Seniority Wages and the Role of Firms in Retirement*

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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 considerations about wage costs and replacement costs, 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 accrual rates or health status, the steepness of the wage profile will have different incentives for workers as compared to firms to maintain the employment relationship. 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 have on average a lower job exit age and a higher incidence of golden handshakes.

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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 focus on – voluntary – labor supply effects, surveys often reveal that a large proportion of workers state their early retirement was not a voluntary decision (Dorn and Sousa-Poza (2010) using ISSP data or Marmot *et al.* (2004) for England). Differentiating between voluntary and involuntary retirement may not be very clear for survey respondents; their answers may reflect retirement regulations, but also a role of firms, which 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 wage schedules opens up important policy channels.

In this paper, we want to study the role of labor demand in individual retirement decisions. In particular, we investigate whether a steep cross-sectional seniority wage profile in a firm leads to a markedly lower job exit age of its workforce, using high-quality administrative data for the universe of Austrian workers and firms. We tackle the endogenous nature of a firm's wage structure by instrumenting with labor market shocks a decade ago. We are also able to study the causal relationship between seniority wage profiles and the incidence of entering into a disability pension, phased retirement or the probability of receiving a golden handshake. Particularly the latter directly points towards an active role of the firm in retirement behavior, because a golden handshake is a voluntary payment to workers and, as such, an exclusive choice variable of the firm.

Previous research on labor demand effects in retirement has been directed towards different issues: Bartel and Sicherman (1993), Bello and Galasso (2015) 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 directly. 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. Hallberg (2011) exploits variations in age-dependent collective fee costs across companies in Sweden to show how non-wage costs affect early retirement probabilities.[1]

Other studies are related to seniority wage profiles in relation to labor demand in general. There is widespread evidence that steep seniority wage profiles have an impact on hiring behavior and retention of workers. Hirsch *et al.* (2000) show that older workers have substantial entry barriers in occupations with steep wage profiles and pension benefits. Zwick (2012) shows that establishments with steeper seniority profiles hire less elderly workers, hire more younger workers and, thus, have a longer job tenure of its workers. Kramarz *et al.* (1996) show that evolutions of employer-specific wage policies are correlated with changes of the workforce in terms of experience and seniority.

In a theoretical life-cycle of workers, firms are indifferent with regard to the retirement age of their workers as long as age-wage profiles correspond to age-productivity profiles. This is not the case otherwise; incentives for firms to lay off older workers may arise, whenever age-wage profiles exceed age-productivity profiles. A prototypical seniority wage profile has been constructed by Lazear (1979): here, workers and firms 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.[2]

In our study, we focus on cross-sectional age-wage profiles in firms rather than such a Lazear-type seniority wage schedule. A steep cross-sectional wage structure in a firm –

[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, which may effectively subsidize workforce reductions similar to non-experience-rated unemployment insurance.

i.e. higher current costs of elderly workers relative to replacement costs – will exert large incentives for firms to reduce the workforce of elderly workers and replace them with younger – cheaper – ones.[3] On the other hand, elderly workers in such firms predominantly have an incentive to stay longer. While higher wages in the worker's late career should increase labor supply, at least for inter-temporal substitution reasons, incentives from pension claims are less clear-cut. A firm effect on individual retirement can only be separated from the individual retirement decision if individual incentives from pensions are addressed properly within the empirical framework.

Our results show that steep wage gradients in firms indeed cause earlier job exit of elderly workers. Such workers leave the firm earlier and typically stay in the unemployment register until finally retiring. Together with an increase in voluntary golden handshakes we can conclude that firms do play a role in retirement behavior of their workers.

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.

Particularly for blue-collar workers, disability insurance is a frequent pathway into retirement. An individual with health problems can access disability pension conditional on having a severe health impairment that lasts for at least 6 months and implies a reduced work capacity of at least 50% relative to a healthy person within comparable education in any occupation[4]. It is difficult for firms to influence the entry into disability retirement,

[3] Such a substitution of older by younger workers is always possible in a frictionless labor market; but also in a market with frictions (long-term contracts or hiring and firing costs) provided that productivity-corrected wage differentials between old and young are large enough.

[4] Above the age-threshold of 57 the same individual qualifies for disability benefits if the ability to work is reduced by more than 50 percent relative to a healthy person with comparable education in a

because eligibility is checked by independent medical doctors.

Concerning the relatively low retirement age in Austria, 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 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 workforce in detail from 1971 to 2012 (Zweimüller *et al.*, 2009). We use all male[5] blue-collar and white-collar workers aged 57 to 65 who retired in the period 2000 to 2009 and worked in private sector firms.[6] We exclude workers from small firms with less than 15 workers and from firms without workers below age 25, because no reliable 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 time between job exit and retirement of maximum 2 years. In fact, this job exit age is the more relevant variable of interest because workers

similar occupation

[5]We do not use female workers, because part-time work is very common among women in Austria, and we have missing working time information.

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

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 job 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.

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 unattractive – Austrians retire the first day possible, we do see large variations in retirement ages. Figure 1 shows boxplots for the distribution of job exit ages for blue-collar workers in firms with the largest number of retirement transitions between 2000 and 2009 in the most relevant sectors of the Austrian economy. The upper-left panel, for example, shows the job exit age distribution of the 21 largest firms in terms of retirement transitions in the steel industry. These firms are relatively homogeneous, but still considerable firm-specific variation in 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 as pronounced as for blue-collar workers (e.g., in the wholesale and energy supply sector).

3 Empirical strategy

We want to study the role of the firm in retirement decisions, so the identification of firms with higher incentives to lay off older workers is pivotal. We argue, that firm incentives depend on wage costs for older workers in particular. In the following, we will describe how we construct those seniority wage profiles and how we account for the productivity of workers. Moreover, we have to control for individual retirement incentives arising from social security considerations. For the identification of the impact of the seniority wage profile on job exit we use 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

How should we measure the incentives of a firm to fire older workers? The most direct measure is wage costs of elderly workers relative to replacement costs – corrected for productivity differences. Firms with a larger cross-sectional difference between old and young workers' wages will have the largest incentive to exchange their personnel. Note that this simple cross-sectional wage differential is unrelated to seniority wage profiles which result from an incentive-based Lazear-type contract. Whether an individual worker had a steep or a flat wage increase over time does not matter considering the essential cost-minimization problem of the firm: produce the currently needed output of the firm with a combination of workers with the lowest possible labor costs. Current wages of young workers are also a good proxy for current replacement costs on the market.

We consider 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 (2000 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 to a full extent. However, the industry profile is composed of the direct competitors who share similar technologies, are of comparable size and share the same collectively bargained minimum wages.[7] A steeper firm wage profile relative to the industry wage profile – a positive wage gradient – is associated with higher seniority wage costs for firms, in particular regarding costs of potential replacement workers. Because firms within

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

an industry are rather homogeneous 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[8].

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 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 ($\beta_{1,ijt}$) 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. 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. The wage gradient is calculated or estimated separately for blue- and white-collar workers in each firm. A more detailed description of the wage gradient calculations can be found in Section A.1 of the Web Appendix.

Nevertheless, using the industry wage profile as a proxy for productivity will incorporate

[8] Note that employers also contribute to the social security system, since health and pension insurance costs are divided between firms and workers. However, these contributions are a fixed percentage of wages (18.12 percent) for every firm with an upper bound. This may lead to an underestimation of the true incentive to lay off workers. Firm pensions are not of high relevance in Austria in our observation period (only 2 to 4 percent of all pension incomes) and are highly concentrated within certain sectors (Url, 2009). Since our wage gradient is defined relative to sectors and we additionally control for sector fixed-effects, these additional pension costs on the firm side should not significantly bias our incentive measure.

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) (AKM wage decomposition), where wages are decomposed into firm- and worker-specific components.[9] Moreover, an instrumental variables strategy – discussed below – will also take care of measurement error problems arising from the use of the industry wage profile.

Finally, our conceptual framework requires us to separate between the role of the firm in retirement decisions and individual's labor supply decision. As already pointed out, steeper wage gradients induce firms to lay off older workers earlier. The relationship between individual retirement incentives and the wage gradient may be ambiguous: on the one hand, steeper wage profiles increase the work incentives for workers, since their expected pension payments increase with higher wages at older age[10], on the other hand, individuals facing steeper wage profiles may also have a higher lifetime wealth, which induces those workers to retire earlier. It is conceptionally impossible to make a clear-cut test whether retirement is a firm's or a worker's decision. Nonetheless, we try to control for individual retirement incentives and wealth effects as good as possible. We control for the number of social security months worked and a person-specific wage fixed effect - both variables should approximate lifetime wealth rather well. Most importantly we include a pension accrual rate to control for social security incentives. These simulated accrual rates are the preferred measure for properly capturing financial retirement incentives; (see for example in Gruber *et al.* (2010))[11] they calculate the marginal incentive for individuals to

[9]Worker fixed-effects are identified within each set of workers and firms that is connected by individual workers moving between different firms. Since the large majority of workers in our sample is observed in more than one firm, these effects are well-identified. For details on the decomposition method see Section A.2 of the Web Appendix.

[10]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 in Austria, 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.

[11]For computing these accrual rates, we use rich information about employment histories and a new source for retirement information from a new survey (esp. the exact amount of income used for retirement calculation, the exact type of employment in the view of the retirement agency, "Verdichtung von Versicherungszeiten und Pensionsberechnung" (VVP)) of all pensioners in Austria to calculate expected retirement benefits for all potential retirement dates (several years before and after their real retirement date). We implement respective legal provisions for different birth cohorts. Since we do not see all of

work one additional year instead of retiring. We are aware that these important control variables may not fully capture wealth and retirement incentive effects of individuals, and so some doubts may remain. However, given our results and additional outcome variables, i.e. golden handshake offers, we are confident that our results can be interpreted as a test for the role of firms in retirement decisions.

3.2 Econometric model

We specify our econometric model in the following way:

$$JE_{nijt} = \alpha_0 + \alpha_1 \cdot WageGrad_{ijt} + \alpha_2 \cdot ACC_{nij} + \alpha_3 \cdot X_{nijt} + \alpha_4 \cdot \eta_n + \epsilon_{nijt} \qquad (1)$$

where JE_{nijt} is the job exit age of worker n in firm i of industry j in year t, and $WageGrad_{ijt}$ corresponds to one of the two firm incentive measures which is calculated for each firm and each year. Equation 1 relates individual worker's job exit age to the firm-level seniority wage gradient and individual pension accrual rate ACC_{nij}. η_n is the personal fixed-effect from the AKM wage decomposition, and the vector X_{nijt} 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 also control for job exit year, region and industry fixed-effects as well as an indicator variable whether the firm is located in a REBP district[12]. Since the individuals are subject to several changes of the pension system, we include a second-order polynomial of quarter-of-birth cohorts. This should control for the gradual change by different quarter-of-birth cohorts of several important legislative pension rules, i.e., the increase of early retirement age or the extension of the pension assessment base (Staubli and Zweimüller, 2013). Finally, local unemployment rates 1 and 5 years ago are included. The key parameter of

our individuals in this new retirement dataset, we use these exactly calculated pensioners and the rich information about employment histories in the data to generate a non-parametric imputation process for the non-observed pensioners. We used a Random Forest Regression model to impute the missing accrual rates to get a final sample of all retirees that we used in our regression. We include a binary indicator for those observations with imputed accrual rates.

[12]REBP stands for "Regional Extension Benefit Program" which extended the duration of UI benefits for a large group of eligible workers in selected regions (Winter-Ebmer (2003) or Lalive *et al.* (2015)).

interest is α_1, which measures the effect of the wage gradient on job exit age. From our theoretical considerations, we expect α_1 to be negative, because a greater gap between firm and industry wage profiles should increase firm's firing incentives and consequently lower the job exit age of their workers. Conditional on pension accrual rates, α_1 should only capture firm effects on job exit age.

3.3 Identification

The identification of the wage gradient's 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 earlier firing of older, highly-paid workers 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 a specific self-selection of workers. Due to these reasons – and also to counter measurement error problems in the wage gradient with respect to firm productivity profiles – we implement an instrumental variable 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 idiosyncrasies at the time of job entry.

Thus, we use the local unemployment rate on the district level for prime-age workers (25-45 years old) 10 years before workers' job exit as the instrumental variable.[13] Our

[13]We also present results for different time lags as a robustness check. An alternative might be to use unemployment rates at the beginning of each worker's career as an instrument. We do not do this for

first stage takes the following form:

$$WageGrad_{ijt} = \gamma_0 + \gamma_1 \cdot UR_{t-10} + \gamma_2 \cdot ACC_{nij} + \gamma_3 \cdot X_{nijt} + \gamma_4 \cdot \eta_n + \mu_{nijt} \qquad (2)$$

with UR_{t-10} as the local unemployment rate for prime age workers 10 years before job exit. We allow observations within the same firm to be correlated and cluster the standard errors on the firm level.

Since we consider job exits between 2000 and 2009, we capture the regional variation of local unemployment rates between 1990 and 1999 in our main specification. Although the Austrian business cycle in these years was relatively modest without severe regional unemployment shocks, there is still ample variation in unemployment rates across districts and years observable. While the average of these local unemployment rates is between 5.5 and 7.65 percent, there are many districts in different parts of the country with unemployment rates above 10 percent. Figure 4 plots the within-district standard deviation of local unemployment rates over time (between 0.23 and 3.2 percentage points) and reveals significant variation across Austria.[14]

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 lower rates. These relatively better hiring conditions in the past will thus reproduce themselves into relatively low wages of the current older workforce. Firms in European countries, in particular Austria, do rarely set individual wage profiles, they typically negotiate and set general wage increases which are then relevant for all workers within a cohort. Knell and Stiglbauer (2012) show that Austrian's unionised wage negotiations are concentrated on wage increases in percent and that these increases can be given to previously negotiated

three reasons: i) such an instrument is closer to a Lazear (1979) career wage argument, ii) we cannot construct this instrument for a large number of our sample and iii) for the remaining workers, such an instrument is somewhat weaker.

[14] On average, around 55% of newly hired workers come from the same district as the firm. We also estimated the main regressions i) using only these workers, and ii) using NUTS3 regions for the instrumentation strategy. Results were largely similar to the ones reported below; they are available on request.

minimum wages as well as to overpayment.

Figure 5 shows the responsiveness of wages to local labor market conditions based on binned scatterplots for blue- and white-collar workers[15]. For both, blue- and white-collar workers, we clearly find a significantly negative relationship between local labor market conditions and entry wages of newly hired workers. In contrast, we do not see any significant response of wages of workers already working in the firm. Hence, the mechanism through which our instrument affects today's firm-specific age-wage profiles is only driven by entry wages of newly hired workers. Since most of our workers in fact did not enter the firm in the year we measure the instrument, the instrument does not directly affect most individual wages, but firm wage profiles only.

In fact, what our IV does, is to shift the level of wages of particular individuals, namely of those who get hired in a certain year. This shift affects the realized wage structure by seniority in the firm in the cross-section. But per se, it does not affect the slope of individuals' wage seniority profile, it just shifts some individual profiles up or down. In this respect, it does not relate to the steepness of the seniority wage profile that would endogenously arise in a model a la Lazear (1979), but to the overall wage costs within the firm conditional on age. So our IV is essentially capturing the effect of wage costs on retirement decisions and not the effect of individual seniority profiles as such.

3.4 Validity of the instrument

Our IV approach identifies a local average treatment effect of a 1€ increase of the wage gradient on job exit age for those who leave the firm because of higher wage gradients

[15]We take all firms of our individuals in our main sample (job exit between 2000 and 2009) and construct a new sample of the active workers of these firms 10 years before, hence at the time the instrument (unemployment rates) was measured. So we end up with a pooled sample consisting of all male blue- and white-collar workers of our firms between years $t = [1990, 1999]$. We define a worker to be a new hire if this individual has a tenure of less than 365 days in year t. A worker is denoted as a current employee if the tenure in year t is greater than 365 days. We calculate the residualized individual wages after controlling for age, sickness days during the last three years, individual productivity (personal fixed effect component from a AKM wage decomposition), experience, overall firm size, industry, year and regional fixed effects as well as a dummy whether the person works in a REBP district. We then plot these residualized wages on the contemporaneous local unemployment rates. The slope parameters shown in the graph are from simple OLS regressions of residualized wages on local unemployment rates, and are statistically significant for new hires, but not for current workers.

due to the past local labor market situation. The validity of the instrument requires $Cov(UR_{t-10}, \epsilon_{nijt}) = 0$, so the identifying assumption is that — conditional on covariates — past unemployment rates are unrelated to any unobserved firm characteristics, worker selection or unobserved individual propensity to retire today.

At this point, 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. The IV affects firms' overall age-wage profiles through entry wages in the past, and shifts the contemporaneous cross-sectional wage profile, but is rather independent from the individual wage profiles of our retirees as long as they are not hired exactly 10 years ago.[16] Nevertheless, there may be some concerns about the exclusion restriction of the instrument, which is fundamentally untestable. In the following, we provide several plausibility checks to support the exogenous nature of past local labor market shocks.

Firstly, studies show that current local labor market conditions and retirement behavior are related. So 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. Figure 4 shows that there is ample variation in unemployment rates in most districts. Nevertheless — to strengthen the validity of our instrument — we allow current local labor market conditions to directly affect individual retirement incentives and control for local unemployment rates 1 and 5 years before workers' job exit.

Secondly, one may argue that past local labor market conditions may be correlated with unobserved firm characteristics which affect labor demand 10 years later. Column (1) in Table 2 first shows that the number of new hires is uncorrelated with contemporaneous labor market conditions: Conditional on incumbent's average characteristics and firm characteristics such as firm size, share of blue-collar workers, industry and regional fixed effects, we find no significant relationship between the local unemployment rate and the firm's hiring rate. Our main hypothesis is that past unemployment has an impact

[16] In a robustness check we restrict our sample to those strictly employed before the time the instrument is measured.

on the wage gradient – and via this wage gradient on retirement behavior of the workforce. In particular, we have to show that past labor market conditions do not influence current firing or leaving policies directly; e.g. by affecting layoff processes. Table A.2 in the Web Appendix reveals clearly that there is no correlation between past local unemployment rates and the incidence of layoffs due to plant closure or mass layoffs – neither for white-collar nor for blue-collar workers. This should support our assumption that past local labor market conditions are not significantly correlated with unobserved firm characteristics

Thirdly, one may also be concerned about the selection of particular workers into firms in situations with different local labor market conditions. A prime suspicion here is that the pool of available workers will be different in a recession. To start with, we always control for a personal fixed-effect as a proxy for individual productivity and other individual characteristics such as individual financial retirement incentives. Also, as columns (II) to (V) of Table 2 reveal, we find no systematic sorting of workers into firms with respect to age, productivity[17] or past sick leave. We only find that hired employees have a bit more work experience (9.2 days or 0.3 percent), which is only a marginal correlation. So we find no evidence for systematic changes in hiring patterns due to worse local labor market conditions. This should support our assumption that past local labor market conditions are not significantly correlated with unobserved firm characteristics[18]. We also do not find a systematic increase in quit rates or take-up rates for disability pensions[19].

Worse labor market conditions could also change the sorting/hiring of workers with different health status. For a sub-sample of individuals working in the province of Upper Austria (20% of Austria), we have information on actual utilization of health care services, i.e. outpatient medical attendance, use of medical drugs or length of hospital stays. Again, we find no systematic selection of workers into firms based on their health and health care utilization (see Table A.5 in the Web Appendix). Based on these plausibility

[17]Individual productivity is proxied by the personal fixed-effect of the AKM wage decomposition.

[18]Table A.3 in the Web Appendix presents another informal test, where we included average characteristics of newly hired employees at the time of the instrument as additional control variables. We find no significant changes to our results.

[19]More details are available in Table A.4 in the Web Appendix. There are only very small and marginally significant results in some sub-cases.

checks we conclude that past local labor market conditions did not significantly alter the selection of workers into firms.

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, i.e. golden handshake, disability retirement, phased retirement, and explores pathways into retirement.

4.1 OLS and first stage results

Tables 3 and 4 summarize the estimation results for OLS, first stage and 2SLS for blue-collar and white-collar workers, respectively. The OLS coefficients of the wage gradient are negative and significant, but relatively 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 will bias the coefficient downwards.

Tables 3 and 4 also report the coefficients for the main covariates. Here, the social security wealth accrual rate and the collected social security contribution months are of particular interest. Both of these indicators of a more favorable pension situation reduce age at job exit, but the effect is small for social security months and negative, but insignificant for the accrual rates.

As the social security wealth and the accrual rate is based on contributions to the social security system, some of these effects may be picked up by our measure of work experience and recent employment behavior: We include both total work experience and job tenure in the actual job as well as a split of the year, the worker turned 55, into weeks worked, on sick leave and out of work. It seems that current sickness has a negative effect on retirement age, whereas total job experience has a positive effect. Firm tenure and firm size do not influence worker's job exit age much. 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 3 and 4 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.[20] The corresponding F-test for weak instruments yields values between 18 to 21, well above conventional critical values for weak instrument problems. These results point towards a clear mechanism: high unemployment ten years ago leads to low hiring wages – with given productivity. Due to constant wage growth scales, these relatively low wages persist for these workers later on. Together with higher hiring wages later on, such a firm will show a relatively low current age-wage profile. One might be worried that 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. Plausibility checks presented in Section 3.4 provide additional evidence against selection into firms.

Likewise, labor market conditions one or five years 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 younger 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 3 and 4 present our causal effects from the 2SLS estimations. For both blue-collar and white-collar workers, the coefficient of the wage gradient be-

[20]Figure A.1 in the Web Appendix is a graphical representation of the first stage relationship for blue- and white-collar workers. The slope of the line in the binned scatterplot corresponds to the coefficients of the unemployment rate in the first-stage regressions.

comes more negative as expected and is statistically significant. For blue-collar workers, a 1€ increase of the wage gradient leads to a 0.88 years lower job exit age. A one standard deviation increase of the steepness of the wage gradient would thus lead to a 5.5-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 6.9 months.

Figure 6 summarizes robustness checks using the other definition of the wage gradient, as well as different time lags of the instrument – from six to 15 years before job exit. 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. All coefficient estimates for the original gradient construct in the upper part of the Figure are very similar: Both for white- and blue-collar workers, all effects are negative; they are in the same quantitative range and – with the exception of lag 6 for white-collar workers – are also statistically significant.

In the lower part of Figure 6, we present results with the alternative wage gradient[21]. Here, an increase of one standard deviation of the alternative wage gradient decreases job exit age by approximately 5.9 months, compared to 5.5 months of the baseline wage gradient. The picture for white-collar workers is very similar: a one standard deviation increase of the alternative wage gradient decreases job exit age by 6.8 months, compared to 6.9 months in the basis. Here, again variations across time lags are very similar to the standard definition of the wage gradient.

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

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

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. Our results are very robust to this test; all estimates are very similar to the ones with the full sample – both in size as well as in statistical significance. We conclude that a potential selection of workers into firms is negligible and does not systematically affect our results.

4.3 Other outcomes and pathways into retirement

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

We do not find convincing effects of wage gradients on the probability that the worker receives a disability pension. The effect for blue-collar workers is negative and insignificant, whereas the one for white-collar workers is positive, but only significant at the 10% level. 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. While blue-collar workers claim disability typically due to physical problems, white-collar workers complain more about mental or psychological illnesses (burnout, etc.) or back pain.

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.[22]

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

[22]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).

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 in firms with high seniority wage profiles 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. Golden handshakes are more common among this type of employees, they may be more institutionalized here (i.e., in the banking sector) and, therefore, may be more independent of the wage structure.

Our discussion centers on distinguishing incentive effects of wage gradients for employers and employees and relies on the possibility of firms and/or workers to influence the retirement date. The social security wealth accrual rate is a good indicator to pinpoint individual's incentives towards retirement: if the accrual rate is high, postponing retirement is particularly valuable for the worker. In the following Table 7 we split the sample by the mean social security wealth accrual rate and present so separate results for workers with high and low work incentives. Results show strong differences between these two groups. In the case of workers with high work incentives, a higher wage gradient led to significantly lower retirement age – as compared to workers with low work incentives. All respective coefficients for low work incentives workers are lower than the ones for high work incentives workers. For white-collar workers with low work incentives we do not even see statistically significant results.[23]

These results are certainly incompatible with the view that the reaction of workers is shaping this pattern: workers with high work incentives would have a stronger incentive to hold on to their job in case of a high wage gradient. The picture we see points towards firm actions: Workers, who are potentially led into early retirement by their firm – the workers with high work incentives – are also more likely to receive a golden handshake and, in the case of white-collars, to enter disability pension. One interpretation of this phenomenon is that firms make it easier for workers to accept such an earlier exit from the job. Less precisely diagnosable illnesses, like burnout, in the case of white-collar

[23]Due to smaller sample sizes, in some of these cases, the instrument is somewhat weaker (F-statistic of 10). But in all cases, following Stock and Yogo (2005), we can rule out a 2SLS bias of more than 15% of the OLS bias.

workers may help with this strategy. The separation between workers with high work incentives and those with low work incentives is positively correlated with the steepness of the seniority wage profiles. Our results that workers with high work incentives react stronger to seniority contracts have to be seen in this context.

So far, we have concentrated on the impact of the wage structure on job exit ages and 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 8 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 Tables 3 and 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 80 percent (100 percent) of the increase in the duration between job exit and retirement for blue-collar workers (white collar workers) can be explained by the increased duration of receiving unemployment benefits. This implies that firms know well 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 5.5 months for blue-collar workers and 6.9 months for white-collars. 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 accrual rates – a steeper wage gradient will stimulate firms to end the work relationship with elderly workers earlier; 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 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 the active role of firms in the retirement process 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.

Other pension reform issues concern an increase in the regular retirement age. Such a reform would also be costly to firms, as older and more expensive workers tend to stay longer. Firms might react by reneging on existing contracts and sending or bribing workers into unemployment benefits or even disability. In a longer term, an increase in the working life may also shift existing seniority contracts and make them flatter. A medium-term policy proposal would ask for additional measures to facilitate the lengthening of the working life: next to a general flattening of the seniority wage profile, subsidies for the extended employment periods seem warranted.

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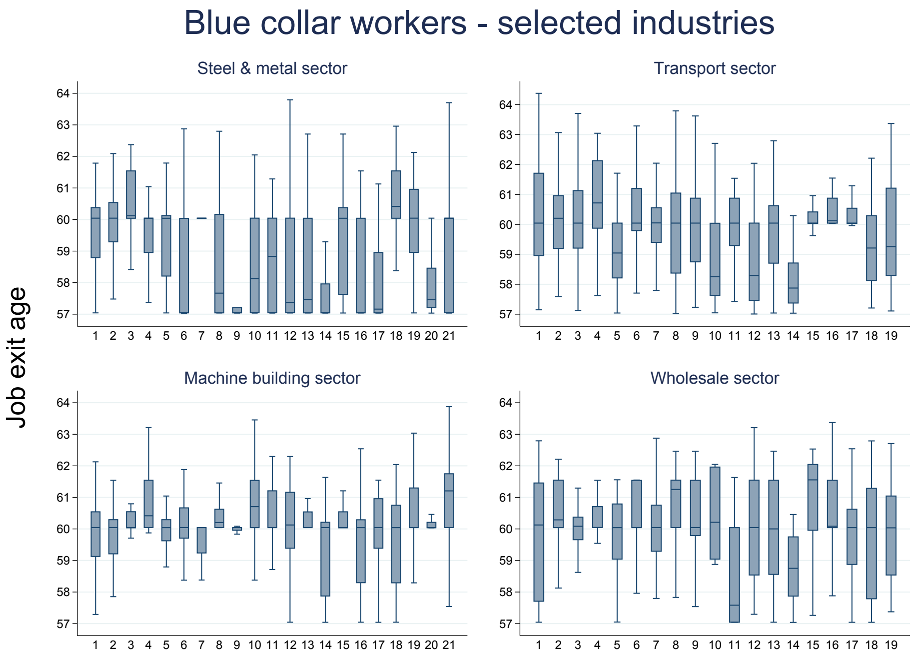
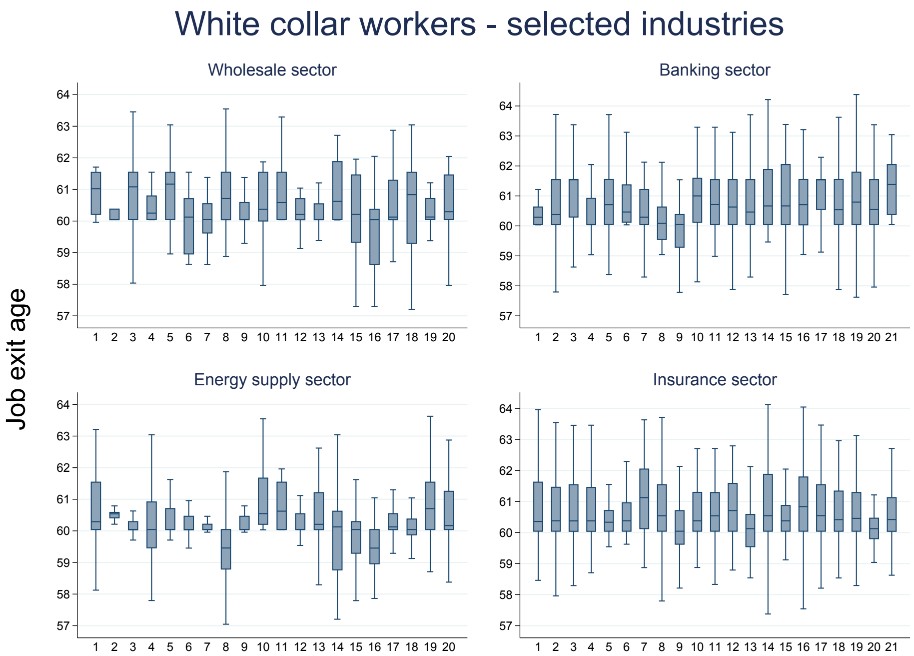
6 Figures and tables

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



Notes: Own calculations based on data from *ASSD*. Job exit age distribution of firms with most retirement transitions in selected sectors for blue-collar workers

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



Notes: Own calculations based on data from *ASSD*. Job exit age distribution of firms with most retirement transitions in selected sectors for white-collar workers

Figure 3: Definition of wage gradients

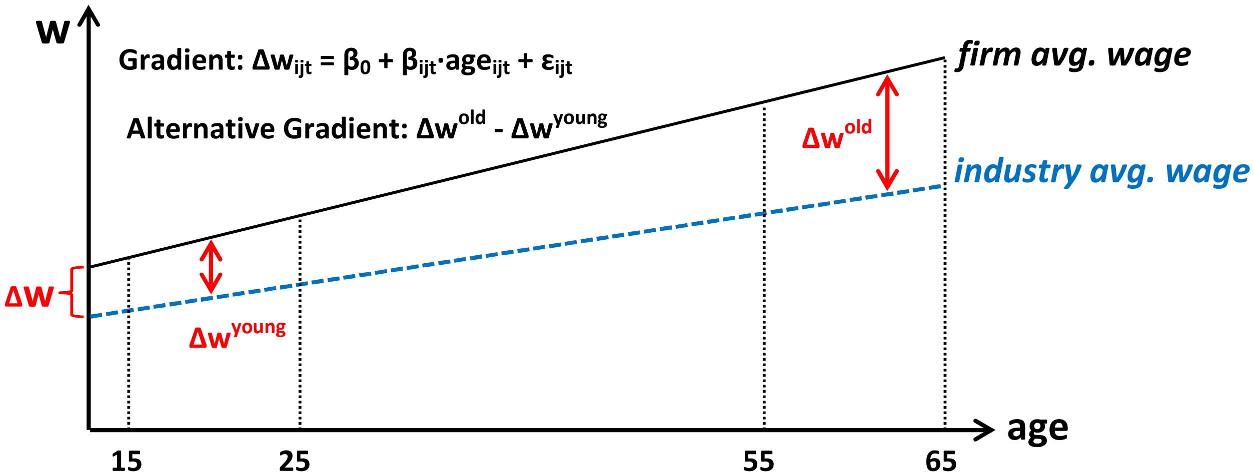
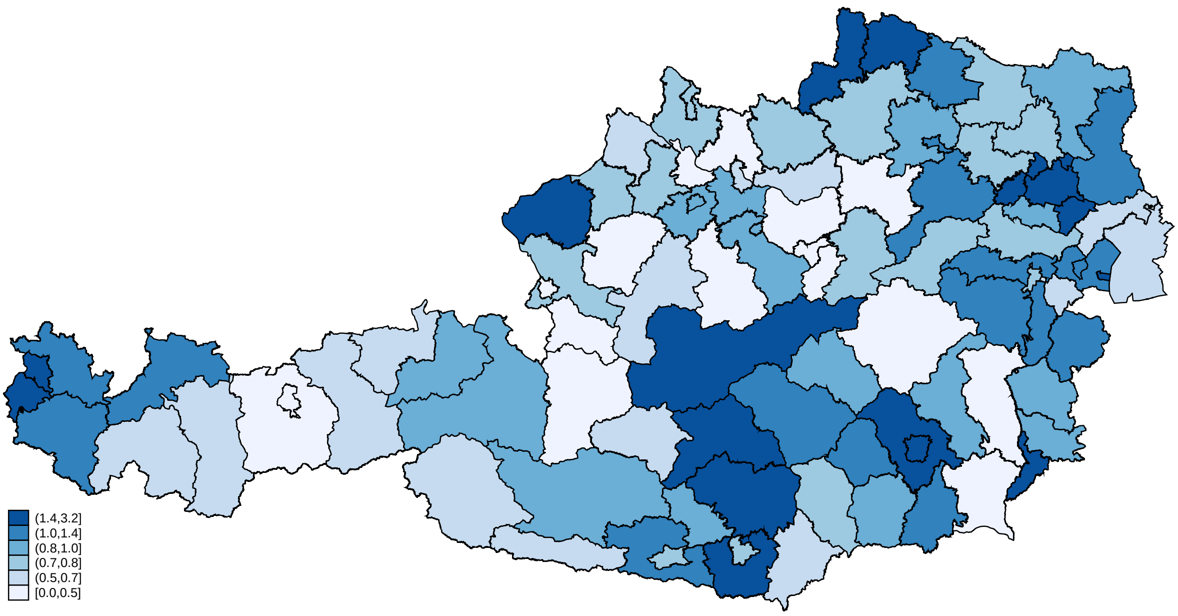
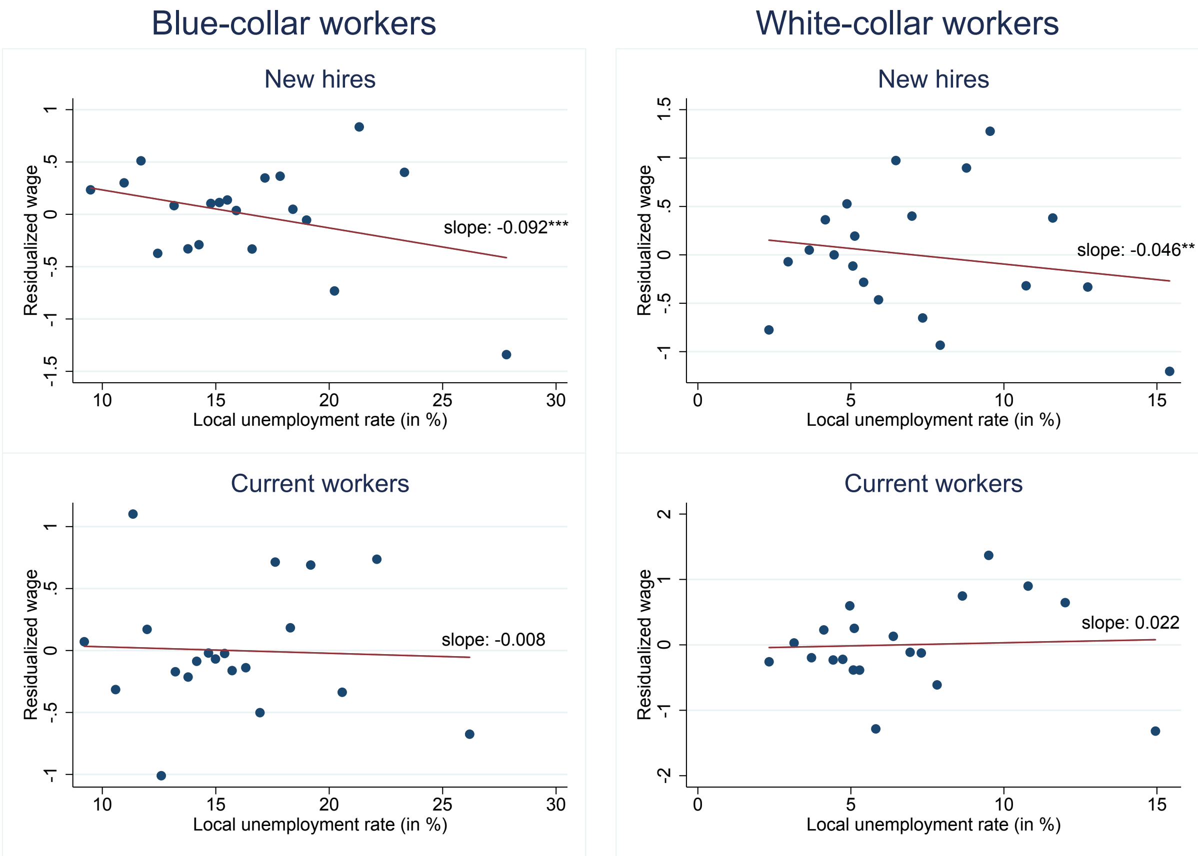


Figure 4: Variation of local unemployment rates 1990-1999



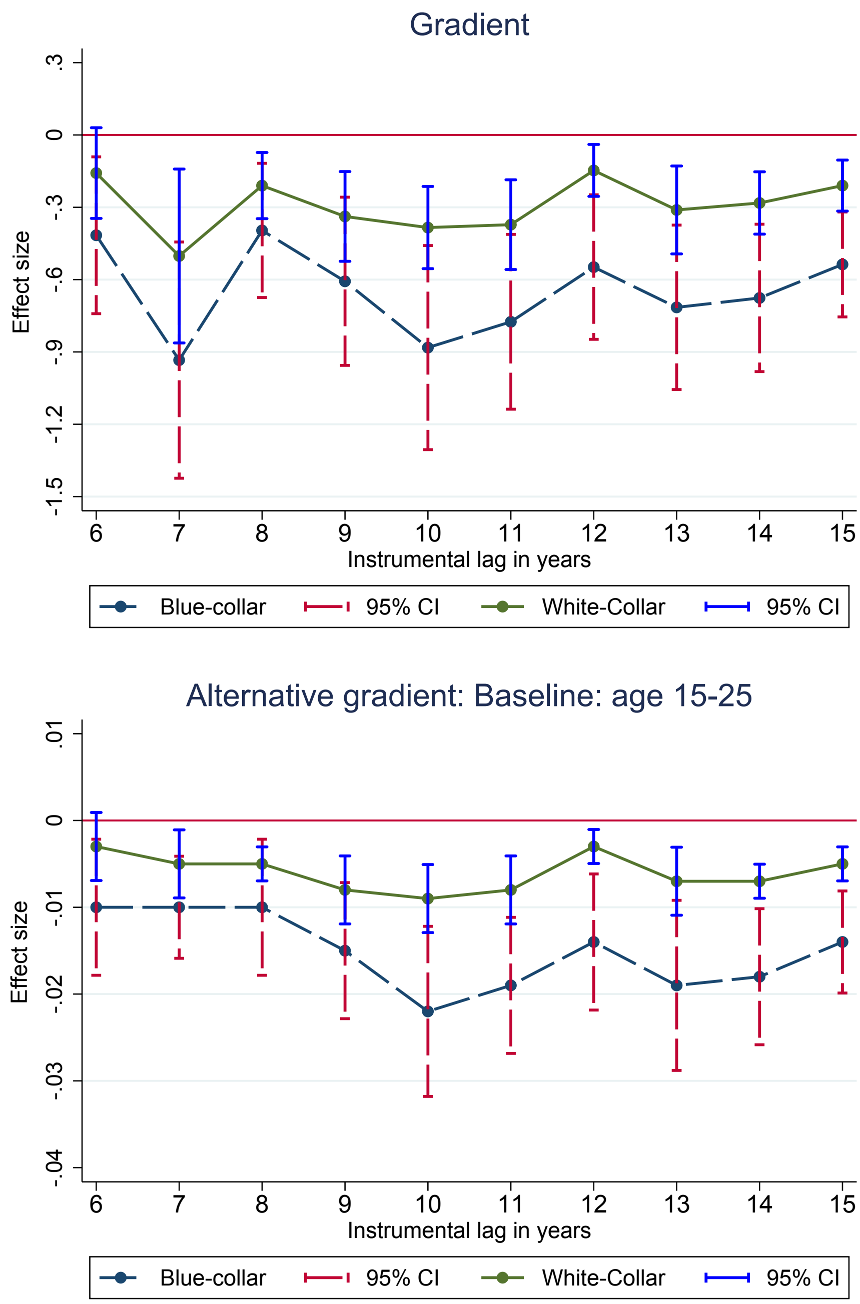
Notes: the map of Austria plots the standard deviation of local unemployment rates and shows the within-district variation of local unemployment rates across the period of 1990 to 1999.

Figure 5: Responsiveness of wages to local unemployment rates between 1989-1999



Notes: Own calculations based on data from *ASSD*. Binned scatterplots show the result of OLS-regressions of wages for newly hired employees and current employees on local unemployment rates as well as individual, regional and firm characteristics, i.e. age, individual productivity, previous sick-leave days, experience, firmsize, indicator for a REBP district, industry, year and regional fixed effects. The slope of the plotted regression line is equal to the estimated coefficient on the local unemployment rates.

Figure 6: The effect of the wage gradient on job exit age: Robustness



Notes: Own calculations based on data from *ASSD*. Figures show estimated 2SLS coefficients of our baseline model for different lags of the instrumental variable as well as an alternative definition of the wage gradient.

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	0.080	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	0.095	-0.062
	(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)
Accrual rate at age 55	0.025	0.021
	(0.095)	(0.043)
Social security contribution months at age 55	359.8	323.5
	(83.45)	(64.96)
Firm size	1,591.5	1,431.7
	(3,866.4)	(3,443.3)
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	-0.008	0.531
	(0.235)	(0.417)
Observations	41,296	45,131

Notes: Standard deviations in parenthesis.

Table 2: Responsiveness of new hires' average characteristics to current local labor market conditions

	Hiring rate	Age	Productivity	Experience	Past sick leave
Local unemployment rate	0.000	0.017	0.000	9.238***	0.002
	(0.000)	(0.011)	(0.000)	(3.182)	(0.068)
Current employees...					
avg. wage	0.002***	-0.004	-0.000	-1.993	0.041
	(0.000)	(0.004)	(0.000)	(1.247)	(0.027)
avg. age	0.002***	0.911***	0.003***	0.430	0.511***
	(0.001)	(0.017)	(0.001)	(4.680)	(0.104)
avg. experience	-0.000***	-0.002***	-0.000***	0.504***	-0.001**
	(0.000)	(0.000)	(0.000)	(0.016)	(0.000)
avg. productivity	-0.106***	-1.885***	0.734***	-720.830***	-16.036***
	(0.011)	(0.342)	(0.014)	(96.344)	(2.103)
REBP district	-0.002	-0.087	-0.005*	-27.451	1.143*
	(0.003)	(0.101)	(0.003)	(30.434)	(0.644)
Firmsize	-0.000***	-0.001***	0.000***	-0.192***	-0.001**
	(0.000)	(0.000)	(0.000)	(0.034)	(0.000)
Blue-collar share	0.072***	0.870***	-0.030***	247.821***	-0.406
	(0.007)	(0.204)	(0.007)	(59.389)	(1.571)
Industry FE	Yes	Yes	Yes	Yes	Yes
Regional FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
Number of obs.	26,079	25,582	25,536	25,569	24,491
R-squared	0.45	0.31	0.59	0.26	0.03
Mean of dep. Var.	0.24	30.78	0.02	3059.67	17.23

Notes: The sample consists of all firms used in our main analysis observed at the time the instrument was measured (between the years 1999-1999). Firm-clustered standard errors in parentheses; * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

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

	(1) OLS	(2) First stage	(3) 2SLS
Wage gradient	-0.029***		-0.882***
	(0.006)		(0.216)
Local unemployment rate (lag 10)		-0.032***	
		(0.007)	
No. of weeks worked at age 55	-0.001	-0.003***	-0.003**
	(0.001)	(0.001)	(0.001)
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***	-0.005***
	(0.001)	(0.001)	(0.002)
Experience (in years)	0.004***	0.006***	0.009***
	(0.001)	(0.002)	(0.002)
Tenure (in years) at age 55	0.000	0.009***	0.009***
	(0.000)	(0.002)	(0.002)
Accrual rate at age 55	-0.021	-0.011	-0.029
	(0.028)	(0.019)	(0.033)
Social security contribution months at age 55	-0.000***	-0.001***	-0.001***
	(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
REBP district indicator	Yes	Yes	Yes
Number of observations	41,296	41,296	41,296
R-squared	0.30	0.20	0.89
Mean of dependent variable	59.78	0.10	59.78
S.d. of dependent variable	1.68	0.53	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			18.71

Notes: This table summarizes OLS estimation results (column 1), first-stage results (column 2) and 2SLS estimation results (column 3) of the effect of the wage gradient on blue-collar worker's job exit age. The local unemployment rate 10 years before job exit serves as an IV for the wage gradient. Standard errors clustered on firms in parentheses, * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 4: White-collar workers: The effect of the wage gradient on job exit age

	(1) OLS	(2) First stage	(3) 2SLS
Wage gradient	-0.006***		-0.384***
	(0.002)		(0.087)
Local unemployment rate (lag 10)		-0.093***	
		(0.020)	
No. of weeks worked at age 55	-0.001	-0.005*	-0.002*
	(0.001)	(0.003)	(0.003)
No. of weeks on sick leave at age 55	-0.001	-0.005*	-0.003**
	(0.001)	(0.001)	(0.001)
No. of weeks out of work at age 55	-0.001	-0.009***	-0.004**
	(0.001)	(0.003)	(0.002)
Experience (in years)	0.001	0.018***	0.008***
	(0.001)	(0.004)	(0.002)
Tenure (in years) at age 55	0.000	0.021***	0.008***
	(0.000)	(0.002)	(0.002)
Accrual rate at age 55	-0.072	0.149	-0.133**
	(0.074)	(0.161)	(0.065)
Social security contribution months at age 55	-0.000***	-0.002***	-0.001***
	(0.000)	(0.000)	(0.000)
Firm size	-0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)
Unemployment rate (lag 1)	-0.054***	0.047**	-0.044***
	(0.004)	(0.022)	(0.008)
Unemployment rate (lag 5)	0.051***	0.059***	0.052***
	(0.004)	(0.019)	(0.008)
Person fixed-effect	0.021**	0.952***	0.382***
	(0.009)	(0.038)	(0.085)
Year of job exit FE	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes
Regional FE	Yes	Yes	Yes
REBP district indicator	Yes	Yes	Yes
Number of observations	45,131	45,131	45,131
R-squared	0.94	0.18	0.84
Mean of dependent variable	60.78	-0.06	60.78
S.d. of dependent variable	1.63	1.51	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			21.08

Notes: This table summarizes OLS estimation results (column 1), first-stage results (column 2) and 2SLS estimation results (column 3) of the effect of the wage gradient on white-collar worker's job exit age. The local unemployment rate 10 years before job exit serves as an IV for the wage gradient.Standard errors clustered on firms 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	-0.738***	-0.509***	-0.344***	-0.173***
	(0.180)	(0.114)	(0.076)	(0.049)
F-test of weak instrument	19.55	24.15	21.81	17.83
Number of observations	27,025	19,989	32,527	25,793

Notes: This table summarizes 2SLS estimation results for our baseline model for a sample of workers who have been employed by the firm before the IV is measured. Each regression includes all covariates of our baseline model. Standard errors clustered on firms in parentheses, * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 6: Alternative outcomes

	Blue-collar	White-collar
Disability	-0.099	0.034*
	(0.088)	(0.018)
Mean of dependent variable	0.31	0.08
Phased retirement	0.006	-0.025
	(0.125)	(0.045)
Mean of dependent variable	0.17	0.22
Golden handshake	0.214***	0.055
	(0.063)	(0.033)
Mean of dependent variable	0.08	0.15
F-test of weak instrument (lag 10)	18.71	21.08
Number of observations	41,296	45,131

Notes: This table summarizes 2SLS estimation results for our baseline model with alternative outcome variables. A 10-year lag of the local unemployment rate serves as an IV for the wage gradient. Each regression includes all covariates of our baseline model. Standard errors clustered on firms in parentheses, * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$;

Table 7: The effect of the wage gradient by work incentives

	High work incentives			Low work incentives		
	Job exit age	Disability	Golden handshake	Job exit age	Disability	Golden handshake
Blue-collar workers						
Wage gradient	-2.191***	-0.126	0.355***	-0.167**	-0.096	0.158***
	(0.667)	(0.124)	(0.121)	(0.079)	(0.103)	(0.060)
F-test of weak instrument	10.83	10.83	10.83	20.29	20.29	20.29
Alternative wage gradient	-0.057***	-0.003	0.009***	-0.004**	-0.002	0.004**
	(0.018)	(0.003)	(0.003)	(0.002)	(0.003)	(0.001)
F-test of weak instrument	10.02	10.02	10.02	23.23	23.23	23.23
Number of observations	20,050	20,050	20,050	21,246	21,246	21,246
R-squared	0.58	0.39	0.32	0.95	0.43	0.02
White-collar workers						
Wage gradient	-0.613***	0.043**	0.109***	-0.055	0.014	-0.013
	(0.133)	(0.021)	(0.041)	(0.044)	(0.027)	(0.043)
F-test of weak instrument	22.65	22.65	22.65	11.19	11.19	11.19
Alternative wage gradient	-0.014***	0.001**	0.002***	-0.001	0.000	-0.000
	(0.003)	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
F-test of weak instrument	25.19	25.19	25.19	14.73	14.73	14.73
Number of observations	23,044	23,044	23,044	22,087	22,087	22,087
R-squared	0.67	0.17	0.08	0.95	0.24	0.04

Notes: This table summarizes 2SLS estimation coefficients for the wage gradients with 10-year lag of the instrument by individual work incentives. The high (low) work incentive sample consists of individuals with above (below) average accrual rates. Each regression includes all covariates of our baseline model except of individual's accrual rate. Standard errors clustered on firms in parentheses; * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$;

Table 8: Pathways into retirement

Outcome variable	Blue-collar	White-collar
Job exit age	-0.882***	-0.384***
	(0.216)	(0.087)
Formal retirement age	-0.175***	-0.081***
	(0.061)	(0.024)
Years between job exit and retirement	0.707***	0.303***
	(0.173)	(0.069)
Mean of dependent variable	0.09	0.08
Years of unemployment after job exit	0.637***	0.275***
	(0.156)	(0.063)
Mean of dependent variable	0.08	0.07
Years of out of labor force after job exit	0.050***	0.025***
	(0.017)	(0.010)
Mean of dependent variable	0.01	0.01
F-test of weak instrument	18.71	21.08
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

Notes: This table summarizes 2SLS estimation coefficients of the wage gradient for a set of outcomes describing the pathways into retirement. Each regression uses a 10-year lag of the IV and includes all covariates of our baseline model. Standard errors clustered on firms in parentheses, * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$;

Web Appendix

This Web Appendix (not for publication) provides additional material discussed in the manuscript 'Seniority Wage and the Role of Firms in Retirement' by Wolfgang Frimmel, Thomas Horvath, Mario Schnalzenberger, and Rudolf Winter-Ebmer.

A.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 do not consider apprentices or marginally employed 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.

A.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[1]:

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

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.[2]

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.

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

[2]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 A.1: 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.

A.3 Graphical representation of the first stage

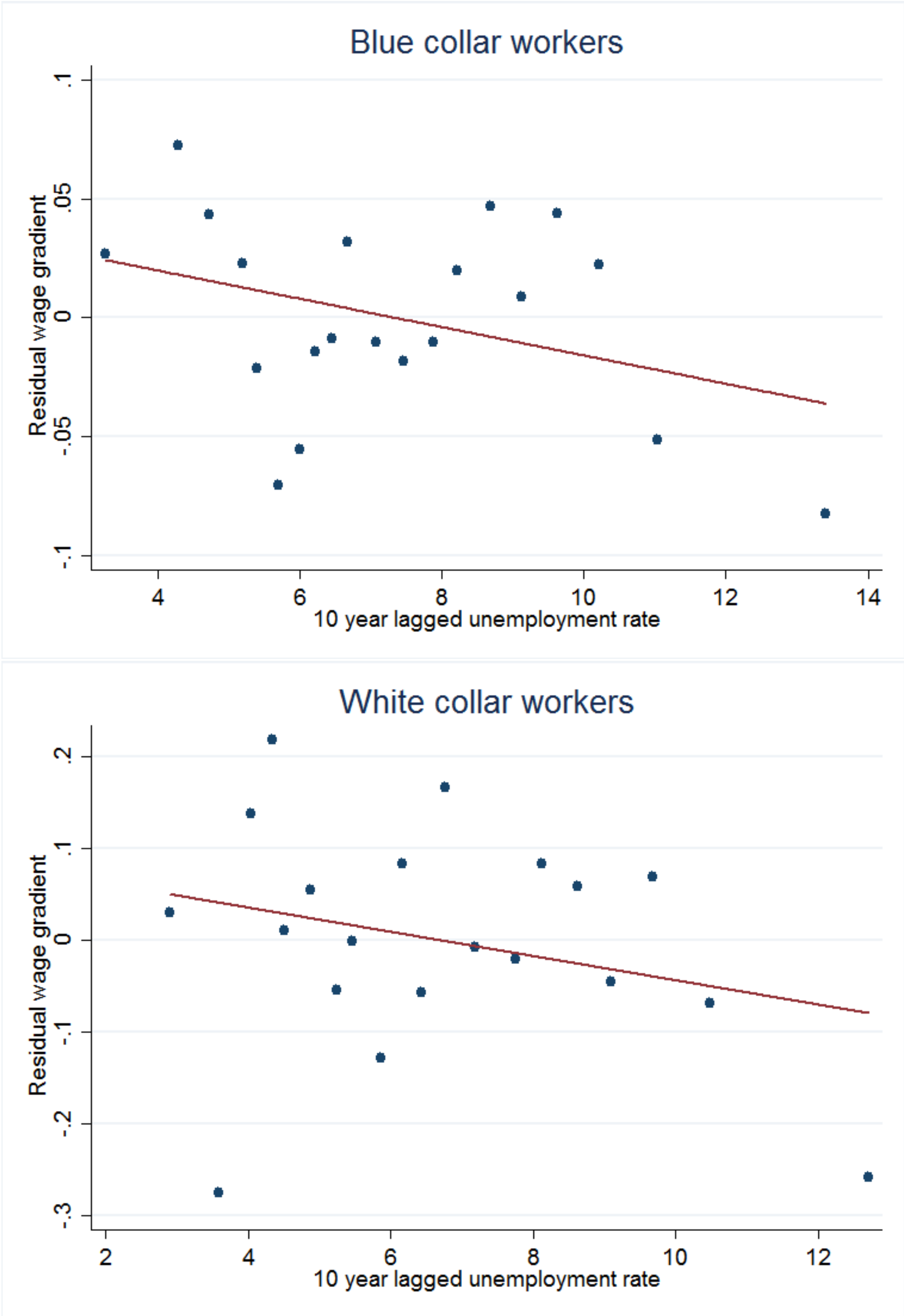


Figure A.1: Wage gradients and lagged local labor market conditions

A.4 Plausibility checks for instrument validity

Table A.2: Job exit by mass layoff or firm closure and past local unemployment rate

	Blue-collar workers		White-collar workers	
	Mass layoff	Firm closure	Mass layoff	Firm closure
Local unemployment rate (lag10)	0.000	-0.001	0.001	-0.001
	(0.003)	(0.001)	(0.003)	(0.001)
Covariates	Yes	Yes	Yes	Yes
Number of observations	41,296	41,296	45,131	45,131
R-squared	0.04	0.01	0.03	0.01
Mean of dependent variable	0.109	0.011	0.084	0.005
Mean of unemployment rate	7.328	7.328	6.677	6.677
S.d. of unemployment rate	2.455	2.455	2.493	2.493

Notes: This table summarizes OLS estimates of a model, where we regress a binary indicator for job exit due to mass layoff or firm closure respectively, on past local unemployment rate and a set of covariates. The indicator for a job exit through a mass layoff is equal to one if the workforce is reduced by at least 10 percent in the year of individual's job exit. The indicator for a firm closure is equal to one if individual's firm closed in the year of the job exit. We use the same set of covariates as in our main estimations. Standard errors clustered on firms in parentheses, * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A.3: Robustness check: Add firm characteristics at the time of the instrument

	Blue-collar Workers		White-collar Workers	
	Gradient 1	Gradient 2	Gradient 1	Gradient 2
Wage gradient	-1.043***	-0.026***	-0.326***	-0.007***
	(0.239)	(0.006)	(0.060)	(0.001)
Baseline covariates	yes	yes	yes	yes
Firm characteristics (lag 10)	yes	yes	yes	yes
Number of observations	41,296	41,245	45,131	44,972
R-squared	0.87	0.86	0.88	0.88
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 growth gradient	0.099	3.44	-0.058	-4.776
S.d. of wage growth gradient	0.529	22.43	1.508	62.809
F-test of weak instrument	20.766	19.608	31.146	33.176

Notes: This table replicates Tables 3 and 4 and additionally includes the following firm characteristics measured at the time of the instrument, hence 10 years before individuals' job exit: firm size, hiring rate, share of blue-collar workers, as well as average characteristics of newly hired employees 10 years ago: average tenure, age, productivity, experience and previous number of sick leave days. Clustered standard errors in parentheses, * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A.4: Responsiveness of quit probability and disability to labor market conditions

	Blue-collar workers		White-collar workers	
	Quit job	Disability	Quit job	Disability
Unemployment rate	0.000299*	-0.000	0.000	-0.000083*
	(0.000)	(0.000)	(0.000)	(0.000)
Age	-0.001***	0.000***	-0.001***	0.001***
	(0.000)	(0.000)	(0.000)	(0.000)
Sickness days (last 3 yrs)	0.000***	0.000***	0.000	0.000***
	(0.000)	(0.000)	(0.000)	(0.000)
Personal FE	0.034***	-0.014***	0.024***	-0.008***
	(0.002)	(0.000)	(0.002)	(0.000)
Experience in days	-0.000***	0.000***	-0.000***	0.000***
	(0.000)	(0.000)	(0.000)	(0.000)
REBP district	0.000	0.001***	0.005**	0.001
	(0.001)	(0.000)	(0.002)	(0.000)
Firm size	-0.000*	0.000	0.000**	0.000
	(0.00)	(0.000)	(0.000)	(0.000)
Industry FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Regional FE	Yes	Yes	Yes	Yes
Number of obs.	1,197,503	1,197,503	511,093	511,093
R-squared	0.04	0.01	0.06	0.007
Mean of dep. var.	0.22	0.004	0.09	0.005

Notes: This table summarizes OLS estimations for the likelihood to quit a job or leave via disability insurance within a year after we measure the local unemployment rate. The sample consists of all blue- and white-collar workers with a minimum tenure of 1 year, who are employed in our firms in the years 1990 to 1999. Employees quit a job if they immediately switch to another job or are not officially recorded as unemployed for two months. Standard errors clustered on firms in parentheses; * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A.5: New hires' health-care expenditures and local labor market conditions

	(1) Doctors	(2) Drugs	(3) Hosp.days	(4) GP	(5) Neuro/Psycho	(6) Radiology	(7) Lab	(8) Psychotherapy
Unemployment rate	-0.026	-0.865	-0.019	-0.179	0.000	-0.036	-0.025	0.023
	(0.356)	(0.564)	(0.021)	(0.202)	(0.027)	(0.040)	(0.026)	(0.021)
Current employees...								
avg. wages	-0.207	-0.203	-0.002	-0.062	0.012	-0.014	-0.002	0.008
	(0.150)	(0.135)	(0.014)	(0.071)	(0.015)	(0.016)	(0.010)	(0.010)
age	-0.203	-0.995	0.009	0.099	-0.036	0.076	0.033	-0.164*
	(0.742)	(0.953)	(0.048)	(0.451)	(0.068)	(0.112)	(0.062)	(0.084)
experience	0.002	0.004	-0.000	0.000	0.000	-0.000	-0.000	0.000
	(0.002)	(0.003)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)
productivity	-9.767	-31.931*	-0.446	-5.778	-1.094	-0.410	0.568	0.345
	(15.050)	(19.319)	(0.704)	(7.519)	(1.085)	(1.751)	(1.218)	(0.782)
New employees...								
avg. entry wages	0.033	-0.269	0.001	-0.047	-0.008	0.005	0.015	-0.002
	(0.145)	(0.198)	(0.007)	(0.079)	(0.017)	(0.015)	(0.010)	(0.010)
age	1.587**	1.840**	-0.012	0.624*	0.117*	0.130	0.054	0.175
	(0.617)	(0.869)	(0.031)	(0.369)	(0.061)	(0.092)	(0.044)	(0.112)
experience	-0.000	-0.001	0.000*	0.001	-0.000	-0.000	0.000	-0.001
	(0.002)	(0.002)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)
productivity	3.323	34.716	-0.941	-0.299	-1.376	-1.501	0.099	-0.475
	(15.529)	(21.245)	(0.657)	(8.395)	(1.255)	(1.629)	(1.285)	(1.003)
Firmsize	0.000	0.001	-0.000*	-0.000	0.000	0.000	0.000	0.000
	(0.001)	(0.001)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)
Blue-collar share	-9.584	-18.123*	0.362	-2.039	-0.857*	-0.895	-0.215	-0.222
	(6.747)	(9.564)	(0.404)	(3.284)	(0.510)	(0.936)	(0.630)	(0.198)
Industry FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Regional FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
REBP district indicator	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Number of obs.	2,815	2,815	2,815	2,815	2,815	2,815	2,815	2,815
Mean of dep. Var.	43.40	24.93	0.78	25.67	0.85	2.59	1.42	0.32

Notes: This table summarizes results for the responsiveness of average health-care expenditures of newly hired employees to local unemployment rates. Outcome variables are average expenditures (in euros) for aggregate outpatient medical attendance (column (1)), aggregate amount for medical drugs (column (2)), number of days spent in hospital (column (3)), as well as expenditures for GPs, neurologists/psychologists, radiologists, laboratory services and psychotherapy (columns (4) to (8)). Data on health-care expenditures is available for the sub-set of firms that are located in the province of Upper Austria in the year 1998, and are measured in the year before employees were hired. Standard errors in parenthesis, * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$