

The Impact of Occupation on Self-Rated Health: Cross-Sectional and Longitudinal Evidence from the Health and Retirement Survey

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Background. The objective of this study is to estimate occupational differences in self-rated health, both in cross-section and over time, among older individuals.

Methods. We use hierarchical linear models to estimate self-reported health as a function of 8 occupational categories and key covariates. We examine self-reported health status over 7 waves (12 years) of the Health and Retirement Study. Our study sample includes 9,586 individuals with 55,389 observations. Longest occupation is used to measure the cumulative impact of occupation, address the potential for reverse causality, and allow the inclusion of all older individuals, including those no longer working.

Results. Significant baseline differences in self-reported health by occupation are found even after accounting for demographics, health habits, economic attributes, and employment characteristics. But contrary to our hypothesis, there is no support for significant differences in slopes of health trajectories even after accounting for dropout.

Conclusions. Our findings suggest that occupation-related differences found at baseline are durable and persist as individuals age.

Key Words: Cohort study—Longitudinal study—Multilevel model—Occupation—Retirement—Self-rated health.

THE impact of occupation on health among older individuals holds particular importance both because of the cumulative effect of an individual's occupational commitment over the lifetime and because much of the decline in health occurs at older ages. Occupation can affect health through direct impacts, such as physical job conditions (e.g., manual labor, exposure to noise and heat), psychosocial job characteristics and stress, and social support (Baker, 1985; Karasek, Baker, Marxer, Ahlbom, & Theorell, 1981; Power, Manor, & Matthews, 1999; Virtanen, Vahtera, Kivimaki, Pentti, & Ferrie, 2002; Walters et al., 1996). Occupation may also affect health through indirect mechanisms via income, health insurance, prestige, and authority that are related to occupation (Bosma et al., 1997; Ferrie, Martikainen, Shipley, & Marmot, 2005; Kubzansky et al., 1997; Marmot, Bosma, Hemingway, Brunner, & Stansfeld, 1997; Stansfeld, Fuhrer, Shipley, & Marmot, 2002). Another indirect effect of occupation may be the influence of peers or workplace characteristics on health habits (e.g., outside work or smoking bans) which in turn may affect health (Cheng, Kawachi, Coakley, Schwartz, & Colditz, 2000; House, Stretcher, Metzner, & Robbins, 1986; Lusk, Kerr, & Ronis, 1995; Sakki, Knuutila, & Anttila, 1998). The Whitehall studies find that occupation has a significant impact on health, with a marked social gradient between British civil service grades and a variety of health outcomes, including coronary heart disease, self-

reported health, and emotional well-being (Bosma et al., 1997; Ferrie, Shipley, Smith, Stansfeld, & Marmot, 2002; Marmot et al., 1991, 1997; Marmot, Ryff, Bumpass, Shipley, & Marks, 1997; Stansfeld, Head, Fuhrer, Wardle, & Cattell, 2003).

How occupation affects health as individuals age is an empirical and open question. The impact on health may widen at older ages if the effects of occupational characteristics described previously (e.g., physical labor, stress, and income) build cumulatively and persist even beyond the working years. Prior research has documented a "cumulative advantage" to higher education, with disparities widening over the life span (Mirowsky & Ross, 2008). Other evidence suggests that chronic stress, which may be systematically related to occupation, results in a physiologic response, leading to overproduction of cortisol, translating into detrimental health effects that may accumulate over time (Miller & O'Callaghan, 2002; Seeman, McEwen, Singer, Albert, & Rowe, 1997; Seeman, Singer, Rowe, Horwitz, & McEwen, 1997). On the other hand, it is possible that differentials could narrow in old age if, for example, the relationship between occupation and health differs after retirement; health could improve for those in manual occupations when retirees are relieved of the physical demands or psychosocial stress of their occupations. Alternatively, the health disparities could persist, but not widen, as shown in an 11-year follow-up period to the Whitehall II study (Ferrie et al., 2002).

Despite the extant evidence, gaps remain. A prominent shortcoming of several studies is that much of the prospective evidence is based on data from two time points. Use of a measure of occupation at a point in time is problematic both because deteriorating health may have prompted occupation to change and because a previous long-term occupation may be most relevant for assessing the cumulative impact of occupation on health. Another shortcoming is that some studies, such as the Whitehall studies, do not control for education, an important factor in the production of health (Grossman, 1972). Thus, it is not possible to separate the effects of occupation and education. Finally, relatively few studies examine older individuals, yet the cumulative impact of long-term occupational exposure may be most evident among older individuals.

This study evaluates whether self-rated health differs by occupation among older individuals. We extend the existing literature in several respects. First, we measure *occupation* as the longest occupation that participants have held over their careers. This approach is better suited to capturing potential cumulative effects of a lifetime of working and permits us to include individuals who are no longer working. Longest occupation also minimizes potential endogeneity between current occupation and health, as individuals may change occupations later in their career in response to a work-related health decline. Second, we examine both the baseline and the trajectory of the relationship between longest occupation and health using seven waves of the Health and Retirement Survey over a 12-year time horizon. In analysis of health trajectories, we control for characteristics at baseline to avoid simultaneity between these characteristics (e.g., income) and health changes. Third, our study sample consists of individuals aged 50–64 years at baseline, an age range during which ill health becomes more common and the potential cumulative effect of occupation may already be present and may widen over time. Fourth, we estimate two specifications. The first allows occupation indicators to capture the full impact of occupation on health. The second specification adds controls for indirect effects of occupation (e.g., income and health insurance). Thus, the coefficients on occupation in this specification capture the direct effect of occupation and any residual unmeasured indirect effects. Finally, we include all available observations of individuals and control for attrition due to death or other dropout status. We also conduct sensitivity analyses to assess the effects of dropout on our results.

In view of the preponderance of support for an inverse gradient underlying the occupation–health relationship, we hypothesize that study participants who hold jobs in occupations with lower social standing will have higher odds of poor health than those in occupations with higher standing and that these differences will widen over time due to the growing impact of the cumulative burden.

METHODS

Data

The data used in this study are from the Health and Retirement Study (HRS), a nationally representative panel survey of community-dwelling older Americans born between 1931 and 1941 and their spouses. It was designed to investigate health and economic consequences of the transition from work to retirement (<http://hrsonline.isr.umich.edu/>). Baseline surveys, first collected in 1992 ($N = 12,652$), were conducted via face-to-face interviews; follow-up interviews, taken biennially, are completed by telephone or mail. Blacks, Hispanics, and Florida residents are oversampled.

We use seven waves of HRS data in this investigation. We define our potential sample (subjects = 11,209, observations = 63,999) as those HRS participants (including spouses) who met the following two selection criteria: (a) were 50–64 years of age or older in 1992 and (b) provided data on self-rated health (the study outcome) at a minimum of one survey wave. Requiring a minimum of only one health observation allows us to maximize sample size and avoid “completer” bias. We eliminate individuals with missing longest occupation ($N = 1,189$) and those 65 years or older. We are thus left with 9,586 subjects and 55,389 observations. The covariates data on these subjects are complete except for 126 (1.3%) subjects who had missing values on employer-provided health insurance coverage, 145 (1.5%) who had missing values on government-provided health insurance, and 288 (3.0%) who were missing on other private health insurance coverage. These values were recoded as not having coverage, so that the entire sample of 9,586 subjects was used in the analyses.

Measures

Outcome Variable.—Self-rated health. Self-rated health is measured by a single-dimension, five-category variable that provides a global assessment of well-being. Values range from excellent (1) to poor (5). Self-rated health has been shown to be a reliable, valid measure of health that has been associated with both functional decline and mortality (Idler and Benyamini, 1997; Idler, Russell, & Davis, 2000). The percentage of individuals in each category at baseline is displayed in Table 1. Our outcome variable uses data on self-rated health from all seven waves of the HRS.

As expected, as people age, the health trajectories of all occupation classes demonstrate steady deterioration. Between Waves 1 and 7, the proportion of subjects who reported excellent health decreased noticeably (22% to about 11%), the proportion of subjects who reported very good health was stable (28% compared with less than 28%), whereas the proportion of subjects who reported good, fair, and poor health increased (29%–32%, 14%–20%, and 7%–8%, respectively).

Table 1. Sample Characteristics at Baseline (N = 9,586)

Variable	Level	Percent
Self-reported health	Excellent	22.1
	Very good	28.2
	Good	28.6
	Fair	13.9
	Poor	7.1
Occupation	Professional	14.2
	Managerial	14.2
	Sales	9.4
	Clerical and administrative	15.3
	Service	14.7
	Farming and fishing	3.1
	Mechanical	12.0
	Operator	17.2
Age (years)	50–54	40.6
	55–59	39.8
	60–64	19.6
Gender	Women	49.7
	Men	50.3
Marital status	Married/partnered	78.1
	Not married/partnered	21.9
Race/ethnicity	White	72.8
	Hispanic	16.6
	Black	8.6
	Other race	2.0
Years of education ^a		12.1 (3.2)
Dropout status	Deceased during study	13.9
	Dropped out during study	16.0
	Completer	70.1
Drinking	No	85.8
	Yes	14.2
Smoking	No	73.1
	Yes	26.9
Wealth (\$)	≤\$7,500	23.1
	>\$7,500 and ≤\$25,000	16.9
	>\$25,000 and ≤\$60,000	16.1
	>\$60,000 and ≤\$150,000	18.5
	Over \$150,000	25.4
Income (\$)	≤15,000	18.6
	>15,000 and ≤30,000	22.8
	>\$30,000 and ≤45,000	19.1
	>45,000 and ≤70,000	21.4
	Over 70,000	18.1
Health insurance	Employer sponsored	70.9
	Government sponsored	10.7
	Other private	12.7
Employment status	Employed full-time	57.9
	Employed part-time	10.8
	Not employed/not retired	16.0
	Retired	15.3

Note: ^a Years of education is a continuous variable.

Explanatory Variables.—Primary explanatory variable.

Occupation is represented by eight categories of job-related affiliation. These eight mutually exclusive and exhaustive categories were created by collapsing data from 17 original categories in the HRS. They are (a) professional and technical support (reference category); (b) managerial; (c) clerical and administrative support; (d) sales; (e) mechanical, construction, and precision production; (f) service (includes private household services, protective services, food preparation, health service, and personal service); (g) operators, fabricators, and laborers; and (h) farming, forestry, and fish-

ing. Occupations are ranked from highest to lowest average education level within the occupation.

In this study, occupation corresponds to the occupational group for the job with the longest reported tenure at baseline. The advantages of this measure, as opposed to current occupation, have been discussed (Moore & Hayward, 1990). We suggest three additional advantages. First, use of longest occupation diminishes the potential for simultaneous determination of health and occupation, a problem common to studies linking current occupation to health. Second, longest occupation proxies the employment with the greatest cumulative exposure to conditions that could influence health. And third, it allows us to include in our sample HRS participants who are out of the labor force at the study baseline, which maximizes sample size and increases statistical power to detect differences among occupation classes.

Covariates.—Variables that represent demographic factors, health habits, employment, and economic attributes are used to account for potential confounding of the relationship between occupational affiliation and health. All variables are measured at the study baseline. Sample means for each independent variable can be found in Table 1.

Most variables are self-explanatory. A short explanation is provided only for those variables which are not self-explanatory. Problem drinking is measured by a dummy variable (problem drinker = 1) that indicates two or more positive responses to the CAGE questionnaire (Ewing, 1984). We control for current employment status, as a nontrivial proportion of sample members are not currently employed, but previously were. Variables measuring age at baseline control for cross-sectional age effects while the longitudinal effects are measured as changes over time (Mendes de Leon, 2007; Shaw, Krause, Liang, & Bennett, 2007). Baseline age cohort is a time-independent factor, whereas the HRS study wave was entered as a time-dependent factor. Dropout due to mortality and dropout due to other or unknown reasons were also included as covariates associated with both the intercept and the slope. Thus, we could investigate the systematic differences in the trajectories between “completers” and “noncompleters” and assess the effects of attrition due to mortality (Shaw et al., 2007).

Statistical Model and Technique

We fit a series of hierarchical linear models (Bryk & Raudenbush, 1987; Casella, 1985; Goldstein, 1987; Laird & Ware, 1982; Ware, 1985) to estimate occupation differences in baseline self-rated health and its rate of change over time. The dependent variable is treated as ordinal, and the cumulative-logit link function is used to relate the odds of being in a poorer health category, versus being in a better health category, to the primary explanatory variable (i.e., longest occupation) and to covariates. Linear change in the cumulative

odds across waves appears consistent with visual inspection of the data. Hierarchical linear models have several advantages over traditional methods of repeated-measures analysis: They use all available data on each subject, are unaffected by randomly missing data, can flexibly model time effects, and allow for estimation of both cross-sectional and longitudinal effects.

The models are described mathematically subsequently. The notation is as follows: y_{ij} refers to self-rated health of subject i at wave j , c_k is the k th category of the response variable, τ_k are threshold parameters for the dependent variable categories, β_{i0} is a subject-specific random intercept, and β_{i1} is a subject-specific random slope. The first-level equation refers to within-subject effects of wave on self-reported health. The second-level equation refers to between-subject effects and assesses how individuals' intercepts (i.e., baseline self-reported health) and slopes (i.e., rate of change in self-reported health over time) depend on longest occupation and baseline covariates. The coefficients of primary interest in this model are the α regression parameters, which, in the first level 2 equation, describe the effect of longest occupation and covariates on the intercept (or baseline health status) and, in the second level 2 equation, describe the effect of longest occupation and covariates on the slope (or linear rate of change in health status) across waves.

Level 1 Model

$$\log\left(\frac{p(y_{ij} \geq c_k)}{p(y_{ij} < c_k)}\right) = \tau_k + \beta_{i0} + \beta_{i1} \text{ wave}.$$

Level 2 Model

$$\beta_{i0} = \alpha_{00} + \sum_{d=1}^7 \alpha_{0d} \text{occupation}_d + \sum \alpha_{0h} \text{covariate}_h + \varepsilon_{i0},$$

$$\beta_{i1} = \alpha_{10} + \sum_{d=1}^7 \alpha_{1d} \text{occupation}_d + \sum \alpha_{1h} \text{covariate}_h + \varepsilon_{i1}.$$

The hierarchical linear model is estimated with two separate specifications. The reduced model allows the coefficient on occupation to capture both the direct and indirect effects of occupation. The expanded model controls for some of the indirect effects, thus allowing occupation to capture the direct effect and any unmeasured indirect impact. The reduced model includes longest occupation and controls for age cohort, demographic variables (gender, race, education, marital status), and dropout status. Our expanded model additionally controls for health habits (problem drinking and tobacco use), economic status variables (categorized total nonhousing wealth and income), and employment controls (not working, working part-time, covered by employment-sponsored health insurance). In both models, the reference categories for each variable are the *a priori*-assumed lowest risk categories. Because we include

only baseline characteristics as covariates, the coefficients on longest occupation can be understood to capture the impact of occupation, conditional on baseline characteristics.

The models are estimated with Version 4.21 of the MPlus software. Model fit is assessed using Akaike Information Criteria. Odds ratios (ORs) and associated 95% confidence intervals (CIs) indicate the magnitude and statistical significance of the relationship between explanatory variables and self-rated health.

RESULTS

In Table 2, we present ORs and 95% CIs for the effect of occupation and covariates on the intercept and slope of self-rated health. The table contains results of reduced and full-model specifications.

The results of the reduced model are presented in columns 1 and 2. They suggest that, compared with the assumed lowest risk group (professional workers), participants in service (OR = 2.33, 95% CI: 1.73–3.13), farming/fishing (OR = 1.92, 95% CI: 1.18–3.12), mechanical (OR = 1.87, 95% CI: 1.36–2.57), and operator (OR = 2.59, 95% CI: 1.92–3.49) have significantly higher odds of reporting in a poorer health category at baseline. In contrast, the slopes of all occupation categories do not differ significantly from the slope of professionals. The results from the full model are presented in columns 3 and 4. In this latter specification (adding additional controls for health habits, employment, and economic characteristics), the cross-sectional associations between occupation and health are typically smaller in magnitude. Also, one of the significant associations found in the reduced model (farming and fishing) is no longer significant at the 5% level. The smaller magnitude of the intercept coefficients is expected, as the additional variables capture some of the effect that would otherwise be attributed to occupation. The health trajectories of all occupation classes demonstrate steady deterioration over the study period ($p < .001$).

Sensitivity logistic regression analysis did not show significant differences in mortality among occupation categories after controlling for the other predictor variables. Moreover, treating death as a sixth, most severe, health category showed similar and slightly stronger occupation effects as our primary analysis. Finally, tests of the interactions between occupation and mortality variables were not significant, suggesting that the likelihood of bias in our results due to mortality-related attrition is small. Results from sensitivity analyses are available on request.

DISCUSSION

Measuring *occupation* as the longest occupation is an approach that helps to capture the accumulation of the potential health impact of work and minimize the endogeneity between occupation and health among older adults. We find

Table 2. Odds Ratios and 95% Confidence Intervals for the Effects of Longest Occupation and Covariates on Intercept (Baseline) and Slope (Rate of Change) of Self-Reported Health (N = 9,586)

Variable	Category	Reduced Model (Akaike Information Criteria [AIC] = 129,438.7)		Full Model (AIC = 128,785.8)	
		Intercept (column 1)	Slope (column 2)	Intercept (column 3)	Slope (column 4)
Occupation	Managerial	0.88 (0.67, 1.14)	1.00 (0.96, 1.04)	0.93 (0.72, 1.20)	0.99 (0.95, 1.04)
	Sales	1.33 (0.97, 1.81)	0.96 (0.92, 1.01)	1.01 (0.76, 1.36)	0.97 (0.92, 1.02)
	Clerical and administrative	1.08 (0.83, 1.42)	0.97 (0.93, 1.02)	0.95 (0.74, 1.23)	0.97 (0.93, 1.01)
	Service	2.33 (1.73, 3.13) ^a	0.96 (0.91, 1.01)	1.33 (1.00, 1.77) ^a	0.98 (0.93, 1.02)
	Farming and fishing	1.92 (1.18, 3.12) ^a	0.94 (0.87, 1.02)	1.44 (0.91, 2.29)	0.94 (0.87, 1.02)
	Mechanical	1.87 (1.36, 2.57) ^a	0.97 (0.92, 1.02)	1.39 (1.03, 1.88) ^a	0.97 (0.92, 1.02)
	Operator	2.59 (1.92, 3.49) ^a	0.96 (0.91, 1.01)	1.58 (1.19, 2.10) ^a	0.97 (0.92, 1.02)
Age (years)	Professional	1.00	1.00	1.00	1.00
	55–59	1.34 (1.14, 1.56) ^a	0.95 (0.93, 0.97) ^a	1.19 (1.03, 1.39) ^a	0.96 (0.94, 0.99) ^a
	60–64	1.34 (1.10, 1.62) ^a	0.98 (0.95, 1.02)	0.90 (0.74, 1.10)	1.02 (0.99, 1.06)
Gender	50–54	1.00	1.00	1.00	1.00
	Men	0.89 (0.76, 1.05)	1.03 (1.01, 1.06) ^a	0.97 (0.82, 1.13)	1.03 (1.00, 1.05)
Marital status	Women	1.00	1.00	1.00	1.00
	Not married	1.89 (1.56, 2.27) ^a	1.00 (0.97, 1.03)	0.91 (0.75, 1.11)	1.02 (0.98, 1.05)
Race/ethnicity	Married	1.00	1.00	1.00	1.00
	Hispanic	1.83 (1.38, 2.42) ^a	0.97 (0.93, 1.02)	1.41 (1.08, 1.84) ^a	0.99 (0.94, 1.04)
	Black	2.51 (2.06, 3.06) ^a	0.95 (0.92, 0.98) ^a	1.88 (1.55, 2.27) ^a	0.96 (0.93, 1.00) ^a
	Other race	1.76 (1.01, 3.06) ^a	1.04 (0.96, 1.12)	1.39 (0.83, 2.33)	1.06 (0.98, 1.14)
Education	White	1.00	1.00	1.00	1.00
	Dropout status	0.76 (0.73, 0.78) ^a	1.01 (1.00, 1.01)	0.84 (0.81, 0.86) ^a	1.00 (1.00, 1.01)
Dropout status	Deceased	6.28 (4.87, 8.11) ^a	1.30 (1.22, 1.38) ^a	3.41 (2.69, 4.34) ^a	1.33 (1.25, 1.42) ^a
	Dropout	0.93 (0.75, 1.14)	1.03 (0.99, 1.08)	0.97 (0.80, 1.19)	1.03 (0.98, 1.08)
	Completer	1.00	1.00	1.00	1.00

Variable	Category	Reduced Model (Akaike Information Criteria [AIC] = 129,438.7)		Full Model (AIC = 128,785.8)	
		Intercept (column 1)	Slope (column 2)	Intercept (column 3)	Slope (column 4)
Drinking	Yes			1.42 (1.16, 1.74) ^a	0.99 (0.95, 1.03)
	No			1.00	1.00
Smoking	Yes			1.15 (0.99, 1.35)	1.08 (1.05, 1.11) ^a
	No			1.00	1.00
Wealth (\$)	Under 7,500			2.66 (2.11, 3.34) ^a	0.97 (0.94, 1.01)
	7,501–25,000			2.06 (1.64, 2.58) ^a	0.96 (0.93, 1.00) ^a
	25,001–60,000			1.34 (1.08, 1.66) ^a	0.98 (0.95, 1.02)
	60,001–1,50,000			0.95 (0.78, 1.17)	1.00 (0.96, 1.03)
	Over 1,50,000			1.00	1.00
Income (\$)	Under 15,000			4.83 (3.54, 6.58) ^a	0.93 (0.88, 0.98) ^a
	15,001–30,000			2.32 (1.80, 2.98) ^a	1.01 (0.97, 1.05)
	30,001–45,000			1.70 (1.34, 2.16) ^a	1.00 (0.97, 1.04)
	45,001–70,000			1.57 (1.26, 1.95) ^a	0.97 (0.94, 1.01)
	Over 70,000			1.00	1.00
Employment insurance	No			1.03 (0.85, 1.24)	1.01 (0.98, 1.04)
	Yes			1.00	1.00
Government insurance	No			0.18 (0.14, 0.24) ^a	1.17 (1.12, 1.22) ^a
	Yes			1.00	1.00
Private insurance	No			0.93 (0.75, 1.14)	1.03 (0.99, 1.07)
	Yes			1.00	1.00
Workforce status	Part-time			0.92 (0.74, 1.15)	1.01 (0.98, 1.05)
	Retired			4.28 (3.42, 5.36) ^a	0.91 (0.87, 0.94) ^a
	Not employed/retired			2.60 (2.10, 3.21) ^a	0.92 (0.89, 0.96) ^a
	Full-time			1.00	1.00

Note: ^aStatistically significant effect at 0.05 level.

substantial variation in self-rated health across longest occupation at baseline in this group of middle-age and older adults. As expected, when we control for additional variables that are related to occupations (e.g., income), the magnitudes of the intercept coefficients on occupation decline.

Contrary to our expectations, we do not find any significant impact of occupation on changes in health over time, nor do the slope coefficients on occupation change as additional variables are included. That is, the occupation-related differences found at baseline are durable and persist as

individuals age, even though some individuals leave their longest occupations because of job changes or retirement. Although the trajectory results show that disparities do not widen over the study period, it is striking that the occupation-related disparities in health persist, given that a considerable share (40%) of sample members transition out of their longest occupation over the study period. Across the survey waves, at baseline 58% of the sample has a current occupation that is the same as their longest occupation, declining to 34% by Wave 4 and 19% by Wave 7. Our results are consistent with a life course view of social determinants of health; health problems accumulate over the lifetime and are systematically different by occupation, but differences appear to be stable in older ages.

Our results are also consistent with, and extend, prior research that finds baseline differences that do not widen over time. For example, Lindenberger and Baltes (1997) find that while there are baseline differences in intellectual capacity by socioeconomic status (SES), the decline in this capacity is equal by SES, thus preserving the baseline differences across age categories. An 11-year follow-up of the Whitehall II study (Ferrie et al., 2002; Marmot et al., 1997) found no change in the physical health disparity observed at baseline, although mental health disparities widened.

By studying civil servants, the Whitehall studies provide particularly compelling evidence that a gradient exists even among a relatively advantaged population. However, it is also important to examine occupation-related differences in health in a representative population and within the age range when potential cumulative effects of employment on health may be more likely to be experienced. Our study examines differences across a representative middle age and older population that reflects the range of occupational experiences in the United States. We also show that these differences persist in the full model that controls for education, income, and wealth, indicating that occupation has a separate effect on health beyond traditional measures of SES. Other studies, including Whitehall, do not control for all of these factors.

There are limitations to this study. As with all observational studies, there is the possibility of omitted variable bias. Nevertheless, the advantages of the HRS, particularly its national sample of older workers and breadth of covariate data, may outweigh shortcomings related to the omission of potentially relevant factors. Another matter is heterogeneity within occupation classes. A rather small set of occupations is a practical requirement for the type of analysis undertaken in this study. However, with such a defined set, there is likely heterogeneity within the occupation categories. Moreover, with a small number of occupational classes in which at least three categories (sales, clerical and administrative, and service) potentially contain a mix of white- and blue-collar workers, our results may reproduce the finding that industrial workers have poorer health than workers in white-collar occupations.

Another possible limitation is that we do not consider time-dependent covariates or mediators such as changes in health habits or employment status. However, changes in health habits, retirement, and other employment status transitions can be as much a consequence of change in health as a mediator of occupation effects on health. For example, highly paid occupations could allow workers to retire relatively early, which in turn could improve health, whereas physically demanding occupations could lead to declining health, which in turn can lead to early retirement. Disentangling these two possible causal chains is a challenging statistical modeling problem that is beyond the scope of this short research note. Our analyses are focused on the specific question of occupation effects when controlling for baseline levels of potential confounding factors, thus avoiding the endogeneity of changes in predictors and health.

We took multiple steps to address dropout, which strengthen the findings of the study. Use of longest occupation allows us to examine nonworking individuals. Individuals with at least one observation are included in our sample even if they drop out or die later in the study period. In so doing, we maximize statistical power and minimize the potential of bias due to sample restriction. In addition, we control for dropout due to death and other reasons by including additional indicator variables in our intercept and slope predictions. Thus, we are able to ascertain that subjects who drop out have poorer health at baseline and faster deterioration over time. Sensitivity analyses show that our results are unlikely to be significantly affected by dropout and death.

In conclusion, our analysis assessing the relationship between longest occupation and health shows that occupation has an effect on health, even after controlling for measures relating to occupation, including education, income, and health insurance. The lasting effect of longest occupation on health even after change in job or retirement indicates the durability of the impact of longest occupation. The persistence of occupation-related health disparities, even when people are no longer working, suggests that longest occupation is a phenomenon worthy of further study.

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