It Feels Like We're Thinking: The Rationalizing Voter and Electoral Democracy

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Abstract

The familiar image of rational electoral choice has voters weighing the competing candidates' strengths and weaknesses, calculating comparative distances in issue space, and assessing the president's management of foreign affairs and the national economy. Indeed, once or twice in a lifetime, a national or personal crisis does induce political thought. But most of the time, the voters adopt issue positions, adjust their candidate perceptions, and invent facts to rationalize decisions they have already made. The implications of this distinction—between genuine thinking and its day–to–day counterfeit—strike at the roots of both positive and normative theories of electoral democracy.

The primary use of party is to create public opinion.

—Philip C. Friese (1856, 7)

Cognitive Consistency, Partisan Inference, and Issue Perceptions[1]

The rise of scholarly interest in "issue voting" in the 1960s and '70s prompted concern about the implications of partisan inference for statistical analyses of the relationship between issue positions and vote choices. The spatial theory of voting (Downs 1957; Enelow and Hinich 1984) cast "issue proximity" as both the primary determinant of voters' choices and the primary focus of candidates' campaign strategies. The proliferation of issue scales in the Michigan (later, National Election Study) surveys provided ample raw material for naïve regressions of vote choices on "issue proximities" calculated by comparing respondents' own positions on these issue scales with the positions they attributed to the competing candidates or parties.

The ambiguity inherent in empirical relationships of this sort was clear to scholars of voting behavior by the early 1970s. Brody and Page (1972) outlined three distinct interpretations of the positive correlation between "issue proximity" and vote choice. The first, "Policy Oriented Evaluation," corresponds to the conventional interpretation of issue voting – prospective voters observe the candidates' policy positions, compare them to their own policy preferences, and choose a candidate accordingly. The second, "Persuasion," involves prospective voters altering their own issue positions to bring them into conformity with the issue positions of the candidate or party they favor. The third, "Projection," involves prospective voters convincing themselves

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that the candidate or party they favor has issue positions similar to their own (and, perhaps, also that disfavored candidates or parties have dissimilar issue positions) whether or not this is in fact the case.

Having laid out persuasion and projection as alternatives to the standard interpretation of issue voting, Brody and Page (1972, 458) wrote:

> The presence of these two 'alternate' processes in the electoral system makes it inappropriate to declare policy-oriented evaluations the cause of the correspondence between issue proximity and voting behavior. We need some means for examining the potential for 'persuasion' and for 'projection' and of estimating them as separate processes.

They proposed simultaneous equation estimation procedures employing "independent causal factors" identified on the basis of our theories of behavior and our knowledge about the act of voting. However difficult it is to specify such causal factors, that is exactly where the problem is. If the estimation of policy voting is important to the understanding of the role of the citizen in a democracy – and theorists of democracy certainly write as if it is – then any procedure which fails to control for projection and persuasion will be an undependable base upon which to build our understanding.

Brody and Page's clear warning was followed by some resourceful attempts to resolve the causal ambiguity they identified (Jackson 1975; Markus and Converse 1979; Page and Jones 1979; Franklin and Jackson 1983). Unfortunately, those attempts mostly served to underline the extent to which the conclusions drawn from such analyses rested on fragile and apparently untestable statistical assumptions. Perhaps most dramatically, back-to-back articles by Markus and Converse (1979) and Page and Jones (1979) in the same issue of the *American Political Science Review* estimated simultaneous equation models relating partisanship, issue proximity, and assessments of candidates' personalities using the same NES data, but came to very different conclusions about the bases of voting behavior. If two teams of highly competent analysts asking essentially similar questions of the same data could come to such different conclusions, it seemed clear that the results of simultaneous equation estimation must depend at least as much on the analysts' theoretical preconceptions and associated statistical assumptions as on the behavior of voters. Pending stronger theory or better data, the search for causal order in voting behavior seemed to have reached an unhappy dead end.

In the face of this apparent impasse, most scholars of voting behavior have adopted a simple expedient—reverting to single-equation models of vote choice, but with sample mean perceptions of the candidates' issue positions substituted for respondents' own perceptions (e.g., Aldrich, Sullivan, and Borgida 1989; Erikson and Romero 1990; Alvarez and Nagler 1998). This approach has the considerable virtue of reducing biases due to projection. On the other hand, it sacrifices a good deal with respect to theoretical coherence, since it is very hard to see how or why voters would compare their own issue positions to sample mean perceptions of the candidates' positions, ignoring their own perceptions of the candidates' positions. Moreover, this approach does nothing to mitigate biases due to Brody and Page's (1972) persuasion effect; to the extent that voters adopt issue positions consistent with those of parties or candidates they support for other reasons, they will still (misleadingly) appear to be engaged in issue voting.

Recent work by Lenz (2006) examining the basis of apparent priming effects suggests that persuasion may play a large role in accounting for observed correlations between issue positions and vote choices. Using panel data from a variety of cases in which previous analysts found (or could have found) apparent priming effects, Lenz showed that increases in the strength of the relationship between issue positions and vote intentions were driven almost entirely by the subset of respondents who learned the candidates' issue positions between survey waves. Moreover, the increased consistency between their own issue preferences and their vote intentions was mostly due to shifts in their issue positions to match their vote intentions, not to shifts in their vote intentions to match their issue positions. For example, in the 2000 presidential campaign, people who supported investing Social Security funds in the stock market and then learned the candidates' positions on that issue became no more likely than they had been to support George Bush; but people who supported Bush and then learned the candidates' positions became significantly more likely to favor investing Social Security funds in the stock market. As with earlier work by Abramowitz (1978), Lenz's work provides much more evidence of vote-driven changes in issue positions—persuasion—than of issue-driven changes in candidate preferences.

In this paper, we take up the topic of voter rationalization, aiming to give it a more nuanced and rigorous foundation by tying it to Bayesian models of voter rationality. In most respects, our theoretical agenda is very much in the spirit of Feldman and Conover (1983), who proposed what they referred to as "an inference model of political perception." They noted that the patterns of rationalization typically interpreted as reflecting cognitive dissonance reduction could also be interpreted as rational inference in the face of uncertainty:

> Rather than being motivated by a need to reduce inconsistency, people may simply learn that certain aspects of the social and political world are, in fact, constructed in a consistent fashion . . . [I]n the absence of information to the contrary, an individual's assumption that certain types of consistency exist may be an efficient way of perceiving the world.

Feldman and Conover (1983, 813) noted that "a theoretical focus on cognitive inference provides more than just a reinterpretation of consistency effects; it suggests a basis for developing a more general explanation of political perception." Their more general explanation involved accounting for perceptions of candidates' issue stands by reference to a variety of plausibly relevant political cues, including respondents' own issue positions and their perceptions of political parties and ideological groups. In subsequent work (Conover and Feldman 1989) they put a similar framework to particularly striking effect in accounting for the crystallization of perceptions of Jimmy Carter over the course of the 1976 presidential campaign. Using panel data gathered over the course of the election year, they showed that most people were quite uncertain of Carter's issue positions during the primary season, but shifted markedly toward associating him with the positions of the Democratic Party after he became the Democratic nominee.

Unlike Feldman and Conover, our focus is on a single potential source of political cues: party identification. On the other hand, we explore the ramifications of partisan inference for a variety of politically relevant perceptions, including matters of fact, perceptions of issue proximity, and people's own positions on specific political issues. Our model of rationalization suggests that all of these politically relevant perceptions should be subject to essentially similar processes of partisan inference.

Our approach also differs from Feldman and Conover's in drawing more explicitly upon the logic of Bayesian updating to structure our model of partisan inference. Feldman and Conover (1983, 817) stressed the importance of prior beliefs and noted that "the adjustment or change in the prior beliefs resulting from the perception of new information may be slight" in the case of well-known candidates and more substantial in the case of candidates who are relatively unknown. However, for any given candidate they represented issue perceptions as a linear function of the various relevant political cues provided by parties, ideological groups, and the respondents' own issue positions. In contrast, we derive a model of partisan inference in which Bayesian updating implies theoretically and politically significant non-linearities.

The resulting non-linear model bears important mechanical similarities to the non-linear model of issue perceptions proposed by Brady and Sniderman (1985). In their model, people attribute policy positions to political groups in an effort to balance two distinct psychological objectives: a desire for accuracy and "a strain to consistency" between perceptions and feelings (Brady and Sniderman 1985, 1068). On one hand, people are assumed to want to minimize the distance between their perception of the group's position and the group's actual position. On the other hand, they are assumed to want to minimize the distance between their perception of the group's position and where they would like the group to stand, given their own policy position and their general attitude toward the group. As a result, their perception represents a weighted average of the group's actual and hoped-for positions.

Perceptions in our model may likewise be interpreted as weighted averages of components representing reality and partisan considerations. However, we differ from Brady and Sniderman in thinking of the latter as reflecting Feldman and Conover's process of "cognitive inference" rather than the sort of "wishful thinking" suggested by Brady and Sniderman's affective language. Our perceivers draw upon partisan considerations in an effort to improve the accuracy of their perceptions, not in an effort "to bring perceptions in line with feelings" (Brady and Sniderman 1985, 1068).

The Model

Our model of voter inference makes the following assumptions (following Achen 1992; Bartels 2002; and others):

- At time n, a citizen is inferring two things —his expected future net utility difference between the parties $\hat{u}_{n+1}$ (which may be interpreted in a stable party system as party identification) and second, his estimated net difference $\hat{\delta}$ between the parties on some new issue, measured on a survey item scale common to all respondents.

- The citizen's current PID $\hat{u}_n$ is a weighted average of k previous issue scale scores δ_j : $\hat{u}_n = \sum_{j=1}^{k} \lambda_j \delta_j$, where the λ_j convert the scale scores to utilities. The convention $\sum_{j=1}^{k} \lambda_j = 1$ sets the utility units.[2] Thus $\hat{u}_n$ corresponds to the citizen's average partisan balance on the first k issues, weighted by the importance of the issue. It is thus scored on the same scale as the issue scales.

- Before considering the new issue, the citizen knows that at the next period his actual new utility will be $u_{n+1} = u_n + \lambda_{k+1}\delta_{k+1}$, and he wishes to estimate this quantity as accurately as possible. Since only λ_{k+1} and δ_{k+1} appear in the following discussion, we denote them simply by λ and δ.

[2] These "issues" might include economic retrospections, parental socialization, and other factors. As an analytic simplification, we treat all the old δ_j ($j \leq k$) as known.

- At time n, the citizen begins with his posterior distribution from the previous period for the utility difference on the old issues. The posterior is normally distributed with mean $\hat{u}_n$ and variance $\omega_n^2 > 0$.

- On the new issue δ, apart from any relationship to his PID, the citizen's prior is $\delta \backsim N(\delta_0, \sigma_0^2)$. This prior may not be entirely uninformative, as when the citizen uses past experience on related issues to forecast. ("I don't know what the current deficit is, but it's usually getting worse, so I'll guess that it's gotten worse lately, too.") The citizen may also have encountered some reported information about this issue $\bar{y}$, with likelihood $\bar{y} \backsim N(\delta, \sigma^2/n)$, where σ^2 is known. We interpret σ^2 as the variance in the reports themselves, while n is the amount of communication the citizen has received.[3] We assume that $\sigma_0^2 >> \sigma^2$, so that if substantial information about the new issue is known to the citizen, it rapidly swamps the prior. However, some issues may be hard to learn about, making the prior relevant for all but the most informed respondents.

- The citizen also has to learn the relevance of the new issue to his partisanship. Let $\gamma = \delta - u_n$. Thus the parameter γ measures "partisan deviance": The larger it is, the less similar is the scale score of the new issue to the citizen's PID. Since political parties organize the political issues, the variance of γ across issues, denoted τ^2, is relatively small. However, the citizen has to learn that. Based on his experience that most topics in life do not correlate with partisanship, he begins with a prior $\tau^2 \backsim \chi_{k_0}^{-2}(s_0^2)$, in which s_0^2 is large. In addition, the citizen may have experience with the deviation of k other issues from partisanship, summarized by the likelihood statistic $s^2 \backsim \chi_{k-1}^2(\ \tau^2)$. By standard Bayesian arguments, this prior and likelihood yield a poste-

[3] Even if the reports are purely factual, subjective variance in the utility of the issue might arise from a variety of sources. The citizen may be concerned that elites with views different from his own are inadvertently or deliberately misleading him, or the "facts" might be urban legends or reporting errors. Reported facts might also be correct but irrelevant to partisan utility calculations, as would occur if WMDs are absent from Iraq but have been hidden in Syria, as some Republican survey respondents currently believe.

rior for $\tau^2 \backsim \chi^{-2}_{k_0+k-1}(\hat{\tau}^2_k)$, where $\hat{\tau}^2_k = [k_0 s_0^2 + (k-1)s^2]/(k_0+k-1)$.[4] Thus as the citizen gets more information k, typically more weight in the posterior will be placed on the smaller number s^2, meaning that political issues are seen as tied more closely to partisanship. Thus the citizen's mean estimate of partisan relevance for issues will rise.

- The citizen may also have some direct personal information $\bar{x}$ about δ, with $\bar{x} \backsim N(\delta,\, s^2/m)$, such as having had an abortion herself when she answers a question about abortion or being gay when the topic is gay marriage. In such cases, s^2 may be very small, and this personal information may swamp everything else. For most citizens thinking about most political issues, however, their only information is derived from the statements of other people and groups, so that they have no direct personal information and $m = 0$. Hence we set aside this information source for now.

- Finally, all these distributions are taken to be jointly independent: Sampling errors on other issues are not correlated with those on the current issue, for instance, and an issue with, say, an unusual true mean does not disturb the citizen's random sampling to learn about it. Similarly, priors are independent across parameters.

Now the citizen needs to estimate what he should think about the utility balance on the new issue δ. Second, he needs to estimate what his new estimated PID $\hat{u}_{n+1}$ should be.

We proceed in four steps:

1. As an estimate of δ, $\hat{u}_n$ is approximately unbiased with a posterior variance of $\omega_n^2 + \hat{\tau}^2_k$, where as before, $\hat{\tau}^2_k = [k_0 s_0^2 + (k-1)s^2]/(k_0+k-1)$. (That is, the first term of the variance is the error in estimating the

[4] This likelihood would result if the citizen has taken a sample of k prior issues, each a draw from a normally distributed sample of issues whose utility is centered at the true partisanship u, and then had computed $s^2 = \sum(\delta_j - \bar{\delta})^2/(k-1)$, where $\bar{\delta}$ is the mean of the δ_j and where $E(\delta_j) = u$. We adopt this approximation, recognizing that for a variety of reasons including parental socialization, partisanship is not identical in practice to the mean of a citizen's issue views.

true u_n, and the second is the variance of δ around u_n. Those two errors are independent and so the variances add.) The statement holds approximately because we have conditioned on the mean of the posterior for τ^2 rather than integrating it out from the joint distribution with u.[5]

2. Hence to this order of approximation and by the usual Bayes normal theory with known variances, the citizen's best estimate of his position on the issue is:

$$\hat{\delta}|\bar{y} \approx \frac{\delta_0/\sigma_0^2 + \hat{u}_n/(\omega_n^2 + \hat{\tau}_k^2) + n\bar{y}/\sigma^2}{(1/\sigma_0^2) + 1/(\omega_n^2 + \hat{\tau}_k^2) + (n/\sigma^2)} \tag{1}$$

With a common prior, and for a fixed level of information and PID strength, this equation gives current issue position as a linear function of the prior issue mean δ_0, the PID $\hat{u}_n$, and issue information $\bar{y}$. Note that if partisan deviance $\hat{\tau}_k^2$ falls quickly with information, the weight on $\hat{u}_n$ will rise more rapidly than that on $\bar{y}$. Hence when the poorly informed prior is neutral but the new information $\bar{y}$ differs from partisanship, the relationship between issue opinion and information will be curvilinear: first neutral, then tending toward the partisan position, then finally turning away from partisanship toward the value of the new information.

3. For the citizen's best estimate of his new PID, we need to incorporate both the weighting λ and the posterior variance of $\hat{\delta}$, and similarly for $\hat{u}_n$. Taking the previous Equation (1) as exact and using standard Bayesian calculations (see appendix) gives:

$$\hat{u}_{n+1}|\bar{y} = \hat{u}_n + \lambda(\hat{\delta}|\bar{y}) + \frac{(\omega_n^2 - \lambda\sigma^2/n)(\bar{y} - \hat{u}_n)}{\omega_n^2 + \hat{\tau}_k^2 + \sigma^2/n} \tag{2}$$

This equation expresses the cross-lagged regression of current PID on

[5] The same result follows to the same degree of approximation from the formal Bayesian approach of considering the joint distribution of $\hat{u}_n$ and $\hat{\delta}$, and then integrating out the marginal distribution for $\hat{\delta}$.

lagged PID and the new issue. Note that even if $\lambda = 0$ (nothing about the new issue itself is incorporated into future PID), the coefficient on lagged PID is not necessarily unity nor the coefficient on the issue zero. Particularly if $\hat{\tau}_k^2$ is small (high partisan relevance), the new issue is informative about partisanship even if it does not affect PID directly.

Furthermore, setting $\Delta\hat{u}_{n+1} = \hat{u}_{n+1}|\bar{y} - \hat{u}_n$, we obviously have:

$$\Delta\hat{u}_{n+1} = \lambda(\hat{\delta}|\bar{y}) + \frac{(\omega_n^2 - \lambda\sigma^2/n)(\bar{y} - \hat{u}_n)}{\omega_n^2 + \hat{\tau}_k^2 + \sigma^2/n} \qquad (3)$$

so that for a fixed level of information and PID strength, the change in PID from the prior period depends linearly on two things—first, the new issue position, and second, the deviation of the new information about the issue from the prior PID.

4. The citizen's best estimate of the old issues is also updated (see appendix).

Some intuition about these mathematical results can be obtained by looking at extreme cases. Assuming that k and n rise with more information, we have the following results, beginning with the least informed voters and proceeding to the most informed:

No PID, no information Here $n = 0$, and $\hat{\tau}_k^2 \approx \infty$. Hence from Equation (1), the voter responds with the vague prior mean δ_0.

PID present, little information or partisan relevance Then k and n are small, making $\hat{\tau}_k^2$ large, and so the prior δ_0 will matter. There will be relatively little rationalization even though the voter needs help knowing what to think about the issue, and PID will be virtually unchanged. Thus sufficiently poorly informed partisans will not differ much in their opinions from similarly uninformed partisans.

PID and partisan relevance, no issue information Then $(\omega_n^2 + \hat{\tau}_k^2)$ is much smaller than σ_0^2, and $n = 0$. It follows that $\hat{\delta} \approx \hat{u}_n$ and $\hat{u}_{n+1} \approx$

$\hat{u}_n$. Thus nearly the entire issue response is rationalization, and PID is almost completely undisturbed.

Strong PID and partisan relevance, some information Then ω_n^2 and $\hat{\tau}^2$ are small, and if they are jointly sufficiently smaller than σ^2/n, then $\hat{u}_n$ will dominate the evaluation of the issue and also the revised PID. Partisanship will be largely retained and rationalization will be substantial, even though the voter is fairly well informed. This case applies particularly to those issues where the partisan relevance is more easily learned than the issue information, e.g., when the name of the president or his party is mentioned as part of the question.

High information Here n and k are both large, but since $\hat{\tau}_k^2$ is bounded below and ω_n^2 is fixed at time n, n eventually dominates. Hence the voter reports something close to $\bar{y}$ as his opinion, and updates his PID toward $\bar{y}$ by an amount dependent on how much he cares about the issue (λ)and the malleability of his PID (ω_n^2).

Very high concern, high information (race?) Then $\lambda \to \infty$ and $n \approx \infty$. It follows that in the limit, $\hat{\delta} = \hat{y}$ and $\hat{u}_{n+1} = \hat{y}$. (Partisan relevance does not matter asymptotically, though it can speed the updating when present.) Thus asymptotically, the only force at work is a (dramatic) rational updating of PID, and no rationalization of the issue position occurs. Less dramatically, people who care more will update PID more, as will those who have more information.

Partisan Inference and Perceptions of Fact

In principle, the processes of inference we have identified should affect perceptions of issues, candidates, and a wide variety of other political objects. However, the workings and implications of our model may be illustrated most clearly in the context of purely factual perceptions, where we have some hope of discerning the impact of a shared reality transcending the partisan inferences that color different individuals' views. Thus, we begin our empirical analysis by applying our model of inference to straightforward perceptions of fact.

It is worth noting that very few politically consequential facts are subject to direct, personal verification. If an ordinary citizen is asked whether the president is a crook, whether the unemployment rate is 4% or 8%, or whether a distant regime possesses weapons of mass destruction, her response will reflect a judgment cobbled together from various more or less pertinent and trustworthy sources, including news accounts, water-cooler conversation, campaign propaganda, and folk wisdom about the way the world works. It will be perfectly rational for her assessment of the inherent plausibility of alternative states of the world to be based, in part, on how well they square with her partisan predispositions.

Put in these terms, partisan inference sounds like a helpful heuristic – and sometimes it is a helpful heuristic. However, we believe it is unwise to jump from the premise that relying on inference processes is "rational in the sense of cutting costs and making a best guess about reality" to the conclusion that "the general contribution of inference processes to vote choice is a positive one" (Feldman and Conover 1983, 837). When partisan inferences pertain to matters of subjective value, it is hard to know how one might weigh the benefits and costs of constructing a logically consistent worldview. By observing the process of partisan inference at work in the realm of purely factual matters, we can see more clearly whether and how it actually contributes to the development of accurate perceptions.

We consider two factual questions included in the 1996 National Election Study survey.[6] One asked respondents whether "the size of the yearly budget deficit increased, decreased, or stayed about the same during Clinton's time as President?" The correct answer was that the budget deficit had declined dramatically during Clinton's first term – by more than 90%. However, as the survey responses summarized in Table 1 make clear, only one-third of the public recognized that the deficit had decreased, while 40% said it had

[6] Data from the NES surveys employed here, along with information about the design and implementation of the studies, are available from the NES website, *http://www.electionstudies.org*.

increased. Republicans were especially clueless: half said that the deficit had increased, while only one-fourth said that it had decreased.[7]

*** Table 1 ***

Responses to the budget deficit question are unusually well-suited to shed light on the processes of political rationalization that are our focus here. First, the question is straightforwardly factual; it would be very hard to argue that Republicans and Democrats have different views about the meaning of the phrase "yearly budget deficit" or different standards for assessing whether the deficit had increased or decreased. Thus, any difference in responses must logically be attributable to some process of rationalization or partisan inference rather than to differences in ideologies or values. Second, the actual trend in the budget deficit was well-publicized, and remarkably clear during this period: after increasing substantially under George H. W. Bush, the deficit shrank steadily and substantially during Clinton's first term – from $255 billion in FY 1993 to $203 billion in FY 1994, $164 billion in FY 1995, $108 billion in FY 1996, and $22 billion in FY 1997.[8] Third, because the 1996 NES survey included some respondents first interviewed in 1992, it is possible to categorize these people, as we have in Table 1, on the basis of partisan predispositions established before Clinton even took office, thus ruling out the possibility that their partisanship was an effect rather than a cause of their perceptions about the budget deficit.

For purposes of comparison, we also examine responses to another factual question in the 1996 NES survey, which asked respondents whether "over the past year the nation's economy has gotten better, stayed the same or gotten worse?" Responses to this question are summarized in Table 2. Here there seems to have been somewhat more consensus than on the budget deficit, with more than three-quarters of the respondents saying that the economy was somewhat better or the same. The responses also seem to be a good deal

[7] Here and elsewhere, we classify "leaners" on the traditional NES 7-point party identification scale as independents rather than as partisans.

[8] The very next question in the 1996 NES survey provides a good example of a factual question for which the correct answer is far from obvious. The question asked whether "the federal income tax paid by the average working person has increased, decreased, or stayed about the same during Clinton's time as President?"

more accurate than for the budget deficit question. Real disposable personal income per capita grew by 1.8% in 1996, while real GNP per capita increased by 2.5%; the unemployment rate was 5.4%. All of these figures represented improvements over the preceding year (1.6% real income growth, 1.4% real GNP growth, and 5.6% unemployment) and over the average figures for the preceding decade (1.3% real income growth, 1.7% real GNP growth, and 6.2% unemployment.) Thus, while it would have been unduly pessimistic to say that the economy had "stayed the same," saying that it was "somewhat better" would seem quite reasonable.

*** Table 2 ***

On the other hand, there is considerable evidence of partisan bias in the responses summarized in Table 2, as there was in Table 1.[9] Whereas half the Democratic respondents said that the nation's economy had improved, only one-third of the Republicans did. Meanwhile, Republicans were almost twice as likely as Democrats were to say that the economy had gotten worse.

Previous research has documented significant partisan biases in a variety of perceptions and evaluations of political figures, issues, and conditions (Fischle 2000; Bartels 2002a; 2002b; Erikson 2004). Thus, the fact that such biases appear in Tables 1 and 2 should not be surprising. What we hope to add here is a more detailed explanation of the nature of those biases derived from our model of partisan inference. Since our model implies specific, non-obvious principles for integrating objective information and partisan cues in formulating judgments about the political world, it offers some promise of providing both a more accurate account and a deeper interpretation of partisan biases.

A primary focus of our analysis is on the complex role of political information in partisan inferences. While it may seem intuitive to suppose that "Rationalization is probably greater for less-informed citizens" (Aldrich, Sullivan, and Borgida 1989, 132), recent work by Shani (2006) has provided a good deal of evidence to the contrary. Her analysis of responses to a variety

[9] As in Table 1, our classification of partisanship in Table 2 is based on responses from the 1992 NES survey. Obviously, it is impossible for these responses to have been influenced by perceptions of economic performance in 1996.

of factual questions produced "a clear bottom line: political knowledge does not correct for partisan bias in perception of 'objective' conditions, nor does it mitigate the bias. Instead, and unfortunately, it enhances the bias; party identification colors the perceptions of the most politically informed citizens far more than the relatively less informed citizens" (Shani 2006, 31).[10]

Our account of partisan inference implies that partisan predispositions and political information are likely to interact in complicated ways in any given case. For example, it suggests that well-informed Republicans should be especially conflicted on the issue of the budget deficit, since they were most likely to be exposed to objective information about the dramatic downward trend in the deficit (larger n), but also most likely to recognize the relevance of their broader political convictions for assessing the plausibility of a dramatic improvement in the deficit under a Democratic president (smaller τ^2). The relative magnitude of these effects is by no means obvious from the model. Either one may dominate at different levels of information. It turns out that they do.

Direct examination of how the responses of Republicans and Democrats varied with levels of political information provides additional grounds for caution. Figure 1 summarizes perceptions of the budget deficit among Republican and Democratic identifiers (classified on the basis of their responses to the 1992 NES survey) with varying levels of political information.[11] The effect of information within each partisan group is clearly non-linear, as is the partisan bias represented by the gap in perceptions between the two

[10] Shani's analysis included eight factual questions in the 2000 NES survey, including the budget deficit and national economy questions examined here. In seven of the eight cases she found substantial (and statistically significant) increases in partisan bias among well-informed respondents. These differences were largely unaffected by the introduction of statistical controls for differing political values or plausible demographic correlates of differing personal experiences.

[11] The curves presented in Figure 1 are derived from locally weighted (lowess) regressions using 30% of the data (50-60 survey responses) at each information level. Our measure of political information cumulates responses to a variety of factual questions (identifying prominent political figures, knowing which party controlled Congress, and so on) in each wave of the 1992-94-96 NES panel. Classifying respondents on the basis of party identification measured in 1996 produces very similar curves, suggesting that parallel analyses with cross-sectional data are unlikely to go too far astray.

groups for any given level of information. This provides a sharp contrast with most discussions of rationalization in the political science literature, which almost uniformly assume monotonic relationships—the more of X, the more of Y. Explicit theorizing demonstrates the limitations of intuition and directs attention to those aspects of the data where surprises can be found.

*** Figure 1 ***

Among the least well-informed respondents, neither objective reality nor partisan bias seems to have provided much structure to perceptions of the budget deficit. Uninformed Republicans and Democrats were slightly, and about equally, more likely to say that the deficit had increased than that it had decreased. Perhaps this tendency reflects a murky understanding that the budget deficit increased at some point in the past; perhaps it is a bit of prejudice based on folk wisdom. In any case, the views of Republicans and Democrats diverge as we move from the bottom to the middle of the distribution of political information; partisan inference seems to dominate throughout this range, since the widening gap owes at least as much to Republicans moving further from the objectively correct answer as to Democrats moving closer to it. The pull of objective reality only begins to become apparent among respondents near the top of the distribution of political information. Among the best-informed 10 or 20% of the public, even Republicans were slightly more likely to say that the deficit had decreased than that it had increased, and Democrats – untroubled by any contradiction between the facts and their partisan expectations – were very likely to recognize at least some decrease.

Figure 2, which summarizes the interaction of partisanship and political information for perceptions of the national economy, provides a rather different picture. As in Figure 1, there appears to be rather little structure in the perceptions of very uninformed people. The average perceptions of the most informed partisans are also fairly similar in the two figures, with Democrats quite likely to recognize an improvement and Republicans close to the neutral midpoint of the scale. However, the patterns between these extremes show little similarity. Perceptions of the national economy generally display less evidence of partisan bias among relatively uninformed people, but as much or more evidence of partisan bias among those in the upper half of the distribution of political information. For Democrats, the most notable learning seems to have occurred around the middle of the information scale, rather than in the upper third of the scale as in Figure 1. For Republicans, the marked non-monotonicity evident in Figure 1 is entirely absent from Figure 2, except for a slight downturn in perceptions at the very top of the information scale.

*** Figure 2 ***

To what extent can these complexities in the responses to the budget deficit and national economy questions be accounted for by our mathematical model of partisan inference? If we take n and k as proportional to Information and $1/\omega_n^2$ as proportional to Age, and if we denote $E(\bar{y})$ (the judgment of informed opinion) by Actuality (measured on the same scale as PID), then the nonlinear regression equation implied by Equation (1) is approximately:

$$\text{Opinion} = \frac{A + \text{PID}/(B_0 + B_1/\text{Age} \ + B_2/\text{Info}) + C(\text{Info})^D \text{Actuality}}{1 + 1/(B_0 + B_1/\text{Age} \ + B_2/\text{Info}) + C(\text{Info})^D} \quad (4)$$

This setup assumes that "no information" is coded zero.

Table 3 presents the results of our non-linear regression analyses of responses to the budget deficit and national economy questions using this specification. Each analysis includes six parameters capturing important aspects of the model of inference set out in Equation (1). The first of these parameters, A, corresponds to the prior belief δ_0 in Equation (1), expressed on the same scale as the observed survey responses.[12] B_0, B_1, and B_2 represent the variance (ω^2+τ^2) of the partisan inference based on $\hat{u}_n$. Since we expect the uncertainty of partisanship, ω^2, to decline with age, we include the reciprocal of age with weight B_1. Similarly, since we expect uncertainty about the relevance of partisanship, τ^2, to decline with information, we

[12] Since multiplying each of the variance terms $\sigma^2, \omega^2, \tau^2$, and σ^2 in Equation (1) by an arbitrary constant would leave δ unchanged, we normalize the model by setting σ^2 equal to 1.0.

include the reciprocal of information with weight B_2.[13]

*** Table 3 ***

The constant weight B_0 is intended to capture other sources of uncertainty in partisan inferences, including prior uncertainty about τ, $k_0 s_0^2$, and any offsets necessitated by our simple operationalizations of the age and information effects.[14] In light of our model, we expect B_1 and B_2 to be positive; in addition, logical consistency requires that the overall variance (B_0 + B_1/Age + B_2/Information) be positive.[15] Finally, the parameters C and D capture the extent to which better-informed people hear and comprehend a greater volume of information about the value of δ. The parameter C represents the greater exposure of better-informed people to the flow of information represented by n (or, more precisely, $n\sigma^2/\sigma^2$) in Equation (1), while the parameter D allows for non-linearity in the relationship between the flow of information on a particular issue and our general measure of political information. We rescale the information scores to range between 0 and 1; thus, the impact of information always ranges from 0 for the least informed people to C for the most informed people, regardless of the value of D. However, lower values of D imply more learning at lower information levels, while higher values of D imply that learning is concentrated near the top of the information scale.

Our estimation strategy also requires us to specify a priori an appro-

[13] We attempted to estimate the functional form of the relationship between political information and the partisan relevance parameter τ^2 using an exponential specification similar to the one employed for the relationship between political information and the learning parameter n. However, our data were uninformative about the precise form of this relationship: the estimated exponent was 1.64 with a standard error of 2.40. In light of this uncertainty, and for the sake of simplicity, we dropped the exponent, leaving us with reciprocal specifications for the effects of both age and information on partisan inference.

[14] For example, our simple reciprocal functional form implies that the uncertainty of partisanship declines by the same amount between the ages of 20 and 25 as between the ages of 50 and 100. If younger people learn more quickly or more slowly than this, relative to older people, the inaccuracy of our specification will be partly absorbed in B_0.

[15] All of the parameter estimates reported below satisfy this logical constraint for every respondent, with one exception. The parameter estimates in the third column of Table 3 imply a slightly negative estimated partisan variance for one respondent. He was in the 99th percentile of the information distribution, 21 years old in 1992, and a strong (Republican) partisan.

priate value for $\bar{y}$, which represents the relevant content of the objective information to which citizens were exposed.[16] In the case of the budget deficit question, the fact that the deficit declined by more than 90% during President Clinton's first term obviously implies that the objectively correct response was "decreased a lot," corresponding to a value of +50 on our budget deficit scale. Thus, our model implies that each respondent's perception of the budget deficit will be some weighted average of the constant (but unknown) prior belief A, her partisan predisposition (ranging from -50 for strong Republicans to +50 for strong Democrats), and the objectively correct value +50.

The parameter estimates presented in the first and third columns of Table 3 are based on the subset of respondents in the 1996 NES survey who were also interviewed in 1992, providing us with a baseline measure of partisanship unclouded by any consideration of Bill Clinton's performance as president. The parameter estimates presented in the second and fourth columns of the table are based on all the 1996 respondents, using their partisanship as measured in 1996. While we doubt that the potential bias in the latter approach is large enough to outweigh the greater precision due to having more than twice as many respondents, we present both sets of parameter estimates for purposes of comparison.

For the question about the budget deficit, the primary difference between the two sets of results presented in the first and second columns of Table 3 is that the weight attached to partisanship varied more with age and information for partisanship measured in 1992 than for partisanship measured in 1996. In other respects, the results are quite similar. In both sets of results, there is a fairly modest but clear negative bias evident in prior beliefs about the budget deficit; absent any other considerations, people's perceptions tended to fall about halfway between the "stayed about the same" and "increased a little" responses. In both sets of results, older and better-informed people seem to have relied more heavily on their partisan predispositions

[16] In principle, we could attempt to estimate $\bar{y}$ along with the other parameters of our model. In practice, however, $\bar{y}$ and C are so nearly collinear that we see little hope of persuading our data to distinguish between them.

to gauge the deficit's trajectory than younger and less-informed people did. And in both sets of results, the actual trajectory of the budget deficit clearly received some weight from well-informed respondents. The estimates of C imply that people who scored at the top of the information scale gave the positive reality (+50 on our 100-point scale) about 50% more weight than the negative prior belief (-10 or -13). However, the large positive estimates for the exponent D imply that the weight of reality increased very slowly over most of the range of our political information scale: for example, the implied weight for people at the midpoint of the scale was less than half of one percent of the implied weight for people at the top of the scale, while the implied weight for people in the 80th percentile of the distribution of information was less than 20% of the implied weight for people at the top of the scale. These results suggest quite strongly that very little real information about the trajectory of the budget deficit reached people below the very top reaches of our information scale.

Figure 3 provides a graphical representation of the extent to which the NES respondents seem to have incorporated the actual trajectory of the budget deficit into their responses to the question asking whether the deficit increased, decreased, or stayed the same during Clinton's first term. For each respondent, the figure shows the relative weight of real information implied by the parameter estimates in the first column of Table 3. For respondents in the bottom two-thirds of the distribution of political information this weight is effectively zero. For those in the upper third of the distribution it ranges upward to almost one-half.[17]

*** Figure 3 ***

Figure 4 provides a similar graphical representation of the extent to which respondents based their perceptions of the budget deficit on their partisan predispositions. As with the weights for reality, the range of weights here is from close to zero to about one-half. However, the distribution of

[17] The variation in weights for respondents at the same information level reflects the impact of age on the complementary weights attached to partisanship through the B_1 parameter. The estimates imply that older respondents at each information level attach more weight to partisanship, and thus less weight to real information about the budget deficit.

weights is quite different. For one thing, the estimated weights are much more variable at any given point on the information scale, reflecting the substantial impact of age on the apparent precision of partisan predispositions. In addition, whereas reality seems to have had virtually no effect on the responses of people in the bottom two-thirds of the information scale, many of these people – especially in the middle third of the scale – attached appreciable weight to partisanship in formulating their views about what had happened to the budget deficit.[18] On the other hand, the average relative weight of partisanship was actually less for people near the top of the information scale – those who responded appreciably to the actual trajectory of the budget deficit – than for those in the upper-middle range. People in the latter group seem to have known enough to recognize the relevance of their partisan predispositions for formulating responses to a question about how the budget deficit changed under President Clinton, but not enough to recognize how the budget deficit actually did change.

*** Figure 4 ***

Finally, we note that our non-linear model accounts for responses to the budget deficit question better than an analogous linear regression model employing the same explanatory variables and the same number of parameters.[19] It also captures much of the non-linearity evident in the relationship between partisanship, political information, and perceptions of the budget deficit in Figure 1. That fact is evident from Figure 5, which compares the average predicted responses implied by the parameter estimates in the first column of Table 3 with the actual average responses of Republicans and

[18] The average estimated weights for people in the bottom third of the information scale are 10% for partisanship and 0.002% for reality. The corresponding estimates for people in the middle third of the information scale are 21% for partisanship and 1.1% for reality. In each case, the remaining weight was attached to the general prior prejudice represented by the parameter A in Table 3.

[19] The standard error of the non-linear regression (with six parameters) presented in the first column of Table 3 is 27.47, and the R^2 statistic is .13; the corresponding average error in the same dependent variable for a linear regression including party identification, age, political information, and interactions between party identification and age and party identification and political information (and a constant, for a total of six parameters) is 28.44, with an R^2 statistic of .09. The other three non-linear regression models presented in Table 3 also produce better fits to the data than analogous linear regression models.

Democrats at each point on the information scale. There is some indication here that our non-linear model understates the extent of partisan inference among Republicans in the middle portion of the information scale and (correspondingly) the steepness of the upturn in the top third of the information scale. However, the model does seem to account with reasonable accuracy for the non-obvious patterns in the data.

*** Figure 5 ***

The parameter estimates presented in the third and fourth columns of Table 3 are derived from applying the same non-linear model to perceptions of the national economy in the 1996 NES survey. Again, we must specify an appropriate value for $\bar{y}$, the content of the objective information about national economic conditions available to the NES respondents. As we suggested above, available economic indicators suggest that the economy in 1996 was "somewhat better" than it had been a year earlier; thus, we set $\bar{y}$ equal to +25.[20]

As with perceptions of the budget deficit, we report separate results using 1992 partisanship (for respondents first interviewed in 1992) and 1996 partisanship (for both panel and fresh cross-section respondents in the 1996 survey). As with perceptions of the budget deficit, using the contemporaneous measure of partisanship reduces the apparent variation among respondents in the inferential weight of partisanship. However, in other respects the two sets of results are generally similar.

As with perceptions of the budget deficit, the estimates of the prior belief parameter A suggest that there was a slight pessimistic bias in perceptions of the state of the economy. However, the parameter estimates for parti-

[20] We examined the implications of this assumption by repeating the analysis reported in the third column of Table 3 with a variety of different values of $\bar{y}$. Higher values (implying that objective economic conditions were better than "somewhat better") improved the fit of the model; but these improvements were so slight (reducing the average error by no more than one-tenth of one percent) that we see no reason to abandon our a priori judgment regarding the substantively appropriate value of $\bar{y}$. For readers who may disagree, we note that the main effect of adopting a higher value of $\bar{y}$ is to reduce the apparent impact of objective information on perceptions of national economic conditions. That should not be surprising, since the perceptions reported in Table 2 are, on average, overly pessimistic even by comparison with our "somewhat better" standard.

san inference suggest a considerably larger information effect (B_2) and a considerably smaller (indeed, slightly negative) age effect (B_1) for perceptions of the national economy by comparison with perceptions of the budget deficit. Finally, and more importantly, the information effects implied by the estimated values of the C and D parameters are markedly different for the two questions. Information had a fairly modest impact on perceptions of the budget deficit, and that impact was highly concentrated among the best-informed respondents. By comparison, information about the actual state of the economy seems to have diffused much more broadly through the public. On one hand, the much larger value of the C parameter suggests that the weight of reality for the best-informed respondents was considerably greater than in the case of the budget deficit. On the other hand, the much smaller value of the D parameter suggests that less-informed respondents absorbed a much larger fraction of available information than in the case of the budget deficit.

The implications of these differences are very evident in Figure 6, which plots the implied weight of reality in perceptions of the national economy by information level using the parameter estimates in the third column of Table 3. The contrast with the analogous pattern in Figure 3 is striking. Whereas the actual trajectory of the budget deficit had virtually no impact on the perceptions of people below the top reaches of the information scale, the actual state of the national economy appears to have had a substantial impact on all but the least-well-informed respondents. For people at the middle of the scale, the estimated weight of reality is almost exactly equal to the estimated weight of uninformed prior beliefs; by comparison, for the same people on the budget deficit question uninformed prior beliefs received more than 100 times as much weight as the actual trajectory of the budget deficit.

*** Figure 6 ***

The apparent weight of reality in Figure 6 increases almost linearly over most of the information scale, but declines noticeably among people in the top quartile. The explanation for that decline is suggested by Figure 7, which provides a similar graphical representation of the extent to which respondents based their perceptions of the national economy on their partisan predispositions. Here there is a noticeable upturn among the best-informed respondents corresponding to the noticeable downturn in Figure 6. Again, the contrast with the pattern for the budget deficit is striking. Figure 4 suggested that the relative weight of partisanship on perceptions of the budget deficit peaked among moderately well-informed respondents, but declined as the weight of reality increased among those at the very top of the scale. For perceptions of the national economy, Figure 7 suggests a generally similar pattern, but with an upturn rather than a downturn among people in the top quartile of the distribution of political information.[21] One plausible explanation for this difference is that the national economy question made no explicit reference to partisan politics or to President Clinton, requiring respondents to supply that connection themselves in order to bring partisan inferences to bear. On the other hand, the budget deficit question asked about "Clinton's time as President," which may have encouraged people below the top reaches of the political information scale to connect their responses to their partisan predispositions.

*** Figure 7 ***

The Ramifications of a Partisan Shock: Reactions to Watergate

Having examined partisan inferences in a particularly simple setting where inferences focus on straightforward matters of fact, we turn next to documenting the impact of partisan inferences on a broader constellation of political perceptions. Our model implies that people's views about a wide range of specific political issues will be significantly influenced by their partisan predispositions. Unfortunately, cross-sectional data can shed little light on this hypothesis, since partisanship may be influenced by more specific

[21] A comparison of Figures 4 and 7 also clearly shows less variation in the estimated weight of partisanship among respondents at any given information level for perceptions of the national economy than for perceptions of the budget deficit. This difference reflects the much smaller impact of age on partisan inferences about the national economy (captured by the parameter B_1 in Table 3).

political views as well as influencing them. Even panel data may be of little help, since both partisanship and specific political views are likely to be quite stable over months or even years, aside from measurement error. And when they do change, considerable care is required to provide cogent causal interpretations for those shifts (Miller 2000).

In an effort to make headway in the face of these inferential difficulties, we focus here on an unusually dramatic sequence of political events that upset the existing equilibrium between partisanship and specific political views – the Watergate scandal. Fortuitously, for our purposes, the scandal was largely unrelated to substantive political issues of the day; there was no obvious reason, aside from partisanship, for people's responses to Watergate to be related to their views about school busing or government employment programs. Equally fortuitously, a large-scale NES panel survey bracketed the major events of the Watergate era, allowing us to observe how a variety of specific political views evolved in response to the escalating scandal, beginning with the run-up to the 1972 presidential election, continuing in the immediate aftermath of President Nixon's resignation in 1974, and ending with the 1976 election cycle.

Our model implies that if PID changed due to some external opinion shock unrelated to opinion on a second issue, then updating on the second issue will occur via the effect of PID on opinion. The latter effect will be small for those with low information (because they did not hear about the shock or did not grasp its partisan relevance). The impact of PID will be larger for those with more information.[22]

Our aim is to demonstrate that the shock to established partisan attachments created by the Watergate scandal reverberated in just the way our model suggests it should have. People's views about a variety of specific issues changed in ways that were statistically related – albeit logically unrelated – to their attitudes about the scandal. Moreover, these effects were concentrated among people who were especially well-informed about politics

[22] Only on issue such as abortion, where many well informed people have substantial personal information, would we expect no issue movement. The issues we consider do not have that character.

– in the top third of the distribution of political information. Those who responded most negatively to Watergate moved significantly to the left, and saw themselves significantly closer to the Democratic Party, on a variety of issues by 1976.

The 1974 NES survey included a variety of questions tapping respondents' reactions to the Watergate scandal, including whether they were pleased or displeased by Nixon's resignation,[23] whether they viewed the House Judiciary Committee's impeachment hearings as fair or unfair,[24] whether the media's coverage of Watergate was fair or unfair,[25] and whether the president's resignation was good or bad for the country.[26] We use responses to these four questions to construct a simple additive scale of Watergate attitudes, with scores ranging from -50 (for the most extreme pro-Nixon responses to all four questions) to +50 (for the most extreme anti-Nixon responses to all four questions). The scale has a mean value of 20.1, a standard deviation of 26.9, and an alpha reliability coefficient of .68.

Not surprisingly, reactions to the Watergate scandal were shaped in significant part by pre-existing partisan attachments. The mean Watergate scale value (in 1974) for people who had called themselves strong Republicans in the fall of 1972, when the origins of the break-in were still quite murky and the broader outlines of the scandal were not yet evident, was 0.6; the corresponding mean value for people who called themselves strong Democrats in 1972 was 29.3. On the other hand, there was also a good deal of variation in responses within each partisan camp. For example, almost one-third of the people who were strong Republican identifiers in 1972 had

[23] "Thinking back a few months to when Richard Nixon resigned from office, do you remember if you were pleased or displeased about his resignation, or didn't you care very much one way or the other?"

[24] "As you probably know, before Richard Nixon resigned, the Judiciary Committee was holding hearings to decide whether he should be impeached, that is, brought to trial in the Senate for possible wrongdoings. Would you say that these hearings were very fair, somewhat fair, somewhat unfair, or very unfair, or didn't you pay much attention to this?"

[25] "How fair would you say that the television and newspaper coverage of the Nixon administration's involvement in the Watergate affair was? Would you say it was very fair, somewhat fair or not very fair, or didn't you follow this very closely?"

[26] "Do you think that President Nixon's resignation was a good thing or a bad thing for the country?"

Watergate scale values below -20 in 1974, while another one-third had scale values above 20. Thus, it should be possible to distinguish the specific effects of reactions to Watergate from more general partisan differences.

We begin by examining the impact of Watergate attitudes on perceptions of relative proximity to the Democratic and Republican parties on a variety of political issues included in the 1972-74-76 NES panel – a summary liberal-conservative scale,[27] government jobs and income maintenance,[28] school busing,[29] rights of accused criminals,[30] and government aid to minorities.[31] We focus on these issues because self-placements and party placements were included in the 1972-74-76 NES panel.[32]

In order to test our assertion that partisan inferences should be concentrated among people sufficiently well-informed to recognize the potential ramifications of their partisan predispositions, the analyses reported in Table 4 are limited to respondents in the upper third of the overall dis-

[27] "We hear a lot of talk these days about liberals and conservatives. I'm going to show you a 7-point scale on which the political views that people might hold are arranged from extremely liberal to extremely conservative. Where would you place yourself on this scale, or haven't you thought much about this?"

[28] "Some people feel that the government in Washington should see to it that every person has a job and a good standard of living. Others think the government should just let each person get ahead on his own. And, of course, other people have opinions somewhere in between. Where would you place yourself on this scale, or haven't you thought much about this?"

[29] "There is much discussion about the best way to deal with racial problems. Some people think achieving racial integration of schools is so important that it justifies busing children to schools out of their own neighborhoods. Others think letting children go to their neighborhood schools is so important that they oppose busing. Where would you place yourself on this scale, or haven't you thought much about this?"

[30] "Some people are primarily concerned with doing everything possible to protect the legal rights of those accused of committing crimes. Others feel that it is more important to stop criminal activity even at the risk of reducing the rights of the accused. Where would you place yourself on this scale, or haven't you thought much about this?"

[31] "Some people feel that the government in Washington should make every possible effort to improve the social and economic position of blacks and other minority groups. Others feel that the government should not make any special effort to help minorities because they should help themselves. Where would you place yourself on this scale, or haven't you though much about it?"

[32] Our research design requires that we be able to compare responses before and after the Watergate scandal. In addition, the fact that these items were included in all three waves of the NES panel facilitates estimation of the statistical reliability of the responses.

tribution of political information.[33] Our parameter estimates are derived from errors-in-variables regression models, using estimates of the reliability of each explanatory variable within this high-information group.[34] To facilitate interpretation of the results, we also present precision-weighted averages of the parameter estimates across all five issues.

*** Table 4 ***

The first row of parameter estimates in Table 4 represents the impact of Watergate attitudes on changes in perceived issue proximity among highly informed respondents. The dependent variable in each case is perceived relative proximity in 1976 (ranging from -50 for people who perceived the Republican Party's position as identical to their own and the Democratic Party's position at the opposite end of the 7-point scale to +50 for people who perceived the Democratic Party's position as identical to their own and the Republican Party's position at the opposite end of the scale). The explanatory variables include the same relative issue proximity in 1972, party identification in 1972, and Watergate attitudes.[35]

The positive parameter estimates for Watergate attitudes indicate that, as expected, people who reacted especially strongly to the scandal tended to see themselves as closer to the Democratic Party, and further from the

[33] This division of the sample partly reflects our sense of the difficulty of the partisan inferences we are attempting to document here. However, it also represents a practical concession to the limitations of the NES data. Less-informed people were less likely to answer the issue questions we are analyzing here, and they were significantly more likely to drop out of the panel between 1972 and 1976. Thus, a more natural-looking division of the sample into two equal halves would leave too few usable cases in the bottom half to provide any realistic hope of finding Watergate effects among less-informed people even if they existed.

[34] For Watergate attitudes, our estimates of reliability are the alpha reliability coefficients derived from the correlations among responses to the four distinct survey items comprising our Watergate scale. For party identification, perceived issue proximity, and respondents' own issue positions, our estimates of reliability are derived from the correlations among responses to each item in the three waves of the NES panel using the measurement error model proposed by Wiley and Wiley (1970).

[35] We include lagged party identification to allow for the possibility that partisan predispositions in place by the time of the 1972 survey produced partisan rationalization on specific issues between 1972 and 1976. However, since our model does not specify the timing of the inferential processes we posit, we have no strong reason to expect such effects. In contrast, the timing of the Watergate scandal virtually ensures that its effects, if any, will be visible within the compass of the four-year NES panel.

Republican Party, on every issue by 1976. On the other hand, people who were relatively sympathetic to President Nixon in the immediate wake of his resignation tended to see themselves increasingly close to the Republican Party and far from the Democratic Party.[36] These estimates are fairly consistent across the five issues for which data are available, and in three of the five cases they are too large to be plausibly attributable to sampling error. Moreover, the implied effects are large enough to be politically consequential. For example, a difference of 35 points on the Watergate scale – roughly the difference between respondents at the 25th and 75th percentiles of the distribution – would correspond to a reduction in perceived distance from the Democratic Party of between two and six points on each of the 100–point issue proximity scales. (By comparison, the average total shifts on these scales from 1972 to 1976, including measurement error, ranged from 11 to 17 points.)

The changes in perceived issue proximity documented in the top panel of Table 4 could be attributable to either or both of the two processes of rationalization distinguished by Brody and Page (1972). On one hand, new (or more committed) Democrats may have projected their own issue preferences onto the party, while viewing Republican positions with a more dispassionate, or even actively critical, eye. On the other hand, they may have been persuaded to change their own issue positions, bringing them into closer alignment with their revised partisan sensibilities. The bottom panel of Table 4 focuses specifically on the latter possibility, estimating the impact of Watergate attitudes on respondents' own positions on the various issue scales included in the 1972-74-76 NES panel. The dependent variable in each case is respondents' issue positions in 1976, coded to range from -50 for the most conservative position on the 7-point scale to +50 for the most

[36] The negative intercepts in these regression models imply that people with scores of zero on the Watergate scale generally saw themselves as increasingly close to the Republican Party by 1976. That may seem odd, given that the Democratic presidential nominee in 1972 was widely viewed as being more ideologically extreme than usual. However, it is worth bearing in mind that a score of zero on the Watergate scale actually represents a relatively sympathetic response; only one-fifth of all respondents, and only half of strong Republicans, had negative scale values.

liberal position. The explanatory variables include the same issue position in 1972, party identification in 1972, and Watergate attitudes.

Here, too, there is surprisingly strong evidence that Watergate attitudes reverberated in seemingly unrelated corners of the political landscape. Those respondents who were most critical of Nixon gravitated to the left on government job guarantees, the rights of accused criminals, and school busing, while those who sympathized with him (or were critical of his critics in Congress and the media) became more conservative on those issues. As with the shifts in perceptions of issue proximity, the magnitudes of these shifts are considerable; a difference of one standard deviation in Watergate attitudes translated into a difference of from two to six points in the various 1976 issue positions. (By comparison, the average total shifts on these scales from 1972 to 1976, including measurement error, ranged from 12 to 25 points.)[37]

Table 5 provides analogous parameter estimates for respondents in the bottom two-thirds of the distribution of political information. In marked contrast to Table 4, there is very little evidence here of partisan inferences in the wake of the Watergate scandal. For perceptions of issue proximity (in the top panel of Table 5), only one of the five separate estimates (for school busing) is comparable in magnitude to the average estimated effect for well-informed respondents, and it is perversely negative. The average estimated effect for all five issues is almost exactly zero. For respondents' own issue positions (in the bottom panel of the table), there is one sizable positive estimate (for aid to minorities), but the average estimated effect across all five issues is only about one-third as large as the corresponding average estimated effect for people in the high-information stratum, and even that effect is too imprecisely estimated to be considered reliable.

[37] As the parameter estimates for 1972 issue positions in Table 4 make clear, well-informed respondents' views about government jobs were considerably less stable than their views about other issue positions between 1972 and 1976. We interpret this instability as reflecting a shift in the debate about whether the government should try to provide every person with "a job and a good standard of living," from McGovern's controversial proposal to give $1000 annual grants to every man, woman, and child in 1972 to discussions of more modest public works programs in 1976.

*** Table 5 ***

In short, just as our formal model suggests, less-informed people seem to have lacked the contextual knowledge necessary to translate the partisan shock of Watergate into partisan inferences about the seemingly unrelated issues we have examined here. Unlike people in the high information group, those in the low information group apparently saw no reason to revise their understanding of specific political issues in response to the unmaking of the president.

The most obvious potential objection to the evidence presented in Tables 4 and 5 is that the same people who were most affected by the Watergate scandal might have become more liberal between 1972 and 1976 for entirely other reasons. Reactions to the scandal were correlated with a variety of characteristics beyond partisanship and ideology; for example, better-educated people were especially pleased to see President Nixon go, whereas southerners were somewhat more critical than non-southerners were of the House Judiciary Committee and the news media. If better-educated people were becoming more liberal during this period, or southerners were becoming more conservative, their views about Watergate may have been only spuriously related to those ideological shifts. To assess that possibility, we replicated the regression analyses presented in Tables 4 and 5 including a variety of demographic characteristics – including age, education, income, race, region, gender, marital status, home ownership, union membership, and church attendance – as additional control variables. The key results of these elaborated analyses are presented in Table 6, along with the parallel results from Tables 4 and 5.[38]

*** Table 6 ***

The results presented in Table 6 generally confirm those presented in Tables 4 and 5. Not surprisingly, the parameter estimates from the elaborated regression models are somewhat less precise than those presented in the earlier tables. Nevertheless, both the magnitude and the consistency of our apparent Watergate effects hold up nicely in the presence of these extensive demographic controls. As in the simpler analyses presented in Table 5,

[38] Complete results of these analyses are available from the authors.

there is rather little evidence here of changes in issue positions or perceived issue proximity among people in the bottom two-thirds of the distribution of information about the political world. Our data are not sufficiently powerful to rule out the possibility that these relatively uninformed people engaged in partisan inference to some modest extent. For the most part, however, the contextual grasp of politics necessary to make an inferential leap from Watergate to economic and social policy seems to have eluded them.

On the other hand, as in the simpler analyses presented in Table 4, there is a good deal of evidence in Table 6 that well-informed people changed both their perceptions of issue proximity and their own views about a variety of logically unrelated issues in response to the Watergate scandal. If anyone had asked these well-informed citizens to explain the changes in their thinking about school busing or government employment programs, we suspect that they would have provided rationalizations of exactly the sort posited by Rahn, Krosnick, and Breuning (1994, 592), "mentioning reasons that sound rational and systematic and that emphasize the object being evaluated, while overlooking more emotional reasons and factors other than the object's qualities." The overlooked factor in this case, we argue, was the exogenous partisan shock of a Republican president's disgrace and forced resignation. The observable ramifications of that exogenous partisan shock among politically attentive people were surprisingly broad and consistent, and thus provide considerable empirical support for the theory of partisan inference we have set out here.

The Dynamics of Abortion Attitudes

In 1973, a divided U.S. Supreme Court ruled that American states could not forbid a woman to have an abortion during the first trimester of her pregnancy. The Court also decided that states could regulate abortion during the second trimester and could forbid it during the final three months. This famous case, *Roe v. Wade*, and related judgments ratified what many states had already done (Rosenberg, 1991), but were well ahead of public opinion in others.

Liberalized abortion laws set off a backlash among cultural and moral traditionalists, including many conservative Catholics, but eventually embracing many Protestant evangelicals as well (Hanna 1979, chap. 5; Balmer 2000). A counter–mobilization by abortion liberals ensued. Bitter struggles in courts and legislatures began, along with struggles to win over public opinion.

Initially, the Democratic and Republican parties were both internally divided on the issue. However, the legal battles began to polarize the leadership and activists of the political parties in the late Seventies and early Eighties (Adams 1997; Carmines and Woods 1997). The 1976 Republican platform, with some waffling, began to lean in the pro–life direction, and the 1980 version clearly declare its opposition to abortion. Subsequent GOP platforms strengthened the language.[39] By the late Nineties, the abortion opinions of ordinary Democrats and Republicans diverged as well. For example, in the Youth–Parent Socialization Panel Study (Jennings and Niemi 1991) of people who were high school seniors in the spring of 1965 , the correlation between abortion attitudes and party identification was only .07 in 1982, but it rose to .22 in 1997. Among the best–informed citizens during the same period, the correlation rose from .04 to .36.[40] (A broad–ranging

[39] 1976: "The Republican Party favors a continuance of the public dialogue on abortion and supports the efforts of those who seek enactment of a constitutional amendment to restore protection of the right to life for unborn children."

1980: "While we recognize differing views on this question among Americans in general–and in our own Party–we affirm our support of a constitutional amendment to restore protection of the right to life for unborn children. We also support the Congressional efforts to restrict the use of taxpayers' dollars for abortion."

1984: "The unborn child has a fundamental individual right to life which cannot be infringed. We therefore reaffirm our support for a human life amendment to the Constitution, and we endorse legislation to make clear that the Fourteenth Amendment's protections apply to unborn children. We oppose the use of public revenues for abortion and will eliminate funding for organizations which advocate or support abortion."

Subsequent years have used language close to that of 1984, with the most recent adding a condemnation of "partial birth abortion."

[40] "Party identification" is the seven–point Michigan scale. "Abortion attitude" is the respondent's choice among four alternatives ranging from forbidding abortion entirely to allowing abortion on demand. "Best–informed" denotes those who scored 5 or 6 on the six–point interviewer's assessment of political knowledge. Lastly, the original coding of the abortion scale was reversed: Here positive correlations indicate that respondents' abortion positions tend to be compatible with their PIDs.

review of the empirical research is Jelen and Wilcox 2003.)

Almost uniquely among issues, abortion attitudes are remarkably stable over time. They easily stand comparison with party identification, the customary gold standard for attitudinal stability. In the Youth–Parent sample, for example, party identification correlates .63 between 1982 and 1997, while abortion attitudes correlate at .59 over the same fifteen–year period. Moreover, among 935 respondents, just nine people lacked an abortion opinion in 1982, and only twelve in 1997, remarkably low for political attitudes. The number who lacked an abortion opinion at both time periods was *zero*. Where abortion is concerned, the overwhelming majority of people know what it means, they know what they think, and drastic change is rare. No surprise, then, that as the parties have polarized on the issue, it has come to play a prominent role in election campaigns. Catholic Democratic politicians often face criticism from local bishops when they embrace pro–choice positions.

Thus the causal link from abortion attitudes to voting and party ID seems obvious to observers of American politics, and the many of the customary tests seem to confirm it (see the review in Jelen and Wilcox 2003, 294-296). However, as we explained earlier, these tests confound issue voting and rationalization. Do people vote Republican because they are conservative on abortion? Or are they conservative on abortion because they are Republicans? No one doubts that there is *some* issue voting where abortion is concerned, but how much after rationalization has been removed? Few have considered the possibility that abortion attitudes are attitudes like other attitudes, and thus are influenced by wishful thinking and cognitive dissonance reduction.

Thus abortion raises an important challenge for those who want to explain the inter-relationship between party identification and issues. Unlike the factual questions we have discussed, abortion attitudes are not novel subjects about which most citizens are only mildly informed. Nor are they perceptions of candidate or party positions, where carefully cultivated ambiguity by the objects of perception make rationalization easy. Instead, abortion attitudes are well formed. Thus if we can find evidence that party membership changes abortion attitudes, the argument of this paper will be confirmed in a challenging case. For that finding would demonstrate that people are taking ethical advice about well known, painfully difficult moral problems from politicians, hardly the customary source for wisdom of that kind. And in turn if that is so, then the optimistic interpretation of learning from party cues (as in Page and Jones 1979 or Feldman and Conover 1983) would need serious rethinking.

Students of public opinion and voting behavior have been documenting patterns of partisan inference for more than half a century. Berelson, Lazarsfeld, and McPhee (1954) showed that "social distance" colored perceptions of group voting norms (chap. 5), that partisanship colored perceptions of candidates' issue positions (chap. 10), and that exposure to like-minded family members, friends, and co-workers reinforced partisan voting predispositions (chap. 7). (For the specific impact of party ID on issues, the corresponding source is Cambell *et al.* 1960, especially chap. 6.) In recent years, abortion attitudes have also been interpreted as caused in part by party membership (Layman and Carsey 2001; Wilcox 2001). However, prior work has used lengthy statistical specifications with all variables having linear, additive effects constant across individuals. While much can be learned from explorations with such specifications, we argue here that the causal effects are different in different groups, not only in size but in shape. Above all, they are not all linear. Hence a different statistical approach is needed, one that relies heavily on the inferential structure provided by model building.

Before beginning, we mention two considerations arising in studying abortion attitudes. The first is that Catholics differ from other Christians on this topic. The Roman Catholic Church has long opposed abortion, and when *Roe v. Wade* was handed down, most Catholics were more conservative on the issue than the average Protestant. As evangelicals have taken up the abortion issue, however, the average Protestant has moved somewhat right and many Catholics well to the left, so that the two groups have now become similar (Jelen and Wilcox 2003, 492). A well defined theological and political left and right wing have developed within Catholicism, a division that goes unmeasured in standard survey questions about denominational membership. Divisions on abortion are correlated with departures from Catholicism, making membership in the Catholic Church endogenous. (See Figure 8.) Pro–life Catholics have dropped out of the Church, whether for that reason or for reasons correlated with it. The consequence for our purpose is that the causal paths for Catholic abortion attitudes are unique in complex ways, and so we study only non–Catholics in the following graphs. Catholics deserve a separate study.[41]

*******Figure 8*******

Second, even a cursory look at abortion attitudes quickly uncovers the greater stability and coherence of women's opinions relative to men's. Even among the best informed, women are more stable. In the 1992-1994-1996 NES panel study, for example, non–Catholic men who fall in the upper 30% of the population for general political knowledge have, according to the Wiley–Wiley model for measurement error in over–time attitudes, abortion attitudes with a reliability of .84. The corresponding reliability for well–informed women is .97.[42] Hence in the following tables, we analyze men and women separately.

How can we sort out the direction of causation between abortion attitudes and PID? Simultaneous equation estimation is not likely to work well here for lack of suitable instruments. Cross–lagged regressions are also unappealing in this context (see Appendix 2). Hence we proceed differently, exploiting the implications of our model to get leverage.

Under our model, for an issue like abortion, where partisan divisions became more salient during the Eighties, we should see greater movement between parties among well informed women than among well informed men, since in that case opinion affects PID in proportion to how much the

[41] We found much weaker effects of this kind for "neo–fundamentalist" denominations, as the Jennings–Niemi study characterizes them. To the extent that these denominations overlap with evangelical churches (and of course the two groups are related but quite distinct theological categories), grouping evangelicals with other non–Catholics is safer statistically than including Catholics. Similarly, although black Democrats sometimes express firm pro–life attitudes, probably for reasons related in part to race, we found that excluding them generally made little difference.

[42] Gender differences are smaller at lower levels of information.

citizen cares about the issue. By contrast, we should see changes to accord with party position more among men than among women when both are moderately informed, since women are better informed and better able to resist the partisan rationalization. There should be little or no gender difference among the poorly informed.[43]

To illustrate our logic, we begin by setting out the simple bivariate relationships between PID and abortion attitudes in the Parent–Child Socialization panel. Abortion opinions were not asked until the 1982 and 1997 waves, so we analyze those years exclusively. Figure 9 shows the percentage of 1982 non–Catholic Republicans (strong or weak identifiers on the classic Michigan scale) who remained Republican in 1997, displayed as a function of their 1982 abortion attitudes. Men and women are shown separately. The two most conservative positions on the abortion scale ("never" and "rarely") have been combined because "never" is chosen by just 4% of this group.[44]

******Figure 9*******

The figure suggests a simple interpretation: Both male and female Republicans were more likely to leave the party if they held liberal abortion views, but the issue was more important to women, they understood the party's stand better, and it affected them more. The impact is not small. Almost half of 1982 pro–choice non–Catholic Republicans had disappeared from the party by 1997.

The importance of political knowledge can be seen in Figure 10, where the same information is displayed, this time limited to those in the upper 30% of the political information scale.[45] Here the sample sizes are small (41

[43] Note that our approach here is consistent with the model set out above. In the Youth–Parent dataset, everyone is the same age, which proxies for stability of PID ω_n^2 (Converse 1969). Hence if we examine homogeneous groups of citizens whose PID $\hat{u}_n$ is the same, and if we take partisan relevance $\hat{\tau}_k^2$ and information about abortion σ^2 as monotonic functions of general political knowledge, then Equation (1) for abortion opinion reduces to just a function of information levels.

[44] There are 98 women and 87 men represented in Figure 13, roughly equally distributed across the three categories of abortion opinion. This implies that as a rule of thumb, differences between cells are statistically significant at conventional levels when they reach seven percentage points or more.

[45] We use the six–point information scale from the 1982 Parent–Child Socialization Study.

men and 18 women), so that the results can be only tentative. However, there is a suggestion here that well informed men act like the average women when it comes to abortion: They see the connection to party and act accordingly. This effect is the first factor making the parties more consistent internally on abortion.

******Figure 10*******

The more interesting question is the reverse effect: How does party membership affect abortion attitudes? In particular, we would like to assess whether moderately informed and well informed men (who are nonetheless more poorly informed than the corresponding women) behave like the rationalizing Republicans assessing Clinton's deficit numbers we discussed above. Women, on the other hand, should be more stable and less persuadable. Note that this is the reverse of the earlier finding, in which women were changing more than men.

With attitudes generally trending to the left in the society, those who are already pro–choice in 1982 will tend to stay put. Republicans will face two balanced forces—the society and their party–while Democrats will be subject only to pro–choice pressures. Neither is likely to move very much.[46] Thus the question can be addressed more easily among 1982 pro–life citizens, where Democrats will tend to move away and Republicans to remain where they were. Independents should be in the middle. And least obviously, if our model is correct, the movement for Democratic men should be larger. Of course, there will be some pseudo–movement due to measurement error, so that no group will be entirely stable. But the basic pattern should hold.

Figure 11 shows opinion change to pro–choice views ("sometimes" or "always") among those with 1982 pro–life opinions ("never" or "rarely"), expressed as a function of their 1982 party ID. The figure shows exactly what one would expect from the model: Democrats move away from their old views the most, then Independent, and Republicans the least. Moreover, the effect is larger for men than for women, just as expected.[47] Again the

[46] There is some movement among 1982 pro–choice Republican women toward pro–life positions by 1997, but we have not yet been able to determine whether this is more than the measurement error in their 1982 views averaging out in subsequent years.

[47] There are few 1982 non–Catholic, non–African–American Democrats in the pro–life

effects are not small. More than half of 1982 male pro–life Democrats had become pro–choice by 1997. This effect is the second factor explaining the growing partisan division on abortion.

******Figure 11*******

In summary, the abortion data from one long–term panel show the patterns expected from our model. Women know more about abortion and care more. Hence when the parties diverge, they will disproportionately tend to change their parties rather than their views. Well informed men will act the same. Men as a whole, however, will have lower levels of information and will be more susceptible to rationalization and thus to influence by their parties. They will disproportionately tend to change their views rather than their parties. Now of course, most people in each gender stay put on both party and abortion views. And plenty of people in each gender exhibit each pattern. But the disproportionate effects are just what one would expect if rationalization plays a large role

The Rationalizing Voter and Electoral Democracy

The literature on "heuristics" in political science is an odd stepchild of the corresponding literature in psychology. Psychologists devote exhaustive attention to the biases in judgment produced by reliance on identifiable heuristics. For example, the classic collection of essays edited by Kahneman, Slovic, and Tversky (1982) includes reports on "belief in the law of small numbers," "shortcomings in the attribution process," "egocentric biases in availability and attribution," "the illusion of control," and "overconfidence in case-study judgments," among other topics. It also includes a series of essays on "corrective procedures" intended to mitigate the effects of these various biases and shortcomings.

Political scientists, on the other hand, typically view "heuristics" as a boon to democracy, helping ordinary people "achieve a kind of rationality

categories—just 14 men and 14 women. Hence percentages for Democrats must be interpreted with care. We hope to develop additional evidence in a subsequent draft, including a modification of our nonlinear model to cope with multiple stages of Bayesian updating.

generally adequate to the tasks of citizens," as Kuklinski and Quirk (2000, 153-154) wrote in a penetrating critique of this literature. What accounts for the disciplinary difference? We suspect that it has much to do with the fact that "the notion of a competent citizenry is normatively attractive. It buttresses efforts to expand citizen participation and credits the citizenry for some of American democracy's success" (Kuklinski and Quirk 2000, 154).

The work we have reviewed here on the bases of political perceptions conveys a good deal of enthusiasm for the "efficiency," "consistency," "reasonableness," and "rationality" of the perceptions resulting from processes of partisan inference. For example, Brady and Sniderman (1985, 1075) conclude their study of attitude attribution by acknowledging that

The mass public surely does not command much abstract knowledge of politics and as a rule does not even pay much attention to it. It seems implausible, therefore, to suppose that the general public is able, or at any rate willing, to assemble complex cognitive hierarchies of political ideas. In contrast, likes and dislikes are easy to form and, even more important, easy to remember. Accordingly, affect can be a quite efficient way of encoding and storing what is after all the most vital political information: who and what one is for or against.

Efficient, perhaps. What is striking, though, is that Brady and Sniderman have very little to say about the implications of this efficiency for the accuracy of people's perceptions of the political landscape. At one point (1985, 107) they assert that "the mass public is remarkably accurate in attributing positions to strategic groups on major issues"; but this assertion appears to refer to the average attributions of the public as a whole rather than to the judgments of individuals. In any case, it is far from clear what contribution, if any, their "likeability heuristic" makes to the accuracy of issue perceptions.

Brady and Sniderman reported estimates of the relative weight of "likeability" from 34 separate regression models, each focusing on perceptions of a particular group on a particular issue. Their results suggest that people's perceptions of disliked groups were relatively accurate, on average, but that perceptions of favored groups were strongly biased by people's desire to see those groups as close to themselves. Indeed, for a typical favored group, people's own issue positions received about one-third as much weight as the group's actual position in shaping people's perceptions of where the group stood.[48]

It is worth emphasizing, too, that Brady and Sniderman's results are based on perceptions of very salient groups (liberals and conservatives, blacks and whites, Republicans and Democrats) on major political issues of the day. It seems safe to assume that projection effects would loom even larger for less familiar groups or candidates and for less prominent issues. Indeed, an analysis along similar lines of perceived issue positions of candidates in presidential primaries (Bartels 1988, 98-107) does suggest even larger projection effects, especially early in the primary season and for candidates who are relatively unknown. For example, Bartels estimated that Democrats in the 1984 primary campaign perceived frontrunner Walter Mondale as about 20% closer, on average, to their own issue positions than he really was; but the corresponding distortion for challenger Gary Hart was about 40% at the beginning of the campaign, only gradually declining to a similar 20% level. For people who were particularly enthusiastic about Hart, for whatever reason, the estimated projection effect was even larger. People who gave Hart the warmest possible rating on the NES "feeling thermometer" at the beginning of the primary season managed to see him as almost 75% closer than he actually was to their own issue positions.

More straightforward evidence of the consequences of partisan inference appears in Figure 12, which provides a simple tabulation of the average perceptions of issue proximity for Republicans and Democrats on the spend-

[48] Brady and Sniderman's model includes two terms measuring projection: a "false consensus" effect applying to all groups regardless of whether they are liked or disliked, and a "more focused, or partisan, effect" pulling perceptions of favored groups toward one's own position and pushing perceptions of disfavored groups away from that position. The average magnitudes of Brady and Sniderman's statistical estimates are .185 for the false consensus effect and .550 for the differential projection effect. Assuming a difference of 25 points between favored and disfavored groups on the NES feeling thermometer (roughly the observed average difference between the groups Brady and Sniderman considered), the combined effect is minimal for disfavored groups (+.05) but considerable for favored groups (+.31) by comparison with the effect of groups' actual positions, which is normalized to 1.0.

ing/services scale in the 2004 NES survey.[49] Obviously, if perceptions of the parties' positions were unbiased, the curves in Figure 8 for Republican and Democratic identifiers would overlap perfectly. Instead, they are markedly divergent, especially for people whose own positions do not happen to fall at the midpoint of the 7-point scale. Even in the absence of any real knowledge about where the parties actually stood on this issue, it is very clear from the figure that someone's heuristics have gone badly astray.

*** Figure 12 ***

Table 7 provides a tabulation of the average perceived proximity for Democrats, Independents, and Republicans at each point on the 7-point scale. The partisan differences between Democrats and Republicans, in the last column of the table, range from one to six points, implying errors of perception for one or both parties' positions amounting to a considerable fraction of the total length of the scale. As a result, even people who should have seen marked conflicts between their own issue positions and their partisan attachments seldom did. For example, Republican identifiers who placed themselves in one of the two most liberal positions on the spending/services scale still saw themselves as closer to the Republican Party than to the Democratic Party, on average, while Democratic identifiers who placed themselves in one of the two most conservative positions nevertheless saw themselves as closer, on average, to the Democratic Party.[50] Are biases of this sort conducive to effective electoral democracy? We see no reason to think so.

*** Table 7 ***

Not every issue produces errors in perception as large as those displayed in Figure 12. It should not be surprising that the issue of abortion provides a notable contrast in this respect. The impact of partisanship on perceptions of issue proximity for that issue is displayed in Figure 13, and the corresponding tabulations of partisan differences appear in Table 8. There

[49] "Some people think the government should provide fewer services, even in areas such as health and education, in order to reduce spending. Other people feel that it is important for the government to provide many more services even if it means an increase in spending. Where would you place yourself on this scale, or haven't you thought much about this?"

[50] The average proximity scores for these groups are +.49 (with a t-statistic of +2.7) and -1.38 (with a t-statistic of -1.9), respectively.

is certainly evidence of partisan bias here, especially for people whose own positions were on the pro-life side of the issue. Nevertheless, most pro-choice Republicans and (by a narrower margin) most pro-life Democrats managed to recognize that their views put them closer to the opposing party than to their own party on this issue.

*** Figure 13 ***

*** Table 8 ***

Lest readers wonder which of these two patterns is more common, we can report that analogous figures for a variety of other issues and years look qualitatively similar to the pattern presented in Figure 12 and Table 7. For example, Figure 14 and Table 9 summarize perceptions of relative proximity for 13,647 NES respondents who placed themselves and the Democratic and Republican parties on the 7-point liberal-conservative scale at some point between 1972 and 2004. Just as in Figure 12, there is a dramatic divergence in perceptions between Republican and Democratic party identifiers on both ends of the ideological spectrum. Liberal Republicans saw the Republican Party's position as no less satisfactory than the Democratic Party's position, while conservative Democrats perceived the Democratic Party as just as close to their own positions as the Republican Party.[51]

*** Figure 14 ***

*** Table 9 ***

We see no reason to doubt that these perceptions are the product of partisan inference processes similar in kind to those we have examined here. However, their patent inaccuracy seems to us to belie the "encouraging" conclusion that "the general contribution of inference processes to vote choice is a positive one" (Feldman and Conover 1983, 837). Voters may indeed "do the best they can," as Feldman and Conover put it, but their "efficient and reasonable response to ambiguity" leaves them depressingly far from having even roughly accurate perceptions of the political world.

[51] Adding year-specific intercepts to allow for movement in the actual positions of the parties on the liberal-conservative scale does little to alter the picture.

Conclusion

Most of the time, the voters are merely reaffirming their partisan and group identities at the polls. They do not reason very much or very often. What they do is rationalize. Every election, they sound as though they were thinking, and they feel as if they were thinking, as do we all. The unwary scholarly devotee of democratic romanticism is thereby easily misled. But in fact, while the voters may be consistent, and while they may be rational in the thin economic sense of the term, they behave in what Lippmann (1922, 10) referred to as a "pseudo-environment" only loosely connected to

> the real environment where action eventuates. If the behavior is not a practical act, but what we call roughly thought and emotion, it may be a long time before there is any noticeable break in the texture of the fictitious world. But when the stimulus of the pseudo-fact results in action on things or other people, contradiction soon develops. Then comes the sensation of butting one's head against a stone wall, of learning by experience, and witnessing Herbert Spencer's tragedy of the murder of a Beautiful Theory by a Gang of Brutal Facts, the discomfort in short of a maladjustment.

For most people thinking about most political topics, including those of us with doctorates, the discomfort of maladjustment never comes, either because we never emerge from the world of political thought and emotion into the world of practical action, or because the concrete consequences of our misperceptions are too indirect for us to apprehend. Are we to be congratulated for living comfortably and efficiently in our pseudo-environments?

It is an unfortunate error, in our view, to confuse "rationality" in the thin mathematical sense of logical consistency with "the notion of a competent citizenry" held up for examination by Kuklinski and Quirk. Competence requires not only logical consistency and cognitive efficiency, but also some modicum of accuracy in perception and receptiveness to new and, perhaps, disconfirming evidence.

Consider, once again, the example of the federal budget deficit. Hansen (1998) provided a detailed analysis of data from a 1995 NES pilot survey in which respondents were invited to favor or oppose a variety of possible departures from current fiscal policy – raising taxes to reduce the budget deficit, increasing the budget deficit to fund increases in spending on domestic programs, and so forth. He found very few logical inconsistencies in responses to these questions (for example, people who wanted to increase the budget deficit in order to increase domestic spending, but also wanted to cut domestic spending in order to reduce the budget deficit). On the basis of his analysis Hansen concluded (1998, 519) that "The public has the ability to make budget policy choices with reasonable discernment. . . . They have well-formed and well-behaved preferences."

One would hardly guess that these are the same people who, one year later, were largely oblivious to the fact that the federal budget deficit had declined by more than 90% over the preceding four years. Could people unaware of such a salient and politically consequential fact "make budget policy choices with reasonable discernment"? Hansen's (1998, 526) assertion that "American democracy does not want for the competence of its citizens" strikes us as subject to doubt.

We have suggested here that the average citizen's perception of the federal budget deficit is constructed of four parts folk wisdom, one part partisan inference, and a trace element of reality. For perceptions of national economic conditions the mix is somewhat more edifying: say, three parts folk wisdom, one part inference, and three parts reality. Nor would it be difficult to guess at similar fractions from the intellectual history of the American professoriate.

We do not contest the notion that ordinary citizens are doing their best to construct consistent, subjectively plausible perceptions of a complex political world. We merely wish to note that their best should be troubling to enthusiasts of democracy – especially when, as Lippmann (1922, 14) put it more than 80 years ago, "these fictions determine a very great part of men's political behavior."

Appendix 1

At time n, the citizen begins with a PID $\hat{u}_n$, an estimate of his true utility difference between the parties on the k issues he has encountered thus far. He is now confronted with a new issue. He wants to estimate the appropriate opinion δ for him on the new issue, and he wants to update his PID by including in it his best estimate of his position on the new issue. He knows that the true relationship is $u_{n+1} = u_n + \lambda_{k+1}\delta_{k+1}$. It will be helpful to define the partisan deviance $\gamma = \delta - u_n$, the degree to which the citizen's appropriate opinion on the new issue differs from his true PID.

At time n, the priors are:

$$\delta \backsim N(\delta_0, \sigma_0^2) \tag{5}$$

$$\gamma \backsim N(0, \hat{\tau}_k^2) \tag{6}$$

$$u_n \backsim N(\hat{u}_n, \omega_n^2) \tag{7}$$

The latter two priors are independent.

The one piece of new data at time n is $\bar{y}$:

$$\bar{y} \backsim N(\delta, \sigma^2/n) \tag{8}$$

Formally, we must integrate out γ, then apply Bayes Theorem to the new information $\bar{y}$, and finally construct the marginal distributions of u_n and δ. However, since all distributions are normal, a simpler approach gives the same answer.

Since the priors in Equations (6) and (7) are independent, they jointly imply that since $\delta = u_n + \gamma$, then $\delta|\hat{u}_n \backsim N(\hat{u}_n,\, \omega_n^2 + \hat{\tau}_n^2)$, where the variances are treated as known. Combing this with Equations (5) and (8) then immediately gives, by the customary Bayesian logic, the posterior distribution for $\delta|\bar{y}$ given in Equation (1) above.

We next wish to estimate $u_n|\bar{y}$. That is, we update the estimate of the utility of the old issues based on the new information $\bar{y}$. We ignore the prior for δ in Equation (3) since the new issue δ does not enter the value of

u_n directly, and since by assumption, σ_0^2 is very large and thus δ_0 contains almost no indirect information about u_n via the relationship $\delta = u_n + \gamma$.

It follows that the relevant information for estimating $u_n|\bar{y}$ is given by Equations (6), (7), and (8), where $\bar{y}$ is informative about u with variance given by the sum of its sampling variance σ^2/n plus the variance of δ around u, namely $\hat{\tau}_k^2$. Then by the usual logic, the posterior is normally distributed with mean:

$$E(\hat{u}_n|\bar{y}) = \frac{(\hat{\tau}_k^2 + \sigma^2/n)\hat{u}_n + \omega_n^2\bar{y}}{\hat{\tau}_k^2 + (\sigma^2/n) + \omega_n^2} \tag{9}$$

Finally, we need an updated estimate of $u_{n+1} = u_n + \lambda\delta$, where λ is fixed. But since $\hat{u}_n|\bar{y}$ and $\delta|\bar{y}$ both have normally distributed posteriors, it follows from Equations (1) and (9) that (ignoring the small amount of information about u in the prior for δ):

$$\begin{aligned} E(\hat{u}_{n+1}|\bar{y}) &= \frac{(\hat{\tau}_k^2 + \sigma^2/n)\hat{u}_n + \omega_n^2\bar{y}}{\hat{\tau}_k^2 + (\sigma^2/n) + \omega_n^2} + \lambda\frac{\sigma^2\hat{u}_n/n + (\omega_n^2 + \hat{\tau}_k^2)\bar{y}}{(\sigma^2/n) + \omega_n^2 + \hat{\tau}_k^2} && (10) \\ &= \frac{[(\hat{\tau}_k^2 + (1+\lambda)\sigma^2/n)]\hat{u}_n + [\hat{\tau}_k^2 + (1+\lambda)\omega_n^2]\bar{y}}{\omega_n^2 + \hat{\tau}_k^2 + \sigma^2/n} && (11) \end{aligned}$$

which after rearrangement yields Equation (2) above.

Appendix 2: Why Not Cross-Lagged Regressions?

Cross–lagged regressions are commonly used to cope with reciprocal causation. Their first difficulty for the study of abortion, however, is that the relationships are not linear. For example, the usual version of cross–lagging with a dummy variable for gender would assume that the two lines in Figure 9 were each straight and parallel to each other. They are not.

Equally importantly, the model we have developed in this paper demonstrates that the cross–lagged relationship between current PID and abortion attitudes with those at a prior period should fail to be linear and additive in just the way that Figure 9 illustrates. The model implies that updating at time 2 will proceed in the same way as described in Equation (1), where now δ_0 and σ_0^2 are interpreted as the one–period lagged opinion and its posterior variance, respectively, and $\bar{y}$ is the new information *that has arrived since the previous period*. That amount of new information will differ for different groups of people, invalidating the cross-lagged specification.

To see the problem at its most drastic, consider those people thinking about abortion at time 2 for whom no new information has arrived since time 1. (Presumably, many well informed middle–aged people fit this description, having heard no new arguments, pro or con, for a decade or more.) Suppose, too, that the time between periods is sufficiently short that PID will not have strengthened with age. Then Equation (1) will determine their attitude at both time periods by assigning partial influence to all three explanatory variables—their prior when they came of age, what they have learned from then to time 1, and their party ID. Whatever the relative weights of these factors, since nothing has changed between time periods 1 and 2, these citizens will express the same opinion at time 2 as at time 1.[52] Hence time 2 opinion is perfectly predictable from time 1 opinion. The cross–lagged regression will therefore give a coefficient of 1.0 to the lagged opinion and 0.0 to their party ID.

When we tried using the Wiley–Wiley model in the 1992-1994-1996 panel to estimate error variances for PID and abortion attitudes, then substituted these into a standard errors–in–variables model for cross–lagged regressions, the effect of PID was negligible (near zero or negative) under all plausible assumptions about its error variance. The lagged value of abortion attitudes, on the other hand, received coefficients depending on the reliabilities. Since these undoubtedly vary somewhat by information level, and since we lack the number of observations needed to estimate them reliably, we cannot be sure precisely what the true lagged coefficient is. However, the most plausible values for its error variance gave lag coefficients of .8 or higher, often near 1.0. We conclude, then, that PID has relatively little impact on changing the average person's abortion views, while true abortion views

[52] We temporarily set aside the measurement error problem for the purposes of this hypothetical argument. The reader concerned about it may insert "after statistical biases due to measurement error are corrected" in each sentence without altering any of the conclusions.

themselves are highly stable over time.[53] All this is conventional wisdom.

Now in one sense, this is the right answer: From time period 1 to period 2, party ID had no *additional* impact on opinion. However, as we have argued, for some individuals in the right situations (for example, many men during 1982-1997), party ID may be influencing their opinion at both time periods. However, and this is the point, it may be doing so no more at time 2 than at time 1. Then PID adds nothing at time 2, but neither does it lose anything: It matters at both time periods. But it will not receive a coefficient from the cross–lag.

Note that this would not be the usual interpretation of a cross–lagged regressions with a zero coefficient for PID Instead, the vanishing coefficient would conventionally be seen as absence of causation. But as Equation (1) shows, this is just wrong. The problem is that the lagged opinion is endogenous but has no real causal effect. Hence it "over-controls," artificially diminishing the apparent causal impact of PID (Achen 2002). Similar concerns apply to Equation (2) for PID.

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[53] That is, people are somewhat conflicted, but their answers vary in a stable pattern around a fixed mean over time.

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