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Fractional (Co)integration and Governing Party Support in Britain

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Recent developments in the analysis of long-memoried processes provide important leverage for analysing time-series variables of interest to political scientists. This article provides an accessible exposition of these methods and illustrates their utility for addressing protracted controversies regarding the political economy of party support in Britain. Estimates of the fractionally differencing parameter, *d*, reveal that governing party support, prime ministerial approval and economic evaluations are long-memoried and non-stationary, and that governing party support and prime ministerial approval are fractionally cointegrated. *Pace* conventional wisdom that party leader images matter little, if at all, analyses of multivariate fractional error correction models show that prime ministerial approval has important short-run and long-run effects on party support. Prospective and retrospective personal economic evaluations are influential but, contrary to a longstanding claim, national economic evaluations are not significant. The article concludes by suggesting that individual-level heterogeneity is a likely source of the observed aggregate-level fractional integration in governing party support and its determinants. Specifying parsimonious models that incorporate theoretically meaningful heterogeneity is a challenging topic for future research.

The history of science testifies that methodological innovation can provide important leverage for addressing theoretical controversies. In this article, we provide an accessible exposition of leading-edge statistical methods for analysing long-memoried time series, and employ these methods to address protracted debates concerning the political economy of public support for parties and party leaders in Britain. After reviewing the controversies at issue, we introduce the concepts of long memory and fractional integration, and consider why one might expect key variables in party support models to be fractionally integrated. Next, we discuss the concepts of fractional cointegration and fractionally integrated error correction mechanisms, and analyse the memory properties of political and economic variables of interest. Multivariate fractional error correction models are used to investigate the determinants of governing party support, and encompassing tests are employed to address disputes about the explanatory efficacy of various kinds of economic evaluations. In the conclusion, we summarize major findings and discuss methodological and theoretical implications for future studies of party support in Britain and elsewhere.

CONTINUING CONTROVERSIES

Following pioneering studies in the 1970s, most studies of the dynamics of party support in Britain have focused exclusively on the effects of objective economic indicators such as inflation and unemployment rates.[1] However, during the past decade, research has

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[1] Charles A. E. Goodhart and R. J. Bhansali, 'Political Economy', *Political Studies*, 18 (1970), 43–106; William L. Miller and William M. Mackie, 'The Electoral Cycle and the Asymmetry of Government and Opposition Popularity', *Political Studies*, 21 (1973), 263–79. For literature reviews, see Michael S. Lewis-Beck, *Economics and Elections: The Major Western Democracies* (Ann Arbor: University of Michigan Press, 1988); Helmut Norpoth, Michael S. Lewis-Beck and Jean Dominique Lafay, eds, *Economics and Politics: The Calculus of Support* (Ann Arbor: University of Michigan Press, 1991).

demonstrated the utility of employing measures of voters' judgements about their own and their country's financial well-being.[2] Although there is now consensus that economic evaluations provide substantial explanatory power, analysts continue to disagree which type of evaluation drives party support.

These debates have a lengthy intellectual lineage. Over three decades ago V. O. Key articulated his famous two-fold argument that voters are rational actors whose electoral decisions are guided by retrospective rather than prospective judgements of government performance.[3] Key's thesis was consistent with a frequently articulated conventional wisdom that incumbents are rewarded or punished for the state of voters' pocketbooks in the recent past.[4] According to this hypothesis, personal retrospective economic evaluations are what matter for party support. However, claims have also been advanced on behalf of personal *prospective* judgements. In Britain, these latter arguments gained considerable credence when Sanders used personal economic expectations as the linchpin in his two-stage forecasting model (the 'Essex model') that predicted the widely unanticipated Conservative victory in the 1992 British general election an impressive sixteen months before that event occurred.[5] Consistent with Sanders's analysis, Figure 1A shows personal economic expectations have a substantial positive correlation ($r = +0.49$) with governing party support over the 1979–96 period studied in this article.

There are also disputes between scholars such as Sanders who emphasize the importance of personal, or 'egocentric', evaluations and analysts such as Kinder and Kiewiet, MacKuen, Erikson and Stimson, Erikson, MacKuen and Stimson, and Norpoth, who favour national, or 'sociotropic', evaluations.[6] The sociotropic camp, in turn, is deeply divided. In research on approval of the president in the United States, MacKuen *et al.* argue that national expectations outperform national retrospections, whereas Kinder and Kiewiet as well as Norpoth make strong claims on behalf of the latter. Still others, such as Clarke and Stewart and Lewis-Beck, find that both national prospections and national retrospections are influential.[7] In their British study, Clarke and Stewart report that national expectations perform as well as personal ones in analyses of governing party support, but that national retrospections outperform other types of economic evaluations as determinants of approval of the prime minister (prime ministerial approval).[8]

[2] See, e.g., Harold D. Clarke and Marianne C. Stewart, 'Economic Evaluations, Prime Ministerial Approval and Governing Party Support in Britain: Rival Models Reconsidered', *British Journal of Political Science*, 25 (1995), 145–70; David Sanders, 'Government Popularity and the Next General Election', *Political Studies*, 62 (1991), 235–61.

[3] V. O. Key Jr, *The Responsible Electorate: Rationality in Presidential Voting, 1936–1960* (New York: Vintage Books, 1968).

[4] See, e.g., D. Roderick Kiewiet, *Macroeconomics & Micropolitics: The Electoral Effects of Economic Issues* (Chicago: University of Chicago Press, 1983), chap. 2.

[5] Sanders, 'Government Popularity'.

[6] See, *inter alia*, Donald Kinder and D. Roderick Kiewiet, 'Sociotropic Politics: The American Case', *British Journal of Political Science*, 11 (1981), 129–61; Michael B. MacKuen, Robert S. Erikson and James A. Stimson, 'Peasants or Bankers? The American Electorate and the US Economy', *American Political Science Review*, 86 (1992), 597–611; Michael B. MacKuen, Robert S. Erikson and James A. Stimson, 'Comment', *Journal of Politics*, 58 (1996), 793–801; Helmut Norpoth, 'Presidents and the Prospective Voter', *Journal of Politics*, 58 (1996), 776–92; Helmut Norpoth, 'Rejoinder', *Journal of Politics*, 58 (1996), 802–5; Robert S. Erikson, Michael B. MacKuen and James A. Stimson, 'Bankers or Peasants Revisited: Economic Expectations and Presidential Approval', *Electoral Studies*, 19 (2000), 295–312.

[7] Harold D. Clarke and Marianne C. Stewart, 'Prospections, Retrospections and Rationality: The "Bankers" Model of Presidential Approval Reconsidered', *American Journal of Political Science*, 38 (1994), 1104–23; Lewis-Beck, *Economics and Elections*, chap. 10.

[8] Clarke and Stewart, 'Economic Evaluations'.

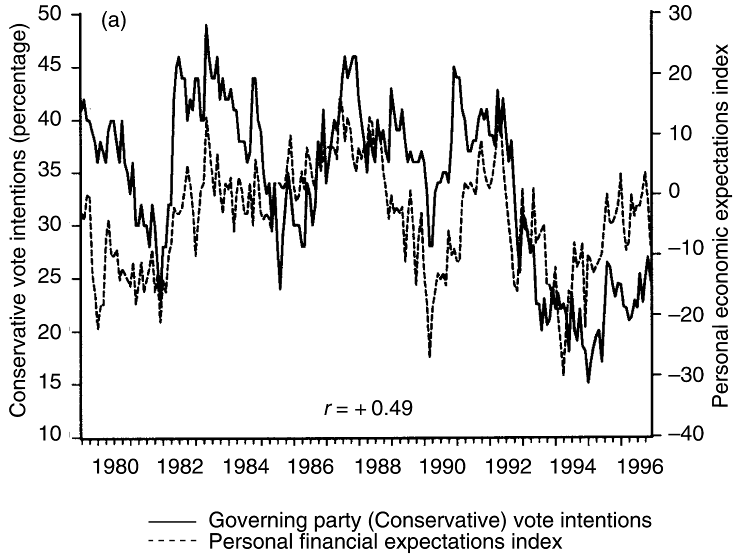

Fig. 1A. Governing party vote intentions and personal economic expectations, July 1979–December 1996

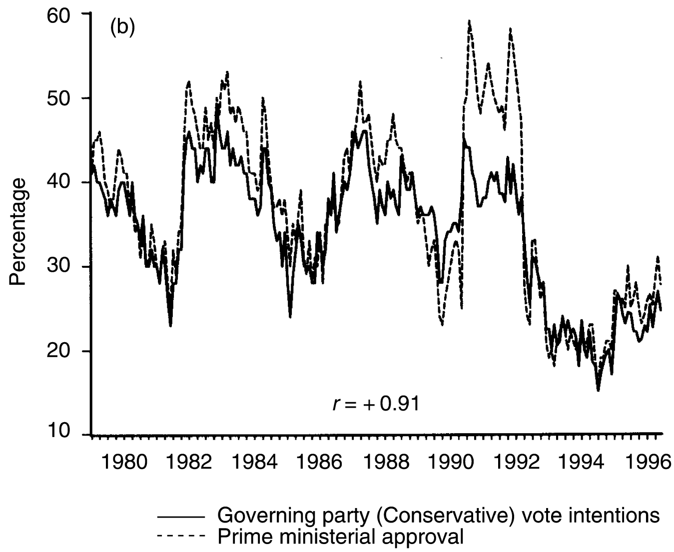

Fig. 1B. Governing party vote intentions and prime ministerial approval, July 1979–December 1996

An additional controversy in British studies concerns the explanatory role of leader images in party support models.[9] Party leader image variables typically are closely related

[9] David Butler and Donald E. Stokes, *Political Change in Britain*, 2nd College edn (New York: St Martin's Press, 1976), chap. 13; Clarke and Stewart, 'Economic Evaluations'; Harold D. Clarke, Marianne C. Stewart and Paul F. Whiteley, 'New Models for New Labour: The Political Economy of Labour Party Support, January 1992–April 1997', *American Political Science Review*, 92 (1997), 559–75; Ivor Crewe and Anthony King, 'Did Major Win? Did Kinnock Lose? Leadership Effects in the 1992 Election', in Anthony Heath *et al.*, eds, *Labour's Last Chance? The 1992 Election and Beyond* (Aldershot, Surrey: Dartmouth, 1994).

to measures of party support in analyses of aggregate time-series data. For example, over the 1979–96 period, prime ministerial approval and governing party support track one other very closely, the correlation between the two series being fully $+0.91$ (see Figure 1B). Impressive statistical relationships such as this have led some analysts to argue that including prime ministerial approval ratings in models of governing party support is tautological,[10] thus providing support for the traditional claim that party leader images matter little, if at all, as determinants of party support in Britain. Clarke and Stewart respond that strong correlations do not necessarily indicate that two variables are measuring the same thing, and that powerful (causal) relationships between conceptually distinct variables are exactly what political scientists wish to discover.[11] Additionally, they note that multivariate analyses of cross-sectional survey data from the British election studies (BES) also consistently indicate that voters' feelings about party leaders have sizeable effects on electoral choice.

This latter controversy has motivated interest in the concept of cointegration and the specification of error correction models that enables analysts to capture possible long-run relationships between party and leader support.[12] Below, we show that the traditional cointegration assumption undergirding these models is too restrictive, and that the relationship between prime ministerial approval and governing party support is best described as *fractionally* cointegrated. Before doing so, we discuss what fractional dynamics are, how they can arise and why one might expect the British time-series data to exhibit this property.

FRACTIONAL DYNAMICS AND PARTY SUPPORT

The concept of fractional integration generalizes the crucial distinction between stationary and non-stationary data. Until recently, time-series analysts have made a 'knife-edge' decision that their data are either stationary I(0) or non-stationary I(1), i.e., *unit-root*, processes.[13] This decision is theoretically consequential. Choosing stationarity implies that one believes a series has a constant mean – stochastic innovations (shocks) are gradually 'forgotten', and the series reverts to the mean. In contrast, the unit-root choice entails the belief that a series has perfect memory – with no shocks, its value at time t is the same as its value at the previous period, $t-1$.[14] However, because the data generating process (DGP) is stochastic, shocks do occur and are not discounted over time. As shocks accumulate, the unit-root series can meander widely, and is said to exhibit a stochastic trend. In the case of party support, such behaviour contradicts theories and empirical evidence concerning the re-equilibriating tendencies of party systems in mature democracies.[15]

[10] David Sanders, Hugh Ward and David Marsh (with Tony Fletcher), 'Government Popularity and the Falklands War: A Reassessment', *British Journal of Political Science*, 17 (1987), 281–313.

[11] Clarke and Stewart, 'Economic Evaluations'.

[12] Clarke and Stewart, 'Economic Evaluations'; Clarke, Stewart and Whiteley, 'New Models'.

[13] Janet M. Box-Steffensmeier and Renée M. Smith, 'The Dynamics of Aggregate Partisanship', *American Political Science Review*, 90 (1996), 567–80.

[14] A weak or wide-sense stationary series has a constant mean, a constant variance and a constant covariance for any lag k, i.e., $E(y_t) = \mu$; $E[(y_t - \mu)^2] = \gamma_0$; $E[(y_t - \mu)(y_{t-k} - \mu)] = \gamma_k$ for all t and k. The simplest unit-root series is the random walk: $Y_t = 1.0{*}Y_{t-1} + \varepsilon_t$, where $\varepsilon_t \sim N(0, \sigma^2)$. Recursive substitution shows that the random walk may be expressed as $Y_t = Y_0 + \Sigma\varepsilon_i$ where Y_0 is the initial value of Y and $\Sigma\varepsilon_i$ is the sum of all shocks from time 1 through time t. See Philip Hans Franses, *Time Series Models for Business and Economic Forecasting* (Cambridge: Cambridge University Press, 1998), p. 68.

[15] See, e.g., Stefano Bartolini and Peter Mair, *Identity, Competition and Electoral Availability 1885–1985* (Cambridge: Cambridge University Press, 1990); Butler and Stokes, *Political Change in Britain*; Donald E. Stokes and Gudmund R. Iversen, 'On the Existence of Forces Restoring Party Competition', in Angus Campbell *et al.*, eds, *Elections and the Political Order* (New York: John Wiley, 1966).

The stationarity–non-stationarity decision also has important consequences for analyses of factors affecting party support. As is well known, the possibility of spurious regressions – concluding that there is a statistically significant relationship between variables that are actually independent of one another – increases dramatically when non-stationary variables are analysed in level form.[16] Recent research indicates that similar threats to inference obtain for near-integrated series, i.e., series for which the DGP is very strongly autoregressive.[17] Assuming non-stationarity can also cause problems. When diagnostics indicate the presence of a unit root, the conventional remedy is to difference the series. However, differencing is not a cost-free option, because one will not be able to detect *long-run* relationships among the variables being analysed.[18] Analysts concerned about this possibility typically include error correction mechanisms (ECMs) in their models, asserting that the ECM specification is warranted by diagnostics indicating that variables of interest travel together in the long run. Such variables are said to cointegrate.[19]

In recent years, researchers have challenged the assumption that it is necessary for time-series variables to have *integer* orders of integration. Influenced by pioneering econometric research by Granger and Joyeux and by Hosking, analyses of American time-series data by Box-Steffensmeier and Smith and by Lebo *et al.* indicate that theoretically important variables such as presidential approval and party identification are *fractionally* integrated.[20] Series exhibiting fractional dynamics may be analysed by generalizing the familiar ARIMA models developed by Box and Jenkins.[21] The DGP of a (fractionally) integrated series Y is:

$$\phi(B)(1-B)^d Y_t = \theta(B)\varepsilon_t \text{ where } \varepsilon_t \sim N(0, \sigma^2) \tag{1}$$

where B is the backshift operator such that $B^k \varepsilon_t = \varepsilon_{t-k}$, $\phi(B)$ represents a stationary autoregressive (AR) process, $\theta(B)$ represents an invertible moving-average (MA) process,[22] and d, the fractional differencing parameter, measures the degree of integration of Y. If one assigns integer values to d, the traditional ARIMA framework is maintained.[23]

16 Clive W. J. Granger and Paul Newbold, 'Spurious Regressions in Econometrics', *Journal of Econometrics*, 2 (1974), 111–20.

17 A near-integrated process is: $Y_t = \beta_1 Y_{t-1} + \varepsilon_t$ where $\varepsilon_t \sim N(0, \sigma^2)$, $\beta_1 = 1.0 + \zeta$, and ζ is a small negative number, e.g., -0.01. See Suzanna DeBoef and Jim Granato, 'Near-Integrated Data and the Analysis of Political Relationships', *American Journal of Political Science*, 41 (1997), 619–40; Suzanna DeBoef and Jim Granato, 'Testing for Cointegrating Relationships with Near-Integrated Data', *Political Analysis*, 8 (2000), 99–117.

18 Renée M. Smith, 'Error Correction, Attractors, and Cointegration: Substantive and Methodological Issues', in John R. Freeman, ed., *Political Analysis*, vol. 4 (Ann Arbor: University of Michigan Press, 1993).

19 Robert F. Engle and Clive W. J. Granger, 'Co-integration and Error Correction: Representation, Estimation and Testing', *Econometrica*, 55 (1987), 251–76.

20 Clive W. J. Granger and R. Joyeux, 'An Introduction to Long-Memory Models and Fractional Differencing', *Journal of Time Series Analysis*, 1 (1980), 15–39; J. R. M. Hosking, 'Fractional Differencing', *Biometrica*, 68 (1981), 165–76; Box-Steffensmeier and Smith, 'The Dynamics of Aggregate Partisanship'; Janet M. Box-Steffensmeier and Renée M. Smith, 'Investigating Political Dynamics Using Fractional Integration Methods', *American Political Science Review*, 42 (1998), 661–89; Matthew Lebo, Robert Walker and Harold D. Clarke, 'You Must Remember This: Dealing With Long Memory in Political Analyses', *Electoral Studies*, 19 (2000), 31–48.

21 George E. P. Box and Gwilym M. Jenkins, *Time Series Analysis: Forecasting and Control*, rev. edn (San Francisco: Holden Day, 1976).

22 The invertibility condition implies that $|\theta| < 1$ so that an MA process can be represented as an infinite autoregression with the sum of the parameters being less than infinity. See T. C. Mills, *The Econometric Modelling of Financial Time Series* (Cambridge: Cambridge University Press, 1993), p. 14.

23 When $d = 0$, a series exhibits mean reversion, finite variance and covariance stationarity (see fn. 14 above), and it can be modelled using combinations of moving-average and autoregressive parameters in an ARMA (p,0,q) specification. When $d = 1$, the variable is a unit-root process characterized by mean, variance and covariance

Fractionally integrated processes relax the assumption that d is an integer and permit its value to lie on the real number line between 0 and 1.[24] Series in this middle ground possess *long memory* because they exhibit statistically significant correlations across long time lags, and they cease to be variance and covariance stationary when $0.5 \leq d < 1$.[25] It is often assumed that when $d \geq 0.5$, a fractionally integrated series remains mean reverting. Phillips and Xiao have recently pointed out that the latter is not true in the conventional sense, although shocks do eventually decay if $d < 1$.[26] As in the unit-root case, a fractionally integrated series can be rendered I(0) by differencing. For a fractionally integrated series, fractional-differencing rather than integer-differencing is employed.

To illustrate the defining long-memory property imparted by fractional integration, we simulate a series with d value of 0.85, and calculate its autocorrelation function across sixty lags. We also calculate autocorrelation functions for a simulated random walk (a non-stationary series) and a simulated stationary, first-order autoregressive series.[27] As Figure 2 shows, although the stationary series is specified to have a large AR1 coefficient ($\beta_1 = 0.85$), its autocorrelation function dampens quickly. In contrast, the random walk's autocorrelation function has extremely large values at low lags, and statistical significance is maintained over the first fifty-six consecutive lags. As intuition would suggest, the fractionally integrated series has intermediate (but decidedly non-trivial) persistence – its autocorrelations are greater (in absolute value) than those for the stationary series through the first thirty-six successive lags, but they are always less than those for the random walk. Indicative of its long memory, the fractional series autocorrelations remain statistically significant through thirty-three consecutive lags – nearly three years, if data were gathered on a monthly basis.

Granger discusses the statistical basis by which a fractionally integrated series can be created and demonstrates how aggregating individual-level heterogeneous behaviour produces fractional dynamics.[28] This heterogeneity exists with respect to varying autoregressive and moving average processes at the micro level. There are many political science time-series variables that would seem to fit this characterization. Of particular interest here are British time-series measures of party support, prime ministerial approval and subjective evaluations of national and personal economic conditions. These series are constructed by aggregating individual-level data from monthly national surveys.

(F'note continued)

non-stationarity. In this case, the series can be rendered stationary by first differencing and then modelled using AR and MA parameters in an ARIMA (p, 1, q) format.

[24] The simplest fractionally integrated process is fractional Gaussian noise, i.e., $(1-B)^d Y_t = \varepsilon_t$, where $\varepsilon_t \sim N(0, \sigma^2)$. $(1-B)^d$ can be expanded (infinitely) as $1 - d\mathrm{L} - (1/2)d(1-d)B^2 - (1/6)d(1-d)(2-d)B^3 - \ldots - (1/j!)d(1-d)(2-d)\ldots((j-1)-d)B^j - \ldots$. See Franses, *Time Series Models*, p. 79.

[25] Richard T. Baillie, 'Long Memory Processes and Fractional Integration in Econometrics', *Journal of Econometrics*, 73 (1996), 5–59; Jan Beran, *Statistics For Long Memory Processes* (New York: Chapman and Hall, 1994).

[26] Peter Phillips and Zhijie Xiao, 'A Primer on Unit Root Testing', in Michael McAleer and Les Oxley, eds, *Practical Issues in Cointegration Analysis* (Oxford: Blackwell, 1999), pp. 33–4.

[27] The three series were simulated using RATS 4.3, with the fractional series being created using the ARFSIM.PRG procedure.

[28] Clive W. J. Granger, 'Long Memory Relationships and the Aggregation of Dynamic Models', *Journal of Econometrics*, 14 (1980), 227–38. See also Janet M. Box-Steffensmeier and Renée M. Smith, 'Heterogeneity and Individual Party Identification' (paper presented at the Annual Meeting of the Midwest Political Science Association, Chicago, 1997); Christopher Wlezien, 'An Essay on Combined Time Series Processes', *Electoral Studies*, 19 (2000), 77–94.

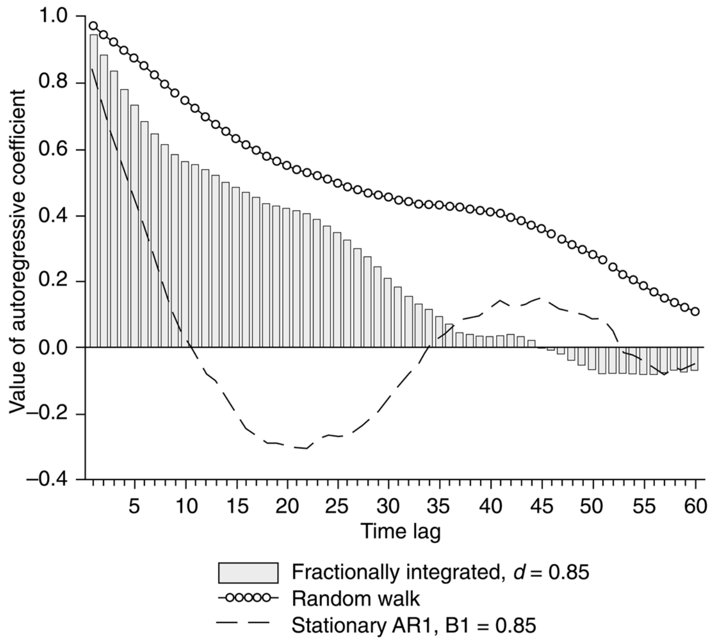

Fig. 2. Autocorrelation functions for fractionally integrated, random walk and stationary first-order autoregressive processes

Recent analyses of public opinion formation suggest that in each such survey there will be individuals with a diverse array of autoregressive tendencies.[29] Although much remains to be learned about the sources of these differences, it may be hypothesized that they are products of varying levels of political information and cognitive capacity, as well as variation in partisan attachments and other core political beliefs and values. For example, Zaller's contention that political knowledge is normally distributed across the electorate implies that voters respond differently to political and economic events.[30] Poorly informed people may repeatedly change their opinions based on very little new information, whereas the opinions of more politically knowledgeable individuals will vary less, barring some major event (such as the Falklands War or Britain's exit from the Exchange Rate Mechanism (ERM) currency crisis).

It may also be hypothesized that the degree of persistence manifested by different aggregate series will vary. For example, a series such as prime ministerial approval may exhibit unit-root behaviour because its dynamics are dominated by the presence of a sizeable group of voters who neither 'forgive nor forget'. Perhaps because the relationship between voters and prime minister has a specificity and interpersonal quality that is absent

[29] See Paul M. Sniderman, Richard A. Brody and Phillip E. Tetlock, eds, *Reasoning and Choice: Explorations in Political Psychology* (Cambridge: Cambridge University Press, 1991); John Zaller, *The Nature and Origins of Mass Opinion* (New York: Cambridge University Press, 1992). See also George A. Krause, 'Voters, Information Heterogeneity, and the Dynamics of Aggregate Economic Expectations', *American Journal of Political Science*, 41 (1998), 1170–200; George A. Krause, 'Testing for the Strong Form of Rational Expectations with Heterogeneously Informed Agents', *Political Analysis*, 8 (2000), 285–305; George A. Krause and Jim Granato, 'Fooling Some of the People Some of the Time? A Test of Weak Rationality With Heterogeneous Information Levels', *Public Opinion Quarterly*, 62 (1998), 135–51.

[30] Zaller, 'Nature and Origins'.

when more amorphous and impersonal entities such as political parties or the economy are considered, some voters' images of a prime minister will constitute a simple accumulation of salient information about that person's behaviour in office. In contrast, aggregate subjective economic evaluation series are likely to be somewhat shorter memoried as a vast majority of voters update their views of the state of the economy and their personal economic circumstances while simultaneously discounting previously acquired information. However, it is unlikely that these processes are homogeneous because voters have different locations in the economy,[31] and 'media, markets and mates' supply differing types and quantities of economic information. The result will be that aggregate economic evaluation series will exhibit long, but imperfect, memory.[32]

More generally, it is plausible that fractional dynamics (of varying degree) may characterize many of the series studied by political economists. It bears emphasis that such series should not be treated as if they were garden variety stationary (i.e., I(0)) processes. In cases where $d \geq 0.5$, failing to account for fractional dynamics will lead to problems in regression analyses, including biased standard errors and Durbin-Watson test statistics, and an enhanced possibility of spurious regressions.[33] How should analysts concerned about the possibility of fractional dynamics in their data proceed? Prior to developing and testing multivariate models, two steps analogous to those employed in a traditional ARIMA modelling exercise should be followed. First, rather than assuming variables are integrated at integer levels, for example, I(0) or I(1), the memory characteristics of each series should be identified empirically. The identification process involves estimating the fractional differencing parameter d and its standard error for each series. Secondly, before employing the data for estimating the parameters in multivariate models, each series should be differenced by its estimated d value. In the next section, we describe the British data and estimate d for each series.

DATA AND THE FRACTIONAL DIFFERENCING PARAMETER

We employ data from 210 consecutive monthly British Gallup polls spanning the July 1979–December 1996 period.[34] The variables are: governing (Conservative) party

[31] For example, variation in extent of concern with different aspects of the macroeconomy, e.g., inflation versus unemployment for middle-class and working-class persons respectively, is a possible source of heterogeneity in economic evaluations. See Douglas A. Hibbs, 'Economic Outcomes and Political Support for British Governments Among Occupational Classes: A Dynamic Analysis', *American Political Science Review*, 76 (1982), 259–79; Douglas A. Hibbs, 'The Dynamics of Political Support for American Presidents Among Occupational and Partisan Groups', *American Journal of Political Science*, 26 (1982), 312–32. Similarly, Sanders observes that one would expect concern with interest rates to vary according to whether a voter is a mortgage holder or a tenant. See Sanders, 'Government Popularity'.

[32] Estimates of the d parameter with aggregate time-series data will not enable one to determine the proportions of voters manifesting various patterns of stability and change in economic and political evaluations. Individual-level analyses of multi-wave panel data are required for this purpose. Promising in this regard are mixed Markov latent class models which allow for multiple 'mover' and 'stayer' chains and take account of random measurement error in observed variables. See, e.g., Harold D. Clarke, Marianne C. Stewart and Paul F. Whiteley, 'New Labour's New Partisans: The Dynamics of Party Identification in Britain Since 1992', in Justin Fisher *et al*., eds, *British Elections & Parties Review*, vol. 9 (London: Frank Cass, 1999).

[33] Lebo, Walker and Clarke, 'You Must Remember This'. See also Janet M. Box-Steffensmeier and Andrew R. Tomlinson, 'Fractional Integration Methods in Political Science', *Electoral Studies*, 19 (2000), 63–76.

[34] In the run-up to the 1997 general election (January 1997), Gallup changed its survey methodology from in-person interviews with quota samples to telephone interviews with random digit dialling sample selection procedures. To avoid possible mode effects artefacts, we terminate the time series in December 1996. The decision is innocuous – only four of a possible 214 observations are lost.

support,[35] prime ministerial approval,[36] and four subjective economic evaluations – personal expectations, personal retrospections, national expectations and national retrospections.[37] Each variable is bounded (between 0 and 100 in the cases of governing party support and prime ministerial approval, and between − 100 and + 100 in the case of the economic evaluations) and each is an aggregation of individual-level data. These characteristics suggest that the series may be fractional integrated.

Estimates of d for each series are consistent with this hypothesis.[38] Although several techniques may be used for this purpose, we employ Robinson's semi-parametric estimator and Sowell's maximum likelihood estimator.[39] To determine the ARFIMA (autoregressive, fractionally integrated, moving average) model that best accounts for both short-term and long-term dynamics of each series, d is estimated for each variable in a $(0, d, 0)$ model and in (p, d, q) models containing up to three autoregressive and three moving average parameters (at lags 1, 2 and 3). The Akaike Information Criterion and the Schwartz Bayesian Criterion are employed to distinguish which of the sixteen rival models

[35] The Gallup vote intention questions are: (a) 'If there were a General Election tomorrow, which party would you support'? (b) [If "don't know"] 'Which party would you be most inclined to vote for'? Conservative vote intentions are calculated as the sum of the percentages of respondents answering 'Conservative' to (a) or (b).

[36] The Gallup prime ministerial approval question is: 'Are you satisfied or dissatisfied with [name] as Prime Minister'? Prime ministerial approval is measured as the percentage saying they are 'satisfied'.

[37] The subjective 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.

[38] Our diagnoses of the dynamic properties of the time series also included standard unit-root tests. As Maddala and Kim emphasize, these tests have limited power in the face of fractional alternatives, and different tests can produce contradictory results. For example, the Dickey–Fuller test indicates that party support is non-stationary. However, Cochrane's variance ratio test rejects the hypothesis that d equals 1.0, and the Kwiatkowski *et al.*'s KPSS test rejects the hypothesis that d equals 0.0. The tests speak more consistently when prime ministerial approval and the economic evaluation series are analysed. For prime ministerial approval, the Dickey–Fuller and variance ratio tests fails to reject the null hypothesis of non-stationarity, and the KPSS test rejects the null of stationarity. For economic evaluations, Dickey–Fuller and variance ratio tests fail to reject the null hypothesis of non-stationarity, and the KPSS test fails to reject the null of stationarity. For these tests, see David A. Dickey and Wayne A. Fuller, 'Distribution of the Estimators for Autoregressive Series with a Unit Root', *Journal of the American Statistical Association*, 74 (1979), 427–31; J. H. Cochrane, 'How Big is the Random Walk in GNP?' *Journal of Political Economy*, 96 (1988), 893–920; D. Kwiatkowski *et al*., 'Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root', *Journal of Econometrics*, 54 (1992), 159–78; G. S. Maddala and In-Moo Kim, *Unit Roots, Cointegration, and Structural Change* (Cambridge: Cambridge University Press, 1998).

[39] Procedures for estimating d include, *inter alia*, semiparametric estimation, approximate maximum likelihood estimation in the frequency domain, exact maximum likelihood estimation in the time domain, Bayesian techniques, bootstrapping and autoregression-based estimation. See John Geweke and Susan Porter-Hudak, 'The Estimation and Application of Long Memory Time Series Models', *Journal of Time Series Analysis*, 4 (1983), 221–38; R. Fox and M. S. Taqqu, 'Large Sample Properties of Parameter Estimates for Strongly Dependent Stationary Gaussian Time Series', *Annals of Statistics*, 14 (1986), 517–32; P. M. Robinson, 'Gaussian Semiparametric Estimation of Long Range Dependence', *Annals of Statistics*, 23 (1995), 1630–61; Gary Koop *et al*., 'Bayesian Analysis of Long Memory and Persistence Using ARFIMA Models', *Journal of Econometrics*, 76 (1997), 149–69; Michael K. Andersson and Mikael P. Gredenhoff, 'Robust Testing for Fractional Integration Using the Bootstrap' (Stockholm School of Economics Working Papers Series in Economics and Finance, No. 218, 1998); Franses, *Time Series Models*, pp. 78–80. Robinson's estimator is computed using the RATS RGSE.SRC procedure; Sowell's is computed using OX. RATS and OX ARFIMA software may be downloaded from www.estima.com and www.nuff.ox.ac.uk/Users/Doornik, respectively.

TABLE 1 *Estimates of Fractional Differencing Parameter* d

	Sowell's d estimate	Standard Error	Robinson's d estimate	Standard Error	t–ratio† ($d = 0$)	t–ratio† ($d = 1$)
Variables						
Governing Party Support	0.82	0.058	0.85	0.059	14.32**	− 2.53*
Prime Ministerial Approval	0.91	0.060	0.96	0.059	16.18**	− 0.67
Economic Evaluations:						
Personal Expectations	0.72	0.059	0.70	0.059	11.80**	− 5.06**
Personal Retrospections	0.77	0.051	0.84	0.059	14.16**	− 2.70**
National Expectations	0.61	0.060	0.63	0.059	10.62**	− 6.24**
National Retrospections	0.83	0.057	0.83	0.059	13.99**	− 2.86**

$N = 210$
*$p < 0.05$
**$p < 0.01$
† Based on Robinson's estimates.

per series is superior.[40] According to both criteria, a (0, *d*, 0) model is best for each of the six variables.

Table 1 presents the values of *d* computed using Robinson's and Sowell's estimators for these (0, *d*, 0) models. The two procedures yield very similar results – *d* does not equal zero for any series ($p < 0.05$), and it does not equal 1.0 for any series except possibly prime ministerial satisfaction. Although the point estimates for *d* of 0.96 and 0.91 for the latter variable indicate fractional integration, each is within two standard errors of 1.0. Thus, one cannot reject the unit-root hypothesis ($p = 0.05$). Based on these results, we conclude that fractional integration characterizes at least five of the six series. In the next section, we use these findings to help specify error correction models of the determinants of governing party support.

FRACTIONAL COINTEGRATION AND FRACTIONAL ERROR CORRECTION MODELS

Just as the likelihood of finding fractional dynamics in political variables motivates a rethinking of the knife-edge distinction between stationarity and non-stationarity, the methodology of cointegration must likewise be rethought in the light of the possibility of long-memory equilibrium relationships between variables. Among the advantages of this new methodology is the ability to specify and estimate equilibrium relationships among variables of interest more precisely. Thus, along with ARFIMA modelling procedures, fractional cointegration techniques improve the chances of resolving longstanding controversies regarding rival models of the dynamics of important variables such as governing party support. The concept of fractional cointegration is most easily understood as a more general form of traditional cointegration.

Cointegration involves a long-term relationship between two or more variables.[41] This relationship is one such that shocks to one variable at any time *t* are re-equilibrated at a constant rate in subsequent time periods.[42] Standard operating procedure for diagnosing cointegration between two variables involves, first, determining if the original series are non-stationary (I(1)) and, secondly, determining if the residuals from a regression of one of the variables on the other are stationary (I(0)). Unit-root tests are employed to assess the order of integration of the individual variables and the regression residuals.

As noted above, the problem with this strategy is that unit-root tests have low power, making it difficult to reject the null of non-stationarity when *d* is large but less than 1. This, and the possibility that many political and economic series may have such *d* values, erodes confidence in the conventional cointegration-testing methodology.[43] However, the requirement that the residuals of a cointegrating regression are I(0) is unnecessary.[44] It is

[40] Franses, *Time Series Models*, p. 59.

[41] Clive W. J. Granger, 'Some Properties of Time Series Data and their Use in Econometric Model Specification', *Journal of Econometrics*, 16 (1981), 121–30; Engle and Granger, 'Co-integration and Error Correction'; Robert F. Engle and Clive W. J. Granger, eds, *Long-Run Economic Relationships: Readings in Cointegration* (Oxford: Oxford University Press, 1991).

[42] For example, if *X* and *Y* are cointegrated and the re-equilibration rate is 0.5, *ceteris paribus*, a ten-point shock to *Y* at time *t* would be eroded by its cointegrating relationship with *X* by 50 per cent in each subsequent time period, i.e., to 5 points at $t + 1$, to 2.5 points at $t + 2$, to 1.25 points at $t + 3$, etc.

[43] John C. Barkoulas and Christopher F. Baum, 'Fractional Differencing Modeling and Forecasting of Eurocurrency Deposit Rates', *Journal of Financial Research*, 20 (1997), 355–72; Michael Dueker and Richard Startz, 'Maximum-Likelihood Estimation of Fractional Cointegration with an Application to US and Canadian Bond Rates', *Review of Economics and Statistics*, 80 (1998), 42–6.

[44] Yin-Wong Cheung and Kon S. Lai, 'A Fractional Cointegration Analysis of Purchasing Power Parity', *Journal of Business and Economic Statistics*, 11 (1993), 193–212.

sufficient that these equilibrium errors are not infinitely persistent. As long as the residuals produced by regressing one fractionally integrated series on another one are of a lower order of integration than are either of the fractionally integrated series, those series can be said to be in dynamic equilibrium and to *fractionally cointegrate*.[45] Hence, one may fuse the methodology of cointegration with the concept of fractional integration.

Extending the methodology of cointegration to include fractional cases enables one to generalize the concept of an error correction mechanism (ECM). Following Engle and Granger, standard practice is to specify an error correction mechanism when two series cointegrate in the traditional sense, thereby enabling one to study long-run relationships among non-stationary variables.[46] This methodology assumes that the residuals of the cointegrating regression, and hence, the ECM, are I(0). Relaxing this assumption as described above to allow for the possibility of fractional cointegration is a natural extension of this methodology. Rather than using (non-)stationarity tests to decide whether the ECM is I(0) or I(1), estimating the d parameter for the ECM permits a more nuanced understanding of the nature of the equilibrium relationship. The range of possible behaviour for the ECM is greatly increased – defining an ECM as possessing long memory simply implies that shocks to one variable are re-equilibrated at a slower rate than the exponential decay of a conventional ARMA process.

The existence of a fractionally integrated ECM does present a problem to the multivariate modeller, however. If one wishes to include a fractionally integrated ECM as an independent variable in a larger equation, one should be wary of threats to inference generated by its long-memory characteristics. Monte Carlo analyses indicate that even when a dependent variable and other independent variables are I(0), a single fractionally integrated independent variable can threaten inference, as both its coefficient estimates and standard errors are biased downwards.[47] Interpretation of an ECM's coefficient as a measure of the strength and significance of the adjustment mechanism will be flawed if one fails to account for fractional cointegration. Hence, it is important to ensure that ECMs are I(0) before including them in one's model. As discussed below in the analysis of British governing party support, this can be done by fractionally differencing the residuals of the cointegrating regression by their estimated value of d. A full multivariate ARFIMA model with an ECM specification thereby can be estimated without worrying about threats to inference posed by the presence of a fractionally integrated component.

Prior to estimating such models with British data, prime ministerial approval and Conservative party support are tested for fractional cointegration. As noted above, previous studies of the relationship between public support for British parties and their leaders have employed an ECM to capture a long-run equilibrium relationship that exists between the two variables. However, given the possibility of fractional cointegration, these studies may have over-estimated the rate at which leader approval and party support return to equilibrium following a shock. To investigate this possibility, a cointegrating regression is performed between Conservative party vote intentions and prime ministerial approval and the residuals are tested for stationarity. The d estimate for the residuals is 0.59

[45] Technically, a bivariate fractional cointegrating relationship may be defined as follows: two variables, Z_t and X_t where Z_t is I(d) and X_t is I(b) are fractionally cointegrated if there is a function g(.) such that $\eta_t = Z_t + g(X_t)$ is I(s) and $s < d$. Thus, the existence of some function $g(X_t)$ that reduces the order of integration for Z_t establishes that Z_t and X_t are fractionally cointegrated. See Karim Abadir and A. M. Robert Taylor, 'On the Definitions of (Co-)Integration' (unpublished paper, University of York, 1998), p. 6.

[46] Engle and Granger, 'Co-inegration and Error Correction'.

[47] Matthew Lebo, 'Fractional Integration and Political Modeling' (unpublished paper, Harvard–MIT Data Center, 2000).

(s.e. = 0.059). This is over four standard errors below the d values for either Conservative vote intentions or prime ministerial approval (see Table 1). Also, one cannot reject ($p = 0.05$) the hypothesis that d for the residual series is less than 0.5, i.e., $0.5 - (0.059 \times 2) = 0.472$. Thus, recalling that series with d less than 0.5 indicate stationarity (see p. 290 above), one cannot reject the possibility that the residuals constitute a stationary series. Taken together, these results are consistent with the hypothesis that the two series are fractionally cointegrated.

As discussed above (pp. 288–9, see also Figure 2), fractionally integrated series are said to be long-memoried because they manifest statistically significant correlations at long lags. The presence of such long memory in the residuals of the governing party support–prime ministerial approval regression tells us that a shock to one of these variables will persist longer than the geometric decay of a standard cointegrating system. Hence, assuming a standard cointegrating ($d = 0$) relationship between prime ministerial approval and governing party support may over-estimate the speed with which the two variables return to their long-term equilibrium following a shock. The implication is that, despite the strong correlation between the two variables, voters' opinions about the governing party and the prime minister can diverge for long periods before a shock dissipates and the variables return to equilibrium. Next, we will show how accounting for this long-run relationship in a multivariate context can improve model specification and enhance confidence in empirical findings.

FRACTIONAL ERROR CORRECTION MODELS OF GOVERNING PARTY SUPPORT

We now specify and estimate fractional error correction models of governing party support. To ensure that threats to inference are minimized, governing party support (Conservative vote intentions), prime ministerial approval and the four subjective measures are rendered I(0) by fractionally differencing each by its own value of d (see Table 1). This procedure is analogous to conventional (full) differencing in ARIMA modelling, and is easily implemented using specialized filtering programs.[48] To capture the long-term equilibrium relationship between Conservative support and prime ministerial approval, the error correction term (lagged one period) is included in the multivariate models. Before doing so, this variable (the residuals of the cointegrating regression of Conservative support on prime ministerial approval) is fractionally differenced by its d estimate (0.59). Other independent variables in the party support models include interventions for the Falklands War, the introduction of the poll tax, the replacement of Margaret Thatcher by John Major as prime minister, Major's re-selection as Conservative leader, the ERM crisis and a summary variable to capture miscellaneous political events.[49]

The error correction model of governing party support is:

$$\mathrm{GOVF}_t = \beta_0 + \beta_1 \mathrm{PMF}_t + \beta_2 \mathrm{EVALF}_t - \alpha \mathrm{FCEM}_{t-1} + \Sigma \beta_{3-k} \mathrm{SHOCKS}_{t-i} + \varepsilon_t \tag{2}$$

[48] The RATS FIF.SRC procedure is used to filter the series by their d values. The filtering operation produces a series of residuals which represent the difference between the original series and the values of the series predicted by their (0, d, 0) ARFIMA models.

[49] The Falklands war, replacement of Thatcher by Major, the ERM crisis and Major reselection interventions are coded as 1 for the month in which the intervention occurred and 0 otherwise. Since the poll tax implementation is hypothesized to exert a temporary one-month effect, it is scored 1 for the month it occurred, -1 for the next month, and 0 otherwise. The miscellaneous intervention variable is scored $+1$ for months in which there was an intervention hypothesized to increase Conservative support, -1 for months in which there was an intervention hypothesized to decrease Conservative support, and 0 for other months. A list of interventions is available upon request.

TABLE 2 *ARFIMA Error Correction Models of Governing Party Support, Two-Step Estimates, September 1979–December 1996*

	Models							
	Personal Expectations		Personal Retrospections		National Expectations		National Retrospections	
	b	s.e.	*b*	s.e.	*b*	s.e.	*b*	s.e.
Predictor Variables								
Constant	0.140	0.134	0.138	0.134	0.131	0.134	0.129	0.135
$\triangle^d$PM approval	0.378***	0.038	0.384***	0.038	0.383***	0.039	0.392***	0.040
Economic Evaluations:								
$\triangle^d$Personal Expectations	0.056*	0.028	†	†	†	†	†	†
$\triangle^d$Personal Retrospections	†	†	0.056*	0.033	†	†	†	†
$\triangle^d$National Expectations	†	†	†	†	0.016	0.013	†	†
$\triangle^d$National Retrospections	†	†	†	†	†	†	0.007	0.014
ECM:								
Adjustment parameter	− 0.272***	0.058	− 0.274***	0.059	− 0.282***	0.058	− 0.281***	0.059
Falklands War	6.625***	1.887	6.891***	1.889	6.795***	1.896	6.751***	1.912
Poll Tax	− 3.174**	1.327	− 3.825**	1.338	− 3.368**	1.330	− 3.473**	1.331
Major Replaces Thatcher	4.980**	1.912	4.765**	1.923	4.941**	1.926	4.989**	1.932
Currency Crisis ($t-3$)	− 3.373*	1.867	− 3.292*	1.870	− 3.198*	1.876	− 3.125*	1.881
Major Reselection ($t-1$)	5.641***	1.881	5.620***	1.889	5.904***	1.887	5.931***	1.892
Miscellaneous Events	1.520***	0.288	1.500***	0.288	1.518***	0.291	1.504***	0.294
Diagnostic Statistics								
Adjusted R^2	0.586		0.583		0.581		0.578	
SEE	1.859		1.864		1.871		1.876	
Durbin–Watson d	1.949		1.952		1.933		1.944	
Serial Correlation	8.980		10.168		8.765		7.880	
Functional Form	0.063		0.005		0.059		0.067	
Normality	13.066***		19.168***		17.313***		19.774***	
Heteroscedasticity	0.482		0.357		0.271		0.380	
ARCH (1)	0.207		0.014		0.053		0.029	

*$p \leq 0.05$; **$p \leq 0.01$; ***$p \leq 0.001$; one-tailed test. †Variable not included in model.
Note: All predictor variables modelled as operating at time t except as indicated.

where: GOVF = fractionally differenced governing party support; PMF = fractionally differenced prime ministerial approval; EVALF = fractionally differenced economic evaluations (personal prospective, personal retrospective, national prospective, national retrospective); FECM = fractional error correction mechanism estimated from first-stage regression of GOVF on PMF; SHOCKS = interventions (Falklands War, etc.); β, α = model parameters (to be estimated); ε = stochastic error term ($\sim N(0, \sigma^2)$.

Four separate analyses, using each of the subjective economic evaluations in turn, are performed.[50] Column 1 of Table 2 shows that prime ministerial approval, personal economic expectations, the error correction mechanism and each of the intervention variables have statistically significant effects ($p \leq 0.05$). Column 2 shows a nearly identical pattern with personal retrospective economic evaluations being significant. However, columns 3 and 4 show that prospective and retrospective judgements of national economic conditions fail to achieve statistical significance (t ratios = 1.16 and 0.49, respectively). In the four models, the estimated coefficient for the error correction mechanism is properly (negatively) signed, and varies by only 0.01 (from -0.27 to -0.28), thereby indicating that shocks to governing party support are slowly eroded by the (fractionally) cointegrating relationship with prime ministerial approval. A variety of such shocks were at work in the period under consideration. As one would anticipate, the Falklands War, the replacement of Thatcher by Major and Major's reselection as party leader boosted the Conservatives' vote intention share, whereas the imposition of the poll tax and the currency crisis diminished it. Miscellaneous events also were influential.[51]

The vote intention analyses indicate that both personal economic expectations and personal economic retrospections exercise significant effects on governing party (Conservative) support. This finding and failure of prospective or retrospective evaluations of the national economy to achieve statistical significance are consistent with the Essex model's emphasis on the importance of personal economic judgements. It also contradicts – for the British case at least – the arguments of the many analysts who have advanced claims on behalf of national prospective or, more frequently, national retrospective economic judgements. But is one of the personal economic evaluation models preferable to the other? Similar to models employing standard cointegration methodology, one may address this question by performing encompassing tests to compare the performance of rival non-nested models.[52] Empirically, a battery of encompassing tests show that neither personal economic evaluation model encompasses the other (Table 3).[53] This result, in turn, suggests that although the fractional error correction models indicate that personal economic evaluations perform better than national ones, the time horizon over which the former operate remains ambiguous.

[50] The models are estimated using the OLS regression program in MICROFIT 4.0. MICROFIT 4.0 provides an extensive battery of regression diagnostics, as well as encompassing tests for rival non-nested models.

[51] With the exception of normality of residuals, all regression diagnostic tests are statistically insignificant. Thus, there is no evidence of residual autocorrelation (first-order or general), heteroscedasticity (ARCH or general) or functional form misspecification. Also, CUSUM and CUSUM squared tests (not shown) are consistent with the hypothesis of parameter constancy.

[52] Wojciech W. Charemza and Derek F. Deadman, *New Directions in Econometric Practice* (Aldershot: Edward Elgar, 1997), chap. 8.

[53] R. Davidson and J. MacKinnon 'Several Tests for Model Specification in the Presence of Alternative Hypotheses', *Econometrica*, 49 (1981), 781–93.

TABLE 3 *Encompassing Tests of Rival Personal Economic Expectations (PE) and Personal Retrospective Economic Evaluation (PR) Models: Two-Step Fractional Error Correction Estimates*

	PE Model v. PR Model	PR Model v. PE Model
Encompassing Tests		
N Test	-1.127 ($p = 0.260$)	-2.200 ($p = 0.028$)
NT Test	-0.523 ($p = 0.601$)	-1.207 ($p = 0.228$)
W Test	-0.522 ($p = 0.602$)	-1.199 ($p = 0.230$)
J Test	0.811 ($p = 0.418$)	1.295 ($p = 0.195$)
JA Test	0.811 ($p = 0.418$)	1.295 ($p = 0.195$)
F Test	0.657 ($p = 0.419$)	1.677 ($p = 0.197$)

One-Step Estimation

Some analysts have argued that considerations of statistical efficiency make it preferable to estimate error correction model parameters using a one-step, rather than a two-step, procedure.[54] The implicit assumption is that all variables of interest, including those in the error correction mechanism, are stationary – either as supplied by nature, or by the application of suitable differencing procedures. As noted, when variables are fractionally integrated, stationarity can be achieved by filtering them using their estimated d values. Once governing party support, prime ministerial approval and the economic evaluations have been filtered by their d values, we may specify the one-step fractional error correction model of governing party support as:

$$\text{GOVF}_t = \beta_0 + \beta_1\text{PMF}_t + \beta_2\text{EVALF}_t - \alpha(\text{GOVF}_{t-1} - \gamma\text{PMF}_{t-1}) + \Sigma\beta_{3-k}\text{SHOCKS}_{t-1} + \varepsilon_t \quad (3)$$

where GOVF, PMF, EVALF, SHOCKS and ε are as described above (p. 297), and the β's, α, and γ are parameters to be estimated.

Parameter estimates (see Table 4) for different variants of this model using the four economic evaluation variables are very similar to those for the two-step models. Personal economic expectations and personal economic retrospections have significant effects (t ratios $= 2.17$ and 1.75, respectively), but neither national economic expectations nor national economic retrospections achieve significance (t ratios $= 1.09$ and 0.35, respectively). Once more, encompassing tests fail to indicate that the personal expectations model is preferable to its personal retrospective rival or vice versa (data not shown). Also as in the two-step models, the adjustment parameters (α) are significant in each model ($p < 0.001$), as are various intervention variables. The γ coefficients capturing the fractional cointegrating relationship between governing party support and prime ministerial approval in the one-step models are significant as well.

The several coefficients in the one-step models all carry the same signs and are very similar in magnitude to their two-step counterparts. Adjustment parameters for the one-step models are slightly smaller than their two-step counterparts (by 0.045 on average),

[54] See, e.g., Nathaniel Beck, 'Comparing Dynamic Specifications: The Case of Presidential Approval', in James A. Stimson, ed., *Political Analysis*, vol. 3 (Ann Arbor: University of Michigan Press, 1992); Suzanna DeBoef, 'Modeling Equilibrium Relationships: Error Correction Models with Strongly Autoregressive Data', *Political Analysis*, 9 (2001), 78–94.

TABLE 4 *ARFIMA Error Correction Models of Governing Party Support, One-Step Estimates, September 1979–December 1996*

	Models							
	Personal Expectations		Personal Retrospections		National Expectations		National Retrospections	
	b	s.e.	*b*	s.e.	*b*	s.e.	*b*	s.e.
Predictor Variables								
Constant	0.197	0.136	0.228	0.143	0.216	0.143	0.212	0.144
$\triangle^d$PM approval	0.364***	0.039	0.367***	0.039	0.367***	0.041	0.377***	0.041
Economic Evaluations:								
$\triangle^d$Personal Expectations	0.063*	0.029	†	†	†	†	†	†
$\triangle^d$Personal Retrospections	†	†	0.061*	0.035	†	†	†	†
$\triangle^d$National Expectations	†	†	†	†	0.015	0.014	†	†
$\triangle^d$National Retrospections	†	†	†	†	†	†	0.005	0.015
ECM:								
Adjustment parameter	− 0.235***	0.060	− 0.227***	0.062	− 0.233***	0.062	− 0.232***	0.062
PM Approval ($t-1$)	0.555***	0.153	0.529***	0.163	0.567***	0.160	0.556***	0.161
Falklands War	6.797***	1.917	7.098***	1.945	7.000***	1.955	6.985***	1.972
Poll Tax	− 3.207**	1.349	− 3.943**	1.381	− 3.448**	1.372	− 3.552**	1.373
Major Replaces Thatcher	4.805**	1.945	4.400*	1.988	4.626**	1.992	4.675**	1.999
Currency Crisis ($t-3$)	− 4.571**	1.908	− 4.531**	1.939	− 4.403*	1.946	− 4.329***	1.952
Major Reselection ($t-1$)	6.032***	1.933	6.120***	1.970	6.384***	1.972	6.438***	1.978
Miscellaneous Events	1.580***	0.298	1.660***	0.314	1.664***	0.317	1.650***	0.321
Diagnostic Statistics								
Adjusted R^2	0.575		0.576		0.572		0.570	
SEE	1.888		1.918		1.928		1.933	
Durbin–Watson d	1.984		1.993		1.978		1.987	
Serial Correlation	9.428		11.098		8.946		8.235	
Functional Form	0.048		0.113		0.280		0.282	
Normality	9.558*		13.555*		12.942*		14.866*	
Heteroscedasticity	0.007		0.010		0.012		0.001	
ARCH (1)	0.145		0.050		0.020		0.020	

*$p \le 0.05$; **$p \le 0.01$; ***$p \le 0.001$; one-tailed test. †Variable not included in model.
Note: All predictor variables modelled as operating at time t except as indicated.

but Wald tests show that the differences are not statistically significant (for example $\chi^2 = 0.38$, df = 1, $p = 0.54$ for the personal expectations model). This also is true for the coefficients (γ) measuring the long-run fractional cointegrating relationships between governing party support and prime ministerial approval – these coefficients are smaller (mean difference = 0.096) than that (0.648) estimated in the two-step procedure, but the differences are not significant (for example $\chi^2 = 0.37$, df = 1, $p = 0.54$ for the personal expectations model). A more general test constraining all parameters in the one-step models to be equal to their counterparts in the two-step ones reveals that the differences are not significant (for example $\chi^2 = 1.53$, df = 11, $p = 1.00$ for the personal expectations model). Thus, alternative estimation procedures for rival models incorporating the fractionally integrated variables tell essentially the same story about the determinants of governing party support in Britain.[55] In the conclusion, we will reprise the highlights of this story, and suggest avenues for future research.

CONCLUSION: NEW METHODS, NEW RESEARCH AGENDAS

The past decade has witnessed substantial theoretical and methodological innovation in studies of the dynamics of party support in Britain and other mature democracies. Sensitive to the serious threats to inference posed by non-stationary data, researchers have tested their data for unit roots and have specified error correction models of the impact of economic evaluations and party leader images on vote intentions. Some critics have been slow to accept the results of these analyses, issuing caveats that unit-root tests have low statistical power, and that variables such as voting intentions, prime ministerial approval and economic evaluations are bounded (either by nature or construction) and, hence, cannot be truly non-stationary at the I(1) level. In this article, we have argued that the concepts of fractional integration and fractional cointegration enable one to generalize the conventional integration–cointegration–error correction methodology to take account of these criticisms. Estimating the fractional differencing parameter d for key variables in a model of the dynamics of British governing party support reveals that threats to inference would arise if these variables were treated as stationary I(0) processes. The d parameters indicate that all of the variables are fractionally integrated and non-stationary. Governing party support and prime ministerial approval are fractionally cointegrated.

These findings warrant the specification of models that generalize the error correction approach used in previous studies by relaxing the assumption of integer orders of integration and cointegration. Estimating the parameters in these new fractional error correction models using either two-step or one-step procedures documents the existence of a governing party vote intention–prime ministerial approval (fractional) error correction mechanism. The size of the ECM coefficient indicates that although these two variables travel in tandem in the long run, shocks caused by economic or political interventions can

[55] Since prime ministerial approval enters the Conservative support model without a lag, the possibility of simultaneity bias is investigated. Test results for both the two-step and one-step estimates of the party support models are consistent with the hypothesis that prime ministerial approval is weakly exogeneous to Conservative vote intentions and, hence, it is possible to perform single-equation estimation of the parameters in the Conservative support model. For example, for the two-step estimates, adding the ECM to a model of the determinants of prime ministerial approval (personal expectations variant) shows that the ECM is statistically insignificant ($t = 1.096$). Adding the residuals of the prime ministerial model to the Conservative support model (personal expectations variant) also yields an insignificant coefficient ($t = 1.295$). On weak exogeneity tests in the context of error-correction models, see Charemza and Deadman, *New Directions in Econometric Practice*, chap. 7. Details of the prime ministerial approval models are available upon request.

drive them apart for lengthy periods of time. Parameter estimates also reveal that evaluations of personal economic conditions are significant predictors of governing party support, but evaluations of the national economy are not significant.

The fractional error correction models thus help to resolve two longstanding controversies in the literature on party support in Britain. *Pace* the (perhaps mythical) era when social class was everything and all else was embellishment and detail,[56] these models forcefully testify that party leader images have been important elements in the calculus of party support in Britain since at least the late 1970s. Although the looseness of the (fractional) cointegrating relationship between prime ministerial approval and governing party support indicates that various economic and political interventions can have sizeable short-term effects, it is equally clear that leader images matter a great deal. Model estimates also testify that personal, not national, judgements count when the *direct* effects of economic evaluations on party support are analysed.[57] However, this latter finding is not an unequivocal endorsement of the widely cited Essex model. Encompassing tests pitting personal expectations against personal retrospections are inconclusive, so therefore we caution that additional inquiry is in order.

When researchers conduct future studies of the dynamics of party support, fractional (co)integration methods will prove helpful. As demonstrated in this article, these methods are easily implemented in the context of generalized versions of familiar ARIMA and error correction models, and they enable analysts to avoid the methodologically contentious and often theoretically problematic assumption that important political and economic variables evolve through time with integer orders of integration. The concepts of fractional integration and fractional cointegration also have important heuristic value in that they encourage researchers to offer conjectures concerning micro-level processes responsible for fractional aggregate-level phenomena. Individual-level heterogeneity is a likely source of aggregate-level fractional integration. Specifying such heterogeneity in parsimonious, theoretically attractive, models of party support poses an interesting and important challenge to political economists in Britain and elsewhere.

[56] Peter Pulzer, *Political Representation and Elections in Britain* (London: Allen & Unwin, 1967), p. 98.

[57] There is evidence that retrospective judgements regarding the national economy have indirect effects on party support via their influence on prime ministerial approval. Encompassing tests applied to models of prime ministerial approval developed to determine if prime ministerial approval is weakly exogeneous to governing party support show that a model incorporating national retrospections encompasses models using other types of national and personal economic evaluations. Details available upon request. See also Harold D. Clarke, Karl Ho and Marianne C. Stewart, 'Major's Lesser (Not Minor) Effects: Prime Ministerial Approval and Governing Party Support in Britain since 1979', *Electoral Studies*, 19 (2000), 255–74.