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journals.sagepub.com/home/eid **Giedo Jansen**

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Abstract

This article addresses the question of whether, and to what extent job flexibility is detrimental to mobilization with regard to the willingness to take part in industrial action. The authors examine the influence of job flexibility ('standard' versus 'non-standard' work) and job instability (changes from one job to another) on employees' willingness to strike. Based on Dutch survey data it is shown that only minor differences exist between 'standard' and 'non-standard' employees in their willingness to participate in a strike. While this study did not establish a major direct effect of job flexibility on strike participation, tests of interaction effects reveal that job flexibility moderates other mobilizing factors, such as union membership and job dissatisfaction. Job instability, on average, has no effect on strike participation.

Keywords

Atypical employment, fixed-term contracts, participation, strikes, temporary employment

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Introduction

This article addresses the question of whether, and to what extent job flexibility and job instability are detrimental to mobilization for the willingness to take part in industrial actions. This question is answered with regard to the Netherlands, a country with a relatively low strike frequency (Vandaele, 2011) and a relatively flexible labor market (Eurostat, 2013). As in many western post-industrialized economies, employees in the Netherlands increasingly less often have a permanent job (Bierings et al., 2013). Consequently, the ‘standard worker’, with a full-time, open-ended contract, is decreasingly the standard. In no other EU country is the share of companies working with temporary and part-time¹ personnel higher than in the Netherlands (Riedmann et al., 2010). Due to the high proportion of employees working on a part-time basis, the Netherlands was already known as the world’s ‘first part-time economy’ (Freeman, 1998; Visser, 2002a), but in recent years other forms of ‘non-standard’, flexible employment have become more widespread. The portion of the workforce with a temporary job grew from 12% in 2001 to 16% in 2012 (Bierings et al., 2013), including nearly 170,000 temporary agency workers and 200,000 on-call workers (CBS, 2013a). Moreover, this trend towards flexibilization is associated with growing *job instability*: employees with flexible employment contracts not only change employers more often than permanent employees do, but the majority of them regularly change from one flexible employment relationship to the other (Bierings et al., 2013).

Although union participation can take various forms (varying in required time and effort, see Klandermans, 1996), most research on trade union participation focuses on underrepresentation of non-standard workers in terms of union membership (Croucher and Brewster, 1998; De Witte, 2005; Gallagher and Sverke, 2005; Gumbrell-McCormick, 2011; Pernicka, 2005; Vandaele and Leschke 2010). Workers in non-standard, and especially temporary, employment are often considered a problematic group for unions as they are traditionally aimed at ‘male, manual workers with a “normal” employment relationship’ (Berntsen and Lillie, 2014; Gumbrell-McCormick and Hyman, 2013: 33). Joining (and staying in) a union, however, are forms of participation requiring relatively low effort (Klandermans, 1996). The differences between atypical and standard employees are likely to be even more pronounced for participation forms for which the costs are higher in terms of time, effort and risk. Participating in industrial action, such as a strike, is a typical form of participation involving high costs and risks.

The question of whether job flexibility reduces the mobilizing capacity of trade unions may therefore best be addressed with regard to participation in industrial action. To our knowledge there are virtually no studies looking into the effects of job flexibility on (the willingness to participate in) industrial action. On an aggregate level, studying companies in Europe, Jansen and Akkerman (2014) produced mixed results: no significant relationship appeared between strike incidence and the proportion of fixed-term contracts in a firm, but a positive – instead of the expected negative – effect was found for the amount of temporary agency workers with regard to strike incidence in a firm. The few individual-level studies addressing the relationship between industrial action and job flexibility look at this relationship only indirectly, for example via feelings of job insecurity (Goslinga, 2005; Van Vuuren et al., 1991) or via the length of tenure (Buttigieg et al., 2008). These studies thus produce inconclusive results, in particular because it is not

clear whether it is *job flexibility* ('standard' versus 'atypical' work) that is responsible for lower willingness to strike, or whether it is the degree of *job instability* (changes from one job to another) associated with flexible work that actually matters. It is important to distinguish the effects of job flexibility from the effects from job instability because the mechanisms explaining the respective effects are fundamentally different, as we examine in the next two sections. The aim of this article is therefore to disentangle the influence of job flexibility and job instability on employees' willingness to take industrial action, both theoretically and empirically.

Empirically, we examine the difference in the willingness to strike between two types of employment, i.e. 'standard' employment on *permanent contracts*, and *flexible employment* on a temporary basis or working fewer than 12 hours a week. In the Dutch case, where part-time work is widespread (Visser, 2002a), we subdivide the group 'standard' employees into separate categories for those working full-time (35 hours or more) and part-time (between 12 and 34 hours). With respect to *job instability*, we distinguish employees who experienced at least one job change over the past three years from employees without any changes in their labor market position.

Job flexibility and the willingness to strike

In this section we argue that job flexibility leads to a different cost–benefit balance for participation in industrial action, even though job flexibility may be associated with different levels of job satisfaction and therefore different (levels) of grievances about work. Participation in a strike may differ between standard and non-standard employees due to the time span of the costs and benefits associated with participation. First, the *benefits* of a strike may be lower for employees with little or insecure career prospects in the organization. Permanent workers fighting for a wage increase or collective agreement may improve their wage levels and working conditions on a long-term basis. But the gains for temporary workers are limited to the duration of their contract. Unlike permanent employees, temporary workers may therefore be less able to compensate the loss of income during a strike. Likewise, employees with short contracts are less able to compensate in the short run, the loss of income during a strike. Moreover, the benefits of participation are not only lower for employees with temporary contracts, the *costs and risks* associated with going on a strike may also be higher. From a power relations perspective (Gumbrell-McCormick and Hyman, 2013: 30–31; Silver, 2003; Wright, 2000), the risks of participating in industrial action are higher for non-standard employees, in particular for those in the most precarious forms of temporary employment. First of all, they generally possess less *structural power* than employees on standard permanent employment contracts. According to power-resource theories structural power may be based on either scarce skills or competences, or on a strategic position within the production or service delivery process. Although some flexi-workers occupy 'strategic' occupations where strikes may have disruptive consequences for production or services (Gumbrell-McCormick and Hyman [2013] mention the example of dock or construction workers who are sometimes hired by the day or workers in just-in-time production systems), most non-standard workers have a weaker bargaining position compared with their 'standard' counterparts. Especially for temporary agency and on-call workers it

holds that their jobs less often require any specialized skills, as a result of which they are more easily replaced by their employer.² Also, temporary employees generally have less *institutional power* as the legislative support is generally weaker in terms of dismissal protection. Although the new Dutch Work and Security Act (Wetsvoorstel Werk en Zekerheid, 2013: Article 673), taking effect in 2015, will ensure the same dismissal (or ‘transitional’) allowance for flexible workers and standard employees, dismissed workers are entitled to the allowance only after two years of temporary contracts. All in all, considering the costs and benefits of strike participation associated with flexible employment, we formulate the following central hypothesis:

Hypothesis 1: Employees working on non-standard flexible contracts (i.e. on a temporary basis or working fewer than 12 hours a week) have a lower willingness to participate in a strike than employees working on standard permanent (full-time or part-time) contracts.

This hypothesis assumes a direct negative effect of job flexibility on the willingness to participate in a strike. We may, however, expect that job flexibility also moderates other mobilizing factors. Next, we therefore formulate two interaction hypotheses that expect that the differences in structural and institutional power between permanent and flexible workers also indirectly mitigate the mobilization for collective action. For one thing, we know that the driving force behind strike participation is the level of grievances. Dissatisfied workers are assumed to be more militant than workers with high job satisfaction (Klandermans, 1986). Temporary contracts and a structural and institutional disadvantaged position lead to less favorable working conditions for temporary workers (Wilkin, 2013). Solely based on their grievances, temporary workers should therefore have more – instead of less – reason to participate (Wilkin, 2013). However, the effect of grievances may be nullified by the higher costs and lower (short-run) benefits associated with participation for temporary workers.

Hypothesis 2: The positive effects of job dissatisfaction on the willingness to participate in a strike are lower for employees working on non-standard flexible contracts than for employees working on standard permanent contracts.

In terms of associational power, the few studies on union membership among temporary workers generally suggest that temporary workers are indeed less likely to join a union (Nätti et al., 2005; Van den Berg, 1995), and, once a member, are more likely to quit (Van der Putte, 1995). Moreover, union membership has often been linked to the ‘capacity to strike’ as the decision to strike is typically made by trade unions, and primarily union members will strike (Jansen, 2014; Kaufman, 1982; Shorter and Tilly, 1974; Snyder, 1975). However, the relationship is not perfect: not all union members are equally likely to join a strike, and also non-union members may sometimes be willing to participate. In the Netherlands, also non-union members may register as a striker, although they normally receive no strike allowance to compensate their loss of income during a strike, or they receive a lower allowance if they join the union during the strike (Akkerman, 2000). The costs of strike participation are therefore higher for non-members than for union members.

Union membership does not guarantee participation in industrial action. Difference in the willingness to participate in a strike between union members with permanent contracts and those with temporary contracts can be attributed to two factors: differences in members' identification with the union and differences in instrumentality of a strike. Identification with the union strongly affects participation in industrial action (Akkerman et al., 2013; Klandermans and Visser, 1995). The mechanism behind union identification refers to the effect of an individual's feeling of belonging to the union on the norms that the individual will adopt (Cialdini and Goldstein, 2004; Terry and Hogg, 1996; White et al., 1994). Thus, identification with the union stimulates the adoption of the union's norm on participation. By nature of their contract, flexible workers change jobs (more) frequently, which urges them to make frequent switches between unions, especially when unions are still sector based. Because identification is based on 'repeated interaction and sympathy' (Opp, 2001: 240), it is plausible that identification with the union is more difficult to establish for flexible workers.

In addition, union members with flexible contracts may evaluate the strike as being less instrumental to their particular interests and needs which might deviate from the traditional members'. Temporary workers are probably more 'passive' members, with a stronger focus on obtaining personal benefits, but without the 'willingness to act' (Gumbrell-McCormick and Hyman, 2013; Offe and Wiesenthal, 1985). Hence, union membership may be associated with the willingness to strike only among permanent employees and less so among temporary employees.

Hypothesis 3: The positive effects of union membership on the willingness to participate in a strike are lower for employees working on non-standard flexible contracts than for employees working on standard permanent contracts.

Job instability and the willingness to strike

The use of more flexible contracts makes it easier for employers to hire and fire workers, implying an increase in workers' mobility on the labor market. It may therefore be no surprise that there is a positive relationship between job flexibility and job instability, i.e. flexible workers display more unstable employment patterns, moving from one job to the other more often than workers with permanent contracts (Bierings et al., 2013). The question therefore arises whether potential detrimental effects of flexible work on the willingness to take industrial action are due to the temporary nature of contracts, or due to effects of job instability. In this section we disentangle the effects on the willingness to strike of job instability (change of job) from job flexibility (type of contract).

There are indications and arguments that job instability, regardless of the contract type, does not account for the effect of job flexibility on (intended) strike participation. Job stability has been associated with stronger organizational commitment and company attachment (Martin and Sinclair, 2001), which probably reduces the willingness to protest against the organization or its management. Buttigieg et al. (2008) found that the willingness to take industrial action among union members in the Australian banking sector was lower for employees with longer tenure, and thus with a more stable

employment history. However, similar to job flexibility, also job instability has been linked to waning union membership. Visser (2002b) has shown that people experiencing job changes, including career interruptions, are more likely to leave a union. He attributes this effect to weakening social cohesion (or 'broken social ties', 2002b: 407) causing membership norms to deteriorate. Temporary jobs and the lack of long-term professional and social interaction diminish workers' perceptions of social cohesion and their sense of collective interest (Cooper and Kurland, 2002; Golden et al., 2008). Solidarity with one's colleagues not only affects union membership (Klandermans and Visser, 1995), but also the willingness to participate in industrial action (Akkerman et al., 2013). Thus, in order to test whether job instability may account for the negative impact of job flexibility on the willingness to strike, we assume that job instability has a negative effect on collectivist norms with regard to strike participation, in other words, norms propagating worker solidarity and free-rider punishment (Akkerman et al., 2013). This leads to the hypothesis that job instability weakens the likelihood of the willingness to strike:

Hypothesis 4: Employees with unstable jobs have a lower willingness to participate in a strike than employees with stable jobs.

Finally – and assuming that Hypothesis 4 holds – we may formulate a similar interaction hypothesis for the moderating effect of job instability as we have done for job flexibility. Again, we assume that job instability and the breakdown of long-lasting professional ties undermine the willingness to act upon one's grievances or the mobilizing capacity of trade union membership. Hence, we expect that union membership and job dissatisfaction are only – or more strongly – associated with the willingness to strike for employees with stable employment, that is without any recent job changes:

Hypothesis 5: The positive effects of job dissatisfaction on the willingness to participate in a strike are lower for employees with unstable jobs than for employees with stable jobs.

Hypothesis 6: The positive effects of union membership on the willingness to participate in a strike are lower for employees with unstable jobs than for employees with stable jobs.

Data and measures

Data

To test our hypotheses we used a survey of more than 1,000 Dutch citizens conducted in June 2013. The survey was funded by the Netherlands Organization for Scientific Research (NWO) and carried out by MWM2, a professional survey company. The questionnaire of this web-based survey was sent to an existing panel group of Dutch citizens. The panel participants are recruited via the internet and consist of individuals who regularly participate in panels organized by the survey company. The panels are regularly

updated to ensure high response rates and sample sizes. The survey was completed by a total of 1,050 respondents. Given our focus on people in different types of employment contracts, we analyzed only respondents who were currently employed. We therefore excluded people outside wage employment (that is, self-employed, retired, unemployed, students or people primarily doing voluntary work and/or household tasks), as well as respondents with missing information on relevant variables. Ultimately, our empirical analyses are based on 493 employees in wage employment. Below we discuss the measurement of all variables in our models; descriptive statistics are presented in the Appendix.

Key variables

Our dependent variable is the *willingness to participate in a strike*. Respondents were asked to rate their willingness to participate in a strike for a wage increase on a Likert-type scale ranging from 1 (not at all willing) to 5 (very willing to participate). On average, the willingness to participate in a strike is rather low, and the distribution of outcomes is skewed to the right (*skewness* = 0.65; *SE of skewness* = 0.11). We therefore coded this variable into four ordinal categories. In doing so, we merged categories 4 and 5 for respondents who are (very) willing to participate in a strike.

The first main independent variable of interest is *job flexibility*. We distinguished between three types of employment. First, there is the group of so-called ‘standard workers’ consisting of employees with full-time jobs working on a permanent contract or with the prospect of a permanent contract. Following the definition of Statistics Netherlands (CBS, 2013a), ‘full-time’ is defined as working 35 or more hours a week.³ Second comes the group of ‘standard part-time workers’, employees working between 12 and 34 hours with (a prospect of) a permanent contract. The group of flexible workers contains all employees with a fixed-term contract without the prospect of a permanent contract, employees working less than part-time (< 12 hours, including zero-hours contracts) or employees working through a temporary staffing agency or ‘payrolling’ arrangement.

The second main independent variable is *job instability*. Respondents were asked to mention the changes in their labor market position over the past three years, up to three or more changes; that is, a new job with another employer or a new job with the same employer. We distinguished employees who experienced at least one change over the past three years from employees without any changes in their labor market position.⁴ Given our sample of wage employees we discarded all movements out of active wage employment (for example, started their own business, retirement or stopped working for other reasons).

To test the interaction hypotheses we included variables for job dissatisfaction and union membership. To measure respondents’ *job dissatisfaction* five items were used. Respondents were asked how satisfied they were with the following aspects of their current job, ranging from 1 (not at all satisfied) to 4 (very satisfied), i.e. (1) how interesting the job is, (2) the work climate, (3) the management, (4) salary,⁵ (5) job security, (6) possibility to work part-time, (7) possibility to work flexible hours and (8) the possibility to work at home. The latter three items, however, were not applicable to all respondents (up to 30% reported these aspects ‘inapplicable’ to their situation). To construct a reliable scale we therefore used the factor scores from a principal factor analysis (Eigenvalue = 2.40; 36.6% of explained variance) based on the remaining five items (Cronbach’s alpha

of 0.73), that are consistently available across respondents. All items were coded in such a manner that a higher value related to more dissatisfaction.

For *union membership* we use a dichotomized measure. Respondents were asked whether, and if so of what union they are a member. The majority of union members in the Netherlands belong to the Dutch Trade Union Federation (FNV). We collapsed the group of FNV members with members of smaller unions, that is, members of the Christelijk Nationaal Vakverbond (CNV), Unie and the Vakcentrale voor Middengroepen en Hoger Personeel (MHP). In our sample, the union membership rate among permanent employees (22.3%) is comparable to that of the national population (approximately 22%; CBS [2013b]). Contrary to the general population, however, for which there is a difference between flexible and permanent employees in unionization rate, in our sample there is no significant difference in unionization between the two types of employees.

Control variables

To account for other differences among employees and their willingness to participate in a strike, we controlled for other potentially relevant factors, namely, perceived employability and social cohesion at the workplace, i.e. variables related to the cost–benefit calculation (including social benefits such as support from colleagues) employees make when assessing the cost and benefits of going on strike. The anticipated returns of a strike – and therefore the willingness to participate – may be higher for an employee who expects not to be able to find a job elsewhere, and with strong ties to his/her colleagues. Finally, we also control for available social and political background characteristics in the survey (e.g. gender, age, education, social class and political orientation). We control for these variables because historically union membership, and, hence, labor conflict, is generally more strongly associated with male, manual workers in lower-educated and working-class occupations, and the labor movement is historically linked with left-wing politics (Western, 1997).

To measure *perceived employability*, respondents were asked to estimate how long it would take them to find a new, equivalent job, if they started looking for a new job: ‘Less than a month’, ‘1–3 months’, ‘3–6 months’, ‘6 months–1 year’, ‘1–2 years’, ‘2 years or more’, or ‘don’t know’. We distinguished respondents who expected to find a new job within three months (coded 1)⁶ from all other respondents, including those reporting ‘don’t know’ (coded as 0). *Social cohesion* at work is measured using respondents’ answers to three statements on a Likert-type scale ranging from 1 (fully disagree) to 5 (fully agree): (1) ‘My colleagues are my most important social contacts’, (2) ‘Some colleagues are also my best friends’, (3) ‘The people at my work are a tight community’. A reliability test yielded a Cronbach’s alpha of 0.76. A scale was constructed by computing the factor scores from a principal factor analysis (Eigenvalue = 2.04; 53.6% of the explained variance).

Next to social background controls for the *gender* (female = 0, male = 1) and *age* (as its natural logarithm, to account for non-linearity) of respondents, we included their highest completed level of *education* in three categories: lower (elementary at most, lower vocational [LBO], lower secondary [MULO, ULO, MAVO]); middle (secondary vocational [MBO], higher secondary [HAVO, MMS, VWO]); and higher (higher vocational [HBO], university [WO, WO+]). Moreover, we controlled for the size of the company where the respondent is working: *company size* is measured in three categories:

small (fewer than 50 employees), medium (50–500 employees) and large (500+ employees).

Social class was measured on the basis of respondent's occupation following a fairly standard procedure. First, respondents were assigned an occupational code (ISCO-08) based on their occupational title, main tasks and supervisory status. Next, we used the conversion tools of Ganzeboom and Treiman (2009) and Güveli (2006) to transform these occupational codes into 11 categories of the EGP class schema of Erikson et al. (1979). By using Güveli's 'adjusted' version of the EGP schema the 'service class' (EGP classes I and II) is further reclassified into subclasses for 'technocrats' (classes Ia and IIa) and 'social-cultural specialists' (classes Ib and IIb). Finally, we collapsed this elaborate class schema into a version containing four classes: the working class, routine non-manual employees, technocrats and social-cultural specialists.

As an indicator of whether respondents hold a *left-wing* political orientation, we used their reported voting behavior in the past two national elections (those of 2010 and 2012), and their voting intention if elections were held today. In the Netherlands, the political parties commonly defined as left-wing are the PvdA (Labor Party), SP (Socialist Party) and GroenLinks (GreenLeft). Respondents were classified as stable left-wing voters (coded 1) if they reported a vote (intention) for either one of these parties in all three instances. Respondents reporting a vote for another party (or reporting abstention) at least once were coded 0. For younger respondents, their voting behavior was used only for elections for which they were eligible to vote.

Analyses and results

Direct effects

Given the ordinal nature of the dependent variable, ordered logit models were used to estimate the willingness to participate in a strike (see Table 1). The coefficients (b) denote the cumulative log-odds for respondents of having a *higher* versus a *lower* willingness to strike. In the first model we included only the control variables, and in second model we added the variables for job flexibility and job instability. By adding our main variables of interest in Model 2 we are able to assess not only their effect (i.e. regression coefficients and standard errors), but also to what extent they contribute to the model fit (improvement of the pseudo R^2).

Model 1 suggests that a number of variables are significantly associated with the willingness to participate in a strike. First, there appears to be a negative effect of (log-)age (−1.80), suggesting that people are less willing to take industrial action as they get older. The log-transformation, however, indicates that this negative effect diminishes as age increases. The willingness to strike is also lower among employees with higher perceived employability (−0.54). Employees who estimate that they will be able to find a new, equivalent job within three months are less willing to join a strike than employees with less favorable perceptions about finding a new job. Second, positive effects, and thus a greater willingness to strike, are found with respect to employees holding left-wing views (0.57), with strong social ties to their colleagues (0.32) and who are employed in a medium-sized or large company (0.53 and 0.77, respectively). Moreover, the ordered logit estimates indicate a greater willingness

Table 1. Ordered logit regression for the willingness to strike (robust standard errors in parentheses).

	Model 1		Model 2	
	b	(SE)	b	(SE)
Age (log)	-1.80***	(0.31)	-1.94***	(0.32)
Gender (male = 1)	0.22	(0.20)	0.19	(0.21)
Education (Lower = ref.)	—	—	—	—
Middle	0.46	(0.29)	0.41	(0.30)
High	0.46	(0.31)	0.44	(0.32)
Social class (working class = ref.)	—	—	—	—
Routine non-manual class	0.23	(0.28)	0.22	(0.29)
Service class: technocrats	0.11	(0.26)	0.03	(0.26)
Service class: social-cultural specialists	-0.22	(0.38)	-0.24	(0.39)
Left-wing political orientation	0.57***	(0.21)	0.58***	(0.21)
Perceived employability (< 3 months = 1)	-0.54**	(0.21)	-0.54**	(0.21)
Social cohesion at work	0.32***	(0.11)	0.32***	(0.11)
Company size (< 50 employees = ref.)	—	—	—	—
50–500 employees	0.53**	(0.21)	0.49**	(0.21)
500+ employees	0.77***	(0.22)	0.72***	(0.22)
Job dissatisfaction	0.56***	(0.12)	0.56***	(0.12)
Union membership	0.81***	(0.21)	0.84***	(0.21)
Contract (permanent full-time = ref.)			—	—
Permanent part-time			-0.07	(0.23)
Flexible contract			-0.45*	(0.25)
Job instability			-0.07	(0.20)
Thresholds				
Cutpoint 1	-5.91	(1.18)	-6.66	(1.26)
Cutpoint 2	-4.94	(1.17)	-5.69	(1.25)
Cutpoint 3	-3.40	(1.16)	-4.14	(1.24)
McFadden pseudo R^2	0.078		0.081	
-2 log likelihood	1198.82		1194.80	
Likelihood-ratio test versus Model 1			4.02	
d.f. (change in d.f. versus Model 1)	17		20 (3)	
N	493		493	

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$ (two-tailed test).

to participate in a strike among employees who are more dissatisfied with their job (0.56) and who are union members (0.81). Interestingly, however, Model 1 shows no significant results with regard to some commonly used social background characteristics, such as gender, education and social class. In our model there is no difference in the willingness to take strike action between employees in working-class occupations relative to employees in routine non-manual or service class occupations.

In Model 2 we tested Hypotheses 1 and 4 on the direct effects of job flexibility and instability. It appears that adding indicators of contract type and recent job changes to the

model contributed only marginally to the model fit (only 0.003 improvement of the pseudo R^2). Tested against 3 degrees of freedom, the reduction of 4.02 in $-2LL$ was not significant. For Hypothesis 1 to hold, we need to find negative coefficients for part-time and flexible contracts relative to standard full-time contracts. Such an effect, however, is found with regard only to flexible contracts. Employees with flexible working arrangements are, all other factors held equal, about 1.5 times ($1/\exp(-0.45)$) less likely to report a higher willingness to strike compared with permanent full-time workers. This finding is consistent with Hypothesis 1, but the effect is fairly modest and with a relatively high degree of uncertainty, achieving significance at a 90% confidence interval level. Therefore, there is some, but no overt support for Hypothesis 1. For Hypothesis 4 to hold, we need to find a negative effect of job instability on the willingness to participate in a strike. But this effect appears to be insignificant. There is no difference in the willingness to strike between employees who experienced recent job changes and those who remained stable in their career over the past three years. Hence, Hypothesis 4 is rejected.

Interaction effects

In Table 2 we tested the interaction hypotheses. These hypotheses state that the positive effects of job dissatisfaction (Hypotheses 2 and 5) and union membership (Hypotheses 3 and 6) on the willingness to strike should be weaker for employees working on non-standard contracts, and for employees with recent job changes. In Model 3 we tested Hypotheses 2 and 3 about the moderating effect of job flexibility. Including interactions with contract type resulted in a significant improvement of the model's fit (reduction of $-2LL = 16.88$, relative to Model 1, Table 1). As expected, we find positive main effects for job dissatisfaction (0.69) and union membership (1.43), combined with negative interactions with flexible contracts (-0.55 and -1.23 , respectively). These interactions imply that union membership and job dissatisfaction are associated positively with the willingness to strike only among standard employees, and not – or at least less so – among non-standard employees. This particularly holds with respect to employees with flexible contracts, and less so for employees with part-time contracts. Part-time work does not influence the effect of job dissatisfaction on the willingness to strike, and only weakens the effect of union membership (-0.87). The latter effect must be interpreted with caution since the interaction is associated with a relatively high degree of uncertainty, achieving significance at a 90% confidence interval level. The significant moderating effects of flexible work (albeit for the effect of union membership only, not job dissatisfaction) and contractual flexibility is plotted in Figure 1. The figure shows the predicted probabilities for respondents to have the highest outcome on the dependent variable, i.e. being (very) willing to participate in a strike. Consistent with Hypotheses 2 and 3, Figure 1 shows that the effects of job dissatisfaction (Figure 1A) and union membership (Figure 1B) are virtually nullified by flexible employment. Figure 1B also shows that part-time work moderates the effect of union membership on the willingness to strike, but this moderating effect is less strong compared to contractual flexibility.

In Model 4 we tested Hypotheses 5 and 6 about the moderating effect of job instability. Despite these hypotheses, we find no significant interactions between job instability, on the

Table 2. Interaction effects for the willingness to strike based on ordered logit regression.

	Model 3		Model 4	
	b	(SE)	b	(SE)
<i>Main effects</i>				
Job dissatisfaction	0.69***	(0.17)	0.69***	(0.15)
Union membership	1.43***	(0.33)	0.99***	(0.26)
Contract (permanent full-time = ref.)				
Permanent part-time	0.09	(0.26)	-0.09	(0.22)
Flexible contract	-0.15	(0.28)	-0.43*	(0.25)
Job instability	0.05	(0.20)		
<i>Interaction effects</i>				
Job dissatisfaction × Permanent part-time	0.13	(0.28)		
Job dissatisfaction × Flexible contract	-0.55**	(0.26)		
Job dissatisfaction × Job instability			-0.37	(0.24)
Union membership × Permanent part-time	-0.87*	(0.47)		
Union membership × Flexible contract	-1.23**	(0.53)		
Union membership × Job instability			-0.47	(0.42)
McFadden pseudo R^2	0.091		0.084	
-2 log likelihood	1181.91		1190.78	
Likelihood-ratio test versus Model 1	16.88**		8.04	
d.f. (change in d.f. versus Model 1)	24(7)		22 (5)	
N	493		493	

Note: All models controlled for age, gender, education, social class, left-wing political orientation, perceived employability, social cohesion at work and company size (see Table 1).

* $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$ (two-tailed test).

one hand, and the effects of job dissatisfaction and union membership, on the other. Hence, job dissatisfaction and union membership appear to have the same effect on the willingness to participate in a strike for employees who experienced recent job changes and those who remained stable in their jobs over the past three years. Not surprisingly, therefore, including these interactions does not contribute to a better model fit. Tested against 5 degrees of freedom, the reduction of 8.04 in -2LL was not significant.

Conclusion and discussion

Returning to the question posed at the beginning of this study, that is whether job flexibility reduces the mobilizing capacity of trade unions, the results show that flexible employment may in fact reduce mobilization for industrial action. One of the most important findings to emerge from this study, however, is that these detrimental effects are more likely to occur indirectly, rather than directly. The empirical analysis showed that, on average, only minor differences exist between 'standard' and 'non-standard' employees in their willingness to participate in a strike. And where differences do occur, they concern only the dissimilarity between employees on flexible contracts versus permanent full-time workers but not part-time workers. Thus, it seems that employees with

Figure 1. Moderating effect of flexible employment.

(A) Moderating effect on job dissatisfaction effect.

(B) Moderating effect on union membership effect.

Notes: Based on marginal effects, all other variables held at mean.

flexible working arrangements are, in general, only slightly less willing to join a strike compared with permanent full-time workers and workers in part-time employment, the most common form of ‘non-standard’ work in the Netherlands, do not substantially differ from full-time employees in their willingness to join a strike.

While this study did not establish a major direct effect of flexible employment on the willingness to strike, tests of interaction effects revealed that job flexibility moderates other mobilizing factors. The results suggest that job dissatisfaction and union membership are

positively associated with the willingness to participate in a strike only among standard (primarily full-time) employees, whereas there is no such association for flexible employees. Being dissatisfied with their jobs does not advance the mobilization of flexible workers for industrial action, while it does for standard employees. Also for part-time employees the positive association between union membership and the willingness to strike is lower compared to full-time employees. The results of this research therefore support the idea that *job flexibility* may, at least indirectly, undermine mobilization for strikes. The same cannot be said about the effect of *job instability* (versus having a more stable job pattern): the current study was unable to detect a substantially lower willingness to strike among employees with recent job changes, neither directly, nor indirectly through union membership or job dissatisfaction.

The empirical findings of this study add to our understanding of how and to what extent flexible work affects the willingness to strike. As mentioned, the few previous studies all deal with more 'low cost' forms of union participation (for example, membership), and arrive at disparate conclusions on this matter. On one hand, the studies of Van den Berg (1995), Van der Putte (1995) and Visser (2002b) concluded that temporary workers in the Netherlands are less likely to be recruited as union members, and – once a member – more likely to leave the union. Goslinga and Sverke (2003), on the other hand, found no differences in union commitment or union turnover between temporary and permanent workers. It was our expectation that such an effect of contract type would be more pronounced in the case of attitudes towards strikes (as a 'high cost' form of union participation). This expectation, however, receives only limited empirical support.

Of course, the generalizability of the current results is subject to certain limitations. First, in comparison to some other European countries, the Netherlands may be considered a low-strike country. Because of the relatively flexible labor force, the Netherlands constitutes a relevant case to test our hypotheses on the effect of flexible work on strike participation, but it would be interesting to assess these effects also in a more strike-prone context. Another point of consideration is the measurement of our dependent variable. It can be criticized that the current study has only examined the willingness to participate in a strike, not actual participation, which is not the same (Gallagher and Strauss, 1991; McClendon and Klaas, 1993). Our data, however, do not contain information on respondents' past strike participation. Although future studies might investigate the effect of job flexibility on actual participation, this will introduce new complexities regarding causality. In cross-sectional studies, there may be a disparity between the time of data collection and the retrospective reports of strike participation. This may be problematic when linking participation to job characteristics, especially with respect to employees on short and temporary contracts.

Goslinga and Sverke have proposed that the lack of (large) differences in union attitudes may be due to changing union strategies towards 'non-standard' workers, in which union policies have become more favorable and open to employees working on temporary and flexible contracts (2003: 308). Although new strategies and services may offer an explanation for union membership among flexible workers, it may be more complex to account for the willingness to take industrial action. Even if flexible workers are unionized, our results suggest that they are still fairly difficult to mobilize in the case of a strike. But specific strategies and resources may help. A recent UK case study has emphasized that successful mobilization of non-standard workers may be strongly dependent on union representatives (Simms and Dean, 2014). According to Simms and Dean, shop stewards, but also union

officials, may play a crucial role in framing collective interests and building solidarity among groups of temporary workers. The current study, however, was unable to evaluate factors related to the union presence at the workplace (Jansen, 2014). Further research is therefore needed to test these and other factors facilitating mobilization among flexible workers, not only in the Dutch context, but preferably also in a comparative perspective.

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Notes

1. The results of Riedmann et al. (2010) on part-time work are based on the European Company Survey 2009. In this survey the proportion of part-time work in a company is estimated by a management respondent responsible for the personnel in the organization. Part-time work is defined as 'all working-time arrangements below the usual full-time level'.
2. In the international literature, it is sometimes suggested that employers even use temporary workers as a weapon against organized labor (Hatton, 2014). Although employers in the Netherlands are able to allocate other employees to do the work of strikers, it is forbidden by Dutch law (article 10 of the Wet Allocatie Arbeidskrachten door Intermediairs, i.e. the Labor Law Allocation by Intermediaries) for employers to hire temporary agency workers as strike-breakers.
3. Statistics Netherlands uses 35 hours as the cut-off between full-time and part-time work, but national definitions for part-time work vary across countries. Sometimes, for international comparisons, 30 hours is recommended as a harmonized cut-off point (Van Bastelaer et al., 1997). Applying a 30 hour cut-off instead of 35, however, does not substantially alter the results presented in this study.
4. Job changes are not necessarily linked to job insecurity and uncertainty, as changes from one job to another may relate to voluntary choices of employees and 'job hopping' behavior. Replacing job instability with a more narrow measure, only including changes between employers, leads to broadly similar results. Using this alternative measure the direct effect in Model 1 is insignificant ($b = 0.53, p = 0.704$), as is the interaction effect in Model 4 with union membership ($b = -1.00, p = 0.772$). Moreover, also the interaction between job instability and job dissatisfaction now loses significance ($b = -1.22, p = 0.467$).
5. Although the dependent variable is specifically about strikes for a wage increase, we here measure general job dissatisfaction. When only wage dissatisfaction is included this leads to a similar, but somewhat weaker effect on the willingness to strike (both in terms of effect size and contribution to explained variance).
6. Under current Dutch law, the so-called 'chain-provision' states that after three years or three consecutive contracts a temporary contract becomes permanent unless there has been an interruption of three months or more (Keizer, 2011). From July 2015, the duration of the 'chain-interruption' will be expanded to six months, and the maximum number of consecutive contracts will be limited to two.

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Appendix

Table A1. Descriptive statistics.

	Frequency	%	Cumulative %	
<i>Dependent variable</i>				
Willingness to strike				
[0] Not at all willing	196	39.8	39.8	
[1] Not willing	101	20.5	60.2	
[2] Not willing, not unwilling	124	25.2	85.4	
[3] (Very) willing	72	14.6	100.0	
	Minimum	Maximum	Mean	SD
<i>Independent variables</i>				
Age (log)	2.89	4.25	3.66	0.32
Gender (male = 1)	0.00	1.00	0.55	0.50
Lower education	0.00	1.00	0.16	0.37
Middle education	0.00	1.00	0.42	0.49
Higher education	0.00	1.00	0.42	0.49
Working class	0.00	1.00	0.19	0.39
Routine non-manual	0.00	1.00	0.25	0.43
Technocrats	0.00	1.00	0.43	0.50
Social-cultural specialists	0.00	1.00	0.13	0.34
Union membership	0.00	1.00	0.23	0.42
Left-wing political orientation	0.00	1.00	0.19	0.39
Social cohesion at work	-1.75	2.39	-0.01	0.90
Job dissatisfaction	-1.32	3.40	0.03	0.86
Perceived employability	0.00	1.00	0.30	0.46
Company size small	0.00	1.00	0.30	0.46
Company size medium	0.00	1.00	0.35	0.48
Company size large	0.00	1.00	0.34	0.47
Standard full-time worker	0.00	1.00	0.48	0.50
Standard part-time worker	0.00	1.00	0.32	0.47
Flexible worker	0.00	1.00	0.20	0.40
Job instability	0.00	1.00	0.35	0.48

N = 493.