Original article

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Sickness absence and mental health: evidence from a nationally representative longitudinal survey

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Objectives Previous studies have consistently reported evidence of large significant associations between measures of psychological health and sickness absence. Some of this association, however, may be confounded by relevant covariates that have not been controlled. By using data with repeated observations from the same individuals, this study aimed to quantify the bias due to unobserved characteristics that are time invariant.

Methods Longitudinal data from the Household, Income, and Labour Dynamics in Australia (HILDA) Survey were used to estimate negative binomial regression models of the number of annual paid sickness absence days. Observations spanning the period 2005–2012, and covering all employed persons aged 15–64 years, were used (56 348 observations from 13 622 individuals).

Results Significant associations between the number of paid sickness absence days taken each year and scores on the mental health subscale of the SF-36 (MHI-5) were found. Inclusion of correlated random effects (which effectively control for unobserved person-specific factors that do not vary over time), however, resulted in a marked decline in the magnitude of this association. For persons with severe depressive symptoms (MHI-5 ≤52), the estimated incidence rate ratios were in the range 1.13–1.14 for men and 1.10–1.12 for women.

Conclusions Poor mental health is a risk factor affecting work attendance, but the magnitude of this effect, at least in a country where the rate of sickness absence is relatively low, is modest.

Key terms absenteeism; Australia; employment; HILDA Survey; panel data; psychological health; sick leave.

Mental health disorders are increasingly recognized as a global health problem, affecting the health and productivity of millions of individuals. According to the Organization for Economic Cooperation and Development, around one in every five persons of working age is suffering from a diagnosable mental disorder (1). Estimates also suggest that the economic cost of mental illness, at least in the USA and Europe, is very large, representing anywhere between 2.5–4% of gross national product (2, 3), with much of this cost attributed to reduced workplace productivity.

One source of this loss in productivity is the impact of mental health on work attendance, with previous studies consistently reporting evidence of large significant associations between measures of mental health and sickness absence (4–7). The review of Duijts and colleagues (4), for example, reported that the presence of psychological problems has consistently been found to be associated with higher likelihoods of sickness absence, with adjusted odds ratios in the range of 1.23–1.31 in the case of short-term absences (3 days duration or less) and between 1.37–2.85 in the case of absences of longer duration. Much of the evidence cited, however, comes from samples that are not representative of the broader population, often drawn from individual employers or patients of health service providers, or restricted to coverage of specific occupation or industry groups. Furthermore, studies utilizing nationally representative population samples have mostly involved cross-sectional designs (8–15). Studies that have employed both nationally representative samples and prospective designs are less common and have involved either (i) data collected at just two points in time, with sickness days during the intervening period regressed against the presence of mental health disorders at baseline (16), or (ii) the linkage of baseline survey responses on mental health measures to follow-up sickness absence data derived from administrative attendance records (17–19).

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The present study revisited the relationship between mental health and sickness absence. Unlike most previous studies, it used survey data that were both collected from a sample representative of a national population (Australia) and came from a panel study where sample members are reinterviewed annually, thus providing repeated observations over time for the same individuals on both sickness absence days and a measure of mental health. Central to the analysis was the estimation of panel data regression models that included a diverse set of covariates and, most importantly, controlled for unobserved heterogeneity (ie, unobserved differences across individuals). In essence, this analysis explicitly confronted the possibility that conventional regression estimates of the effect of mental health on sickness absence may be biased because of unobserved characteristics that are correlated with both mental health and sickness absence. The panel data methods employed in this paper correct for this bias.

The data used in this analysis came from Australia, a country where industrial regulations require employers to guarantee full pay for all workers (except those hired on a casual basis) when sick for a minimum of ten days each year. Further, these entitlements accumulate with each year of service. The exclusion of casual employees is important, with different sources suggesting that, over the last decade or so, casual employees have accounted for about one in every five Australian workers (20).

Other forms of protection against loss of income due to illness or injury (if work-related) is provided by workers compensation insurance, which is compulsory and funded by employer contributions. Benefit levels vary across states, with payment levels related to earnings and dependent on the length of the period out of work, and are capped.

Finally, the federal government provides a means-tested, flat-rate income support payment for workers who are temporarily unable to work. The level of payment is very low (set to the level of the unemployment benefit, and hence well below the national minimum wage) and only available for ≤13 weeks. This, together with the stringent income and assets test and the complexity of the application process, means that uptake among workers is very low – over the period covered by this study the number of sickness allowance recipients at any point in time averaged just over 7000 persons (21) or <0.1% of the employed workforce.

Methods

Data and sample

The data source for this analysis was the Household, Income and Labour Dynamics in Australia (HILDA) Survey, which has been collecting data on an annual basis from members of a nationally representative sample of Australian households first surveyed in 2001 (22).

The initial sample of households was selected using a three-stage area-based design, and resulted in a responding sample of 13 969 persons aged ≥15 years from 7682 households. Further interviews were then sought every subsequent year with these persons, along with anyone else (aged ≥15 years) who was living with them. Sample attrition saw the responding sample size decline to 12 408 individuals by wave 4, but sample growth as a result of changes in household composition together with high annual reinterview rates (which have stabilized at around 96%) has seen respondent numbers rise since.

The sample used in this analysis was restricted to persons who were aged 15–64 years and employed during the week preceding interview. Since the HILDA Survey only began collecting information about paid sick days (and other forms of leave) in wave 5, the sample was also restricted to cover observations from waves 5 to 12 (ie, 2005–2012), but with information from wave 4 used in the construction of explanatory variables, and information from wave 13 used in the construction of a control for attrition bias. This provided an initial sample comprising 62 790 observations from 14 365 persons. The mental health measure, however, is administered as part of a separate self-completion instrument, which is associated with additional non-response; 10.3% of observations could not be used because of the absence of a completed self-completion instrument. The final sample for analysis comprised 56 348 observations from 13 622 different individuals.

A summary of selected sample characteristics is provided in table 1.

Measures

The principal outcome variable was a self-reported count of the number of days absent from work while on paid sick leave during the previous 12 months. When averaged across all available years, the mean number of annual paid sick leave days is just 2.9. While low, this figure was not unexpected given the presence of both the self-employed, who represent almost 15% of the sample, and casual employees, who as noted earlier, do not have paid sick leave entitlements. We also considered an alternative outcome variable that supplemented this variable with the number of days away from work while in receipt of workers' compensation insurance payments.

Mental health was measured with the five-item Mental Health Inventory (MHI-5), a sub-scale of the SF-36 (23) that has been shown to be an effective screening instrument for persons with mental health problems in large populations (24, 25). The five items assess fre-

Table 1. Selected (pooled, unweighted) sample characteristics.

Characteristic	Men (51.3%)		Women (48.7%)		Persons	
	Mean	%	Mean	%	Mean	%
Mental health (MHI-5 score)	76.62		74.51		75.59	
≤52		9.1		11.4		10.2
53–60		7.7		8.5		8.1
61–75		18.2		20.3		19.2
76–80		20.4		21.4		20.9
81–88		26.5		24.3		25.5
89–100		18.0		14.1		16.1
Employment status						
Employee (sick leave entitlement)		64.0		62.2		63.1
Employee (no sick leave entitlement)		16.8		26.7		21.7
Self-employed		18.9		10.6		14.8
Other		0.2		0.5		0.4
Proportion of past year in employment	0.959		0.934		0.947	
Usual workdays/week	4.92		4.26		4.59	
Usual weekly hours of work						
Full-time (≥35)		83.8		50.6		67.5
Part-time (<35)		16.2		49.4		32.5
Age group (years)						
15–24		17.2		19.3		18.2
25–34		20.3		19.2		19.8
35–44		24.6		24.5		24.6
45–54		23.9		24.6		24.3
55–64		13.9		12.3		13.2
Country of birth						
Australia		80.9		81.9		81.4
Overseas (English-speaking country)		10.0		8.4		9.2
Overseas (other country)		9.1		9.7		9.4
Indigenous Australian						
Yes		1.5		1.7		1.6
No		98.5		98.3		98.4
Marital status						
Married / partnered		69.5		64.9		67.3
Single		30.5		35.1		32.7
Highest level of education						
Post-graduate qualification		10.3		12.3		11.3
Undergraduate degree / diploma		23.3		28.5		25.9
Certificate Level III / IV		28.0		15.7		21.9
Year 12 (secondary school)		15.9		17.7		16.8
Year 11 or lower (secondary school)		22.5		25.8		24.1
Long-term illness condition/disability						
No		84.1		84.0		84.0
Yes, but not work limiting		8.0		6.8		7.4
Yes, and work limiting		7.9		9.2		8.6

quency of symptoms of anxiety and mood disturbance over the four weeks preceding the interview. Raw scores on each item are summed and scale values transformed to range from 0–100, with relatively low scores indicative of a poor mental health state. Since the distribution of scores was highly skewed, and following Bültmann et al (17), we opted against including a linear specification of MHI-5 in our regression analyses. Instead, we divided scores (roughly) into quintiles and included dummy variables identifying the category to which an individual belonged. We further divided the bottom quintile in two, with the dividing point being a score of 52, a level often chosen by researchers using the MHI-5 when seeking to identify persons at high risk of severe depression (17, 26–28).

To control for variation across individuals in exposure to work, we included the proportion of the previous 12 months spent in employment, the number of days usually worked in a week, and whether the respondent usually worked full-time (≥35 hours/week) or part-time hours. We similarly controlled for exposure to paid sick leave by including dummy variables identifying a respondent's employment status at the time of interview.

The data also enabled the inclusion of controls for a large number of other individual and job-related characteristics. Specifically, we included controls for: age (ten-year age bands); marital and relationship status; the number of persons living in the household; educational attainment; region of birth; whether an indigenous Australian; English-language speaking ability; physical health (as measured by both the physical functioning scale of the SF-36 and the presence of a work-limiting disability); location (state plus a measure of remoteness); smoking status; the log of real annual equalized household income; length of tenure with the current employer (and its square); shift work; union membership; sector (private or public); occupation (8 groups); industry (19 groups); and a measure of the relative socioeconomic advantage / disadvantage of the region in which the respondent lived.

We also included a variable indicating whether the sample member was a non-respondent at the next survey wave. This variable acted as both a check and control for attrition bias. Persons that subsequently drop out of the sample, for example, tend to have worse mental health than those who stay (mean MHI-5=74.5 versus 76.6; t=3.74), and thus it might be expected that these dropouts would also be more likely to be absent from work.

A more detailed description of all variables included in the analysis, along with definitions and summary statistics, is provided in a supplementary data appendix (www.sjweh.fi/data_repository.php).

Statistical analysis

Our outcome variable is both a count and characterized by over dispersion; ie, the variance in the annual number of paid sick leave days (41.9 days) is large relative to its mean (2.9 days). Consequently, we estimated negative binomial regression models. Further, we made use of the longitudinal nature of the data and controlled for unobserved heterogeneity through the estimation of models with random effects. These models differ from conventional regression models by including an individual-specific term into the error structure. Random effects were preferred to the alternative – conditional fixed effects – given that conditional fixed effects when applied to negative binomial models does not truly control

for fixed covariates (29). Conventional random effects estimation, however, has the problem that it maintains the unrealistic assumption that the errors be uncorrelated with the observed covariates. This was dealt with here, in the spirit of Mundlak (30), by including the means of all time-varying covariates as additional regressors. The inclusion of these person-specific means effectively controls for between-person effects, and as a result the coefficients on all time-varying variables can be interpreted as within-person effects. This is why this estimator is sometimes referred as to as the "within-between" estimator (31). The exponeniated coefficients are reported, which have the interpretation of incidence rate ratios.

As noted previously, the outcome variable measures absence days over a one-year period prior to interview (ie, between *t-1* and *t*). Most explanatory variables (including the MHI-5) were therefore measured at the time of the prior interview (ie, at *t-1*). The exceptions here were the controls for work exposure and attrition.

Our broad approach is similar to that of Milner et al (32), who used these same data to examine the impact of psychosocial job quality on sickness absence. There are, however, a number of important differences. First, our models included controls for the amount of exposure to work. Second, our sample included all employed persons. Third, given widespread evidence that women are absent from work more often than men, and – more importantly – that attendance behavior of men and women responds differently to different covariates (33–35), separate models for men and women were estimated. Fourth, our analysis focused on associations with mental health, whereas the analysis of Milner et al did not include any measures of health.

Results

Table 2 reports the estimated incidence rate ratios (IRR) for the main variables of interest from negative binomial regression models before including random effects. These ratios rise more or less in a linear fashion as MHI-5 scores decline. As expected, the lower the mental health score, the greater the number of sickness days taken. The IRR indicate that those employed persons who were most likely to be experiencing severe depressive symptoms (those with MHI-5 scores ≤52) have sickness absence rates that are around 1.3–1.4 times greater than the quintile of workers with the best mental health scores. The size of these ratios does not appear to be much affected by choice of outcome measure – when time on workers' compensation is included, the ratio falls slightly for men while rising slightly for women (but these differences are not statistically significant).

These IRR are consistent with results from previous research reviewed by Dujits et al (4). They are, nevertheless, much lower than that found in previous studies using both national samples and prospective designs. Thorsen et al (19), for example, reported in their Danish sample that a MHI-5 score of ≤60 was associated with an adjusted hazard ratio of 2.56 when compared with persons with MHI-5 scores >90. This is not surprising given Thorsen et al only considered absence spells lasting a minimum of four weeks whereas our analysis considered all paid absence spells, which will be dominated by short-term spells (ie, 1–2 days).

While results on most control variables have not been reported, it is worth noting that estimates mostly accord with expectations and prior research. Thus absence rates declined with worker age, were very sensitive to a worker's physical health, were higher among smokers than non-smokers, rose with household income, rose with length of job tenure, were higher among union members than non-union members, were much higher among persons working in large firms (>500 employees) than smaller firms, were much lower for persons employed in private for-profit businesses than for workers in other types of enterprises, and were relatively sensitive to a worker's occupation and industry of employment. There was also no evidence that subsequent attrition from the sample was associated with either a relatively lower or higher absenteeism rate, suggesting that the estimates are not much affected by any attrition bias.

Once we controlled for unobserved heterogeneity, through the inclusion of correlated random effects, and focus on within-person changes, the magnitudes of the estimated IRR declined markedly (see table 3). The coefficients of most interest – those on the dummy for MHI-5 ≤52 – were obviously very different in the two models given the estimated coefficients from the simple (biased) model (table 2) did not lie within the 95% confidence intervals (95% CI) reported in table 3. Indeed, among women, the correlated random effects results suggested there was no significant association between changes in mental health and changes in sickness days when scores lie in the range of 53–100 (as it does for 89% of the female sample). It was only when MHI-5 scores fell to quite low levels (≤52) that there was a significant rise in sickness absence rates. Further, even at this very low level of mental health, the IRR – at around 1.10 – were much smaller than estimated in the absence of random effects.

Among men, the pattern of IRR showed that absence rates still rise gradually as MHI-5 scores fall, but as with women, the magnitudes of these IRR were relatively small. A within-person decline in MHI-5 from the highest category (the healthiest quintile) to the lowest category (the least healthy decile) was now associated with a rise in the rate of paid sick leave of 13% (and an increase of 14% in the rate of days on paid sick leave or workers compensation).

Table 2. Incidence rate ratios (IRR) from negative binomial regressions of number of sickness absence days in past 12 months. [95% CI=95% confidence interval]

	Outcome = Paid sick leave [a]						Outcome = Paid sick leave + workers' compensation [b]					
	Men (N=6739) Observations= 27 775			Women (N=6688) Observations= 26 721			Men (N=6737) Observations= 27 764			Women (N=6685) Observations= 26 706		
	IRR	95% CI	Wald test (χ^2)	IRR	95% CI	Wald test (χ^2)	IRR	95% CI	Wald test (χ^2)	IRR	95% CI	Wald test (χ^2)
MHI-5 scores [c]			7583.9			11 329.2			6142.75			9824.07
≤52	1.38	1.27–1.50		1.31	1.22–1.41		1.33	1.21–1.46		1.35	1.24–1.46	
53–60	1.22	1.12–1.33		1.18	1.09–1.27		1.27	1.16–1.40		1.17	1.07–1.27	
61–75	1.22	1.14–1.31		1.18	1.11–1.25		1.21	1.12–1.30		1.10	1.02–1.17	
76–80	1.14	1.06–1.21		1.08	1.02–1.15		1.27	1.18–1.37		1.06	0.99–1.13	
81–88	1.09	1.03–1.16		0.99	0.94–1.06		1.06	0.99–1.13		0.93	0.87–0.99	
Selected controls												
Usual workdays/week	1.09	1.06–1.13		1.30	1.27–1.33		1.15	1.12–1.19		1.34	1.31–1.37	
Proportion of past year in employment	3.13	2.63–3.73		2.50	2.22–2.83		1.99	1.66–2.37		2.44	2.15–2.77	
Full-time employed	1.46	1.34–1.59		1.05	1.00–1.10		1.44	1.33–1.57		1.05	0.99–1.10	

[a] Model I controls for: employment status; access to sick leave entitlements; age; marital / relationship status; household size; educational attainment; region of birth; whether of Indigenous origin; English language speaking ability; physical health; and location; and a dummy variable indicating whether the individual responded to the survey at t+1.

[b] Model II includes all controlled variables in Model I as well as smoking status; real annual equivalized disposable household income; job tenure; shift work; union membership; sector; occupation; industry; and a measure of the neighborhood's relative socio-economic / disadvantage.

[c] Reference MHI-5 >88.

Sensitivity analyses

One potential weakness of the analysis is that, by focusing on paid forms of sickness absence, we have ignored some forms of non-attendance at work. Re-estimation of all models but using as the outcome variable the number of unscheduled days off work that were not paid for, however, did not produce evidence of any statistically significant associations with the MHI-5 measure. (Detailed results may be obtained upon request from the first author.)

Another potential criticism is that despite the inclusion of controls for employment status, the effects of mental health on sick days may be attenuated by the presence of workers without paid leave entitlements within the sample. This would be expected if access to paid leave is correlated with mental health. Mean MHI-5 scores are indeed lower among employees without paid sick leave entitlements than among employees with such entitlements (73.7 versus 76.0; t=13.56). Nevertheless, restricting the sample to employees with paid leave entitlements (and thus excluding both casual employees and the self-employed) made very little difference to the magnitudes of the estimated coefficients on the MHI-5 dummies, especially in the models that incorporated correlated random effects. The coefficient on the dummy representing MHI-5 scores ≤52, for example, increased by <1.3% in the female specifications and <0.7% in the male specifications. (Again, detailed results of these supplementary analyses may be obtained upon request from the first author.)

Discussion

Previous research has consistently reported evidence of large significant associations between indicators of mental health status and sickness absence, but all of this body of work has suffered from at least one serious limitation: the inability to deal with the potential bias that could arise from unobserved heterogeneity. Using an ongoing panel survey dataset, we found that the association between a measure of the mental health of the employed population and the number of annual paid sickness absence days was much smaller once we took account of unobserved heterogeneity and focused on within-person differences. Changes in mental health states were still significantly associated with changes in sickness absence days, but the magnitudes of these associations were arguably quite small.

Despite the strengths of the data – its nationally representative coverage and use of repeated observations over time for the same individuals – this study is also not without weaknesses. First, the measure of absence days is self-reported and thus subject to recall biases. In particular, comparisons with estimates of total sickness absence (including both paid and unpaid leave) generated from the Australian National Health Survey (36), which uses a two-week recall period, suggested that absence days have been under-reported in the HILDA Survey. Second, the mental health measure used – the MHI-5 – is also based on self-reported data, and thus also subject to reporting errors and biases. Nevertheless, the presence of measurement bias does not undermine the key conclusion of this analysis – that the magnitude

Table 3. Incidence rate ratios from correlated random effects negative binomial regressions of number of sickness absence days in past 12 months. [IRR = Incidence rate ratio; 95% CI = 95% confidence interval]

	Outcome = Paid sick leave [a]								Outcome = Paid sick leave + workers' compensation [b]							
	Men (N=6739) Observations=27 775				Women (N=6688) Observations=26 721				Men (N=6737) Observations=27 764				Women (N=6685) Observations=26 706			
	IRR	95% CI	Wald test (χ^2)	Ratio test [c] (χ^2)	IRR	95% CI	Wald test (χ^2)	Ratio test [c] (χ^2)	IRR	95% CI	Wald test (χ^2)	Ratio test [c] (χ^2)	IRR	95% CI	Wald test (χ^2)	Ratio test [c] (χ^2)
MHI-5 scores [d]			7200.0	2900.6			8681.2	2345.1			7095.43	4670.15			8567.13	3858.72
≤52	1.13	1.03–1.23			1.10	1.02–1.19			1.14	1.04–1.24			1.12	1.03–1.22		
53–60	1.14	1.05–1.24			1.07	0.99–1.15			1.16	1.06–1.26			1.07	0.99–1.16		
61–75	1.09	1.01–1.16			1.05	0.98–1.12			1.09	1.01–1.17			1.06	0.99–1.13		
76–80	1.05	0.99–1.12			1.01	0.95–1.08			1.06	0.99–1.13			1.02	0.96–1.09		
81–88	1.06	1.00–1.12			1.02	0.97–1.08			1.06	1.00–1.12			1.03	0.98–1.10		
Selected controls																
Usual workdays/ week	1.06	1.02–1.10			1.19	1.16–1.23			1.06	1.02–1.10			1.19	1.16–1.22		
Proportion of past year in employment	7.24	5.70–9.20			4.32	3.69–5.07			6.53	5.17–8.25			4.49	3.83–5.27		
Full-time employed	1.30	1.16–1.45			1.07	1.02–1.13			1.31	1.17–1.46			1.09	1.03–1.15		

[a] Model I controls for: employment status; access to sick leave entitlements; age; marital / relationship status; household size; educational attainment; region of birth; whether of Indigenous origin; English language speaking ability; physical health; and location; and a dummy variable indicating whether the individual responded to the survey at t+1.

[b] Model II includes all controlled variables in Model I as well smoking status; real annual equivalized disposable household income; job tenure; shift work; union membership; sector; occupation; industry; and a measure of the neighborhood's relative socio-economic / disadvantage.

[c] Likelihood ratio test for individual effects (χ^2).

[d] Reference MHI-5 >88.

of the association between mental health and absenteeism is smaller in the presence of correlated random effects. Third, like all longitudinal surveys, our results are subject to the potential for attrition bias. That said, we specifically tested whether the number of absence days of sample members who subsequently dropped out of the sample was any different from those who stayed and found no evidence that such differences mattered.

It also needs to be borne in mind that the outcome examined in this study was total absence days, which will be dominated by absence spells of very short duration. It is thus possible that the conclusions of this study do not extend to analyses of long-term absences. Evidence often suggests that covariates have different effects on absence rates depending on whether or not the outcome comprises short or long absence spells (37–38). Most critical for the analysis reported on here, measures of health status are usually found to be much more strongly associated with long than short absence spells (38–39). Short-term absences are relatively more likely to be associated with temporary or mild health complaints (40) and are often assumed to be more a function of attitudes and behaviors (41). We would thus expect long-term absences to be much more responsive to measures of both underlying physical and mental health status.

Finally, the results reported here were based on Australian data and may not be generalizable to populations in other countries. Previous research, for example, has tended to focus on populations in Northern European countries where rates of sickness absence are relatively high, and much higher than in Anglo-Saxon countries such as Australia, but also Canada, the UK, and the US (42). Relatedly, institutional arrangements that impact on sickness absence levels can vary greatly across countries (43). Entitlements to employer-provided sick pay in Australia are determined by industrial awards and agreements, with the number of days provided being very modest; the norm is just ten days per year (though these do cumulate with years of service). And there is no government-mandated sickness insurance scheme. Further, casual employees, who represent a relatively large fraction of the Australian workforce, do not have any paid sick leave entitlements. Australian workers may thus face greater pressures to attend work when sick than workers in some other developed nations, suggesting perhaps that problems associated with presenteeism may be relatively more acute in Australia. Alternatively, mental health problems might be more likely to be associated with withdrawal from the workforce and ultimately greater levels of welfare dependency in Australia. While there is little convincing evidence for Australia linking mental health to presenteeism, previous research using the same data source as used here has reported evidence showing that declines in mental health are followed by declines in employment rates, most of which was the result of an increase in the number of workers who quit

their jobs rather than an increase in the number of workers who are fired (44).

In summary, the findings reported here suggest that, as a result of omitted variables bias, previous research may have overstated the magnitude of the association between poor mental health and work-related sickness absence. But clearly the approach used here needs to be replicated in datasets from other countries, and especially countries where rates of sickness absence are much higher and where the institutional arrangements that determine sick pay / benefits are very different. It would also be interesting to analyze whether the results obtained here would be replicated in datasets where long-term sickness absence spells were more prevalent.

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