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From male dominance to sharing: partner's class and female political party identification 1964–2010

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ABSTRACT

Has the dominating influence of husband's/male partner's class position on his wife's/partners' political party identification declined in Britain? Contrary to predictions, previous research did not reveal a decline in male dominance. We claim a more accurate test by using a theoretical-based cohort design and more appropriate models. To investigate the relative impact of women's and their men's class position, we analyse married and partnered women in the British Election Surveys and distinguish four cohorts with a 1888–1991-birth range and model the relative impact of spouse's class positions with adjusted logistic diagonal reference models allowing the absolute association to change over time. The results show that in case the husband is self-employed, a skilled labourer/foreman or an unskilled/semiskilled labourer, there are no cohort changes in the relative association and women weight their own class position equal to that of husband's class position. However, there is a substantial cohort effect in case the husband has a salariat or lower white-collar class position. In such cases, there is a male dominance class association, but this disappeared for the most recent (i.e. 1961–1991) birth-cohort. For most classes, a sharing-model (both partners equally important) is for the youngest cohort the most appropriate description.



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KEYWORDS Male dominance; political party identification; class; spouse's social class; cohort changes; diagonal reference models

1. Introduction

In this paper, we ask whether there has been a decline of male dominance and a shift towards a more egalitarian situation within the family with respect to women's political preferences in Britain. Previous research

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had found an asymmetric pattern, with men's political preferences being determined only by their own occupational positions not by their partner's position (Marshall *et al.*, 1995: 9). In contrast, in the case of married or partnered women, their political preferences were shaped by their partner's rather than by their own occupation. Moreover, research was unable to find any decline in male dominance in this sphere: at the end of the twentieth century in Britain, compared to wife's class position, husband's class position was still dominant both for wife's political preference (De Graaf and Heath 1992) and class identification (Sobel *et al.*, 2004). However, by using a cohort design with a longer time range and more recent data, and by employing more appropriate models than in previous research, we show that for the 1961–1991 birth-cohort, male dominance disappeared with regard to women's political party identification in Britain.

To investigate the relative impact of women's and their partner's class positions on her political party identification we use the British Election Surveys (BES), covering almost five decades from 1964 to 2010. By comparing party identification across birth cohorts, we are able to extend the time span to cover over a century, distinguishing four cohorts entering the electorate from 1918 to 2009. We include both married and partnered women in the analysis. Throughout this paper, therefore, we will use the term 'partner' rather than husband or wife (except when we refer to research that specifically referred to married men or women). We focus exclusively on married or partnered women, not the male partners. This is partly for pragmatic considerations – the earliest British Election surveys did not ask men about their partner's occupations. Moreover, a focus on women also makes sense in that men have always, according to previous research, based their political choices on their own occupational positions not on their partner's. We have less reason to expect change among men although it is certainly possible that men may have shifted towards sharing within more gender-egalitarian families. This should be explored in a separate paper.

We model the relative impact of the two partners' class positions with adjusted logistic diagonal reference models (DRMs). DRMs provide a powerful technique for comparing the relative weights of the two partners on a specified outcome. Previous research employing these models did not, however, consider the observed decline in the association between class and political preferences over the period covered by our study (see Evans and De Graaf 2013). Although the class/party association as such is not the focus of this paper, a decline in the class

association might potentially affect conclusions about the relative impact of partners' classes and we therefore need to control for this. We adjust the DRMs accordingly by allowing the absolute association between class and political party identification to change over time. Further details about DRMs are given in Section 4 below.

2. Background and earlier findings

The classical idea was that one's social class position would affect one's political interests and in turn political party preferences (Lipset 1981 [1960]). Broadly speaking, classic approaches assumed that the working class, who tend to have lower current incomes, poorer prospects of career progression and higher risks of unemployment, would have a political interest in supporting parties advocating redistributive and welfare-oriented policies. In contrast, the managerial and professional classes with their greater security, material prosperity and prospects for advancement would have a political interest in supporting parties opposed to redistribution and favouring lower taxes instead of expenditure on welfare services (Heath *et al.* 1985). Overall, citizens themselves appear to share this logic. Measures of the public's left–right political attitudes, focussing on the distinction between redistributive and pro-market positions have been shown to be strongly linked to social class (Evans and De Graaf 2013: 4). The link between social class and political values remains strong among the public in Britain (Evans and Tilley 2017, chapter 4), although the movement of the Labour Party towards the centre of the political spectrum in the 1990s, embracing free market policies, weakened the attraction of the party to working-class voters, who turned instead towards non-voting (Weakliem and Heath 1999; Heath 2016). In contrast, support among the middle classes (and especially the petty bourgeoisie) for the Conservatives remained relatively strong (Evans and Tilley 2017, chapter 7).

The classical model also assumed that class interests were shared by members of a family, and were shaped in particular by the occupational position of the male 'head of household'. The conventional practice was to treat the family as the unit of stratification and the male breadwinner as the head of household. The other members of the family were assumed to share similar life chances deriving from the economic resources generated for the family by the (male) head of household. 'The family is, in other words, the unit of class "fate"' (Erikson and Goldthorpe 1992a: 233). Similarly, early studies of class voting, such as Butler and Stokes'

(1969) pioneering study of class voting in Britain, did not ask married men about their wife's occupation. In line with the basic logic of class voting outlined above, it was held to be in the interest of both spouses/partners to follow the political interest of the main earner. In addition, politics was largely seen as part of the 'male sphere' (Verba *et al.* 1997). If you wanted to know the political preference of a wife, you therefore only needed to know the class position of her husband. In general, wherever social class was an important explanatory concept for understanding behaviour, knowledge of the husband's occupation was held to be sufficient. We can term this the 'male dominance' hypothesis (for an overview up until the early 1990s, see Sorensen, 1994).

However, the conventional 'male dominance' assumption came to be heavily criticized (Acker 1973; Britten and Heath 1983; Baxter 1994). Structural changes such as women's increasing participation in the labour market alongside men's declining participation, meant that the man's position was less frequently the main economic articulation between the family and the class structure. Women's growing participation in the labour market (De Graaf and Wiertz 2019: 165) thus called into question the assumptions of the conventional approach. In Britain, for example, the participation of married women in the labour market increased from 9.6% in 1911 to 21.7 in 1951, 53.1% in 1991 (table 8.7 in Gallie 2000), reaching 76.5% in 2013 (own calculation from Labour Force Survey 2013) for married women of working-age. To be sure, in Britain a high proportion (45% in 2013) of the married or partnered women who were economically active held part-time jobs but alongside this increase in married women's participation in the labour market was a long-term decline in men's participation (in Britain down from 95% in 1971 to 83% in 2010, for men of working age, 16–64). So there has been a major decline over time in the extent to which the 'male breadwinner' model holds true in Britain.

Alongside their increased participation in paid work, women have also increasingly entered the more advantaged and secure professional and managerial occupations of the salariat, thus generating considerable resources for the family and more often than before holding occupations that are of higher socio-economic standing than that of their husband or partner. There has thus been an increase in the proportion of 'cross-class' families where women have higher-level occupations than their husbands (Paterson and Iannelli 2007). In the most recent wave of the United Kingdom Longitudinal Household Study (wave 9 conducted in 2018/19), we find that in 27% of the households, women are in higher positions

than their partner is, in 43% women are in similar class positions to those of their partner, leaving only 30% of households where men have higher class positions than their female partners.

Because of these changes, it has therefore become more common in class analysis to test whether a gender-neutral 'class dominance' hypothesis applies, in which whoever of the two partners – either the male or female – makes the greater contribution to family livelihood is treated as the effective head of household.¹ As Erikson explains: 'we can rank the individual positions [of the partners] in a dominance order and let the family position be equal to the individual code that is highest in this order' (1984: 503). However, while this approach avoids the charge of intellectual sexism, its superior explanatory power has not been convincingly demonstrated.

In contrast, Britten and Heath (1983), Heath and Britten (1984) and Beller (2009) have shown that joint approaches offer better explanatory power than either the male dominance or class dominance models among younger generations (see also Korupp *et al.* 2002). This suggests that a 'sharing' hypothesis might stand the tests better than either the male or class dominance hypothesis. A shift to sharing would also be in line with the growing tendency in Britain (and elsewhere) of men and women alike to support equal gender roles (Scott 2008). There has been a marked decline in the belief that 'a man's job is to earn money; a woman's job is to look after the home and family' – down from over 40% in 1984 to less than 10% in 2017 – a change that has probably been driven both by changing attitudes about gender equality and by the changing participation of women in the labour market and hence their increasing contributions to household income (NatCen 2018).

There is also a separate, more theoretical literature suggesting the hypothesis that there has been a decline of community and rising individualism, with Beck's well-known (but rarely empirically tested) thesis of the risk society. Beck emphasizes the way in which certain institutional features of post-modern society, especially the organization of the labour market and the welfare state, undermine class cohesion and ensure that life chances depend on decisions made by individuals. As Beck puts it 'as soon as people enter the labour market ... they are forced to take charge of their own life. The labour market ... reveals itself as a driving force behind the individualization of people's lives' (Beck and Beck-

¹Thaning & Hällsten (2020) investigated whether a more general 'dominance approach' still applies in stratification research by evaluating various measures of socio-economic background.

Gernsheim 2001: 32–33). In such a world, according to Beck's famous aphorism, 'Community is dissolved in the acid bath of competition' (1992, 94). The individualizing role of the labour market is likely to have been given added impetus, at least in Britain, by the Thatcherite reforms of the 1980s which deregulated the labour market. In one of the few empirical tests of Beck's thesis, Heath *et al.* (2009) find gradual trends between 1964 and 2005 in line with Beck's claims, notably a gradual disconnect between class origins, class identities and political allegiances. This suggests that an individualistic approach to understanding family members' class-related behaviour may have become more prevalent. In our case, the individualism thesis would imply that individuals would increasingly behave more according to their own class position and less according to that of their spouse or partner.

These structural and cultural changes, which are doubtless closely inter-linked, suggest then that there will have been a decline in male dominance and a shift towards either class dominance (where the woman's occupation becomes important only if it is higher than that of the male partner), sharing (where both partners' positions are more or less equally important whether or not they have a higher or lower-class position) or individualism (where women as well as men follow their own occupational interests rather than those of their spouse or partner).

However, previous research has failed to show evidence that these changes had impacted on partners' effects. For example, De Graaf and Heath (1992), comparing results from the 1974 to 1987 British Election Surveys found that men's party preferences were unaffected throughout by their partners' occupational positions. They also found little evidence that the conventional approach (i.e. that the male partner's occupational position has more weight on the women's party preferences than her own occupation) had given way to class dominance over this period. However, this research is now quite old and covered only a short time span. With continuing social change, and more data available, the question arises as to whether and when the conventional 'men matter more' hypothesis can be refuted and whether for recent cohorts the class dominance, sharing or individualistic hypothesis better reflects the relative impact of women's class and male partner's class on her political orientation.

To allow for changes in the impact of the woman's class relative to male partner's class on her party identification, we use a cohort approach in order to enable us to cover an even longer effective time span. Changes across cohorts are frequently employed (cf. Inglehart 1971) when there is reason to expect a generational basis for change. We distinguish between four

theoretically based cohorts: (1) respondents born from 1888–1929; (2) 1930–1942; (3) 1943–1960; (4) 1961–1991. We have chosen these four cohorts instead of arbitrary periods of time (such as decades) for several reasons. Our four birth cohorts came of political age – the age at which they could first vote – between 1918–1950, 1951–1963, 1964–1978 and 1979–2009, respectively. Note that universal suffrage for women was not finally introduced until the Equal Franchise Act 1928 although around two-thirds of women had received the right to vote at the age of 30 in 1918. The minimum age for voting was then lowered from 21 to 18 in 1970.²

The earliest of these four periods covered the upheavals of the depression and the two world wars, with generally low peacetime participation of married women in the labour force – 10% or less in the censuses of 1911, 1921 and 1931 (Gallie 2000: table 8.7) and in politics (as measured by the percentage of women MPs – Butler 2000: table 11.9). It is to be expected that, within couples, men's class dominated the political identity of women.

The second period (1951–1963) was in many respects a continuation of the previous one, but with greater political and economic stability. This post-war period is often thought of as a golden age of conventional nuclear families and stable marriages, although in fact married women's participation in the labour force moved up to 21.7 in 1951 and 29.4% in 1961 while divorce rates were considerably higher than in the pre-war period. In a sense, then, this was a transition period from the conventional male head of household model to more egalitarian models of the family.

In the third period, there were radical developments (1964–1978). This was a period when Labour had two periods in power, there were major reforms in a range of social fields (notably divorce, abortion and homosexuality), the rise of the feminist movement in Britain, and the 1975 Equal Pay Act. Emancipation suggests that women would become less likely to regard politics as men's domain. Women also overtook men in education, married women's participation in the labour market further increased (to 42% in 1971), and divorce rates continued their rise although there was no rise in the number of women MPs. Hence, we can expect women to have more interest in politics, to be less likely to follow their male partner's lead and to weight their own class position more than in the first or second periods.

²At most 139 of the 1769 respondents in the oldest cohort might not have been eligible to vote until 1928 although many of these 139 respondents, if they met the requisite property qualification, would have had the right to vote in 1918. Excluding this group, however, does not change the results and the group is too small to be considered as a separate cohort.

Our youngest cohort entered adulthood from 1979, when Britain had its first female Prime Minister and experienced the Thatcherite reforms moving Britain in a markedly more market-oriented direction – often typified as one of increasing individualism – with much greater inequality. Meanwhile married women's participation in the labour force increased further, as did divorce and, to some extent, representation in Parliament. Hence, this was a period potentially leading women to weight their own class position even more.

So, these four periods represent very different social, political and economic environments for our four birth cohorts, women gaining substantial economic independence (through increased economic participation) and social independence (in the sense of increased freedom to choose divorce or abortion) over the course of the four cohorts. Following Inglehart's (1971) theory of generational change, we anticipate that women coming of age and entering the electorate in these different 'formative' environments will have had different political and social orientations. This could be expected to have led to an increasing political awareness of women and hence less impact of their husband's or partner's class relative to their own class for their political identification.

Given our cohort approach, we therefore focus on party identification as our dependent variable. We focus on identification with the Conservative party, since support for non-Conservative parties has fluctuated between a range of opposition parties while the Conservatives have largely maintained their traditional class basis in the petty bourgeoisie and among managers. New Labour's move in the 1990s to the centre ground in order to (successfully) win middle-class votes means that a focus on Labour partisanship has lesser equivalence of meaning over time. There have also been important changes over time among the parties of the centre (the Liberals, the split of the Social Democratic Party (SDP) from Labour and finally their alliance to form the Liberal Democrats). It therefore makes more sense in the British political context to focus on support for the Conservative party.

3. Data

Our main data source is the series of post-election British Election Surveys conducted from 1964 until 2010. These are all high-quality probability samples based either on the electoral registers (in the earlier surveys) or the postal address files. They are nationally representative of the electorates in Great Britain that is of England, Wales and Scotland,

although Scotland north of the Caledonian Canal was typically excluded on cost grounds because of the sparse population. Northern Ireland was also excluded throughout because of its very different party political structure; hence, the surveys are representative of Great Britain rather than of the United Kingdom of Great Britain and Northern Ireland. The surveys that we use were all based on probability sampling and were conducted through face-to-face interviewing during the weeks following the relevant election. (Some of these surveys included mail-back self-completion supplements while others included a panel component, but we do not make use of either the self-completion supplements or the panel components.) Response rates declined somewhat over time (from around 70% to around 60%), as they have in other European countries, but they compare favourably with those achieved in, for example, the highly regarded European Social Survey. Full technical details, and the original data, are available through the UK Data Service (see for example <https://beta.ukdataservice.ac.uk/datacatalogue/studies/study?id=7529>). A replication syntax-file can be obtained at DOI.org/10.17605/OSF.IO/V4689.

In total, we analyse 12 BES datasets (in each case the post-election wave) starting in 1964 and ending in 2010. Unfortunately, later surveys do not measure partner's occupations in sufficient detail to be included (Hence we do not cover the most recent period in Britain when the cross-cutting cleavage of Brexit, driven largely by chauvinist and authoritarian value dimensions, associated with low education rather than with class per se, complicate the picture; see Sobolewska and Ford 2020). The data refer to the elections of 1964, 1966, 1970, 1974, 1979, 1983, 1987, 1992, 1997, 2001, 2005 and 2010. The great advantage of these data is that both political party identification and occupational class of the respondents are measured consistently throughout. Spouse or partner's occupation was also asked throughout with the exception of the 1964, 1966 and 1970 surveys which did not ask men for their spouse or partner's occupation. Our analysis sample consists of married or partnered women. This leaves us with 9005 cases, and after dropping missing cases for political party identification and occupational class, 7764 female respondents remained.

3.1. Dependent variable – party identification

Throughout the British Election Surveys, respondents have been asked 'Do you generally think of yourself as Conservative, Labour, Liberal, or

what?’ We restrict ourselves to the contrast Conservative (1) versus any other party (0). The other main political parties in Britain were the Labour Party, the Liberal Party and the Social Democratic Party (which later formed an alliance with the Liberal Party and subsequently merged to form the Liberal Democrat Party), Plaid Cymru (a Welsh nationalist party) and the Scottish National party. Respondents (about 10% of the overall sample) who did not report a party identification are excluded.

Occupational class: Respondents were asked for details of their current or last job and that of their male partner. The occupational (and employment) data were then coded into the Erikson/Goldthorpe (1992a) class scheme, which has been widely used in research on class outcomes, including research on class politics in Britain. We collapsed the original 11 categories into 5 categories:

- (1) Salariat (professional or higher technical work plus manager or senior administrator)
- (2) Lower white-collar (clerical plus sales or service)
- (3) Self-employed (small business owner, farmer, plus non-professional own account workers prior to 2005)
- (4) Skilled manual and foremen (manual foreman or supervisor of other workers plus skilled manual work)
- (5) Semiskilled and unskilled manual

As Erikson and Goldthorpe (1992b) explain, the class schema is not strictly hierarchical, although the salariat (class 1) clearly comes highest in terms of access to social and economic resources, while the semi- and unskilled class (class 5) clearly ranks lowest. The other three classes are broadly at the same socio-economic level, although they vary considerably in their political partisanship, the petty bourgeoisie being particularly notable for high levels of support for the Conservatives, while the skilled manual class has typically shown high levels of support for the Labour party. Within the salariat, it would have been preferable to make a distinction between managers in the private sector (who tend to be more Conservative) and public-sector employees and the social and cultural specialists (who tend to be more likely to support Labour or the Liberals; cf. Güveli *et al.* 2007; Oesch 2008) but this could not be done consistently across the surveys. Given our long-time frame, we have to realize that several occupations within classes will have changed (Oesch 2008) and that this might have affected the association

between class and party identification. We also therefore control for changes in this association so that this should not affect the relative impact of men and women's class on her party identification.

Birth cohorts: As described above, we distinguish between four theoretically based cohorts: (1) respondents born from 1888–1929; (2) 1930–1942; (3) 1943–1960; (4) 1961–1991. We have chosen these four cohorts instead of arbitrary periods of time (such as decades) for several reasons. Our four birth cohorts came of political age between 1918–1950, 1951–1963, 1964–1978 and 1979–2009, respectively. We will interact these cohorts on the relative impact of women's class and her partner's class position.

Control variables: we include age and period (the year in which the survey was conducted) as covariates in the models. We control for each election year, since we are not interested in any difference that might be caused by the popularity of a certain political party in a certain year (i.e. period effect). We also control for age because of potential life cycle effects. Since we control for both age and period, we also control for the main effect of cohort, since cohort is linearly dependent on period and age (cohort = period – age). Hence, we have no collinearity issues with regard to the interaction of our four specific cohorts with the weight coefficient that compares the relative importance of the woman's class and that of her male partner on her political identification.

Table 1 shows the distribution of our selected female respondents across cohorts and election years (periods). While the earlier birth cohorts are inevitably more frequent in the earlier election years, and vice versa, we can see that there is substantial overlap of birth cohorts in the (relatively large) 1987, 1992 and 1997 election surveys.

4. Methods – diagonal mobility models and formalization of hypotheses

In this paper, we use diagonal reference models (DRMs) to investigate the relative impact of the occupational position of male partner and that of the female partner on her party identification (Sobel 1981). The model is particularly well-adapted to estimating the relative weight of two partners on an outcome. The underlying idea of the DRM is that homogamous families (i.e. spouses or partners having the same class position as each other) represent the reference for each social class. These diagonal reference classes are perceived as showing the typical party identification of those classes, since they are not biased by partners having a different

Table 1. Number of cases in each cohort and survey year (BES 1964, 1966, 1970, 1974, 1979, 1983, 1987, 1992, 1997, 2001, 2005 and 2010): $n = 9005$.

	64	66	70	74	79	83	87	92	97	01	05	10
Coh1	173	164	150	305	116	378	280	193	88	40	40	9
%	76.5	71.6	57.0	46.2	31.2	28.7	23.2	16.7	9.3	4.9	3.8	1.2
Coh2	51	57	79	190	115	342	277	256	194	145	186	101
%	22.6	24.9	30.0	28.9	30.9	25.9	22.9	22.2	20.4	17.8	17.6	13.5
Coh3	2	8	34	165	141	561	541	479	377	332	367	262
%	0.9	3.5	12.9	25.0	37.9	42.6	44.7	41.5	39.6	40.7	34.7	34.9
Coh4	0	0	0	0	0	37	111	226	292	299	464	378
%	0	0	0	0	0	2.8	9.2	19.6	30.7	36.6	43.9	50.4
	226	229	263	660	372	1318	1209	1154	951	816	1057	750

class position. In other words, the diagonal cells in a table cross-classifying one's own class and one's partner's class are assumed to be the core positions establishing the norms and values which set the reference for persons with a heterogamous partner (that is, one having a different class position from the other). The partisanship of these class-heterogamous persons will be intermediate between that of the respective homogamous categories. Cox (1990) described how the DRM as a statistical model bridges theoretical and empirical concerns while Hendrickx *et al.* (1993) provide a comparison of these models with other mobility models. A recent overview of some applications is given by Zang *et al.* (2022).

To model the trend of the relative importance of women's and male partner's class for her party identification realistically we adjust the DRMs. We propose a model that allows both the absolute association between class and political party identification (i.e. the estimates of the party identification on the diagonal cells) to change over time as well as the relative weight coefficient. Allowing the absolute association to vary over time is an important development of our model, given the declining strength of association between class and party – in part due to New Labour's move to the centre and possibly due to changing occupations within a class.

Formally, within DRMs, the political party identification of female respondents in the ij cell of the class table is modelled as a function of the party identification of the homogamous women in the woman's class (cell ii) and of the homogamous women in the partner's class (cell jj). The additive DRM with covariates for a dependent interval variable is given by:

$$Y_{ijk} = p\mu_{ii} + (1 - p)\mu_{jj} + \sum_b \beta_b x_{ijb} + \varepsilon_{ijk} \quad (\text{Model 1})$$

where ε_{ijk} is a stochastic term with expectation 0, and μ_{ii} and μ_{jj} are the population means of the ii th and jj th cells of the mobility table, and there are k observations. The parameter p indicates the salience or weight of one's own social class relative to that of the partner for the dependent variable in question and should theoretically lie in the $[0,1]$ interval. It can be interpreted as the relative weight, or importance, of the respondent's class (and $1 - p$ the relative weight of the partner's class) for the explanation of the dependent variable Y_{ijk} , in our case party identification. If p is significantly smaller than 0.5, partner's class has a stronger relative impact on the dependent variable than the

woman's (the female respondent's) class position. This is what we would expect to find under the male dominance model.

The covariates are expressed by different x_{ijb} variables and the corresponding covariate parameters by β_b . As covariates, we include age of the respondent as well as dummy variables for each election year with the year 2010 as the reference category.

We know that the strength of class voting has declined over time (Evans and De Graaf 2013). In previous international research, this was not considered when employing DRMs. However, in order to model the trends realistically, we have to allow the absolute association between the reference classes and partisanship to change. In other words, the differences in levels of Conservative identity across the diagonal cells of Table 1 are likely to become smaller over time. To do this parsimoniously we model a linear trend.³ In model 2, we therefore fit a model estimating the relative impact of both partners' class positions on her political partisanship, while controlling for a possible weakening of class partisanship over time.

$$Y_{ijk} = p(\mu_{ii} + t_{ii}\text{year}) + (1 - p)(\mu_{jj} + t_{jj}\text{year}) + \sum_b \beta_b x_{ijb} + \varepsilon_{ijk} \quad (\text{Model 2})$$

In model 2 ($\mu_{ii} + t_{ii} \cdot \text{year}$) and ($\mu_{jj} + t_{jj} \cdot \text{year}$) are the population means of the ii th and jj th cells of the mobility table where the parameters t_{ii} and t_{jj} are multiplied by year of survey in order to allow for varying trends in partisanship for each class. For this interaction term, we code the years (i.e. election survey years) from 0 to 46.

We also allow the weight p to have a different value for each class of the female respondent in order to test whether different class positions have different weights. Hence, we allow p to vary according to the class of the female respondent, giving separate estimates (p_i) for each class.

$$Y_{ijk} = p_i(\mu_{ii} + t_{ii}\text{year}) + (1 - p_i)(\mu_{jj} + t_{jj}\text{year}) + \sum_b \beta_b x_{ijb} + \varepsilon_{ijk} \quad (\text{Model 3})$$

This model implies that, for a given class of the female partner, each class of the male partner has the same weight. For example, for women with salariat occupations, we estimate a separate class weight while the class weight of her spouse is the same irrespective of his class position. In most cases, scholars model class dominance in this way, but this is a rather strong assumption. An alternative and theoretical equally

³In line with Evans and Tilley (2013), who showed that there is more or less a linear decline in class-based Conservative partisanship, various tests allowing non-linear changes did not improve the fit of the model.

defensible model, because it also models class dominance, is to define weights to be dependent on husband's/male's class, i.e. p_i in model 3 becomes p_j . For the same example, for the male partner being a salariat worker, we estimate a single weight coefficient and therefore a single weight coefficient for the female partner irrespective of her class position.

Theoretically, if male dominance applies, one could argue that a model based on the class position of the male partner should result in a better fit. In contrast, class dominance implies that the weight coefficients should be higher for the spouse with the better class position, irrespective whether it is the husband or the wife. It is of course possible that there is indeed class dominance but that a certain amount of male dominance also exists.

Another possibility is Weakliem's (1992: 158) suggestion of a gender-neutral mobility model that allows the weight coefficient to depend on both wife's and husband's class position, but this did not lead to any significant improvement.

Finally, we introduce a model in order to test whether there is a cohort increase of the class weights (p_i) of the woman at the cost of a decline of the relative weight of her partner ($1 - p_i$). This model allows any change across cohorts to vary across classes.

$$Y_{ijk} = (p_i + c_{ai}\text{cohort})(\mu_{ii} + t_{ii}\text{year}) \\ + (1 - p_i - c_{ai}\text{cohort})(\mu_{jj} + t_{jj}\text{year}) + \sum_b \beta_b x_{ijb} + \varepsilon_{ijk} \quad (\text{Model 4})$$

In this model, we allow each weight coefficient p_i to be different for each cohort c_a , where 'a' stands for three different cohorts and the first cohort is the reference category. For testing the hypotheses on shifts toward sharing and towards individualism, the weight coefficient should head into the direction of 0.5 and for a trend towards individualism, the weight coefficient should increase towards a maximum of 1, or at least become significantly larger than 0.5.

Since we have a binomial dependent variable, the errors might be heteroscedastic and the predicted values are not bounded within the 0–1 interval. We therefore apply a logistic DRM. Model 4 can be rewritten as:

$$Y_{ijk} = 1 / (1 + \exp - ((p_i + c_{ai}\text{cohort})(\mu_{ii} + t_{ii}\text{year}) \\ + (1 - p_i - c_{ai}\text{cohort})(\mu_{jj} + t_{jj}\text{year}) + \sum_b \beta_b x_{ijb})) \quad (\text{Model 5})$$

5. Analysis

First of all, we present an overview of the proportion identifying with the Conservatives for each cell in the heterogamy table. From the proportions in Table 2, we can see that especially the homogamous self-employed and small employers (often termed the petty bourgeoisie) identify with the Conservatives, almost 70%. At the other extreme, among the homogamous semiskilled and unskilled manual workers only 21% identify with the Conservatives. Among the homogamous members of the salariat, the Conservative partisans are a relatively modest 43%, reflecting the strength of the Liberals in this class.⁴ In line with our assumption about core reference classes, the variations along the diagonal are in general greater than those to be discovered in the marginal cells of the table.

In order to test our expectations, we present next in Table 3 a comparison of the various DRMs described above. Model 1, corresponding to the first model introduced in the equations, uses 18 degrees of freedom: 1 for the weight, 5 for the diagonal cells, and 12 for the covariates. The fit amounts to 19023.4 χ^2 . The weight coefficient p estimate is 0.41 and therefore the weight of the male partner's class is 0.59 (this difference is statistically significant: $p < .01$). In other words, in the case of political partisanship, women weight their own class position somewhat less than their male partner's position. Hence, this result suggests a significant but modest degree of male dominance.

In model 2, we allow the association between class and partisanship in the diagonal cells to vary over time and therefore we use another 5 degrees of freedom, resulting in a significant improvement of fit, $\delta\chi^2(5) = 18.3$, $p = 0.0026$. Controlling for changes in the association does not however affect the value of the weight coefficient p . In the following models, we continue to allow the class/political identification association (i.e. the diagonal estimates) to vary in this way. We also investigated at this stage whether for this relatively simple model cohort differences are to be detected for p . However, the fit of this model is poor (2.5 χ^2 improvement against 4 df) and the estimates for the cohorts are not significantly different from each other. We will show later that there are cohort differences, but only when we allow the weight to vary among classes.

In model 3.1 (see Equation (3) in the description of the models), we allow the weights to differ according to the female respondent's class

⁴For a detailed discussion, see Heath et al (1991), chapter 6 on 'the new middle class', which analyses the strength of the centre parties in the salariat among university graduates, public-sector employees and socio-cultural specialists.

Table 2. Female conservative party identification by own and husband's occupational class (BES 1964–2010; $n = 7764$).

Woman's class		Male partner's class					Average total
		1	2	3	4	5	
Salariat	1	0.43 1152	0.46 80	0.47 151	0.36 288	0.33 183	0.42 1854
Lower white-collar	2	0.57 780	0.52 164	0.62 211	0.37 511	0.35 274	0.49 1941
Self employed	3	0.65 75	0.55 11	0.67 172	0.39 41	0.42 36	0.60 335
Skilled manual and Foremen	4	0.56 106	0.50 22	0.46 54	0.26 206	0.24 188	0.34 576
Semiskilled and unskilled manual	5	0.47 537	0.38 128	0.41 265	0.23 1073	0.20 1055	0.28 3058
Average		0.49	0.46	0.53	0.29	0.25	0.38
Total		2650	405	854	2119	1736	7764

Table 3. Logistic diagonal reference models for effects of own and occupational class of male partner on her party identification.

Model	df used	χ^2	δ df (between models)	$\delta \chi^2$
1. Baseline model: $p * m(Fclass) + (1 - p) * m(Mclass)$	18	19023.4	–	–
2. Baseline and varying diagonal estimates over time: $p * (m(Fclass) + t(Fclass) * year) + (1 - p) * (m(Mclass) + t(Mclass) * trend)$	23	19005.1	(2-1)	5 18.3**
3.1. Female class-specific dominance: $(p_i) * (m(Fclass) + t(Fclass) * year) + (1 - p_i) * (m(Mclass) + t(Mclass) * year)$	27	18968.4	(3.1-2)	4 36.7***
3.2. Male class-specific dominance: $(p_j) * (m(Fclass) + t(Fclass) * year) + (1 - p_j) * (m(Mclass) + t(Mclass) * yea)$	27	18941.1	(3.2-2)	4 64***
4.1. Female class-specific dominance and cohort-specific: $(p_i + c_c * cohort) * (m(Fclass) + t(Fclass) * year) + (1 - p_i - c_c * cohort) * (m(Mclass) + t(Mclass) * year)$	42	18927.3	(4.1-3.1)	15 41.1***
4.2. Male class-specific dominance and cohort-specific: $(p_j + c_c * cohort) * (m(Fclass) + t(Fclass) * year) + (1 - p_j - c_c * cohort) * (m(Mclass) + t(Mclass) * year)$	42	18911.7	(4.2-3.2)	15 29.4***
4.3. Male class-specific and similar youngest cohort effect if male in salariat (1) or lower white-collar (2) class: $(p_j + cl_{1\&2} * cohort_{61-91}) * (m(Fclass) + t(Fclass) * year) + (1 - p_j - cl_{1\&2} * cohort_{61-91}) * (m(Mclass) + t(Mclass) * year)$	28	18924.9	(4.3-4.2)	–14 13.2

Note. Fclass = Female's class; Mclass = class of Male partner. Each model includes covariates for age and for each year of survey with the last survey year as a reference.

** $p \leq .01$. *** $p \leq .001$.

position (irrespective of male partner's class position). This uses a further four degrees of freedom. In model 3.2, weights are allowed to vary according to the class position of the male partner. Both Model 3.1 and 3.2 result in clear improvements of fit compared to model 2, with Model 3.2 showing

a notably large improvement. The estimated weights in Model 3.2 are largest for men in the salariat (class 1), followed by that for men in class 2, but are only around 0.5 for the three other classes. As we noted above, we also tested a gender-neutral model that allows the weight coefficient to vary across both the woman's and her partner's class position, but this did not lead to any significant improvement.

Next, we investigate the hypothesis that the relative weight of women's own class position has increased across birth cohorts at the cost of the relative weight of her partner's class. For this purpose, we allow the weight coefficients to vary according to the four cohorts described in Equation (4) above. In model 4.1, the weights are conditioned on the woman's class position and are interacted with birth cohort (the earliest cohort being the reference category) and in model 4.2 the weights are conditioned on the male partner's class position. The fit of both models is clearly an improvement compared to both models 3.1 and 3.2, respectively, with model 4.2 giving a better fit. Again, the male-specific model results in the better fit. This implies that for her political identification in case the male partner belongs for example to the salariat and she belongs to 'any' other class, she will weight his class position the same, irrespective of her class position. This of course, also applies when the male partner has for example an unskilled- (semiskilled) class position. However, it is important to note that this does not imply anything regarding the relative importance of the husband's class position, which is the focus of this paper. Hypothetically, in this model, the male's relative weight could be anything within the 0–1 interval. It is important to note though that the model implies a specific sort of dominance in the sense that his class position determines the relative weight for his partner, irrespective of her class position, but again it presupposes nothing about the relative weight.

The key question is: which of the cohort changes are significant and in which direction? The class-specific weights for the four cohorts show little variation across the first three birth cohorts, with weights being lowest for class 1, somewhat higher for class 2, and close to 0.5 (a sharing model) for classes 3, 4 and 5.⁵ In the fourth cohort, there is

⁵The relative weights for women are for each cohort:

	Coh 1	Coh 2	Coh 3	Coh 4
Class 1	.135	.059	.021	.505
Class 2	.473	.363	.264	1.00
Class 3	.561	.450	.552	.339
Class 4	.487	.584	.472	.491
Class 5	.598	.678	.491	.459

then a big increase in the size of the weight for classes 1 and 2, but little change for classes 3, 4 and 5. We applied a step-by-step procedure to test the significance of each cohort difference in each class-specific weight. We found that the only significant cohort differences in the weight coefficients were for the salariat and lower white-collar classes (classes 1 and 2) in the youngest (1961–1999) birth cohort. We also tested for alternative cohort distinctions (more arbitrarily chosen and even distinguishing seven different cohorts), but we failed to find a significant improvement in fit.

This suggests that a more parsimonious model, 4.3, which fits a single cohort interaction term for classes 1 and 2 in the youngest cohort, might give an acceptable fit to the data. This model saves 14 degrees of freedom, and yields a fit that is not significantly poorer than that of model 4.2. The parameter estimates of our preferred model 4.3 are presented in [Table 4](#).⁶

[Table 4](#) reveals various striking findings. First of all, in cases where the male partner held a salariat occupation, women's own class, was given very little weight among those born between 1888 and 1960. However, for the 1961–1991 cohort, it becomes a sharing model, where women's own class has a weight of 0.525 ($=0.060 + \text{interaction parameter } 0.465$) and her partner's class weight is therefore 0.475. In cases where the male partner was in a lower white-collar position, the weight of a woman's own class was 0.392 and that of her husband 0.608 for the cohort born 1888–1960 and this, given the standard errors, is not significantly different from a sharing model with a weight of 0.5. However, for the cohorts born after 1960, her party identification is largely driven by her own class position (i.e. $0.402 + \text{interaction parameter } 0.465 = 0.867$). In other words, in cases where the male partner has a lower white-collar position, we notice a marked shift towards individualism.

These results show the importance of estimating class-specific weights, since this reveals that we cannot conclude that there is a general trend towards individualism or sharing. The most important conclusion, however, is that even when men have high-class positions, male dominance vanished completely in the most recent cohort.

Second, we notice that in cases where the male partner belongs to the self-employed, skilled manual or unskilled manual class, women weight their own class about equally to that of their partner, and have done so throughout all four birth cohorts. The slightly greater weight for women

⁶We checked some alternative options for model 4.3, but even adding a separate weight if the woman is a Salariat worker does not improve the fit, also not when we restrict the class weights of classes 3, 4 and 5 to be similar.

Table 4. Parameter estimates of the logistic diagonal reference model 4.3 (standard errors in parentheses).

Man's class	Weight coefficients p (and $1 - p$)				Diagonal estimates	
	Born 1888–1960		Born 1961–1991		Parameters diagonal cells	Trend parameters diagonal cells
	Female partner	Male partner	Female partner	Mail partner		
Class 1 Salariat	0.060 (0.071)	0.940*** (0.071)	0.525*** (0.145)	0.475** (0.145)	0.531*** (0.123)	−0.030*** (0.004)
Class 2 Lower white-collar	0.402*** (0.129)	0.598*** (0.129)	0.867*** (0.161)	0.133 (0.161)	0.747*** (0.235)	−0.039*** (0.009)
Class 3 Self employed	0.501*** (0.072)	0.499*** (0.072)	0.501*** (0.071)	0.499*** (0.071)	1.148*** (0.286)	−0.031** (0.010)
Class 4 Skilled manual	0.504*** (0.070)	0.496*** (0.070)	0.504*** (0.070)	0.496*** (0.070)	−1.032*** (0.204)	−0.014 (0.008)
Class 5 Semiskilled manual	0.559*** (0.066)	0.441*** (0.066)	0.559*** (0.068)	0.441*** (0.068)	−1.166*** (0.155)	−0.026*** (0.006)
B age	0.015*** (0.002)					
For classes 1 & 2: interaction effect on weight p for 1961–1991 cohort	0.465* (0.182)					

Note: Estimates of the survey years are not shown. * $p \leq .05$, ** $p \leq .01$, *** $p \leq .001$.

in cases where the partner belongs to the unskilled manual class (.559) is not significantly different from 0.5. In other words, for all cohorts alike we find a sharing model where the male partner occupies one of these three classes and there is clearly therefore no male dominance.

In order to test the robustness of our findings, we also tested whether period effects showed similar or different patterns of change over time and whether such models lead to improvements in fit. For this test of a hypothesized decline in male dominance, it is not in itself problematic whether the effect is due to cohort replacement or to period effects; yet it is of course clear that cohort effects provide information about replacement of generations. Furthermore, party identification tends to be rather stable over the lifecycle and the latter makes a period effect less plausible. Tests for period effects instead of cohort effects resulted in a worse model fit.⁷ Given our enormous time range, namely those born between 1888 and 1960, we have a substantial age range within each cohort. However, also linear and non-linear age effects, either including the cohort effects or without the cohort effects, did not lead to any improvement in fit. Finally, a selection of more fine-grained cohort groupings did not change our conclusions about the cohorts in which the decline in male dominance occurred.

The finding that four of the five class weight coefficients are not significantly different from .5 in the most recent cohort, suggests that a sharing model offers the best general description. Interestingly, in cases where the husband/male partner has a lower white-collar position, her own class, irrespective of her class position, is dominant for her party identification. However, Table 2 reveals that the percentage of husbands being a lower white-collar worker is very low (i.e. 5.2%) compared with 25.0% for the women. For each cohort, we have 7.3%, 4.8%, 5.0% and 3.5% lower white-collar husbands and about three-quarters of these husbands are partnered either with a woman in the salariat or in a lower white-collar class position. Hence, the results of change for the most recent cohort have empirically a small basis, although the findings are statistically significant.

6. Conclusions

Our findings based on the coverage of 47 years of surveys have several important implications. With regard to understanding married or partnered

⁷The years of the surveys represent the periods and interacting those on the weight coefficients even leads to a worse fit than baseline model 2 in Table 4 ($\chi^2 = 18996.3$ using 34 *df* against $\chi^2 = 19005.1$ and 23 *df* for baseline model 2).

women's political attitudes and partisanship, it no longer suffices to gather data on the socio-economic position of their husband or partner only as had been the practice in the earlier surveys. In this paper, we have shown that for each class position of the male partner, the class position of the woman is of at least equal importance compared to that of her partner for understanding her political identification (which is in turn a strong predictor of voting behaviour). For the first time, we are able to show clearly the decline in male dominance with respect to the relationship between class and party identification for women in Great Britain. The fact that the decline happened for the cohort born after 1960 explains why previous research such as by De Graaf & Heath (1992) did not find a change over time when they analysed only the BES surveys 1979 till 1987.⁸

Secondly, our findings clearly suggest an important step in the emancipation process due to the generational replacement of cohorts. The cohort born after 1960 is clearly different with respect to political identification. It is not only that the typology of the male breadwinner is becoming less relevant, and that more women have a better class position than their male partner. Our results reveal that, on top of these trends, even if the male partner is in a salariat class position, the woman's 'lower' class position has become of equal importance for her party identification if she is born in the most recent cohort. Since the class-specific weight coefficients showed a sharing model for four of the five classes, the class dominance model also no longer applies. Hence, this is an important change on top of the structural changes.

Compared to the older cohorts, only for the youngest cohort there is both a substantial increase of the group of women being a salariat worker with a partner in another class position (from 9.7% for the 1943–1960 cohort to 14.0% in the most recent cohort) on the one hand, and a substantial decline of the group of women with a lower-class position with a partner in a salariat class position on the other (from 20.9% to 16.0%). This would have been an important conclusion in case the result would have revealed that a class dominance model would apply for the recent cohort. However, these structural changes might also explain that such groups, i.e. where the woman is in the better class position (i.e. salariat worker), become much more common and that as part of the emancipation process, women apply more or less a sharing model with their partner for their political party

⁸Another relevant difference is that we allow the class party identification 'association' to change over time, which is not only novel but substantially important.

identification. Hence, there is no overall trend towards individualism regarding partisanship and we have therefore not found support for Beck's hypothesis. The most important conclusion, however, is that male dominance definitely has come to an end due to cohort replacement.

Although, as reported, we applied robustness tests and checked various alternative models, one might argue whether one can properly compare data over such a long-time range. Perhaps for the older cohorts women worked more part-time (although this was less likely the case during World War II). This might partly be an explanation for the changes in the relative increasing impact of women's own class position. However, this does not undermine our conclusions, which is also confirmed by similar conclusions if we leave out the 1960s surveys. We discussed why we expected a change and even if there are other reasons, we still uphold the important conclusion that the male dominance model does not apply anymore. Furthermore, regarding the comparability, the important change happened for youngest cohort of women compared to all other cohorts. Given that the changes happened relatively recently, suggests that the conclusions do not depend on the oldest surveys.

Our findings on cohort differences are in line with the changes in women's attitudes towards gender equality within the family that we noted earlier (Scott 2008). Recent research has also suggested that attitudes towards equality for women showed a marked change in the 1980s, alongside marked shifts towards other social issues (Scott and Clery 2013). The 1980s also corresponds with the period in which women overtook men in their schooling and the 1990s in higher education. Our interpretation therefore is that the spread of egalitarian attitudes is likely to be an important part of the explanation.

Another possibility is that the youngest birth cohort has moved away from seeing politics as men's domain. There is a dearth of direct evidence on this issue in Britain, but European research using 2009 European Election Study data (including the UK as one of the participating countries) has found significant gender and age differences in political knowledge (Fraile 2014). The author suggests that generational differences are a plausible candidate for understanding the effect of age, although there are the usual caveats about inferring generational processes from age. It is also the case in the UK that, though still unequal, women's membership of Parliament showed a step change upwards after 1992 (HOCL 2022).

It is however of considerable interest that a sharing model was apparent among the self-employed and working classes throughout all birth

cohorts. This runs counter to an attitudinal explanation since in general higher levels of education, typically associated with membership of the salariat, tend to encourage 'progressive' views on social issues. This suggests that additional structural processes may be at work within the salariat. One possibility is that there was greater within-class inequality between men and women in the salariat among earlier generations. Throughout most of the twentieth century (although somewhat less so at the end of the period) women were markedly under-represented in higher professional occupations but greatly over-represented in lower professional, intermediate and clerical positions (Gallie 2000, table 8:10). (Thus up until 1981 women's representation among higher professionals was less than a third of what would be expected given their participation in the labour market but among lower professionals and technicians was around one-and-a-half times the expected level.) This imbalance will surely have entailed marked inequalities between men's and women's economic contributions to family income within the salariat and hence affected perceptions of class interests. In other words, unobserved within-class economic heterogeneity could well explain the presence of male dominance within the salariat among members of the earlier birth cohorts. The data suggest that women's under-representation was much smaller in the case of manual occupations.

Due to lack of comparable class measures in the Election Surveys later than 2010, we only analysed data till 2010. The question is whether in the most recent although small 92-01 cohort will show the same result as our 61-91 cohort. The political context has changed substantially with the new Brexit cleavage and the corresponding rise of higher education as a driving force for political preferences (Evans and Tilley 2017; Sobolewska and Ford 2020). It is possible that Brexit was a temporary aberration and the cleavage will lose its importance because all main parties have accepted the outcome. In this case, social class might return to be the dominant cleavage with the sharing model consolidating further. Alternatively, the failure of Brexit to deliver on its promises may keep divisions over Brexit alive as the committed Brexiteers will continue to fight to achieve greater national independence in order to reach the promised land. In this alternative scenario, education might gain political prominence at the expense of social class. This will lead to a different analytical focus, namely the relative importance of women's and their male's partners educational levels.

Given that the modernization process, emancipation and attitude change occurred in other Western nations as well, we might expect

that the findings for the British case might apply to these countries as well. However, class plays a less dominant role in political preferences in other Western societies and other social cleavages such as education might well be more relevant. But even in the case of educational cleavages the claim of a trend towards an increasing relative importance of women's own level of education compared to that of their partner on their political identification can be tested. Future research should reveal whether similar trends are indeed occurring and whether they too have a generational basis.

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No potential conflict of interest was reported by the author(s).

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