The Dynamics of Aggregate Partisanship

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Despite extensive research on party identification, links between partisanship at the individual and aggregate level have largely been ignored. This leaves a gap in our understanding of the dynamics of aggregate partisanship. To remedy this, we identify a set of ideal types that capture the essential arguments made about individual-level party identification. We then combine the behavioral assumptions for each type with existing results on statistical aggregation to deduce the specific temporal pattern that each type implies for aggregate levels of partisanship. Using new diagnostic tests and a highly general time series model, we find that aggregate measures of partisanship from 1953 through 1992 are fractionally integrated. Our evidence that the effects of a shock to aggregate partisanship last for years—not months or decades—challenges previous work by party systems theorists and students of "macropartisanship." Our arguments and empirical evidence provide a conceptually richer and more precise basis for theories of issue evolution or endogenous preferences—in which partisanship plays a central role.

An enduring issue within the literature on party identification focuses on the extent to which individual- and aggregate-level measures of party identification exhibit stability, or more accurately persistence or permanence, over time.[1] The recurring nature of this question is due to the centrality of the concept of party identification within the study of politics. Debates about persistence are quite pronounced among those who study micro-level partisanship (e.g., Allsop and Weisberg 1988, Campbell et al. 1960, Converse 1962, Fiorina 1981, Franklin 1984, Franklin and Jackson 1983, Green and Palmquist 1990, Jackson 1975), but questions about its magnitude also arise in analyses of aggregate-level partisanship (e.g., Abramson and Ostrom 1991, 1992; MacKuen, Erikson, and Stimson 1989, 1992; Miller 1991; Weisberg and Smith 1991).

At the individual level, controversies over the degree to which party identification is permanent are stirred by theoretical debates about the formation and causes of change in partisanship. According to traditional theories of individual party identification (e.g., Abramson 1975, 60; Beck 1974, 398; Campbell et al. 1960, 120–5), an individual's party affiliation represents a "lasting attachment" that forms early in life through a process of political socialization. These psychological attachments are expected to persist over time and to work as exogenous influences on voting behavior. Revisionist research (e.g., Achen 1992, 203; Fiorina 1981, 80–90; Franklin 1984, 461; Franklin and Jackson 1983, 958; Jackson 1975, 162; Markus and Converse 1979, 1077–8; Page and Jones 1979, 1059), in contrast, suggests that an individual's party identification is not necessarily permanent but is both a cause of and can be caused by political factors, such as the distance of individual preferences from party policies (Franklin and Jackson 1983; Jackson 1975), candidate evaluations (Page and Jones 1979), prior presidential voting behavior (Markus and Converse 1979), and evaluations of incumbent performance (Fiorina 1981). Yet, Green and Palmquist (1990) call these revisionist claims into question with their empirical evidence that, after accounting for measurement error, political factors have no effect on individual party identification. Green and Palmquist (1990, 872) conclude that "the outstanding characteristic [of party identification] is its persistence over time."

In analyses of aggregate data, the magnitude of persistence in partisanship is also an important issue. For instance, realignment and other party systems theories (e.g., Burnham 1970; Cavanagh and Sundquist 1985; Clubb, Flanagan, and Zingale 1980; Sundquist 1983) are often predicated on an assumption of the durability of individual party identification over long periods. Persistent individual party affiliations are hypothesized to produce a fixed aggregate partisan distribution. Yet, MacKuen, Erikson, and Stimson's (1989) evidence that "macropartisanship" responds to changes in aggregate economic evaluations and approval of the incumbent president (see also Allsop and Weisberg 1988; Weisberg and Smith 1991) fails to support party system theories.[2] In contrast, Miller (1991), who disaggregates the national partisan balance between Republicans and Democrats by region, argues that among whites "the evidence of pervasive, long-term, aggregate stability outside the

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[1] The data and all RATS and MicroTSP computer programs used in this paper may be obtained by anonymous ftp from the Inter-university Consortium for Political and Social Research's Publication-Related Archive at ftp.icpsr.umich.edu/pub/PRA/outgoing/s1119.

[2] MacKuen, Erikson, and Stimson (1989) define macropartisanship as the aggregate percentage Democratic of all party identifiers. They analyze data from 1953 to 1987.

South is so dramatic" that a reexamination of change and persistence in party identification is essential.

The degree of permanence in partisanship has important implications for the way we think about individual partisan attitude change and shifts in aggregate partisanship. For instance, the magnitude of individual partisan persistence has ramifications for explanations of political behavior, such as voting, abstention, and political participation in general. Likewise, the degree of persistence in aggregate partisanship has significant consequences for our interpretations and predictions of local, state, and national election outcomes and trends in public opinion.

Dalton and Wattenberg (1993, 206) note that the extent of persistence in party identification remains an open question. In fact, they suggest that analysts of party identification place this issue at the top of their research agenda. We agree. In addition, we contend that a necessary step toward resolving questions about macro-level partisan persistence involves investigating the relationship between individual- and aggregate-level party identification.[3] Hence, the purpose of our paper is to explore links between individual and aggregate behavior and to use these hypothesized connections to guide our investigation into the magnitude of persistence in aggregate partisanship.

The paper is organized as follows. In part one, we address the question of micro/macro links. To do so, we identify common theoretical assumptions about individual partisan behavior and combine them with results on statistical aggregation to deduce the specific temporal pattern that each set of behavioral assumptions implies for aggregate levels of partisanship. In part two, we test these deductions about aggregate partisanship using diagnostic tests and a general time series model for fractional integration, both of which are new to political scientists. Our use of a time series model that imposes few restrictions on the data allows us to gain leverage over questions surrounding the rate at which shocks to macro-level partisanship dissipate over time. We find that previous conclusions about the degree of persistence in aggregate partisanship must be revised. In part three, we discuss the substantive import of our conclusions and suggest avenues for future research.

LINKING INDIVIDUAL AND AGGREGATE BEHAVIOR

A key difference between revisionist and traditional accounts of party identification revolves around the specification of causal forces in the individual decision calculus. Franklin and Jackson (1983) argue that the theoretical debates between the two sets of theories can be nested within the following individual decision mechanism:

$$x_{it} = \theta_0 + \alpha_i x_{it-1} + \delta z_{it} + \varepsilon_{it}, \qquad (1)$$

where x_{it} is the ith individual's current party identification, x_{it-1} is last period's partisan affiliation for the ith individual, and z_{it} denotes other systematic factors, such as evaluations of incumbent presidential performance (Fiorina 1981) or evaluations of issue proximity (Jackson 1975) that might influence an individual's party identification. Within the context of equation 1, traditionalists generally assume that $\delta = 0$, so that past party identification is the primary determinant of current party identification. In contrast, revisionists typically assume $\delta \neq 0$, which implies that other systematic factors can cause shifts in party affiliation. Both sets of researchers agree, however, that an individual's party identification is characterized by persistence via "an inertial element . . . that cannot be ignored" (Fiorina 1981, 102).

Characteristics of Individual-Level Behavior

We begin our analysis by thinking of the extent of permanence in party identification as ranging along a continuum from complete or perfect permanence to no persistence. We then consider a behavioral dimension anchored at opposite ends by the categories of fully homogeneous versus heterogeneous behavior across individuals.[4] Considering these two dimensions jointly results in the four cases of individual behavior depicted in Figure 1. We view these four cases as ideal types or approximations whose purpose is to facilitate our empirical research.[5] Beginning in the upper left-hand corner of Figure 1 and continuing clockwise, let us discuss each case in turn.[6]

Case 1: Complete persistence with homogeneous behavior. Here we consider the situation in which all individuals exhibit completely persistent partisan behavior from one period to the next. Within the context of equation 1, this assumption becomes $\alpha_i = \alpha = 1$. As revisionists, Franklin and Jackson (1983) argue that traditional theories of party identification imply exactly this set of assumptions—complete persistence by all individuals and no effects from short-run factors on party identification. Green and Palmquist's (1990) empirical evidence strongly supports these assumptions.

Case 2: Less than complete persistence with homogeneous behavior. In this case, we emphasize the possibility that individual party identifications exhibit a high but less than complete degree of permanence across all individuals. Relaxing the assumption of complete persistence produces conditions that can be expressed as $0 \leq \alpha < 1$.[7]

[3] Connections between the two levels of analysis have largely been ignored, although Converse's (1966) discussion of the normal vote contains an attempt at formalizing links between the two. MacKuen, Erikson, and Stimson (1989) also speculate that aggregation gain may explain the linkage.

[4] That is, we take seriously the arguments of Rivers (1988), who suggests that researchers too frequently have assumed that all individuals behave in the same way, and of Achen (1992), who argues assumptions of homogeneity are both highly prevalent and problematic within the literature on individual-level party identification.

[5] Our definition of ideal types relies heavily on Smelser's (1976, 55–6) and Burger's (1976, 133) discussions of Weberian ideal types.

[6] Although it may be possible to identify other ideal types, these four are mutually exclusive and exhaustive with respect to the two dimensions of individual behavior—extent of persistence and degree of homogeneity—that we deem essential to theories of micro-level partisanship.

[7] The existing empirical evidence (e.g., Franklin and Jackson 1983;

FIGURE 1. Four Ideal Types

Degree of heterogeneity	Degree of Permanence: Complete	Degree of Permanence: Less than Complete
Homogeneous	Case 1 $\alpha = 1$	Case 2 $0 \leq \alpha < 1$
Heterogeneous	Case 4 $0 \leq \alpha_i \leq 1$	Case 3 $0 \leq \alpha_i < 1$

In this situation, behavior does not vary across individuals, but individual party affiliations are no longer completely permanent. Some prominent theories and analyses that relax the assumption of complete persistence but do not make allowances for differences across individuals include Fiorina (1981), Markus and Converse (1979), Page and Jones (1979), and Weisberg and Smith (1991). Moreover, Achen (1992, 196–7) argues that a majority of the existing empirical work done by traditionalists falls into this category since a homogeneity assumption is built into the linear regression models most traditionalists use to test their hypotheses.

Case 3: Less than complete persistence with heterogeneous behavior. In this case, we continue to focus on situations in which individuals exhibit less than complete partisan persistence, but we now lift the restriction of homogeneous individual behavior. Within the context of equation 1, these assumptions imply that $0 \leq \alpha_i < 1$, thus ruling out complete partisan persistence while allowing for variation across individuals. Research which posits that the degree of partisan persistence varies for those of different ages or with different amounts of political experience (e.g., Beck 1974, Campbell et al. 1960, Converse 1969, Franklin 1984, Franklin and Jackson 1983, Shively 1979) is consistent with these assumptions. Converse's (1962) and Achen's (1992) hypotheses about the way in which partisan persistence varies as a function of new information also fit these assumptions.

Case 4: Complete persistence with heterogeneous behavior. Here we refer to the case in which the degree of partisan persistence varies across individuals over a range that includes instances of complete persistence. Within the context of equation 1, these assumptions can be expressed as $0 \leq \alpha_i \leq 1$. Miller's (1991) empirical evidence that the net balance of party identification for white nonsoutherners is completely persistent, while that for southerners is not, is potentially consistent with these assumptions.[8]

Implications for Aggregate Partisanship

Notice that in some instances the assumptions associated with each of these four cases fit existing explanations closely, while in others the fit is less tight. Our purpose in forming these ideal types is to use them to shed light on the magnitude of persistence in macrolevel party identification. Testing rival hypotheses about individual-level behavior is not our goal. Instead, we emphasize the way in which individual- and aggregate-level behavior are linked and the way in which these

Green and Palmquist 1990) suggests that the persistence parameters have a lower bound of zero. Existing theories do not suggest that party affiliations are explosive processes (i.e., $\alpha_i > 1$), so we do not consider such alternatives.

[8] We say "potentially" because, as Miller notes, it could also be the case that countervailing individual shifts in party identification could cancel one another and produce no change in the net balance of party identification.

micro-level assumptions provide insight into the degree of permanence in macro-level party identification.

To formalize the connections between individual- and aggregate-level partisan behavior, we will rely on existing theorems and results on statistical aggregation due to Theil (1954) and Granger (1980).[9] Combining these aggregation results with the assumptions for cases 1 through 4 will allow us to deduce the temporal path that aggregate measures of party identification will follow for all cases. To begin, assume that there are N_j individuals who state that they identify with one of the $j = \{1, 2\}$ major political parties. That is, we want to aggregate separately for Republican and Democratic identifiers. Furthermore, assume that the behavioral decision mechanism for each individual (regardless of party identification) is given by equation 1. Aggregating over the individual equations for one party's identifiers will yield $x_{jt} = \Sigma_{i=1}^{N} x_{it}$, which is an aggregate value of partisanship at a single point in time for the jth political party.[10] Performing this aggregation at multiple points for each set of identifiers produces a macro-level time series for each party. Using this approach, we now consider the degree of persistence that we can expect aggregate measures of party identification to exhibit given the behavioral assumptions associated with each ideal type.

Assumption A: $\delta = 0$

Since the values of α_i provide direct information about the degree of persistence in individual-level partisanship, we will initially assume that $\delta = 0$ in equation 1. In the next section we relax this assumption. By setting $\delta = 0$, we can focus solely on the persistence parameter.[11]

Case 1. In this instance behavior is assumed to be homogeneous across all individuals, and the degree of persistence is complete. The assumption that behavior is the same for all individuals allows us to assume that the coefficients in the micro-level equation are constant and independent of x_{it-1}, and this makes aggregation straightforward.[12] Granger (1980) and Theil (1954) show that aggregating over the behavior depicted in equation 1, when coefficients are constant, produces a macro-level time series that is *integrated of order one*.[13] Time series that are integrated of order one, denoted I(1), are also called unit root processes. Such time series exhibit long stochastic swings up or down and do not return to a constant mean level. In addition, the effects of a shock to these processes persist permanently.[14]

Case 2. In this case, behavior is homogeneous, but the degree of persistence is less than complete. This situation represents a case in which the α coefficient is constant across all individuals, and hence exact aggregation is possible. Under the case 2 assumptions, Theil's (1954) results imply that the aggregate time series will be *integrated of order zero*, denoted I(0).[15] Such time series are also called stationary processes. The effects of a shock to a stationary process are transitory and dissipate quickly over time as the series returns to its constant mean level.

Case 3. This case presents a situation in which behavior is no longer homogeneous, and the degree of persistence is less than complete, such that $0 \leq \alpha_i < 1$. Because of the heterogeneity, aggregation becomes more complex. Yet, Granger (1980) proves that, under these circumstances, if N is large and the α_i are drawn from a beta distribution on the range (0,1) with $p,q > 0$,[16] then aggregating over the behavior depicted in equation 1 when $\delta = 0$ will yield a macro-level time series that is *fractionally integrated* (Granger 1980; Granger and Joyeux 1980; Hosking 1981).[17]

Fractionally integrated time series, denoted I(d) with $0 < d < 1$, differ from stationary processes because they exhibit significant dependence between observations—even those that are separated by very long time spans. Fractionally integrated series also differ from unit root processes because the effects of a shock do dissipate over time as the series reverts to its mean. The rate at which the effects dissipate, however, is slower for a fractionally integrated time series than for a stationary process.

Case 4. In this situation behavior is heterogeneous over a range that includes complete persistence. Based on the algebra of integrated series (e.g., Granger and Newbold 1977), Granger (1980) argues that aggregating under these conditions produces a time series that is *integrated of order one*. That is, the aggregate time series is equal to the sum of a fractionally integrated time series (resulting

[9] A key assumption underlying these theorems is that individuals answer questions about party identification sincerely rather than strategically so that individual decision mechanisms are independent across all individuals. We later relax the independence assumption.

[10] We follow Kramer's (1983) practice of thinking of the aggregation as over a latent, continuous variable ranging from $-\infty$ to $+\infty$ that maps into discretely measured observations depending upon where x_{it} falls with respect to threshold values along the continuum. This latent index tells us about the strength of party affiliation for each individual.

[11] As does Granger (1980), we also set $\theta = 0$, which implies that each aggregate series will contain no drift.

[12] If we believe party identification is measured with random error, then we need to assume $E(\alpha \mid x_{it-1}) = 0$ for all individuals for consistent aggregation (Theil 1954, Rivers 1988).

[13] The intuition behind this aggregation result is that since each individual's behavior is characterized by an autoregressive process with $\alpha_i = 1$, then summing them up and dividing by N produces the same average macro-level relationship.

[14] Additional details about unit root processes can be found in Durr (1993) and Ostrom and Smith (1993).

[15] The intuition here is the same as in case 1: Because each individual's behavior follows an I(0), stationary autoregressive process with the same degree of persistence, so will its macro-level counterpart.

[16] The beta distribution is given by:

$$f(\alpha | p, q) = \frac{\alpha^{p-1}(1-\alpha)^{q-1}}{\frac{\Gamma(p)\Gamma(q)}{\Gamma(p+q)}}$$

The shape of the distribution depends on the values of p and q. When $p = q$, the distribution is symmetric; when $p, q > 1$, the distribution is concave; when $p, q < 1$, the distribution is convex. Given the flexibility of the beta distribution, this assumption seems noncontroversial. In our work, the q parameter indicates how micro-level partisan persistence maps into macro-level persistence. All else equal, the smaller the value of q, the larger is the mean of the beta distribution of persistence parameters, and the more persistent will be the aggregate series.

[17] Specifically, $x_t \sim I(1 - q/2)$, where q is a parameter in the beta distribution. Because the proof of this result is fairly technical, interested readers are referred to the appendix and to Granger (1980).

from aggregation over the individuals for whom persistence is less than complete) and a unit root process (due to aggregation over individuals for whom $\alpha_i = 1$). Because the variance of an integrated series grows as N gets large, the higher order unit root dynamics dominate the lower order dynamics from fractional integration; thus, the resulting macro-level partisanship series will contain a unit root.

Assumption B: $\delta \neq 0$

The previous aggregate-level predictions were based on the assumption that $\delta = 0$, implying that no political factors or events affect individual party identification or the extent of macro-level persistence. Yet, under the revisionist conceptualization of party identification, individual partisan affiliations are influenced by current political forces. Hence, the question we now address is whether and how the magnitude of persistence in aggregate partisanship may be affected by the features of the z_{it} in equation 1; that is, what linkages arise when we relax the $\delta = 0$ restriction?

To answer this question, consider the following case (Granger 1980):

$$x_{it} = \alpha_i x_{it-1} + \delta z_t + \varepsilon_{it}, \tag{2}$$

where the ε_{it} are independent for all individuals, the z_t represent other political factors whose values are taken to be common across individuals, and all else is as defined in equation 1. Technically, the assumption that the z_{it} are common to all individuals implies that $z_{it} = z_t$ for all i, which in turn makes the x_{it} dependent across individuals. This assumption allows the properties of the aggregate z_t time series to influence the time series characteristics of the aggregated partisanship series. In addition, this assumption allows factors like unemployment or inflation, battle deaths, or presidential performance ratings to influence an individual's party identification. Given these conditions, we now reexamine our predictions for the degree of persistence in aggregate levels of partisanship.

Cases 1 and 4. Granger (1980) argues that, under the assumptions that $\alpha = 1$ and $\delta \neq 0$ (i.e., case 1), aggregating over the behavior in equation 2 continues to yield a macro-level time series that is integrated of order one regardless of the time series properties of the z_t political factors. Similarly, aggregating in case 4 under the assumptions that $0 \leq \alpha_i \leq 1$ and $\delta \neq 0$ will also yield an aggregate partisanship series that contains a unit root. Once again, the intuition is that the unit root dynamics contributed by those individuals for whom $\alpha_i = 1$ dominate the time series properties contributed by the z_t.

Cases 2 and 3. When we reconsider cases 2 and 3, however, we find that different predictions result. Recall that for case 2, under the assumptions that $0 \leq \alpha < 1$ and $\delta = 0$, aggregate party identification will be stationary. But if $\delta \neq 0$ and the z_{it} are common across individuals and are generated by a process with higher order dynamics (e.g., fractionally integrated or integrated of order one), the resulting macro-level party identification series will be integrated of the same order as the highest order of integration among the components contributing to it (Granger 1980). Thus, in case 2, assuming that $0 \leq \alpha < 1$ and $\delta \neq 0$, aggregate partisanship could be a stationary series, a fractionally integrated process, or a nonstationary unit root process. Finally, consider case 3 when the assumptions $0 \leq \alpha_i < 1$ and $\delta \neq 0$ hold. If we continue to assume that the α_i follow a beta distribution on (0,1), then the aggregate partisanship series, x_t, will be dominated by the higher degree of integration contributed by either the z_t or ε_{it} (Granger 1980).[18] Hence, for case 3 under assumption B, macro-level party identification could be either fractionally integrated or integrated of order one.[19]

FIGURE 2. Predicted Dynamic Properties of Aggregate Partisanship

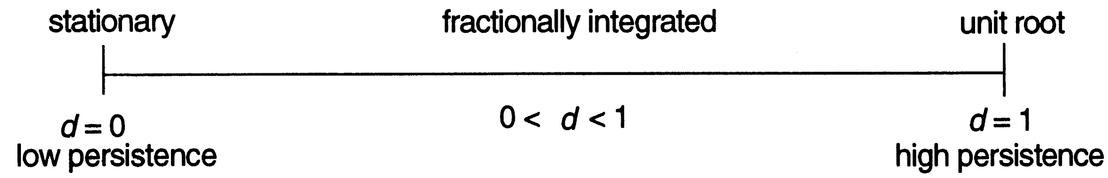
	Assumption A $\delta = 0$	Assumption B $\delta \neq 0$
Case 1 $\alpha = 1$	Unit Root	Unit Root
Case 2 $0 \leq \alpha < 1$	Stationary	Stationary, Unit Root, or Fractionally Integrated depending on dynamics of z_t
Case 3 $0 \leq \alpha_i < 1$	Fractionally Integrated	Unit Root or Fractionally Integrated depending on dynamics of z_t
Case 4 $0 \leq \alpha_i \leq 1$	Unit Root	Unit Root

By combining assumptions about individual behavior with existing aggregation results we have deduced the temporal path of aggregate partisanship for each of the four ideal types under two different assumptions about the effects of other exogenous political factors. Figure 2

[18] These degrees of integration are denoted $I(d_z + 1 - q)$ and $I(1 - q/2)$, respectively.

[19] When other exogenous factors that are independent across individuals, say, w_{it}, are included in equation 2, Granger (1980) shows that x_t will be $I(d_x)$, where d_x is the largest of three components: $1 - q/2 + d_w$ (from the w_{it}), $1 - q + d_z$ (from the z_t), and $1 - q/2$ (from the ε_{it}).

FIGURE 3. Persistence Properties of Time Series



summarizes our predictions. Here we see that evidence of stationarity would be consistent only with the case 2 assumptions. Evidence of fractional integration would be consistent with the case 3 assumptions or with the case 2 assumptions when $\delta \neq 0$. Finally, empirical evidence of a unit root would be consistent with cases 1 and 4 or with cases 2 and 3 when $\delta \neq 0$. Although these predictions for macro-level partisanship overlap in some instances, empirical tests of the degree of persistence in aggregate party identification can, nonetheless, provide us with important insight into the nature of macro-level partisan persistence and its links to individual-level heterogeneity and permanence.

TESTING THE ALTERNATIVE PREDICTIONS

From a technical viewpoint, distinguishing among the predictions in Figure 2 requires that we estimate the degree of persistence in aggregate partisanship. Figure 3 shows the empirical patterns of persistence associated with various types of time series. For a stationary series, shocks will show little persistence, while shocks to a first-order autoregression process [i.e., AR(1)] with a unit root will exhibit high levels of persistence. For fractionally integrated time series, the degree of persistence varies such that the higher the value of d, the greater the degree of persistence. To investigate the degree of persistence in aggregate partisanship and test the predictions in Figure 2, we use two approaches. First, we employ diagnostic tests of the general class into which a time series falls (i.e., unit root, variance ratio, and KPSS tests). Second, we use a maximum likelihood estimator to obtain point estimates of d, the order of integration. These point estimates provide direct information about the degree of persistence.

The fact that we focus on empirical estimates of the degree of persistence in aggregate partisanship becomes very important once we take up issues of measurement. We measure aggregate levels of partisanship from the first quarter of 1953 through the fourth quarter of 1992 as the percentage of Republican identifiers, the percentage of Democratic identifiers, and the Democratic percentage of the major party identifiers. This latter measure is MacKuen, Erikson, and Stimson's (1989) "macropartisanship" series.[20] That is, we use percentages, which are bounded by 0 and 100, as proxies for the aggregate number of party identifiers.[21] Some analysts question the meaningfulness of unit root tests in bounded data. Since unit root processes have asymptotically infinite conditional variances, these researchers argue that we cannot expect a bounded variable to have a unit root (DeBoef and Granato n.d.; Williams 1993).

In our analysis, however, we emphasize the degree of persistence in a finite sample rather than the asymptotic distribution of the OLS estimator; hence, the observations of Hamilton (1994) apply. As Hamilton points out, when the question is one of persistence, unit root tests are indeed relevant. According to Hamilton (1994, 447), "we can measure whether innovations have much persistence over a fixed interval [with a unit root process as the true model] or very little persistence over that interval [with a stationary model as the true model]." Thus, we contend that, even for bounded data, unit root tests provide meaningful information about the degree

TABLE 1. Dickey-Fuller Tests for Unit Roots

	DF	ADF[a]	Joint Test
% Republican	−2.475	NA	3.175
% Democrat	−1.942	−1.130	2.821
Macropartisanship	−2.198	−1.598	2.423
Critical values:			
α = .01	−3.47[b]	−3.47	8.43[c]
α = .05	−2.88	−2.88	6.34
α = .10	−2.58	−2.58	5.39

Null hypothesis for DF/ADF tests: unit root process, d = 1
Alternative hypothesis: stationary process, d = 0

Null hypothesis for joint test: unit root process with no linear trend
Alternative hypothesis: stationary process

[a] When we detected autocorrelation in the residuals from the DF regression, we estimated ADF regressions using enough lagged first differences to make the residuals white noise.
[b] The DF and ADF statistics have a nonstandard distribution. The appropriate critical values (shown above) are given in MacKinnon (1991).
[c] The statistic for the joint hypothesis has a nonstandard distribution. The appropriate critical values (shown above) are given in Dickey and Fuller (1981).

[20] The macropartisanship series ends with the second quarter of 1992. We thank Jim Stimson for these data. The macropartisanship series has been corrected for potential biases from telephone polling, while the others have not. These biases could make long-run persistence more difficult to detect. All data are taken from responses to the Gallup Poll's party identification question. The summary statistics for each series are: % Republican—mean = 27.90, std. dev. = 3.92, N = 160; % Democratic—mean = 42.90, std. dev. = 5.42, N = 160; macropartisanship—mean = 61.34; std. dev. = 4.68; N = 158.

[21] In the best case scenario, we would want to use the number of people identifying with a political party. These numbers could grow and decay over time. Unfortunately, the only data available are from samples of individuals, and hence we must use percentages to account for varying sample sizes. These percentages can grow or decay but are admittedly bounded by 0 and 100.

TABLE 2. Variance Ratio Tests

	VR (k)				
	VR(2)	VR(4)	VR(8)	VR(16)	VR(32)
% Republican	1.212	1.593	1.791	2.650	2.051
% Democrat	1.508	1.845	2.549	2.997	3.033
Macropartisanship	1.159	1.360	1.516	2.062	1.872
Critical values:					
α = .05	1.157	1.328	1.609	2.170	3.578
α = .10	1.122	1.254	1.465	1.891	2.909

Null hypothesis: random walk with drift, $d = 1$
Alternative hypothesis: pure fractional noise, $d < 1$

Recommended differencing interval (Diebold 1989:) $k = 4$ or $k = 8$

of persistence in a time series, helping analysts determine where a series falls along the continuum in Figure 3. In addition, we rely on multiple tests of persistence, which help to ensure that our conclusions are robust.

Results of Diagnostic Tests

We begin our empirical analysis by using diagnostic tests to investigate the time series properties of the aggregate partisanship data. Although these tests cannot always pinpoint the order of integration for a time series, the evidence they provide can be used to determine whether a time series is "closer" to being stationary, with $d = 0$ and low persistence, or to exhibiting unit root behavior, with $d = 1$ and high persistence. In general, we agree with Lo (1991, 1296) that these tests can be used to complement evidence obtained from point estimates of d. We employ unit root, variance ratio, and KPSS tests as our diagnostic tools.[22]

Table 1 presents the results of the conventional Dickey-Fuller (DF) tests for unit roots. As the results in column 1 show, the null hypothesis that each of these series has a unit root cannot be rejected. As shown in column 2, this conclusion continues to hold when Augmented Dickey-Fuller (ADF) tests are applied to the macropartisanship and Democratic series.[23] Furthermore, a joint test of the null hypothesis of a unit root but no time trend indicates that these series are not deterministic functions of time. To summarize, the evidence in Table 1 suggests that each of these series exhibits unit root behavior and is highly persistent.

Dickey-Fuller tests, however, have low power in the face of fractionally integrated alternatives.[24] To compensate for this, we also used variance ratio tests (Cochrane 1988; Diebold 1989) of the null hypothesis of a random walk with drift, i.e., $d = 1$, versus an alternative of pure fractional integration with $d < 1$. The results of these variance ratio tests appear in Table 2 for various choices of k, the differencing interval. As these results show, the null hypothesis of a unit root is rejected for both the Republican and Democratic series when $k = 2, 4, 8$, and 16. When applied to macropartisanship, the variance ratio test rejects the null hypothesis at the $\alpha = 0.05$ level for $k = 2$ and 4 and at the $\alpha = 0.10$ level for $k = 2, 4, 8$, and 16.[25] On the whole, the evidence in Table 2 suggests that each of these three series is fractionally integrated with $d < 1$ and with a degree of persistence that varies depending upon the value of d. This implies that aggregate partisanship exhibits a lower degree of persistence than is found in an AR(1) process with a unit root.

Finally, in Table 3, we present the results from KPSS tests of the null hypothesis that each series is a strong mixing process.[26] We present these results for various choices of ℓ (lag truncation parameter), which affects the number of lagged autocovariances used to compute the long-run variance of the series.[27] As the empirical evidence in Table 3 shows, the null hypothesis that these series are strong mixing can be rejected at the $\alpha = 0.05$ level for all values of the lag truncation parameter from 0 through 6. These results suggest that these partisanship time series are not stationary processes and that $d > 0$ for each. This implies that aggregate partisanship exhibits a higher degree of persistence than is found in a stationary AR(1) process.

Taken together, the evidence in tables 1 through 3

[22] Since unit root tests are only one of three diagnostic tests that we employ, and since we estimate the order of integration directly, the consistent empirical results obtained from these multiple tests suggests that our conclusions are robust to the type of test employed.

[23] Dickey-Fuller tests are biased when the residuals are serially correlated. When we detected autocorrelation in the residuals of the DF regressions, we estimated ADF regressions using enough lagged first differences to make the residuals white noise.

[24] DeBoef and Granato show that DF tests also have low power when the alternative hypothesis is one of near integration. See DeBoef and Granato (n.d.) for more information on near-integrated time series.

[25] Diebold's (1989) simulations suggest that for our sample size, the variance ratio test is most powerful when $k = 4$ or $k = 8$. Diebold's Monte Carlo results show that when $d = 0.70$, the power of the test is 0.889 for a 95% confidence interval and is 0.95 for a 90% confidence interval. When $d = 0.80$, the power of the test is 0.56 for a 95% confidence interval and is 0.70 for a 90% confidence interval.

[26] Intuitively, a time series is strong mixing if the rate at which dependence between past and future observations goes to zero as the distance between them grows is "fast" enough (Lo 1991). Stationary series, which decay at a geometric rate, are strong mixing processes; fractionally integrated series, which decay at a hyperbolic rate, and unit root processes, which do not decay, are not strong mixing. Hence, determining whether aggregate party identification measures are strong mixing provides general information about rate of decay and degree of persistence.

[27] Simulations by Kwiatkowski et al. (1992) show that $\ell = 4$ provides the most powerful test, given our sample size. Their Monte Carlo results show that when $\ell = 4$, the power of the test is greater than or equal to 0.92 for a 95% confidence interval in all typical circumstances.

TABLE 3. KPSS Tests for Strong Mixing

	Lag Truncation Parameter (ℓ)				
	0	2	4	6	8
% Republican	2.852	1.011	.629	.465	.374
% Democrat	8.784	3.112	1.932	1.425	1.145
Macropartisanship	3.159	1.124	.706	.529	.431
Critical values:					
α = .01	.739	.739	.739	.739	.739
α = .05	.463	.463	.463	.463	.463
α = .10	.347	.347	.347	.347	.347

Null hypothesis: strong mixing process
Alternative hypothesis: not strong mixing
Recommended lag truncation (Kwiatkowski et al. 1992): $\ell = \text{int}[4(T/100)^{.25}] = 4$

suggests that aggregate measures of Republican and Democratic identifiers exhibit neither stationary nor unit root behavior. Similarly, the macropartisanship series is not stationary. In the case of macropartisanship, however, the evidence against the unit root hypothesis is not quite as strong as it is for the other two series., The pattern of hypothesis rejections for the Republican and Democratic time series is consistent with the conclusion that these series are fractionally integrated. The evidence for macropartisanship suggests that this series is fractionally integrated, although high persistence and unit root behavior cannot be ruled out conclusively.

Point Estimates of the Order of Integration

While they are useful tools, these diagnostic tests do not allow us to draw precise inferences about the degree of persistence in aggregate partisanship. To determine which of the predictions in Figure 2 is consistent with the aggregate-level evidence, we need point estimates of d, the order of integration, for each of these time series. To obtain estimates of d, we use the full information maximum likelihood (ML) estimator derived by Sowell (1992a, 1992b).[28]

The ML estimates of d appear in Table 4.[29] Here we see that each of the three partisanship series is fractionally integrated, since for each series the hypothesis that $d = 0$ (i.e., stationarity) as well as the hypothesis that $d = 1$ (i.e., unit root) can be rejected at conventional levels of statistical significance.[30] In addition, we see that each of these partisanship series exhibits a high degree of persistence, since the estimates of d are 0.804 for the Republican series, 0.698 for the Democratic series, and

TABLE 4. Maximum Likelihood Estimates of Degree of Fractional Integration (*d*)

	Parameter Estimates[a]
% Republican	.804[b]
	(.057)
% Democrat	.698
	(.058)
Macropartisanship	.839
	(.064)

[a] The standard errors of the estimates are shown in parentheses.
[b] The ML "t-ratios" for tests of the null hypothesis that $d = 1$ are −3.44, −5.21, and −2.52 for the Republican, Democratic, and macropartisanship series, respectively. The ML "t-ratios" for tests of the null hypothesis that $d = 0$ are 14.12, 12.03, and 13.11 for the Republican, Democratic, and macropartisanship series, respectively.

28 Sowell's (1992a) FIML estimator of d is always applied to first-differenced data so that stationarity is forced to hold. Then assuming that $X_T = \{x_1, x_2, \ldots, x_T\}$ is a normally distributed sample of T observations, the likelihood function of the ARFIMA(p,d,q) process is given by:

$$L(X_T|\Sigma) = (2\pi)^{-T/2}|\Sigma|^{-1/2}\, e^{(-1/2X'\Sigma^{-1}X)}$$

where Σ, the $T \times T$ Toeplitz covariance matrix of X_T, is a function of $d, \phi_1, \ldots, \phi_p, \theta_1, \ldots, \theta_q$, and σ^2. This function must be maximized with respect to Σ, d, and the p autoregressive and q moving-average parameters. To do so, the spectral density of x_t needs to be written in terms of the model parameters, and then the autocovariance function is calculated by:

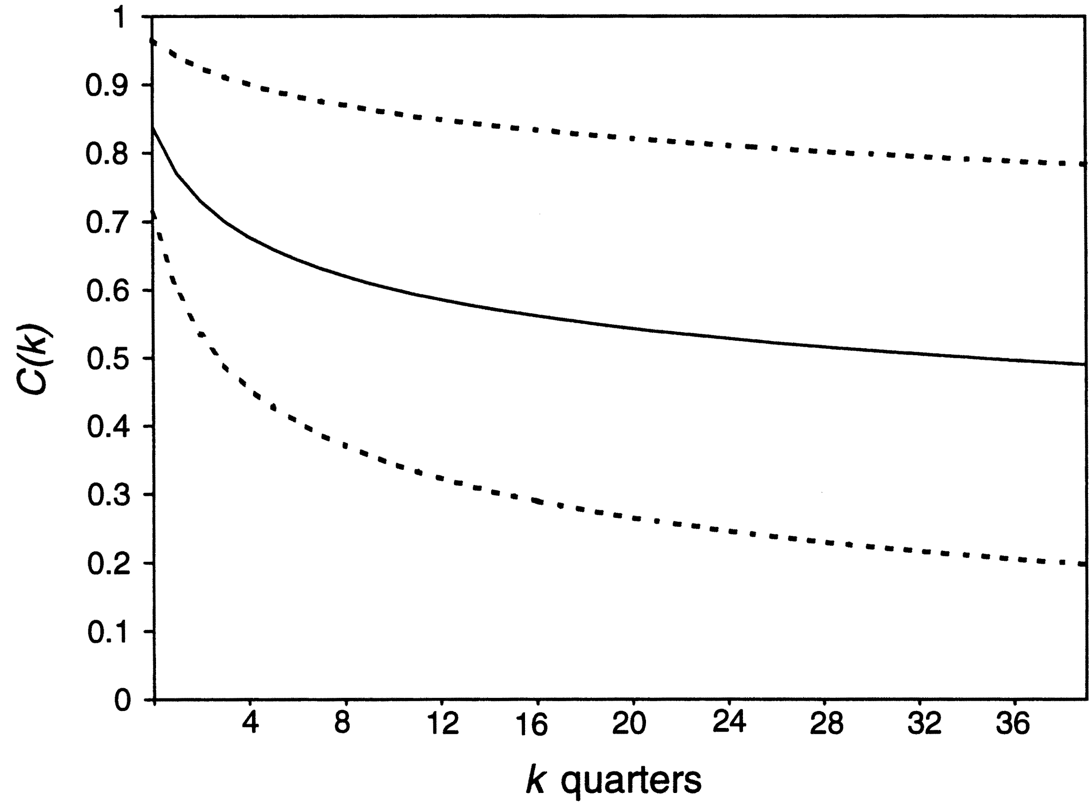
$$\gamma(s) = \frac{1}{2\pi}\int_0^{2\pi} f_x(\lambda)e^{i\lambda s}\, d\lambda$$

which Sowell proves is well approximated for various lag lengths by sums of a hypergeometric function. Thus, the assumption of a beta distribution never enters into the likelihood but is used only to establish micro/macro links.

29 We also estimated d using a popular nonparametric, spectral regression-based procedure, called the GPH estimator after its developers, Geweke and Porter-Hudak (1983). Our GPH results indicate that for each series the null hypothesis of stationarity can be rejected, but the GPH estimates of d cannot be statistically distinguished from a value of 1.0 due to limited degrees of freedom and large standard errors. Sowell's (1992a) simulations show that his ML estimator generally has less bias and a smaller MSE in small samples than does the GPH estimator. Hence, we use Sowell's time domain ML estimator.

30 It is possible for a time series to be fractionally integrated (FI) and to contain short-run autoregressive (AR) or moving average (MA) components. Such series are called ARFIMA(p,d,q) processes. The results in Table 4 are based on estimation of ARFIMA(0,d,0) models with no AR or MA components. To test our restrictions, we estimated ARFIMA(p,d,q) models with up to three AR and three MA parameters each. Based on likelihood ratio tests comparing the various unrestricted models to the restricted model, we conclude that the restricted ARFIMA(0,d,0) cannot be rejected.

FIGURE 4. Impulse Response Function for Macropartisanship Series



0.839 for the macropartisanship series.[31] The closer the absolute value of d to 1, the more persistent are the effects of shocks to these series.

The Magnitude of Aggregate Persistence

Additional insight into the degree of permanence in macro-level partisanship can be obtained by examining the cumulative impulse response function for each series.[32] By considering the entire sequence of cumulative impulse responses, we can investigate the effect on aggregate partisanship of a shock to each series after one quarter, one year, one presidential administration, or even one hypothesized party system period.

The cumulative impulse response functions, denoted $C(k)$, and their associated confidence intervals are presented separately for each partisanship series in figures 4 through 6.[33] In these figures, the solid lines indicate the degree to which the effect of a one-standard-deviation shock to each series persists over time.[34] For instance, consider the response function for the macropartisanship series shown in Figure 4. Here we find that the degree to which a shock persists drops from 0.84 to 0.70 after four quarters (one year). After 16 quarters (four years), the degree of persistence has only been cut by about one-third, since $C(16) = 0.57$. In fact, as the confidence intervals make clear, the degree of persistence in macropartisanship remains statistically different from zero even after 40 quarters (ten years). As shown in figures 5 and 6, similar patterns of persistence are found in the Republican and Democratic series.[35] The slow decay of these cumulative impulse response functions

[31] The ML estimates were obtained through the Davidon-Fletcher-Powell algorithm. We follow Sowell (1992a) and use the GPH estimates of d as our starting values. We checked numerous other values, however, and found that our ML estimates of d are highly robust to the starting value used. The standard errors of d for the Republican and macropartisanship series are robust to the starting value used for d, although the standard error of d for the Democratic series varied somewhat when starting values near zero were used. To ensure we had convergence of the standard error of d, we restarted the DFP algorithm at exactly the point it stopped and ran additional iterations.

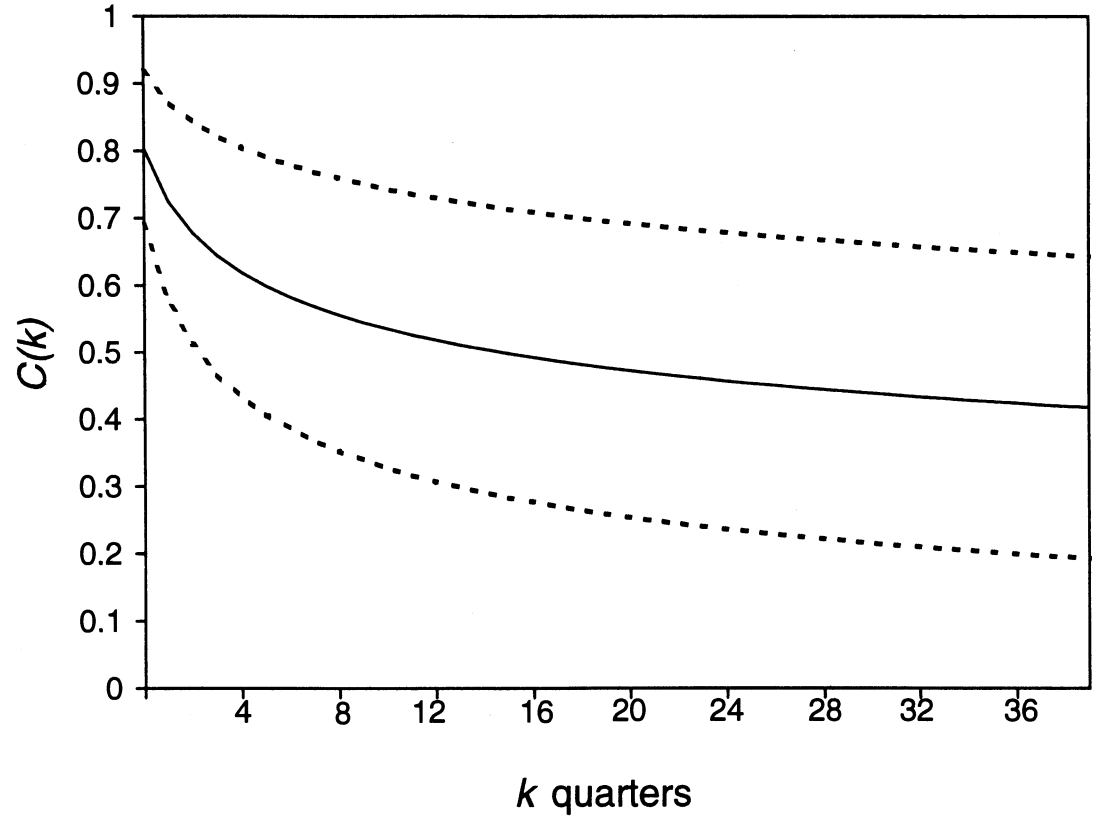
[32] The cumulative impulse response function represents the sum of the coefficients of the moving-average lag-operator polynomial of the differenced series. See Diebold and Rudebusch (1989) for a useful discussion of the impulse response function for fractionally integrated time series.

[33] The confidence intervals (represented by dashed lines) are constructed from ± 2 standard errors of $C(k)$. As Hamilton (1994, 336–40) points out, there are alternative approaches to calculating these standard errors, especially for the orthogonalized impulse response function of a vector autoregression (e.g., Runkle 1987). Since we are working in the univariate context, many of these approaches are not relevant to our work. Hence, we follow Diebold and Rudebush (1989) and Hamilton (1994) and use a numerical approach for nonorthogonalized impulse responses to calculate the standard errors.

[34] More precisely, these figures show the effect of a one-standard-deviation shock to the level of each partisanship series for up to k future periods, where k ranges from 1 to 40 quarters (see also Cheung 1993).

[35] We note that the degree of persistence in the Democratic series is significantly different from the extent of permanence in the other two series. Substantively, this implies an asymmetry in the degree of partisan persistence for Democrats and Republicans. This, in turn,

FIGURE 5. Impulse Response Function for Republican Series



indicates that increases and decreases in aggregate levels of partisanship are "permanent" in the sense that they last for years.[36]

Ruling out a Rival Hypothesis

While we have focused on individual-level heterogeneity and the features of the z_{it} political factors as the sources of the dynamics of aggregate partisanship, it is also important to rule out the hypothesis that generational replacement is responsible for the high degree of persistence found in our data.[37] The issue is whether population changes affect the distribution of new entrants' party identifications such that these distributions change over time. For instance, there is some evidence that individuals who entered the electorate in recent years are more Republican than in previous years (Abramson, Aldrich, and Rohde 1995).

Whether such population changes are the source of the dynamic patterns of fractional integration in the aggregate partisanship data is an empirical question that can be addressed by dividing the data into subsamples and reestimating the persistence parameter, d. For instance, in case 3, under the assumption that $\delta = 0$, the estimated value of d is mathematically related to the mean of the beta distribution of α such that the higher the estimate of d, the higher is this mean. In all cases, stability of the estimates of d across periods is inconsistent with the hypothesis that population changes are the primary cause of the degree of permanence in aggregate measures of party identification.

To test the stability of the coefficient estimate of d for the macropartisanship series, we used subsamples of approximately two decades in length, one from the second quarter of 1953 through the fourth quarter of 1972 and the other from the first quarter of 1973 through the second quarter of 1992. Our choice of these subsamples was based on Abramson's (1983, 54) observation that "during the postwar era, nearly half the white electorate has been renewed every two decades." The estimate of d for macropartisanship during the period from the second quarter of 1953 to the fourth quarter of 1972 is 0.746, while the estimate for the period from the first quarter of 1973 to the second quarter of 1992 is 0.953. Each of these coefficient estimates falls into the 95% confidence interval of (0.714, 0.964) around the original estimate of $d = 0.839$ for the entire sample.[38] The fact that these coefficients are not statistically different from each other suggests that they are stable during periods of population change, and this evidence

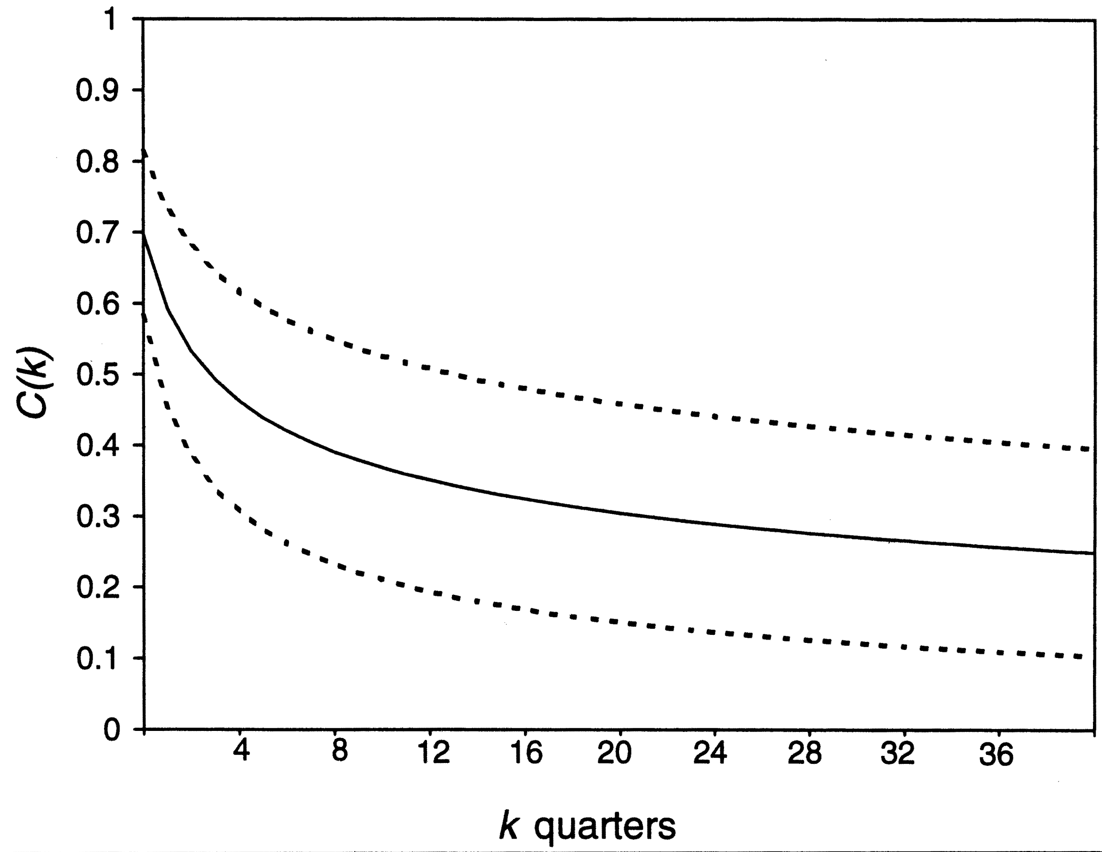
could imply that the individual-level sources of heterogeneity exert a slightly different influence on Democrats than on Republicans.

[36] The degree of permanence in aggregate partisanship continues to weaken after the ten-year period. For instance, after twenty years, $C(k) = 0.365$ for the Republican series, $C(k) = 0.204$ for the Democratic series, and $C(k) = 0.439$ for the macropartisanship series.

[37] We thank a reviewer for raising this point.

[38] The results for Democratic and Republican identifiers are similar—0.713 and 0.677 for Democrats and 0.560 and 0.884 for Republicans.

FIGURE 6. Impulse Response Function for Democratic Series



refutes the idea that such changes are the primary cause of aggregate partisan persistence. We thus reject this rival hypothesis.

DISCUSSION AND IMPLICATIONS

The empirical evidence in tables 1 through 4 can be used to resolve certain questions about the nature and degree of persistence in aggregate-level party identification and its links to individual partisanship. The evidence shows that aggregate levels of partisanship are fractionally integrated. Since this empirical result disconfirms the predictions associated with cases 1 and 4, we can conclude that theories in which all individuals are assumed to maintain completely persistent party identifications provide an inadequate micro-level explanation of the dynamics of aggregate party identification. We also see that the stationary and unit-root predictions from case 2 and the unit-root prediction from case 3 fail to be supported by the empirical evidence.

Instead, our evidence that aggregate measures of party identification are fractionally integrated is consistent with only two instances—case 3, in which individuals exhibit heterogeneous and incompletely persistent party identifications, or case 2, in which individuals exhibit homogeneous behavior and less than complete persistence when $\delta \neq 0$ and the z_t follow a fractionally integrated process. In case 3, the persistence in macro-level partisanship is solely a function of individual-level heterogeneity. In case 2, the degree of macro-level persistence can be a function of an aggregate z_t series that is fractionally integrated or of a mixture of individual-level heterogeneity and long memory time series properties for z_t.

Implications for Micro-Level Party Identification

Our evidence of fractional integration in aggregate levels of party identification along with our discussion of potential micro/macro links help us refocus the attention of analysts of individual party identification on the issues of the heterogeneity of the α coefficients and on the degree of persistence expected over time in the exogenous z_{it} factors. Given our evidence, analysts of micro-level party identification must now ask such questions as: Are issue proximities persistent enough to cause the kind of aggregate-level permanence we find? Do evaluations of the incumbent president's performance contain a permanent component that would explain the fractionally integrated dynamics of aggregate partisanship? Furthermore, these researchers must open their theoretical and empirical analyses to the possibility of behavioral heterogeneity.

Finally, it seems most likely that the above issues cannot be investigated separately. Instead, future research into individual-level party identification will need to specify both the nature of behavioral heterogeneity (through the α_i) and the dynamic patterns of any political factors (the z_{it}) thought to influence party identification. Only those researchers who do so will be able to provide a solid microfoundation for aggregate partisanship.

Implications for Aggregate-Level Partisanship

Our empirical evidence that aggregate measures of party identification as well as MacKuen, Erikson, and Stimson's (1989) macropartisanship series are fractionally integrated also has serious implications for conclusions about the magnitude of persistence in macro-level party identification. Based on their analyses, MacKuen, Erikson, and Stimson (1989, 1139) reject the long-cycle hypothesis of party system theorists and conclude that partisan "gains and losses are 'permanent' on a scale of months, not decades."

But as shown in figures 4 through 6, our analysis sheds new light on the nature of the "mid-range dynamics" of aggregate partisanship.[39] As the evidence shows, more than one-half the effects of any shock to aggregate partisanship persist for as long as a four-year presidential term. More than one-quarter of the effects remain after eight years. Thus, based on our evidence, we also reject the party system theorists' arguments about decades-long cycles of partisanship. But, in contrast to MacKuen, Erikson, and Stimson (1989), we conclude that shocks to aggregate partisanship are "permanent" on a scale of years—not months.

The disparity between our conclusions and those of MacKuen, Erikson, and Stimson arises from their assumption that macropartisanship is stationary with $d = 0$. Our evidence shows that the stationarity hypothesis can be rejected in each of our three aggregate time series. Because MacKuen, Erikson, and Stimson (1989) impose the restriction $d = 0$ in their statistical analyses, they automatically constrain the rate of partisan decay to the quicker, short-term types associated with stationary time series. In contrast, we use a more general model that allows for a range of decay patterns.[40] Our use of the more general model allows us to *estimate*, rather than to assume, the degree of persistence in aggregate levels of partisanship.[41]

What are the implications of aggregate partisan behavior that is persistent over a period of years but not months or decades? As one example of the consequences of our conclusion, consider Carmines and Stimson's (1989) work on issue evolution in the U.S. electorate. As they argue, the temporal path of an issue evolution is intimately intertwined with the dynamics of aggregate partisanship. If issue evolution does indeed depend on the dynamics of partisanship, as Carmines and Stimson suggest, then our empirical evidence implies that the complete evolution of a partisan cleavage on a particular issue could be accomplished by the party of a two-term president and need not take "a very long time to complete" (Carmines and Stimson 1989, 193). Thus, our evidence that shocks to partisanship persist for years, in conjunction with our look at potential micro-level assumptions accounting for these aggregate patterns of growth or decay, provides a conceptually richer and empirically more precise basis for the concept of issue evolution.

Similarly, our empirical evidence that shocks to aggregate partisanship persist for years, but not months or decades, is consistent with Gerber and Jackson's (1993) evidence about endogenous preferences. They found changing relationships between partisanship and preferences over an eight-year period (for civil rights) and a four-year period (for the Vietnam War). We, too, find evidence of a years' long dynamic within the macro-level party system between 1953 and 1992.

In addition, our work has implications for controversies about the effects of question wording on conclusions about the degree of persistence in aggregate partisanship (Abramson and Ostrom 1991, 1992, 1994; Bishop, Tuchfarber, and Smith 1994; MacKuen, Erikson, and Stimson 1992). Using aggregate measures of party identification based on Gallup surveys, we find a high degree of persistence in aggregate partisanship. Thus, our results disconfirm the hypothesis that aggregate partisanship as measured by Gallup surveys will exhibit high volatility in response to shocks.

In our analysis, we attempt to bridge the gap between individual- and aggregate-level partisan behavior. In doing so, we have obtained important insights into the degree of persistence in macro-level partisanship. What we now know is that there is a medium-term dynamic of years—not months or decades—within the party system. We also know that this dynamic is potentially caused by one of three factors: behavioral heterogeneity in the electorate, persistence in exogenous political factors that influence individual party identification, or the interaction of heterogeneity and persistence. In future research, we plan to investigate these issues in greater detail by undertaking analyses of individual-level partisan behavior.

APPENDIX

In this appendix, we repeat Granger's (1980) proof for aggregation under the conditions associated with case 3, in which behavior is not homogeneous and the degree of persistence is incomplete. Assuming $\delta = 0$, the individual decision mechanisms is given by:

$$x_{jt} = \alpha_j x_{jt-1} + \varepsilon_{jt}, \tag{A-1}$$

where x_{jt} is the jth individual's current party identification, x_{jt-1} is the last period's partisan affiliation for the jth individual, and ε_{jt} is a mean zero, white noise process. (We use j as the subscript for individuals

[39] The term "mid-range dynamics" is due to MacKuen, Erikson, and Stimson (1989). We remind the reader that our criticisms of MacKuen, Erikson, and Stimson's (1989) work apply only to their conclusions about the *magnitude* of aggregate persistence.

[40] It appears that Nardulli (1995, 10), who finds evidence that critical realignments are regional phenomena that "endure for decades," imposes a unit root in his analyses. This restriction would constrain him to find higher degrees of persistence than those associated with fractionally integrated processes.

[41] It is possible to approximate a fractionally integrated process using an autoregressive model with a *large* number of AR(p) parameters (Granger 1980, 231). MacKuen, Erikson, and Stimson (1989), however, use only an AR(1) transfer function to model macropartisanship. They thus omit a large number of lagged AR(p) terms, which seriously biases their empirical estimates of the effects of shocks to macropartisanship.

because we later use i to denote complex numbers.) According to Granger, when N is large and the α_i are drawn from a beta distribution on the range (0, 1) with p, $q > 0$, then aggregating over the behavior shown in A-1 yields an aggregate time series that is fractionally integrated.

Granger's (1980) proof proceeds as follows. First, consider the power spectrum of the aggregate series:

$$\bar{f}(\omega) = \sum_{j=1}^{N} f_j(\omega),$$

where

$$f_j(\omega) = \frac{1}{|1 - \alpha_j z|^2} \cdot \frac{var(\varepsilon_{jt})}{2\pi}, \text{ and } z = e^{-i\omega}. \quad \textbf{(A-2)}$$

Assuming that the α_j are random variables drawn from $F(\alpha)$ and assuming that the var (ε_{jt}) are random variables that are independent of the α_j, Granger (1980, 231) states that the power spectrum of the aggregate series can be approximated as:

$$\bar{f}(\omega) \simeq \frac{N}{2\pi} E[var(\varepsilon_{j,t})] \cdot \int \frac{1}{|1 - \alpha z|^2} dF(\alpha). \quad \textbf{(A-3)}$$

After some manipulation of the denominator in the integral, Granger (1980, 232) shows that when $F(\alpha)$ is given by the beta distribution, then equation A-3 can be reexpressed so that the coefficient of z^k in the power spectrum is

$$\frac{2}{B(p, q)} \int_0^1 \alpha^{2p+k-1}(1 - \alpha^2)^{q-2} d\alpha. \quad \textbf{(A-4)}$$

Granger (1980, 232) notes that this coefficient equals $\bar{\mu}_k$, the kth autocovariance of the aggregate series, which for large k (and with $q >$ 1) can be approximated by

$$\bar{\mu}_k = A_1 k^{1-q}. \quad \textbf{(A-5)}$$

Granger (1980) notes that the differential equation in A-5 has the same form as the differential equation that defines a fractionally integrated process:

$$\rho_j \simeq A_1 j^{2d-1}. \quad \textbf{(A-6)}$$

Thus, equating coefficients, we see that $d = 1 - q/2$, and the aggregate series is fractionally integrated.

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