Pension Reforms and their Implications for Establishment Downsizing☆

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Abstract

While the empirical literature on the effects of pension reform on workers is broad, less is known about the impact on employers. Yet reforms that create incentives to postpone retirement may have extensive effects on employer labor demand and labor costs, especially in settings where there are strict legal protections against age discrimination in employment. Although public pension system reforms generally are structured to treat all workers within the same birth cohort similarly, the impact on employers may vary substantially due to differences in the age composition of their employees. Using this variation as a source of identification, we examine whether the differential impact of pension reform leads to differences in the incidence of workforce downsizing, a sign of possible financial distress. To ensure estimates are not biased due to attrition, we also model associations between the impact of pension reform and establishment closures and find no association. Results for downsizing consistently show establishments with a higher share of older workers are more likely to experience downsizing. When we segment workers within establishments by age, the absolute changes in downsizing probabilities are highest for younger workers. Preliminary results indicate works councils may increase the risk of downsizing for older workers and protect employment for young and prime workers.

Keywords: Survival Analyses, Pension Reform, Downsizing, Admin Data

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1. Introduction

Workforces are aging across the globe and a majority of OECD nations are raising ages of eligibility for public pension benefits currently or will in the near future (OECD, 2011). These changes incentivize later claiming of pension benefits, making retirement at younger ages less financially attractive. The political debate over consequences of these reforms focuses mostly on the employee. However, employers are affected by these changes as well. Changes in pensionable age are generally phased in gradually across birth cohorts so the timing and extent of the reform's impact will vary with the cohort composition of the employer's workforce. This means idiosyncratic differences in the shares of workers affected by the reform can create variance in the degree of disruption of normal retirement patterns the policy creates across employers. Employers experiencing policy impacts may be more likely to actively manage workforce aging through incentives to retire earlier (e.g. buyouts) or may counterbalance postponed retirements with layoffs of less senior employees, who are generally younger. Both will result in downsizing of the establishment workforce.

In this paper, we investigate workforce downsizing associated with pension reforms. We hypothesize downsizing will be more likely in establishments that experience larger impacts of pension reform, and will be more likely to impact younger workers because there are fewer legal barriers to their dismissal. Finally, we anticipate any downsizing effects may be mitigated by works councils, who are likely to advocate for cost saving strategies that preserve employment.

Prior literature clearly establishes pension reforms influence worker retirement behavior in expected directions, though the magnitude of the effects can be small in settings where private savings are a large component of overall retirement wealth (Berkel and Borsch-Supan, 2004; Atalay and Barrett, 2015; Gustman and Steinmeier, 2009; Maestas and Zissimopoulos, 2010). Postponed retirements have the potential to impact firm profitability directly through increased labor costs and indirectly through possible productivity effects. Because the acquisition of human capital is related to age and tenure within establishments, shifts in the labor force participation of older workers due to pension reform may change the composition of the workforce in ways that affect firm performance and the risk of downsizing. For example, research on firm productivity shows that firm age and human capital are key determinants of firm productivity and profitability (Vandenberghe, 2013; Audretsch and Fritscht, 1994; Barron et al., 1994; Dunne et al., 1988; Lane et al., 1999; Mahlberg et al., 2013; Schnabel and Wagner, 2012). This potential linkage is also supported by evolutionary economics, which proposes that establishments make decisions under constraints and that the strategies firms adopt vary with these constraints (Alchian, 1950). Firms adopting strategies poorly suited to the conditions of their external environment will be eliminated through competition. Changing pensionable ages represents a shift in the constraints firms face. Assuming firms had optimized their workforce size prior to a pension reform, downsizing may be a necessary strategy to counteract the incentive the reform created for postponed retirements. Therefore, establishments that are differently affected by pension reforms should adopt different strategies.

Currently there is little known about the potential effects of pension reforms on firms' labor demand. There are at least two channels through which firms could adjust their workforce composition after a pension reform takes place: hiring and downsizing. In the context of this German reform, a working paper by Eckrote (2019) finds that establishments with larger shares of older workers reduce their hiring when pensionable age is raised, with the reduction mostly impacting young workers. A recent working paper by Bovini and Paradisi (2019) investigates layoffs after a change in pensionable ages in Italy. They find delayed retirements increase layoffs among workers of all ages. The German context may differ from the Italian context due to the importance of works councils and additional protections in place for workers, especially more senior workers (Bhankaraully, 2019). In a related study that supports this assertion, Muñoz-Bullón and Sánchez-Bueno (2014) find downsizing among Spanish firms is associated with both labor law and behavior of industry peers.

Studying the effect of pension reform on employers is challenging because there are few data sources that contain demographic information needed to infer pensionable age for all workers across many firms. Where these data do exist, they often do not encompass a long enough time series to estimate the effects of a gradual increase in pensionable age or contain a large enough sample of firms to examine heterogeneity. Finally, the reforms themselves are often phased in so slowly that they do not create enough variation to convincingly separate employer responses from other factors.

We address these problems using over two decades of administrative data from German social security notifications encompassing a 1992 reform that raised pensionable ages by 5 years. This reform was fully phased in over a span of 19 years and 8 birth cohorts. For comparison, the increase in age of eligibility for Social Security benefits in the US from age 65 to 67 was announced in a 1983 amendment, did not begin to bind on the first affected cohorts until 2003, and included an 11 year hiatus during which the age remained constant at 66 for cohorts born in 1943 through 1954 (Social Security Administration, 2019). The full two year increase will not bind for US workers until the 1960 birth cohort reaches age 67 in 2027. So, while the relationships between pension reform and downsizing we hypothesize are also relevant in other countries, the rich administrative data available in Germany coupled with larger and more expedient changes in pensionable age provide a uniquely advantageous study setting. Our data follow a representative sample of West German establishments existing in 1990 through 2010. We find in establishments where the reform lead to a larger share of workers over the age of 58 than otherwise similar establishments, downsizing is more likely. The effects are largest for establishments without works councils. Although the percentage point changes in the probabilities of downsizing are similar across age segments of the workforce, this means the relative risk increases far more for older workers who generally have the lowest risk of downsizing.

2. The German Pension System and the 1992 Reform

The German pension system is designed as a pay-as-you-go scheme, providing pension benefits for all private and public sector employees entitled to social security.[1] It covers about 90 percent of the German workforce (Richter and Himmelreicher, 2008) and accounts for approximately 85 percent of retirement income (Börsch-Supan, 2000). Public pension accrual is a function of one's wages relative to countrywide average wages, years of service, and age, calculated every year. In 2005, estimates indicate that less than five percent of households headed by older workers had private pensions, despite incentives for private savings introduced in the 2001 Riester Reform (Börsch-Supan, 2000). The German pension system has historically offered a "window" of ages at which workers can begin claiming pension benefits, beginning as early as age 58 if combining early retirement with the unemployment insurance (Börsch-Supan and Wilke, 2004). Statistics show that in years 1993 to 1995, at age 60, approximately 60 percent of German men had retired and 45 percent were receiving pension benefits.

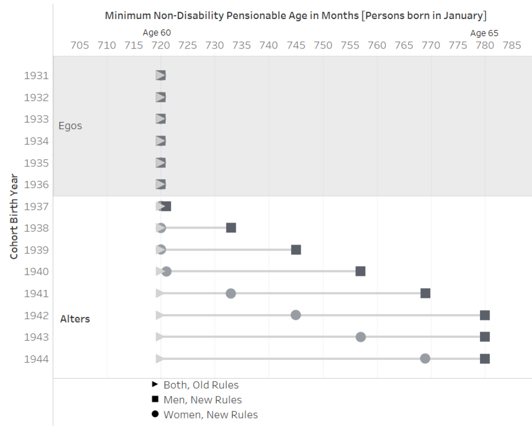
In 1992, a reform gradually increased ages of eligibility for full benefits to 65. This reform was the first in a 15 year period of pension reform in Germany (Bonin, 2009). The goal was to stop access to full pension benefits at age 60 for persons born after 1936. Figure 1 provides a graphical overview of the minimum pensionable ages of the historical pension insurance compared to after the 1992 reform. As explained above, workers were able to effectively retire up to two years before the ages depicted in Figure 1 if they claimed unemployment benefits to bridge the gap between the end of employment and claiming. So, raising the age of claiming above 60 should lead to an increase in labor supply at age 58 and above.

Prior research finds, as of 2004, the 1992 reform lead to a two-year increase in the average retirement age among men and a nine-month increase among women

[1] Self-employed workers and civil servants are excluded from the pension system.

Figure 1: Summary of 1992 Reform of Pensionable Ages

By Cohort



Source: Illustration taken from Hamman et al. (2019)

(Berkel and Borsch-Supan, 2004). The 1992 reform first began to postpone claiming benefits as early as 1 year after it was announced. All changes in eligibility for full benefits were phased in between the 1937 and 1944 birth cohorts and thus were fully implemented by 2011. Overall, the 1992 reform created differences in pensionable ages of 1 to 12 months across adjacent birth cohorts and differences of 6 to 12 months for men and women within the 1940 through 1941 cohorts (Börsch-Supan and Wilke, 2004).

This outlined heterogeneity in incentives to retire among older workers creates variation in the impact of the reform across employers. Small differences in the age distribution may lead to large differences in retirement patterns. In total, the reform should lead to an increase in the share of workers who continue to work past age 58, and this increase should be largest in establishments that employ more workers from the affected cohorts, and during the later years in the phase in period. We use this heterogeneity as a source of identifying variation to estimate the impact of pension reforms on establishment survival (as outlined in Section 5).

3. Analyses Sample

We use the Linked-Employer-Employee Data (LIAB) [cross-sectional model 2 1993-2014 (LIAB QM2 9314)] from the Institute for Employment Research which is provided for academic research use.[2] The LIAB matches administrative employment records to establishment survey information. Baseline for the sampling is the IAB Establishment Panel (IABBP), which collects data on about 15,500 establishments per year (Fischer et al., 2009). In the LIAB QM2 9314 all individuals who work in these surveyed establishments on June 30th in each year are sampled. For each of these workers we know their employment state on June 30th, as well as a rich set of variables describing the employment characteristics, including wages, detailed occupations, and industry. Socio-demographic variables including sex, age and education attainment are included as well (Klosterhuber et al., 2016). Having all the workers in each establishment in a given year allows us to aggregate individual information on an establishment level and describe the entire workforce in an establishment. These administrative records can be combined with information from the IABBP, which allows us to add information on establishments legal entities, personnel policies and operating strategies to the analyses.

The LIAB data begin in 1993, which is after the 1992 reform was introduced. Thus, to construct a measure of policy impact that is exogenous to any policy response, we need data that includes pre-policy information. We use a custom extract from the Employment History data (BeH) provided by the FDZ[1], which contains age distributions by gender for each establishment in the LIAB QM2 9314 that existed in 1990. Thus, our analytic sample is based on all establishments which are part of the LIAB QM2 9314, but also existed in

[2]Data access was provided via on-site use at the Research Data Centre (FDZ) of the German Federal Employment Agency (BA) at the Institute for Employment Research (IAB) and subsequently remote data access.

[1]We thank Andreas Ganzer for sampling the data for us and supporting us with deidentification of the data.

1990. This restricts our sample to West German establishments, because data on East German establishments is lacking prior to 1993. Overall, we can follow 74,985 establishments during the time period from 1993 to 2010. However, this is an unbalanced sample (cross sectional yearly sample) as we only have an observation for the years the establishments participated in the survey.

4. Indicators of Establishment Downsizing

We are measuring downsizing by relying on the wide literature of displacement studies. There are different ways that this strand of literature constructs measures, such as layoffs, outflow, turnover, etc (Lengermann and Vilhuber, 2002; Jacobson et al., 1993; Dustmann and Meghir, 2005; Bowlus and Vilhuber, 2002; Abowd et al., 2009). This paper relies on establishment wide and age specific net employment following Flaaen et al. (2017).

We measure overall downsizing on the establishment level as follows, whereas the downsizing rate D in year t at establishment j is defined as the count of total workers (EMP(t)) in establishment j at the end of June in year t, divided by the number of total workers (EMP(t+1)) in establishment j at the end of June in the following year. Because small changes in the workforce in small establishments can result in large percentage changes in employment, we require the total workforce in an establishment in year t has to be at least 50 workers to be included in the analysis.

$$D_{jt} = \frac{EMP(t)_{jt}}{EMP(t+1)_{jt}}$$

As for our main definition, we also define an age specific downsizing indicator for the alternative measure for younger, prime and older workers:

$$D_{jta} = \frac{EMP(t)_{jta}}{EMP(t+1)_{jta}}$$

The downsizing rate D in year t of workers in age group a at establishment j is defined as the count of total workers (EMP(t) in age group a in establishment j at the end of June in year t, divided by the number of total workers (EMP(t+1)) in age group a in establishment j in the following year. The total workforce in an establishment in year t has to be at least 50 workers in order to be able to experience downsizing.

Conceptually, these measures reflect a segmentation of establishment employees by age and allow us to capture cases where one age segment experienced a large outflow of workers relative to their age group's total employment, but the firm as a whole may not appear to have downsized because that segment's share of total employment is relatively small.

After calculating D_{jt}, respectively, D_{jta} we construct outcomes to study the effect of pension reform at different thresholds of downsizing: 10%, 20%, and 30%. In each of the cases the outcome y takes on the value 1 if the downsizing measure (D_{jt} or D_{jta}) is higher than the thresholds, and is zero otherwise. We end up with 12 different outcome variables.

5. Empirical Strategy

5.1. Estimation of Reform Impact on Downsizing

We measure the impact of the reform via the share of workers 58 years and older in an establishment, as the pension reform in 1992 leads to a higher share of older workers in establishments. Using variance in the share of workers over 58 attributable to the reform, we hypothesize establishments with more workers working past the old effective retirement age of 58 will consequently show a higher risk of workforce downsizing. Thus, we construct yearly shares of employees in each establishment age 58 and older, $share58_{jt}$. We use age 58 as the threshold because, as explained, this was the earliest age in the pre-reform retirement window that workers could finance retirement through a combination of pension and unemployment benefits. The estimated OLS equation will then be

$$y_{jta} = \beta_1 share58_{jt} + \beta_2 X_{jt} + u_t + \epsilon_{jt}, \tag{1}$$

Our main outcomes y_{jta} are binary indicators for whether an establishment j in year t experiences a downsizing of workers in age group a of at least 10%, 20% or 30%. $share58_{jt}$ is the share of workers in establishment j who are age 58 and older in year t. X_{jt} is a vector of establishment controls including, industry, inflows, outflows, establishment size, legal form, existence of collective bargaining agreements and work councils. Furthermore, we include year fixed effects u_t.

Our estimates of β_1 in Equation 1 could be attenuated if our hypothesized linkage between policy induced postponement of retirements and firm financial distress is correct and some firms opt to fully shut down rather than downsize. To check for this possibility, we also model establishment closures using the same functional form as Equation 1 above and substituting a binary indicator of closure for the dependent variable y_{jta}.

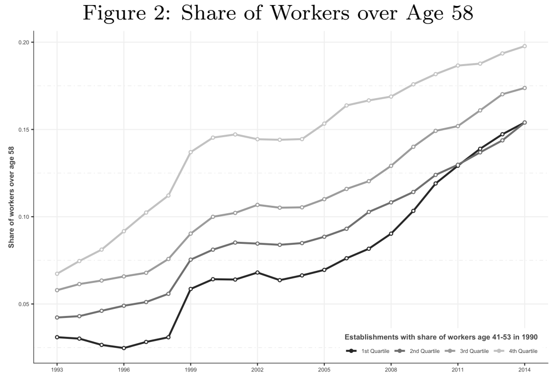
Our estimates of β_1 in Equation 1 could be also be biased if the employers who have higher shares of older workers are more likely to engage in practices, like buyouts, that also impact the probability of downsizing. In the case of buyouts, if buyouts are more common among employers with more older workers (or in years when the share of older workers is high), and if buyouts of the oldest workers reduce the probability of downsizing in other age groups, then our OLS estimate of β_1 would be negatively biased. The importance and durability of firm specific human capital is another source of possible bias. Employers where firm specific human capital is important and highly durable tend to retain their older workers. They also may, by virtue of their specific human capital, have competitive advantages that reduce the likelihood of buyouts. If so, these unobserved factors would also lead to negative bias in β_1. The overall desirability of the employer to workers would also lead to negative bias. To address these potential sources of bias, we instrument the share of workers age 58 and older with what we call our ingap measure z_ingap_{jt}. It represents the number of workers in the gap between the old pensionable age and the lowest new pensionable age, based on projected workforce estimates. It is constructed following a shift share approach which is outlined in the following section.

5.2. Shift Share Instrument Construction

We use the 1990 BeH custom data extract containing pre-policy information to construct a shift share instrument. Shift-share instruments, sometimes called "Bartik instruments" after Bartik (1991), have been widely used in the immigration and the regional growth literature but have many other applications (Goldsmith-Pinkham et al., 2018).

We first predict differences in the shares of employees eligible to retire in each establishment, relative to industry-average shares in each post policy year, that are attributable only to the differences in pre-policy employment of cohorts affected by the reform using pre-reform employment information. Specifically, we construct counts of workers in each affected cohort by sex in each of the establishments in our analytic sample. These counts comprise the "share" portion of the instrument.

Figure 2: Share of Workers over Age 58



Source: Authors' calculations

To demonstrate the relevance of these shares for predicting future workforce aging, Figure 2 shows the correlation between 1990 shares of workers across all cohorts who will be affected by the reform by 2014 (those aged 41 to 53 in 1990) and actual shares of workers over age 58 across the subsequent years. Whereas all establishments experience a growing older workforce, we can see that establishments that employ fewer affected workers before the reform do so to a lower extent, even 24 years later. The more detailed shares we construct by sex and single year birth cohort are also strongly correlated with the employment of older workers in subsequent years.

The shifts are computed from the fitted values after estimating the following two regressions using 1993-2014 data separately for each of 11 industry sectors by sex.

$$begin_{ijt} = \beta_0 + \beta_1 age_{ijt} + \beta_2 year_t + \beta_3 age_{ijt} * year_t + \epsilon_{ijt},$$

$$end_{ijt} = \beta_0 + \beta_1 age_{ijt} + \beta_2 year_t + \beta_3 age_{ijt} * year_t + \epsilon_{ijt},$$

Where $begin_{ijt}$ is equal to 1 for employees in their first year of employment with establishment j in year t and equal to 0 for all subsequent years. end_{ijt} is equal to 1 in the last year of employment with establishment j, which is indicated when the employer files an end of employment notification. age_{ijt} is a vector of age dummy variables for ages 19 through 67 with age 18 as the omitted age group, and $year_t$ is a vector of dummy variables for years 1994 through 2014 with 1993 as the omitted year. After estimating each equation for men and for women by industry, we obtain fitted values $\widehat{begin_{ijt}}$ and $\widehat{end_{ijt}}$ at each age for each year from each equation.

We then take the averages of these fitted values for each age in each year. This yields estimates of the probabilities of being in the first year of employment with establishment j in year t conditional upon working for establishment j and of ending employment with establishment j in year t conditional upon working for establishment j in each of the 11 industry sectors for men and for women. We use these probabilities to "age" the 1990 workforce for each establishment as follows:

$$workers_{at} = workers_{a-1,t-1} * [1 - \widehat{end_{a-1,t-1}} + \widehat{begin_{at}}]$$

Where the number of workers age a in year t is equal to the number of workers at age a-1 from the prior year adjusted by the probabilities of ending employment in the prior year at age a-1 and beginning employment at age a in year t. $workers_{at}$ is computed for each age separately for men and women by industry sector.

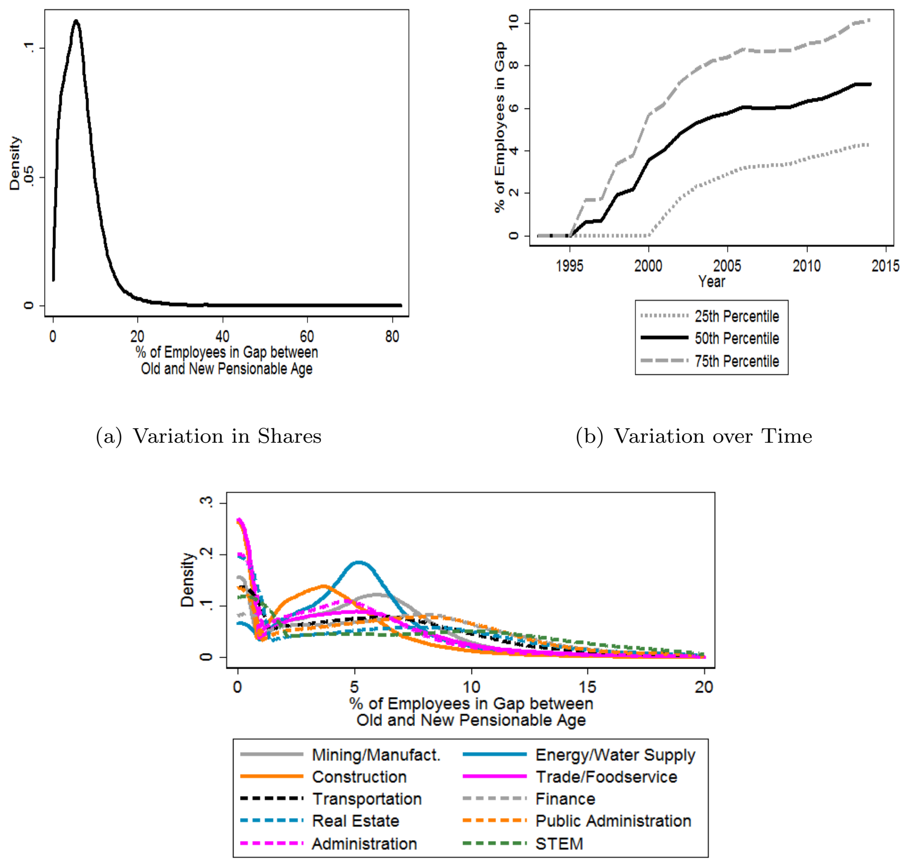
Using $workers_{at}$, we calculate the number of workers in the gap between the old pensionable age and the lowest new pensionable age where full benefits can be claimed without disability for each establishment j in each year t. We then divide those counts by the size of the establishment workforce in 1990. The resulting z_ingap_{jt} is our instrument.

5.3. Validity of the Instrument

Recent studies raise concerns about the validity of these instruments. First, to meet the exclusion restriction, the initial shares used to construct the instruments must be exogenous. In our setting, this means the shares of employees in each establishment who are affected by the change in pensionable age must be exogenous to future survival probabilities of the establishment. To ensure this, we measure the shares of employees in affected cohorts before the policy is announced.

Second, there must be sufficient variation in initial shares to ensure the instruments for units receiving the same shift will be different. In our setting, this means the distributions of workers from different birth cohorts and of different sexes within the same cohort must vary across establishments in the same industry. Figure 3 provides visual proof of sufficient variation of our instrument.

Figure 3: Validity of Instruments



(a) Variation in Shares

(b) Variation over Time

(c) Variation across Industry

Panel a) shows our projected measure puts most establishments between 0 and 20 percent of workers in the gap between the old and new pensionable age. Panel b) focuses on the 0 to 20 percent range and displays variation separately by industrial sector. Agriculture is omitted because the density at 0 is over 1.5, far above other sectors. Finally, Panel c) plots the 25th, 50th and 75th percentiles of the in gap distribution across all establishments by year. All three figures indicate there is substantial variation in the in gap measure across establishments, even establishments in the same industrial sector, and over time. The between establishment standard deviation in the in gap measure from 1996 forward is 2.79 percentage points and the within is 4.26 percentage points.

Overall, these statistics are convincing that we have a valid instrument to deal with the endogeneity introduced in the OLS estimates. Thus, we use these predicted retirement eligibilities to instrument our contemporaneous policy measure, using two-stage least square models. Estimates of OLS and 2SLS are discussed in Section 7.

6. Descriptive Results

This section provides a descriptive overview on establishment downsizing and sample statistics. Our sample period comprises the years 1993 to 2010. Starting in 1993 we can follow establishments and see how they progress. As shown in Table 1 we observe at a minimum 2,691 (in 1998) and at a maximum 5,821 establishments (in 2001) in our sample. The sample sizes differ each year as our sample is a non balanced cross section and not every establishment participates in the survey every year. Some establishments are joining the panel later when the IABBP did refresher samples due to panel attrition. This explains certain jumps in specific years. However, establishments are only in our sample if they existed in 1990. Table 1 also shows that the average establishment size decreases each year, from 730 employees in 1993 to 265 employees in 2010. This is also due to the structure of the IABBP. Large establishments were over sampled when the survey started in 1993. We account for this in the regression by adding all the sample strata as control variables. As expected we can see that the share of workers 58 and older rises on average every year. Whereas in 1993, on average 5% of the workforce in our sample was 58 and older, in 2010 this number increased to 14%.

Table 1: Sample Descriptives

year	observations (count)	estab. size (mean)	share 58 (mean)
1993	3,594	730	0.0509
1994	3,307	661	0.0535
1995	2,908	629	0.0558
1996	2,719	578	0.0596
1997	2,353	580	0.0643
1998	2,691	494	0.0704
1999	2,820	446	0.0913
2000	5,166	296	0.0989
2001	5,821	282	0.1003
2002	5,842	270	0.1012
2003	5,505	258	0.0991
2004	5,409	298	0.1005
2005	5,216	307	0.1056
2006	4,840	302	0.1121
2007	4,570	278	0.1178
2008	4,292	284	0.1232
2009	4,190	287	0.1327
2010	3,742	265	0.1432

Figure 4 shows a Kaplan Meier Survival probability for the establishments in our sample to get an impression how many establishments are experiencing a downsizing event. For illustration purposes we choose to display the probability of downsizing at the 30% threshold. For this purpose we balanced our panel by computing the number of years until the downsizing event happens starting from 1990. We can see that over the sample period more than half of the observed establishments experience a downsizing event of at least 30% of the workforce. Downsizing happens more frequently within the first five years in our observation window. Looking at the sub sample of establishments that experienced a downsizing event at the 30% threshold we can see in Figure 5 that most of these establishments experience this event about 6 years into our study period, which is when the pension reform started hitting the first worker cohorts.

Figure 4: Survival rate of Downsizing at the 30% threshold

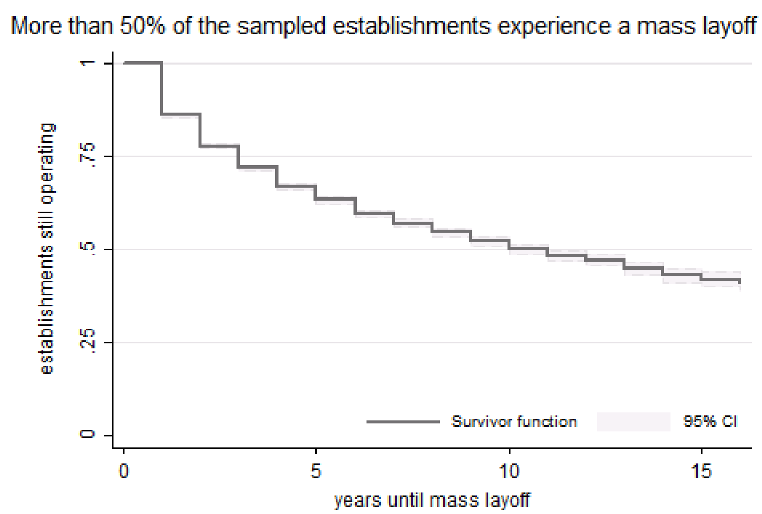
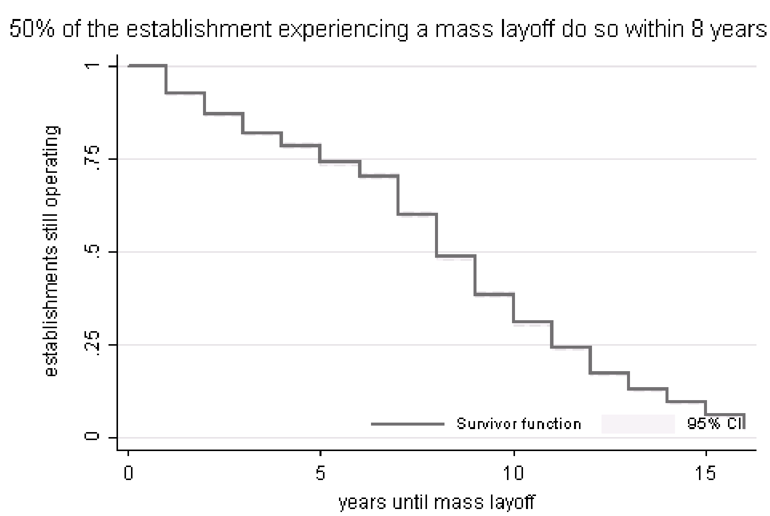


Figure 5: Survival rate of Downsizing at the 30% threshold - downsizing sample



7. Main Results

Table 2 displays the results of the first stage for our IV estimates. As the endogenous regressor, $share58$, and control variables are the same across all regressions and the sample of establishments is consistent across specifications, the first stage estimate is identical for all results reported. The estimated coefficient for the instrument z_ingap_{jt} has the expected positive sign and implies a one percentage point increase in the projected share of workers in the gap the reform created between the old and new pensionable ages is associated with a 0.18 percentage point increase in the share of the establishment's workforce age 58 or older.

Table 2: First Stage Estimates

	Coefficient	Std. Error
Shift-Share Instrument (on share 58)	0.184***	0.016
N	74,985	

The unit of observation is the establishment-year. Standard errors are clustered at the establishment level. Each regression includes a set of establishment characteristics (workforce demographics, wage bill, size, flows, existence of a work council and any industry agreements, location) and year dummies as controls. Three stars denote statistical significance at the 1-percent confidence level.

Table 3 contains the estimated effects of policy induced workforce aging on the likelihood of downsizing events overall, and by age group. Ignoring possible heterogeneity across age groups, overall our IV estimates indicate establishments with a higher share of workers over the age of 58 are more likely to experience downsizing at each of the thresholds we considered. Specifically, a one percentage point increase in the share of workers over age 58 is associated with a 0.3 percentage point increase in the likelihood of a downsizing event involving 10 percent or more of the establishment's workforce (Table 3 - All Workers - column

(1)). The OLS estimates, are negative, which is consistent with the expected bias discussed above.

When we segment establishment workforces by age and consider separations within subpopulations, the IV estimates again consistently reveal positive associations between policy induced workforce aging and the likelihood of downsizing within each demographic segment at each threshold considered. As different age groups have different baseline probabilities of downsizing, we report the coefficient estimates alongside the percentage change they imply relative to the baseline probability for each age group.

The coefficient estimates are highest for the youngest age group at all three thresholds. However, because downsizing events are most common in the younger segment of the workforce and least common among the oldest, in some cases these estimates imply a larger percentage change in the probability of downsizing among the oldest workers, here for the 10% and 30% threshold. For example, the impact of a 1 percentage point increase in the share of workers over age 58 is associated with a nearly 7 percent increase in the likelihood of downsizing events involving 10 percent or more of the older workers in an establishment, whereas we observe an slightly over 5 percent increase in the younger worker segment. For downsizing events involving 20 percent or more of the age group, both the percentage point change in the likelihoods and the percentage changes relative to baseline probabilities are larger for the younger rather than the older workers.

The reported estimates could be attenuated if establishments with higher shares of workers over age 58 close down during our study period. To address this, we estimate the same regression models using a closure indicator as outcome. The IV and OLS results, as well as the First Stage estimate are presented in the Appendix in Table A.3 and Table A.2 respectively. We find no significant associations between the share of workers over age 58 and the probability of closure, and the point estimate for the full sample is negative.

Table 3: Downsizing at different thresholds and age groups

	10% cutoff		20% cutoff		30% cutoff	
	(1)	(2)	(3)	(4)	(5)	(6)
Model	IV	OLS	IV	OLS	IV	OLS
			Older Workers			
Share 58	0.004***	-0.000***	0.001**	-0.000***	0.001*	-0.000***
	(0.001)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)
%	0.067	0.000	0.043	0.000	0.077	0.000
baseline	0.060		0.023		0.013	
			Prime Workers			
Share 58	0.002**	-0.001***	0.001	-0.001***	0.000	-0.000***
	(0.001)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)
%	0.032	-0.016	0.048	-0.048	0.000	0.000
baseline	0.062		0.021		0.011	
			Younger Workers			
Share 58	0.008***	-0.002***	0.004***	-0.001***	0.002***	-0.001***
	(0.001)	(0.000)	(0.001)	(0.000)	(0.001)	(0.000)
%	0.055	-0.014	0.065	-0.016	0.069	-0.034
baseline	0.146		0.062		0.029	
			All Workers			
Share 58	0.003***	-0.001***	0.001**	-0.000***	0.001**	-0.000***
	(0.001)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)
%	0.058	-0.019	0.053	0.000	0.100	0.000
baseline	0.052		0.019		0.010	
N	74,985	74,985	74,985	74,985	74,985	74,985

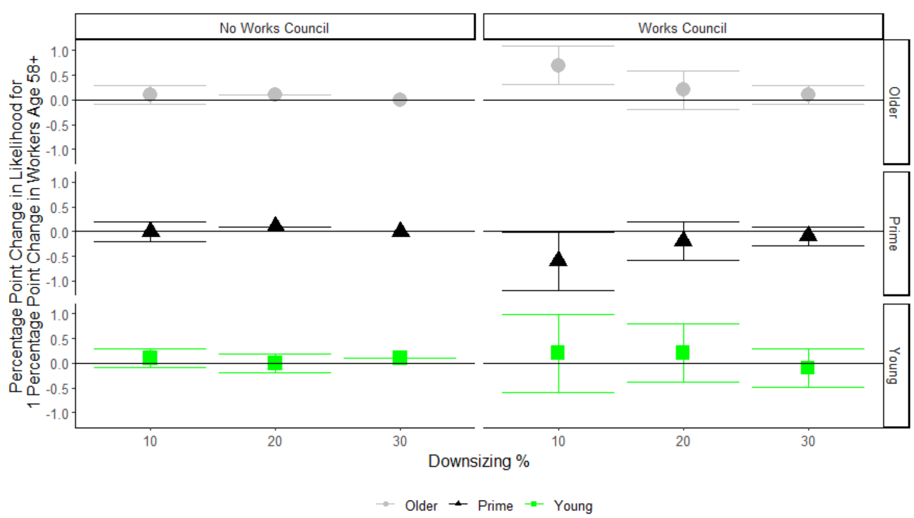
Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Each regression includes a general set of controls: location, share of younger and share of older people at workplace, and a marker to indicate missing years in the layoff calculation. In addition, each regression includes a set of establishment characteristics: workforce demographics, wage bill, size, existence of a work council and any industry agreements. Furthermore year fixed effects are included. The instrumental variable regressions are estimated by two-stage least squares. One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively. Baseline represents the mean of share 58.

8. What is Driving the Main Results?

To better understand the role of the institutional mechanisms that may influence the downsizing events we are studying, we separate our establishment sample into two groups: establishments with and without works councils. These absolute results are summarized graphically in Figure 6, whereas the percentage change in the probability is illustrated in Figure 7

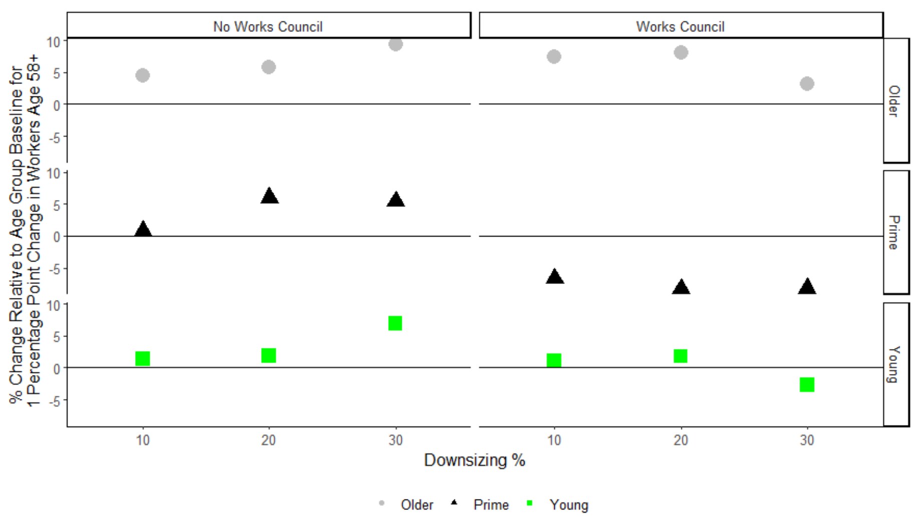
These results reveal that the positive relationships between policy induced workforce aging and downsizing events of 10% or greater in the prime-aged segment of the workforce shown in Table 3 was driven entirely by the establishments without works councils. For prime workers, works councils appear to mitigate the effects of workforce aging on downsizing probabilities, and they may reduce the likelihood of downsizing events involving over 30% of the workforce for all age segments.

Figure 6: Coefficients Age Group Estimates by Work Council



In all establishments, we find the largest percentage increases in downsizing probabilities occur in the oldest segment of the workforce, workers aged 50 and above at the 10% and 30% threshold. This is notable because many of these workers are in the earliest cohorts affected by the pension reform and had the least amount of time to adjust to the increase in their pensionable ages. In establishments with works councils, it appears older workers are the only segment of the workforce that experiences a statistically significantly higher likelihood of downsizing as the share of the workforce over the old effective retirement age increases. Among prime aged workers in establishments with works councils, the likelihood of downsizing appears to be decreasing in the share of workers over age 58. This relationship is not simply mechanical. Our measures of downsizing within each age segment are relative to the number of workers within that age segment, not total establishment employment. These negative relationships could be the result of works council efforts to protect employment of prime aged workers in the negotiation of a social plan to facilitate downsizing as required under German law.

Figure 7: Relative to Baseline Age Group Estimates by Work Council



We note the results presented in this section should be interpreted with caution because works council information is missing for some establishments in our sample, and is not likely missing at random.

9. Conclusion

This paper finds positive associations between workforce aging and the likelihood of downsizing events. Using an increase in pensionable age as a source of exogenous variation in establishments' employment of workers over age 58 (the old effective retirement age), we find a one percentage point increase in the share of workers aged 58 and older is associated with as much as a 10 percent increase in the likelihood of downsizing events. Also, this impact appears to vary across age segments within establishments, and differs between establishments with and without works councils. Whereas, the coefficient estimates are highest for the youngest age group at all three thresholds, in some cases the estimates however imply a larger percentage change in the probability of downsizing among the oldest workers. Our results suggest that work councils seem to take on a protective function for prime age workers.

The welfare implications of our findings for older workers are unclear. On the one hand, older workers appear to bear both the burden of a shorter planning horizon in which to adjust to increases in pensionable age under the law, and a greater risk of employment separation due to the postponed retirement incentives the law created. If, however, the downsizing we observe in this segment is primarily the result of buyouts that include compensation, it is possible older workers are able to achieve retirements on a similar timeline as was feasible before the reform with little or no loss of retirement wealth. If instead, the downsizing we observe reflects layoffs, older workers may be reliant upon public income support programs which were becoming less generous over this period. Unfortunately our data do not contain information about buyout payments and we cannot empirically investigate these possibilities.

Our current analysis is limited in its ability to convey the importance of the downsizing events for the labor force as a whole because we measure downsizing relative to the size of the establishment workforce and our estimates are unweighted. This means that an establishment with 100 employees and an establishment with 10,000 employees that each layoff 10 percent of their workers are treated the same, yet these events have very different implications for the well-being of the population and for social programs that support displaced workers. In future iterations of this work, we intend to produce weighted estimates to account for differences in establishment size.

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Appendix

Addressing Sample Attrition

To make sure that our main downsizing estimates are not biased through attrition of establishments in our panel because they completely go out of business we also produce a set of estimates on how workforce aging induced by pension reform relates to the death of an establishment. In a non biased sample we would expect the coefficients to not be statistically significant. To test this hypotheses we construct a closure indicator from the administrative establishment data. These contain an indicator for each establishment telling denoting the kind of closure that happened in case an establishment closed down. The different closures are outlined in Table A.1. We see that out of the establishments that close a bit more than 60% experience a death closure.

Table A.1: Construction of Closure Indicator

Label	absolute	percent
ID Change	2,121	7.98
Take-over	1,563	5.88
Spin-off	3,286	12.36
Small Death	2,640	9.93
Atomized Death	7,042	26.49
Chunky Death	8,467	31.85
Reason unclear	1,462	5.50

Thus, we define an establishment to be closed in a given year if we observe one of the three categories of death (small, atomized, chunky). These categories represent firm closures to differing degrees of impact (for detailed information on how this was constructed please see Hethey-Maier and Schmieder (2013)). Thus, the closure outcome for the regressions shown below takes on the value 1 if an establishment experiences a closure in year t and 0 otherwise. We estimate the effect of workforce aging on closure in the full sample and stratified by works councils. Table A.2 shows the First Stage estimates of all the models, whereas

Table A.3 displays the IV and OLS results. As expected we find no significant effects.

Table A.2: First Stage: Shift-Share Instrument (on share 58)

	Coef.	Std. Error	N
Full Sample	0.184***	0.016	74,985
Employer has Work council	0.150***	0.022	39,770
Employer has no work council	0.230***	0.021	28,930

The unit of observation is the establishment-year. Standard errors are clustered at the establishment level. Each regression includes a set of establishment characteristics (workforce demographics, wage bill, size, flows, existence of a work council and any industry agreements, location) and year dummies as controls. Three stars denote statistical significance at the 1-percent confidence level.

Table A.3: Closures, different subsamples, establishment controls

	Full Sample		Works Council		No Works Council	
	(1)	(2)	(3)	(4)	(5)	(6)
Model	IV	OLS	IV	OLS	IV	OLS
Share 58	-0.000	0.000	-0.001	-0.000	0.001	0.000
	(0.001)	(0.000)	(0.001)	(0.000)	(0.001)	(0.000)
%	-0.008	0.017	-0.104	-0.012	0.041	0.013
baseline	0.009		0.005		0.015	
N	74,985	74,985	39,770	39,770	28,930	28,930

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Each regression includes a general set of controls: location, share of younger and share of older people at workplace, and a marker to indicate missing years in the layoff calculation. In addition, each regression includes a set of establishment characteristics: workforce demographics, wage bill, and size. Furthermore year fixed effects are included. The instrumental variable regressions are estimated by two-stage least squares. One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively. Baseline represents the mean of share 58.