The Effects of Judicial Review in American Politics

Joseph Daniel Ura

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Georg Vanberg, Chair

John Aldrich, Reader

Kevin McGuire, Reader

Macro Steenbergen, Reader

James Stimson, Reader

ABSTRACT
JOSEPH DANIEL URA: The Effects of Judicial Review in American Politics.
(Under the direction of Georg Vanberg.)

At least since Robert Dahl's (1957) pathbreaking essay, "Decision-Making in a Democracy: The Supreme Court as a National Policy-Maker," political scientists have invested tremendous energy in exploring linkages between the Supreme Court and its environment. This dissertation joins those efforts by offering studies of three such paths of influence flowing from the judiciary. In turn, I investigate the effects of Supreme Court decisions on public opinion, the role of judicial review in legislative decision-making, and the responsiveness of the media's agenda to Supreme Court decision-making. In all cases, the evidence suggests that the Supreme Court's influence extends beyond the legal world and into the larger political system.

To my parents, Jay and Suzanne.

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Along those same lines, I owe many debts to my advisor, Georg Vanberg. I can't neatly sum up his contribution to this dissertation or to my professional development. Georg has encouraged my interests, talked me through tough theoretical and technical problems, trained me as a formal theorist, helped me join my work to bigger theoretical issues, and generally been an available, sympathetic, and thoughtful mentor and friend. I am certain that I have had the much better end of our collaboration, and I hope that this work and my work yet to come can do some justice to his efforts.

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Chapter 1

The Supreme Court and Public Opinion Reconsidered

At least since Robert Dahl's (1957) seminal essay on the role of the Supreme Court in the American political system, political scientists have been concerned with the linkages between the nation's highest court and its external political environment. Much of the literature that has followed in Dahl's footsteps has concerned itself with the relationship between the Court's policy choices and public opinion. Specifically, scholars have often asked: *Does the mass public react systematically to Supreme Court decision-making?*

There are numerous reasons to suspect that this is the case. First, many of the Supreme Court's most prominent cases involve "easy issues" (Carmines and Stimson 1989): race, abortion, separation of church and state, and free speech are prominent examples. Even relatively unsophisticated observers can make political sense of many Supreme Court decisions. Second, the Supreme Court can focus the media's attention to specific issues, altering the national political agenda (Flemming, Bohte, and Wood 1997; Flemming, Wood, and Bohte 1999). There are also important policy consequences attached to Supreme Court decisions. As the Supreme Court speaks men and women are executed or not (*Furman v. Georgia* 1972; *Gregg v. Georgia* 1976; *Atkins v. Virginia* 2003), affirmative action programs are renewed or repealed (*Regents of the University of*

California v. Bakke 1978; *Gratz v. Bollinger* 2003; *Grutter v. Bollinger* 2003), abortions are provided or banned (*Roe v. Wade* 1973; *Harris v. McCrae* 1980; *Planned Parenthood v. Casey* 1992), and votes in Florida are recounted or not (*Bush v. Gore* 2000). Finally, even in cases of noncompliance with Supreme Court decisions (Rosenberg 1991), the Supreme Court's decisions may have a powerful symbolic effect. For example, many people are clearly moved by the fact that the Court has banned a great deal of comingling between church and state. Given all this, it may not be surprising to see predictive effects flowing from the Supreme Court to public opinion.

Yet, most scholarly inquiry into this question has concluded that there is no general relationship between Supreme Court decision-making and national public opinion (e.g. Bass and Thomas 1984; Caldeira 1991, Hoekstra 2000, 2003; Marshall 1989; Johnson and Martin 1998). This conclusion is built on the dearth of evidence that Supreme Court decision-making affects individuals' *absolute* policy preferences. When the Supreme Court makes policy on some issue, there is little systematic evidence that, on a national scale, people who had previously opposed a position selected by the Court change their minds and adopt the Court's position as their own. In the jargon of this literature, the Court does not generally "legitimize" policy positions in the mass public.

Despite this relatively well established conclusion, there is at least one other pathway from Supreme Court decision-making into public opinion which deserves consideration. It is possible that, as the Court makes public policy, it affects the public's *relative* preferences for the direction of future policy change. As I will discuss in greater detail, as the Supreme Court moves policy in one ideological direction or another, it either satisfies or exacerbates some public demand for future policy change, even though the Court's action may have had no effect on the underlying distribution of absolute preferences in the mass public.

In this chapter, I argue that Supreme Court decision-making has important effects for relative public opinion, that is, for public opinion about the direction of future policy changes. I argue that the American mass public responds systematically to Supreme

Court decision-making in a manner consistent with a thermostatic model of public opinion. Specifically, I demonstrate that there is a significant, negative relationship between Supreme Court decision-making and public policy *Mood* (Stimson 1999). In short, the public responds in an orderly and predictable way to Supreme Court decision-making.

1.1 The Supreme Court and Public Opinion: The Legitimation Hypothesis

In 1957, Robert Dahl (1957) speculated that a major function of the Supreme Court in the American political system was to legitimize the policy choices of the national lawmaking majority. Dahl argued that the Court's special relationship with the Constitution, combined with the public's apparently broad and deep support for the judiciary, would imbue policies upheld by the Supreme Court with legitimacy they would not otherwise enjoy. In other words, Dahl argued that the Supreme Court *legitimizes* public policies: when the Court speaks on an issue public support should swing toward the position it adopts. Since Dahl's essay appeared, political scientists have spent considerable time and energy examining his legitimation hypothesis. However, most scholars who have examined that conjecture have rejected it.

Marshall (1989), for example, examines each of 18 instances over a 45-year period where survey data exist on aggregate public opinion on issues shortly before and after relevant Supreme Court decisions. He finds no systematic relationship between Supreme Court decisions and the public's issue preferences, let alone a consistent pattern supporting the legitimation hypothesis. Indeed, Marshall finds that the mean shift in public opinion surrounding a Supreme Court decision is 0.06%. Likewise, Franklin and Kosaki (1989) find that the Supreme Court's decision in *Roe v. Wade* (1973) homogenized within group preferences on abortion, though it did not affect the aggregate distribution of public opinion on that issue (See also Johnson and Martin 1998). These findings are supported by experimental results that indicate that the Supreme Court does not have a legitimizing effect (Bass and Thomas 1984; But see Mondak 1990, 1994). Gregory Caldeira (1991) summarizes the null result dominance in this field: "[W]e have relatively few well-documented instances when the Supreme Court has shaped the aggregate distribution of public support for this or that policy... We [will] do better, I think, to look for shifts within segments of the population" (1991, 312).

Valerie Hoekstra and Jeffrey Segal (1996; Hoekstra 2000, 2003) explicitly accept Caldeira's critique of the literature. They argue that students of the relationship between the Supreme Court and public opinion should narrow their focus, looking for evidence of legitimation within subpopulations, such as residents of particular geographic areas or demographic groups, resulting from Supreme Court decisions in which a group may have a special interest. In particular, Hoekstra's (2000, 2003) finding that Supreme Court decisions significantly affect attitudes in the communities from which cases originate supports this theoretical position.

The search for general support for the legitimation hypothesis is a quest for evidence of *positive feedback* from Supreme Court decision-making to public opinion. For the legitimation hypothesis to be valid, the Supreme Court must move the *absolute* preferences of the public on any given issue that it decides towards its own preferences. And while the legitimation hypothesis remains a natural theoretical extension of Dahl's claims about the role of the Supreme Court in the American political system, evidence of legitimation has been difficult to find in practice.

1.2 The Supreme Court and Public Opinion: A Thermostatic Model

While most evidence suggests that the Supreme Court does not generally legitimize policy attitudes, it is possible that the Supreme Court's ability to make public policy may have an

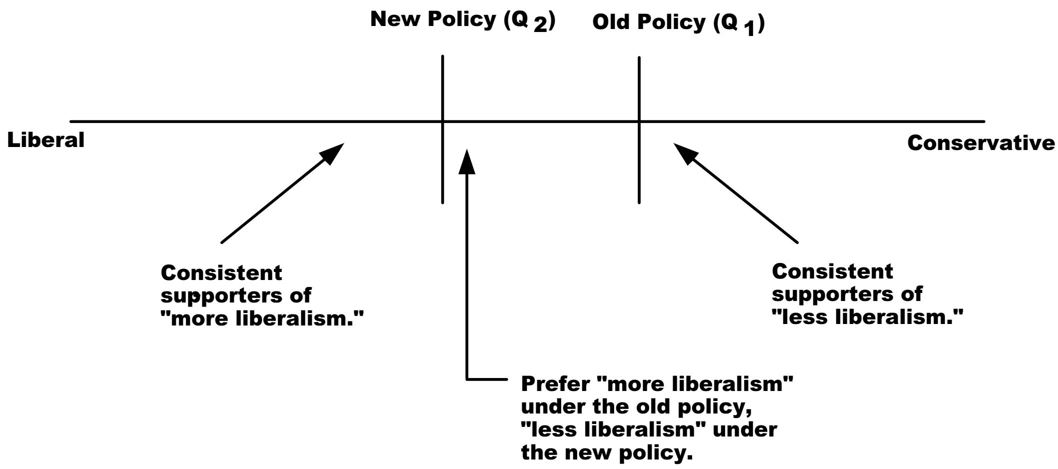

Figure 1.1: The Supreme Court and Relative Policy Preferences

effect on the public's *relative preferences* for policy change. To see this dynamic, consider a hypothetical uniform distribution of citizens in a unidimensional, liberal-conservative policy space where there exists some *status quo* policy, Q_1 (Figure 1.1). Each individual in this space has some absolute preference, a policy ideal point. And, each also has a relative preference for the direction of future policymaking—those to the left of the status quo prefer "more liberalism" in the future; those to the right would prefer "less liberalism."

Now, suppose the Supreme Court changed public policy in a liberal direction, creating a new *status quo* at Q_2. If the Court does not legitimize policy attitudes, there would be no redistribution of citizen's ideal points as a result of the change in the *status quo*. However, the change in policy would produce a change in the proportion of citizens expressing a preference for "more liberalism" in future policymaking. Citizens whose ideal points fall between the old policy and the new policy would have expressed a preference for "more liberalism" before the Court acted and "less liberalism" afterward. So, while the Court may have no effect on the aggregate distribution of absolute preferences, it may have an impact on the distribution of relative preferences. And, generally, as the Court makes more and more liberal or conservative decisions over time, individuals will continue to exhibit changes in their relative preferences for future policy changes in the opposite direction of policymaking. In other words, there will be *negative feedback* from policymaking by the

Supreme Court in the public's relative policy preferences.[1]

This micro theory has a clear macro implication: the accumulation of public policy changes emanating from the Court should produce dynamic, negative responses in the public's preference for the direction of future policy change. This line of thinking amounts to a thermostatic model of public opinion.

Thermostatic models of public opinion posit that the mass public behaves as a political thermostat: when policy deviates from the public's ideal position, it signals policymakers to adjust policy in a corrective direction. These models have been applied broadly to study the national government's responsiveness to changes in public opinion (Erikson, Stimson, and MacKuen 2002; McGuire and Stimson 2004; Stimson, MacKuen, and Erikson 1995; Wlezien 1996) and the public's responsiveness to policymaking by Congress and the President (Erikson, Stimson, and MacKuen 2002). Yet, their application to the study of public responsiveness to the Supreme Court, or the judiciary more generally, has been highly limited. Again, the Supreme Court's power to make policy by overturning and interpreting acts of Congress and the state legislatures represents a potentially powerful ability to affect the public's relative policy preferences.

Nevertheless, there is some evidence of *negative feedback* from judicial decision-making in public opinion. Page, Shapiro, and Dempsey (1987) find that the direction of federal courts' decision-making has a negative relationship with public opinion on several issues. They conclude that, "When their [the federal courts'] statements and actions push in one

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[1]This hypothetical reflects the political reality surrounding the Court. Consider *Lawrence v. Texas* (2003), in which the Supreme Court struck down state laws that banned homosexual sodomy. Existing evidence against the legitimation hypothesis suggests that the Court's decision was unlikely to engender greater support for gay rights at either the individual or aggregate levels. However, the decision amounted to a major practical and symbolic change in policy on the treatment of homosexuals. Thus, a person with a moderate absolute preference on gay rights may have expressed a relative preference for more expansive gay rights before the decision and for less expansive gay right afterwards. Aggregating across individuals, relative preference change of this type would produce a reduction in macro support for more expansive gay rights. This type of effect might have important political implications, such as shifting marginal support toward a political party that advocated a more narrow scope of gay rights. In any event, while the Court failed to legitimize policy attitudes, i.e. change absolute preferences, it may still have a politically relevant effect on relative public opinion.

direction... public opinion tends to move in the opposite direction" (1987, p. 32). These findings contradict the authors' priors, derived from the legitimation hypothesis. Rationalizing these results, they explain that, "the federal courts served as negative referents in the 1970's and early 1980's because of their unpopular actions on such issues as busing and capital punishment" (1987, p. 32).[2]

In keeping with Page, Shapiro, and Dempsey, Wlezien and Goggin (1993) find that public opinion in support of abortion policy "as it is now" increased during the 1980s as the Supreme Court permitted states to regulate abortion more strictly during the course of that decade. Contrary to the explanation of Page and his coauthors, Wlezien and Goggin attribute their findings to a thermostatic model of public opinion. Wlezien and Goggin conclude that the public processed information about the Court's decisions through the media and subsequently exhibited changed relative preferences about abortion as the *status quo* policy changed.

Along these lines, Durr, Martin, and Wolbrecht (2000) find a thermostatic-like relationship between the public's support for the Supreme Court and the level of consonance between liberalism in Supreme Court decision-making and public opinion liberalism. As the ideological divergence between the Court's policymaking activities and the public's preferences for the direction of policymaking grows, Durr, Martin, and Wolbrecht argue that support for the Court falls. The authors conclude, "despite the supposed imperceptibility of the Court, the research suggests that in some way, the public perceives... and evaluates the Court" (p. 775).

Page, Shapiro, and Dempsey (1987) and Wlezien and Goggin's (1992) research indicates that their is *negative feedback* from Supreme Court decision-making in public opinion on a number of unique issues. Durr, Martin, and Wolbrect (2000) show that

[2]While their results are not inconsistent with their speculation about the judiciary's role as a political reference point, the results Page, Shapiro, and Dempsey (1987) report are equally supportive of a thermostatic model of public opinion.

the public translates information about the ideological divergence of the Supreme Court's policymaking activities from its preferred course into a dynamic evaluation of the Court. Taken with previous evidence of general ideological responsiveness in public opinion to policymaking by other institutions of national government (e.g. Erikson, MacKuen and Stimson 2002), these results indicate that the public may produce a general thermostatic response to judicial policymaking. Specifically, as the Supreme Court moves the *status quo* policy in one ideological direction, the proportion of the public that prefers future policy changes in the opposite direction should grow.

1.3 Modeling Public Reaction to Supreme Court Decisions

Through this theoretical framework, it is straightforward to model the effects of Supreme Court decisions on aggregate public opinion, controlling for other factors that may influence public sentiment. First though, it is critical to define public opinion conceptually and to attach a metric of public sentiment to that concept. Fortunately, James Stimson's (1999) study of American public opinion provides guidance on these points, and I draw on his work here.

Stimson defines a general concept of *public policy mood*, or *Mood*, for short: "It connotes shared feelings that move [together] over time and circumstance... [M]ood here captures the idea of changing general dispositions"(1999, 20). He uses this conceptual definition as the starting point to develop a measure of *Mood*. Stimson's *Mood* measure is generated from more than 100 unique national survey items asked in identical form in two or more years from the mid-1950's through the present. Stimson codes responses to each item as liberal or conservative and statistically extracts the common dimensional variance across aggregated responses to every item. This produces a macro level measure

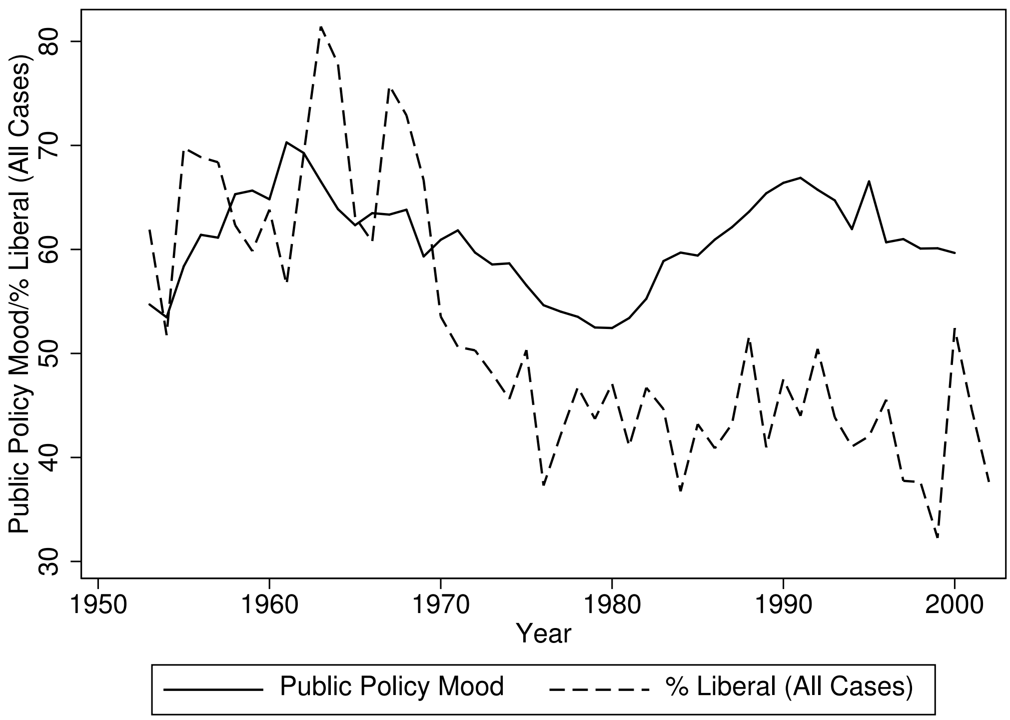

Figure 1.2: Policy Mood and Spaeth % Liberal (All Cases): 1954 to 1996

of relative opinion liberalism in the American mass public. Stimson (1999; see also Erikson, MacKuen, and Stimson 2002) interprets *Mood* as the public's preference for more or less government—that is, more or less policy liberalism. The estimated *Mood* index is scored such that higher values correspond to greater relative liberalism.

Erikson, MacKuen, and Stimson's (2002) extensive study of the relationship between public opinion and public policy points to numerous causal influences on *Mood*, each of which represents an important control for modeling the Supreme Court's influence on public opinion. Erikson and his colleagues find that acts of Congress and macroeconomic conditions produce movement in *Mood* consistent with a thermostatic model of public opinion. They show that there is negative feedback in public opinion from public policy, measured discretely as the net number of important liberal laws (*Laws*) using Mayhew's

(1991) list of important legislation.[3] As Congress creates more conservative public policy (*Policy*)[4], the country exhibits increased liberalism, and *vice versa*. Likewise, they find that Americans demonstrate increased conservatism as inflation (*Inflation*) rises and increased liberalism with rising unemployment (*Unemployment*).

Incorporating Supreme Court decision-making into this model creates a test of the public's ability to respond thermostatically to changes in judicial policy-making: Supreme Court decisions (or patterns of Supreme Court decision-making) should produce negative feedback in *Mood* in-keeping with a thermostatic model of public opinion. In other words, liberal Supreme Court decisions should produce conservative movement in *Mood*, and conservative decisions should produce liberal movement. This conjecture can be subjected to empirical scrutiny.

The thermostatic theory of public opinion advanced here suggests that the value of *Mood* is set by the public's relative preferences for changes in public policy given the macroeconomy and the current state of public policy—whether that policy has emanated from the judiciary or from another institution of governmeny. To the extent that collective judgements about the economy and policy as well as collective expectations about the consequences of policy choices are accurate, *Mood* should adjust quickly to changes in the political or economic *status quo*. Of course, there is reason to suspect that this may not be the case. It is likely that some information will take longer to filter into the political system and that the implementation of public policy will not always meet the public's expectations. Moreover, information about the effects of public policy changes

[3]Mayhew (1991) identifies important legislation in two sweeps. The first sweep includes legislation viewed as important at the time of its passage. The second sweep includes legislation viewed as important some time after passage. Like Erikson, MacKuen, and Stimson (2002), I exclude legislation collected in the second sweep. This list was extended through 1996 by Jay Greene (reported in Erikson, MacKuen, and Stimson 2002).

[4]*Policy* is constructed by aggregating the detrended value of *Laws* over time, beginning, arbitrarily, at zero.

and changes in economic conditions may not spread through the public equally or instantaneously. As such, some of the influence of changes in public policy or the macroeconomy on *Mood* should be effective in the short run, while the full impact of a policy change, for example, may appear only in the long run.

Consider the Supreme Court's decision in *Lawrence v. Texas* (2003). At the time of the decision, many conservative commentators argued that *Lawrence* would serve as the basis for undoing state laws which criminalized or regulated other consensual sexual activity, including prostitution. A citizen observing the decision and the commentary might reasonably have believed that *Lawrence* represented a major change in constitutional law which would generally abridge or forbid state regulation of some licentious activities. However, several federal and state courts have since rejected these arguments, construing the Court's decision in *Lawrence* narrowly. Seeing this application, or at least observing the continued enforcement of laws against prostitution, a citizen might update her beliefs about the location of the policy status quo created by the decision in *Lawrence*. Collectively, reactions along these lines should produce a conservative movement in public opinion in the short term and a corrective movement in a liberal direction in the long run.

Complex dynamics may also be present in cases where the import of a policy change is not felt immediately. Consider the Patriot Act, which passed the Congress with broad, bipartisan support following September 11. Yet, many have subsequently argued that the law has been applied too aggressively by the Bush administration. The passage of the Patriot Act, a conservative proposal, should have produced an initial liberal movement in *Mood*. This shift should have been complemented with additional increase in public opinion liberalism as it became clear that the law would be applied strictly and vigorously.

Patterns of casual flow from Supreme Court decision-making, congressional policy, and macroeconomic conditions such as these suggest that a simple dynamic regression model of public opinion liberalism may be inadequate. To capture the potential for short run and long run influences flowing from political events to public opinion, I employ a series of single equation error correction models (ECM) of *Mood* to operationalize the empirical analysis of the thermostatic theory advanced above.[5]

This is not a traditional application of an ECM, which is typically reserved for the analysis of cointegrated relationships. Yet, established analytical results and recent empirical work indicate that this is an appropriate model specification with which to assess the potentially complex dynamic relationships between public opinion and Supreme Court decision-making.

The Bardsen (1989) single equation ECM, which I employ in these analyses, is an alternative parameterization of the autoregressive distributed lag model (ADL). The ADL takes the bivariate form:

$$Y_t = \alpha_0 + \alpha_1 Y_{t-1} + \beta_0 X_t + \beta_1 X_{t-1} + \epsilon_t \tag{1.1}$$

The notation ADL(p, q) identifies the number of lags of the dependent variable (p) and the number of lags of the independent variable (q). It can be proven that the single equation ECM is a special case of the ADL(1, 1) (Bannerjee *et al* 1993; Davidson and MacKinnon 1993; DeBoef and Keele 2005). The Bardsen ECM differs from the ADL in that an error correction parameter as well as estimates of short run and long run effects appear directly in the model. In the bivariate case, the Bardsen ECM takes the form:

$$\Delta Y_t = \alpha_0 + \alpha_1^* Y_{t-1} + \beta_1^* \Delta X_t + \beta_2^* X_{t-1} + \epsilon_t, \tag{1.2}$$

where α_1 indicates the speed of the reiquilibration of Y to a deviation from its equilibrium with X, β_2 reflects the long run effect of changes in X on Y, and β_1 indicates the contemporaneous relationship between a change in X and a change in Y (Davidson and MacKinnon 1993; DeBoef and Keele 2005). DeBoef and Keele (2005) also note that, since

[5] Volume 4 of *Political Analysis*, published in 1993, contains an excellent series of introductory articles on error correction models (Beck 1993; Durr 1993a, 1993b; Smith 1993; Williams 1993).

the dependent variable of the Bardsen ECM is differenced, any danger of estimating a spurious regression with near integrated data is eliminated. In addition to these attractive analytic properties, Monte Carlo experiments indicate that the ECM implemented through OLS capably recovers the data generating process even in small samples (DeBoef and Keele 2005). In short, the ECM stands out as a useful and powerful model specification for testing dynamic relationships in time series data.

Both the analytic and empirical results support the use of an ECM to examine complex dynamic relationships generally, but there is additional theoretical justification in this case. The thermostatic theory of *Mood* advanced by Erikson, MacKuen, and Stimson (2002) and others implies an equilibrium, supply and demand-like relationship between government activity and public opinion. *Mood* reflects the public's relative demand for more or less policy liberalism from the government. If the economy stood still and policy remained constant, *Mood* should maintain a steady value over time, a value determined by the distance between the public's aggregated absolute preferences for policymaking and the *status quo*. If the government were to disturb this steady state, by producing a single new liberal policy, a simple thermostatic theory of public opinion liberalism would predict a contemporaneous conservative change in *Mood* indicating the extent to which some individuals in the mass public had become satisfied or dissatisfied with the new policy.

Yet, it is clear that, in the real world, the effects of a policy change are not immediately evident nor even predictable. Thus, it would also be reasonable to expect that addition information about the policy change would flow into the public as additional time passes. Uncertainty of this variety may be especially important when the Supreme Court is involved, as the implementation of decisions may occur with long and variable lags (Rosenberg 1991). This would produce some additional movement in *Mood* in the long run. As I discussed in more detail above, this sort of complex dynamic relationship may well play out between the public and the Supreme Court. Thus, the analytical problem of assessing the hypothesized relationships—and determining whether they exist as instantaneous adjustments, long run equilibrations, both, or neither—strongly suggests the use of an ECM.

I conduct and present this analysis in two stages. First, treating Supreme Court decisions as discrete events, I develop and estimate two single equation error correction models of *Mood* from 1954 to 1996 as a function of the macro economy (Changes in *Unemployment* and *Inflation*.),[6] public policy (defined as *Laws* and *Policy*),[7] and the annual net number of liberal *salient* Supreme Court decisions (Figure 1.3)[8]. Following Epstein and Segal (2000), I define salient cases as those for which the decision is mentioned on the front page of *The New York Times*. This measurement approach will capture the effects of the Supreme Court's decisions that are most readily available for mass consumption through the media, rather than the decisions of the uncensored Court. The salient cases time series has a maximum value of 26, produced in 1964, and a minimum value of -8, reached in 1989. The mean of the series is 3.14.

In the second stage, I estimate the same model using alternative measures of the Supreme Court decision-making: the proportion of all Supreme Court decisions decided in a liberal direction annually, the proportion of salient Supreme Court decisions decided in a liberal direction annually, the proportion of all cases in which the Court found for the appellate decided in a liberal direction, and Martin and Quinn's (2002) ideal point

[6]Here, *Unemployment* and *Inflation* are annual national unemployment rates and the U.S. Consumer Price Index, respectively, as reported by the U.S. Census Bureau (2003).

[7]The fist difference of *Policy* is, by definition, exactly equal to the number of *Laws* created in the previous time period, i.e. $\Delta Policy_t = Laws_{t-1}$. To resolve the emergent collinearity issue, I do not model a long run effect for *Laws*.

[8]Ideological coding for case outcomes are drawn from Spaeth's (2004) United States Supreme Court Database.

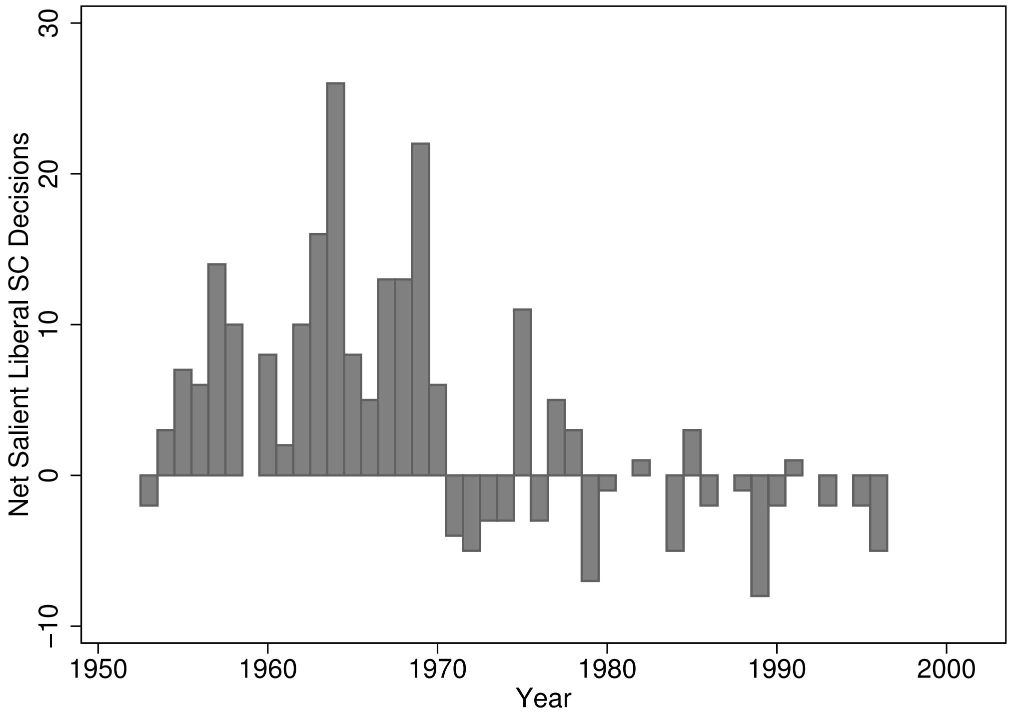

Figure 1.3: Net No. Liberal Salient Supreme Court Cases: 1953 to 1996

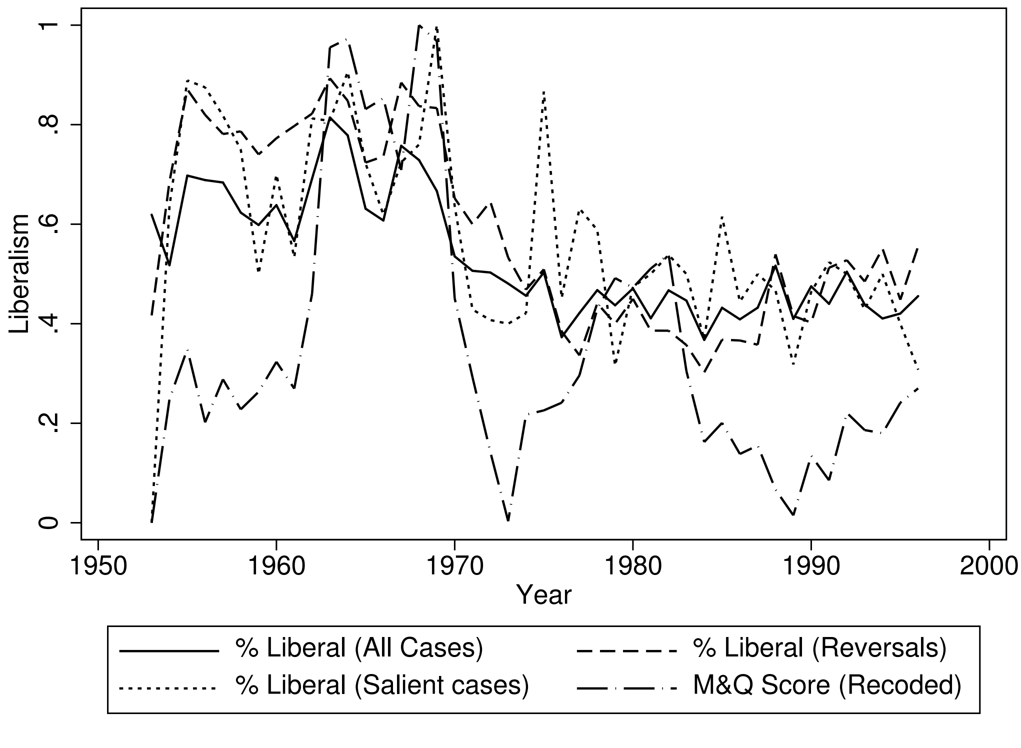

Figure 1.4: Alternative Measures of Supreme Court Liberalism: 1953 to 1996

estimates for the median justice on the Court (Figure 1.4).[9,10] One should note that three of these measurement approaches operationalize the Supreme Court decisions directly, rather than through the filter of media presentation. These operationalized measures imply a more robust flow of information about the Court reaching the public. However, a cursory visual inspection reveals, all of these series correlate very highly (bivariate correlations from 0.54 to 0.90). Yet, there are several years where one or another metric of Court liberalism varies highly from the others. Thus, testing the model under various measurement specification is important.

1.4 Results and Analysis

The results indicate that Supreme Court decision-making (measured as salient, liberal decisions) has a significant, negative effect on public opinion liberalism (Table 1.1). In other words, the data indicate that there is negative feedback from Supreme Court decisions to public opinion consistent with a thermostatic model of public opinion. These effects occur both in the short run and in the long run.

In this specification, each additional salient liberal decision produces a conservative movement in *Mood* of -0.15 points in the short run, while producing an additional long

[9]Scholars of judicial decision-making typically distinguish between Supreme Court decision-making in different issue areas, such as criminal procedure, civil rights, and economic regulation. This choice reflects a theoretical proposition that substantively different areas of law may be governed by different decision-making processes. This may be the case, but, in terms of actual case outcomes—the policy products consumed by the public and other political actors—there is a strong, unidimensional structure to aggregate patterns of Supreme Court decision-making. For example, time series of the proportion of cases decided in a liberal direction among all criminal procedure, civil rights, and economics cases, using Spaeth's (2004) issue coding, correlate at 0.70 or better. Moreover, a simple principal components analysis reveals that a single component, with an Eigenvalue of 2.43, explains more that 80% of the variance in the three series (See also Martin and Quinn 2002). Whatever causal regime may influence Supreme Court decision-making in a given issue area, it appears that the Court's policy outputs may be reasonably well represented by a single dimension. As such, I employ aggregate measures of the Supreme Court's liberalism on a single, left-right ideological dimension here.

[10]The Martin and Quinn (2004) ideal point estimates were recoded to be bounded between 0 and 1 and such that higher values indicated greater liberalism.

run movement of -0.15 during the following years. The long run effect filters into *Mood* at a rate determined by the coefficient on the first lag of the undifferenced dependent variable, in this case -0.48. This means that approximately half of the long run movement of *Mood* to its new equilibrium value will occur at time point $t + 1$, and that half of the remainder will occur at $t + 2$, and so on. Asymptotically, the total long run effect of an additional, liberal Supreme Court decision is -0.15.[11] Thus, an additional salient, liberal decision has, functionally, two effects on public *Mood*, one that is instantaneous (within the same year) and another which filters into public opinion gradually as the decision is implemented and becomes the object of political debate. Considering the mean value of the salient cases time series is just over 3, a typical year's Supreme Court decisions produce a short run conservative movement in *Mood* of roughly the same magnitude as a point increase in *Inflation* and an additional long-run effect of the same magnitude.

At the extremes, the model predicts that the Supreme Court caused a short run movement of -3.9 points in *Mood* in 1964, when the Court produced 26 net liberal salient decisions, and a liberal movement of 1.2 points in 1989 when the Court produced 8 net conservative salient decisions. The short run predicted effect in 1964 is roughly equal to the predicted effect a 8% increase in *Inflation* or a 3% decrease in *Unemployment*, respectively. The long run, equilibrium effect of Supreme Court decision-making in the years following these extreme values are predicted to be roughly the same magnitude.

Additionally, the parameter estimates confirm the expected effects of macroeconomic conditions on public opinion. *Unemployment* has significant causal influence on *Mood* in the short term, indicating that the public quickly assimilates information about the changing condition of the economy. Conversely, *Inflation*'s political influence manifests itself in the long run. Growing *Unemployment* produces liberal movement in *Mood*; increased *Inflation* produces conservative movement in public opinion.

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[11] Given the error correction parameter estimate of -0.48, roughly 90% of the long run effects of Supreme Court decision-making should be in place in 5 years.

Table 1.1: Δ $Mood_t$ as a Function of Supreme Court Liberalism

Predictors	Effects
Long Run Effects	
Mood_{t-1}	-0.48***
	(0.09)
Inflation_{t-1}	-0.54***
	(0.20)
$\text{Unemployment}_{t-1}$	0.07
	(0.32)
Policy_{t-1}	-0.17**
	(0.06)
Laws_{t-1}	-0.38**
	(0.15)
Supreme Court Liberalism_{t-1}	–0.15***
	(0.06)
Short Run Effects	
$\Delta\text{Inflation}_t$	-0.26
	(0.20)
$\Delta\text{Unemployment}_t$	1.27***
	(0.45)
ΔPolicy_t	0.25
	(0.18)
ΔSupreme Court Liberalism_t	-0.15**
	(0.05)
Constant	32.63***
	(5.95)
R^2	0.63
N	42

Note: OLS Estimates. Standard Errors in Parentheses.
*** $p < 0.01$; ** $p < 0.05$; *$p < 0.1$

Laws also have a sharp long run effect on public opinion. Each additional piece of important liberal legislation produced by Congress yields a negative change in *Mood* of 0.38. Likewise, *Policy*, on produces a significant, long run change in *Mood* in a direction supportive of the thermostatic model. These findings indicate that public policy changes have powerful effects on public opinion liberalism, consistent with previous research in this field (Erikson, MacKuen, and Stimson 2002). Finally, the fit of the model is satisfactory; the R^2 of Model 1 is 0.63.

These results also hold substantively across different metrics of Supreme Court liberalism. Models 2, 3, 4, and 5 share the same structure as Model 1, but they are estimated with alternative measures of the macro policy outputs of the Court. Model 2 represents Supreme Court liberalism as the proportion of all orally argued cases decided in a liberal direction in each year. Model 3 utilizes the proportion of all cases in which the Court found for the petitioning party decided in a liberal direction. This measure takes into account the strategic thinking of appellants approaching the Court, who do generally seek a hearing they expect to lose (McGuire and Stimson 2004). Model 4 follows Durr, Martin, and Wolbrecht (2000) by measuring Supreme Court liberalism as proportion of salient Supreme Court cases decided in a liberal direction in each year. Finally, Model 5 uses Martin and Quinn's (2002) ideal point estimates for the median justice on the Court.[12] The results are presented in Table 1.2.

The estimates produced by these models are substantively identical to those produced by Model 1 and hold for both the direct and indirect (media filtered) measures of Supreme Court liberalism. The coefficients of the Supreme Court variables are also very similar in these four models, though this is not surprising given the high correlation between the Supreme Court liberalism time series. In all cases, Supreme Court liberalism maintains its significant, negative relationship with public opinion liberalism. As the Court becomes

[12]These ideal point estimates are calculated by Supreme Court term, not by calendar year. Thus, these figures are on an eccentric calendar with respect to the other measures.

more liberal, *Mood* indicates greater conservatism. Also, the estimates for the macroeconomic and policy variables are very similar across these models and to the estimates produced using the salient cases series. The overall fit of these models remains high. The R^2s for these models range from 0.58 to 0.62.[13] Figure 1 overlays the most parsimonious of these proportional metrics, the proportion of all orally argued cases decided in a liberal direction, with the *Mood* time series.

1.5 Conclusions and Directions for Future Research

Here, I have asked, *Does the mass public react systematically to Supreme Court decision-making?* The data indicate that it does.

Utilizing various measures of Supreme Court liberalism, the data consistently indicate a dynamic relationship between Supreme Court decision-making and public opinion. The relationship takes the form of negative feedback: increased Supreme Court liberalism produces mass conservatism and *vice versa*. This finding holds across models estimated using media related indicators of Supreme Court liberalism and direct measures of Supreme Court policymaking. Moreover, the size of these effects are of the same order of magnitude as the predicted effects of important macroeconomic changes and legislative policymaking on public opinion liberalism. These results suggest that the Supreme Court occupies a larger place in the public's understanding of politics than has been previously thought.

These results should not be taken to mean that Supreme Court decision-making is important to all individuals at all times. The models predict mass opinion change of approximately 5 to 10 points in *Mood* resulting from the observed range of Supreme Court decision-making, depending on the measure employed. While the absolute magnitude of this marginal effect is modest, the relative magnitude of public responsiveness to Supreme

[13]Simple LDV regression models of *Mood* produce substantially similar results. Supreme Court decision-making has a sizeable effect on public opinion in a direction consistent with the causal theory advanced here.

Table 1.2: Δ $Mood_t$ as a Function of Supreme Court Liberalism (Various Metrics)

	Supreme Court Liberalism Metrics			
Predictors	Prop. Lib. All	Prop. Lib. Rev.	Prop. Lib. Salient	M&Q Score
Long Run Effects				
Mood_{t-1}	-0.50***	-0.48***	-0.56***	-0.49***
	(0.09)	(0.09)	(0.09)	(0.09)
Inflation_{t-1}	-0.55**	-0.65***	-0.63***	-0.38
	(0.20)	(0.21)	(0.20)	(0.23)
$\text{Unemployment}_{t-1}$	-0.22	-0.33	0.07	-0.05
	(0.39)	(.43)	(0.32)	(0.34)
Policy_{t-1}	-0.20**	-0.17**	-0.17***	-0.16**
	(0.07)	(0.07)	(0.06)	(0.06)
Laws_{t-1}	-0.35**	-0.38**	-0.38**	-0.38**
	(0.16)	(0.16)	(0.16)	(0.17)
$\text{Supreme Court Liberalism}_{t-1}$	-11.55**	-7.78**	-7.17***	-3.56**
	(4.60)	(3.26)	(2.48)	(1.47)
Short Run Effects				
$\Delta\text{Inflation}_t$	-0.29	-0.40*	-0.35*	-0.25*
	(0.21)	(0.22)	(0.21)	(0.22)
$\Delta\text{Unemployment}_t$	1.02**	1.08**	1.27***	0.89*
	(0.48)	(0.47)	(0.45)	(0.49)
ΔPolicy_t	0.232	0.18	0.24	0.36
	(0.19)	(0.18)	(0.18)	(0.21)
$\Delta\text{Supreme Court Liberalism}_t$	-9.15*	-10.55**	-5.85***	-2.91
	(4.62)	(4.22)	(1.91)	(2.10)
Constant	40.87***	39.42***	41.84***	32.85***
	(7.66)	(7.39)	(5.95)	(6.63)
R^2	0.58	0.59	0.62	0.58
N	42	42	42	42

Note: OLS Estimates. Standard Errors in Parentheses.
*** $p < 0.01$; ** $p < 0.05$; *$p < 0.1$

Court decision-making is surprisingly large; it is easily on par with public reactions to congressional policymaking and important macroeconomic changes. In sum, an attentive subset of Americans react systematically to Supreme Court decision-making and this reaction is sufficient to produce a sizeable, significant, and negative response to the Court in the public's *relative* preferences for future policymaking.

These results raise raise questions with respect to electoral politics. Mass opinion shifts of the size predicted in these models by Supreme Court policymaking may have implications for the outcomes of elections. Yet, since the Court is itself an unelected institution, the public has no way to directly influence the judiciary; it's political response to Supreme Court decision-making must flow through the electoral process. This suggests that Supreme Court decision-making may have interesting effects for American national elections which might be limned, for example, by the prominence of the issue of Supreme Court nominations in recent presidential elections. Obviously, though, this speculation should be explored with future research.

More broadly, these results speak to a larger issue of the substantive role of the Court in the American political system. Democracy implies a reciprocal relationship between the governed and the government. Government must be responsive to the will of the public; the public must be aware of and responsive to the activities of the government. Much of the scientific study of American politics is geared, directly or indirectly, to assessing the quality of one of these two implied relationships. Taken with results that indicate that the Supreme Court, or at least a subset of its members, responds positively to public opinion on a liberal-conservative dimension (McGuire and Stimson 2004; Mishler and Sheehan 1993), this evidence of negative feedback from judicial policymaking in relative public opinion liberalism may indicate a cybernetic relationship—a closed "feedback loop"—between public opinion and the Court, much like the demonstrated reciprocal relationships between the elected branches of government and the mass public (e.g. Erikson, MacKuen, and Stimson 2002). Again, this issue deserves additional attention by scholars.

Chapter 2

"If We Are Wrong. . . The Courts Will Correct It": Legislative Voting under Judicial Review

Scholars of the legislative process and interinstitutional relationships have increasingly recognized the strategic concerns that the anticipation of judicial review may create for legislators. Indeed, judicial review provides numerous opportunities for strategic interactions between courts and legislatures.[1] The most simple of these evolve from the veto player tradition (e.g. Tsebelis 2001), in which legislatures are thought to regard courts as obstacles which must be avoided in the policy-making process (Volcansek 2001). Marks (1989) offers an account of Congress's inability to pass new legislation overturning *Grove City College v. Bell*. And, Pickerill (2004) assesses the extent to which members of Congress shape legislation in anticipation of judicial review, concluding: the "threat of judicial review. . . [is] analogous to veto threats communicated to Congress by presidents" (p. 68).

[1]Despite the pioneering nature of Brian Marks (1989) account of Congress's inability to overturn *Grove City College v. Bell* (1984), most "separation of powers" models have focussed on the limits that legislatures impose on courts (Ferejohn and Weingast 1992; Martin 2001; Spiller and Gely 1992; but see Segal 1997).

Veto player models certainly capture an important dynamic in legislative-judicial relations. But, in an important sense, these models are limited by the narrow view of the politics of legislative-judicial relations. While courts are often roadblock to policy change, policymakers may employ courts to advance a number of policy and political objectives they could not achieve in the absence of judicial review. More broadly, we might ask: *How do legislatures and legislators* use *judicial review to serve some political purpose?*

Political scientists and scholars of public law have offered a number of answers to this question in both comparative and American contexts. For example, Vanberg (1998) shows how judicial review can be used as a tool of legislative coalition building in systems where constitutional challenges must originate in the legislature. Helmke (2002) illustrates how strong governments (dictatorial and democratic) in developing political systems use judicial review to legitimize policy choices, which in turn yields strategic defections by courts when governments weaken. In the American case, Rogers (2001) argues that legislatures may make use of "informational judicial review" to ensure that policy is enacted as intended. Whittington (2005) argues that, "elected officials may actively seek to turn over controversial political questions to the courts" (p.592; See also Graber 1993).

In important respects, the claims of Rogers, Whittington, and others advancing similar arguments about congressional interactions with the Supreme Court have resurrected James Bradley Thayer's (1893) classic critique of judicial review. Thayer argues that judicial review ameliorates the serious consideration of constitutional issues by a legislature since they may delegate that function to the judiciary. He writes, "No doubt our doctrine of constitutional law has had a tendency to drive out questions of justice and right... [I]f we [legislators] are wrong... the courts will correct it" (pp. 155-56).[2]

[2]Thayer's perspective, judicial review serves interests of politicians by providing them with a chance to take a popular action that employs the judiciary as a "backstop" to protect their privately preferred policy outcomes. Thayer's claim echoes in contemporary writings over the role of judicial review and claims of "popular constitutionalism" (e.g. Kramer 2004) and parallels important relatively recent developments in the study of presidential decision-making which examine the balance struck by presidents cross-pressured

This paper refines, expands, and offers some empirical support for Thayer's argument. I explore this issue employing decision theoretic models of legislative voting.[3] The models show that as the probability that a court will annul popular legislation increases, a proposal is likely to garner growing support in a legislature. This dynamic emerges as the institution of judicial review creates opportunities for legislators to vote on proposals with diminished expected policy consequences—that is, policy consequences discounted by the probability that a constitutional court will annul them. I test this prediction through comparative analyses of congressional votes surrounding the issue of flag burning in 1989 and 1990. I conclude that judicial review transforms the incentives facing legislative actors, leading to a legislative decision-making that may privilege political expediency over "questions of justice and right."

2.1 Thayer and the Political Consequences of Judicial Review

The bulk of Thayer's 1893 essay, "The Origin and Scope of the American Doctrine of Constitutional Law," is a case for extreme judicial restraint. He argues that, though judicial review is a legitimate component in American constitutionalism, legislatures have an independent ability to interpret the constitutionality of their own actions and that a judgment that legislation is constitutional is implicit in the decision to enact a law. Because of the close relationship between representatives and their constituents—the sovereign People—courts owe much deference to those judgments about the constitutionality

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between their constituents' preferences and their private judgments, information, and interests (Canes-Wrone, Herron, and Schotts 2001; Canes-Wrone and Schotts 2004; Groseclose and McCarty 2000).

[3]While the following discussion and models are applicable to almost any elected policymaker who interacts with constitutional review, the special relevance of position-taking interests to members of the United States Congress (Mayhew 1974) makes the legislative case an appealing starting point for modeling the complex interactions between elected policymakers, public opinion, and courts.

of legislation. Thus, Thayer argues, courts should adopt a "reasonable doubt" standard for overturning legislative actions: laws should be allowed to stand unless there can be no reasonable doubt that they are unconstitutional.

Like many advocates of judicial restraint before and since, Thayer bemoaned his contemporary judiciary's tendency to step beyond its appropriate role in the American constitutional system. But, his commentary did not stop with a plea for courts to cease and desist in aggressive judicial review. Thayer concluded his essay by outlining some political consequences he expected to flow from courts overstepping their bounds. Thus, his critique of judicial review is tied to a critique of legislative decision-making under judicial review. Tushnet distills Thayer's key point:

> [B]ecause legislatures have mistakenly come to rely on judicial review to correct their "legal" errors, and have abandoned concern for "questions of justice and right," they actually make such [constitutional] judgments less often than they should. Further, even if legislatures make constitutional determinations, many times their decisions will not be reviewed by the courts... [or] courts may turn out to be "broken reeds," failing to exercise their powers of judicial review appropriately (1993, p. 24.)

The danger of emaciated legislative judgment is two-fold: the implementation of unconstitutional policies and the erosion of independent constitutional judgment in legislatures.

Unfortunately, Thayer's brief conjecture about the nature of legislative decision-making under judicial review is just that: brief conjecture. He does not offer any specifics about what mechanisms might cause legislatures to pass final authority over important policy questions onto courts, and he does not offer any evidence to support his speculation. While legal scholars have and will continue to spill vats of ink arguing about the proper degree of deference due legislative choices in constitutional courts, political scientists are uniquely suited to evaluate Thayer's claims about the political consequences of judicial review. Our rich literature on legislative decision-making offers powerful theoretical tools to begin to explore the dynamics that might emerge in a Thayerian legislature.[4] In the following sections, I explore the dynamics of legislative decision-making under judicial review, by offering a simple model of legislative decisions that emerges readily from the congressional literature, then complicating the model by introducing the element of constitutional review by an independent court. Briefly, the models show how judicial review may cause an important change in the relative weights legislators attach to to position-taking considerations—political expediency—versus policy concerns—"questions of justice and right."

2.2 Legislative Decision-Making

Legislators generally, and members of Congress in particular, pursue a variety of goals, including securing reelection, changing public policy, gaining authority in their, and, perhaps, seeking higher office (Fenno 1973). Nevertheless, David Mayhew's (1974) simple observation that, for members of Congress, "[r]eelection underlies everything else" (p. 16), rings true. This formulation was suggested even earlier in the work of Anthony Downs (1957) in the context of party behavior, "Since none of the appurtenances of office can be obtained without being elected, the main goal of every party is the winning of elections. Thus, all its actions are aimed at maximizing votes, and it treats policies merely as means towards this end" (p. 35). Conceptualizing legislators as "single-minded seekers of reelection" (Mayhew 1974, p. 17) has led to tremendous insight into a variety important phenomena and behaviors.

For present purposes, the most important consequence of this electoral connection is

[4]Not all legal scholars who take up Thayer's critique of judicial review in context of legislative decision-making ignore political science scholarship on legislative behavior. In particular, Garrett and Vermeule (2001) propose a set congressional procedural reforms as remedies for the deficiencies of the "Thayerian Congress," which they develop out of assumptions about legislative behavior grounded in empirical studies of Congress.

the implication that a legislator is constrained in her ability to act on preferences that do not match those of her constituency. In the case of the United States Congress, this hypothesis has found at least two strong pieces of supporting evidence. First, members of Congress are held electorally accountable by their constituents for moving "out of step" with their district's preferences (Ansolabehere, Snyder, and Stewart 2001; Canes-Wrone *et al* 2002; Erikson 1971; Erikson and Wright 2000). Secondly, and perhaps more importantly, decades of interviews with members of Congress indicate that these men and women are deeply concerned with reelection and that they actively believe that their voting records affect their chances of reelection (See e.g. Clausen 1973; Fenno 1978; Kingdon 1989; Matthews and Stimson 1975). As Morris Fiorina comments, "No close observer of Congress can doubt that members are highly attuned to the electoral consequences of their actions" (1989, p. 103). This indicates that members of Congress are especially attuned to the electoral consequences of their roll call votes, which are permanent and public records of their actions to change or protect policy.

Nevertheless, there are many circumstances under which legislators might hold personal policy preferences that diverge from those of their constituents. New issues might emerge. Issues that had previously been unimportant achieve new salience. Legislators may have private information about which policy is most likely to produce favorable results. Or, as the Founders may have suspected, legislators, in their capacity as elites, may have a sufficiently different world-view from ordinary citizens that they become out of step with their constituents. Under these or other conditions where legislators might hold personal views that diverge from their constituents' preferences, strategic officeholders must balance the position taking and policymaking costs and benefits of their votes. These cross-pressured legislators are the object of this inquiry and the motor of the Thayerian legislature.

Thus, in deciding how to vote on legislation, legislators must balance at least two competing considerations. They must consider the political consequences of their votes as well as the policy effects of legislation. In other words, when members of a legislature consider their vote choices, they balance the position taking value of their vote against the value of policy change resulting from that vote. As the popularity of a proposal, or as an individual legislator's valuation of the popularity of a proposal, increases so does the position taking value of supporting the proposal. Likewise, as the position of a proposal approaches the ideal point of a legislator in a policy space, the policy value of supporting the legislation also increases.

These decision-making elements facing members of Congress, and legislators generally, can be captured in a decision theoretical model of decision-making that I call the *simple legislative choice*. In the simple legislative choice, legislators face a decision between some proposal (P) that changes public policy and that is associated with some utility from position taking and the status quo ($\sim P$).[5] More formally, the considerations of the simple legislative choice are:

$$\begin{aligned} x_i &= \text{Utility from Policy Change; } x_i \in \Re && (2.1) \\ y &= \text{Utility from Position Taking; } y \in \Re && (2.2) \\ \alpha_i &= \text{Individual Position Taking Parameter; } \alpha \in (0,1) && (2.3) \end{aligned}$$

Thus, under any legislative decision rule, the utilities to a pivotal legislator for choosing P or $\sim P$ are, respectively:

$$\begin{aligned} U_p(P) &= x_p + \alpha_p y && (2.4) \\ U_p(\sim P) &= 0 && (2.5) \end{aligned}$$

This choice is represented in Figure 1.[6]

[5] I assume that passed proposals are fully implemented.

[6] Quantities indicate the utility to the pivotal legislator.

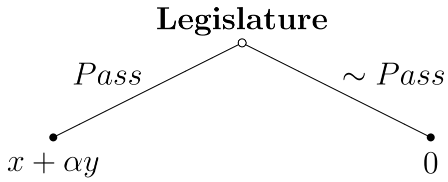

Figure 2.1: Simple Legislative Choice

In a state of the world without judicial review, which might be analogous to a parliamentary vote in a Westminster system or a vote on a constitutional amendment in the American context, these considerations are the only factors that legislators need to incorporate into their decision calculus over a legislative choice. This state implies a decision rule; a pivotal legislator should choose P if:

$$x_p + \alpha_p y \geq 0 \tag{2.6}$$

In terms of the pivotal legislator's policy preferences, the model implies that she should choose P if:

$$x_p \geq -\alpha_p y \tag{2.7}$$

Thus, when $y > 0$ and $\alpha > 0$, the model predicts a circumstance under which a legislator would support a legislative proposal *even for some negative values of* x. In other words, a legislator might rationally choose to support a proposal that violates her private policy preferences if the position taking gains associated with that support are sufficiently large.[7]

[7] Obviously, the inverse is true as well. The model equally predicts that legislators may be willing to accept negative utility for position taking (supporting unpopular legislation) if the gains from policy associated with a particular proposal are sufficiently large. Accepting a policy loss to gain from position taking is usually referred to as pandering; ignoring a proposal's unpopularity to support it is a "Profile in Courage." Regardless, this particular implication of the model is not addressed here.

2.3 The Legislative Choice under Judicial Review

Now, consider the legislature's decision calculus under an institution of constitutional review by an independent judiciary. The institution of judicial review forces legislators to subject their consideration of a proposal to an expected utility calculation with respect to the probability that a court will annul a passed proposal. Considering the case of the pivotal legislative voter, if a court upholds legislation, (U), her payoff from supporting legislation remains identical to supporting legislation under the simple legislative choice. If the court annuls the legislation $(\sim U)$ however, the legislator's utility from policy change is attenuated, while her gains from position taking remain intact.[8]

Treating the court's decision to annul or uphold legislation as a random process such that it will uphold legislation with some known probability,[9] the additional considerations of the legislative choice under judicial review can be formalized as:

$$\lambda = \text{Policy Implementation Parameter; } \lambda \in [0,1], \tag{2.8}$$

$$p = \Pr(U), \tag{2.9}$$

$$(1-p) = \Pr(\sim U) \tag{2.10}$$

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[8]The limited enactment of legislation that may precede judicial review is not necessarily trivial. For example, the Violence Against Women Act of 1994 created a federal civil cause of action for women who were the target of violence motivated by their gender. This provision was struck down by the Supreme Court in *U.S. v. Morrison* (2000). Prior to the Court's decision, at least twelve civil claims were filed and pursued in federal district courts (Masters 1999). These actions encumbered respondents with expenses and provided a basis for those who might have been influenced by the law to observe its enforcement. Also, legislation may have symbolic value. In the case of the Violence Against Women Act, the Supreme Court's annulment of the legislation's civil liability provision may not have detracted from the symbolic value of Congress addressing the incidence of violence aimed at women. Judicial annulment of legislation therefore limits, but does not necessarily remove, the policy effects of legislation. As such, I represent the policymaking considerations in the history (Pass, Annul) as a parametrization, rather than an elimination, of the policymaking utility of a proposal.

[9]Obviously, courts do not uphold or annul legislation by some stochastic process. However, this representation of judicial review is reasonable from the perspective of analyzing legislative behavior. Since legislators cannot know *for certain* how a court will react to legislation, they must act on some probabilistic prior belief about the likelihood a statute will be annulled. As such, the model maps reasonably well onto the actual decision calculus of legislators operating under judicial review.

Thus, under any legislative decision rule, the utilities to a pivotal legislator for choosing P or $\sim P$ are, respectively:

$$U_p(P) = p(x_p + \alpha_p y) + (1-p)(\lambda x_p + \alpha_p y) \tag{2.11}$$

$$U_p(\sim P) = 0 \tag{2.12}$$

This choice is represented in Figure 2.[10]

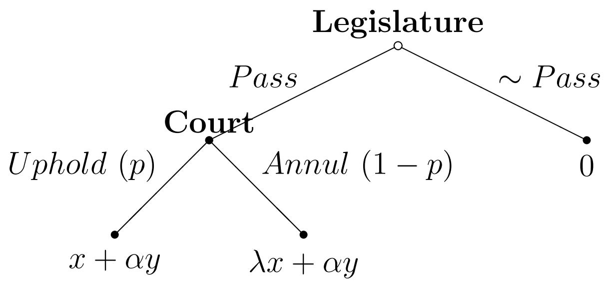

Figure 2.2: Legislative Choice under Judicial Review

Again, this situation implies a decision rule for a pivotal legislator; choose P if:

$$p(x_p + \alpha_p y) + (1-p)(\lambda x_p + \alpha_p y) \geq 0 \tag{2.13}$$

And, in terms of her policy preferences, this reduces to a choice of P if:

$$x_p \geq \frac{-\alpha_p \; y}{p(1-\lambda)+\lambda} \tag{2.14}$$

Like the simple legislative choice, the decision model predicts that a legislator would support a proposal *even for some negative values of* x when $y > 0$ and $\alpha > 0$. And,

[10]Quantities indicate the utility to the pivotal legislator.

because $p \in [0, 1]$ and $\lambda \in [0, 1]$ the model of legislative choice under judicial review predicts a choice of P at a lower value of x than for the simple legislative choice. Thus, for all admissible values of p and λ, the threshold value of x, the cut point, at which a legislator would vote "No" is reduced by the potential of judicial review. The possibility of judicial annulment of legislation raises the the threshold of lost policy utility necessary to overcome a positive gain from position taking necessary to induce legislative opposition. To the extent that a legislator anticipates negative judicial review, she is relatively free to vote for a policy change that enjoys the support of her constituents, even if she opposes the policy personally. In other words, the model shows that judicial review transforms legislators' incentives in precisely the manner Thayer suggests, inducing some legislators to vote for proposals that which they would have opposed in the absence of the possibility of a judicial veto. More plainly, judicial review creates opportunities for legislators to have their cake and eat it, too.

2.4 The Consequences of Judicial Review

The decision models proposed above lead to predictions of comparative voting behavior across various levels of the likelihood that a court will annul a statute; that is, across values of p. Thus, we can use the models to make predictions about how legislators' voting behavior might change as a result of variance in the decision elements of the models, which, in turn, might be used to assess the model empirically. The decision rule at (2.14) implies that legislators' decision-making on a particular proposal will be in part a function of their expectations about judicial review and their valuation of position taking relative to their the proposal's policy utility. Thus, under different conditions on p and α, individuals legislators voting choices on a given proposal may vary. I discuss each of these in turn.

2.4.1 Judicial Review

Assume that a proposal is associated with some positive position taking payoff and that legislators place some uniform positive value on that payoff, i.e. $y > 0$ and $\alpha_i > 0 \; \forall \; i$. On the dimension x, along which legislators may be arrayed by their policy payoffs resulting from the enaction of a given proposal, there exists a cut point, Θ_p, at:

$$\frac{-\alpha y}{p(1-\lambda)+\lambda} \tag{2.15}$$

such that legislators for whom $x_i < \Theta_p$ (those to the left of the cut point) will vote "No" on a proposal and those for whom $x_i \geq \Theta_p$ (those at or to the right of the cut point) should vote "Yes." This cut point will move along the dimension x as p varies, holding all else constant. The cut point may exist anywhere in the range from $-\alpha y$ (where $p = 1$) to $-\frac{\alpha y}{\lambda}$ (where $p = 0$).

Between any two potential cut points, there exists an interesting interval in which legislators' voting behavior on a particular proposal is contingent on their belief of the likelihood that a court will strike down the proposal. Thus, in a circumstance where a legislature considered the same proposal twice, once where $p_1 \leq 1$ and again where $p_2 < p_1$, the model predicts that legislators to the right of Θ_1 line will vote "Yes" twice, those to the left of Θ_2 will vote "No" twice, and those legislators in the interval between Θ_1 and Θ_2 will "flip" on the proposal, voting "No" at p_1 and voting "Yes" at p_2 (Figure 3).

The properties of this flip interval can be illustrated clearly by considering the special cases where $p = 1$ and $p = 0$. In the case $p = 1$, a court will uphold a proposal with certainty. This situation is equivalent to legislative voting in the absence of judicial review, analogous to a congressional vote on a proposed constitutional amendment. The expected utility for the pivotal legislator reduces to the utility associated with P under the simple legislative choice, $x_p + \alpha_p y$. The cut line, Θ_1 falls at $-\alpha y$. Substantively, on a given

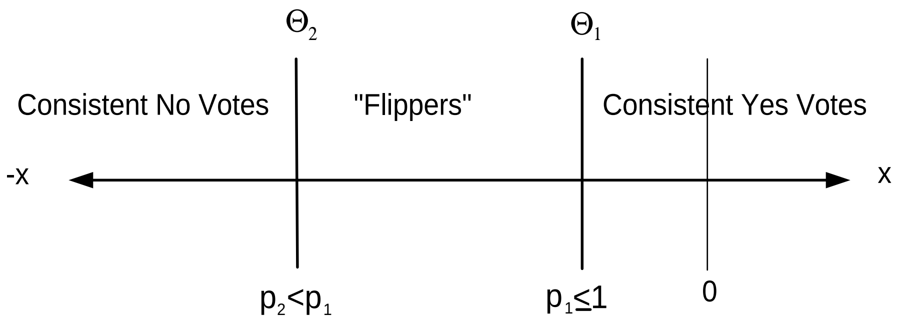

Figure 2.3: Comparative Statics over Judicial Review

vote under these circumstances, some marginal policy losers will vote to pass a proposal (those for whom $x_i \geq -\alpha_i y$), but the bulk of the proposal's policy opponents are willing to absorb the lost position taking opportunity to prevent the policy change (those for whom $x_i < -\alpha_i y$).

Now, consider a second vote on the same proposal under the same general conditions, but where a court is certain to annul the proposal, $p = 0$. This situation is comparable a vote on a statute that faced a certainty of being annulled by the Supreme Court. Substantively, a legislative vote in this case is an exercise in symbolic politics; a court will annul the enacted proposal with certainty. More formally, the expected utility to the pivotal legislator for choosing P reduces to $\lambda x_p + \alpha_p y$. Thus, the cut point, Θ_2, falls at $-\frac{\alpha y}{\lambda}$, which is to the left of Θ_1. Thus, the model predicts that any legislator who falls between Θ_1 and Θ_2 should flip on the proposal, voting "Yes" when $p = 0$ and having voted "No" when $p = 1$.

2.4.2 Position Taking

In addition to expectations about judicial review, p, the weight attached to position taking, α, also contributes to the location of a cut point, Θ_α, along the policy dimension holding all else constant. Recall, the cut point is defined as:

$$\frac{-\alpha y}{p(1-\lambda)+\lambda} \tag{2.16}$$

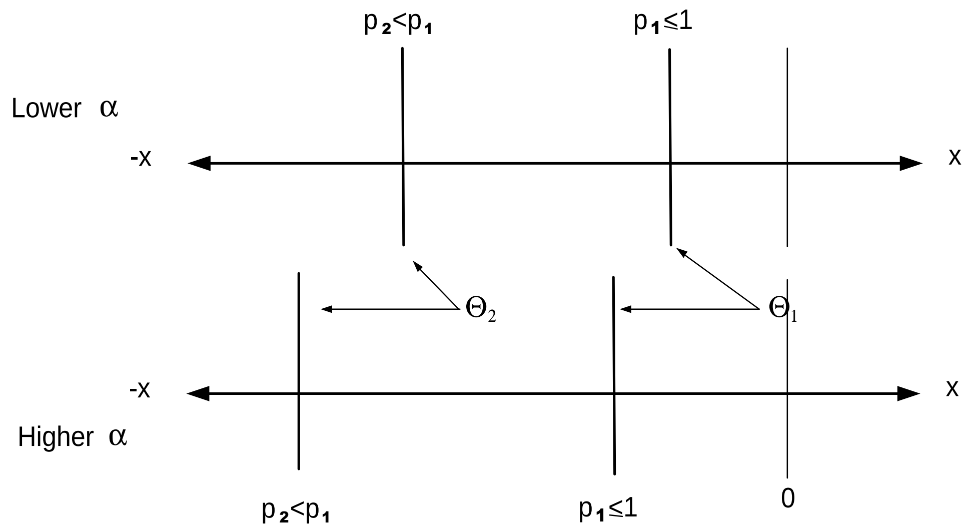

Figure 2.4: Comparative Statics on Position Taking

and $\alpha \in (0, 1)$. Thus, as a proposal's position taking utility increases for legislators from 0 to 1, the value of the numerator in the cut point function grows, moving Θ_α to a lower value on the dimension x (i.e. to the left). In terms of practical politics, this result is intuitive. As a legislative proposal becomes more popular, it will become increasingly difficult for election-seeking legislators to justify voting against a proposal because of their private policy reservations. Thus, the value of x needed to induce a "No" vote is negatively related to the position taking value of a proposal. So, legislators in a vulnerable electoral position would be more likely to vote "Yes" on a popular proposal than their like-minded colleagues who might be in safer seats.

Additionally, these position taking effects are interactive with expectations about judicial review in driving the location and size of the flip interval along x. As I have shown, the flip interval for two legislative votes grows as the difference between p increases between the votes. And because λ and α are bound between 1 and 0, a given unit increase in in α will move the cut line further at lower values of p than at higher values, expanding the size of the interval and moving it to the left as α grows. This interactive effect is depicted in Figure 4.

2.5 An Empirical Assessment

Thayer argued that judicial review erodes legislative responsibility since legislators may come to rely on courts to correct their errors—"If we are wrong. . . the courts will correct it." The preceding models have introduced a mechanical explanation for this theory; judicial review introduces an element of risk associated with the policy consequences of a legislative choice. The model, in turn, leads to clear predictions of the behavior of legislators in various political circumstances, which may be used to assess the Thayerian theory of the effects of judicial review.

An appropriate test of the theoretical model in the context of the United States Congress would be an analysis of votes on a popular statute that is likely to be overturned by the Supreme Court and a popular constitutional amendment concerning the same substantive issue considered by the same session of Congress. Under these conditions, the theory advanced here predicts support for the statute from all but the strongest policy opponents, since position taking benefits would outweigh the proposal's sharply reduced policy considerations given the likely negative reaction from the Court. However, support for the amendment would be less widespread, since that would reflect a balance of both policy concerns and position taking considerations. In other words, legislators whose personal policy ideal points fall in the interval between $\Theta_{Statute}$ and $\Theta_{Amendment}$ should vote differently on the same substantive policy question on the basis of changed expectations about judicial review.

In 1989 and 1990, both houses of Congress cast a series of votes on proposed bans of flag burning. The House and the Senate voted on a statutory ban, the Flag Protection Act of 1989, as well as a constitutional amendment, which the Senate considered twice. These votes were precipitated by a Supreme Court decision in the summer of 1989, *Texas v. Johnson* (1989), to strike down state laws banning flag burning. This created a strong expectation that the Court would strike down a statutory flag burning ban, creating a relatively high constant value for p across individuals. A statutory ban on flag burning would be a "free vote." An amendment, though, was almost certain to be ratified in the states and carried strong policy consequences. This condition creates a quasiexperimental control: most members of Congress cast at least two votes on the same substantive policy question, in more or less the same political conditions, over short period of time: one in the presence of judicial review (with a relatively low value of p) and one in its absence ($p = 1$). Analyzing the preferences and political circumstances of the legislators involved in these votes provides an opportunity to perform a simple, direct test of the theory of legislative voting under judicial review advanced here.

2.5.1 Expectations

Under these conditions, the models lead to clear predictions about what factors might lead to a member of Congress to cast a particular set of votes on the flag issue, that is, to vote "No" on both the statute and the amendment, to vote cast two "Yes" votes, or to flip on the issue, voting "Yes" on the statute while later opposing the amendment. In particular, the models suggest two important independent variables in assessing their empirical validity: legislators' preferences (which corresponds to their location along the policy dimension, x) and variance in legislators' weight on the proposals' position taking value (α). These variables should be significantly related to the legislators' individual cut points on the two votes. Using observed patterns of voting across the flag burning votes as a dependent variable, these hypotheses can be tested in relatively straightforward models of flipping.

As for preferences, the theoretical model suggests a negative relationship between support for civil liberties and legislative support for proposed the flag burning bans. Those members of Congress with the highest levels of support for civil liberties (those with the lowest value on x) should be the most likely to oppose the ban in both statutory and amendment form. Those with the lowest level of support for civil liberties (those for whom x is relatively high) should be the most likely to support the ban on both votes. Those with intermediate values of x, should be the most likely to flip on the flag issue, supporting a flag desecration ban in statutory form and opposing it in amendment form. Individual legislator's relevant policy preferences can be measured using American Civil Liberties Union (ACLU) scores.

In addition to policy preferences, position taking is the key concept in the decision models—it is the interaction between this benefit and the likelihood of a judicial veto that drives the location and size of the flip interval. As such, operationalizing the position taking incentives facing legislators is critical for assessing the validity of the theoretical model. The fact that position taking needs are strongly related to electoral concerns provides leverage on this problem. Position taking incentives are likely to be especially important for legislators for whom electoral concerns are more pronounced than others. For the Senate, whose members are elected on a rotating basis, a dummy variable for Senators seeking reelection in 1990 can be used as an indicator of variance on the weight attached to gains from position taking, α, since these Senators were the first to face the electorate after the flag votes, giving them less time and fewer opportunities to distance themselves from unpopular votes. In the House, where members are elected on a uniform schedule, the weight on position taking can be operationalized with the Representatives' margin of victory in the previous election and a dummy variable for those in competitive seats.[11]

I should note here that there were relatively few members of Congress, especially in the Senate, who opposed the statutory ban on flag burning. The Senate vote on the Flag Protection Act was 91-9; it was 380-38 in the House. Moreover, as I will discuss in more detail, a number of these "No" votes came from conservatives who wished to express their

[11]The "winning margin" is the percent difference of the two party vote share of the Representative over his or her opponent in the 1988 election. I count competitive seats as those won by less than a 10% share of the two-party vote.

preference for a constitutional solution over a statutory remedy. As such, there are only a relatively small number of members of Congress for whom the cut point between voting "No" twice and flipping is relevant. In the Senate, the numbers involved are sufficiently small to motivate an ordinary logit model to assess the likelihood of moving from a "Yes" vote on the statute to a "No" vote on the amendment. In the House, the larger number of observations justifies a an ordered logit approach to model behavior across the three categories.

In both cases, the theoretical models lead to clear predictions about the sign and significance of the coefficient estimates in these models of flipping. Indicators of policy preferences, where higher commitment to civil liberties are coded with higher values, should be positively and significantly related to stronger categorical opposition to the flag desecration ban, i.e. moving from casting two "Yes" votes to flipping, moving from flipping to casting two "No" votes. Substantively, those who have the most to lose in a policy space should be the most likely to vote against a proposal when it will count. Conversely, indicators of electoral sensitivity should be negatively and significantly related to opposing the flag burning ban. Those most in need of the proposal's position taking value should be most apt to vote on the popular side of the issue whenever possible.

2.5.2 The Flag Burning Controversy: Background

In August, 1984, Gregory Johnson was arrested for violating Texas's flag desecration statute by burning an American flag outside the Republican National Convention in Dallas.[12] He was convicted and sentenced to a year in prison and a $2,000 fine. Johnson appealed his conviction, and the Texas Court of Criminal Appeals overturned his conviction on First Amendment grounds. The state of Texas appealed to United States Supreme

[12] Readers interested in a more complete narrative of the flag burning controversy that played out in 1989 and 1990 should consult Goldstein (1996), and for an empirical analysis of House votes on the 1990 flag protection amendment, see Clark and McGuire (1996).

Court, arguing that its statute protected a unique national symbol. In March, 1989, the Supreme Court heard arguments in *Texas v. Johnson* (1989); in June, the Court rendered its decision.

By a 5 to 4 vote, the Supreme Court held that flag burning is expressive speech protected by the First Amendment. Writing for the majority, Justice Antonin Scalia concluded, "If there is a bedrock principle underlying the First Amendment, it is that the Government may not prohibit the expression of an idea simply because society finds the idea itself offensive or disagreeable." The Court's decision in *Texas v. Johnson* catalyzed the "flag burning" issue.

The public's response to *Texas v. Johnson* was decisive. A poll conducted the day after the Court handed down its decision found that 71% of adults would "support a new constitutional amendment to make flag burning illegal." Only 24% opposed the amendment.[13] Goldstein (1996) writes,

> The... *Johnson* decision touched off what *Newsday* characterized as a "firestorm of indignation" and *Newsweek* termed "stunned outrage" across the United States. Certainly, no Supreme Court decision within recent memory, if ever, was so quickly, bitterly, and overwhelmingly denounced by the American public and political establishment (p. 113).

The support for a response to the Supreme Court's ruling prompted swift action from the Congress.

Republican leaders, supported a constitutional amendment to empower Congress to prohibit flag burning. Many Democratic leaders campaigned to create a statutory flag burning ban. Republicans countered that the Supreme Court's decision in *Texas v. Johnson* made it clear that it would not tolerate a statutory ban. Democrats counselled a "wait

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[13] Gallup Poll conducted June 23, 1989. N=500 Adults.

and see" approach, advocating a statutory solution until the Supreme Court could reconsider the issue. However, the Court's forceful decision made the prospect of a legislative solution dim.[14]

That October, congressional responses to the flag issue took two forms: passage of a statute banning flag burning, the Flag Protection Act of 1989, and consideration of a constitutional amendment to empower Congress to prohibit the desecration of the American flag. The Senate voted on both the flag protection statute, which passed, and the constitutional amendment, which failed. The House considered only the statute, which passed; Democratic leaders were successful in preventing a vote on the constitutional amendment at that time.

The Flag Protection Act took effect on October 30, 1989. At midnight, protesters across the country burned American flags. Three protesters, Shawn Eichman, Dave Blalock, and Dread Scott, were arrested on the steps of the United States Capitol. In March 1990, the protesters pled guilty in federal district court but argued that the Flag Protection Act was unconstitutional. The district court agreed and set aside the protesters' convictions.

Under a provision of the Flag Protection Act, the Supreme Court heard an expedited appeal of the district court's decision. While some observes may have had doubted the case's outcome, there appears to have been a consensus that the Supreme Court which had acted to protect flag burning in *Johnson* would not have changed its mind in a matter of months. These suspicions were confirmed on June 11, 1990 when the Supreme Court handed down its decision in *United States v. Eichman* (1990). By the same 5 to 4 vote as *Texas v. Johnson*, the Court held that the Flag Protection Act unconstitutionally prohibited expressive speech. Concluding for the majority, Justice Brennan wrote, "the

[14] Senator Robert Dole wrote, "it [proposed statutory remedy] flunks the test...because the statute itself is unconstitutional...[T]he Supreme Court has already ruled that the government can't protect the flag as a 'symbol'" (Dole 1989).

[Flag Protection] Act suffers from the same fundamental flaw as the Texas law [at issue in *Johnson*]."

Following the Court's decision, Congress again took up the issue. In June, both houses of Congress voted on a flag protection amendment. Though the proposal received majority support in both houses, it failed to achieve the two-thirds support necessary to send the amendment to the states for ratification.

2.5.3 Comparative Vote Analysis: The Senate

In October, Senators cast their first vote on flag burning, a vote for passage of the Flag Protection Act of 1989 which had previously passed the House of Representatives. Given public support for a flag burning ban and the high probability that the Supreme Court would not tolerate a federal flag burning ban when it had just struck down similar state laws, the theoretical models predict broad support for the statute. Only those Senators most committed to a to civil liberties (those who had the largest policy losses under the statute) should have voted against the bill.

The Senate vote was 91 for the statute with 9 against. Three conservative Senators, Robert Dole (R-KS), Charles Grassley (R-IA), and Orrin Hatch (R-UT) voted against the statute to demonstrate their preference for a constitutional amendment. Senator Gordon Humphrey (R-NH), who was about to retire from the Senate, called the flag issue "an exercise in silliness" and voted against the bill (Lewis 1989).[15] The remaining no votes came from civil libertarians: John Chafee (R-RI), Edward Kennedy (D-MA), Bob Kerrey (D-NE), Howard Metzenbaum (D-OH), and Daniel Patrick Moynihan (D-NY).

Sincere "No" votes on the flag burning statute came from among those who have

[15]Speaking on the floor of the Senate, Senator Humphrey argued that the rush to create a flag burning prohibition demeaned the national legislature's upper house. He remarked, "In this present circumstance, namely the pendency of this resolution, one might ask: What distinguishes the Senate from the House of Representatives, except that we serve for 6 years?...Where is the cool and the calm and . . . to expect of the Senate? Perhaps it will yet prevail. . . But, certainly, for a while, it was a very near thing and not the least bit becoming to the Senate, in my view" (Congressional Record 1989).

embraced the highest levels of support for a broad definition of civil liberties.[16] Aside from the pro-amendment "No" votes, along with Senator Humphrey's esoteric "Nay", all opposition to the statute came from the most libertarian cohorts, in keeping with the theory of legislative voting discussed above. In Table 1, I report the votes of Senators on the Flag Protection Act of 1989 tabulated by ACLU score quintiles.

Table 2.1: Senate Flag Burning Votes

	Statute		89 Amend't		90 Amend't	
ACLU Score Quintle	Yes	No	Yes	No	Yes	No
1 (0-14)	23	4	26	1	26	1
2 (15-31)	13	0	11	1	13	0
3 (32-67)	20	0	11	9	14	6
4 (68-80)	18	2	3	17	5	15
5 (81-100)	17	3	0	20	0	20
Total	91	9	51	48	58	42

The second Senate vote on the flag issue came roughly two weeks later, when it considered a constitutional amendment to empower Congress to ban flag burning. Recall, the theoretical models predicts that voting on a constitutional amendment would reflect a balance between position taking needs and sincere policy concerns. Given a relatively uniformly large and positive position taking gain to be had by supporting a flag burning ban, the models predict that Senators would have been increasingly likely to oppose the ban as their commitment to civil liberties grew. Again, the data indicate that this prediction holds.

[16]The "No" votes in the first quintile are from Senators Dole, Grassley, Hatch, and Humphrey.

The Senate voted 51 to 48 for the amendment, which failed to achieve the two-thirds support required for passage. Forty-two Senators switched sides, voting against the constitutional ban on flag burning despite casting a vote in favor of the Flag Protection Act. Senator Pete Wilson (R-CA) did not vote on the amendment. In Table 1, I report the votes of Senators on the 1989 Flag Burning Amendment tabulated by ACLU rating quintiles. The pattern of voting on the flag burning amendment support the notion that Senators, relieved of the threat of judicial review, were substantially more sanguine about limiting the scope of free speech. Indeed, opposition to the amendment is positively correlated with higher ACLU scores (0.79).

This interpretation is reinforced by the results of an estimated logit model of "flipping" among those 90 Senators who cast a "Yes" vote on the statute. I report the results in Table 2. The estimates also confirm the intuition of the theoretical models. Higher ACLU scores, indicating increased support for civil liberties, predict an increased propensity to flip. As Senators' policy loss grows under the proposal, as their position along the policy dimension x moves left, they are more likely to be on the left side of the relevant cut point.

Conversely, a forthcoming reelection bid, placing a premium on being on the right side of an issue, predicts a propensity for remaining supportive of the flag burning ban. Recall, an elevated value of α will shift the flip cut point to the left. *Ceteris paribus* a legislator with a relatively high value on position taking is more likely to be to the left of the relevant cut point along x than a legislator with a lower value on α. Both this result and the model's estimates for Senators preferences confirm the model's predictions.[17]

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[17] An estimated logit model of flipping predicted by ACLU scores, a 1990 reelection dummy, first dimension DW-NOMINATE scores (Poole and Rosenthal 1997), and a dummy for Democratic partisanship produces substantively identical results with respect to the theoretically derived variables. ACLU scores produce a significant positive effect; 1990 reelection produces a significant negative effect. See the Appendix for the results of the extended control model.

Table 2.2: Logit Analysis of Senate Flipping
(1 = "Flip")

Predictors	89 Amend't	90 Amend't
ACLU Rating	0.11***	0.13***
	(0.01)	(0.03)
1990 Reelection	-1.53**	-2.07**
	(0.83)	(0.91)
Constant	-5.03***	-6.74***
	(1.17)	(1.60)
χ^2	79.60	82.98
Pseudo R^2	0.64	0.68
PRE	0.79	0.78
N	90	91

Note: Logit Estimates. Standard Errors in Parentheses.
*** $p < 0.01$; ** $p < 0.05$; *$p < 0.1$
One tailed tests.

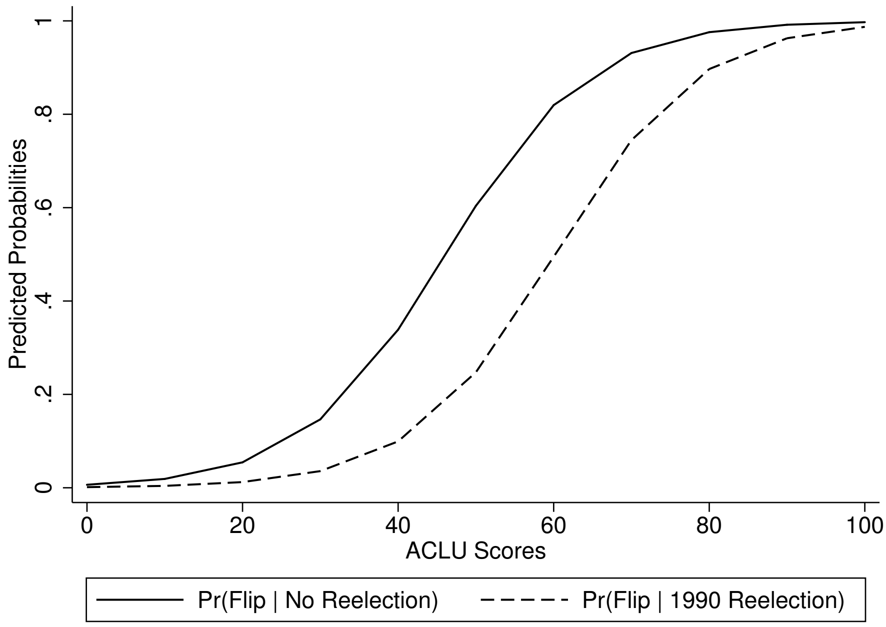

Figure 2.5: Predicted Probabilities of Senate Flipping

Figure 7 shows the marginal effects of changes in support for civil liberties on a Senator's propensity to flip under two different electoral situations. This figure also supports the intuition of the theoretical models. Senators who place a higher value on civil liberties are more likely to change from voting *for* a statutory ban on flag burning to voting *against* a constitutional ban on flag burning. The only substantive difference in these situations is the potential for judicial review in the statutory case. The presence of judicial review over the statute acts as a quasiexperimental control; the observed changes in voting behavior are the result of the influence of judicial review on members' of Congress decision calculus.

A Senator might have plausibly, if naively, believed that the Flag Protection Act was sufficiently distinguishable from the state laws overturned in *Texas v. Johnson* as to escape a judicial veto. But, following the Supreme Court's actions in *U.S. v. Eichman*, a straightforward statutory remedy was not available in the absence of constitutional change. Preventing flag burning unambiguously demanded a constitutional amendment.

This additional vote on the amendment provides another chance to assess the models' comparative predictions on Senators' votes with the policy stakes of the vote more clear than in the preceding October. The Senate vote was 58 to 42, failing to meet the constitutional requirement of two-thirds majority support to send a proposed amendment to the states. Ninety-one Senators cast votes on both 1989 vote on the Flag Protection Act and the 1990 proposed constitutional amendment; 36 Senators flipped on these votes. In Table 2, I report the results of a logit analysis of this second set of flips.

The results confirm the findings of the prior model of flipping. Higher support for civil liberties produced a heightened propensity to switch sides on an issue once the possibility of judicial review was removed. Likewise, a higher valuation of position taking on the issue predicts a reduced probability of switching sides. The intervening actions of the Supreme Court only marginally affect the estimates in the two cases, producing slightly larger predicted effects for the independent variables and a somewhat better overall fit.[18]

2.5.4 Comparative Vote Analysis: The House

The House of Representatives also held two roll call votes on the flag issue: a vote on the Flag Protection Act of 1989 in October and a vote on a flag burning amendment in June 1990. The statute passed by a vote of 380 to 38. The amendment garnered 253 votes; there were 175 votes against, failing to achieve a two-thirds majority. Among the members of the House who voted on both items, 229 voted "Yes" twice, 149 "flipped," voting yes on the statute and no on the amendment, and 20 Representatives were consistent "No" voters. [19] These votes are tabulated by ACLU rating quintiles in Table 3.

[18]The estimates produced by an extended control model as described in footnote 12 remain substantively the same.

[19]Eighteen Representatives, all Republicans, also voted "No" on the statute and "Yes" on the amendment. Though an explanation of their behavior is immediately evident from the political context, these votes are beyond the predictive facility of the models advanced above and are excluded from the empirical analyses reported here.

Table 2.3: House Flag Burning Votes

	Statute		Amendment	
ACLU Score Quintle	Yes	No	Yes	No
1 (0-12)	97	11	107	1
2 (13-31)	57	7	59	9
3 (32-57)	85	1	65	22
4 (58-85)	88	3	21	72
5 (86-100)	53	16	1	71
Total	380	38	253	175

The eyeball-level patterns of flipping comport with those observed in the Senate. House members became more likely to oppose the flag burning bans as their commitment to civil liberties grew. This interpretation is supported by the results of an ordered logit analysis of the three categories of House voters: consistent opponents (coded high), flippers (coded middle), and consistent supporters (coded low). I report these estimates in Table 4.[20, 21] Among members of the House who voted on both items, commitment to civil liberties, thus a larger policy loss under an enacted flag burning ban, is significantly, positively related to opposing the flag burning bans. At the same time, Representatives of competitive districts, those with an higher value of taking a popular stand on a salient issue, were significantly more likely to more consistently support the flag burning ban.

Figures 6 and 7 show the marginal effects of changes in support for civil liberties

[20] Again, the estimates produced by an extended control model as described in footnote 12 remain substantively the same for the theoretically important variables (Appendix).

[21] Including the 18 members of Congress excluded for voting "No" on the statute and "Yes" on the amendment as a fourth, ordered category of "reverse flippers," however, produces substantively the same results as the included analyses. ACLU scores have a significant positive effect; an indicator for members serving competitive districts yields a significant negative effect.

Table 2.4: Ordered Logit Analysis of House Voting Cohorts
(Consistent Opponents=2, "Flippers"=1, Consistent Supporters=0)

Predictors	Effects
ACLU Rating	0.09***
	(0.01)
1988 Win Margin	0.00
	(0.01)
Competitive District	-1.37**
	(0.70)
Cut Line 1	4.86***
	(0.54)
Cut Line 2	9.47***
	(0.76)
χ^2	316.06
Pseudo R^2	0.48
PRE	0.55
N	396

Note: Logit Estimates. Standard Errors in Parentheses.
*** $p < 0.01$; ** $p < 0.05$; *$p < 0.1$
One tailed tests.

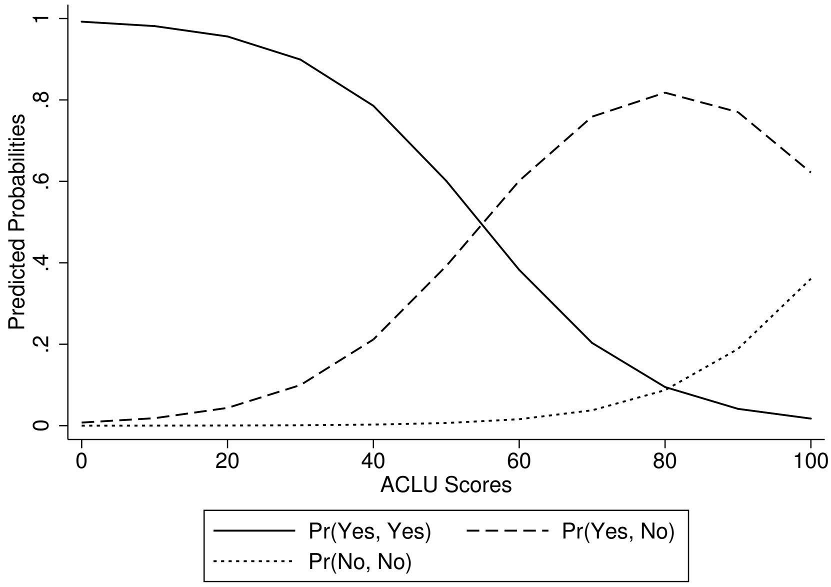

Figure 2.6: Predicted Probabilities of Flag Vote Categories (House-Safe Seats)

on a Representative's voting records under different electoral situations. In both safe and competitive districts, the probability being a consistent opponent of flag burning is maximized as general support for civil liberties is minimized. The probability of being a flipper is maximized among middling support of civil liberties, and the probability of being a consistent opponent of the flag burning ban is maximized at the highest level of support for civil liberties. Also, moving from a safe district to a competitive district shifts the cut points in the predicted direction. In other words, the estimates indicate that the increased commitment to civil liberties necessary to induce an increasingly anti-ban voting record is elevated in the case of a competitive district over the case of a safe district. These results also support the predictions of the theoretical models; indeed, they follow directly from the comparative predictions of the decision theoretical models presented earlier (Figure 3).

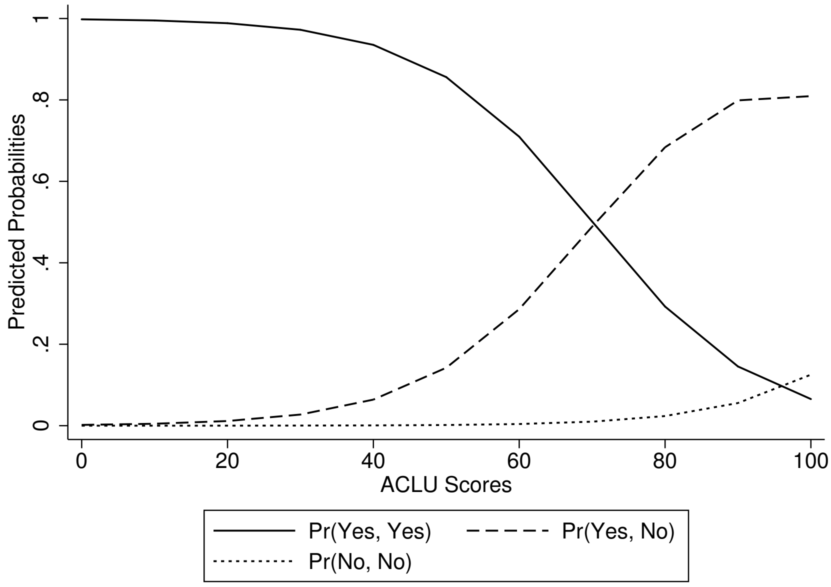

Figure 2.7: Predicted Probabilities of Flag Vote Categories (House-Competitive Seats)

2.6 Concerns

The process for ratifying a constitutional amendment complicates the interpretation of these results. While the intense public support for the flag burning amendment in 1989 and 1990 was likely an indicator of the amendment's ability to achieve approval in the states, a flag burning ban would not necessarily have been implemented despite a congressional vote "in the absence of judicial review." However, this should strengthen the interpretation of the results offered above. Since the expected utility from policy change under the amendment must have been attenuated by the small possibility that it would not have been ratified, legislators would have been less likely generally to have flipped on the flag issue at all. Thus, any observed variance in voting behavior has occurred despite the discount on policy change members of Congress may have applied to policy change in the amendment case.

Also, the observed patterns of vote changing analyzed here might be the result of some exogenous preference against amending the Constitution regardless of the policy associated with that act. This would produce some flips on any issue considered by Congress in both amendment and statutory form. While this may be the case, it is likely that a sincere personal preference for constitutional protectionism that is systematically related to policy preferences. In particular, preferences for constitutional protectionism, assuming that are not merely attitudinal pretenses for fighting a particular policy change, are not likely to be meaningfully correlated with the predictors in the empirical models presented here. In other words, while opponents of constitutional change are apt to invoke the sanctity of the Constitution to rally support their cause (even if they would willingly change the Constitution to protect their preferred policy outcome), it is unlikely that sincere constitutional protectionism would be collinear with both preferences over civil liberties and an elevated value for position taking.

In particular, the observed relationship between flipping and electoral sensitivity is a compelling result in this respect. Recall, Senators seeking reelection in 1990 are less likely to flip than others, as are members of the House in competitive districts. The constitutional protectionism counterargument has no facility to explain this result, which is directly predicted by the model advanced here. So, if constitutional protectionism has some effect on congressional voting behavior, it is most likely that this effect coexists with, rather than detracts from, the effects of judicial review on members' of Congress voting calculus.

Lastly, it is important to consider to what extent the model presented here is generalizeable, particularly in light of the unusual political circumstances that surrounded the case study used to assess the model. While the expectations about judicial review were unusually strong in the flag burning case, producing especially clear indications of the competing considerations facing members of Congress, the dynamic of legislative voting under judicial review is likely to emerge on any proposal that is constitutionally questionable. For example, many proposals involving campaign finance, abortion, religion, domestic terrorism investigations, and affirmative action fall into this category. The flag controversy of 1989 and 1990 is not a unique case of constitutionally questionable legislating, it is merely unique in the juxtaposition of viable statutory and constitutional proposals that emerged in Congress, providing leverage to assess the effects of this dynamic empirically. In principle, this model could be assessed in any circumstance where a two such votes have taken place.

2.7 Conclusions

The models and analyses reported here indicate that Thayer's critique of judicial review offers an apt account of legislative decision-making in cases where legislators may anticipate a judicial veto. This dynamic emerges because the institution of judicial review creates the possibility for legislators to cast "free votes," that is, to vote with reduced expected policy consequences. Specifically, the models predict that as the probability that a court will annul popular legislation increases, a proposal is likely to garner growing support in a legislature. Thus, judicial review is capable of altering the expected payoffs to legislators for supporting a proposal even in cases where a great deal of uncertainty over the likely outcome of litigation is present. This dynamic might affecting individual legislative voting behavior and, perhaps, the outcomes of relevant votes.

This theoretical result is supported by empirical analyses of congressional votes surrounding the issue of flag burning in 1989 and 1990. The results hold across all relevant pairs of votes and in both chambers of Congress. Together, these results indicate that judicial review creates opportunities for Congress to enact legislation that does not reflect its members' best policy judgements. Indeed, this study of legislative voting under judicial review demonstrates important but under-appreciated dimensions of judicial influence in the American political system.

Thayer's (1893) admonition that judicial review "drive[s] out questions of justice and right" in legislatures appears as valid in the modern period as in his own. Faced with the opportunity to shirk and delegate an unpopular task to the judiciary, members of Congress face tremendous political incentives to forgo voting on the basis of their private, policy judgments and follow the popular currents. Again, this dynamic is likely to emerge as part and parcel of the consideration of proposals in a number of policy domains that regularly induce judicial review: campaign finance, abortion, religion, domestic terrorism investigation, and affirmative action, for example. While these issues make up only a fraction of the business of Congress, they are among the most salient and important issues in contemporary politics. This study provides insight into an important dynamic of legislative decision-making on these issues, generated by judicial review, which has not received much attention from political scientists.

While this analysis is aimed at understanding a particular dynamic of legislative-judicial relations, it also speaks to a more general issue in the study of legislative behavior. Legislators, and perhaps members of Congress especially, pursue a variety of goals in their office. These goals may often pressure legislators in a competing fashion. Strategic politicians can and will take advantage of institutional structures, such as judicial review and, perhaps, the presidential veto, to pursue multiple goals simultaneously. Uncovering the mechanisms available to legislators to balance their complex policy and political motives is an important but underdeveloped topic.

Chapter 3

The Supreme Court and Media Attention to Homosexuality

Scholars have invested tremendous resources in examining the links between the Supreme Court and the larger political system. While some prominently conclude that the Supreme Court has little broad influence on public policy or public opinion (Rosenberg 1991), a growing literature has identified numerous paths of influence from the Court into the external political environment. For example, Franklin and Kosaki (1989) find that the Supreme Court's decision in *Roe v. Wade* polarized public opinion on abortion. Valerie Hoekstra (2000, 2003) finds that Supreme Court decision may influence the aggregate distribution of political attitudes in communities with strong attachments (and thus large information flows) to particular cases, such as the communities from which cases originated. And, of particular importance here, Flemming, Bohte, and Wood (1997; Flemming, Wood, and Bohte 1999) show that some Supreme Court decisions have drawn the media's attention to political issues involved in those cases.

In particular, Flemming, Bohte, and Wood (1997) offer a critical exploration of the relationship between Supreme Court decision-making and the media. They show that a subset of the Court's decisions on school desegregation, free speech, and religion increased the media's attention to those issues. Specifically, Flemming, Bohte, and Wood find the cases that "markedly rearranged the prior distribution of political benefits, either material or symbolic, for various segments of the population" (1997, p. 1247) produced the observed agenda-setting effects. In other words, Flemming and his coauthors generalize that cases which substantially change a *status quo* policy are likely to draw the media's attention to the issue involved in the case. That inductive claim can serve as a general theory of media responsiveness to the Supreme Court that can be tested out-of-sample-sample.

Here, I employ that theory of agenda-setting to develop hypotheses about the media's responsiveness to the Supreme Court's gay right's cases from 1990 to 2005. Specifically, I argue that those cases which expanded the scope of gay rights, i.e. those that "rearranged the prior distribution of political benefits," should accompany increases in the media's attention to homosexuality. Conversely, those cases that confirm an existing policy should have no effect on media coverage.

I test these hypotheses using monthly indicators of the level of attention paid to homosexuality by two stylistically divergent newspapers, *The New York Times* and *USA Today*, employing Box-Tiao (1975) intervention analyses to assess the influence of relevant court actions on these media time series. The results indicate that changes in the media's coverage of homosexuality largely conform to theoretical predictions. Significant increases in attention to homosexuality accompanied Supreme Court decisions that expanded gay rights, though the effects varied somewhat across media outlets, while the remaining cases had no effect on media coverage. The data confirm that the Supreme Court can be systematically influential in raising the media profile of an issue in American politics though media outlets may respond somewhat differently to these stimuli.

3.1 The Supreme Court, the Media, and Agendas

The idea that agendas and agenda control matter for determining political outcomes may be among the most widely held ideas in modern political science. Moving an issue or an alternative onto the agenda, that is to say, to the attention of those who are empowered to make policy changes, is a necessary condition for policy change to occur (Bachrach and Baratz 1964). For American national politics, numerous studies suggest, the national media's agenda—the issues chosen to fill pages of print and minutes of broadcast time—have a strong effect on which proposed policy changes receive serious consideration in the elected branches of government, and which are dead on arrival (e.g. Cook *et al* 1983; Dalton, Beck, and Huckfeldt 1998; Kingdon 1989; Iyengar and Kinder 1987). Thus, political scientists have frequently asked, *who sets the media's agenda*?

In part, the answer may be the Supreme Court. Flemming, Bohte, and Wood (1997) investigate the role of the Court in elevating the media's attention to some issues—specifically civil rights, free speech, and public displays of religion (1997) and civil rights, civil liberties, and environmental policy (See also Flemming, Wood, and Bohte 1999). Generally, the authors find that, of the Court's substantively important decisions in these issue areas (those listed in the *CQ Guide to the United States Supreme Court*), a small number have a significant influence on the media's systematic attention to the issues each case involved. The cases identified by Flemming, Bohte, and Wood are shown in Table 3.1, along with information about the case's date, subject matter, and the duration of the media effect attached to each case.

Explaining these results, i.e. why these cases produced changes in the media's attention to issues while other legally important cases did not, the authors argue that this subset of cases created important changes in the distribution of constitutional rights:

> Each decision markedly rearranged the prior distribution of political benefits, either material or symbolic, for various segments of the population. The

Table 3.1: Agenda-Setting Cases: 1947-1992

Case	**Decision Date**	**Issue**	**Effect Type**
Brown v. Board of Education	May 1954	Civil Rights	Permanent
Cooper v. Aaron	September 1958	Civil Rights	Permanent
Griffin v. County School Board of Prince Edward County	May 1964	Civil Rights	Transient
Texas v. Johnson	June 1989	Free Speech	Permanent
Illinois ex. rel. McCollum v. Board of Education	March 1948	Establishment	Permanent
Engel v. Vitale	June 1962	Establishment	Permanent
Lynch v. Donnelly	March 1984	Establishment	Transient

> issues involved in in all of these decisions were also highly affective. As a result, the decisions were extremely controversial at the time they were announced. The media participated in expanding the scope of system-wide conflict conflict by publicizing the initial decision and its implications. In each case, the Supreme Court's decisions sparked intense national debates that drew in new participants and expanded the scope of conflict through time. . . The issues involved in each decision opened wide ideological cleavages among political actors that remain until this day.

Other cases, while substantively important, may have confirmed an existing *status quo* or offered a marginal refinement to a previous "landmark" case that established or redistributed political benefits, thus failing to generate enough "news" to draw the media's attention.

This explanation of observed results amounts to a general theory of the media's responsiveness to the Supreme Court. Like all such claims generated from exploring patterns in one set of observations, though, Flemming, Bohte, and Wood's (1997) theory of media responsiveness to the Supreme Court requires out-of-sample validation. As Beck, King, and Zeng (2000) note, "[A]ll statistical analysts must be concerned about whether they are taking advantage of some idiosyncratic features of their data [when drawing inferences]. . . To guard against this problem. . . out-of-sample forecast accuracy is considered the gold standard for model assessment." In this case, an out-of-sample test requires identifying a previously unexamined issue space in which the Court has acted to both confirm existing group rights and "rearrange" political benefits and then examining media coverage of that issue to see if its behavior over time conforms to the theory's predictions.

I take up this task through study of the Supreme Court's influence on the media's attention to homosexuality. In particular, I examine the extent to which the Court's gay rights decisions drew systematic attention to homosexuality from 1990 through 2005.

"Gay rights" is a quintessential "affective issue" and fits directly into the theoretical paradigm advanced by Flemming and his colleagues. In the last decade and a half, homosexuality's role as an issue has grown in importance, emerging alongside abortion as a central "social" or "cultural" issue in national and state-level politics. Moreover, homosexuality is a policy domain in which courts are presumed to have been instrumental in determining policy outcomes, though no academic studies have investigated this conjecture.

Also, to capture the potentially divergent streams of coverage that exist in different media outlets, I have chosen two distinct print news sources, *The New York Times* and *USA Today*, to represent the range of news coverage of the relevant political issue, homosexuality. These newspapers both enjoy a wide, national readership though they embrace diverse styles and intended audiences. *The New York Times* is often regarded as the nation's newspaper of record, at least by political and social elites. Its content is heavily and consciously dominated by "hard news" and its writing is geared toward a relatively well-educated audience. *USA Today*, on the other hand, is famously written and illustrated to appeal to a mass public audience—human interest stories frequently share space with prominent news items. In many respects, these two newspapers represent poles in style and substance available in the American print media. These least similar cases should reveal any heterogeneity in various media outlets attention to a given political issue and thus, taken together, provide indicators for a robust test of Flemming, Bohte, and Wood's theory.

3.2 Identifying Relevant Cases

Not all Supreme Court decisions are likely to draw the media's attention to a given issue. Many cases involve obscure issues, legal technicalities, or controversies without broad policy consequences. On the other hand, a subset of cases may have important consequences for the media's attention to an issue. "[D]ecisions that overturn long-standing precedents, create new precedents or establish new rights, push the Court into new areas of constitutional law, or alter political relationships between individuals, groups, or institutions are the kind of decisions that presumably reshape national dialogues" (Flemming, Bohte, and Wood 1997, pp. 1230-31).

Defining the political rights and the social standing of gays, lesbians, and other sexual minorities is a relatively new issue in American national politics. Almost by definition, then, any Supreme Court decision that touches on the scope of gay rights creates a new precedent—either including or excluding homosexuals from some constitutional or legal protection. The relevant task for this study, then, is to identify the set of Supreme Court cases which involve some aspect of gay rights.

The subset of relevant cases can be identified through the application of a relatively simple criteria to the most widely used database of Supreme Court decisions. I analyze any Supreme Court decision that involves as a litigant a homosexual person or group or agency which advocates for gay rights which was decided between 1990 and 2005, the dates for which media coverage data are available. Specifically, I admit those cases recorded in Spaeth's United States Supreme Court Database (Spaeth 2005) which identify one or another parties to the case as a "homosexual person or organization."[1] In this instance, the Spaeth database is especially useful, since it identifies parties based on references in the Court's writings on the case. Parties are described using "terminology which places them in the context of the litigation in which they are involved" (Spaeth 2005, p. 23). Thus, database entries that identify one or another party as a "homosexual person or organization" do so because this identification is relevant to the issue before the Court. In other words, cases involving a party identified as a "homosexual person or organization" are cases involving an issue that makes the sexual identity of the party relevant: a gay

[1]Specifically, I use those orally argued cases for which the variables "PARTY_1" or "PARTY_2" are coded as "GAY."

right's case.[2]

This search reveals four relevant cases in the observed period: *Hurley et al v. Irish-American Gay, Lesbian, and Bisexual Group* (1995), *Romer v. Evans* (1996), *Boy Scouts of America (BSA) v. Dale* (2000), and *Lawrence v. Texas* (2003). The cases are summarized in Table 3.2. In two of these cases, *Hurley* and *BSA*, the Court held that the First Amendment protected private organizations from state laws that barred the exclusion of homosexual from "public accommodations," retracting state protections against discrimination. On the other hand, the remaining cases acted to protect or expand gay rights. *Romer* invalidated a voter-created amendment to Colorado's constitution which forbade identification of homosexuals as a protected class under antidiscrimination laws. *Lawrence* invalidated state sodomy prohibitions, overturning *Bowers v. Hardwick* (1986).

Flemming, Bohte, and Woods theory of media response to the Supreme Court leads to clear predictions about how coverage of homosexuality should respond to three of the four decisions. Clearly, the theorey predicts that *Lawrence*, which overturned a relatively recent precedent and greatly expanded the substantive and symbolic rights of homosexuals, should draw the media's attention to the issue of homosexuality. And, those cases that ratified the right of private organizations to exclude homosexual, *Hurley* and *BSA*, should have no effect on media coverage.

The fourth case, *Romer v. Evans*, is more complicated, though. Superficially, the Supreme Court created a policy change, invalidating Colorado's 1992 Amendment 2. That situation should produce a change in the media's attention to homosexuality. However, that voter-created clause had never gone into effect; its implementation had been enjoined by a federal district court almost immediately after its ratification by Colorado's voters. Also, the amendment had been declared unconstitutional by Colorado's Supreme Court in

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[2]This procedure is superior to using Spaeth's "ISSUE" variable. This variable might be used to identify subsets of cases which might include those relevant for this study, such as privacy and due process, but these identifiers would include many other cases that are not relevant, i.e. the "ISSUE" variable does not identify an exclusive class of gay rights cases.

Table 3.2: Gay Rights in the Supreme Court: 1990-2005

Case	**Decision Date**	**Holding**
Hurley et al v. Irish-American Gay, Lesbian, & Bisexual Group	June 1995	"The... application of the Massachusetts public accommodations law to require private citizens who organize a parade to include among the marchers a group imparting a message that the organizers do not wish to wish to convey violates the First Amendment."
Romer v. Evans	May 1996	"[Colorado's] Amendment 2 [prohibiting laws which provide antidiscrimination protection to homosexuals] violates the Equal Protection Clause."
Boy Scouts of America v. Dale	June 2000	"Applying New Jersey's public accommodations law to require the Boy Scouts to admit Dale [a homosexual] violates the Boy Scouts' First Amendment right of expressive association."
Lawrence v. Texas	June 2003	"The Texas statute making it a crime for two persons of the same sex to engage in certain intimate sexual conduct violates the Due Process Clause."

1993 and 1994 on Fourteenth Amendment grounds. Thus, the state of Colorado's appeal in *Romer* amounted to a final attempt to force the implementation of a policy change that had never taken effect. Given the murky nature of defining the *status quo* with respect to the case, there is no clear theoretical prediction about its relationship with media change.

3.3 Measuring Homosexuality in the Media's Agenda

The central measurement task in this study is to establish the level of media attention to homosexuality. Fortunately, previous studies of the role of national political institutions' influence on the media provide ample guidance in this task. In particular, Flemming, Bohte, and Wood's (1997) study of the judiciary's role in agenda setting is particularly germane, and it is in this study's mold, with an important modification, that I proceed.

Flemming, Bohte, and Wood (1997) show that a small set of important Supreme Court decisions have drawn the media's attention to issues, such as desegregation and prayer in public schools, that were not previously on the national political agenda, at least at the pitch at which they appeared after the Court acted.[3] They reach this conclusion through an analysis of a monthly count of news stories listed in *The Readers' Guide to Periodical Literature* containing at least one of a set of keywords that indicate articles that relate to the issues under analysis. In their own words, these authors choose *The Readers' Guide* as a data source because:

> [I]t surveys a wide assortment of general interest and specialized publications with a combined readership far greater than the circulation of any single newspaper...[and] *because of the size and diversity of markets served by*

[3]This finding contrasts with the well-known conclusion reached by Rosenberg (1991), who argued that Supreme Court decision-making does not significantly alter the media's coverage of particular issues. Flemming, Bohte, and Wood (1997) conclude that Rosenberg's result is a product of his reliance on an overly broad measure of media coverage of race-related stories, which failed to capture sharp increases in media coverage of segregated public education and racial segregation generally (p. 1229, n6).

> *the periodicals, the combined editorial emphases. . . can be assumed to be more representative of concerns on the systematic agenda than measures resting on selective, narrow samples of...media* [Emphasis added] (p. 128).

For this reason, *The Readers' Guide* has been a recurring source of article count data for political scientists studying the effects of institutional activity on the media's agenda (e.g. Baumgartner and Jones 1993; Flemming, Wood, and Bohte 1997; Rosenberg 1991).

This methodological choice, has important theoretical implications, though. The authors explicitly choose to examine agenda setting affects common to the broad spectrum of American media. This choice consciously, and for good reason in the context of a particular research question, ignores potentially important variance in issue attention that may occur within particular publications or sets of publications over time. Many differences exist among media with respect to geographic dispersion, audience sophistication, and competitive standing. These differences place unique economic pressures on each media outlet. Moreover, these influences may interact with the journalistic and ideological predispositions of editors and publishers to yield interesting variance across different media with respect to the type and volume of coverage offered to different issues at different times. For studies of agenda-setting, it may be advantageous to examine issue coverage across individual media outlets rather than to treat the media as a single unit.

I measure the media's attention to homosexuality, through a simple count of the mean number of daily stories in each month in *The New York Times* and *USA Today*, archived at Lexis-Nexis, that mention a set of keywords which indicate content related to homosexuality.[4] The two series stretch from January 1990 to December 2005 (Figure 3.1).

A cursory inspection of these time series indicates at least two differences between

[4]The keywords are: gay, gays, lesbian, lesbians, homosexual, and homosexuals. Interestingly, the most regular contamination of this procedure is annual surge in the number articles mentioning the keyword "gay" surrounding the anniversary of the American bombing of Hiroshima at the close of World War II. As such, articles that include the term "Enola" are excluded.

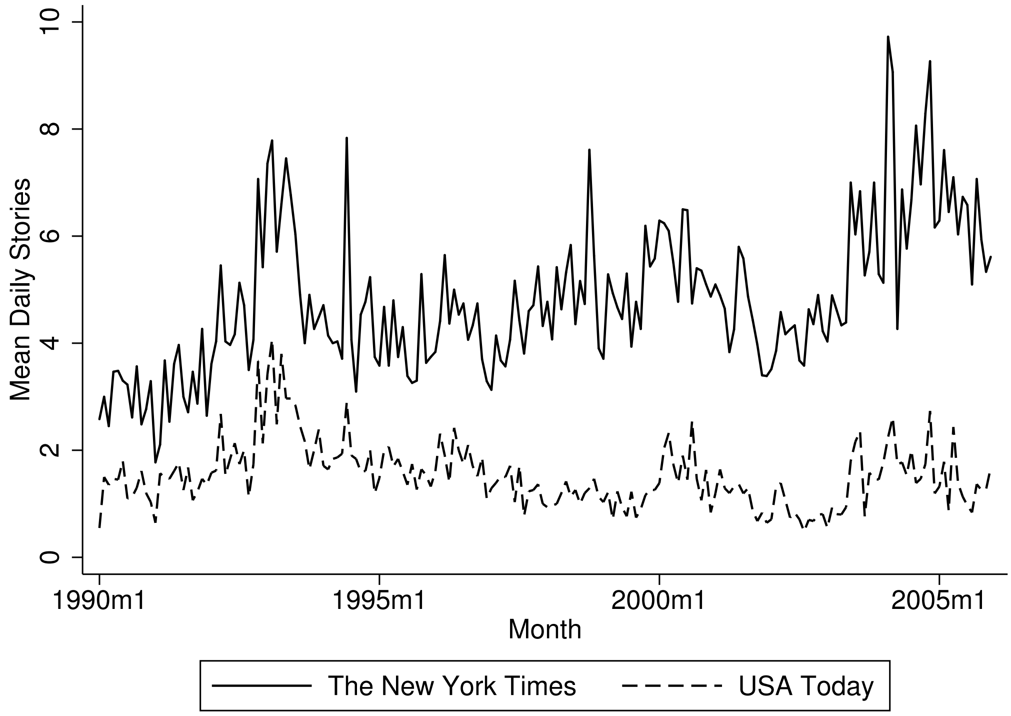

Figure 3.1: Media Coverage of Homosexuality: Monthly Mean Daily Stories

them. First, the absolute difference in magnitude between the two series is seemingly large. While *The New York Times* reports an average of approximately 4.7 stories per day that make some mention of homosexuality, *USA Today* reports an average of only 1.5; this difference is statistically significant ($p < 0.01$). In fact, *USA Today*'s attention to homosexuality exceeds the observed minimum of *The New York Times*' coverage, 2.1 stories per day, in only 20 of 187 months observed. More plainly, on any given day, *The New York Times* is likely to have published about 3 more stories involving homosexuality than *USA Today*.

Aside from the difference in the level of attention to homosexuality evident between these two series, there is also a marked difference in trends. *The New York Times* series has a relatively strong upward trend of approximately 0.013 daily stories per month.[5]

[5]The time series' trends are diagnosed by the estimation of the OLS model, $y_t = \alpha_0 + \alpha_1 t + \epsilon_t$, where α_1 indicates the estimated trend component. The detrended time series used here are obtained by removing

This amounts to an expected difference of nearly two stories per day from t_1 to t_{192}. In contrast, the *USA Today* time series exhibits a small, but significant, downward trend of roughly 0.003 stories per month. This reflects a reduction of roughly one half story per day over the observed time period. Substantively, these divergent trends reflect a growing heterogeneity in the media's treatment of homosexuality. In *The New York Times*, homosexuality has become an increasingly prominent issue over the 1990s and into the new century. Conversely, *USA Today* has moved (very) modestly away from coverage of homosexuality in terms of absolute attention to the issue.

These superficial differences, however, mask underlying similarities. First, absolute differences in coverage of homosexuality in the two papers fail to show that the coverage of this issue relative to other important policy topics is quite similar in each. Using the same article count procedure describe above, I produced a monthly time series indicating *The New York Times*' and *USA Today*'s coverage of the economy.[6] The resulting monthly time series, presented in Figure 3.2, correlate at 0.70. In turn, these series provide a comparative baseline for *The New York Times*' and *USA Today*'s coverage of homosexuality. The ratio of each paper's coverage of homosexuality to its coverage of the economy is presented in Figure 3.3.

USA Today's and *The New York Times*' coverage of homosexuality relative to their respective coverage of the economy are quite similar. In fact, the series correlate at 0.72. Of course, the data generating process in these time series is a joint function of the newspapers' coverage of both the economy and homosexuality, so it would be ill-advised to

---

the over-time change attributable to $\alpha_1 t$.

[6]The keywords are: unemployment and inflation. This keyword set almost certainly undercounts the coverage of the economy in each newspaper, leaving out stories on trade deficits, market performance, home sales, and other economic indicators too numerous to count. Thus, they provide a conservative baseline estimate of the coverage of general economic conditions in each of the observed newspapers. I have selected these keywords from among the wide-ranging possibilities because of the special role of these two indicators in short-term political economy vis-a-vis the Phillips Curve (e.g. Erikson, MacKuen, and Stimson 2002).

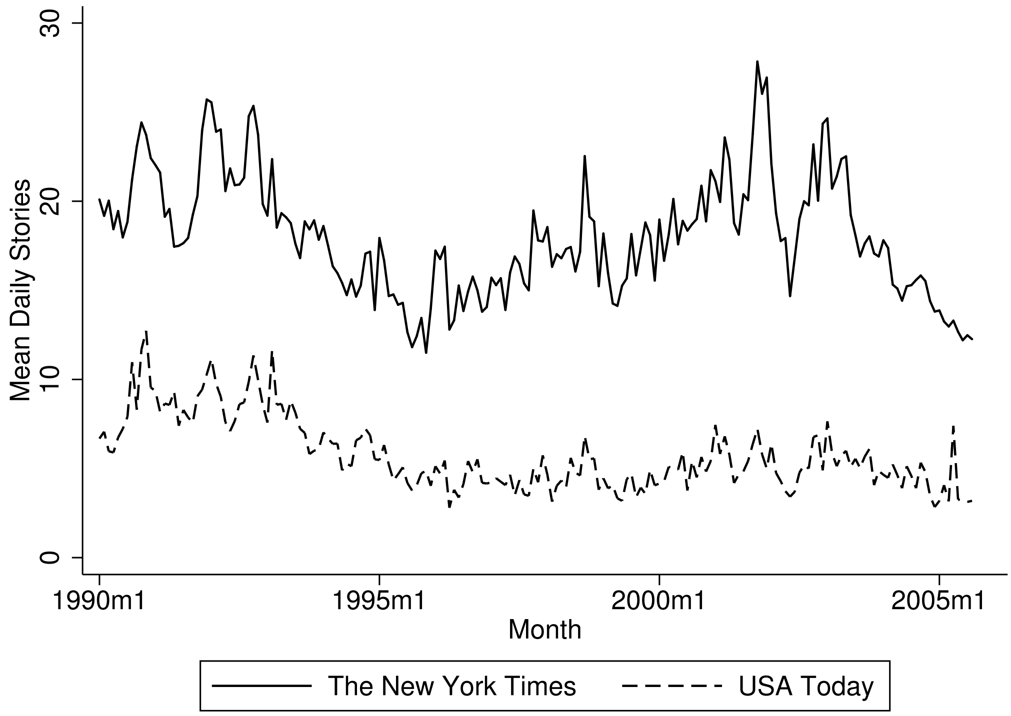

Figure 3.2: Media Coverage of the Economy: Monthly Mean Daily Stories

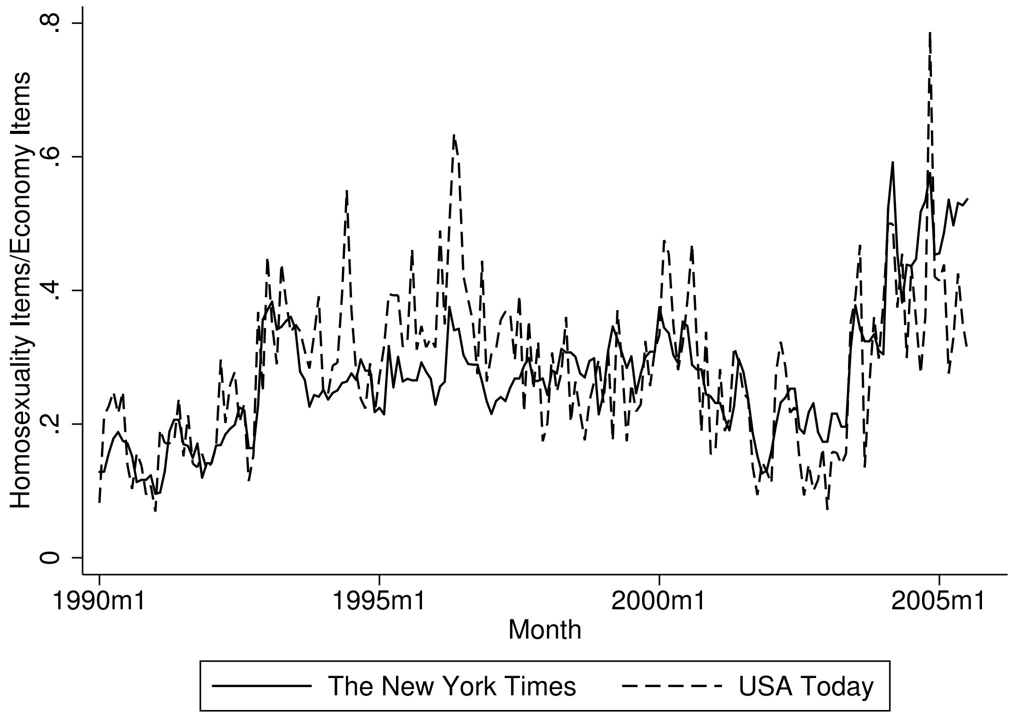

Figure 3.3: Ratio of Homosexuality Stories to Economy Stories

attempt to read too much into a superficial inspection of a complex system. Nevertheless, these data point to the fact that the newspapers examined here, which are explicitly constructed to appeal to different audiences and which are often taken by scholars and elites as embracing distinct points of view, share, at least in the period observed, similar patterns of coverage with respect to homosexuality.[7]

In addition to these emergent similarities in the newspaper's relative coverage of homosexuality, the trends in the raw time series also belie a common set of "high points" in the two newspapers coverage of homosexuality. In other words, a shared set of salient events appears to have drawn the media's attention generally to homosexuality. This common responsiveness to news-making events for *The New York Times* and *USA Today* may be seen more clearly in Figures 3.4, which reflect the first differences in the detrended monthly time series of mean daily news items involving homosexuality in each newspaper. Overlaid, the two series show a common systemic pattern: brief, common periods of large increases in coverage generally immediately followed by a short period of decreases in coverage. These patterns suggest a model of media coverage of homosexuality in which salient news events draw the media's attention to homosexuality, which produces the observed increases in media coverage. However, these increases appear, for the most part, short lived as the media moves on to cover other events and issues.

Of course, this does not imply that the observed trends in the data are unrelated to the the causal events listed above, or others. The increases in coverage over time may be related to the accumulated effects of identifiable events over time. Identifying these effects, their decay, or their accumulation, is an empirical matter, though.

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[7] I have made no attempt to code the content of the stories beyond their object. It may well be the case that these outlets cover the same issue in importantly different ways. Additional research would be necessary to investigate this speculation.

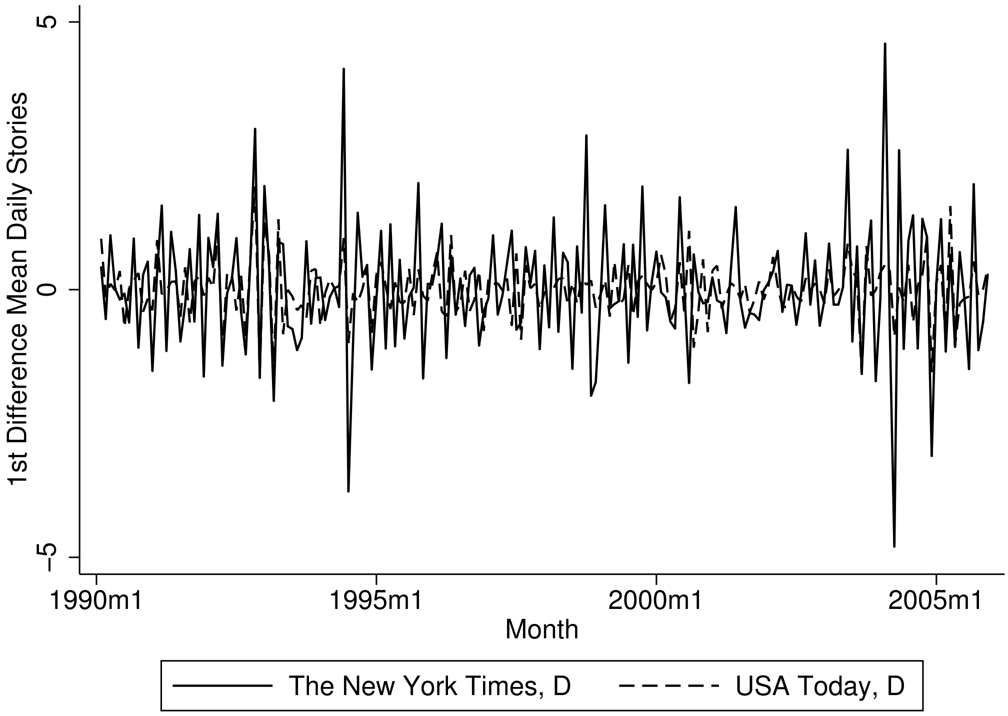

Figure 3.4: Media Coverage of Homosexuality: ΔMonthly Mean Daily Stories

3.4 Modeling Influence on the Media

This theory of agenda-setting can be assessed using Box-Tiao (1975) "intervention analyses," again the strategy employed by Flemming, Bohte, and Wood's (1997) study of the effects of Supreme Court decision-making on media attention, though applied to new data on a new issue. Box-Tiao analyses are used to estimate the effects of an event or events on a set of time serial observations of some quantity of interest. Specifically, Box-Tiao methods estimate the equation:

$$
\begin{aligned}
Y_t &= f(I_t) + N_t \\
& \quad where: \\
Y_t &= \text{ Media coverage time series,}
\end{aligned}
$$

$$I_t = \text{The effect of judicial "interventions," and}$$
$$N_t = \text{ARIMA error process.}$$

This design has a number of attractive properties. First, as a matter of internal validity, intervention analyses—Box-Tiao or some other formulation—are methods for assessing natural experiments. In this case, the "experiment" examines the effect of judicial decisions on media coverage of issues relating to homosexuality. While "the rival hypothesis exists that. . . some more or less simultaneous event may have produced the shift" in the dependent series (Campbell and Stanley 1963, p. 37), only the identification of a rival causal event threatens inferences drawn from the assessment of the impact of the event of interest.

Moreover, Box-Tiao analyses explicitly model ARIMA error processes (the model component N_t) by estimating the autoregressive and moving average components of the dependent time series (Box and Jenkins 1976). Thus, the effects of interventions on a time series are assessed once other sources of dynamic errors in that series have been accounted for with an appropriate ARIMA model, i.e. the dependent series is reduced to white noise prior to the estimation of intervention effects.

Thus, to begin, I develop an appropriate ARIMA model of each media coverage series, which have been logged to ensure variance stationarity. Next, I model the effects of each judicial decision as a dummy variable where the incidence of a judicial decision involving gay rights is indicated by a dichotomous variable that is set to zero prior to the relevant event and "switches on," changing to one, in the month that decision is announced. Pulse interventions—those which produce transient changes in the dependent series—are coded "on" for the month of the intervention and "off" for all other time periods. Step interventions—those which produce permanent effects in the time series—are coded zero for time periods prior to the event and one for all other time periods. Finally, I estimate transfer functions to assess the effects of the interventions on the media time series.

The flexibility of Box-Tiao analyses, though, produce difficult modeling choices for time series analysts. As Flemming, Bohte, and Wood (1997) point out, "[I]n Box-Tiao modeling, there is always some duplicity between specifications involving steps and pulses. For example, a first order pulse specification with a very slow rate of change parameter is virtually identical to a zero order step." These types of modeling choices are ubiquitous in Box-Jenkins-type analyses and decisions between alternative specifications must ultimately balance theoretical priors and model performance. This reflects the origins of the Box-Jenkins approach to time series analysis in applied forecasting problems. Here, final model specifications were derived in a stepwise approach. First, all events were modeled as zero order pulses. Events that demonstrate a significant influence were subsequently modeled with increasing complexity, adding decay parameters and treating the interventions as "steps" to the extent that the data supported such specifications. Events without significant influence, or added complexities that were unsupported by the data, were eliminated from the models. The final model estimates for *The New York Times* and *USA Today* are presented in Tables 3.3 and 3.4, respectively.

3.5 Results and Analysis

The data support the hypotheses derived from Flemming, Bohte, and Wood's (1997) theory of media responsiveness. Consistent with the hypotheses, the Court's decision in *Lawrence v. Texas* (2003), which invalidated a Texas statute and other similar state laws that criminalized homosexual sodomy, produced a significant, lasting (step) increase in coverage of homosexual content for both *The New York Times* and *USA Today.* Conversely, *Hurley* and *BSA* failed to produce a significant increase in coverage of homosexuality in either paper. Finally, though there were no clear expectations for this case, the Supreme Court's decision in *Romer v. Evans* (1996) had a temporary (pulse) influence on *USA Today*'s attention to the issue, though it had no effect on *The New York Times.*

I discuss the results for each media outlet more fully below.

3.5.1 *The New York Times*

The final model for *The New York Times'* coverage of homosexuality (Table 3.3) includes one of the four gay rights cases heard by the Supreme Court in the observed period, *Lawrence v. Texas* (2003). This case was best modeled as a step intervention; that is, the data indicate that *Lawrence* had a permanent effect on *The New York Times'* coverage of homosexuality. In the month preceding the Court's decision, *The New York Times* printed an average of 4.4 stories a day involving homosexuality in some respect. In the month the decision was announced, it printed 7.0 homosexuality related stories. This immediate increase was followed by asustained elevation in the paper's attention to that issue. Thus, even though there was a small recession of media coverage of homosexuality—following an immediate flurry of attention to homosexuality accompanying *Lawrence*—post-intervention coverage is significantly higher than pre-intervention coverage.

The magnitude of this effect should be assessed by transforming the logged coefficients into a quantity of substantive interest. One such quantity, the percent change due to the first order intervention, is given by:

$$\%\ Change = 100\left[exp\left(\frac{\omega_0}{1-\delta_1}\right)-1\right] \text{ (McCleary and Hay 1980, 174)} \qquad (3.1)$$

Using this transformation, the transfer function estimates predict that *Lawrence* produced a permanent 39% increase in *The New York Times* articles involving homosexuality. The negative first-order transfer function denominator for the *Lawrence* intervention indicates that the effect of the case on the media coverage series oscillates before settling into its equilibrium effect.

Table 3.3: SCOTUS Decisions and *The New York Times* Coverage of Homosexuality

Model Component	Parameter	Estimate
Lawrence v. Texas (Step)	ω_{12}	0.48*
		(0.15)
	δ_{11}	-0.45†
		(0.31)
First-Order Autoregressive (AR1)	ϕ_1	0.93*
		(0.06)
First-Order Moving Average (MA1)	θ_1	-0.67*
		(0.07)
Mean (Constant)	μ	1.51*
		(0.07)

*$p < 0.05$; $^{\dagger}p < 0.1$ (One-tailed tests). N=192

3.5.2 *USA Today*

The final model for *USA Today* appears in Table 3.4. The results indicate that two cases, *Romer v. Evans* (1996) and *Lawrence v. Texas* (2003) had a significant influence on the paper's coverage of homosexuality. The first case, however, was best modeled as a pulse intervention while the second was best modeled as a step intervention.

The Supreme Court's decision in *Romer v. Evans* (1996)—which invalidated a voter-created amendment to Colorado's constitution that forbade the state or municipalities from extending antidiscrimination protection on the basis of sexual orientation—temporarily increased *USA Today's* attention to homosexuality. In the month preceding the announcement of the *Romer* decision, *USA Today* ran an average of 1.4 stories a day that involved homosexuality in some respect. In the month of the decision, that figure increased to 2.4. The δ parameter of 0.75 indicates that *Romer*'s influence on this paper decayed at a relatively brisk pace. The model predicts that logged intervention effect of 0.42 decayed by slightly less than half in two months. In five months, the intervention effect had approximately a quarter of its initial influence, and the effect of *Romer* decayed almost completely by the eighth month. So, while *Romer* had a temporary influence on *USA Today*'s coverage of homosexuality, the case does not appear to have made a lasting influence on the prominence of homosexuality as an issue.

On the other hand, *Lawrence v. Texas* (2003) had a permanent effect on *USA Today*'s attention to homosexuality. In the month preceding the decision, the paper offered an average of 0.9 stories per day involving homosexuality; there were an average of 1.8 stories per day in the the month of the decision. Using the percent change transformation implemented above, the model predicts a permanent 37% increase in *USA Today*'s coverage of homosexuality. As with *The New York Times'* model, the negative first-order transfer function denominator suggests that *Lawrence* caused some oscillation in coverage of homosexuality before the equilibrium effect was achieved.

Table 3.4: SCOTUS Decisions and *USA Today* Coverage of Homosexuality

Model Component	Parameter	Estimate
Romer v. Evans (Pulse)	ω_{01}	0.42* (0.25)
	δ_{11}	0.75* (0.36)
Lawrence v. Texas (Step)	ω_{12}	0.55* (0.24)
	δ_{11}	-0.75* (0.19)
First-Order Autoregressive (AR1)	ϕ_1	0.90* (0.04)
First-Order Moving Average (MA1)	θ_1	-0.50* (0.08)
Mean (Constant)	μ	0.31* (0.11)

$*p < 0.05$; $^{\dagger}p < 0.1$ (One-tailed tests). N=192

3.5.3 Conclusions: The Supreme Court and Media Attention to Homosexuality

Here, I have sought to confirm and extend existing research that has demonstrated an agenda setting affect in the national media flowing from the Supreme Court by analyzing media coverage of homosexuality. In a period of fifteen years, stretching from 1990 to 2005, the Supreme Court heard, by the criteria employed here, four cases involving gay rights. Examining two stylistically distinct newspapers, *The New York Times* and *USA Today*, reveals that these media outlets' attention to homosexuality has been, in part, influenced by the Court. These results are an out-of-sample verification of Flemming, Bohte, and Wood's (1997) generalization that the media increases its attention to issues in which the Court's decisions produce important policy changes.

Both papers demonstrated a lasting (step) response to *Lawrence v. Texas* in 2003; each permanently increased coverage of homosexuality-related items by 37-39%. *Lawrence* substantially expanded the scope of gay rights in the United States. By invalidating homosexual sodomy statutes, the case "markedly rearranged the prior distribution of political benefits, either material or symbolic, for various segments of the population" (Flemming, Bohte, and Wood 1997, p. 1247). *Lawrence* also overturned a prominent privacy decision (*Bowers v. Hardwick* 1986), yielding a landmark change in the national legal status of gays and decriminalizing homosexual sexual. Moreover, the case hinted that the Supreme Court might be amenable to other claims of privacy rights by homosexuals. These rearrangements of political benefits are not only news in their own, yielding extensive news coverage of the decision itself, it provoked the mobilization of political resources to oppose further expansions of gay rights, particularly in the case of *Lawrence*, producing the changes in media coverage of homosexuality. *Hurley et al v. Irish-American Gay, Lesbian, and Bisexual Group* (1995) and *Boy Scouts of America v. Dale* (2000), on the other hand, produced no such effects. In both *Hurley* and *BSA*, the Court merely upheld a *status quo*, affirming the right of private organizations to exclude homosexuals. Because these cases yielded no change in the distribution of political benefits, these cases yielded less news value and provoked less reallocation of political resources to the issue of homosexual's political and social status than *Lawrence*.

Finally, the murky theoretical status of *Romer* is reflected in the data. While *USA Today* temporarily increased its coverage of homosexuality in the wake of the decision, this effect was short-lived while *The New York Times*' coverage of homosexuality was unmoved. This observed heterogeneity in the newspapers' responsiveness to *Romer* may be a function of their respective audiences. *The New York Times*, though it has a national audience, is, in the end, a local paper for New York City. Municipal concerns, New York state news, and regional news from the mid-Atlantic, are over-represented in its pages. *USA Today*, on the other hand, is an explicitly national newspaper. Though *Romer v. Evans* had implications for that status of homosexuals in Colorado and in other states likely to adopt similar constitutional measures, the case had little direct influence on New York, where such a provision had little chance of success. Though surely there are other plausible explanations for the observation that *USA Today* responded to *Romer* in a more significant way than *The New York Times*.

Still, the results reported here also offer a refinement to existing agenda-setting research by treating the media as a set of independent news outlets rather than as an undifferentiated whole. The heterogenous responses to some Supreme Court cases shown here demonstrate that, for a variety of economic and editorial reasons, various news outlets may not respond to the same stimuli in the same way. While evidence of systematic media responsiveness to some newsmaking events is important, to the extent that scholars might expect some media outlets' coverage to be more influential in a given political context, it is worthwhile to examine media content in specific providers or sets of providers rather than to assume away potentially interesting variance. And while modest differences of this sort emerge in this study of two print media outlets, more substantial and important variance might be uncovered across mediums—print, broadcast, and electronic. Exploring this heterogeneity, particularly in terms of reassessing established media effects findings in one or another medium, will undoubtedly prove fruitful for scholars.

Appendix A

Control Models of Legislative Voting

The independent variables included in the choice models of legislative voting contained in Chapter 2 are are limited to those derived directly from the theoretical models. However, the inclusion of additional control variables, reflecting legislators' partisanship and overall ideologies, does not affect the estimates of these theoretically relevant variables. In models estimated including party and first dimension DW-NOMINATE Scores (Poole and Rosenthal 1997), the parameter estimates on the explanatory variables included in the original models maintain their sign and significance. Moreover, for both the the House and the Senate estimates neither the overall fit of the models nor their predictive power are greatly increased by the inclusion of these additional variables. The full results of these extended control models are presented below in Tables A.1 and A.2.

Table A.1: Logit Analysis of Senate Flipping (Appendix)
(1 = "Flip")

Predictors	89 Amend't	90 Amend't
ACLU Rating	0.09***	0.16***
	(0.04)	(0.06)
1990 Reelection	-1.54**	-2.21**
	(0.89)	(0.97)
Party (Democrat=1)	-4.33**	1.53
	(2.36)	(2.26)
1^{st} Dimension DW-NOMINATE	-7.24	4.46
	(6.24)	(6.58)
Constant	-2.47	-8.75***
	(2.42)	(3.50)
χ^2	85.02	83.48
Pseudo R^2	0.68	0.68
PRE	0.81	0.75
N	90	91

Note: Logit Estimates. Standard Errors in Parentheses.
*** $p < 0.01$; ** $p < 0.05$; *$p < 0.1$
One tailed tests.

Table A.2: Ordered Logit Analysis of House Voting Cohorts (Appendix)
(Consistent Opponents=2, "Flippers"=1, Consistent Supporters=0)

Predictors	Effects
ACLU Rating	0.08***
	(0.01)
1988 Win Margin	0.00
	(0.01)
Competitive District	-1.33**
	(0.72)
Party (Democrat=1)	0.97*
	(0.69)
1^{st} Dimension DW-NOMINATE	2.06*
	(1.43)
Cut Line 1	3.87***
	(0.84)
Cut Line 2	8.55***
	(0.66)
χ^2	318.26
Pseudo R^2	0.48
PRE	0.56
N	396

Note: Logit Estimates. Standard Errors in Parentheses.
*** $p < 0.01$; ** $p < 0.05$; *$p < 0.1$
One tailed tests.

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