Peer Effects and Retirement Decisions: Evidence from Pension Reform in Germany

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Abstract

Peers are hypothesized to influence a wide range of economic behaviors, but identification of causal relationships is difficult. We use employee-employer matched administrative data to estimate the impact of peers on retirement decisions. Using exogenous variation in pensionable ages created by a reform that affected particular birth cohorts, we find robust evidence of spillovers from the retirement behavior of the affected cohorts to the unaffected cohorts. The magnitude of spillovers is large but limited to cohorts within approximately four to five years of age. In these cohorts, a one percentage point reduction in the share of peers who are affected by the reform and retire is associated with (at least) a 0.25 percentage point reduction in the share of peers retiring who are unaffected by the reform. We conclude peer effects amplify the effects of changes in retirement incentives. Our estimates imply reforms that encourage later retirements may produce changes in the share of older workers who retire that are at least 27% larger than the estimated response to changes in own incentives.

Keywords: *Retirement, Peer Effects, Pension Reform, Administrative Data*

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1. Introduction

The majority of developed countries are facing a demographic transition that threatens the solvency of their social security programs (Gruber and Wise, 2009). Private pension plans are under similar pressure as employers manage an aging workforce and combat legacy costs. As policymakers and employers consider options to address these issues, it is important to understand how individuals make retirement decisions. The literature has focused primarily on examining the impact of changes in individuals' own pension incentives on their retirement behavior (Mastrobuoni, 2009). Much less is known about the impact of peer retirements on individual retirement behavior. Peer effects could operate through, for example, information sharing and coordinated retirement decisions. If peer effects are present, changes in the retirement incentives facing one's peers may spillover and amplify the magnitude of individual responses to changes in their own retirement incentives. A social multiplier in the context of retirement could have important implications for social security fund balances in the future.

In this study, we use employee-employer matched administrative data to make four noteworthy contributions to the retirement literature. First, we produce estimates of peer effects generalizable to a wide range of industries and occupations by using data administrative employment records for West German establishments with 100 or more employees. These establishments employ approximately 50% of the German workforce. Previous studies examined retirement peer effects among school teachers in the Los Angeles School District (Brown and Laschever, 2012) and non-federal public-sector workers in Oregon (Chalmers et al., 2008). Both studies find that individual-retirement decisions are positively affected by peer retirements, but it is impossible to know whether these findings generalize to the broader economy or the private sector. Second, we study peer effects in a setting in which the identifying change in pension incentives is a relatively straightforward increase in pensionable age, rather than a complex change in pension accrual. In our setting, any peer effect is likely the result of social ties rather than information sharing. Third, we examine more narrowly defined peer groups than those used in prior work. Brown and Laschever (2012) estimate peer effects at the school level, and Chalmers et al. (2008) estimated peer effects among all employees with the same employer eligible to retire in the same month. Our narrowest peer effect estimates are based on workers born within one year of one another and working in the same occupation within the same establishment. Our broadest peer effect estimates include all workers born in 1931 through 1944 and employed in the same occupation within the same establishment. Fourth, existing studies of peer effects on retirement decisions use differences in pension incentives across peers to identify peer effects, but the changes in pension incentives over their study periods affected all members of the peer group to some degree. Thus, previous identification strategies are unable to circumvent reflection problem (Angrist, 2014; Manski, 1993). Our approach examines the impact of retirements among the first cohorts affected by the reform on the retirement behavior of peers whose pensionable ages were not altered by the reform, which circumvents the reflection problem. Importantly, pensionable ages for affected cohorts depend on both birth date and sex, which adds variation in the policy impact and associated behavior across peer groups with similar age profiles.

Our identification strategy is based on stepwise changes in pensionable ages affecting persons born after 1938 that gradually increased ages of eligibility for full pension benefits from age 60 to age 65. During the phase-in, adjacent single-year birth cohorts and persons of different sexes born in the same year had pensionable ages six or more months apart. The policy changes did not directly alter incentives for workers born before 1938. The reform provides a source of exogenous variation with which to identify the causal impact of peer retirements on individual retirement decisions, as it incentivized later retirements among an identifiable demographic group but did not impact incentives of all workplace peers. The reliance on changes in pensionable ages, which vary by birth cohort and sex, circumvents the potential endogeneity problems associated with using changes in benefits, which are a function of past earnings and, as a result, could affect old-age labor-supply decisions through accumulated wealth (Mastrobuoni, 2009).

Using the reform as identification, we find for each percentage point decrease in share of peers in the 1938 cohort who retire, there is a 0.26 percentage point reduction in the share of peers in the 1931 through 1937 cohorts who retire. In analysis restricted to only the youngest unaffected cohorts (1935 through 1937), we find peer effects in response to retirements from 1938 through 1940 cohorts as large as 0.82 percentage points. As expected, estimates are generally largest among cohorts closest in age. In total, we conclude workplace peers have an important impact on retirement timing. Peer and own retirements are positively correlated, and policies that encourage later retirements spillover to adjacent cohorts. These peer effects lead the total shares of peers retiring to be approximately 27% larger than the estimated response in cohorts directly affected by the reform.

2. Pension Reform in Germany

Like most European countries, Germany has a pay-as-you-go pension system. The system covers approximately 85 percent of the German population (Berkel and Börsch-Supan, 2004). Private retirement savings is uncommon in Germany for the cohorts we study. In the 1990s, public pension benefits accounted for approximately 80 percent of income among households headed by persons aged 65 or older (Börsch-Supan, 2000). 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; Börsch-Supan et al., 2008). Public pension accrual is a simple function of one's own wages, years of service, age, and national average wages and benefits in each year. Benefits are based only on one's own work history. There are no spousal benefits, only survivor benefits.

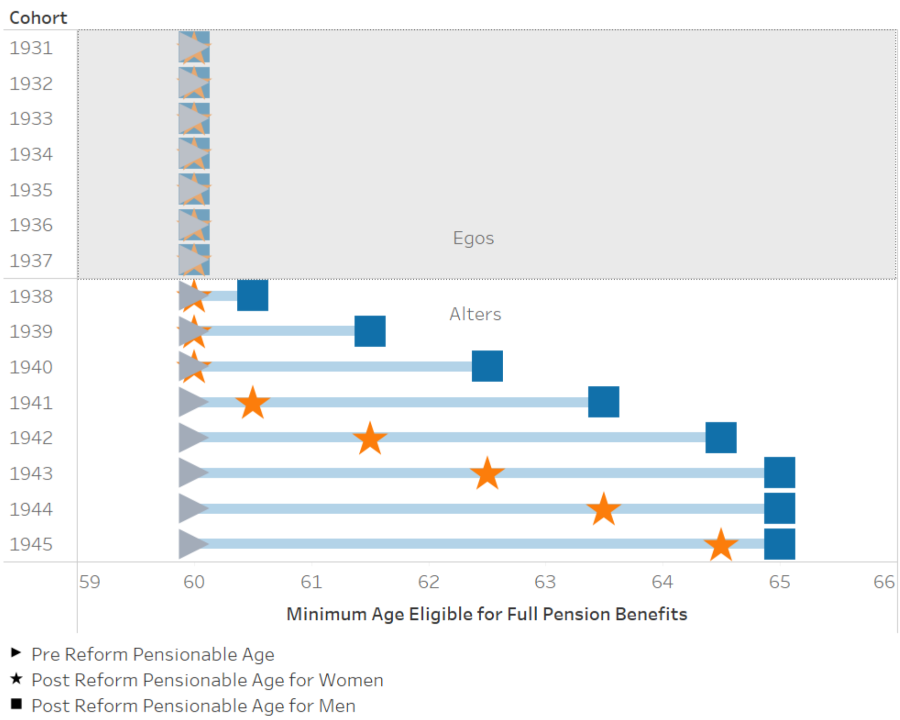
Since 1972, Germany's pension system has contained many "pathways" to claiming old-age pension benefits (Boersch-Supan and Wilke, 2004). Like in the

U.S., Germans can retire before reaching "normal pensionable age" and receive actuarially adjusted old-age pension benefits. Workers born before 1938 who claimed pensions before age 65 had lower disbursements because they had fewer years of contributions, but their benefits were not otherwise actuarially adjusted (Berkel and Börsch-Supan, 2004). In 1992, Germany introduced reforms that gradually eliminated pathways to retirement that pay full benefits (without actuarial adjustment) prior to age 65. These reforms affected cohorts born in 1938 and later, as documented in Berkel and Börsch-Supan (2004, p. 397) and Boersch-Supan and Wilke (2004, p. 28). Because the early retirement pathways were highly popular, simulations in prior research predicted the reform would lead to a 2 year extension of working life (Berkel and Börsch-Supan, 2004).

The pre-reform and the post reform pensionable ages are plotted by cohort in Figure 1. As shown, differences in pensionable age across adjacent birth cohorts ranged from 6 month to 12 months and there were differences within cohorts by sex. Differences by sex arose because under the pre-reform rules women could claim old-age pension benefits at age 60 whereas men could not until age 65 unless they were unemployed or disabled. The 1992 reform increased the pensionable age for unemployed workers more rapidly than the pensionable age for women. This created differences in men's and women's eligibility for full pension benefits *within* the 1938 through 1945 birth cohorts of 6 months to 36 months.

In practice, many German workers born before 1938 had exited the labor force at age 57 or 58, claimed unemployment benefits and then old age pension benefits at age 60. Less than 25% of men retiring in the early 1990s were 65 years old at retirement and 45% called themselves retired at age 59 (Boersch-Supan and Wilke, 2004). This practice continued even as pensionable ages for unemployed workers were increasing. Figure A.1 displays the ages when each cohort could have entered an unemployment spell leading directly into retirement and claiming of full pension benefits. The cross-cohort and within cohort patterns by sex are the same as those shown in Figure 1, but the ages of labor force exit into unemployment are two years earlier than the ages of

Figure 1: Summary of 1992 Reform of Pensionable Ages By Cohort



pension claiming shown in Figure 1. This means changes in labor force exit patterns attributable to the 1992 reform may have been evident in 1996 (when the 1938 cohort reached age 58), or even earlier.

3. Data

3.1. Data Sources and Sample Construction

Our data are a sample from administrative records of the German pension system. We create a custom extract from the Establishment History Panel (BHP) matched to Integrated Employment Biographies (IEB) for all workers within each establishment.

From the BHP population, we first select all West German establishments recorded at least once between 1990 and 2002 that employed at least 100 workers at any point in time. We exclude East German establishments because records were incomplete during our study period. Smaller establishments are excluded from the analysis because they generally have too few workers affected and unaffected by the reform to identify a peer effect. Also, we exclude peer groups in the 90th percentile of the size distribution. After these restrictions, our data set includes 7,833 establishments and 14,739 peer groups. The majority of observed establishments belong to the mining and manufacturing sectors, followed by trade and food services, finance and real estate. The numbers of workers and peer groups vary across specifications as discussed below.

For each establishment in our sample, we obtain IEB data for all employees aged 50 to 65. Establishments without any workers over age 50 in *any* year from 1993 through 2002 are excluded, but establishments may have some years without workers over age 50. These individual level data include workers' employment histories, necessary information to determine the pensionable age for each individual. The IEB data include employment information from the notification process of the social security system and the internal procedures of the Federal Employment Agency. We use these employment histories to determine workforce exit for each worker in the sample.

To obtain characteristics of all workers within the same occupation at each establishment (not just those who are age 50 and older), we obtained another custom data extract similar to the BHP. The Institute for Employment Research constructed a data set that included annual aggregated workforce statistics on the workforce composition on the establishment-occupation level. The information is identical to the BHP (Gruhl et al., 2012), however, we have all the information by establishment and 3-digit occupation within establishments. We refer to this as our Occupational Group Characteristics File.

3.2. Definition of Peer Groups

We define peers as workers age 50 to 65 working for the same establishment and in the same occupation. Occupation is defined according to the Blossfeld classification, a 12 category occupational grouping (Blossfeld, 1987). The 12 occupation groups are agricultural jobs, simple manual jobs, simple services, simple sales jobs, medium-skilled manual jobs, medium-skilled services, technicians, medium-skilled sales jobs, engineers, semi-professionals, professionals, and managers.

3.3. Key Measures

Our key outcome measure is retirement of peers from birth cohorts unaffected by the reform, expressed as a share of total peer group employment. The numerator is constructed from an individual level dummy variable equal to one in the year when an end of employment notification is filed for the employee (i.e. the last year of work). We then sum all retirements at the peer-group level by birth cohort and divide by total peer group employment. We prefer to measure peer retirements as shares because we hypothesize the causal effect of one peer retiring likely varies considerably across large and small peer groups, but the causal effect of a one percentage point change in the share of peers retiring should similar.

The key explanatory variable is the share of peers retiring who are born to birth cohorts affected by the reform. This variable is constructed in the same manner as the dependent variable with the same denominator, but the retirements in the numerator are based on peers born to the 1938 and later cohorts.

4. Identification Strategy

The identification of peer effects has three main challenges: simultaneity, correlated unobservables, and endogenous group membership (Manski, 1993). To convincingly address these issues in an observational study, an instrument that is strongly correlated with endogenous group outcome and uncorrelated with unobservable determinants of retirement decisions is required. However, more is needed to identify the causal effect of peer behavior on individual outcomes: the instrument must vary within the peer group (Brock and Durlauf, 2001). In addition, the instrument must exhibit between-peer-group variation. In the peer-effects literature, such instruments are referred to as Partial Population Interventions (PPI) because they affect some but not all individuals in the peer group (Moffitt, 2001). [1]

The typical research design for estimating peer effects includes a sample of individuals who are subjects as well as peers. However, even with a valid PPI, estimating peer effects via 2SLS is problematic in the sense that the first-stage regression of the group behavior on the instruments is identical to a reduced-form regression of the individual behavior on the instrument, as every individual in the sample is also a different individual's peer. Thus, the instrument appears in both the first- and second-stage equations. (Angrist, 2014). [2] As such, one may attribute behavioral responses to one's peers when in fact the estimated

[1] Brown and Laschever (2012) extend the PPI-approach via the use a policy intervention that affects members of peer groups differently, which they refer to as a Differential Population Intervention (DPI). The German pension reform we exploit has features of both the PPI and DPI approaches, as workers born before 1938 were unaffected and workers born in 1938 and after were differentially affected (the reform also affects men and women differently.

[2] Angrist (2014) shows that with this research design estimated peer effects are approximately equal to the ratio of the 2SLS to OLS estimates.

peer effect could simply be due to differences in 2SLS and OLS estimates that do not reflect a response to peer behavior.

In order to obtain valid estimates of peer effects, the mechanical link between individuals and their peers must be broken (Angrist, 2014). Our research design breaks the mechanical link by studying the behavioral responses of workers unaffected by the pension reform to peers born to cohorts directly affected by pension reform. To distinguish between these two subgroups within the peer group, we adopt terminology commonly used in this literature and refer to the peers born after 1938 as the "alters" and the peers who we expect to respond to the alters' policy induced changes in retirement behavior as "egos".

Angrist (2014) also states OLS and 2SLS estimates should be equivalent when peer effects are absent. Random assignment to peer groups would create an expectation of equivalence, but randomized group assignment is rare. Although random assignment to groups is absent in our setting, the cutoff between affected and grandfathered cohorts (alters and egos) in the 1992 reform was arbitrary and changes began to bind as early as 1995. Furthermore, job changes and moving to different establishments are rare among older workers in Germany, making selection into and out of peer groups in response to the reform unlikely. To further support these assertions, we present descriptive statistics in Table 1 that illustrate little systematic variation in the shares of alters from each birth cohort across peer groups.

Table 1: Correlation between birthcohorts

	Cohort 1938	Cohort 1939	Cohort 1940	Cohort 1941
Cohort 1938	1.0000	-	-	-
Cohort 1939	0.2079	1.0000	-	-
Cohort 1940	0.1797	0.1557	1.0000	-
Cohort 1941	0.1342	0.1133	0.1062	1.0000

The variables used to calculate the correlation matrix are the share of workers of each birthcohort in the peer group relative to all workers in the peer group in year 1993. The correlation matrix is based on 24,795 entities.

4.1. Estimation

To estimate peer effects, we specify a system of two peer-group level equations, which are estimated via two stage least squares (2SLS).

$$AlterRetires_{g,t} = \delta_0 + \delta_1 P_{g,t} + Z'_{g,t}\delta_2 + \phi_g + \phi_t + \eta_{g,t} \tag{1}$$

$$EgoRetires_{g,t} = \beta_0 + \beta_1 Alter\widehat{Retires}_{g,t} + Z'_{g,t}\beta_2 + \phi_g + \phi_t + \epsilon_{g,t} \tag{2}$$

In Equation (1), the first stage equation, $AlterRetires_{g,t}$ is the share of alters (peer group members born in years 1938 through 1944) who retire in year t; $P_{g,t}$ is the share of alters employed in the peer group at year t who are eligible to retire with full benefits under the post-reform pension rules in year t, which is our instrument. $Z_{g,t}$ is a vector of occupational group characteristics including shares of employees in the same occupation within the establishment who are female, low skill, highly skilled, under 30 years old, and under 50 years old. These shares include workers of all ages and are from the custom Occupational Groups Characteristics File described in Section 3.1. The same variables are also constructed at the establishment level and included in $Z_{g,t}$ along with the share of establishment employees in each of the 12 occupational groups. [3] We also control for the median peer age, the interquartile range of peer ages, and the share of peers who are foreign born (because nationality may affect pension eligibility). The ϕ_g and ϕ_t are group and year fixed effects, respectively. Equation (1) is used to produce the vector of fitted values, $Alter\widehat{Retires}_{g,t}$, that appears in Equation (2).

Equation (2) is also estimated at the peer group level, and the dependent variable $EgoRetire_{g,t}$ represents the share of egos (peer group members born in years 1931 through 1937) who retire in year t. All other variables are as defined above.

[3] These variables are available in the BHP

$\hat{\beta_1}$ is the estimated peer effect and is interpreted as the percentage point change in the share of egos who retire in response to a one percentage point increase in the share of alters who retire. A negative (positive) estimate of $\hat{\beta_1}$ would imply egos become less (more) likely to retire as the share of alter retirements increases.

Equations (1) and (2) constrain the peer effect of all alter retirements to be the same, but it is reasonable to expect peer effects to be strongest among egos and alters who are closest in age. To examine this possibility and better understand the identifying variation underlying the estimates in Equations (1) and (2), we expand our initial empirical model as follows:

$$AlterRetires_{c,g,t} = \delta_0 + P'_{c,g,t}\delta + Z'_{g,t}\delta 2 + \phi_g + \phi_t + \eta_{g,t} \tag{3}$$

$$EgoRetires_{g,t} = \beta_0 + Alter\widehat{Retires}_{c,g,t}{}' \beta_1 + Z'_{g,t}\beta_2 + \phi_g + \phi_t + \epsilon_{g,t} \tag{4}$$

Equation (3) represents four first-stage equations, one for each alter birth cohort, c, from 1938 through 1941. We are unable to model retirements among cohorts born after 1942 in this exercise because they do not become eligible for retirement until after 2002. As a result, the instruments, $P_{c,g,t}$, which represent the share of peers in that cohort eligible to retire are all equal to zero for cohorts born in 1942 and later. Equation (4) is an expansion of Equation (2) and includes the four instrumented $AlterRetires_{g,t}$ variables and the coefficients vector β_1 contains the estimates for the associated peer effects.

Both sets of equations are estimated via 2SLS with heteroskedasticity robust standard errors clustered at the establishment level.

4.2. Variation in the Instrument

As explained, we construct instruments $P_{g,t}$ and $P_{c,g,t}$ as the shares of peers eligible to retire with full pension benefits in each year, peer group, and birth cohort where applicable. Individual eligibility is based both birth year, age, and sex. Thus, the instruments vary over time and with the birth year and gender composition of the peer group. To investigate balance across other observables, we regressed the pooled cohort instrument $P_{g,t}$ on year and group fixed effects, ϕ_t and ϕ_g, and our group and establishments controls, $Z_{g,t}$. The regression results are reported in Appendix Table A1. We find the expected significant relationships with year, measures of the group age distribution, and share of female employees, nearly all other covariates have t-statistics below 1.96. Establishment size is significantly associated with the instrument, but because peer groups tend to be larger in larger establishments the shares of peers eligible to retire will be smaller in larger establishments by construction. Conditioning on these observables, 52% of the residual variance in the instrument is attributable to the group fixed effects. The estimated standard deviations in the group component of the error term is 1.9 percentage points, and of the idiosyncratic component is 1.8 percentage points. These are both large relative to the mean of $P_{g,t}$, which is 1.9%.

To further understand the within peer group identifying variation, we compute the shares of peer group members in each alter birth cohort (1938 through 1941) in the year 1993. We report correlations between these shares in Table 1. These cohort shares are a key component of the variation in the pooled instrument $P_{g,t}$ and must exhibit sufficient within peer group variation to separately identify the peer effects by alter cohort, β_1, in Equation 4. As shown in Table 1, the correlations between these shares are all positive but are at or below 0.21.

5. Results

Table 2 contains the descriptive statistics for our alter and ego retirement variables and peer group characteristics. On average, 3.2% of egos and 3.1% of alters in each peer group retire per year, but only 1.9% of alters are eligible for full pension benefits. The standard deviations in each of these variables are at least as large as the means, and the shares range from zero to between 0.4 and 0.5. Peer groups are 37.6% female on average, though this percentage varies from 0% to 100% and has a standard deviation of 31.1 percentage points. This

Table 2: Summary Statistics

	Mean	Std. Dev.	N
Peer Retirement Variables			
Ego Retires	0.032	0.035	127,161
Alter Retires	0.031	0.038	87,092
$P_{g,t}$	0.019	0.049	87,092
Occupational Group Characteristics			
Share Female	0.376	0.311	86,225
Share Low Skilled	0.241	0.251	86,225
Share High Skilled	0.091	0.210	86,225
Share Under 30 Years Old	0.188	0.119	86,225
Share Over 50 Years Old	0.719	0.124	86,225

variation is useful because it ensures peer groups with identical age distributions will differ in pension eligibility so long as they differ in gender composition. Our sample selection criteria required establishments to contain workers over age 50 at least one time during the study period, and peer groups to contain workers in every alter cohort. Not surprisingly, the share of workers in the peer groups who are over age 50 is high (71.9%).

In Table 3, we report results of estimating Equations (1) and (2) by 2SLS and Equation (2) by OLS. The first stage shows a large positive relationship between the share of alters eligible to retire and the share of alters who actually retire. The magnitude indicates that a one percentage point increase in the share eligible to retire increases the share who actually retire by 0.16 percentage points, and the first-stage F-statistic is 317.27 (Kleibergen-Paap Robust).

The OLS estimate of the peer effect indicates a one percentage point increase in the share of alters who retire is associated with a 0.042 percentage point increase in the share of peer group members who retire and are egos. The IV estimate is -0.001 percentage points, and not statistically significantly different from zero but is not precisely estimated enough to conclude it is smaller than

Table 3: Main Results (Equation (1) and (2))

	Ego Retires Pooled*		Share Alter Retires Pooled*
	(1)	(2)	(3)
Model	OLS	IV	First Stage
Share Alter Retires Pooled	0.042***	-0.001	-
	(0.006)	(0.035)	-
Share Alter Eligible to Retire Pooled	-	-	0.157***
	-	-	(0.009)
N	88,309	86,225	86,225

* Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is peergroup-year. Each regression includes a set of establishment characteristics (median age and interquartile age range, share of: females, low-skilled, high-skilled, part-time, under 30, over 50, foreigners, agricultural occupations, simple/qualified manual occupations, simple/qualified service occupations, simple/qualified administrative occupations, technical occupations, managerial occupations, engineering occupations, semi/professional occupations), and peer group characteristics (share of: females, low-skilled, high-skilled, part-time, under 30, over 50, foreigners), and year and peer group fixed effects. 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.

the OLS estimate.

When we allow the peer effect to vary by alter cohort, we find the expected heterogeneity underlying the pooled estimate. Table 4 contains the the OLS and IV estimates of Equation (4).

The first stage estimates (Equation 3) are reported in Appendix Figure 2. They exhibit large positive relationships between each cohorts' own share eligible to retire and share who actually retire.[4] The OLS and second stage IV estimates reported in Table 4 both indicate peer effects are largest between egos and alters closest in age. The IV estimates now indicate the share of egos who retire rises by 0.258 percentage points in response to a one percentage point increase in the share of peer group members alters born in 1938 but does not change or even

[4]The Kleibergen-Paap Wald F statistic for the weak identification test is 29.17, the Kleibergen-Paap LM statistic for the underidentification test is 85.53, and the Anderson-Rubin Wald F test of joint significance of the endogenous regressors is 10.03

Table 4: Second Stage - Equation (4)

	Ego Retires Pooled*	
	(1)	(2)
	OLS	IV
Share Alter Retires 1938	0.103***	0.258***
	(0.009)	(0.051)
Share Alter Retires 1939	0.039***	0.000
	(0.010)	(0.056)
Share Alter Retires 1940	0.006	-0.272***
	(0.011)	(0.076)
Share Alter Retires 1941	0.009	-0.167
	(0.012)	(0.114)
N	130,070	127,161

* Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is peergroup-year. Each regression includes a set of establishment characteristics (median age and interquartile age range, share of: females, low-skilled, high-skilled, part-time, under 30, over 50, foreigners, agricultural occupations, simple/qualified manual occupations, simple/qualified service occupations, simple/qualified administrative occupations, technical occupations, managerial occupations, engineering occupations, semi/professional occupations), and peer group characteristics (share of: females, low-skilled, high-skilled, part-time, under 30, over 50, foreigners), and year and peer group fixed effects. 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.

falls in response to retirements of younger alters. The OLS estimates exhibit a similar pattern but are smaller in magnitude. In total, these estimates indicate ego and alter retirement behavior is positively correlated among those who are close in age but uncorrelated or even negatively associated among more distant peers.

5.1. Robustness Checks

The estimates presented in Tables 3 and 4 are based on models that include many time varying group and establishment control variables $Z_{g,t}$. However, given that our instruments appear balanced across these observable characteristics, our results should be highly similar in models that omit these controls. Omitting these controls may improve the efficiency of the estimates and could eliminate other potential sources of endogeneity bias. When we do this, the pooled peer effect IV estimate (Equation 2) is 0.464 and the standard error falls to 0.043. Omitting control variables from Equation (4) leads to the same pattern of peer effects estimates across the 1938 through 1941 cohorts but now the estimates for the 1938 and 1939 cohorts are 0.646 and 0.347, and these estimates are statistically significant at conventional levels.

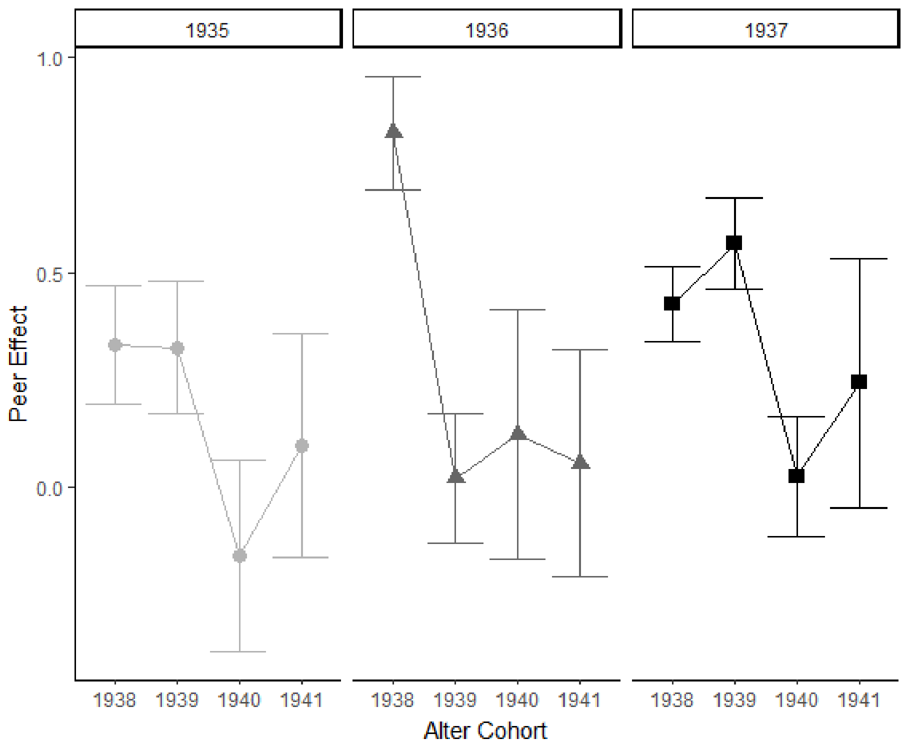
Because the estimates indicate peer effects are strongest among peers closest in age, it is sensible to expect peer effects will be strongest on the youngest egos and weakest among the oldest egos. To test this hypothesis, we re-estimate Equations (3) and (4) three times restricting the sample to the following ego cohorts: 1935, 1936, and 1937. The IV estimates are reported in Figure 2 and Appendix Table A.2. This hypothesis is somewhat supported by the IV estimates and fully supported by the OLS estimates. The IV estimates indicate there are large positive peer effects for 1938 alter retirements for the three ego cohorts, but the 1996 ego cohort estimate is the largest. Estimates for the peer effect associated with 1939 alter retirements are largest among the 1937 egos.

The aforementioned findings indicate peer effects are likely most important among peers who are close in age, but estimating peer effects at the group level does not allow us to examine variation by other individual characteristics. Moreover, although we control for time varying establishment and peer group covariates and use an instrumental variables estimator, there may still be bias associated with variation in individual worker characteristics that our models do not capture. For these reasons, we estimate Equation (5), which is an expansion of Equation (4), at the individual level for ego cohorts.

$$Retire_{i,e,g,t} = \beta_0 + Alter\widehat{Retires}_{c,g,t}{}' \beta_1 + Z'_{g,t}\beta_2 + Z'_{i,g,t}\beta_3 + \phi_e + \phi_g + \phi_t + \epsilon_{i,e,g,t} \tag{5}$$

Equation (5) includes a vector of of individual covariates, $Z_{i,g,t}$ that contains

Figure 2: IV Estimates of Peer Effects by Ego Cohort



: Peer effects are percentage point changes in the share of peer members who are egos and retire individual measures of full and part-time work experience, tenure within the peer group, wages, German nationality, and educational attainment. We also add ego birth year fixed effects phi_e. The dependent variable is now $Retire_{i,e,g,t}$, which is a binary variable equal to one in the year t when ego i, born in ego cohort e, in peer group g retires, zero when she is still working, and missing after retirement. This means the coefficient estimates can be interpreted as changes in the retirement hazard rate. $Alter\widehat{Retires}_{c,g,t}$ are the fitted values after estimating Equation (3) at the peer group level for each alter cohort c.[5]

Appendix Table 3 reports the results from estimating Equation (5) for all egos and for men and women separately. We find peer effects of each alter cohort are positive and significant. However, they do not exhibit the same decay as peers become increasingly distant in age observed when estimated at the peer group level. The separate estimates by gender help to explain why and point to another potentially important observation about peer effects in retirement. The only large and statistically significant peer effect for women egos is with the 1941 alter cohort. Returning to Figure 1, we can see 1941 is the first cohort of women affected by the reform. Results among men instead exhibit positive associations with all alter cohorts, and although the point estimates rise they are too imprecisely estimated to conclude they are different from one another. Together, these estimates point to potential gender differences in the responsiveness to peers.

6. Discussion and Conclusion

We examine the impact of peers on retirement decisions using a pension reform that raised pensionable ages for some, but not all, peer group members.

[5]Because of the two levels of analysis standard IV packages will not produce these estimates. Instead, the 2SLS estimates are derived explicitly. In the final version of this paper, correct standard errors will be achieved through block-bootstrapping. We have postponed this task because it is a non trivial exercise for such a large sample, and is exacerbated by data use restrictions.

Using a sample containing a large number and variety of establishments, we find evidence of spillover effects in the context of retirement. The retirement decisions of peers affected by the pension reform (who we refer to as "alters") affect the retirement behavior of peers unaffected by the reform ("egos"). We show the range of cohorts considered "peers" matters; pooling across all workers born in 1931 through 1944 produces an estimated peer effect of zero. However, underlying that estimate, we find peer effects of 0.258 percentage points among peers closest in age and -0.272 percentage points among more distant peers (in terms of age). Negative peer effects are not robust across specifications, but positive estimates among the closest cohorts are consistently found across specifications.

Our estimates indicate that pension reforms, such as the 1992 reform in Germany, may generate important social multiplier effects, which could impact the solvency of social security programs. For example, our first stage estimates indicate a one percentage point reduction in the share of peer group members who are eligible to retire is associated with a 0.15 to 0.20 percentage point reduction in the share who actually retire, depending on the cohort. Using the 0.15 percentage point estimate, a second stage peer effect estimate of 0.25 would indicate the change in behavior among affected cohorts triggered an additional 0.04 percentage point reduction in the share of peers retiring through spillover to unaffected cohorts. Thus, the total change in the share of retirees across all cohorts is approximately 0.19 percentage points, of which the peer effect comprises 21%. Many of our peer effect estimates are larger than 0.25. This implies the total change in retirement behavior, including spillover to peers, is at least 27% larger than the estimated response to changes in own incentives. However, spillovers appear to be limited to peers within about 4-5 years of age.

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Appendix

Figure A.1: Eligibility to Begin Unemployment Spell Leading to Pension Claim, by Birth Cohort and Sex

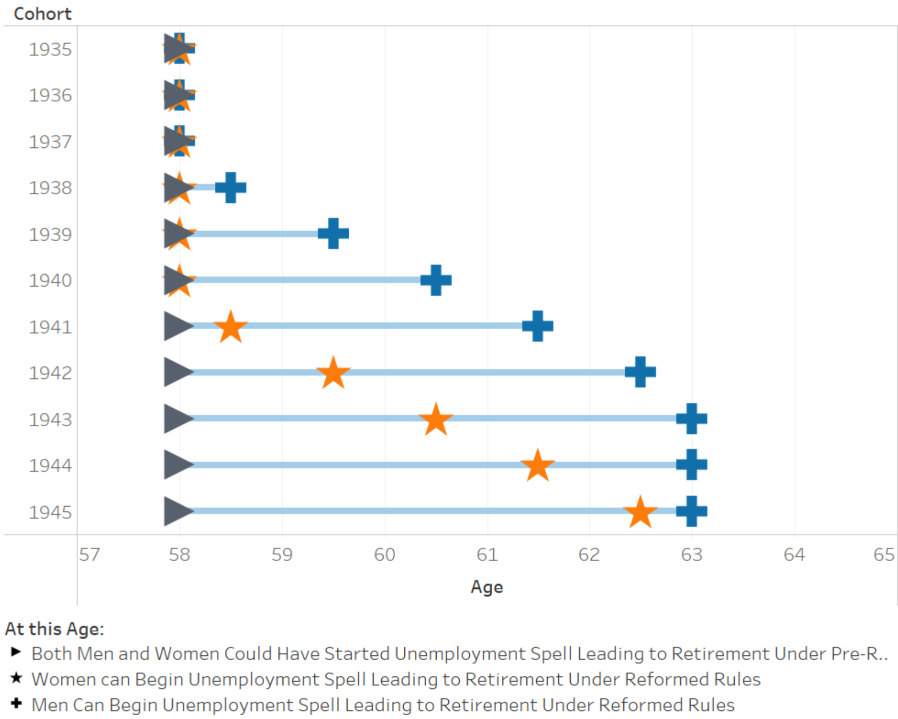

Figure A.2: First Stage Estimates by Alter Cohort

Share Eligible	Share Retire 1938	Share Retire 1939	Share Retire 1940	Share Retire 1941
1938	0.193*** (0.007)	-0.005 (0.007)	0.006 (0.005)	-0.009** (0.004)
1939	0.015*** (0.006)	0.175*** (0.008)	-0.025*** (0.006)	-0.009** (0.005)
1940	0.048*** (0.009)	0.016** (0.007)	0.163*** (0.009)	-0.020** (0.009)
1941	-0.009 (0.010)	0.035*** (0.013)	-0.013 (0.010)	0.164*** (0.016)

Table A.1: Regression of IV on Covariates

Covariate	**Coeff.**	**Std. Err.**
1994	-0.000	0.000
1995	-0.001**	0.000
1996	-0.002***	0.001
1997	-0.002**	0.001
1998	-0.002**	0.001
1999	0.038***	0.001
2000	0.085***	0.001
2001	0.138***	0.002
2002	0.188***	0.002
	peer group controls	
share female	0.005	0.005
share low qualified	0.007	0.005
share high qualified	-0.007	0.007
share under 30	-0.000	0.005
share over 50	0.054***	0.006
share foreigner	-0.015**	0.009
	establishment controls	
median age	0.000***	0.000
interquartile age range	0.000***	0.000
number of employees	-1.92e-06***	4.99e-07
share female	0.021**	0.008
share low qualified	-0.008	0.007
share high qualified	-0.023**	0.008
share part-time	-0.012*	0.005
share under 30	-0.012*	0.007
share over 50	0.036***	0.008

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Table A.1 – *Continued from previous page*

Covariate	**Coeff.**	**Std. Err.**
share foreigner	0.022*	0.011
share agricultural occ.	-0.005	0.030
share simple manual occ.	-0.024	0.016
share simple service occ.	-0.010	0.016
share simple admin occ.	-0.014	0.017
share qualified manual occ.	-0.017	0.016
share qualified service occ.	-0.032*	0.019
share qualified admin occ.	-0.011	0.016
share technical occ.	-0.032*	0.019
share semi prof. occ.	-0.057***	0.017
share engineering occ.	-0.012	0.022
share professional occ.	-0.026	0.018
share managerial occ.	-0.027	0.020
cons	-0.245***	0.021
N	86,286	
$corr(u_i, Xb)$	-0.189	
sigma u	0.019	
sigma e	0.018	
rho	0.517	

Notes: Standard errors, clustered at the peergroup level, are in parentheses. The unit of observation is peergroup-year. The regression includes peergroup fixed effects. One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels.

Table A.2: Second Stage - Allowing the peer effect to vary by alter cohort (Equation (4))

	Ego Retires 1937*		Ego Retires 1936*		Ego Retires 1935*	
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	IV	OLS	IV	OLS	IV
Share Alter Retires 1938	0.123***	0.426***	0.083***	0.824***	0.057***	0.332***
	(0.006)	(0.045)	(0.007)	(0.067)	(0.007)	(0.071)
Share Alter Retires 1939	0.073***	0.568***	0.055***	0.021	0.028***	0.326***
	(0.007)	(0.055)	(0.007)	(0.077)	(0.007)	(0.079)
Share Alter Retires 1940	0.050***	0.025	0.012	0.123	0.007	-0.158
	(0.008)	(0.071)	(0.008)	(0.148)	(0.010)	(0.114)
Share Alter Retires 1941	0.021***	0.244*	0.003	0.057	0.024**	0.096
	(0.008)	(0.148)	(0.009)	(0.134)	(0.011)	(0.133)
N	108,080	105,404	95099	92,338	80650	77,690

* Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is peergroup-year. Each regression includes a set of establishment characteristics (median age and interquartile age range, share of: females, low-skilled, high-skilled, part-time, under 30, over 50, foreigners, agricultural occupations, simple/qualified manual occupations, simple/qualified service occupations, simple/qualified administrative occupations, technical occupations, managerial occupations, engineering occupations, semi/professional occupations), and peer group characteristics (share of: females, low-skilled, high-skilled, part-time, under 30, over 50, foreigners), and year and peer group fixed effects. 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.

Table A.3: Second Stage - Individual Level

	Ego Retires	Std. Err.	N
	(1)	(2)	(3)
	Full Sample*		
Alter Retires 1938	2.340***	0.466	1,245,107
Alter Retires 1939	2.112***	0.624	1,245,107
Alter Retires 1940	1.634**	0.950	1,245,107
Alter Retires 1941	5.289***	1.377	1,245,107
	Men*		
Alter Retires 1938	1.146**	0.457	879,389
Alter Retires 1939	1.211*	0.657	879,389
Alter Retires 1940	2.444**	1.059	879,389
Alter Retires 1941	4.907**	1.590	879,389
	Women*		
Alter Retires 1938	0.948	2.028	364,053
Alter Retires 1939	-0.349	1.458	364,053
Alter Retires 1940	-0.213	1.494	364,053
Alter Retires 1941	7.264***	2.048	364,053

* Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is peergroup-year. Each regression includes a set of establishment characteristics (median age and interquartile age range, share of: females, low-skilled, high-skilled, part-time, under 30, over 50, foreigners, agricultural occupations, simple/qualified manual occupations, simple/qualified service occupations, simple/qualified administrative occupations, technical occupations, managerial occupations, engineering occupations, semi/professional occupations), peer group characteristics (share of: females, low-skilled, high-skilled, part-time, under 30, over 50, foreigners), individual characteristics (experience, wage, nationality, education), and year and peer group fixed effects. 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.