SIX

The Dynamics of Party Identification

For nearly half a century, party identification has been a central concept in the study of electoral choice. As discussed in Chapter Two, the concept was originally developed by Angus Campbell and his colleagues at the University of Michigan in their research on American voting behaviour in the 1950s (Campbell, Gurin, and Miller 1954; Campbell et al. 1960, 1966). Imported to Britain by Butler and Stokes (1969), party identification was designated as a key element in public political psychology. According to Butler and Stokes, the process by which the political effects of a deeply divided class society were transmitted to the electoral arena was rooted in class self-identifications that were powerful, pervasive, and primordial. The vast majority of people recognized themselves as members of either the middle or the working class. These class identifications prompted the development of durable identifications with political parties that were thought to represent the economic interests of those classes. Working class people tended to be Labour identifiers who regularly voted Labour, and middle class people tended to be Conservative identifiers who regularly voted Conservative. When voters cast their ballots, they expressed their 'tribal loyalties'. These loyalties were stable products of deeply ingrained, class-based partisan allegiances.

This conception of party identification has prompted protracted debates among students of voting and elections in Britain. A fundamental issue in these controversies is the *stability* of partisan attachments. If party identification lacks stability at the aggregate and individual levels, then theories of voting behaviour and election outcomes premised on the durability of partisan attachments are problematic. Hence, we begin our analysis by investigating long-run aggregate trends in party identification using data from several British Election Study (BES) surveys. Then, these data are employed to address longstanding debates about partisan dealignment. We map trends in the strength of the relationship between social class and party identification, and in the intensity of party identification in the electorate as a whole as well as within various age groups from the time at which they first reached the age of majority.

Next, we use data from several BES and British Election Panel Study (BEPS) panel surveys to gauge the extent to which individual voters change their party identifications over time. Recently developed statistical techniques enable us to determine whether observed levels of instability persist once random measurement error in survey responses is taken into account. We also consider the possibility that much of the apparent individual-level instability in partisanship reflects difficulties with the traditional BES party identification question. For this purpose, we examine data gathered in the 2001 BES, the British Household Panel Surveys (BHPS), the British Social Attitudes Surveys (BSA), and other sources. The BES and BEPS panel data then are used to investigate factors affecting the individual-level dynamics of party identification. A final set of analyses returns to the aggregate level. Using data gathered in monthly Gallup surveys, we model the aggregate dynamics of party identification between January 1992 and December 2002. The chapter concludes by arguing that partisanship in Britain may be usefully conceptualized as a storehouse of party and party leader performance evaluations.

AGGREGATE TRENDS

When the first BES post-election survey was conducted in 1964, slightly over 40 per cent said that they generally thought of themselves as Labour identifiers and an equal number said they were Conservative identifiers[1] (see Figure 6.1). About one in ten was a Liberal identifier and about one in twenty did not identify with a party. Four decades later, the picture was somewhat different. An overwhelming majority of those interviewed in the 2001 BES post-election survey stated that they identified with a party. Although the combined percentage of Conservative and Labour partisans had fallen from 87 per cent to 68 per cent, virtually all of the decrease is attributable to a rapid post-1992 decline in the number of Conservatives. The percentage of Liberal Democrats in 2001 was almost exactly the same as the percentage of Liberals in 1964. Although the size of the group not professing a partisan attachment had doubled, it was only 11 per cent. In sum, as in 1964, a large majority of those interviewed in 2001 claimed to be party identifiers, and most said that they were either Conservatives or Labour.

One might be tempted to argue that these data are compatible with the traditional Michigan theoretical perspective that says that partisan change typically is evolutionary, not revolutionary. According to Campbell et al. (1966), partisan realignments are rare, but when they occur, they involve rapid, substantial, and enduring changes in levels of identification with various parties. Realignments are driven by voters' reactions to big events, such as depressions or world wars, or the greatly enhanced salience of an issue that divides the supporters of one or more of the parties.

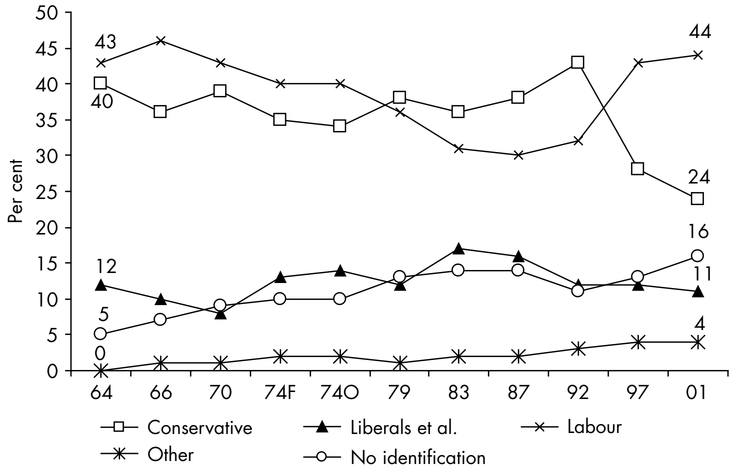

Figure 6.1 Party identification, 1964–2001

Source: BES post-election surveys from 1964 to 2001.

However, in Britain, there is an anomaly. As just observed, the number of Conservative identifiers fell precipitously after the 1992 election, and there was an accompanying large increase in Labour identifiers. To reconcile these rapid, sizeable movements in Conservative and Labour partisan shares with the Michigan perspective, one must argue that a realignment of partisan forces occurred after 1992. The Conservatives' bitter internecine conflicts in the 1990s about the Euro and other aspects of Britain's relationship with the European Union suggest that the realignment hypothesis is plausible if one accepts the Michigan model's proposition that individual-level partisan attachments are generally very stable. As observed in Chapter Two, the premise of individual-level stability in party identifications, except in the face of realigning forces, implies that swift, sizeable shifts in the aggregate balance of such identifications *must* be a product of such forces. However, as demonstrated below, partisan attachments in Britain exhibit much more individual-level movement than the Michigan account allows. This individual-level instability means that the observation of a change in the aggregate balance of party identifiers by itself is not an indicator of realignment. Thus, the concept of realignment is primarily a tool for *post-hoc* description.

Dealignment

Another aspect of the debate about the aggregate-level stability of party identification concerns dealignment. In the British context, this refers to changes in the intensity of partisan attachments and in the strength of the relationship between these attachments and social class. Soon after *Political Change in Britain* appeared, Crewe, Sarlvik, and Alt (1977) announced that BES surveys conducted in the 1970s revealed a weakening of party identification, and an erosion of the correlation between party identification and social class. As noted in Chapter Two, these findings fueled a long-lived controversy. Here, we assess changes in the strength of the relationship between party identification and social class[2] since 1964 by computing the consistency index (CI).[3] The results echo those for the correlation between social class and voting behaviour discussed in Chapter Three. As Figure 6.2 illustrates, the CI declines almost continuously across the entire 1964–2001 period. The magnitude of this long-term trend may be calibrated by regressing the CI on time while controlling for a step-shift downward in the social class × party identification relationship between 1966 and 1970. Using a 0–1 dummy variable[4] to capture this step-shift, we specify the following model:

$$CI_t = f(\beta_0 + \beta_1 \times \text{Time} + \beta_2 \times \text{SHIFT6670})$$

where CI_t, consistency index score at time *t*; Time, time in years; SHIFT6670, permanent step shift in the CI beginning in 1970. The estimated coefficients, with standard errors in parentheses, are:

$$CI_t = \underset{(1.27)}{72.32} \underset{(0.06)}{-0.79} \times \text{Time} \underset{(1.78)}{-11.84}\text{SHIFT6670}$$

Adjusted $R^2 = 0.98$, D.W. $= 2.06$

With an adjusted R^2 of 0.98, the model's explanatory power is very impressive. The linear trend and the 1966–70 step-shift coefficients have the expected negative signs and are statistically significant ($p < 0.001$). The coefficient (-0.79) for the trend term indicates that over the nearly four decades separating the 1964 and 2001 BES, the CI lost about three-quarters of a point per year in its value (on a 200-point scale). Thus, there is strong evidence of a gradual, but ultimately sizeable, erosion of the correlation between social class and party identification. This process has been going on since at least the 1960s and it has continued into the twenty-first century.

This long-term downward trend does not mean that the relationship between social class and party identification was originally strong. It bears emphasis that class defection in party identification has been widespread since at least the 1960s. Over two-fifths of middle class respondents in the 1964 BES did not identify with the Conservatives, and an equal proportion of working class respondents did not identify with Labour. Similar patterns are found in every subsequent BES survey. For example, at the high tide of Conservative electoral success in 1987, over half of the middle class respondents did not identify with that party, and nearly three-fifths of those in the working class did not identify with Labour. In 2001, nearly

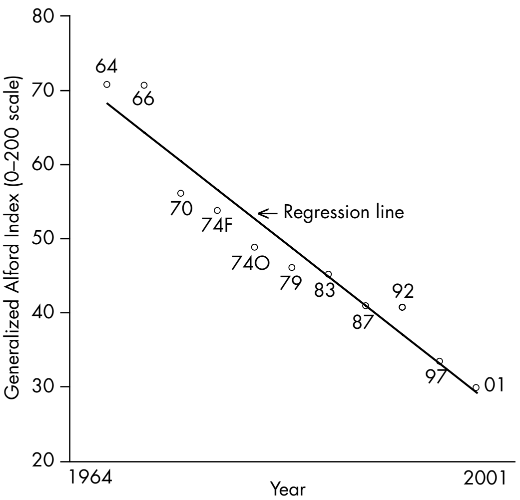

Figure 6.2 Trends in strength of relationship between social class and party identification
Source: BES post-election surveys from 1964 to 2001.

seven of ten middle class respondents did not identify with the Conservatives, and nearly five in ten of those in the working class did not identify with Labour.[5] The prolonged downward trend has loosened a relationship between class and party identification that already had considerable 'play' forty years ago.

A second aspect of the dealignment debate concerns the strength of party identification. In Chapter Three, we present data showing that the percentage of very strong party identifiers has decreased in every successive BES post-election survey. This pattern might be a product of a general weakening of attachments with all parties or, alternatively, a gradual weakening of attachments with some parties. Regarding the latter, one might imagine a scenario in which groups of older identifiers with a given party are not replaced by younger ones with equally strong attachments to it. Although plausible, this scenario is not what has occurred in Britain. To map the dynamics of the strength of party identification, we score 'very strong' identifiers +3, 'fairly strong' identifiers +2, 'not very strong' identifiers +1, and nonidentifiers 0. We then compute the average strength of party identification on this scale by party for each BES survey. Figure 6.3(a) shows that the erosion in the intensity of partisan attachments is generalized—identifications with the Conservatives, Labour, and the Liberals and their predecessors all decreased in strength between 1964 and 2001. And, with the exception of a relatively sharp drop in the intensity of Conservative and Liberal identifications between 1970 and 1974,[6] these long-term decreases are approximately linear.

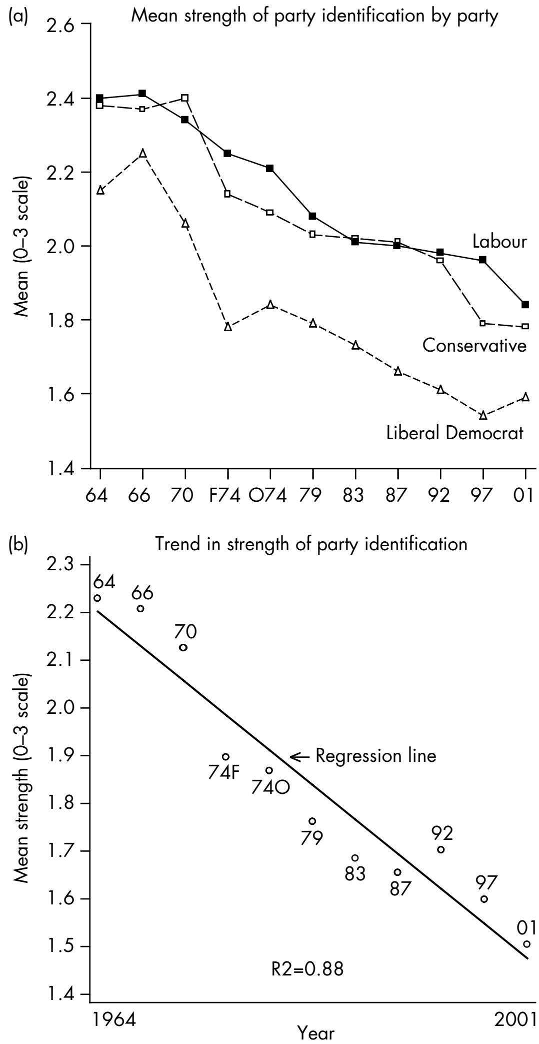

Figure 6.3 Trends in strength of party identification, 1964–2001
Source: BES 1964–2001 post-election surveys.

The overall decrease in the average strength of party identification since 1964 is depicted in Figure 6.3(b). Using a 0–1 dummy variable to allow for a possible downward shift between 1970 and 1974,[7] we specify the following model:

$$\text{SPID}_t = \text{f}(\beta_0 + \beta_1 \times \text{Time} + \beta_2 \times \text{SHIFT7074})$$

where, $SPID_t$ is the average strength of party identification at time t; Time is time in years; SHIFT7074, permanent step-shift in SPID beginning in 1974.

Estimating coefficients with OLS regression (standard errors in parentheses) yields:

$$SPID_t = 3.02 - 0.01 \times \text{Time} - 0.24\text{SHIFT7074}$$
$$\text{(s.e.)}\ (0.11)(0.002) \qquad (0.04)$$
$$\text{Adjusted } R^2 = 0.97, \text{D.W.} = 1.65$$

The linear trend and the 1970–4 step-shift collectively explain virtually all of the variance in the strength of party identification over time. The trend term's coefficient (−0.01) indicates that the average intensity of party identification, as measured on the 0–3 point scale, decreased by about a tenth of a point per year. The magnitude (0.24 points) of the downward step-shift between 1970 and 1974 suggests that its effects were relatively small.

These estimates may be placed in perspective by comparing the strength of partisanship in 1964 and in 2001. In 1964, the average intensity of party identification was 2.2 points, just slightly above the score assigned to people who labelled themselves fairly strong identifiers. In 2001, the average intensity was 1.5 points, midway between fairly strong and not very strong identification. Thus, the protracted decline in the strength of partisan attachments has moved the electorate somewhat less than one-fifth of the way across the entire four-point (0–3) scale. Rather than documenting a wholesale abandonment of party identifications, the 1964–2001 BES data show that the erosion of partisan attachments has been ongoing, but incomplete—a dealignment of degree.

Of Age and Partisan Intensity

Dealignment of degree does not speak strongly against a Michigan-style conceptualization of party identification. However, the several BES surveys do provide evidence that the strength of party identification has not behaved in accordance with what the Michigan theorists would predict. Particularly important in this regard is an assessment of Converse's argument (1969, 1976) that, at the individual level, party identifications normally strengthen the longer they are held. Two processes are said to explain this. First, behavioural reinforcement of a psychological orientation occurs as party identifiers repeatedly vote for 'their' party, and subsequently justify their behaviour to themselves. Second, party identification performs a 'perceptual screening' function (Campbell et al. 1960; see also Bartels 2002). Incoming information about parties' issue positions, performance in office, leaders, and the political world more generally is neither accurate nor unbiased; rather it is selected on the basis of, and filtered through, a 'lens of partisanship'. The result is that voters possess information favourable to their parties and unfavourable to other parties. This combination of positively biased information about the party voters identify with, and negatively biased information about other parties, feed back to strengthen existing party identifications. These reinforcement processes are repeated throughout a voter's political life cycle. The results is a positive correlation between the strength of party identification and age.

Positive relationships between party identification and age are present in the several BES surveys. However, cross-sectional correlations between party identification and age are not conclusive evidence of life-cycle effects such as those proposed by Converse. It is possible that these correlations reflect generational differences, that is, for whatever reasons, voters in different age cohorts vary in the strength of their partisan attachments. Age cohorts may undergo distinctive socialization experiences that affect the strength of their party identifications in the long-run. The picture is further complicated by the possibility of 'period' effects, whereby at a particular point in time the strength of party identification in several age groups is affected either positively or negatively by events. Although life-cycle, generational, and period effects are conceptually distinct, they cannot be separated empirically in a single cross-section; this requires very long-term panel surveys (Glenn 1977). Unfortunately, such long-term panels are unavailable.[8] There are several multi-wave BES panel surveys, but they cover only years—a modest fraction of the political life cycle.

An alternative way of obtaining a long-term perspective on partisan change in various age groups is to construct 'pseudo-panels' (e.g. see Abramson 1976, 1979). This involves dividing respondents in successive cross-sectional surveys into several age groups. The strength of party identification in each of these groups can be tracked in subsequent surveys. For example, if one wished to monitor the strength of party identification for a group aged 18–21 in the 1992 BES, then one would examine a group aged 23–26 in the 1997 survey. These groups comprised different individuals, but they are samples of the larger population of people who, by ageing, have moved from the first of these groups to the second. If the number of cases available for analysis is reasonably large, then this approach provides useful estimates of changes in the strength of party identification in various age cohorts.

We compute the percentages of very strong party identifiers in groups first eligible to vote in successive general elections held between 1970 and 2001.[9] We also do this for three older groups who arguably constitute distinct political generations. These are the Atlee generation (who entered the electorate between 1945 and 1950), the Churchill–Macmillan generation (1951, 1955, or 1959), and the Wilson generation (between 1964 and 1966). These analyses are presented in Table 6.1. Overall, there is little evidence of a tendency for the proportion of very strong party identifiers to grow within age cohorts over time. Over the entire 1970–2001 period, only two cohorts have more very strong party identifiers, and seven have fewer very strong identifiers. This result is not an artefact of choosing 2001 as the end-point for the analysis. Using 1997 as the end-point, two cohorts have at least 1 per cent more

Table 6.1 Percentages of 'Very Strong' party identifiers among groups of voters first entering the electorate between 1945 and 2001

		Election at which 18–21 year-old age groups entered electorate							
		1970	*Feb 1974*	*1979*	*1983*	*1987*	*1992*	*1997*	*2001*
A. Percentages of 'Very Strong' party identifiers among successive 18–21 year-old age groups entering the electorate between 1970 and 2001[a]									
Years after initial entry	0	18.2	18.8	12.4	8.9	6.1	8.0	9.4	8.2
	4		16.9	13.6	12.4	10.7	8.6	7.2	7.5
	9			16.5	15.5	14.2	13.6	9.9	7.8
	13				13.7	11.2	10.2	8.8	9.0
	17					16.2	15.9	14.9	10.6
	22						17.5	13.5	7.0
	27							16.0	10.9
	31								10.4

	Election							
	1970	*Feb 1974*	*1979*	*1983*	*1987*	*1992*	*1997*	*2001*
B. Percentage of 'Very Strong' party identifiers among 'Post-War', 'Macmillan', and 'Wilson' era age groups[b]								
Age group (dates first entering electorate)								
Wilson era (1964–6)	26.1	18.7	15.8	17.2	15.4	17.9	14.5	15.7
Macmillan era (1951–9)	36.2	25.7	20.4	21.7	22.5	23.6	19.8	17.8
Attlee era (1945–50)	39.8	24.1	20.7	24.0	23.8	25.6	29.3	21.7

Note:
[a] Read table *diagonally* to determine percentage of very strong party identifiers among a specific age group at time of subsequent elections. For example, the percentage of very strong party identifiers among the 18–21 year-old age group entering the electorate in 1970 was 18.2. In 1974 the percentage of very strong party identifiers among *that* group was 16.9, in 1979, it was 16.5, etc.
[b] Read table *horizontally* to determine percentage of very strong party identifiers among specific age group at time of subsequent elections. For example, the percentage of very strong party identifiers among the Wilson era group was 26.1 in 1970, 18.7 in 1974, 15.8 in 1979, etc.

Source: BES post-election surveys 1970–2001.

very strong identifiers and five have at least 1 per cent fewer very strong identifiers. The results when 1992 and 1987 are used as end-points are identical. Also, recalling the 1970–4 downward shift in the strength of party identification documented earlier, we note that there has been no general tendency for the percentage of very strong identifiers to increase, regardless of when the analysis *begins*.

We next compute the average strength of party identification on the 0–3 scale described above for the several age groups. Again, there is little evidence to indicate that party identification strengthens over the life cycle (see Table 6.2A and B). Consider the people first eligible to vote in 1970. At that time, the average strength

Table 6.2 Average strength of party identification among groups of voters first entering the electorate between 1945 and 2001

		Election at which 18–21 year-old age groups entered electorate							
		1970	*Feb 1974*	*1979*	*1983*	*1987*	*1992*	*1997*	*2001*
A. Mean strength of party identification on 0–3 scale[b] for successive 18–21 year-old age groups entering the electorate between 1970 and 2001[a]									
Years after initial entry	0	1.62	1.54	1.39	1.31	1.20	1.38	1.45	1.30
	4		1.71	1.58	1.42	1.43	1.51	1.31	1.34
	9			1.66	1.52	1.46	1.58	1.36	1.23
	13				1.50	1.59	1.56	1.46	1.36
	17					1.57	1.58	1.63	1.37
	22						1.65	1.64	1.38
	27							1.65	1.42
	31								1.64
		Election							
		1970	*Feb 1974*	*1979*	*1983*	*1987*	*1992*	*1997*	*2001*
B. Mean strength of party identification on 0–3 scale[b] for Wilson, Macmillan, and Post-War era age groups entering the electorate between 1945 and 1966[c]									
Age group (dates first entering electorate)									
Wilson era (1964–6)		1.80	1.65	1.60	1.64	1.61	1.76	1.60	1.60
Macmillan era (1951–9)		2.00	1.84	1.78	1.74	1.74	1.77	1.69	1.73
Attlee era (1945–50)		2.14	1.89	1.79	1.85	1.84	1.86	1.90	1.78

Note:

[a] Read table *diagonally* to determine mean strength of party identification for a specific age group at time of subsequent elections. For example, the mean strength of party identification for the 18–21 year-old age group entering the electorate in 1970 was 1.62. That group's mean strength of party identification was 1.71 in 1974, 1.66 in 1979, etc.

[b] Strength of party identification measured as: non-identification = 0, weak = 1, fairly strong = 2, very strong = 3.

[c] Read table *horizontally* to determine mean strength of party identification for a specific age group at time of subsequent elections. For example, the mean strength of party identification for the Wilson era group was 1.80 in 1970, 1.65 in 1974, 1.60 in 1979, etc.

Source: 1970–2001 BES post-election surveys.

of party identification for this group was 1.62. Thirty years later, it was essentially unchanged at 1.64. More generally, using 2001 as the end-point for the analysis, there was only one group whose average strength of party identification increased by at least one-tenth of a point, and six groups for which it decreased by at least that much. Using other years as end-points also indicates that there has not been a generalized tendency for party identification within particular age cohorts to strengthen over time. Over the three decades encompassed by the 1970–2001 BES surveys, the intensity of partisanship within age cohorts has varied—sometimes increasing, but more often decreasing. The partisanship-reinforcement mechanisms hypothesized by Converse are not apparent in these data. If they exist, their effects have been overwhelmed by other forces that have decreased the overall intensity of partisanship.

The pseudo panel data suggest what these other forces might be. First, it is evident that there are generational differences—successive cohorts of young people entering the electorate have not been as strongly partisan as once was the case. For example, Table 6.1 shows that 18 per cent of first-time voters in 1970 were very strong identifiers. In 2001, this figure had fallen to 8 per cent. The pattern is the same when one considers the average strength of party identification for first-time voters (Table 6.2). Even if, *ceteris paribus*, there are tendencies for party identification to strengthen over the political life cycle, the fact that successive groups of newly eligible voters have been less partisan than once was the case helps to explain the overall decline in the intensity of party identification. Second, period effects are also evident. For example, Table 6.2 shows that partisanship diminished by at least a tenth of a point in eight of ten age groups between 1997 and 2001. In contrast, between 1987 and 1992, there was an equally clear tendency for partisanship to increase. Data on the percentage of very strong identifiers show the same patterns for these 1997–2001 and 1987–92 comparisons (see Table 6.1). These results indicate that events and conditions associated with particular elections affect the strength of partisanship across multiple age groups.

INDIVIDUAL-LEVEL INSTABILITY

Aggregate-level data presented above do not make a conclusive case for or against the Michigan model of party identification. Data on individual-level dynamics constitute the 'acid test' for understanding the nature of party identification. As noted above, it would be ideal to have very long-term panels that track individual-level partisanship over a span of several decades. Although such data do not exist, several multi-wave panel surveys have been conducted since the 1960s. These panels are particularly useful for statistical analyses when they span four or more points in time. Figure 6.4 summarizes the individual-level dynamics of party identification in six four-wave panels. Contradicting what the Michigan model predicts, there is abundant evidence of substantial individual-level instability in party identification. In the most recent (1998–2001) panel, less than two-thirds of the respondents consistently identified with the same party. An additional, but very small, group (2 per cent) consistently denied that they were party identifiers. The remaining respondents—over one-third of all those interviewed—exhibit partisan instability of various kinds. Twenty per cent switched from one party to another, and an additional 14 per cent moved between being an identifier and being a non-identifier. These numbers are typical of those for other four-wave panel surveys conducted in the 1990s (see Figure 6.4).

Although the 1990s panel data reveal impressive dynamics in individual-level party identification, it might be conjectured that they reflect the dealigning trend discussed above. The proposed story is one in which substantial instability in party identification is a recent phenomenon and, if one examines panel data gathered in earlier decades, then the incidence of partisan instability is much smaller. However, this story is not true. Over one-third of the respondents in the 1974–9 BES panel reported changing their party identifications one or more times. Moreover, although British voters may once have lived in a 'golden age' of Michigan-style party identification, the 1964–70 BES panel data strongly contradict this idea. Figure 6.4 shows that the percentage of unstable identifiers in the 1964–70 panel approaches two-fifths of all those interviewed, with 28 per cent switching identifications one or more times and an additional 10 per cent moving between

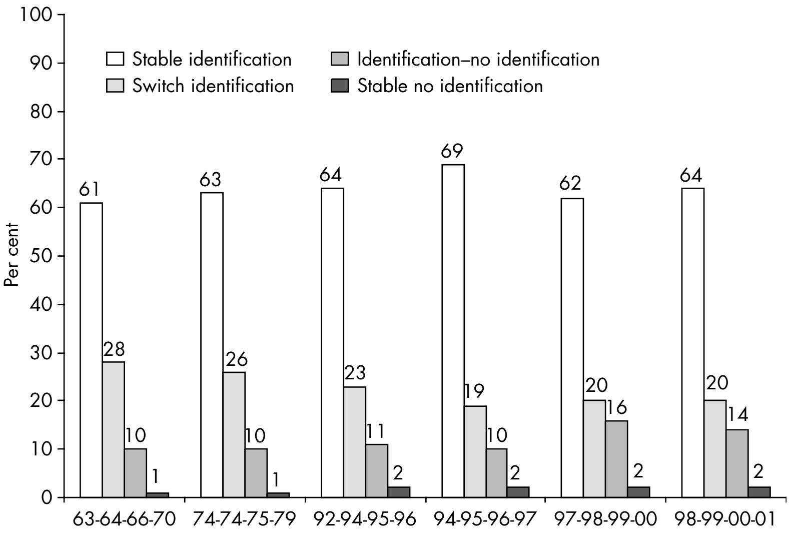

Figure 6.4 Dynamics of party identification in four-wave panels, 1963–2001
Source: BES and BEPS surveys of during 1963–2001.

identification and nonidentification. Indeed, by a small margin, the percentage of unstable partisans in this panel exceeds that for any of the subsequent four-wave panels. Thus, partisan attachments in Britain have demonstrated considerable individual-level dynamism since at least the time that Butler and Stokes imported the concept of party identification from Ann Arbor to Nuffield.

Some analysts have expressed the idea that being a supporter of the Liberals or one of the minor parties is the British equivalent of American voters saying that they are 'independents' (e.g. Clarke and Zuk 1989). If so, then one might conjecture that partisan instability is largely confined to identifiers with the smaller parties who shift their declared partisan attachments in response to a changing mixture of short-term forces operating in the electoral arena at particular points in time. In contrast, those reporting that they are Conservatives or Labour are 'true identifiers' whose partisan attachments manifest impressive durability. Figure 6.5 shows that, in fact, identifiers with the Liberals or other smaller parties consistently report higher levels of instability than Conservative or Labour identifiers. However, rates of instability among the latter two groups are substantial. Across the 1963–70, four-wave panel, about three Conservative identifiers in ten abandoned their party, almost exactly the same proportion as in more recent panels. In the 1992–7 panel, instability in Conservative identification reached fully 44 per cent. Rates of instability in Labour identification are also substantial, ranging from 27 per cent to 31 per cent across various panels. Thus, instability has long been characteristic of major, as well as minor, party identifiers.

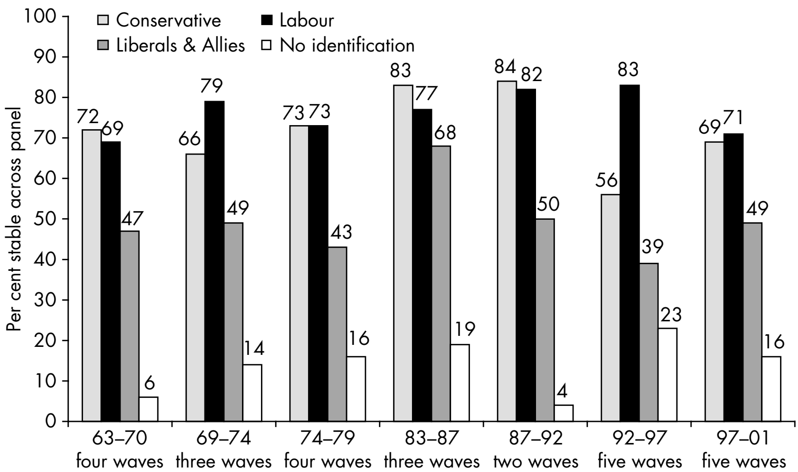

Figure 6.5 Stability of party identification across panel by initial party identification

Note: Switching patterns involving other parties included in calculations, but not shown in figure.

Source: BES and BEPS panel surveys during 1963–2001.

Latent Dynamics

The data in Figures 6.4 and 6.5 testify that party identification in Britain is much more mutable than many analysts have assumed. But, can we believe what these panel data are saying? As noted in Chapter Two, Green and his colleagues have argued that the *observed* instability in panel surveys is misleading because researchers fail to consider the effects of random measurement error (e.g. Green, Palmquist, and Schickler 2002). When random measurement error is taken into account, the dynamics of party identification in panel surveys are reduced to the low levels hypothesized by the Michigan model. Although most of the analyses by Green et al. have used American data, they report the same results using data from other countries including Britain (see Schickler and Green 1997).

Green et al. are not the first to advocate models of survey responses that incorporate random measurement error (Zaller 1992; Alvarez and Brehm 2002). Perhaps best known is Converse's (1964) famous 'black–white' model. Converse's core conjecture is that some people ('blacks') have real and stable orientations towards particular political objects, whereas others ('whites') do not have any real orientations. When asked about these orientations, people in the first group report them accurately, whereas members of the second group answer randomly. Applied to party identification, the 'black–white' model implies that some respondents do not have partisan attachments but, when queried, they invent one. These 'door-step' or 'top-of-the-head' party identifications are chosen at random from the options provided by the interviewer. When interviewed in the subsequent wave of a panel survey, people without real party identifications again answer randomly. Assuming several response categories (e.g. 'Conservative, Labour, Liberal Democrat, or what' in the traditional BES party identification question) are available, there is only a small probability that a respondent without a real party identification will select the same party on multiple occasions.[10] Assuming also that there is a sizeable number of random choosers in the surveys, one will observe considerable instability in party identification across several waves of interviewing. But, most of this observed instability is artefactual. This hypothesis deserves careful scrutiny. If party identification really is stable net of random measurement error, then the panel data reviewed above are misleading, and partisan attachments in Britain resemble what the Michigan model predicts.

We begin our assessment of the random measurement error hypothesis with a simple construct-validation analysis of observed individual-level dynamics of party identification in conjunction with the dynamics of other measures of partisan orientations. If movements in party identification are mostly random, then we would not expect them to correlate consistently and strongly with movements in these other measures. But, consistently strong correlations are exactly what we find. In BES and BEPS panels conducted over the past four decades, individual-level, over-time variations in party identification have strong and predictable relationships with over-time variations in the affective and evaluative components of party images.[11] For example, between 1997 and 2000, the mean change in feelings about the Conservative Party on a 1–5 point scale among people adopting a Conservative identification is positive (0.57) (see Table 6.3). Among those abandoning a Conservative identification, the mean change in such feelings is negative (−0.41). The correlation is strong and highly significant (eta = 0.46, $p < 0.001$). The pattern for comparable groups of Labour identifiers

Table 6.3 Patterns of changing party identification and changing party images in panel surveys

Panel	
A. 1966–70	Mean change in number of likes and dislikes about party
Party identification	
To Conservative	1.16
From Conservative	0.09
$\eta = 0.30, p = 0.000$	
To Labour	−0.08
From Labour	−1.31
$\eta = 0.29, p = 0.000$	
B. 1992–4	Mean change in feelings about party
Party identification	
To Conservative	−0.01
From Conservative	−1.64
$\eta = 0.46, p = 0.000$	
To Labour	1.14
From Labour	−0.64
$\eta = 0.62, p = 0.000$	
C. 1997–8	Mean change in feelings about party
Party identification	
To Conservative	0.57
From Conservative	−0.41
$\eta = 0.46, p = 0.000$	
To Labour	0.32
From Labour	−0.94
$\eta = 0.52, p = 0.000$	
D. 1997–8	Mean change in overall party image score
Party identification	
To Conservative	0.90
From Conservative	−0.39
$\eta = 0.39, p = 0.000$	
To Labour	0.19
From Labour	−0.89
$\eta = 0.34, p = 0.000$	

Source: 1966–70 BES panel survey; 1992–4 and 1997–2001 BEPS panel surveys.

is similar. Once more, the correlation is strong and significant (eta $= 0.52, p < 0.001$). Similar relationships obtain in other panels, including the early 1966–70 one. The patterns are inconsistent with the hypothesis that partisan change is wholly random.

Although these analyses are suggestive, mixed Markov latent class models (MMLC) enable us to control for random measurement error and to analyse directly the stability of party identification at the latent variable level (see van der Pol Langeheine, and Jong 1999; Hagenaars and McCutcheon 2002). Unlike the structural equation models used by Green, Palmquist, and Schickler,[12] MMLC does not require that the analyst impose an ordering on the categories of party identification. Thus, MMLC is useful in polities such as Britain where there are multiple parties, and ordering parties on a single underlying continuum may require problematic assumptions.[13] MMLC is also attractive because it is explicitly designed for estimating the parameters of 'mover-stayer' models in which some, but not all, of the participants in a panel survey change their responses one or more times. Converse's black–white model is a special case of such mover-stayer models where the movers shift randomly across the choice set provided by the interviewer. Here, we employ MMLC to investigate the latent-level dynamics of party identification in Britain using data for the six four-wave panels displayed in Figure 6.4.

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We begin by specifying a general mover–stayer model for party identification in the 1997–2000 BEPS panel. Unlike Converse's black–white model, the general mover–stayer model does not assume that movers (unstable partisans) have an equal probability of shifting across various party identification categories. To ensure sufficient cases for analysis, we consider three party identification categories, Conservative, Labour, and others (Liberals/Liberal Democrats, Nationalists, minor party identifiers, and nonidentifiers). Controlling for random measurement error, this general mover–stayer MMLC model estimates that the proportions of unstable and stable party identifiers (called the 'mixture proportions' in Table 6.4) are 0.34 and 0.66, respectively, in the 1997–2000 panel. Stated otherwise, these numbers indicate that 34 per cent of the panel respondents had a non-zero probability of changing their party identifications, and 66 per cent had a zero probability of doing so, at some time between 1997 and 2000. The analysis shows that no type of party identifier is immune from change. Specifically, 43 per cent of those in the mover chain (group) were initially other party identifiers, 37 per cent were Conservatives, and 20 per cent were Labour.[14]

Although the mover chain in the 1997–2000 panel is sizeable, the transition probabilities for the movers indicate that identifiers with a given party were more likely to stay where they were than to go elsewhere. For example, Conservative identifiers in wave one of the panel had a 0.78 probability of being Conservative but a 0.14 probability of moving to Labour in wave two. Wave-one Labour identifiers had a 0.79n probability of remaining Labour and a 0.21 probability of shifting to the Liberal Democrats or another smaller party. For respondents in the 'other' category, the wave one-wave two probability of staying in that category was slightly less, 0.69. Thus, the MMLC analysis of the 1997–2000 panel data clearly

Table 6.4 Mixed Markov latent class model of the dynamics of party identification, 1997–2000 four-wave panel

		Movers	*Stayers*	
Mixture proportion (π)		0.34	0.66	
Initial state (δ)				
Conservative		0.37	0.28	
Labour		0.20	0.52	
Other		0.43	0.20	
Response probability (ρ)		*True*	*True*	
Conservative		0.78	0.99	
Labour		0.88	0.96	
Other		0.84	0.97	
Transition probabilities (τ) *for movers*				
		Conservative	Labour	Other
Conservative:	1997–8	0.78	0.14	0.08
	1998–9	0.92	0.04	0.04
	1999–2000	1.00	0.00	0.00
Labour:	1997–8	0.00	0.79	0.21
	1998–9	0.08	0.92	0.00
	1999–2000	0.08	0.88	0.04
Other:	1997–8	0.10	0.21	0.69
	1998–9	0.00	0.02	0.98
	1999–2000	0.00	0.00	1.00
Latent turnover table—1997–2000 (row percentages)				
			2000	
		Conservative	Labour	Other
1997	Conservative	73.9	14.6	11.5
	Labour	11.9	63.9	24.2
	Other	12.1	18.2	69.6

Model fit: $\chi^2 = 92.448$, $df = 55$, $p = 0.001$, AIC $= 10886.919$.

Source: 1997–8–9–2000 BEPS panel survey.

indicates that, although the dynamics of party identification in the mover group were substantial, they do not resemble the pattern that would be expected if movements between categories were random, as in Converse's black–white model. For Converse's model with three partisan categories, the transition probability for each would be 0.33. The latent-level turnover table tells the same story.[15] Across wave one to wave four of the panel, 74 per cent of the Conservative identifiers remained Conservative, 64 per cent of Labour identifiers remained Labour, and 70 per cent of those in the 'other' category stayed there.

MMLC analyses of other four-wave panels yield similar results. For example, the estimated proportion of respondents in the mover chain in the 1992–6 panel is quite sizeable, 0.37 (see Table 6.5). However, the transition probabilities are again far from equal levels (0.33) that would obtain if Converse's black–white model held. Also, in keeping with the fact that this panel spans the period when the aggregate share of Conservative identifiers fell sharply (see Figure 6.1), the initial state estimates indicate that a majority of those in the mover chain are Conservatives. And the transition probabilities for the Conservatives across the first two waves of the panel (1992–4) suggest that a Conservative identifier *in the mover group* in 1992 had only a slightly better than 50–50 probability of being Conservative in 1994. Similarly, the latent turnover table across panel waves one to four indicates that only slightly over one-half of Conservative identifiers in the mover group in 1992 were Conservatives

Table 6.5 Mixed Markov latent class model of the dynamics of party identification, 1992–6 four-wave panel

		Movers	*Stayers*	
Mixture proportion (π)		0.37	0.63	
Initial state (δ)				
Conservative		0.55	0.38	
Labour		0.25	0.37	
Other		0.20	0.25	
Response probability (ρ)		*True*	*True*	
Conservative		0.98	0.97	
Labour		0.84	0.99	
Other		0.94	0.94	
Transition probabilities (τ) *for movers*				
		Conservative	Labour	Other
Conservative:	1992–4	0.56	0.15	0.29
	1994–5	0.74	0.05	0.21
	1995–6	0.95	0.04	0.01
Labour:	1992–4	0.00	1.00	0.00
	1994–5	0.00	0.99	0.01
	1995–6	0.02	0.98	0.00
Other:	1992–4	0.00	0.67	0.33
	1994–5	0.19	0.22	0.59
	1995–6	0.25	0.00	0.75
Latent turnover table—1992–6 (*row percentages*)				
			1996	
		Conservative	Labour	Other
1992	Conservative	52.2	25.5	22.3
	Labour	1.7	97.8	0.5
	Other	11.9	73.3	14.8

Model fit: $\chi^2 = 68.015$, $df = 55$, $p = 0.128$, AIC $= 5547.173$.

Source: 1992–4–5–6 BEPS panel survey.

four years later. Even greater change is evident for the 'other' category. Among those in this category who were in the mover chain in 1992, only slightly over one in ten was in the 'other' category in 1996. In sharp contrast, and consonant with the surge in Labour partisanship in the mid-1990s, virtually all (98 per cent) of the 1992 Labour identifiers in the mover chain were Labour identifiers in 1996.

The MMLC analyses show substantial latent-level dynamics in party identification in recent years. This is inconsistent with the high levels of partisan stability as argued by Greene et al. and other proponents of Michigan-style partisanship. But, what about earlier time periods? Particularly interesting in this regard are the results of an MMLC analysis of the 1963–4–6–70 panel data. This analysis decisively rejects the conjecture that the 1960s were an era of Michigan-style partisan stability in Britain. As Table 6.6 shows, the proportion of panelists with a non-zero probability of changing party identification (the mover chain) was 0.31. There are substantial numbers of Conservative, Labour, and other types of partisans in the mover group. Once more, the transition probabilities for this 1963–70 panel are inconsistent with the idea that 'black–white' dynamics describe movements across adjacent panel waves. Between 1963 and 1964, the probabilities that Conservatives and Labour identifiers would abandon their party were nontrivial. Perhaps even more impressive are the latent level wave one to wave four mobility figures. These show that slightly over half of the Conservatives in the mover group changed their party identification between 1963 and 1970. Similarly, almost two-thirds of the Labour identifiers in the mover group changed their identification over the seven years of the panel.

MMLC analyses of other BES and BEPS four-wave panels tell similar stories. Figure 6.6 summarizes data on the estimated sizes of the mover and stayer chains in

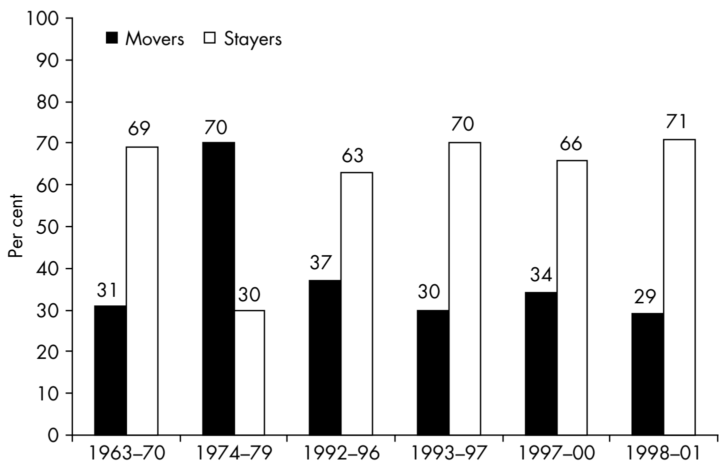

Figure 6.6 Sizes of mover and stayer groups estimated by mixed Markov latent class models
Source: 1963–2001 BES and BEPS panel surveys.

Table 6.6 Mixed Markov latent class model of the dynamics of party identification, 1963–70 four-wave panel

		Movers	*Stayers*	
Mixture proportion (π)		0.31	0.69	
Initial state (δ)				
Conservative		0.22	0.45	
Labour		0.36	0.38	
Other		0.43	0.17	
Response probability (ρ)		*True*	*True*	
Conservative		0.97	0.94	
Labour		0.94	0.99	
Other		0.88	0.25	
Transition probabilities (τ) *for movers*				
		Conservative	Labour	Other
Conservative:	1963–4	0.55	0.23	0.23
	1964–6	0.55	0.41	0.05
	1966–70	0.72	0.17	0.10
Labour:	1963–4	0.29	0.63	0.08
	1964–6	0.11	0.72	0.16
	1966–70	0.47	0.53	0.00
Other:	1963–4	0.00	0.00	1.00
	1964–6	0.03	0.02	0.96
	1966–70	0.22	0.07	0.71
Latent turnover table—1963–70 (*horizontal percentages*)				
			1970	
		Conservative	Labour	Other
1963	Conservative	48.4	28.3	23.3
	Labour	48.2	35.7	16.1
	Other	23.4	08.4	68.2

Model fit: $\chi^2 = 82.618$, $df = 52$, $p = 0.004$, AIC = 4725.681.

Source: 1963–4–6–70 BES panel survey.

all of the four-wave panels. As the figure shows, with one exception, the mover group varies between about three in ten (1998–2001 panel) to about four in ten (1992–6 panel). The exception is the 1974–9 panel where the size of the mover group reaches fully seven in ten. This is the era of dealignment as studied by Crewe, Sarlvik, and Alt (1977) (see also Sarlvik and Crewe 1983). Thus, the general conclusion of the several MMLC analyses is straightforward. Controlling for random measurement error, party identification in Britain has always exhibited impressive individual-level dynamics. There have been more stayers than movers, but the latter has consistently exceeded what would be expected if partisan attachments were stable.

There is more. The general mover–stayer models can be compared with rival models of the (non)dynamics of partisanship. One rival is Converse's black–white model which, as noted, imposes equal probabilities on the transition matrices for the mover group. Another rival is a core version of the Michigan model, which does away altogether with the mover chain and specifies that everyone is a stayer. Table 6.7 summarizes the results of estimating the alternative models for the six four-wave panels. As indicated by its smaller chi-square values, the general mover–stayer model

Table 6.7 Rival mixed Markov latent class models of the dynamics of party identification

Panel A.
Goodness of fit

		Chi-square	*df*	*p*
1963–70:	Mover–stayer: general	82.62	52	0.004
	Mover–stayer: black–white	223.02	66	0.000
	All stayer	158.79	55	0.000
1974–9:	Mover–stayer: general	48.56	52	0.610
	Mover–stayer: black–white	100.16	66	0.004
	All stayer	89.37	55	0.002
1992–6:	Mover–stayer: general	67.02	55	0.128
	Mover–stayer: black–white	295.97	66	0.000
	All stayer	190.48	55	0.000
1994–7:	Mover–stayer: general	36.15	55	0.977
	Mover–stayer: black–white	137.76	66	0.000
	All stayer	124.08	54	0.000
1997–2000	Mover–stayer: general	92.45	55	0.001
	Mover–stayer: black–white	306.67	66	0.000
	All stayer	205.75	54	0.000
1998–2001:	Mover–stayer: general	68.75	54	0.085
	Mover–stayer: black–White	256.21	67	0.000
	All stayer	234.50	54	0.000

B. Akaike information criterion (smaller is better)

		Mover–stayer Models	
Panel	General	Black–white	All stayer
1963–70	4725.68*	4824.08	4783.86
1974–9	3367.23*	3376.83	3390.04
1992–6	5547.17*	5734.12	5652.64
1994–7	4935.62*	4995.23	5005.54
1997–2000	10886.92*	11059.14	10982.22
1998–2001	8274.93*	8420.39	8422.68

* Smallest value.

Source: BES and BEPS four-wave panel surveys during 1963–2001.

consistently has a better fit with the data than its rivals. The Akaike Information Criterion (AIC), which discounts model fit in terms of the number of parameters estimated (Burnham and Anderson 1998; see also ch. 4), also suggests the superiority of the general mover–stayer model. In all cases, the AIC for the general mover–stayer model is smaller than those for its rivals. In sum, analyses of the several BES and BEPS panel surveys strongly indicate that party identification in Britain has dynamic properties. These dynamics are not novel—they have existed since at least the first BES panels were conducted in the 1960s.

ASKING THE WRONG QUESTION?

The evidence presented above testifies that party identification in Britain is not the unmoved mover of Michigan lore. Impressive individual-level instability in responses to the standard BES party identification questions is apparent both in simple turnover tables, as well as in more sophisticated analyses that take into account random measurement error in survey responses. However, before accepting the conclusion that there are substantial individual-level dynamics in party identification, a critic might protest that the problem with the preceding analyses is that they are based on responses to the wrong question (e.g. see Bartle 2001; Blais et al. 2002). The critique is similar to the argument motivating Converse's black–white model. The hypothesis is that the standard BES question—'Generally speaking, do you think of yourself as Conservative, Labour, Liberal Democrat, or what?'—cues respondents to accept a party label and, thus, it inflates estimates of the proportion of party identifiers in the electorates and the proportion who switch their identifications over time. More specifically, some respondents are not true identifiers, but when prompted to select a party, they oblige the interviewer and do so. When asked in a subsequent interview about their party identification, these people again provide one, but because they are not true identifiers, there is a substantial probability that they choose a different party than the first time. Wording a party identification question in a way that does not cue respondents to choose a party would reduce the number of identifiers *and* reveal that partisan stability is much more prevalent than responses to the traditional question lead one to believe.

AQ: We have changed as Blais et al as per the ref list. Pls check.

Data gathered in successive waves of the British Household Panel Survey (BHPS) enable us to test the wrong-question hypothesis. In the BHPS, respondents are asked if they think of themselves as 'party supporters'.[17] Answers to this question sequence show that the proportion of self-acknowledged party supporters at any point in time is much lower than the proportion of party identifiers as measured by the BES question. However, analyses of responses to the BHPS question do not suggest an absence of dynamics. Although the proportion of stable non-supporters (approximately one-third of those participating in three-year rolling BHPS panels conducted between 1991 and 1997) is much larger than the proportion of nonidentifiers in the BES and BEPS panels (see Figure 6.4), overall levels of instability are very substantial among BHPS panelists. The percentages changing their answers to the party supporter question range from 46 per cent in the 1995–7 panel to 52 per cent in the 1991–3 panel (Sanders, Burton, and Kneeshaw 2002).

The British Social Attitudes (BSA) surveys provide another opportunity to study the impact of using alternative question wording to measure partisan attachments.[18] The BSA question sequence is: 'Some people think of themselves as usually being a supporter of one political party rather than another. Do you usually think of yourself as being a supporter of one particular party or not? [If yes] Which party is that'? The BSA surveys consistently reveal that large numbers of people do not think of themselves as party supporters. In 15 such surveys conducted between 1983 and 1999, the percentage stating that they are not party supporters ranges from a low of 49 per cent (1989, 1990) to a high of 61 per cent (1999), and it averages 54 per cent. Also similar to the BHPS, there is considerable individual-level instability in answers to the party supporter question, with slightly over 40 per cent of those in a 1983–6 BSA panel changing their responses one or more times. Rather than shifting directly between parties, nearly nine in ten of these people moved from being a party supporter to being a nonsupporter or vice versa.

To provide a direct comparison of alternative measures of partisan attachments, we asked the BSA party supporter question in the 2001 BES. To prevent mutual contamination of the two question sequences, they were widely separated in the survey instrument and their ordering was randomized. Since both sequences were asked in the pre- and the post-election surveys, it is possible to investigate differences in the dynamics of responses to them.

Answers to the 2001 BES party supporter questions are similar to those in the BHPS and BSA surveys. Table 6.8 shows that, in both the pre- and post-election 2001 BES surveys, the percentage that stated that they were party supporters is much lower than the percentage selecting a party in response to the traditional party identification question. In the pre-election survey, approximately 50 per cent said they were not party supporters, but only 21 per cent did not accept a party when answering the traditional question. In the post-election survey, the comparable figures are 41 per cent and 16 per cent. However, as in the BHPS and BSA studies, there is considerable over-time instability in answers to the party supporter questions in the 2001 BES. Over the brief period between the pre- and post-election waves of interviewing, slightly over 26 per cent of the panelists changed their answers to the party supporter question, with an overwhelming majority of these people moving between being a supporter and being a nonsupporter. The total instability in responses to the traditional question is almost identical and, again, a large majority of switchers indicated that they had moved between identification and nonidentification.

Thus, alternative question wordings do not produce levels of partisan stability commensurate with those predicted by the Michigan model. Data gathered in a

Table 6.8 Responses to traditional party identification question and alternative party supporter question

	Pre-election survey		*Post-election survey*	
	Party identification	*Party supporter*	*Party identification*	*Party supporter*
A. Responses in pre- and post-election cross-sectional surveys				
Labour	41.6%	28.4%	44.4%	33.3%
Conservative	24.9	16.9	24.4	18.8
Liberal Democrat	8.9	3.1	11.0	5.3
Scottish Nationalist	1.6	1.0	1.6	1.2
Plaid Cymru	0.3	0.2	0.4	0.2
Greens	0.6	0.2	0.8	0.3
Other party	1.1	0.4	1.4	0.2
None, D.K./not supporter	21.0	49.8	16.0	40.7
(N)	(3163)	(3191)	(2989)	(2994)
B. Patterns of partisanship in pre-post-election panel survey				
Stable partisans				
Labour	37.1%	24.7%		
Conservative	20.1	13.8		
Liberal Democrat	6.4	2.3		
Scottish Nationalist	1.2	0.6		
Plaid Cymru	0.2	0.1		
Greens	0.4	0.2		
Other Party	0.1	0.1		
Total stable partisans	65.5	41.8		
Stable non-identifier/ non-supporter	8.9	32.0		
Total stable	74.4		73.8	
To<–>From identification/supporter		16.7		24.0
Switch party identification/supported		9.9		2.2
(N)		(2266)		(2282)

Source: BES pre- and post-election cross-sectional and panel surveys in 2001.

variety of panel studies conducted over the past two decades indicate that substantial minorities switch back and forth between claiming to be a party supporter and claiming not to be one. As predicted by critics of the traditional party identification question, the BHPS, BSA, and BES surveys do reveal that levels of partisanship in Britain are lower than responses to the traditional question indicate. This finding also characterizes data gathered in several surveys conducted in 1998 and 1999 (Sanders, Burton, and Kneeshaw 2002). But, given that responses to the party supporter question also indicate that partisan attachments have considerable mutability, the lower incidence of party supporters is hardly good news for the Michigan model. One cannot use data generated by party supporter-type questions to portray partisan attachments as widespread, stable elements in the psychology of the British electorate. Rather, the overall impression conveyed by answers to party supporter questions is one of stable partisans, being the exception, not the rule. In the next section, we develop models of that forces that affect the dynamics of partisanship over time.

MODELLING DYNAMICS

The Individual Level

It is not novel to argue that party identification has dynamic properties. As discussed in Chapter Two, over the past two decades, a number of analysts have rejected the idea that party identification does not change, and they have developed models of its individual-level dynamics. Although differing in detail, these models propose that voters are at least rough-and-ready utility maximizers who update their existing party identifications using information about the actual or anticipated performance of political parties. *Pace* Bartels (2002), these models do not require that the information used to update party identification be unbiased in the sense that it is totally free from the influence of existing party identification. It may be sensible for voters to use their existing party identification as a heuristic device to help them simplify their search for and interpretation of new information. Deliberation costs are thereby reduced (Conlisk 1996). This new information is not wholly redundant to existing party identifications; it is used to update these identifications. Since the updating process is ongoing and may reinforce or erode an existing partisan attachment, party identification becomes, in Fiorina's (1981) widely used phrase, a running tally of current and past party performance evaluations. The process may be represented as:

$$\mathrm{PID}_{it} = \mathrm{f}(\beta_0 + \lambda_1 \mathrm{PID}_{it-1} + \Sigma\beta_k \mathrm{X}_{kt})$$

where, PID_{it} is the party identification for voter i at time t; X_{kt}, party performance evaluations weighted by parameters β_k; λ, the effect of time $t-1$ party identification on time t party identification; β_0 is a constant.

In this model, a voter i begins with an initial party identification at time $t = 0$. Some analysts (e.g. Achen 2002) have interpreted this PID_{i0} as reflecting the effects of parental and other early-life socialization experiences. The λ parameter measures the rate at which voters discount information encapsulated in their existing party identifications. If λ equals 1.0, then existing information is never discounted, and all current and past party performance evaluations, as well as various random shocks, simply accumulate at their full values through time. If one hypothesizes that voters rely less on older information, then it follows that the λ parameter is less than 1.0. The β_k's capture the effects of current party performance evaluations.

β_0 may be interpreted as a baseline indicator of partisanship in the electorate at time *t*. This would occur if voters had perfectly balanced evaluations of current and prior party performance.[19]

The BES and BEPS panels contain the data needed to estimate the parameters in this model. Here, we focus on the recent 1997–8 and 1992–4 panels and the earlier 1966–70 panel. We measure party identification in two ways. In one set of analyses, we construct seven-point measures of the direction and strength of Labour, Conservative, and Liberal Democrat party identification. For example, the Conservative measure ranges from very strong Conservative identifier (scored $+3$) to nonidentifier (scored 0) to very strong other party identifier (scored -3). In a second set of analyses, we focus strictly on the direction of identification with these parties. In this set, we use three-point measures, ranging, for example, from Conservative (scored $+1$), to nonidentifier (scored 0), to other party identifier (scored -1). Measures of party performance evaluations vary depending on the set of questions asked in particular surveys. These questions concern evaluations of national and/or personal economic conditions, evaluations of government performance in various policy areas, and judgements of which party leader would make the best prime minister.[20] Variables measuring social class and age cohort are also included in the model to control for the possibility that any observed effects of time $t - 1$ party identification and time t party performance evaluations are, in fact, spurious—that is, products of voters' class locations or cohort-specific socialization experiences. Model parameters are estimated using an instrumental variables regression procedure.[21]

The results are presented in Table 6.9. Parameter estimates accord well with expectations. In every analysis, time $t - 1$ (prior) party identification has significant, positive effects on time t (current) party identification and all coefficients are less than 1.0 in magnitude. This indicates that previous information stored in an existing party identification is progressively discounted as new information becomes available. Variables measuring the effects of new information also operate as anticipated. For example, in the 1997–8 panel, voters who judged economic conditions favourably, and evaluated government performance positively, were more likely to identify with the governing Labour Party than voters who judged economic conditions unfavourably and evaluated government performance negatively. Net of these effects, voters who thought that Tony Blair would make the best prime minister were more likely to identify with Labour, and those who favoured William Hague were less likely to do so. The results for the 1997–8 analyses of Conservative identification also make sense. As expected with a Labour government in power, favourable economic and government performance evaluations lowered the likelihood of being a Conservative identifier. Similarly, voters who thought Blair would make the best prime minister were less likely to be Conservative identifiers, and those who judged that Hague would make the best prime minister were more likely to be Conservatives.

Table 6.9 Models of the individual-level dynamics of party identification

	Conservative		*Labour*		*Liberal Democrat*	
Predictor variables	*7pt*	*3pt*	*7pt*	*3pt*	*7pt*	*3pt*
A. 1997–8 BEPS panel						
			1998 Party identification			
Party identification ($t-1$)	0.77***	0.81***	0.75***	0.81***	0.44*	0.42*
Government performance(t)	−0.02***	−0.01**	0.03***	0.02***	−0.02*	−0.01*
Economic evaluations(t)	−0.02*	−0.01	0.03***	0.01*	−0.01	−0.00
Best Prime Minister:						
Blair(t)	−0.12***	−0.06***	0.17***	0.08***	−0.10***	−0.05***
Hague(t)	0.10***	0.06***	−0.04*	−0.02*	−0.04*	−0.03**
Ashdown(t)	−0.05***	−0.03***	−0.02	−0.02*	0.16**	0.11***
Social class	0.05***	0.03***	−0.04**	−0.02*	0.01	0.01
Age cohort:						
Blair	−0.16*	−0.09	0.22*	0.15*	0.32*	0.09
Thatcher/Major	−0.05	−0.02	0.16*	0.09*	0.14	0.01
Wilson/Callaghan	−0.00	0.02	0.13*	0.07*	0.13	0.03
Macmillan et al.	0.05	0.03	−0.07	−0.00	0.11	0.01
Constant	0.58***	0.14	−1.33***	−0.55***	−0.46	−0.37*
Adjusted R^2 = ($N = 2677$)	0.67	0.62	0.63	0.56	0.44	0.39
B. 1992–94 BEPS panel						
			1994 Party identification			
Predictor variables						
Party identification ($t-1$)	0.63***	0.64***	0.76***	0.83***	0.52***	0.50***
Government performance(t)	0.03***	0.02***	−0.02**	−0.01***	0.02**	0.01
Economic evaluations(t)	0.05***	0.02***	−0.04***	−0.01**	−0.00	−0.00
Best Prime Minister:						
Smith/Blair(t)	−0.12***	−0.07***	0.12***	0.06***	−0.14***	−0.05***
Major(t)	0.13***	0.09***	−0.02	−0.02	−0.13***	−0.07***
Ashdown(t)	−0.10***	−0.07***	−0.04*	−0.02*	0.18***	0.12***
Social class	0.04**	0.03**	−0.05*	−0.02*	0.06***	0.02*
Age cohort:						
Thatcher/Major	−0.06	−0.05	0.16*	0.10*	0.10	−0.03
Wilson/Callaghan	0.00	−0.02	0.03	0.00	0.05	−0.05
Macmillan et al.	0.13*	0.04	−0.06	0.01	−0.00	−0.04
Constant	−1.36***	−0.67	0.93***	0.50***	−0.83***	−0.42***
Adjusted R^2 ($N = 2159$)	0.66	0.60	0.59	0.54	0.39	0.33

Table 6.9 Continued

Predictor variables	Conservative 7pt	Conservative 3pt	Democrat 7pt	Democrat 3pt	Liberal Labour 7pt	Liberal Labour 3pt
A. 1966–70 BES Panel						
			1970 Party identification			
Party identification(t − 1)	0.48***	0.47***	0.43***	0.37***	0.54***	0.49***
Party closest on issues(t)	0.42***	0.16***	−0.42***	−0.18***	0.01	0.01
Gov. Ec. performance(t)	−0.23***	−0.10***	0.24***	0.09***	0.04	0.02
Best Prime Minister						
Wilson(t)	−0.15***	−0.05***	0.16***	0.05***	−0.03	−0.01
Heath(t)	0.13***	0.04**	−0.14***	−0.05***	0.00	−0.00
Social class	0.09**	0.03*	−0.13***	−0.06***	−0.00	0.01
Age cohort						
Wilson	0.07	−0.04	−0.07	0.03	0.38**	0.07
Macmillan et al.	0.09	0.04	−0.00	0.01	0.29**	0.06
Post-war	−0.03	−0.02	−0.07	−0.01	0.27**	0.04
Constant	0.41	0.16	−0.48	−0.16	−1.05***	−0.46***
Adjusted R^2	0.70	0.61	0.71	0.65	0.33	0.26
(N = 1066)						

Note: All models estimated with instrumental variables for party identification at $t - 1$, except for Liberal party identification 1966–70 which is estimated using OLS.
*** $p \leq 0.001$; ** $p \leq 0.01$; * $p \leq 0.05$, one-tailed test.

Source: BES two-save panel survey, and 1992–4, 1997–8 BEPS two-wave panel surveys during 1966–70.

Similar results obtain for the 1966–70 panel. In this case, current party identification again reflects evaluations of the performance of parties and their leaders net of prior party identification and the effects of social class and age cohort. Nor is the case that the models using the 1966–70 panel data are pale reflections of those for the 1990s. The 1966–70 models explain large percentages of the variance in 1970 Conservative and Labour party identifications, and all of the evaluation variables are correctly signed and statistically significant. The 1966–70 models do less well in accounting for the dynamics of Liberal identification, but this is also true for the 1990s panels. In this regard, the percentage of variance explained for the Liberal Democrat model is largest for the 1997–8 panel, which is the only case when a Liberal Democrat party leader evaluation variable is available.

These analyses of the individual-level dynamics of party identification contradict the spirit of the original Michigan model. Rare party realignments aside, the canonical *American Voter* version of the model says that party identification at time $t - 1$ should drive party identification at time t. Current (time t) evaluations of parties and their leaders are strongly influenced by $t - 1$ party identifications and, when the latter is controlled, party and leader evaluations have little, if any, effect on time t party identification. The results also contradict the spirit of the Butler–Stokes version of the Michigan model. In the core Butler and Stokes model, all variables, except social class, should fail to affect current (time t) party identification. Even prior (time $t - 1$) party identification should fail. The statistical relationship between party identification at times t and $t - 1$ derives from the fact that partisan attachments are stable products of enduring class identities. If these identities are controlled, then the statistical relationship between party identification at different points in time should disappear. But, analyses of panel data gathered since the 1960s strongly gainsay these predictions. Neither version of the Michigan model can account for the individual-level dynamics of party identification in Britain.

The Aggregate Level

Although individual-level analyses are highly information, aggregate-level data are especially useful to assessing the dynamics of party identification. To generate the time series data required for such analyses, we have included the standard BES party identification question sequence in monthly Gallup surveys conducted since January 1992. By using the BES question battery, we are able to avoid the question-wording controversy that has bedevilled research on the aggregate dynamics of party identification in the United States (see, e.g. Abramson and Ostrom 1992, 1994; MacKuen, Erikson, and Stimson 1992a). The resulting time series enables us to track short-term movements in party identification over an 11-year interval, and to investigate factors that influence those movements.

The Gallup time series data on identifications with the Conservative, Labour, and Liberal Democrat parties for the January, 1992–December, 2002 period are displayed in Figure 6.7. As the figure shows, the percentage of Conservative identifiers fell below the Labour percentage shortly before the September 1992 ERM crisis. Subsequently, the Conservatives always had fewer identifiers than Labour. Over the entire 11-year period, the Conservative share averaged 27 per cent, as compared to 41 per cent for Labour, and 12 per cent for the Liberal Democrats. It is also noteworthy that identification with all three parties varied substantially over time. The Conservative percentage moved by 16 points, ranging from a high of 37 per cent in February, 1992 to a low of 21 per cent in May 1996. As for Labour, its share varied by fully 24 points, with the cohort of Labour identifiers ranging from slightly below 30 per cent in February 1992 to nearly 54 per cent in October 1997. After that high point, Labour partisanship receded, albeit irregularly, to 36 per cent in December 2002. Liberal Democrat identification also varied considerably—climbing from a low of 7 per cent in January 1997 to a high of 19 per cent in September 1993. In December 2002, the percentage of Liberal Democrats (12 per cent) was exactly what it had been 11 years earlier.

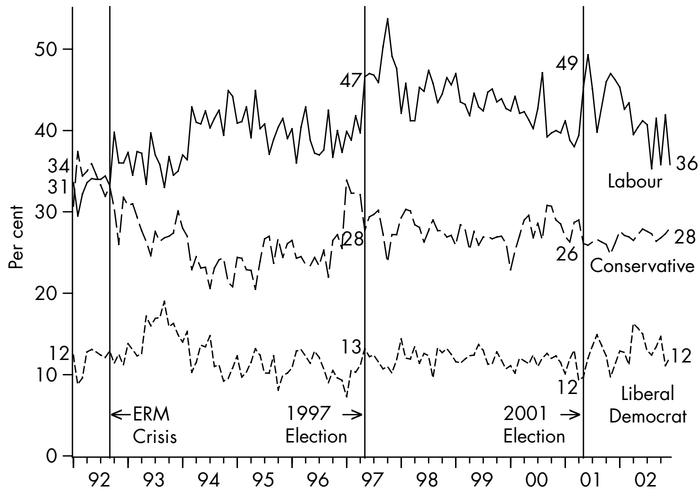

Figure 6.7 Dynamics of party identification, January 1992–December 2002
Source: PSCB and P&D surveys during 1992-2002.

To see trends in the Gallup data more clearly, we employ the nearest neighbour regression technique,[22] regressing Labour, Conservative, and Liberal Democrat party identification on time. The results show that Labour enjoyed a prolonged surge in party identification from 1992 onward, peaking shortly after its resounding general election victory in 1997 (see Figure 6.8). Afterward, although there was substantial variation in the monthly percentage of Labour identifiers, the overall pattern shows a slow downward trend. Conservative identification followed a very different course—falling sharply between 1992 and 1995, then partially recovering in the run-up to the 1997 election. Since shortly after that election, the underlying pattern has been quite stable, with the Conservative percentage of party identifiers varying between 25 and 30 per cent. The Liberal Democrat pattern was different again; although there were upward and downward swings, the overall trend line has been quite flat.

Can these aggregate-level movements in party identification be explained by changes in evaluations of the performance of political parties and their leaders? The Gallup surveys are useful for answering this question because they inquire about voters' national and personal economic evaluations, as well as their opinions about which party leader would make the best prime minister. These data enable us to specify models of the aggregate dynamics of party identification. However, before doing so, it is necessary to consider whether the data are characterized by trends in the technical sense of that term. In the language of time series analysis, trending variables are said to be *nonstationary*, that is, they lack constant means and/or variances.

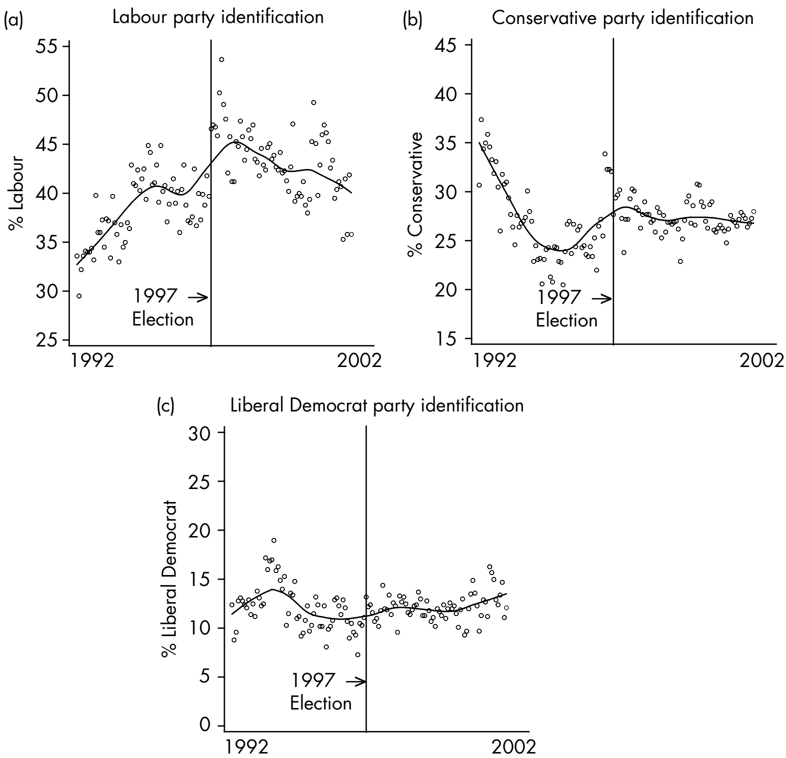

Figure 6.8 Trends in party identification, January 1992–December 2002

Note: Trend lines are generated using nearest neighbour regression, with polynomial degree 1 and bandwidth 0.3

Source: January 1992–December 2002 PSCB and P&D surveys.

As Granger and Newbold (1974) demonstrated, analysing nonstationary variables[23] frequently produces 'spurious regressions' that pose serious threats to inference.

Some analysts (e.g. Box-Steffensmeier and Smith 1996) contend that nonstationary processes are unlikely to characterize time series data on variables, such as party identification,[24] because they imply that support for particular parties could increase or decrease without limit. In substantive terms, this would mean that party systems lack mooring—no forces would e at work to restore party competition (Stokes and Iversen 1966; see also Bartolini and Mair 1990). It is more likely that the aggregate-level evolution of party support is driven by 'long memory' or 'fractionally integrated' processes (e.g. Beran 1992). If so, then levels of support for particular parties may drift upward or downward for lengthy periods of time, but will eventually move back

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toward a very long-run central tendency. Thus, a party's cohort of identifiers may increase for several years but not forever. Long-memory processes in time series data on party identification also are interesting because, as Granger (1980) has shown (see also Box-Steffensmeier and Smith 1997; Wlezien 2000), long-memory may result from aggregating heterogeneous individual-level data. In this regard, analyses presented above consistently indicate the presence of mover and stayer groups of party identifiers in the individual-level BES and BEPS panel data.

The simplest model for a long-memory process is: $(1-L)^d Y_t = \epsilon_t$ where ϵ_t is a random shock at time t, L is a backshift operator, and d is a fractional differencing parameter (Franses 1998). If d equals 1.0, then the process is a random walk and, hence, nonstationary. Long-memory processes remain nonstationary, although they are ultimately 'mean reverting' if d is greater than or equal to 0.5, but less than 1.0. As noted above, nonstationary processes invite so-called 'spurious regressions' and, hence, it is important to determine whether time series variables are nonstationary. One way of doing this is to estimate the value of their d parameters.[25] Here, we do so for the party identification series, as well as the variables measuring the monthly balance of different kinds of positive versus negative economic evaluations and monthly judgements of who would make the best prime minister.[26]

Table 6.10 shows that the d values for all of these series are greater than 0.5, thereby indicating that they are nonstationary. However, in most cases, d is significantly less than 1.0 ($p < 0.05$). As per the discussion above, such data are said to possess long memory. Substantively, the values of d for the party identification series confirm the visual impressions of trends in these series in Figure 6.8. Between 1992 and 2002, the party identification variables moved, albeit at different rates, in one direction for lengthy periods, but they did not do so indefinitely. As noted above, the presence of long memory in the party identification time series data also is consistent with the data being generated by mover–stayer processes at the individual level.

Given that the time series variables of interest have long memory and are nonstationary, we analyse the determinants of party identification by specifying fractional error correction (FEC) models (Clarke and Lebo 2003). These models enable us to study both the short- and long-term effects of various independent variables on party identification. A general FEC model for the dynamics of party identification is:

$$\Delta^d \mathrm{PIDt} = f(\beta_0 + \beta_1 \Delta^d \mathrm{BESTPM}_t + \beta_2 \Delta^d \mathrm{EVAL}_t - \alpha(\mathrm{PID}_{t-1} - \gamma_1 \mathrm{BESTPM}_{t-1} - \gamma_2 \mathrm{EVAL}_{t-1}) + \Sigma\beta_{3-k}\mathrm{EVENTS}_{t-i})$$

where: PID_t = party identification t;
BESTPM_t = best prime minister at time t;
EVAL_t = economic evaluations (personal prospective, personal retrospective, national prospective, national retrospective) at time t;
EVENTS = various important events;
Δd is the fractional differencing operator;[27]
d, β, α, and γ are parameter to be estimated.

Table 6.10 Estimates of the fractional differencing parameter, *d*

Time series variables	*d*	*s.e.*
Party identification		
Conservative	0.63	0.07
Labour	0.54	0.07
Liberal Democrat	0.52	0.08
Best Prime Minister		
Major/Hague/Duncan-Smith	0.89[a]	0.07
Kinnock/Smith/Blair	0.85[a]	0.08
Ashdown/Kennedy	0.76	0.08
Economic evaluations		
Personal retrospective	0.74	0.07
Personal prospective	0.61	0.06
National retrospective	0.93[a]	0.08
National prospective	0.82[a]	0.09

[a] Fails to reject null hypothesis that $d = 1.0$, $p = 0.05$.

Source: Monthly PSCB and P&D monthly surveys during January 1992 to December 2002.

In this model, β_1 and β_2 measure the short-term effects best prime minister judgements and economic evaluations. The term $(PID_{t-1} - \gamma_1 PM_{t-1} - \gamma_2 EVAL_{t-1})$ is an error correction mechanism (e.g. Hendry 1995) which captures the long-term effects of these variables on party identification. The α parameter measures the rate at which shocks to party identification from whatever source are eroded by the long-run relationship between party identification on the one hand, and party leader performance judgements and economic evaluations on the other.[28] If α equals 0, then this means that reactions to economic conditions and party leaders do not have long-term effects on party identification. Any effects of the former variables on the latter are short term and are captured by the β_1 and β_2 parameters.

To address controversies regarding the impact of different kinds of economic evaluations on party support (see Chapter Two), we specify four FEC models of Labour and Conservative party identification. One of the four types of economic evaluations—personal retrospections, personal prospections, national prospections, and national retrospections—appears in each model. With respect to the best prime minister judgements, the percentage selecting the Conservative leader is included in the analysis of Conservative party identification, and the percentage selecting the Labour leader is used in the analysis of Labour party identification. A parsimonious set of variables measuring the impact of important events is also included in the models. In the Labour model, these include the ERM crisis (September 1992), the change in Clause Four of Labour's constitution (April 1995), the 1997 and 2001 elections, the petrol protest (September 2000), the terrorist attack on the World Trade Center (September 2001 or 9/11), and the ensuing war in Afghanistan

(September 2001–January 2002).[29] Events in the Conservative model are the ERM crisis, Clause Four, and the 1997 election.

Table 6.11 presents parameter estimates for the Labour party identification models. The model using personal retrospective economic evaluations performs

Table 6.11 Error correction models of Labour party identification, April 1992–December 2002

	β value for the Models			
Predictor variables	*Personal retrospective* β	*Personal prospective* β	*National retrospective* β	*National prospective* β
ΔBest Prime Minister(*t*)	0.25***	0.28***	0.29***	0.26***
Economic evaluations				
ΔPersonal retrospections(*t*)	0.14**	*x*	*x*	*x*
ΔPersonal prospections(*t*)	*x*	0.04	*x*	*x*
ΔNational retrospections(*t*)	*x*	*x*	0.02	*x*
ΔNational prospections(*t*)	*x*	*x*	*x*	0.04[a]
Error correction mechanism				
Party identification($t-1$)	−0.41***	−0.41***	−0.35***	−0.38***
Best Prime Minister($t-1$)	0.31***	0.28***	0.24***	0.15*
Personal retrospections($t-1$)	0.17**	*x*	*x*	*x*
Personal prospections($t-1$)	*x*	0.12**	*x*	*x*
National retrospections($t-1$)	*x*	*x*	0.08*	*x*
National prospections($t-1$)	*x*	*x*	*x*	0.04
Shocks				
ERM crisis($t-1$)	3.28[a]	3.90[a]	3.42[a]	3.86[a]
Clause four($t-1$)	−3.25[a]	−3.39[a]	−3.33[a]	−3.05
1997 election(*t*)	4.25*	4.69*	4.36*	3.99[a]
Petrol crisis($t-1$)	−2.23***	−2.55***	−1.64*	−2.43***
2001 election(*t*)	2.29	2.60	3.57[a]	3.06
911 and Afghan war	2.20*	3.27**	2.96**	3.83***
Constant	13.14***	14.49***	13.01***	15.89***
Adjusted R^2	0.47	0.43	0.42	0.41
AIC	−293.27	−298.12	−299.71	−300.65
DW	1.69	1.59	1.75	1.64
Serial correlation (12 lags)	17.08	21.82*	14.05	16.85
Normality	3.53	7.78**	9.34**	7.37*
Heteroscedasticity	0.30	0.34	1.18	1.53

x—variable not included in model.
*** $p \leq 0.001$; ** $p \leq 0.01$; * $p \leq 0.05$.
[a] $p \leq 0.10$, one-tailed test.

Source: Monthly PSCB and P & D surveys from April 1992 to December 2002.

best. In addition to having the best fit as measured by the adjusted R^2 and the best AIC value (see Chapter Four), all coefficients for the long- and short-term effects of the economic and party leader evaluation variables are statistically significant and properly signed. In the other three models, either the short- or the long-term effects of economic evaluations do not achieve statistical significance ($p \leq 0.10$). Party leader evaluation variables are significant in all four models. Several salient events also affected Labour partisanship. The ERM crisis, the 1997 general election, 911, and the subsequent war in Afghanistan had positive effects on the percentage of Labour identifiers. In contrast, the September 2000 petrol protests diminished Labour's partisan share, as did the decision to change Clause Four of the party's constitution.

As should be the case for error correction models, the α coefficients in the Labour models are statistically significant, negatively signed, and less than 1.0. The α (-0.41) for the personal retrospections model indicates that shocks to Labour party identification were eroded at the rate of 41 per cent per month by the long-term relationship between Labour identification on the one hand, and personal economic evaluations and best prime minister judgements on the other. Thus, although various events and conditions affected the size of Labour's partisan share, assessments of one's personal economic condition and the performance of the Labour leader eventually offset these effects. This is what one would expect if party and party leader performance evaluations are key factors in the process that updates party identification over time.

Models of Conservative party identification are summarized in Table 6.12. Although these models are generally similar to those for Labour, there are noteworthy differences as well. First, the performance of all four models is very similar, although the adjusted R^2 and AIC values are slightly better in the case of the national prospective economic evaluation model. Second, unlike the Labour case, the short- and long-term effects of the economic and party leader evaluation variables are statistically significant and properly signed in all four models. Third, although α coefficients are significant in the four Conservative models, their values (-0.16 to -0.23) are considerably smaller than those for the Labour models. This suggests that Conservative party identification was tethered in the long run to economic evaluations and judgements of party leader performance. However, that tether was considerably looser than was the case for Labour. Net of these several effects, Conservative party identification was influenced by important events. As anticipated, the Conservative partisan share dropped in the aftermath of the ERM crisis (see Chapter Three). It also fell in the wake of the party's crushing defeat in the 1997 general election, and in response to Labour's decision to jettison Clause Four. The latter result suggests that both Labour and the Conservatives lost partisans as a result of New Labour's decision to 'modernize' the party's constitution.[30]

Diagnostics indicate that we can believe the stories that the time series models are telling us. With few minor exceptions, the models generally perform well on standard diagnostic tests for autocorrelation, heteroskedasticity, and normality (see Tables 6.11 and 6.12). Regarding which model is preferable, encompassing tests (Charemza and Deadman 1997; see also Chapter Four) confirm that the personal retrospections model provides a better account of the dynamics of Labour party identification than its rivals. The personal retrospections model encompasses all of

Table 6.12 Error correction models of Conservative Party identification, April 1992–December 2002

	β value for the Models			
Predictor variables	*Personal retrospective* β	*Personal prospective* β	*National retrospective* β	*National prospective* β
ΔBest Prime Minister(t)	0.21***	0.22***	0.21***	0.22***
Economic evaluations:				
ΔPersonal retrospections(t)	0.07*	*x*	*x*	*x*
ΔPersonal prospections(t)	*x*	0.04*	*x*	*x*
ΔNational retrospections(t)	*x*	*x*	0.04*	*x*
ΔNational prospections(t)	*x*	*x*	*x*	0.05***
Error correction mechanism				
Party identification($t-1$)	−0.21***	−0.23***	−0.16**	−0.16**
Best Prime Minister($t-1$)	0.75***	0.69***	0.78***	0.87***
Personal retrospections($t-1$)	0.11*	*x*	*x*	*x*
Personal prospections($t-1$)	*x*	0.10*	*x*	*x*
National retrospections($t-1$)	*x*	*x*	0.02	*x*
National prospections($t-1$)	*x*	*x*	*x*	0.05
Shocks				
ERM crisis($t-1$)	−2.85*	−2.77*	−3.64*	−3.95**
Clause four($t-1$)	−2.65[a]	−2.93*	−3.03*	−2.96*
1997 election(t)	−2.89*	−3.26*	−2.42[a]	−2.40[a]
Constant	3.87*	4.07**	2.69*	2.45*
Adjusted R^2	0.30	0.30	0.30	0.33
AIC	−250.23	−250.12	−250.95	−247.89
DW	1.92	1.97	1.92	1.94
Serial correlation (12 lags)	17.85	15.00	19.44	19.46
Normality	5.42	7.67*	3.34*	0.87
Heteroskedasticity	2.84	1.75	4.05*	7.54***

x—variable not included in model.
*** $p \leq 0.001$; ** $p \leq 0.01$; * $p \leq 0.05$.
[a] $p \leq 0.10$ one-tailed test.

Source: Monthly PSCB and P & D surveys from April 1992 to December 2002.

the alternative models, but none of them encompass it. In contrast, encompassing tests of the rival models of Conservative party identification are inconclusive, thus providing additional testimony that the performance of the four models is very similar.

A final set of diagnostic tests investigates the exogeneity of leader effects in the party identification models. Such tests address the question of whether, at any time *t*, judgements about party leaders reflect as well as cause party identification. If so, then the party identification analyses suffer from simultaneity bias and, as a result, one cannot believe the parameter estimates (Greene 2003: 396). As in the analyses of voting behaviour in Chapter Four, exogeneity tests may be performed to determine whether simultaneously bias is, indeed, a problem. These procedures show that party leader evaluations are weakly exogenous to party identification. As required, the error correction mechanisms for the party identification models are not significant ($p > 0.05$) in the leader models. As also required, the residuals from the leader models are not significant in the party identification models. This means that the coefficients are free of simultaneity bias. A related point deserves emphasis. These results do *not* mean that party identification at time *t* never affects evaluations of party leader performance at some later time $t + i$. Indeed, it is plausible that party identification serves as a heuristic mechanism when voters seek to acquire and process information. The result is that party identification plays a kind of inter-temporal 'tennis match' with party leader evaluations, with effects flowing back and forth between them over time. Sorting out that process statistically raises the question of 'strong exogeneity' (Charemza and Deadman 1997), which is not something that needs to be investigated in order to believe the testimony provided by the time series models of party identification.

CONCLUSION: VALENCED PARTISANSHIP

The analyses presented in this chapter are inconsistent with a sociological approach to party support. Rather, the evidence presented here speaks strongly on behalf of 'valenced partisanship' in the British electorate. By valenced partisanship, we mean that voters' party identifications can be thought of as a storehouse of accumulated party and party leader performance evaluations. Moreover, we show that this has always been the case in Britain. To begin with, partisan attachments have considerable dynamism. Some of these dynamics involve long-term aggregate trends. Consonant with claims first advanced by Crewe et al. in the late 1970s, BES data gathered over the past four decades reveal downward trends both in the strength of party identification and in the strength of its relationship with social class. These trends are evident for the entire 1964–2001 period. However, when interpreting this finding, it would be incorrect to assume that 1960s' Britain was a nation of very strong partisans. Nor were the 1960s a time overwhelming majorities of working class people identified with Labour, and overwhelming majorities of middle class people identified with the Conservatives. Rather, the 1960s BES data show that many voters did not claim to be intense partisans, and large minorities of working and middle class people did not identify with the 'natural' party of their class. Thus, an impressive dealignment of degree has occurred in an electorate where the intensity of partisan attachments and their linkages with social class have always been weaker than the popular image of tribal politics would suggest.

Other analyses indicate that partisan attachments do not behave in accordance with the Michigan model of party identification. Converse's (1969) conjecture concerning the strengthening of party identification across the life cycle does not fare well in Britain. BES evidence showing tendencies for party identifications to strengthen over the life cycle is very weak, and if such tendencies do exist, then it is clear that they have been overwhelmed by age-cohort and period effects. Moreover, party identifications do not manifest the individual-level stability assumed by proponents of the Michigan model. Panel data from the 1960s to the present clearly show that very sizeable minorities change their party identifications. Some move from being a party identifier to being a nonidentifier or vice versa, whereas others switch between parties. This individual-level partisan instability is not an artefact of random measurement error. Rather, mixed Markov latent class analyses show that there are large mover groups in all of the BES and BEPS panels. Nor is it the case that the apparent mutability of partisanship reflects flaws in the wording of the BES party identification question. Responses to alternative party supporter questions indicate lower levels of partisanship, but they also show considerable individual-level instability.

Evidence of large-scale, ongoing dynamics in partisan attachments accords well with the conception of valenced partisanship. This model has it origins in the updating models of party identification proposed by Fiorina (1981) and others. In these models, as voters acquire new information, they use it to reassess their existing party identifications, with previous information being progressively discounted over time. The intensity and direction of partisanship can change when new information strongly contradicts that which is stored in an existing party identification. Analyses of panel data show that such updating models of party identification perform well. Evaluations of the performance of the economy, political parties, and party leaders influence the direction and intensity of party identification in predictable ways.

Additional evidence of the nature of British voters' partisan attachments is provided by time series data gathered in monthly Gallup surveys. The long-memory dynamics of party identification in the Gallup data are consistent with the mixture of movers and stayers found in the individual-level panel data. Once more, economic evaluations and party leader performance judgements have important effects on the dynamics of party identification.

In sum, the evidence of substantial, ongoing dynamics in partisanship in Britain is strong. Party identification is not now and, at least since the 1960s, has not been an unmoved mover. Its movements are explicable in term of a simple information-updating model. This valenced partisanship model does not make strong assumptions about voter rationality but, rather, contends that voters behave sensibly. Reminiscent of the boundedly rational economic agents described by Simon (e.g. 1987) (see also Conlisk 1996), voters use current information about the performance of parties and their leaders in sensible, albeit rough and ready, ways as guides to political action, and they reduce deliberation costs by storing that information in updated partisan attachments. This ongoing process leads some voters to stay where they are, but prompts others to move on.

NOTES

1. The 2001 BES party identification sequence is: (i) 'Generally speaking, do you think of yourself as Labour, Conservative, Liberal Democrat, Scottish Nationalist [in Scotland only], Plaid Cymru [in Wales only], or what? (ii) [If "None", "Don't Know", or "Refused"] Do you generally think of yourself as a little closer to one of the parties than the others?' (iii) [If 'Yes'] Which party is that? (iv) 'Would you call yourself very strong [PARTY GIVEN in (i) or (iii)], fairly strong or not very strong?' See Appendix B for the wording of the 1964–97 BES party identification questions.

2. Social class is measured using the Market Society classification. People in categories A, B, and C1 are middle class, and those in categories C2, D, and E are working class.

3. The Consistency Index (CI) is a generalized version of the Alford index (see Chapter Three). The CI is computed by cross-tabulating social class (dichotomised as discussed in Note 2 above) with party identification categorized as Conservative, Labour, Liberal/Liberal Democrat, other party, none. The index is the sum of the absolute values of the differences in working and middle class respondents supporting the Labour and Conservative parties.

4. The variable is scored 0 for 1964 and 1966, and 1 for subsequent years.

5. It is not the case that simply calculating the percentages of people who fail to identify with 'their proper party' obscures strong correlations between class membership and party identification. Using the 2001 BES data, a multinomial logit analysis of party identification with social class as the predictor variable yields an estimated R^2 of only 0.02. Nor is it the case that other socio-demographic variables are powerful predictors that have been overlooked by scholars preoccupied with social class. A multinomial logit analysis of the 2001 data using age, employment sector, ethnicity, gender, home ownership, and region, as well as social class, as predictors produces an estimated R^2 of 0.08.

6. This decrease in the strength of party identification may, in part, reflect a slight change in question wording (see Appendix B).

7. The variable is scored 0 for 1964, 1966, and 1970, and 1 for subsequent years.

8. Although the British Household Panel Surveys (BHPS) will eventually provide very long term, individual-level data on party support, they do not ask the standard BES party identification question.

9. Since the age of majority was lowered from 21 to 18 in 1969 (Leonard and Mortimore 2001: 14), the main portion of the analysis starts with the 1970 BES data. In addition, the October 1974 data are not used since the cohort first entering the electorate at that time would be quite small, given that the preceding general election was only seven months earlier.

10. For example, if there are four equiprobable choices (Conservative, Labour, Liberal Democrat, other) and a voter chooses at random, then there is a 0.25 probability of selecting a particular party. If he or she is asked a party identification question and again chooses at random in a second survey, then probability of selecting the same party twice is $0.25 \times 0.25 = 0.063$.

11. Variables differ from one panel to the next depending on the mix of questions asked in the panel surveys. For example, for the 1997–8 panel, feelings about the Conservative and Labour parties were measured using variables ranging from 'strongly in favour' (scored +5) to 'neither/don't know' (scored +3) to 'strongly against' (scored +1). Changes in feelings about the parties were measured by subtracting a 1997 variable from its 1998 counterpart. Variables measuring party images included: (*a*) 'good for one class, good for all classes'. Responses that a party was good for all classes were scored +2, responses that a party was good for one class were scored 0, and 'don't knows'were scored 1; (*b*) 'capable of being a strong government, not capable of being a strong government'. Responses that a party was capable of strong government were scored +2, responses that a party was not capable of strong government were scored 0, and 'don't knows' were scored +1. The overall 1997 and 1998 party image variables were constructed by summing (*a*) and (*b*). Changes in party images were measured by subtracting the 1997 overall image variable from its 1998 counterpart. Details concerning construction of variables used in analyses of the 1992–4 and 1966–70 panels are available upon request.

12. The structural equation modelling procedures used by Green, Palmquist, and Schickler (e.g. 2002: ch. 3) require at least an ordinal ordering of categories. In their study of the stability of party identification in the United States, they follow a traditional practice of ordering identifiers along a seven-point scale from very strong Democrats to very strong Republicans, with independents (nonidentifiers) in the middle category. The structural equation models used by Greet et al. also require multi-wave panels to have sufficient data for the purpose of identifying parameters of interest.

13. British parties might be ordered along a general 'left-right' continuum with Conservatives on the right, Liberal Democrats in the middle, and Labour on the left. However, as discussed in Chapter Four, there is evidence that the Liberal Democrats are now seen as being further to left than are Labour. The Manifestos Project data (Budge et al. 2001) echo this finding (see Chapters Three and Eight). Placement of minor parties and nonidentifiers poses additional problems.

14. Note also that all of the response probabilities are less than 1.0. This result is consistent with Green et al.'s conjecture that party identifications are measured with error.

15. A party identification 'turnover table' analysis uses panel data and cross-tabulates party identification at time *t1* with party identification at time *t2*. All cases on the main (upper-left to lower-right) diagonal of the table are stable identifiers. Respondents in the off-diagonal cell are unstable identifiers.

16. The 'initial state' is the estimated proportions of Conservatives, Labour, and others in the mover and stayer groups in the first wave of the four-wave panel.

17. The BHPS party supporter question sequence is: (*a*) 'Generally speaking, do you think of yourself as a supporter of any particular party?' [If 'Yes', go to (c)]; (*b*) 'Do you think of yourself as a little closer to one party than to the others?' (*c*) 'Which one?' (*d*) 'Would you call yourself a very strong supporter of [PARTY], fairly strong or not very strong?'

18. The BSA party supporter question sequence is: (*a*) 'Some people think of themselves as usually being a supporter of one political party rather than another. Do you usually think of yourself as being a supporter of one particular party or not?' [If 'Yes'] 'Which party is that?'

19. Fiorina (1981) states that the constant may also be interpreted technically as a scaling device that takes into account differences in measurement of dependent and independent variables.

20. The sets of predictor variables vary from one panel to another depending upon the availability of relevant questions. For example, for the 1997–8 panel, the predictor variables are: (*a*) evaluations of how good a job the party leaders would do as prime minister, with categories ranging from 'very good' (scored +5) to 'don't know' (scored +3) to 'not at all good' (scored +1); (*b*) retrospective and prospective evaluations of national and personal economic conditions, with categories ranging from 'got/get a lot better' (scored +5) to 'stay(ed) the same/don't know' (scored +3) to 'got/get a lot worse' (scored +1). The four kinds of economic evaluation are summed to produce overall economic evaluations indices; (*c*) government performance evaluation indices based on judgements regarding how well the government has done in the areas of health, education, youth employment, taxation, spending controls, and corruption. For each area, the categories range from 'very successful' (scored +5) to 'don't know' (scored +3) to 'not at all successful' (scored +1). These evaluations are summed to produce the overall indices; (*d*) social class—the standard six-category measure, ranging from A (scored 6) to E (scored 1); (*e*) age: the 18–23, 24–41, 42–49, and 60–68 age-cohorts are 0–1 dummy variables with people over 69 as the reference category. The set of instruments for lagged party identification include: social class, the age cohort dummies, 1997 and 1998 economic evaluation indicies, 1997 and 1998 government performance indices, and the 1997 and 1998 party leader evaluations. Details regarding variables in the analyses of the 1992–4 and 1966–70 panels are available upon request.

21. The instrumental variables for party identification at time $t - 1$ include lagged versions of the economic evaluations, party performance evaluations, and party leader evaluation variables, as well as the measures of social class and age cohort. The set of these variables varies depending upon their availability in particular panels (details available upon request). Note also that analyses that do not use instruments for lagged party identification produce substantially similar results.

22. Nearest neighbour regression of a time series variable (e.g. the percentage of Conservative party identifiers) on a time counter variable is a convenient way of mapping nonlinear trends in the data (see Fox 2000).

23. If political variables such as party identification are nonstationary, then it is likely that the data generating processes at work involve stochastic trends. The simplest stochastic trend model is the random walk, that is, $Y_t = Y_{t-1} + \epsilon_t$. In this process, the value of Y at time t is equal to its value at $t - 1$ plus the value of any random shock that might occur at time t. Since the coefficient for Y_{t-1} the variance of the series grows without limit (Franses 1998). If the data generating process also includes a constant, that is, $Y_t = \beta_0 + Y_{t-1} + \epsilon_t$, then it is called a 'random walk with drift'. In addition to accumulating shocks, the random walk

with drift has a deterministic component such that, *ceteris paribus*, Y changes by β_0 in each successive time period.

24. To see this, note that $(1 - L)^1 Y_t = \epsilon_t$ can be rewritten as $Y_t - Y_{t-1} = \epsilon_t$, which can be rewritten as $Y_t = Y_{t-1} + \epsilon_t$.

25. Unit-root tests are a popular way of testing for nonstationarity. However, these tests have been criticized for their lower power in the face of fractionally integrated and near-integrated alternatives (see, e.g. Maddala and Kim 1998).

26. The Gallup economic evaluation questions are: (*a*) personal prospections—'How do you think the financial situation of your household will change over the next 12 months?' (*b*) personal retrospections—'How does the financial situation of your household now compare with what it was 12 months ago?' (*c*) national prospections—'How do you think the general economic situation in this country will develop over the next 12 months?' (*d*) national retrospections—'How do you think the general economic situation in this country has changed over the last 12 months?' The response categories are: 'get(got) a lot better', 'get a little better', 'stay the same', 'get a little worse', 'get a lot worse'. The economic evaluation variables are constructed by subtracting the percentage offering negative responses from the percentage offering positive ones. These economic evaluation variables are multiplied by -1 starting in May 1997 to take account of the change in government from Conservative to Labour. The question measuring which party leader would make the best prime minister is: 'Who would make the best prime minister—Mr. Hague, Mr. Blair, Mr. Kennedy?' with the names of the party leaders changing as appropriate (see also Chapter Three). In the Labour analysis, the percentage selecting Smith/Blair is used; in the Conservative analysis, the percentage selecting Major/Hague is used.

27. $(1 - L)^d$ can be expanded (infinitely) as $1 - dL - (1/2)d(1 - d)L^2 - (1/6)d(1 - d)(2 - d)L^3 - \ldots - (1/j!)d(1 - d)(2 - d)\ldots((j - 1) - d)L^j - \ldots$ See Franses (1998: 79).

28. For example, assume a one-time shock increases Labour party identification by 10 points in month *t*. If $\alpha = -0.5$, this means that the effect of the shock will decrease at the rate of 50 per cent per month. Thus, the impact of the shock on party identification will be 5 points in month $t + 1$, 2.5 points in month $t + 2$, etc.

29. There is not a canonical list of politically important events and, ultimately, degrees of freedom will limit how many can be modelled. In the interest of parsimony, we confine attention to a small group of events that preliminary analyses suggest may have exerted significant effects on Conservative or Labour identification. With one exception, the events variables are coded 1 for the month in which they occurred, and 0 otherwise. The exception is the September 2000 petrol protest variable. Preliminary analyses suggest that this variable marked a thus far enduring downward 'break' in Labour partisanship. Hence, the variable is coded 1 for each month from September 2000 onwards, and 0 otherwise.

30. A plausible interpretation of the negative Clause Four parameters in the Conservative and Labour models is that its abandonment persuaded some Conservative identifiers that Tony Blair and New Labour really were different from their predecessors. Hence, they defected. But Labour paid a price as well, namely the support of some Labour's core supporters for whom Clause Four symbolized the party's historic commitment to socialism. The analyses suggest that, in net terms, Labour lost more identifiers than it gained.