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Do good working conditions make you work longer? Analyzing retirement decisions using linked survey and register data

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ABSTRACT

We analyzed the role of adverse working conditions and new management practices in the determination of employees' retirement behavior. The combined data contain both comprehensive information on perceived job disamenities, job satisfaction, and intentions to retire from two nationally representative cross-sectional surveys and information on employees' actual retirement decisions from longitudinal register data that can be linked to the surveys. Using a trivariate ordered probit model, we find that job dissatisfaction arising from adverse working conditions is significantly related to intentions to retire and that this, in turn, is related to actual retirement during an extensive follow-up period.

Introduction

Populations in industrialized countries are aging rapidly. This structural change puts pressure on public finances, social support programs that target retired persons, and the sustainability of pension systems. In the related policy discussion, two broad approaches to addressing these challenges have been promoted. First, there are "hard" measures. A popular policy measure to improve the sustainability of the pension systems has been to increase the mandatory retirement age and/or to cut pension benefits to force people to retire later in life. Second, there are "soft" measures, which refer primarily to improvements in perceived working conditions. The goal of these policy designs is to encourage people to lengthen their working careers voluntarily, thus avoiding the need to change regulations.

Perceived well-being at work is important for employees because job satisfaction is a key domain of employees' overall well-being in life (Oswald, 2010). Job satisfaction and productivity at the firm level are also positively related (Böckerman and Ilmakunnas, 2012; Oswald et al. 2015). Consequently, investments in better working conditions and improvements in employee well-being can be mutually beneficial for both employees and employers and to society more broadly.

This paper examines the links between measures of working conditions and new management practices (the so-called "high involvement management") regarding actual retirement decisions. Workers' satisfaction with their work and their subsequent retirement decisions are likely connected not only with physical working conditions but also with how they are treated by management and supervisors. Hence, it is important to consider both of these aspects of work. We contribute to the literature by modeling the complete chain from perceived working conditions and new management practices to job satisfaction, retirement intentions, and actual retirement. To accomplish this, we used two sets of comprehensive survey data on perceived well-being at work as well as administrative data on actual retirement during an extensive follow-up period. Our survey data contain very detailed information on working conditions (perceived harms and hazards) at the individual level, and the linked survey and register data are nationally representative for the working age population in Finland. We found that adverse working conditions are negatively correlated with job satisfaction, positively correlated with retirement intentions and negatively correlated with continuing work beyond the official retirement age. New management practices have effects that are opposite those of the work disamenities. Section "Related literature" briefly summarizes the key aspects of the earlier literature on working conditions and retirement. Sections "Data" and "Modeling approach" describe the linked survey and register data and our modeling approaches. Section "Results" presents the baseline estimation results with a battery of sensitivity analysis. Section "Conclusions" concludes.

Related literature

Beehr (1986) analyzed the process of retirement and argued that personal characteristics and the work environment influence a person's

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preference for retirement. Preferences determine the intention (or decision) to retire, which then manifests as actual retirement behavior.[1]

Karasek (1979) presented the seminal model concerning the determinants of employee well-being at work. He stressed the balance between job demands and job control. The combination of job demands and job control affects employees' intentions to quit. It also affects actual retirement decisions, as retirement is an "extreme form" of employee quitting behavior. Therefore, high job demands coupled with poor job control increase both the intention to quit and actual retirement.[2]

The rational decision approach stresses the role of benefits and costs. The personal preference for leisure time (or the disutility of work) and consumption opportunities provided by income determine labor supply decisions at the extensive margin. Retirement leads to an increase in leisure time, but opportunities to consume decrease because the pension is lower than the prior wage. Therefore, workers face a tradeoff. In the standard model, adverse working conditions, since they increase the disutility of work, reinforce the preference for retirement if all else is equal. If there is a compensating wage differential for perceived adverse working conditions, the outcome is not clear (Filer and Petri, 1988). However, the monetary compensation for adverse working conditions in terms of higher pay is only rarely complete (e.g., Böckerman and Ilmakunnas, 2006, 2009). As a result, perceived working conditions tend to have an economically significant influence on retirement decisions.

There is empirical literature on the relationship of working conditions to retirement intentions or actual retirement in the fields of industrial relations, labor economics, epidemiology, management, and organizational psychology. The working condition variables range from general indicators of having physically or psychologically demanding work to more specific indicators of stress, repetitive work, etc. Positive work aspects include, for example, management support and recognition. Much of the research is cross-sectional and studies either retirement intentions (or planning) or actual retirement (or retirement decision) but not both (for surveys, see Topa et al. 2009; Wang and Shultz, 2010; Fisher et al., 2016). The analyses of intentions are usually based on cross-sectional surveys, and most early studies of actual retirement use cross-sectional information on retirement status (e.g., Quinn, 1978).

To examine the relationships of working conditions to both retirement intentions and actual retirement, one would need both subjective and objective information on behavior. So far, the research is quite limited.[3] However, some recent studies have examined the connection of working conditions to subjective and objective retirement indicators side-by-side, using longitudinal surveys such as HRS, SHARE and ELSA (e.g., Schnalzenberger et al. 2014; van Solinge and Henkens 2014; dal Bianco et al. 2015; Angrisani et al. 2015; Munnell et al. 2015; Carr et al. 2016). A key drawback of this research is that intentions and actual retirement are not integrated. An alternative approach is to combine survey and register data. For example, Solem et al. (2016) used this kind of data to compare retirement intentions and actual retirement, but they did not have information on working conditions. Additionally, Blekesaune and Solem (2005), for example, analyzed survey-based measures of working conditions and register-based information on actual retirement, but they did not have information on intentions.[4]

[1] We provide a more comprehensive discussion of the earlier empirical literature in the working paper version.

[2] Retirement intentions are important *per se* because they are a form of worker discontent, which may lead to lower performance at work.

[3] A separate literature explicitly considers the rationality of retirement expectations using information on both expected and realized retirement age (e.g., Bernheim, 1989; Benítez-Silva and Dwyer, 2005). However, this research has not examined the potential role of perceived working conditions and management practices in the decision process.

[4] Longitudinal survey or register data makes it possible to model actual retirement using various alternative outcomes, such as retirement age (an

Thus, the literature has analyzed retirement intentions and retirement choices as separate processes. We took a different approach by integrating them as proposed in Beehr's (1986) retirement model and Ajzen's (1991) theory of planned behavior. A deeper understanding of the connections between retirement intentions and choices is also necessary to draw policy-relevant insights into the pertinent mechanisms. An important gap in terms of modeling approaches is that unobservable individual- or job-specific characteristics, such as personality traits and job attitudes, are not observed in the data and therefore not accounted for. These unobservable traits can jointly determine the outcome variables of interest. Unobservable traits are problematic for the identification of effects if the same characteristics (or different but mutually strongly correlated characteristics, such as satisfaction at work and retirement intentions) affect actual retirement decisions and their determinants. Even with longitudinal data, estimation is based mostly on cross-sectional variation because actual retirement occurs only once. This makes it impossible to use fixed effects estimation to eliminate unobservable time-invariant traits. Our modelling strategy partially addresses this issue.

Data

Our empirical analysis used linked survey and register data. The data on perceived working conditions and intention to retire originated from the 2003 and 2008 Quality of Working Life Surveys (QWLS) of Statistics Finland (Lehto and Sutela, 2005, 2009). The initial sample for QWLS is from the Labor Force Survey, where a random sample of the working age population is selected for a telephone interview. The respondents are wage and salary earners between 15 and 64 years old with a normal weekly working time of at least 5 h. The response rate of the surveys was 77.9% in 2003 and 67.6% in 2008. The sample sizes were 4101 in the 2003 survey and 4392 in the 2008 survey.

The QWLS is a repeated cross-sectional data set that does not contain information on actual retirement choices. For this reason, we linked the QWLS data to comprehensive longitudinal register data for the same persons. These included Finnish Longitudinal Employer–Employee Data (FLEED) from Statistics Finland and the pension records of the Central Pension Institute. FLEED records each employee's employer during the last week of each year. FLEED contains rich background information on both employees and their employers. The Central Pension Institute keeps comprehensive administrative records of actual retirement for the payment of pensions. We linked the data using unique personal identifiers, i.e., ID codes for the persons. We could follow *all* employees in the QWLS data up to 2013. Using information from the Central Pension Institute, we observed actual retirement choices during the follow-up period (2004–2013 for QWLS 2003 and 2009–2013 for QWLS 2008).

The QWLS contains information about intention to retire. It includes a question about having thoughts of retirement before the official retirement age and uses the following alternatives: 'often' (coded as 3), 'sometimes' (2) or 'never' (1). Because this is ordered qualitative information, we formed a similar variable for the timing of retirement. Actual retirement could occur after (coded as 3), at (2) or before (1) the official retirement age.

Pension reform occurred in Finland in 2005, before which the normal retirement age in the private sector was 65 years; however, it was possible to retire earlier at ages 60–64 with a lower pension. Most state or municipal employees and some special occupations have had a lower retirement age. Pension reform made old-age retirement flexible between the ages 63 and 68, with earlier retirement with a lower pension possible only for those who were 62 years old. The retirement

(*footnote continued*)
indicator of retiring early), indicators for exit through alternative exit channels in a multinomial model, or the hazard of exit in a duration model.

ages in the private and public sectors were harmonized for new employees, but existing public sector employees were given a personal retirement age based on age and tenure. The pension reform led to an increase in the average retirement age. At the same time, there has been a clear concentration of retirements at the age of 63, which has become a social norm or default option (van Erp et al., 2014). Therefore, we treated 63 as the official retirement age referred to by retirement intention and actual retirement variables. However, for public sector employees we used personal retirement ages when they were below 63. There was another reform in 2017, but our follow-up period does not extend that far. There is some measurement error in the retirement age, as we know only the month and year of birth. Thus, we rounded the measured retirement age to the nearest integer years. Retirement at age 63.5 was therefore defined as retirement after the lowest official retirement age.

In addition to the old-age pension, there are other early exit routes: disability, part-time retirement, and unemployment. Disability pensions require medical verification but have no specific age limit. Part-time pensions can be granted to an employed person at least 61 years old who continues working part-time. In addition, disability retirement is possible on a part-time basis. Older employees who become unemployed can use extended unemployment benefits to bridge the time until old-age pension age. The lower age limit for this system has gradually increased to 59 years. In our analysis, we concentrated on full-time retirement, either old age or full disability. Some of the QWLS survey participants were actually already on part-time retirement. We left out the unemployment route, as it is a separate system and not officially treated as retirement. Full-time retirement is relevant because it is the policy variable that has gained most attention in debate.

In the estimations, we concentrated on those who were 53 or older in the 2003 survey and those who were 58 or older in the 2008 survey. These employees reached age 63 by the end of the follow-up period, and we observed whether they retired at age 63.[5] Moreover, we required the participants to be younger than 63 years during the survey, because for them the choice to retire before, at, or after age 63 is still relevant. During the 2003 survey, the pension reform was not yet in force, but the upcoming changes were public information. In fact, the respondents of the QWLS were explicitly reminded of the reform. However, to account the fact that some of the 2003 survey participants could retire with lower pensions before the new rules came into force, we included an indicator variable for the 2003 survey. In robustness analysis, we also studied the two surveys separately. Furthermore, we left out those who were already fully retired but still doing some work, persons who died during the follow-up period before reaching official retirement age, and a few inconsistent answers (high retirement intentions although already above retirement age). The sample size in the estimations was 1217.

It is notable that the likelihood of a job switch is quite low for individuals who are approaching the (official) retirement age. Among the QWLS 2003 participants in our estimation sample, 18.7% have switched jobs between the survey and the end of the follow-up. In QWLS 2008, the corresponding share is only 8.6%, as the follow-up period is shorter and the sample participants older. In robustness checks we estimate the baseline model using also a sample that includes only those who have not changed employer.

Fig. 1 illustrates the structure of the linked data using the so-called Lexis diagram. The calendar time is on the horizontal axis, and age is presented on the vertical axis. Each upward-sloping line in the figure depicts the increasing age of a birth cohort and, importantly, the ages at which the cohorts were observed in the data. These lines start from the survey year, 2003 (left panel) or 2008 (right panel), and end at the end of the follow-up period in 2013. The horizontal line is at age 63, which is the lowest official retirement age in Finland for most of the persons in the dataset. The figure includes only those cohorts that reached age 63 during the follow-up, i.e., those used in the analysis. The lowest upward-sloping line in both parts of the figure shows the youngest birth cohort in the analysis, i.e., those who were born in 1950. The highest upward-sloping lines are the oldest cohorts, i.e., those who were born in 1941 (in the 2003 survey) or 1946 (in the 2008 survey). Lines that were above them in the figure correspond to older cohorts that were not used because they were already above age 63 during the survey years. Their share of all respondents was 0.6% in both surveys. On the other hand, lines that were below those in the figure correspond to younger cohorts that were not used because they did not reach age 63 during the follow-up period. Because the surveys included persons of working age, 16–65 years, many of them are still very far from retirement. In the 2003 survey, the share of those who were 52 or younger was 78.9%, and in the 2008 survey those who were at most 57 years old accounted for 90.0% of all respondents.

Our interest is in retirement intention, actual retirement and their background factors, especially job satisfaction, working conditions and management practices. The question about job satisfaction contained the following responses: 'very satisfied,' 'quite satisfied,' 'rather unsatisfied,' 'very unsatisfied,' and (in the 2008 survey) 'difficult to say'. Most respondents were satisfied with their work. We combined the lowest satisfaction levels (rather unsatisfied, very unsatisfied or difficult to say) into a group called 'unsatisfied.' We therefore have three groups: 'very satisfied' (coded as 3), 'quite satisfied' (2), and 'unsatisfied' (1).

The key working condition variables capture perceived harms and hazards. We focused on the physical aspects of the workplace because self-reported information on working conditions could reflect social desirability and justification. For example, workers can report worse working conditions in response to failing performance at work, bad health or other personal characteristics. The potential bias of self-reported information should be much smaller for the concrete measures of physical working conditions compared with psychosocial working conditions.

For perceived harms, the highest category corresponds to the perception by a worker that a certain feature of working conditions is 'very much' (on a five-point scale) an adverse workplace factor. The harms included 19 factors such as heat, cold and dust, among other things. For perceived hazards, the highest category among three possibilities was the one in which the respondent considered a certain feature at the workplace to be 'a distinct hazard.' The hazards included 10 factors, such as accident risk, risk of strain injuries and risk of grave work exhaustion, among other things.[6] We aggregated the responses to the questions about adverse working conditions by forming a dummy variable that equals one if there is at least one clearly adverse factor (the variable 'harms') and a dummy that equals one if there is at least one distinct hazard (the variable 'hazards'). These measures have been used earlier in Böckerman and Ilmakunnas (2006, 2009) and Böckerman et al. (2012a). The correlation of the harms and hazards variables was 0.39 (significant at a better-than-1% level). This means that they are related but not fully correlated and therefore measure different aspects of work.

We also leveraged detailed self-reported information on the quality of management practices from the QWLS as an additional aspect of perceived working conditions. We used a binary indicator, new management practices, to signify having more than one of the following new management practices: incentive pay, employer-provided training, self-managed teams and information-sharing by the employer

[5] Specifically, the included persons either reached age 63 and retired during the follow-up period or they had not yet retired but reached age 63.5 years in 2013 and were classified to have retired after age 63. Persons who reached age 63, but not 63.5, in 2013 and who had not retired were left out (36 observations).

[6] Appendix (Table A1) documents all of the aspects of working conditions that were included in the harms and hazards variables.

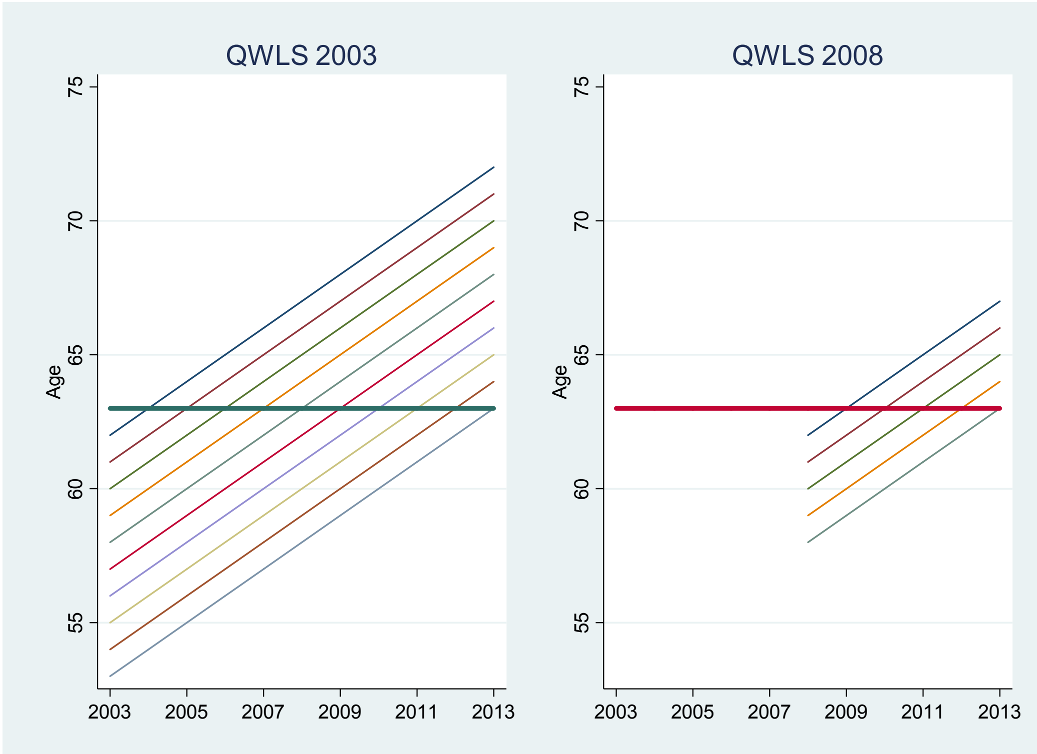

Fig. 1. Lexis diagram of the birth cohorts included in the estimations. Notes: The horizontal line is the age of 63, which is the lowest official retirement age in Finland.

(Böckerman et al., 2012a, 2013). Incentive pay is an indicator for those who are personally subject to performance-related pay; training is relevant for employees who have participated in employer-provided training during the past 12 months; self-managed teams refer to individuals who work in a team that selects its own leader and decides on the internal division of responsibilities; and information sharing involves employees who are informed about changes at work at the planning stage rather than shortly before the change or at the time of its implementation. These measures correspond to the crucial pieces of a high-performance workplace from the perspective of employees, as outlined in Appelbaum et al. (2000). Becker and Gerhart (1996) maintain that the four most common components of high involvement management systems are self-managed teams, quality circles, employer-provided training, and contingent pay (cf. Böckerman et al., 2012b). We captured all of these, except quality circles, in our measure of management practices using the QWLS. The management practices variable is negatively correlated with the adverse working conditions variables. The correlations with both harms and hazards were −0.13 (both significant at a better-than-1% level). New management practices are somewhat more common in workplaces with less adverse working conditions, but it is not possible to judge the causal relationships between the variables.

We used individual characteristics from QWLS and FLEED, measured during the survey year, as the standard control variables. These included age (in years), gender (an indicator for females), education (indicators for secondary and tertiary education, with basic education as the reference group), and income level (log of annual earnings) as indicators of a person's socioeconomic status. Income theoretically has opposing effects on retirement. Higher income increases the cost of retiring (substitution effect), but it also makes it more affordable to retire earlier (income effect). The income variable also takes into account the fact that the pension level depends on pre-retirement earnings. QWLS also contains information on self-assessed capacity to work on a scale from 0 to 10. We expected that individuals with good working capacity would be more inclined to work longer. Furthermore, we controlled for the size of the employer (indicators for size classes in terms of the number of employees: 10–49, 50–249, and 250-, with below 10 as the reference category).

Using information from the comprehensive registers of the Central Pension Institute, we controlled for some key retirement-related covariates that included an indicator for being in part-time retirement, an indicator for whether the person has a retired spouse, and the (age-dependent) pension accrual rate (i.e., the percentage of annual income by which one additional work year increases the pension). These were measured during the survey year (the accrual rate is the average of two years following the survey). Those already on partial retirement may be more inclined to enter full-time retirement earlier than those who have not taken the part-time option. It is important to account for the spouse's labor market status because it affects the utility of leisure time. Thus, the return on a person's leisure time is higher if the spouse is also retired and they can spend their leisure time together. The accrual rate creates incentives to work longer. After the pension reform, there was a higher accrual rate for those at age 63 or higher. The accrual rate before year 2005 was 1.5% until age 60 and 2.5% at higher ages; in the years 2005–2016, it was 1.5% until age 52, 1.9% at ages 53–62, and 4.5% at ages 63–68. We included an indicator of experiencing unemployment during the follow-up period until 2013 as an additional variable from FLEED. Those who become unemployed at older ages may have difficulty returning to work and are therefore more likely to retire at the (minimum) official retirement age. Actually, the institutional system encourages this, as the older unemployed can receive extended unemployment benefits until the official retirement age.

There may be selectivity of certain kinds of individuals to workplaces with certain kinds of working conditions. This may create a correlation of the working condition variables with the equation error. Because we have three working conditions variables, instrumenting them is cumbersome. We therefore included as additional controls variables that reflect past working career and that may be correlated with selection to, e.g., unfavorable working conditions. They were indicators for having had more than two clearly different kinds of occupations over the working career and for having switched jobs more than twice during the past five years before the survey. These indicators were based on the QWLS. Descriptive statistics of the variables are reported in Table A2.

Modeling approach

Our empirical application had three endogenous variables (job satisfaction, intention to retire and actual retirement) measured with an ordinal scale (1, 2, or 3). We assumed that there are latent continuous variables behind the observed ordinal variables. We modeled the relationships as a trivariate ordered probit model. This implies the presence of some unobservable variables, such as personality traits, that may affect all three dependent variables. Thus, the equations' errors are correlated with each other.[7] We used the extended regression model framework in Stata (StataCorp, 2017; Roodman, 2011) to estimate the parameters of the model.

The identification of the parameters of the model is based on the idea that there is a triangular structure between the variables of interest. Thus, based on the earlier literature summarized in Section "Related literature" we assumed that job satisfaction affects the intention to retire and that retirement intentions affect the timing of actual retirement, but there are no backward effects. On the right-hand side of the estimated equations, these variables appear as indicators for categories 2 and 3 (using the category 1 as the reference) of the job satisfaction and retirement intentions variables. In addition, we used exclusion restrictions on the explanatory variables. Research using equation systems with binary dependent variables and endogenous dummy regressors has shown that exclusion restrictions are required to correctly identify the parameters (Mourifié and Méango, 2014; Han and Vytlacil, 2017). There are no corresponding results for ordered variables, but presumably a similar principle holds.

We assumed that individual characteristics such as age, gender and education affect all outcomes. Earnings is a measure of socioeconomic status and was included in all models. Perceived working conditions and management practices influence job satisfaction, but we assumed that they have an effect on the intentions to retire only through job satisfaction. Working capacity and firm size influence both job satisfaction and intentions to retire, but they do not directly influence actual retirement choices because the working capacity and the firm in which a person is employed could have changed after the QWLS survey. Additionally, we assumed that part-time retirement and a spouse's retirement status influence the intention to retire and actual retirement timing, but not job satisfaction. Finally, unemployment measured after the QWLS had an influence only on the actual retirement decisions, as it was not yet known during the surveys.

Potential biases are relevant to the interpretation of the estimation results. First, there may be unobserved heterogeneity that is correlated with the dependent variables and not captured by modeling the correlation structure between the equations. Therefore, we are cautious to interpret the estimated relationships as causal effects. Second, there are issues related to sample selection. As the data included only people who were (still) working in 2003 or 2008, those who had particularly good working conditions were likely to be overrepresented in the estimation sample. On the other hand, retirement before the official retirement age is more likely observed for those who are exposed to adverse working conditions and, consequently, retire early. These two sample selection biases have opposite effects on the estimates. In an ideal situation, they cancel each other out. Third, the selection of employees to workplaces with different working conditions or management practices is not random.

[7] The model includes the standard assumption that the error terms are normally distributed.

Results

Descriptive evidence

Tables 1 and 2 present cross-tabulations of the variables of interest. The tables show the expected relationships between the variables. Job satisfaction was negatively correlated with having retirement intentions. Of those who are unsatisfied with their work, 61.6% often think about retirement before the lowest official retirement age and 19.8% never do. Of those who are very satisfied, only 21.4% think often about retirement and 49.8% never do. Retirement intention was, on the other hand, negatively correlated with the timing of retirement. Among those who never think about early retirement, 59.7% retire after the lowest official retirement age and only 7.6% before it. Among those who often think about retirement, the distribution of early exits and late retirements is notably different: 27.7% delay retirement to an age above the lowest official retirement age, and 24% retire before it. The rank correlations were −0.288 for retirement age and retirement intentions and −0.252 for retirement intentions and job satisfaction. Importantly, the rank correlation of retirement age and job satisfaction was only 0.078, which supports the assumption of our three-equation model that job satisfaction is related to retirement age via retirement intentions, but not directly.

Baseline estimates

The estimation results of the trivariate ordered probit model that partly account for the endogeneity concerns are presented in Table 3.[8] We found that perceived harms and hazards were negatively correlated with job satisfaction, whereas new management practices were positively correlated to it. These results are consistent with earlier findings in the relevant literature. Job satisfaction was negatively correlated with retirement intentions, and retirement intentions were negatively correlated with actual retirement age, as was expected based on the cross-tabulations. Furthermore, the correlations of the equation errors were consistent with the view that, on the one hand, the unobservable factors behind actual retirement decisions and retirement intentions are related and, on the other hand, the unobservable factors behind retirement intentions and job satisfaction are related.[9]

The estimates of the control variables show that self-assessed working capacity was positively correlated with job satisfaction and negatively correlated with retirement intentions. The level of education was negatively correlated with retirement intentions, and the highest level of education was negatively correlated with job satisfaction. The pension accrual rate was negatively correlated with retirement intention, supporting the idea that higher accrual at higher ages encourages delayed retirement. Already being on part-time pension during the survey was positively correlated with having retirement intentions, as

[8] We do not report the estimated thresholds of the ordered models because they have no clear interpretation. As the model has no constant terms, the thresholds are numbers between minus infinity and infinity.

[9] The raw (unconditional) correlation between job satisfaction and retirement intentions is negative in Table 1. However, the correlation between the relevant unobservable factors is positive in Table 3. Similarly, the residual correlation of unobservable factors in the retirement intentions and retirement timing equations is positive despite the negative raw (unconditional) correlation. It should be noted that these are the residual correlations after we have already taken into account the negative relationship between job satisfaction and retirement intentions by including job satisfaction in the retirement intentions equation and similarly the negative correlation of retirement intentions and retirement timing by including retirement intentions in the retirement equation. Interestingly, different signs for raw correlations and residual correlations have been obtained also in studies that have used a trivariate probit model for racial harassment, job satisfaction and quit intentions (Shields and Price 2002; Antecol and Cobb-Clark, 2009).

Table 1
Job satisfaction and retirement intentions.

Retirement intention	Job satisfaction 1 Unsatisfied	2 Rather satisfied	3 Very satisfied	Total
1 Never	17	203	212	432
	3.94	46.99	49.07	100.00
	19.77	28.79	49.77	35.50
2 Sometimes	16	242	123	381
	4.20	63.52	32.28	100.00
	18.60	34.33	28.87	31.31
3 Often	53	260	91	404
	13.12	64.36	22.52	100.00
	61.63	36.88	21.36	33.20
Total	86	705	426	1217
	7.07	57.93	35.00	100.00
	100.00	100.00	100.00	100.00

Notes: In each cell, the first entry is the number of observations, the second is the percentage share of the row total and the third is the percentage share of the column total.

Table 2
Retirement intentions and actual retirement.

Actual retirement	Retirement intention 1 Never	2 Sometimes	3 Often	Total
1 Before lowest official retirement age	33	42	97	172
	19.19	24.42	56.40	100.00
	7.64	11.02	24.01	14.13
2 At lowest official retirement age	141	181	195	517
	27.27	35.01	37.72	100.00
	32.64	47.51	48.27	42.48
3 After lowest official retirement age	258	158	112	528
	48.86	29.92	21.21	100.00
	59.72	41.47	27.72	43.39
Total	432	381	404	1217
	35.50	31.31	33.20	100.00
	100.00	100.00	100.00	100.00

Notes: In each cell, the first entry is the number of observations, the second the percentage share of the row total and the third the percentage share of the column total.

expected. However, neither pension accrual nor part-time pension were significantly correlated with actual retirement, which may follow from the fact that actual retirement can occur several years after the survey. Earnings were not significantly correlated with any of the dependent variables, possibly because the models included several controls that are likely to be correlated with earnings. Experiencing unemployment in the follow-up period had a negative connection with retirement age. The indicator for past job changes was positively related to job satisfaction, whereas the indicator for several different occupations was not significant.

Among the demographic variables, age obtained a positive coefficient in the equations for actual and intended retirement. This is natural because those who have already continued to work until old age may be more likely to keep working after the official retirement age. Females had less frequent retirement intentions, but in actual retirement there are no gender differences. Finally, having a retired spouse was negatively correlated with continuing to work but is insignificant in the equation for retirement intentions.

Table 4 reports the average marginal effects of these variables on the three dependent variables. A variable can have a marginal effect on the dependent variable of the equation where it appears and on the dependent variables of subsequent equations but not on the dependent variables of previous equations. Furthermore, a working condition variable for example, which appears in the job satisfaction equation,

Table 3
Three-equation ordered probit model.

	Job satisfaction	Retirement intentions	Actual retirement
Retirement intentions = 2			−0.520***
			(0.161)
Retirement intentions = 3			−1.156***
			(0.267)
Job satisfaction = 2		−0.967***	
		(0.268)	
Job satisfaction = 3		−1.859***	
		(0.490)	
Age	0.018	0.049**	0.070***
	(0.015)	(0.021)	(0.024)
Female	0.078	−0.127*	−0.090
	(0.071)	(0.070)	(0.070)
Secondary education	−0.056	−0.137	0.017
	(0.086)	(0.083)	(0.083)
Tertiary education	−0.209**	−0.299***	0.133
	(0.092)	(0.086)	(0.093)
Log (Earnings)	0.002	−0.037	−0.011
	(0.028)	(0.038)	(0.034)
Pension accrual rate		−0.348***	0.002
		(0.065)	(0.067)
Part-time retirement		0.380***	0.015
		(0.144)	(0.150)
Spouse retired		0.003	−0.191**
		(0.076)	(0.080)
Unemployment			−0.676***
			(0.090)
Working capacity	0.170***	−0.149***	
	(0.029)	(0.044)	
Past job changes	−0.519***		
	(0.195)		
Different occupations	−0.002		
	(0.069)		
Harms	−0.249***		
	(0.083)		
Hazards	−0.235***		
	(0.077)		
New management practices	0.407***		
	(0.088)		
Establishment size 10–49	0.212**	0.123	
	(0.091)	(0.094)	
Establishment size 50–249	0.082	−0.036	
	(0.099)	(0.099)	
Establishment size 250-	0.125	0.311***	
	(0.122)	(0.121)	
Survey year 2008	−0.180**	−0.220***	−0.121
	(0.080)	(0.088)	(0.088)
Correlations of the errors			
		Retirement intentions	Actual retirement
Job satisfaction		0.336*	−0.037
		(0.174)	(0.045)
Retirement intentions 0.214			0.214*

Notes: N = 1217. Significance: *** 1%, ** 5%, * 10%.

has a marginal effect on retirement timing through a channel of effects. Working conditions affect job satisfaction, job satisfaction affects retirement intentions, and intentions affect retirement timing. Therefore, the marginal effects tend to decrease the later the dependent variable is in the three-equation system. Table 4 presents the average marginal effects on the probabilities of the highest categories (i.e., those coded as 3) of the ordered variables, i.e., high job satisfaction, frequent thoughts about early retirement, and retirement after the official retirement age.[10]

[10] The average marginal effects on the probabilities of categories 1 and 2 are shown in Table A3. In the three-category ordered model, the sign of a marginal effect of the lowest category is opposite the sign of the corresponding marginal

Table 4
Average marginal effects.

	Average marginal effect on the probability of:		
	Job satisfaction = 3	Retirement intentions = 3	Actual retirement = 3
Age	0.006	0.015**	0.022***
	(0.005)	(0.007)	(0.008)
Female	0.027	−0.048**	−0.022
	(0.024)	(0.022)	(0.025)
Secondary education	−0.019	−0.041	0.013
	(0.029)	(0.027)	(0.029)
Tertiary education	−0.071**	−0.082***	0.062*
	(0.031)	(0.027)	(0.033)
Log (Earnings)	0.001	−0.012	−0.002
	(0.010)	(0.011)	(0.013)
Pension accrual rate		−0.115***	0.023
		(0.021)	(0.024)
Part-time retirement		0.125***	−0.019
		(0.047)	(0.053)
Spouse retired		0.001	−0.069**
		(0.025)	(0.028)
Unemployment			−0.243***
			(0.033)
Working capacity	0.058***	−0.063***	0.013***
	(0.010)	(0.014)	(0.003)
Past job changes	−0.177***	0.041***	−0.013***
	(0.067)	(0.015)	(0.005)
Different occupations	−0.001	0.0001	−0.0001
	(0.023)	(0.005)	(0.002)
Harms	−0.085***	0.020***	−0.006***
	(0.028)	(0.007)	(0.002)
Hazards	−0.081***	0.019***	−0.006***
	(0.026)	(0.006)	(0.002)
New management practices	0.139***	−0.032***	0.010***
	(0.029)	(0.007)	(0.002)
Establishment size 10–49	0.073**	0.024	−0.003
	(0.031)	(0.030)	(0.006)
Establishment size 50–249	0.028	−0.018	0.004
	(0.034)	(0.031)	(0.006)
Establishment size 250-	0.043	0.093**	−0.017**
	(0.042)	(0.039)	(0.008)
Survey year 2008	−0.062**	−0.058**	−0.034
	(0.027)	(0.028)	(0.031)

Notes: N = 1217. Significance: *** 1%, ** 5%, * 10%.

Negative work aspects, harms and hazards, decrease the probability of being very satisfied with work by over 8 percent points and increase the probability of thinking often about retirement before the official age by 2 percent points. Likewise, they decrease the probability of retiring after the official age by 0.6 percent points. Being exposed to new management practices increases the probability of high job satisfaction by 14 percent points, decreases the probability of frequent early retirement intentions by 3 percent points and increases the probability of late retirement by 1 percent points. The associations are relatively small for the retirement variables, as the unconditional shares of those who are very satisfied, have frequent retirement thoughts, and retire late are 35.0%, 33.2%, and 43.4%, respectively (Tables 1 and 2).

Among the control variables, an age increase of one year increases the probability of late retirement by 2.2 percent points while females are 4.8 percent points less likely than males to have frequent thoughts of early retirement. Those with tertiary education have a 7-percentage-point lower probability of high job satisfaction, an 8-percentage-point lower probability of early retirement intentions, and a 6-percentage-point higher probability of late retirement than those with basic education. The variables related to pension and unemployment have relatively high marginal effects. A one-percentage-point increase in the accrual rate decreases the probability of frequent early retirement thoughts by 11.5 percent points and being on part-time pension by 12.5 percent points, but their effects on actual retirement are insignificant. Having a retired spouse decreases the probability of late retirement by 7 percent points and experiencing unemployment by 24 percent points. Having had several job changes in the past is related to early retirement. It decreases the likelihood of very good job satisfaction by over 17 percentage points, increases the likelihood of frequent retirement thoughts by 4 percentage points and decreases the likelihood of late retirement by one percentage point.

(footnote continued)
effect of the highest category, whereas in the middle category the sign can go either way. Table A3 shows that in the cases of job satisfaction and actual retirement the marginal effects on the probabilities of categories 1 or 2 are smaller in absolute value than the marginal effects on the probability of category 3. In the case of retirement intentions, the marginal effects on the lowest and highest categories are symmetric (but of opposite sign) and the marginal effect is practically zero in the middle category.

Sensitivity analyses

The exclusion restrictions in the model are based on our *a priori* reasoning and insights from the empirical literature summarized in Section "Related literature". We analyzed the robustness of the baseline results to the changes in the set of variables in the equations. We have also examined the influence of sample definition on the results. Thus, we estimated several variants of the baseline model. First, when job satisfaction was added as an explanatory variable for actual retirement, it was insignificant (the estimation results are not shown). This is consistent with the unconditional correlations between the variables.

Second, we estimated single-equation ordered probit models for each of the dependent variables, where we included (besides retirement intentions in the actual retirement model and job satisfaction in the intentions model) all of the variables. The only exception was future unemployment, which was, for reasons of timing, included only in the actual retirement model. In only four cases was a variable that was not included in an equation of our baseline three-equation model a statistically significant explanatory factor in the corresponding single-equation model. The indicator for having had several different occupations had a negative coefficient in the actual retirement equation (significant at the 1% level), the largest plant size category had a negative coefficient in the actual retirement model, the indicator for hazards had a negative coefficient in the retirement intentions model (significant at the 5% level), and the indicator for a retired spouse had a positive coefficient in the job satisfaction model (although significant only at 10% level). In particular, the harms, hazards, and new management practices variables were not significant in the actual retirement equation, and harms and new management practices were not significant in the retirement intentions equation. The average marginal effects of these variables on the highest categories of the dependent variables are shown in Panel B of Table 5. (For comparison, Panel A shows the marginal effects from the baseline model). The only statistically significant marginal effects are those on job satisfaction and the relationship of hazards with the probability of having frequent thoughts of retirement. This supports the use of the multivariate model to study the connection of working conditions to actual retirement.

Third, we re-estimated the three-equation model including the four significant variables in the corresponding equations (in the actual retirement equation we included all three plant size dummies). In the retirement intentions equation, the coefficients of the indicators for job satisfaction categories decreased when the hazards variable was included. Panel C of Table 5 shows the key average marginal effects. For job satisfaction, the main change compared with Table 4 is that the new management practices variable has a somewhat higher marginal effect, and the hazards variable a slightly lower one in absolute value. For retirement intentions, the hazards variable has a much higher marginal effect, 6 percent points. On the other hand, the marginal effects of harms and new management practices on retirement intentions are

Table 5
Average marginal effects of the key variables in robustness analyses.

	Average marginal effect on the probability of:		
	Job satisfaction = 3	Retirement intentions = 3	Actual retirement = 3
A. Baseline trivariate model (N = 1217)			
Harms	−0.085***	0.020***	−0.006***
	(0.028)	(0.007)	(0.002)
Hazards	−0.081***	0.019***	−0.006***
	(0.026)	(0.006)	(0.002)
New management practices	0.139***	−0.032***	0.010***
	(0.029)	(0.007)	(0.002)
B. Univariate models with all variables (N = 1217)			
Harms	−0.083***	0.035	0.005
	(0.028)	(0.026)	(0.029)
Hazards	−0.068***	0.058**	0.0001
	(0.026)	(0.024)	(0.027)
New management practices	0.154***	0.016	−0.012
	(0.025)	(0.024)	(0.026)
C. Trivariate model with additional variables (N = 1217)			
Harms	−0.084***	0.006***	−0.003***
	(0.029)	(0.002)	(0.001)
Hazards	−0.069***	0.064***	−0.011***
	(0.026)	(0.028)	(0.004)
New management practices	0.151***	−0.011***	0.005***
	(0.026)	(0.002)	(0.001)
D. Trivariate model, survey 2003 only (N = 786)			
Harms	−0.081**	0.018**	−0.002**
	(0.035)	(0.008)	(0.001)
Hazards	−0.061	0.013*	−0.002*
	(0.038)	(0.008)	(0.001)
New management practices	0.143***	−0.031***	0.004***
	(0.046)	(0.011)	(0.001)
E. Trivariate model, survey 2008 only (N = 431)			
Harms	−0.080*	0.013*	−0.009*
	(0.048)	(0.008)	(0.005)
Hazards	−0.111***	0.018**	−0.013**
	(0.043)	(0.007)	(0.005)
New management practices	0.132***	−0.022***	0.015***
	(0.041)	(0.007)	(0.005)
F. Trivariate model, 3-year follow-up (N = 344)			
Harms	−0.135**	0.055**	−0.028**
	(0.056)	(0.025)	(0.011)
Hazards	−0.044	0.018	−0.009
	(0.047)	(0.019)	(0.010)
New management practices	0.106**	−0.043**	0.022**
	(0.045)	(0.018)	(0.010)
G. Trivariate model, job switchers excluded (N = 960)			
Harms	−0.088***	0.019***	−0.006***
	(0.032)	(0.007)	(0.002)
Hazards	−0.061**	0.014**	−0.004**
	(0.031)	(0.007)	(0.002)
New management practices	0.147***	−0.032***	0.011***
	(0.035)	(0.008)	(0.003)
H. Trivariate model, retirement age rounded towards zero (N = 1044)			
Harms	−0.113***	0.029***	−0.011***
	(0.031)	(0.008)	(0.003)
Hazards	−0.070**	0.018**	−0.007**
	(0.029)	(0.007)	(0.003)
New management practices	0.140***	−0.035***	0.013***
	(0.032)	(0.008)	(0.003)

Notes: Significance: *** 1%, ** 5%, * 10%.

lower, which is due to the lower coefficients of the job satisfaction indicators. For actual retirement, the results are similar; the hazards variable has a larger (in absolute value) marginal effect, but harms and new management practices have lower ones.

Fourth, we used the two surveys separately with the specification of the baseline model. Using only the data from the QWLS 2003 survey resulted in qualitatively similar results as the combined data from 2003 and 2008, but the marginal effects of hazards were significant only at the 10% level (Panel D of Table 5). Using only the QWLS 2008 survey, all of the key variables had higher marginal effects on actual retirement (Panel E of Table 5), which may result from the (on average) shorter follow-up period.

Fifth, to study more explicitly the influence of the length of the follow-up, we kept both surveys but restricted the follow-up to three years (until 2006 for the QWLS 2003 and until 2011 for the QWLS 2008). The marginal effects of harms and new management practices on retirement intentions and actual retirement were higher than in the baseline model (Panel F of Table 5). This supports the view that due to possible future job switches or changes in work organization after the survey, the current working conditions variables have a stronger impact on retirement with a short follow-up period. However, the hazards variable was no longer significant in this sample. This may be due to the very small number of observations in the analysis (344). With a longer, 5-year, follow-up (until 2008 for the 2003 survey and 2013 for the 2008 survey) the results were closer to those with the original follow-up period (the results not shown).

Sixth, to account for the influence of job switches during the follow-up period on the relationship between working conditions and retirement behavior, we estimated the model with a sample that included only those who have not changed employer (Panel G of Table 5). The estimated marginal effects were close to the baseline estimates and therefore not affected by job switches. In any case, the share of job switchers is fairly low in our data, as reported in Section “Data”.

Seventh, we experimented with using alternative definitions of retirement. Instead of rounding retirement age to the nearest integer, we rounded it downwards to the nearest integer (e.g., 63.9 was rounded to 63). These results were qualitatively similar to our baseline estimates (Panel H of Table 5).[11]

Eighth, we examined the potential endogeneity problems. The standard tests for endogeneity or instrument validity cannot be applied in the multivariate ordered probit context. Therefore, we made a strong assumption that the dependent variables are cardinal ones and estimated the retirement equations using 2SLS. In the actual retirement equation, additional instruments for retirement intentions were the variables in the intentions equations, excluded from the retirement equation (except for job satisfaction). In the retirement intentions equation, additional instruments for job satisfaction were the variables in the job satisfaction equation excluded from the intentions equation.

In the test of endogeneity of retirement intentions in the actual retirement equation exogeneity of the intentions variable was clearly accepted (the Wooldridge’s robust score chi-squared test had a p-value of 0.11). In the test of endogeneity of job satisfaction in the retirement intentions equation, exogeneity of job satisfaction was accepted at the 5% level (p-value of 0.060). In the test of instrument relevance, the robust F-test statistics (F-test for the joint significance of the coefficients of the additional instruments in the first stage equation) were clearly above the commonly recommended value 10 (26.248 when retirement intentions were explained in the first stage and 15.116 when job satisfaction was explained). In the test of instrument validity, the overidentifying restrictions were accepted in the 2SLS estimation of the actual retirement equation (the Wooldridge’s robust score chi-squared

[11] We also used an alternative definition, 64 years, for the official retirement age. Because 63 is the most common retirement age in Finland, this changed the distribution of observations between the categories of actual retirement by increasing the share of retirements before the official age to 54% and decreasing the shares of the other categories. There are also more censored observations, as fewer individuals reach the official retirement age during the follow-up period. Again, the results were qualitatively similar to the baseline results (the results not shown). Increasing the definition further to 65 years of age would shrink the category “retirement after official age” to 9%, as very few individuals in Finland retire after 65 years of age.

test had a p-value of 0.114). The restrictions were accepted in the retirement intentions model at the 5% level (p-value of 0.051).[12] Although it may be difficult to generalize the results from the linear estimations to the trivariate ordered probit model, it seems that endogeneity is not necessarily a great concern.

We also estimated the three-equation linear model with 3SLS (the results not shown). This produced marginal effects of harms, hazards, and new management practices on the dependent variables that had the same signs as in the ordered probit model and a similar declining pattern (in absolute value) when the effects on job satisfaction, retirement intentions, and actual retirement are considered. The magnitudes are naturally not comparable to our other results.

Finally, there is no evidence for compensating wage differentials using our measures of poor working conditions (i.e., perceived harms and hazards). Cross-sectional regression of log earnings on standard human capital variables and working conditions gives the coefficients 0.018 (standard error 0.048) and 0.068 (0.044) to harms and hazards, respectively. Neither of the coefficients is significant at the 10% level.[13] Furthermore, the log earnings variable is not a significant explanatory variable for retirement intentions or actual retirement. For these reasons, we conclude that compensating wage differentials are not an important element for understanding the links between perceived working conditions and retirement behavior, at least not in our empirical context.

Conclusions

People spend much of their time at work. Consequently, it is not surprising that working conditions are an important aspect of overall well-being. Using nationally representative linked survey and register data from Finland, we found that perceived working conditions and management practices are important for retirement intentions and decisions.

Most of the empirical literature focuses on a narrow set of industries and firms. A key problem is that the motives of retirement are most likely significantly heterogeneous across organizations. Thus, the specific organizations that have garnered researchers' focus may be those in which the effects are anticipated and/or occur according to prior theoretical considerations.

Interpreted from a broader perspective, our results support the idea that inequality in perceived working conditions leads to inequality in retirement choices. Adverse working conditions are statistically significantly related to early retirement. Early retirement, on the other hand, leads to lower pensions. Our results support the idea that "soft" measures may be effective in delaying retirement and may thereby indirectly equalize post-retirement incomes. Therefore, policies to encourage investments in working conditions and the adoption of better management practices can help to increase the effective retirement age, which is an important policy target in many countries.

Some caveats are relevant to the interpretation of our results. First, the estimates are not necessarily causal. Our analysis reveals links between the variables of interest. Second, it is difficult to quantify the estimated relationships. To provide quantitative estimates, validated quantitative rather than qualitative measures of perceived working conditions are needed. Third, there may be practical challenges to implementing the "soft" measures, i.e., decreasing the level of perceived harms and hazards at work by investing resources in better working environments. Fourth, the cost-effectiveness of investments in better working conditions remains an open issue from the firms' point of view. Future research should address these issues.

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Appendix A

See Tables A1–A3

Table A1
Exact definitions of harm/hazard variables.

Harm
At least one adverse factor that affects work "very much" (includes heat, cold, vibration, draught, noise, smoke, gas and fumes, humidity, dry indoor air, dust, dirtiness of work environment, poor or glaring lighting, irritating or corrosive substances, restless work environment, repetitive, monotonous movements, difficult or uncomfortable working positions, time pressure and tight time schedules, heavy lifting, lack of space, or mildew in buildings) = 1; otherwise = 0

Hazard
At least one factor is experienced as "a distinct hazard" (includes accident risk, becoming subject to physical violence, hazards caused by chemical substances, radiation hazard, major catastrophe hazard, hazard of infectious diseases, hazard of skin diseases, cancer risk, risk of strain injuries, risk of succumbing to mental disturbance, risk of grave work exhaustion, risk of causing serious injury to others, or risk of causing serious damage to valuable equipment or product) = 1; otherwise = 0

Table A2
Descriptive statistics.

Variable	Mean	Std. dev.	Explanation
Actual retirement	2.293	0.700	Ordered indicator, values 1, 2, or 3
Retirement intention	1.977	0.829	Ordered indicator, values 1, 2, or 3

(*continued on next page*)

[12] As the univariate ordered probit models suggested that hazards could be included in the retirement intentions model, we re-estimated the linear 2SLS model including this variable directly instead of as an instrument. In this case, the overidentifying restrictions and exogeneity of job satisfaction were clearly accepted (p-values 0.428 and 0.768, respectively).

[13] Naturally, this result is subject to the standard problems of estimating compensating differentials using cross-sectional data (Hwang et al. 1992).

Table A2 (*continued*)

Variable	Mean	Std. dev.	Explanation
Job satisfaction	2.279	0.586	Ordered indicator, values 1, 2, or 3
Age	57.388	2.511	Age in full years
Female	0.574	0.495	Dummy
Basic education	0.293	0.455	Reference group
Secondary education	0.356	0.479	Dummy
Tertiary education	0.351	0.477	Dummy
Log (Earnings)	10.240	1.156	Log of annual earnings in euros
Pension accrual rate	2.249	0.855	Rate (%)
Part-time pension	0.077	0.267	Dummy
Spouse retired	0.281	0.450	Dummy
Unemployment	0.158	0.365	Dummy
Working capacity	8.054	1.453	Values from 0–10
Past job changes	0.019	0.136	Dummy
Different occupations	0.574	0.495	Dummy
Harms	0.273	0.446	Dummy
Hazards	0.394	0.489	Dummy
New management practices	0.392	0.488	Dummy
Plant size 0–9	0.214	0.411	Reference group
Plant size 10–49	0.372	0.484	Dummy
Plant size 50–249	0.267	0.443	Dummy
Plant size 250-	0.142	0.349	Dummy

Notes: N = 1217.

Table A3
Average marginal effects of key variables on categories 1 and 2 of the dependent variables.

	Average marginal effect on the probability of:		
	Job satisfaction = 1	Retirement intentions = 1	Actual retirement = 1
Harms	0.030***	−0.020***	0.003***
	(0.010)	(0.007)	(0.001)
Hazards	0.029***	−0.019***	0.003***
	(0.010)	(0.006)	(0.001)
New management practices	−0.049***	0.033***	−0.006***
	(0.011)	(0.007)	(0.001)
	Average marginal effect on the probability of:		
	Job satisfaction = 2	Retirement intentions = 2	Actual retirement = 2
Harms	0.055***	0.0004	0.003***
	(0.018)	(0.0005)	(0.001)
Hazards	0.052***	0.0004	0.003***
	(0.017)	(0.0005)	(0.001)
New management practices	−0.090***	−0.001	−0.004***
	(0.019)	(0.001)	(0.001)

Notes: Baseline trivariate model, (N = 1217). Significance: *** 1%, ** 5%, * 10%.

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