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Abstract

In general, retirement is seen as a pure labor supply phenomenon, but firms can have strong incentives to send expensive older workers into retirement. Based on the seniority wage model developed by Lazear (1979), we discuss steep seniority wage profiles as incentives for firms to dismiss older workers before retirement. Conditional on individual retirement incentives, e.g., social security wealth or health status, the steepness of the wage profile will have different incentives for workers as compared to firms when it comes to the retirement date. Using an instrumental variable approach to account for selection of workers in our firms and for reverse causality, we find that firms with higher labor costs for older workers are associated with lower job exit age.

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1 Introduction

Retirement decisions are typically seen as a labor supply phenomenon and most scholars have focused on individual retirement incentives. There is a large literature on the influence of health (e.g., Currie and Madrian (1999)), individual productivity (Burtless, 2013), working conditions (Schnalzenberger *et al.*, 2014), the generosity of social security systems in terms of pensions (Van Soest and Vonkova, 2013) or retirement age regulations (Mastrobuoni (2009) or Staubli and Zweimüller (2013)).

In spite of this research concentration on – voluntary – labor supply effects, surveys often reveal that a large proportion of workers state they did not retire voluntarily so early (Dorn and Sousa-Poza (2010) using ISSP data or Marmot *et al.* (2004) for England). Differentiating between voluntary and involuntary retirement may not be completely clear for survey respondents, when it comes to the potential role of firms. In this paper, we want to explore the role of labor demand in retirement outcomes. Using high-quality administrative data for the universe of Austrian workers and firms, we investigate whether a particularly steep seniority wage profile in a firm leads to a markedly lower retirement age of its workforce. We identify the role of a firm's wage structure by instrumenting with labor market shocks a decade ago.

Looking at the role of firms in retirement decisions is important in several respects: Leaving out labor demand in retirement processes is unwise given the big policy problem of early retirement rates across Europe; in particular, investigating the role of wage costs and wages schedules opens up important policy channels. Moreover, distinguishing voluntary from involuntary retirement may shed light on well-being in retirement and may also help explaining the retirement-consumption puzzle (Smith, 2006).

Previous research on labor demand effects in retirement has been scarce. Bartel and Sicherman (1993), Bello and Galasso (2014) and Bellmann and Janik (2010) explore the role of technology and trade shocks on retirement. The role of seniority wage profiles in retirement decisions has not been studied before. Hakola and Uusitalo (2005) and Hallberg (2011) are related to our topic, as they study the impact of non-wage labor costs on retirement age. Hakola and Uusitalo (2005) analyze the introduction of an experience-rating of early retirement benefits in Finland and find a significant reduction of early job exits of older workers. This implies a firm's impact on retirement, as workers need to be laid off before obtaining early retirement benefits at all. For Sweden, Hallberg (2011) shows how exogenous variation in non-wage costs affects early retirement probabilities.[1]

Firms are indifferent with regard to the retirement age of their workers if age-wage profiles correspond to age-productivity profiles. This is not the case otherwise, firm incentives to lay off older workers arise, whenever age-wage profiles exceed age-productivity profiles. Our theoretical approach is based on an implicit contract model (Lazear (1979) and Lazear (1983)). In order to discourage employee shirking and malfeasance, a firm and its workers may adhere to an implicit contract, whereby workers' wages are below their marginal product at the beginning and higher at the end of their career with the firm. While such a contract eliminates the shirking incentives of the workers, it opens up moral hazard problems from the side of the firm: A steep seniority wage schedule - on the one hand - stimulates workers to stay longer with the firm; on the other hand, firms may want to terminate the contract prematurely to reduce wage costs. Lazear (1979) solves this problem by referring to reputation costs, reneging firms will have.

Under what conditions will steep seniority wage profiles induce firms to send workers into early retirement? At first, reputation costs are less severe, once workers are not fully informed or aware of firms' opportunistic behavior or when there is no infinite horizon of the firm. Moreover, the possibility of severance payments and actuarially unfair social security pensions may ameliorate such early retirement transitions by reducing the costs to workers.[2]

It turns out that the steepness of a firm's seniority wage profile relative to productivity development is the key to differentiate between firms' and workers' decisions for early retirement. Ceteris paribus, a steeper profile will increase the incentive for the firm, but

[1]Other studies implicitly related to the wage structure look at firing penalties or subsidies of older workers, e.g., Behaghel *et al.* (2008) or Schnalzenberger and Winter-Ebmer (2009).

[2]See also Hutchens (1999), who models the firm's impact on early retirement decisions of its workers by emphasizing the role of the social security system, effectively subsidizing workforce reductions similar to non-experience-rated unemployment insurance.

at the same time, individual retirement incentives will decrease due to higher expected social security benefits induced by higher wages close to retirement.[3] A firm effect on individual retirement can only be separated from the individual retirement decision if individual incentives are addressed properly within the empirical framework.

2 Institutional background and data

Compared to other OECD countries, Austria shows a relatively low effective retirement age and high net replacement rates. The average pension in Austria for men is 76.6 percent of an average worker's earnings (compared to the total OECD average of 54.5 percent, values for 2012). With a statutory retirement age of 65, Austrian men retire on average at age 60.6 (value for 2014), taking advantage of early retirement options due to long periods of social security contributions and disability pensions.

Hofer and Koman (2006) conclude that the low labor force participation among the elderly can be attributed to some extent to disincentives of the Austrian pensions system, which provides too many incentives to retire early. Hanappi (2012) computed the social security wealth and accrual rates for Austria. He finds that the social security wealth peaks at age 63 for men, hence creating strong disincentives to work longer than 63.

The generosity of the Austrian pension system also appears in other relevant dimensions: In order to smooth the transition into retirement, there are old-age part-time schemes for older employees, where working time reductions of elderly workers are subsidized – often leading to early retirement altogether (Graf *et al.*, 2011). Special severance payments (golden handshakes) paid to the worker in case of leaving the job bring along tax advantages to the employer and the employee.

For our analysis we use data from the Austrian Social Security Database (ASSD) containing comprehensive information on all employment and income data necessary to calculate

[3] The calculation of pension payments provides an additional incentive for workers to stay with the firm in case of a steep seniority wage profile. At this time, pensions were not calculated out of the sum of lifetime contributions, but out of the best 15 years of contributions. Higher wages at the end of the career, i.e., higher contributions, would thus increase the incentive to hold on to the job.

pensions – and the social security wealth at each point in time. It covers the universe of Austrian workers together with firm identifiers, which allows the construction of a firm's work force in detail from 1971 to 2012 (Zweimüller *et al.*, 2009). We currently use all male[4] blue-collar and white-collar workers aged 57 to 65 who retired in the period 2000 to 2009 and worked in private sector firms.[5] We exclude workers from small firms with less than 15 workers and from firms without workers below age 25, because no sensible seniority wage schedule can be constructed in such firms.

When we define our "retirement age" we do not explicitly look at the age at actual retirement, but consider the age of the worker when he exits from the last job before retirement – and restrict ourselves to a maximum time between job exit and retirement of 2 years. In fact, this job exit age is the more relevant variable of interest because workers might become unemployed and receive unemployment benefits for 52 weeks before retiring and terminating a job in such a pre-retirement phase could, thus, be a firm strategy (see also Staubli and Zweimüller (2013)). We also condition on a firm tenure of at least 2 years, leaving us with approximately $41,300$ blue-collar and $45,100$ white-collar retirees. Table 1 provides some descriptive statistics. Compared to white-collar workers, blue-collar workers retire on average one year earlier, have a higher incidence of disability, but a lower incidence of phased retirement and golden handshakes. They also have lower tenure and social security wealth at age 55.

While some studies (Hofer and Koman, 2006) claim, that – due to an actuarially unfair social security system, where staying longer in the workforce is financially punished – Austrians retire the first day possible, we do see large variations in retirement ages. Figure 1 respectively shows boxplots for the distribution of job exit ages for blue-collar workers in the largest firms in the most relevant sectors. The upper-left panel, for example, shows the job exit age distribution of the 21 largest firms in the steel industry, where firm size is measured by the number of retirement transitions in that firm between 2000 and 2009. These firms are relatively homogeneous, but still considerable firm-specific variation in

[4]We do not use female workers for the time being, because of missing working time information.

[5]We do not go beyond the year 2009 in our analysis to exclude any potential impact of the economic crisis on retirement.

the job exit patterns can be observed. This variation is also very pronounced in the transport or machine building sector. As these firms in each sector are comparable in size and production technologies, it is doubtful whether these patterns are exclusively created by a selection of workers in firms. Instead, at least some variation in retirement behavior across firms is probably due to different firm policies with respect to retirement. Figure 2 is the equivalent picture for white-collar workers, where firm-specific differences are similarly pronounced compared to blue-collars (e.g., in the wholesale and energy supply sector).

3 Empirical strategy

The identification of firms with higher incentives to lay off older workers is pivotal. As argued, such firm incentives depend on wage costs for older workers in particular. In the following, we will describe how we construct seniority wage profiles and how we proxy for productivity. Moreover, we have to control for individual retirement incentives arising from social security considerations. The identification of the impact of the seniority wage profile on retirement entry is achieved via an instrumental variables strategy: to control for reverse causation problems associated with hiring and firing processes of a firm, we use labor market conditions in the past as an instrument.

3.1 Constructing the wage gradient

We define the wage gradient as an incentive measure for firms to dismiss older workers. Clearly, the "true" wage gradient would be the difference between wage and productivity profile by seniority. There are two possibilities to construct a seniority wage profile. First, individual wage profiles for each worker could be calculated using the wage history starting with the entry into the firm. Second, a cross-sectional wage profile for wages paid in a firm at a specific point in time uses only current wages. While the first approach is closer to a Lazear-type contract, we use the cross-sectional wage structure which corresponds to a mark-to-market valuation disregarding historical costs. This approach is closer to the idea of substituting expensive elderly workers with young ones: the actual wages paid to these workers some 20 years ago would not matter much, but current replacement costs, i.e., the wages of young workers, will matter.

We look at wages paid to workers from age 15 to a maximum of 65 years and construct a cross-sectional wage profile for each firm and each year (1997 to 2009) separately. As age productivity profiles are not observable, we use the corresponding industry wage profile as a proxy. It is clear that an industry wage profile does not reflect productivity and we are not able to construct the "true" wage gradient. However, we can derive similar firm incentives by also looking at the differences between firm and industry wage profiles. First, the industry profile is composed of the direct competitors who share similar technologies, are of comparable size and share the same minimum collectively bargained wages.[6] A steeper firm wage profile relative to the industry wage profile – a positive wage gradient – is associated with increasing costs for firms (also relative with respect to costs for substitutes), and because of certain homogeneity with respect to collectively bargained wages and technology, a positive wage gradient is likely to reflect a seniority wage scheme rather than a pure marginal product payment scheme. Nevertheless, using the industry wage profile instead of productivity will incorporate potential measurement errors of the true magnitude of the firm incentive to dismiss older workers. To tackle this problem we add a person fixed-effect as an additional covariate in order to control for the individual productivity of a worker. These person fixed-effects are derived following Abowd *et al.* (1999), where wages are decomposed into firm- and worker-specific components.[7] Moreover, an instrumental variables strategy – discussed below – will also take care of measurement error problems.

Figure 3 provides a schematic representation of the wage gradients. Assume that the black solid line represents the firm wage profile of one particular firm and the dotted blue line is the corresponding industry wage profile. We propose two comparable definitions of

[6] In fact, within-industry wage profile heterogeneity across firms comes from firm-specific wage settings above the collectively bargained wages.

[7] Worker fixed-effects are identified within each set of workers and firms that is connected by individual workers moving between different firms. Since the majority of workers in our sample is observed in more than one firm- these effects are to be well-identified. For details on the decomposition method see the Appendix.

the wage gradient. We compute the wage gradient within a regressions-based framework and we regress the difference between firm and industry wage (Δw) on age for each firm and year separately. The resulting age coefficient for each firm can be interpreted as the wage gradient. A positive coefficient (β_{ij}) means that the firm wage profile is steeper than the industry wage profile and higher coefficients are associated with higher incentives.

Similarly, we test our results with an alternative wage gradient definition which is simply the difference between firm and industry wage profile at ages 55 to 65 (Δw^{old}) subtracted by the difference at ages 15 to 25 (Δw^{young}) in a given year. If this value is positive, then the firm wage profile is steeper than the industry wage profile and the firm is associated with a higher incentive for layoff. Note that the wage gradients measure the deviations between firm and industry wage profiles in euros.[8] The main difference between these two definitions is the time period. A 1€ increase of the wage gradient reflects an annual increase of firm wages over industry wages, whereas a 1€ increase of the alternative wage gradient implies that firm wages increase relative to industry wages by 1€ over 40 years. A more detailed description of the wage gradient calculations can be found in the Appendix (Section 7).

3.2 Identification

The identification of the wage gradient impact on retirement age is plagued by potential endogeneity problems: Quite automatically, the measured steepness of the wage gradient may depend on the amount and structure of hiring and firing patterns in the firm. In particular, the firing of older, highly-paid workers early on may lead to a flat measured wage gradient in a firm – thus, reverse causality. Moreover, wage gradients in a firm as well as a particularly low "firm retirement age" might initiate the specific self-selection of workers. Due to these reasons – and also to counter measurement error problems in the wage gradient with respect to productivity profiles – we suggest an instrumental variables

[8] Alternatively, we also test a third version where we only focus at the difference between firm and industry wages for older wages. Referring to Figure 3, the wage gradient would only be Δw^{old}. We do not find significant differences in the results compared to the other two definitions. Results are available upon request.

strategy.

We suggest to instrument the wage gradients by past local labor market conditions. It has been shown, that wages depend on the business cycle and higher unemployment rates enable firms to pay lower wages (e.g., Bils (1985), Blanchflower and Oswald (1994), Gregg *et al.* (2014)). Empirical evidence also suggests that wages of job movers or entrants are pro-cyclical, whereas wages of job stayers do not react much to the business cycle (Haefke *et al.* (2013), Devereux and Hart (2006)). As a consequence, past labor market conditions should have a certain explanatory power in the determination of the wage structure of the firm today (Beaudry and DiNardo (1991), Hagedorn and Manovskii (2013)), because individual wage profiles are shaped by idiosyncracies at the time of job entry.

We use local unemployment rates on the district level for prime-age workers (25-45 years old) and calculate them 10 years before workers' job exit.[9] We expect, ceteris paribus, higher local unemployment rates 10 years ago to reduce the current cross-sectional wage gradient, as firms in districts with higher unemployment rates may have been able to hire more cheaply as compared to firms in districts with better local labor market conditions. These relatively better hiring conditions in the past will thus reproduce themselves into relatively low wages of the current older workforce.

The local average treatment effect is the effect of a 1€ increase of the wage gradient on job exit age for those who leave the firm because of higher wage gradients due to the past local labor market situation. This should be completely unrelated to any unobserved firm characteristics, worker selection or unobserved individual propensity to retire today. It is particularly noteworthy, that our instrument is not firm-specific, as any firm-specific characteristic might be related to firm personnel policies in general. Current local labor market conditions and retirement behavior are related, though. In case of worsening labor market conditions, retirement becomes more attractive to older workers (Coile and Levine, 2007). If local unemployment rates are persistent within districts, past unemployment rates may also capture a potential direct effect through serial correlation. To strengthen the validity of our instrument even further, we also allow current local labor

[9]We also present results for a time lag of 15 years as an additional robustness check.

market conditions to directly affect individual retirement incentives and control for local unemployment rates 1 and 5 years before workers' job exit.[10]

We specify our empirical model in the following way:

$$JE_{nij} = \alpha_0 + \alpha_1 * (WageGrad_{ij}) + \alpha_2 * SSW_{nij} + \alpha^3 * X_{nij} + \epsilon_{nij} \quad (1)$$

where JE_{nij} is the job exit age of worker n in firm i of industry j, and $WageGrad_{ij}$ corresponds to one of the two firm incentive measures. Equation 1 relates individual worker's job exit age to the firm-level seniority wage gradient and individual social security wealth SSW_{nij}. The vector X_{nij} contains further individual characteristics measured at age 55, i.e., collected social insurance months, job tenure, experience, firm size, the number of sickness and employment days. We further control for job exit year, region and industry fixed-effects as well as the personal fixed-effect from the wage decomposition. Finally, local unemployment rates 1 and 5 years ago are included. The key parameter of interest is α_1, which measures the effect of the wage gradient on job exit age. From theory we expect α_1 to be negative, because a greater gap between firm and industry wage profiles should increase firm incentives and consequently lower the job exit age of their workers. Conditional on the social security wealth, α_1 should only capture firm effects.

The first stage is:

$$WageGrad_{ij} = \gamma_0 + \gamma_1 * UR_{t-10} + \gamma_2 * SSW_{nij} + \gamma_3 * X_{nij} + \mu_{nij} \quad (2)$$

with UR_{t-10} as the local unemployment rate for prime age workers 10 years before job exit.

The validity of the instruments requires $Cov(UR_{t-10}, \epsilon_{nij}) = 0$. We cluster the standard errors on the district level to allow for within-district correlations of the observations.[11]

[10] One may argue that past local labor market conditions may induce plant closures or mass layoffs in the future (e.g., plant closures). To rule out such competing explanations, we also tested whether past unemployment rates directly affect plant closures or mass layoffs and find no significant effects. Results available upon request.

[11] In a further robustness check we also allow observations in the same firm to be correlated, but calculated standard errors remained unchanged.

4 Results

At first, we briefly discuss our results from OLS regressions and the first stage results. Section 4.2 provides our main estimation results on job exit age for both definitions of the wage gradient and blue- and white-collar workers separately. Section 4.3 expands our analysis to alternative outcome variables like golden handshake, etc.

4.1 OLS and first stage results

Tables 2 and 3 summarize the estimation results for OLS, first stage and 2SLS for blue-collar and white-collar workers respectively. The OLS coefficient of the wage gradient is insignificant and very small in size for both types of workers. As discussed before, these coefficients are likely to be biased by reverse causality of job exit age and the wage gradient. If the true causal effect is negative, the reverse causality issue biases the coefficient downwards and, in this case, even to zero.

Tables 2 and 3 also report the coefficients for the main covariates. Here, the social security wealth and the collected social security contributions months are of particular interest. As expected, higher social security wealth at age 55 reduces job exit age. The amount of social security contribution months is also negative for white-collar workers, but positive for blue-collar workers. Moreover, more experienced workers tend to leave much earlier. That might indicate, that the amount of working years is relevant for individual retirement incentives, whereas a higher number of non-working contribution years (months of unemployment, unpaid leave for training, parental leave,...) tend to increase overall working life for blue-collars, as these years hardly contribute to the expected social security wealth. Firm tenure, firm size and the number of weeks on sick leave or out of work at age 55 do not significantly influence worker's job exit age. The recent unemployment rates only seem to be relevant for white-collar workers. For both groups, persons with a high personal fixed effect, i.e., high productivity tend to retire later.

The second columns of Tables 2 and 3 report the results for the first stage regressions. In line with our expectations, a higher local unemployment rate 10 years before job exit reduces the wage gradients significantly for both blue-collar and white-collar workers; the quantitative effect on the wage gradient is much higher for white-collar workers, as these workers have larger career opportunities in general. The corresponding F-test for weak instruments yields values between 23 to 27, well above conventional critical values for weak instrument problems. The negative coefficient for local unemployment rates shows that the gap between firm and industry profile narrows with worse labor market conditions. This indicates that firms are able to hire workers for a lower wage, given workers' productivity. On the other hand, selection effects might hamper the analysis, when more productive workers are hired in such worse labor market conditions. As we are able to control for personal fixed-effects in the regressions, such selection effects should play no role.

Labor market conditions one year before job exit are positively correlated with the wage gradient, mainly due to new hirings. A higher recent unemployment rate increases the wage gradient, as new and typically young workers can be hired for a lower wage, which increases the steepness of the wage profile in return.

4.2 Job exit age

The third columns of Tables 2 and 3 present our causal effects from the 2SLS estimations. For both blue-collar and white-collar workers, the coefficient of the wage gradient becomes negative as expected and statistically significant. For blue-collar workers, a 1€ increase of the wage gradient leads to a 0.926 year lower job exit age. A one standard deviation increase of the steepness of the wage gradient would thus lead to a 5.8-months reduction of the job exit age, which is 59.8 years on average. For white-collar workers, the coefficient of the wage gradient appears to be much smaller. In terms of standard deviations, the results are fairly comparable, though: A one standard deviation increase of the wage gradient reduces the job exit age of white-collar workers by approximately 4.9 months. Note, that all other estimated coefficients do not change from OLS to 2SLS, which is reassuring.

Table 4 summarizes robustness checks using the other definition of the wage gradient, as well as different time lags of the instrument. The latter serves as an additional robustness check for the validity of the instrument, on the one hand, as higher time lags of unemployment rates should be less relevant for retirement intentions today and, on the other hand, different time lags affect different workers within the firm. A comparison of results across time lags enables us to exclude potential selection effects of workers into firms affected by these labor market conditions. Overall, it turns out that our results remain very stable and robust. For blue-collar workers, a standard deviation increase of the *alternative wage gradient*[12] decreases job exit age by approximately 6.2 months, compared to 5.8 months of the baseline wage gradient. The effects become somewhat smaller when using 15 years as the relevant time lag of the instrument (approximately -4.8 months per standard deviation) but remain significant. The picture for white-collar workers is very similar: a one standard deviation increase of the *alternative wage gradient* decreases job exit age by 4.5 months, compared to 4.9 months in the basis. Using a higher time lag for the instruments also yields quantitatively very comparable results.

Workers' selection into firms might depend on local labor market prospects as well, and one could argue that our instrument could have changed the selection of workers into firms. We tackle this potential concern by focusing on workers who entered the firm already well before the date we measure the local unemployment rates. This restriction avoids the problem that the wages of the workers in our sample could be directly affected by the instrument. The results are reported in Table 5. In this exercise, we lose observations, which leads to a loss in the precision of our estimates. Nevertheless, our results are very robust to this test; the point estimates are very similar to the general case and are also statistically significant in three out of four cases. We conclude that a potential selection of workers into firms is negligible and does not systematically affect our results

[12]Note the different dimension of the alternative wage gradient definition.

4.3 Alternative outcomes

In this section, we report estimation results for alternative outcomes. Table 6 summarizes estimation results for the probability of leaving via disability pension or a publicly subsidized phased retirement scheme and the probability of receiving a golden handshake.

We do not find any significant effects of wage gradients on the probability that the worker receives a disability pension. Access to disability pensions requires severe health problems which need to be confirmed by firm-independent public health officers at the request of the worker. This result fits well into the general pattern we found so far: a steep wage gradient points towards a role of the firm in retirement behavior; as the firm cannot influence entry into disability pensions, there is no causal impact of the wage gradient.

Remarkably, we also do not find any effects on the probability of entering into a phased retirement scheme. Firms have a direct impact on the access to these job exit programs, as their approval is necessary. Subsidies for such phased retirement require mandatory new hirings of a younger worker as a substitute; potential cost saving effects might thus be offset for the firm.[13]

Finally, the probability of receiving a golden handshake increases significantly in firms with a steep wage gradient – at least for blue-collar workers. As there is no legal claim to these additional voluntary severance payments, the decision to offer a golden handshake is purely driven by firms. The fact that the probability of such offers increases with firm costs for older workers strengthens our interpretation of the effects of wage gradients as a firm effect. While the point estimate is positive as well, the coefficient for white-collar workers is not statistically significant, although golden handshakes are more common among this type of employees. Golden handshakes may be partly more institutionalized there (i.e., in the banking sector) and, therefore, may be more independent of the wage structure.

So far, we have concentrated on the impact of the wage structure on job exit ages and

[13]This seems to be in line with conversations with personnel managers, that phased retirement programs are perceived as more expensive to firms, and that they are typically demanded by the employees rather than the firms (Graf *et al.*, 2007).

found a negative causal effect. Does a steeper wage gradient, thus, also lead to an earlier retirement age? Not necessarily. Workers may leave the job permanently; but instead of entering formal retirement immediately, they may bridge the time until formal retirement with some other benefits.

Table 7 focuses on such pathways into retirement, i.e., the time between job exit and actual retirement start. For comparison reasons, the first row replicates the coefficients for job exit age from Table 4. When looking at formal retirement age directly, it turns out that a steep wage gradient does not significantly reduce formal retirement age. Instead, the significant reduction in job exit age is absorbed by a corresponding increase of the duration between job exit and formal retirement. Most of the effect comes from longer spells of unemployment. Almost 90 percent of the increase of the duration between job exit and retirement can be explained by the increased duration of receiving unemployment benefits. This suggests that firms well know that an early layoff of older workers is mostly compensated by the public unemployment insurance system and they take this into account when optimizing their firing (and hiring) policies.

5 Conclusions

Steep wage gradients in firms may cause earlier job exit of elderly workers. Using a decade of Austrian retirement entries and an instrumental variables approach, we find that a one standard deviation increase of the wage gradient in a firm leads to an earlier job exit of approximately 6 months for blue-collar workers and 5 months for white-collar ones. These effects are substantial in size and stable across a variety of robustness checks concerning definitions of wage gradients and time lags of the instrument. Steep wage gradients also lead to a higher incidence of golden handshakes.

Our interpretation of these results is that firms play an active role in the determination of their workers' retirement age. Given individual retirement incentives – represented by detailed social security wealth calculations – a steeper wage gradient will stimulate firms to get rid of elderly workers prematurely; although the workers have an incentive to hold on to these good jobs even longer. Following this interpretation, it turns out that firms try to lay off elderly workers with the help of golden handshakes and an – unintended – assistance of the unemployment insurance system: These laid off workers do not enter formal retirement earlier, but rather bridge the gap until formal retirement with unemployment benefits.

Recognizing and quantifying an active role of firms in the retirement processes is a major step in discussions about early retirement problems and potential remedies. From a policy perspective, our results suggest that decreasing firm incentives by reducing seniority wage profiles – i.e., flattening the wage profiles at higher ages – can increase employment at older ages. Since early labor market exit is associated with higher cost for the social security systems (via prolonged receipts of unemployment benefits or pension payments), firms could also be obliged to bear parts of these costs directly, e.g., via some sort of experience rating.

The construction of deferred compensation schedules via steep age-wage profiles (Lazear, 1979) typically requires some form of mandatory retirement age. Increasing the regular retirement age – as is discussed in many countries – would, thus, also be costly for firms: Given a seniority wage contract, a later retirement age would require firms to keep older and expensive workers longer on the payroll. While in the long run new contracts will take this longer working life into account, in the short run incentives for firms to renege on these contracts may increase.

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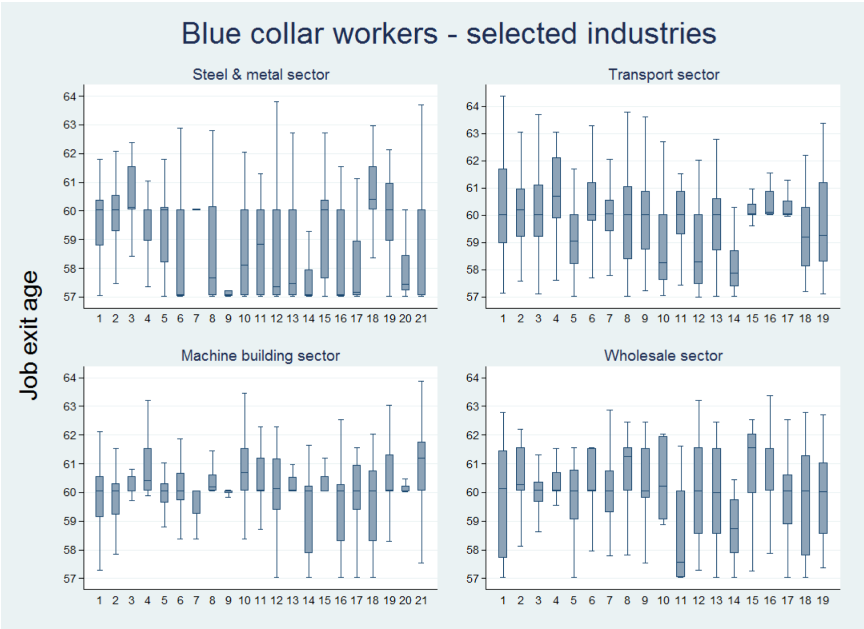
6 Figures and tables

Figure 1: Job exit age of male blue-collar workers[a]



[a] Own calculations based on data from *ASSD*. Job exit age distribution of largest firms in selected sectors for blue-collar workers

Figure 2: Job exit age of male white-collar workers[a]

[a] Own calculations based on data from *ASSD*. Job exit age distribution of largest firms in selected sectors for white-collar workers

Figure 3: Definition of wage gradients

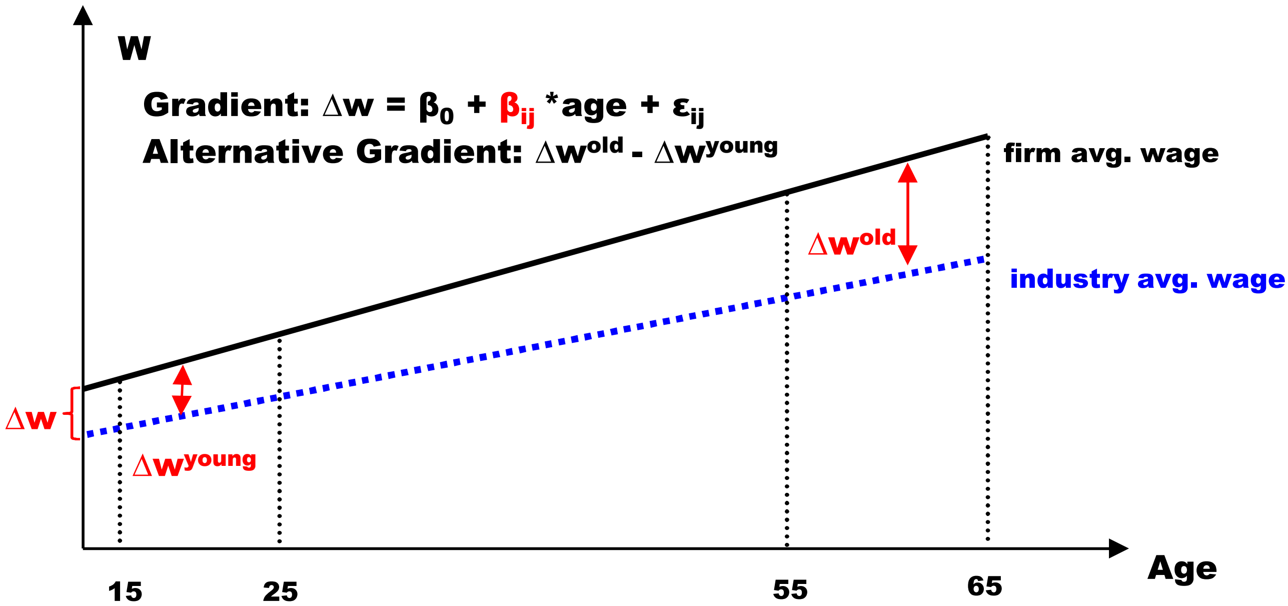


Table 1: Descriptive statistics

	Blue-collar workers	White-collar workers
Job exit age	59.78	60.78
	(1.683)	(1.634)
Disability	0.309	0.081
	(0.462)	(0.273)
Golden handshake	0.0797	0.149
	(0.271)	(0.356)
Phased retirement	0.170	0.224
	(0.376)	(0.417)
Years between job exit and retirement	0.088	0.083
	(0.286)	(0.282)
Years of unemployment after job exit	0.079	0.069
	(0.269)	(0.253)
Years being out of labor force after job exit	0.006	0.012
	(0.064)	(0.107)
Wage gradients		
Wage gradient	0.0947	-0.0623
	(0.514)	(1.700)
Alternative wage gradient	3.443	-4.776
	(22.43)	(62.81)
Additional covariates		
No. of weeks worked at age 55	49.87	51.15
	(9.289)	(7.045)
No. of weeks on sick leave at age 55	1.216	0.449
	(6.830)	(4.532)
No. of weeks out of work at age 55	0.783	0.361
	(4.610)	(3.700)
Experience (in years)	25.60	25.39
	(5.769)	(4.520)
Tenure (in years) at age 55	11.76	13.54
	(10.11)	(10.49)
Social security wealth (in 1,000) at age 55	397.4	505.8
	(100.7)	(98.29)
Social security contribution months at age 55	359.8	323.5
	(83.45)	(64.96)
Firm size	1591.5	1431.7
	(3866.4)	(3443.3)
Unemployment rate (lag 1)	8.314	8.006
	(2.879)	(3.060)
Unemployment rate (lag 5)	8.519	8.105
	(2.877)	(2.990)
Person fixed-effect	-0.00813	0.531
	(0.235)	(0.417)
Observations	41,296	45,131

Table 2: Blue-collar workers: The effect of the wage gradient on job exit age

	(1) OLS	(2) First stage	(3) 2SLS
Wage gradient	0.017		-0.926**
	(0.024)		(0.471)
Local unemployment rate (lag 10)		-0.029***	
		(0.006)	
No. of weeks worked at age 55	-0.005	-0.002***	-0.008
	(0.004)	(0.001)	(0.005)
No. of weeks on sick leave at age 55	-0.007*	-0.002**	-0.009*
	(0.004)	(0.001)	(0.005)
No. of weeks out of work at age 55	-0.002	-0.004***	-0.006
	(0.004)	(0.001)	(0.005)
Experience (in years)	-0.153***	-0.003*	-0.156***
	(0.009)	(0.002)	(0.009)
Tenure (in years) at age 55	0.002	0.010***	0.011**
	(0.002)	(0.001)	(0.005)
Social security wealth (in 1,000) at age 55	-0.003***	0.001***	-0.002***
	(0.000)	(0.000)	(0.000)
Social security contribution months at age 55	0.004***	-0.000***	0.004***
	(0.000)	(0.000)	(0.000)
Firm size	-0.000	0.000***	0.000
	(0.000)	(0.000)	(0.000)
Unemployment rate (lag 1)	-0.008	0.017**	0.002
	(0.010)	(0.007)	(0.012)
Unemployment rate (lag 5)	0.012	0.010	0.005
	(0.010)	(0.006)	(0.011)
Person fixed-effect	1.069***	0.614***	1.646***
	(0.059)	(0.058)	(0.318)
Year of job exit FE	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes
Regional FE	Yes	Yes	Yes
Number of observations	41,296	41,296	41,296
R-squared	0.30	0.20	0.23
Mean of dependent variable	59.78	3.75	59.78
S.d. of dependent variable	1.68	24.02	1.68
Mean of wage gradient	0.10		0.10
S.d. of wage gradient	0.53		0.53
Mean of unemployment rate (lag 10)		7.33	
S.d. of unemployment rate		2.46	
F-test of weak instrument			22.79

Standard errors clustered on district in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 3: White-collar workers: The effect of the wage gradient on job exit age

	(1) OLS	(2) First stage	(3) 2SLS
Wage gradient	-0.014		-0.273***
	(0.009)		(0.096)
Local unemployment rate (lag 10)		-0.093***	
		(0.018)	
No. of weeks worked at age 55	0.015***	-0.005*	0.013***
	(0.003)	(0.002)	(0.003)
No. of weeks on sick leave at age 55	0.011***	-0.005*	0.010***
	(0.003)	(0.003)	(0.003)
No. of weeks out of work at age 55	0.017***	-0.010***	0.014***
	(0.004)	(0.003)	(0.004)
Experience (in years)	-0.180***	-0.013*	-0.184***
	(0.011)	(0.007)	(0.011)
Tenure (in years) at age 55	0.003***	0.021***	0.009***
	(0.001)	(0.002)	(0.002)
Social security wealth (in 1,000) at age 55	-0.002***	0.002***	-0.001***
	(0.000)	(0.000)	(0.000)
Social security contribution months at age 55	-0.002***	-0.000	-0.002***
	(0.000)	(0.000)	(0.000)
Firm size	0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)
Unemployment rate (lag 1)	-0.033***	0.044*	-0.026***
	(0.008)	(0.022)	(0.009)
Unemployment rate (lag 5)	0.027***	0.061**	0.028***
	(0.007)	(0.025)	(0.008)
Person fixed-effect	1.033***	0.797***	1.240***
	(0.039)	(0.051)	(0.095)
Year of job exit FE	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes
Regional FE	Yes	Yes	Yes
Number of observations	45,131	45,131	45,131
R-squared	0.45	0.18	0.40
Mean of dependent variable	60.78	-3.61	60.78
S.d. of dependent variable	1.63	65.09	1.63
Mean of wage gradient	-0.06		-0.06
S.d. of wage gradient	1.51		1.51
Mean of unemployment rate (lag 10)		6.68	
S.d. of unemployment rate		2.49	
F-test of weak instrument			26.63

Standard errors clustered on districts in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 4: The effect of the wage gradient on job exit age: Robustness

	Blue-collar workers		White-collar workers	
	Wage gradient	Alternative	Wage gradient	Alternative
A: IV: Unemployment rate (lag 10)				
Wage gradient	-0.926**	-0.023**	-0.273***	-0.006***
	(0.471)	(0.011)	(0.096)	(0.002)
F-test of weak instrument	22.797	25.487	26.631	30.007
Number of observations	41,296	41,245	45,131	44,972
B: IV: Unemployment rate (lag 15)				
Wage gradient	-0.749**	-0.020**	-0.220**	-0.005**
	(0.349)	(0.009)	(0.112)	(0.003)
F-test of weak instrument	32.382	31.523	14.756	14.940
Number of observations	40,144	40,102	43,300	43,146
Mean of dependent variable	59.78	59.78	60.78	60.78
S.d. of dependent variable	1.68	1.68	1.63	1.63
Mean of wage gradient	0.10	3.44	-0.06	-4.78
S.d. of wage gradient	0.53	22.43	1.51	62.81

Standard errors clustered on districts in parentheses;
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 5: Restrict samples to workers with minimum tenure of instruments' lag

	Blue-collar workers		White-collar workers	
Lag of instrument:	Lag 10	Lag 15	Lag 10	Lag 15
Minimum tenure:	10 years	15 years	10 years	15 years
Wage gradient	-0.760*	-0.713**	-0.280***	-0.194
	(0.391)	(0.305)	(0.095)	(0.121)
F-test of weak instrument	20.67	23.16	26.22	14.92
Number of observations	27,025	19,989	32,527	25,793

Standard errors clustered on districts in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 6: Alternative outcomes

	Blue-collar	White-collar
Disability	-0.005	0.023
	(0.126)	(0.018)
Mean of dependent variable	0.31	0.08
Phased retirement	-0.080	-0.027
	(0.166)	(0.048)
Mean of dependent variable	0.17	0.22
Golden handshake	0.199***	0.055
	(0.069)	(0.034)
Mean of dependent variable	0.08	0.15
F-test of weak instrument (lag 10)	22.79	26.63
Number of observations	41,296	45,131

Standard errors clustered on districts in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 7: Pathways into retirement

Outcome variable	Blue-collar	White-collar
Job exit age	-0.926**	-0.273***
	(0.471)	(0.096)
Formal retirement age	-0.126	0.031
	(0.467)	(0.074)
Years between job exit and retirement	0.800***	0.304***
	(0.212)	(0.065)
Mean of dependent variable	0.09	0.08
Years of unemployment after job exit	0.718***	0.276***
	(0.194)	(0.060)
Mean of dependent variable	0.08	0.07
Years of out of labor force after job exit	0.057***	0.025**
	(0.019)	(0.010)
Mean of dependent variable	0.01	0.01
F-test of weak instrument	22.79	26.63
Mean of wage growth gradient	0.10	-0.06
S.d. of wage growth gradient	0.53	1.51
Number of observations	41,296	45,131

Standard errors clustered on districts in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

7 Appendix

7.1 Calculation of wage gradients

We use the daily wages of all male workers between 15-65 years between 1997 and 2009 in order to compute a firm wage profile and the respective industry wage profile of all firms. The wage gradients are computed separately for blue-collar and white-collar workers, the wage gradients for blue-collar workers only consider firm wages for blue-collar workers; equivalently for white-collar workers.

We split the (blue-collar/white-collar) workforce into 10 age groups, where age group 1 refers to ages 15-20 and age group 10 to ages 61-65 respectively. Let w_{naijt} be the daily wage of worker $n = 1, ..., N$ in age group $a = 1, ...10$, firm $i = 1, ..., I$, industry $j = 1, ..., J$ and year $t = 1997, ..., 2008$.

Let

$$w_{aijt} = \frac{1}{N} \sum_{n=1}^{N} w_{naijt} \tag{3}$$

be the average daily wage of age group a, firm i in industry j and year t, and

$$w_{ajt} = \frac{1}{I} \sum_{i=1}^{I} w_{aijt} \tag{4}$$

the average daily wage of age group a in industry j and year t.

In the regression-based approach, we compute the wage gradient directly by estimating the following OLS regression separately for each firm and year:

$$(w_{aijt} - w_{ajt}) = \beta_{0,ijt} + \beta_{1,ijt} * age_{ijt} + \epsilon_{ijt} \tag{5}$$

with age_{ijt} as the mean age of each age group. Equation 5 gives us an age coefficient for each firm and year and, thus, $\beta_{1,ijt}$ can be directly defined as our wage gradient.

A gradient 1€ higher means that the firm age-wage profile is higher by 1€ per year relative to the industry age-wage profile.

The alternative definition of the wage gradient uses the difference between average firm daily wage and average industry daily wage:

$$Grad_{ijt}^{Alternative} = \frac{1}{2} \sum_{a=9}^{10} (w_{aijt} - w_{ajt}) - \frac{1}{2} \sum_{a=1}^{2} (w_{aijt} - w_{ajt}) \tag{6}$$

Here, a gradient higher by 1€ implies that a firm age-wage profile is higher compared to the industry profile by 1€ over 40 years.

7.2 Derivation of worker fixed-effects

Our proxy measure of workers' productivity is based on a decomposition procedure developed by Abowd *et al.* (1999) that separates individual workers' wages into one part that is explained by observable time-varying productivity characteristics of the worker (such as age, labour market experience or tenure), as well as time-invariant worker fixed and firm fixed wage components (AKM model henceforth).

Formally, the calculation of worker and firm fixed wage components takes the form[14]:

$$y_{ijt} = \phi_j + \theta_i + X'_{ijt}\beta + \epsilon_{ijt} \tag{7}$$

where

$$E\left[\epsilon_{it}|\theta_i, \phi_j, t, X_{ijt}\right] = 0. \tag{8}$$

The parameters ϕ_j and θ_i provide the firm and person fixed wage components, respectively, while X_{ijt} controls for observable time varying productivity characteristics of the worker (tenure and experience). While the firm-fixed effect measures the average deviation in wages paid to its employees irrespective of their individual productivity ("firm rent"), the person fixed effect can be interpreted as an indicator of worker's individual productivity. Identification of the AKM model requires that workers' are mobile between firms and that this mobility is "exogenous"; thus, any form of assortative matching between ("good") workers and ("good") firms must be ruled out.

Several tests have been proposed to test for assortative matching in the context of the AKM model. A first, albeit imperfect test is the correlation between worker and firm fixed-effects. This correlation is slightly negative (-0.012) in our sample implying - if any - weak negative assortative matching. A more elaborated test analyzes the movement of workers between firms. If there is no assortative matching between workers and firms, then workers who move from a "high firm-wage" firm should experience wage losses on average while those who move from "low firm-wage" firms should experience corresponding wage gains (Card *et al.* (2013)). Additionally, the effects of moving "up and down the ladder" should roughly be symmetrical, implying that associated wage changes should be roughly symmetrical.

We follow Card *et al.* (2013) and Flabbi *et al.* (2014) and classify the origin and destination firm for all job movers in our data by the quartile of the estimated firm effect. We then calculate for all workers within the 16 origin and destination cells the average wages for the two years prior to and after job change.[15]

As the following table shows, the assumption of exogenous mobility is well-supported by our data. Workers who move from "low firm-effect" firms to "high firm-effect" experience wage increases that are roughly symmetrical to the wage losses of those who move from "high firm-effect" firms to lower quartile firms.

[14] We use Ouazad's (2008) Stata module.

[15] We perform this calulation for all workers in our sample that are employed in the origin and destination firm for at least two consecutive years. Overall, our sample consists of 713,400 workers

Table 8: Wages before and after job change

Origin / destination Quartile	Number of movers	2 years before	1 year before	1 year after	2 years after	change from 2 years before to 2 years after Raw	adjusted
1 to 1	69084	3,60	3,72	3,84	3,89	0,30	0,00
1 to 2	44855	3,64	3,79	4,13	4,21	0,56	0,26
1 to 3	31136	3,70	3,87	4,32	4,40	0,70	0,40
1 to 4	19809	3,41	3,62	4,37	4,48	1,06	0,76
2 to 1	31840	3,84	3,94	3,79	3,83	-0,01	-0,29
2 to 2	52812	3,91	4,02	4,13	4,19	0,28	0,00
2 to 3	51040	4,04	4,16	4,33	4,40	0,37	0,08
2 to 4	51773	4,18	4,34	4,56	4,66	0,48	0,20
3 to 1	19187	4,01	4,09	3,69	3,74	-0,27	-0,53
3 to 2	33586	4,04	4,13	4,15	4,21	0,17	-0,09
3 to 3	57601	4,16	4,26	4,36	4,42	0,26	0,00
3 to 4	56201	4,25	4,37	4,55	4,63	0,37	0,12
4 to 1	14288	4,26	4,34	3,60	3,65	-0,61	-0,88
4 to 2	17503	4,19	4,27	4,13	4,18	-0,01	-0,28
4 to 3	36553	4,30	4,39	4,41	4,46	0,16	-0,11
4 to 4	126172	4,44	4,54	4,64	4,71	0,27	0,00

NOTES: Mean log of daily wages of workers by origin and destination firm. Firms are classified by the quartile of estimated firm effects. Adjusted: Mean wage change for origin destination group minus the mean change for job movers from the same origin quartile who move to a firm in the same destination quartile.