Received: 6 November 2018 | Revised: 1 February 2021 | Accepted: 22 February 2021

DOI: 10.1111/irj.12324

ORIGINAL ARTICLE

WILEY

Men, women and unions

Getinet Astatike Haile[1,2]

[1]Department of Industrial Economics, Nottingham University Business School, University of Nottingham, Nottingham, UK

[2]IZA Institute of Labor Economics, Bonn, Germany

Correspondence
Getinet Astatike Haile, Department of Industrial Economics, Nottingham University Business School, The University of Nottingham, South Building, Jubilee Campus, Wollaton Road, Nottingham NG8 1BB, UK.
Email: getinet.haile@nottingham.ac.uk

Abstract

The paper examines whether workplace gender dynamics contributed to the decline of unions. To this end, it reviews relevant literature and proposes three hypotheses, which it then tests using alternative empirical analyses and data from Workplace Employment Relations Survey (WERS) and British Social Attitudes Survey (BSAS). The results from employee-level analysis reveal that, compared with women, (i) men were significantly less likely to have never been union members and (ii) they were also significantly more likely to have been union members in the past. In addition, workplace-level analysis using WERS reveals that there is an inverse link between union membership and the share of women in workplaces, which is also found to have a non-linear form. The paper ponders if unions may need to encompass broader agenda than those informed by the median voter to improve their fate.

1 | INTRODUCTION

There is extensive literature on the decline of unions in Britain since the 1970s. Some of the key reasons underpinning the sweeping decline include unions' failure to organise in new establishments, particularly outside of their traditional manufacturing base; increased competitive pressures; legislative changes; and changes in the composition of the workforce (Blanchflower & Bryson, 2008; Blanden et al., 2006; Brown & Nash, 2008; Bryson, 2008; Bryson & Gomez, 2003; Disney et al., 1995, 1996; Freeman & Pelletier, 1990; Machin, 2000, 2003; Willman et al., 2007).

A related feature of the labour market over this period has been the increase in the share of women in workplaces—dubbed as the 'influx of women' (Parker, 2002, p. 23)—which changed the gender composition of workplaces and unions.[1]

The increase in the share of working women took the form of precarious employment predominantly (Fredman, 2004; Pollert & Charlwood, 2009). However, more recent evidence points to a significant increase in women employment also taking the form of full-time employment, which has increased from 29% in 1985 to 44% in 2017 (Roantree & Vira, 2018). Notwithstanding the perspective of 'core–periphery' or 'insider–outsider' explanation in union membership (Benassi, 2013; Benassi & Dorigatti, 2015; Pulignano et al., 2015), it is reasonable to expect more and more women to have been joining unions over this period given unions' well-recognised role in addressing the challenges women face at work.

If there has been an 'influx of women' to the labour market in recent decades and if more and more of them were likely to have joined unions, then a fall in men's membership must be a key factor behind the observed decline in union membership over the period. Indeed, recent official statistics are in support of this view (BEIS, 2020). The reason why men might have abandoned unions can be linked to unions' utility function. As Booth (1994) and Bryson et al. (2019) explain, unions' preferences are likely to be dictated by the median voter, which women have become as more and more of them join unionised workplaces. If unions were to advance causes such as the gender pay gap and the promotion of work–life balance, which are predominantly woman-centric, men's demand for union membership could fall as per the union demand/supply framework (Bryson & Gomez, 2003; Farber & Krueger, 1993).

This paper attempts to test empirically three hypotheses relating to possible gender differentials in union entry and exit probabilities, which forms an employee-level analysis, as well as a workplace-level analysis examining whether there is a link between union density and the share of women in workplaces. To this end, we use data from the British Workplace Employment Relations Survey (WERS) and the British Social Attitudes Survey (BSAS). The literature review points to potentially divergent interests between men and women, which may be informed by the median voter argument. The empirical results obtained suggest that men were significantly unlikely to have never been union members vis-à-vis women and significantly more likely to have been past union members, thus suggesting a gender differential in union entry/exit probabilities. They also reveal a significant negative relationship between union density and the share of women in workplaces.

It is widely acknowledged that unions have struggled to establish in non-traditional sectors. However, maintaining the size of their existing membership where they are already established seems to be an achievable goal. This they may do through changing their membership preferences, by encompassing broader interests within than just those informed by the median voter. The remainder of the paper is organised as follows. Section 2 provides a review of the literature and develops testable hypotheses. Section 3 describes the data. Section 4 sets out the empirical framework used. Section 5 discusses the results obtained before the final section concludes the paper.

2 | REVIEW OF LITERATURE

Unions are voluntary organisations with the traditional role of organising workers for collective voice and bargaining power (Freeman & Medoff, 1984). They also provide insurance against various risks in the employment relationship such as poor workplace practices (including, among others, payment of unfair wage, unfair dismissal and bullying) and poor health and safety standards. Such working conditions are generally likely to be workplace specific, and unionisation might signal to workers the availability of insurance.[2]

There is a consensus that employees' demand for union representation increases with the extent of problems they face at work (Bryson, 2016; Bryson & Freeman, 2013). Generally, demand for unionisation is a function of the expected benefits (wage and nonwage benefits) from unionisation and the costs associated with joining unions. These benefits and costs are likely to vary across workers, firms, industries and even geographic areas. Individual membership status is also conditional on union supply, which can also vary by firm, industry and geographic area, and may entail queueing for union jobs (Abowd & Farber, 1982), thereby potentially leading to frustrated demand for unionisation, or may lead to its over-supply (see, e.g., Bryson & Gomez, 2003; Willman et al., 2007).

Following the market segmentation literature (see, e.g., Bryson & Gomez, 2003), two workplace-related factors can be identified as key in explaining gaps in unionisation between men and women. First, there is evidence of gender discrimination in Britain (Arulampalam et al., 2007; Berthoud & Blekesaune, 2007; Booth, 2009; Riach & Rich, 2006; Wajcman, 2000). Unions do play a major role in providing insurance against discrimination and disadvantage at work, particularly for women and minority groups (Metcalf, 2000; Waddington & Whitston, 1997). They do so by narrowing the wage differential between women and men (Metcalf et al., 2001) or generally through the collective voice they accord to women (Freeman & Rogers, 1999), which may help them overcome some of the challenges they face including individual behavioural issues thought to explain some of the pay gap they experience (see Artz et al., 2018, for a recent discussion on this). It is well recognised that women's labour market histories are characterised by more interruptions due to childcare and domestic responsibilities than men's (Blau & Kahn, 2000; Hotchkiss & Pitts, 2007; Phipps et al., 2001). Such interruptions—or perceptions thereof—may make women more vulnerable to workplace problems. This is likely to prompt women to embrace unions better than men do, as they seek unions' safeguard against potential vulnerabilities. Available evidence suggests this to be the case, with unions now often reported to be majority-female membership organisations and recent years witnessing significant progress in gender representations within the union power structure (Bryson et al., 2019; Healy & Kirton, 2000; Schnabel & Wagner, 2007).[3]

Second, workplace health and safety matters may also prompt women to embrace unions better than men do. Historically, unions have played a vital role as providers of insurance against poor health and safety at work through promoting union endorsed occupational health and safety standards (Donado & Walde, 2012). In Britain, there has been a general decline in workplace accidents and musculoskeletal disorders, but reported stress and mental health problems continue to escalate (Health and Safety Executive [HSE], 2020; Vickerstaff et al., 2012). Women are likely to take the brunt of such health and safety problems for two main reasons. First, and related to the interrupted labour market histories mentioned above, they are likely to face more stress and mental health problems than men do as they try to readjust to their jobs following such career interruptions. Second, they also tend to play more caring role at home than men do, which may leave them with more difficulties in balancing work and family life. Such difficulties may entail stress and mental health problems among women. Unions have been shown to reduce job-related anxiety even for the most constrained of women with domestic caring responsibilities. Bryson and Forth (2017) report that 'women in the private sector report less anxiety in unionised workplaces, whether or not they have caring responsibilities. And caring for the ill, disabled, or aged is much more strongly linked to higher job-related anxiety in the non-union sector than the union sector—in the union sector the association disappears in the case of women' (p. 2).[4] Given the well-recognised role unions play in addressing workplace health and safety matters, therefore, it is to be expected that women embrace unions more than men would do.

As argued earlier, if (i) the share of women in the labour market and workplaces has shown significant increase in recent decades and (ii) they were more likely to embrace unions, as the literature review above indicates, then the decline in union density must be related to a fall in union membership by men. A recent Labour Force Survey (LFS)-based official statistics covering trade union membership in the United Kingdom over the period 1995–2019 does support this view (BEIS, 2020). Accordingly, 35% of male employees had a trade union membership in 1995 compared with just under 30% for females, but by 2019, male membership has dropped to 20.1% whereas female membership remained comparatively stable at 27.1%. Unlike the LFS, WERS and BSAS monitor union membership histories of respondents, which will allow us to shed some new light on gender disparities in union exit probabilities. Moreover, WERS may offer further insights into the relationship between workplace union density and %female employees.

If unions were stronger in membership terms when men (women) had a significantly higher (lower) labour market participation rate, hence, a higher (smaller) likelihood of being union members (BEIS, 2020; Office for National Statistics [ONS], 2014; Roantree & Vira, 2018), then it must be that men (women) were significantly more (less) likely to have been union members in the past. Therefore, we hypothesise:

Hypothesis 1. Men, compared with women, are less likely to have never been union member in the past.

Also, following Booth (1994) and Bryson et al. (2019), we argue that unions' preferences are likely to be dictated by the median voter, which women have become in recent years as more and more of them join unionised workplaces. If unions were to promote causes such as the gender pay gap or the promotion of work–life balance, which are predominantly woman-centric, men may see relatively less value in maintaining their membership. In other words, men's demand for union membership falls. Therefore, we hypothesise:

Hypothesis 2. Men are likely to have a higher union exit probability than women.

Hypotheses 1 and 2 will be tested by examining if men or women were more likely to have been members in the past and whether they have left unions, respectively, based on employee-level analyses of union membership histories, which we conduct using both WERS and BSAS data sets.

If we accept Hypothesis 2, then men's departure from unions may be more significant where women are the workplace majority. As argued earlier, workplaces with majority-female employees are more likely to represent majority-female union membership workplaces. If so, unions' preferences are likely to be dictated by women, in which case men's membership may fall. Therefore, we hypothesise:

Hypothesis 3. Workplace union density and %female are negatively related.

Hypothesis 3 will be tested based on multivariate analyses that regress workplace union density and workplace %female. We observe workplace %female only in the WERS sample;

hence, the analyses here rely on the WERS samples, both the cross-sections and the panel. If there is a negative relationship between membership and %female as proposed, it is likely that the relationship is non-linear in nature depending on whether women have a simple or an absolute majority in a workplace. This is because, where women have an absolute majority, fewer men would be there to join unions in the first place and even fewer of them to be leaving them. Taking this into account, we measure %female both discretely (0%–25%, 26%–50%, 51%–75% and 76%–100%) and continuously.

3 | DATA AND VARIABLES

The analyses we conduct use data from two different surveys, both of which monitor respondents' union membership history. First, we use data from the two most recent British WERSs (2004 and 2011). Second, we also use data from the BSAS series for the period 2010–2018. The selection of the BSAS years is informed by proximity to the most recent WERS to allow comparisons between the two data sets.

3.1 | WERS 2004 and 2011 data

The WERS data are the most authoritative linked employer–employee data on employment relations in Britain representative of all workplaces with five or more employees. That the data cover large number of demographically varied workplaces provides ample scope for examining union decline and possible links with the changes in workforce gender composition observed. The data cover a range of issues relating to both employees and employers as well as geographic information, which allow controlling on a battery of individual- and workplace-level characteristics. The employer surveys used management questionnaires, which were completed via face-to-face interviews with managers in charge of the day-to-day task of employment relations. The employee surveys, on the other hand, used self-completion employee questionnaires, which were completed by up to 25 employees in participating workplaces (Kersley et al., 2006; van Wanrooy et al., 2013). The 2004 and 2011 WERSs monitored 2295 and 2680 workplaces, respectively, each with 22,451 and 21,981 employees in them; 989 of the establishments were surveyed in both waves, so offering a two-wave panel data on a subset of the WERS establishments.[5] The employee-level analysis uses 21,779 and 20,870 employees, respectively, from the 2004 and 2011 surveys after the exclusion of employees with missing values in some of the key variables. The workplace-level analysis yielded 2050 and 2330 workplaces with non-missing information on key variables in 2004 and 2011, respectively, of which 773 workplaces were surveyed in both waves yielding a panel. Both the employee-level and workplace-level analyses use the WERS employee and workplace weights.

3.2 | BSAS data

The BSAS cover a representative sample of adults in Britain, who are aged 18 years or over, living in private households. Most of the BSAS data are collected via face-to-face interviews, which are supplemented by a self-completion questionnaire.[6] Our analyses use data from the 2010, 2014, 2017 and 2018 waves of BSAS.[7] The BSAS 2010 was chosen to allow comparisons with the most recent WERS (WERS 2011). The remaining BSAS waves thenceforth were chosen because respondents' union histories were monitored only in these waves, BSAS 2018 being the most recent wave available. We achieved 3175, 2861, 3873 and 3728 employees for each of the waves, respectively. The analysis uses BSAS weights, which were intended to account for complex survey designs.

3.3 | Outcome variables

The outcome measures for the WERS-based employee-level analysis come from employees' responses to the membership history question: 'Are you a member of a trade union or staff association?' Employees provided one of the following three responses: (i) 'yes', (ii) 'no, but have been in the past' and (iii) 'no, have never been a member', which yielded the corresponding multichotomous outcome measure with three categories: 'current member', 'past member' and 'never member'. In addition, the workplace-level analysis (WERS) uses *union density* as an outcome measure, which is derived from employers' responses to the following two questions: (i) 'How many employees at this establishment are members of a trade union or independent staff association—whether recognised by management or not?' (${n^u}_j$) and (ii) 'Currently, how many employees do you have on the payroll at this establishment?' (N_j). Based on the responses provided to these questions, a percentage measure of workplace union density (UD) has been generated for each workplace as $UD = \left(n_j^u / N_j\right) \times 100$.

The outcome measure for the BSAS-based employee-level analysis comes from respondents' answers to two membership history-related questions. First, employees were asked: '(May I just check) are you now a member of a trade union or staff association?' to which employees would respond one of the following: 'yes, trade union', 'yes, staff association' or 'no'. If they answered 'no', then they would be asked the second question: 'Have you ever been a member of a trade union or a staff association?' to which they would once again respond one of the following: 'yes, trade union', 'yes, staff association' or 'no'. We used these responses to generate multichotomous outcome measure with three categories: 'current member', 'past member' and 'never member', like the outcome from WERS.[8]

3.4 | Control variables

The WERS employee-level analysis controls for a range of employee, workplace and geographic characteristics, whereas the workplace-level analysis controls for workplace and geographic characteristics, which include workplace size, industry, ownership type and a count measure of workplace equality practices. The '%female' control variable is obtained from employers' responses to the following questions: (i) 'How many women work full-time?' (${n^{f1}}_j$) and (ii) 'How many women work part-time?' (${n^{f2}}_j$). Combining the responses to these two questions yields the total number of female employees in a workplace (${n^f}_j = {n^{f1}}_j + {n^{f2}}_j$). '%female' is then obtained as $\%female = \left(n_j^f / N_j\right) \times 100$, where N_j represents the total number of employees in a workplace. As discussed in Section 2, we measure %female both continuously and discretely to account for potential non-linearities in the US–%female relationship.

The BSAS-based employee-level analysis controls for a range of demographic (age, marital status and whether young children in the household), human capital (respondent's qualification), job (occupation and industry) and region characteristics.[9]

4 | ANALYTICAL FRAMEWORK

We use three empirical approaches in this paper. The first approach is a three-choice multinomial probit (MNP) model relating to the three union membership status categories of 'current member', 'past member' and 'never member', which we observe in both the WERS and BSAS data sets. The MNP model is motivated by the latent variable approach.[10] Accordingly, employees are assumed to choose one of the three union membership categories, where the utility associated with the jth membership status category is given by $U_j = V_j + \varepsilon_j$, $j = 1, 2, 3$, with the errors assumed to be joint normally distributed. The probability that an employee i chooses the jth union membership status category is then modelled as

$$p_{ij} = pr[y_i = j] = F_j\left(\boldsymbol{X}_i'\boldsymbol{\beta}\right), \tag{1}$$

where F is the standard normal distribution function and $\boldsymbol{X}$ represents the vector of time variant and invariant employee and workplace characteristics including geographic region. Marginal effects are computed from the estimated coefficients to determine if the probabilities of being 'never member' and 'past member' are significantly different between men and women, which will allow us to test Hypotheses 1 and 2 using each of the WERS and BSAS cross-sections.

The second empirical analysis relies on the WERS data to regress UD on %female to test Hypothesis 3 using each of the WERS cross-sections separately as well as pooled. Strictly speaking, UD is a proportion best modelled using fractional response models (Baum, 2008; Papke & Wooldridge, 1996; Williams, 2017).[11] Taking this into account, we estimate generalised linear models (GLMs) using STATA's 'GLM' suits (StataCorp, 2019). To make comparisons with the third approach below, we assume Gaussian distribution for UD and the identity link function to estimate the model:

$$E(UD_k) = \boldsymbol{X\beta},\ UD_k \sim \text{Normal}, k = 1, \ldots, M, \tag{2}$$

where UD represents union density, $\boldsymbol{X}$ represents the vector of time variant and invariant workplace characteristics and k indexes workplaces. As noted earlier, we use discretised measure of %female to account for possible non-linearities in the relationship between union membership and %female.

Finally, we also implement panel data regressions using the panel of 773 workplaces, which were surveyed in both the 2004 and 2011 waves of WERS. The estimated model in this case has the following general form:

$$UD_{kt} = \alpha + \boldsymbol{\beta}'\boldsymbol{X}_{kt} + \mu_k + \varepsilon_{kt};\ k = 1, \ldots, N \text{ and } t = 2\,(2004, 2011), \tag{3}$$

where, as in Equation 2, UD represents workplace union density and k indexes workplaces ($N = 773$), t represents time, $\boldsymbol{X}$ represents the vector of time variant workplace characteristics including %female, μ_k is the time invariant unobserved workplace characteristics and ϵ_{kt} is the idiosyncratic error term. We make alternative assumptions on the nature of μ_k and estimate pooled ordinary least squares (OLS) and random-effect (RE) and fixed-effect (FE) regressions ruling out censoring in UD.[12] As in the second approach above, we measure %female discretely to address potential non-linearities. In addition, we estimate FE models with %female measured continuously, which can provide the sternest tests for Hypothesis 3 as does the FE model with the discretised %female measure. The models are estimated using STATA's 'reg' (pooled OLS) and 'xtreg' suits (StataCorp, 2019).

5 | RESULTS AND DISCUSSION

The results from the employee-level analysis using the MNP model are reported in Tables 1 and 2 for both the WERS and BSAS samples. Table 1 reports estimated probabilities of current, past and never membership from the MNP model for each of the WERS and BSAS samples (left panel), together with the raw proportions for each membership categories (right panel). The estimated probabilities do strikingly resemble the raw proportions, which is very reassuring. Comparing the estimated probabilities and/or raw proportions within and between the data sets reveals that there is similar spread of the three membership categories within each data set. On the other hand, there is a significant variation in the three membership categories between WERS (roughly 37%, 17% and 46%) and BSAS (roughly 19%, 27% and 54%).

Table 2 reports estimated gender-specific marginal effects for men from the MNP estimation, which are meant to test Hypotheses 1 and 2.[13] Accordingly, the most consistent results from the two data sets indicate that, compared with women, men were (i) significantly less likely to have never been union members and (ii) significantly more likely to be past union members. The WERS samples suggest men to be about 3 percentage points less likely to have never been union members and between 2 and 5 percentage points more likely to be past members vis-à-vis women. Estimates from the BSAS samples are even larger in magnitude

TABLE 1 Estimated probabilities of union membership from MNP and raw membership proportions (WERS 2004 and 2011 and BSAS 2010–2018)

	Probability estimates			Raw data (%)			
	Current member	Past member	Never member	Current member	Past member	Never member	*N*
	b (s.e.)	b (s.e.)	b (s.e.)				
WERS 2004	0.367*** (0.01)	0.167*** (0.00)	0.466*** (0.00)	36.8	16.6	46.6	21,779
WERS 2011	0.369*** (0.01)	0.170*** (0.00)	0.462*** (0.00)	37.2	17.0	45.8	20,870
BSAS 2010	0.207*** (0.01)	0.236*** (0.01)	0.557*** (0.01)	21.0	26.3	52.8	3175
BSAS 2014	0.185*** (0.01)	0.230*** (0.01)	0.585*** (0.01)	18.8	26.8	54.4	2861
BSAS 2017	0.194*** (0.01)	0.231*** (0.01)	0.575*** (0.01)	19.3	26.2	54.5	3873
BSAS 2018	0.180*** (0.01)	0.242*** (0.01)	0.578*** (0.01)	18.7	27.1	54.1	3728

Note: Standard errors are in parentheses. WERS and BSAS employee survey weights have been used.
Abbreviations: BSAS, British Social Attitudes Survey; MNP, Multinomial Probit; WERS, Workplace Employment Relations Survey.
$^{***}p < 0.01$. $^{**}p < 0.05$. $^{*}p < 0.1$.

TABLE 2 Estimated marginal effects of men's union membership (WERS and BSAS).

	WERS 2004	WERS 2011	BSAS 2010	BSAS 2014	BSAS 2017	BSAS 2018
Male						
Current member	−0.017* (0.01)	0.008 (0.01)	0.046** (0.02)	0.028 (0.02)	0.021 (0.01)	0.023 (0.01)
Past member	0.052*** (0.01)	0.021*** (0.01)	0.086*** (0.02)	0.087*** (0.02)	0.066*** (0.01)	0.071*** (0.02)
Never member	−0.035*** (0.01)	−0.029*** (0.01)	−0.131*** (0.02)	−0.116*** (0.02)	−0.086*** (0.02)	−0.094*** (0.02)
Wald's χ^2	4486.62	4262.62	790.98	808.28	797.98	771.70
Prob > χ^2	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
N	21,779	20,870	3175	2861	3873	3728

Note: Reference category: female. Standard errors are in parentheses. WERS and BSAS employee survey weights have been used.

Abbreviations: BSAS, British Social Attitudes Survey; WERS, Workplace Employment Relations Survey.

$^{***}p < 0.01$. $^{**}p < 0.05$. $^{*}p < 0.1$.

suggesting men to be between 9 and 13 percentage points less likely to have never been union member and between 7 and 9 percentage points more likely to have been past members. The results are all significant at the 99% level, thus providing strong evidence in support of Hypotheses 1 and 2.[14]

Figures 1 and 2 depict plots of the estimated marginal effects for men from the WERS and BSAS samples, respectively; thus, they correspond to the marginal effects reported in Table 2. In each case, the plots clearly depict the estimated probabilities of being ‘never member’ and ‘past member’ for men on either side of the vertical line, which represents zero value, thus indicating significant negative and positive associations, respectively. In contrast, the estimated probabilities of ‘current member’ fall on the zero line for the most part, suggesting no statistically significant difference between men and women in this respect except in BSAS 2010 as noted above.

Table 3 reports marginal effects from the workplace-level analysis based on WERS samples, which are obtained from the GLM estimations on the 2004 and 2011 cross-sections separately as well as pooled.[15] The marginal effects relating to the discretely measured %female reveal that UD and %female are negatively related. What is more, we find strong evidence of non-linearity in this relationship. Specifically, we find the consistent result that compared with the base category of workplaces with up to 25% female employees, workplaces with 26%–50% and 51%–75% female employees exhibit statistically significant negative relationship with UD. In other words, it appears that workplaces with female majority (or near majority), which are highly likely to be majority-female union membership workplaces where unions' preferences may be dictated by the median voter, have experienced a decline in union membership. These results thus provide some evidence in support of Hypothesis 3. Given the evidence in BEIS (2020) and our results from the individual-level analyses, it is reasonable to attribute a good part of the membership decline to the fall in men's membership. If so, median voter-based explanation may be a candidate explanation here even though Bryson et al. (2019) have ruled this out for the United Kingdom.

Table 4 reports estimation results from the panel data analysis on the subset of 773 panel workplaces surveyed in both WERS 2004 and 2011. They include pooled OLS (left panel) and

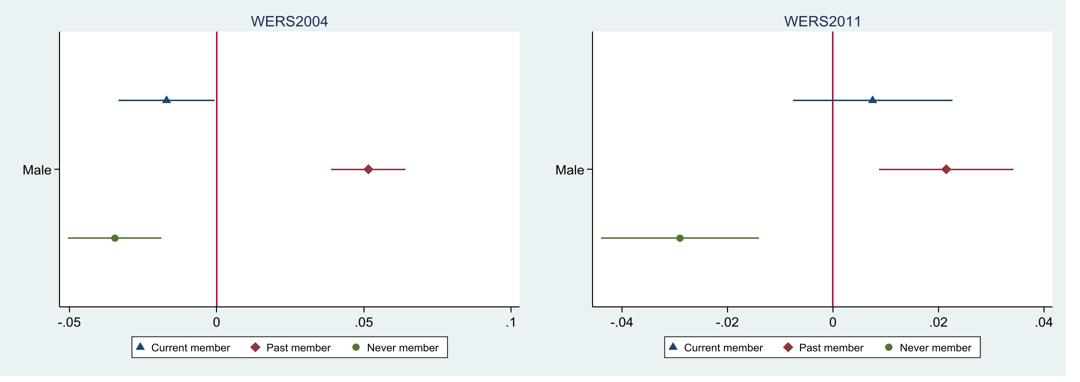

FIGURE 1 Plots of estimated marginal effects of men's union membership (Workplace Employment Relations Survey [WERS] 2004 and 2011)

linear panel data models (RE and FE, middle and right panels, respectively).[16] Once again, we used discretely measured %female to account for non-linearities initially. The OLS and RE results do consistently reinforce the earlier findings on the inverse link between UD and % female we found.[17] Even what can be regarded as the most restrictive of the models (FE) reveals that compared with workplaces with up to 25% women employees, workplaces with 51%–75% female employees exhibit significantly negative relationship with UD. Broadly, therefore, the results we found from the panel data analysis provide additional evidence supporting Hypothesis 3, as did the earlier results from the cross-sectional and pooled analysis on WERS. Once again, therefore, workplaces with female majority (or near majority), which are highly likely to be majority-female union membership workplaces, seem to have experienced a decline in UD. As we argued earlier, this seems to be largely driven by the fall in men's membership.

Finally, Table 5 reports FE estimates from the panel workplaces as in Table 4, but with % female measured continuously and with and without other time-varying workplace controls as alternative specifications. As noted earlier, this (and the FE specification in Table 4) is likely to be the sternest of tests for Hypothesis 3. The results indicate a negative and statistically significant link between UD and %female. Thus, once more, the results in Table 5 provide evidence in support of Hypothesis 3.

Overall, therefore, our analyses do lend clear evidence in support of the three hypotheses proposed. The employee-level analyses did support Hypotheses 1 and 2 on gender disparities in union entry/exit probabilities strongly, whereas the workplace-level analyses using both cross-sectional and panel data regressions also lent evidence in support of Hypothesis 3. The evidence we found from the panel data analysis may not be as overwhelming as the cross-sectional evidence but fairly strong evidence nevertheless bearing in mind that these are based on the much-reduced sample size and such restrictive estimators as the FE estimator.

6 | SUMMARY AND CONCLUSION

The paper attempted to examine whether there is a link between the increase in the workplace share of women in Britain in recent decades and the decline of unions over the same period. It

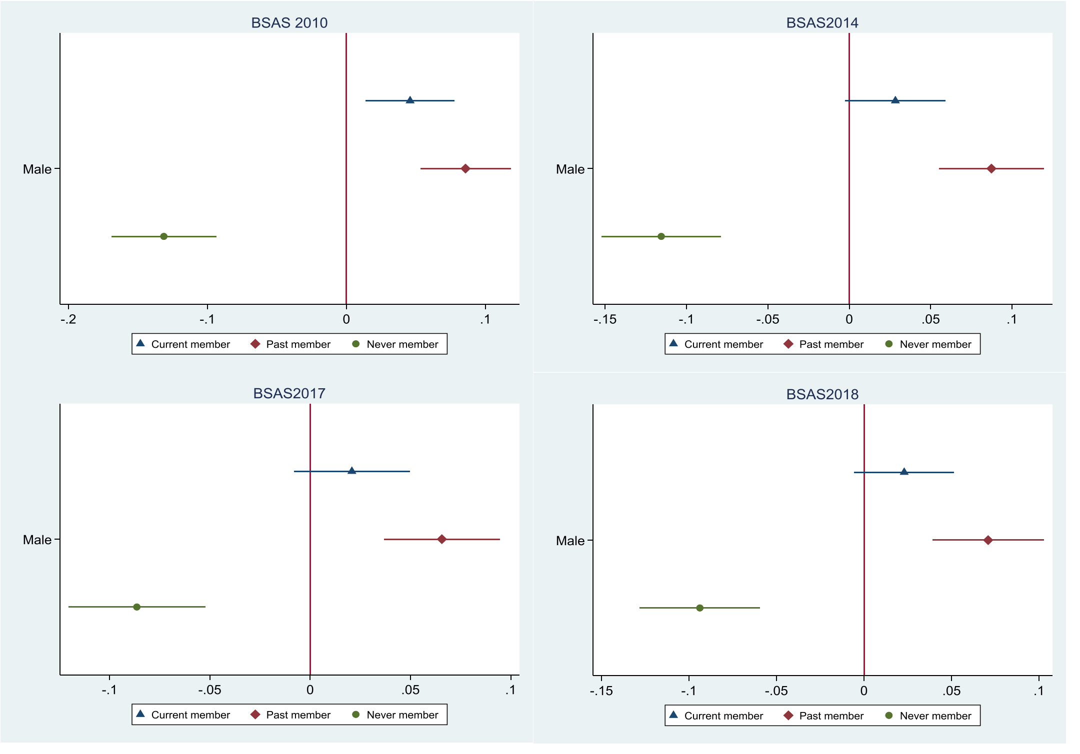

FIGURE 2 Plots of estimated marginal effects of men's union membership (British Social Attitudes Survey [BSAS] 2010, 2014, 2017 and 2018)

argued that unions should have been boosted when there has been an 'influx of women'—unions' more instinctive allies than men—into workplaces in Britain. The literature review highlighted that employees' demand for union representation depends on (i) the problems they face at work and (ii) the net benefit they expect to get from being union member. Membership is also a function of union supply, and both the demand for and the supply of union membership vary across firms, industries and geographic regions.

The paper sat up three testable hypotheses on gender differential in union entry/exit probabilities and the link between union density and %female in a workplace setting. The hypotheses were based on (i) the observed differences in labour market participation between men and women, hence the scope for union membership (BEIS, 2020; ONS, 2014; Roantree & Vira, 2018), and (ii) the argument that unions' preferences are likely to be dictated by the median voter (Booth, 1994; Bryson et al., 2019), which women have become in recent years as more and more of them have joined workplaces. We argued that women are likely to be better allies for unions than men, given their well-recognised labour market challenges such as the gender pay gap and interrupted labour market histories. If more and more women were to join unions and if unions were likely to promote causes that are predominantly woman-centric, then men might have lost the drive to embrace unions, with their demand for unionisation falling.

To test the three hypotheses developed, we used data from the two most recent British WERSs (WERS 2004 and 2011) and the four most recent waves of the BSAS (BSAS 2010, 2014,

TABLE 3 Union density and %female, marginal effects from Generalised Linear Models (cross sectional and pooled 2004 and 2011).

	Pooled WERS 2004 and 2011		WERS 2004		WERS 2011	
%female (0%–25% base)						
26%–50%	−0.1305***	−0.1289***	−0.1337***	−0.1326***	−0.1269***	−0.1251***
	(0.014)	(0.011)	(0.021)	(0.016)	(0.019)	(0.016)
51%–75%	−0.0604***	−0.1339***	−0.0806***	−0.1335***	−0.0438**	−0.1326***
	(0.014)	(0.012)	(0.020)	(0.018)	(0.019)	(0.016)
76%–100%	−0.0023	−0.0738***	0.0038	−0.0557**	−0.0080	−0.0938***
	(0.015)	(0.015)	(0.021)	(0.022)	(0.021)	(0.020)
Workplace characteristics	No	Yes	No	Yes	No	Yes
2011		−0.0328***				
		(0.007)				
Observations	4380	4380	2050	2050	2330	2330

Note: Standard errors are in parentheses. WERS employer survey weights have been used.
Abbreviation: WERS, Workplace Employment Relations Survey.
$^{***}p < 0.01$. $^{**}p < 0.05$. $^{*}p < 0.1$.

TABLE 4 Union density and %female, estimates from linear panel data models (WERS panel establishments).

	Pooled OLS		Random effects		Fixed effects	
%female (0%–25% base)						
26%–50%	−0.1730***	−0.1726***	−0.0943***	−0.1027***	−0.0357	−0.0340
	(0.024)	(0.022)	(0.021)	(0.020)	(0.025)	(0.024)
51%–75%	−0.1212***	−0.1636***	−0.1011***	−0.1106***	−0.0580**	−0.0538**
	(0.024)	(0.021)	(0.022)	(0.021)	(0.029)	(0.027)
76%–100%	−0.0747***	−0.0554**	−0.0345	−0.0192	0.0250	0.0315
	(0.025)	(0.005)	(0.027)	(0.025)	(0.039)	(0.038)
Workplace characteristics	No	Yes	No	Yes	No	Yes
2011		−0.0300**				
		(0.015)				
Constant	0.3807***	0.0413	0.3413***	0.0933***	0.3020***	0.2321***
	(0.020)	(0.030)	(0.020)	(0.031)	(0.021)	(0.050)
R^2	0.036	0.205				
$N \times 2$	1546	1546	1546	1546	1546	1546
N	773		773	773	773	773

Note: Robust standard errors are in parentheses. Workplace Employment Relations Survey employer panel survey weights have been used.
Abbreviation: OLS, ordinary least squares.
$^{***}p < 0.01$. $^{**}p < 0.05$. $^{*}p < 0.1$.

TABLE 5 Union density and continuously measured %female, estimates from fixed-effects model (WERS panel establishments).

%female	−0.0033* (0.002)	−0.0032** (0.002)	−0.0033** (0.001)	−0.0033** (0.001)
Establishment size		0.0116 (0.027)	0.0155 (0.022)	0.0161 (0.024)
Ln workplace age			−0.0351 (0.030)	−0.0349 (0.030)
Summative gender equality HPWS				−0.0032 (0.014)
Constant	0.5857*** (0.090)	0.4903*** (0.181)	0.5959*** (0.209)	0.6020*** (0.203)
R^2	0.020	0.022	0.047	0.047
$N \times 2$	1546	1546	1546	1546
N	773			

Note: Robust standard errors are in parentheses. Workplace Employment Relations Survey employer panel survey weights have been used.
***$p < 0.01$. **$p < 0.05$. *$p < 0.1$.

2017 and 2018), both of which monitored the union membership histories of respondents. Similar union membership histories had been monitored by BSAS waves earlier than 2010. However, the proximity of BSAS 2010 to WERS 2011 in terms of timing has informed the decision to consider BSAS 2010 as the earliest wave to allow validating results from WERS. As well as using alternative data sets, the paper also used alternative empirical approaches, thus checking the robustness of our results.

The results obtained provided strong evidence in support of the hypotheses proposed. The employee-level analyses using WERS and BSAS revealed that, compared with women, men were (i) significantly less likely to have never been union members and (ii) significantly more likely to have left unions (or have been union members in the past), which are both strongly significant. The workplace-level analysis using WERS, which involved cross-sectional (GLM) regressions on each of the 2004 and 2011 samples separately and pooled, revealed workplace union density and %female to be significantly negatively related. The results also uncovered a non-linear pattern in the union density–%female relationship. In particular, the negative link found appears to increase in magnitude and statistical significance as the share of women increased from '25% or less' to between '26%–50%' and '51%–75%'. A further increase in the share of women to '>75%' was not found to have a significant link with union density. As we argued earlier, the decline in union density found is in workplaces with female majority (or near majority). These are likely to be majority-female union membership workplaces, where unions' membership preferences might have been dictated by the median voter, which women have become in recent years, to address causes that are women-centric in the main. The results from the employee-level analysis indicated that such declines in membership are driven to a significant extent by the fall in men's membership, which is also in line with the evidence elsewhere (BEIS, 2020). On the other hand, where women have an absolute majority (such as >75%), there must have been fewer men to be union members in the first place and even fewer of them to be abandoning their membership due to the shift in unions' membership preferences.

The workplace-level analysis also implemented alternative panel data regressions using data on the reduced subsample of panel workplaces surveyed in both WERS 2004 and 2011. The results found from the panel data analysis reinforced those obtained from the cross-sectional analyses in both the pattern of statistical significance and the non-linear link between union density and % female we found. In this regard, the sternest test was the FE regression with %female measured both discretely and continuously, which nonetheless lent additional support for the hypothesis proposed regarding the negative link between workplace union density and %female.

The paper is rigorous in many respects including the review of the literature, the use of two different data sets, the employee- and workplace-level analyses and the implementation of alternative empirical approaches to test the hypotheses proposed. As pointed earlier, unions have struggled to establish in new settings in non-traditional sectors. That may remain a challenge for the foreseeable future. However, they may be able to maintain, if not necessarily boost, membership where they are already established. This may be achieved if unions change their preference for membership by appealing to broader interests within and going beyond the median voter.

ACKNOWLEDGEMENTS

The author would like to extend special thanks to the editors and an anonymous referee for useful comments and suggestions on an earlier version of this paper. The author also acknowledges the Department of Business, Innovation and Skills, the Economic and Social Research Council, the Advisory, Conciliation and Arbitration Service and the National Institute of Economic and Social Research as the originators of the 2011 Workplace Employee Relations Survey data and the Data Archive at the University of Essex as the distributor of the data.

CONFLICT OF INTERESTS

There are no conflict-of-interest issues to report in connection with this work, which did not benefit from any funding. The usual disclaimer applies.

ORCID

Getinet Astatike Haile https://orcid.org/0000-0001-6274-971X

ENDNOTES

1 Recent official figures reveal that the gap in the labour market participation rate between men and women has declined from 14.5 (1994) to 8.8 (2014) percentage points in favour of the latter (ONS, 2014) whereas the proportion of women in employment has increased from 57% in 1975 to a record high of 78% in 2017 (Roantree & Vira, 2018).

2 This is ruling out the possibility that some workers might sort into workplaces with poor practices.

3 Translating improvements in such representations into actions directed at enhancing women's experience within unions and workplaces seems to remain a challenge, however, which may explain why, in Britain, there is a gender gap in the benefits accruing from unionisation, with only men benefiting from unionisation even in settings where women are the median voters (Bryson et al., 2019).

4 Bryson (2016) has also showed the link between the management of workplace health and safety matters and the health and safety risks employees face. In particular, he finds that having on-site worker representation to deal with health and safety matters leads to a lower health and safety risk than having direct consultation between management and employees.

5 However, WERS does not offer a panel sample on employees.

6 See Park et al. (2002) for further details of the survey and Bryson (2005) for published work using the data.

[7] Except in 1988 and 1992, the BSAS has been conducted annually since 1983, often achieving at least 60% response rate. Union membership histories had been monitored in waves prior to 2010. However, the motivation here is the validation of results from the WERS, hence the selection of BSAS waves nearer WERS 2011. BEIS (2020) provides long-term union membership trends of men and women in the United Kingdom.

[8] As noted earlier, the second question was asked only in 2010, 2014, 2017 and 2018, hence our selection of these years.

[9] A summary of these characteristics is provided in the supporting information.

[10] See Chapter 15 of Cameron and Trivedi (2005) for a discussion of multinomial probability models. We cluster standard errors by workplaces for the WERS sample to account for employees sharing employer characteristics.

[11] Baum (2008) emphasises on this point that '... it is not appropriate strategy [to use censored normal regression] as values outside the [0, 1] interval are not feasible for proportions data' (p. 302).

[12] In addition, we have also estimated panel data tobit models using STATA's 'metobit' suit considering the preponderance of 0 s in *UD* (40% of the panel workplaces).

[13] The corresponding full regression tables are provided in the supporting information.

[14] In terms of 'current membership', the estimated marginal effects suggest similar probabilities for men and women except around 2010, when men were found to be significantly more likely to have been 'current member'. This seems a blip around then, which BEIS (2020) also picks, which we suspect might have to do with the great recession even though studies elsewhere (e.g., Addison et al., 2016) report the economic crisis left membership 'largely' unaffected.

[15] The corresponding full regression table is provided in the supporting information.

[16] The corresponding full regression table is provided in Table A4 of the supporting information. Considering the zero values for the union density measure, which amounts to about 40% in the combined WERS 2004 and 2011 data, we estimated tobit models for the 2004 and 2011 cross-sections and the panel. The estimated results are very similar to the GLM estimates for the cross-sectional analysis while the panel tobit yielded estimates that are between the pooled OLS and RE estimates we obtained in terms of magnitude.

[17] Average marginal effects from panel tobit ('metobit') are found to be remarkably similar.

REFERENCES

Abowd, J., & Farber, H. (1982). Job queues and the union status of workers. *Industrial and Labor Relations Review*, *35*, 354–367.

Addison, J. T., Portugal, P., & Vilares, H. (2016). Unions and collective bargaining in the wake of the Great Recession: Evidence from Portugal. *British Journal of Industrial Relations*, *55*(3), 551–576.

Artz, B., Godall, A., & Oswald, A. (2018). Do women ask? *Industrial Relations*, *57*(4), 611–636.

Arulampalam, W., Booth, A., & Bryan, M. L. (2007). Is there a glass ceiling over Europe? Exploring the gender pay gap across the wage distribution. *Industrial and Labor Relations Review*, *60*(2), 163–186.

Baum, C. (2008). Modelling proportions. *The Stata Journal*, *8*(2), 299–303.

BEIS. (2020). *Trade union membership, UK 1995–2019: Statistical bulletin: National Statistics*. London, UK: Department for Business, Energy & Industrial Strategy (BEIS).

Benassi, C. (2013). *Political economy of labour market segmentation: Agency work in the automotive industry, ETUI working paper 2013*. Brussels: ETUI.

Benassi, C., & Dorigatti, L. (2015). Straight to the core—Explaining union responses to the casualization of work: The IG Metall campaign for agency workers. *British Journal of Industrial Relations*, *53*(3), 533–555.

Berthoud, R., & Blekesaune, M. (2007). *Persistent employment disadvantage*. London: Department for Work and Pensions Research. Report No. 416

Blanchflower, D., & Bryson, A. (2008). *Union decline in Britain, CEP discussion paper no 846*. London: London School of Economics.

Blanden, J., Machin, S., & Van Reenen, J. (2006). Have unions turned the corner? New evidence on recent trends in union recognition in UK firms. *British Journal of Industrial Relations*, *44*(2), 169–190.

Blau, F., & Kahn, L. (2000). Gender differences in pay. *Journal of Economic Perspectives*, *14*(4), 75–99.

Booth, A. (1994). *The economics of the trade union*. Cambridge: Cambridge University Press.

Booth, A. (2009). Gender and competition. *Labour Economics*, *16*, 599–606.

Brown, W., & Nash, D. (2008). What has been happening to collective bargaining under New Labour? Interpreting WERS 2004. *Industrial Relations Journal*, *39*(2), 91–103.

Bryson, A. (2005). Union effects on employee relations in Britain. *Human Relations*, *58*, 1111–1139.

Bryson, A. (2008). Union free-riding in Britain and New Zealand. *Journal of Industrial Relations*, *50*(1), 5–24.

Bryson, A. (2016). Health and safety risks in Britain's workplaces: Where are they and who controls them? *Industrial Relations Journal*, *47*(5–6), 547–566.

Bryson, A., Dale-Olsen, H., & Nergaard, K. (2019). Gender differences in the union wage premium: A comparative case study. *European Journal of Industrial Relations*, *26*(2), 173–190.

Bryson, A., & Forth, J. (2017). *Work/life balance and trade unions: Evidence from the Workplace Employment Relations Survey 2011*. London: TUC.

Bryson, A., & Freeman, R. (2013). Employee perceptions of working conditions and the desire for worker representation in Britain and the US. *Journal of Labor Research*, *34*(1), 1–29.

Bryson, A., & Gomez, R. (2003). *Segmentation, switching costs and the demand for unionization in Britain, Centre for Economic Performance discussion paper no. 568*. London: London School of Economics.

Cameron, A. C., & Trivedi, P. K. (2005). *Microeconometrics: Methods and applications*. New York: Cambridge University Press.

Disney, R., Gosling, A., & Machin, S. (1995). British unions in decline: An examination of the 1980s fall in trade union recognition. *Industrial and Labor Relations Review*, *48*, 403–419.

Disney, R., Gosling, A., & Machin, S. (1996). What has happened to union recognition in Britain? *Economica*, *63*, 1–18.

Donado, A., & Walde, K. (2012). How trade unions increase welfare. *Economic Journal*, *122*, 990–1009.

Farber, H., & Krueger, A. (1993). Union membership in the United States: The decline continues. In B. Kaufman & M. Kleiner (Eds.), *Employee representation: Alternatives and future directions* (pp. 70–79). Madison, Wise: IRRA.

Fredman, S. (2004). Women at work: The broken promise of felxicurity. *Industrial Law Journal*, *33*(4), 299–319.

Freeman, R., & Medoff, J. (1984). *What do unions do?*. New York: Basic Books, Inc., Harper TorchBooks.

Freeman, R., & Pelletier, J. (1990). The impact of industrial relations legislation on British union density. *British Journal of Industrial Relations*, *28*(2), 141–164.

Freeman, R., & Rogers, J. (1999). *What workers want*. Ithaca, New York: Cornell University Press.

Healy, G., & Kirton, G. (2000). Women, power, and trade union government. *British Journal of Industrial Relations*, *38*(3), 343–360.

Hotchkiss, J., & Pitts, M. (2007). The role of labor market intermittency in explaining gender wage differentials. *The American Economic Review*, *97*(2), 417–421.

HSE. (2020). *Historical picture statistics in Great Britain, 2020: Trends in work-related ill health and workplace injury*. London: Health and Safety Executive, November 2020.

Kersley, B., Alpin, C., Forth, J., Bryson, A., Bewley, H., Dix, G., & Oxenbridge, S. (2006). *Inside the workplace: Findings from the 2004 Workplace Employment Relations Survey*. London, UK: Routledge.

Machin, S. (2000). Union decline in Britain. *British Journal of Industrial Relations*, *38*(4), 631–645.

Machin, S. (2003). Trade union decline, new workplaces and new workers (Chapter 2). In H. Gospel & S. Wood (Eds.), *Representing workers: Trade union recognition and membership in Britain*. London: Routledge.

Metcalf, D. (2000). 'Fighting for equality', Centrepiece, The Magazine for Economic Performance.

Metcalf, D., Hansen, K., & Charlwood, A. (2001). Unions and the sword of justice: Unions and pay systems, pay inequality, pay discrimination and low pay. *National Institute Economic Review*, *176*, 61–75.

Office for National Statistics. (2014). *Participation rates in the UK labour market: 2014*. London: Office for National Statistics.

Papke, L., & Wooldridge, J. (1996). Econometric methods for fractional response variables with an application to 401(k) plan participation rates. *Journal of Applied Econometrics*, *11*(6), 619–632.

Park, A., Curtice, J., Thomson, K., Jarvis, L., & Bromley, C. (2002). *British social attitudes: 19th report*. Aldershot: Ashgate.

Parker, J. (2002). Women's groups in British unions. *British Journal of Industrial Relations*, *40*(1), 23–48.

Phipps, S., Burton, P., & Lethbridge, L. (2001). In and out of the labour market: Long-term income consequences of child-related interruptions to women's paid work. *The Canadian Journal of Economics*, *34*(2), 411–429.
Pollert, A., & Charlwood, A. (2009). The vulnerable worker in Britain and problems at work. *Work, Employment and Society*, *23*(2), 343–352.
Pulignano, V., Meardi, G., & Doerflinger, N. (2015). Trade unions and labour market dualisation. *Work, Employment and Society*, *29*(5), 808–825.
Riach, P., & Rich, J. (2006). An experimental investigation of sexual discrimination in hiring in the English labor market. *The B.E. Journal of Economic Analysis and Policy*, *6*(2), 1–20.
Roantree, B., & Vira, K. (2018). *The rise and rise of women's employment in the UK, IFS Briefing Note BN234*. London: The Institute for Fiscal Studies.
Schnabel, C., & Wagner, J. (2007). Union density and determinants of union membership in 18 EU countries: Evidence from micro data, 2002/03. *Industrial Relations Journal*, *38*(1), 5–32.
StataCorp. (2019). *Stata Statistical Software: Release 16*. College Station, TX: StataCorp LLC.
van Wanrooy, B., Bewley, H., Bryson, A., Forth, J., Freeth, S., Stokes, L., & Wood, S. (2013). *Employment relations in the shadow of recession: Findings from the 2011 Workplace Employment Relations Study*. London: Palgrave Macmillan.
Vickerstaff, S., Phillipson, C., & Wilkie, R. (2012). *Work health and wellbeing*. Bristol: Policy Press.
Waddington, J., & Whitston, C. (1997). Why do people join unions in a period of membership decline? *British Journal of Industrial Relations*, *35*(4), 515–546.
Wajcman, J. (2000). Feminism facing industrial relations in Britain. *British Journal of Industrial Relations*, *38*(2), 183–201.
WERS. (2011). *Workplace Employee Relations Survey, Department for Business, Innovation and Skills; Advisory, Conciliation and Arbitration Service, and National Institute of Economic and Social Research, Workplace Employee Relations Survey, 2011* (6th ed.). London: UK Data Service (SN: 7226).
Williams, R. (2017). *Analysing proportions: Fractional response and zero one inflated beta models*. Indiana: University of Notre Dame. https://www3.nd.edu/~rwilliam/stats3/FractionalResponseModels.pdf (last accessed 02/11/2020)
Willman, P., Bryson, A., & Gomez, R. (2007). The long goodbye: New establishments and the fall of union voice in Britain. *International Journal of Human Resource Management*, *18*(7), 1318–1334.

SUPPORTING INFORMATION

Additional supporting information may be found online in the Supporting Information section at the end of this article.

How to cite this article: Haile GA. Men, women and unions. *Ind. Relat*. 2021;1–17. https://doi.org/10.1111/irj.12324