The relationship between job tenure and work disability absence among adults: A prospective study

F. Curtis Breslin [a,b,*], Emile Tompa [a], Ryan Zhao [a], Jason D. Pole [a],
Benjamin C. Amick III [a,c], Peter M. Smith [a,d], Sheilah Hogg-Johnson [a]

[a] *Institute for Work & Health, Toronto, Canada*
[b] *Department of Public Health Sciences, University of Toronto, Toronto, Canada*
[c] *School of Public Health, University of Texas Health Science Center at Houston, Houston, TX, USA*
[d] *Institute of Medical Science, University of Toronto, Toronto, Canada*

Abstract

Little population-based, prospective research has been conducted to examine the demographic and work-related determinants of occupational injury or illness. This study examined the relative contribution of sociodemographic characteristics and work factors to the likelihood of a work-related disability or illness. In a representative sample of adult Canadians 25–70 years old from a prospective survey, a hazard modelling approach of time to work disability absence from the start of a new job was estimated with the following predictors: age, gender, type of job (manual, non-manual, and mixed), hours worked, highest education achieved, multiple concurrent job, job tenure, school activity, union membership and living in a rural or urban area. Workers holding manual or mixed jobs and having a low education level were factors independently associated with the increased likelihood of a work disability absence. Gender was not independently associated with work disability absences. A strong job tenure gradient in the unadjusted work disability absence rates was virtually eliminated when controlling for demographic/individual and other work factors. In multivariate analyses, work-related factors remained predictors of work disability absence whereas individual characteristics such as gender did not. The exception was workers with less education who appeared to be particularly vulnerable, even after controlling of physical demands on the job. This may be due to inadequate job training or increased hazard exposure even in the same broad job category.
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1. Introduction

Work-related injury and illness remains a significant source of morbidity in developed countries. In the U.S., a total of 4.3 million injuries were reported in private industry in 2004, representing 4.8 injuries per 100 full-time workers (Bureau of Labor Statistics, 2005). In Canada work-related injury/illness are also a public health concern, with slightly less than 1 million occupational injuries and illnesses being reported in 2004, representing approximately 1 work-related event for every 38 Canadian workers (Association of Workers' Compensation Boards of Canada,

2006). The purpose of the current study was to examine the relative contribution of sociodemographic and work factors to work-related injury/illness, with an emphasis on job tenure (i.e., number of months on the job) as a predictor.

Conceptual frameworks and quantitative research suggests that several sociodemographic and work factors are associated with work-related injuries and illnesses (DeJoy and Southern, 1993; Amick and Lavis, 2000). Subgroups within the working population often found to be at elevated risk are males, younger workers, and those with low education, even in multivariate analyses (Breslin and Smith, 2005; Strong and Zimmerman, 2005; Oh and Shin, 2003; Leigh, 1986).

The type of job and the related exposure to hazards clearly play an important role. Workers in manual jobs (e.g., labourers and handlers) and jobs in the goods producing sectors (e.g., agriculture, manufacturing, construction) have higher injury rates than workers in sales, administration and professional jobs

* Corresponding author at: Institute for Work and Health, 481 University Ave., Suite 800, Toronto, Ont., Canada M5G 2E9. Tel.: +1 416 927 2027x2225; fax: +1 416 927 4167.
E-mail address: cbreslin@iwh.on.ca (F.C. Breslin).

(Breslin and Smith, 2005; Strong and Zimmerman, 2005). Given that musculoskeletal disorders represent a large majority of work injuries (Mustard et al., 2003), it appears that different physical demands contribute substantially to this pattern of risk.

With the current labor market shifting towards more temporary employment, it is increasingly important to accurately estimate the relationship between work injury risk and job tenure. Many cross-sectional studies have shown that newly hired workers are more likely to be injured than those with longer job tenures (Breslin and Smith, 2006; Cellier et al., 1995; Butani, 1988). This high risk for newly hired workers could be explained by a combination of unfamiliarity with identifying and controlling hazards, a lack of effective safety training, and exposure to more hazards at the start of a job (Breslin and Smith, 2006).

Consistent with this notion, a 12-month prospective study of workers in manual jobs found an association between job tenure and work injuries when individual (e.g., age and gender) and work factors (e.g., type of manual job) were controlled (Leigh, 1986). Whether this relationship is more broadly applicable remains to be determined.

Working longer hours is also associated with work-related injury/illness risk (Leigh, 1986), possibly due to fatigue or simply a function of increased exposure time (Frone, 1998). Another variable related to time exposure is holding multiple jobs concurrently. Paradoxically, multiple job holding has been associated with decreased likelihood of work-related injury/illness (Strong and Zimmerman, 2005), even with number of work hours controlled. This finding may be due to primarily healthier workers taking on more than one job. In addition, this finding may indicate that individuals in different situations may have different thresholds for taking time off due to an injury/illness.

Union membership may also influence the risk of a work-related injury/illness. For example, unionized workers report receiving more safety training than their nonunionised counterparts (Vector Research and Workers Health and Safety Centre, 2002). Union protection from reprisals may either make one more likely to refuse unsafe work, or be more likely to take time off for an injury/illness than unionized workers. However, a recent prospective study showed no significant association between union membership and work-related injury/illness rates (Strong and Zimmerman, 2005).

Several methodological issues limit the conclusion of previous studies. First, many studies have examined correlates of work injury and disability in particular industries and identified possible risk factors in those industries (Bennett and Passmore, 1984; McCurdy et al., 1991; Cellier et al., 1995), but these studies may be limited in their generalization to other industries.

Second, for studies using administrative data sources, underreporting may be an issue. For example, studies which rely on workers' compensation claims could fail to capture all work injuries, especially if filing a claim might affect a firm's insurance premium or increase their risk of being inspected (Mustard et al., 2003). Underreporting of claims could bias risk factor estimates if the likelihood of reporting injuries to the compensation system differs by industry or by worker characteristics. Furthermore, in many jurisdictions not all industries are covered by workers' compensation, therefore only a subset of the working population is captured by such studies. Relying on health records (i.e., health care visits) as a data source can also be problematic and lead to reporting bias, since not all injured workers seek health care services. Research shows that only 34% of occupational injuries are treated in emergency departments (Centers for Disease Control and Prevention, 1998). These biases towards underreporting in administrative sources underscore the need for self report data as an important additional source of work-related injury/disability data.

Finally, very few cross-industry studies have utilized longitudinal data to examine risk factors for work-related injury (for exceptions see Strong and Zimmerman, 2005; Oh and Shin, 2003; Leigh, 1986). The use of longitudinal data is particularly important when investigating time varying predictors such as job tenure (i.e., months since the start of the job).

We used longitudinal data from a nationally representative sample of adult workers to examine the relationship between individual, job, and temporal factors and self-reported work absence due to a work-related injury/illness. Specifically, we investigate the role of job tenure on health risk exposure at work using a hazard modelling approach of time from starting a job to first exit due to a work-related injury/illness.

2. Methods

2.1. Data

The Survey of Labour and Income Dynamics (SLID) consists of a series of 6-year overlapping panels (each of the 6 years is considered a wave within a panel), with a new panel being introduced every 3 years to replace the oldest panel (Statistics Canada, 1997). Available data from these panels cover the time period from 1993 to 2003 inclusive. Respondents are interviewed annually about their work experiences over the previous calendar year in up to six jobs and report monthly on the characteristics of each job. These secondary data analyses were approved by the Health Sciences I ethics committee at the University of Toronto.

2.2. Population

The target population for SLID is all persons living in Canada, excluding people in the Yukon or Northwest Territories, residents of institutions, persons living on Reserves, and full-time members of the Canadian Armed Forces living in barracks. The initial representative sample of each panel is drawn from respondents of Statistics Canada's Labour Force Survey. For each panel, the sample is approximately 15,000 households, comprising about 31,000 persons aged 16 years and over.

This analysis focused on respondents who were between 25 and 70 years old and began a new job at some point during the observation period.

2.3. Outcome: work disability absence

The outcome measure (event) was the first occurrence of a work disability absence (WDA) for a person in a given job. A

WDA was defined as any absence for 1 week or greater due to a work-related illness or disability. While the nature of the illness/disability is not specified in the questionnaire, this self-report WDA question was found to produce similar time trends when compared to workers' compensation claim data in a Canadian province (Mustard et al., 2003), with the main types of claims coming from injuries (e.g., falls) and musculoskeletal disorders. This finding provides evidence of concurrent validity for our WDA measure.

For each respondent, every new job that started over the six waves of the panel provided a new person-job episode to be used for the analysis. However, person-job episodes that started prior to the respondent entering the panel were not included in the analysis due to difficulties in determining exposure time and because a WDA could have occurred prior to panel entry. A total of 64,299 person-job episodes were excluded due to this type of censoring. A logistic regression examining the likelihood of this type of censoring showed women were slightly more likely to be excluded compared to men. Also, as age increased the likelihood for selection of the job into the final model decreased. The age effect is expected due to older workers being engaged in the workforce when the observation period started.

WDAs may not be attributable to exposures in the particular job from which an absence was taken. For example, if an individual held multiple concurrent jobs at the time of the WDA, an absence could be recorded in one or more of the multiple jobs. Given this, we attributed the work absence of respondents who held multiple concurrent jobs at the time of a WDA to the job identified as the main-job for that month (the job with highest usual hours for that month).

For person-job episodes with a WDA, all months of observation after the month of the WDA were excluded. For respondents who held multiple concurrent jobs at the time of a WDA, observed months in all concurrent jobs after the month of the WDA were excluded.

2.4. Covariates

Covariates for each person-job episode included in the model were age, gender, job type, hours worked, highest education achieved, having multiple concurrent jobs, job tenure, school activity, living in an urban versus rural area, union membership, and a flag indicating what panel the person-job episode occurred.

Age was considered a time-varying covariate, and was categorized into six broad groups, 25–29, 30–34, 35–39, 40–44, 45–49 and 50+ years old, with the oldest group serving as the referent.

Jobs were classified into three categories of physical demands, manual, mixed, and non-manual (referent group), using the Institut de recherche Robert-Sauvé en santé et en sécurité du travail (IRSST) occupation coding system (Herbert et al., 1996). The coding system is based on observations in workplaces and agreement among experts in the occupational health and safety field on the typical frequency of handling loads and the weight of the load.

Hours worked in a given month for a particular job were assessed by multiplying the average number of paid hours in a week by the number of weeks worked in that month. Total hours were categorized into four groups (0–60 [referent], 61–120, 121–160 and 161 or more hours per month), with the largest category representing 40+ hours per week.

Highest level of educational attainment was identified with a three-category variable (no high school diploma, high school diploma, post secondary degree), based on educational attainment as of December of each year.

Job tenure was calculated as the number of months between the start date of the person-job and the current month of observation. To examine nonlinear relationships, job tenure was divided into six groups (0–2, 3–4, 5–6, 7–12, 13–24, 25+ [referent] months)

A flag indicating current full-time school activity in a given month of observation was used to assess the differences between those respondents who were balancing work and school commitments compared to those who were not.

Given the potential differences in employment patterns and the varying types of jobs held by adults living in urban environments compared to living in rural environments, a flag indicating urban residence was used in the model.

An indicator of the panel (1–4) from which the data arose was also used in the model to control for any difference that may exist between the different panel cohorts.

2.5. Analysis

We used a complementary log–log model for continuous time processes, as described by Allison (1995). This model is based on the assumption that work absences are generated by a proportional hazards model. Each person-job episode was broken down into monthly intervals. Therefore, a single person-job episode could contribute a maximum of 72 months of observation over the 6-year observation period. Within these monthly intervals, the assumption is of an exponential hazard. Over the complete set of observations for each person-job episode, however, there is no assumption about the overall shape of the hazard. This modelling framework allows for time-dependent covariates.

A list-wise deletion was employed, deleting any observation that did not provide complete responses to all covariates. All covariates were entered in the analysis model simultaneously.

Given the complex nature of the survey sampling, we calculated robust variance estimates using a weighted bootstrap method with 1000 replicates to adjust variance estimates. The bootstrap technique assumes independence across respondents when estimating the coefficients but accounts for the true non-independence of respondents when estimating the variance of coefficients (Yeo et al., 1999). The weights take into account sampling probabilities and non-response. The weight from the first wave of each panel was used for all subsequent waves of data.

3. Results

The adult respondents who had complete covariate information and met the study inclusion criteria provided 841,133 person-months of observation from 64912 jobs. During the

Table 1
Number of jobs, person-months of observation and work disability absence rate, weighted

	Jobs	Person-months		Work disability absence	
	N	N	%	N	Rate per 1000 person-months
Total	64912	841133	100.00	894	1.06
Gender					
Female	32148	419622	49.89	384	0.92
Male	32764	421511	50.11	510	1.21
Age (years)					
25–29	13147	139995	16.64	162	1.16
30–34	11178	153700	18.27	167	1.09
35–39	11212	154267	18.34	154	1.00
40-44	10036	134754	16.02	176	1.31
45–59	7701	107456	12.78	99	0.92
50+	11638	150962	17.95	136	0.90
IRSST job type					
Manual	27828	275582	32.76	480	1.74
Mixed	15760	206500	24.55	226	1.09
Non-manual	21324	359051	42.69	188	0.52
Total hours worked (month)					
0–60	9872	113157	13.45	30	0.26
61–120	10802	132518	15.75	130	0.98
120–160	31063	448006	53.26	548	1.22
162+	13175	147452	17.53	186	1.26
Highest education					
Less than high school	11975	112738	13.40	212	1.88
High school	17139	216975	25.80	267	1.23
Post-secondary	35798	511420	60.80	415	0.81
Multiple concurrent job					
No	50284	683213	81.23	814	1.19
Yes	14628	157920	18.77	80	0.51
Tenure (month)					
0–2	12053	24350	2.89	107	4.40
3–4	9384	39063	4.64	78	2.01
5–6	6536	36154	4.30	79	2.19
7–12	15790	164684	19.58	194	1.18
13–24	11262	212995	25.32	248	1.16
25+	9887	363887	43.26	190	0.52
School activity					
No	61386	795604	94.59	869	1.09
Yes	3526	45529	5.41	25	0.55
Living in urban area					
No	19828	162201	19.28	199	1.22
Yes	45084	678932	80.72	696	1.02
Unionized					
No	51976	653004	77.63	565	0.87
Yes	12936	188129	22.37	329	1.75

observation period there were a total of 894 WDAs observed. The number of jobs, person-months of observation, number of WDAs and the WDA rate per 1000 person-months is provided in Table 1 for sociodemographic and work-related factors. Particularly striking is the large risk gradient for job tenure. The work disability absence rate at 0–2 months on the job was over eight times the rate at greater than 25 months on the job (4.40 and 0.52 per 1000 person-months, respectively).

Table 2 provides the number of jobs, person-months of observation, number of WDAs and the WDA rate per 1000 person-months for each age and gender category, while Table 3 provides the same information for each age and tenure category. With the exception of the 40–49-year old age category, males consistently have a higher WDA rate compared to females in the same age category. When examining tenure by age, the overall tenure gradient noted in Table 1 is pervasive in each age category.

3.1. Multivariate regression results

Results of the fully adjusted multivariate regression analysis examining the antecedents to WDA are presented in Table 4 including the hazard ratio (HR) and the corresponding 95% con-

Table 2
Number of jobs, person-months of observation and work disability absence rate by age and gender, weighted

Age (years)	Gender	Jobs	Person-months		Work disability absence	
		N	N	%	N	Rate per 1000 person-months
		64912[a]	841133[a]	100.00[a]	894	1.06[a]
25–29	Female	6469	69901	8.31	56	0.81
	Male	6678	70094	8.33	106	1.51
30–34	Female	5610	74432	8.85	79	1.06
	Male	5568	79268	9.42	88	1.11
35–39	Female	5764	78291	9.31	65	0.83
	Male	5448	75975	9.03	89	1.17
40–44	Female	5236	71855	8.54	83	1.15
	Male	4800	62898	7.48	93	1.49
45–49	Female	3934	55436	6.59	57	1.03
	Male	3767	52020	6.18	42	0.80
50+	Female	5135	69707	8.29	44	0.63
	Male	6503	81255	9.66	92	1.13

[a] Total.

Table 3
Number of jobs, person-months of observation and work disability absence rate by age and job tenure, weighted

Age (years)	Tenure (months)	Jobs	Person-months		Work disability absence (rate per 1000)	
		N	N	%	N	Person-months
		64912[a]	841133[a]	100.00[a]	894	1.06[a]
25–29	0–2	2527	5814	0.69	24	4.10
	3–4	2107	9094	1.08	21	2.36
	5–6	1332	7769	0.92	15	1.99
	7–12	3036	29475	3.50	42	1.43
	13–24	2311	37576	4.47	31	0.82
	25+	1834	50267	5.98	29	0.57
30–34	0–2	2127	4318	0.51	17	3.95
	3–4	1711	7639	0.91	14	1.84
	5–6	1139	6637	0.79	14	2.13
	7–12	2722	30608	3.64	43	1.42
	13–24	1873	39068	4.64	43	1.09
	25+	1606	65430	7.78	35	0.54
35–39	0–2	2065	4009	0.48	12	2.96
	3–4	1665	7193	0.86	9	1.27
	5–6	1146	6491	0.77	21	3.24
	7–12	2603	28036	3.33	34	1.21
	13–24	1986	39239	4.67	42	1.07
	25+	1747	69298	8.24	36	0.52
40–44	0–2	1815	3541	0.42	20	5.71
	3–4	1409	5478	0.65	17	3.15
	5–6	983	4903	0.58	16	3.18
	7–12	2412	25180	2.99	24	0.96
	13–24	1770	32276	3.84	62	1.93
	25+	1647	63376	7.53	37	0.58
45–49	0–2	1375	2636	0.31	17	6.52
	3–4	1009	3846	0.46	6	1.68
	5–6	757	4016	0.48	6	1.45
	7–12	1913	19603	2.33	19	0.97
	13–24	1375	27233	3.24	29	1.07
	25+	1272	50122	5.96	21	0.42
50+	0–2	2144	4032	0.48	17	4.25
	3–4	1483	5815	0.69	10	1.73
	5–6	1179	6338	0.75	7	1.16
	7–12	3104	31782	3.78	28	0.89
	13–24	1947	37601	4.47	41	1.10
	25+	1781	65393	7.77	32	0.48

[a] Total.

Table 4
Estimates of hazard ratios for work disability absence[a]

	Parameter estimate	Standard error	Hazard ratio	95% CI	Percent change
Intercept	−8.625	0.335			
Gender					
Females			1.00		
Males	−0.069	0.098	0.93	0.77–1.13	−6.70
Age (years)					
25–29	0.133	0.169	1.14	0.82–1.59	14.26
30–34	0.089	0.166	1.09	0.79–1.51	9.33
35–39	0.121	0.157	1.13	0.83–1.53	12.82
40–44	0.260	0.183	1.30	0.91–1.86	29.67
45–49	0.152	0.170	1.16	0.83–1.62	16.39
50+			1.00		
IRSST job type					
Non-manual			1.00		
Manual	1.002	0.132	2.72	2.10–3.53	172.40[b]
Mixed	0.731	0.140	2.08	1.58–2.73	107.72[b]
Total hours worked					
0–60			1.00		
61–120	1.151	0.263	3.16	1.89–5.29	216.20[b]
120–160	1.362	0.246	3.91	2.41–6.32	290.52[b]
161+	1.400	0.258	4.06	2.45–6.72	305.52[b]
Highest education					
Less than high school			1.00		
High school	−0.237	0.130	0.79	0.61–1.02	−21.12
Post-secondary	−0.482	0.129	0.62	0.48–0.80	−38.23[b]
Multiple concurrent job					
No			1.00		
Yes	−0.418	0.171	0.66	0.47–0.92	−34.14[b]
Tenure (months)					
0–2	−0.095	0.164	0.91	0.66–1.25	−9.08
3–4	−0.165	0.167	0.85	0.61–1.18	−15.19
5–6	−0.003	0.181	1.00	0.70–1.42	−0.29
7–12	0.051	0.138	1.05	0.80–1.38	5.23
13–24	0.165	0.140	1.18	0.90–1.55	17.94
25+			1.00		
School activity					
No			1.00		
Yes	−0.206	0.251	0.81	0.50–1.33	−18.62
Living in urban area					
No			1.00		
Yes	−0.056	0.109	0.95	0.76–1.17	−5.47
Union					
No			1.00		
Yes	0.634	0.100	1.89	1.55–2.30	88.57[b]

[a] In addition, we controlled for which of the four panels the observations were drawn from.
[b] $p < .05$.

fidence interval (CI). The final column of the table represents the percentage deviation of the hazard ratio from the null value (1.00).

Notable in terms of temporal factors was that the unadjusted elevated risk for new workers was virtually eliminated when other risk factors were controlled. Other temporal factors such as total hours worked per month (160 h/month = 40 h week) showed a dose–response relationship with increasing hours and increasing risk. Jobs with monthly hours between 61 and 120 h had

a 216% increase in the likelihood of a WDA compared to jobs with hours between 0 and 60 (HR: 3.16, CI: 1.89, 5.29). A 291% increase was seen for jobs in the 120–160 h range (HR: 3.91, CI: 2.41, 6.32) with a 306% increase for jobs working greater than 160 h per month (HR: 4.06, CI: 2.45, 6.72) compared to jobs with hours between 0 and 60.

Manual occupations were 172% more likely to experience a WDA compared to non-manual occupations (HR: 2.72, CI: 2.10, 3.53) while mixed occupations had a 108% increase (HR:

2.08, CI: 1.58, 2.73). Respondents who had more than one job concurrently had lower odds of a WDA than respondents working just one job (HR: 0.66, CI: 0.47, 0.92). Being a member of a union was associated with increased odds of experiencing a WDA (HR: 1.89, CI: 1.55, 2.30).

With regard to demographic/individual characteristics, respondents who had completed high school or post-secondary education were seen to decrease their risk of a WDA by 21–38% (HR: 0.79, CI: 0.61, 1.02 and HR: 0.62, CI: 0.48, 0.80, respectively). Although unadjusted work absence rates for men were higher than women, when other risk factors were controlled, no gender differences were observed.

4. Discussion

This study of work injury/illness has several methodological strengths. Our study was based on a nationally representative sample that was followed for a 6-year period, allowing for the temporality of exposures and outcomes to be specified—a methodological necessity for identifying risk factors (Côté et al., 2001). Indeed the specificity of month-to-month information on work and health measures for each job over this time period is unmatched compared to previous longitudinal studies. In addition, our analyses included a comprehensive range of demographic/individual and work-related variables whereas previous research included only subsets of these variables (e.g., Cellier et al., 1995).

Job characteristics were the strongest predictors of work disability absence. Factors such as the type of job and the associated physical exposures have been identified as important variable mediating the risk of occupational injury/illness in other prospective studies (Strong and Zimmerman, 2005; Oh and Shin, 2003; Leigh, 1986). As would be expected, there was a positive association between hours worked per month and occupational injury/illness absence, with over time work (>40 h/week) showing a 300% increase in risk. However, we also observed that working more than one job significantly reduced the odds of having an occupational injury/illness absence. The fact that work hours and multiple job holding worked in different directions suggest that risk is not simply a function of on-the-job exposure (Strong and Zimmerman, 2005). There may be some selection effects driving these results, for example individuals choosing to take on multiple concurrent jobs may be in better health or have a higher threshold for absence taking. There may also be differences in the job tasks performed by multiple job holders not captured by our occupational categories (i.e., residual confounding). Our ability to clearly interpret this finding is limited, however, because of the survey's lack of specificity regarding which of the multiple concurrent jobs led to the work absence.

Factors such as work experience are recognized in the literature as being risk factors for occupational injury/illness. Consistent with cross-sectional, descriptive studies (Breslin and Smith, 2006; Cellier et al., 1995), we found the unadjusted WDA rates show workers starting new jobs to be at substantially higher risk than their counterparts with longer job tenure. However, this association was eliminated in our multivariate analysis, suggest-

ing that the elevated rates of short-tenure reflects confounding with other risk factors such as type of job. It should be noted however, that the outcome for previous studies were lost-time claims that may not have included a full week work absence as in the present study – making generalizability to less severe occupational injury/illness a potential issue. Another useful topic for future research to clarify this relationship would be to examine which factors influence job tenure.

Our results also showed that human capital, as indicated by education, is predictive of work disability absence (Oh and Shin, 2003). Since this analysis controlled for other demographic and work factors, it is more than finding that better educated people end up with safer jobs. For example, within our broad job categories, people with more education may have less hazard exposures than people with low education levels. Also, people with more education may reduce the risk of injury/illness by having better access to training and information regarding work hazards, or be more able to refuse unsafe work. Consequently, future research might investigate the role of residual confounding since we controlled only for work hours and one of several job characteristics (physical demands) that may be associated with both injury risk and level of education.

We found males had a higher unadjusted work disability absence rate compared to females. However, the adjusted odds of a work disability absence showed no significant gender differences. The lack of significance in the multivariate model suggest that gender differences are due to differential work exposures of male and female workers.

Our findings need to be interpreted in the light of certain methodological limitations. We cannot rule out the possibility that a WDA was a recurrence of a condition that arose in a previous job prior to entry into the panel. In addition, error could have been introduced into the reporting of WDAs by employers' policies or financial pressures on the worker to not take a work absence. For example, temporary workers have been found to have a higher risk of occupational injuries compared to permanent employees, but their work absence taking was lower (Virtanen et al., 2005). Having other factors besides the occurrence of an injury/illness affecting work absence taking is a measurement issue that also exists for compensation claim data (Shannon and Lowe, 2002). Given that our unadjusted work disability absence rates were highest at the beginning of a job episode, however, suggests that this absence taking threshold issue was not a major influence on our indicator of work injury/illness in this population-based sample.

4.1. Implications and summary

Using prospective data on a representative sample of workers, we found that the significance of characteristics such as gender and job tenure were eliminated when we controlled for job and other sociodemographic characteristics. Our findings support the notion that a means by which to reduce WDA among adult workers is by focusing on reducing the physical hazards they encounter at work, potentially through technical safety improvements in the equipment and work environment as well as monitoring overtime hours to decrease fatigue.

Workers with less education appear to be a particularly vulnerable population, possibly due to inadequate job skills and or because they engage in more dangerous tasks within similar jobs (i.e., residual confounding of hazard exposures). Our results suggest that education operates both through occupational characteristics (e.g., physical demands of job) and directly. Consequently, organizations should consider specific job training and injury prevention programs targeting this subgroup of workers may be warranted.

In summary, this prospective study indicates that job characteristics are a key risk factor in determining occupational health among adult workers. This study also contributes methodologically to the occupational injury/illness literature by modelling the complex work patterns arising from increases in temporary work and multiple jobs.

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