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Long working hours and psychiatric treatment: A Danish follow-up study

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ABSTRACT (ENGLISH)

Objective This study aimed to estimate prospective associations between long working hours and (i) redeemed prescriptions for psychotropic drugs and (ii) psychiatric hospital treatment due to mood, anxiety or stress-related disease, among full-time employees in Denmark. Methods Full-time employees who participated in the Danish Labor Force Survey sometime in the period 2000-2013 (N=131 321] were followed for up to five years in national registers for redeemed prescriptions for psychotropic drugs and psychiatric hospital treatment due to mood, anxiety or stress-related disease. Rate ratios (RR) were estimated for 41-48 versus 32-40 and >48 versus 32-40 working hours a week. The analyses were controlled for sex, age, night shift work, calendar time of the interview and socioeconomic status (SES). Prevalent cases were excluded in primary analyses. Results The RR for psychotropic drugs were estimated at 0.94 [99% confidence interval (CI) 0.88-1.01] for 41-48 versus 32-40 working hours a week and 1.08 (99% CI 0.99-1.18) for >48 versus 32-40 working hours a week. The corresponding RR for psychiatric hospital treatments were estimated at 0.90 (95% CI 0.75-1.08) and 0.96 (95% CI 0.76-1.21). We did not find any statistically significant interaction between weekly working hours and age, sex, SES or night shift work. Conclusion Long working hours as they occur in in the general working population of Denmark are not an important predictor of mental ill health.

FULL TEXT

Headnote

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Conclusion Long working hours as they occur in in the general working population of Denmark are not an important

predictor of mental ill health.

Key terms anxiety; mood disorder; occupational health; prescription drug; psychiatric hospital treatment; psychotropic medicine; stress-related disorder.

The governments of Japan, Taiwan and South Korea recognize extra-long working hours as a compensable risk factor for mental disorders (1). In a meta-analysis by Virtanen et al (2), the relative risk of developing depressive symptoms among workers with long hours (most often defined as >55 weekly hours) versus standard hours (most often 35-40 weekly hours) in Asian countries was estimated at 1.50 [95% confidence interval (CI) 1.13-2.01]. In a recent questionnaire-based study of employees at training hospitals in Japan, the relative risks of having developed depressive symptoms at a three-months follow-up were estimated at 2.83 for 80-99.9 versus <60 working hours a week and at 6.96 for >100 versus <60 working hours a week (3).

In western countries, long working hours are most often not as excessive as in some Asian countries and do not appear to be a similarly serious public mental health problem. In the aforementioned meta-analysis by Virtanen et al (2), the risk ratios (RR) for development of depressive symptoms between workers with long versus standard working hours were estimated to be 1.11 (95% CI 1.00-1.22) in Europe, 0.97 (95% CI 0.70-1.34) in North America, and 0.95 (95% CI 0.70-1.29) in Australia.

In a previous study, we used RR of redeemed prescription for psychotropic drugs to estimate prospective associations between long working hours and mental illhealth among employees in the general working population in Denmark during 99 018 person-years at risk (4). RR for the contrasts 41-48 versus 32-40 and >48 versus 32-40 working hours a week were estimated with and without stratification by age, gender, socioeconomic status (SES) and shift work, respectively. None of the results was statistically significant after adjustment for multiple comparisons. The study suggested, however, that overtime work that exceeds the limit of the EU working time directive (>48 working hours a week) (5) might be associated with a slightly increased risk among employees in the general population (RR 1.15, 95% CI 1.02-1.30). It suggested, moreover, that >48 working hours a week may be an important risk factor among shift workers (RR 1.51, 95% CI 1.15-1.98).

The aim of the present study was firstly to replicate the estimations of our previous study in a larger data material and secondly to supplement the examination with estimated RR for psychiatric hospital treatment due to mood, anxiety or stress-related disease.

Methods

In the present paper, we will only give a brief description of the material and methods of the study. A detailed description can be found in our study protocol (6), which was peer-reviewed and published before we commenced with the analysis. The protocol defines two major studies, one of them focuses on effects of long working hours (reported here) while the other focuses on effects of night shift work (results to be reported elsewhere).

The data material

Individual participant data on usual weekly working hours and night shift work were extracted from the Danish Labor Force Survey (DLFS). The DLFS data were thereafter linked to person-based data on redeemed prescriptions, psychiatric hospital treatments, industry, socioeconomic status, migrations and deaths, from national registers, which cover the entire population of Denmark. The DLFS is based on random samples of 15-74 years old inhabitants of Denmark that have been drawn each quarter of each calendar year since 1994. During the time-period spanned by the present study, the data were collected by means of telephone interviews. The selected participants were invited to be interviewed four times during a period of approximately 1.5 years (7). The response rates have decreased with time, from 70% in 2002 to 53% in 2013. The primary analyses of the present study are based on the participants' first interview in the calendar period 2000-2013.

Inclusion criteria

The study included people who (i) responded to DLFS sometime during the calendar years 2000-2013; (ii) were 20-59 years old at the start of the follow-up period; (iii) were employed with 32-100 usual working hours a week at the time of the interview; and (iv) did not receive psychiatric hospital treatment for any type of mental disorder (ICD-10: F00-F99), as principal diagnosis and did not redeem a prescription for any type of psychotropic drug (ATC: N05-

N06) during a one-year period preceding the start of the follow-up period.

Clinical endpoints

The following endpoints were regarded: (i) redeemed prescriptions for any type of psychotropic medicine, ie, drugs in the ATC-code category N05 (psycholeptica) or N06 (psychoanaleptica); and (ii) psychiatric hospital treatment with a mood, anxiety or stress-related disorder (ICD-10: F30 - F41 or F43) as principal diagnosis.

Weekly working hours

The participants of DLFS were asked first how many hours a week they usually work in their primary job and then (if relevant) how many hours they usually work in each of their secondary jobs. In the present study, we added the hours worked in the primary and secondary jobs in order to form the exposure variable "weekly working hours". Further details are given in our study protocol (6).

Follow-up

The follow-up started at the end of the calendar year of the participant baseline interview and ended whenever any of the following events occurred: (i) five years had elapsed since start of follow-up; (ii) the subject reached the clinical endpoint of the analysis; (iii) the subject emigrated; (iv) the subject died; (iv) the study period ended (31 December 2014 for psychotropic medicine; 31 December 2017 for psychiatric hospital treatment).

Statistical analyses

Poisson regression was used to estimate rates of redeemed prescriptions for psychotropic medicine and psychiatric hospital treatment due to mood, anxiety or stress-related disorders, separately, as a function of weekly working hours [32-40 (reference); 41-48; >48 hours/week]. The analyses were controlled for night shift work (yes versus no), sex, age (10-year classes), calendar time of the interview (2000-2004; 2005-2009; 2010-2013) and SES (legislators, senior officials and managers; professionals; technicians and associate professionals; workers in occupations that require skills at a basic level; workers in elementary occupations; and gainfully occupied people with an unknown occupation). The logarithm of person years at risk was used as offset. Likelihood ratios were used to test for statistical significance.

For redeemed prescriptions of psychotropic medicine, we tested the following effects, each at the significance level 0.01: (i) main effect of weekly working hours; (ii) effect of interaction between age and weekly working hours; (iii) effect of interaction between sex and weekly working hours; (iv) effect of interaction between SES and weekly working hours; and (v) effect of interaction between night shift work and weekly working hours.

For psychiatric hospital treatment due to mood, anxiety or stress-related disorders, we tested for a main effect of weekly working hours at the significance level 0.05. We did not have the statistical power necessary to test for interaction effects.

RR for the main effects of weekly working hours were estimated for redeemed prescriptions of psychotropic drugs as well as for psychiatric hospital treatments. The RR for redeemed prescriptions were, moreover, firstly stratified by sex, age, night shift work and SES and thereafter pooled with the corresponding results from our previous study on associations between weekly working hours and redeemed prescriptions for psychotropic drugs in the general population of Denmark 1995-2010 (4).

The pooling procedure is described in our study protocol (6) as follows: "The pooled results will be obtained through inverse-variance weighting. Since the present study has the same target population as our previous study and the study periods are overlapping, it is likely that some of the participants in our previous study also have participated in the DLFS. Based on the number of participants in our previous study in relation to the number of people in the target population, we expect that approximately one percent of the participants of the present study also participated in our previous study. This overlap will be taken into account in the pooling of the results by use of the following strategy: Before the results are pooled, the standard error of the present study will be multiplied by the square root of (1/(1-x)), where x=0.01 is the proportion of the participants in the present study that are likely to have participated in our previous study."

Sensitivity analyses were conducted to: (i) compare results obtained with and without exclusion of former and current cases of psychiatric treatment, (ii) compare results obtained with and without control for industrial sector, and

(iii) find out if the estimated strength of the association between weekly working hours and redeemed prescriptions for psychotropic drugs would increase when exposure is more stable over time. The methods and results of the sensitivity analyses are given in the supplementary material (www.sjweh.fi/ show_abstract.php?abstract_id=3936). In the sensitivity analyses, we estimate a series of RR and present these with 99% CI. Here, it should be noted that we do not regard the sensitivity analyses and their CI as statistical significance tests. The results of the sensitivity analyses will therefore never be regarded as statistically significant. They may, however, strengthen, weaken or invalidate statistical conclusions of the primary analyses.

Results

A total of 341 482 persons participated in the DLFS sometime during the time period 2000-2013. Of these, we excluded: 99 736 for not being 20-59 years old, 63 914 for not being employed, 33 571 for working <32 hours per week, 292 for reporting to work >100 hours per week, 436 for emigration or death during the calendar year preceding start of follow-up, 10 for not being found in national registers, 668 for missing data on night-time work, and 11 534 for redeemed prescriptions for psychotropic drugs or use of psychiatric hospital treatment during the calendar year preceding start of follow-up. The remaining 131 321 participants were included in the primary analysis. Among the included, we observed a total of 15 826 cases of redeemed prescriptions for psychotropic drugs in 521 976 person years at risk and 1480 cases of psychiatric hospital treatment due to mood, anxiety or stress-related disease in 636 673 person years at risk. Of the hospital treatment cases, 22% were inpatients, 53% were outpatients and 25% were emergency ward patients. The diagnoses among the cases were distributed as follows: F30 manic episode, 0.5%; F31 bipolar affective disorder, 2.4%; F32 depressive episode, 25.8%; F33 recurrent depressive disorder, 10.2%; F34, F38, F39 persistent, other or unspecified affective mood disorders, 0.5%; F40 phobic anxiety disorders, 3.0%; F41other anxiety disorders, 8.6%; F43 reaction to severe stress, and adjustment disorders, 48.9%.

We found a statistically significant relationship between weekly working hours and redeemed prescriptions for psychotropic drugs (P=0.0031), with RR estimated at 0.94 (99% CI 0.88-1.01) for 41-48 versus 32-40 hours a week and 1.08 (99% CI 0.99-1.18) for >48 versus 32-40 hours a week. Here it should be noted that the P-value rejects the null hypothesis of no association between weekly working hours and redeemed prescriptions for psychotropic drugs. In other words, it rejects the null hypothesis, which states that there are no rate differences among the three working hour categories. The statistical significance is due to the difference between the categories >48 versus 41-48 working hours a week (RR 1.15, 99% CI 1.03-1.28).

The association between weekly working hours and psychiatric hospital treatment due to mood, anxiety or stress-related disease was statistically non-significant (P=0.52), with RR estimated, at 0.90 (95% CI 0.75-1.08) for 41-48 versus 32-40 hours a week and 0.96 (95% CI 0.76-1.21) for >48 versus 32-40 hours a week.

We tested interaction effects on redeemed prescriptions for psychotropic drugs but not on psychiatric hospital treatment. We did not find any statistically significant interaction between weekly working hours and age (P=0.86), sex (P=0.56), SES (P=0.24), or night shift work (P=0.26).

The result of the present primary analysis of redeemed prescriptions for psychotropic drugs as well as results obtained after pooling with our previous study (4) are shown in table 1. The result of the analysis of psychiatric hospital treatment are shown in table 2.

The results of a series of pre-specified sensitivity analyses are given in the supplementary material. None of the sensitivity analyses altered the statistical conclusions of the primary analysis. However, the analyses with stratification into industries may suggest increased risks of redeemed prescription for psychotropic drugs for >48 hours a week in the industrial groups of agriculture, forestry, hunting and fishing (estimated RR 1.49, 99% CI 0.95-2.36), construction (estimated RR 1.47, 99% CI 1.00-2.15) and human health and social work activities (estimated RR 1.27, 99% CI 1.02-1.58).

As part of the sensitivity analyses, we estimated the RR of relapse among employees with a past record of psychiatric treatment (Population 3, table S2), because we hypothesized that this group might be more vulnerable to long working hours than groups without past records of psychiatric treatment. To our knowledge, it is the first time ever that relapse rates have been estimated as a function of weekly working hours. The results did, however, not

support the hypothesis of increased risk of working long hours within this group. With 32-40 working hours per week as reference, the estimated relapse RR were 0.94 (99% CI 0.79-1.13) for 41-48 working hours and 1.03 (99% CI 0.82-1.29) for >48 working hours.

Discussion

In the present study, we found a statistically significant association between weekly working hours and redeemed prescriptions for psychotropic drugs with a U-shaped rate ratio pattern. The U-shaped pattern was present in all of the examined gender and age group strata, but not in all SES-groups or among night shift workers. It may, however be difficult to detect U-shaped relationships in small groups due to random errors. The U-shaped pattern was furthermore present in the analysis of psychiatric hospital treatment for mood, anxiety or stress-related disease.

Statistically significant U-shaped relationships between weekly working hours and adverse health outcomes have previously been observed or estimated for cardiovascular disease in Korea (8), cardiovascular mortality in Italy (9) and all-cause mortality in Denmark (10). A U-shaped relationship between weekly working hours and mental ill health makes sense. On one hand, prolonged workweeks may increase family incomes, which are known to be inversely associated with the risk of developing mental health problems (11-13). On the other hand, it is known that excessively long working hours may lead to short sleep and fatigue due to insufficient recovery between work shifts (14-19), which have been associated with an increased risk of developing mental health problems (20-26). From this viewpoint, we may expect the risk of mental ill health to decrease with weekly working hours up to a point where they become too long to allow sufficient sleep and recovery. Another explanation may be that employees in the reference group (32-40 hours per week) who work less than the standard 37 hours per week in Denmark, do so for a reason, which could in some cases be a lower threshold of stress, depression or anxiety.

In the present study, we tested the null hypothesis, which states that the expected RR for incident use of psychotropic drugs in the three exposure categories are equal versus the alternative hypothesis that they are not. The null hypothesis was rejected (P=0.003). The relationship between the estimated RR with 32-40 hours a week as reference was U-shaped, and the RR between the exposure categories >48 versus 41-48 hours a week was estimated at 1.15 (99% CI 1.03-1.28). Since the alternative hypothesis did not specify any particular shape of the relationship between weekly working hours and the examined rates (eg, monotonically increasing, monotonically decreasing, U-shaped), we may regard the test as an exploratory hypothesis test. To back up the U-shaped association, it should ideally be statistically tested in an independent data set with an a priori research hypothesis that explicitly states that the expected pattern is U-shaped.

It has previously been suggested that the association between long working hours and mental ill health depends on age, gender and SES (27). It has, moreover, been hypothesized that the effect of long working hours on the risk of mental ill health is stronger among employees with shift work than among employees without shift work (4). We could not confirm any of these interaction hypotheses; neither in the present study alone nor when data were pooled with our previous study.

In the present study, we used two different proxy measures for mental ill health. The RR for the contrast >48 versus 32-40 hours a week were estimated at 1.08 for redeemed prescription of psychotropic drugs and at 0.96 for psychiatric hospital treatment due to mood, anxiety or stress-related disease. These RR align quite well with the results obtained among western samples in the meta-analysis by Virtanen et al, where the pooled RR for development of depressive symptoms among workers with >54 versus 35-40 hours a week was estimated to be 1.11 in Europe, 0.97 in North America and 0.95 in Australia. The estimated RR of the present study are, however, substantially lower than the pooled risk ratio for development of depressive symptoms among long versus standard working hours in Asian countries, which Virtanen et al estimated to be 1.50. The difference between the Asian and Danish RR may due to the low prevalence of workers with very long hours in Denmark. Only 6.5% of the study population worked >48 hours a week. Approximately half of these worked <55 hours a week and only 20% of them worked >60 hours a week (cf. supplementary table S6).

We wanted the clinical endpoint of the psychotropic drug analysis to be the same as the one used in our previous study (12) (the study that we aimed to replicate). Thus, we included the entire group N05, which contains the

subcategories N05A (antipsychotics), N05B (anxiolytics) and N05C (hypnotics and sedatives). We also included the entire group N06, which covers N06A (antidepressants), N06B (psychostimulants), N06C (antidepressants in combination with psycholeptics) and N06D (antidementia drugs).The relevance of psychostimulants and antidementia drugs in the study of associations between long working hours and mental ill health may be questionable. However, in our previous study, only 0.4% of the cases were due to psychostimulants and there were no cases of antidementia drugs, which indicates that it is unlikely that the inclusion of these two types of drugs have altered any statistical conclusions. The reason for including antipsychotics is that they often are used to treat depressive disorders (28).

It is possible that some types of mental disorders are associated with long working hours while others are not. It is, moreover, possible that an association between long working hours and mental disorders will be missed if the case definition includes too many diagnoses that have nothing to do with weekly working hours. To decrease the risk of missing an association due to dilution, we excluded psychiatric diagnoses that we, a priori, deemed unlikely to have long weekly working hours in their causal path. We included manic and bipolar disorders (F30, F31) in the case definition, and this choice may be questioned. The numbers of cases in these categories were, however, too low to affect the results of the study in any serious way.

The sensitivity analyses suggested a slightly elevated risk among employees with a stable exposure to >48 working hours per week (table S1). They also suggested that the association between weekly working hours and development of mental ill health might depend on industrial sector (table S4). In particular, they suggested a weak positive association between weekly working hours and incident use of psychotropic drugs in the following industrial groups: agriculture, forestry, hunting and fishing; construction; and human health and social work activities. It is possible that the high RR in these industries were due to high prevalences of work-related musculoskeletal injuries and disorders (29-32). It has been shown that musculoskeletal pain is an important predictor of depression, and it has been estimated that a third of the observed cases of redeemed prescriptions for antidepressants or psychiatric hospital treatment due to depression in the Danish labor force could be attributed to musculoskeletal pain (33). It is, moreover, reasonable to believe that the risk of incurring a work-related musculoskeletal injury or disorder depends on how many hours one spends at work. Another possible explanation for the high RR is that they were due to chance. In the present study, we estimated and reported a total of 91 RR and CI and it is therefore not surprising that some of the CI do not contain unity.

We followed the participants in national registers, which cover the entire target population, and thereby eliminated bias from missing follow-up data (except for ten persons not identified in registers). We were able to investigate effects of former cases of psychiatric treatment and could thereby rule out bias from pre-existing mental health problems (table S2). Another advantage of the present study is that the number of participants was large enough to allow us to: (i) differentiate between overtime work within and beyond the limit (48 working hours a week) of the EU Working Time Directive (5), (ii) supplement the analysis of psychotropic drug usage with estimated RR for psychiatric hospital treatment due to mood, anxiety or stress-related disease, and (iii) supplement the analysis of incident use of psychotropic drugs with an analysis of relapse rates among employees with a past record of psychiatric treatment.

The analyses were governed by a study protocol (6) that was accepted for publication before we linked the exposure data of the project to its outcome data. The analyses were, however, not completely blinded since the exposure data of the project have previously been analyzed in relation to ischemic heart disease (34), stroke (35), injuries (36) and all-cause mortality (10). There were two protocol violations. Firstly, the participants who were interviewed in 2013 were followed for a maximum period of four instead of the stipulated five years. The reason for this violation is that the project only had access to hospital treatment data up to 31 December 2017, which we had overlooked when we wrote protocol. The first violation is negligible since it only decreased the total number of person years at risk with approximately 1.4%. Secondly, we did not abide by the following protocol stipulation: "We will moreover exclude all participants who were registered in the employment classification module (ECM) as unemployed or otherwise not economically active during the main part of the calendar year preceding the start of the follow-up." The purpose of

this stipulation was to delete participants with unstable employment status, and the reason for the violation is that our project only had access to employment status according to the DLFS but not according to the ECM, which we had overlooked when we wrote the protocol. Our sensitivity analysis where we only included participants with stable employment status and stable weekly working hours (table S1) suggest, however, that the effect of the second violation was small.

The major drawback of the present study is that it is observational. Since it is not a randomized controlled trial, we cannot rule out the possibility that the results have been influenced by detection bias, referral bias, prescription bias and bias due to self-selection into working hour categories. The response rates have declined during the time of inclusion to 53%. It cannot be ruled out that employees with subclinical or emerging mental health problems could have a lower response rate than others, which may cause an underestimation of the true effect.

Concluding remarks

To allow for possible effects of selection bias, misclassifications and uncontrolled confounding, Monson's guide to strength of association recommends that a RR obtained in an observational cohort study should be interpreted as "no association" (too weak to be detected by epidemiologic methods) if it lies within the CI 0.9-1.2 (32). Although the main test for an association between weekly working hours and redeemed prescriptions for psychotropic drugs was statistically significant in the present study, we note that all RR that were estimated in the primary analyses of the study (tables 1 and 2) lie within Monson's "no association" region. We also note that all of the RR obtained after pooling with our previous study lie within this region. Moreover, not only the RR but also the 99% CI of the pooled non-stratified analyses lie within the no association region. We may therefore safely conclude that, in this large prospective study, we found that long working hours are not an important predictor of mental ill health in the general working population of Denmark where long hours are less prevalent and extensive than in some of the countries reporting stronger associations.

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Conflict of interest

The authors declare no conflicts of interest. The writing of the report and the submission of the paper for publication in a peer-reviewed journal was promised in the grant application. The funder has, however, not been involved in the writing of the report nor in the interpretation of the results.

Sidebar

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DETAILS

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Effect of long work hours and shift work on high-sensitivity C-reactive protein levels among Korean workers

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ABSTRACT (ENGLISH)

Objective We aimed to investigate the association between low-grade inflammation as indicated by highsensitivity C-reactive protein (hsCRP) level and organizational factors, such as work hours and shift work. Methods We evaluated 7470 young and middle-aged workers who participated in the Korea National Health and Nutrition Examination Surveys from 2015-2018. Work hours were determined from self-reported questionnaires. Shiftwork was defined as a non-daytime fixed work schedule. An interaction effect between shiftwork and long work hours on the hsCRP level was estimated using relative excess risk due to interaction (RERI) and attributable proportion (AP) with 95% confidence intervals (CI). Results Increased hsCRP levels were prevalent in 25.2% of the study population. There was a significant association between long work hours and increased hsCRP, especially among middle-aged men [odds ratio (OR) 1.50 (95% CI 1.20-1.87) for moderately increased hsCRP and OR 1.62 (95% CI 1.14-2.30) for highly increased hsCRP]. There was a significant interaction effect between long work hours and shiftwork on increased hsCRP among middle-aged workers. The RERI were 0.03 (95% CI 0.02-0.04) and 0.56 (95% CI 0.45-0.68) among middle-aged men and women, respectively. The AP were 0.02 (95% CI 0.01-0.03) and 0.36 (95% CI 0.31-0.40) among middle-aged men and women, respectively. Conclusions There was no significant association between shiftwork and the level of hsCRP. Long work hours were related to low-grade inflammatory processes, but only in middle-aged workers. There was an interaction effect between long work hours and shiftwork for increased hsCRP, especially in middle-aged women.

FULL TEXT

Headnote

Objective We aimed to investigate the association between low-grade inflammation as indicated by highsensitivity C-reactive protein (hsCRP) level and organizational factors, such as work hours and shift work.

Methods We evaluated 7470 young and middle-aged workers who participated in the Korea National Health and Nutrition Examination Surveys from 2015-2018. Work hours were determined from self-reported questionnaires. Shift work was defined as a non-daytime fixed work schedule. An interaction effect between shift work and long work hours on the hsCRP level was estimated using relative excess risk due to interaction (RERI) and attributable proportion (AP) with 95% confidence intervals (CI).

Results Increased hsCRP levels were prevalent in 25.2% of the study population. There was a significant

association between long work hours and increased hsCRP, especially among middle-aged men [odds ratio (OR) 1.50 (95% CI 1.20-1.87) for moderately increased hsCRP and OR 1.62 (95% CI 1.14-2.30) for highly increased hsCRP]. There was a significant interaction effect between long work hours and shift work on increased hsCRP among middle-aged workers. The RERI were 0.03 (95% CI 0.02-0.04) and 0.56 (95% CI 0.45-0.68) among middle-aged men and women, respectively. The AP were 0.02 (95% CI 0.01-0.03) and 0.36 (95% CI 0.31-0.40) among middle-aged men and women, respectively.
Conclusions There was no significant association between shift work and the level of hsCRP. Long work hours were related to low-grade inflammatory processes, but only in middle-aged workers. There was an interaction effect between long work hours and shift work for increased hsCRP, especially in middle-aged women.
Key terms cardiovascular disease; hsCRP; inflammation; KNHANES; Korea; shift worker; working hour.
Although an adequate systemic inflammatory response is essential for successful recovery from injuries and infections, chronic low-grade or subclinical inflammation may lead to health problems (1). Several reports have described the relationship between inflammation and cardiovascular diseases (CVD) (2), arrhythmias (3), stroke (4) and neurological disorders (5). A complex inflammatory response mediates the pathogenesis of atherosclerosis, which is the underlying pathology of CVD (6). Among numerous biomarkers related to systemic inflammation, the high-sensitivity C-reactive protein (hsCRP) level, which is easily accessible in the clinical setting (7), is associated with the risk of CVD and other vascular pathologic changes (8, 9).
Occupational factors, such as working hours and shift work, may be associated with various health problems (10, 11). The disruption of the circadian rhythm, exposure to artificial light, psychological stress, and other social and lifestyle factors associated with work patterns can promote chronic systematic inflammation (12-14). Factors associated with organizational work may also influence systemic inflammation. In a recent review, Virtanen &Kivimäki (15) concluded that long working hours were a risk factor for CVD. The plausible pathways leading from shift work to CVD are related to psychosocial, behavioral, and physiological mechanisms (16). Shift work may cause physiological stresses, such as inflammation, blood coagulation, and altered cardiac autonomic functions, which are related to atherosclerotic changes. Puttonen et al (17) reported that irregular working hours, such as 2- and 3-shift work schedules, were associated with an increased risk of inflammation among airline employees. Results of other studies indicate that shift work is associated with elevated hsCRP levels among female workers (18) and a higher prevalence of hypertension among male workers with elevated serum ferritin levels (19). However, information on the interactive effects of work-related factors, such as long working hours and shift work, on systemic inflammation in the working population is still scarce.
Thus, this study aimed to investigate the association between hsCRP and organizational factors such as working hours and shift work in a representative sample of Korean workers.

Methods

Study design and data collection

We used data from the Korea National Health and Nutrition Examination Surveys (KNHANES), 2015-2018, that comprise a series of nationally representative population-based surveys on the health and nutritional status of Korean citizens conducted by the Korea Centers for Disease Control and Prevention. The KNHANES were based on self-questionnaires that gather information on different aspects such as demographics, socioeconomic status, dietary habits, and medical history. The surveys were conducted by a trained interviewers' assistant. Blood and urine sampling and anthropometric examination are also performed by trained professionals either at the home of respondents or in mobile examination centers (20). All KNHANES data are publicly available at the KNHANES website (http://knhanes.cdc.go.kr).
The participants of the KNHANES are identified annually through systematic sampling of Korean citizens using multistage clusters based on age, sex, and household registries. A total of 31 649 participants were included in the 2015-2018 KNHANES (7380 in 2015; 8150 in 2016; 8127 in 2017; and 7992 in 2018).
The present study evaluated young and middleaged workers. We excluded 22 030 individuals of the non-working population, 1649 individuals aged <20 or >59 years, and 500 individuals with missing data or who refused to share

data. In total, 7470 participants were included in the analysis. The participant inclusion flowchart is shown in figure 1.

Shift work and long work hours

Shift work was defined as a non-daytime fixed work schedule identified with the question "Do you mainly work weekly (06:00-18:00 hours)? Or do you work during other time slots?" An answer other than "daytime fixed working schedule (06:00-18:00 hours)" indicated a shift-work schedule. The shift-work schedule included evening (14:00-24:00 hours) work, night-time (21:00-08:00 hours) work, 24-hour shift, and split or irregular shifts.

Working hours were based on the total number of hours worked per week ascertained from the selfreported questionnaire that included the following question: "on average, excluding meal times, how many hours do you work at your job per week including overtime?"

The definitions of working schedules and long working hours reflect social, cultural, or economic circumstances and are changeable. We defined long working hours on the basis of the Korean labor standards and descriptions in previous studies. The Labor Standard Act in Korea allows weekly working hours of up to 52 hours, although the weekly working hours can be extended with the agreement of all parties (21). Previous studies from Korea indicated that if the work hours per week exceed 52 hours, the risk of illness because of long working hours increased (22, 23). Thus, we dichotomously categorized work hours into two groups on the basis of the responses "No" and "Yes" to long working hours (weekly work hours <52 and >52, respectively).

High-sensitivity C-reactive protein

The hsCRP level was determined from blood samples collected in 3-mL ethylenediaminetetraacetic acidcoated tubes (BD Vacutainer, Franklin Lakes, NJ, USA), stored at 2-8°C in refrigerated containers, and analyzed within 24 hours of sample collection. The hsCRP level was measured using an immunoturbidimetric method (Cobas, Roche, Germany). Human CRP agglutinates with latex particles coated with monoclonal anti-CRP antibodies, and the aggregates obtained were quantified turbidimetrically. The Food and Drug Administration has approved this method for clinical use in Korea and other developed countries (24, 25). In this study, the participants were categorized into three groups based on the hsCRP level in accordance with the criteria of the American Heart Association and Centers for Disease Control and Prevention: normal <1.0; moderately increased 1.0-3.0; and highly increased >3.0 mg/L (26).

Other covariates

We used age, sex, educational status. and household income level as socioeconomic variables. The educational level was stratified into three levels: graduate to middle school, high school, and college or higher. The household income level was categorized by quartiles [lowest (quartile 1) to highest (quartile 4)] based on the yearly household income level. Past smoking indicated smokers who had stopped tobacco smoking for >1 month. Individuals who had smoked <100 cigarettes in their lifetime were placed in the "none" category. Moderate alcohol drinking for women and men was defined as the consumption of <5 and <7 glasses of alcohol <2 times per week, respectively. Severe alcohol drinking was classified as the consumption of more than the moderate alcohol intake level. Metabolic syndrome was diagnosed according to the recommendations of the International Diabetes Federation (27) and was defined as the presence of >3 of the following 5 abnormalities: (i) central obesity (waist circumference >90 cm among men or >80 cm among women); (ii) hypertension (blood pressure >130/85 mmHg or antihypertensive drug treatment); (iii) hyperglycemia (fasting glucose level of serum >100 mg/dL or use of antidiabetic medication); (iv) high triglyceride (TG) levels (TG >150 mg/dL or drug treatment for dyslipidemia); or (v) low high-density lipoprotein cholesterol (HDL-C) levels (<40 mg/dL among men and <50 mg/dL among women).

Statistical analysis

The differences in general characteristics according to shift work or long work hours were calculated for each variable using the chi-square test. Odds ratios (OR) and 95% confidence intervals (CI) were calculated using logistic regression models to evaluate the association between the hsCRP level and shift work or long work hours with sex and age group stratification. We calculated the age-standardized prevalence ratio of abnormal hsCRP (>1.0 mg/L) according to work conditions with reference to the same-aged general population in the KNHANES. The interaction effect between shift work and long work hours on hsCRP level was estimated with the P-value. In subgroup analysis

with sex and age stratification, an interaction effect was demonstrated based on the P-value, and a relative excess risk due to interaction (RERI) and attributable proportion (AP) were represented by the 95% CI. All of the interaction effects were estimated with "no long work hours" and "no shift work" group as a reference. All statistical analyses were performed using SAS (version 9.4; SAS Institute, Cary, NC, USA). Two-tailed P-values <0.05 were considered statistically significant.

Results

The baseline characteristics according to hsCRP levels are presented in table 1. The prevalence of extremely high and moderately high hsCRP levels was higher among men than women. A lower household income level was significantly associated with a higher prevalence of increased hsCRP levels. Current smokers and participants with metabolic syndrome showed significantly higher prevalence of increased hsCRP levels. The long work hours group also had a higher prevalence of increased hsCRP levels.

Table 2 presents the results of the logistic regression analysis for the risk of increased hsCRP levels with respect to shift work or long work hours based on sex and age group stratification. Significantly elevated risks of moderately and highly increased hsCRP levels were identified among men in the long work hours group after adjustments for age, educational level, income level, smoking, alcohol consumption, and metabolic syndrome [OR 1.21 (95% CI 1.02-1.44) and OR 1.33 (95% CI 1.01-1.76), respectively]. After sex and age group stratification, the association between long work hours and increased hsCRP levels was enhanced among middle-aged men [40-59 years: OR 1.50 (95% CI 1.201.87) and OR 1.62 (95% CI 1.14-2.30), respectively]. However, there was no significant association between shift work and increased hsCRP levels.

Table 3 presents the results of the interaction effect of long working hours and shift work on the risk of increased hsCRP levels (>1.0 mg/L). Overall, there was no interaction between long working hours and shift work with the hsCRP level (P-value for interaction = 0.1648). After sexspecific stratification, the strata with only women showed a statistically significant interaction effect on increased hsCRP levels (P-value for interaction = 0.0001).

Figure 2 shows the results of the interaction analysis of long work hours with shift work and the influence on higher hsCRP level with regard to sex and age group stratification. There was an interaction effect between long work hours and shift work on increased hsCRP levels among middle-aged men and women. The RERI were 0.03 (95% CI 0.02-0.04) and 0.56 (95% CI 0.45-0.68) among middle-aged men and women, respectively. The AP were 0.02 (95% CI 0.01-0.03) and 0.36 (95% CI 0.31-0.40) among middle-aged men and women, respectively. However, only middle-aged women showed a statistically significant result (P-value for interaction = 0.0365)

Discussion

This study investigated the association of work conditions, such as long work hours and shift work, with the hsCRP level, which was considered a systematic proinflammatory factor. Male workers with long work hours had a higher likelihood of increased hsCRP levels. Despite the attenuation of the relationship between the risk of increased hsCRP levels and shift-work schedule after controlling for confounders, we found a significant interaction for an increased prevalence of elevated hsCRP levels between long work hours and the shift work schedule, especially among the middle-aged working population.

We found that long work hours were associated with elevated hsCRP levels, consistent with the findings of previous studies that showed a close association between long work hours and elevated hsCRP levels, especially among older workers (28). Long work hours could lead to sleep disturbance that is, in turn, directly associated to a lack of time for physical and neuropsychiatric recovery (21, 29). Previous studies have indicated that sleep disturbance is associated with increased levels of markers of systemic inflammation, including hsCRP (30). Extended work hours could activate the stress-response system that would influence the hypothalamic-pituitaryadrenal axis to increase the levels of glucocorticoids, which is considered a response to increased systematic inflammation (31). Moreover, a higher prevalence of smoking, obesity, and other adverse behaviors has been reported in those who work long hours (32).

A recent study showed that shift work influenced some immunological biomarkers (33). Thus, we hypothesized that hsCRP is also associated with shift work; however, we found no significant association between shift work and

elevated hsCRP levels. This finding is inconsistent with that of an earlier report (34). However, the interaction between long work hours and shift work and elevated hsCRP levels was statistically significant. This result is consistent with previous studies that reported that the association between shift work and elevated inflammatory indices is markedly evident among those with long work hours or those working in prolonged shifts (35, 36). Zeler et al (35) investigated the oxidative and inflammatory status according to physical workloads of employees. They found that the levels of various inflammatory markers, including total peroxides, total antioxidant capacity, adrenocorticotropic hormone, oxidative stress index, and hsCRP, were significantly higher among heavy workers (slaughterhouse employees) than among office workers, especially those who worked longer hours (ie, 12- versus 8-hour shifts).

Buyukhatipoglu et al (36) also reported increased oxidative stress indices after prolonged work hours among healthcare workers who were on 24-hour on-call shifts compared with non-healthcare staff who worked 8-hour shifts. Thus, the hypothesis that both long work hours and shift work are more likely to induce systemic inflammation with an interaction effect, as shown in this study, seems feasible. A plausible explanation for this is the alterations in circadian rhythms that exacerbate chronic systemic inflammation by triggering proinflammatory activation of macrophages and inflammatory signaling pathways (37).

It remains unclear why the interaction effect was only significant among women, although sex differences in the circadian system may be a possible explanation. Duffy et al (38) reported that the intrinsic circadian period was shorter among women than men, and a greater proportion of women had intrinsic circadian periods <24 hours (38). This shorter intrinsic circadian period leads to circadian misalignment associated with alteration in physiologic processes, including elevated levels of markers of systemic inflammation. Wright et al (39) observed significantly increased CRP levels in the circadian misalignment group and decreased CRP levels in the synchronized group after weeks of circadian entrainment.

There may be another possible explanation considering the difference in total working hours between men and women. The total working hours of women may be longer than those of men because women do more domestic labor. The Statistics Korea 2019 nationwide Time Use Survey showed that men (husbands) of dualincome households spent 8 hours and 46 minutes for inevitable daily activities, including 5 hours and 50 minutes for work and 54 minutes for domestic labor, whereas women (wives) of dual-income households spent 9 hours and 24 minutes for inevitable daily activities, including 4 hours and 37 minutes for work and 3 hours and 7 minutes for domestic labor (40). Men and women in single-income households spent very similar time for inevitable daily activities compared with those in dual-income households. Although only average statistics are available and such statistics cannot be directly linked to this study data, women apparently do more domestic labor than men, even when they have a job. Women who work long hours with shift work may have less time for rest than men with the same work characteristics because women do more housework. This may be one of the possible explanations why the interaction effect between long work hours and shift work on the increase of hsCRP levels was found only among women.

Given the nature of chronic diseases, one of the best approaches to prevent CVD is primary prevention. Emphasis should be placed on controlling risk factors and reducing unsafe behaviors and conditions to prevent CVD. From this viewpoint, reducing work hours or maintaining an appropriate level of work hours for young and middle-aged workers may be helpful to reduce the risk of diseases that occur due to chronic inflammation. Special attention should be focused on shift workers to emphasize the need to control other risk factors for the prevention of diseases when shift work is inevitable.

The present work contributes to the existing knowledge base of the association between work conditions and inflammation by providing an interaction analysis with sex- and age group-stratified analysis that was adjusted for confounders. In support of the evidence from previous studies that were conducted with relatively small sample sizes, we report a significant association between long work hours and shift work and increased hsCRP levels in a large representative sample of Korean workers.

This study had some limitations, the major being its cross-sectional design. This made it difficult to determine a

causality association between work conditions and hsCRP. Further longitudinal studies are required to ascertain the causality in this regard. Moreover, we could not describe other work condition-related factors of inflammation besides the work hours and working schedule due to a lack of information about workplace hazardous factors in the KNHANES. Systematic inflammation could have been increased because of exposure to various hazardous factors such as organic chemicals, dust, or stress at the workplace. The findings of this present work cannot be extrapolated to all work conditions in the general population. We could not describe the duration of career or the duration of exposure to shift work schedule or long work hours associated to the level of hsCRP because of the lack of information on these aspects in the KNHANES. Further studies are needed to identify the dose-response relationship between shift work or long work hours and the risk of elevated hsCRP levels.

In conclusion, long work hours were significantly associated with increased hsCRP levels among middleaged men. Shift work itself was not a risk factor for low-grade inflammation. However, there was a significant interaction effect between long work hours and shift work on low-grade inflammation, especially among middle-aged women. Overall, working hours were related to low-grade inflammatory processes, but only among middle-aged workers.

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Conflicts of interest

The authors declare no conflicts of interest.

Sidebar

Lee W, Kang S-K, Choi W-J. Effect of long work hours and shift work on high-sensitivity C-reactive protein levels among Korean workers. Scand J Work Environ Health. 2021;47(3):200-207. doi:10.5271/sjweh.3933

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Two-year neurocognitive responses to first occupational lead exposure

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ABSTRACT (ENGLISH)

Objectives Lead exposure causes neurocognitive dysfunction in children, but its association with neurocognition in adults at current occupational exposure levels is uncertain mainly due to the lack of longitudinal studies. In the Study for Promotion of Health in Recycling Lead (NCT02243904), we assessed the two-year responses of neurocognitive function among workers without previous known occupational exposure newly hired at lead recycling plants. Methods Workers completed the digit-symbol test (DST) and Stroop test (ST) at baseline and annual follow-up visits. Blood lead (BL) was measured by inductively coupled plasma mass spectrometry (detection limit 0.5 jj.g/ dL). Statistical methods included multivariable-adjusted mixed models with participants modelled as random effect. Results DST was administered to 260 participants (11.9% women; 46.9%/45.0% whites/Hispanics; mean age 29.4 years) and ST to 168 participants. Geometric means were 3.97 and 4.13 ^g/dL for baseline BL, and 3.30 and 3.44 for the last-follow-up-to-baseline BL ratio in DST and ST cohorts, respectively. In partially adjusted models, a doubling of the BL ratio was associated with a 0.66% [95% confidence interval (CI) 0.03-1.30%; P=0.040] increase in latency time (DST) and a 0.35% (95% CI -1.63-1.63%; P=0.59) decrease in the inference effect (ST). In fully adjusted models, none of the associations of the changes in the DST and ST test results with the blood lead changes reached statistical significance (P>0.12). Conclusions An over 3-fold increase in blood lead over two years of occupational exposure was not associated with a relevant decline in cognitive performance.

FULL TEXT

Headnote

Objectives Lead exposure causes neurocognitive dysfunction in children, but its association with neurocognition in adults at current occupational exposure levels is uncertain mainly due to the lack of longitudinal studies. In the Study for Promotion of Health in Recycling Lead (NCT02243904), we assessed the two-year responses of neurocognitive function among workers without previous known occupational exposure newly hired at lead recycling plants. Methods Workers completed the digit-symbol test (DST) and Stroop test (ST) at baseline and annual follow-up visits. Blood lead (BL) was measured by inductively coupled plasma mass spectrometry (detection limit 0.5 jj.g/ dL). Statistical methods included multivariable-adjusted mixed models with participants modelled as random effect. Results DST was administered to 260 participants (11.9% women; 46.9%/45.0% whites/Hispanics; mean age 29.4 years) and ST to 168 participants. Geometric means were 3.97 and 4.13 ^g/dL for baseline BL, and 3.30 and 3.44 for the last-follow-up-to-baseline BL ratio in DST and ST cohorts, respectively. In partially adjusted models, a doubling of the BL ratio was associated with a 0.66% [95% confidence interval (CI) 0.03-1.30%; P=0.040] increase in latency time (DST) and a 0.35% (95% CI -1.63-1.63%; P=0.59) decrease in the inference effect (ST). In fully adjusted models, none of the associations of the changes in the DST and ST test results with the blood lead changes reached statistical significance (P>0.12). Conclusions An over 3-fold increase in blood lead over two years of occupational exposure was not associated with a relevant decline in cognitive performance.

Key terms digit-symbol test; neurocognitive function; occupational exposure; Stroop test.

Lead is a ubiquitous environmental toxicant. The Global Burden of Disease study assumed a causal association

between intellectual disability and lead exposure in children (1), mainly justified by a participant-level meta-analysis involving 1333 children enrolled in seven population-based studies and followed up from birth or infancy until 5-10 years of age (2). The IQ point decrements associated with blood lead increments from 2.4-10, 10-20, and 20-30 qg/dL were 3.9, 1.9, and 1.1, respectively (2). The lead-associated intellectual decrement in children with a maximal blood lead level <7.5 pg/dL was greater than that observed in those with a maximal blood lead level of >7.5 pg/dL (P=0.015). These counterintuitive findings might be a product of residual confounding, falling exposure levels over time or a decreasing vulnerability for cognitive impairment with higher age (2). Turning to adults, the literature relating neurocognitive function to lead exposure in studies of the general population (3-9) or workers (10-13) with a cross-sectional (3-5, 7, 8, 10), case-control (11, 13) or longitudinal design (6, 9, 12) is contradictory. Similarly, two systematic reviews (14, 15), including 22 studies of exposed and unexposed workers but using different statistical methods, concluded that there was an inverse (14) or a null (15) association between neurocognition and occupational lead exposure. Unexposed and exposed blood lead levels in workers were unavailable in over ten studies (15). None of the studies compared blood lead levels before and after exposure (15). None of the individual studies was conclusive. Lack of true measures of the pre-occupational exposure and observer and publication bias were other issues obscuring the true relation between neurocognitive function and lead exposure for blood lead levels <70 pg/dL (15). Given the contradictory results of individual studies (3-13) and literature reviews (5, 14, 15), we identified a great need for prospective studies that would account for variability between people by comparing test results before and after lead exposure In the Study for Promotion of Health in Recycling Lead (SPHERL; NCT02243904) (16), we assessed the association between neurocognitive function and blood lead in young workers prior to (17) and up to two years after starting first occupational exposure.

Methods

Participants

SPHERL is a longitudinal study of newly hired lead workers at battery manufacturing and lead recycling plants in the United States (16). SPHERL complies with the Helsinki Declaration for investigations in humans (18). The Ethics Committee of the University Hospitals Leuven (Belgium) approved the study protocol. The health of the labor force was protected in compliance with the US Occupational Safety and Health Administration Standard (www.osha.gov/laws-regs/regulations/ standardnumber/1910/1910.1025), which includes regular health check-ups, proper workplace ventilation, and the obligatory use of personal protective equipment. The two-year neurocognitive responses to first occupational lead exposure were a predefined secondary study endpoint (16).

Of 746 newly hired workers invited to participate, 601 (80.6%) consented. However, in the interval between consent and the planned baseline examination (median, 19 days; 5-95th percentile interval, 9-59 days), 95 laborers left the workplace or withdrew. From 25 January 2015 until 19 September 2017, 506 workers underwent the baseline examination, of whom 289 (57.1%) had one and 236 (46.6%) had two follow-up visits (figure 1). Of 289 participants with at least one follow-up visit, 22 were disqualified for analysis because blood lead had not been measured at baseline (N=3) or follow-up (N=1), because both the digit-symbol test (DST) and Stroop test (ST) had not been administered (N=2), or because workers were on neuropsychiatric medications (N=16), including antidepressants, amphetamines, sedatives, recreational drugs and/or opioids. Of the 267 analyzed participants, 7 were excluded from the DST cohort, because of missing baseline DST; 99 were excluded from the ST cohort, because of missing ST at baseline (N=2) or follow-up (N=1), missing congruent trials at follow-up (N=89), or because they had achromatopsia (N=7). The statistical analysis therefore included 260 participants in the DST cohort and 168 in the ST cohort with both a baseline and at least one follow-up assessment of their cognitive function and simultaneous blood lead measurements.

Clinical measurements

At the study sites, trained nurses measured the workers' anthropometric characteristics and applied current guidelines to measure office blood pressure at the brachial artery. After the workers had rested for five minutes in the sitting position, the nurses obtained five consecutive blood pressure readings to the nearest 2 mm Hg by auscultation of the Korotkoff sounds, using standard mercury sphygmomanometers. For analysis, the five readings

were averaged. Blood pressure was categorized according to the 2017 ACC/AHA guideline (19). If systolic and diastolic blood pressure were in different categories, the highest value was used to classify participants. Heart rate was counted over 15 seconds. Body mass index (BMI) was body weight in kilograms divided by height in meter squared. The study nurses administered a validated (20) questionnaire at baseline and follow-up to collect information about each worker's medical history, exposure to heavy metals, smoking and drinking habits, intake of medications, and educational attainment. Alcohol consumption was categorized as absent, light, moderate or heavy. The thresholds for the daily alcohol consumption self-reported by questionnaire were <6, >6-14 and >14 gram in women and <12, >12-28 and >28 gram in men.

Neurocognitive function tests

The neurocognitive examination was conducted in a quiet air-conditioned room. We administered the computerized version of the DST and ST as published by Xavier Educational Software Ltd, Bangor, Wales, UK, using a laptop with touch screen. A video uploaded at the journal's website illustrates the administration of these tests.

The DST measures processing speed, working memory, visuospatial processing, and attention (21). Participants performed the DST test at baseline and follow-up to assess the impact of lead exposure on general cognitive functions such as processing speed, working memory, visuospatial processing, and attention. A row of nine symbols paired vertically with nine digits was displayed at the top of the computer screen. The same symbols were also presented at the bottom of the screen but in a different order. The task is to touch as fast as possible the symbol at the bottom of the screen that is paired with the displayed digit. During the test, 36 digits appear one after one in the center of the screen. The worker performing the test has to provide the correct response before a new digit is presented. The time needed to complete the test, called latency, and the total number of errors served as measures of performance.

The ST was used to measure the impact of lead exposure on the Stroop effect, which is related to selective attention. Workers saw the printed name of a color and four buttons displayed in yellow, red, blue and green on the laptop screen. In congruent trials, the name of the color was printed in the matching color (eg, "yellow" was printed in yellow). In incongruent trials, the name of the color was printed in a different color (eg, "yellow" was printed in red). The task consists of touching the screen button with the color matching the printed color name as fast and accurately as possible, ignoring the color of the printed color name. The ST consisted of 4 congruent and 12 incongruent trials. Before the test, participants completed four practice trials. The mean reaction time (ms) is the average time that passed between the appearance of the color name and touching the correct button in congruent and incongruent trials, respectively. The main outcome measure in the Stroop test is the inference effect, calculated as the ratio of the mean reaction time for the incongruent to the congruent trials, which is equivalent to the antilog of the difference between the log transformed reaction times. The inference score is defined as the proportion of the correct answers in congruent trials minus the proportion of correct answers in incongruent trials.

Biochemical measurements

Venous blood samples were obtained after 8-12 hours of fasting. Blood lead levels were determined on whole blood by inductively coupled plasma mass spectrometry at an analytical laboratory certified for blood lead analysis in compliance with the provisions of the OSHA Lead Standard, 29CFR 1910.1025 (Occupational Safety and Health Administration [www.osha.gov]). Prior to analysis, the specimens were digested by nitric acid and spiked with an iridium internal standard. The detection limit was 0.5 gg/dL. The accuracy of the lead tests was verified by use of proficiency samples purchased from the College of American Pathologists (CAP) and the Pennsylvania Department of Blood Lead Programs (22). Proficiency testing was performed in six separate trial runs, including in total 30 test samples annually. All survey materials were handled in the same manner as the study samples and processed with the normal workflow, utilizing the same repeat/dilution protocols and calibration and quality control frequency (22). Compliance with the Clinical Laboratory Improvement Amendments (CLIA), CAP and the New York State accreditation and regulatory requirements was verified routinely with test level review of the laboratory services by external auditors. Calibrators with certified accuracy (National Institute of Standards and Technology [www.nist.gov]) were included in each batch of study samples and spanned the range of the analytical measurement range.

Accuracy was evaluated by Westgard Rules (23) and defined within the total allowable error established with review of the CAP, Centers for Disease Control and Prevention, CLIA 88 (24), and OSHA guidelines. Accuracy, defined as the deviation from known lead standards ran along with the study samples, was within 10% (22). The bias determined according to the Bland &Altman approach (25) in 30 spilt blood samples with blood lead concentrations (average in duplicate samples) ranging from 0.70-27.9 gg/dL, was 0.08 gg/dL [95% confidence interval (CI) -0.01-0.18, P=0.078; supplementary material www.sjweh.fi/show_abstract.php?abstract_ id=3940, figure S1) (17). The repeatability coefficient, defined as twice the SD of the signed differences between duplicate measurements (25), was 0.52. Expressed as a percentage of the mean blood lead concentration or as a percentage of near maximal variation in blood lead (four times the SD of the logarithmically transformed distribution), the repeatability coefficients were 6.7% and 1.9%, respectively. Lower values indicate better repeatability.

Serum total and high-density lipoprotein (HDL) cholesterol, serum creatinine, and blood glucose were measured by automated enzymatic methods and serum insulin by ELISA. Over three evaluations, the laboratory obtained a proficiency score of 100% for blood lead and 100% for routine biochemistry. Diabetes mellitus was a self-reported diagnosis, a fasting blood glucose of 126 mg/dL (7.0 mmol/L) or higher, or use of antidiabetic drugs.

Statistical analysis

For database management and statistical analysis, we used the SAS software, version 9.4, maintenance level 5 (SAS Institute Inc, Cary, NC, USA). Departure from normality was evaluated by the Shapiro-Wilk statistic. Skewness and kurtosis were computed as the third and fourth moments about the mean divided by the cube of the standard deviation. We applied a logarithmic transformation (base 10) to normalize the distributions of latency time (DST), mean reaction time and interference effect (ST), and blood lead. We reported the central tendency and spread of continuously distributed variables as mean with standard deviation (SD) or for logarithmically transformed variables as geometric mean with interquartile range (IR) or with the 5-95th percentile interval. To compare means and proportions, we applied the t-statistic or ANOVA for continuous variables, and the Fisher exact test for categorical variables, respectively.

In exploratory analyses, we assessed the results of DST and ST across fourths of the follow-up-to-baseline blood lead concentration ratio. Changes in DST and ST were correlated with the corresponding changes in log blood lead using a random intercept mixed model, accounting for the correlations between repeated observations within the same participant. A compound symmetry correlation structure was assumed and variance parameters were estimated using restricted maximum likelihood. The model included change in log blood lead as a fixed effect. Neurocognitive responses to the changes in blood lead were expressed for a doubling of the followup-to-baseline blood lead concentration ratio. For each outcome, unadjusted, partially and fully adjusted models were constructed. Partially adjusted models included sex, age and the neurocognitive function test at baseline as covariables. Fully adjusted models additionally accounted for ethnicity (white versus other), change in age, baseline BMI, changes in body weight, educational attainment, baseline blood lead and the baseline values and changes during follow-up in smoking status, and the total-to-HDL serum cholesterol, and alcohol consumption (light, moderate and heavy drinkers). Covariables were selected on the basis of their associations with both neurocognitive function and blood lead in previous publications (26-29). In sensitivity analyses, we stratified the study participants according to median age, the median baseline blood lead level and the median cumulative blood lead index (CBLI) (30). We also checked the performance of the mixed models by relating changes in neurocognitive function and blood lead separately for the 1- and 2-year visits by means of linear regression.

Results

Characteristics of participants

Of 260 participants, 229 (88.1%) were men, 122 (46.9%) were white, 117 (45.0%) were Hispanic, and 21 (8.1%) had other self-reported ethnicities. At baseline, age averaged 29.4 years (supplementary figure S2), BMI 28.8 kg/m2, serum creatinine 0.93 mg/dL, total and HDL serum cholesterol 171.8 mg/dL and 46.8 mg/dL, the total-to-HDL cholesterol ratio 3.90, and blood glucose 93.8 mg/dL (supplementary table S1). The cohort included 6 women and 63 men, who were current smokers (N=69; 26.5%); 11 women and 102 men (N=113; 43.4%) reported alcohol

intake, of whom 9 and 65, 2 and 23, 0 and 14 were light, moderate and heavy drinkers, respectively. The baseline characteristics of the 168 workers in the ST cohort were similar (supplementary table S2). The characteristics of 267 workers included in the DST or ST cohort or both and the 239 workers not analyzed were largely similar (supplementary table S3).

Blood lead

Median follow-up was 2.0 [5-95th percentage interval (PI) 1.0-2.2] years. In the DST cohort, the geometric mean blood lead concentration was 3.97 (PI 0.90-14.3) pg/dL at baseline, 13.4 (PI 3.70-30.3) pg/dL and 12.8 (PI 2.80-29.2) pg/dL at the first and second follow-up visits, respectively. The corresponding blood lead levels in the ST cohort were 4.13 (PI 1.20-13.0) pg/dL, 14.4 (PI 4.60-30.3) pg/dL and 16.1 (PI 5.40-31.5) pg/dL. The last-follow-up-to-baseline blood lead concentration ratio averaged 3.30 (PI 0.79-14.9) and 3.44 (PI 1.0113.8) in DST and ST cohorts, respectively (figure 2 and supplementary figure S4). The increase in the blood lead concentration was fully observed at the 1-year follow-up visit (supplementary figure S3).

Digit-symbol test

Workers with a completely correct DST numbered 153 (58.9%) at baseline and 160 (61.5%) at the last followup examinations. Among all participants, the geometric mean test duration was 108.9 [interquartile range (IQR) 95.8-120.8] seconds at baseline and 107.6 (IQR 91.4122.6) seconds at last follow-up. The number of errors and the mean latency time were similar at baseline and last follow-up (table 1). Across fourths of the distribution of blood lead changes, trends in baseline (P>0.076; supplementary table S4), follow-up (P>0.38; supplementary table S5) and the longitudinal changes (P>0.079; table 2) of errors and of the mean latency time were not significant. In mixed models (supplementary figure S5) only accounting for clustering within participants (P=0.0033; table 3) and in partially adjusted models (P=0.040) also adjusted for sex, age and the baseline test result, the change in latency time increased with the follow-up-to-baseline blood lead concentration ratio. However, in a fully adjusted model, this association size weakened to 0.55% (95% CI -0.33-1.42; P=0.22; table 3 and supplementary figure S5). Moreover, in unadjusted, partially adjusted and fully adjusted model, none of the odd ratios for an increasing error rate reached significance (P>0.12; table 3).

We ran stratified analyses using fully adjusted models to evaluate the consistency of the changes of neurocognitive function among workers aged <26.4 and >26.4 years (supplementary table S6), baseline blood lead <4.20 and >4.20 gg/dL (supplementary table S7), and CBLI <32.5 and >32.5 gg/dL x year (supplementary table S8), respectively. In these stratified analyses, an increasing error rate in the high baseline blood lead subgroup was the only measurement, which tended to be associated with the follow-up-to-baseline blood lead concentration ratio: odds ratio, 1.68 (95% CI 0.99-2.86; P=0.056 in the high baseline blood lead stratum vs 1.00 in the low exposure group (95% CI 0.64-1.58; P=0.99) with a nonsignificant interaction (P=0.34; supplementary table S7). The results of the linear regression analyses correlating changes in latency time and blood lead separately at the 1- and 2-year follow-up visits largely confirmed the findings obtained by mixed models (supplementary table S9).

Stroop test

The Stroop test with incongruent trials was completed error free in 145 (86.3%) workers at baseline, with no difference between baseline and follow-up in these proportions (P=0.55; table 1). The mean reaction time for incongruent trials increased from 1606 ms at baseline to 2088 ms at the last follow-up visit in all participants and from 1608 ms to 2077 ms, if only the correct responses were considered (P<0.0001). The changes from baseline to follow-up averaged 30.1% (CI 22.7-37.9; P<0.0001) and 29.8% (CI 22.3-37.8; P<0.0001), respectively (table 1). Supplementary tables S11 and S12 show the mean reaction time and blood lead levels at baseline and follow-up, and overall, in the workers tested and broken down by the attending observer. Supplementary table S11 illustrates the effect of the observer on test performance and supplementary table S12 reflects the unpaired distribution of observers between the baseline and follow-up examinations.

Across fourths of the distribution of the blood lead changes, there was a trend towards smaller increases in mean reaction time with larger increases in blood lead (P<0.015; table 2). For congruent trials, 168 (100%) were completed with fully correct answers at baseline and 164 (97.6%) at follow-up visit with an increasing mean reaction

time from 1485 ms to 1979 ms (P<0.0001, table 1). The geometric means of the interference effects were 1.08 at baseline and 1.06 at the last follow-up visit in all workers. The distributions of inference score were similar at baseline and last follow-up visits (P=0.14; table 1). For congruent trials, in unadjusted and in partially and fully adjusted models, irrespective of whether all trials or only the error-free trials were analyzed, there was no association between the changes in mean reaction time and those in blood lead (P>0.072; table 3). For incongruent trails, in the unadjusted models only accounting for clustering within participants and in the partially adjusted models, the longitudinal change in mean reaction time decreased as the blood lead increasing. However, in the fully adjusted models, the association sizes for a doubling of blood lead were -0.83% (95% CI -3.20-1.59; P=0.49) in all trials and -1.26% (95% CI -3.59-1.13; P=0.29) in error-free trials, respectively (table 3 and supplementary figure S5). Moreover, there was no association between the changes in interference effect and those in blood lead (P>0.29; table 3). In the analyses stratified by median age (27.0 years; supplementary table S5), median baseline blood lead (4.30 gg/dL; supplementary table S6), or median CBLI (33.3 gg/dL x year; supplementary table S7), associations were all nonsignificant (P>0.096; interaction P>0.22). Linear regression analysis of the 1-year data was confirmatory (supplementary table S9).

Consistency between baseline and last follow-up data

Supplementary table S10 lists the associations between blood lead level and the performance of participants in the neurocognitive tests at baseline and the last followup separately. None of the association sizes (slopes) in unadjusted or adjusted analyses reached significance (P>0.14), with no differences between baseline and last follow-up in the association sizes (P slope>0.23).

Discussion

In a real-world experiment, among workers without known previous occupational exposure and taking up new jobs in lead recycling and battery manufacturing plants in the United States, an over threefold increase in the blood lead concentration over the 2-year follow-up was not associated with worsening of cognitive function, as assessed by the DST and ST. These longitudinal findings are in keeping with the cross-sectional analysis of the baseline SPHERL data (supplementary table S10), which did not show any association between cognitive performance as assessed by the same tests and blood lead prior to occupational exposure (17). The longitudinal changes in mean reaction time in incongruent ST trials tended to correlate inversely with the corresponding changes in blood lead, similar to congruent trials (tables 2 and 3). To what extent training effects (31) or the interaction between observers and test takers (supplementary table S11) influenced the ST test performance cannot be ascertained. To exclude an effect of the cumulative lead dose, we ran analyses stratified by the medians of age, baseline blood lead or CBLI in both cohorts, which confirmed the main analysis.

Lead is a cumulative toxicant, which is for 90-95% stored in bone, from where it is recirculated with a half-life of 20-25 years (32, 33). Blood lead, for 95% carried by red blood cells, reflects recent exposure over the past 1-2 months and the amount of lead released and recirculated from bone (32). Bone lead correlates with blood lead (33, 34) and explains around 20% of the variance in blood lead, depending on seasonality (33) and hormonal and other endogenous and environmental stimuli, influencing the balance between bone formation and resorption (34). Recirculation of lead from bone explains why there is a lag time when occupational (32) or environmental (35) lead exposure drops. These lead toxicokinetics are important in the interpretation of our current results. A narrative review on the association of neurocognitive function and lead exposure compiled evidence from 21 studies published from 1996-2006. All studies had assessed bone and blood lead as biomarkers of internal exposure, 15 in occupational studies and 6 in environmental settings (5). At exposure levels representative of contemporary environmental exposure, associations of cognitive function with biomarkers of cumulative dose (mainly lead in tibia) were stronger and more consistent than associations with blood lead levels as assessed by concurrent, cumulative or peak blood lead levels (36). Conversely, studies of currently exposed workers generally found associations that were more apparent with blood lead levels (36). Given the persistence of lead in the human body, both bone and blood lead increase with advancing age (33, 34). Consequently, with advancing age, the blood lead concentration reflects environmental exposure levels stretching further back in time. In the United States

(scienceprogress.org/2008/10/a-brief-history-of-lead-regulation), lead-containing paint was only effectively banned in 1976 and leaded gasoline was completely phased out only in 1995 (37). Mean blood lead levels in the United States decreased from 13.1 gg/ dL in the National Health and Nutrition Examination Survey II (NHANES II; 1976-1980) (38) to 2.76 gg/dL in NHANES III (1988-1994) (39) and to 1.64 gg/dL in NHANES IV (1999-2002) (28). Our study moves the field forward because the cumulative lead dose in our young participants must reflect present-day environmental exposure levels and, as suggested in a systematic literature review (15), we addressed variability between people by comparing neurocognitive test results before and after occupational exposure.

Neurocognitive functions are integrated cognitive processes linked to multiple particular cerebral areas, neural pathways or cortical networks in the brain (40, 41). In this study, we evaluated the neurocognitive function, using two complementary tests, which are sensitive to detect mild cognitive impairment under lead exposure (31, 42). On the one hand, the DST assesses complex attention, motor speed, visual-perceptual functions and executive function (42). Functional magnetic resonance imaging (fMRI) studies in young healthy adults (43) and octogenarians (44) showed that taking the DST activated the frontal parietal cortical network, probably reflecting visual search and working memory processes (43, 44). The ST provides information on processing speed, selective attention, automaticity and parallel distributed processing (45-47). In fMRI studies, taking the ST activated the anterior cingulate, insula, premotor and inferior frontal brain regions (48).

Strengths and limitations

The strong points of our study are its longitudinal design (15), the young age of its participants the starting blood lead level representative of current environmental exposure levels, and the stringent quality control of the blood lead concentration. However, our study also has limitations. First, the attrition rate among the 506 workers who participated in the baseline examination, but defaulted from follow-up amounted to 217 (42.9%). However, according to the SPHERL protocol (16), the anticipated attrition rate was estimated to be 50% and >500 workers had to be enrolled. We met these numbers. The baseline characteristics of workers included or not included in the analyses were to a large extent similar (supplementary table S3), so that it is unlikely that attrition significantly biased the study results, although bias due to unmeasured confounders can never be excluded. Second, the study was primarily powered for blood pressure and renal outcomes, while neurocognitive function was among the secondary outcomes. However, the association sizes between the changes in the neurocognitive indexes and blood lead were small and sample size does impact on significance levels, but has no direct link with estimates of association size. Third, due to a software error, the ST with congruent trails was only administered to 168 (62.9%) participants at follow-up. However, the results of the congruent and incongruent tests were consistent. Fourth, the observerparticipant pairing was not standardized throughout the study (supplementary tables S11 and table S12), which might have introduced bias in the observed baseline to follow-up changes in neurocognitive function. Fifth, the median 2-year follow-up might have been too short for neurocognitive effects associated with lead exposure to become evident. For this reason, as anticipated (16) the cohort will be kept in follow-up for an additional two years. Finally, although the ethnic distribution of the workers was representative for the population at the recruitment sites, women were under-represented. Only 11.6% of 267 analyzed participants were female, which precluded analyses stratified by sex.

Concluding remarks

At the exposure level in our study, we failed to demonstrate a consistent and significant association of changes in neurocognitive function in the workers with an over threefold increasing blood lead concentration over the 2-year follow-up. Lead exposure represents an occupational and environmental health hazard that should be addressed worldwide. Our findings and the contradictory literature (15) suggest that in adults there is no causal link between neurocognitive impairment and low-level lead exposure as reflected by blood lead levels below 30 pg/dL. In weight-of-the-evidence approaches, policy makers might account for our findings in setting thresholds for occupational and environmental lead exposure levels, so that the prevention resources of neurocognitive function are dedicated to the more important drivers of cognitive impairment, in particular low educational attainment, socio-economic deprivation, abuse of alcohol and recreational drugs, discriminating based on ethnicity, and not providing opportunities for

immigrants to assimilate the skills necessary for social integration (29).

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DETAILS

Subject:	Cognitive ability; Statistics; Mass spectrometry; Blood; Cognition; Longitudinal studies; Recycling; Questionnaires; Blood levels; Statistical analysis; Statistical methods; Cognitive tasks; Nurses; Workers; Confidence intervals; Blood pressure; Occupational exposure; Exposure; Occupational health; Mass spectroscopy; Lead content; Latency; Lead poisoning; Inductively coupled plasma mass spectrometry
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Industry mobility and disability benefits in heavy manual jobs: A cohort study of Swedish construction workers

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ABSTRACT (ENGLISH)

Objectives This study aimed to investigate whether change from the construction industry to work in other industries at age 45-55 years lowered risks of disability benefits (DB) later in life (60-64 years of age). We hypothesized that risks would be lowered the most among those changing from the heaviest occupations. Methods The study included men employed in the construction industry during 1971-1993. We selected workers from the largest occupational groups in heavy (concrete workers and painters) and less heavy (drivers, electricians and foremen) occupations. The occurrence of DB in 1990-2015 was retrieved from national registers. Regression analyses were used to calculate relative risks (RR) of DB at 60-64 years, comparing those working in other industries to those still in the construction industry at the age of 45, 50 and 55 years. Results Mobility away from the construction industry was related to lowered DB risks at 60-64 years in all selected occupations. Effects were most pronounced among those who, at 55 years of age, worked in an industry other than construction, with significantly reduced RR for DB among concrete workers [RR 0.63, 95% confidence interval (CI) 0.51-0.77], electricians (RR 0.61, 95% CI 0.47-0.77) and foremen (RR 0.78, 95% 0.63-0.96). Conclusions Risks for DB at 60-64 years of age were reduced among those who changed from construction work to other industries. Notable reductions were observed among workers originating from both heavy and less heavy occupations, and future studies should explore other factors, in addition to heavy workload, as motivators for leaving the construction industry.

FULL TEXT

Headnote

Objectives This study aimed to investigate whether change from the construction industry to work in other industries at age 45-55 years lowered risks of disability benefits (DB) later in life (60-64 years of age). We hypothesized that risks would be lowered the most among those changing from the heaviest occupations. Methods The study included men employed in the construction industry during 1971-1993. We selected workers from the largest occupational

groups in heavy (concrete workers and painters) and less heavy (drivers, electricians and foremen) occupations. The occurrence of DB in 1990-2015 was retrieved from national registers. Regression analyses were used to calculate relative risks (RR) of DB at 60-64 years, comparing those working in other industries to those still in the construction industry at the age of 45, 50 and 55 years. Results Mobility away from the construction industry was related to lowered DB risks at 60-64 years in all selected occupations. Effects were most pronounced among those who, at 55 years of age, worked in an industry other than construction, with significantly reduced RR for DB among concrete workers [RR 0.63, 95% confidence interval (CI) 0.51-0.77], electricians (RR 0.61, 95% CI 0.47-0.77) and foremen (RR 0.78, 95% 0.63-0.96). Conclusions Risks for DB at 60-64 years of age were reduced among those who changed from construction work to other industries. Notable reductions were observed among workers originating from both heavy and less heavy occupations, and future studies should explore other factors, in addition to heavy workload, as motivators for leaving the construction industry.

Key terms construction industry; heavy work; Sweden; work ability.

Longer life expectancy and the growing amount of elderly in proportion to active workers have created a need to expand work life duration (1). One way is to raise the statutory retirement age, but longer life expectancy does not necessarily equal delayed age-related disabilities (2). Above 60 years of age, most persons have some chronic disease. A study of Finnish municipality workers showed that 74% were diagnosed with a chronic disease around the age of 60 (3). Old age and chronic diseases are thus, unsurprisingly, the strongest predictors for premature labor market exit through disability benefits (DB) (4, 5).

As health and physical capacity decrease by age, older persons with heavy physical jobs should be more at risk of DB than others. The construction industry is a large occupational sector in many countries. In Sweden it employs about 6% of the workforce. The work is mostly physically demanding, including heavy physical workload, repetitive movements and working in demanding postures. Due to a high prevalence of musculoskeletal disorders, construction workers tend to leave the labor market earlier than others, often through DB programs (6-8).

Work modification - a process aimed at enhancing the match between job conditions and a workers resources - may promote a sustainable work life in those with reduced work ability. In persons with lower back pain (9) or injuries (10, 11), reduced work demands or tasks reassignments appear to be most effective. Work time control, eg, over breaks or flexible working hours, seem to benefit continued work at an older age in general (3). When such modifications are not possible in a current job setting, changing jobs may be the only option for improved job conditions and remaining at work. There is evidence that older persons in physically heavy occupations with musculoskeletal disorders, foremost lower back or spine disorders, change jobs more often than others (12-14). However, whether this expands working life is unknown, partly since such associations cannot be investigated by experimental studies, but has to rely on observational studies, which require a large study population and a long follow-up.

This study aimed to investigate whether changing from the construction industry to work in other industries, evaluated at the age of 45-55 years, lower the risks of DB at 60-64 years. We hypothesized that DB risks would be lower among those who changed to other industries, with the largest reductions occurring in those originating from the physically heaviest jobs.

Methods

This study aimed to evaluate industry change and DB using the Swedish Construction Worker Cohort and Swedish national register data. The Construction Worker Cohort consists of 389 132 men and women, who were employed in the construction industry and attended health examinations during 1971-1993 through Bygghälsan (the Foundation for Occupational Safety and Health in the Swedish Construction Industry). The examinations were free of charge as part of the occupational health services, and all workers were invited on a routine basis at 2- to 5-year intervals. About 80% of those employed in the Swedish construction industry during that time have participated at least once. The cohort has been described in detail elsewhere (15).

DB are included in the Swedish sickness benefit welfare program and provide financial support for people with long-lasting reduced work ability. All residents aged 19-64 years, including unemployed persons, are covered. Eligibility requires >25% reduced work ability for at least one year as assessed by a physician. Benefits are either granted for

a limited period of time with reassessments or, if the work ability is assessed as permanently reduced, benefits can be granted until entering old age retirement.

For the analyses, we selected the largest occupational groups with either heavy or lighter physical work. The job title used was determined by the title in the employment contract held at the time of the first medical examination. The division of heaviness was based on physical long-term cardiovascular load as a measure of work intensity, which was available through a job exposure matrix of heart rate measurements (16). The division also corresponds with practical knowledge of the construction industry. Two of the most common occupations - concrete workers and painters - were selected as the heaviest physical jobs based on cardiovascular load. Three of the other largest occupations were carpenters, electrician and foremen. We included electricians who have a lighter cardiovascular load than concrete workers and painters, even though they sometimes work in demanding postures. The foremen are mostly previous blue-collar workers, ie, they are from a similar socioeconomic group as the other three jobs, but with less manual tasks and therefore less load. We did not include carpenters as their work conditions are varied, and we wanted to include groups that could be generalized into heavy or lighter physical load. Instead, we included drivers of trucks, cranes and heavy equipment. Electricians and foremen have less load but typically also higher qualifications. By including drivers, we could observe a group with a similar level of qualifications to concrete workers and painters but who work in sitting postures and on average have a low cardiovascular load.

We compared the occurrence of DB at age 60-64 years among men in the selected occupations, who worked in another industry the year they turned 45, 50 and 55 years of age, to those who remained in the construction industry. Change of industry could have occurred at any time between inclusion and the evaluated ages, but since many, especially in younger ages, change back and forth between industries, we defined "change of industry" as not working in construction the calendar year that the individual turned 45, 50 or 55 years of age. Hence, change of industry was defined as yes/no at the evaluated age, ie, persons who left construction at age 45 could be counted as being in the industry at age 55 years if they returned to construction. Information on type of industry after inclusion and timing of granted DB between 1990-2015 was available through the national register the longitudinal integrated database for health insurance and labor market studies (LISA), provided by Statistics Sweden, which covers all Swedish residents aged >16 years. The LISA register was established in 1990. Due to the Sweden's unique personal identification number, the register data could be matched to all studied men on an individual level.

Comparing occurrence of DB between those who had changed industry with those who remained in the construction industry, evaluated at the ages of 45, 50 and 55 years, meant that three sub-cohorts were used (table 1). Each sub-cohort included all uncensored persons from inclusion at clinical examination until age of evaluated industry change (1990-2010). The ages 45, 50 and 55 were chosen since physical capacity and health starts to decline at these ages. The follow-up started in the calendar age when the men were 60 years, if they lived in Sweden, currently did not uphold DB and were registered in LISA as having a job (full or part-time). Each individual was followed until: (i) the first occurrence of DB, (ii) the calendar year they turned 65 years, (iii) death, (iv) emigration, or (v) 31 December 2015. For example, a person who was 45 years in 1991 was followed until 2006-2010 at the latest. The analyses were restricted to men as there were too few women in the selected occupations to make analyses of women feasible. The age window 60-64 years constitutes an age span in close proximity to statutory retirement age where chronic diseases and DB are most common.

Relative risks (RR) and 95% confidence intervals (95% CI) for DB between those who changed from the construction industry and those who remained were estimated by a negative binomial regression model of incidence rates using log link. Persons who stayed in the construction industry were the reference. The analyses were adjusted for age between 60-64 years (1-year intervals), smoking habits at health examination [non-, ex-, moderate, and heavy smokers (15+ cigarettes)] and body mass index (BMI) (18.5-24.9, 25-29.9, and 30-34.9 kg/m2). Persons with unknown smoking habits (6.7%) and unknown BMI or BMI <18.5 or >35 kg/ m2 were excluded (3.1%). A 95% CI not including 1 was considered as statistically significant. The Ethical Review Board at Umeå University approved this study (2016/308-31).

We carried out sensitivity analyses by expanding change of industry to the age intervals 44-46, 49-51 and 54-56

years of age, but these analyses displayed similar results as those presented and are not included in this paper. To study a possible influence of poor health on risks of DB at 60-64 years of age, additional analyses were restricted to persons that had not been hospitalized around the evaluated ages of change in industry, eg, not been hospitalized between 44-46 years for those observed at 45 years of age. The results were similar as in the main analyses and are not presented.

Results

Table 1 shows the number of DB cases by sector mobility, stratified by age cohorts. The proportion of workers who, at the age of evaluation, worked in another industry was lowest among painters, while the other job categories displayed similar proportions of industry mobility (43-50%). Characteristics of the sub-cohorts at the first clinical examination and at follow-up are available in the supplementary material (www.sjweh.fi/show_abstract.php?abstract_id=3932), appendix 1a-c.

In figure 1, RR are displayed for different age groups. Although several results were non-significant, all analyses displayed reduced RR for DB at 60-64 years of age among those who had shifted out of the construction industry compared to those who had remained.

Results also revealed differences between the selected occupations and ages at evaluated industry change. Among concrete workers, those who worked in another industry had statistically significantly lowered DB risks at 60-64 years in analyses that examined a move from the construction industry at age 50 (RR 0.74, 95% CI 0.58-0.94) and 55 (RR 0.63, 95% CI 0.51-0.77), compared to those who remained in the industry. Similarly, results for electricians showed lowered estimates for workers who changed industry at age 50 (RR 0.66, 95% CI 0.51-0.87) and 55 (RR 0.61, 95% CI 0.47-0.77). Despite overall lower DB risks among foremen transferring to other industries, only results of industry change evaluated at 55 years were statistically significant (RR 0.78, 95% 0.63-0.96).

For persons that worked in another industry at age 45, all risk estimates were <1 but non-significant and CI were wide. Industry change evaluation at age 50 displayed lowered DB risks than those who remained in construction, which was further lowered at aged 55 compared to 50 years. The age effects of lower DB risks in older ages were observed in all occupations except drivers and appeared most conspicuous in concrete workers, electricians and foremen.

Discussion

This study showed that movement out of the construction industry to other industries was related to lowered risks for DB at 60-64 years of age. Results were most pronounced among those whose industry change was evaluated at age 55, with considerably lowered RR among concrete workers, electricians and foremen.

We investigated the study objectives in five occupational groups of varied physical workload heaviness. Industry change was related to lowered DB risks in all groups, a finding which concurs with similar studies (12-14), as mobility away from heavy occupations presumably reduces work load, improves health, or at least enhances the match between job demands and a worker's capacity. The reference group, who remained in heavy jobs, face higher risks for onset of or worsened chronic diseases.

In contrast to our hypothesis, that change from the heaviest jobs would be the most beneficial: the largest DB risk reductions were found in both heavy (concrete workers) and less heavy occupations (electricians and foremen). In a study based on the same cohort, concrete workers and painters represented occupations with the most lost working years due to DB, while foremen and electricians displayed the least lost working years (8). Thus, it seemed probable the largest DB reduction would occur among concrete workers and painters. Other similar studies have not analyzed occupations according to variations of workload (12, 13), and it is unknown if DB risks generally are the most lowered in those leaving the heavier occupations.

Results among the heaviest occupations also displayed deviations as significant lowered risks of DB among concrete workers, but rather small and nonsignificant effects among painters, were observed. A suggested explanation is that painters have better opportunities for self-employment, through which they remain in the same industry but with better control possibilities over job offers, work hours and breaks, which enhance sustainable work ability (3). Painters had the lowest percentage of industry mobility, providing some support to this theory.

Among the lighter occupations, electricians displayed the most lowered DB risks. Among foremen, industry change was also related to statistically lowered risks of DB, but effects were smaller and only statistically significant in those evaluated at age 55. We lack information on reasons for industry change and can only speculate. Electricians and foremen typically not only have lighter work tasks but also higher qualifications. If leaving the construction industry, less straining jobs may be available, eg, electricians may move to maintenance and foremen become instructors. Industry mobility may then, at least partly, be determined by pull factors to attractive jobs rather than poor health and heavy work. Among drivers, an occupation with lighter work tasks but lower qualification requirements, DB risk were small and non-significant in all age groups. The different patterns between the selected occupations indicate that several factors other than heavy workload, are important. Since DB is likely determined by conditions in the new job, we examined which industries the workers changed to (supplementary appendix 2a-c). However, since we only had access to industry, not occupation, there is little precision on new work conditions. Most notably, there was high industry mobility to the financial sector among foremen (around 43%), likely since they have higher qualifications than other occupations. Otherwise, patterns were similar for most occupational groups and added little explanation for differences in DB risks.

The results displayed age effects, as those whose industry mobility were evaluated at 45 years of age generally had the least reduced DB risks, while those working outside the construction industry at age 55 displayed the most reduced risks. Industry change was defined as not working in the construction industry the calendar year the worker turned 45, 50 or 55 years of age. It could be, foremost in younger ages, that workers change back and forth and the industry change is not permanent. We could only study industry change from 1990 when the LISA register started. Some men changed industry a few times during the study period. Therefore, we also studied men who had the same job during 3-year periods around 45, 50 and 55 years, but the findings were similar.

Most occupations displayed high numbers of workers changing industry, even among those evaluated at age 45. In some, industry change may relate to the period (ie, the 1970s and 1980s), when it was quite common for construction workers to be employed in short-term contracts that ended when the building project ended. If no other construction work was available, the worker would seek employment in other industries requiring similar skill levels. Repeated changes, foremost in older ages, may be a sign of trying other jobs due to health problems, but also a sign of skills and health as such workers have more possibilities to find other jobs with better pay. If health is a decisive factor, the time between change and follow-up is of importance. Those evaluated at age 55 had the most reduced risks and were in closest proximity to follow-up, which started at 60 years of age. As health starts to decline sharply around age 55 in physically demanding jobs (12), industry change in this age group would be most driven by health and thus benefit the most. Contrary, if good health is positively correlated to industry change, it will also have a larger influence for reduced DB risk in this age group, as poor health is less common among younger workers. Driver was the only occupation in which no age effects appeared, perhaps because the work conditions are lighter but advancement outside the construction industry is limited.

Results may also relate to selection effects. Workers who succeeded in changing jobs may represent a selection of individuals with better health and a personality type, who want a long working life and choose to invest in a more fulfilling and stimulating job. After changing industry, additional positive effects may have followed and reignited the motivation to remain at work. The presence of such persons will over-estimate positive effects of changing industry. In contrast, if changing industry due to chronic diseases, the effect could be underestimated due to higher prevalence of poor health. The presented results are similar to those in the sensitivity analyses, in which all hospitalized persons were excluded, and poor health may have less effect on underestimations. However, most disorders among construction workers are treated in outpatient care. There are no national registers of outpatient care that cover our observation period.

Our observation period stretches over a large time span, during which incidence of DB has varied greatly in this cohort. Welfare legislations have also varied over this time, and - during some periods - eligibility criteria for DB, reassessment of work ability, and return to work among persons >60 years have either been more generous or stricter (17). Since our follow-up is determined by age and not calendar year, it is difficult to determine the effects of

welfare legislations. Furthermore, during times of recession, it is common that workers with lower capacity and chronic diseases are pushed out of their jobs, underestimating the effects of industry mobility (14). But in periods of recession, short-term sick leave tends to decrease due to fear of losing one's job (18). As a consequence, long-term sick-leave and DB increase, as many with poor health continue working until onset of chronic illnesses. Meanwhile, in times of economic boom and more work opportunities, many may change industry for reasons unrelated to health. Variations in societal economics can therefore lead to both under- and overestimation of DB risk among those changing industry.

Limitations

The cohort consisted of male blue-collar workers in the construction sector, and results may not be generalized to women or other occupational sectors. This is a major limitation since there is evidence that construction workers have a higher risk for DB, even compared to other blue-collar workers in eg, metal and chemical industries, or mining (19). We were also unable to examine risk reduction of DB from mobility between employers while remaining in the same industry or if industry change resulted in changed job conditions. Another limitation is that we did not assess if persons granted DB at 60-64 years returned to work. Given the heaviness of construction work, it is unlikely that someone >60 years of age would resume work if entering DB programs, but it cannot be ruled out.

Strengths

This study examined the effects of moving out of the construction industry using a large cohort with a long follow-up time and individual information on several important covariates.

The cohort size allowed division into sub-cohorts by occupational groups with variations of physical workload. Since age is an important covariate in this context, we also took this into consideration by division into age-specified cohorts. Due to Sweden's usage of unique personality numbers and high-quality national registers, we also have accurate data on timing of DB and industry. Finally, considering that low socioeconomic status and less education consistently has been identified as a strong determinant for DB and lost working years (20), it was also beneficial to have access to a large cohort of only construction industry workers.

Concluding remarks

Our results found reduced DB risks at 60-64 years of age among those who changed from construction work to other industries. Notable reductions were observed in both heavy and less heavy jobs, indicating that factors other than physically demanding work could be important for industry mobility and DB; this needs to be explored further. Still, supportive functions that facilitate job mobility close to statutory retirement age may increase work participation in older workers.

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Sidebar

Söderberg M, Stattin M, Robroek SJW, Burdorf A, Järvholm B. Industry mobility and disability benefits in heavy manual jobs. A cohort study of Swedish construction workers. Scand J Work Environ Health. 2021;47(3):217-223. doi:10.5271/sjweh.3932

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DETAILS

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Shift work, work time control, and informal caregiving as risk factors for sleep disturbances in an ageing municipal workforce

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ABSTRACT (ENGLISH)

Objectives This study aimed to examine the contribution of shift work, work time control (WTC) and informal caregiving, separately and in combination, to sleep disturbances in ageing employees. Methods Survey data were obtained from two prospective cohort studies with repeated measurements of working conditions, informal caregiving, and sleep disturbances. We used fixed-effect conditional logistic regression analysis to examine whether within-individual changes in shiftwork, WTC and informal caregiving were associated with changes in sleep. Secondary analyses included between-individuals comparison using standard logistic regression models. Results from the two cohorts were pooled using meta-analysis. Results Low WTC and informal caregiving were associated with sleep disturbances in within-individual analyses [odds ratios (OR) ranging between 1.13 (95% confidence interval 1.01-1.27) and 1.48 (95% CI 1.29-1.68)] and in between-individuals analyses [OR 1.14 (95% CI 1.03-1.26) to 1.33 (1.19-1.49)]. Shiftwork alone was not associated with sleep disturbances, but accumulated exposure to shiftwork, low WTC and informal caregiving was associated with higher risk of sleep disturbances (OR range 1.21-1.76). For some of the sleep outcomes, informal caregiving was related to a higher risk of sleep disturbances when WTC was low and a lower risk when WTC was high. Conclusions Informal caregiving and low WTC are associated with risk of sleep disturbances among ageing employees. The findings also suggest that low WTC in combination with informal caregiving may increase the risk of sleep disturbances whereas high WTC may alleviate the adverse impact of informal caregiving on sleep.

FULL TEXT

Headnote

Objectives This study aimed to examine the contribution of shift work, work time control (WTC) and informal caregiving, separately and in combination, to sleep disturbances in ageing employees.

Methods Survey data were obtained from two prospective cohort studies with repeated measurements of working conditions, informal caregiving, and sleep disturbances. We used fixed-effect conditional logistic regression analysis to examine whether within-individual changes in shift work, WTC and informal caregiving were associated with changes in sleep. Secondary analyses included between-individuals comparison using standard logistic regression

models. Results from the two cohorts were pooled using meta-analysis.
Results Low WTC and informal caregiving were associated with sleep disturbances in within-individual analyses [odds ratios (OR) ranging between 1.13 (95% confidence interval 1.01-1.27) and 1.48 (95% CI 1.29-1.68)] and in between-individuals analyses [OR 1.14 (95% CI 1.03-1.26) to 1.33 (1.19-1.49)]. Shift work alone was not associated with sleep disturbances, but accumulated exposure to shift work, low WTC and informal caregiving was associated with higher risk of sleep disturbances (OR range 1.21-1.76). For some of the sleep outcomes, informal caregiving was related to a higher risk of sleep disturbances when WTC was low and a lower risk when WTC was high.
Conclusions Informal caregiving and low WTC are associated with risk of sleep disturbances among ageing employees. The findings also suggest that low WTC in combination with informal caregiving may increase the risk of sleep disturbances whereas high WTC may alleviate the adverse impact of informal caregiving on sleep.
Key terms insomnia; municipal employe; older employee; shift worker; working hour.
Sleep disturbances are common among working populations (1), and ageing is associated with many negative changes in sleep, such as more frequent arousals, shorter sleep duration and decreased slow-wave sleep (2, 3). Sleep disturbances increase the risk of various health problems, including the metabolic syndrome, type 2 diabetes, and premature mortality (4-6). Among the employed, sleep disturbances and insufficient sleep have also been associated with impaired work performance, occupational injuries, absenteeism, work disability, and increased healthcare costs (6-8). The high prevalence of sleep disturbances and the related burden on the individual, the workplace, and the healthcare system underlines the need to find new targets for prevention.
Time-related factors that may make life irregular or hectic have gained increasing attention in studies of factors that may contribute to sleep disturbances. These include, for example, shift work (ie, work schedule including irregular or unusual working hours, even during night) and informal caregiving (ie, caring for ill or disabled child, spouse or other member of a person's social network). Shift work can disrupt the normal sleep-wake cycle and may increase the risk of health problems due to sleep deprivation, impaired sleep and fatigue (9, 10). Furthermore, the prevalence of informal caregiving increases with age and research has shown associations of informal caregiving with sickness absence (11) and cardiovascular diseases (12), which may be partially mediated through high levels of stress and sleep disturbances (13-16). However, the studies on the association between informal caregiving and sleep have rarely focused on participants in employment and their working conditions (16, 17) although, eg, in the Europe and US, more than 50% of caregivers are employed (18, 19). The few longitudinal studies on the association between informal caregiving and sleep disturbances among employees have produced inconsistent findings (20, 21).
In addition, it is unclear whether having influence on working hours at the workplace, that is, work time control (WTC), might alleviate the risk of sleep disturbances among ageing employees. WTC refers to, for example opportunities to self-determine the length of the workday, the start and end times of a duty period, taking of breaks, running of private errands during work, and the scheduling of vacations and other type of leave from work (22-24). WTC is thus flexibility based on employees' rather than employers' needs and the positive effect on well-being has been assumed to result from greater possibilities to balance one's resources and timeuse to better cope with job- and home-based demands. There are at least three factors underlying the positive effects of employee-oriented flexibility in working life: (i) sustainable healthy work obtained by increased opportunities for employees to recover from demands at work; (ii) improved balance between work and family life in a changing working life; and (iii) benefits for the organizations in terms of attaining and maintaining a committed, skilled and healthy workforce (24). Low WTC has previously been associated with sleep disturbances (22, 25-27), depressive symptoms (27-29), work accidents (30) and sickness absence (23), and high WTC has been shown to be associated with extended working life among older employees (31). WTC may therefore be one of the modifiable factors that protect older employees from sleep disturbances.
In the present study, we utilized data from two prospective cohort studies to examine whether shift work, low WTC and informal caregiving, separately or in combination, are associated with increased risk of sleep disturbances among older employees. To minimize confounding by between-individual differences, we used repeated measurements of shift work, WTC, and informal caregiving and applied the fixed-effect method (within-individual

analysis), which uses individuals as their own controls, thus controlling for all time-invariant measured and unmeasured confounders.

Methods

Study populations

The participants were from the Finnish Public Sector (FPS) study (32) and the Finnish Retirement and Ageing (FIREA) study (33), ongoing cohort studies of municipal employees. In FPS, we used survey responses from waves 2012, 2014 and 2016. We restricted the data to those aged >50 who had responded to a minimum of two surveys (N=24 418) to be eligible for the fixed-effect (withinindividual) analysis (see supplementary material, www.sjweh.fi/show_abstract.php?abstract_id=3937, for the final number of participants in each fixed-effect analysis). For a secondary analysis examining onset of sleep disturbances with repeated data (a between-individuals approach), we used the data of those participants who had survey data of two subsequent time points (2012-14 or 2014-16 or both; N=23 613). For this analysis, we restricted the data to those without sleep disturbances at the baseline survey (N=17 615).

The FIREA study cohort consists of municipal employees whose personal statutory retirement date (ie, individual old age pension date) was between 2014-2019. The first survey was sent to the eligible population 18 months before the statutory retirement date. In Finland, the public sector employees' retirement ages are regulated by the Public Sector Pensions Act and the personal statutory age is calculated by the Pension Institute for Public Sector Employees (Keva). Thereafter, annual questionnaires have been sent to the respondents, and the present data include up to six surveys collected between 2013-2019. We included all study waves during which the participant was still employed and excluded study waves after the participant retired. The participants who were employed on at least two survey points (N=2838) formed the analytical sample for the fixed effect (within-individual) analysis. For between-individuals analysis, we used the data of those participants who had survey data of two subsequent time points (N=2771). We restricted the data to those without sleep disturbances measured at baseline (N=2084).

The FPS and FIREA studies were conducted according to the Declaration of Helsinki and approved by the Ethics Committees of the Hospital Districts of Helsinki and Uusimaa and Southwest Finland, respectively.

Measures

The FPS and FIREA studies included identical information regarding the variables used in the current study. Work factors included shift work, based on survey responses and categorized in three groups: day work, shift work without night work, shift work with night work. WTC was also based on survey responses and measured with a 7-item Likert scale from 1 (very little) to 5 (very much) regarding the respondent's ability to influence the following aspects of working hours: duration of work day, beginning and ending times, taking breaks, attending personal affairs, scheduling work shifts, taking vacations/days off, and taking unpaid leave (22, 23). The mean of 7 items was dichotomized at the median (2.60 in FPS and 2.75 in FIREA), thus 'high WTC' indicates both intermediate and high scores of WTC. Informal caregiving was assessed with the following question: "Do you provide care to a family member or relative who is unable to take care of himself or herself because of age, illness or disability?" (yes versus no; giving care for own healthy children was not accounted for).

Sleep disturbances were assessed by the Jenkins Sleep Problem Scale (34), which is a self-report scale including the following four types of insomnia symptoms: difficulty in falling asleep, frequent awakenings (ie, difficulty in maintaining sleep), early awakenings and non-restorative sleep (ie, feeling tired and worn out after normal amount of sleep). Participants reported the frequency of sleep disturbances during the past 4 weeks. As previously, the response scale from 1=never to 6=almost every night was dichotomized into no sleep disturbances (0-4 nights/week) and sleep disturbances (5-7 nights/week) for each type of symptoms (22, 35). Any sleep disturbances were indicated as reported having any of the four sleep disturbance types.

Covariates were age, sex, SES, and self-rated health. Age, sex and SES were based on employers' registers, SES including occupational titles that were coded according to the International Standard Classification of Occupations (ISCO) and categorized into three groups: high (ISCO classes 1-2, eg, teachers, physicians), intermediate (ISCO classes 3-4, eg, registered nurses, technicians), and low (ISCO classes 5-9, eg, cleaners, maintenance workers).

Self-rated health was a single-item perception of one's current health dichotomized into optimal (good, rather good) and non-optimal (average, rather poor, poor) health. Additionally, the following health behaviors were included and they were based on survey responses: height and weight from which body mass index (BMI) was calculated and dichotomized into BMI >25 kg/m2 (overweight/obesity) versus less than that; current smoking versus non-smoking; alcohol consumption (weekly average consumption of beer, wine and spirits transformed into grams of alcohol and further dichotomized as risky; >288g among men; >192g among women) (36); and physical activity, which was measured as metabolic equivalent task (MET) (37) hours per week, dichotomized as <14 hours (low physical activity) versus >14 hours (moderate/high physical activity).

Statistical analysis

We used fixed-effect method (38) with conditional logistic regression analysis in which the analyses are performed within individual. This design enables data from longitudinal cohort studies to be used in a casecontrol design where each individual serves as his/her "case" (disturbed sleep) and "control" (undisturbed sleep). This study design allows individuals to serve as their own control and therefore bias arising from all time invariant factors, such as genetics, response style and personality, is controlled for. To be eligible to the fixedeffect (within-individual) analysis the participants had to respond to at least two study waves. The basic requirement in the fixed-effect method is that the respondent have a change in the outcome (ie, are at one point/study wave cases and another time point controls in relation to sleep disturbance), and there is also variation in the exposure (shift work, WTC, or informal caregiving) in the selected population. The analyses estimate whether working shift work, giving informal care and reporting WTC varied when the participants reported sleep disturbances compared to time(s) when they reported no sleep disturbances.

To estimate a possible dose-response pattern between risk factors and sleep disturbances, we calculated a cumulative exposure variable that described the number of risk factors [0, 1, 2-3 risk factors out of (any form of) shift work, informal caregiving and low WTC].

As a secondary analysis, we undertook repeated measures logistic regression analysis with generalized estimating equations (GEE) method (39) to predict the onset of sleep disturbances among those without sleep disturbances at the baseline survey (a between-individuals approach). The repeated measurements were nested within participants (ie, one individual can contribute to more than once to the prospective design), thus, we used GEE that takes into account the interdependence of observations of the participants that have multiple data points.

Results were presented as odds ratios (OR) with 95% confidence intervals (CI) and the models were adjusted for socio-demographic factors (age, sex, and SES) and health-related factors (overweight/obesity, smoking, risky alcohol use, low physical activity, and self-rated health). In the within-individual fixed-effect analyses, the health-related covariates were used as time-dependent covariates and taken from each measurement point. In the between-individuals analysis with GEE, the covariates were taken from the baseline survey.

We also calculated variables combining the presence or absence of high versus low WTC and other risk factors (any shift work, informal caregiving) to examine whether WTC as a modifiable factor at work plays a role in the potential associations. To examine whether there was statistical interaction between variables, ie, whether there were the joint exposure effects of WTC with shift work or informal caregiving on sleep disturbances, we calculated relative excess risk due to interaction (RERI) which provides a formal assessment of departure from additivity of effects on a relative risk scale (40).

After conducting the analyses separately for FPS and FIREA cohorts, we performed fixed-effect metaanalyses (41) to obtain pooled estimates for the associations. I2 statistics was used to test the heterogeneity of results between cohorts. Study-specific analyses were performed with SAS version 9.4 (SAS Institute, Cary, NC, USA) and meta-analyses with Stata 15 (Stata Corp, College Station, TX, USA).

Results

Descriptive statistics of the participants at their first survey point are presented in supplementary table S1. In FPS, the participants were younger (mean age 55 years) than in FIREA (62.5 years). FIREA participants had in general a better health and sleep profile than FPS participants, but otherwise the cohorts were rather similar with the majority

of participants being women (81% in FPS, 80% in FIREA). Of the FPS and FIREA participants, 18.8% and 24.1% of women and 13.7% and 9.0% of men reported shift work with or without night work, respectively, and 16.6% and 17.4% of women and 11.6% and 13.5% of men reported informal caregiving, respectively. Any sleep disturbances were reported by 22.0-30.4% of the participants. The most commonly reported sleep disturbance was frequent awakenings during the night (18.2-25.1%) and the least commonly reported sleep disturbance was difficulty falling asleep (2.5-5.8%). Women reported all types of sleep disturbances more often than men.

Figure 1 presents results from the pooled withinindividual associations between shift work, informal caregiving and WTC, as well as different forms of sleep disturbances. Overall, shift work with or without night work was not associated with any of the examined sleep outcomes. Informal caregiving and low WTC, in turn, were associated with a greater likelihood of all types of sleep disturbances (OR range 1.13-1.48; except the association between informal caregiving and frequent awakenings did not reach statistical significance). When cumulative exposure to 1->2 (ie, shift work, low WTC and informal caregiving) was assessed, the findings suggested a relatively consistent dose-response pattern; the greater the number of risk factors, the greater the odds of sleep disturbances (see also supplementary material). Study-specific estimates, number of participants in each analysis and I2 heterogeneity indices are presented in supplementary table S2. The I2 heterogeneity estimates suggested no significant heterogeneity between FPS and FIREA data.

The corresponding results from the between-individual analyses are shown in figure 2 and supplementary table S3. Again, shift work was not associated with the onset of sleep disturbances with one exception; shift work with and without night work were both associated with onset of difficulty in falling asleep (OR 1.48 and 1.45, respectively). Otherwise, the results replicated those obtained from the within-individual analyses; low WTC and informal caregiving were relatively consistently associated with all types of sleep disturbances, and, in many outcomes, there was a dose-response pattern with a greater number of risk factors. Of the 29 I2 heterogeneity estimates, 6 suggested significant heterogeneity between FPS and FIREA data (supplementary table S3).

Results from the within-individual analyses of the association of shift work and informal caregiving with sleep disturbances by the level of WTC are presented in figure 3. Generally, compared with a situation with high WTC and no shift work or informal caregiving, there was a small or non-significant association with sleep disturbances when the participant reported high WTC in combination with shift work or informal caregiving. Respectively, the likelihood of sleep disturbances was greater at times when the participant reported a combination of low WTC and shift work or informal caregiving. However, when we formally tested the interaction by RERI, only two of the RERI estimates were close to statistical significance; the combined association between WTC and informal caregiving predicting difficulty in falling asleep (RERI 0.52, 95% CI -0.004-1.048, P=0.052) and early awakenings (RERI 0.29, 95% CI -0.03-0.61, P=0.076). Study-specific estimates, number of participants in each analysis and I2 heterogeneity statistics are presented in supplementary table S4. The I2 heterogeneity estimates suggested significant heterogeneity between FPS and FIREA data only in one of the 30 meta-analyses.

Figure 4 shows results from between-individuals analyses on the relation of shift work and informal caregiving with new-onset sleep disturbances by the level of WTC. The findings largely replicate those obtained from within-individual analyses. RERI analyses suggested no departure from additivity. Study-specific data and heterogeneity estimates are presented in supplementary table S5, showing significant heterogeneity in one of the 29 meta-analyses.

Results from the models adjusted for time-dependent health-related covariates (within-individual analyses) and health-related covariates at the baseline survey (between-individuals analyses) are presented in supplementary figures S1-6. There were no major changes in the estimates although they attenuated to some degree.

Discussion

Results from two cohort studies of older employees showed that low WTC and informal caregiving were associated with various forms of sleep disturbances. We found evidence of a dose-response pattern suggesting a greater likelihood of sleep disturbances with increasing number of the three examined risk factors; shift work, low WTC and informal caregiving. Furthermore, our withinindividual analyses showed that informal caregiving was related to a

higher risk of sleep disturbances when WTC was low, and to a lower risk when WTC was high.
Shift work alone was generally not associated with sleep disturbances, which is consistent with a recent systematic review of longitudinal studies on shift work and sleep (42). That review found a small non-significant association between shift work and insomnia symptoms (RR 1.16). We found an association between shift work and one symptom of sleep disturbances; compared to day workers, those working in shifts without or with night work had 1.45 and 1.48-fold increased risk of new-onset difficulties in falling asleep in between-individuals analysis. Previous studies suggest that within-individual changes in shiftwork are associated with sleep problems (10), and that a change from non-night to night work increases the risk of common mental disorders (CMD), a correlate of sleep problems, while moving back from night to non-night work increases recovery from CMD (43). However, the study populations in previous studies were different as our cohorts included older employees only. In our study, those who are well-adapted to shift work may continue shift work longer, thus the association between shift work and sleep disturbances may be different among them due to selection. Municipal sector also offers possibilities for employees to transfer from shift to day work, thus our findings are not generalizable to other sectors, especially to those where quitting shift work is not possible. Previous research also suggests acute but not chronic sleep loss among shift workers and that the adverse effects of shift work on sleep may be linked to schedule-related aspects, such as speed of rotation (9). The association found between shift work and difficulty in falling asleep in our study may be linked to various mechanisms such as shift type or length, but also to circadian rhythm disruption, lifestyle changes, or job strain. Because our assessment of shift work only addressed shift work with and without night work and we did not have detailed information on the other shift work characteristics, such as speed of rotation, further research is warranted to examine in greater detail the contribution of different types of shift work.
Our findings on the relation between informal caregiving and sleep disturbances are in agreement with one of the two previous prospective studies which included female long-term care employees characterized as double- or triple-duty (caring for older adults and caring for older adults plus childcare) caregivers because of their occupational roles as nurses (20). In that study, within-individual analysis suggested a weak association with impaired sleep quality for double-duty but not for triple-duty whereas the between-individuals analysis suggested the reverse. In the other longitudinal study with two data points, giving up informal caregiving between baseline and follow-up was associated with reduced risk of sleep disturbances at follow-up (21).
We found an association between low WTC and practically all types of sleep disturbances. In a small longitudinal study of 39 employees, an increase in WTC was associated with improvements in psychomotor vigilance tasks (25). In another study, removal of low WTC could not explain the improved sleep following retirement (35). In our previous prospective study using the FPS data (22), the outcome was onset of sleep disturbances with the total Jenkins score and the used method was a standard between-individuals approach. Here we applied a within-individual approach, examined WTC in combination with shift work and informal caregiving, and analyzed each type of sleep disturbances separately. The mechanisms through which WTC may affect health and well-being include work-non-work balance and job satisfaction although evidence from intervention studies is still sparse (26).
We examined whether low levels of WTC could exacerbate the risk and whether high levels could alleviate the potential adverse impacts of shift work or informal caregiving on sleep disturbances. Although the RERI estimates indicating departure from additivity were not statistically significant, only cautious conclusions can be made of the joint associations. However, the results from the fixed-effect approach give some support to the hypothesis low WTC may increase the risk and high WTC may protect the employee from sleep disturbances when engaged in informal caregiving.
The specific strengths of our study are its large sample size and repeated measurements of exposures and outcomes, which enabled us to apply an advanced analysis method and compare it with a standard prospective approach. The two analytical approaches in our study produced relatively similar findings. We also had the possibility to adjust the models for several timevarying confounders. There were no major changes in the estimates although they somewhat attenuated. Selfrated health was strongly associated with sleep disturbances, thus, adjusting for self-rated health may have been conservative. This is because in addition to poor health status causing

sleep disturbances, the direction of association may be bidirectional; ie, poor sleep may also lead to a perception of impaired health (44).

The limitations of our study include that work factors and sleep disturbances were based on self-reports and sleep disturbances were not based on clinically determined sleep disorders. Although the Jenkins scale includes the factor 'non-restorative sleep' (ie, feeling tired and worn out after normal amount of sleep), we did not have a measurement of daily functioning due to sleep disturbances. Objective assessment of sleep would be the next step in this field, as in one study where improvements in objective sleep indicators were observed during a year with an improvement of WTC (25). When calculating the cumulative risk score, we used the number of risk factors for each participant as a measure of risk factor burden and did not weight them according to their assumed importance although some of the risk factors may contribute more to sleep disturbances than others. The validity of this assumption should be assessed in future studies. Our assessment of informal caregiving was coarse, therefore future studies are needed to evaluate, for example, the content and impact of tasks in caregiving as well as the difference between child and adult caregiving. A further aspect to which older employees' control over working hours may have an effect is caring of grandchildren. Regarding WTC, the incongruence between personal need for and access to WTC should be considered in future studies (45). Finally, the female-dominated public sector study population limits the generalizability of our findings to other populations, and we cannot exclude the possibility that some time-varying confounders that were not included (eg, stressful life events) could have biased the results.

In conclusion, this study suggests that informal caregiving and low WTC are associated with risk of sleep disturbances among ageing employees. The findings also suggest that low WTC in combination with informal caregiving may increase the risk of sleep disturbances whereas high WTC may alleviate the adverse impact of informal caregiving on sleep. Intervention studies are needed to confirm whether improving employees' influence over working hours at the workplace would reduce sleep disturbances and improve quality of sleep.

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Conflict of interest

None declared.

Sidebar

Virtanen M, Myllyntausta S, Ervasti J, Oksanen T, Salo P, Pentti J, Kivimäki M, Ropponen A, Halonen JI, Vahtera J, Stenholm S. Shift work, work time control, and informal caregiving as risk factors for sleep disturbances in an ageing municipal workforce. Scand J Work Environ Health. 2021;47(3):181-190. doi:10.5271/sjweh.3937

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DETAILS

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Amendments &corrections

Anonymous

FULL TEXT

Re: Sasaki N, Kuroda R, Tsuno K, Kawakami N. The deterioration of mental health among healthcare workers during the COVID-19 outbreak: A population-based cohort study of workers in Japan. Scand J Work Environ Health. 2020;46(6):639-644. doi:10.5271/ sjweh.3922

This erratum concerns table 1. There was an error in the results reported for Education among healthcare workers. The results were incorrectly reported as:

Education

<16 years 39 (35.1)

*16 years 72 (64.9).

The correct numbers are reported in the table to the left.

Sidebar

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Effects of changes in early retirement policies on labor force participation: the differential effects for vulnerable groups

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ABSTRACT (ENGLISH)

Objectives This study investigated the effects of a national early retirement reform, which was implemented in 2006 and penalized early retirement, on paid employment and different exit pathways and examined whether these effects differ by gender, income level and health status. Methods This study included all Dutch individuals in paid employment born six months before (control group) and six months after (intervention group) the cut-off date of the reform (1 January 1950) that fiscally penalized early retirement. A regression discontinuity design combined with restricted mean survival time analysis was applied to evaluate the effect of penalizing early retirement on labor force participation from age 60 until workers reached the retirement age of 65 years, while accounting for secular trends around the threshold. Results The intervention group postponed early retirement by 7.41 months [95% confidence interval (CI) 6.11-8.72], and partly replaced this by remaining 4.87 months (95% CI 3.60-6.24) longer in paid employment. Workers born after the threshold, annually earning €25 000-40 000, spent 1.24 months (95% CI 0.31-2.18) more in economic inactivity than those born before. The working months lost to unemployment increased by 1.50 months (95% CI 0.30-2.71) for female workers and 1.99 months (95% CI 0.06-3.92) for workers reporting multiple chronic diseases. Conclusions The national reform successfully prolonged working lives of older workers. However, workers with a middle income, female workers, and workers with chronic diseases were more vulnerable to premature exit from the labor market through unemployment or being without any income or benefit.

FULL TEXT

Headnote

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Key terms chronic disease; gender; older worker; regression discontinuity.

High and increasing youth unemployment rates in the 1970s and early 1980s (1) led to the implementation of an array of labor market policies to lower - or at least contain - these rates in most OECD countries. Policies included early retirement schemes for older workers that encouraged retirement before the statutory age (2). Early retirement

schemes made it financially unattractive for older people to remain in the labor force, resulting in a drop in the average effective retirement age for male workers from 68.4 in 1970 to 61.5 years in the mid1990s in Europe (3). Since the late 1980s, it became apparent that the widespread use of early retirement schemes was no longer financially sustainable. In most OECD countries, these early retirement schemes were operated on a payas-you-go (PAYG) basis where the working population contributes to the benefits obtained by retirees (4). However, population ageing has expanded the fraction of occupationally inactive population (5) and, therefore, increased the overall dependency rate, which placed a great financial burden on a relatively smaller working population. This so-called old-age dependency ratio will continue to rise in OECD countries reaching 53.2% in 2050, up from 19.5% in 1975 (4). Consequently, reform-ing public pension systems has become a central policy issue across developed countries since the early 1990s. This is also the case in The Netherlands, where the effective retirement age in the mid-1990s was even below the average retirement age in other OECD countries (61.0 versus 63.2 years, respectively) (3). In response to the financial concerns about the provision of pension income, the Dutch Government announced a reform in the summer of 2005 that would abolish the favorable tax deductions of contributions to early retirement schemes from 1 January 2006 onwards for workers born after 1 January 1950. While a pension reform was expected in the light of the pension debates in The Netherlands and other countries, the speed of the implementation and the sharp distinction between those born just before or after 1 January 1950 was unforeseen when the reform was announced.

Recent studies have concluded that the early retirement reform did prolong the working lives of Dutch workers, and this effect was stronger for males (6, 7). Using a similar basic dataset and the same policy measure as in the current paper, Boot et al (6) showed that a larger proportion of workers born in 1950 left paid employment through unemployment benefits. Our study contributes to the literature by building on existing studies and provides a more thorough and nuanced methodological evaluation of the reform. First, we combine a sharp regression discontinuity (RD) design with restricted mean survival time (RMST) analysis to estimate the causal effect of the policy on the time in months spent on the labor rather than changes in proportions in different work states. In this sense, the reform acted as 'a natural experiment' with a clear cut-off of the treatment assignment and provides estimates that are the closest to a randomized controlled trial in observational data (8). The combination with time analysis is rather scarce, and the current study serves as an example for future studies using panel data. This leads not only to a methodological improvement but provides us with more nuanced policy-relevant results by presenting a precise measure of the additional number of months spent in paid employment encouraged by the reform.

Second, we provide a more substantive contribution by exploring the heterogeneous effects of the reform on labor force participation by groups of workers categorized by gender, socioeconomic status and health. While previous studies have evaluated the impact of this reform by gender, they have ignored the other dimensions that have been reported as important determinants of early retirement (9, 10). For example, workers with poor health have an increased risk to leave paid employment through early retirement (11, 12). As early retirement, disability and unemployment benefits might act as communicating vessels to exit paid employment (13), it could be hypothesized that vulnerable workers choose or are forced to leave paid employment through other pathways with reduced social benefits when the pathway of early retirement becomes less attractive. Thus, exploring effect heterogeneity also among socioeconomic status and health is of great policy relevance.

Against this background, the objectives of this study were to (i) investigate the effects of the reform on duration of paid employment and the timing to other pathways out of paid employment and (ii) examine whether these effects differed across groups of workers.

Methods

Dutch pension system and reform

The Dutch pension system consists of three pillars. The first pillar is the old-age pension, which is the basic old age pension provided by the government to all Dutch residents when they reach the statutory retirement age. The second pillar comprises supplementary collective pension schemes, which are linked to specific companies or industries and most employees are entitled to benefits from this pillar. Early retirement schemes are determined by

negotiations between labor unions and employer organizations at the sectoral level. Participation in these collective schemes is mandatory for each individual employee, ensuring coverage by the sector pension, which allows workers to retire earlier. As contributions to the sectoral pensions were tax deductible, early retirement was the social norm among Dutch workers before the policy reform. On average, the first and second pillar both make up half of the pension entitlements (14). The third pillar consists of individual voluntarily built-up savings supplementary to the public and sector pensions.

The reform in this study relates to the second pillar of the Dutch pension system. Since 1 January 2006, early retirement schemes were integrated into the capital funded occupational pension system. As a consequence, early retirement benefits - which could be applied for from the age of 60 - were reduced to discourage early retirement. To exemplify how much the new system penalized early retirement compare a person born before 1950, who could retire at the age of 62 years and three months with 70% of their gross wage, to a worker born in 1950 or later who would have to work an additional 13 months to retire with the same percentage of gross wage or retire at 62 years and three months with only 64% of their gross wage as retirement income: a 6 percentage point drop in retirement income (7). This penalty applied for any individual born after 1950 who retired between the ages of 60 and 65 - the full legal retirement age.

Datasets and study population

Register data of Statistics Netherlands were used, cov- ering the entire Dutch population. The study sample consisted of individuals born six months before (control group) and six months after (intervention group) 1 January 1950. Individuals at the age of 60 years were included in the study because early retirement is possible from this age onwards. Of the entire population, 49.0% of the control group and 51.2% of the intervention group were working at baseline and thus included in our study.

These data were enriched with data on dispensed medicines and personal gross income from additional register data of Statistics Netherlands. Additionally, monthly information was included on main income components, social benefit pensions and gross wages, derived from the Dutch tax registers and stored in the social statistical database (SSB) (15). Individuals were followed up until the age of 65, the Dutch statutory retirement age at the time. Finally, information regarding whether and when a participant died during follow-up was also included. The final study sample consisted of 102 617 participants (49 501 and 53 116 in the control and intervention groups, respectively).

Employment status

Monthly employment status was defined based on the most important source of income and categorized into paid employment, disability benefits, unemployment benefits, early retirement benefits, and economically inactive. Employed individuals were defined as having their main source of income from paid employment. Selfemployment was not included in the current study as the regulations regarding social benefits differ substantially from those in paid employment. Early-retired persons were defined as those who received a company pension as their main source of income but had not reached the Dutch statutory retirement age yet. Unemployed persons received either unemployment benefits due to having lost their job or social security benefits. Receiving disability benefits was defined as when disability benefits represented >50% of total monthly income. Finally, economically inactive persons did not have personal income or benefits because they stopped for reasons such as being a homemaker or retired without receiving early retirement benefits. If an individual reached a specific exit pathway for >3 months, then this pathway was considered an actual event. Only the first event over time was considered in this study. Additionally, whether and when a participant died was added as a competing event.

Chronic diseases

The presence of a chronic disease at baseline was based on the database of dispensed medicines in 2006 from Statistics Netherlands, which contains information on purchased drugs that were reimbursed by the health care insurances. Dispensed medicines are classified into anatomic, therapeutic, chemical (ATC) classification codes, according to the WHO (World Health Organization) drug classification system of 2010 (16). Previous research has used these ATC codes to identify specific chronic diseases and estimate the prevalence among the working population (17, 18). Following these definitions, six chronic diseases were identified: cardiovascular diseases,

respiratory diseases, inflammatory diseases, diabetes, mental disorders (depressive, anxiety and sleep disorders) and psychotic disorders. From this, workers were categorized into having 0, 1 or >2 chronic diseases.

Income level

Income level was based on the database of annual personal gross income of Statistics Netherlands. Since retirement decisions rely on personal income and savings, we use personal gross income at the age of 54 as a measure of economic wellbeing. Hence, annual personal gross income was measured five years before the reform and, thus, the reform is less likely to have influenced income: 2003 for those born in 1949 (control group) and 2004 for those born in 1950 (intervention group). Income was adjusted for inflation between both birth cohorts and then divided into four categories: <€25 000, €25 000-40 000, €40 000-55 000, and >€55 000.

Statistical analyses

To evaluate the effect of disincentivizing early retirement a RD approach with a RMST design was used to estimate the causal effect of penalizing early retirement on labor force participation between 60-65 years (19). Under this approach, two populations are compared where the assignment to treatment depends only on whether an individual observed assignment variable (ie, running variable) exceeds a cut-off point. In the current study, the intervention group - those who lose the tax incentives for early retirement - was defined as those who are born just on or after 1 January 1950. The benefits of using an RD approach is threefold. First, one can use the observed data to verify whether the assumption of exogeneity is likely to hold. This is explained in the following paragraph. Second, RD takes advantage of a natural experiment setting whereby the variation isolated by the cut-off can be considered as good as random due to the inability of individuals in the study to influence the treatment allocation (20). Finally, it represents a transparent and understandable approach to estimate the causal effects of a policy measure.

The intuition behind this RD approach is that the treatment is exogenous in the vicinity around the threshold and, thus, offers an estimation of the causal effect of the policy. Two assumptions are important to verify. First, individuals should not be able to manipulate the assignment variable to treatment (8), which is the case in our study by using date of birth. The second condition is that individuals on either side of the age threshold introduced by the reform are statistically comparable (8). To verify this, we show graphically that individuals around the threshold are statistically the same while account for potential secular trends. The combination of both conditions entails that for individuals in the vicinity of the threshold there is only variation in the treatment status that is uncorrelated with any baseline characteristics.

The statistical analyses consisted of three steps. In the first step, monthly employment records from age 60 until 65 were used to calculate the observed number of months (up to 60 months) that each worker spent in paid employment before exiting the labor force. Then for each individual, we identified the specific exit route for that individual (ie, disability benefits, unemployment benefits, early retirement, economic inactivity or death). The monthly employment records were also used to identify censored individuals. Workers were censored at 60 months if they were still working at the age of 65 (ie, statutory retirement age) or at their latest observation in paid employment when they (i) left paid employment through any of the pathways other than those of interest, (ii) became self-employed, or (iii) were missing (eg, due to emigration). A descriptive figure of the Kaplan-Meier cumulative incidence curves for each specific exit route is constructed based on this data.

In the second step, the monthly employment records were used to estimate the RMST in number of months for each specific exit pathway, which was the outcome in our RD equations (21). In practice, this was done using the approach by Andersen et al (22) where the causespecific RMST is calculated with a "leave-one-out" technique in combination with Kaplan-Meier survival curves starting at the age of 60 and until the age of 65. This approach enables us to account for right-hand censoring in each of the exit pathways. The advantages of this RMST approach are the ease in interpretation of the coefficients (ie, actual months) and the possibility to include the RMST estimates in other statistical models.

In the final step, the estimated RMST for each individual was included as the outcome in a RD analysis by comparing the differences in the mean employment time and time in the different pathways between the intervention group and control group for each exit route separately, while accounting for secular trends around the threshold. All

RD models were controlled for covariates (ie, gender, income level, chronic disease). For a full description of the RD model, refer to the supplementary material (www.sjweh.fi/show_abstract. php?abstract_id=3946, mathematical model). As death was not considered a primary outcome, the effect of the reform on death is not further presented in this paper. To evaluate heterogeneity, we estimated the equation above stratified by gender, income level, and the presence of a chronic disease. To verify the robustness of the results, sensitivity analyses were conducted using quadratic functions in both sides of the threshold. In the main analysis, a bandwidth of six months around the threshold was used, and bandwidths of three and twelve months were used in the sensitivity analyses. The trade-off is such that, as the bandwidth is increased, it is likely that individuals in the treatment and control group are less balanced, thus, reducing the internal validity. However, with a smaller bandwidth, the sample size decreases and statistical power will be reduced.

Statistical analyses were performed in RStudio statistical software, version 1.1.463, using Rdrobust package for the RD estimation.

Results

Table 1 presents the characteristics of the total study population (N=102 617), 49 501 workers (61.2% male) constituted the control group, while 53 116 workers (61.1% male) composed the intervention group. Approximately half of the respondents had >1 chronic disease (49.9% and 49.7% in the control and intervention groups, respectively). Baseline characteristics, as required for the RD approach, were similar between the intervention and control groups. Supplementary figure S1 shows the graphical balance in baseline characteristics for those in the treatment and control group. When stratified for gender, income level or chronic disease, slightly more female workers in the intervention group were in the €40 000-55 000 income bracket compared to the control group [ß 0.09 (95% CI 0.01-0.17)].

Figure 1 shows a lower probability to exit from paid employment through early retirement in the intervention compared to control group, while the probability to exit from paid employment through other pathways was slightly higher for the intervention compared to control group (absolute probabilities for each pathway are presented in supplementary table S1). Almost half of the workers born in 1949 (49.8%) left paid employment through early retirement compared to less than one third of the workers born after the policy threshold (29.4% as shown in table 1). Figure 2 shows the effect heterogeneity of reform eligibility on the proportion of workers who exit paid employment through early retirement by gender, income and health. This figure shows important discontinuities but also differences in magnitudes for each group.

Table 2 shows that working months lost due to early retirement decreased by 7.41 months (95% CI 6.118.72) in the intervention compared to control group, coinciding with prolonged paid employment by 4.87 months (95% CI 3.60-6.24). Furthermore, workers in the intervention group left paid employment through other pathways more often than the control group, but this was not statistically significant for any specific pathway. With regard to gender, male workers in the intervention group postponed early retirement by 8.50 months (95% CI 6.81-10.23) compared to the control group and replaced this by spending additional months in paid employment (6.32 months (95% CI 3.79-8.84). After the reform, women postponed early retirement by 5.63 months (95% CI 3.65-7.60). They replaced this by more months in unemployment (1.50 months, 95% CI 0.30-2.71) but also by spending additional months in paid employment (2.09 months, 95% CI -0.09-4.29), albeit the latter was not statistically significant.

The reform significantly delayed early retirement for all income levels, ranging from -5.82 months (95% CI -8.20- -3.45) for the lowest income level up to -8.08 months (95% CI -10.86-5.30) for workers annually earning €25 000-40 000 (table 2). Workers from all income levels replaced this largely by additional months in paid employment, with the highest gain for workers with the highest income [6.32 months (95% CI 3.79-8.84)]. Even though workers in all income groups left paid employment earlier through other pathways, this was only significant for workers with an income of €25 000-€40 000, as they lost 1.24 working months (95% CI 0.31-2.18) through economic inactivity.

Additionally, the reform postponed early retirement among workers with one chronic disease by 6.86 months (95% CI 4.53-9.19), while this increased to 9.12 months (95% CI 6.10-12.14) for workers with multiple chronic diseases (table 2). This postponement in early retirement was partly replaced by prolonging paid employment (3.45 months

(95% CI 1.02-5.88) for workers with one chronic disease and 4.99 months (95% CI 1.77-8.20) for workers with multiple chronic diseases). Both, workers with 1 and >2 chronic diseases in the intervention group lost more working months through disability and unemployment benefits than their counterparts in the control group. This was, however, only significant for workers with multimorbidity in the intervention compared to control group, as they lost 1.99 months (95% CI 0.06-3.92) more due to unemployment benefits.

In general, sensitivity analyses conducted with varying bandwidths and with a quadratic polynomial form showed similar results (supplementary tables S2-4).

Discussion

This study shows that the Dutch policy reform was effective in delaying exit from paid employment through early retirement and prolonging paid employment among workers, in line with previous research (6, 23). However, differences were found across groups of workers. Workers with the highest income level and male workers in the intervention group fulfilled this postponement largely by remaining in paid employment compared to their counterparts in the control group. Workers annually earning €25 000-40 000 spent more months in economic inactivity and female workers spent more months in unemployment. Additionally, workers with chronic diseases in the intervention group were only partly able to replace their postponement with prolonging paid employment than workers in the control group, resulting in the fact that workers in the intervention group lost more paid employment months due to disability and unemployment benefits.

The reform aimed to discourage early retirement by disincentivizing rather than forcing workers to prolong their working careers. This was reflected by the fact that both women and men postponed early retirement, by 5.63 and 8.51 months respectively, which is below the 13 months required to receive the same amount of benefits as people born before 1950. In line with a previous study (6), female workers born from 1950 onwards replaced the postponement of early retirement by spending more months in unemployment and - to a lesser extent - economic inactivity, while corresponding male workers spending more months in paid employment. This difference in effect on the reform between gender needs to be explained by the different labor market attachment across genders in The Netherlands. Females mainly worked part-time, which was also reflected by the fact that almost 80% of the female workers were in the lowest two income levels. They have probably a lower need to prolong their working careers than men.

With regard to income level, the largest impact of the reform was found among workers with the highest income who replaced early retirement with paid employment, while workers with lower incomes exited paid employment more often through other pathways. The latter is undesirable because receiving disability benefits or becoming unemployed might have negative health consequences for individuals (24). In light of interpreting the effect of the reform, it needs to be highlighted that workers with a higher income level - and probably more savings - could also counter the economic consequences induced by disincentivizing early retirement more easily (7, 25). The Health Council of The Netherlands has recommended the government to monitor these socioeconomic (health) differences with respective to prolonging working careers (26).

Furthermore, workers with 1 or >2 chronic diseases postponed early retirement by 6.86 and 9.12 months, respectively. This is in line with a study of Pedersen et al (27) that showed a decline in disability pensions after a structural reform of the Danish Disability Pension Act. The interpretation of our effect is twofold. Previous studies showed that workers with multimorbidity (28) or poor health (29) had an increased risk to leave paid employment through early retirement, which could be seen as a health protection mechanism. When the reform discouraged exiting paid employment through early retirement, these workers left the labor force earlier through involuntary pathways (ie, disability benefits and unemployment benefits). Thus, early retirement and these social security pathways act as communicating vessels (13). On the other hand, workers with chronic diseases partly compensated the postponement of early retirement by prolonging their working lives. However, keeping workers with a chronic disease in paid employment longer could have negative consequences for their productivity and long-term health (30).

As many countries are currently discussing reforms in their pensions systems that penalize early retirement, policy-

makers could benefit from the conclusions drawn in the current study. In particular, our results show that discouraging early retirement does lead to longer working lives. However, the study also showed that reforming the pension system could lead to effects in other social programs, emphasizing the interconnection among social benefits, which need to be paid by the society. Additionally, although such reforms can indeed delay retirement, the current study offers a word of caution for policy-makers. They need to be aware of the societal and individual side effects of these kind of reforms across different groups when implementing reforms prolonging working lives.

This study demonstrates that the RD design combined with survival data is a valuable approach to evaluate interventions with a strictly defined cut-off date. Taking the assumptions into account, the RD design can resemble a randomized controlled trial and estimate causal effects (31). As these methods are relatively new in the field of occupational health, researchers need to be encouraged to report the study quality and to promote transparency on the validity of the results by reporting five key elements (19): (i) a discussion of the RD validation conditions within the specific context, (ii) a clear presentation of the assignment rule, (iii) a test of the distribution of baseline covariates between the intervention and control group, (iv) a graph of the assignment variable to show no bunching of the data around the cut-off, and (v) several sensitivity analyses with varying bandwidth and function forms to check for robustness.

Major strengths of the current study include its large study population and administrative data on all variables. This enabled not only the in-depth evaluation of the effect for specific demographic subgroups of the population but also the ability to distinguish workers with chronic conditions from those free of disease. The current study, however, is not without limitations. First, as applying a RD design requires more statistical power, it was not possible to further stratify, for instance for both gender and income level. Second, the RD design estimates local average intervention effects and the results cannot be extended simply to other generations, eg, workers who were <50 years at the time of the reform. Third, we used register data on prescribed medication use to identify the five most common diseases. The register did not include over the counter medication such as painkillers that may be used for particular chronic diseases. Also, a chronic disease that requires treatment without medication was not included. Hence, we were unable to identify prevalent chronic diseases such as musculoskeletal disorders or less severe psychological health problems (eg, burnout) that are related to premature exit out of paid employment (32). Likewise, less common diseases, such as cancer, were also not included because the prevalence in the workforce was too low.

To conclude, the reform to discourage early retirement was effective in prolonging working lives. However, differences were found across groups of workers. Workers with the highest income and male workers largely replaced the postponement of early retirement by prolonging their working lives. Workers with a middle income lost more paid employment months through disability benefits, whereas female workers lost more paid employment months due to unemployment benefits. Workers reporting 1 or >2 chronic diseases postponed their early retirement but were unable to fulfill this entirely with paid employment.

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Sidebar

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The burnout enigma solved?

Schaufeli, Wilmar, PhD

ABSTRACT (ENGLISH)

The international expert panel chose a pragmatic way of cutting the Gordian burnout knot by looking for the common denominator in the definitions used so far. Not surprising--and in fact quite predictably--work-related (emotional and physical) exhaustion appeared on the scene. Here, Schaufeli examines why the definition of burnout of the panel is problematic.

FULL TEXT

In the current issue of the Scandinavian Journal of Work, Environment and Health, an international expert panel presented a consensual definition on burnout that reads as follows: "In a worker, occupational burnout or occupational physical AND emotional exhaustion state is an exhaustion due to prolonged exposure to work-related problems." (1, p95) Does this definition solve the burnout enigma? Obviously, the quest for an agreed-upon definition has haunted burnout research from the moment the concept was introduced by Freudenberger in the mid 1970s (2). Needless to say, the absence of consensus on what burnout 'is' has hampered the accumulation of our knowledge on its epidemiology, antecedents, and consequences and interventions to reduce it (3).

But is it really true that scholars completely disagree about burnout - as the panel assumes? The answer to this question is affirmative as well as negative, just like quantum particles that can be charged both positively and negatively at the same time. On the one hand, virtually all scholars agree that exhaustion is the core, constituting component of burnout. Tellingly, exhaustion is mentioned in 12 of the 13 definitions that are listed in Table 1 of the panel's paper. Also, the Maslach Burnout Inventory (MBI), which prominently includes an exhaustion subscale, is used almost universally in about 90% of all studies on burnout (4). On the other hand, it is not clear what kind of exhaustion is supposed to be typical for burnout: emotional, physical, mental, or cognitive exhaustion? Or all of these? And how do they differ, for that matter? To the best of my knowledge the nature and the different kinds of exhaustion have not been discussed systematically in the burnout literature so far. In addition, although most scholars would agree that exhaustion is a necessary condition for burnout, they disagree on whether it is a sufficient condition as well. For instance, for some, burnout is equivalent to exhaustion, and they argue that the two other dimensions that are included in the MBI are unnecessary as (i) mental distancing (ie, cynicism or depersonalization) refers to a way of coping and (ii) reduced personal accomplishment is a consequence of burnout (5). In contrast, for others such as myself, exhaustion is a necessary yet insufficient component for burnout as will be explained below.

The panel chose a pragmatic way of cutting the Gordian burnout knot by looking for the common denominator in the definitions used so far. Not surprising - and in fact quite predictably - work-related (emotional and physical) exhaustion appeared on the scene. In my opinion, the definition of burnout of the panel is problematic for three reasons. First, the nature of work is not specified. It seems that the definition is restricted to formal employment relationships as the exhaustion is termed as 'occupational'. That would mean that - quite incongruously - the first burnout cases Freudenberger described, who happened to be volunteers working with drug addicts, would not have suffered from the syndrome. And what about students, athletes and parents, for instance? Can they burn out? According to the panel's definition they cannot, despite numerous papers on the subject. Or does the adjective 'occupational' in the definition imply that there are other non-occupational forms of burnout as well? Be as it may, should we abandon the work-related nature of burnout altogether? Not quite, if we understand work psychologically and not in terms of labor relations - that is when all structured, goal-oriented activities that are compulsory in nature and aim to transform the physical or social environment and/or the self are considered 'work'. Seen from that psychological perspective, volunteers, students, athletes and parents do 'work' because they are engaged in such activities.

Second, in the definition burnout is equated with a combination of emotional and physical exhaustion: but what

about the other kinds of exhaustion? An what exactly is emotional exhaustion? Is it different from mental exhaustion? Does the "... exhaustion due to prolonged exposure to work-related problems" include cognitive exhaustion as well, which manifests itself in typical burnout complaints such as lack of concentration and inability to think clearly? Unfortunately, these questions remain unanswered. In short, the exact nature of exhaustion and the domain in which it reveals itself remains unclear.

Last but not least, because of the pragmatic approach of the panel, their burnout definition lacks a proper theoretical underpinning. Essentially, exhaustion is a state of being extremely fatigued. A long tradition exists on work-related fatigue beginning with the seminal work of Edward Thorndike (1874-1949), who maintained that the basic tenet of fatigue is "the intolerance of any effort" (6, p104). In his view, fatigue is both the inability and the unwillingness to spend effort at work, which is reflected by its energetic and motivational component, respectively. The unwillingness to perform is expressed in increased resistance, reduced commitment, cynicism, lack of interest, disengagement, and so on - in short, mental distancing. Building on the work of Thorndike, Schaufeli &Taris (7) have argued that inability (exhaustion) and unwillingness (distancing) constitute two sides of the same burnout coin. Hence, based on a theoretical analysis mental distancing should be recognized in addition to exhaustion as the second constituting dimension of burnout (8). This meshes with the results of a recent review of 12 burnout questionnaires, 8 of which assessed burnout as a multidimensional concept; all of which included an exhaustion as well as a mental distance subscale (the remaining 3 questionnaires reduced burnout to mere exhaustion) (8).

So, taken together,it seems that - notwithstanding the panel's laudable endeavor - the burnout enigma has not yet been solved. This means that, despite the agreement about exhaustion as a key element of burnout, the discussion about the definition of burnout continues.

Sidebar

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The influence of chronic diseases and multimorbidity on entering paid employment among unemployed persons - a longitudinal register-based study

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ABSTRACT (ENGLISH)

Objectives This study aimed to investigate the influence of chronic diseases and multimorbidity on entering paid employment among unemployed persons. A secondary objective was to estimate the proportion of persons not entering paid employment that can be attributed to specific chronic diseases across different age groups. Methods Data linkage of longitudinal nationwide registries on employment status, medication use and sociodemographic characteristics was applied. Unemployed Dutch persons (N=619 968) were selected for a three-year prospective study. Cox proportional hazards analyses with hazard ratios (HR) were used to investigate the influence of six common chronic diseases on entering paid employment, stratified by age. The population attributable fraction (PAF) was calculated as the proportion of all persons who did not enter paid employment that can be attributed to a chronic disease. Results Persons with chronic diseases were less likely to enter paid employment among all age groups. The impact of a chronic disease on maintaining unemployment at population level was largest for common mental disorders (PAF 0.20), due to a high prevalence of common mental disorders (6%), and for psychotic disorders (PAF 0.19), due to a high likelihood of not entering paid employment (HR 0.21), among persons aged 45-55 years. Multimorbidity increased with age, and the impact of having multiple chronic diseases on remaining unemployed increased especially among persons aged ≥45 years. Conclusion Chronic diseases and multimorbidity are important factors that reduce employment chances among all age groups. Our results provide directions for policy measures to target specific age and disease groups of unemployed persons in order to improve employment opportunities.

FULL TEXT

Headnote

Objectives This study aimed to investigate the influence of chronic diseases and multimorbidity on entering paid employment among unemployed persons. A secondary objective was to estimate the proportion of persons not entering paid employment that can be attributed to specific chronic diseases across different age groups.
Methods Data linkage of longitudinal nationwide registries on employment status, medication use and sociodemographic characteristics was applied. Unemployed Dutch persons (N=619 968) were selected for a three-year prospective study. Cox proportional hazards analyses with hazard ratios (HR) were used to investigate the influence of six common chronic diseases on entering paid employment, stratified by age. The population attributable fraction (PAF) was calculated as the proportion of all persons who did not enter paid employment that can be attributed to a chronic disease.
Results Persons with chronic diseases were less likely to enter paid employment among all age groups. The impact of a chronic disease on maintaining unemployment at population level was largest for common mental disorders (PAF 0.20), due to a high prevalence of common mental disorders (6%), and for psychotic disorders (PAF 0.19), due to a high likelihood of not entering paid employment (HR 0.21), among persons aged 45-55 years. Multimorbidity increased with age, and the impact of having multiple chronic diseases on remaining unemployed increased

especially among persons aged ≥45 years.
Conclusion Chronic diseases and multimorbidity are important factors that reduce employment chances among all age groups. Our results provide directions for policy measures to target specific age and disease groups of unemployed persons in order to improve employment opportunities.
Key terms PAF; population attributable fraction; register data; medicine; unemployment.
Employment status is an important determinant of health, as demonstrated by a lower mental and physical health status among unemployed persons (1). These health inequalities can be explained by the causation and selection hypothesis. The causation hypothesis states that unemployment causes persons to deteriorate in health, whereas employment causes persons to improve in health. The selection hypothesis states that persons with poor health have a higher risk of leaving paid employment, as well as remaining unemployed for a longer period (2, 3).
Many studies have shown that chronic diseases, such as psychological disorders, cardiovascular diseases and diabetes, are typically more common among unemployed individuals (4-7) and are associated with a higher risk of being unemployed (8). However, most studies address loss of paid employment in the workforce rather than health as a barrier for (re)entering paid employment (9). Many qualitative studies have shown that entering paid employment is often challenging for persons with mental health problems (10, 11). Some quantitative studies also have shown that persons with psychological or physical health conditions have a lower likelihood of entering paid employment (12-15). Less is known about the influence of having a specific chronic disease or multiple chronic diseases on entering paid employment. Especially those who have multiple chronic diseases seem to be less likely to enter paid employment. In fact, multimorbidity - the occurrence of at least two chronic diseases within an individual - has been associated with a higher risk of unemployment (16). One study showed that unemployed individuals with two health conditions had a 24% decreased likelihood of entering paid employment compared to those with no health conditions (14).
So far, most studies on associations between chronic diseases and entering paid employment have mainly focused on self-reported health outcomes and some have focused on certain health aspects such as mental health only (6, 17). However, self-reported health outcomes are prone to reporting bias and justification bias. This often occurs when unemployed individuals over-report their level of disability or work limitations to justify that they are not in paid employment (18). Therefore, in order to minimize bias, it is important to use more objective data. One way to objectively investigate the presence of chronic diseases is by using pharmacy data (19). Pharmacy data on dispensed drugs provide a reliable information source, often covering a broad range of health aspects in the general population (20). Linkage of objective pharmacy data with employment registers can create large study populations that facilitate precise estimations of associations between health and employment.
Only few studies have used pharmacy data as a health indicator in relation to entering employment. One Danish study found that prescription medicine purchases for both physical illnesses (including cardiovascular diseases, chronic lung diseases and type 2 diabetes) and mental illnesses had a negative influence on the likelihood of entering paid employment (21). Another study showed that prescribed medication purchases for mental illnesses significantly reduced employment chances in a random population sample (22). However, these studies have focused on single chronic diseases and their associations with entering paid employment rather than the impact of an array of (multiple) chronic diseases on entering paid employment at population level. In addition, there is little insight in the age-specific influence of chronic diseases on entering paid employment. The sheer size of the present study facilitates to study multimorbidity and age-specific patterns in the impact of chronic diseases on employment, which is a unique feature of this study. This study does also not suffer from any selection bias into the study population as the total Dutch population of unemployed persons was included.
Age is closely related to the presence of chronic diseases as well as the likelihood of entering paid employment (14, 23, 24). In general, older persons often encounter more barriers in the labor market than younger persons, resulting in a lower likelihood of entering paid employment (25, 26). Chronic diseases may impact individuals differently according to the phase in their working life course. For instance, younger persons are more likely to aim for a successful career compared to older persons who are almost at the end their careers (27). Therefore, diseases may

have a larger impact on employment among younger compared to older persons.
However, little is known about how age and chronic diseases may interact to create or increase barriers to enter paid employment. Therefore, the influence of chronic diseases on entering paid employment should ideally be explored among different age groups. It is of particular interest to know how the influence of specific chronic diseases on paid employment varies across disease groups and whether the presence of multiple chronic diseases has synergistic or independent effects. This is especially of high importance for interventions targeting health and improving employment opportunities for unemployed persons.
Most longitudinal studies have focused on associations of chronic diseases with entering paid employment, interpreting these associations at individual level, ie, the reduced risk for an individual person with a chronic disease to enter or maintain paid employment. This information can only be generalized to the total population, when the study population is a random sample of the total population. A population-based approach offers the opportunity to estimate the contribution of chronic diseases to being without paid employment in the population. In this sense, the population attributable fraction (PAF) is the proportion of persons without paid employment that can be attributed to the presence of chronic diseases. The PAF is an integrated measure that takes into account both the strength of association between the risk factor and the outcome as well as the prevalence of the risk factor in the population. This measure reflects the theoretical reduction in unemployment in the population that would occur if the prevalence of chronic diseases were reduced to zero, or chronic diseases no longer had any effect on likelihood of paid employment.
In the present study, nationwide data on employment status, medication, and sociodemographic characteristics were used to conduct a longitudinal study with a threeyear follow-up. The primary aim of the study was to investigate the influence of chronic diseases and multimorbidity on entering paid among unemployed persons. The secondary aim was to estimate the proportion of persons not entering paid employment that can be attributed to specific chronic diseases across different age groups.

Methods

Study design and population

A longitudinal study in The Netherlands with a three-year follow-up was conducted on register data covering information on residents between January 2015 and December 2017. Statistics Netherlands provided secure access to individual-level databases on demographics, educational level, labor market status, and dispensed medication. All Dutch residents were pseudonymized using a personal unique number. Data registries were linked at the individual level using these pseudonymized numbers. In the present study, persons who received social security benefits or unemployment benefits on 1 January 2015 (baseline), were not employed for at least the previous three months, and were aged 18-65 years (N=625 691) were selected. No informed consent was needed to conduct this study as Dutch legislation allows authorized research institutes to use pseudonymized register-based data for research purposes. To report the findings of this study, the STROBE guidelines were followed.

Employment status

The database on socioeconomic position per month (SECM) provided the main source of income for each consecutive month between 2014 and 2017. Persons who received a social security or unemployment benefit were defined as unemployed. Persons who were employees or self-employed were classified as employed. In this longitudinal study, the transition from unemployment into paid employment was investigated.

Chronic diseases and multimorbidity

The database on medication (Medicijntab) provided information on purchased drugs that were reimbursed by health insurance companies. Since no data was available on the precise dates of the drugs being purchased during the year, medication use in 2014 was used to identify the presence of chronic diseases among unemployed persons (19), based on the World Health Organization Anatomical Therapeutic Chemical (ATC) classification codes (28). For instance, common mental disorders were identified by ATC codes assigned to drugs such as antidepressants and anxiolytics, whereas respiratory illness was identified by the ATC code assigned to inhalants and other drugs for airway diseases. The presence of a specific chronic disease was dichotomized into having or not having a chronic

disease. The most prevalent chronic diseases investigated in this study were: inflammatory conditions, common mental disorders, psychotic disorders, cardiovascular diseases, diabetes, and respiratory illness (supplementary material, www.sjweh.fi/ show_abstract.php?abstract_id=3942, table S4).

Multimorbidity was investigated as the number of chronic diseases. The total number of chronic diseases was computed for each participant, based on the 21 different chronic diseases identified by medication. This measure of multimorbidity was then categorized into four groups: 0, 1, 2, >3 chronic diseases. Having >2 chronic diseases was defined as multimorbidity.

Sociodemographic variables

The databases of Statistics Netherlands on sociodemographic characteristics provided information on age, sex, education, and migration background in 2014. Educational level was categorized into three educational groups: high (higher vocational training or university), intermediate (higher secondary and intermediate vocational training) and low education (pre-primary education, primary education, and lower secondary education). Age was categorized into four age groups (18-30, 30-45, 45-55, 55-65 years). Migration background was categorized into six groups: native Dutch, Moroccan, Turkish, Surinamese and Antillean, other Western, other non-Western.

Analyses

Descriptive statistics were used to describe baseline characteristics. Cox proportional hazards analyses were performed to investigate the influence of each chronic disease on entering paid employment. The dependent variable was the first event of entering paid employment for at least three consecutive months during the three-year follow-up period. Persons were censored at the transition from unemployment to disability or (early) pension, after three months of missing data on employment status, and at the end of the follow-up period (December 2017). The same was done to investigate the influence of multimorbidity on entering paid employment. Stratified analyses were performed to estimate associations within age groups. All analyses were adjusted for sex, educational level and migration background.

Since this study was conducted on all unemployed persons in The Netherlands, the discriminatory power is high and traditional testing of statistical significance has little relevance as we do not have a sample but the entire population. Hence, we present confidence intervals (CI) but refrain from formal statistical tests when deemed not informative.

In order to investigate whether the effect estimates were significantly different between the four age groups, an interaction term of chronic disease^age was added to the cox proportional hazards models on the nonstratified population (N=619 968). An interaction term of multimorbidityxage was also included in the Cox proportional hazards analysis.

The PAF can be defined as the proportion of unfavorable outcomes (remaining unemployed) that would have been prevented if the exposure of interest (chronic diseases) was eliminated from the population. The following formula was used: PAF=Pe(HR-1)/(1+Pe(HR-1), where Pe is the prevalence of a chronic disease among a particular age group (29). The hazard ratio (HR) in this formula is the likelihood to remain unemployed, which is the inverse of the HR to enter paid employment. A PAF close to 1 indicates that remaining unemployed is almost completely attributed to the chronic disease. A PAF close to 0 indicates that chronic diseases do not play a role in remaining unemployed. The PAF was also calculated at each age between 25 and 60 years, based on age-specific prevalence and HR. All statistical analyses were performed using SPSS v.22.0 (IBM Corp, Armonk, NY, USA).

Results

Approximately half of the unemployed persons were older than 45 years (45.9%), female (52.2%), lower educated (45.1%), and native Dutch (54.3%). The most prevalent chronic diseases were inflammatory conditions (24.8%), cardiovascular diseases (18.2%), and common mental disorders (17.3%). Almost 60% of the study population had at least one chronic disease. The prevalence of having at least one chronic disease was higher among unemployed persons (59.1%) compared to employed persons (39.8%) in The Netherlands (table 1, and supplementary table S1)

The proportion of persons entering paid employment during the follow-up period was highest for persons aged 18-30 years (46.0%), followed by the age groups 30-45 (39.8%), 45-55 (33.3%), and 55-65 (15.3%) years. The likelihood to enter paid employment reduced with increasing age. The likelihood to enter paid employment also reduced with

increasing number of chronic diseases. The prevalence of multimorbidity was highest among the older age groups, with the highest prevalence of >3 chronic diseases (30.9%). Persons with >3 chronic diseases had the lowest likelihood to enter paid employment, with the lowest likelihood among persons aged 30-45 years (HR 0.43, 95% CI 0.42-0.44). The presence of multiple chronic diseases showed no synergistic effects on entering paid employment, indicating chronic diseases to be largely independent of each other. The fraction of persons staying unemployed due to having at least three chronic diseases increased from 0.04 in the youngest age group to 0.22 in the oldest age group (table 2, supplementary table S2)

Persons with common mental disorders, psychotic disorders, cardiovascular diseases, inflammatory conditions, diabetes and respiratory illness were all less likely to enter paid employment among all age groups. Persons with psychotic disorders had the lowest chances to enter paid employment, followed by persons with common mental disorders. Persons with inflammatory conditions had the highest likelihood to enter paid employment. Except for persons with psychotic disorders, the likelihood to enter paid employment was comparable across all age groups. Among persons with psychotic disorders, those aged >45 years had the lowest chance to enter paid employment. The fraction of remaining unemployed due to having psychotic disorders was highest for those aged 45-55 years (PAF 0.19), due to a stronger association between psychotic disorders and entering paid employment (table 3, supplementary table S3)

For persons with common mental disorders, the PAF increased with age to 0.20 until the age of 50 years, and declined onwards to 0.09. The same pattern was observed for inflammatory conditions. The decline in these PAF was primarily caused by a lower prevalence of common mental disorders and inflammatory conditions from middle age onwards. The PAF for cardiovascular diseases increased from 0 at the age of 18 years until 0.10 at the age of 61 years old. The same pattern was observed for respiratory illnesses (figure 1, supplementary figure S1, table 3)

Discussion

This longitudinal register-based study showed that unemployed persons with chronic diseases were less likely to enter paid employment among all age groups. In particular, persons with psychotic disorders had the lowest likelihood to enter paid employment at all ages, followed by common mental disorders. For psychotic disorders, the impact on maintaining unemployed on population level (PAF 0.20) was primarily due to psychotic disorders being the strongest barrier to enter paid employment. Multimorbidity increased with age, and the impact of having multiple chronic diseases on remaining unemployed increased especially among persons aged >45 years. The PAF for common mental disorders and inflammatory conditions increased with age but declined from middle age onwards, whereas the PAF for cardiovascular disorders and respiratory illness gradually increased with age.

The findings of this study corroborate the findings of other studies focusing on the negative influence of poor health on entering paid employment (15, 17, 30, 31). Our study does not suffer from selection bias as the register contains all unemployed persons in The Netherlands. The size of the study population also allowed a detailed insight into the age-specific impact of chronic diseases and multimorbidity on entering paid employment across working age, expressed as the PAF by year of age. The PAF was highest for common mental disorders (PAF 0.20) and psychotic disorders (PAF 0.19) among persons aged 45-55 years. This implies that respectively 20% and 19% of the unemployed persons would have entered paid employment if common mental disorders respectively psychotic disorders were eliminated among this particular age group. This knowledge is particularly valuable for re-integration policies, since targeting persons with common mental disorders or psychotic disorders aged 45-55 years could result in the highest impact on labor force participation at population level compared to other age and disease groups.

This study also adds new insights into the role of chronic diseases and multimorbidity on entering paid employment across different age groups. The current study showed that among younger persons, the presence of one or two chronic diseases had a stronger effect on entering paid employment compared to older persons with one or two chronic diseases. In general, younger persons are more likely to enter paid employment compared to older persons (25). However, our finding can be explained by differences in the type of chronic diseases between younger and older persons. For instance, this study showed that younger persons predominantly suffered from common mental disorders, whereas older persons more often had long-term age-related diseases such as cardiovascular diseases.

It may be that agerelated diseases such as cardiovascular diseases or diabetes are easier manageable through (modifiable) lifestyle changes and appropriate medication than mental health problems. In this sense, a study by Claussen et al suggested that mental and physical illnesses predict re-employment differently (32, 33).
Many studies investigating the influence of chronic diseases on entering paid employment have used selfreported health measures. Although self-reported diseases may have some advantages, such as being able to include chronic diseases that do not require treatment in healthcare (34, 35), its associated risk of justification bias and reporting bias may be problematic among unemployed persons. To minimize these biases, objective measures of health (eg, register-based data) were used to estimate the effect of chronic conditions on entering paid employment. Moreover, the register contains all unemployed persons in The Netherlands, thus, there is no selection bias. Our results are in line with a Danish register-based study showing that prescription medicine purchases for both physical illnesses (including cardiovascular diseases, chronic lung diseases and type 2 diabetes) and mental illnesses had a negative influence on the likelihood of entering paid employment (21). In accordance with our study, they also found a stronger negative influence (ie, a lower likelihood to enter employment) of mental illnesses on entering paid employment compared to physical health problems such as cardiovascular diseases.
One of the major strengths of this study is the use of objective register-based data. This minimized the risk of reporting and justification bias which is more likely to be present among studies using self-reported health outcomes. In addition, the use of register-based data enabled to include a large study population and thus provided enough statistical power to perform age-specific analyses. Also, use of register-based data allowed a longitudinal design without experiencing challenges such as a low response or loss-to-follow-up, often occurring in other longitudinal studies. Another strength of this study was the calculation of the PAF, which provides insight into the impact of chronic diseases at the population level, combining information on the association of chronic diseases with entering employment and the prevalence of chronic diseases in the population.
The use of register-based data has also limitations. First, registered medication includes merely registrations of individuals who fulfill three criteria: (i) they receive a prescription by their general practitioner or specialist, (ii) they purchase the prescribed medicine at the pharmacy, and (iii) the costs of the medicines are reimbursed by health insurances. It is possible that persons who have certain chronic diseases but are not treated fall outside these criteria. For instance, persons with common mental disorders who are treated with cognitive behavioral therapy will not be covered by registered medication prescriptions. There was no data available on the severity of health conditions. However, health conditions were severe enough for the physician to prescribe medication. A second limitation is that several specific chronic conditions are treated with generic medicines such as paracetamol or other pain-relieving medicines. Therefore, it is not possible to identify specific conditions such as back pain or musculoskeletal disorders.
A third limitation relates to the classification of multimorbidity. The present study operationalized multimorbidity as having at least two different chronic diseases. However, multimorbidity can also be present when an individual suffers from multiple conditions within a particular chronic disease, such as having both rheumatism and eczema as inflammatory conditions. Therefore, it may be that persons with a single chronic disease category may have more conditions and therefore should actually be classified as having multimorbidity. In this study, this limitation may have led to an underestimation of persons with chronic diseases and multimorbidity.
Lastly, the stressful economic and social circumstances of unemployed persons may cause higher utilization of the health care system (36). However, empirical evidence suggests that the higher frequency of visits to a general practitioner among lower educated persons is primarily the result of a higher prevalence of disease rather than earlier care seeking (37). Hence, we do not think that the reported prevalence of medications for chronic diseases, as prescribed by a physician, is influenced by differential care seeking.
The PAF are calculated based on precise estimates of the prevalence of chronic diseases and the association of chronic diseases and entering paid employment using register data on medication and employment status. There are two assumptions underlying a correct interpretation of the PAF (38). First, our longitudinal study design with objective measures of both dependent and independent variables is less sensitive for reversed causality

mechanisms and, thus, supports a causal relationship between chronic diseases at baseline and entering paid employment during follow-up. Second, we have adjusted for factors that are known to have a strong influence on the association between chronic diseases and entering employment. However, some other factors may have influenced the likelihood of entering paid employment, such as language skills and specific qualifications. Unfortunately, it was not possible to capture these factors by the registers.

Several recent studies have shown that entering paid employment is associated with improvement of physical and mental health outcomes (39, 40). The positive health effects of paid employment are argued to be mediated by a range of benefits, such as a higher income, having a social role and purpose, and access to and social networks and social support (41, 42). Therefore, there is a need for re-integration policies and interventions that facilitate employment among persons with chronic diseases, particularly among persons with common mental disorders and psychotic disorders. Unemployed persons with chronic diseases may have a reduced work capacity and may experience limitations in performing work (43). It is important that interventions target not only target disease-related factors, but also personal (eg, increasing empowerment and self-management skills) and environmental (eg, facilitating work accommodations) factors that are critical for entering and sustaining employment (44).

It should also be mentioned that stigma in mental health problems may have an impact on employment opportunities for unemployed persons, and their expectations to enter paid employment (45). For instance, studies have shown that employers believed that persons with mental illnesses lack competences to meet the demands of work, and that working is not healthy for them (46, 47). In order to improve employment opportunities for these individuals, the importance of paid employment for persons with mental health problems should be emphasized among clinicians and employers, as it has been shown that employment can contribute to recovery of mental health problems (39).

In conclusion, this register-based study provides detailed insight in the impact of specific chronic diseases on remaining unemployed among different age groups in the population. The age-specific patterns of PAF were highest for mental disorders, due to a high prevalence of common mental disorders, and psychotic disorders, due to a strong association between this chronic disease and remaining unemployed, especially among persons aged 45-55 years old. The impact of multimorbidity on remaining unemployed increased with age. Our results provide directions for interventions and re-integration policies to target chronic diseases among specific age groups in order to reach highest impact on labor force participation at population level.

Sidebar

Yildiz B, Burdorf A, Schuring M. The influence of chronic diseases and multimorbidity on entering paid employment among unemployed persons - a longitudinal register-based study. Scand J Work Environ Health. 2021;47(3):208-216. doi:10.5271/ sjweh.3942

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DETAILS

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Changes in precarious employment in the United States: A longitudinal analysis

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ABSTRACT (ENGLISH)

Objective This longitudinal study aimed to measure precarious employment in the US using a multidimensional indicator. Methods We used data from the National Longitudinal Survey of Youth (1988-2016) and the Occupational Information Network database to create a longitudinal precarious employment score (PES) among 7568 employed individuals over 18 waves (N=101 290 observations). We identified 13 survey indicators to operationalize 7 dimensions of precarious employment, which we included in our PES (range: 0-7, with 7 indicating the most precarious): material rewards, working-time arrangements, stability, workers' rights, collective organization, interpersonal relations, and training. Using generalized estimating equations, we estimated the mean PES and changes over time in the PES overall and by race/ethnicity, gender, education, income, and region. Results On average, the PES was 3.17 [standard deviation (SD) 1.19], and was higher among women (3.34, SD 1.20), people of color (Hispanics: 3.24, SD 1.23; non-Hispanic Blacks: 3.31, SD 1.23), those with less education (primary: 3.99, SD 1.07; high school: 3.43, SD 1.19), and with lower-incomes (3.84, SD 1.08), and those residing in the South (3.23, SD 1.17). From 1988 to 2016, the PES increased by 9% on average [0.29 points; 95% confidence interval (CI) 0.26-

0.31]. While precarious employment increased over time across all subgroups, the increase was largest among males (0.35 points; 95% CI 0.33-0.39), higher-income (0.39 points; 95% CI 0.36-0.42) and college-educated (0.37 points; 95% CI 0.33-0.41) individuals. Conclusions Long-term decreases in employment quality are widespread in the US. Women and those from racialized and less-educated populations remain disproportionately precariously employed; however, we observed large increases among men, college graduates and higher-income individuals.

FULL TEXT

Headnote

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Methods We used data from the National Longitudinal Survey of Youth (1988-2016) and the Occupational Information Network database to create a longitudinal precarious employment score (PES) among 7568 employed individuals over 18 waves (N=101 290 observations). We identified 13 survey indicators to operationalize 7 dimensions of precarious employment, which we included in our PES (range: 0-7, with 7 indicating the most precarious): material rewards, working-time arrangements, stability, workers' rights, collective organization, interpersonal relations, and training. Using generalized estimating equations, we estimated the mean PES and changes over time in the PES overall and by race/ethnicity, gender, education, income, and region.

Results On average, the PES was 3.17 [standard deviation (SD) 1.19], and was higher among women (3.34, SD 1.20), people of color (Hispanics: 3.24, SD 1.23; non-Hispanic Blacks: 3.31, SD 1.23), those with less education (primary: 3.99, SD 1.07; high school: 3.43, SD 1.19), and with lower-incomes (3.84, SD 1.08), and those residing in the South (3.23, SD 1.17). From 1988 to 2016, the PES increased by 9% on average [0.29 points; 95% confidence interval (CI) 0.26-0.31]. While precarious employment increased over time across all subgroups, the increase was largest among males (0.35 points; 95% CI 0.33-0.39), higher-income (0.39 points; 95% CI 0.36-0.42) and college-educated (0.37 points; 95% CI 0.33-0.41) individuals.

Conclusions Long-term decreases in employment quality are widespread in the US. Women and those from racialized and less-educated populations remain disproportionately precariously employed; however, we observed large increases among men, college graduates and higher-income individuals.

Key terms employment condition; employment quality; health disparity; non-standard work; precarious work; precarity.

Employment conditions in the US have drastically changed since the 1970s, due to changes in class relations, social policies, recessions, globalization, and technology (1, 2). The number of workers in highquality, full-time employment, with adequate wages and benefits, has decreased over the last 40 years, while the number with low-quality, precarious employment (PE) has increased (2-7). Such trends have implications for myriad social and economic challenges facing society, including growing wealth and health inequities (7, 8). Researchers agree that multidimensional measures of PE can best capture trends in employment conditions (8-12). Yet, prior US estimates are unidimensional, cross-sectional, and/or rely on aggregated data, the latter of which may mask individual-level trends (2, 3, 13). Inadequate evidence about longitudinal, multidimensional trends in US employment conditions hinders our understanding of the causes and consequences of PE, as well as the development of effective policy interventions.

PE is best conceived as the accumulation of multiple unfavorable facets of employment quality (10). Although there is no single definition of PE, we build on a number of recent studies that have identified seven key dimensions of employment quality that capture PE's multidimensionality: material rewards, workingtime arrangements, employment stability, workers' rights, collective organization, interpersonal relations, and training opportunities (10, 14-16). These dimensions represent the most widely used multidimensional operationalization of PE, drawn from a thorough review of the literature and expert consultations via the Employment Conditions Network (EMCONET) advisory group. Importantly, this conceptualization of PE captures both the contractual (eg, wages) and relational (eg, power relations) aspects of employment that play a critical role in worker health and well-being.

These seven dimensions are described in detail by Julia and colleagues (10). Briefly, material rewards denotes the wage and non-wage benefits afforded by an employment arrangement. Workingtime arrangements denotes the length and intensity of working hours, underemployment, and schedule predictability. Employment stability captures continuity of employment, contractual temporariness and/or organizational changes (eg, downsizing), whereas the workers' rights dimension denotes welfare state provisions associated with employment, such as access to health insurance or pensions. Collective organization refers to the possibilities (or lack thereof) for employee representation, most commonly operationalized though union representation. Interpersonal relations refers to the power of employees relative to management (eg, their ability to make decisions or control their schedule), and can include exposure to discrimination. Finally, training opportunities denotes workers' opportunities for promotion or to enhance their skills. Prior studies employing these seven dimensions suggest that up to half of the European workforce experiences some precarity, including those in permanent full-time employment (15-22). However, most studies have employed cross-sectional data. Moreover, their estimates may not generalize to the US, given differences in labor markets and strength of social safety-nets (23).

Much of the empirical work in the US has narrowly focused on contract type alone (3). This approach is limited because it fails to capture the holistic experience of workers and the interrelation of dimensions, such as hours and wages (10). Notable exceptions include three US-based studies that have used multidimensional indicators. Peckham and colleagues employed a latent class analysis (LCA) approach to investigate the construct of employment quality using cross-sectional General Social Survey (GSS) data (24). Using similar indicators and the aforementioned seven dimensions, the authors identified eight employment quality types (16); five of these, constituting 58% of the sample, were deemed lowerquality or precarious (24). Cho and colleagues also applied an LCA approach to GSS data, although their PE indicators were subjective (eg, whether respondents described their benefits as "good") rather than objective (eg, types of benefits respondents had access to) (25). These authors identified four employment types, two of which (40% of the sample), were deemed precarious. Finally, Eisenberg-Guyot and colleagues utilized the Panel Study of Income Dynamics and a sequenceanalysis approach to identify mid-career employment quality trajectories based on dimensions similar to those employed in European-based analyses (26). They found heterogeneity between men and women with 46% of women and 20% of men in precarious or lower-quality employment trajectories (26). These studies improved upon prior US studies that employed unidimensional measures of precarity; however, they did not describe trends over time. Although Kalleberg's seminal study on PE in the US did document time trends, he used national data to analyze separate trends for each indicator of PE, rather than analyzing individual-level data using a single multidimensional indicator (2).

Characterizing trends in PE using a multidimensional indicator is critical given that employment quality is increasingly recognized as a social determinant of health (8, 27). Scholars hypothesize three main mechanisms through which PE may be related to adverse physical and mental health (8). First, PE can cause deprivation (eg, insufficient income) and compromise access to necessities. Second, precariously employed individuals may have greater exposure to adverse physical (eg, toxic exposures) and psychosocial (eg, low control) working conditions. Third, precariously employed individuals may have limited control over both their professional and personal lives, leading to psychosocial stress. Importantly, poor employment quality may be contributing to widening health inequities, as women, people with lower education levels, and those from racialized groups tend to have a higher prevalence of PE (8, 13, 15, 17, 25-28).

To our knowledge, no prior studies have used a multidimensional indicator to estimate longitudinal trends of PE in the US. We addressed this gap using data from the National Longitudinal Study of Youth (NLSY) (1988-2016) and the Occupational Information Network (O·NET) database. The objectives of this study were to: (i) create a multidimensional continuous measure of PE in the US and (ii) describe changes in precarious employment over time, both overall and within subgroups. A better understanding of longitudinal trends is a critical first step for informing future policies aimed at improving PE and population health in the US.

Methods

Data sources

We used data from the 1988-2016 NLSY, which began collecting data in 1979 among a sample of 12 686 men and women (aged 14-21 years). The cohort's profile and survey's sampling methodology have been described elsewhere (29, 30). Briefly, the sample was designed to represent the civilian population of the US in 1979. In addition, this cohort oversampled Black, Hispanic and economically disadvantaged people. Individuals were interviewed annually through 1994 and biennially thereafter. About 55% (N=6912) of the original sample was retained in 2016. Respondents were consistently asked labor-market-related and demographic questions. To ascertain employment characteristics beyond those captured in NLSY, we linked individual-level NLSY data to occupation-level data from the O·NET database (2019 version 24.2). O·NET contains information on occupation-level characteristics for approximately 1000 occupations (31).

The Bureau of Labor Statistics permitted the use of the NLSY79 restricted-use data, which contains state-level geographic identifiers, and the University of Illinois Chicago Institutional Review Board deemed this secondary data analysis non-human-subjects research.

Sample

Our original data set contained 12 686 individuals followed for an average of 8 waves between 1979 and 2016. We first restricted our analyses to 11 052 individuals who participated in at least two waves between 1988 (aged 23-31 years) and 2016 (aged 51-59 years). Then, to align with prior literature, we restricted to 10 281 individuals who were employed at least once (13, 16, 17, 20, 24, 25, 28). We further excluded individuals who had <40% of the survey items observed across their employed years combined (N=1730); this exclusion included selfemployed individuals, who lacked data on many PE indicators. Finally, we excluded individuals with incomplete information on key demographic characteristics (eg, gender, race/ethnicity, education) (N=983). The final analytic sample contained 101 290 employed observations of 7568 individuals over an 18-wave or ~30-year period (7 waves annually from 1988-1994 and 11 waves biennially 1996-2016). Each individual contributed 13-waves of data, on average (interquartile range 7-17).

Precarious employment score construction

We based our scale off European research conceptualizing PE as a multidimensional construct of accumulated adverse employment conditions and measured seven dimensions of the employment relationship, as described above: material rewards, working time arrangements, employment stability, workers' rights and social protections, collective organization or empowerment, interpersonal relations, and training and employability opportunities (10-12).

Variable selection and inclusion

Guided by prior studies on the topic (10, 15), alignment with the seven PE dimensions and data availability, we identified 13 survey indicators to operationalize our PE score (table 1). The promotion indicator was not available in 1989-1994; to partially address missingness, we carried the 1988 values (including missing values) forward to 1989-1993 and the 1996 values (including missing values) backward to 1994. In addition, we used single multivariate imputation to address missingness in the workers' rights (9% missing), empowerment (10% missing), and opportunities (10% missing) dimensions using age, gender, race/ethnicity, marital status, region, education, occupation, industry, and auxiliary items within the same dimension (eg, using health insurance to impute for retirement plan). Combined, all other dimensions had less than 4% missingness.

Scoring

Consistent with previous research, we had no a priori rationale for upweighting any of the seven dimensions (13, 15). Therefore, each dimension was worth a maximum of 1 point, regardless of the number of variables in the dimension. For example, each of the three variables in the working time arrangements dimension was worth 0.33 points. To create the total PE score (PES) for each respondent in each year, we summed their scores for each of the dimensions, yielding a maximum score of 7, with 7 indicating the most precarious (range 0-7). We then adjusted the PES to age 30, using a linear regression of PES on age, age-squared, and year dummies, so that aging within our sample did not drive the observed time trends. To assess the prevalence of PE in various subgroups, we also

created tertiles of PE pooled across all survey waves.

Statistical analyses

Analyses were conducted in Stata/IC 14.2 (StataCorp LP, College Station, TX, US), incorporating NLSY survey weights and sample design parameters for descriptive statistics to account for clustered sampling, attrition, and oversampling (32).

Statistical analysis proceeded in three steps. First, we estimated the age-adjusted mean PES in the full sample, as well as by race/ethnicity, gender, educational attainment, income and region. Second, we estimated the proportion of respondents in each PES tertile overall and by sociodemographic characteristic. Finally, we used linear generalized estimating equation (GEE) models to quantify the overall change in the PES over time. We also examined heterogeneity in change over time within and between subgroups (race/ethnicity, gender, educational attainment, income, and region). We tested between-group differences in the change over time using an interaction term (eg, time x Hispanics). All GEE models included an unstructured correlation structure, categorical indicators of time (quartiles: 1988-1993, 1994-2000, 2002-2008, 2010-2016), and robust standard errors, clustered at the respondent level. For all GEE models, we used the Stata margins command to estimate the predicted average PES values for each time period and subgroup.

Sensitivity analyses

We conducted several sensitivity checks. First, we used a data-driven principal component analysis (PCA) approach to construct the PES, which allowed us to assess the validity of our a priori choice of variable groupings and equally weighted domains. Second, to assess concordance between the theory-driven and datadriven scores, we computed percent agreement and Cohen's Kappa across quartiles of the continuous measures. Third, we combined the materials rewards (relative wages, paid vacation) and workers' rights (health insurance, retirement plan) dimensions to address ambiguity about whether health insurance belonged in the rights or material rewards dimensions (table 1). For example, Julia et al (10) classified "additional" insurance plans as a fringe benefit or material reward, but classified health insurance a workers' right; however, nationalized healthcare is common in the European context, unlike in the US, where it is largely employersponsored. Fourth, we examined whether the estimated PES trends changed when including unemployed observations in our sample (N=22 154). In these analyses, we assigned unemployed individuals a PES of 7 (most precarious) since precariously employed individuals frequently move in and out of unemployment. Fifth, we used an external income cutoff - the age- and yearspecific mean income as measured in the American Community Survey - to dichotomize people as higheror lower-income, which ensured the cutoff was nationally-representative. In our primary analyses, we used the sample-specific state-year mean wage/income as a cutoff (see table 1 and supplementary material, www.sjweh.fi/show_abstract.php?abstract_id=3939, figure S1). Both approaches allowed our measure to incorporate the increases in income inequality over follow-up; as income inequality increased, a smaller proportion of the sample had wages above the sample mean (given the variables' increasingly skewed distribution). Sixth, in our regression-based analyses, we tested whether including survey wave fixed-effects rather than categorical year indicators affected our estimates. Finally, we assessed the sensitivity of our results when using additional correlation structures (independent, exchangeable, auto-regressive).

Results

Descriptive statistics

Overall, the average age-adjusted PES was 3.17 [standard deviation (SD) 1.19]; the average PES was 2.96 (SD 1.23) in 1988, compared to 3.43 (SD 1.15) in 2016 (table 2, supplementary table S1). In the overall sample (ie, pooling data across years), tertiles of PE were defined as: low precarity: <2.62, medium precarity: 2.63-3.74, high precarity: >3.75. The proportion of respondents in high-precarity jobs increased from 27% in 1988 to 35% in 2016. Moreover, between 1988 and 2016, 31% of individuals went from low to medium or from medium to high precarity, with an additional 10% of the sample transitioning from low to high employment precarity over time.

The average age-adjusted PES and the prevalence of PE were significantly higher among women (3.34, SD 1.20), people of color (Hispanics: 3.24, SD 1.23; non-Hispanic Blacks: 3.31, SD 1.23), those with less education (primary:

3.99, SD 1.07; high school: 3.43, SD 1.19), those with lower-incomes (3.84, SD 1.08), and those residing in the South (3.23; SD 1.17).

Time trends in precarious employment

From 1988 to 2016, the overall PES increased by 0.29 points [95% confidence interval (CI) 0.26-0.31) or by 9% from 3.22 (95% CI 3.20-3.24) to 3.51 (95% CI 3.48-3.53) (table 3, figure 1). Except for the interpersonal relations dimension, all dimensions showed increasing precarity over time (data not shown).

Between 1988-1993 and 2010-2016, the PES increased by about 9% among non-Hispanic Whites (difference: 0.28 points; 95% CI 0.25-0.32), Hispanics (difference: 0.29 points; 95% CI 0.24-0.35); and non-Hispanic Blacks (difference: 0.30 points; 95% CI 0.25-0.35). However, the increase observed among nonHispanics Whites did not significantly differ from the increase among Hispanics (interaction: 0.01 points; 95% CI -0.06-0.08) or non-Hispanic Blacks (interaction: 0.02 points; 95% CI -0.05-0.08). Meanwhile, employment quality improved among people of "Other" races relative to the change among non-Hispanic Whites (interaction: -0.24 points; 95% CI -0.43- -0.05). The 11% increase observed among males (difference: 0.35 points; 95% CI 0.33-0.39) was significantly larger than the 6% increase among females (difference: 0.21 points; 95% CI 0.18-0.25) (interaction: 0.15 points; 95% CI 0.20-0.10).

The PES remained persistently high among the lowest-education subgroup over follow-up, but the precision was poor and the increase over time was not statistically significant (difference: 0.21 points; 95% CI -0.16-0.57). The PES significantly increased among all other education subgroups; we observed an 11% (0.38 point; 95% CI 0.34-0.42), 12% (0.37 points; 95% CI 0.33-0.41), and 6% (0.17 points; 95% CI 0.09-0.24) increase in the PES among those with a high school, college, and graduate level of education, respectively. Employment precariousness worsened over time among both lower- (difference: 0.22 points; 95% CI 0.19-0.25) and higher-income (difference: 0.39 points; 95% CI 0.36-0.42) individuals; however, workers with higher wages experienced a 14% increase in PE over time, compared to 6% among workers with lower wages (interaction: 0.17 points; 95% CI 0.13-0.21).

The score increased by about 9% (0.30 points) within most regions. Compared to the South, the change over time was only different in the West (interaction: -0.11 points; 95% CI -0.18- -0.04).

Sensitivity analyses

Overall, the score was robust to our sensitivity checks. First, although the variables grouped in our PCA approach differed from those grouped in our theoretical approach (supplementary figure S2, panel A), the qualitative trend in the scores was similar between 1988 and 2016 and there was evidence for the two scores' agreement (supplementary table S2). Second, when we combined the material rewards and workers' rights domains, trends were similar to our primary specification (supplementary figure S2, panel B). Third, as in our primary sample, we observed increases in employment precarity over time when we included the unemployed observations of the same respondents; however, the increases were larger in magnitude (supplementary fig- ure S2, panel C, table S3). For example, when including the unemployed, the overall PES increased by ~13% (difference: 0.50 points; 95% CI 0.46-0.54), including by 12% among non-Hispanic Whites (difference: 0.44 points; 95% CI 0.38-0.49), 14% among Hispanics (difference: 0.54 points; 95% CI 0.44-0.63), and 15% among non-Hispanic Blacks (difference: 0.60 points; 95% CI 0.52-0.67). Additionally, the increase was larger among non-Hispanic Blacks (interaction: 0.16; 95% CI 0.06-0.25) compared to non-Hispanic Whites. Overall trends were also similar when we used an alternative definition of lower- and higher-income (supplementary figure S2, panel D, table S4). When we used year fixedeffects rather than categorical indicators, results were larger in magnitude; there was a 17% increase in the PES comparing 1988 to 2016 (difference: 0.52 points; 95% CI 0.48-0.56) (supplementary figure S3, tables S5 and S6). Finally, results were similar in magnitude and significance when employing independent, exchangeable, and auto-regressive correlation structures (data not shown).

Discussion

To our knowledge, this is the first US-based study to use longitudinal, individual-level data, and a multidimensional indicator to describe employment precariousness over time. We also measured PE on a continuum, which we believe allowed for a more nuanced approach to assessing employment quality than the binary indicators predominately used in prior research. We found that the average PES throughout follow-up was significantly higher

among people of color, women, people with lower levels of education, and people with lower income. Moreover, between 1988 and 2(H6, the; overall PES significantly increased indicating wosnening rmpleymrnt quality ever time. Finally, we observed the largest increases in PE among males, people with a college education and higherincome individuals, suggesting long-term decreases in employment quality are widespread in the US, rather than confined to marginalized segments of the labor market.

Prior US-based trend estimates are based on unidimensional indicators and/or repeated-cross-sectional data, making comparisons with the present study challenging. However, our finding that PE increased in the US over the last 3E-yeaıs is generally consistent with repeated-cross-sectional data from the OSS, which suggested that non-standard work arrangements (temporary, on-call, contract workers) increased by about 11% from 2006 to 2010 (3). Our findings are also consistent with Kalleberg, who reported an increase in PE between 1970 and 2002, characterized by increased workloads, time pressures, hours worked, and insecurity (2).

Ba se d oe priod literature documenting differential opp ortunities for workers across the labor market due to factors like discrimination and sexism, we hypothesized that people of color, women, and individuals with lower levels of income and education would experience the largest increases in PE (13, 15, 17, 25, 26, 28). Women may face employment-related discrimination and are often responsible for childcare, which may keep them less attached to the labor force (27, 33). People of color are more likely than White people to experience job insecurity, leave their jobs involuntarily, and face employment-related discrimination (34-37). Meanwhile, higher education levels may afford workers more favorable employment conditions, like autonomy, stability and opportunities for advancement (27). Oonsistent with prior studies, we also found that OP was persistently higher and increased over follow-up among women and those from racialized and less-educated populations (13, 15, 17, 25, 26, 28). However, contrary to our hypothesis, we observed the largest change over time in precarity among males and college-educated and higherincome individuals. There are a number of plausible explanations for this finding. Larger increases for the most advantaged groups could, in part, reflect the fact that their PES was lower in 1988 and thus, there was a greater opportunity for their employment quality to worsen. The relatively large increase in PE among males may also be due to the declining rate of union membership in the US, as union membership is associated with better employment quality and historically, was more common among males (38). For example, the proportion of male workers covered by labor union contracts decreased from 25% to 11% from 1983 to 2018, whereas women's union participation only decreased from just 15% to 10% over that time period (39). The increasing precarity among those with college degrees could reflect increasing educational attainment in the US; college degrees may no longer afford workers additional bargaining power and prestige as they have become more common.

More broadly, large increases in PE over time among males and those with higher-educations or incomes suggests that PE is a more structural phenomenon, affecting large segments of the population. Our findings are generally consistent with scholars' hypothesis that employment quality in the US has declined even for individuals employed in "good", permanent, or Standard Employment Relationship (ie, permanent, full-time, regularly scheduled work with secure wages) jobs (2, 13, 17, 39). Kalleberg reports that overall, employment has gotten harder, more insecure, and largely lacks benefits or opportunities for advancement (2). Nonetheless, our findings do contrast with several US Bureau of Labor Statistics studies that have examined only contract type and found a relatively stable prevalence of non-standard employment arrangements (3, 13). However, we believe these studies are limited, as contract type is only one dimension of employment that impacts worker wellbeing. Moreover, recent evidence suggests such studies may be undercounting emerging forms of precarious work (eg, gig employment) (7, 40, 41).

Importantly, our findings may have implications for worker health and well-being. For example, material deprivation, occupational hazards, and stressful environments, could prevent people from obtaining necessities (eg, health care and housing), limit health-promotion (eg, leisure time physical activity), and increase psychological distress (eg, anxiety, depression) (8). Thus, this widespread increase in PE could harm public health in the US. Moreover, the persistently greater level of PE among marginalized populations may contribute to health disparities. Furthermore, in our specification that included the unemployed in our sample, there was a greater increase in PE among non-Hispanic Blacks and Hispanics than among non-Hispanic Whites, suggesting that changes in PE may exacerbate

racial/ethnic health disparities.

Strengths and limitations

Key strengths of this study included our use of a multidimensional, longitudinal measure of PE, which allowed us to describe employment precariousness over time and among subgroups in the US. We conducted our analyses over a 30-year time-period, using more than 100 000 observations and conducted many sensitivity checks, which supported our overall inferences. Although some scholars suggest that PE is best measured from surveys that are developed specifically to capture the health relevant aspects of precarity, we also demonstrated the utility of using secondary data to longitudinally measure PE in the US with a multidimensional indicator (10).

Nevertheless, our study had limitations. First, the NLSY does not sample with replacement and in any given wave the age distribution was narrow (seven years). Although the score was age-adjusted, the trends over time may be most generalizable to the 1957-1964 birth cohort sampled. The labor force is now more diverse and with the worsening employment quality over time, it is possible that younger generations are experiencing more precariousness earlier in their careers. Because we have seen growth in both low- and highly precarious jobs, younger generations' potential for improvement in employment quality may be more limited over time (ie, they may have a more difficult time closing the gap). Thus, PE may be contributing to increasing inequality. Second, after 2008, wages in our sample grew faster than the national average (data not shown). This may be because workers in our sample were older than the national average after 2008 (42), and because more-precarious workers were more likely to drop out of the sample during followup. Thus, we may have underestimated the extent to which PE increased over time. Nonetheless, we used age-adjustment and sampling weights, as appropriate, to address these potential biases. Third, we did not examine self-employed respondents because they lacked data on certain PE indicators and may have a different social class than non-self-employed respondents (eg, business owners versus workers). Therefore, these results are not generalizable to the 8% of the US population that is self-employed. Fourth, we relied on data from O·NET to operationalize interpersonal power relations; data from O·NET are time-fixed and only changed for an individual in our sample if they changed jobs during followup. This is likely why we did not see large increases in the dimension over time. Fifth, our carrying forwards and backwards of the promotion variable could have introduced measurement error in the training opportunities dimension in the early 1990s. Finally, although a multidimensional indicator has many strengths, it could mask heterogeneous trends within dimensions. Despite these limitations, the NLSY is one of the longest running studies in the US of a racially, ethnically and economically diverse sample of adults, spanning early adulthood through age 60, and thus, was the best individual-level longitudinal dataset available for our research question.

Concluding remarks

This is the first study to describe trends in PE in the US using a multidimensional indicator and longitudinal measurement. We found that PE increased between 1988 and 2016 by 9% overall and with variation among all subgroups. The PES was persistently higher and increased over time among women and those from racialized and less-educated populations, but the largest increases were observed among male, college-educated and higher-income individuals. Future work should strongly consider multidimensional employment quality indicators and longitudinal measurement, given that these widespread increases in PE could deleteriously affect public health.

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Sidebar

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DETAILS

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Hannerz, H., FilLic, Albertsen, K., PhD., Nielsen, M. L., PhD., & Garde, A. H., PhD. (2021). Long working hours and psychiatric treatment: A danish follow-up study. Scandinavian Journal of Work, Environment & Health, 47(3), 191-199. doi:https://doi.org/10.5271/sjweh.3936

Objective This study aimed to estimate prospective associations between long working hours and (i) redeemed prescriptions for psychotropic drugs and (ii) psychiatric hospital treatment due to mood, anxiety or stress-related disease, among full-time employees in Denmark. Methods Full-time employees who participated in the Danish Labor Force Survey sometime in the period 2000-2013 (N=131 321] were followed for up to five years in national registers for redeemed prescriptions for psychotropic drugs and psychiatric hospital treatment due to mood, anxiety or stress-related disease. Rate ratios (RR) were estimated for 41-48 versus 32-40 and >48 versus 32-40 working hours a week. The analyses were controlled for sex, age, night shift work, calendar time of the interview and socioeconomic status (SES). Prevalent cases were excluded in primary analyses. Results The RR for psychotropic drugs were estimated at 0.94 99% confidence interval (CI) 0.88-1.01] for 41-48 versus 32-40 working hours a week and 1.08 (99% CI 0.99-1.18) for >48 versus 32-40 working hours a week. The corresponding RR for psychiatric hospital treatments were estimated at 0.90 (95% CI 0.75-1.08) and 0.96 (95% CI 0.76-1.21). We did not find any statistically significant interaction between weekly working hours and age, sex, SES or night shift work. Conclusion Long working hours as they occur in in the general working population of Denmark are not an important predictor of mental ill health.

Lee, W., M.D., Kang, S., M.D., & Choi, W., M.D. (2021). Effect of long work hours and shift work on high-sensitivity C-reactive protein levels among korean workers. Scandinavian Journal of Work, Environment & Health, 47(3), 200-207. doi:https://doi.org/10.5271/sjweh.3933

Objective We aimed to investigate the association between low-grade inflammation as indicated by highsensitivity C-reactive protein (hsCRP) level and organizational factors, such as work hours and shift work. Methods We evaluated 7470 young and middle-aged workers who participated in the Korea National Health and Nutrition Examination Surveys from 2015-2018. Work hours were determined from self-reported questionnaires. Shiftwork was defined as a non-daytime fixed work schedule. An interaction effect between shiftwork and long work hours on the hsCRP level was estimated using relative excess risk due to interaction (RERI) and attributable proportion (AP) with 95% confidence intervals (CI). Results Increased hsCRP levels were prevalent in 25.2% of the study population. There was a significant association between long work hours and increased hsCRP, especially among middle-aged men odds ratio (OR) 1.50 (95% CI 1.20-1.87) for moderately increased hsCRP and OR 1.62 (95% CI 1.14-2.30) for highly increased hsCRP]. There was a significant interaction effect between long work hours and shiftwork on increased hsCRP among middle-aged workers. The RERI were 0.03 (95% CI 0.02-0.04) and 0.56 (95% CI 0.45-0.68) among middle-aged men and women, respectively. The AP were 0.02 (95% CI 0.01-0.03) and 0.36 (95% CI 0.31-0.40) among middle-aged men and women, respectively. Conclusions There was no significant association between shiftwork and the level of hsCRP. Long work hours were related to low-grade inflammatory processes, but only in middle-aged workers. There was an interaction effect between long work hours and shiftwork for increased hsCRP, especially in middle-aged women.

Yu, Y., M.D., Thijs, L., M.Sc, Saenen, N., PhD., Melgarejo, J. D., M.D., Wei, D., M.D., Yang, W.,M.D.PhD., . . . Yuzhang, Z.,M.D.PhD. (2021). Two-year neurocognitive responses to first occupational lead exposure. Scandinavian Journal of Work, Environment & Health, 47(3), 233-243. doi:https://doi.org/10.5271/sjweh.3940

Objectives Lead exposure causes neurocognitive dysfunction in children, but its association with neurocognition in adults at current occupational exposure levels is uncertain mainly due to the lack of longitudinal studies. In the Study for Promotion of Health in Recycling Lead (NCT02243904), we assessed the two-year responses of neurocognitive function among workers without previous known occupational exposure newly hired at lead recycling plants. Methods Workers completed the digit-symbol test (DST) and Stroop test (ST) at baseline and annual follow-up visits. Blood lead (BL) was measured by inductively coupled plasma mass spectrometry (detection limit 0.5 jj.g/ dL).

Statistical methods included multivariable-adjusted mixed models with participants modelled as random effect. Results DST was administered to 260 participants (11.9% women; 46.9%/45.0% whites/Hispanics; mean age 29.4 years) and ST to 168 participants. Geometric means were 3.97 and 4.13 ^g/dL for baseline BL, and 3.30 and 3.44 for the last-follow-up-to-baseline BL ratio in DST and ST cohorts, respectively. In partially adjusted models, a doubling of the BL ratio was associated with a 0.66% 95% confidence interval (CI) 0.03-1.30%; P=0.040] increase in latency time (DST) and a 0.35% (95% CI -1.63-1.63%; P=0.59) decrease in the inference effect (ST). In fully adjusted models, none of the associations of the changes in the DST and ST test results with the blood lead changes reached statistical significance (P>0.12). Conclusions An over 3-fold increase in blood lead over two years of occupational exposure was not associated with a relevant decline in cognitive performance.

Söderberg, M., PhD, Stattin, M., PhD., Robroek, S. J. W., PhD., Burdorf, A., PhD., & Järvholm, B., PhD. (2021). Industry mobility and disability benefits in heavy manual jobs: A cohort study of swedish construction workers. Scandinavian Journal of Work, Environment & Health, 47(3), 217-223. doi:https://doi.org/10.5271/sjweh.3932

Objectives This study aimed to investigate whether change from the construction industry to work in other industries at age 45-55 years lowered risks of disability benefits (DB) later in life (60-64 years of age). We hypothesized that risks would be lowered the most among those changing from the heaviest occupations. Methods The study included men employed in the construction industry during 1971-1993. We selected workers from the largest occupational groups in heavy (concrete workers and painters) and less heavy (drivers, electricians and foremen) occupations. The occurrence of DB in 1990-2015 was retrieved from national registers. Regression analyses were used to calculate relative risks (RR) of DB at 60-64 years, comparing those working in other industries to those still in the construction industry at the age of 45, 50 and 55 years. Results Mobility away from the construction industry was related to lowered DB risks at 60-64 years in all selected occupations. Effects were most pronounced among those who, at 55 years of age, worked in an industry other than construction, with significantly reduced RR for DB among concrete workers RR 0.63, 95% confidence interval (CI) 0.51-0.77], electricians (RR 0.61, 95% CI 0.47-0.77) and foremen (RR 0.78, 95% 0.63-0.96). Conclusions Risks for DB at 60-64 years of age were reduced among those who changed from construction work to other industries. Notable reductions were observed among workers originating from both heavy and less heavy occupations, and future studies should explore other factors, in addition to heavy workload, as motivators for leaving the construction industry.

Virtanen, M., PhD., Myllyntausta, S., PhD., Ervasti, J., PhD., Oksanen, T., M.D., Salo, P., PhD., Pentti, J., B.Sc, . . . Stenholm, S., PhD. (2021). Shift work, work time control, and informal caregiving as risk factors for sleep disturbances in an ageing municipal workforce. Scandinavian Journal of Work, Environment & Health, 47(3), 181-190. doi:https://doi.org/10.5271/sjweh.3937

Objectives This study aimed to examine the contribution of shift work, work time control (WTC) and informal caregiving, separately and in combination, to sleep disturbances in ageing employees. Methods Survey data were obtained from two prospective cohort studies with repeated measurements of working conditions, informal caregiving, and sleep disturbances. We used fixed-effect conditional logistic regression analysis to examine whether within-individual changes in shiftwork, WTC and informal caregiving were associated with changes in sleep. Secondary analyses included between-individuals comparison using standard logistic regression models. Results from the two cohorts were pooled using meta-analysis. Results Low WTC and informal caregiving were associated with sleep disturbances in within-individual analyses odds ratios (OR) ranging between 1.13 (95% confidence interval 1.01-1.27) and 1.48 (95% CI 1.29-1.68)] and in between-individuals analyses OR 1.14 (95% CI 1.03-1.26) to 1.33 (1.19-1.49)]. Shiftwork alone was not associated with sleep disturbances, but accumulated exposure to shiftwork, low WTC and informal caregiving was associated with higher risk of sleep disturbances (OR range 1.21-1.76). For some of the sleep outcomes, informal caregiving was related to a higher risk of sleep disturbances when WTC was low and a lower risk when WTC was high. Conclusions Informal caregiving and low WTC are associated with risk of sleep disturbances among ageing employees. The findings also suggest that low WTC in combination with informal caregiving may increase the risk of sleep disturbances whereas high WTC may alleviate the adverse impact of informal caregiving on sleep.