model as a PAR(q) model (Brandt and Williams 2001) indicates there is no statistically significant serial correlation—i.e., $\hat{\rho}$ is not statistically distinguishable from 0.

[12]It is important to note here that the analysis with the GSS measure uses only a very small number of observations. Maximum likelihood estimators, such as that used here, are generally weaker with small samples than more traditional estimators, such as OLS. For this reason, it is important to understand the inherent limitations of maximun likelihood estimates derived from a small sample.

FIGURE 2 Substantive Effect of Court Curbing



Notes: Estimates based on coefficients from empirical model (1). Dashed line shows the actual number of laws struck down; solid line shows the number of laws that would have been struck down with minimum Court curbing; dotted line shows predicted number of laws that would have been struck down with maximum Court curbing.

Indeed, the substantive effect of these results is considerable. Figure 2 shows the actual number of laws held unconstitutional during each decade as well as the predicted number of laws that would have been struck down with minimal Court curbing each year. It also shows the predicted number of laws that would have been struck down under a higher level of Court curbing each year. In particular, the figure shows the predicted number of laws held unconstitutional each decade with $CourtCurbing_{t-1}$ for each year set at the observed minimum and the observed maximum for that decade. This figure shows that Court curbing has deterred between 10 and 20 constitutional invalidations each decade—as many as two invalidations per year in some instances. Given that only about two laws are held unconstitutional each year, this finding is indeed substantial. Between 1877 and 2006, 284 laws were held unconstitutional; the model predicts that an additional 143 laws would have been struck down had there been no Court curbing, an over 50% increase in the frequency of judicial review. By contrast, we see that additional Court curbing in any given decade might not have deterred many more laws; under the observed maximum level of Court curbing each decade, only an additional few laws would have been deterred. This suggests Congress generally uses Court curbing to the fullest extent it can.

Finally, I note briefly that there do not appear to be any systematic patterns in the estimated fixed effects. Most of the estimates are not statistically distinguishable from 0, though there may be a slight upward trend in the point estimates over the long (127-year) period. This potential, but weak, trend further suggests that any possible time-series dynamics are in fact corrected by the inclusion of fixed effects.

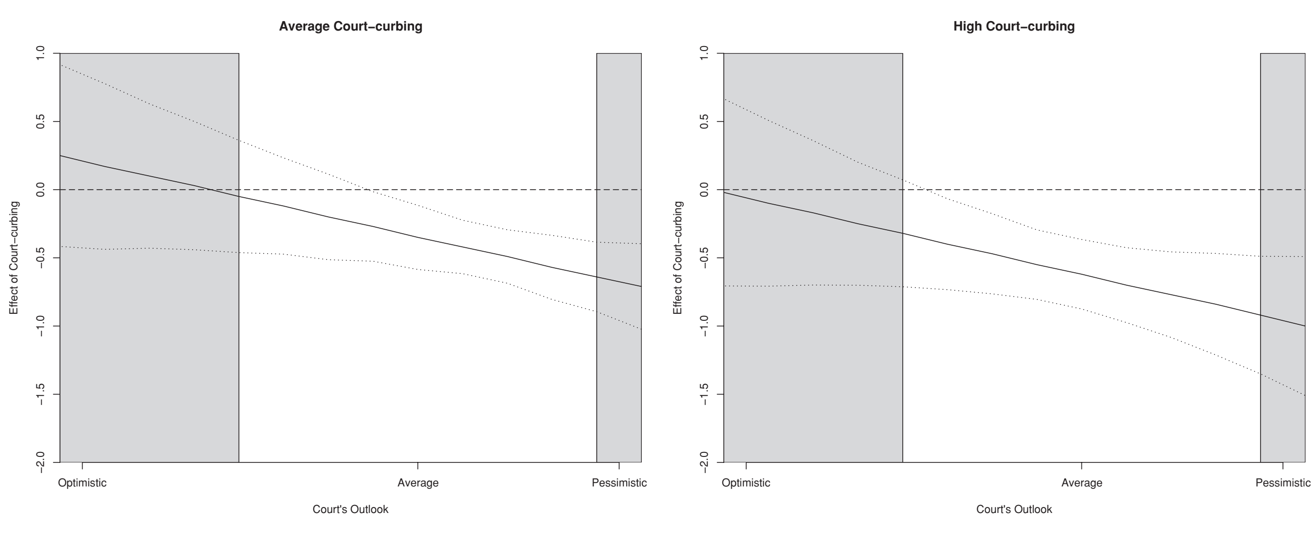
Interactive Model

Estimation. While the preceding results provide evidence in support of the two equilibrium-independent predictions derived from the theoretical model, we can press the data much harder. In particular, we can test the model's interactive predictions to more rigorously assess the theory's explanatory power. In this section, I extend model (1) to capture the interactive relationships posited by the theoretical model. Recall that Hypothesis 3 predicts the constraining effect of Court curbing should weakly decrease as policy divergence between the Court and Congress increases. Thus, I include an interaction between $CourtCurbing_{t-1}$ and $OppositePartyCongress_t$. For the coefficient on this interactive term, we should estimate $\hat{\beta} \geq 0$. By contrast, Hypothesis 4 predicts that the constraining effect of Court curbing should (weakly) increase as the Court's outlook becomes more pessimistic. Thus, I include an interaction between the measures of public support for the Court and $CourtCurbing_{t-1}$. For the coefficient on this interactive term, we should estimate $\hat{\beta} \leq 0$. To help account for any heterogeneity that may be correlated with the composition of the Court, I again estimate this model with standard errors clustered on the natural court.

Results. The results of this estimation are reported in the final two columns of Table 1. First, notice that the main effect of $CourtCurbing_{t-1}$ is negative and statistically significant in both of the interactive models. However, in light of the interactive hypotheses, we must consider the net effect of Court curbing under various levels of the interacted covariates. First, consider the interaction between $CourtCurbing_{t-1}$ and $CourtOutlook_t$. In neither case is the estimated coefficient statistically distinguishable from 0, but in order to properly assess the effect of this interaction, we must evaluate the net effect at various levels of $CourtOutlook_t$. Figure 3 does just this. The left panel here shows the net effect of the average level of Court curbing as a function of the Court's outlook; the right panel shows the net effect of the maximum observed amount of Court curbing as a function of the Court's outlook. As the figures make clear, when the Court is sufficiently optimistic, Court curbing has no statistically significant effect on the use of judicial review. However, as the Court grows increasingly pessimistic, Court curbing begins to exert an increasingly strong and statistically significant effect on the use of judicial review. Indeed, once the Court reaches its "average" level of pessimism, even a small amount of Court curbing can have a statistically significant effect on the use of judicial review. Thus, these data strongly indicate that the effect of Court curbing on the Court is attenuated by an optimistic judicial outlook but is exacerbated by a pessimistic judicial outlook. This finding provides direct support for Hypothesis 3.

Finally, we must consider the effect of Court curbing under the two configurations of Court-Congress alignment. Again, to do this, we must consider the net effect of Court curbing by taking the linear combination of the coefficients from Table 1. When we do this, we find that the net effect of Court curbing when only one Court-curbing bill is introduced is $\hat{\beta}^{SameCongress}_{Bills=1} = -0.15,\ t = 4.70$ when the Court and Congress are aligned and $\hat{\beta}^{OppositeCongress}_{Bills=1} = 0.33,\ t = 0.43$ when the Court and Congress are not aligned. We see a similar pattern when many Court-curbing bills (25) are introduced. In this case we find $\hat{\beta}^{SameCongress}_{Bills=25} = -0.77,\ t = 4.70$ when the Court and Congress are aligned and $\hat{\beta}^{OppositeCongress}_{Bills=25} = -0.57,\ t = 0.84$ when the Court and Congress are not aligned. What does this tell us? Very clearly, this result indicates that the effect of Court curbing on judicial decision making is strongest when the Court and Congress are ideologically aligned and weakest when the Court and Congress are ideologically opposed. This is

FIGURE 3 Net Effect of Court Curbing on Constitutional Invalidations as a Function of $CourtOutlook_t$ (Poisson Coefficient with 95% Confidence Intervals Shown)



Notes: Grey areas show range of $CourtOutlook_t$ more than one standard deviation above or below average; left panel shows effect of average amount of Court curbing; right panel shows effect of maximum amount of Court curbing; effect estimated using coefficients from interactive model with Divergence operationalization of the Court's outlook.

precisely what was predicted by the theoretical model (Hypothesis 4). Because Congress has an increased incentive to "bluff" when it disagrees with the Court over policy, the Court will have a weaker incentive to respond to Court curbing from its ideological opponents than from its ideological allies.

Thus, as predicted by Hypotheses 3 and 4, we see that (1) the effect of Court curbing is exacerbated by a pessimistic outlook with respect to public support for the Court, and (2) the effect of Court curbing is attenuated by ideological divergence between the Court and Congress. These are precisely the interactive relationships predicted by the theoretical model and Hypotheses 3 and 4.

TABLE 2 Replication of Specifications Using Bailey (2007) Continuous Measures of Ideological Divergence between the Court and Congress

	Public Opinion			No Public Opinion		
	β	*s.e.*	$\|z\|$	β	*s.e.*	$\|z\|$
$Court\text{-}Curbing_{t-1}$	−1.32	0.13	10.40	−1.40	0.17	8.07
GSS_t	−0.11	0.05	2.28	–	–	–
$Distance_t$	−3.54	2.04	1.73	−0.74	0.64	1.15
N	30			52		

Note: Dependent variable is the number of federal laws held unconstitutional in year t.

Robustness

Finally, I report the results of two robustness analyses. First, I reestimate the empirical models using the continuous measure of judicial ideology developed by Bailey (2007). However, because this analysis uses continuous measures of ideology, the specification is slightly different. In particular, I specify the empirical model using $Distance_t$ to replace $OppositePartyCongress_t$. $Distance_t$ is the ideological distance between the Court median and the nearest of either the House or Senate medians, in the Bailey space. If the Court median is in between the two chamber medians, $Distance_t$ equals 0.

The primary coefficients of interest from the analyses are reported in Table 2. These results demonstrate the robustness of the findings reported above. All of the coefficients of interest remain signed as expected. Moreover, all of the coefficients remain statistically significant where expected. These results, then, provide evidence that the estimates reported above are robust to alternative specifications using more sophisticated measures of judicial ideology.

Second, I consider an alternative specification. Because the data are time-series data, one might be concerned that the results reported above are dependent on the Poisson model. An alternative approach to dealing with the data is to employ a time-series model and transform the dependent variable. Here, I transform the dependent variable by taking the log of the number of laws struck down (plus 1, because the log of 0 is undefined). Then, I estimate the model specified as a Prais-Winsten regression. A Prais-Winsten regression is similar to an OLS regression, except it allows for serial correlation in the errors. This model is appropriate when the variables are themselves stationary (as they are here) but one suspects possible time-series dynamics, though it is also appropriate in the absence of time-series dynamics. The results of this estimation are reported in Table 3.

We find here further evidence of the robustness of the results reported above. The estimated coefficient for the main effect of $CourtCurbing_{t-1}$ is consistently negative, and in all the models except those using the GSS measure, it remains significant at conventional levels. In the GSS models, however, most likely due to the small number of observations, the estimates become very noisy ($t = 1.03$ in the additive model; $t = 1.54$ in the interactive model). Nevertheless, that the same pattern appears in the data with this alternative specification, which imposes a crude transformation on the dependent variable, provides evidence to remain confident in the robustness of the findings reported above. Even with a less ideal way of modeling the data, we still find evidence of the constraining effect of Court curbing on the Court's use of judicial review.

Discussion and Conclusion

The theoretical model and empirical analyses presented in this article provide a new interpretation of the separation-of-powers model that has been the focus of much scholarship in the area of judicial-congressional relations. The evidence from interviews with Supreme Court justices and former law clerks suggests students of Court-Congress relations must account for the role of judicial legitimacy in the Court's decision calculus. Judicial legitimacy is an important mechanism that drives judicial sensitivity to congressional preferences. Moreover, it can be a

TABLE 3 Effect of Court Curbing on Frequency of Constitutional Invalidations of Acts of Congress

	GSS Measure	Proxy Measure	No Prior Measure	GSS Measure	Proxy Measure
CourtCurbing$_{t-1}$	−0.85	−1.10	−0.98	−5.24	−1.20
	(0.83)	(0.53)	(0.53)	(3.39)	(0.54)
CourtOutlook$_t$	−0.05	−0.52	–	−0.17	−0.39
	(0.02)	(0.35)		(0.12)	(0.78)
OppositePartyCongress$_t$	0.18	0.35	−0.03	−0.05	0.28
	(0.86)	(1.01)	(0.41)	(1.53)	(1.24)
CourtCurbing$_{t-1}$ ×	–	–	–	0.31	−0.29
CourtOutlook$_t$				(0.26)	(2.07)
CourtCurbing$_{t-1}$ ×	–	–	–	0.40	0.16
OppositePartyCongress$_t$				(2.18)	(1.16)
Constant	2.58	1.93	1.89	4.30	1.97
	(0.37)	(0.23)	(0.18)	(1.53)	(0.25)
N	32	50	127	32	50
Serial Correlation	1.97	2.01	2.08	1.92	2.01

Notes: Dependent variable is number of Acts of Congress invalidated in year *t*; Prais-Winsten regression coefficients shown (standard errors clustered on natural court in parentheses); natural court fixed effects not shown; serial correlation statistic is transformed Durbin-Watson Statistic.

condition that gives rise to constrained judicial decision making. Indeed, scholars have long recognized the importance of institutional legitimacy for the Supreme Court (Baum 2006; Caldeira 1987; Caldeira and Gibson 1992; Lasser 1988; see also Staton 2006; Vanberg 2005); however, this study unites this literature with scholarship on congressional constraints on judicial behavior in a previously unappreciated way.

By recasting the separation-of-powers model as a strategic interaction in which responses to Supreme Court decisions are not limited to congressional overrides but also include consequences for the Court's institutional integrity, this model of judicial independence presents a fuller, more nuanced and dynamic interpretation of the judicial decision-making environment. The analysis of judicial-congressional relations as an interaction in which concerns for institutional legitimacy are integral to the Court's decision-calculus unites two important bodies of judicial politics scholarship, and may reorient empirical scholarship that has focused largely on the relative explanatory power of the two dominant models of judicial decision making (the attitudinal model and the SOP model).

As a first step in this empirical direction, the analysis of Court curbing and its relationship to the use of judicial review to invalidate federal legislation provides promising evidence. As the analysis above demonstrates, the relationship between the frequency of judicial review and congressional hostility provides strong, direct support for the theoretical model. When the Court fears it will lose public support, it will adjust its behavior in light of congressional signals about the Court's level of public support. However, the magnitude of that effect is mediated by the political context in which those signals are sent. Instead of responding to Court curbing more strongly when it is facing its ideological opponents, the Court responds most strongly when the Court curbing comes from its ideological allies. Moreover, the constraining effect of Court curbing increases as the Court becomes more pessimistic about its public support. Notably, these interactive relationships run against the intuition following from the conventional wisdom that Court curbing's effect on the Court is due to its threat of enactment. They are, however, predicted by the public-Congress-Court interaction analyzed here.

Nevertheless, there remains more work to be done. Future research can focus more directly on the substance and content of Court-curbing movements. The historical tradition of Court-curbing research has been to focus on specific substantive areas of conflict, but the current research has moved towards a more systematic analysis and as such has not been able to fully treat the nuanced substantive content of Court curbing. In addition, this article has not been able to fully address the determinants of Court curbing. Some preliminary evidence suggests that Court curbing is not necessarily driven by prior instances of federal judicial review (Clark 2008), but such an inquiry is beyond the scope of this study. It is most surely the case, though, that there is more to the decision to introduce Court curbing than the simple strategic signaling game analyzed in this study.

This research has implications for both positive and normative studies of the courts and Congress. On the positive side, future scholarship should further examine ways in which public opinion interacts with the separation of powers and affects the use of checks and balances across the branches of government. A body of literature has been concerned with assessing whether or not public opinion has a direct effect on judicial decision making (Flemming and Wood 1997; Giles, Blackstone, and Vining 2008; McGuire and Stimson 2004; Mishler and Sheehan 1993). This article suggests a framework for thinking about how institutional structures influence the effect of public opinion on the Court. Research can also use the theoretical and empirical groundwork laid here to examine statutory interpretation, which has been at the center of much separation-of-powers scholarship. On the normative side, scholars should consider the implications of the Court's responsiveness to threats to its institutional legitimacy. Judicial self-restraint in the face of institutional signals from democratically elected bodies may potentially ameliorate concerns with the "countermajoritarian" difficulty. It remains an open question the extent to which such an institutional design promotes or hinders a normatively desirable judicial function.

In addition, this study presents a new framework for thinking about Court curbing. Previous studies of Court curbing (Nagel 1965; Rosenberg 1992) have been primarily concerned with the determinants of congressional hostility towards the Court and the conditions under which such legislative proposals are "successful." In this study, I have offered a definition of successful Court curbing that differs from these previous studies; in particular, the definition of success offered here derives from the alternative theory of Court-curbing motivation. Interpreting Court curbing as a primarily position-taking enterprise, rather than an effort to enact legislation, reconceptualizes the role of such activity in the political system. Moreover, this alternative interpretation reframes the theoretical and empirical questions that can motivate studies of Court curbing.

Finally, this study provides some context for interpreting and analyzing contemporary political events. While conservatives have long been the leading force behind legislative hostility towards the Court, the Court has moved in a markedly conservative direction during the Bush administration. Recent media coverage highlights the growing concern among liberals about the Court's jurisprudence, and some early public opinion surveys indicate that some of the Court's conservative decisions during the 2006 term have begun to affect the Court's perception among the public (Barnes and Cohen 2007; Lazarus 2007). The theoretical and empirical analyses presented here provide a framework that can be useful for interpreting these developments and their implications for future judicial-congressional interactions.

Technical Appendix

Proof. *Proof of Proposition 1.* Suppose C plays a separating strategy. J's posterior beliefs are given by $\Pr(\Omega = H \mid \omega = h) = 1$ and $\Pr(\Omega = L \mid \omega = l) = 1$. J's expected utility from playing $d = c$ is 0. J's expected utility from playing $d = u$ upon observing $\omega = h$ is $\Pr(\Omega = H \mid \omega = h) \cdot (-b_c) + (1 - \Pr(\Omega = H \mid \omega = h)) \cdot (b_c) = -\, b_c$, and upon observing $\omega = l$ is $\Pr(\Omega = H \mid \omega = l) \cdot (-b_c) + (1 - \Pr(\Omega = H \mid \omega = l)) \cdot (b_c) = 1$. Therefore, upon observing $\omega = h$, J strictly prefers to play $d = c$, and upon observing $\omega = l$, J strictly prefers to play $d = u$. Now, consider C. If the state of the world is $\Omega = H$, and if C plays $\omega = l$, it expects to receive $-b_c - \varepsilon$; if it plays $\omega = h$, it expects to receive $b_c + \varepsilon$. If the state of the world is $\Omega = L$, and if C plays $\omega = l$, it expects to receive $\varepsilon - b_c$; if C plays $\omega = h$, it expects to receive $b_c - \varepsilon$. Thus, when $b_c \leq \varepsilon$, C has no incentive to deviate from the separating strategy.

For proof of uniqueness, see Clark (2008). ■

Proof. *Proof of Proposition 2.* Suppose that $b_c > \varepsilon$ and $p > \frac{1}{2}$ and that C is playing a pooling strategy. Upon observing $\omega = h$, J's posterior belief is that $\Pr(\Omega = H \mid \omega = h) = p$. Therefore, J's expected utility from playing $d = u$ is given by $(-b_j) \cdot (p) + (b_j) \cdot (1 - p) = b_j \cdot (1 - 2p)$. J's expected utility from playing $d = c$ is 0. $p > \frac{1}{2}$ implies that J prefers to play $d = c$. Now, to see that this is an equilibrium, consider C. C's expected utility from playing $\omega = h$ when $\Omega = H$ is given by $b_c + \varepsilon$; if it plays $\omega = l$, it expects $b_c - \varepsilon$. Therefore, when $\Omega = H$, C prefers to play $\omega = h$. Now, suppose that $\omega = l$. If C plays $\omega = h$, it expects to receive $b_c - \varepsilon$. If it plays $\omega = l$, it expects to receive $\varepsilon - b_c$. $b_c > \varepsilon$ implies that C will strictly prefer to play $\omega = h$ regardless of Ω whenever $b_c > \varepsilon$ and $p > \frac{1}{2}$.

For proof of uniqueness, see Clark (2008). ■

Proof. *Proof of Proposition 3.* Suppose C is playing a semi-separating strategy, in which it sends message $\omega = h$ whenever $\Omega = H$ and plays $\omega = h$ with probability q if $\Omega = L$ and $\omega = l$ otherwise. Further assume that J plays its prescribed strategy,

$$d(\omega) = \begin{cases} u & \textit{if } \omega = l \\ u & \textit{with probability } m = \dfrac{2\varepsilon - b_c}{b_c} \textit{ if } \omega = h \\ c & \textit{else} \end{cases}$$

J's posterior beliefs are given by $\Pr(\Omega = H|\omega = h) = \frac{\Pr(\omega=h|\Omega=H)\cdot\Pr(\Omega=H)}{\Pr(\omega=h)} = \frac{p}{p+q\cdot(1-p)}$. Therefore, $EU_J(d = u|\omega = h; q) = b_j \cdot (1 - \frac{2p}{p+q\cdot(1-p)})$ and $EU_J(d = c) = 0$. Thus, J is indifferent between $d = u$ and $d = c$ when $q(p) = \frac{p}{1-p}$. Notice that for $q \in (0, 1)$, it must be the case that $p \leq \frac{1}{2}$, which is true by assumption.

Now, to see that $q = \frac{p}{1-p}$ is an optimal strategy for C, observe that C is indifferent between playing its prescribed mixing strategy and another pure strategy. First, C may deviate and play $\omega = h$ when $\Omega = L$. C's expected payoff then is $m \cdot b_c - \varepsilon$. If it plays $q = \frac{p}{1-p}$, then its expected utility is $q \cdot ((1 - m) \cdot b_c - m \cdot b_c - \varepsilon) + (1 - q) \cdot (\varepsilon - b_c)$. Therefore, C prefers to play its prescribed strategy rather than deviate and play $\omega = h$ whenever $b_c < \varepsilon,\ p \leq \frac{1}{2}$ and J plays its prescribed strategy, $m = \frac{2\varepsilon - b_c}{b_c}$. Now, consider the deviation where C plays $\omega = l$ whenever $\Omega = L$. C's expected utility is given by $\varepsilon - b_c$. Thus, it prefers to play q whenever $b_c > \varepsilon,\ p \leq \frac{1}{2}$ and J plays its prescribed strategy, $m = \frac{2\varepsilon - b_c}{b_c}$. Therefore, C has no incentive to deviate.

For proof of uniqueness, see Clark (2008).

This equilibrium, however, cannot exist whenever $b_c > 2\varepsilon$. In that case, no mixing strategy for the Court can make Congress indifferent between playing $\omega = h$ and $\omega = l$. However, this case is substantively extreme, and, as noted in the article, I do not analyze it here. If the assumption is relaxed, though, a pooling equilibrium exists when $b_c > 2\varepsilon$ and $p \leq \frac{1}{2}$. ■

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