Notes and Comments

Trends in Social Capital: Membership of Associations in Great Britain, 1991–98

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This Note uses the British Household Panel Study (BHPS) to consider the changing volume and distribution of voluntary association membership (and hence social capital) in Great Britain. We aim to supplement Hall's study of trends in social capital published in this *Journal* with more recent and longitudinal data.[1] This allows us to show that whilst the volume of social capital is not declining, it is becoming increasingly class specific, and that its relative aggregate stability masks considerable turnover at the individual level. These findings are significant for current debates on social capital.

MEMBERSHIP OF ASSOCIATIONS

The BHPS was designed as an annual survey of each adult (16 +) member of a nationally representative sample of more than 5,000 households, making a total of approximately 10,000 individual interviewees.[2] The same individuals were re-interviewed in successive waves and, if some had become separated from their original households, all adult members of their new households were also interviewed. Children were interviewed once they reached the age of 16; there was also a special survey of 11–15 year old household members from Wave 4 onwards. Thus the sample was designed to remain broadly representative of the population of Great Britain as it changed through the 1990s.

Membership and Activism

Table 1 shows that levels of involvement in different types of association vary considerably. However, only 'other civic' and 'other' groups recorded fewer members in 1995 than in 1991, partly a statistical artefact arising from the recording of an additional set of organizations after 1993. Thus, total social capital (in Putnam's sense, as measured by associational membership) appeared to increase markedly between 1991 and 1995.[3] Moreover, there were appreciable increases in

[1] Peter A. Hall, 'Social Capital in Britain', *British Journal of Political Science,* 29 (1999), 417–61.

[2] M. Taylor with J. Brice, N. Buck and E. Prentice-Lane (eds), *British Household Panel Survey User Manual, Volume A & B* (Colchester: University of Essex, 1999).

[3] Robert Putnam, 'Tuning In and Tuning Out: The Strange Disappearance of Social Capital in America', *Political Science and Politics* (December 1995), 664–83.

TABLE 1   *Frequencies of Associational Membership and Activism, Waves 1–5 and 7 (Percentages)**

		W1	W2	W3	W4	W5	W7
Political party	Member	3.1	3.4	3.5	3.4	3.1	2.8
	Active	1.4	1.5	1.8	1.6	1.5	1.5
Trade unions	Member	17.5	17.2	17.2	16.6	17.7	15.9
	Active	3.1	3.0	3.7	3.9	4.3	3.7
Environment	Member	2.9	3.7	4.0	4.3	4.2	3.5
	Active	1.5	1.8	2.1	2.4	2.3	1.8
Parent/School Association	Member	3.6	4.4	4.7	4.5	4.3	3.3
	Active	5.0	5.1	5.8	5.8	5.8	4.9
Tenants/Residents	Member	6.8	7.0	9.3	10.0	10.3	8.9
	Active	3.1	3.1	4.0	5.0	5.3	4.3
Religious Group	Member	10.1	12.2	12.7	12.6	12.7	11.1
	Active	10.1	12.2	12.7	12.7	13.1	11.3
Voluntary Service	Member	3.5	4.4	4.4	4.7	5.1	4.0
	Active	3.8	4.3	4.4	4.9	5.0	4.1
Other civic group	Member	3.5	3.2	2.9	1.8	1.7	1.5
	Active	3.4	3.1	2.8	1.7	1.7	1.6
Social/Working Men's Club	Member	12.5	12.9	13.2	13.2	13.5	11.3
	Active	8.8	9.4	10.2	10.0	10.4	9.2
Sports Club	Member	16.1	16.4	18.0	18.3	18.5	18.1
	Active	17.1	16.9	19.5	19.4	19.7	19.4
Women's Institute	Member	1.9	1.9	1.9	1.6	1.6	1.6
	Active	1.9	1.8	2.0	1.6	1.8	1.6
Women/Feminist org.	Member	0.8	1.1	1.1	1.1	1.2	1.0
	Active	0.8	1.0	1.0	1.1	1.2	1.0
Professional	Member			1.4	2.4	2.0	8.7
	Active			0.8	1.5	1.4	3.6
Pensioner	Member			0.5	0.7	0.7	2.1
	Active			0.3	0.7	0.7	2.1
Scouts	Member			1.1	0.9	1.1	1.6
	Active			1.1	0.8	1.1	2.3
Other group	Member	10.1	10.6	9.3	7.8	7.3	6.5
	Active	8.8	8.8	7.8	6.5	6.1	5.8

* Data is not available for Wave 6 as no information was collected on associational membership in that wave.

membership (and activism) in environmental, religious, residents, voluntary, social, sports and women's groups. The increase in environmental and women's groups is particularly sharp, though from a low base.

Table 1 also records a fall in aggregate levels of membership between Wave 5 (1995) and Wave 7 (1997) to the point where levels of memberships were similar to that in 1991. This is confirmed by Table 2, which examines the proportion of the population who were in at least one association. This appears to be other than a statistical artefact, though a change in the nature of the data collection makes it difficult to interpret. Generally, people who remained within the sample are more likely

TABLE 2   *Respondents Claiming to be a Member of at least One Organization, All Waves, Including Longitudinal Weights (Percentages)*

W1	W2	W3	W4	W5	W7
54.3	55.0	56.4	57.4	57.1	55.9

Note: These percentages deal with all proxy responses – usually between 3 and 4 per cent of all responses – by assuming that the person has no memberships, thus offering a conservative estimate of the proportions with any associational memberships. Data on Wave 6 is not available.

to be members of associations than those who are newly recruited, although the use of longitudinal weights on the panel data should satisfactorily compensate for this.

The overall pattern indicates a gradual but significant increase between 1991 and 1995 and then a fairly rapid loss of membership in the years 1995–97. It is not easy to explain this change and we must be circumspect in our interpretation. It could be that the (impending) election of a Labour government, with some expectation of political change, led to decline in other forms of participation, but that seems unlikely.

Multiple Membership

Table 3 shows the number of types of association of which respondents indicated membership between 1991 and 1997. The proportion that were members of no association fell by 2 or 3 percentage points between 1991 and 1995, from 42 to 39, but then increased to almost 45 per cent in 1997. This might perturb a follower of Putnam. The proportion without affiliation in the 1990s is considerably greater than that recorded by Parry *et al.* for the years 1984–85, which indicated that only 27 per cent of the British population had no affiliation.[4]

Multiple memberships are comparatively small. Between 20 and 25 per cent of people claimed

TABLE 3   *Number of Types of Organization of which Respondents Claimed to be Members, All Waves, Longitudinally Weighted (Percentages)*

	W1	W2	W3	W4	W5	W7
None	42.3	41.1	38.8	39.7	40.0	45.0
1	30.6	30.0	30.6	31.0	30.3	30.3
2	15.4	15.7	15.6	15.6	16.0	14.1
3	5.8	6.4	6.2	6.4	6.3	5.6
4	1.7	1.9	2.3	2.1	2.2	1.7
5	0.5	0.6	0.6	0.8	0.9	0.6
6	0.2	0.2	0.3	0.2	0.2	0.1
7	0.0	0.1	0.1	0.1	0.1	0.0
8	0.0	0.0	0.0	0.0	0.0	0.0

Note: Because of the nature of the derived variable, the number of respondents recorded as being without any membership do not accord exactly with the data in Table 2. No data is available for Wave 6.

[4] Geraint Parry, George Moyser and Neil Day, *Political Participation and Democracy in Britain* (Cambridge: Cambridge University Press, 1992), p. 91.

to be members of two or more organizations in the early 1990s and less than 4 per cent were members of four or more. These data show fewer memberships than previous studies have done. However, it is not possible to make a direct comparison with the evidence in either Goldthorpe *et al*. or Parry *et al*. because they asked how many *organizations* people were members of, whereas the BHPS records *categories* or types of association.[5] The maximum that can be indicated is one of each type.

Where comparison is possible, a marked change between 1984 and the 1990s is suggested. The most significant is that in 1984 only 27 per cent of respondents had no memberships and 73 per cent had one or none. The comparable BHPS figures in 1991 were 42 per cent and 55 per cent. Some of the discrepancy might be attributed to a fall in party and trade-union membership – from 6.8 and 27 per cent in 1984 to 3.1 and 18 per cent in 1991 respectively. Moreover, Parry *et al*.'s questionnaire explicitly named several types of organization absent from BHPS including Hobby Club, Armed Forces Association, Evening class or study group, 'Art, literary or cultural group' and 'Self help group'.[6] The naming of such would substantially increase the rate of recall – as was demonstrated within the BHPS when it explicitly listed some types that in earlier waves had appeared as 'Other groups'. Indeed, these additional categories in the Parry *et al*. study accounted for 705 out of a total of 1,750 recorded memberships of voluntary groups.[7] The potential error arising could easily account for the entire discrepancy and allow us even to conclude that membership of voluntary groups had increased to compensate for losses in party and union membership.

Changing Volume of Associational Social Capital

Best evidence, then, suggests that the total volume of associational capital circulating in Britain was fairly constant with a small increase in the early 1990s, which fell back in 1997 to the level of 1991. Levels in the 1990s were not much different from those of previous decades.[8] However, one significant trend between 1991 and 1997 concerns gender. The number of women who said that they were members of an organization increased by 5 percentage points between Wave 1 and Wave 5, the corresponding figure for men being just over 1 percentage point.[9] There was then a sharper fall for women between Waves 5 and 7. It might then seem that the explanation of changing participation in the 1990s was mostly the result of the changing behaviour of women, though we have no obvious explanation of this.

SOCIO-DEMOGRAPHIC PROFILES OF PEOPLE WHO JOIN ASSOCIATIONS

Our analysis confirms that there is considerable variation in the likelihood of people with different socio-demographic characteristics joining associations.[10] Gender, ethnicity, age, class, education, income and having children in the household were all significantly associated with the claim to be a member of any (one or more) of the categories of organization mentioned in the BHPS. Because of potential problems of colinearity between age, education and income, we ran separate models with each attempting to examine more closely the effects of these variables on propensity to join. We applied logistic regression analysis to explore the effects of socio-demographic characteristics. Overall, the results were in line with previous expectations and a typical model, applied to

[5] J. H. Goldthorpe, with C. Llewellyn and C. Payne, *Social Mobility and Class Structure in Modern Britain*, 2nd edn (Oxford: Clarendon Press, 1987); Parry *et al*., *Political Participation and Democracy in Britain*.

[6] Parry *et al*., *Political Participation and Democracy in Britain,* p. 456.

[7] Parry *et al*., *Political Participation and Democracy in Britain*, p. 93.

[8] Hall, 'Social Capital in Britain'.

[9] Respondents reporting membership of one or more organizations, by sex: W1, men 63.5%, women 50.0%; W2, men 63.4%, women 52.0%; W3, men 65.3%, women 54.9%; W4, men 64.9%, women 55.5%; W5, men 64.8%, women 55.0%; W7, men 63.0%, women 52.3%.

[10] Goldthorpe *et al*., *Social Mobility and Class Structure in Modern Britain;* Parry *et al*., *Political Participation and Democracy in Britain.*

TABLE 4   *Propensity to Join One or More Associations, by Socio-Demographic Variables, Wave 1, Logistic Regression*

Parameter	Estimate	Sig.	Odds ratio
White	0.444	0.000	1.559
Female	− 0.607	0.000	0.545
Child 0–4 years of age	− 0.361	0.000	0.697
Child 5–11 years of age	0.112	0.062	1.119
Class: Unskilled & semi-skilled (reference)			
Skilled manual	0.205	0.010	1.228
Foreman, technician	0.359	0.000	1.432
Small propietors, farmers, smallholders	0.050	0.560	1.051
Routine non-manual, professional	0.523	0.000	1.686
Service	0.994	0.000	2.702
Age: Under 20 years (reference)			
20–40	0.481	0.000	1.617
40–60	0.708	0.000	2.030
Over 60 years	0.213	0.016	1.238
Intercept	− 0.686	0.000	0.503

Note: − 2 loglikelihood = 12,583.753; Cox and Snell $R^2 = 0.069$; Nagelkerke $R^2 = 0.093$.

respondents in Wave 1, is reported in Table 4, exploring the impact of gender, ethnicity, presence of children in the household, age and class.

This table indicates the differences in odds of persons with various characteristics, controlled for each of the other variables, becoming a member of at least one association. If the odds ratio is greater than one, a person with that feature is more likely to be a joiner, if it is less than one then they have a proportionate disinclination to join. Thus, the odds of a respondent being a joiner were increased by 56 per cent if the respondent was white. Being female or having a child under 5 reduced the odds. Class was significant and there was a monotonic relationship, the higher the class, the more likely the respondent was a joiner, with the marked exception of those who were self-employed. As in many other models, the petty bourgeoisie shows reluctance to join. A member of the service class was 2.7 times more likely than an unskilled manual worker to be a joiner. The relationship with age shows an increase, as people get older, with a peak in the age group 40–60. Those aged 40–60 are twice as likely as those under 20 to be a joiner. The ratio is somewhat less for the oldest age group, those over 60.

Other models of this kind, which included class and education and class and income showed independent and significant associations with the propensity to be a joiner.

We also examined whether there were any significant socio-demographic predictors of the tendency to join more than one organization. Being male, being white, having more education, being of a higher social class, having greater personal income and having more educational qualifications all significantly increased the likelihood of membership of more organizations.

The Distribution of Social Capital: The Decline of Class?

To determine whether there was any change in the socio-demographic basis of involvement between 1991 and 1997, we examined the impact of gender, age, class, income, education and household composition on the propensity to join at least one type of organization, in each wave both separately and in a combined model using logistic regression. All variables proved significant in all waves.

The weighted panel data showed one intriguing tendency for change over time, increased
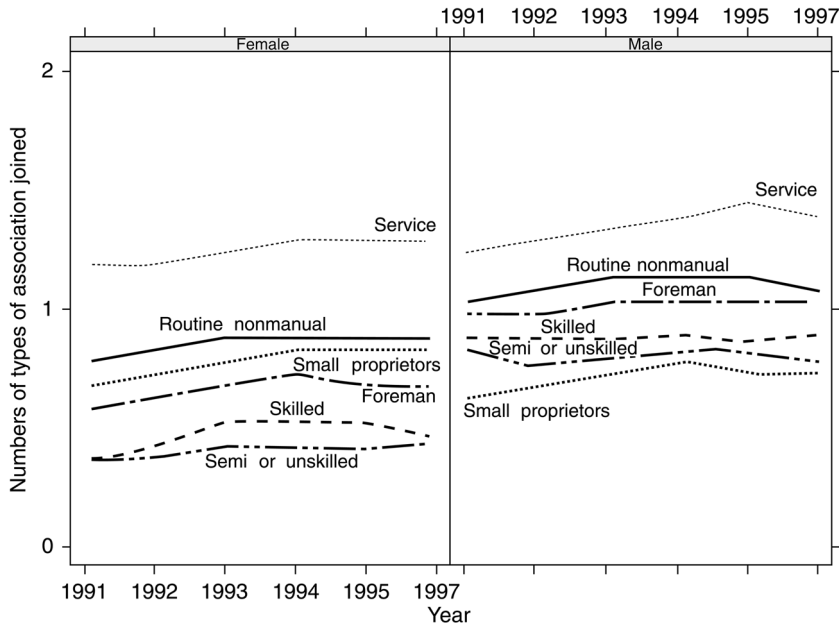
Fig. 1. Mean number of associational memberships claimed by respondents of different classes, women and men, 1991–97

differences in levels of participation between the service class and all other classes. While class has a statistically significant effect in every wave, the basis of that effect changes over time. Figure 1 shows, for men and women separately, that the tendency for the service class to participate disproportionately appeared to increase during the 1990s.[11] This tendency is slightly more pronounced for women, with the main difference between women and men being that among the self-employed the former participate more. Regression analysis on different waves supports this interpretation (see Table 5). In Waves 1 and 3 all classes except the petty bourgeoisie were more involved than the unskilled and semi-skilled. In Wave 5, the skilled manual workers and routine white-collar workers' behaviour was not significantly different from that of the unskilled and semi-skilled. However, there were still differences *vis-à-vis* supervisory workers and technicians and respondents in the service class. By Wave 7, the only statistically significant difference was between the service class and the rest. So the importance of class remains, but the main differences in behaviour come to be detectable only at the boundary between the service class and the remainder of the population.[12] Differences between all other classes, i.e. manual, routine white-collar and petty bourgeois, were gradually eroded between 1991 and 1997. [13]

Comparison with the findings of Parry *et al.* suggests some new features. Parry found disproportionate involvement by the service class, but engagement among the manual working class

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[11] The inter-class pattern was the same when men and women were considered together.

[12] We were able to confirm this finding also when we analysed the unweighted samples.

[13] BHPS data present some technical problems arising not only from item non-response and unit non-response, but also from attrition from the panel. These issues are typically ignored or dealt with in an *ad hoc* manner in the literature. Additionally, in this instance, the response variable, membership in voluntary associations, is not normally distributed. With a dataset containing over 10,000 individuals size is a non-trivial problem for transformation using logit and analysis using the generalized linear mixed model. Acknowledging these difficulties, which we are exploring within the Bayesian framework using Markov Chain Monte Carlo, we claim only that the evidence in Table 5 is supportive of our conclusions about class and gender effects.

TABLE 5   *The Changing Influence of Gender and Social Class on Organizational Membership, Waves 1, 3, 5 and 7 (Longitudinally Weighted)*

	Wave 1			Wave 3			Wave 5			Wave 7		
		12,737			5,828.310			5,499.035			5,494.997	
−2 log likelihood												
Cox and Snell R^2		0.054			0.041			0.034			0.048	
Nagelkerke R^2		0.073			0.056			0.047			0.065	
Parameter	Estimate	Sig.	Odds ratio	Estimate	Sig.	Odds ratio	Estimate	Sig.	Odds ratio	Estimate	Sig.	Odds ratio
Female	−0.589	0.000	0.555	−0.391	0.000	0.677	−0.325	0.000	0.723	−0.433	0.000	0.648
Class: Unskilled (ref.)												
Skilled manual	0.188	0.017	1.207	0.467	0.073	1.595	0.249	0.073	1.282	0.137	0.338	1.140
Foreman	0.386	0.000	1.471	0.426	0.001	1.531	0.480	0.001	1.616	0.103	0.459	1.109
Small propietors	0.112	0.187	1.118	0.136	0.001	1.145	−0.077	0.527	0.926	0.092	0.469	1.097
Routine non-manual	0.529	0.000	1.697	0.412	0.255	1.510	0.133	0.178	1.142	0.144	0.154	1.155
Service	1.041	0.000	2.832	1.035	0.000	2.814	0.821	0.000	2.274	0.952	0.000	2.590
Intercept	0.134	0.000	1.144	0.396	0.000	1.486	0.473	0.000	1.605	0.330	0.000	1.391

in proportion to their presence in the population, largely sustained by membership of trade unions and the Labour party.[14] In the mid-1980s it was the intermediate classes (and most especially the petty bourgeoisie) who were under-represented among activists. The position was slightly different among the 1991 BHPS sample, with intermediates other than the self-employed more active than skilled or unskilled manual workers. The trends in the later 1990s increase the plausibility of Hall's concern that the working class might become increasingly marginalized.[15] The social and political influence of the service class appears to have been consolidated. Since associational membership not only increases political participation, and hence political influence, but also delivers benefits of sociability including instrumental networks, we have evidence of a further polarization in British society where the most advantaged class is increasing its privileges relative to the rest. The decline of trade-union membership may be one of the key reasons for declining involvement of both routine white-collar and manual workers in formal associations. Since trade unions had at least a partly political function, this probably reduces the impact of the working class on contemporary politics. Also, the fact that union membership and membership of professional associations have proved most stable among members of the service class further skews the pattern of influence. Given that the other type of association which particularly attracts the working class – social clubs – are one of the least likely to entail discussions about politics, the political impact of the working class may be further reduced.[16]

POLITICAL CAREERS

Parry *et al.* speculate about the future of voluntary group membership and its effect on political participation.[17] They emphasize the importance of membership for activism: 'affiliation to groups has been shown to be one of the best predictors of a propensity to participation'.[18] They argue, in support of Barnes *et al.*, that 'since actual political behaviour is contingent on political events, it is more important to study the readiness of people to be mobilised'.[19] BHPS panel data permit analysis of careers, separating out 'stocks' and 'flows' of social capital. Estimates of the volume of social capital circulating in any one year can fruitfully be complemented by an appreciation of the levels of volatility in memberships. We inquire into volatility, as a means to explore potential for 'sporadic intervention',[20] and then into what factors encourage respondents, year on year, to join, leave or remain a member of an association.

    We examined movement in and out of different associations over the five years covered by Waves 1 to 5. An example of the pattern is offered in the first row of Table 6, which considers joining and leaving a political party during the years 1991–1995. This shows that while the percentage of members at any one time was never greater than 3.5 (in Wave 3), some 6.8 per cent of panel respondents were at one time or another a party member. This more than doubles the numbers who have shown a preparedness to join a party and gives a more optimistic impression of the extent of party membership. The data also show that only 1.6 per cent of people were members continuously in all five years, indicating a significant degree of volatility of party membership.

[14] Parry *et al.*, *Political Participation and Democracy in Britain.*
[15] Hall, 'Social Capital in Britain', p. 455.
[16] See Parry *et al.*, *Political Participation and Democracy in Britain,* pp. 100, 103.
[17] Parry *et al.*, *Political Participation and Democracy in Britain*, pp. 422–4.
[18] Parry *et al.*, *Political Participation and Democracy in Britain,* p. 419.
[19] S. Barnes, M. Kaase *et al., Political Action: Mass Participation in Five Western Democracies* (London: Sage, 1979); Parry *et al., Political Participation and Democracy in Britain.* p. 423.
[20] See Robert Dowse and John Hughes, 'Sporadic Interventionists', *Political Studies*, 25 (1977), 84–92.

TABLE 6   *Turnover of Membership of Organizations, Waves 1–5 (1991–95), Percentages*

	Member throughout	Member in at least one year	Maximum members in any year
Political party	1.6	6.8	3.5
Trade union	9.3	28.1	17.7
Environmental group	1.4	8.5	4.3
Parents association	1.5	10.3	4.7
Tenants association	2.0	21.7	10.3
Religious group	7.5	19.9	12.7
Voluntary service	1.0	12.1	5.1
Other civic	0.1	9.8	3.5
Social club	4.4	26.4	13.5
Sports club	6.0	32.1	18.5
Women's Institute	1.2	3.5	1.9
Women's group	0.2	3.4	1.2

TABLE 7   *Patterns of Joining and Leaving a Political Party, Waves 1–5*

Pattern	Frequency	Valid percentage
OOOOO	6,190	93.2
OOOOX	33	0.5
OOOXO	24	0.4
OOOXX	16	0.2
OOXOO	33	0.5
OOXOX	7	0.1
OOXXO	10	0.2
OOXXX	17	0.3
OXOOO	31	0.5
OXOOX	4	0.1
OXOXO	3	0.0
OXOXX	8	0.1
OXXOO	8	0.1
OXXOX	1	0.0
OXXXO	8	0.1
OXXXX	18	0.3
XOOOO	33	0.5
XOOOX	3	0.0
XOOXO	5	0.1
XOOXX	4	0.1
XOXOO	6	0.1
XOXOX	1	0.0
XOXXO	2	0.0
XOXXX	10	0.2
XXOOO	14	0.2
XXOOX	1	0.0
XXOXO	8	0.1
XXOXX	5	0.1
XXXOO	12	0.2
XXXOX	6	0.1
XXXXO	16	0.2
XXXXX	103	1.6
Total valid	6,640	100.0
Missing	3,624	
Total	10,264	

Table 7 indicates the pattern of flows in and out of party membership during the period. An 'X' indicates that a respondent was a member in a particular year, an 'O' that s/he was not. Thus the pattern 'OOOOO' indicates that the respondent was never a member, 'XXOOO' that s/he was a member in the first two waves but had departed never to return by Wave 3, while 'XOXOX' shows someone whose membership lapsed before Wave 2 and Wave 4 but who rejoined in the intervening years. The most common condition is never to be a member, the next most frequent to be a member throughout.

As Table 8 shows, year on year, about 2.5 per cent of the sample renewed a political party membership, while about half that proportion either joined or ceased to be members. This is a significant turnover rate. As Table 7 indicates, these were not necessarily people who let their membership lapse for a year – perhaps forgetting to pay a subscription – as there were many different patterns of arrival and departure. The proportion that simply ceased to be members, i.e. who were members in Wave 1 but never again, was 0.5 per cent of the sample; a similar proportion left after Wave 1 but re-joined.

TABLE 8   *Turnover of Party Membership, Year on Year, Waves 1–5 (Percentages)*

	Wave 1–2		Wave 2–3		Wave 3–4		Wave 4–5	
	n	*%*	*n*	*%*	*n*	*%*	*n*	*%*
Stayed both waves	165	2.5	172	2.6	184	2.8	181	2.7
Joined by later wave	81	1.2	86	1.3	73	1.1	56	0.8
Left by later wave	64	1.0	74	1.1	74	1.1	76	1.1

The results of analysis of turnover for all the types of organization mentioned in the questionnaire and patterns are summarized in Table 6 above. The proportion of people who were members in at least one of five years outweighed those present in any given year by between 1.5 and 3 times. Thus for example, 28.1 per cent of people held trade-union membership during the period, but even in the best year showed an aggregate membership of only 17.7 per cent (a ratio of 1.5 : 1). For environmental groups the ratio was about 2 : 1, for women's groups about 3 : 1.

Also with respect to all organizations, a much larger proportion of the population were likely to join at some point during the five years than were likely to remain a member throughout. The ratio for parties was approximately 4 : 1 (i.e. 6.8 per cent to 1.6 per cent) that for unions 3 : 1, for environmental groups 6 : 1, for women's groups 17 : 1. A similar range can be seen among the civic and religious groups with tenants and residents associations exhibiting a ratio of 11 : 1, religious groups 2.5 : 1. Given that our measure of membership identifies only categories of association (i.e. the response is to the question: are you a member of any, for example, religious group?), which makes it impossible to determine that continued membership means attachment to the same church, or even a church of the same denomination, this level of volatility is remarkable. Trade unions, religious groups and sports clubs were the only types of association to command continued membership greater than 5 per cent over a 4–5 year period.

Such levels of turnover of membership of political parties, associations and movements indicate patterns of behaviour that have been little remarked upon. On one hand, this might be seen as evidence of low levels of attachment to associations and an almost flighty attitude towards activism. On the other hand, it might signify that people reflect carefully on their attachments at regular intervals in order better to direct their energies towards currently important and pressing activities.[21]

Inspection of the careers of individuals over time conveys a significantly different impression of the volume of social capital in circulation. Another instance concerns the likelihood of being a member of any association. While in any one year at least two out of five persons claim to have no memberships, only 15.6 per cent never held any during the 1990s. Comparatively few people

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[21] Data collected in later phases of our study will enable us to comment further on this issue.

have never experienced the supposed benefits of associational membership. Though many are not active at a given time, they are probably neither hostile to membership nor unavailable for mobilization through organizations should an appropriate occasion arise. Secondly, the weak ties developed during a bout of membership probably will not immediately erode, such that individuals will be better connected than might be imagined from considering only the total volume at any one time.

CONCLUSIONS

Great Britain, unlike the United States as described by Putnam, is not undergoing a decline in associational social capital. The trends in the 1990s suggest continued stability, even some increase in voluntary group membership, when compared to the data collected by Hall for the previous four decades. There is, however, an increasingly unequal distribution of membership, and by extension, of political participation. There is some increase in women's participation, but there is a growing divide between the service class and all other classes. Any benefits of social capital are increasingly going to professional and managerial workers relative to other social class groups. Other studies show that formal associational membership is a good predictor of political involvement and that therefore possession of this type of social capital is likely to increase the influence of its holders. Thus the tendency for it to become relatively more concentrated in the service class, and its continuing to be disproportionately the province of men, raises concerns about the democratic character of Britain.

An optimistic interpretation of participation is, however, much enhanced by the analysis of the panel aspect of the BHPS data. In the light of the large numbers of people who move in and out of organizations from year to year, Britain appears to be a more participatory and active society than might otherwise be imagined. Very few people were never members of any association in the 1990s, a mere 15.6 per cent. Most associations, including political parties and social movements, had people move in and out of membership at a substantial and steady rate. Associational and political experience touches many more people over, say, a five-year period than cross-sectional data would suggest. This suggests both that a great many people, and the community as a whole, obtain the benefits of membership at frequent intervals, and that a large proportion of the population might be considered available for mobilization even if they are not currently active. Of course, volatility might be a sign of limited or transient commitment. Also membership is by no means equivalent to regular active involvement. Nevertheless, the picture is probably more positive once we take account of people's careers.

Electoral Institutions, Unemployment and Extreme Right Parties: A Correction

MATT GOLDER*

In their 1996 article in this *Journal*, Robert Jackman and Karin Volpert analyse the systematic conditions that influence the electoral success of extreme right parties in sixteen West European

countries from 1970 to 1990.[1] In particular, they focus on the effects of unemployment, electoral thresholds and multi-partism. By specifying and estimating a Tobit regression model, they conclude that:

(1) higher rates of unemployment increase the electoral support of extreme right parties;
(2) increasing electoral thresholds dampen support for the extreme right as the number of parliamentary parties expands;
(3) multi-partism increasingly fosters parties of the extreme right with rising electoral proportionality.

They argue that these conclusions are encouraging since unemployment, electoral thresholds and multi-partism can be adjusted through policy intervention (pp. 517–18).[2] Without questioning the theoretical justification of their model or changing any of the statistical techniques that they use, I illustrate that their second and third conclusions are open to question. A critical re-analysis of their model's theoretical justification suggests that their first conclusion may also be misleading.

There are four sections to this research note. The first very briefly outlines Jackman and Volpert's model. I then note some of the important methodological and theoretical contributions that their article makes for the study of extreme right parties. In the third section, I replicate and interpret their results. I show that only their conclusion about unemployment is actually supported by their analysis. Finally, I raise further concerns about Jackman and Volpert's conclusions by calling the theoretical justification of their model into question.

THE MODEL

I focus on Jackman and Volpert's most preferred model specification. This is given below:

$$\ln ERPS = \beta_0 + \beta_1 THRESH + \beta_2 ENPP + \beta_3(THRESH \times ENPP) + \beta_4 UNEMP + \beta_5 COUNTRY_1$$
$$+ \dots \beta_{19} COUNTRY_{15} + \varepsilon \tag{1}$$

where Greek letters are parameters to be estimated; $\ln ERPS$ is the natural log of the voteshare received by extreme right parties;[3] $THRESH$ is the effective electoral threshold;[4] $ENPP$ is the effective number

of parliamentary parties;[5] *UNEMP* is the level of unemployment at the national level; and *COUNTRY* is simply a country dummy variable.

Jackman and Volpert argue that higher levels of unemployment provide a favourable environment for extreme right parties. Since extreme right parties focus their attacks on immigrants and foreign workers, they expect that their campaigns are more successful 'when jobs are scarce than when they are plentiful' (p. 507). The implication is that extreme right voters blame the visible immigrant population for job losses. They also predict that higher electoral thresholds 'dampen the prospects for smaller parties of the extreme right' (p. 506). This is because voters do not want to waste their vote.[6] Their decision to include an interaction term between the effective threshold and the effective number of parties is based on Lijphart's conclusion that disproportionality and multi-partism are 'loosely inter-dependent'.[7] They anticipate that 'the dampening effect of disproportionality on extreme right support will increase with multi-partism, while the positive effect of multi-partism diminishes with rising disproportionality' (p. 507).

THEORETICAL AND METHODOLOGICAL CONTRIBUTIONS

Jackman and Volpert make several fruitful contributions to the study of extreme right parties. The first is their attempt to test hypotheses about extreme right parties through a cross-national statistical analysis. Much of the literature on extreme right parties is dominated by single and small-*N* case studies. Although this research is very helpful in generating hypotheses, it is less useful for hypothesis testing. This is because it is difficult to draw valid causal inferences from this type of work.[8] This shift towards the systematic testing of general hypotheses should be applauded. A second contribution is their emphasis on the institutional constraints posed by electoral systems. There are historical narratives that emphasize how changes to electoral laws have influenced the success of extreme right parties. There is also some more analytical work investigating the relationship between the type of electoral system and the success of these parties.[9] On the whole, though, the institutional constraints posed by electoral systems have been relatively understudied compared to the roles played by other factors such as unemployment and immigration. This issomewhat surprising given the enormous amount of interesting and successful work on electoral rules that exists in the party system literature more generally.[10]

(F'note continued)

(50%)/(M + 1) + (50%)/(2M), where *M* is the district magnitude. See Arendt Lijphart, *Electoral Systems and Party Systems: A Study of Twenty-Seven Democracies, 1945–1990* (Oxford: Oxford University Press, 1994), p. 27. For more information on electoral thresholds, see Rein Taagepera, 'Effective Magnitude and Effective Threshold', *Electoral Studies*, 17 (1998), 393–404; Rein Taagepera, 'Nationwide Inclusion and Exclusion Thresholds of Representation', *Electoral Studies*, 17 (1998), 405–17; Rein Taagepera, 'Nationwide Threshold of Representation'(unpublished paper, Department of Political Science, University of California, Irvine).

[5] The effective number of parties is calculated as $1/\Sigma s_i^2$, where s_i is the percentage of seats won by the *i*th party. This is the reciprocal of the Hirschman–Herfindahl index used in economics to measure industrial concentration. See Markku Laakso and Rein Taagepera, 'Effective Number of Parties: A Measure with Application to Western Europe', *Comparative Political Studies*, 12 (1979), 3–27.

[6] Maurice Duverger, *Political Parties: Their Organization and Activity in the Modern State* (New York: Wiley, 1963).

[7] Lijphart, *Electoral Systems and Party Systems*, p. 77.

[8] See James Fearon, 'Counterfactuals and Hypothesis Testing in Political Science', *World Politics*, 43 (1991), 169–95; Gary King, Robert Keohane and Sidney Verba, *Designing Social Inquiry: Scientific Inference in Qualitative Research* (Princeton, NJ: Princeton University Press, 1994); Stanley Lieberson, 'Small *N*'s and Big Conclusions: An Examination of the Reasoning in Comparative Studies Based on a Small Number of Cases', *Social Forces*, 70 (1991), 307–20.

[9] See Herbert Kitschelt, *The Radical Right in Western Europe: A Comparative Analysis* (Ann Arbor: University of Michigan Press, 1997).

[10] See William Riker, 'The Two-Party System and Duverger's Law: An Essay on the History of Political Science', *American Political Science Review*, 76 (1982), 753–66; Rein Taagepera and Matthew Shugart, *Seats and Votes: The Effects and Determinants of Electoral Systems* (New Haven, Conn.: Yale University Press, 1989);

Jackman and Volpert's most important contribution is in raising the issue of selection bias in empirical analyses of extreme right parties (p. 513).[11] They note that it is problematic to analyse the factors that influence the electoral success of extreme right parties when these parties do not exist in all countries. This is because there is good reason to believe that these factors might also be systematically related to whether an extreme right party exists in the first place. They recognize that their dependent variable (the electoral support for extreme right parties) is left-censored at 0 because it cannot be observed in those countries where there is no organized extreme right party. In fact, thirty-five out of 103 observations in their sample are censored at 0. In these circumstances it would be wrong to drop all the censored observations since this leads to biased and inconsistent estimates. It be equally wrong to retain the censored observations and simply code them as 0.[12] In response, Jackman and Volpert use a Tobit model that utilizes a maximum likelihood estimator for (left- or right-) censored variables. As King notes, 'the result is a much more realistic model of the process generating censored data and may be interpreted as if from a linear Normal regression with no censoring.'[13] The estimated coefficients represent the marginal effect of the independent variables on the underlying support for extreme right parties.

RESULTS AND INTERPRETATION

Although these contributions are significant, a replication and critical reanalysis of their model indicates that their results and inferences are open to question. The first thing to note is that I was able to replicate their results with only minor difficulties.[14] This is somewhat of an achievement given how difficult (or impossible) it often is to replicate empirical research.[15] These results are shown in Model 1 in Table 1.[16]

The impact of unemployment can be interpreted directly from Model 1 in Table 1. Jackman and Volpert's conclusion that higher levels of unemployment increase the support of extreme right parties is clearly supported. However, the effect of electoral thresholds and the effective number of parties cannot be interpreted so easily. This is because the results of interaction models cannot

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(F'note continued)

Gary Cox, 'Centripetal and Centrifugal Incentives in Electoral Systems', *American Journal of Political Science*, 34 (1990), 903–35; Gary Cox, *Making Votes Count: Strategic Coordination in the World's Electoral Systems* (New York: Cambridge University Press, 1997); Gary Cox, 'Electoral Rules and Electoral Coordination', *Annual Review of Political Science*, 2 (1999), 145–61.

[11] The fact that selection bias can seriously damage one's ability to make valid causal inferences is increasingly well understood and accepted. See James Heckman, 'Sample Selection Bias as a Specification Error', *Econometrica*, 47 (1979), 153–61; Barbara Geddes, 'How the Cases you Choose Affect the Answers You Get: Selection Bias in Comparative Politics', in James Stimson, ed., *Political Analysis* (Ann Arbor: University of Michigan Press, 1990), pp. 131–50; Adam Przeworski and Fernando Limongi, 'Selection, Counterfactuals and Comparisons', (unpublished paper, University of Chicago, 1992); King, Keohane and Verba, *Designing Social Inquiry*.

[12] Paul Kennedy, *A Guide to Econometrics* (Cambridge, Mass.: MIT Press, 1998), p. 251.

[13] Gary King, *Unifying Political Methodology: The Likelihood Theory of Statistical Inference* (New York: Cambridge University Press, 1994), p. 210.

[14] The default settings for the maximum likelihood optimizers used by STATA did not converge. However, if the tolerance for the log-likelihood is changed to ltolerance (1e-5) then full convergence is achieved. Convergence is declared when the relative change in the log-likelihood from one iteration to the next is equal to ltolerance (·).

[15] William Dewald, Jerry Thursby and Richard Anderson, 'Replication in Empirical Economics: The *Journal of Money, Credit and Banking* Project', *American Economic Review*, 76 (1986), 587–603; Gary King, 'Party Platforms, Mandates, and Government Spending', *American Political Science Review*, 87 (1993), 744–80; Gary King, 'Replication, Replication', *PS: Political Science and Politics*, 28 (1995), 443–99.

[16] Model 1 is equivalent to Model 2 in Jackman and Volpert's article (p. 514). They actually calculate three models. I focus purely on Model 2 since this is their most preferred specification and it is the model from which their conclusions are drawn.

TABLE 1   *The Estimated Causal Effect of Electoral Institutions and Unemployment on Extreme Right Parties*

Regressor	Model 1	Model 2	Model 3
Threshold	0.276*	0.165*	− 0.027
	(0.059)	(0.058)	(0.019)
Effective number of parties	1.172*	0.837	—
	(0.382)	(0.455)	
Threshold × Effective number of parties	− 0.103*	− 0.065*	—
	(0.019)	(0.019)	
Unemployment	0.066*	0.091*	0.097*
	(0.016)	(0.019)	(0.021)
Country dummies	—	—	—
Constant	− 4.943*	− 3.212	0.403
	(1.480)	(1.666)	(0.752)
R^2	0.731	0.637	0.601
Standard error	0.373	0.458	0.499
Log likelihood	− 40.17	− 53.87	− 59.23
N	103	102	102

* Indicates $p < 0.05$.
Note: Standard errors given in parentheses.

be interpreted as in regular additive models since the coefficients are conditional.[17] For example, the marginal effect of *THRESH* on the dependent variable (ln*ERPS*) is determined by taking the derivative of Equation 1 with respect to *THRESH*. This is:

$$\frac{\partial \ln ERPS}{\partial THRESH} = \beta_1 + \beta_3 ENPP. \tag{2}$$

It is clear from this that the effect of *THRESH* depends on the value of *ENPP*. The coefficients for electoral thresholds and the effective number of parties that appear in Model 1 are only informative in the special case when the relevant modifying variable is actually 0. In the case of electoral thresholds, this means that an increase in the threshold helps extreme right parties when the effective number of parties is 0. Although some limited information can be learned from these figures, better inferences can be made when the full range of conditional coefficients and standard errors are estimated.

To some extent, Jackman and Volpert recognize this since they provide point estimates for the conditional coefficients at selected percentile values of the modifying variable. The problem is that it is difficult to interpret them since they do not provide the corresponding conditional standard errors. Nor do they provide the information necessary to calculate them. The standard error of *THRESH* conditional on *ENPP* is:

$$S_{\beta_1 + \beta_3 ENPP} = \sqrt{var(\beta_1) + var(\beta_3)ENPP^2 + 2ENPPcov[\beta_1, \beta_3]}. \tag{3}$$

The figures found in Table 1 of their article allow one to calculate the first two terms on the right-hand side of Equation 3. These are simply the squares of the standard errors on the relevant

[17] For a good discussion of interaction effects in linear and non-linear models see Robert Friedrich, 'In Defense of Multiplicative Terms in Multiple Regression Equations', *American Journal of Political Science*, 26 (1982), 797–833, and Jeff Gill, 'Interpreting Interactions and Interaction Hierarchies in Generalized Linear Models: Issues and Applications' (paper presented at the Annual Meeting of the American Political Science Association, San Francisco, 2001).
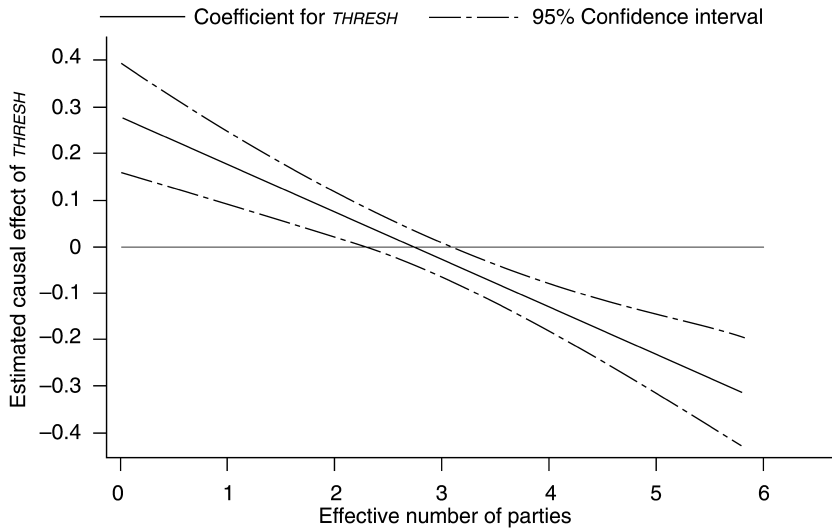
Fig 1. The Impact of effective thresholds (THRESH) on support for extreme right parties conditioned on the effective number of parties

variables. The problem is that the covariance in the last term cannot be derived from the information in their article. This means that it is impossible to know from their article whether their conclusions about electoral thresholds and multi-partism are valid.

Given access to their data, I was able to calculate the full range of conditional coefficients and standard errors during the replication process. These are shown graphically in Figures 1 and 2. The solid sloping lines indicate how the value of the estimated causal effect of THRESH or ENPP changes across the full range of the relevant modifying variable. For example, Figure 1 indicates that electoral thresholds have a positive effect on the electoral success of extreme right parties when the effective number of parties is low, but a negative effect when the effective number of parties is high. One can see whether these conditional coefficients are statistically significant by considering the 95 per cent confidence intervals (dashed lines) that are drawn around them. The coefficients are not significant when the lower bound of the confidence interval is below the zero line and the upper bound is above it. In other words, the estimated causal effect of THRESH or ENPP is indistinguishable from 0 at these points. The coefficients are only significant when the upper and lower bounds are both above or below the zero line.

Figure 1 shows that electoral thresholds have no effect on extreme right parties when the effective number of parties ranges between 2.3 and 3.1.[18] About one third of the observations fall in this range. The figure also implies that increases in electoral thresholds help extreme right parties when there are few parties. This is at odds with Jackman and Volpert's theoretical argument that raising electoral thresholds should decrease the electoral support of extreme right parties. If their argument were correct, then one would expect the conditional coefficient for electoral thresholds always to be negative. This is not the case here. Figure 2 illustrates that the effective number of parties has no effect on the support of extreme right parties when the effective threshold ranges between 4.7 per cent and 19.1 per cent. Over half of the observations fall in this range. Jackman and Volpert provide no theoretical reason as to why electoral thresholds and the effective number of parties might fail to have an effect on extreme right parties in these ranges. So far I have made no modification in the data, methods or model used in the original article.

[18] One can see this because the upper bound of the confidence interval is above the zero line in this range whereas the lower bound is not.
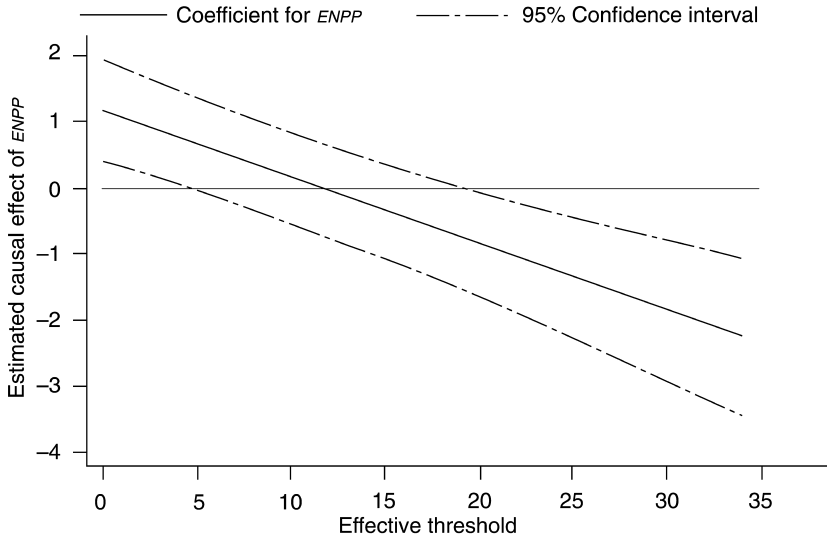
Fig. 2. The impact of the effective number of parties (ENPP) on support for extreme right parties conditioned on the effective threshold

Jackman and Volpert make some data collection errors, though. For example, data for the 1971 Danish election were omitted. They also included results for elections in Ireland in 1990 and the United Kingdom in 1989, although there were no legislative elections in those years. Incorrect figures were also used for the effective number of parties and effective threshold in the Austrian election of 1970.[19] The results from the statistical model when these data errors are corrected are shown in Model 2 in Table 1. It is clear that there is still no reason to question the conclusion that higher levels of unemployment help extreme right parties. However, the conditional coefficient for the effective number of parties shown in the table is now no longer significant.[20] This means that the number of parties in a system has no effect on the electoral success of extreme right parties when the electoral threshold is 0. More information about the impact of electoral thresholds and the effective number of parties can be gathered by calculating the full range of conditional coefficients and standard errors using the corrected data. These are shown graphically in Figures 3 and 4. Again, the solid sloping lines represent the estimated causal effect of THRESH or ENPP on the success of extreme right parties across the full range of the relevant modifying variable. The 95 per cent confidence intervals indicate when this effect is significant and when it is not.

Figure 3 illustrates that electoral thresholds have no effect at all on the extreme right when the effective number of parties ranges between 1.6 and 3.3. Now almost 50 per cent of the observations fall in this range. Thus, Jackman and Volpert's conclusion that 'increasing electoral thresholds dampen support for the extreme right as the number of parliamentary parties expands' is extremely misleading (p. 501). Figure 4 illustrates that the effective number of parties only has an impact on extreme right parties in single-member, plurality systems (those with electoral thresholds over 31 per cent). Less than 10 per cent of the observations fall into this category. Moreover, the coefficient on multi-partism is always negative whenever it is significant. This means that multi-partism never increases the support for extreme right parties. Thus, there is no evidence to support the conclusion that 'multi-partism increasingly fosters parties of the extreme right with rising electoral

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[19] The correct figures were taken from Lijphart, *Electoral Systems and Party Systems*, pp. 33, 160, which is the same source that Jackman and Volpert cite for their data.

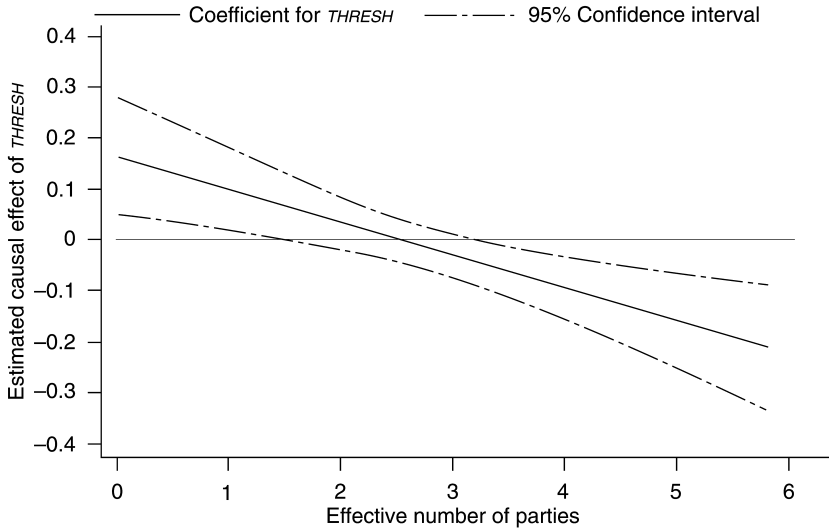[20] The coefficient on the constant is also no longer significant.

Fig. 3. The impact of effective thresholds (THRESH) on support for extreme right parties conditioned on the effective number of parties using corrected data

proportionality'(p. 501). It is clear that Jackman and Volpert's conclusions concerning electoral thresholds and the effective number of parties need to be seriously qualified. This is the case if one uses the original data. It is even more the case when the data is amended to take account of data collection errors. The only conclusion that remains totally valid relates to unemployment.

THEORETICAL JUSTIFICATION OF MODEL

Clearly the empirical evidence presented above provides reasons to doubt Jackman and Volpert's conclusions. The validity of their conclusions can be further called into question once one evaluates
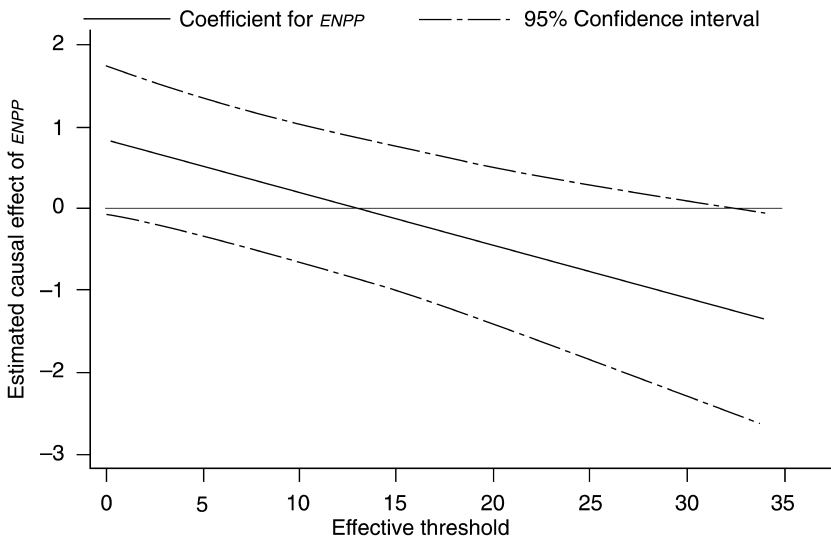

Fig. 4. The impact of the effective number of parties (ENPP) on support for extreme right parties conditioned on the effective threshold using corrected data

the theoretical justification for their model. In particular, there seems little theoretical reason to include the interaction term (*THRESH* × *ENPP*) in the first place. Jackman and Volpert state that this interaction term would not be needed if there were a perfect relationship between electoral proportionality and multi-partism (p. 506). However, they believe that this relationship is far from perfect. As evidence for this, they cite Lijphart's simple correlation between disproportionality and the effective number of parties (− 0.45).[21] They claim that 'while this correlation has the expected sign, the association cannot be characterized as strong' (p. 507). As a result, Jackman and Volpert use Lijphart's conclusion that disproportionality and multi-partism are 'loosely inter-dependent' to justify the inclusion of their interaction term.

The problem is that this reasoning is based on the premise that electoral system features such as disproportionality directly affect the number of parties at the national level. Whilst this is a common argument in comparative electoral studies, Cox has shown that it is theoretically unsubstantiated.[22] Indeed, Cox has illustrated that electoral systems primarily have their political impact at the district level, where district magnitude is the principal determinant of the number of parties. Strong empirical evidence for this comes from numerous analyses of different electoral systems.[23] The effective number of parties that exists at the national level depends on how well the various local party systems are 'linked' across districts.[24] There is no automatic reason why the effective number of parties at the national and district levels should be the same. Thus, Lijphart's failure to find a strong correlation between disproportionality and the number of parties at the national level should come as no surprise. One would only expect to find a strong correlation between disproportionality and the number of parties at the district level. As a result, the number of parties at the national level is not especially relevant to an investigation of the effects of electoral institutions on the success of extreme right parties. This means that Jackman and Volpert should not have added an interaction term between electoral thresholds and the number of parties at the national level in their model; instead they should only have included effective thresholds.[25] The results of such a model can be seen in Model 3 in Table 1. Although the coefficient on electoral thresholds is negative as one would predict, it is not significant. This means that there is no evidence that electoral thresholds actually influence extreme right parties at all.[26]

A second theoretical issue concerns the inclusion of an unconditional unemployment variable in their model. This assumes that unemployment directly causes individuals to vote for extreme right

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[21] Lijphart, Electoral Systems and Party Systems, p. 76.

[22] Cox, *Making Votes Count*; Cox, 'Electoral Rules and Electoral Competition', pp. 145–61.

[23] See Steven Reed, 'Structure and Behaviour: Extending Duverger's Law to the Japanese Case', *British Journal of Political Science*, 29 (1991), 335–56; Gary Cox, 'Strategic Voting Equilibria under the Single Non-Transferable Vote', *American Political Science Review*, 88 (1994), 608–21; Pradeep Chhibber and Ken Kollman, 'Party Aggregation and the Number of Parties in India and the United States', *American Political Science Review*, 92 (1998), 329–42; John Hsieh and Richard Niemi, 'Can Duverger's Law be Extended to SNTV? The Case of Taiwan's Legislative Yuan Elections', *Electoral Studies*, 18 (1999), 101–16; Neil Jesse, 'Candidate Success in Multi-Member Districts: An Investigation of Duverger and Cox', *Electoral Studies*, 18 (1999), 323–40.

[24] Empirical evidence suggests that the extent to which local party systems are 'linked' in national party systems is related to factors such as cleavage structures and the degree of fiscal and political centralization. See Octavio Amorim Neto and Gary Cox, 'Electoral Institutions, Cleavage Structures, and the Number of Parties', *American Journal of Political Science*, 41 (1997), 149–74; Chhibber and Kollman, 'Party Aggregation and the Number of Parties in India and the United States', pp. 329–42; Allen Hicken, 'Political Parties and Linkage: Strategic Coordination in Thailand' (paper presented at the Annual Meeting of the American Political Science Association, Atlanta, 1999).

[25] Since the electoral system endogenously determines the number of parties at the district level, Jackman and Volpert could conceivably use the average effective number of parties at the district level instead of effective thresholds. However, it is unclear how this would be an improvement.

[26] I also calculated two other variants of Model 3. In one I replaced electoral thresholds with the effective number of parties. In the other I included both electoral thresholds and the effective number of parties, still without an interaction term. There was no evidence that either of these factors ever have an influence on extreme right parties. These results are not shown, but are available from the author on request.

parties. However, a causal story that implies a direct relationship such as this is not especially convincing. It is undeniable that economic conditions shape electoral outcomes.[27] The problem is that economic voting theories focus on how incumbent political parties are rewarded or punished for their economic performance. They do not explain why voters who wish to punish incumbent parties should vote for extreme right parties over any other opposition party. Since people traditionally think of left-wing parties as more competent to deal with unemployment, it might actually be more reasonable to assume that these parties are better positioned to benefit from high unemployment levels. Although, there is little theoretical support for assuming that unemployment would have a direct and unconditional effect on the electoral success of extreme right parties, this does not mean that unemployment does not matter. Much depends on why voters think that unemployment is high. There may be little reason for people to vote for extreme right parties if they think that tight monetary policy or rigidities in the labour market cause unemployment. However, it is less difficult to see why they might do this if they think immigration is the cause of unemployment. This is likely to happen when there are large numbers of foreigners in the country. Thus, it might be the case that voters turn to extreme right parties if they think that immigrants cause unemployment.

In fact, this is exactly the type of causal story implied by Jackman and Volpert (pp. 507–8). The problem is that they do not actually test this causal argument. Their model investigates whether unemployment increases the support of extreme right parties in an unconditional manner, even though their reasoning implies that this should only happen when voters think immigrants cause unemployment. A better test of their argument would be to include an interaction term between immigration and unemployment since this allows one to test the prediction that unemployment only helps extreme right parties when immigration is high. I have examined this possibility using a similar Tobit model.[28] I distinguish between populist and neofascist parties on the extreme right and use a larger dataset based on 163 national elections in nineteen West European countries between 1970 and 2000. I find (i) that unemployment only increases the electoral strength of populist parties when there is a large number of foreigners in the country (more than 6.3 per cent of the total population) and (ii) that unemployment never increases the electoral support of neofascist parties. This suggests that Jackman and Volpert's conclusion regarding unemployment may be misleading.

CONCLUSION

Although Jackman and Volpert make important contributions to the study of extreme right parties, their empirical conclusions are open to question. Their claim that higher levels of unemployment help extreme right parties is valid given their model. However, it may be somewhat misleading. This is because there is theoretical reason to believe that the effect of unemployment on extreme right parties depends on the number of foreigners in a country. My reanalysis shows that their claims relating to multi-partism and electoral thresholds are not supported by their own data. It also implies that it would be much harder to contain the electoral growth of extreme right parties through policy intervention than Jackman and Volpert believe. This is because there is no convincing evidence that changing electoral rules would influence their electoral success in one way or another.

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[27] Michael Lewis-Beck and Mary Stegmaier, 'Economic Determinants of Electoral Outcomes', *Annual Review of Political Science*, 3 (2000), 183–219.

[28] Matt Golder, 'Explaining Variation in the Electoral Success of Extreme Right Parties in Western Europe' (paper presented at the Annual Meeting of the Midwest Political Science Association, Chicago, 2001).