What Moves Macropartisanship?
A Response to Green, Palmquist, and Schickler

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Contrary to the claim by Green, Palmquist, and Schickler (1998), macropartisanship is largely shaped by presidential approval and consumer sentiment. It is not the case, however, that macropartisanship mirrors the ever-changing levels of current presidential popularity and prosperity. Rather, macropartisanship reflects the cumulation of political and economic news that shapes approval and consumer sentiment. Using ECM technology, we show that, far from being the weak force that Green et al. suggest, the cumulation of innovations in presidential approval and consumer sentiment largely account for the long-term trends in macropartisanship. For forecasting macropartisanship in the near future, it is better to predict from the fundamentals represented by the history of approval and consumer sentiment up to a given moment than from current values of macropartisanship itself.

When "Macropartisanship" was written almost a decade ago (MacKuen, Erikson, and Stimson 1989), we saw our contribution as the following basic points. First, we showed that macrolevel party identification undergoes considerable movement over time. Second, we showed that much of this movement tracks changes in presidential approval and the economy as monitored by the Michigan Index of Consumer Sentiment. In retrospect, "Macropartisanship" is remarkable for what it did *not* say. Apart from chiding the then conventional view that national partisan sentiment seldom moves except when shaken by rare "realignments," the article offered few specifics about the nature of the macropartisanship time series. And it offered no specific viewpoint about microlevel party identification, other than dutifully citing competing "schools" ("psychological attachment" versus "running tally"). The motivating theoretical argument of that article, evidently original for its time, was that the relative stability of microlevel party identification did not logically require that macrolevel partisanship be a constant. These empirical results merely put partisanship back in the realm of politics.

Green, Palmquist, and Schickler (1998) charge that our earlier work seriously overestimated the effects of consumer sentiment and presidential approval on macropartisanship. More may be at stake than the size of a

few coefficients. As they tell it, their revisionary interpretation helps to vindicate both the microlevel theory of party identification as a stable psychological attachment and the macrolevel theory of a partisan division largely immune from major political shocks. In their preferred model, macropartisanship is a stationary series oscillating in long cycles around its long-term mean, propelled largely by shocks of unknown origin. Curiously, they offer no discussion of the possible implications of this model, which suggests that party fortune is little affected by party performance.

Green, Palmquist, and Schickler are able researchers, but in this instance they manage to overlook effects that are hiding in plain sight. We welcome this opportunity to set the record straight. In our response, we emphasize two major points.

First, *macropartisanship is driven largely by the same political and economic forces that drive presidential approval and consumer sentiment.* When political and economic events, cause approval and sentiment to change, macropartisanship also changes. The changes are small but permanent and cumulating, so that macropartisanship at any moment is largely the sum of the preceding economic and political shocks. Changes in macropartisanship from other sources that do not register in the approval and sentiment series (such as election campaigns and their aftermath) are mostly transient and of little long-term consequence.

Second, *the perceived inconsistency between macrolevel movement and microlevel stability is nothing more than a statistical illusion.* Our macrolevel evidence and interpretation are in harmony with the evidence and interpretation (e.g., Campbell et al. 1960; Green and Palmquist 1994) that microlevel party identification behaves as a stable psychological attachment.

Putting the parts together, we offer the following model for reconciliation of macrolevel and microlevel findings. (1) The time series of individual-level party identification represents an initial partisan disposition that is slowly modified by events, including the national political and economic forces that shape party fortune. (2) The macropartisanship time series is dominated by

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FIGURE 1.   Consumer Sentiment, Presidential Approval, and Macropartisanship over Nine Administrations

M = Macropartisanship; A = Approval; C = Consumer Sentiment
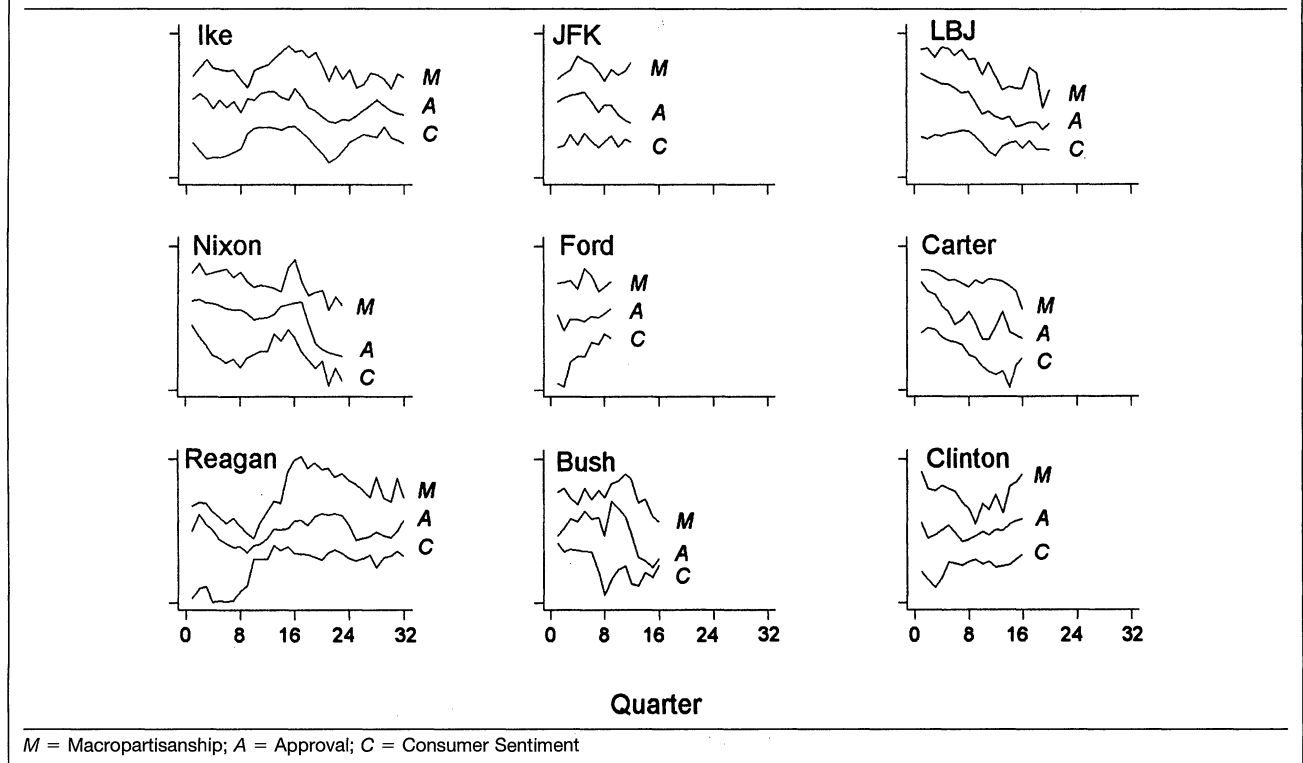
the cumulation of these national political and economic forces.

THE STATISTICAL ARGUMENT OF GREEN AND COLLEAGUES

Figure 1 presents the time series of *Macropartisanship* (here, percentage of partisans for the presidential party), Gallup's *Presidential Approval*, and the *Michigan Index of Consumer Sentiment* for nine administrations. The graphed lines represent the raw data, unadulterated except for the standardization of scales to enhance the visual display.[1] As we view these data, party identification tracks approval, which tracks consumer sentiment. Not all approval trends stem from consumer sentiment. But virtually all detectable shifts in macropartisanship reflect the movement of approval and (perhaps indirectly) consumer sentiment.

Does this informal assessment stand up to statistical scrutiny? Green et al. say "no." While accepting our assessment of the effect of consumer sentiment on approval (MacKuen et al. 1989), they claim that we exaggerate these variables' influence on macropartisanship. They believe we "overfit" the data, particularly

by our use of dummy variables for presidential administrations as controls. They prefer to report results without these controls, due to the presidential dummies' alleged lack of theoretical relevance. Their imposition of this statistical rule appreciably diminishes the coefficients for consumer sentiment and presidential approval and by itself accounts for much of the present disagreement.[2] Virtually all the ambiguity that they generate regarding our earlier claims is due to their omission of controls for presidential administration.

Table 1 presents two versions of a partial adjustment model predicting macropartisanship (percentage of Democrats among Democratic and Republican partisans) from current (Gallup) approval, one with and one without the presidential dummies.[3] (Consumer

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[1] In Figure 1, the Michigan Index of Consumer Sentiment (MICS) and presidential approval are scaled in standard deviation units, while macropartisanship is scaled in 1.5 standard deviation units. This calibration allows approximately identical variances to the change scores in the three variables. Each series is depicted as centered around the within-administration mean, with the approval series centered two units higher than MICS and macropartisanship centered two units higher than approval.

[2] Not a matter of contention is use by Green et al. of *all* available Gallup surveys rather than the sample we used in our 1989 work. Their better measure of macropartisanship allows them to revise downward the coefficients for consumer sentiment and presidential (or "political") approval, while revising upward the coefficients for lagged macropartisanship. As they explain, the revised coefficients follow from the improved measurement of the control variable of lagged partisanship. We accept the revised numbers from their replications over our original coefficients when applied to our original models.

[3] Table 1 includes an MA(1) term to replicate the statistical practice of our replicators. The statistical analysis is virtually unchanged when the MA(1) term is omitted. The justification for the MA(1) by Green et al. is that it serves as an adjustment for sampling error, following Beck (1989). Yet, this adjustment is valid only when the underlying model is an AR(1) process, which we see as invalid for the macropartisanship data (discussed later). If the model *were* AR(1), then the

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sentiment is omitted here because its effect is largely indirect.) As the table shows, the approval coefficient is virtually cut in half when the presidential dummies are eliminated.[4] How should one interpret this result?

The larger approval coefficient with the dummies represents the averaging over nine separate *within-administration* time series (Eisenhower through Clinton I), controlling away between-administration differences. For each administration, support for the presidential party is highest when approval is highest: about one point of macropartisanship gained for every ten points of approval. Green et al. would have us mingle different presidents of the same party, leaving between-administration differences uncontrolled and thus leaving open the door to unwanted noise and potential biases. Does this matter?

The presidential dummies are collectively significant at the .00006 level, a magnitude that may surprise readers of Green et al. (1998).[5] These dummies represent different equation intercepts for different presidents, reflecting the variation in party history that each president inherits. Statistically, these controls can only reduce bias rather than add bias, and at the trivial cost of a few degrees of freedom. The claim of theoretical irrelevance by Green et al. not withstanding, on strictly statistical grounds the higher estimate of an approval effect with the administration controls is decisively better.

Still, Green et al. have a point. Why should the presidential administration matter? Rather than deny the existence of this effect, as Green et al. prefer, let us try to understand it. The dummy coefficients in Table 1 do not represent presidential characteristics but the partisan nature of the times. The base categories are

TABLE 1. Partial Adjustment Models Predicting Macropartisanship with and without Presidential Dummies

Variables	With Dummies	Without Dummies
Approval$_t$ × Party	0.092	0.044
	(0.013)	(0.010)
Party (D = 1, R = −1)	−5.156	−2.347
	(0.859)	(0.528)
Macropartisanship$_{t-1}$	0.637	0.904
	(0.061)	(0.028)
Intercept	22.358	5.844
	(3.770)	(1.720)
Presidential Dummies (Ike, JFK = 0)		
LBJ	0.734	
	(0.521)	
Nixon	−0.132	
	(0.394)	
Ford	1.119	
	(0.632)	
Carter	2.863	
	(0.692)	
Reagan	−1.531	
	(0.375)	
Bush	−2.150	
	(0.525)	
Clinton	−1.451	
	(0.652)	
MA(1)	−0.179	−0.225
	(0.099)	(0.081)
Significance of Presidential Dummies	0.000063	
Adjusted R^2	.889	.865
SEE	1.569	1.729
(N)	(175)	(175)

Note: Standard errors are in parentheses.

implied proportion of the variance in the dependent variable due to measurement error would be (using Green et al. symbols) $-(1 - \phi L)/\gamma_0$, or minus the ratio of the MA(1) coefficient to the autoregressive coefficient. This ratio is persistently .2 or higher in the Green et al. analysis, suggesting an implausible degree of sampling error in political surveys. Based on average quarterly Ns exceeding 4,000 (Gallup) and 2,500 (CBS/NYT) and sampling theory, we estimate the reliability of macropartisanship to be .98 in the Gallup series and .96 in the CBS/NYT series. This comfortable level of accuracy does not call for debatable attempts at correction. To avoid making the analysis any more complicated than necessary, we do not include an MA(1) term in our original analysis which follows. Fortunately, all substantive findings are virtually unaffected by the presence or absence of an MA(1) term in the equation.

[4] The Table 1 equation without dummy variables is virtually identical to the first equation in Table A-1 of Green et al. In this and following data analyses, we use the Green et al. data set, except that we do not mingle approval ratings for different presidents for the same quarter. We use Kennedy's approval numbers for 1963:4; and Nixon's approval numbers for 1974:3.

[5] Green et al. appear to argue that administration dummies are not statistically significant. They report that our "slew" of 24 control variables, including the administration dummies, is not collectively significant for the period of our 1989 study. But that is without MICS or approval in the equation! Indeed, administration dummies achieve no more than borderline significance when the dummies and lagged macropartisanship are alone on the right-hand side of the equation. The reason the dummies gain dramatically in significance when when approval is entered into the equation will become clear as the discussion progresses. Presidential dummies are proxies for the equilibrium level of macropartisanship derived from cumulative political and economic approval over time.

zero for Eisenhower (for Republican administrations) and Kennedy (for Democratic administrations). The coefficients are near zero for Johnson and Nixon but then turn sharply Democratic for Ford and Carter, as a post-Watergate legacy, and sharply Republican for Reagan, Bush, and Clinton, after partisan fortune turned again in the 1980s and 1990s. The dummy variable coefficients remind us that partisanship is a function not only of political and economic conditions of the moment but also of the political and economic history leading up to that moment.

The contrast between the Carter and Clinton I administrations provides a useful illustration. Within each administration, higher approval (along with consumer sentiment) is associated with higher Democratic partisanship, just as one might expect. The story changes when we look *between* administrations. On average, Clinton enjoyed better approval numbers (and

better economic numbers) than Carter, yet Americans were more Democratic in the late 1970s than in the mid-1990s. This should surprise no one who knows the political history preceding Carter (the Watergate debacle) and Clinton (the Reagan-Bush successes). Any analysis that fails to incorporate that recent history would mistakenly conclude that good economic management and political approval produce partisan failure.

It is often said that categorical dummy variables of the sort represented by presidential administrations represent theoretical ignorance. If they matter statistically, then they control for something important, even if we do not know precisely what it is. Our next task is to identify that something. Calling it history is only the beginning. We refine our model to demonstrate the economic and political forces affecting macropartisanship, while rendering unnecessary a resort to presidential dummy variables, and we learn something important about the nature of partisan change.

A REFORMULATION

Consider again the three time series depicted in Figure 1. Statistically, consumer sentiment and presidential approval can be modeled as AR(1) *stationary* series, where current values are a function of values lagged one quarter:[6]

$$Approval_t = 3.792 + 0.917\ Approval_{t-1} + a_t, \quad (1)$$

and

$$MICS_t = 8.998 + 0.897\ MICS_{t-1} + e_t. \quad (2)$$

The terms e_t and a_t represent the innovations in the two series. Given the autoregressive coefficients of about .90, the innovations dissipate at a fairly rapid rate, about 10% per quarter, which limits the ability to forecast several quarters ahead. In effect, the electorate "forgets." It weighs recent inputs most heavily, discounting the past about .10 more each passing quarter. All this is reasonable. It makes sense for consumer sentiment to mirror the current economic circumstances more than those of the past. It makes sense for the president's approval rating to mirror current successes and failures more than those of the past.[7]

Macropartisanship, we argue, is different. Let us consider the implications of a partisanship that is not

only the enduring trait emphasized by the traditionalists (e.g., Campbell et al. 1960; Miller and Shanks, 1996) but also changeable in response to events as emphasized by revisionists (e.g., Fiorina 1981; Franklin and Jackson 1983; Weisberg and Smith 1991). According to the original portrayal of partisanship as a stable attachment, people's fundamental feelings about the parties rarely change. If, in the extreme, virtually all partisanship is learned by early adulthood, then aggregate partisanship could change only glacially, as a function of the shifting partisanship of successive birth cohorts that enter and exit the system. Call this model I. Macropartisanship would be essentially a constant. Consider the other extreme: routine shifts in partisanship according to current evaluations of the president and the economy, along with perhaps other variables. Aggregate partisanship would reflect aggregate values of these predictor variables, with a memory no longer than the predictor variables themselves. Call this model II. Macropartisanship would move as a stationary series, reflecting the recent political and economic environment. We could aggregate the two models, making the near-constant birth-cohort component of model I the long-term equilibrium, while political and economic events generate a stationary series around it. Call this hybrid model III. While it anchors the equilibrium component of the stationary series by attributing it to aggregated early partisan socialization, its empirical prediction is indistinguishable from that of model II.

Our challenge is that none of these models fit empirically. Model I cannot be true by the simple test that macropartisanship moves too much. Models II and III do not fit the data well (at least without adding time-dependent dummy variables), which we know from Green et al. (1998). This verdict presents the following puzzle. How may partisanship (1) be largely shaped by early political learning and remain essentially stable during adulthood, (2) show real change in the aggregate, and (3) be effectively modeled as a function of its political and economic environment?[8]

Our solution is a revised blending of the two models of partisanship as permanence and change. To the extent the partisanship of mature adults is constrained by early socialization, in effect there must be no "forgetting" of partisan information learned during early adulthood, adolescence, and before. Of course, we also expect partisan learning to continue during adulthood. We suggest that the retention rate for partisan learning should not depend on the location in the partisan life cycle. *If early socialization leaves a permanent imprint, then any later changes in partisan disposition must also be permanent.* It makes no sense to argue that people discount their recent past but cannot shake off their distant past.[9] Thus, we expect the equilibrium value of partisanship (macro and micro) to

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[6] For the estimated innovations in approval, only the lagged innovations for the current president are used. While this rule allows us to estimate innovations for the first quarter of the Johnson and Ford administrations, it means we must omit each presidential transition in the inaugural quarter following a presidential election.

[7] The approval and MICS series can be validated as stationary instead of unit-root by the conservative Dickey-Fuller test. The hypothesis that MICS is unit-root can be rejected at the .05 level ($|t| = 3.03$). Perhaps surprisingly, the verdict for approval is more ambivalent. The hypothesis that approval is unit-root can be rejected at the .10 level but just misses the .05 level ($|t| = 2.79$) using the quarterly series. When approval is measured monthly, however, the unit-root hypothesis is rejected at the more stringent .05 level. For earlier discussions, see Ostrom and Smith 1992 and Beck 1992. The approval and MICS series each show the autocorrelation pattern consistent with an AR(1) model.

[8] One resolution is that macropartisanship moves a lot, but for reasons that have little to do with the observable political and economic environment.

[9] By the same logic, if recent effects are transient, then those from the distant past are even more so. For most people, the effects of childhood socialization would be long forgotten.

represent the cumulation of its various causes, with no "forgetting."

This reasoning implies that while partisanship changes only slowly, its changes are permanent rather than transitory. When political and economic events affect presidential approval and/or consumer sentiment, the corresponding effect on macropartisanship will be small. But while these influences on approval or consumer sentiment wear off quickly, the small effects on macropartisanship stay, so that macropartisanship represents the accumulation of political and economic news.

Crucial to our understanding is that the stable, integrated series represents partisanship at equilibrium rather than its observed (macro or micro) realization. At the individual level, this formulation is consistent with Green and Palmquist's (1990, 1994) statistical portrayal of individual-level stable party attachments that rarely change. Most adults work out a social and psychological niche in which they encounter information, or interpret information, in ways that rarely threaten their understanding of the party system. While people can and do see reasons for changing their party loyalty, other factors usually "bring them home." These include psychological mechanisms, such as selective perception, internal counterarguing, and a need for psychological order; they also include such social elements as the issue considerations that frame political argument, the weight and direction of the frequently consulted mass media, and, of course, the political views of relevant others. For most people most of the time, these conditions and mechanisms will remain in balance. Thus, partisanship itself is not merely a verbal self-description but is the product of a self-regulating social and psychological system. Crucially, partisanship is a marker for the equilibrium conditions generated by a constellation of social and psychological forces. The equilibrating nature of partisanship (macro and micro) is important for politics because the resulting stability anchors the future to provide predictability.[10]

Equilibrium partisanship implies a distinctive dynamic: When party identification is at equilibrium, it will not change without outside surprises. Simply speaking, a Democrat with Democratic friends and liberal policy views who lives during good times presided over by a Democratic government will remain a Democrat unless something remarkable happens. Of course, unusual events may change the mix—the individual may move into new friendship circles, the national issue debate may induce conservatism, or the performance of the parties may fluctuate. To the extent to which these factors constitute the equilibrating mechanism, equilibrium partisanship will change in response to these shocks. Thus, we have $\Delta PID_{it}^* = u_{it}$, where PID_{it}^* is an individual's equilibrium parti-

sanship, and u_{it} is the shock or innovation at any time. This implies that equilibrium partisanship is a cumulation of all previous innovations in the individual's lifetime (plus an initial partisanship)[11]:

$$PID_t^* = (u_{it} + u_{i,t-1} + u_{i,t-2} + \ldots) + PID_0^*.$$

In addition, we expect some people to stray temporarily from their equilibrium partisanship—perhaps attracted by particular personalities or moved by specific events. For instance, the Democratic partisan of our example may be seduced by a charismatic candidate to support the Republicans but will sooner than later "return home" to equilibrium partisanship. This produces:

$$\Delta PID_{it} = -\gamma(PID_{i,t-1} - PID_{i,t-1}^*) + v_{it},$$

where $(PID_{i,t} - PID_{i,t}^*)$ is the temporary deviation from equilibrium partisanship, and v_{it} represents short-term forces that deflect voters from their partisan equilibrium. These temporary deviations are based entirely in the present and are eventually "forgotten." Thus, equilibrium partisanship integrates history, whereas the temporary deviations disappear. These two parts imply that, to the extent to which equilibrium partisanship itself is unaffected by the personal environment and national forces, all movement will represent temporary deviation, and individual partisanship itself should be a stationary series.

At the macrolevel, the individual-level u_{it} and v_{it} shocks cancel out, except for their mean values, which we label α_t and ε_t, respectively. The macrolevel extension is straightforward. We have an observed macropartisanship, M_t, and its equilibrium value, M_t^*, where:

$$\Delta M_t^* = \alpha_t,$$

$$M_t^* = (\alpha_t + \alpha_{t-1} + \alpha_{t-2} + \ldots),$$

and

$$\Delta M_t = -\gamma(M_{t-1} - M_{t-1}^*) + \varepsilon_t.$$

The α_t inputs for macropartisanship represent the cumulated political and economic shocks to macropartisanship.[12] But we need not represent these shocks as mere white noise. We model the cumulated α_t as a function of the cumulated shocks to presidential approval and consumer sentiment. Our argument is that to know the cumulated political and economic shocks to approval and consumer sentiment is largely to know the equilibrium value of partisanship. Next, we put this argument to the test.

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[10] This equilibrium partisanship is consistent with partisanship as self-identification as well as partisanship as a running tally, where the citizen accumulates evidence regarding the parties' relative advantages. These two equilibrating mechanisms may be less distinct in reality than in abstract theory.

[11] Note that $\Delta PID_{it}^* = u_{it}$ can be written $PID_{it}^* = PID_{i,t-1}^* + u_{tt}$. We may take this last expression, substitute $PID_{i,t-1}^* = PID_{i,t-2}^* + u_{i,t-2}$ for $PID_{i,t-1}^*$, and repeat substituting recursively to generate a model of PID_{it}^* as an accumulation of innovations.

[12] A potential complication is that the γ term should be subscripted for individuals, as if different people have different autoregressive terms for the short-term component (see Box-Steffensmeier and Smith 1996, 1998). The possibility that the short-term component of partisanship may be fractionally integrated is beyond the scope of our attention here.

FROM THEORY TO ANALYSIS

In practice, macropartisanship decidedly does not behave as a stationary series. As Box-Steffensmeier and Smith (1996, 1998) observe, macropartisanship is long-memoried.[13] Box-Steffensmeier and Smith diagnose the series as "fractionally integrated," meaning as a composite representing a density of different AR(1) series. We find it satisfactory to model it more simply, as a cumulative series (integrated of order 1) plus a short-term stationary component. The cumulative series represents the aggregated partisan dispositions of citizens—in effect the collective equilibrium that itself changes in response to economic and political shocks. The short-term series represents the short-term variation around equilibrium.

We are now ready to reexamine the responsiveness of macropartisanship to economic and political forces. We want to model M_t^*, the equilibrium condition of macropartisanship (M_t). For M_t^* to be statistically integrated (no "forgetting"), current innovations must be uncorrelated with lagged values. This condition is readily accomplished by modeling macropartisanship not as a function of consumer sentiment and approval per se, but as a function of the *innovations* in these variables, e_t and a_t from equations 1 and 2. In this formulation, changes in consumer sentiment and presidential approval do not directly cause macropartisanship; rather, the economic and political events responsible for these changes also cause macropartisanship to change. The innovations in consumer sentiment and approval represent surprises, that is, changes in the economic or political climate unanticipated by the previous levels of consumer sentiment and approval.

Before we proceed, there are a few additional preliminaries. To gauge how the dynamics evolve, it is helpful to separate the flows of political and economic influences on M_t^*. Consumer sentiment reflects the workings of economic forces that operate largely but not entirely via approval. Complicating matters is the greater delay in the effects of economic news compared to political news. Consequently, we purge the innovations in approval (a_t) of the portion due to economic news (e_t). We do this by regressing a_t on e_t and e_{t-1}, taking residuals. These new residuals represent the political innovations (p_t) in approval, purged of the economic innovations. We also must adjust both sets of innovations to take into account the party controlling the presidency. We do so by multiplying each by a party variable scored $+1$ when the president is a Democrat and -1 when the president is a Republican, while scaling macropartisanship as the percentage of Democrats. In this way, good news adds to Democratic identification, given a Democratic president, but sub-

tracts from it in the case of a Republican president.[14] Finally, we measure the cumulated versions of the p_t and e_t series. We start each cumulative series at 0 in 1953:1 and add scores.[15]

MACROPARTISANSHIP, POLITICS, AND THE ECONOMY

Armed with theory and data, we test the plausibility of our dynamic equilibrium formulation. In the Green et al. model, macropartisanship responds minimally and gradually to political and economic innovations, eventually trending toward a constant long-term equilibrium. In our proposed model, macropartisanship cumulates the effects of political and economic forces, thus undergoing permanent changes over time. We construct an error correction model (ECM) of change in macropartisanship as a function of the innovations in political and economic news plus the reequilibration of previous "errors" toward the equilibrium level of macropartisanship. We explicitly specify the equilibrium level to be a cumulative function of political and economic conditions. In abstract form, the model is:

$$\Delta M_t = b_E E_t + b_P P_t - \gamma(M_{t-1} - M_{t-1}^*) + u_t, \quad (3)$$

and the equilibrium is:

$$M_{t-1}^* = c_0 + c_E CumE_{t-1} + c_P CumP_{t-1}, \quad (3a)$$

where E_t and P_t are current economic and political innovations (capitalized to signify their status as independent variables), and $CumE_{t-1}$ and $CumP_{t-1}$ are the lagged *cumulative* economic and political innovations. (For an expanded discussion of the error correction model and its application to the problem at hand, see the Appendix.)

This model incorporates both of the dynamic formulations we wish to test. If the equilibrium M_t^* were a constant or driven by forces other than political and economic news, then we should estimate zero coefficients for c_E and c_P. Such a condition would reduce equation 3 to a model in which macropartisanship is a function of disturbances in the political economy that dissipate over time. Thus, the key test lies in the coefficients c_E and c_P. If they are zero, then we can adequately model the political part of macropartisanship as a gradual transitory process. If they are not zero, then the effects of the political economy persist in the form of equilibrium macropartisanship. (See also the discussion in the Appendix.) Using the Gallup macropartisanship series, the estimates are:[16]

$$\Delta M_t^G = 0.055E_t + 0.111P_t - 0.198(M_{t-1}^G - M_{t-1}^{G*}),$$
St. Err. = (0.027)   (0.029)   (0.044)   (4)

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[13] One sign is that the over-time correlations are far stronger than they should be, given the autocorrelations at short lags. For instance, considering that the Gallup autocorrelation is .92 over one quarter and assuming an AR(1) model, the autocorrelation should be $.92^{12}$ after 12 quarters, or a mere .37. The observed autocorrelation after 12 quarters is .58, indicating a long-memoried process. By contrast, after 12 quarters the autocorrelations for MICS and approval (same president) drop to .18 and .08.

[14] An alternative to cumulating simple residuals would be to cumulate "recursive" residuals. See the Appendix for a discussion.
[15] For inaugural quarters we have no measure of political innovations because we have no lagged approval scores for the new president. For these quarters we substitute the value of zero for generating cumulative political innovations. Given the lags involved in the creation of the independent variables, the estimation equation is based on a set of observations that do not start until 1953:3.
[16] These estimates are from the "one-step" procedure.

FIGURE 2. Gallup Macropartisanship, 1952–96

Dots = M; Line = M^*. Solid Dots = Presidential Elections

and
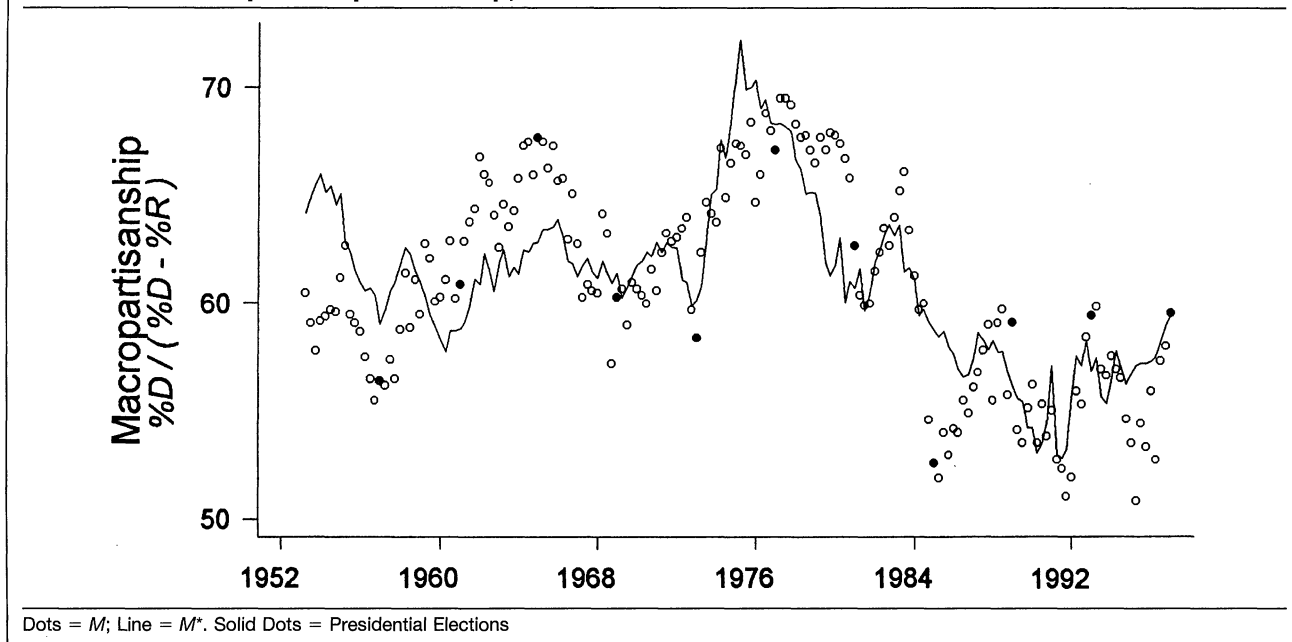
$$M^{G}_t{}^* = 64.17 + 0.169 CumE_t + 0.126 CumP_t,$$

St. Err. = (1.41) (0.036)   (0.029)

Adj. $R^2 = 0.168$; $SEE = 1.712$; $N = 175$.

We see that c_E and c_P are decisively nonzero (a joint test produces $p = 0.00000$). The equilibrium M_t^* is a dynamic function of the political and economic environment. Furthermore, if we add administration dummy variables, they neither are statistically significant nor affect the other coefficients in an appreciable way. In short, by explicitly modeling the enduring component of macropartisanship in terms of its underlying cumulative economic and political causes, we account for the historical "effects" otherwise represented by administration dummies and resolve the empirical dispute raised by Green et al. More important, we affirm that the crucial part of macropartisanship—its equilibrium condition—is driven by political and economic information.

More is revealed in equation 4 than the likely existence of a moving equilibrium. We observe from the $(M_{t-1} - M_{t-1}^*)$ coefficient that in the absence of M_t^* change, macropartisanship moves toward its equilibrium value to fill about 20% of the gap per quarter. We also observe a distinctive dynamic response of macropartisanship to politics compared to economics. Political shocks (P_t) produce an immediate reaction in macropartisanship of 0.111 points per unit shift. Thus, when the equilibrium shift is in the form of a changing political input, we obtain nearly the full movement toward the new equilibrium (0.111 of 0.126 points). For economics, in contrast, a one-point innovation in E_t yields only an 0.055 immediate response, which is only one-third of the

way toward the new equilibrium (0.169 points higher than before). In subsequent periods, both deviations will reequilibrate at a rate of 0.198, but the major movement in the political part will have already occurred. Thus, we see that the political part of equilibrium partisanship reacts almost immediately to exogenous shocks, while the economic part reacts only partially, that is, gradually. This makes sense in that people understand quickly the partisan meaning of political news, but they take longer to appreciate the partisan implications of economic change.[17]

Figure 2 graphs observed (Gallup) M_t along with M_t^* equilibrium values. With M_t and M_t^* correlated at 0.77, more than half the variance in M_t is accounted for by M_t^*, the moving long-term equilibrium. It is important to note that the largest part of macropartisanship's range, from its post-Watergate peak to its Gulf War trough, is captured by the fundamentals.[18] Meanwhile, the $(M_t - M_t^*)$ "residuals" reveal a pattern of systematic short-term variation. Most of these short-term surges center around landslide national elections, such as the short-term Democratic

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[17] Consistent with macropartisanship's gradual response to economic change, macropartisanship moves only when economic realizations are felt in the pocketbook. This is unlike presidential approval, which responds as economic change is anticipated in the form of expectations (MacKuen et al. 1992). For details, see the Appendix.

[18] We also can observe the M_t^* economic and political components separately over time. The economic component (with a slightly larger variance) has generally favored the Democrats. The political component has more consistently favored the Republicans. The two largest reversals of this Republican trend occurred during Watergate and late in the Bush administration.
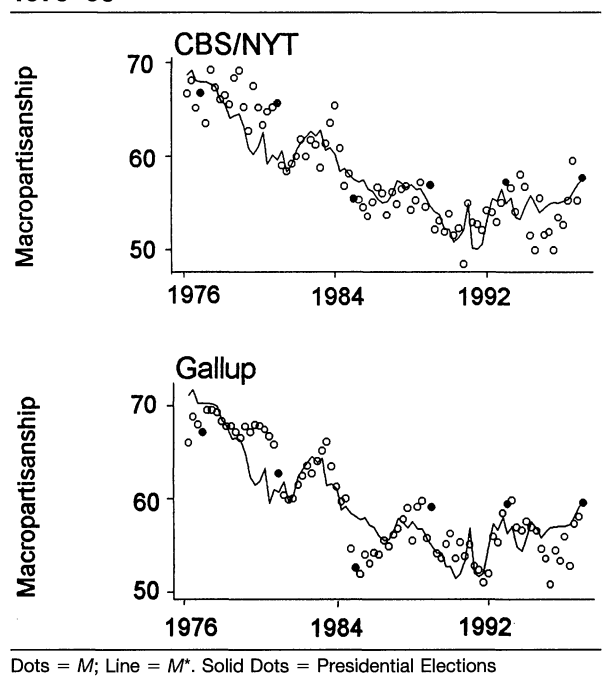
FIGURE 3. Two Macropartisanship Series, 1976–96

Dots = M; Line = M*. Solid Dots = Presidential Elections

TABLE 2. Predicting Future Macropartisanship from Current *Observed* Values (M_t) plus Current *Equilibrium* Values (M^*_t), 1976–96

Dependent Variable	Gallup		CBS/NYT	
	M_t	M^*_t	M_t	M^*_t
M_{t+1} (N = 83)	0.70 (9.43)	0.35 (3.57)	0.49 (5.08)	0.50 (4.49)
M_{t+2} (N = 82)	0.55 (6.20)	0.50 (4.22)	0.31 (3.05)	0.67 (5.75)
M_{t+3} (N = 81)	0.38 (3.49)	0.65 (4.47)	0.23 (2.15)	0.74 (6.12)
M_{t+4} (N = 80)	0.24 (1.95)	0.78 (4.84)	0.26 (2.41)	0.67 (5.41)
M_{t+5} (N = 79)	0.04 (0.30)	0.96 (5.62)	0.19 (1.61)	0.71 (5.12)
M_{t+6} (N = 78)	−0.12 (0.60)	1.01 (5.34)	0.09 (0.69)	0.77 (5.45)

Note: Entries are regression coefficients, with *t*-values in parentheses.

surge in 1964 and a short-term Republican surge in 1984.[19]

We can repeat the ECM exercise using the CBS/*NYT* measure available for 1976–96:

$$\Delta M^C_t = 0.019E_t + 0.096P_t - 0.536(M^C_{t-1} - M^C_{t-1*}),$$

St. Err. = (0.053)   (0.048)   (0.098)

                                        (5)

and

$$M^{C*}_t = 64.52 + 0.174CumE_t + 0.156CumP_t,$$

St. Err. = (1.83) (0.031)   (0.026)

     Adj. $R^2 = 0.254$; $SEE = 2.281$; $N = 83$.

For comparison, the Gallup version of the equation for the same time period is:

$$\Delta M^G_t = 0.029E_t + 0.127P_t - 0.294(M^G_{t-1} - M^G_{t-1}*),$$

St. Err. = (0.039)   (0.035)   (0.070)

                                        (6)

and

$$M^{G}_t* = 64.29 + 0.221CumE_t + 0.143CumP_t,$$

St. Err. = (2.50) (0.044)   (0.035)

     Adj. $R^2 = 0.265$; $SEE = 1.168$; $N = 83$.

Figure 3 compares the tracking of the Gallup and CBS/*NYT* series as a function of their M^*_t predictions. Even though the latter series looks noisier (due to its inherently greater measurement error), on average it is just as adequately predicted by M_t^*. The apparent differences in the two series are largely a matter of statistical illusion.[20] The two measures of M_t^* correlate at .99 with each other and at .87 with the two observed M_t series, while the two M_t measures correlate at .91 with each other.

There can be little doubt that the forces driving long-term change in macropartisanship are largely the same forces driving consumer sentiment and approval. The proof is in the predictive accuracy. Table 2 compares the ability to predict future values of macropartisanship from current realizations (M_t) and the current equilibrium (M_t^*) of macropartisanship, for 1976–96. The coefficients represent the OLS coefficients for M_t and M_t^* predicting future macropartisanship at various leads. To predict future Gallup realizations as immediate as three quarters ahead, the current equilibrium value provides more information than current realizations. For predicting five or more quarters ahead, current realizations are not even statistically significant when the current equilibrium value is controlled. For predicting six quarters ahead, current macropartisanship has the wrong sign. For the CBS/*NYT* series, which Green et al. claim is so unresponsive to economic and political inputs, the predictive dominance of M_t^* is even stronger. For predicting one quarter ahead, M^C_t and M^C_t* perform about equally. For longer leads (even a mere two quarters), M^C_t* predicts better than M^C_t. *It is more important to know the fundamentals represented by cumulative innovations in consumer sentiment and approval than to know the current value of macropartisanship.* The equilibrium value represents the fundamentals; any current deviation from this value will be temporary. Moreover, the

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[19] This transient component is exactly that envisioned by Converse (1976). Much (but not all) of the evident partisan fluctuation before elections that seems to presage the election outcome is merely temporary.

[20] Since the CBS/*NYT* series reequilibrates faster than the Gallup series, a case can be made that the Gallup series (with its "in politics today" wording) is more affected by short-term forces unrelated to M_t^*. The crucial point is that the two series respond identically to the long-term component of macropartisanship represented by M_t^*.

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fundamentals provide stability, correlating at .97 with its lagged values.[21]

The key to understanding the macropartisanship series is its long memory. The response of macropartisanship to new economic and political inputs may be imperceptibly small at the time of occurrence, but it will be long-lasting. For instance, the gain in Democratic macropartisanship during Watergate was far less noticeable than the decline in Nixon's popularity. Yet, as Nixon left the scene, the Democratic gain from Watergate persisted to provide an otherwise inexplicable level of Democratic identification during the Ford and Carter presidencies. Similarly, the more visible long-term Republican gain of the 1980s provided a new base of Republican strength that limited the Democratic recovery starting in the late Bush administration. (See the Appendix for further discussion of M^* as an integrated series.)

In time-series terminology, the long-term or fundamental component of macropartisanship is cointegrated with the cumulative economic and political innovations that influence consumer sentiment and presidential approval. At the same time, a short-term component of macropartisanship (most visible in the Gallup version) represents the transitory response to momentary events. This is very different from the Green et al. model of macropartisanship as a stationary series propelled largely by its own lagged values, with only small transitory effects for consumer sentiment and presidential approval.[22]

THE MICRO-MACRO CONNECTION

To say that the fundamental part of macropartisanship derives from the permanent cumulation of economic and political shocks implies that partisan changes for *individuals who comprise the series* are permanent.[23] Here, we address the plausibility of our model, given what we know about individual-level partisanship. Are the equilibrium values of individuals' party identifications integrated? Is there enough microlevel change in party identification to account for the macrolevel movement?

Actually, plentiful side-evidence supports an integrated individual-level partisanship. For instance, it is established lore that individuals become more partisan with time (Converse 1976). The growth of the variance of individual partisanship with age implies that individual partisanship is an integrated rather than stationary series. We can also cite the distinctively Republican partisanship of the pre-Depression generation, famous for its long memory, even as this older generation approaches extinction.[24]

Much of what we know about the stability of party identification at the individual level is due to the sophisticated statistical analysis by Green and Palmquist (1990, 1994) themselves. Using Wiley-Wiley methodology (Wiley and Wiley 1970), they estimate that most of the observed change in the survey responses to the party identification question is actually error, while "true" party identification is remarkably stable. Corrected for "error," Green and Palmquist's microlevel equation is:

$$PID_{it}^* = \beta PID_{i,t-1}^* + u_{it}, \qquad (7)$$

where PID_{it}^* is the "true" value of party identification, and u_t is its normally distributed disturbance. The variables are mean-centered to omit the intercept. Using the two NES four-year panel studies, Green and Palmquist's (1994) mean estimate of β is 0.972 for eight quarters, which prorates to 0.997 for one quarter.[25]

For all practical purposes, Green and Palmquist's (1994) "true" partisanship is unit-root. In our terminology, their "true" partisanship is *equilibrium* partisanship at the microlevel, while their measurement error is a composite of actual survey error plus short-term movement around the respondent's equilibrium partisanship. Just as the long-term component of M_t is integrated, so is the long-term component of individual party identification.

Now we restore the intercept to equation 7:

$$PID^*_{it} = \beta PID_{i,t-1}^* + \alpha_t + u_{it}, \qquad (8)$$

where, as we now see, $\beta = 1$. The intercept α_t represents a uniform shock shared by all citizens at time t due to a common reaction to shared events, while u_{it} is the individual variation around α_t due to variation in individual experience and evaluation. At the macrolevel, the u_{it} terms cancel out, so that only national forces matter. The result is that $\alpha_t = \Delta \overline{PID}^*_t = k\Delta M_t^*$, so that microlevel change and macrolevel

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[21] When the Gallup series is extended to include observations as early as 1957 (the start of Eisenhower's second term), we get nearly as strong a verdict in favor of the forecasting power of $M^G_t{}^*$. This result falters, however, when the first Eisenhower term is included. As Figure 2 shows, the M^G_t versus $M^G_t{}^*$ gap is unusually pronounced for 1953–56. The explanation is uncertain and could lie in weaker data (with more missing values) during this period or perhaps omitted variables.

[22] Green et al. object to the possibility that macropartisanship could be an integrated variable. For our response, see the Appendix.

[23] Given individual mortality, effects are permanent only for the life of the person experiencing them. How can effects be "permanent" if voters are mortal? The time series can be regarded as an amalgam of parallel time series for different birth cohorts that enter and exit. The "slopes" of these series are parallel random walks with cohort-specific intercepts. Our M_t^* captures the slopes but not the mean intercepts that reflect the electorate's shifting cohort composition. Compared to the response of macropartisanship to political and economic events, cohort replacement makes only a small contribution. We estimate that the loss of the pre-Depression generation may have cost the Republicans a few percentage points from 1952 to 1980, with early departures from the post-Depression generation causing a mild reversal of the arrow in more recent years. Including estimates of the contribution of cohort replacement adds to our prediction of M_t^* without altering our ECM model in any appreciable way.

[24] Voters from the pre-Depression generation respond to contemporary stimuli in modern polls, but they remain less Democratic than the post-Depression era generation.

[25] To prorate estimates of β for the quarterly time frame, we take β reported over c quarters to the $(1/c)$ power. In addition to the two NES panels with waves two years apart, Green and Palmquist (1994) also present estimates for four additional adult panels involving shorter time intervals. If we estimate β from pooling six separate estimates from six panels, we obtain a sample mean of 1.00, with a range (from panel to panel) of 0.99 to 1.01.

change are now equivalent, except for a constant k to correct for different scales.

This directs us to the crucial puzzle. How can we reconcile the "large" macrolevel movement in macropartisanship (ΔM_t^*) with the relatively "small" amount of microlevel change represented by the variance in u_{it}? This discrepancy exemplifies the typical illusion of a large change at the macrolevel appearing small at the microlevel. We demonstrate by rescaling ΔM_t^* in party identification units and comparing its variance with the variance in quarterly individual-level movement in equilibrium partisanship, u_{it}. For a plausible result, there should be less variance in the common movement represented by $\alpha_t = k \Delta M_t^*$ than in the individual deviation around this movement represented by u_{it}, since most movement of microlevel partisanship probably depends on idiosyncratic reasons rather than national forces.

Green and Palmquist (1994) provide four estimates of the variance in u_{it} from the four-year NES panels, which we prorate for quarterly time intervals to obtain the estimate of 0.032.[26] To compare this variance with that for ΔM_t^* we need the conversion rate k to transform ΔM_t^* into microlevel seven-point-scale units. We can approximate k using the 23 NES surveys, 1952–96, comparing the standard deviation of aggregate party identification on the 1–7 scale with the standard deviation of macropartisanship $[(\%D/\%(D + R)]$ from the same NES surveys. (The two measures correlate at 0.95, reassuring that they measure the same thing.) The ratio of the standard deviations is 24.1:1.[27] Using this conversion table to translate ΔM_t^* variance into party identification units, we obtain relative variances of u_{it} and α_t as 0.032 to 0.0017, or a ratio of 19:1.[28] Thus, by our estimates, there is far less partisan change due to national conditions than to idiosyncratic factors. The amount of macrolevel movement in partisanship is quite compatible with Green and Palmquist's understanding of microlevel partisanship's stability.

What of Green and Palmquists' (1994) measurement error in observed individual partisanship? Measured at two-year intervals, all short-term variation in individual partisanship around equilibrium values would be statistically indistinguishable from classical measurement error, with its instantaneous decay rate. Aggregated, this short-term change is the observed variation in M_t

$- M_t^*$. We compare the observed variances of short-term macropartisanship ($M_t - M^*_t$) with Green and Palmquist's (1994) error ($PID_{it} - PID^*_{it}$). Using our conversion rule, and comparing the variance of $M_t - M_t^*$ (8.86) with the mean Green-Palmquist (1994) estimate of the error variance (0.599), we obtain a 1:40 ratio of variances. Thus, if only about .025 of Green and Palmquist's "error variance" were induced by national short-term forces, we can readily account for the observed short-term deviations of M_t around M_t^*.

CONCLUSION

Macropartisanship responds to the political and economic forces represented by presidential approval and consumer sentiment. To see this, one must know where to look. An obvious starting point is to investigate the short-term connection between presidential approval and consumer sentiment, on the one hand, and macropartisanship, on the other. But as Green, Palmquist, and Schickler (1998) demonstrate, such a search is only partially fruitful. As we have shown here, macropartisanship incorporates not only the political and economic news of the present but also the cumulation of news from the past. By adopting an "equilibrium" model of *micro*partisanship, we see that the major turns of partisan fortune found in the macropartisanship series can be explained by changes in the economic and political environment. This result, we believe, should be reassuring rather than controversial: Politics does matter. And that confirmation does not require upending current understanding about the stability of microlevel party identification.

APPENDIX

The Error Correction Model and Macropartisanship

The error correction model (ECM) is a standard formulation in which current change has two components: one due to current inputs, and one due to a reequilibration or a return to a sustainable value. In our application, the first part is straightforward: People react to political and economic information. The idea of the second part is that if M_{t-1} were out of equilibrium—for example, if people overreact to good news—then people would move back toward the partisanship that is consistent with their psychological and social environment, hence "error correction."

The classic formulation of this model is by Davidson et al. (1978), and its use is now commonplace (e.g., Bannerjee et al. 1993; Hendry 1995). Beck (1991) provides an instructive discussion of ECMs in political science. The following equations, repeating equations 3 and 3a from the text, illustrate how the dynamics work.

$$\Delta M_t = b_E E_t + b_P P_t - \gamma(M_{t-1} - M_{t-1}^*) + u_t \quad \textbf{(A-1)}$$

and the equilibrium is:

$$M_{t-1}^* = c_0 + c_E \, CumE_{t-1} + c_P \, CumP_{t-1} \quad \textbf{(A-1a)}$$

If $b_P = c_P$, for example, the immediate adjustment in M_t due to P_t would exactly equal the new long-run equilibrium M_t^* due to P_t, so that the path of M_t and M_t^* would correspond.

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[26] Given $\beta = 1$, the reported disturbance variance of u_{it} can be converted to quarterly intervals by dividing the reported variance of u_{it} by the number of quarters between waves.

[27] That is, a change of 24 percentage points in macropartisanship is required for the party identification mean to move one full unit on the seven-point scale.

[28] This calculation assumes the Gallup version, $M^G_t.^*$ Another interpretation is as follows. Suppose we could follow typical citizens over 44 years, measuring their true (equilibrium) partisanship over this span and at the same time represent the long-term component of macropartisanship in 7-point-scale units over the same span. The individual citizen's equilibrium partisanship would display 20 (19 + 1) times the variance of M_t^*. In 1–7 scale units, movement in macropartisanship has moved only in the range between "pure-Independent" and "Independent, leaning Democrat" (or less than one point) since 1953.

If the immediate effects were substantially higher than the long-run implications, $b_P > c_P$, public overreaction to political news, the public would then readjust to the more modest change in equilibrium partisanship. Finally, if the immediate influence is only a fraction of the long-run implication, $b_P < c_P$, then we have evidence for a partial adjustment to the new dynamic equilibrium. The short-run dynamics are a function both of the ratio b_P/c_P (which is the proportion of the equilibrium effect felt immediately) and of the "return to equilibrium" parameter γ, which shows the proportion of the remaining disequilibrium $(M_t - M_t^*)$ is adjusted away each quarter.

If c_E and c_P are zero in equations 3a and A-1a, then the equilibrium M_t^* becomes the constant c_0, and equations 3 and A-1 reduce to $\Delta M_t = b_E E_t + b_P P_t - d(M_{t-1} - c_0) + u_t$. Were this the case, the change in M_t would be a function of a shock in the political economy and a return to a constant equilibrium, making the political or economic component a transient. If c_E and c_P are nonzero, then the effects of political or economic shocks carry forward in the form of an updating M_t^*.

The estimates for equations 3 and 3a (A-1 and A-1a) are obtained by algebraically substituting the latter into the former and using a nonlinear least squares algorithm. Alternatively, one could use a two-step procedure (Engle and Granger 1987): estimate equation 3a and then substitute the estimated M_{t-1}^* into equation 3, all in OLS. Although there is a lively literature on whether the two-step or one-step procedure is preferrable, the matter is moot here. The two-step estimate correlates with our one-step estimate at .998. (We have also estimated the ECMs with MA and AR terms modeling the disturbances; the reequilibration rate estimates diminish slightly, but the rest of the model remains the same.)

Macropartisanship as an Integrated Series

By construction, our cumulative series ($CumE_t = \Sigma E_s$; $CumP_t = \Sigma P_s$; $s = 0, 1 \ldots t$) represent the strict addition of estimated political and economic shocks. Because the shocks are statistically independent over time, the cumulated series are statistically integrated. $CumE_t$ and $CumP_t$ (and therefore M_t^{G*} and M_t^{c*}) all easily pass the statistical challenge of the Dickey-Fuller unit root tests; with nonsignificant $|t|$ values of 1.65, 1.57, 1.80, and 2.31, respectively, the null hypothesis of a random walk cannot be rejected for any of these variables. Dickey-Fuller tests on the residuals ($M_t^G - M_t^{G*}$) and ($M_t^C - M_t^{C*}$) show $|t|$ values of 4.44 and 4.36, respectively, which are significant beyond the .01 level. Thus, standard diagnostic tests reassure that M^* is unit-root with stationary residuals.

The Alternative of Stationarity. Of course, to construct integrated series from political and economic shocks is not by itself sufficient evidence of the correct specification for the explanation of macropartisanship. An alternative construction would be a stationary series where effects decay geometrically ($CumE_t = \Sigma E_s \lambda^{t-s+1}$; $CumP_t = \Sigma P_s \lambda^{t-s+1}$) according to a λ parameter between 0 and 1. In effect, we chose $\lambda = 1.00$. Other possibilities, even in the $\lambda = .99$ range, produce lesser correlations between predicted and actual M_t. For this test, we search for the value of λ that maximizes the R^2 predicting M_t from the λ-specific $CumE$ and $CumP$ measures. We start the cumulations in 1953 and use them to predict M_t for 1976–96. Actually, the λ that maximizes the R^2 is 1.002 for Gallup and 1.006 for CBS/*NYT*.

Thus, we model macropartisanship's equilibrium as a

random walk and observe that empirically this model appears to defeat all alternatives. Still, our analysis does not depend on the knife-edge assumption that the autoregressive parameter is precisely 1.0000. If in truth M^* is merely "near integrated" (see Banerjee et al. 1993; DeBoef and Granato 1997) then the implications would be no different for practical politics.

The Plausibility of Integration. Green et al. reject as fundamentally implausible the possibility that a bounded series such as macropartisanship could be integrated or unit-root. Their objection is that a random walk implies an explosive variance and a run to the boundary conditions (0% or 100%). In theory, this objection carries weight; in practice the concern is misplaced.

First, macropartisanship is an unbounded variable that is indexed as a bounded percentage only due to measurement limitations. The concept itself (the electorate's mean relative attraction for the Republican or Democratic Party is unbounded. For comparison, consider the plausible argument that the stock market is unit-root. If we were forced to measure its performance as the percentage of stocks above some arbitrary value-per-unit threshold, then the truth of the statistical properties of the underlying variable would remain unchanged. In any case, we could make measured macropartisanship unbounded by the simple expedient of recalibrating it by the logit transformation. Such a maneuver is unnecessary for the observed data that range from about 50% to 70% Democratic. M_t and its logit transformation [log($M_t/(1 - M_t)$)] correlate at 0.9998 for the actual M_t observations.

Second, M_t^* avoids the calamity of a run to the 0% and 100% boundaries because of its slow speed of movement. In theory, the net change in a unit-root variable will have zero mean with a variance equal to $m\text{Var}(u_t)$, where m is the number of periods and u_t is the change over interval t. The variance in ΔM_t^* is a shade under 1.0, so the expected variance over m quarters is about equal to the number of quarters. For instance, following a century of change, the expected variance is about 400, and the expected standard deviation is about 20. With a standard deviation of 20, half the cases are within 16 points of the mean. Thus, the median absolute value of ΔM_t^* over several runs of 100 years is a mere 16 percentage points.

Because the bounded variable works well enough and because we want to keep the measure easily interpretable ("percentage Democratic" rather than "log of the odds that a partisan is a Democrat"), we chose to follow convention and use percentages, even though we expect an integrated behavioral process.

Recursive Residuals

One interpretation of the innovations in approval and economic sentiment is that these represent the degree of positive/negative surprise conditional upon the lagged levels of the variables. To estimate these innovations, we merely regress the current value on the previous value and take the residuals, with the intent that these represent actual surprises. A plausible alternative is to generate the "recursive" residuals. In that case, the "surprise" is estimated for every time point using only the historical data available up to that time, as if the public could perform these regressions using history when they form their judgments.

The obvious attraction of recursive residuals is that people know history but not the future; for instance, people living in 1956 could not have known the course of events for the next 40 years and thus could not have used those data to estimate persistence functions. It is awkward to assume, however, that everyone's historical monitoring began precisely at the time

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when our statistical monitoring began, in 1953. In defense of our use of regular residuals we assume that citizens early in our time span used pre-1953 information to estimate the persistence functions and that these estimates are similar to those using later information as well.

If the persistence function changes over time, then there may be an advantage to using the recursive residuals. If the persistence function is relatively stable, however, then this advantage disappears. Furthermore, the recursive residuals approach produces poor estimates at the beginning of the series because the data are too slim for the task.

We compared estimates of our model using both methods. For the recursive residuals model we follow the exact same procedures as we did for the "simple residuals" estimates of the innovations in economic sentiment and political approval, this time substituting the recursive residuals for the simple residuals. In the end, the two procedures produce statistically indistinguishable estimates.

Decomposing Economic Shocks

Our main analysis shows equilibrium macropartisanship as a function of cumulated economic and political shocks, where the economic shocks are drawn from the Michigan Index of Consumer Sentiment (MICS). Although the central question about economics has been framed as whether consumer sentiment matters, we can probe further to ask which ingredients of the index actually stir mass partisanship. If the question is which ingredient affects presidential approval, the answer is business expectations (MacKuen et al. 1992). The effect on macropartisanship is different. Consistent with its long-term character, macropartisanship's economic response is to collective pocketbook realizations.

To see this, from MICS we extract the separate components of pocketbook realizations (whether family fortunes improved over the past year) and long-term business expectations (the five-year forecast of business conditions). Then, we repeat our ECM exercise with *two* economic components plus the political component (approval unexplained by economics). With two economic components (R_t = shocks to personal retrospections; F_t = shocks to business expectations; and P_t redefined as presidential approval not accounted for by the pair of economic series), we obtain:

$$\Delta M_t^G = 0.031R_t + 0.018F_t + 0.115P_t - 0.271(M_{t-1} - M_{t-1}^*)$$

St. Err. = (0.024)   (0.019)   (0.028)   (0.053)

(A-2)

and

$$M_t^{G*} = 61.89 + 0.119CumR_t - 0.017CumF_t + 0.113CumP_t$$

St. Err. = (1.50) (0.024)   (0.025)   (0.020)

(A-2a)

Adj. $R^2 = 0.222$; $SEE = 1.693$; $N = 174$.

The result is decisive. M^* is driven by aggregate pocketbook considerations and the political component of approval but not by business expectations. Partisan self-identification, quite unlike approval of a president, is a matter of personal experience rather than informed speculation. Thus, while economic expectations can affect a president's popularity, any long-term effect on macropartisanship depends on whether the expectations are followed by their predicted realizations.

For the presidential party's economic policies to alter the electorate's collective partisan memory, anticipation is not enough. The effects must be felt by the voters.

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