Essays on long-term labor market developments and retirement in Germany

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General Introduction

The welfare state, its institutions and the labor market are subject to constant change in Germany. At all times, their current state is an outcome of an ongoing process of adjustment. First elements of insurance against labor market risks were implemented as early as 1260 by miners' associations (Bingener et al, 2009). The evolution of welfare state and according legislation until today was heavily influenced by regime changes, wars, the zeitgeist, as well as social and political necessities.[1] This cumulative dissertation comprises of four chapters and focusses on the development of employees' working careers under Germany's current regime, which emerged following World War 2. Since then, working careers are strongly influenced by major political changes like the German reunification but also by smaller changes in regulatory framework and socioeconomic environment.

The first two chapters deal with labor market earnings, which Barr (2012) lists as the most important source of welfare aside from governmental activities. Both chapters show how working lives and earnings trajectories of current West German employees differ from those of previous generations. The main contribution is an intragenerational comparison of cohorts' earnings inequality and volatility to scrutinize long-term differences and the evolution of labor market outcomes. An example is the increase in unemployment over the time frame considered, which in turn affects distinct cohorts at different ages and, therefore, in different ways. The analyses reflect challenges that German workers face through labor market adjustments caused by e.g. deregulation, deunionization, globalization, and skill biased technological change. The underlying data enables the comparison of complete working life cycles of older cohorts to early and middle stages of younger cohorts' careers. By taking this cohort perspective both studies show how different generations fared at identical ages, uncovering long-term trends and impacts of various labor market developments. Further, such an approach complements more common methods of using annual data or short panels to measure earnings inequality and volatility.

Another great challenge Germany faces is population aging, which exerts financial pressure on the public pension system of the German welfare state. To counteract, Germany introduced reforms that aim to keep persons employed for a longer time and limit pension growth. The last two chapters deal with questions related to this financial pressure and effects on the financial well-being of affected individuals. Therefore, those chapters complement the analysis of the labor market by looking at the end of employees' active working lives – at what happens when workers transition into retirement and how they fare when retired. The chapters concentrate on cohorts that are already retired and analyze questions concerning the German statutory pay-as-you-go (PAYG) pension system. Since old-

[1] For a thorough overview see Bartels (2014).

age security provided by the welfare state is the most important source of old-age income in Germany, trends and results found in these studies provide important evidence for subsequent cohorts' retirement behavior, pension provisions and possible financial problems of the monetarily less fortunate. Central to both papers is a reform that introduces disincentives for early retirement, effectively reducing pensions for early retirees. The underlying data consists of cohorts that are not affected by the reform, cohorts that are partially affected, and cohorts that are completely affected. This allows for disentangling reform effects from time effects.

Outline

The first chapter, *Lifetime Earnings Inequality in Germany*, joint work with Timm Bönke and Giacomo Corneo, is published in the *Journal of Labor Economics* 33(1), 2015. Each author contributed one-third to the article. In the paper, we investigate for the first time the magnitude, pattern and evolution of lifetime earnings inequality and mobility of West German men. We employ German administrative data, the Insurance Account Sample (Versicherungskontenstichprobe, VSKT), and observe complete earnings life cycles for cohorts 1935 through 1949 and initial earnings life cycles for cohorts 1950 through 1969. When looking at cohorts with completed earnings careers, we find that inequality of cross-sectional earnings is u-shaped over the life cycle. Mobility is rather high at the beginning of the life cycle, decreases afterwards and is basically nonexistent after age 40. We then calculate net present values (NPVs) of earnings up to various ages to compare the distributions of lifetime earnings across cohorts. The intragenerational comparison of inequality of NPVs up to age 40 reveals our main result: a strong secular rise of intragenerational inequality in lifetime earnings. West German men born in the 1960s are likely to experience about 85% more lifetime inequality than their statistical fathers. On the contrary, long- and short-term mobility are rather stable across cohorts. We further identify that an intragenerational rise of absence times from the labor market, most notably times of unemployment, especially affects lifetime earnings of workers at the bottom of the distribution. This rise accounts for about 20%-40% of the increase in lifetime earnings inequality. The remaining increase is caused by growing wage dispersion.

The second chapter, *The Dynamics of Earnings in Germany: Evidence from Social Security Records*, is joint work with Timm Bönke and Matthias Giesecke, who each contributed one-third to the paper. This chapter aims to uncover trends in idiosyncratic earnings volatility across generations. It complements the previous chapter by not looking at amounts of earnings, but rather at how the trajectories differ on which those earnings were attained. Employing the same dataset as in the previous chapter, we look at earnings dynamics of West German men born 1935 through 1974. The underlying model is based on a decomposition of residual earnings auto-covariances into a permanent and a transitory component. The transitory component covers short-term earnings fluctuations or earnings insecurity, induced by e.g. temporary job loss, and the permanent component represents long-term earnings divergences, depicting *inter alia* different career paths. Over the period covered, 1960 to 2009, the German labor market undergoes a heavy transformation, i.e. strong deregulation, deunionization and a shift in employment from industrial to service sector. Therefore the study's findings on increases in both components are no surprise. Still, particular trends tend to mirror distinct phases of the transformation process, like recessions followed by deregulation. In terms of magnitude, the transitory component increases most strongly in the early 1970s and 1990s for young workers. This implies that labor market entries become increasingly more unstable, induced by e.g. an increased likelihood of job change, short-term contracts and brief periods of unemployment for labor market entrants. The permanent component increases most for older workers in the early 1980s and the 2000s. Thus, when workers are established at the labor market and the earnings trajectories become more stable, variation between these trajectories increases across generations, i.e. the earnings paths differ more and wage differences widen.

Chapter 3, entitled *Rates of Return and Early Retirement Disincentives: Evidence from a German Pension Reform,* forthcoming in the *German Economic Review*, focuses on earnings after active labor market participation ends: pensions. Statutory pensions make up the greater majority of old age income for most German employees and are therefore a central feature of the German welfare state. As in the next chapter, the motivation for looking at pensions relates to pressure on solvency of pension systems of most modern welfare states, with Germany providing an example of time trends and reform effects. Here I compute the profitability of contributions to the German statutory pay-as-you-go pension system for cohorts 1935-1945. Since the system is of Bismarckian variety, pensions in Germany are strongly related to prior contributions, which in turn relate to employment biographies and wages earned over the life cycle. The profitability indicates the generosity of the pension system and therefore indicates e.g. if low income earners need to worry about their standard of living after retirement. A high profitability would imply that even low contributions potentially translate into sufficient old-age income. The underlying datasets again are excerpts of the VSKT, the SUFs. As in the first two chapters, I use complete biographies and therefore observe all contributions made to the pension insurance. Based on pension claims, I then calculate statistical pensions. The analysis takes place against the background of the introduction of disincentives for early retirement. In contrast to the following chapter, however, this study does not address disincentives in general, instead focusing on the particular disincentive level actually implemented by the reform. On the other hand, this chapter provides evidence for the whole population, including women. To measure profitability, these contributions and pensions are then used to compute internal rates of return (IRR) for each individual. For men, the IRR declines over observed cohorts from about 2.4% to 1.2% and for women from 5.2% to 3.7%. Counterfactual scenarios suggest that about three-quarters of the trend are caused by increased pension contributions and not by the disincentives introduced by the reform. This means that the trend of a declining IRR will most likely continue. This potentially increases old-age poverty risk for future cohorts and might also decrease the willingness to contribute to the public pension system.

The forth chapter, *Effectiveness of early retirement disincentives: individual welfare, distributional and fiscal implications*, is joint work with Timm Bönke and Daniel Kemptner, who each contributed one-third to the paper. The study looks at transitions from employment into retirement. In particular, we scrutinize effects of early retirement disincentives on retirement behavior. Since many modern welfare states face aging societies, information how to influence retirement timing is essential to policymakers. After looking at the generosity of old-age provisions in chapter 3 in a more general manner, this chapter provides an in-depth analysis of a particular reform within the German welfare state. Still, since this study analyses the important transition period from wage earners into transfer recipients, it also complements the first two chapters: it shows how former employees fare when employment ends and how their transition is influenced by the regulatory framework of the welfare state. We focus on male employees with a strong labor market attachment. This allows us to tackle the choice of retirement behavior apart from unemployment or health shocks. We employ Scientific-Use-Files (SUFs) of the VSKT (25% samples) and a detailed model of the German tax and social security system to estimate a structural retirement model. With this model, we examine to what extent disincentives are able to steer retirement behavior and look at financial, fiscal and individual welfare effects. We find that labor market participation and retirement behavior in general are strongly influenced by the level of disincentives. High levels of disincentives even have a prohibitive effect and basically abolish early retirement completely. We further find the net public returns to be about five times higher than individual welfare losses. Still, inequality in remaining lifetime consumption increases. For similar net public returns, individual welfare losses are at least twice as high when generated by indiscriminating pension cuts. In sum, the transition of workers into retirement is strongly influenced by disincentives and when trying to reach a particular net public return with a retirement reform, disincentives are to be preferred over indiscriminating pension cuts.

References

Barr, N. (2012): Economics of the Welfare State (5th ed.), Oxford University Press, Oxford, United Kingdom.

Bartels, C. (2014): Versicherung und Anreize im deutschen Wohlfahrtsstaat, Springer Gabler: Wiesbaden.

Bingener, A., Bluma, L., Sulzer, M., Montz, M. and J. Streb (2009): Vergangenheit und Zukunft sozialer Sicherungssysteme am Beispiel der Bundesknappschaft und ihrer Nachfolger. Ein Forschungsprojekt der Leibniz-Gemeinschaft Past und Future of the Knappschaft-Bahn-See. A Research-Project of the Leibnitz Community, Economic History Yearbook 50(2), 195–217

Chapter 1:
Lifetime Earnings Inequality in Germany*

Abstract: We employ German social security records to investigate intragenerational lifetime earnings inequality and mobility of yearly earnings for 35 cohorts, starting with the birth year 1935. Our main result is a striking secular rise of intragenerational inequality in lifetime earnings: West German men born in the early 1960s are likely to experience about 85% more lifetime inequality than their fathers. In contrast, both short-term and long-term intragenerational mobility are stable. Longer unemployment spells of workers at the bottom of the distribution of younger cohorts contribute to explaining 20%–40% of the overall increase in lifetime earnings inequality.

Keywords: Earnings Distribution, Lifetime Inequality, Intra-generational Mobility.
JEL Classification: D31, D33, H24.

*This chapter is joined work with Timm Bönke and Giacomo Corneo. A similar version is published in the Journal of Labor Economics, see page II of this dissertation. doi: http://dx.doi.org/10.1086/677559

Introduction

Labor income inequality is usually described in terms of a distribution of yearly earnings and such earnings distributions have become more unequal in many advanced economies during the past three decades (e.g., Atkinson and Piketty 2010, Autor et al. 2006, Card and DiNardo 2002, Goos et al. 2009, Lemieux 2008). However, the labor market generates patterns of earnings dynamics that vary across individuals, so that the evolution of inequality of long-term earnings might differ considerably from the evolution of inequality of yearly earnings. A life-cycle perspective recognizes that some levels of earnings are transient and not representative of an individual's position in the long-term distribution, e.g. low earnings during college years and when unemployed, or high earnings thanks to temporarily skyrocketing bonuses. In that perspective, it is the inequality of lifetime earnings that is crucial in order to assess how much inequality is generated by the labor market.

This paper uses a sample of high-quality administrative data to study actual lifetime earnings, their dispersion, and the mobility of individuals in the earnings distribution. We take a cohort perspective and investigate the distributions of earnings of individuals who were born in the same year. Intra-generational inequality of lifetime earnings is important not only because it mirrors long-term disparities in labor-market outcomes. Given the prominence of earnings as a determinant of the lifetime resources available to agents, intra-generational inequality of lifetime earnings is suggestive of inequality of permanent incomes. In turn, inequality of permanent incomes speaks of consumption inequality and is closely related to the social welfare of generations. Intra-generational inequality also matters because individuals tend to compare their earnings with those of people of similar age (Pérez-Asenjo 2011). Finally, a cohort-based analysis of the mobility experienced over the life cycle can help us to better understand the drivers of growing cross-sectional inequality and the ways in which labor markets have changed during the last decades.

Our empirical investigation targets the largest European economy, Germany. We exploit data on earnings biographies from social security records to shed light on the following issues: What is the magnitude of lifetime earnings inequality and how does it compare to measures of inequality of annual earnings? How do cohort-specific inequality and mobility evolve over the life cycle? How is lifetime inequality for individuals who are currently in the middle of their career going to compare with the one experienced by their parents?

In order to answer those questions we analyze the earnings histories of thirty-five birth cohorts in Germany, ranging from individuals who were born in 1935 to those born in 1969. The dataset we scrutinize is a highly representative sample of the employee population in West Germany. We define lifetime earnings as the present value of an individual's earnings until the individual reaches age sixty. For the fifteen oldest cohorts in our dataset we observe all annual earnings until they reach age sixty, so that we can compute their lifetime inequality as well as their mobility in the intra-generational distribution of annual earnings during their entire active life cycle. We observe younger cohorts' earnings only for an initial part of their life cycle and can compute measures of earnings inequality and mobility up to some age. Using both the information about cohorts that have completed their labor-market life cycle and the information about the still-active cohorts, we attempt to gauge how lifetime inequality is evolving across generations in Germany.

We find that the Gini coefficient of the intra-generational distribution of lifetime earnings is about two-thirds of the Gini coefficient of annual earnings. Age-specific annual earnings inequality follows a U-shaped pattern over the life cycle, with a minimum reached around age thirty-five. Even controlling for age, measures of inequality of annual earnings substantially overestimate the inequality of lifetime earnings, the difference between the two measures being due to individuals' mobility in the distribution over time. Within cohorts, mobility in the distribution of yearly earnings is substantial at the beginning of the life cycle, decreases afterwards and virtually vanishes after age forty. Age-earnings profiles are concave and steeper for better-educated individuals.

Our main finding concerns the evolution of lifetime inequality across cohorts. We detect striking evidence of a dramatic secular rise of intra-generational inequality in lifetime earnings: West German men born in the early 1960s are likely to experience about 85% more lifetime inequality than their fathers. In stark contrast, both short-term and long-term intra-generational mobility are rather stable.[2]

The rise of intra-generational lifetime earnings inequality has affected both the bottom and the top of the distribution, but the rise has been stronger at the bottom. We find that some 20–40% of the rise of lifetime inequality can be attributed to an increase in the duration of unemployment for individuals at the bottom of the earnings distribution. The rest is due to an increase of intra-generational wage inequality.

Our paper is related to various strands of literature. Firstly, it relates to the literature on the long-run evolution of earnings inequality. Our finding of a secular rise of intra-generational lifetime earnings is, to the best of our knowledge, a novel one. There seem to be no other studies that attempt to pin down the evolution of the inequality of lifetime earnings. Closest to the current paper is probably the article by Kopczuk et al. (2010) about earnings inequality in the United States. Using social security data, they compute Gini coefficients of cohort-specific long-term earnings distributions since 1937. Long-term earnings are defined as earnings over a twelve-year period and three benchmark periods are considered: from age twenty-five to age thirty-six, from age thirty-seven to age forty-eight, and

[2] Lifetime earnings inequality appears to be on the rise also for women. However, their labor-market behavior tremendously changed in Germany during the last sixty years. As a consequence, our sample of women born in the 1930s is quite different from the one of women born in the 1960s and sample representativeness varies across cohorts. Therefore, we present here only an analysis of men's earnings distributions. Our findings for women are presented in the Online Appendix II of this paper. The aforementioned changes in sample composition are documented in Online Appendix I.6.

from age forty-nine to age sixty. For cohorts born after the mid-1930s, all three measures of long-term earnings exhibit an upward trend of cohort-specific inequality. Our finding that intra-generational inequality of lifetime earnings has increased in Germany points to a remarkable common trend in the two countries.

Second, this paper complements various analyses of how inequality has evolved in Germany over the last three decades. That literature has mainly focused on the cross-sectional distribution of wages and has found that it has become more unequal over time (Gernandt and Pfeiffer 2007, Dustmann et al. 2009, Bach et al. 2009, Card et al. 2013). As shown by Fuchs-Schündeln et al. (2010), similar trends can also be observed for the cross-sectional distributions of household income and consumption, although they find the trend of consumption inequality to be rather flat.

Our paper adds to that literature by establishing how lifetime earnings inequality has changed across cohorts, which is necessary in order to assess how increases in cross-sectional wage inequality translate into inequality experienced over the life cycle. Our findings suggest that the burden of adjusting the German labor market to changing conditions was mainly carried by the low-skilled of the younger cohorts. They also suggest that measures of cross-sectional consumption inequality might underestimate the increase of consumption inequality in Germany. Furthermore, our investigation of age-earnings profiles confirms the importance of controlling for the age composition of the workforce when evaluating long-run changes in cross-sectional distributions.[3]

Third, our work is related to the literature on the relationship between annual and lifetime income inequality and the extent of intra-generational mobility. We contribute to that literature by offering findings based on high-quality data drawn from a sample that is much larger than those analyzed in earlier work. The main previous study is Björklund (1993), who exploits Swedish tax registers to compute the lifetime income before taxes of cohorts of men born between 1924 and 1936. He finds that the Gini coefficient of the distribution of lifetime earnings is around 35%–40% lower than the one for cross-sections of annual incomes and that there is substantial intra-generational mobility during the early stages of the life cycle.[4]

Fourthly, our paper adds to the literature on the life-cycle variation in the association between annual and lifetime earnings.[5] We confirm Björklund's (1993) result that the correlation between

[3] OECD (2008) gives an overview of the impact of demographic change on the income distribution. Almas et al. (2011) provide evidence that changes in the age structure of the workforce had a significant impact on the Gini coefficient of annual earnings in Norway in the period 1967-2000.

[4] Burkhauser and Poupore (1997) compare the distribution of annual earnings with the one of earnings over a six-year period from 1983 to 1988. Using the SOEP, they find that when the Gini coefficient is computed over six years, its level falls by less than ten percent. See also Maasoumi and Trede (2001). Trede (1998) analyzes short-run earnings mobility between 1983 and 1993 using the SOEP. He finds that mobility declines with age until age thirty-five and does not change thereafter.

[5] Implications of that variation for regression models are discussed by Jenkins (1987) and further worked out by Haider and Solon (2006). Böhlmark and Lindquist (2006) apply Haider and Solon's model to Swedish data. An

annual income and lifetime income is high after age thirty-five. With respect to age-earnings profiles, our finding that they are much steeper for university graduates than for uneducated workers is in line with standard models of human capital investment. It also accords well with recent findings by Bhuller et al. (2011) based on Norwegian earnings biographies.

The next Section describes our dataset and defines the variables of interest. Section III quantifies lifetime earnings inequality and compares it with annual earnings inequality. Section IV is devoted to the pattern of earnings mobility during the life cycle. The core of the paper is Section V where we analyze the evolution of intra-generational lifetime inequality and dissect its main driving forces. Section VI concludes.

Data and Methodology

Our analysis is based on administrative data of the German social security. Most employees in Germany mandatorily participate in its national pay-as-you-go pension system which, being of the Bismarckian variety, carefully records all contributors' earnings biographies. The dataset we analyze is based on the Insurance Account Sample (*Versicherungskontenstichprobe*, VSKT for short) of the Federal Pension Register.[6] The VSKT is a stratified random sample of individuals who live in Germany, have at least one entry in their social security record and are aged between thirty and sixty-seven in the reference year of the sample. VSKT waves of reference years 2002 and 2004 to 2009 form the basis of our study.[7] Each sample contains the earnings biographies of the observed individuals up to the reference year. The data are collected following individuals over time so as to form a panel. For each individual, a monthly history of employment, unemployment, sickness, and contributions to the pension system is recorded. It starts when the individual reaches age fourteen and it ends when the individual turned sixty-seven in case of complete biographies. Information about the contributions made to the pension system allows one to recover the earnings received by that individual in each month.

The current investigation focuses on German citizens – including naturalized immigrants with complete earnings biographies in Germany and excluding ethnic Germans that immigrated to Germany after having worked in their country of origin. Because of insufficient comparability of

application of their methodology to correct for the life-cycle bias that uses German earnings data is Brenner (2010).

[6] The final dataset we work with (*FDZ-RV – VSKT2002, 2004-2009_Bönke*) is provided to researchers by the Data Research Centre of the German Federal Pension Insurance. It is accessible through controlled remote computing.

[7] A detailed description of the data is given by Himmelreicher and Stegmann (2008). We use all seven samples in our analysis. Information on birth cohorts 1935 and 1936 is picked from the 2002 sample; cohort 1937 stems from the 2004 sample, cohort 1938 from the 2005 sample, cohort 1939 from the 2006 sample, cohort 1940 from the 2007 sample and cohort 1941 from the 2008 sample. Later birth cohorts are covered using the 2009 sample.

earnings information and wage levels in the FRG and the GDR, we restrict the attention to individuals who have only been working in West Germany.[8] Furthermore, we exclude contributors for whom a consistent earnings biography cannot be reconstructed.[9] In this way we exclude contributors who worked also as self-employed or civil servants, or who emigrated abroad at some point in time, and who may thus have substantial earnings that are not recorded in the Federal Pension Register. After elimination of those observations, we are left with a number of individuals for each cohort that oscillates between 1,000 and 1,600 - see Appendix B, Table B1.[10]

While the dataset we use is virtually free from measurement errors, three adjustments were necessary in order to prepare the earnings data for the analysis. The first one concerns the imputation of one-time payments. Those payments were not included in the social security data before 1984 while they are included from that year onwards. In order to obtain a time-invariant definition of earnings, we exploit the panel structure of our data and estimate each individual's earnings path so as to identify spurious growth between 1983 and 1984. Conditional on an individual's age and position in the earnings distribution we then adjust his earnings before 1984.[11]

Our second adjustment is the addition of the employers' social security contributions (to pension, unemployment, health, and nursing care public insurances) to the individuals' gross earnings. In first approximation, those contributions represent the value of insurance that the employees would have purchased if it had not been provided by the government. Adding those elements of pay is warranted in order to take into account the heterogeneity of insurance protection offered to the various subgroups of the working population - subgroups whose relative weights in the working population have substantially changed across cohorts.[12] Thus, the earnings measure we employ is a measure of the market value of labor. As a major robustness check, we have repeated the entire analysis when the employer contributions are excluded. As shown in Online Appendix III.2, all findings remain qualitatively unaltered - in particular the rise of lifetime earnings inequality retains the same order of magnitude when employer contributions are excluded.

Third, we deal with the issue of top-coded earnings. In Germany, employees contribute a share of their gross wage to the mandatory pension system up to a wage ceiling. As a result, our social

[8] West-East migration was almost inexistent before reunification; after reunification it affected a tiny share of the labor force from West Germany, see Fuchs-Schündeln and Schündeln (2009).

[9] More precisely, we only allow for an average of one month of missing information per year after the age of thirty. For further details see Online Appendix I.4.

[10] In Online Appendix I.5.we document how many individuals are originally included in the dataset and how many remain after eliminating individuals that do not satisfy our selection criteria.

[11] See Online Appendix I.3 for further details and a robustness check. Our method to correct for the 1984 break extends the one proposed by Fitzenberger (1999) and used by Dustmann et al. (2009) and Card et al. (2013) in a cross-sectional setting so as to make it suitable for a longitudinal analysis. While also those papers investigate social security records, their datasets stem from the Employment Register of the Federal Labor Office.

[12] Otherwise, it would be highly problematic to include in the analysis some categories of employees like miners, sailors and distinctive employees of the federal railways that have special social security arrangements. In Online Appendix I.1 we relate the evolution of contribution rates and contribution ceilings.

security data is right-censored as individuals whose wages exceed that ceiling are recorded as if their wages were equal to the ceiling. On average over all years and cohorts, censoring concerns about seven percent of the recorded earnings of men.[13] In order to better approximate the true distribution of top earnings, we impute them to the individuals affected by top coding. Our imputation method rests on the assumption that the upper tail of the earnings distribution behaves according to the Pareto law. We posit that the top ten percent of individual earnings below the contribution ceiling are Pareto-distributed. Then, we estimate the corresponding Pareto-coefficient by OLS. The estimation is conducted separately for all years and birth cohorts. The estimated Pareto-coefficients are then used to determine the distribution of the unobserved earnings above the contribution ceiling. The assignment of estimated earnings to individuals is done so as to preserve the individual rankings in the distribution of annual earnings. Thereby, the rank of an individual is based on the last observable rank in relation to all individuals at or above the contribution ceiling in the cohort-specific earnings distribution. We also explore the implications of two alternative imputation methods: an imputation of the estimated mean income above the ceiling to all individuals with top-coded earnings and a maximum mobility scenario where the ranking order is reversed every year. Results from those alternative imputations are reported in Online Appendix III.3. They do not differ much from those obtained under our preferred rank-preserving assumption.[14]

In order to validate the earnings data we work with, we have compared it with the corresponding earnings data from the SOEP, i.e. earnings data that concern the same population in terms of gender, age, region, and employment status as the one we investigate. The SOEP is based on an annual survey of private households and is constructed so as to be highly representative of the total population in Germany. As shown in Appendix A, the cross-sectional earnings distributions obtained from the VSKT reproduce remarkably well those obtained from the SOEP for the same years and the two are statistically undistinguishable. Furthermore, the SOEP data reveal that the VSKT represents about 80% of the total male labor force in West Germany.

[13] Further information about how censoring affects our sample is provided in Online Appendix I.2. There we also provide additional information on our imputation procedure.

[14] In Online Appendix III.7 we also present a robustness check concerning the bottom of the distribution. Legislated exemptions from social security contributions may lead to an underrepresentation of very low earnings in some years. As it turns out, simulating a constant exemption regime over time generates qualitatively the same results as the ones reported here.

I. Inequality of Lifetime Earnings

We compute lifetime earnings from the monthly earnings an individual has received from age seventeen to age sixty.[15] Given that age limit, we can determine the complete lifetime earnings of fifteen cohorts, born between 1935 and 1949. When computing lifetime earnings, we discount yearly earnings to the year the individual turned seventeen and then determine the corresponding present value of earnings. We set the discounting rates equal to the average nominal returns on German government bonds, obtained from an official time series provided by the German central bank.[16] As a robustness check, we discount earnings using the consumer price index.

Results about the Gini coefficient of the cohort-specific distribution of lifetime earnings for men are displayed in Figure 1. The lowest curve represents the Gini coefficient of lifetime earnings when annual earnings are discounted using the rate of returns on German federal bonds. The Gini coefficient reaches a minimum of 0.156 for the cohort born in 1935 and peaks at 0.212 for the one born in 1949. The curve in the middle of Figure 1 obtains when annual earnings are discounted using the consumer price index. The discounting method affects the level of lifetime inequality but not its evolution. A lower discount rate increases intra-generational inequality because of the steeper rising age-profile of earnings for better educated workers, who are also those with the higher lifetime earnings. We display age-earning profiles in the next section.

Because of earnings mobility, inequality in lifetime earnings is smaller than inequality in annual earnings. The curve in the upper part of Figure 1 helps to compare yearly inequality with lifetime inequality. It depicts the average of the Gini coefficients of the distribution of yearly earnings for each cohort. That average Gini coefficient ranges from a minimum of 0.262 for the 1938 cohort to a maximum of 0.336 for the 1949 cohort. Hence, Gini coefficients of lifetime earnings distributions are somewhat less than two-thirds of the corresponding average Gini coefficients of annual earnings distributions. Inequality measured from annual earnings substantially overestimates the inequality of lifetime earnings, but the latter is by no means negligible.

[15] For months during which no earnings are recorded (e.g. in case of unemployment or schooling) individuals are assigned zero earnings; see Online Appendix I.4 for further details.
[16] Details on the methodology used to compute the time series are available at http://www.bundesbank.de/statistik/statistik_zeitreihen.php?lang=de&open=zinsen&func=row&tr=WU0004.

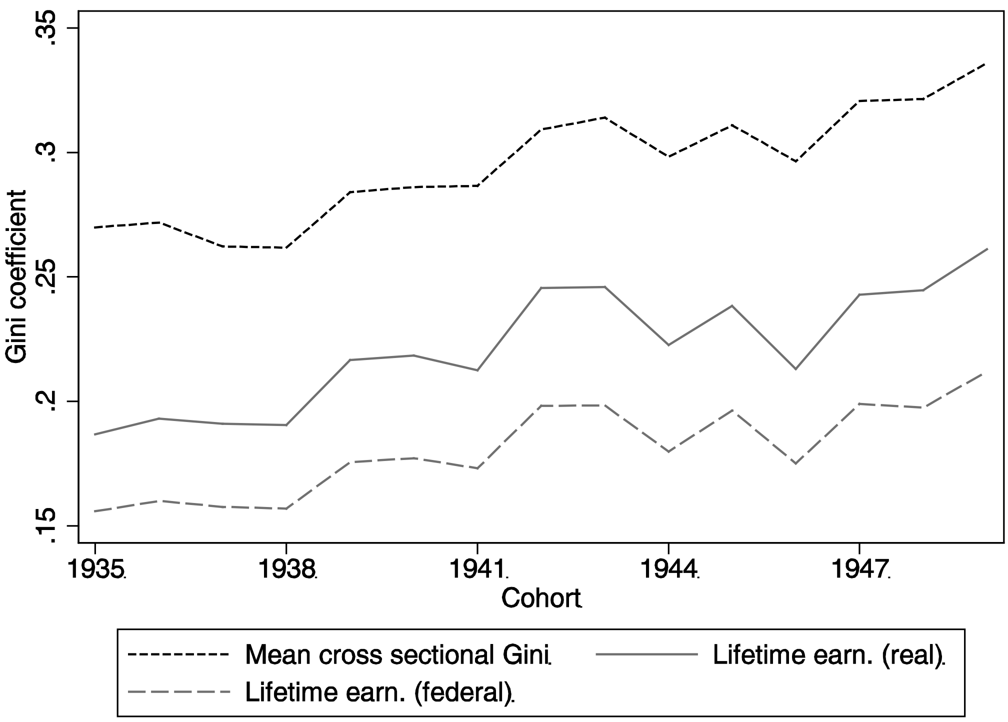

Fig. 1.– Means of annual Gini coefficients and Gini coefficients of lifetime earnings for cohorts 1935 – 1949.
Note.– "real" denotes CPI discounting, "federal" denotes federal bond discounting.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

II. Inequality and Mobility over the Life Cycle

We are now in a position to assess how intra-generational inequality develops along the whole life cycle and how it relates to lifetime inequality. Figure 2 shows for selected cohorts the evolution of the Gini coefficient of annual earnings as a cohort grows older. A U-shaped pattern clearly emerges from the data. Inequality is maximal when the cohort is below twenty because many individuals have not yet entered the labour market and thus have zero earnings. Inequality then declines and reaches a minimum when the cohort is in its mid-thirties. After that, a period of rising inequality of annual earnings sets in.[17] At the time individuals are sixty-years old the distribution of their annual earnings exhibits about the same Gini coefficient as the distribution that prevailed when they were twenty-years old. This pattern is consistent with the presumption that better educated workers have a steeper age-earnings profile, something to which we return below. The sudden and short-lived rise of

[17] Models of stochastic earnings dynamics focus on employed individuals and predict that, for any cohort, earnings inequality grows with age. See e.g. Deaton and Paxson (1994) and Huggett et al. (2011).

annual inequality for men in their early twenties born in 1938 and thereafter can be attributed to mandatory military and civil service which entail a temporary lack of earnings.[18]

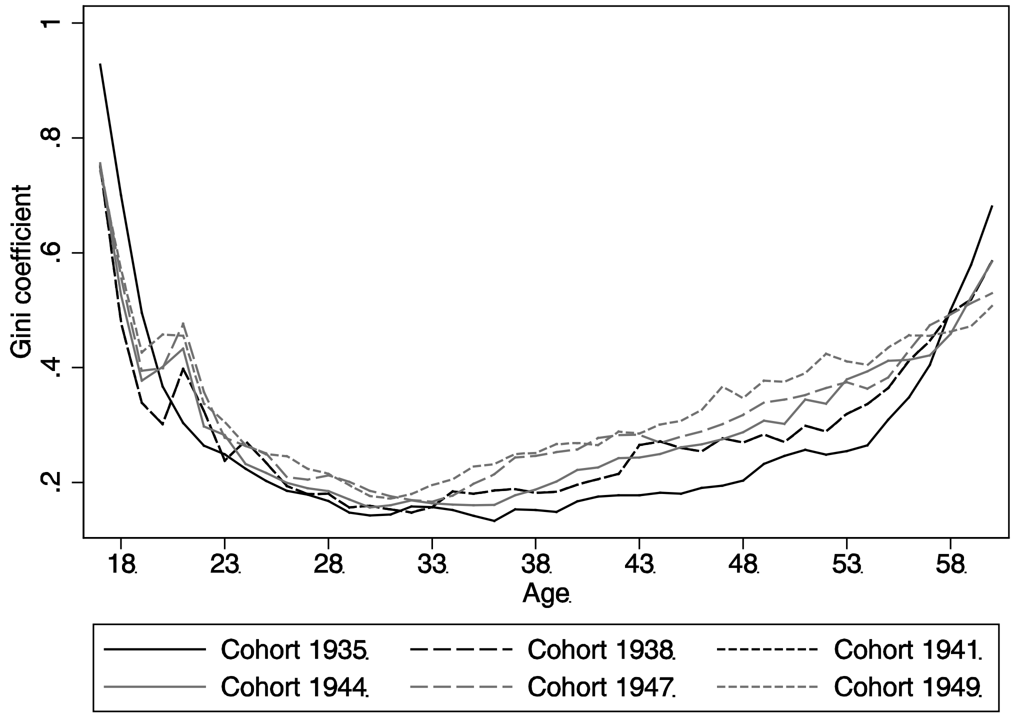

Fig. 2.– Annual Gini coefficients from age 17 to age 60 for cohorts 1935 – 1949

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

Figure 3 shows for selected cohorts the correlation of individuals' ranks in the earnings distributions of two consecutive years. The displayed correlation coefficients are inversely related to the short-run mobility of individuals in the cohort-specific earnings distribution: the lower is that coefficient, the higher is their mobility. As shown by Figure 3, some intra-generational mobility always exists during the life cycle and that mobility decreases with age.[19] While there is significant mobility when the cohort is in its twenties, mobility virtually vanishes when the cohort enters its forties.

[18] Individuals in our sample who were born before Juli 1937 were not affected by drafting. The effect on subsequent cohorts is heterogeneous because of changes in the mandatory serving time.

[19] The drop of the rank correlation for the 1935 cohort when it reaches age fifty-five is due to early-retirement. Changes in legislation and workforce composition entailed a reduced incidence of early retirement for subsequent cohorts.

Fig. 3.– Earnings rank correlations between consecutive years for cohorts 1935-1949

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

Further details on mobility are provided by the rank correlation between annual and lifetime earnings. As shown by Figure 4, there is a distinctive age pattern. When adulthood begins, annual earnings contain virtually no information about lifetime earnings as their mutual correlation is close to zero. The correlation between annual and lifetime earnings then rapidly increases with age. A correlation coefficient of 0.9 is reached when the cohort is at the end of its thirties and such a high level persists until the mid-fifties. In that period of the life cycle the level of individuals' annual earnings can be considered as a good proxy of their respective lifetime earnings.[20]

[20] Unless stated otherwise, we shall always present the findings obtained when using the German federal bond rate as the discount rate. Online Appendix III.4 contains the corresponding findings obtained when using the CPI.

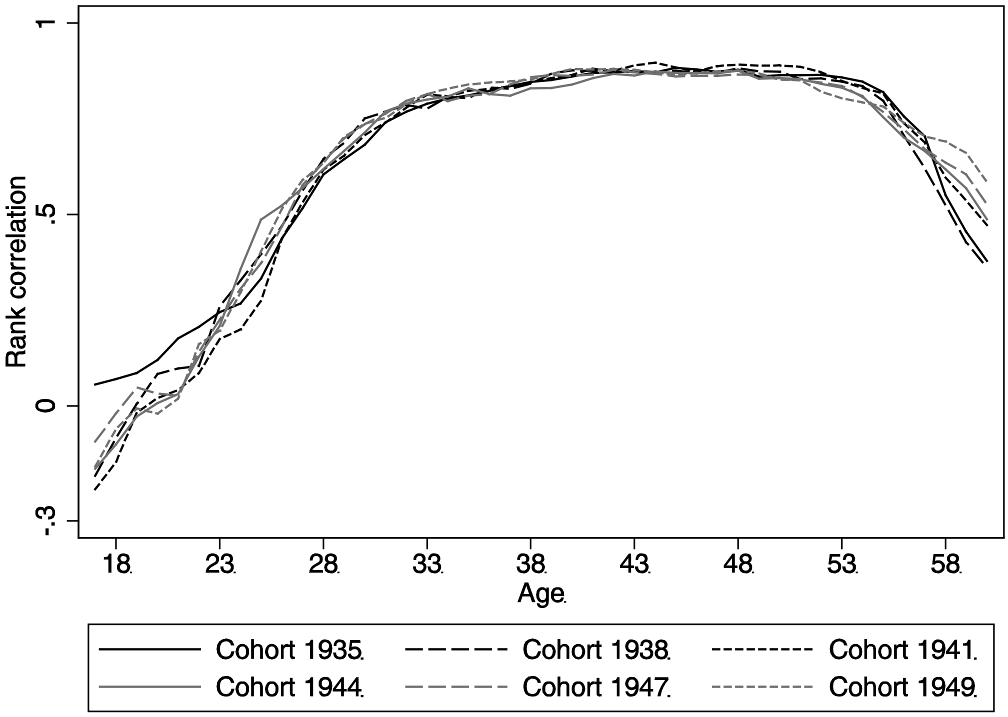

Fig. 4.– Rank correlation of annual and lifetime earnings for cohorts 1935-1949

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

The role of mobility in shaping long-term inequality can be assessed by computing the effect of rank changes in the earnings distribution over a small number of years on the inequality of the present value of earnings received up to certain a age. For that purpose, we employ the concept of "up-to-age-X" earnings, UAX for short. For a given individual, UAX is the present value of all his earnings before the individual becomes X-years old. The higher X, the closer that earnings measure to lifetime earnings, and the two concepts coincide if $X = 60$.

In order to measure the impact of mobility on the UAX distribution, we decompose the change in the Gini coefficient of the UAX distribution into two components, one that mirrors the growth of earnings in different parts of the distribution, and one that mirrors the re-ranking of individuals in the UAX distribution. Our decomposition method follows the one developed by Jenkins and Van Kerm (2006) in a related framework.

Let $G_{X,c}$ denote the Gini coefficient of the UAX distribution for a cohort c. We are interested in decomposing the change $\Delta_{X,c} = G_{X+5,c} - G_{X,c}$, i.e. the change in the Gini coefficient of the present value of earnings at a given age and five years later. From the covariance definition of the Gini coefficient (Lerman and Yitzhaki, 1985), we have:

$$G_{X,c} = \frac{2\operatorname{cov}\left(W_{X,c}, F\left(W_{X,c}\right)\right)}{E\left[W_{X,c}\right]}$$

where $W_{X,c}$ represents the present value of earnings that members of cohort c have received between age 17 and age X. Furthermore, $E[W_{X,c}] = \mu_{X,c}$ denotes the mean of those earnings and $F(W_{X,c})$ their cumulative density function.

If one keeps the ranking of individuals in the original UAX distribution when computing the Gini coefficient of the UAX distribution five years later, the following concentration coefficient obtains:

$$C_{X+5}^{(X)} = \frac{2\,\mathrm{cov}\big(W_{X+5}, F(W_X)\big)}{\mu_{X+5}}$$

where we have suppressed the cohort index for notational simplicity. Hence, the difference between G_{X+5} and $C_{X+5}^{(X)}$ captures the re-ranking effect, while the remaining portion of the change in the Gini coefficient of the UAX distribution is due to heterogeneous earnings growth at the various ranks. This invites one to partition the change in the Gini coefficient as

$$\Delta_X = \underbrace{\left[G_{X+5} - C_{X+5}^{(X)}\right]}_{\equiv R_X} - \underbrace{\left[G_X - C_{X+5}^{(X)}\right]}_{\equiv P_X}$$

where

$$R_X = \frac{2}{\mu_{X+5}}\left[\mathrm{cov}\big(W_{X+5}, F(W_{X+5})\big) - \mathrm{cov}\big(W_{X+5}, F(W_X)\big)\right]$$

is the re-ranking effect and $R_X = 0$ if no re-ranking occurs. Furthermore, the term

$$P_X = \frac{2}{\mu_X \mu_{X+5}}\left[\mathrm{cov}\big(W_X, F(W_X)\big)\,\mu_{X+5} - \mathrm{cov}\big(W_{X+5}, F(W_X)\big)\mu_X\right]$$

captures the relative average earnings growth between the two periods, where the growth is weighted by the earnings hierarchy in the initial distribution. Following Jenkins and Van Kerm (2006), P_X measures the progressivity of earnings growth: $P_X > 0$ ($P_X < 0$) indicates that earnings growth is concentrated at the lower (upper) end of the distribution, which leads to decreasing (increasing) inequality over time.

We now employ the above framework to decompose the changes in the inequality of UAX measured between the age of 25 and 30, 26 and 31, and so on, up to age 55 and 60. Figure 5 plots our decomposition results for the cohort of 1944. The continuous line, indicating the change of the Gini coefficient in each five-year interval, shows that the UAX distribution becomes more equal during the initial part of the life cycle and that inequality starts increasing when the cohort enters its forties. The two dashed lines describe the progressivity effect and the re-ranking effect, as of Eq. (3). Figure 5 shows that the change in UAX inequality as the cohort grows older is mainly driven by the progressivity effect: earnings growth is markedly pro-poor before the cohort enters its forties and switches to pro-rich thereafter. The effect from re-ranking peaks at the beginning of the life cycle and declines afterwards. Its influence on the development of UAX inequality becomes negligible in the second half of the life cycle, which means that five-year mobility in that earnings ladder is nearly non-existing during the second half of the life cycle. As shown in Online Appendix III.5, the pattern revealed by Figure 5 carries over to the remaining cohorts.

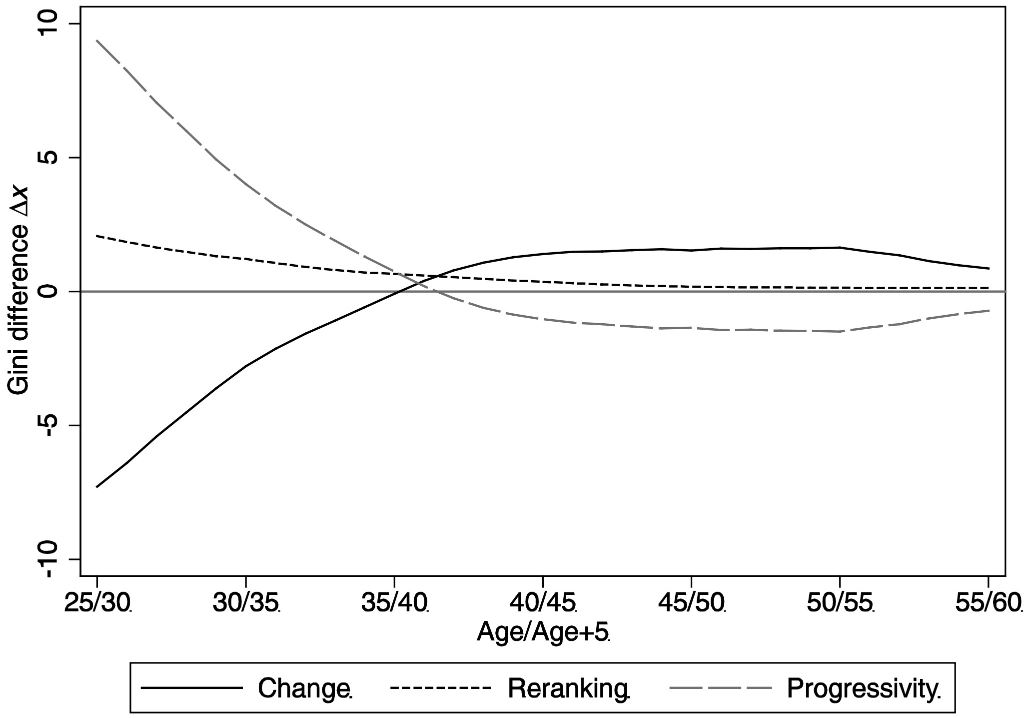

Fig. 5.– Decomposition of changes in inequality as of Eq. (3) for cohort 1944

Note.– Accumulated discounted earnings refer to the age in the abscissa as compared to accumulated earnings five years later, as in Eq. (3). Coefficients are multiplied by 100.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

It is interesting to relate the various mobility patterns detected above to the age-earnings profiles of individuals with different educational attainments. In Figure 6 we plot those profiles for three levels of education for the pooled cohorts from 1935 to 1949. The horizontal lines depict the annualized value of the corresponding present value of lifetime earnings. All earnings are in real terms, on the basis of prices in 2000, and expressed in logs. For each educational group, its profile has a mainly rising, concave shape. However, the higher educated individuals experience more rapid earnings growth through the entire life cycle. This is consistent with the kind of earnings dynamics suggested by standard human-capital theory.

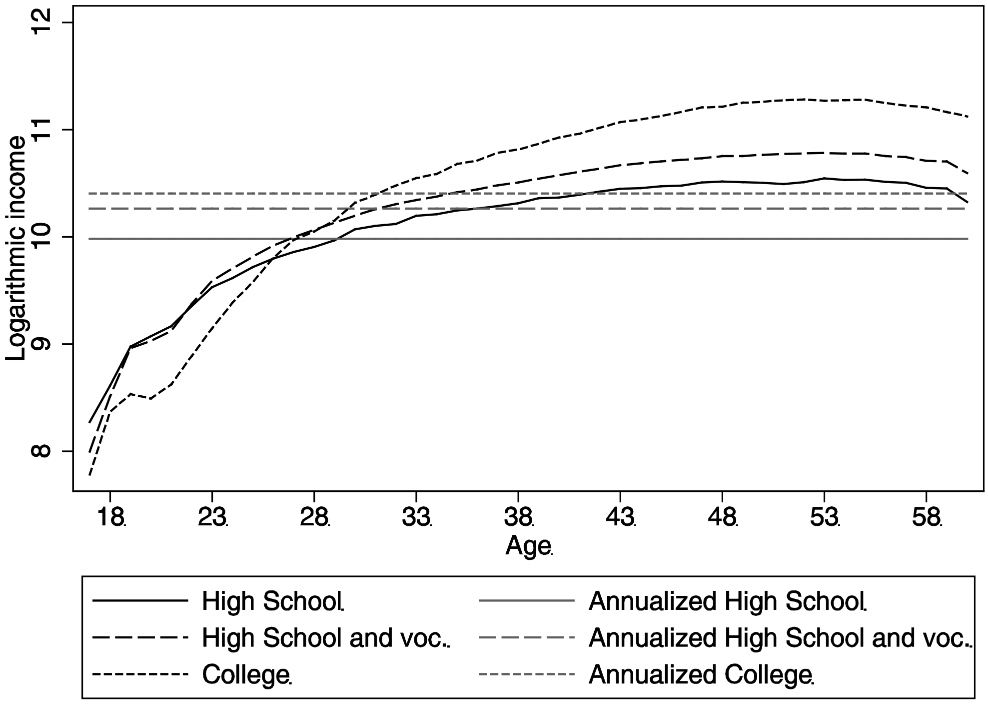

Fig. 6.– Age-earning-profiles by highest educational attainment for pooled cohorts 1935-1949

Note.– voc. abbreviates vocational training.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

III. Evolution of Lifetime Inequality

Are cohorts in Germany becoming more or less equal in terms of their lifetime earnings? This question cannot be satisfactorily answered by examining just the cohorts born between 1935 and 1949 for which lifetime earnings can be computed. We now exploit also the data available for younger cohorts in order to uncover patterns of the long-run evolution of lifetime earnings inequality.

A. Main Finding

We resort to the concept of "up-to-age-X earnings", UAX for short. As already mentioned, UAX is the present value of an individual's earnings before the individual becomes X-years old. For each cohort, the Gini coefficient of the distribution of UAX can be computed for different values of X. Establishing how the Gini coefficient of the distribution of UAX has evolved over successive cohorts can provide valuable hints about the underlying evolution of lifetime earnings inequality. If younger cohorts display higher Gini coefficients for the same X and if this applies to all X, that would strongly suggest that there is a trend of increasing lifetime earnings inequality.

The results in the previous section indicate that mobility in the earnings distribution is significant until about age forty. Therefore, we focus on the distribution of UAX for $X \geq 40$. The data allows us to compute UAX for $X \geq 40$ for all thirty-five cohorts born between 1935 and 1969. For each cohort and each definition of X, we then compute the Gini coefficient of the distribution of UAX.

Representative results are displayed in Figure 7 for earnings up to the ages of 40, 45, 50, 55, and 60 (lifetime earnings). They show that Gini coefficients trend upwards for each value of X. This indicates that younger generations are likely to experience more intra-generational lifetime economic disparity than their statistical parents. [21]

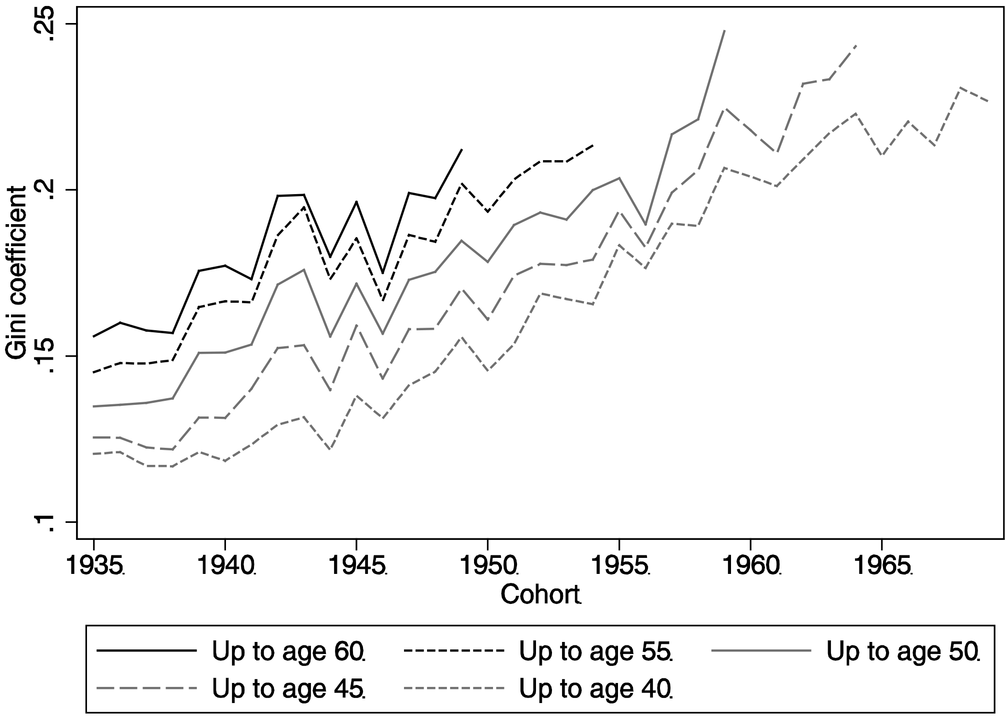

Fig. 7.– Gini coefficients of UAX for cohorts 1935-1969

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

The overall increase in intra-generational earnings inequality is remarkable. To illustrate, compare the cohort of men born in 1935 with the cohort born in 1963, which may respectively be seen as "fathers" and "sons". When they reached age forty-five, the fathers' generation was characterized by a distribution of accumulated earnings with a Gini coefficient of about 0.126. At the same age, their

[21] Statistical inference shows that the observed trend of increasing inequality is significant. Confidence intervals for UAX Ginis are provided in Online Appendix III.1.

sons' generation was characterized by a distribution of accumulated earnings with a Gini coefficient of about 0.233, an increase of inequality by roughly 85 %.

A similar finding obtains if we replace the Gini coefficient with an interquantile ratio. Figure 8 plots the evolution of the ratio between the UAX at the 85th quantile and the one at the 15th quantile.

Figures 7 and 8 show that the finding that inequality of accumulated earnings increases with age after age forty holds true for all cohorts. As indicated by the decomposition analysis in Section IV, cohort members who by age forty have received larger earnings tend to experience a stronger earnings growth at a later age. Furthermore, inequality comparisons across cohorts tend to be rather unaffected by the age at which they are made. By way of an example, relative to its neighbouring cohorts, the cohorts of 1942 and 1943 are characterized by a large inequality of UAX and that is true for all $X \geq 40$. This suggests that the evolution of inequality of lifetime earnings is likely to mirror the evolution of inequality of earnings up to age forty.

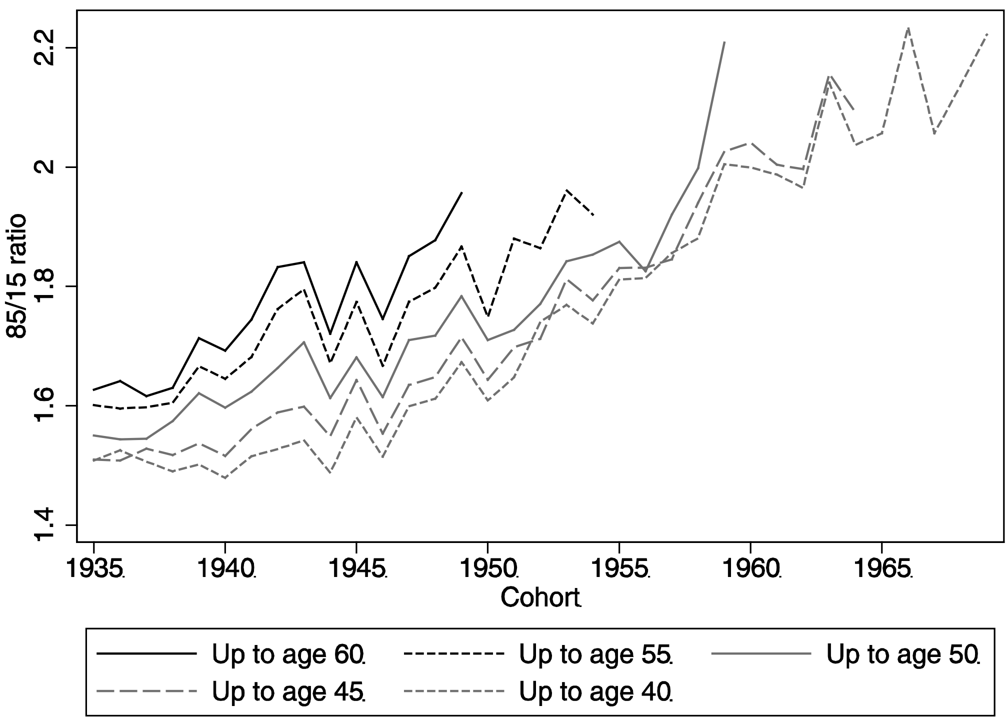

Fig. 8.– 85th / 15th ratio of UAX- earnings for cohorts 1935-1969

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

Our finding of a rising intra-generational inequality does not hinge on the fact that younger generations enter the labor market at a later age. The same pattern as in Figure 7 obtains if UAX are computed starting with a higher age so that virtually all individuals in the sample participate in the labor market in all years when their earnings are taken into account.[22]

The dramatic rise of intra-generational inequality manifests itself also in the distributions of annual earnings received by the various cohorts at a common age. Figure 9 is based on the earnings distributions at ages 40, 45, 50 and 55 as earnings at those ages are good proxies of lifetime earnings. The figure shows that at any given age the Gini coefficient of annual earnings tends to be higher for the younger cohorts.

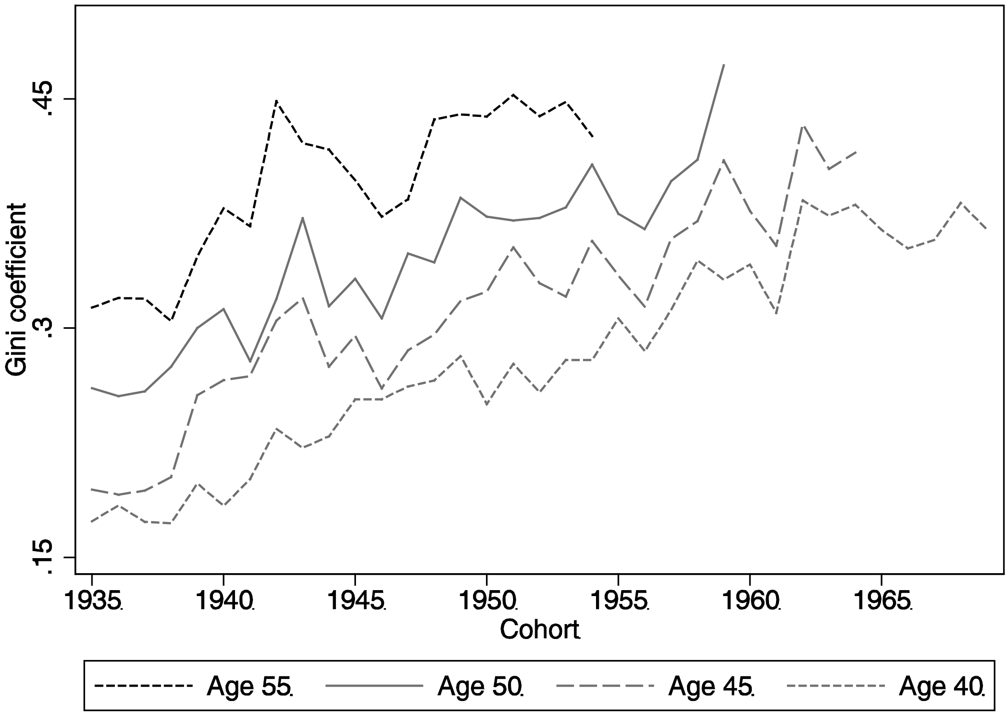

Fig. 9.– Gini coefficients of annual earnings at various ages for cohorts 1935-1969

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

The rise of intra-generational inequality concerns all education groups that can be identified within our dataset. As shown in Appendix D, within-group inequality of the UAX distribution is systematically higher for the younger cohorts. This suggests that the increase in lifetime inequality is not simply driven by the expansion of tertiary education.[23]

Further insights into the evolution of intra-generational inequality come from an analysis of the evolution of mobility after age forty. For each cohort, we compute the correlation between the individuals' ranks in the distribution of UAX for $X = 40$ with their ranks in the distribution of UAX for $40 < X \leq 60$. Representative findings for $X = 45, 50, 55$, and 60 are plotted in Figure 10. No

[22] See Online Appendix III.6.

[23] This finding should be taken with some caution as the VSKT fails to report the educational attainment of about 40% of the sample and the share of missing information is especially high in the case of older cohorts.

major change in mobility across generations can be detected. By way of an example, the rank correlations observed for the cohort born in 1935 are virtually undistinguishable from those observed for the 1963 cohort for the same X.

In Figure 10 we also plot the rank correlation of UA-35 with UA-40, which is distinctively affected by the dynamics of earnings in that period of the life cycle in which most individuals settle into stable employment. Also that correlation varies little across cohorts.

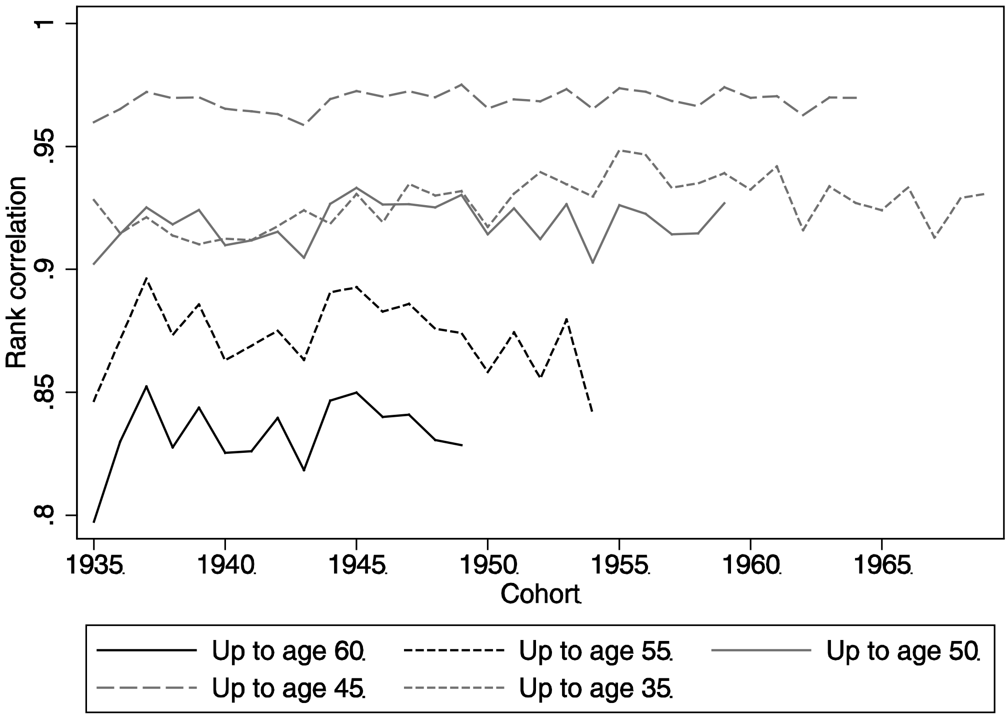

Fig. 10.– Rank correlation of UA-40 with selected UAX for cohorts 1935-1969

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

B. Proximate causes

The aim of the remaining part of our paper is to get some insight into the proximate causes of the rise of lifetime earnings inequality in Germany. As a first step, we are interested in how lifetime earnings inequality for men has evolved at various parts of the distribution. This can be assessed by means of generalized entropy inequality indices that are more sensitive to distinctive parts of the distribution. Results for the Theil index, the mean logarithmic deviation and half the squared coefficient of variation are reported in Online Appendix III.5. They suggest that intra-generational lifetime inequality has significantly increased both at the bottom and at the top of the distribution. Here, we merely present the evolution of two interquantile ratios of the UAX distribution that respectively capture inequality at the bottom and at the top of the distribution. In Figure 11, the left panel plots the 50th / 15th ratio while the right panel plots the 85th / 50th ratio. They show that while lifetime earnings inequality has increased both at the bottom and at the top of the distribution, the increase has been stronger at the bottom of the distribution.

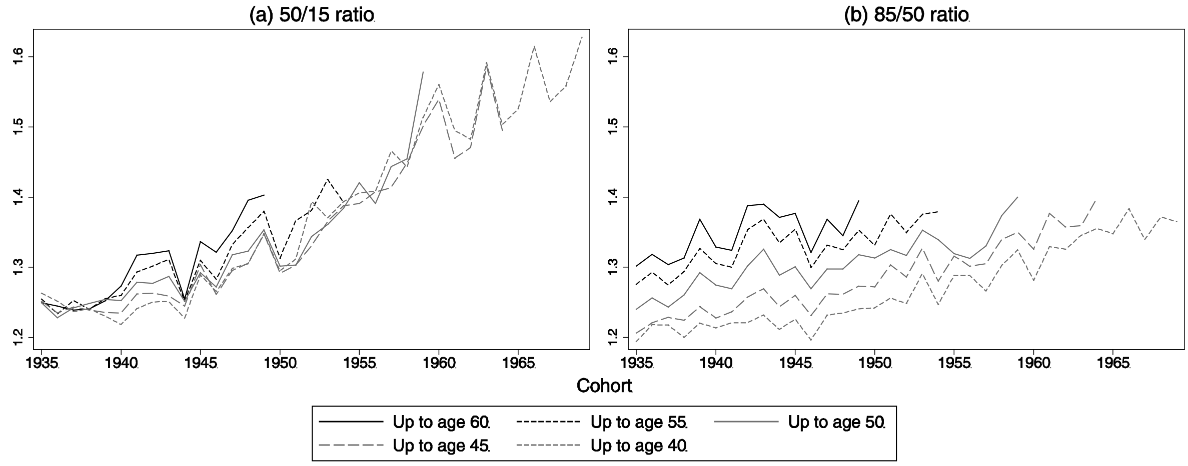

Fig. 11.– 50th / 15th and 85th / 50th ratio of selected UAX for cohorts 1935-1969

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

The second step of our analysis is a decomposition of the inequality increase into a part due to increased wage dispersion and one due to longer unemployment spells. The first part refers to months with strictly positive earnings while the second one refers to months with zero earnings. The interest of this decomposition lies in the distinctive temporal pattern exhibited by the unemployment rate in West Germany. Until the first oil shock, almost full employment prevailed. Then, a strong stepwise increase of the unemployment rate set in which lasted about three decades. Individuals with a low educational attainment were severely hit by the rise of unemployment.[24]

Figure 12 plots for each cohort the average number of months spent in employment, registered unemployment, and other ways during the life span that goes from age seventeen to age forty. The residual category "Other" mainly includes periods of education and of community or military service as well as periods of missing information. Within each cohort, individuals have been ranked into quartiles according to their lifetime earnings up to age forty.

Over time, there has been a substantial increase in the number of months of unemployment for the bottom quartile, a moderate increase for the next quartile, and virtual stability for the upper half of the distribution. Individuals in the bottom quartile of the earnings distributions of cohorts born in the mid-1930s spent on average about 5 months in unemployment before reaching age forty. By contrast, their statistical children born in the early 1960s spent about 41 months in unemployment

[24] During the last three decades, the unemployment rate of individuals with a low educational attainment has usually been at least twice the average unemployment rate (Reinberg and Hummel 2007).

before reaching age forty. For individuals in the upper half of the distribution, no comparable rise of unemployment incidence for the younger cohorts can be observed.[25]

The findings shown in Figure 12 fit well with the notion that the rise of unemployment after the first oil shock severely hit workers with low skills. Moreover, the Figure reveals that low-skilled unemployment was very unevenly distributed across cohorts, with the younger generations carrying most of the burden. This is consistent with the view that hiring and firing costs entail a higher unemployment risk for the entrants in the labor market than for the incumbents.

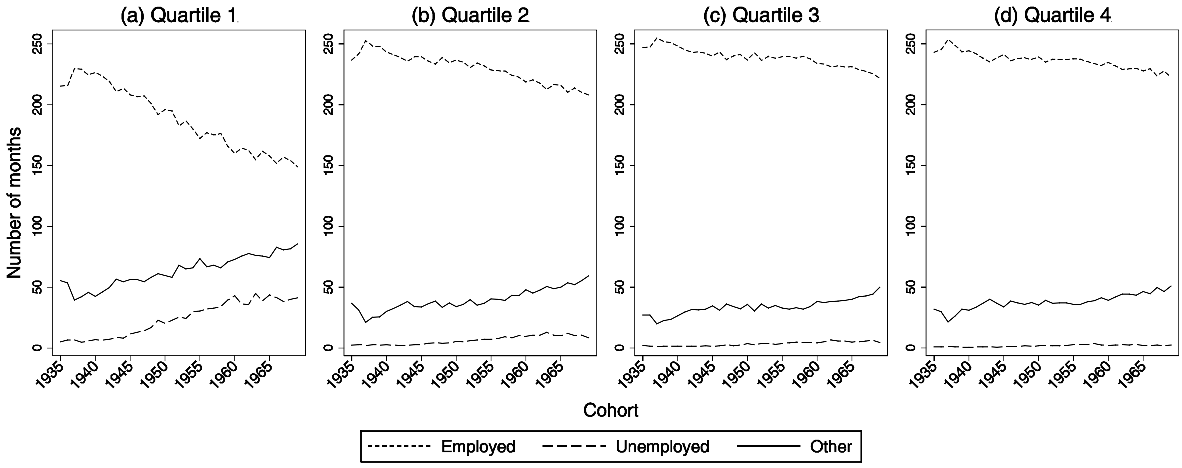

Fig. 12.– Months of employment status up to age forty by quartile of UA-40 for cohorts 1935-1969

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

In order to disentangle the effect on lifetime earnings inequality due to changes in the distribution of unemployment spells from the one due to changes in the wage structure, we simulate the evolution of lifetime inequality under the counterfactual of full employment. In this way, we estimate the intergenerational change of lifetime inequality that had occurred in a hypothetical labor market without unemployment. In a first approximation, a situation of full employment characterized the oldest cohorts in our sample. Hence, the rise of lifetime inequality computed under the counterfactual of full employment is a first approximation of the rise of lifetime inequality due to changes in the wage structure, while the difference between actual and hypothetical inequality rise captures the effect from changes in unemployment spells.

Based on the actual earnings distribution, we construct full-employment scenarios by imputing earnings when individuals have none in the original data. The imputed value for an individual is the

[25] The same striking difference obtains if one only considers the spells of unemployment after age twenty-five. See Online Appendix III.6.

last level of strictly positive monthly earnings that is observed for that individual.[26] Two full-employment scenarios are considered. In one, earnings are imputed only for the months during which an individual was registered as unemployed. In the other, earnings are imputed for all months in which an individual was not in employment. This is based on the notion that protracted periods of education and in the military and periods of missing information may mirror the inability to find a job.

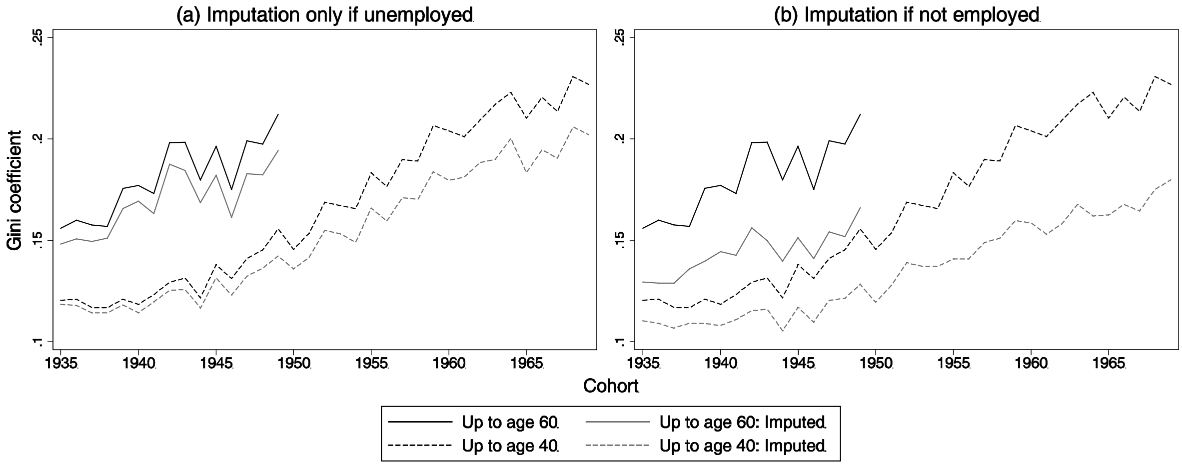

Fig. 13.– Gini coefficients of UA-40 and UA-60 with and without earnings imputation in case of unemployment

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

Figure 13 compares the inequality of lifetime earnings and UA-40 for the various cohorts with the corresponding inequality under the counterfactual of full employment.[27] It shows that the unequal evolution of unemployment spells goes some way in explaining the rise of lifetime earnings inequality. While imputing earnings in case of unemployment has a relatively small impact on the Gini coefficients of UAX for the older generations, it substantially lowers them for the younger generations.

To illustrate our results, one may again consider the cohort born in 1935 and the one of their statistical children born in 1963. Under the counterfactual of no unemployment underlying panel (a) of Figure 13, at the time parents reached age forty-five their accumulated earnings were distributed with a Gini coefficient of about 0.123. At the same age, their children's generation was characterized by a distribution of accumulated earnings with a Gini coefficient of about .207 - an increase of inequality by about 68 %. In the scenario covered by panel (b), the same comparison yields an

[26] In cases where no previous strictly positive individual earnings are observed, we impute retrospectively the first level of strictly positive earnings observed for that individual. In an additional scenario, we reversed our imputation procedure and imputed the level of earnings observed when the individual exits unemployment. Results were similar to those based on our preferred imputation and can be obtained upon request.

[27] The results for the inequality of UA-45, UA-50 and UA-55 are presented in Appendix C, Figure C1.

increase of the Gini coefficient by about 52 %. In both cases, the Gini coefficient increases by much less than 85 %, the actual growth rate of UA-45 inequality between the two cohorts. This suggests that the unequal evolution of unemployment spells for individuals at different points of the earnings distribution contributes to explain some 20 to 40 percent of the secular rise of lifetime earnings inequality.[28]

Using the same imputation method to compute interquantile ratios of UAX distributions under the counterfactual of full employment gives some insight into the effect of unemployment on lifetime inequality at bottom versus top of the distribution. As we report in Appendix C, imputation has little impact on the 85th / 50th ratio while it substantially decreases the 50th / 15th ratio. By way of an example, the 50th / 15th ratio of the UA-45 for the two cohorts considered above increases from 1.25 to 1.59 without imputation while it goes from 1.24 to 1.45 in the case of imputation for registered unemployment. Thus, the rise of unemployment contributes to explain increasing lifetime inequality at the bottom of the distribution but not at the top.[29]

The remaining 60 to 80 percent of the secular rise of intra-generational lifetime earnings inequality can be attributed to the evolution of the cohort-specific wage structure, i.e. the distribution of strictly positive monthly earnings received by a cohort. Unfortunately, our dataset does not contain information about working time, so that we cannot distinguish between the role played by the inequality in hourly wages and the one played by the inequality in hours worked. Cross-sectional evidence from other sources suggests that both types of inequality increased during the last decades but it remains to be seen to what extent this holds true for cohort-specific distributions.[30]

Cohort-specific unemployment and wage structure may be related to their cross-sectional counterparts. As mentioned in the Introduction, several studies have found that cross-sectional wage rates have become more unequal in West Germany during the last decades. According to Dustmann et al. (2009), skill-biased technological change is the best explanation for the widening of the dispersion of wage rates at the top of the distribution. Changes in labor market institutions – in particular, declining union power – and labor supply shocks – in particular, immigration waves – are seen as key drivers of the growth of wage inequality at the bottom. Labor market institutions are also frequently blamed for the rise of unemployment in West Germany since the mid-1970s, although views differ on the relative importance of shifts in unemployment compensation, employment protection and union power (Hunt 1995, Nickell et al. 2005).

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[28] In the case of full employment described by panel (a) the share of the inequality increase approximately attributed to the rise of unemployment is (85-68)/85 = 0.2. In the case of panel (b) we have (85-52)/85 = 0.39. Online Appendix IV shows that this approximation is exact if a plausible symmetry assumption is made.

[29] Further evidence is provided in Online Appendix IV. There, we show that the bulk of the inequality-reducing effect from our imputation exercise stems from imputation in the lowest quartile.

[30] As reported by Fuchs-Schündeln et al. (2010), per-capita hours worked by male employees have been rather stable since 1984, while the correlation between hours and wages has slightly increased, from about -0.2 to approximately zero.

The way in which those factors may enter an explanation of the rise of intra-generational lifetime earnings inequality is a priori unclear and merits an in-depth investigation that is beyond the scope of this paper. For instance, it remains to be seen whether skill-biased technological change significantly increased cohort-specific inequality in spite of the expansion of education. As we document in Appendix D, West German baby boomers have benefitted from substantially more schooling than their parents' cohorts.

Cohort size is another dimension with respect to which the cohorts in our sample substantially differ, which makes it a natural candidate for explaining the rise of cohort-specific inequality. In Appendix B we show that cohort size in the year the cohort turns forty displays a non-monotonic pattern. It displays a local maximum for the 1940-cohort, a global minimum for the 1945-cohort, and a global maximum for the 1964-cohort. There is no one-to-one relationship between cohort size and cohort-specific inequality. The Gini coefficient of UA-40 strongly increases both during the years 1940-1945, when cohort size shrinks, and during the years 1945-1964, when cohort size grows. It does not change much neither during the years 1935-1940 (of growing cohort size) nor during the years 1964-1969 (of shrinking cohort size).

IV. Conclusion

We have documented, for the first time, the magnitude, pattern, and evolution of lifetime earnings inequality in Germany. Based on a large sample of earnings biographies from social security records, we have shown that the intra-generational distribution of lifetime earnings has a Gini coefficient that amounts to about two-thirds of the value of the Gini coefficient of annual earnings. Within cohorts, mobility in the distribution of yearly earnings is substantial at the beginning of the life cycle, decreases afterwards and virtually vanishes after age forty.

A comparison of earnings mobility across cohorts has not revealed noticeable differences. The pattern of mobility within a cohort's earning distribution is similar across all the cohorts we have scrutinized, from the one born in 1935 to the one born in 1969. Hence, changes in intra-generational mobility cannot be held responsible for the increase of cross-sectional earnings inequality in the German labor market.

The main novel finding from our investigation is the secular rise of intra-generational inequality in lifetime earnings: West-German men born in the early 1960s are likely to experience about 85 % more lifetime inequality than their fathers.

Our analysis has begun to shed some light on the proximate causes of the rise of intra-generational inequality in lifetime earnings. Longer unemployment spells, mainly affecting workers at the bottom of the distribution of younger cohorts, account for some 20 to 40 percent of the overall increase in lifetime earnings inequality. The remaining 60 to 80 percent is due to an increase in cohort-specific wage dispersion. While our decomposition of the rise of intra-generational inequality is just a first pass, we believe that the results of this paper convincingly demonstrate the benefits in terms of insights into the workings of the labor market that can be gained from following a cohort-based approach.

From the generation born immediately before World War II to the baby boomers of the 1960s, the German labor market has generated much more lifetime earnings inequality. The potential implications of this fact are far-reaching. By itself, such an increased heterogeneity in terms of labor-market outcomes might have a significant impact on cultural and political attitudes by weakening people's feeling of sharing a common fate. Through its effect on the distribution of lifetime consumption, the increase in lifetime earnings inequality might substantially affect the social welfare of generations. Examining those potential implications in detail is an important task of future research.

Appendix

A. Representativeness of VSKT as assessed through the SOEP

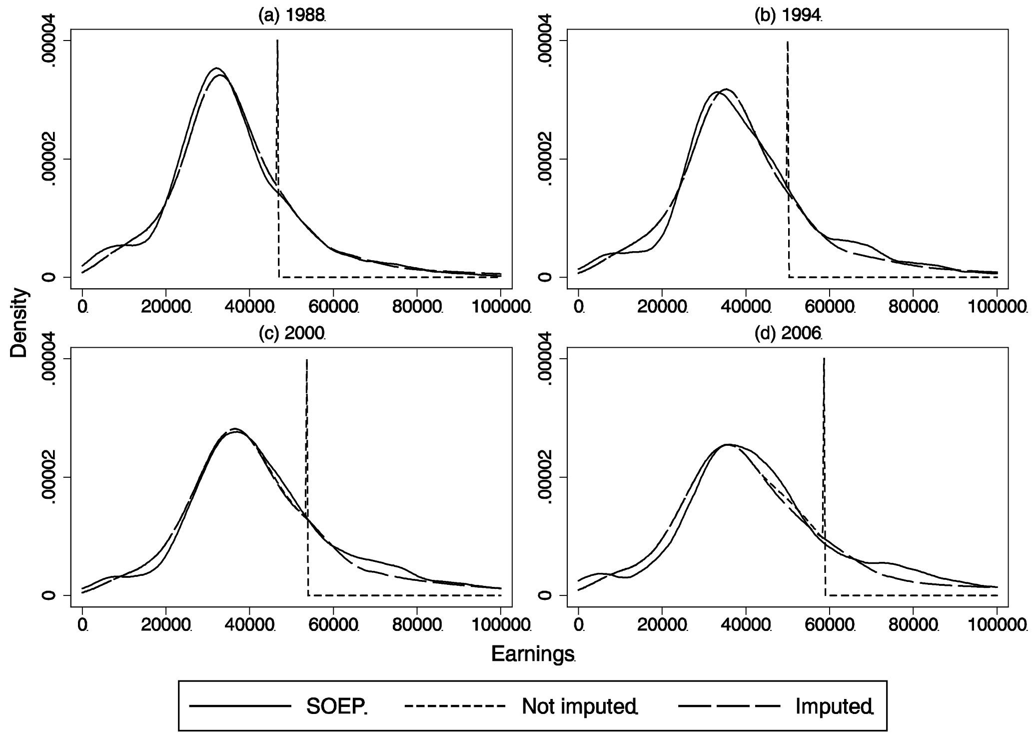


Fig. A1.− Comparison of Kernel density estimates for annual earnings distributions of men

Note.− "Not imputed" denotes estimates based on original VSKT data, "imputed" denotes estimates based on the VSKT after applying our imputation method; all earnings include employer's social security contributions. Population composition of the SOEP mirrors the one of the VSKT with respect to age, gender, region of residence, citizenship, and employment status; see Table A1 for further details.
Source.− FDZ-RV − VSKT2002, 2004-2009_Bönke, SOEP v28, own calculation using weighted data.

Table A1
Male labor force in West Germany for selected years, SOEP

Year	1988		1994		2000		2006	
Age range	20-53		25-59		31-59		37-59	
Labor force status	Observations	%	Observations	%	Observations	%	Observations	%
Employed[A]	10,078,221	70.07	11,343,612	70.35	9,871,416	72.87	7,733,104	71.4
Unemployed[A]	716,579	4.98	1,145,635	7.1	732,898	5.41	808,084	7.46
Apprentice[A]	462,953	3.22	76,840	0.48	22,960	0.17	3,727	0.03
Miner[A]	84,576	0.59	101,913	0.63	90,558	0.67	19,925	0.18
Com./Military service[A]	304,337	2.12	17,443	0.11	0	0	0	0
Sum of items above[A]	11,646,666	80.99	12,685,443	78.67	10,717,832	79.12	8,564,840	79.08
Civil servant	1,592,497	11.07	1,778,165	11.03	1,125,649	8.31	876,367	8.09
Self-employed	1,143,363	7.95	1,661,736	10.31	1,703,570	12.58	1,388,851	12.82
Total	14,382,526	100.00	16,125,344	100.00	13,547,051	100.01	10,830,058	100.00

Note.− Sample selection mirrors the respective birth cohorts in our deployed FDZ-RV − VSKT2002, 2004-2009_Bönke data.
[A]Labor force covered in the VSKT.
Source.− SOEP v28, own calculations using weighted data.

B. Number of Observations

Table B1
Number of observed men with valid UAX-biographies

Cohort	Up to 40	Up to 45	Up to 50	Up to 55	Up to 60
1935	1,114	1,091	1,073	1,022	1,000
1936	1,067	1,042	1,019	974	955
1937	1,081	1,079	1,061	1,021	981
1938	1,104	1,099	1,090	1,053	1,023
1939	1,207	1,165	1,140	1,081	1,049
1940	1,095	1,084	1,080	1,046	1,022
1941	1,121	1,118	1,116	1,084	1,070
1942	1,109	1,087	1,082	1,042	1,032
1943	1,107	1,101	1,084	1,048	1,025
1944	1,087	1,067	1,054	1,005	978
1945	1,154	1,143	1,140	1,113	1,090
1946	1,172	1,143	1,133	1,094	1,057
1947	1,175	1,154	1,137	1,089	1,051
1948	1,189	1,167	1,151	1,106	1,056
1949	1,163	1,132	1,110	1,062	1,016
1950	1,202	1,175	1,152	1,101	
1951	1,228	1,206	1,175	1,127	
1952	1,212	1,168	1,145	1,101	
1953	1,223	1,195	1,171	1,120	
1954	1,271	1,230	1,202	1,144	
1955	1,293	1,261	1,230		
1956	1,311	1,268	1,236		
1957	1,295	1,255	1,236		
1958	1,322	1,292	1,256		
1959	1,345	1,316	1,277		
1960	1,377	1,336			
1961	1,417	1,389			
1962	1,481	1,435			
1963	1,494	1,444			
1964	1,437	1,411			
1965	1,493				
1966	1,507				
1967	1,511				
1968	1,531				
1969	1,622				
Total	44,517	36,053	28,550	21,433	15,405

Note.– Number of observations for a cohort changes with age because of the selection criterion for valid biographies (see details in Online Appendix I.5).
Source: FDZ-RV − VSKT2002, 2004-2009_Bönke, own calculations using unweighted data.

Table B2
Weighted number of observations with valid UA-40 biographies, men

Birth cohort	Observations with valid UA40-biographies (weighted)	Actual cohort size at age 40			
		without foreigners		including foreigners	
		size	coverage[A]	size	coverage[A]
1935	214,783	431,149	0.4982	474,200	0.4529
1936	217,551	436,191	0.4988	481,363	0.4519
1937	207,309	438,432	0.4728	484,576	0.4278
1938	221,022	463,038	0.4773	512,694	0.4311
1939	245,519	489,539	0.5015	541,907	0.4531
1940	233,767	491,013	0.4761	548,271	0.4264
1941	216,453	454,854	0.4759	505,586	0.4281
1942	172,882	366,390	0.4719	419,750	0.4119
1943	175,621	374,491	0.4690	423,065	0.4151
1944	173,017	361,344	0.4788	405,798	0.4264
1945	126,931	263,183	0.4823	308,797	0.4110
1946	162,292	308,837	0.5255	358,143	0.4531
1947	178,106	348,759	0.5107	400,945	0.4442
1948	188,304	372,573	0.5054	425,099	0.4430
1949	201,483	398,952	0.5050	450,614	0.4471
1950	210,781	.	.	455,050	0.4632
1951	202,075	.	.	453,496	0.4456
1952	207,547	.	.	466,666	0.4447
1953	198,846	.	.	462,634	0.4298
1954	218,223	.	.	480,666	0.4540
1955	218,160	.	.	491,565	0.4438
1956	232,274	.	.	512,988	0.4528
1957	237,176	.	.	526,243	0.4507
1958	242,756	.	.	535,051	0.4537
1959	258,979	.	.	559,580	0.4628
1960	267,044	.	.	578,547	0.4616
1961	267,736	.	.	593,879	0.4508
1962	279,379	.	.	607,311	0.4600
1963	276,530	.	.	629,334	0.4394
1964	280,680	.	.	636,891	0.4407
1965	282,497	.	.	628,727	0.4493
1966	283,604	.	.	624,951	0.4538
1967	288,091	.	.	608,938	0.4731
1968	277,011	.	.	593,330	0.4669
1969	261,663	.	.	562,571	0.4651

Note.– Cohorts 1935 – 1949: West Germany includes West Berlin; cohorts 1950 – 1969: West Germany includes Berlin.
[A]Coverage equals the number of observations with valid UA-40 biographies (weighted) divided by actual cohort size at 40.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke. Actual cohort size at age 40 according to Federal Statistical Office, own calculations using weighted data.

C. Effects of Earnings Imputation in Case of Unemployment

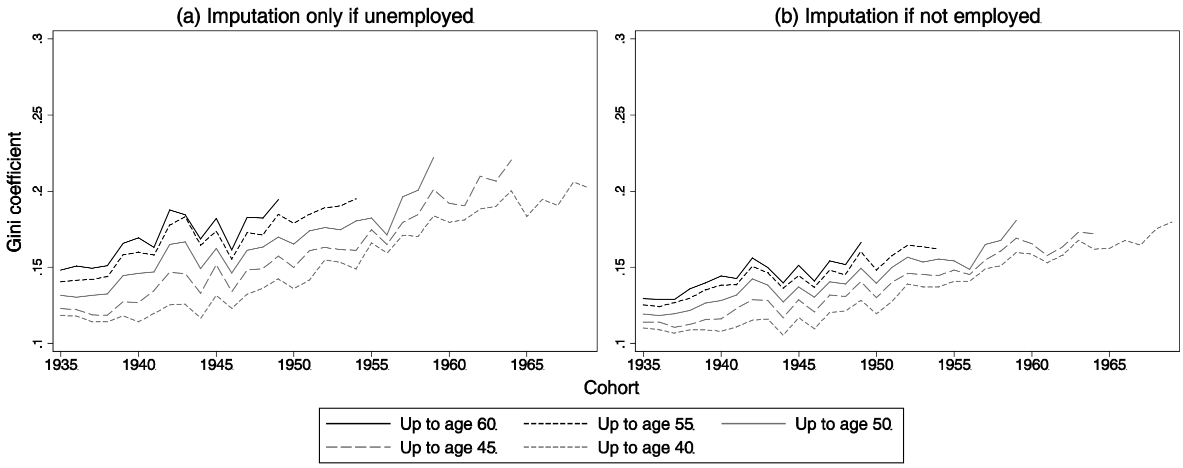


Fig. C1.– Gini coefficients of various UAX with earnings imputation if individual is not employed

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

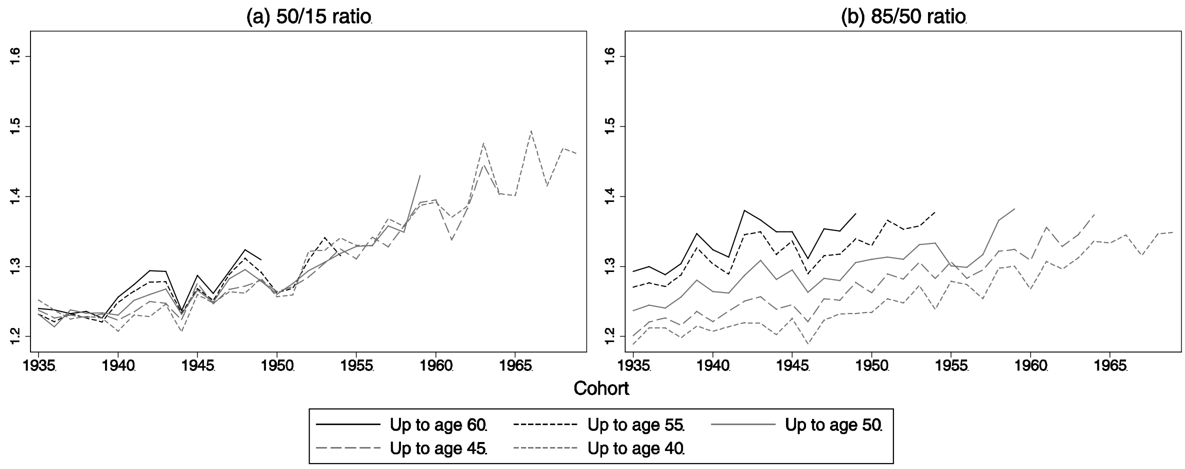


Fig. C2.– 50th / 15th and 85th / 50th ratio of UAX with imputation for registered unemployment

Note.– UAX based on federal bond discounting.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

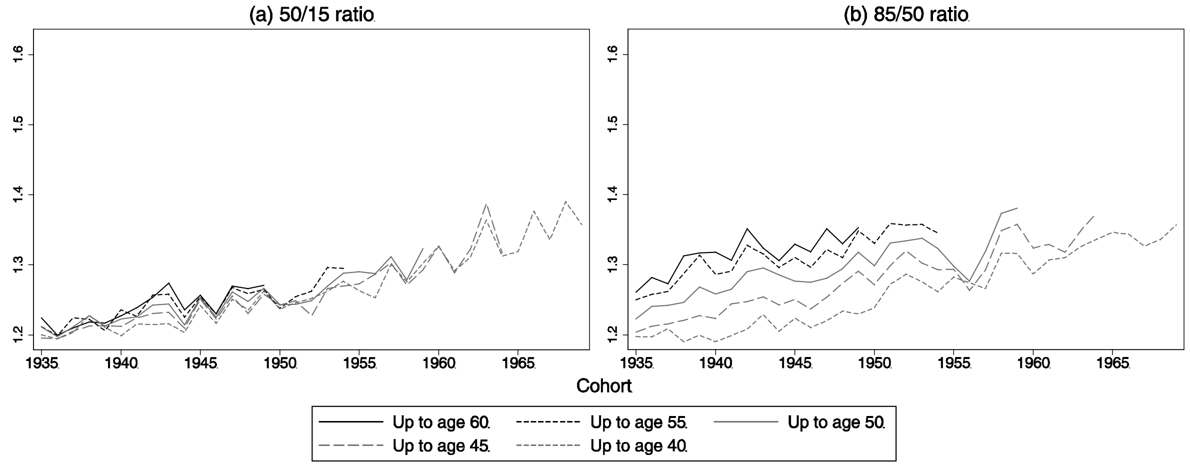


Fig. C3.– 50^{th} / 15^{th} and 85^{th} / 50^{th} ratio of UAX with imputation if not employed

Note.– UAX based on federal bond discounting.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

D. Educational Attainment and Inequality across Cohorts

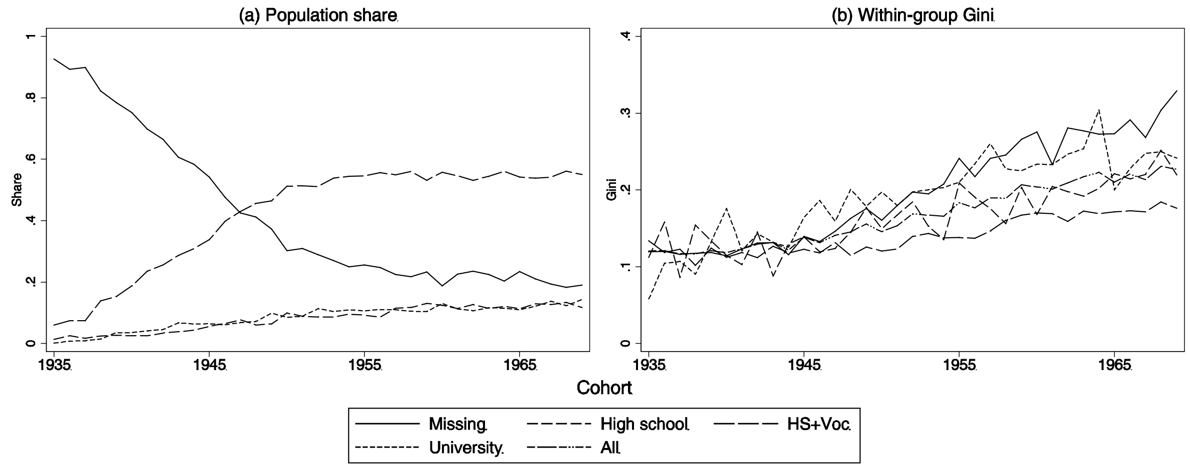


Fig. D1.– Educational attainment and inequality in our sample

Note.– Within-group Gini coefficients refer to the distributions of UA-40 with federal bond discounting.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

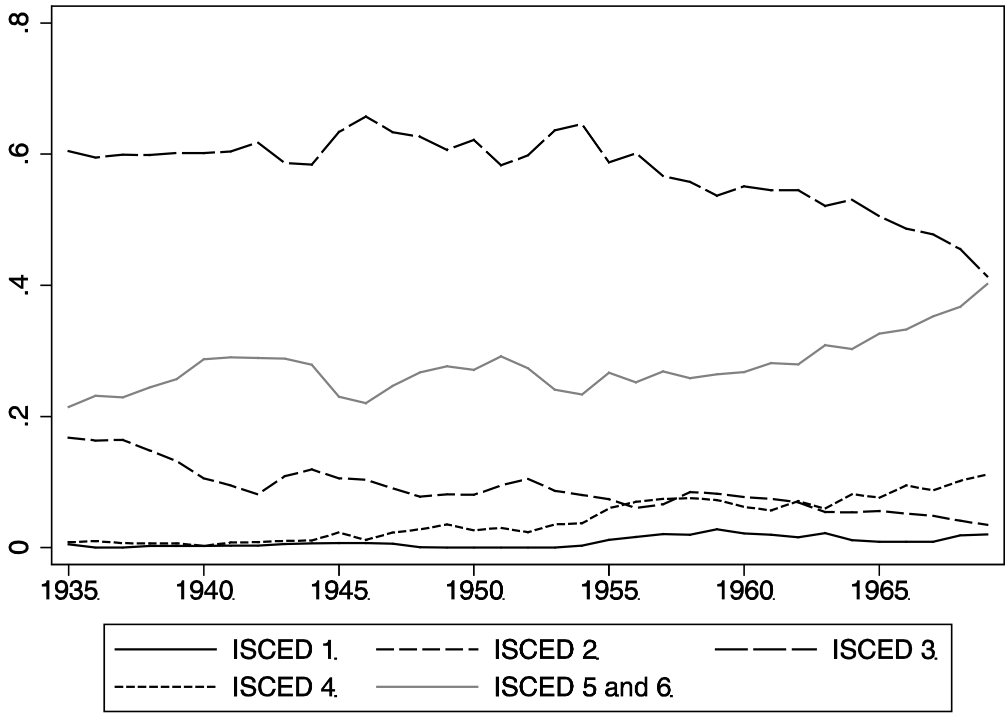


Fig. D.2.– Educational attainment of cohorts of West German men according to the SOEP

Note.– The education groups are defined according to the International Standard Classification of Education 1997 (ISCED-97): ISCED 1: Primary education; ISCED 2: Lower secondary education; ISCED 3: Upper secondary education; ISCED 4: Post-secondary non-tertiary education; ISCED 5/6: Tertiary education. All cohorts born after 1944 are analyzed at age 40. Since the SOEP starts in 1984, older cohorts are analyzed at the closest distance to age 40, e.g. age 45 for those born in 1939.
Source.– SOEP v28, own calculations using weighted data.

References

Almas, Ingvild, Tarjei Havnes, and Magne Mogstad. 2011. Baby booming inequality? Demographic change and earnings inequality in Norway, 1967-2000. *Journal of Economic Inequality* 9:629-650.

Atkinson, Anthony B., and Thomas Piketty. 2010. Top Incomes – A Global Perspective. Oxford: Oxford University Press.

Autor, David H., Lawrence F. Katz, and Melissa S. Kearny. 2006. The polarization of the U.S. labor market. *American Economic Review (Papers & Proceedings)* 96:189-194.

Bach, Stefan, Giacomo Corneo, and Viktor Steiner. 2009. From bottom to top: The entire income distribution in Germany, 1992-2003. *Review of Income and Wealth* 55:303- 330.

Bhuller, Manudeep, Magne Mogstad, and Kjell G. Salvanes. 2011. Life-cycle bias and the returns to schooling in current and lifetime earnings. Discussion Paper no. 5788, Institute for the Study of Labor, Bonn.

Björklund, Anders. 1993. A comparison between actual distributions of annual and lifetime income: Sweden 1951-89. *Review of Income and Wealth* 39:377-386.

Böhlmark, Anders, and Matthew Lindquist. 2006. Life-cycle variations in the association between current and lifetime income: Replication and extensions for Sweden. *Journal of Labor Economics* 24: 879-896.

Brenner, Jan. 2010. Life-cycle variations in the association between current and lifetime earnings: Evidence for German natives and guest workers. *Labour Economics* 17:392-406.

Burkhauser, Richard V., and John G. Poupore. 1997. A cross-national comparison of permanent inequality in the United States and Germany. *Review of Economics and Statistics* 79:10-17.

Card, David, and John E. DiNardo. 2002. Skill-biased technological change and rising wage inequality: Some problems and puzzles. *Journal of Labor Economics* 20:733-83.

Card, David, Jög Heining, and Patrick Kline. 2013. Workplace heterogeneity and the rise of West German wage inequality. *Quarterly Journal of Economics* 128:967-1015.

Deaton, Angus, and Christina Paxson. 1994. Intertemporal choice and inequality. *Journal of Political Economy* 102:437-467.

Dustmann, Christian, Johannes Ludsteck, and Uta Schönberg. 2009. Revisiting the German wage structure. *Quarterly Journal of Economics* 124:843-881.

Fitzenberger, Bernd. 1999. *Wages and employment across skill groups: An analysis for West Germany.* Heidelberg, New York: ZEW Economic Studies, Physika-Verlag.

Fuchs-Schündeln, Nicola, Dirk Krueger, and Mathias Sommer. 2010. Inequality trends for Germany in the last two decades: A tale of two countries. *Review of Economic Dynamics* 13:103-132.

Fuchs-Schündeln, Nicola, and Matthias Schündeln. 2009. Who stays, who goes, who returns?.

Economics of Transition 17: 703-738.

Gernandt, Johannes, and Friedhelm Pfeiffer. 2007. Rising wage inequality in Germany. *Journal of Economics and Statistics* 227:358-380.

Goos, Maarten, Alan Manning and Anna Salomons. 2009. Job polarization in Europe. *American Economic Review (Papers & Proceedings)* 99: 58-63.

Haider, Steven, and Gary Solon. 2006. Life-cycle variation in the association between current and lifetime earnings. *American Economic Review* 96: 1308-1320.

Himmelreicher, Ralf, and Michael Stegmann. 2008. New possibilities for socio-economic research through longitudinal data from the Research Data Centre of the German Federal Pension Insurance. *Journal of Applied Social Science Studies* 128: 647 - 660.

Huggett, Mark, Gustavo Ventura, and Amir Yaron. 2011. Sources of lifetime inequality. *American Economic Review* 101:2923-2954.

Hunt, Jennifer. 1995. The effect of unemployment compensation on unemployment duration in Germany. *Journal of Labor Economics* 13: 88-120.

Jenkins, Stephen P. 1987. Snapshots versus movies: "Lifecycle biases" and the estimation of intergenerational earnings inheritance. *European Economic Review* 31:1149-1158.

Jenkins, Stephen P., and Philippe Van Kerm. 2006. Trends in income inequality, pro-poor income growth and income mobility. *Oxford Economic Papers* 58:531-548.

Kopczuk, Wojciech, Emmanuel Saez, and Jae Song. 2010. Earnings inequality and mobility in the United States: Evidence from social security data since 1937. *Quarterly Journal of Economics* 125:91-128.

Lemieux, Thomas. 2008. The changing nature of wage inequality. *Journal of Population Economics* 21:21-48.

Lerman, Robert I., and Shlomo Yitzhaki. 1985. Income inequality effects by income Source: A new approach and applications to the United States. *Review of Economics and Statistics* 67:151-156.

Maasoumi, Esfandiar, and Mark M. Trede. 2001. Comparing income mobility in Germany and the United States using generalized entropy mobility measures. *Review of Economics and Statistics* 83:551-559.

Nickell, Stephen, Luca Nunziata and Wolfgang Ochel. 2005. Unemployment in the OECD since the 1960s. What do we know? *Economic Journal* 115:1-27.

Organization for Economic Cooperation and Development. 2008. *Growing Unequal? Income Distribution and Poverty in OECD Countries.* Paris: Organization for Economic Cooperation and Development.

Pérez-Asenjo, Eduardo. 2011. If happiness is relative, against whom do we compare ourselves?

Implications for labour supply. *Journal of Population Economics* 24:1411-1442.

Reinberg, Alexander, and Markus Hummel. 2007. Der Trend bleibt – Geringqualifizierte sind häufiger arbeitslos. *IAB Kurzbericht* no. 18.

Trede, Mark M. 1998. The age profile of mobility measures: An application to earnings in West Germany. *Journal of Applied Econometrics* 13:397-409.

Online Appendix to Chapter 1: “Lifetime Earnings Inequality in Germany”

Appendix I: Data

I.1 Earnings concept

The earnings information provided in the Insurance Account Sample is based on the employee's gross wage. In order to obtain the market value of earnings w_G, the social security contributions paid directly by the employer have to be added to the gross wage w according to equation (I.1):

$$w_G = w + \min(z_p, w)\, r_p + \min(z_h, w)\,(r_h + r_l) + \min(z_u, w)\, r_u \qquad \text{(I.1)}$$

In (I.1), z_s denotes the contribution ceiling and r_s the employer's contribution rate in the various branches $s = p(ension), h(ealth), u(nemployment), l(ongterm\ care)$ of the German social security scheme. The respective contribution ceiling and rate are provided in Table I.1.

Some categories of employees like miners, sailors and distinctive employees of the federal railways have special social security arrangements (*knappschaftlich Versicherte*). For these employees the contribution ceiling and contribution rate of the pension insurance, z_p and r_p, differ. Both are higher than in the regular scheme (see Table I.2) and mirror the historically higher risks of employees in these industries. Due to the higher health risks, this pension scheme includes an additional pension component used for earlier retirement entries, additional invalidity care and higher pensions (§§ 40, 45, 85, 238, 239 and 242; Social Code VI [*Sozialgesetzbuch VI*]). This pension scheme is especially relevant for male employees of older cohorts. For example an average of 10% to 15% born between 1935 and 1940 are subject to these special social security arrangements where it is negligible for the younger cohorts.

In order to provide the most accurate picture of the comparison between younger and older employees, we want to include employees with have special social security arrangements (*knappschaftlich Versicherte*), as their share is non-negligible in older cohorts. The differences in the contributions ceiling could be accounted for with by imputations methods. However, the differences in the contribution rates range from 6% to 10%, depending on the year (see Table I.1 below). This could potentially bias the analysis of earnings inequality and to account for that matter, we turn to the concept of market value of labor. As our robustness-section III shows, our results are nonetheless robust regardless the treatment of employer's social security contributions.

Table I.1
Key parameters of German social security (default case for regular insured)

Year	Average earnings[A]	Pension insurance ceiling	Pension insurance rate	Health insurance ceiling	Health insurance rate[B]	Unemployment insurance ceiling	Unemployment insurance rate	Long-term care insurance rate[C]
1952	3,852	7,800	5.000	6,000	3.000	6,000	2.000	
1953	4,061	9,000	5.000	6,000	3.000	6,000	2.000	
1954	4,234	9,000	5.000	6,000	3.100	6,000	2.000	
1955	4,548	9,000	5.375	6,000	3.100	6,000	1.630	
1956	4,844	9,000	5.500	6,000	3.100	6,000	1.500	
1957	5,043	9,000	6.750	6,480	3.900	9,000	1.083	
1958	5,330	9,000	7.000	7,920	4.200	9,150	1.000	
1959	5,602	9,600	7.000	7,920	4.200	9,150	1.000	
1960	6,101	10,200	7.000	7,920	4.200	9,150	1.000	
1961	6,723	10,800	7.000	7,920	4.700	9,150	1.000	
1962	7,328	11,400	7.000	7,920	4.800	9,150	0.775	
1963	7,775	12,000	7.000	7,920	4.800	9,150	0.700	
1964	8,467	13,200	7.000	7,920	4.850	9,150	0.650	
1965	9,229	14,400	7.000	8,880	4.950	9,150	0.650	
1966	9,893	15,600	7.000	10,800	5.000	9,150	0.650	
1967	10,219	16,800	7.000	10,800	5.050	10,650	0.650	
1968	10,842	19,200	7.500	10,800	5.100	15,600	0.650	
1969	11,839	20,400	8.000	11,250	5.250	18,000	0.650	
1970	13,343	21,600	8.500	14,400	4.100	21,600	0.650	
1971	14,931	22,800	8.500	17,100	4.100	22,800	0.650	
1972	16,335	25,200	8.500	18,900	4.200	25,200	0.850	
1973	18,295	27,600	9.000	20,700	4.600	27,600	0.850	
1974	20,381	30,000	9.000	22,500	4.700	30,000	0.850	
1975	21,808	33,600	9.000	25,200	5.200	33,600	1.000	
1976	23,335	37,200	9.000	27,900	5.600	37,200	1.500	
1977	24,945	40,800	9.000	30,600	5.700	40,800	1.500	
1978	26,242	44,400	9.000	33,300	5.700	44,400	1.500	
1979	27,685	48,000	9.000	36,000	5.600	48,000	1.500	
1980	29,485	50,400	9.000	37,800	5.700	50,400	1.500	
1981	30,900	52,800	9.250	39,600	5.900	52,800	1.500	
1982	32,198	56,400	9.000	42,300	6.000	56,400	2.000	
1983	33,293	60,000	9.083	45,000	5.900	60,000	2.300	
1983	34,292	62,400	9.250	46,800	5.700	62,400	2.300	
1985	35,286	64,800	9.454	48,600	5.900	64,800	2.150	
1986	36,627	67,200	9.600	50,400	6.100	67,200	2.000	
1987	37,726	68,400	9.350	51,300	6.300	68,400	2.150	
1988	38,896	72,000	9.350	54,000	6.500	72,000	2.150	
1989	40,063	73,200	9.350	54,900	6.500	73,200	2.150	
1990	41,946	75,600	9.350	56,700	6.300	75,600	2.150	
1991	44,421	78,000	8.980	58,500	6.100	78,000	3.090	
1992	46,820	81,600	8.850	61,200	6.400	81,600	3.150	
1993	48,178	86,400	8.750	64,800	6.700	86,400	3.250	
1994	49,142	91,200	9.600	68,400	6.600	91,200	3.250	
1995	50,665	93,600	9.300	70,200	6.600	93,600	3.250	0.500
1996	51,678	96,000	9.600	72,000	6.700	96,000	3.250	0.850
1997	52,143	98,400	10.150	73,800	6.800	98,400	3.250	0.850
1998	52,925	100,800	10.150	75,600	6.800	100,800	3.250	0.850
1999	53,507	102,000	9.850	76,500	6.800	102,000	3.250	0.850
2000	54,256	103,200	9.650	77,400	6.800	103,200	3.250	0.850
2001	55,216	104,400	9.550	78,300	6.800	104,400	3.250	0.850
2002	28,626	54,000	9.550	40,500	7.000	54,000	3.250	0.850
2003	28,938	61,200	9.750	41,400	7.200	61,200	3.250	0.850
2004	29,060	61,800	9.750	41,856	7.200	61,800	3.250	0.850
2005	29,202	62,400	9.750	42,300	7.100	62,400	3.250	0.850
2006	29,494	63,000	9.750	42,756	6.500	63,000	3.250	0.850
2007	29,951	63,000	9.950	42,756	6.800	63,000	2.100	0.850
2008	30,625	63,600	9.950	43,200	6.900	63,600	1.650	0.850
2009	30,879	64,800	9.950	44,100	7.000	64,800	1.400	0.975

Note.– Average earnings and contribution ceilings denoted in current prices and currency (1952 - 2001 in DM, 2002 - 2009 in Euro), reported rates are employer's contribution rates. [A]Subject to social security contributions. [B]Average contribution

rate. Employees with high earnings who are eligible to opt for private health insurance (*Versicherungsfreigrenze*) are considered to remain in the public health insurance. [c] The contribution ceilings of the long-term care and the health insurance coincide.
Source.– Appendices 1 and 2 of Social Code VI (*Sozialgesetzbuch VI*), Federal Ministry of Labour and Social Affairs.

Table I.2 Regulations for the old age pension schemes of miners

Year	Average earnings[A]	Pension insurance ceiling	rate
1952	3,893	12,000	15.500
1953	4,104	12,000	15.500
1954	4,279	12,000	15.500
1955	4,596	12,000	15.500
1956	4,895	12,000	15.500
1957	5,096	12,000	15.200
1958	5,386	12,000	15.000
1959	5,661	12,000	15.000
1960	6,165	12,000	15.000
1961	6,794	13,200	15.000
1962	7,405	13,200	15.000
1963	7,857	14,400	15.000
1964	8,556	16,800	15.000
1965	9,326	18,000	15.000
1966	9,997	19,200	15.000
1967	10,327	20,400	15.000
1968	10,957	22,800	15.000
1969	11,965	24,000	15.000
1970	13,485	25,200	15.000
1971	15,090	27,600	15.000
1972	16,508	30,000	15.000
1973	18,489	33,600	15.000
1974	20,597	37,200	15.000
1975	22,039	40,800	15.000
1976	23,582	45,600	15.000
1977	25,209	50,400	15.000
1978	26,520	55,200	15.000
1979	27,979	57,600	15.000
1980	29,798	61,200	15.000
1981	31,228	64,800	15.000
1982	32,540	69,600	14.750
1983	33,646	73,200	15.170
1983	34,655	76,800	16.000
1985	35,660	80,400	15.300
1986	37,015	82,800	15.350
1987	38,125	85,200	15.100
1988	39,307	87,600	15.100
1989	40,486	90,000	15.100
1990	41,946	93,600	15.100
1991	44,421	96,000	14.645
1992	46,820	100,800	14.600
1993	48,178	106,800	14.500
1994	49,142	112,800	15.900
1995	50,665	115,200	15.400
1996	51,678	117,600	15.900
1997	52,143	121,200	16.750
1998	52,925	123,600	16.750
1999	53,507	124,800	16.380
2000	54,256	127,200	15.950
2001	55,216	128,400	15.850
2002	28,626	66,600	15.850
2003	28,938	75,000	16.150
2004	29,060	76,200	16.150
2005	29,202	76,800	16.150
2006	29,494	77,400	16.150
2007	29,951	77,400	16.450
2008	30,625	78,600	16.450
2009	30,506	79,800	16.450

Note.– Average earnings and contribution ceilings denoted in current prices and currency (1952 - 2001 in DM, 2002 - 2009 in Euro), reported rates are employer's contribution rates. [A]Miners subject to social security contributions. Source.– Appendices 1 and 2 of Social Code VI (*Sozialgesetzbuch VI*)

I.2 Imputation of top-coded earnings

The imputation of incomes for top-coded observations assumes that top incomes are distributed according to the Pareto law. Several studies investigating income distributions in various countries indicate that this is a good assumption.

Assume that individual earnings w_i exceeding $\widetilde{w}$ are Pareto-distributed. Then, the probability to observe an income greater or equal to $w_i > \widetilde{w}$ is given by

$$1 - F(w_i) = \left(\frac{w_i}{\widetilde{w}}\right)^{-\alpha} \tag{I.2}$$

where $F(w_i)$ denotes the cumulative probability density function. Consider n to be the number of earners with $w_i > \widetilde{w}$ and $i = 1, \dots, n$. Furthermore, earners i are ranked in ascending order according to their income. From equation (I.2) each individual's rank r_i in the income distribution is determined as

$$r_i = nF(w_i) = n\left(1 - \left(\frac{w_i}{\widetilde{w}}\right)^{-\alpha}\right) \tag{I.3}$$

In top-coded data, individual earnings are available up to a contribution ceiling, z. If an individual earns more, reported earning is $w_i = z$. Consider m out of the n earners to receive an income above the contribution ceiling $z > \widetilde{w}$. Since for m earners neither r_i nor w_i is observable, we estimate the parameters of the Pareto-distribution by exploiting earnings data from the interval $[\widetilde{w}, z]$. Rearranging equation (I.3) yields

$$\ln\left(1 - \frac{r_i}{n}\right) = -\alpha \ln\left(\frac{w_i}{\widetilde{w}}\right) \tag{I.4}$$

We employ equation (I.4) to estimate the Pareto-coefficient α. Suppose at least the top 10% of individual earnings w_i in the interval $[0, z)$ to be Pareto-distributed. Accordingly, $\widetilde{w}$ is assigned the value of the 90th percentile in the respective distribution of earnings below z. The Pareto-coefficient is estimated by means of an OLS regression without constant. The regression is conducted separately for all years t and birth cohorts c. Hence, the cohort and year specific Pareto-coefficient $\hat{\alpha}_{c,t}$ is derived for $c = 1935, \dots, 1969$ and $t = 1952, \dots, 2009$ distributions. With the estimated Pareto-coefficient at hand, unobserved earnings above the contribution ceiling z can be estimated by rearranging (I.3):

$$\widehat{w}_i = \widetilde{w}\left(1 - \frac{\hat{r}_i}{n}\right)^{-\frac{1}{\hat{\alpha}}} \tag{I.5}$$

where $\boldsymbol{\widehat{w}_i}$ denotes the estimated earned income and $\boldsymbol{\hat{r}_i}$ the assumed rank. The conjectures regarding $\boldsymbol{\hat{r}_i}$ have an immediate effect on measures of income mobility and, therefore, are crucial when investigating earnings dynamics. In our preferred imputation, we choose $\boldsymbol{\hat{r}_i}$ under the minimal mobility assumption. Thereby, the rank $\boldsymbol{\hat{r}_i}$ is based on the last observable rank in relation to all individuals at or above the contribution ceiling in the cohort-specific earnings distribution.[31] This imputation procedure leads to plausible annual earnings distributions. Comparing the obtained annual earnings distributions to (almost) uncapped survey-based micro data reveals a good fit (see Figure A1).

[31] For illustration consider two earnings distributions in subsequent periods $t-1$ and t made out of three individuals a, b and c. Suppose the following ordering of earnings in $t-1$: $w_{a,t-1} < w_{b,t-1} < z_{t-1} < w_{c,t-1}$ and resulting ranks $r_{a,t-1} = 1$, $r_{b,t-1} = 2$ and the estimated rank $\hat{r}_{c,t-1} = 3$ since c's earnings exceed z_{t-1}. In t individual a has earnings above the contribution ceiling such that $w_{b,t} < z_t$ and $w_{a,t}, w_{c,t} > z_t$ where it is not observable whether a or c earns more. Then, the ranking order in t is $r_{b,t} = 1, \hat{r}_{a,t} = 2$ and $\hat{r}_{c,t} = 3$ because of $r_{c,t-1} > r_{a,t-1}$. Thus, the relative ordering of a and c remains unchanged for future years unless either a's or c's earnings fall below the contribution ceiling. To establish whether mobility results are robust, two alternative mobility scenarios are calculated: an equal ranking with imputation of estimated average earnings above the contribution ceiling and a maximum mobility scenario. In the maximum mobility scenario, the ranking order is reversed between years t and $t+1$. All alternative results and a scenario without imputation are provided in Online Appendix III.3.

Table I.3
Cohort-specific means of shares of censored spells

Cohort	Men	Women
1935	8.00	0.37
1936	7.47	0.37
1937	7.95	0.30
1938	7.98	0.38
1939	9.35	0.47
1940	9.09	0.62
1941	9.31	0.42
1942	10.63	0.54
1943	10.45	0.79
1944	9.68	0.65
1945	9.03	0.60
1946	7.84	0.48
1947	7.98	0.75
1948	7.58	0.55
1949	7.88	0.65
1950	7.64	0.45
1951	7.94	0.48
1952	7.53	0.64
1953	6.97	0.49
1954	7.59	0.63
1955	6.42	0.31
1956	5.96	0.56
1957	6.60	0.69
1958	6.47	0.41
1959	6.41	0.58
1960	6.14	0.39
1961	5.47	0.60
1962	5.43	0.65
1963	5.11	0.58
1964	5.37	0.46
1965	4.88	0.88
1966	4.37	0.49
1967	3.99	0.60
1968	3.77	0.65
1969	3.24	0.49
Mean	7.07	0.54

Note.– Means are calculated as share of censored spells on all annual spells in all years for each cohort. Differences in shares of censored spells across cohorts are due to changes in the contribution ceiling (see Table I.1).
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke.

I.3 Correction of structural break 1983/1984

Starting with 1984, one-time payments (e.g. holiday and Christmas allowances or bonuses) are subject to social security contributions and included in the basis of assessment and hence the earnings measure. This leads potentially to an artificial increase in inequality in the annual earnings distributions after 1983. Facing the same problem for comparable but cross sectional data, Fitzenberger (1999) suggests fixing this structural break by estimating quantile specific deviations from the median growth rate between 1983 and 1984.[32] A similar strategy is adopted by Dustman et al. (2009) and Card et al. (2013). In order to meet our data requirements we adjust Fitzenberger's (1999) strategy to panel data.

The imputation of one-time payments for observations before 1984 is accomplished as follows. First, we generate a variable containing the average individual rank in the cohort specific earnings distribution between age 35 and 40. This variable serves as an approximation for the individual's permanent position in the earnings distribution and reflects the finding by Fitzenberger (1999) that spurious growths due to one-time payments is more pronounced for higher earnings. The earnings position is coded as dummy d_q for $q = 1, \dots, 20$ quantiles and $d_q = 1$ if the individuals average rank in the annual earnings distributions between 35 and 40 falls into the respective quantile. Furthermore, we define earnings growth between t and $t + 1$ as $\Delta w_{i,t} = ln(w_{i,t}) - ln(w_{i,t-1})$. Earnings growth is estimated with a generalized least squares random effects regression in an unbalanced panel restricted to prime age individuals from 26 to 59 according to equation I.6:

$$\begin{aligned}\Delta w_t = \alpha_0 + \ \alpha_1 d_{1984} + \alpha_2 age_t + \alpha_3 age_t^2 \ + \alpha_4 age_t^3 \\ + \alpha_5 d_{1984} age_t + \alpha_5 d_{1984} age_t^2 + \alpha_5 d_{1984} age_t^3 + \boldsymbol{\beta} \mathbf{d}'_q \ + \boldsymbol{\gamma} d_{1984} \mathbf{d}'_q \\ + \ \boldsymbol{\delta} age_t \mathbf{d}'_q + \ \varepsilon\end{aligned} \tag{I.6}$$

To identify spurious growth between 1983 and 1984, we include a dummy variable marking the structural break with $d_{1984} = 1$ in 1984 and zero else. Furthermore, we model age-earnings profiles by including age as a third order polynomial function and the vector of average earnings rank dummies $\mathbf{d}_q$ as well as interactions for all rank dummies with age respectively the structural break dummy d_{1984}.

Regression results confirm the cross sectional pattern reported in Fitzenberger (1999) with higher spurious growth rates for above median annual earnings. Figure I.1 displays the spurious growths pattern for selected quantiles. Depending on their positions in the cohort specific permanent earnings distribution, individual earnings are corrected by the quantile and age specific excessive growth factor for years predating 1984. Due to top coding, we assume the 17th quantile's for earnings in the 18th, 19th and 20th quantile (see Fitzenberger 1999).

[32] Fitzenbergers (1999) study is based on the *IAB Beschäftigtenstichprobe*, also obtained from social security administration data.

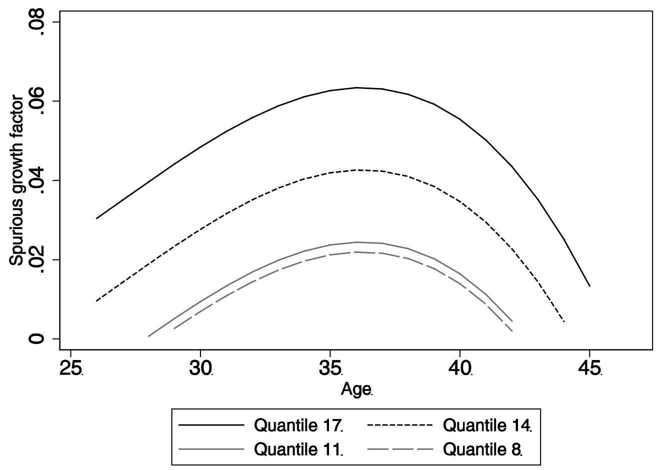


Fig. I.1.– Spurious growth rates for selected quantiles, men

Note.– The estimation is based on prime age males. Displayed is the relevant age range only. Spurious growth is identified as excessive growth between 1983 and 1984.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke.

I.4 Sample selection: General

To ensure that our lifetime earnings capture all relevant labor market activities, we restrict our sample to individuals with "valid biographies". To construct valid biographies, we measure the time where we cannot eliminate the possibility that an individual earned income besides what is recorded in our data. In order to achieve this, we exploit the fact that our data provides information on times apart from regular employment. Depending on the information on labor market activities we distinguish three cases:

(1) The individual is a regular employee subject to social security contributions (times of regular employment). Here we observe all relevant labor market activities and do not alter the earnings recorded in our data.

(2) The individual is not an employee subject to social security contributions but accumulates times relevant for old age pension. In this sense, the information provided in our data excludes the possibility of income earned on the regular labor market (e.g. times of educational training, care, sickness, unemployment, community or military service, disability or retirement). In this case we treat the respective monthly earnings as zeros.

(3) The individual is not an employee subject to social security contributions and we cannot exclude the possibility of income earned apart from what is recorded in our data. This is where times of self-employment, working as civil servant or labor market withdrawal cannot be distinguished and we do not have the necessary information to conclude that we observe all relevant labor market activities. Therefore we recode these monthly earnings into times of missing information.

In order to select individuals with complete occupational biographies, we now exclude all individuals who display more than one month of missing information per year after the age of 30. Hence, for an individual to be included in the analysis, the up-to-age 40 concept allows for up to 10 month of missing information before an individual is excluded, the up-to-age 41 concepts for up to 11 month and so forth.

I.5 Sample selection: Men

Table I.4
Number of observed men

Cohort	Original number of observations	Sample size after exclusion of: Foreigners	East Germans	Repatriates (*Aussiedler*)	Self-employed craftsmen	Rejoiners (marriage-law)[A]	Incomplete biographies (up to age 40)
1935	3,236	2,195	1,746	1,661	1,606	1,606	1,114
1936	3,214	2,210	1,746	1,671	1,623	1,623	1,067
1937	3,201	2,253	1,751	1,680	1,626	1,626	1,081
1938	3,198	2,269	1,768	1,683	1,653	1,653	1,104
1939	3,228	2,295	1,791	1,733	1,678	1,678	1,207
1940	3,230	2,259	1,735	1,666	1,623	1,623	1,095
1941	3,328	2,330	1,802	1,716	1,677	1,677	1,121
1942	3,289	2,335	1,799	1,723	1,686	1,686	1,109
1943	3,320	2,382	1,772	1,697	1,653	1,653	1,107
1944	3,324	2,376	1,724	1,673	1,623	1,623	1,087
1945	3,334	2,406	1,818	1,761	1,725	1,725	1,154
1946	3,280	2,311	1,852	1,803	1,767	1,767	1,172
1947	3,380	2,427	1,837	1,771	1,741	1,741	1,175
1948	3,472	2,461	1,913	1,840	1,806	1,806	1,189
1949	3,514	2,517	1,903	1,816	1,775	1,775	1,163
1950	3,706	2,624	1,934	1,839	1,802	1,802	1,202
1951	3,988	2,787	2,034	1,913	1,869	1,869	1,228
1952	4,087	2,806	1,976	1,850	1,813	1,813	1,212
1953	4,122	2,832	2,100	1,958	1,927	1,927	1,223
1954	4,215	2,789	2,103	1,979	1,949	1,949	1,271
1955	4,497	2,911	2,135	2,029	2,003	2,003	1,293
1956	4,505	2,877	2,178	2,056	2,029	2,029	1,311
1957	4,806	2,914	2,210	2,094	2,052	2,052	1,295
1958	5,130	2,948	2,250	2,133	2,094	2,094	1,322
1959	5,510	3,027	2,305	2,194	2,156	2,156	1,345
1960	6,174	3,117	2,339	2,242	2,211	2,211	1,377
1961	7,013	3,259	2,465	2,360	2,318	2,318	1,417
1962	7,338	3,380	2,527	2,423	2,394	2,394	1,481
1963	7,488	3,436	2,590	2,506	2,464	2,464	1,494
1964	7,595	3,325	2,522	2,463	2,428	2,428	1,437
1965	7,646	3,305	2,555	2,496	2,460	2,460	1,493
1966	7,750	3,347	2,606	2,558	2,516	2,516	1,507
1967	7,699	3,307	2,567	2,519	2,476	2,476	1,511
1968	7,830	3,254	2,528	2,488	2,453	2,453	1,531
1969	8,044	3,355	2,641	2,608	2,577	2,577	1,622
Total	168,691	96,626	73,522	70,602	69,253	69,253	44,517

Note.− Second column: Original sample size. Columns three to eight: All numbers denote observation after the stepwise exclusion of the respective groups. [A]Insured who left and rejoined the statutory pension system due to the law of marriage refunds (valid until 1967), which was possible until 1995. The earnings reported of these insured do not correspond to those actually earned.
Source.− FDZ-RV − VSKT2002, 2004-2009_Bönke, own calculations using unweighted data.

Table I.5
Weighted shares of groups from the initial dataset, men

Cohort	Foreigners	East Germans	Repatriates (*Aussiedler*)	Self-employed craftsmen	Rejoiners (marriage-law) [A]	"Incomplete" Up to 40	"Complete" Up to 40
1935	0.1540	0.1761	0.0317	0.0235	-	0.1959	0.4187
1936	0.1593	0.1703	0.0281	0.0197	-	0.2166	0.4061
1937	0.1696	0.2016	0.0259	0.0211	-	0.1997	0.3822
1938	0.1704	0.1994	0.0255	0.0124	-	0.2083	0.3839
1939	0.1697	0.2090	0.0178	0.0212	-	0.1792	0.4031
1940	0.1786	0.2093	0.0220	0.0168	-	0.1976	0.3758
1941	0.1719	0.2138	0.0266	0.0144	-	0.1989	0.3743
1942	0.2090	0.2013	0.0222	0.0128	-	0.2006	0.3541
1943	0.1967	0.2169	0.0228	0.0148	-	0.1975	0.3514
1944	0.1964	0.2218	0.0160	0.0172	-	0.1940	0.3547
1945	0.2566	0.1899	0.0160	0.0125	-	0.1927	0.3323
1946	0.2492	0.1527	0.0140	0.0130	-	0.2021	0.3689
1947	0.2357	0.1787	0.0199	0.0107	-	0.1981	0.3570
1948	0.2233	0.1775	0.0230	0.0109	-	0.2095	0.3557
1949	0.2000	0.1982	0.0280	0.0139	-	0.2050	0.3549
1950	0.1920	0.2086	0.0282	0.0130	-	0.1970	0.3611
1951	0.1752	0.2255	0.0350	0.0143	-	0.2015	0.3486
1952	0.1745	0.2273	0.0375	0.0128	-	0.1991	0.3488
1953	0.1697	0.2220	0.0395	0.0102	-	0.2244	0.3342
1954	0.1724	0.2184	0.0348	0.0099	-	0.2121	0.3525
1955	0.1738	0.2305	0.0310	0.0085	-	0.2120	0.3443
1956	0.1743	0.2098	0.0367	0.0087	-	0.2160	0.3545
1957	0.1648	0.2067	0.0364	0.0134	-	0.2258	0.3529
1958	0.1625	0.2112	0.0353	0.0131	-	0.2237	0.3542
1959	0.1545	0.2202	0.0326	0.0133	-	0.2209	0.3585
1960	0.1606	0.2186	0.0279	0.0103	-	0.2230	0.3595
1961	0.1467	0.2337	0.0278	0.0133	-	0.2271	0.3515
1962	0.1529	0.2221	0.0275	0.0095	-	0.2282	0.3598
1963	0.1562	0.2302	0.0206	0.0120	-	0.2364	0.3445
1964	0.1607	0.2230	0.0160	0.0103	-	0.2433	0.3467
1965	0.1658	0.2151	0.0146	0.0115	-	0.2385	0.3545
1966	0.1649	0.2051	0.0124	0.0131	-	0.2458	0.3587
1967	0.1657	0.2031	0.0130	0.0131	-	0.2304	0.3747
1968	0.1788	0.1990	0.0105	0.0111	-	0.2322	0.3685
1969	0.1890	0.1984	0.0090	0.0096	-	0.2295	0.3644

Note.− The numbers denote observation share of the respective groups before excluding any observations. [A]Insured who left and rejoined the statutory pension system due to the law of marriage refunds (valid until 1967), which was possible until 1995. The earnings reported of these insured do not correspond to those actually earned.
Source.− FDZ-RV − VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

Table I.6
Weighted number of observations with valid UAX-biographies, men

Birth cohort	Up to 40	Up to 45	Up to 50	Up to 55	Up to 60
1935	214,783	210,073	206,947	197,408	193,415
1936	217,551	212,263	207,186	196,663	191,881
1937	207,309	206,856	203,374	195,114	186,527
1938	221,022	218,897	217,169	211,985	209,564
1939	245,519	236,111	231,068	223,601	219,909
1940	233,767	230,358	228,571	227,105	224,172
1941	216,453	214,801	213,377	210,465	209,591
1942	172,882	169,064	168,109	164,225	164,017
1943	175,621	174,271	171,203	168,750	166,712
1944	173,017	168,663	166,300	161,142	159,641
1945	126,931	125,355	124,422	123,337	122,304
1946	162,292	157,618	155,222	152,719	149,049
1947	178,106	174,483	171,523	167,621	164,812
1948	188,304	183,558	180,554	177,935	173,946
1949	201,483	194,494	189,937	186,256	182,587
1950	210,781	205,003	200,438	197,303	
1951	202,075	198,195	192,300	188,904	
1952	207,547	198,705	194,186	191,816	
1953	198,846	193,264	188,495	185,667	
1954	218,223	210,309	204,987	199,045	
1955	218,160	212,973	207,454	205,115	
1956	232,274	223,581	217,471		
1957	237,176	229,484	225,704		
1958	242,756	236,871	228,939		
1959	258,979	252,939	245,655		
1960	267,044	258,361			
1961	267,736	261,133			
1962	279,379	270,243			
1963	276,530	267,379			
1964	280,680	275,448			
1965	282,497				
1966	283,604				
1967	288,091				
1968	277,011				
1969	261,663				
Total	7,926,092	6,370,753	4,940,591	3,932,176	2,718,127

Note.– Number of observations for a cohort changes because of the selection criterion for valid biographies.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

I.6 Sample selection: Women

Table I.7
Female labor force in West Germany for selected years, SOEP

Year	1988		1994		2000		2006	
Age range	20-53		25-59		31-59		37-59	
Labor force status	Observations	%	Observations	%	Observations	%	Observations	%
Employed[A]	7,688,808	76.48	8,571,636	78.89	7,935,596	80.88	7,112,982	77.5
Unemployed[A]	733,902	7.3	859,543	7.91	550,099	5.61	887,568	9.67
Apprentice[A]	448,453	4.46	71,369	0.66	9,951	0.1	16,838	0.18
Miner[A,B]	0	0	0	0	30,991	0.32	796	0.01
Sum of items above[A]	8,871,163	88.24	9,502,548	87.45	7,980,085	81.33	8,018,184	87.36
Civil servant	484,884	4.82	621,621	5.72	520,966	5.31	465,570	5.07
Self-employed	697,280	6.94	741,731	6.83	760,714	7.75	694,739	7.57
Total	10,053,327	100.00	10,865,900	100.00	9,811,864	100.01	9,178,493	100.00

Note.– Sample selection mirrors the respective birth cohorts in our deployed FDZ-RV – VSKT2002, 2004-2009_Bönke data.
[A]Labor force covered in the VSKT, [B]not weighted cell size < 5.
Source.– SOEP v28, own calculations using weighted data.

Table I.8
Weighted number of observations with valid UA-40 biographies, women

Birth cohort	Observations with valid UA40-biographies (weighted)	Actual cohort size at age 40			
		without foreigners		including foreigners	
		size	coverage[A]	size	coverage[A]
1935	214,783	422058	0.1247	442171	0.1190
1936	217,551	429765	0.1242	450445	0.1185
1937	207,309	432061	0.1166	453370	0.1111
1938	221,022	456876	0.1355	480458	0.1288
1939	245,519	483157	0.1244	508572	0.1182
1940	233,767	486274	0.1330	515466	0.1254
1941	216,453	449906	0.1592	477378	0.1500
1942	172,882	362910	0.1667	394090	0.1535
1943	175,621	371086	0.1970	402179	0.1818
1944	173,017	361942	0.1928	394009	0.1771
1945	126,931	265921	0.2093	299934	0.1856
1946	162,292	309634	0.2393	346937	0.2136
1947	178,106	341762	0.2418	381039	0.2169
1948	188,304	360960	0.2321	404417	0.2072
1949	201,483	385314	0.2593	432152	0.2312
1950	210,781	.	.	440548	0.2118
1951	202,075	.	.	441156	0.2061
1952	207,547	.	.	452801	0.2179
1953	198,846	.	.	450526	0.2135
1954	218,223	.	.	464567	0.1989
1955	218,160	.	.	469187	0.2146
1956	232,274	.	.	486236	0.2007
1957	237,176	.	.	499854	0.2105
1958	242,756	.	.	508413	0.1921
1959	258,979	.	.	533458	0.2056
1960	267,044	.	.	553228	0.1863
1961	267,736	.	.	568443	0.1906
1962	279,379	.	.	577470	0.2007
1963	276,530	.	.	596598	0.1959
1964	280,680	.	.	602147	0.1963
1965	282,497	.	.	593553	0.2189
1966	283,604	.	.	595330	0.2034
1967	288,091	.	.	583619	0.2185
1968	277,011	.	.	569535	0.2093
1969	261,663	.	.	542073	0.2639

Note.– Cohorts 1935 – 1949: West Germany including West Berlin, cohorts 1950 – 1969: West Germany including Berlin.
[A]Coverage equals the number of observations with valid UA-40 biographies (weighted) divided by actual cohort size at 40.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke. Actual cohort size at age 40 according to Federal Statistical Office, own calculations using weighted data.

Table I.9
Number of observed women

Cohort	Original number of observations	Sample size after exclusion of: Foreigners	East Germans	Repatriates (*Aussiedler*)	Self-employed craftsmen	Rejoiners (marriage-law)[A]	Incomplete biographies (up to age 40)
1935	4,678	3,359	2,546	2,390	2,387	2,072	344
1936	4,753	3,589	2,596	2,433	2,428	2,049	354
1937	5,004	3,766	2,777	2,563	2,553	2,144	381
1938	4,784	3,634	2,674	2,498	2,485	2,078	407
1939	4,900	3,705	2,635	2,458	2,446	2,016	373
1940	5,004	3,783	2,711	2,570	2,564	2,157	425
1941	4,923	3,608	2,575	2,440	2,432	2,051	433
1942	4,879	3,668	2,637	2,491	2,482	2,155	468
1943	4,777	3,572	2,593	2,485	2,472	2,228	519
1944	4,614	3,499	2,481	2,375	2,364	2,181	479
1945	4,423	3,249	2,442	2,333	2,316	2,196	531
1946	4,090	3,005	2,378	2,300	2,286	2,209	525
1947	4,071	3,049	2,275	2,175	2,157	2,110	539
1948	4,076	2,987	2,263	2,162	2,136	2,124	508
1949	3,969	2,892	2,165	2,054	2,040	2,033	521
1950	4,104	3,010	2,133	1,982	1,972	1,972	504
1951	4,259	3,138	2,211	2,047	2,035	2,035	522
1952	4,283	3,115	2,198	2,049	2,030	2,030	551
1953	4,381	3,161	2,196	2,011	1,995	1,995	528
1954	4,495	3,182	2,185	1,983	1,964	1,964	499
1955	4,591	3,196	2,239	2,066	2,055	2,055	558
1956	4,684	3,158	2,242	2,053	2,044	2,044	503
1957	4,812	3,209	2,330	2,167	2,150	2,150	567
1958	4,973	3,172	2,321	2,144	2,129	2,129	501
1959	5,218	3,261	2,379	2,222	2,207	2,207	552
1960	5,579	3,381	2,446	2,300	2,288	2,288	545
1961	5,886	3,533	2,555	2,412	2,399	2,399	565
1962	6,538	3,804	2,675	2,558	2,542	2,542	654
1963	6,550	3,723	2,672	2,537	2,519	2,519	647
1964	6,769	3,673	2,685	2,577	2,564	2,564	657
1965	6,995	3,799	2,711	2,621	2,605	2,605	692
1966	6,949	3,706	2,732	2,649	2,624	2,624	657
1967	7,002	3,689	2,695	2,624	2,609	2,609	700
1968	7,197	3,683	2,749	2,674	2,663	2,663	698
1969	7,390	3,725	2,785	2,718	2,703	2,703	888
Total	181,600	119,683	86,887	82,121	81,645	77,900	18,795

Note.− Second column: Original sample size. Columns three to eight: All numbers denote observation after the stepwise exclusion of the respective groups. [A]Insured who left and rejoined the statutory pension system due to the law of marriage refunds (valid until 1967), which was possible until 1995. The earnings reported of these insured do not correspond to those actually earned.
Source.− FDZ-RV − VSKT2002, 2004-2009_Bönke, own calculations using unweighted data.

Table I.10
Weighted shares of groups from the initial dataset, women

Cohort	Foreigners	East Germans	Repatriates (*Aussiedler*)	Self-employed craftsmen	Rejoiners (marriage-law) [A]	"Incomplete" Up to 40	"Complete" Up to 40
1935	0.0618	0.2086	0.0362	0.0008	0.0897	0.5902	0.1024
1936	0.0635	0.2055	0.0356	0.0014	0.1059	0.5927	0.1013
1937	0.0697	0.2568	0.0446	0.0027	0.1004	0.5318	0.0944
1938	0.0724	0.2481	0.0370	0.0031	0.1034	0.5299	0.1094
1939	0.0736	0.2610	0.0355	0.0025	0.1066	0.5277	0.0997
1940	0.0822	0.2819	0.0271	0.0010	0.0920	0.5020	0.1058
1941	0.0842	0.2710	0.0279	0.0016	0.0851	0.4905	0.1249
1942	0.1082	0.2605	0.0270	0.0017	0.0758	0.4749	0.1276
1943	0.1082	0.2611	0.0238	0.0020	0.0563	0.4547	0.1503
1944	0.1130	0.2521	0.0221	0.0020	0.0417	0.4650	0.1457
1945	0.1557	0.2197	0.0221	0.0028	0.0292	0.4483	0.1513
1946	0.1580	0.1719	0.0177	0.0028	0.0175	0.4725	0.1772
1947	0.1526	0.2077	0.0253	0.0030	0.0124	0.4386	0.1727
1948	0.1543	0.1853	0.0244	0.0045	0.0027	0.4665	0.1650
1949	0.1456	0.1976	0.0311	0.0029	0.0012	0.4413	0.1815
1950	0.1492	0.2207	0.0420	0.0017	0.0000	0.4232	0.1632
1951	0.1368	0.2327	0.0404	0.0026	0.0000	0.4273	0.1602
1952	0.1397	0.2390	0.0341	0.0030	0.0000	0.4142	0.1700
1953	0.1315	0.2352	0.0445	0.0028	0.0000	0.4201	0.1659
1954	0.1289	0.2470	0.0468	0.0029	0.0000	0.4192	0.1552
1955	0.1267	0.2356	0.0395	0.0016	0.0000	0.4292	0.1673
1956	0.1198	0.2250	0.0469	0.0019	0.0000	0.4482	0.1582
1957	0.1116	0.2265	0.0387	0.0033	0.0000	0.4535	0.1664
1958	0.1126	0.2328	0.0422	0.0023	0.0000	0.4584	0.1517
1959	0.1079	0.2338	0.0389	0.0030	0.0000	0.4551	0.1614
1960	0.1113	0.2381	0.0355	0.0023	0.0000	0.4649	0.1479
1961	0.1032	0.2379	0.0370	0.0024	0.0000	0.4685	0.1510
1962	0.1087	0.2530	0.0254	0.0028	0.0000	0.4499	0.1603
1963	0.1094	0.2432	0.0322	0.0029	0.0000	0.4548	0.1575
1964	0.1148	0.2413	0.0256	0.0024	0.0000	0.4578	0.1581
1965	0.1196	0.2258	0.0193	0.0025	0.0000	0.4552	0.1776
1966	0.1212	0.2137	0.0193	0.0037	0.0000	0.4760	0.1660
1967	0.1241	0.2138	0.0178	0.0027	0.0000	0.4621	0.1796
1968	0.1362	0.2032	0.0161	0.0019	0.0000	0.4699	0.1727
1969	0.1500	0.1897	0.0131	0.0024	0.0000	0.4283	0.2166

Note.− The numbers denote observation share of the respective groups before excluding any observations. [A]Insured who left and rejoined the statutory pension system due to the law of marriage refunds (valid until 1967), which was possible until 1995. The earnings reported of these insured do not correspond to those actually earned.
Source.− FDZ-RV − VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

Table I.11
Number of observed women with valid UAX-biographies

Birth cohort	Up to 40	Up to 45	Up to 50	Up to 55	Up to 60
1935	344	332	318	311	313
1936	354	349	349	329	336
1937	381	366	348	336	346
1938	407	403	387	360	362
1939	373	360	361	356	346
1940	425	420	419	432	440
1941	433	420	422	440	438
1942	468	463	472	479	476
1943	519	506	496	500	505
1944	479	481	481	478	472
1945	531	544	537	528	519
1946	525	502	492	496	500
1947	539	521	517	527	517
1948	508	501	512	514	500
1949	521	526	531	523	514
1950	504	515	507	512	
1951	522	535	538	539	
1952	551	539	533	528	
1953	528	523	526	522	
1954	499	503	514	511	
1955	558	569	565		
1956	503	527	550		
1957	567	567	581		
1958	501	526	541		
1959	552	567	578		
1960	545	553			
1961	565	561			
1962	654	654			
1963	647	646			
1964	657	661			
1965	692				
1966	657				
1967	700				
1968	698				
1969	888				
Total	18,795	15,140	12,075	9,221	6,584

Note.– Number of observations for a cohort changes because of the selection criterion for valid biographies.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using unweighted data.

Table I.12
Weighted number of observations with valid UAX-biographies, women

Birth cohort	Up to 40	Up to 45	Up to 50	Up to 55	Up to 60
1935	52,622	50,353	48,384	47,316	48,175
1936	53,356	52,572	53,091	49,557	51,276
1937	50,362	47,803	45,417	43,887	45,512
1938	61,887	61,498	58,893	55,125	55,975
1939	60,105	58,912	58,727	57,776	57,050
1940	64,653	64,182	64,474	66,696	69,552
1941	71,620	68,661	69,313	72,244	70,475
1942	60,512	60,087	60,867	61,635	62,068
1943	73,109	71,529	69,801	70,128	71,327
1944	69,795	70,204	70,192	69,223	68,046
1945	55,662	56,701	54,976	54,022	53,185
1946	74,109	70,441	68,875	68,793	70,248
1947	82,645	79,152	79,550	80,563	79,427
1948	83,779	83,093	83,629	83,924	83,742
1949	99,915	99,882	100,330	97,999	96,264
1950	93,323	95,009	95,676	95,400	
1951	90,938	93,035	93,411	93,860	
1952	98,684	97,226	96,700	96,101	
1953	96,190	95,506	95,110	95,126	
1954	92,391	93,927	96,216	94,149	
1955	100,683	103,887	103,634		
1956	97,569	101,439	105,649		
1957	105,207	107,349	109,690		
1958	97,649	99,741	101,691		
1959	109,672	111,108	113,736		
1960	103,066	103,457			
1961	108,367	107,331			
1962	115,887	116,899			
1963	116,896	116,696			
1964	118,184	117,276			
1965	129,950				
1966	121,063				
1967	127,507				
1968	119,226				
1969	143,057				
Total	3,199,640	2,554,956	1,998,032	1,453,524	982,322

Note.– Number of observations for a cohort changes because of the selection criterion for valid biographies.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

Appendix II: Results for women

This section replicates graphs 1-10 from the paper for women and provides information on their educational attainment.

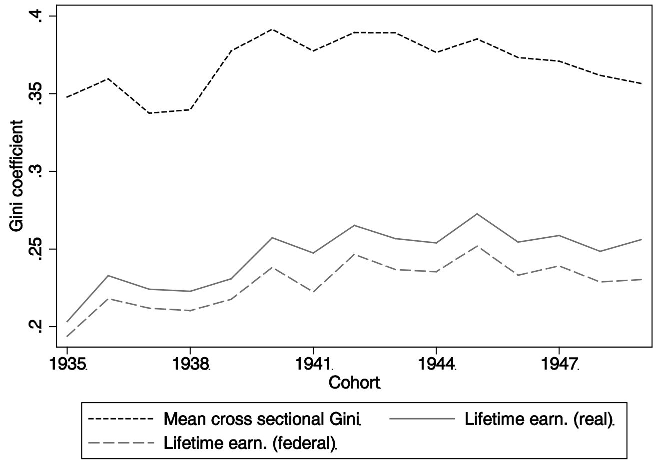


Fig. II.1.– Means of annual Gini coefficients and Gini coefficients of lifetime earnings for cohorts 1935 - 1949

Note.– real denotes CPI discounting, federal denotes federal bond discounting.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

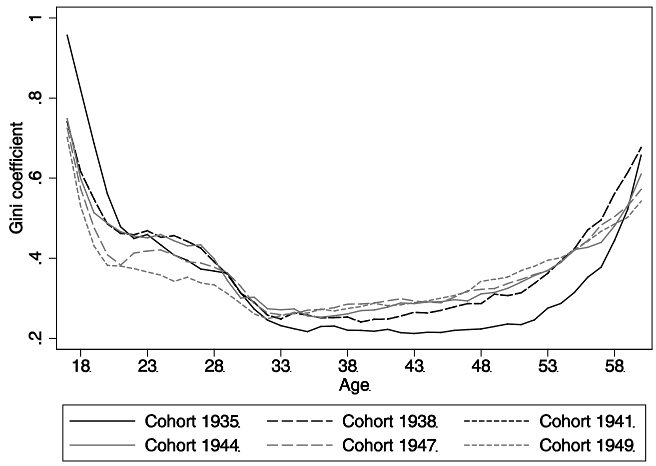


Fig. II.2.– Annual Gini coefficients from age 17 to age 60 for cohorts 1935 - 1949

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

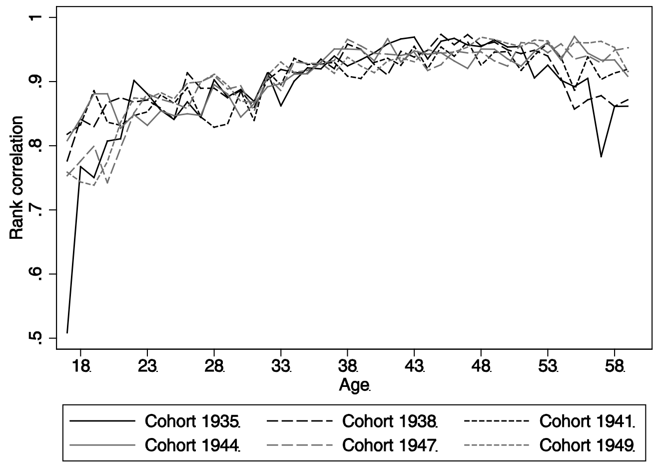


Fig. II.3.– Earnings rank correlations between consecutive years for cohorts 1935-1949

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

Fig. II.4.– Rank correlation of annual and lifetime earnings for cohorts 1935-1949

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

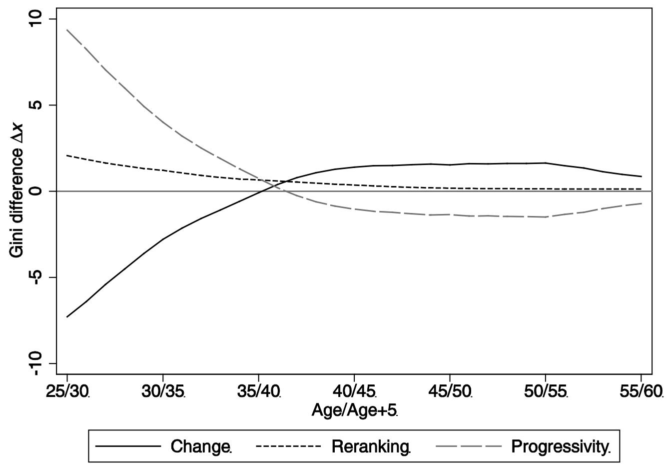


Fig. II.5.– Decomposition of changes in inequality as of Eq. (3) for cohort 1944

Note.– Accumulated discounted earnings refer to the age in the abscissa as compared to accumulated earnings five years later, as in Eq. (3) in the paper. Coefficients are multiplied by 100.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

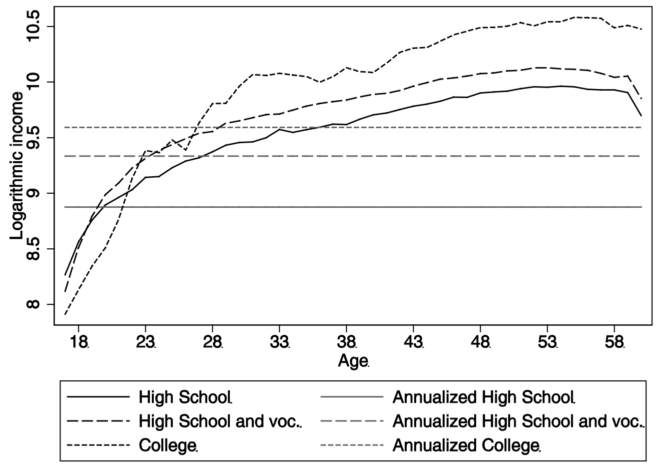


Fig. II.6.– Age-earning-profiles by highest educational attainment for pooled cohorts 1935-1949

Note.– voc. abbreviates vocational training.

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

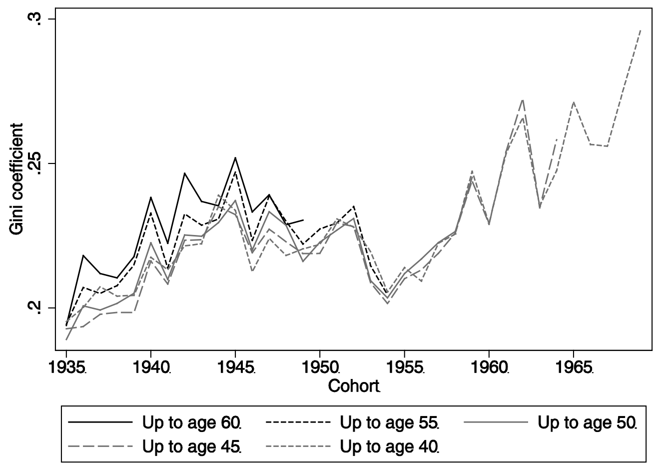


Fig. II.7.– Gini coefficients of UAX for cohorts 1935-1969

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

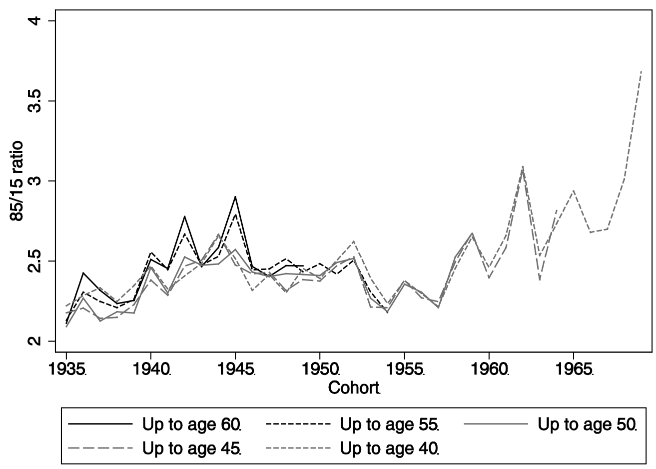


Fig. II.8.– 85th / 15th ratio of UAX- earnings for cohorts 1935-1969

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

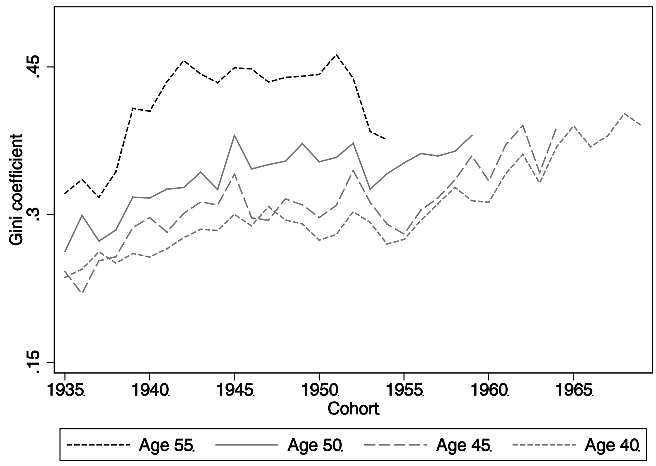


Fig. II.9.– Gini coefficients of annual earnings at various ages for cohorts 1935-1969

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

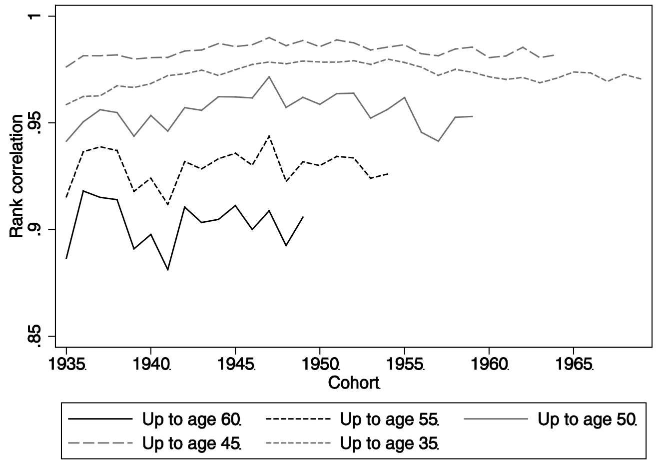


Fig. II.10.– Rank correlation of UA-40 with selected UAX for cohorts 1935-1969

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

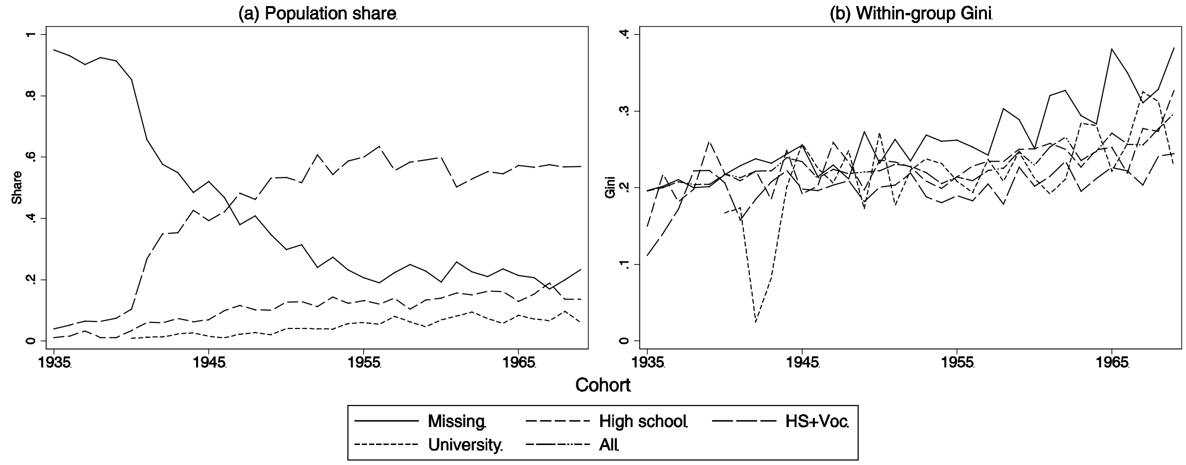


Fig. II.11.– Educational attainment and inequality in our sample, women

Note.– Within-group Gini coefficients refer to the distributions of UA-40 with federal bond discounting.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

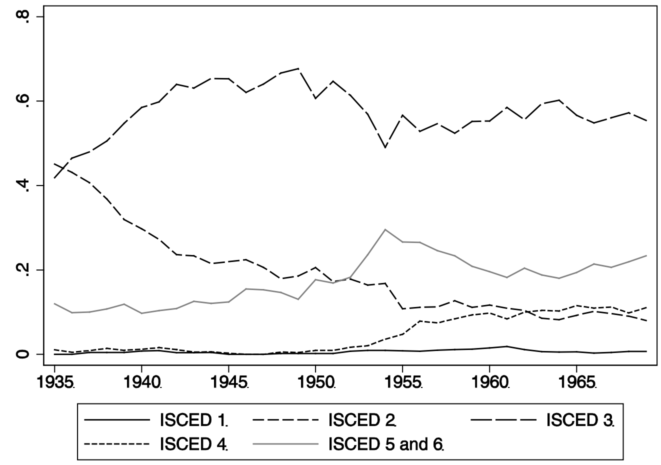


Fig. II.12.– Educational attainment of cohorts of West German women according to the SOEP

Note.– The education groups are defined according to the International Standard Classification of Education 1997 (ISCED-97): ISCED 1: Primary education; ISCED 2: Lower secondary education; ISCED 3: Upper secondary education; ISCED 4: Post-secondary non-tertiary education; ISCED 5/6: Tertiary education. All cohorts born after 1944 are analyzed at age 40. Since the SOEP starts in 1984, older cohorts are analyzed at the closest distance to age 40, e.g. age 45 for those born in 1939.
Source.– SOEP v28, own calculations using weighted data.

Appendix III: Robustness and supplementary graphics

III.1 Confidence intervals for UAX-earnings

Table III.1
UAX Ginis for selected cohorts, men

Cohort	Up to 40	Up to 45	Up to 50	Up to 55	Up to 60
1935	0.121	0.125	0.135	0.145	0.156
	(0.114; 0.128)	(0.119; 0.134)	(0.127; 0.144)	(0.137; 0.155)	(0.145; 0.167)
1940	0.118	0.131	0.151	0.166	0.177
	(0.111; 0.126)	(0.123; 0.141)	(0.141; 0.164)	(0.155; 0.182)	(0.165; 0.192)
1945	0.138	0.159	0.172	0.185	0.196
	(0.130; 0.147)	(0.149; 0.172)	(0.160; 0.186)	(0.173; 0.203)	(0.183; 0.214)
1950	0.146	0.161	0.178	0.193	
	(0.138; 0.156)	(0.151; 0.173)	(0.167; 0.196)	(0.179; 0.210)	
1955	0.183	0.194	0.204		
	(0.173; 0.195)	(0.182; 0.208)	(0.191; 0.220)		
1960	0.204	0.218			
	(0.192; 0.218)	(0.205; 0.234)			
1965	0.210				
	(0.200; 0.223)				
1969	0.227				
	(0.215; 0.239)				

Note.– The UAX are based on federal bond discounting. Bias corrected and accelerated bootstrap confidence intervals at the 95%-level in brackets.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

Table III.2
UAX Ginis for selected cohorts, women

Cohort	Up to 40	Up to 45	Up to 50	Up to 55	Up to 60
1935	0.195	0.193	0.189	0.194	0.194
	(0.183; 0.210)	(0.180; 0.211)	(0.176; 0.203)	(0.180; 0.211)	(0.181; 0.212)
1940	0.218	0.216	0.223	0.233	0.238
	(0.203; 0.233)	(0.200; 0.233)	(0.207; 0.241)	(0.217; 0.248)	(0.221; 0.257)
1945	0.234	0.232	0.237	0.247	0.252
	(0.219; 0.249)	(0.218; 0.246)	(0.223; 0.254)	(0.233; 0.266)	(0.237; 0.270)
1950	0.222	0.219	0.223	0.227	
	(0.208; 0.240)	(0.204; 0.235)	(0.210; 0.238)	(0.214; 0.244)	
1955	0.214	0.210	0.212		
	(0.200; 0.229)	(0.196; 0.225)	(0.199; 0.228)		
1960	0.229	0.229			
	(0.215; 0.246)	(0.214; 0.248)			
1965	0.271				
	(0.254; 0.293)				
1969	0.297				
	(0.282; 0.313)				

Note.– The UAX are based on federal bond discounting. Bias corrected and accelerated bootstrap confidence intervals at the 95%-level in brackets.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

III.2 Alternative earnings concepts

Annotation: The calculations in this section are based on federal bond discounting unless stated otherwise. In order to obtain the earnings concept reported in the paper we apply three changes to the original earnings: the imputation of top coded earnings, the inclusion of the employers' social security contributions and the correction of the structural break. In this section we provide results for four alternative earnings concepts (see Table II.3):

(a) Original: Earnings as recorded in the dataset with no changes applied.

(b) Imputation: Original earnings with imputation of top coded earnings.

(c) Imputation, market wage: Original earnings with imputation of top coded earnings plus employers' social security contributions.

(d) Imputation, break: Original earnings with imputation of top coded earnings and correction of structural break.

Table III.3
Alternative earnings concepts

Earnings concept	Imputation of top-coded earnings	Including employers' social security contributions	Correction of the structural break
(a) Original			
(b) Imputation	X		
(c) Imputation, market wage	X	X	
(d) Imputation, break	X		X
Main concept in the paper	X	X	X

Note.– X marks if the change is included in the respective earnings concept.

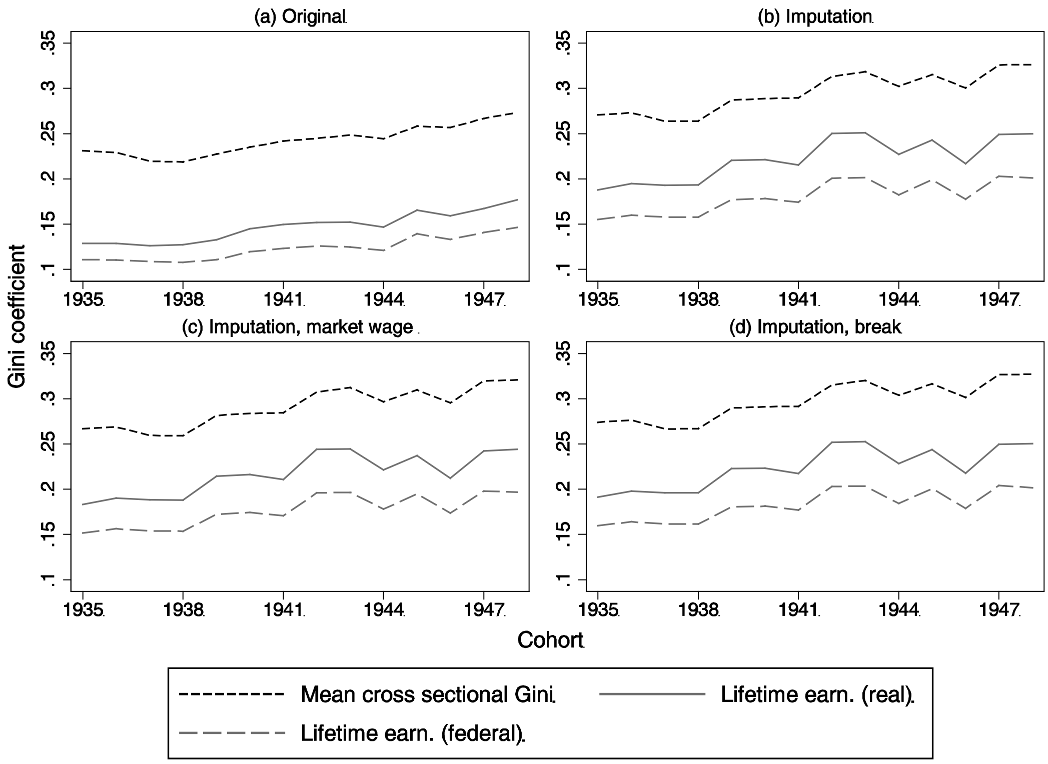


Fig. III.1.– Means of annual Gini coefficients and Gini coefficients of lifetime earnings, men

Note.– "real" denotes CPI discounting, "federal" denotes federal bond discounting. Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

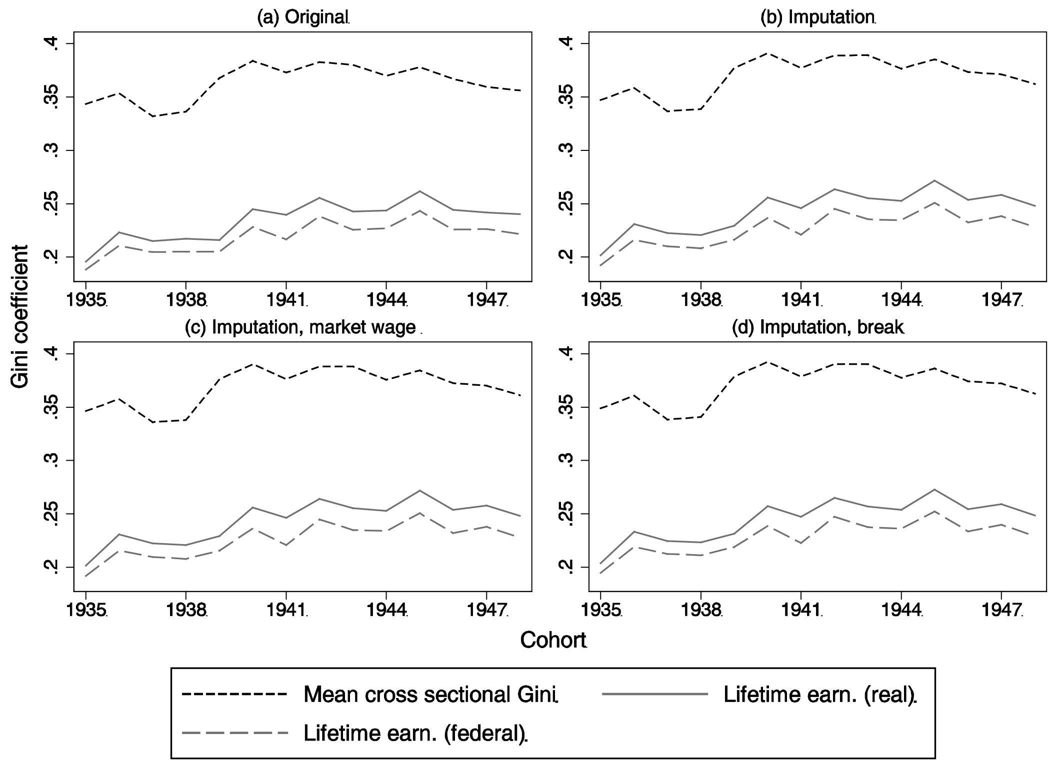


Fig. III.2.– Means of annual Gini coefficients and Gini coefficients of lifetime earnings, women

Note.– "real" denotes CPI discounting, "federal" denotes federal bond discounting.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

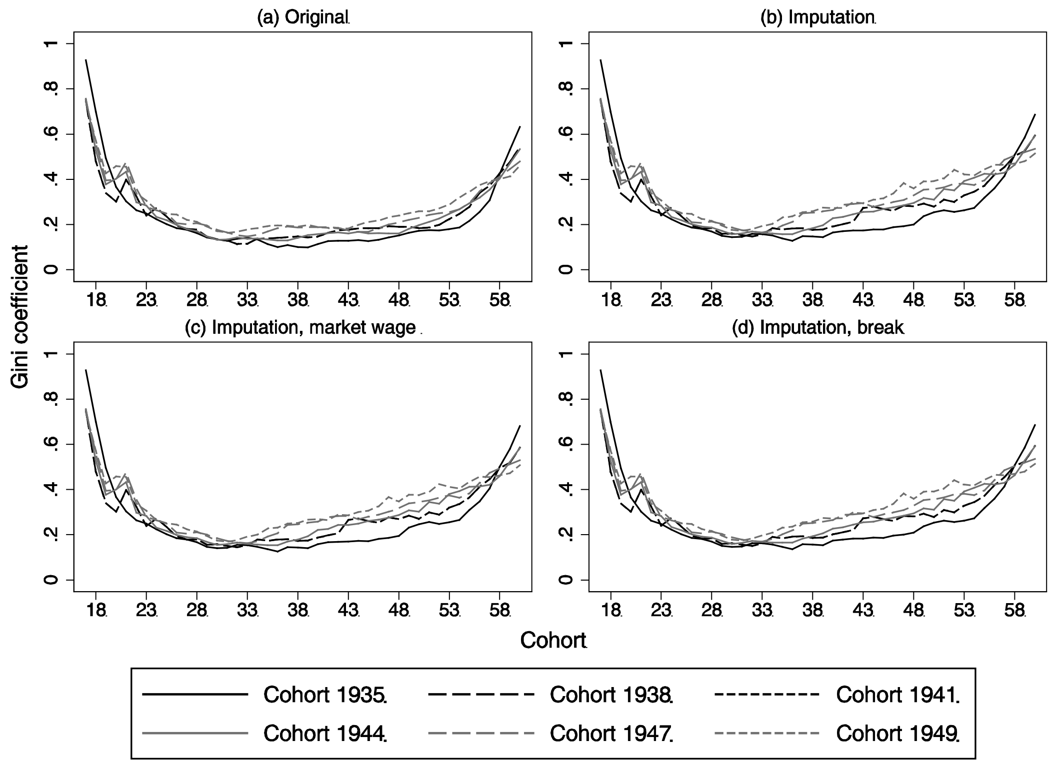


Fig. III.3.– Annual Gini coefficients from age 17 to 60 for cohorts 1935-1949, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

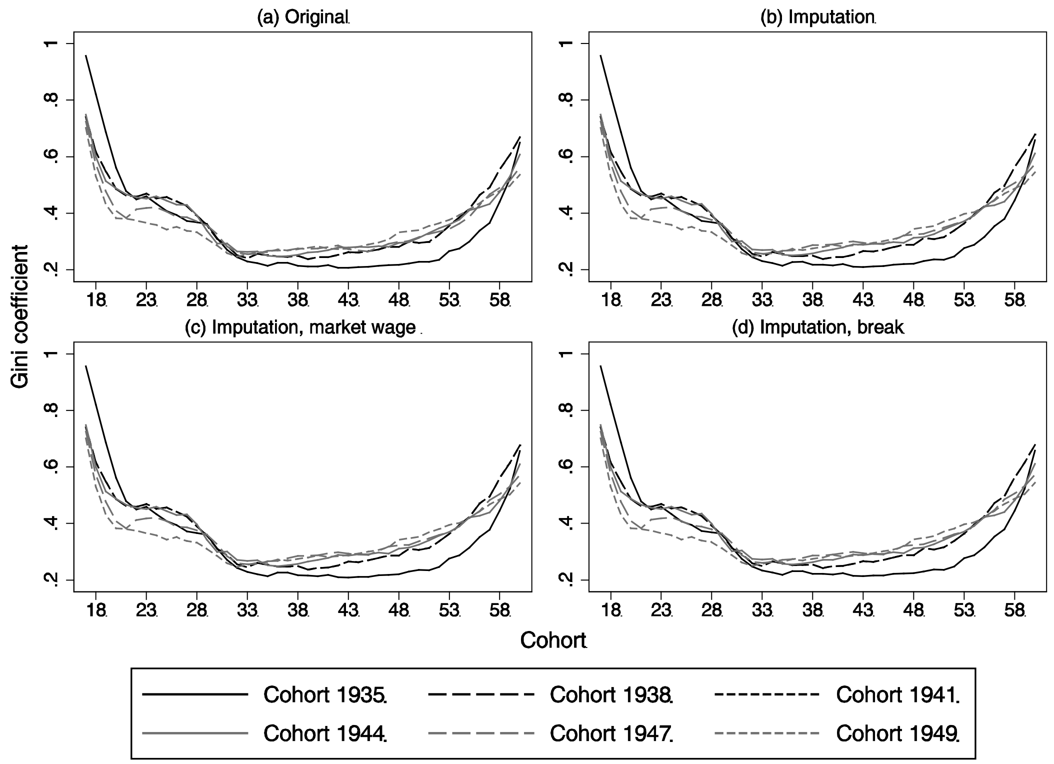


Fig. III.4.– Annual Gini coefficients from age 17 to 60 for cohorts 1935-1949, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

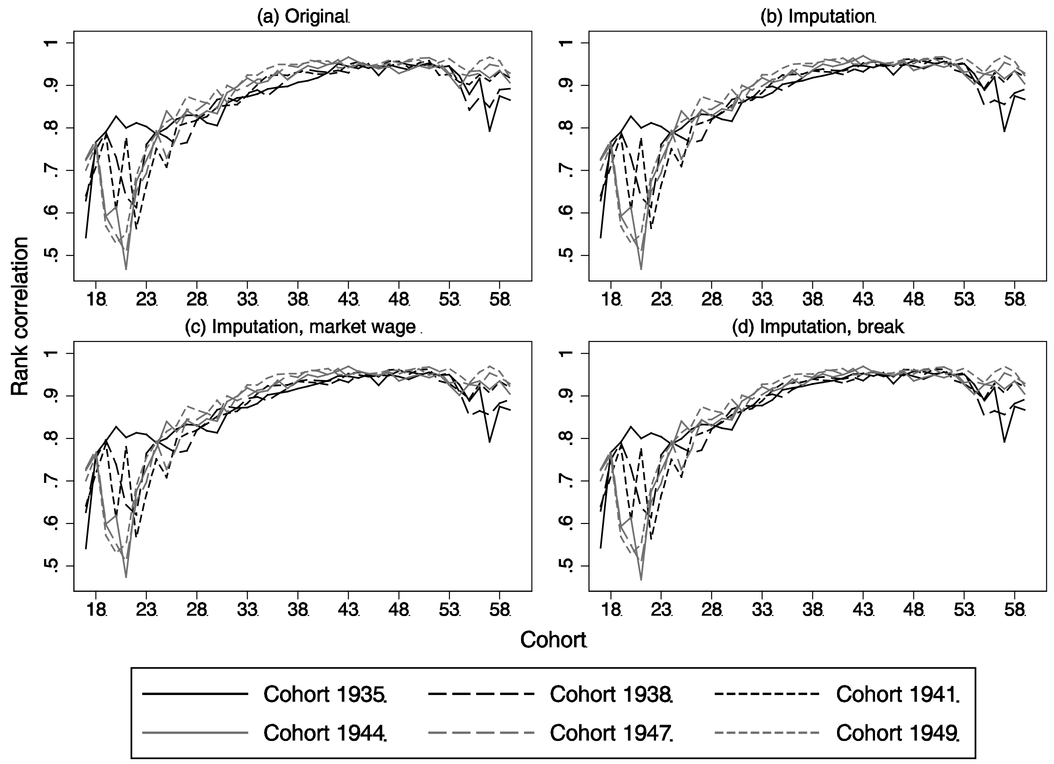


Fig. III.5.– Earnings rank correlations between consecutive years for cohorts 1935-1949, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

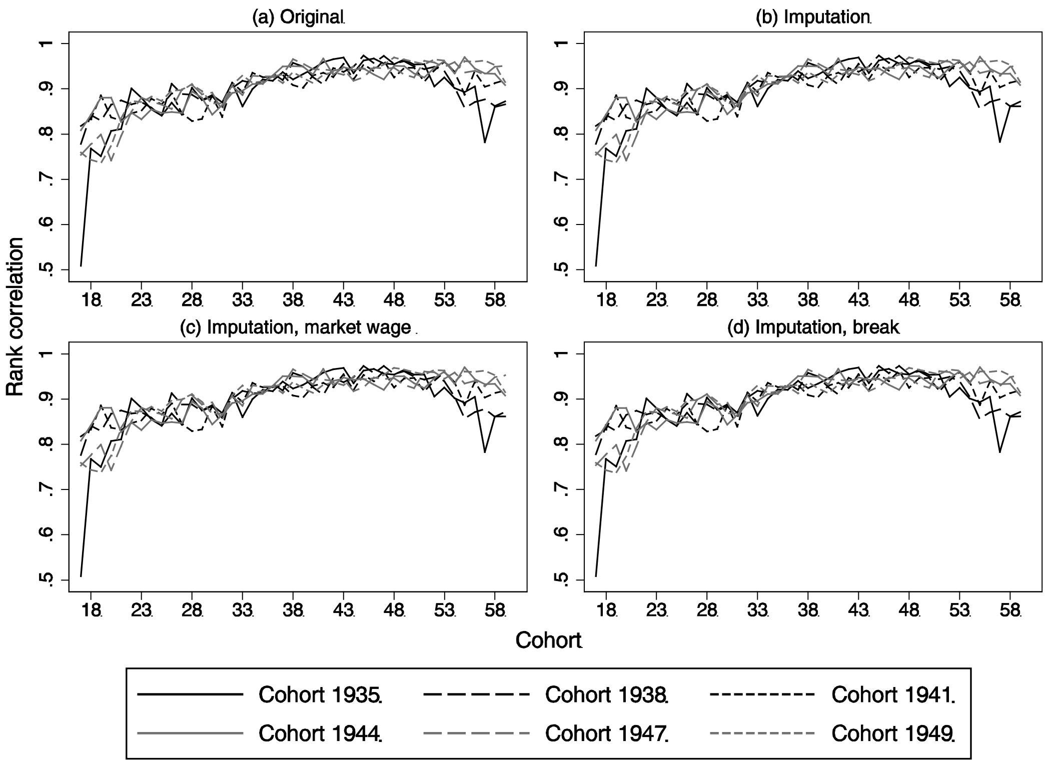


Fig. III.6.– Earnings rank correlations between consecutive years for cohorts 1935-1949, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

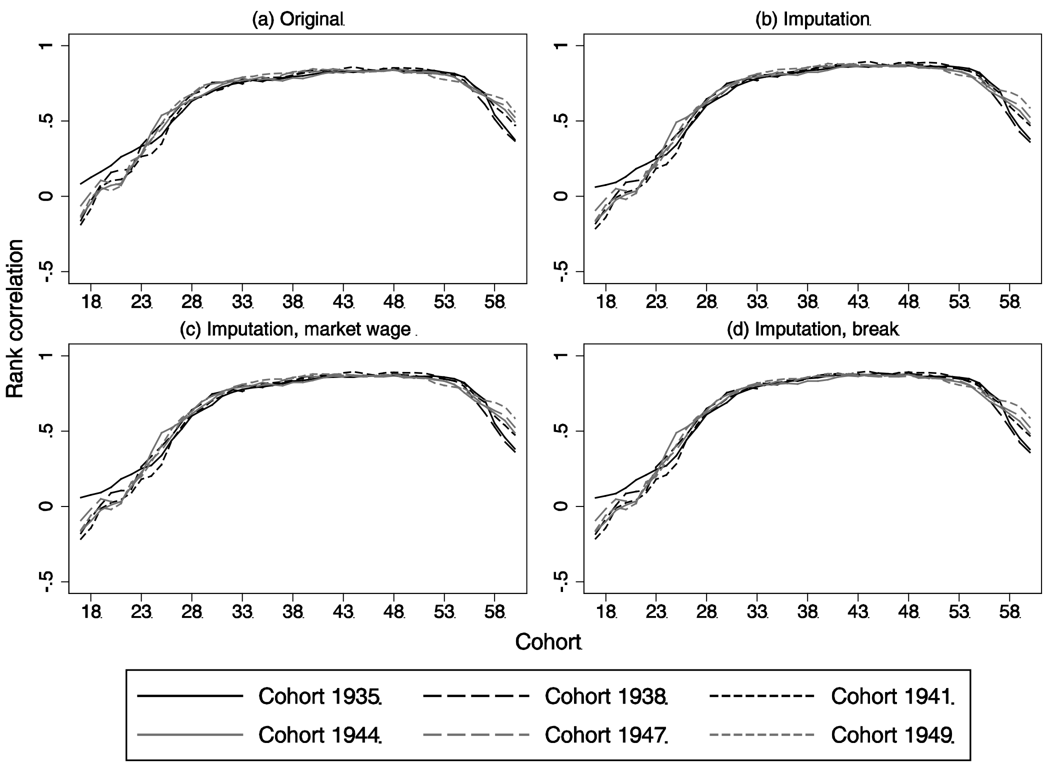


Fig. III.7.– Rank correlation of annual and lifetime earnings for cohorts 1935-1949, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

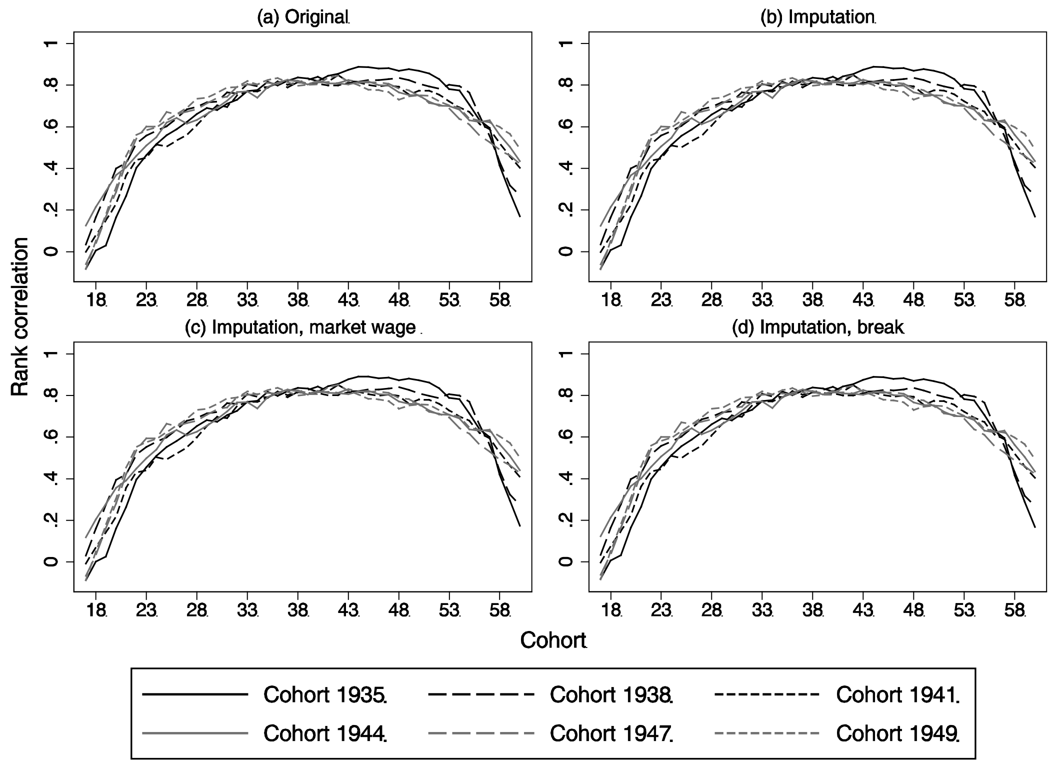


Fig. III.8.– Rank correlation of annual and lifetime earnings for cohorts 1935-1949, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

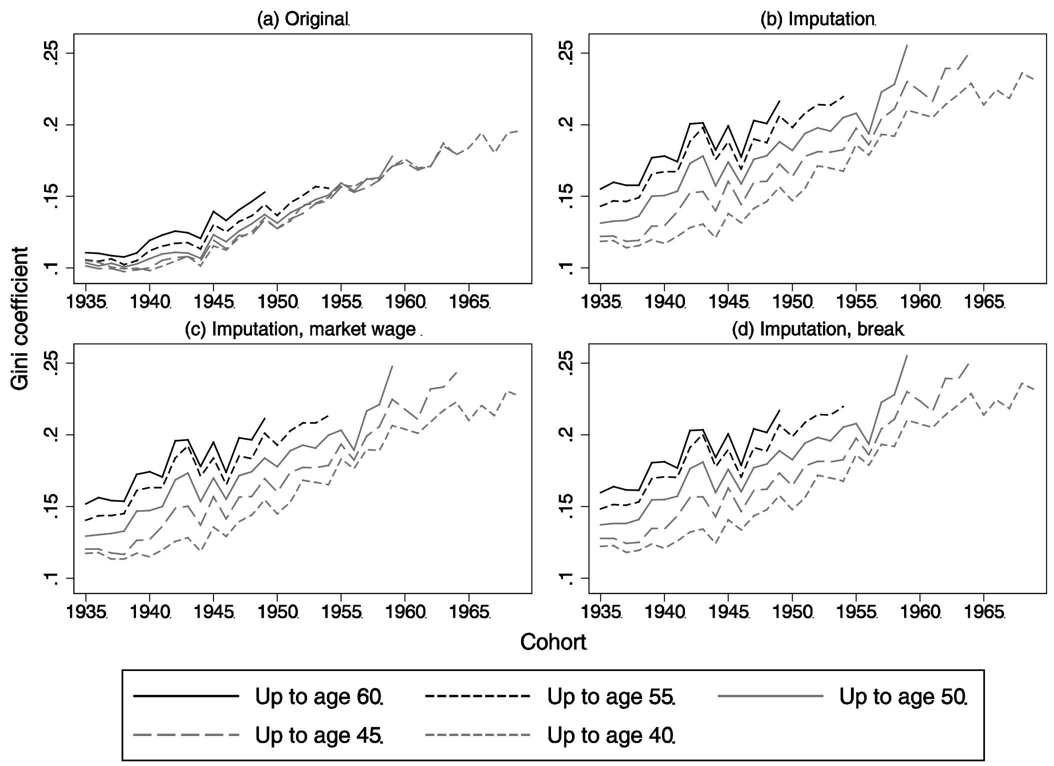


Fig. III.9.– Gini coefficients of UAX for cohorts 1935-1969, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

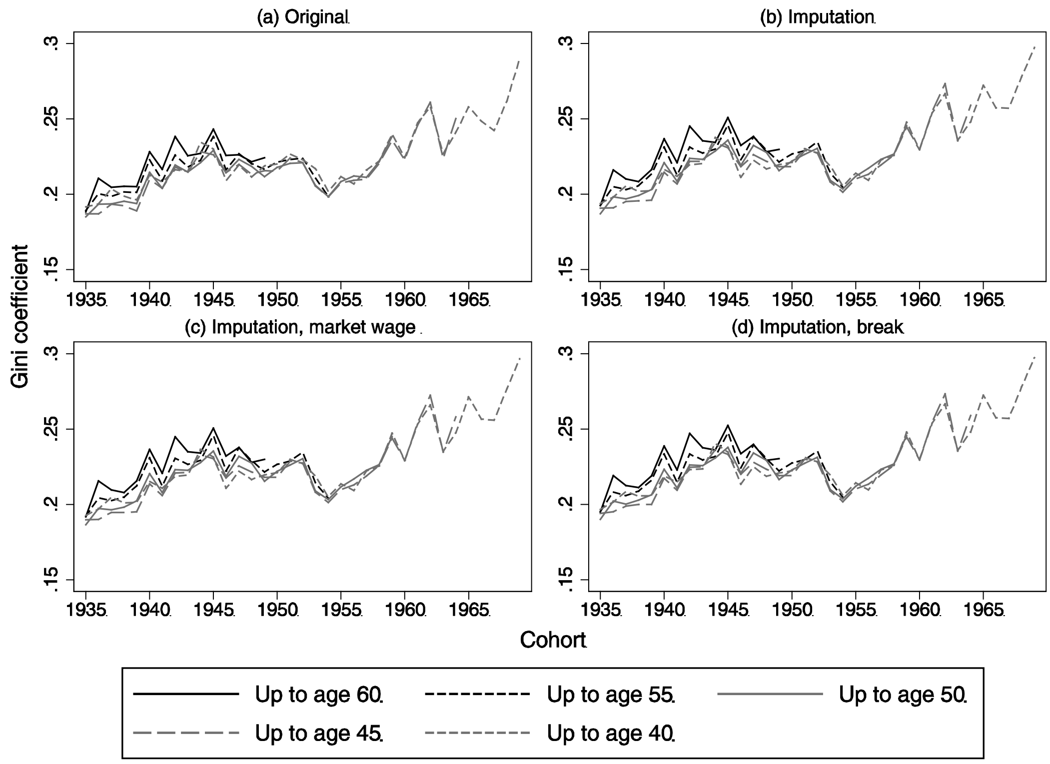


Fig. III.10.– Gini coefficients of UAX for cohorts 1935-1969, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

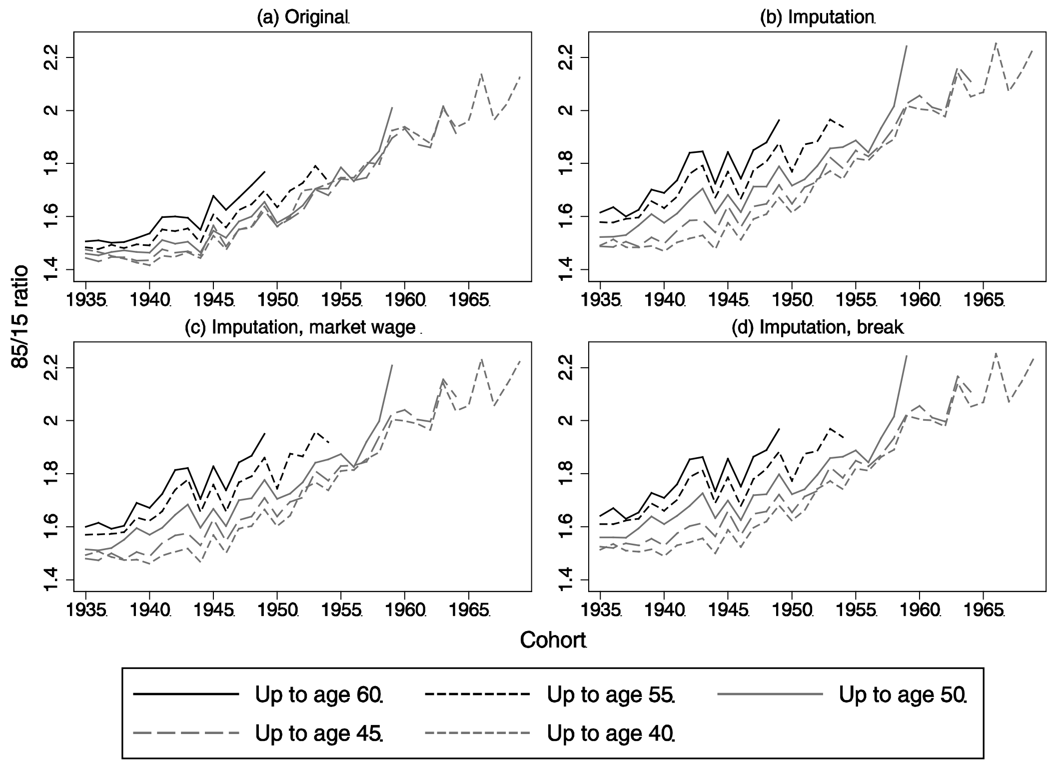


Fig. III.11.– 85th / 15th ratio of UAX- earnings for cohorts 1935-1969, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

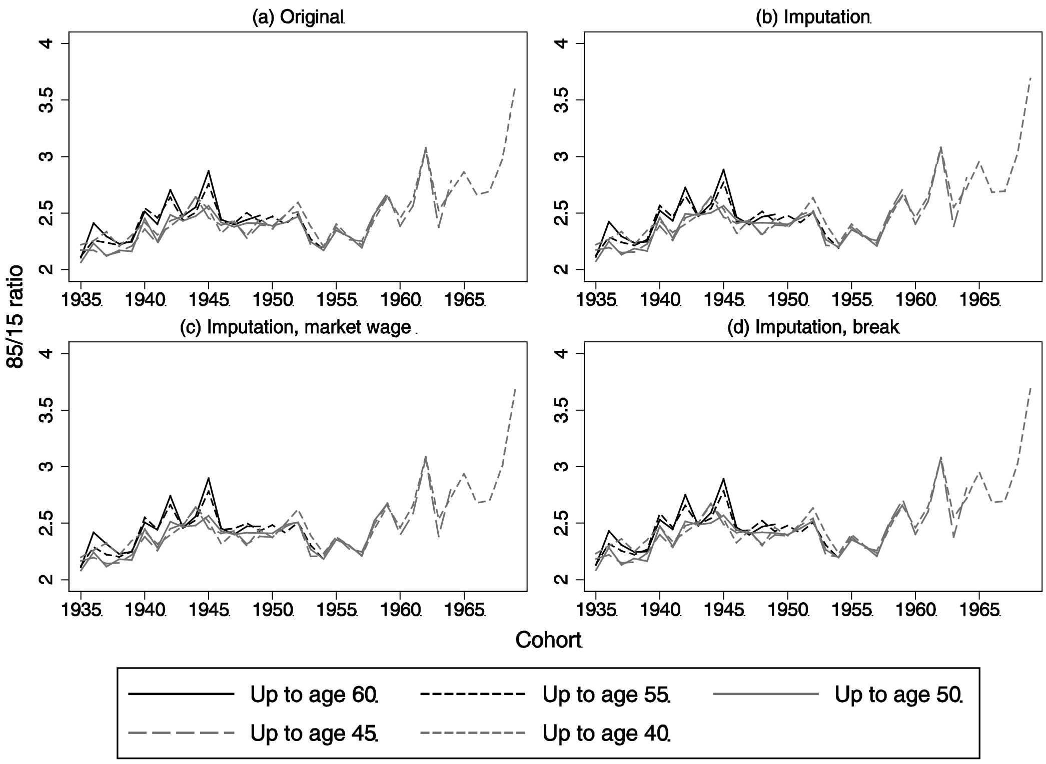


Fig. III.12.– 85th / 15th ratio of UAX- earnings for cohorts 1935-1969, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data

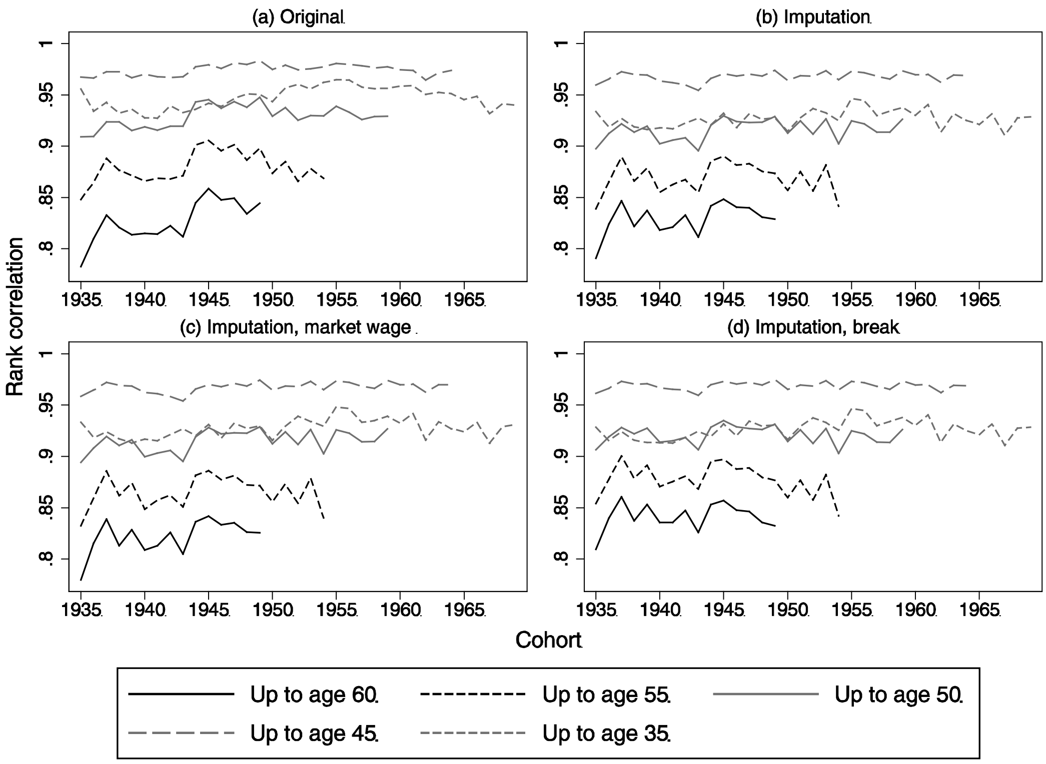


Fig. III.13.– Rank correlation of UA-40 with selected UAX, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

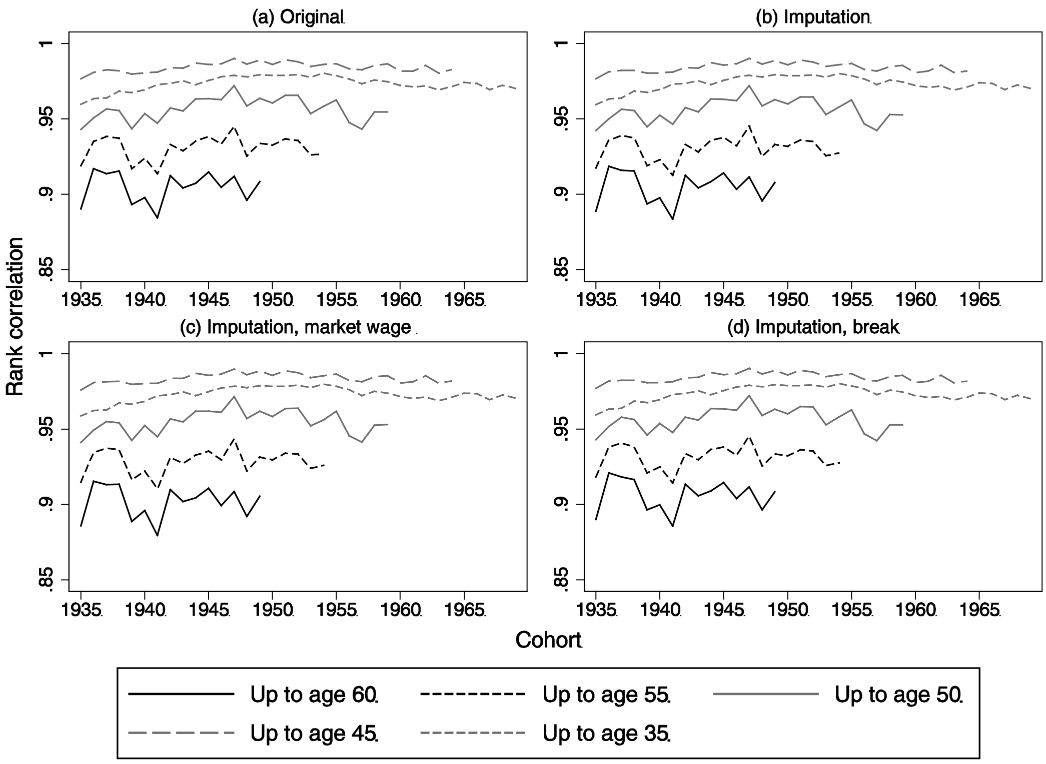


Fig. III.14.– Rank correlation of UA-40 with selected UAX, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

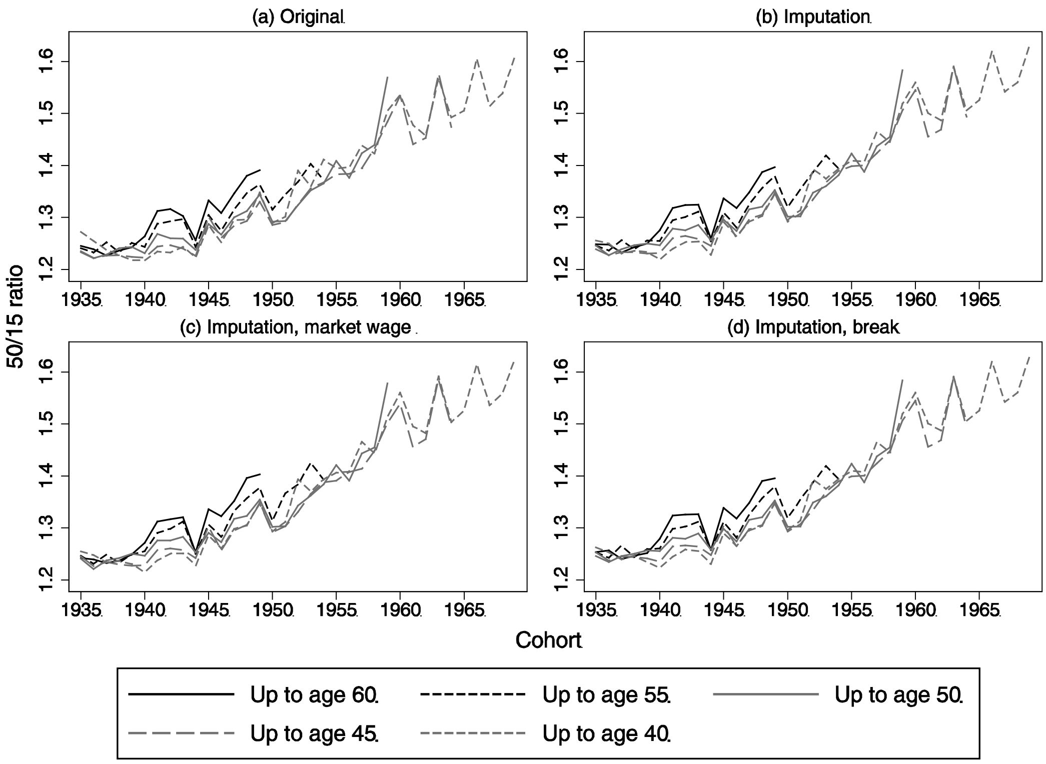


Fig. III.15.– 50th / 15th ratio of selected UAX for cohorts 1935-1969, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

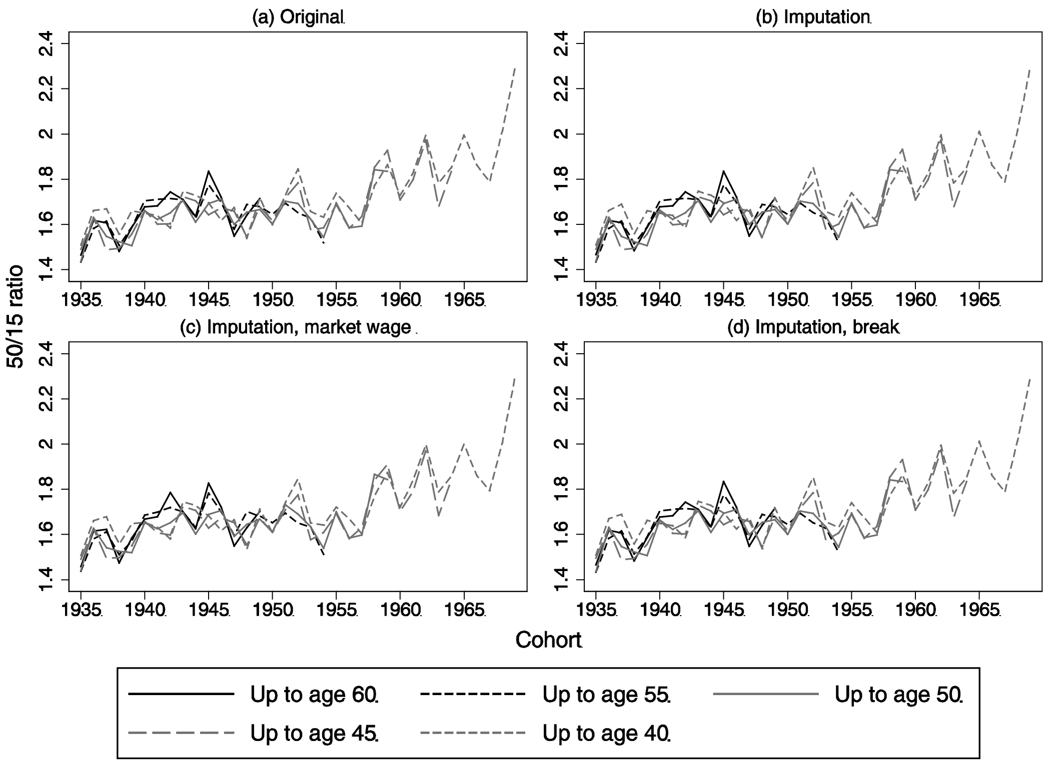


Fig. III.16.– 50^{th} / 15^{th} ratio of selected UAX for cohorts 1935-1969, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

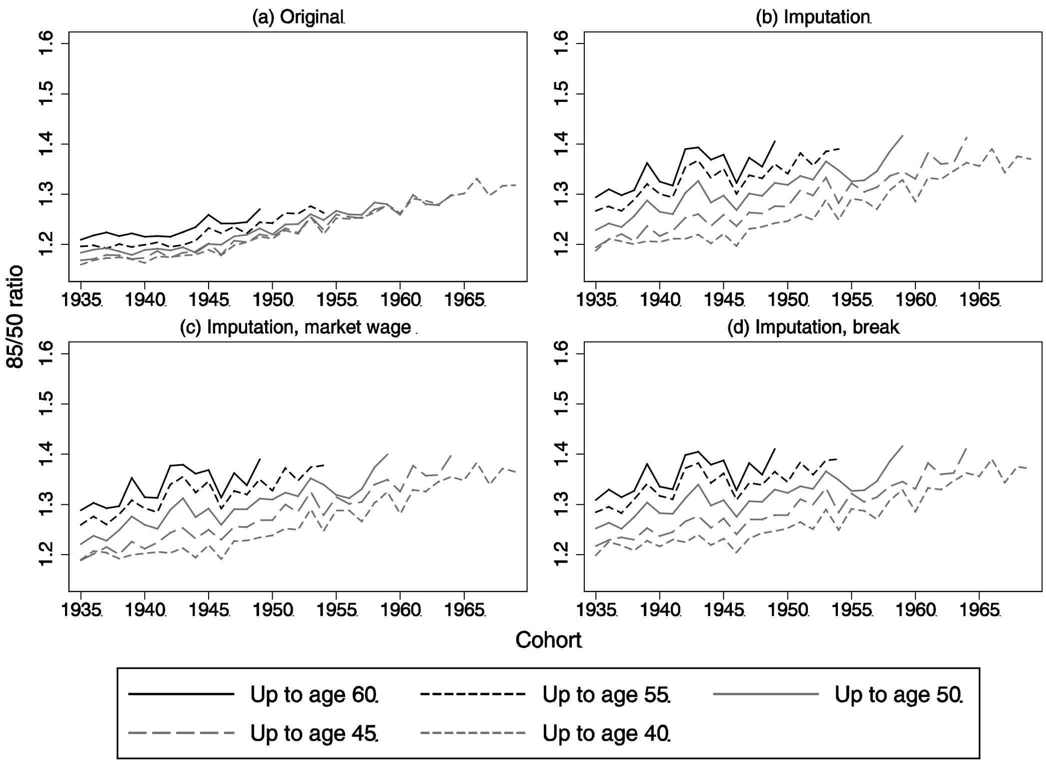


Fig. III.17.– 85^{th} / 50^{th} ratio of selected UAX for cohorts 1935-1969, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

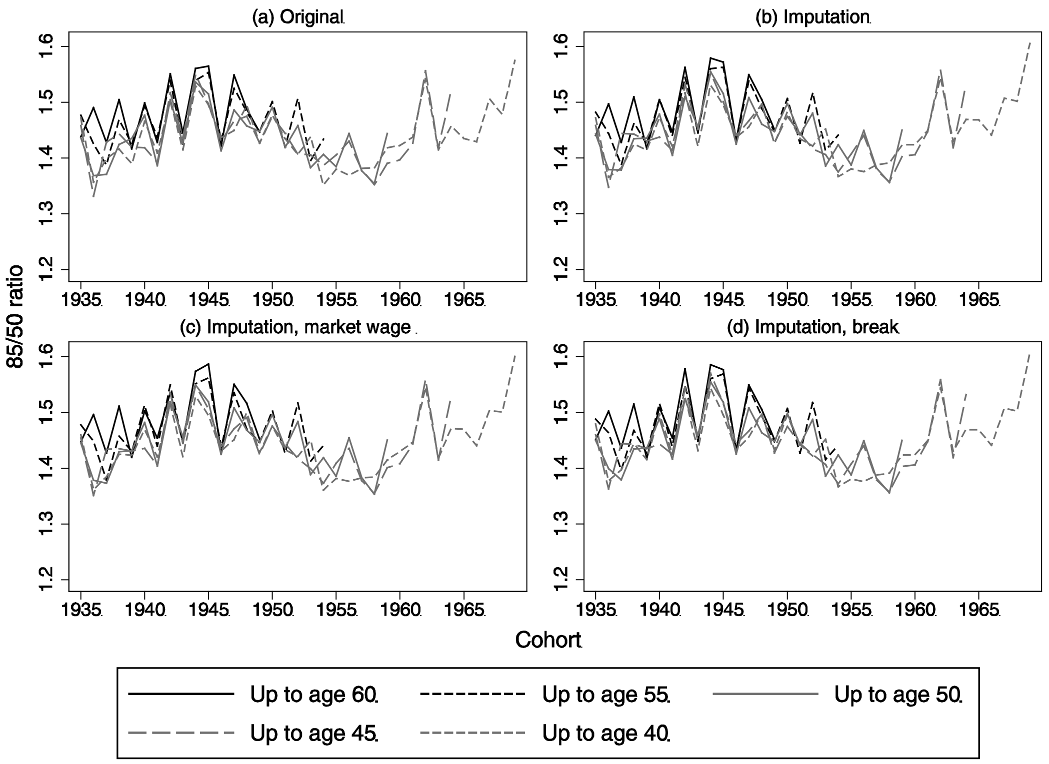


Fig. III.18.– 85th / 50th ratio of selected UAX for cohorts 1935-1969, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

III.3 Alternative imputation assumptions

Annotation: The calculations in this section are based on federal bond discounting unless stated otherwise. Each graph shows three different imputation assumptions as described in Online Appendix I:

(a) Minimal mobility depicts our main concept of minimal mobility of the imputed earnings.

(b) Maximal mobility depicts perfect mobility of the imputed earnings.

(c) Mean imputation wage assigns the average imputed wage to everyone above the contribution ceiling.

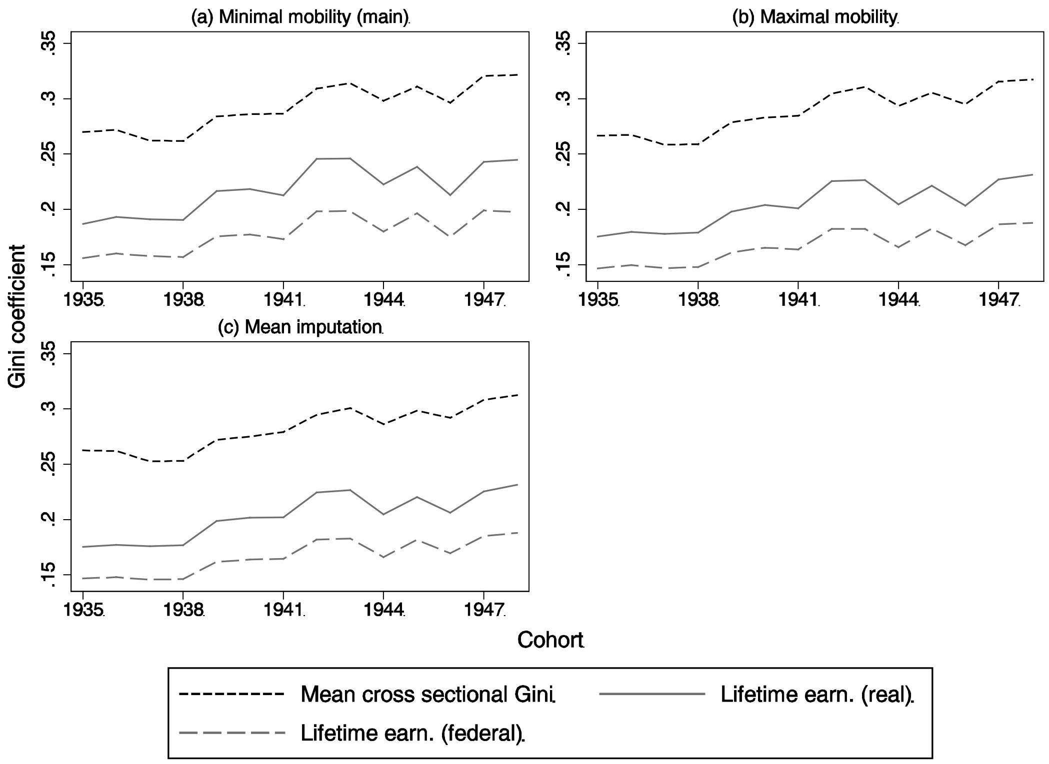


Fig. III.19.– Means of annual Gini coefficients and Gini coefficients of lifetime earnings, men

Note.– “real” denotes CPI discounting, “federal” denotes federal bond discounting.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

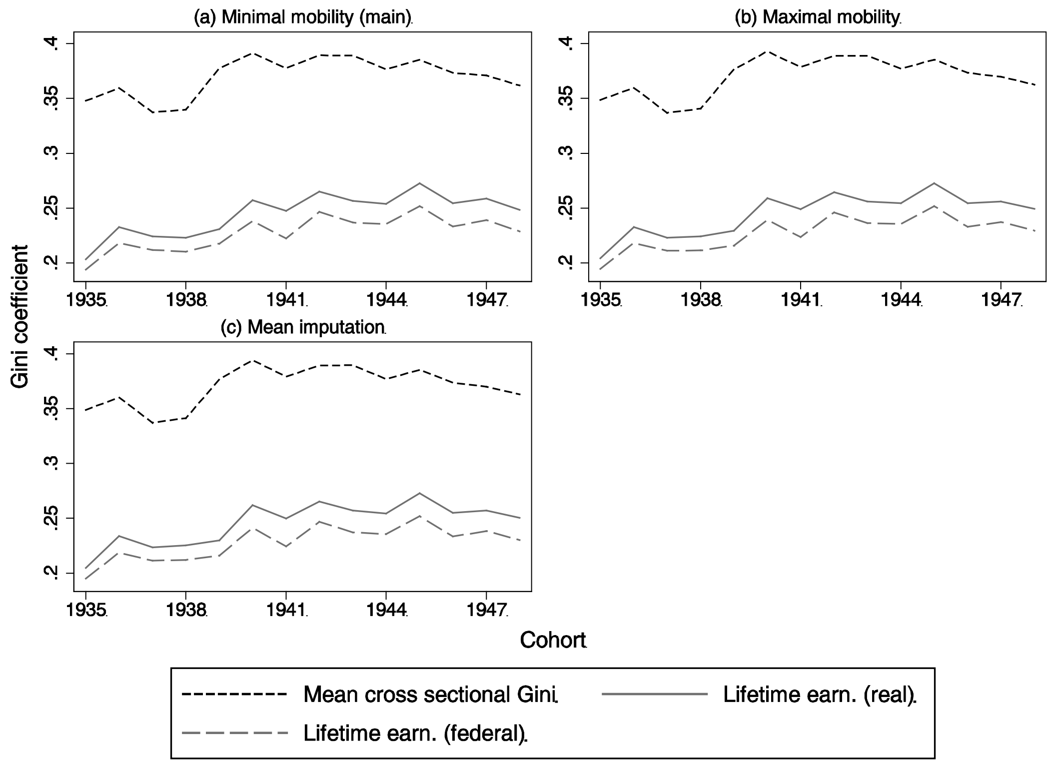


Fig. III.20.– Means of annual Gini coefficients and Gini coefficients of lifetime earnings, women

Note.– “real” denotes CPI discounting, “federal” denotes federal bond discounting.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

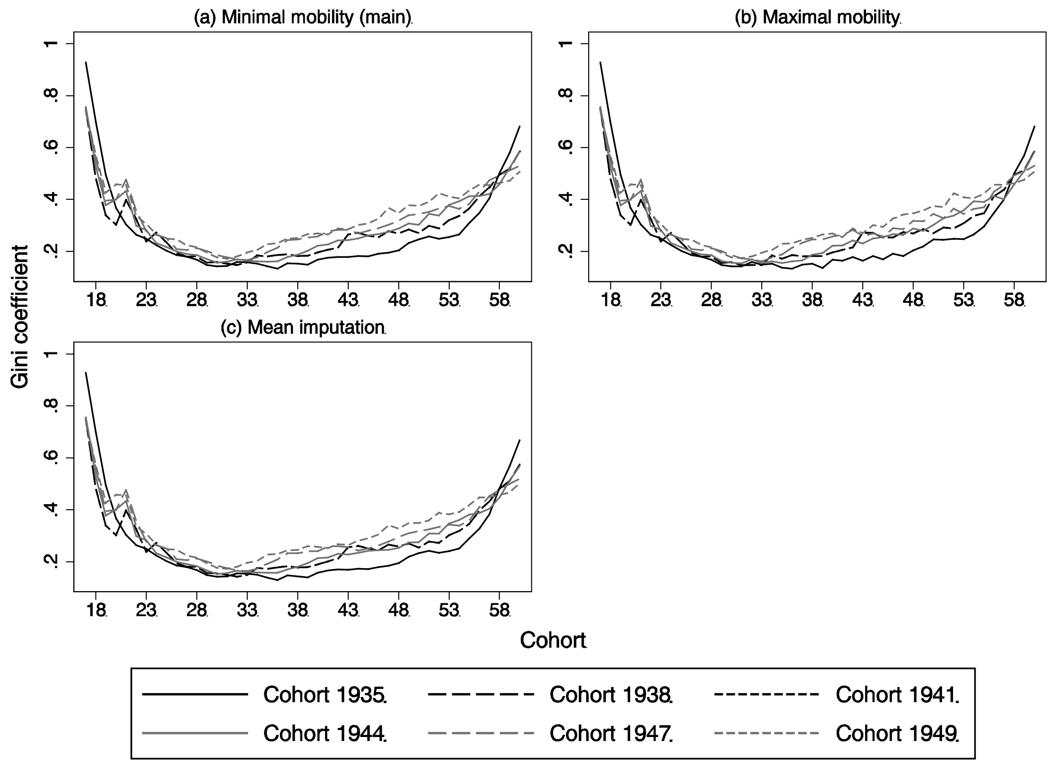


Fig. III.21.– Annual Gini coefficients from age 17 to 60 for cohorts 1935-1949, men

Note.– “real” denotes CPI discounting, “federal” denotes federal bond discounting.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

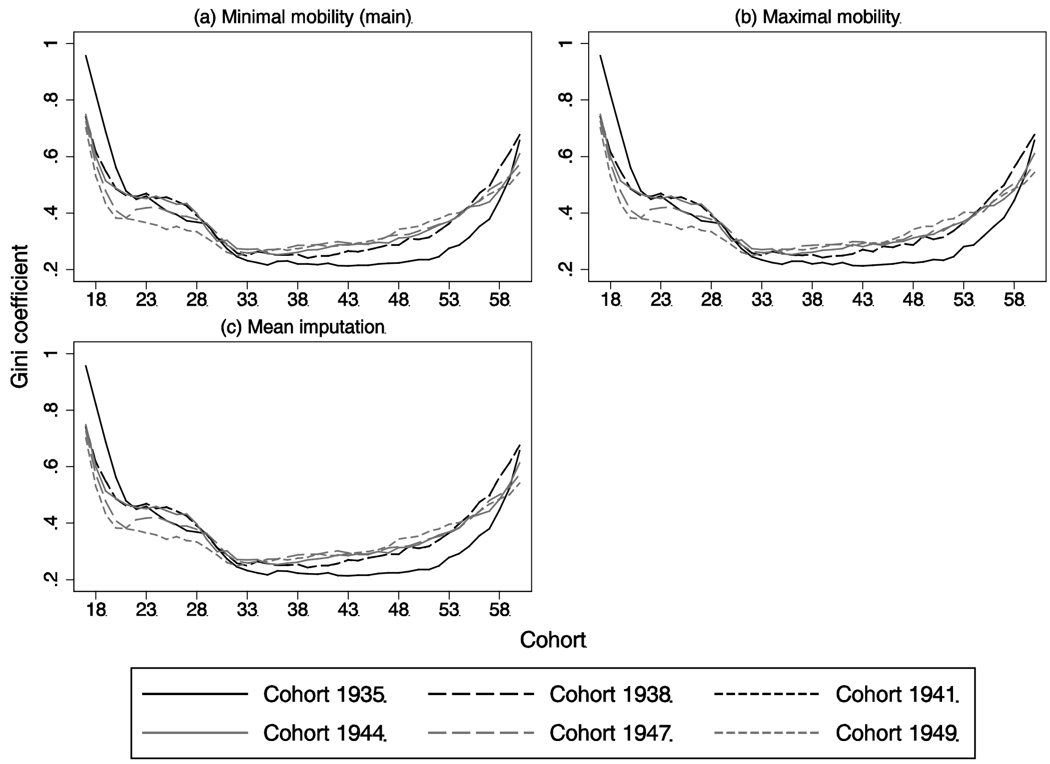


Fig. III.22.– Annual Gini coefficients from age 17 to 60 for cohorts 1935-1949, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

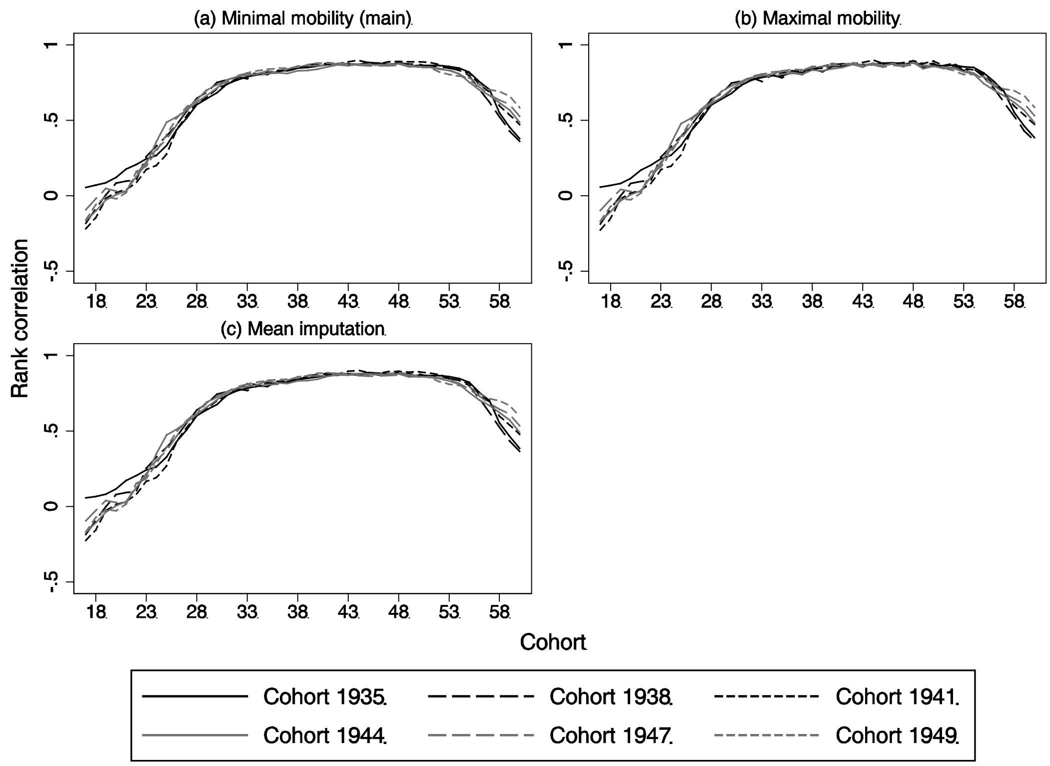


Fig. III.23.– Earnings rank correlations between consecutive years for cohorts 1935-1949, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

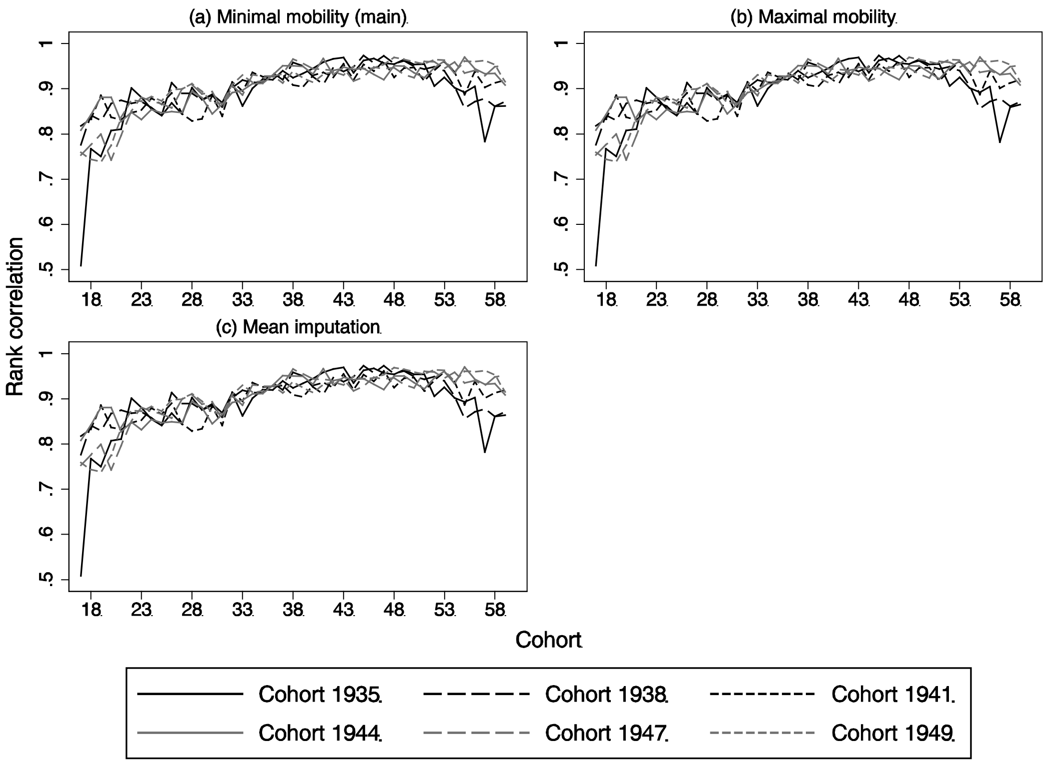


Fig. III.24.– Earnings rank correlations between consecutive years for cohorts 1935-1949, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

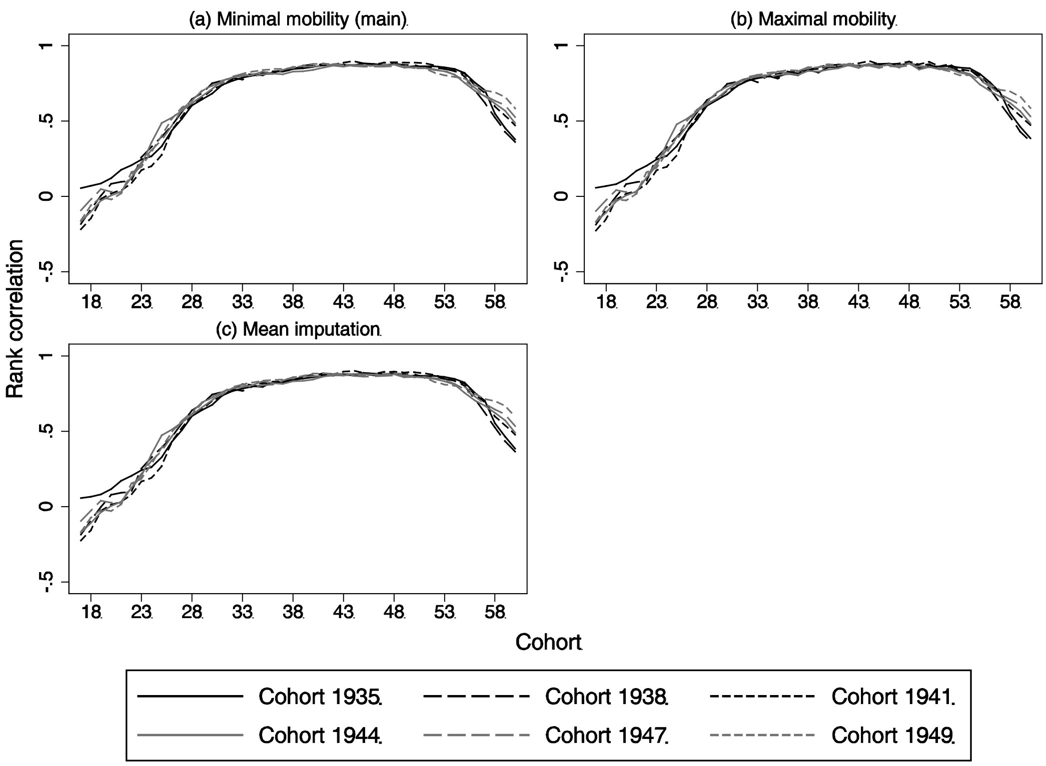


Fig. III.25.– Rank correlation of annual and lifetime earnings for cohorts 1935-1949, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

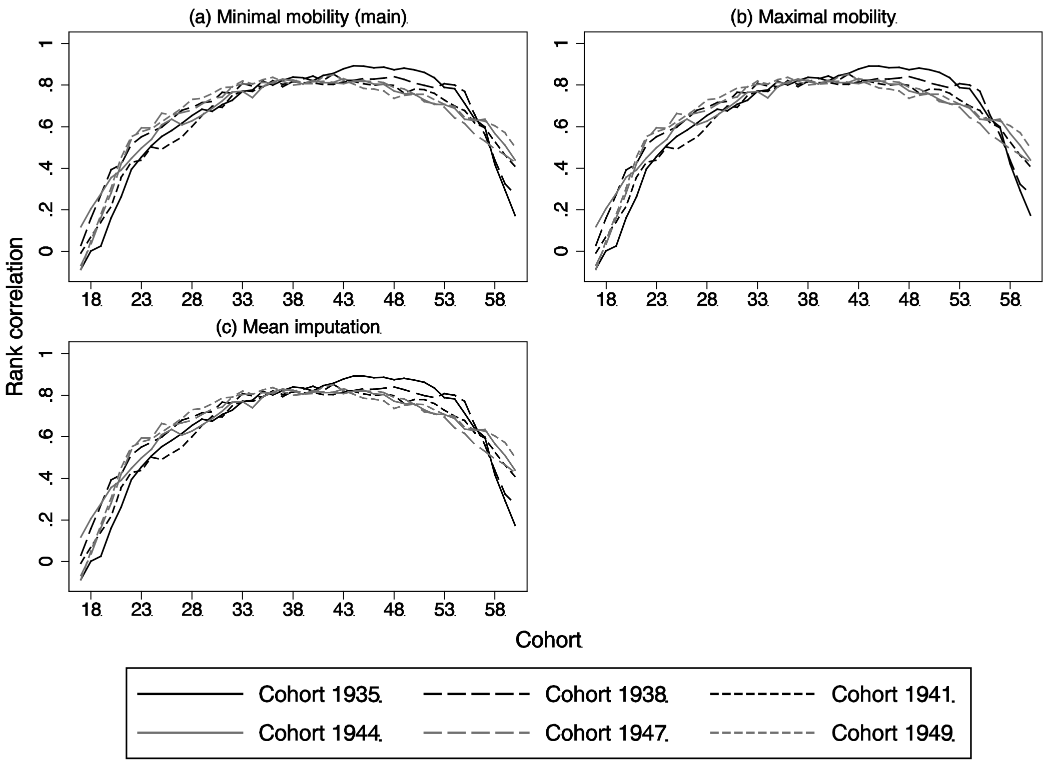


Fig. III.26.– Rank correlation of annual and lifetime earnings for cohorts 1935-1949, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

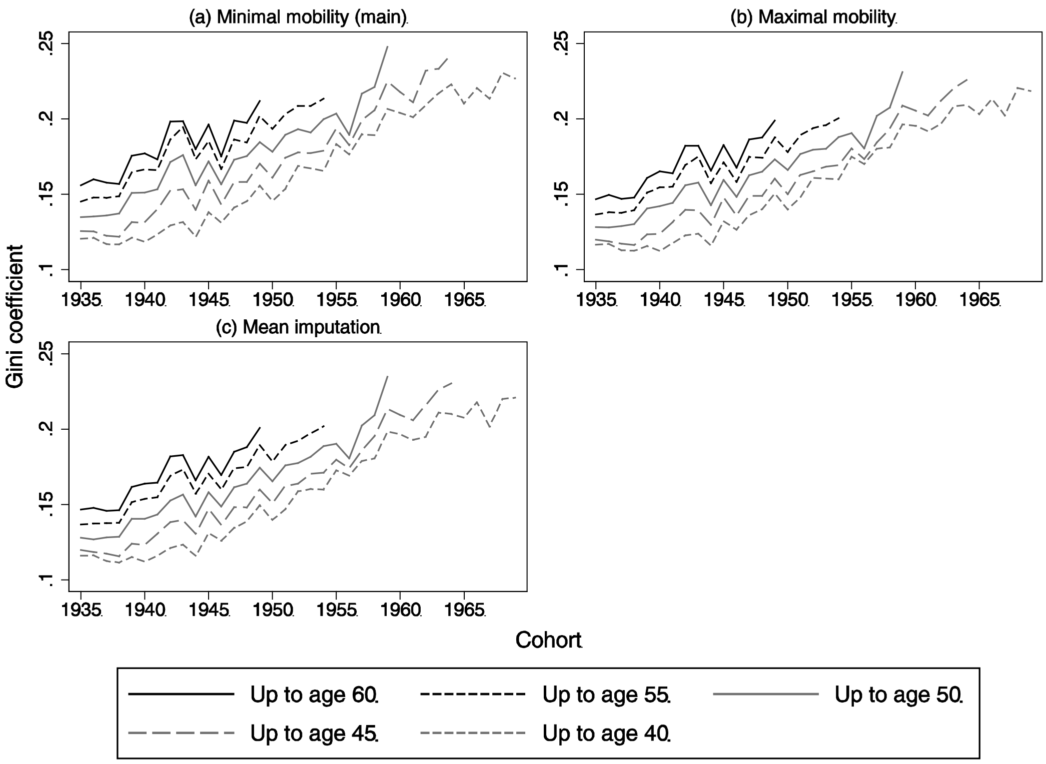


Fig. III.27.– Gini coefficients of UAX for cohorts 1935-1969, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

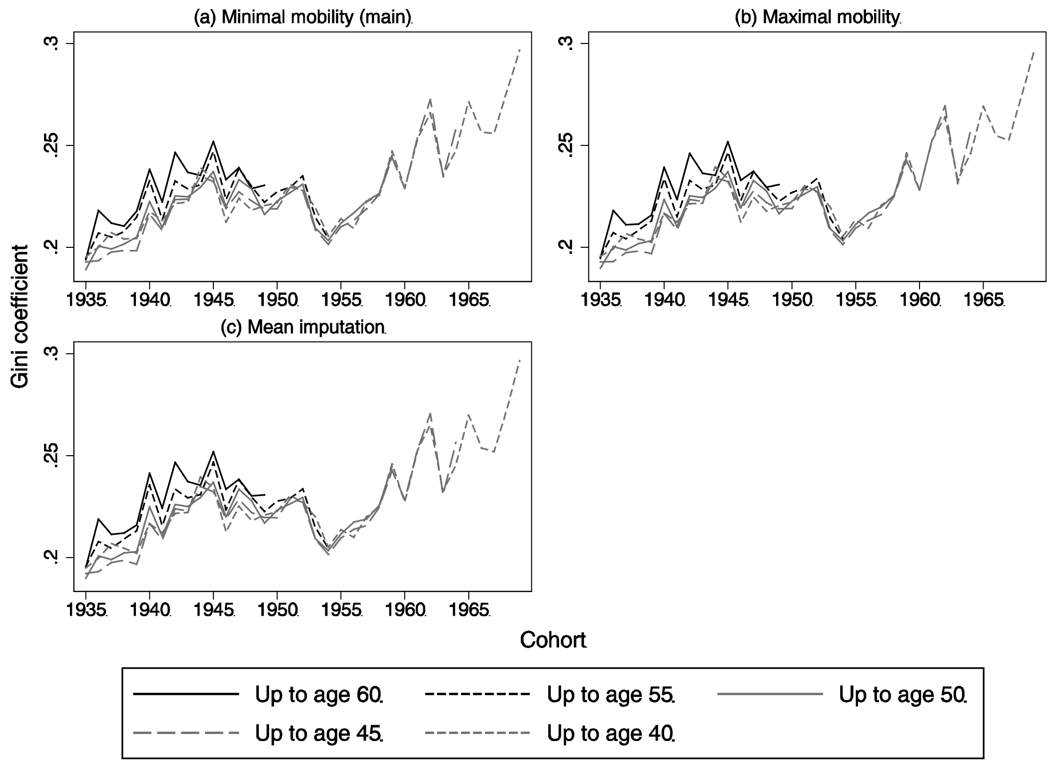


Fig. III.28.– Gini coefficients of UAX for cohorts 1935-1969, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

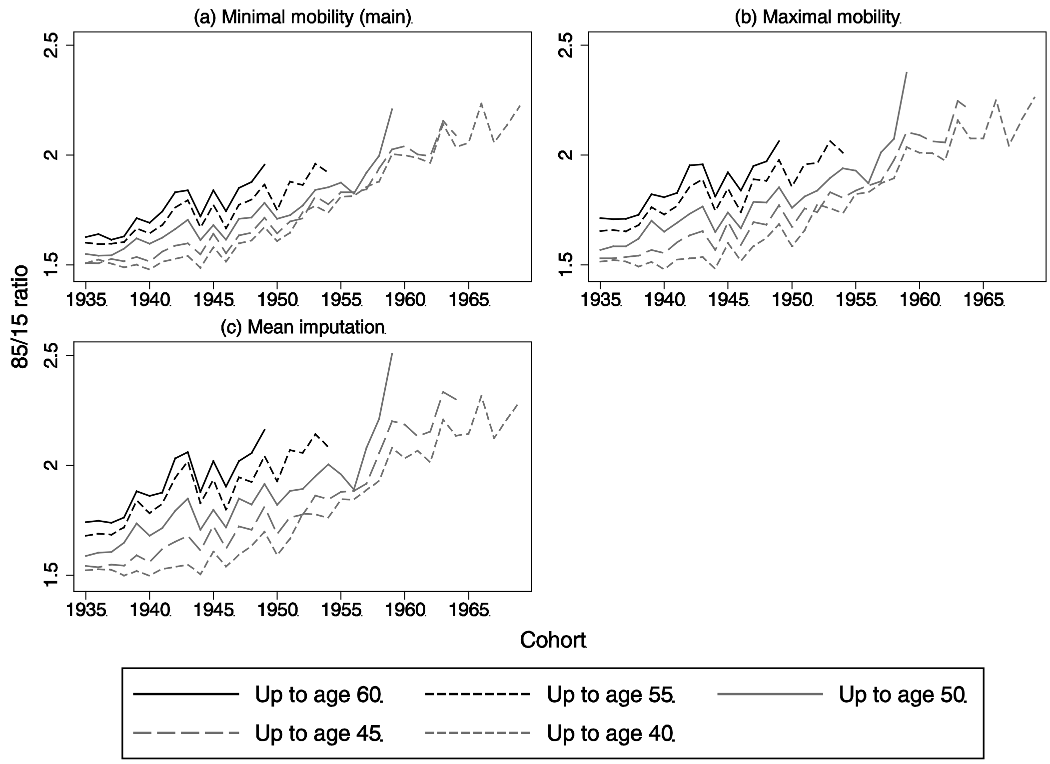


Fig. III.29.– 85^{th} / 15^{th} ratio of UAX- earnings for cohorts 1935-1969, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

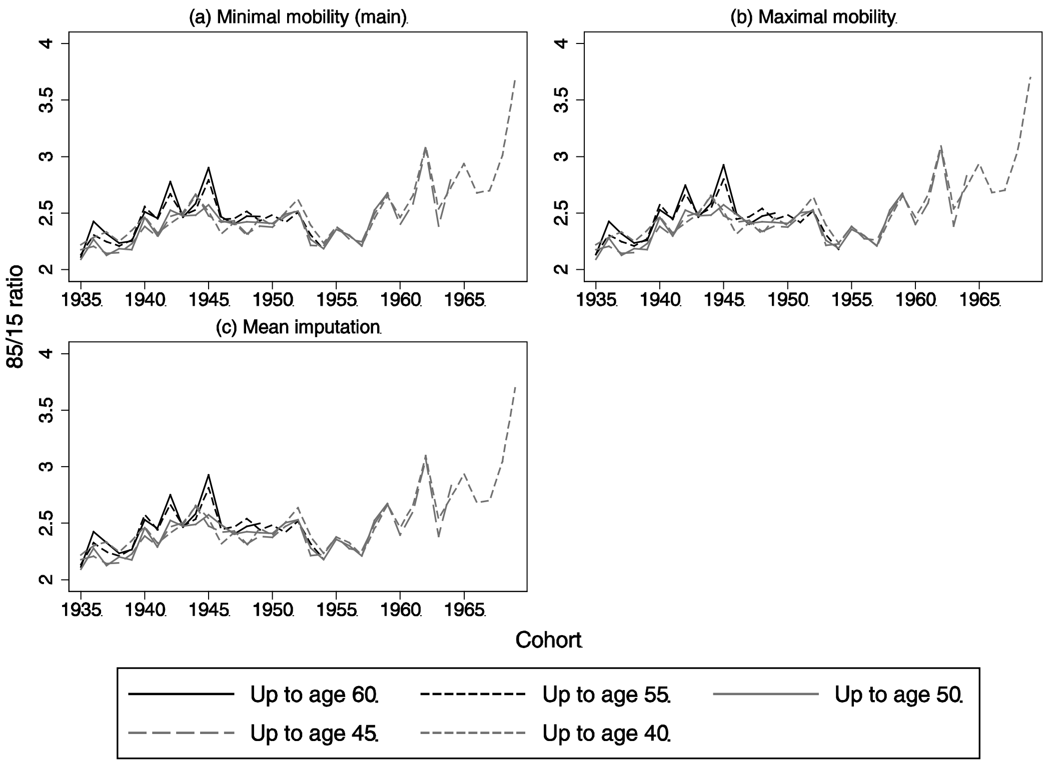


Fig. III.30.– 85th / 15th ratio of UAX- earnings for cohorts 1935-1969, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

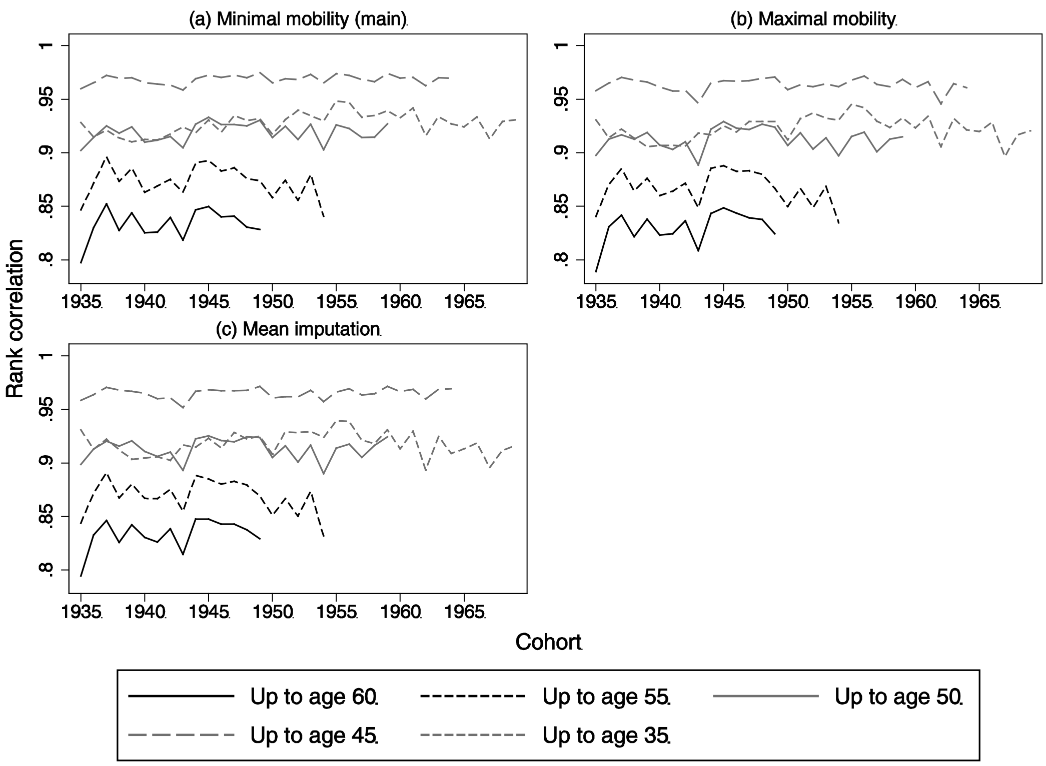


Fig. III.31.– Rank correlation of UA-40 with selected UAX, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

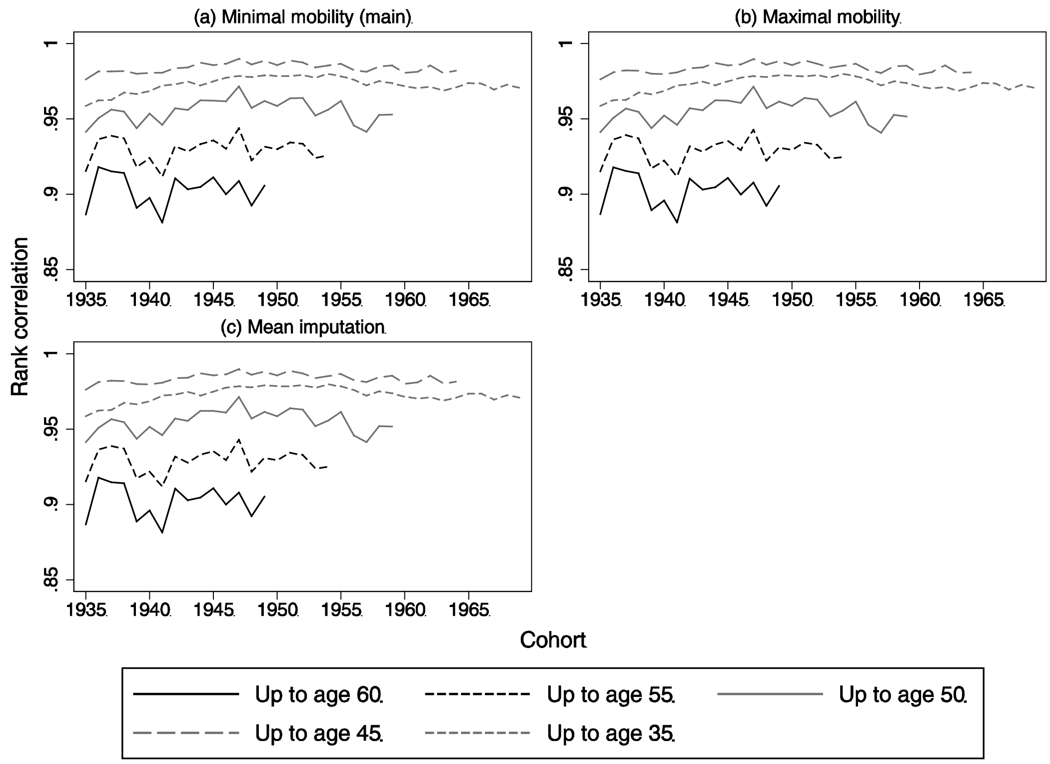


Fig. III.32.– Rank correlation of UA-40 with selected UAX, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

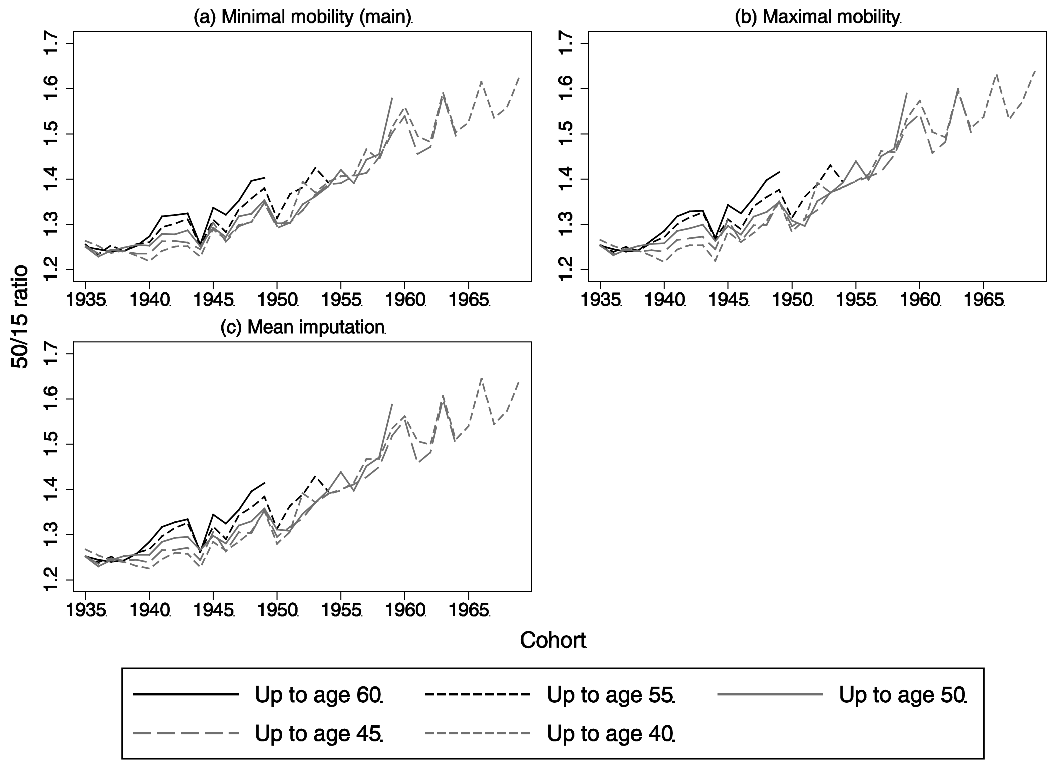


Fig. III.33.– 50^{th} / 15^{th} ratio of selected UAX for cohorts 1935-1969, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

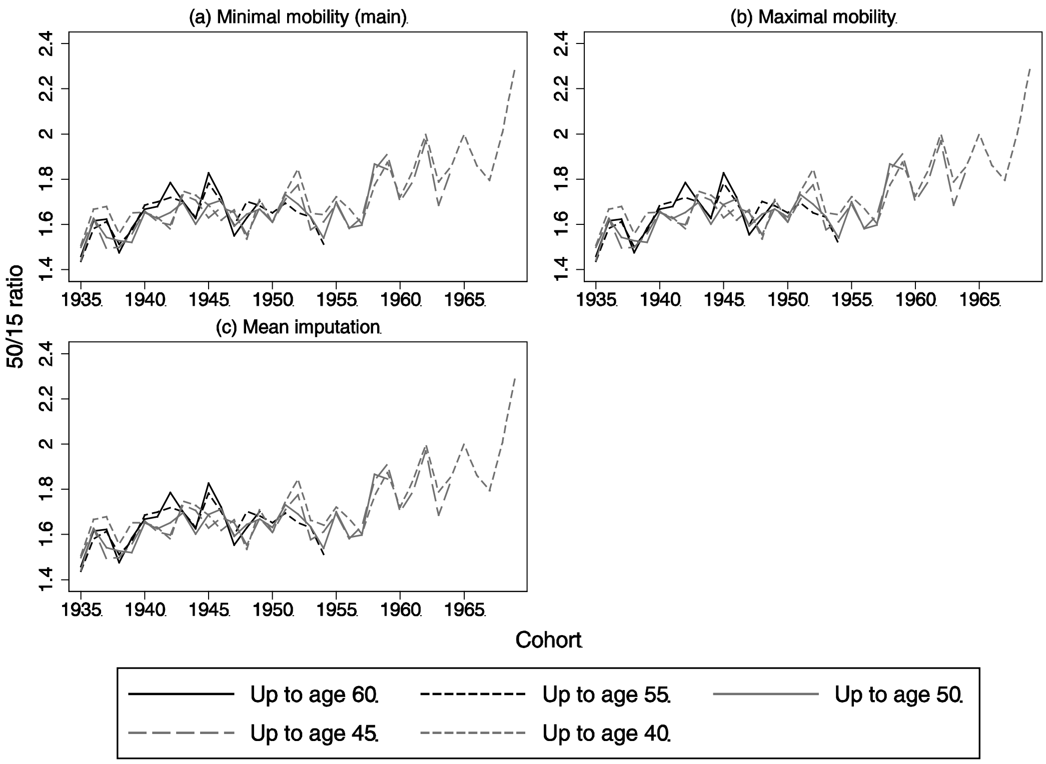


Fig. III.34.– 50th / 15th ratio of selected UAX for cohorts 1935-1969, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

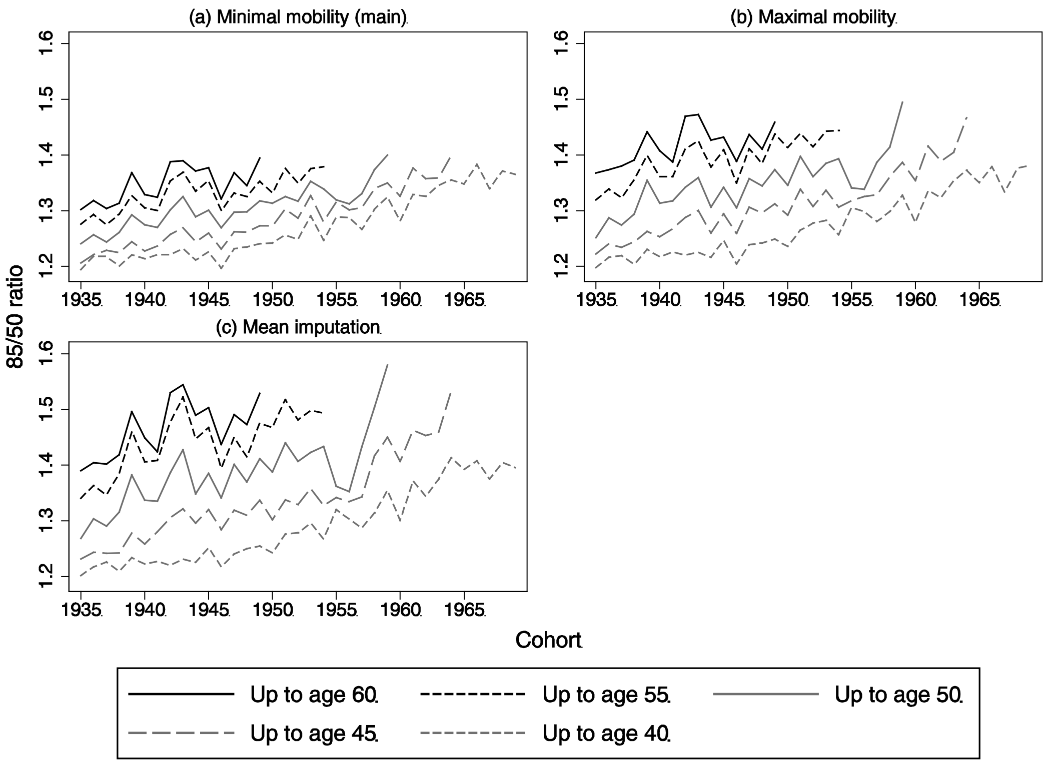


Fig. III.35.– 85th / 50th ratio of selected UAX for cohorts 1935-1969, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

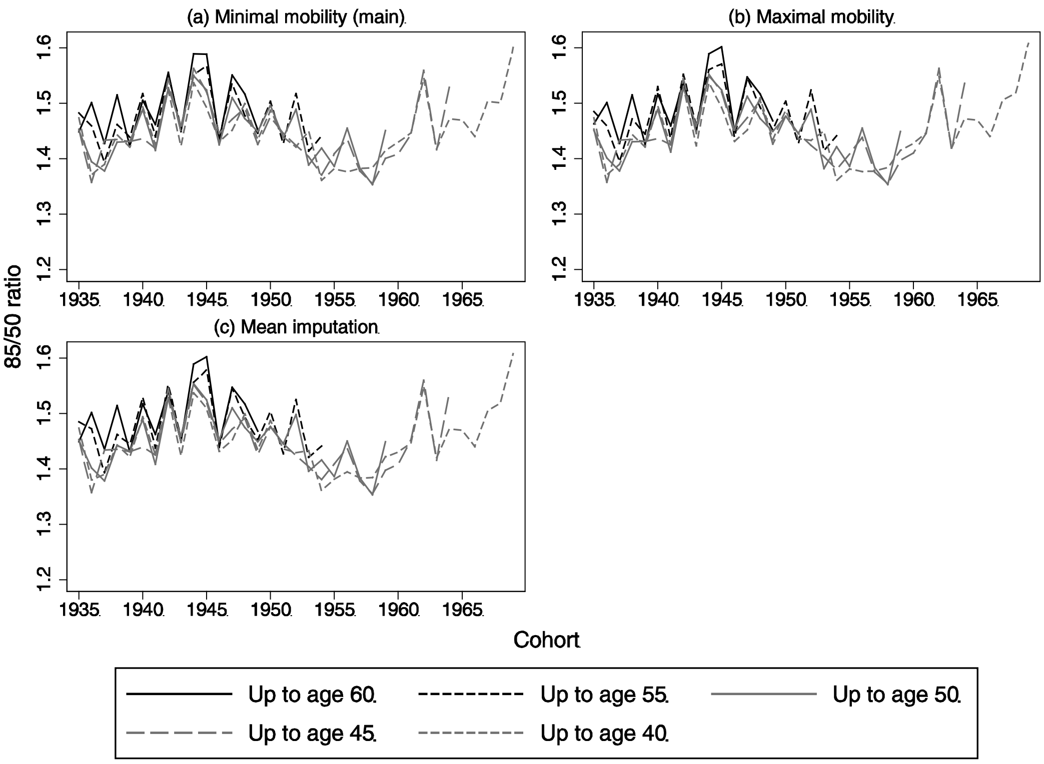


Fig. III.36.– 85th / 50th ratio of selected UAX for cohorts 1935-1969, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

III.4 Alternative discounting method – real earnings

Annotation: Each graph in this section uses real earnings instead of federal bond discounted earnings. The section mimics the relevant graphs in the paper.

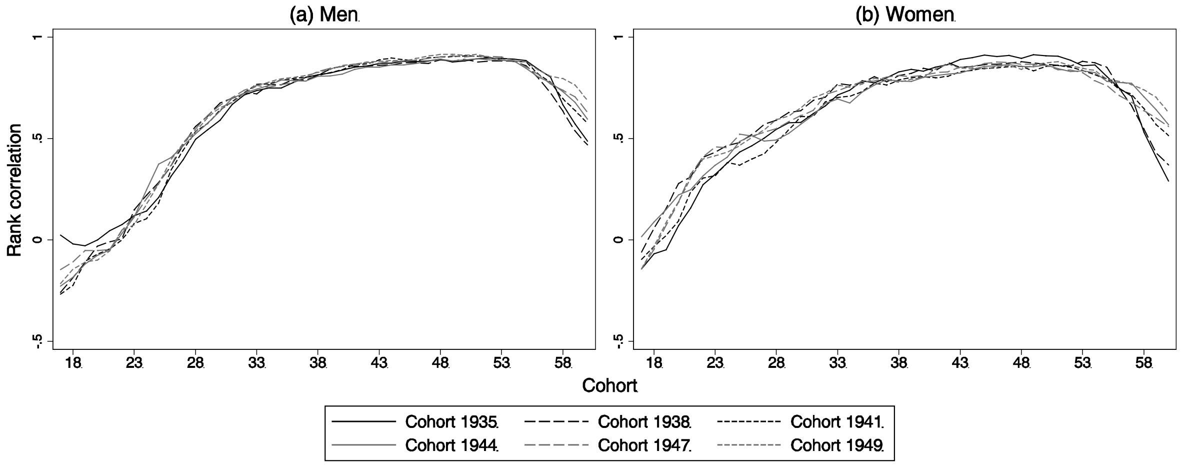


Fig. III.37.– Rank correlation of annual and lifetime earnings with for cohorts 1935-1949, men and women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

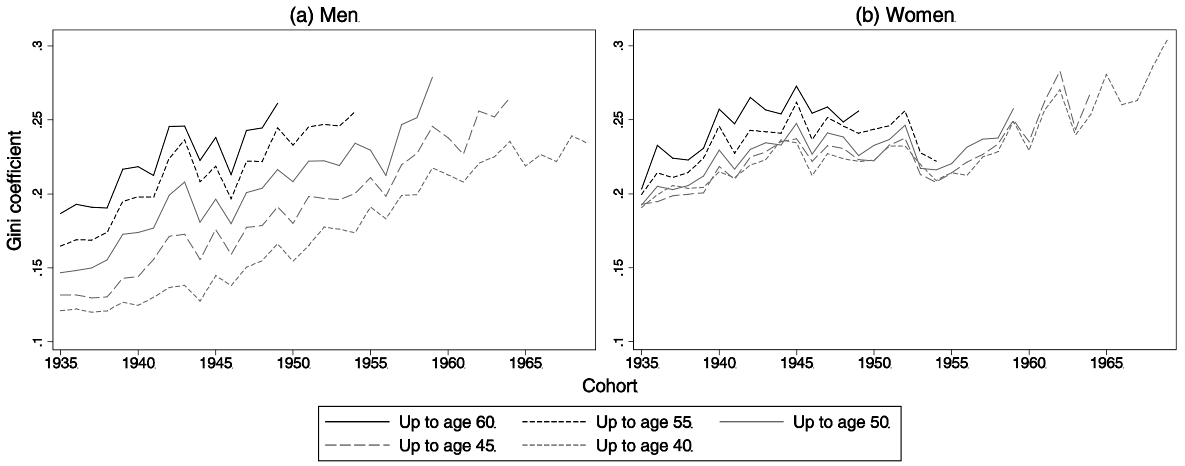


Fig. III.38.– Gini coefficients of UAX for cohorts 1935-1969, men and women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

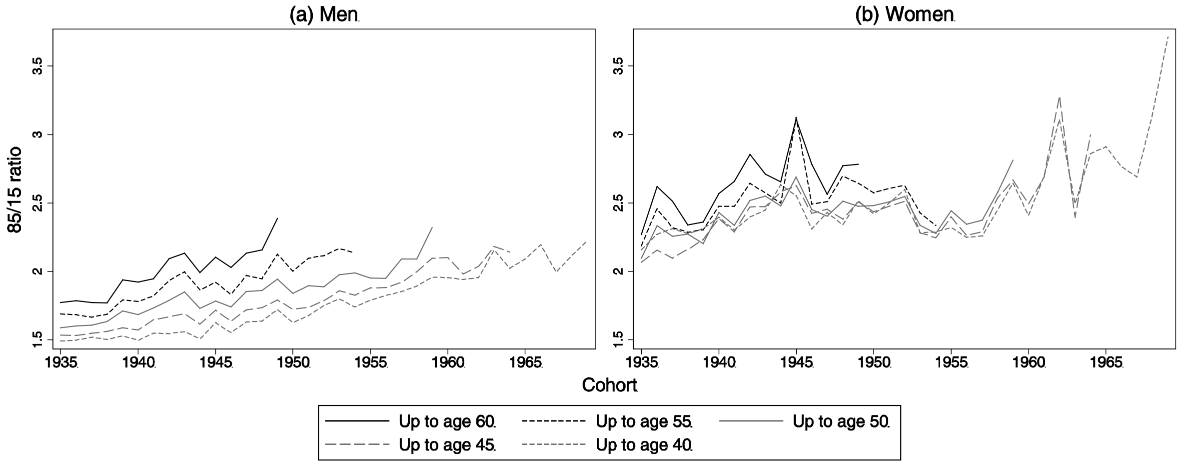


Fig. III.39.– 85th / 15th ratio of UAX- earnings for cohorts 1935-1969, men and women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

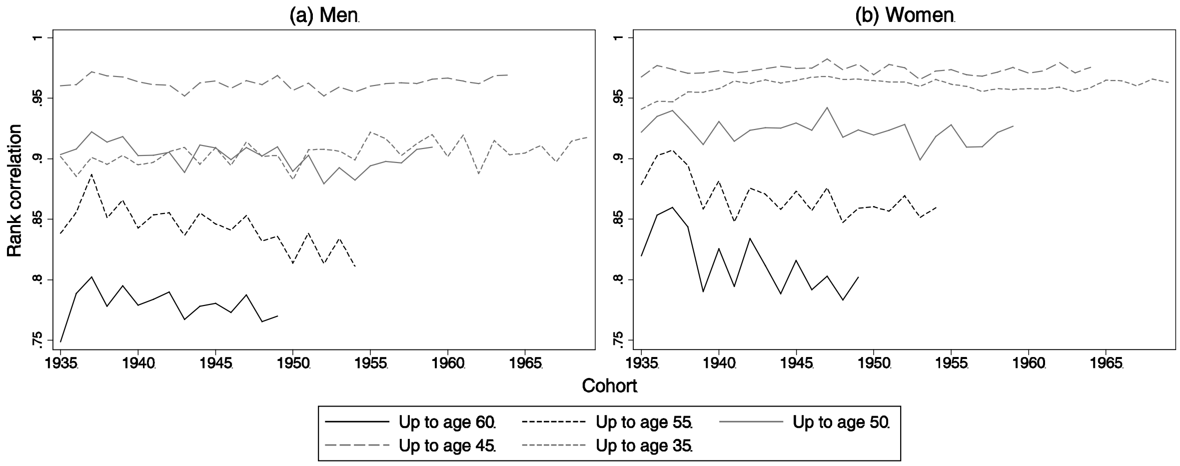


Fig. III.40.– Rank correlation of UA-40 with selected UAX, men and women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

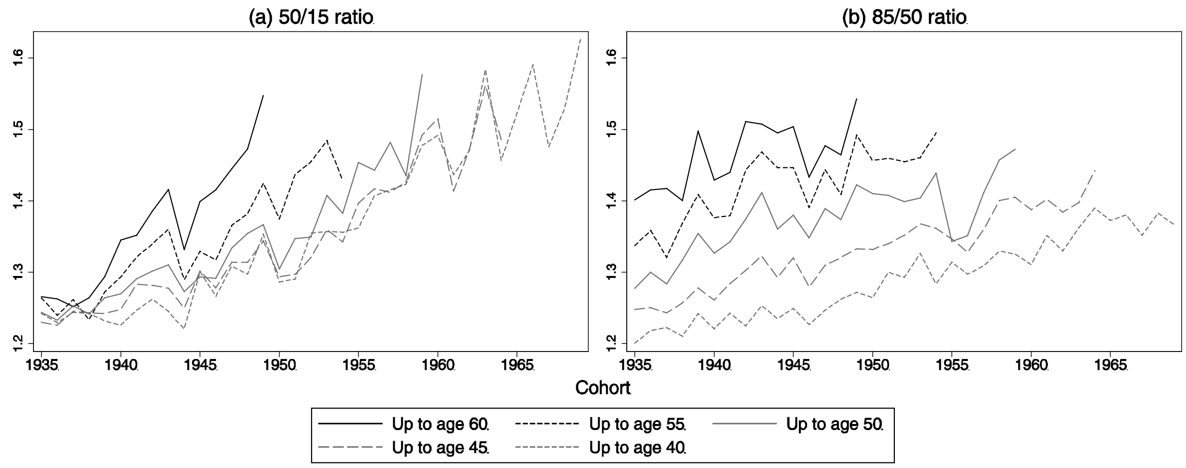


Fig. III.41.– 50th / 15th and 85th / 50th ratio of selected UAX, federal bond discounting, men cohorts 1935-1969.

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

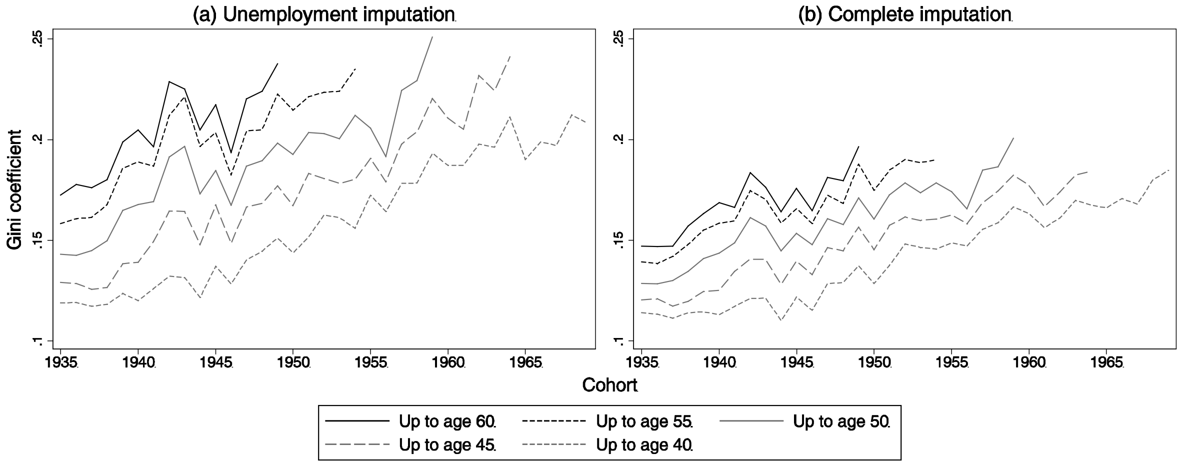


Fig. III.42.– Gini coefficients of UAX for cohorts 1935-1969 with earnings imputation if individual is not employed, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

III.5 Generalized entropy measures and further results from the Gini-decomposition

Annotation: This sections first depicts the UAX-earnings results for the GE[0], the GE[1] and the GE[2]. Then it shows further results from Gini-decompositions. All graphs in this section are based on federal bond discounting.

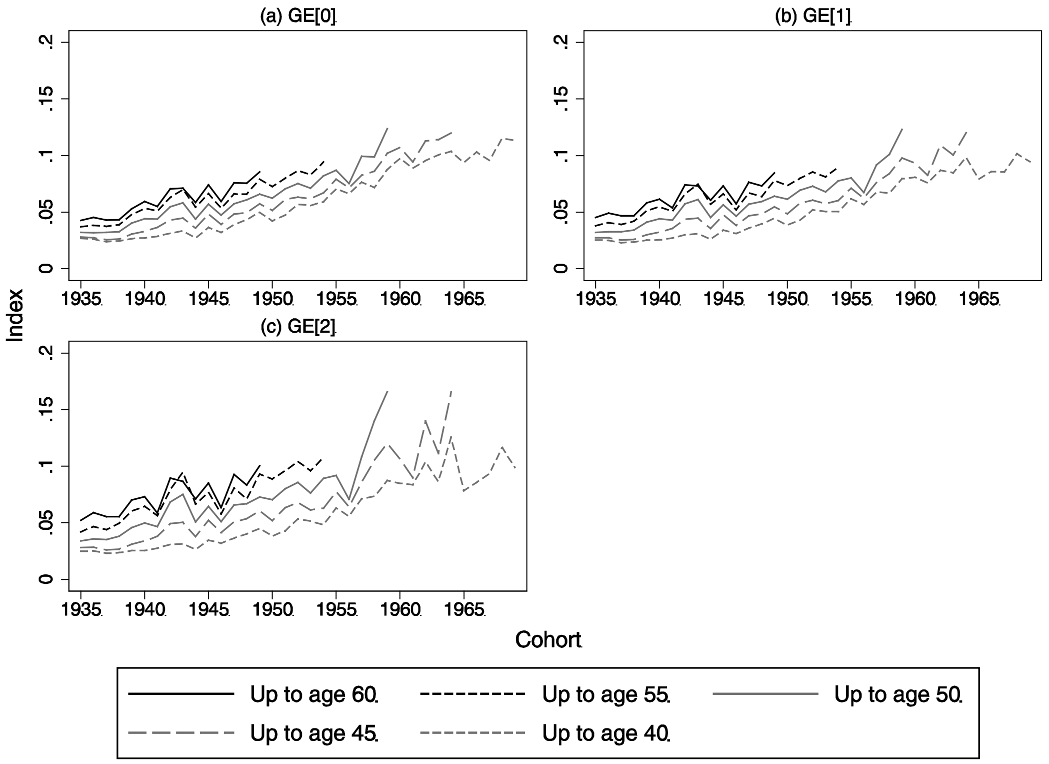


Fig. III.43.– Generalized entropy measures of UAX for cohorts 1935-1969, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

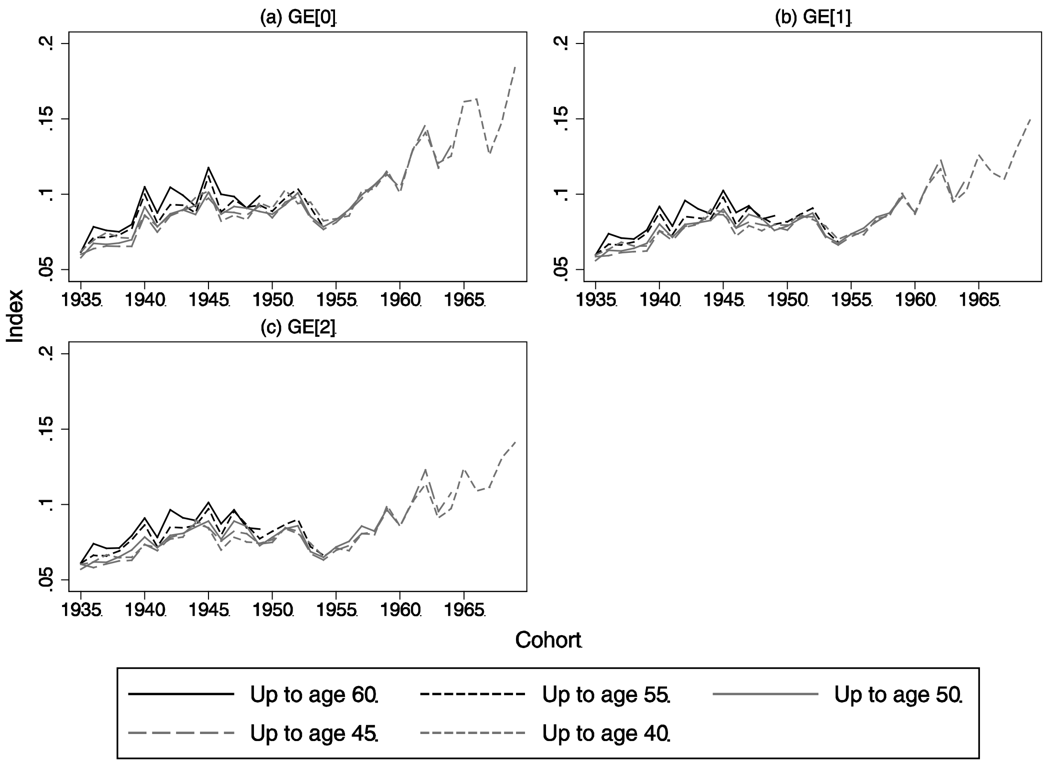


Fig. III.44.– Generalized entropy measures of UAX for cohorts 1935-1969, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

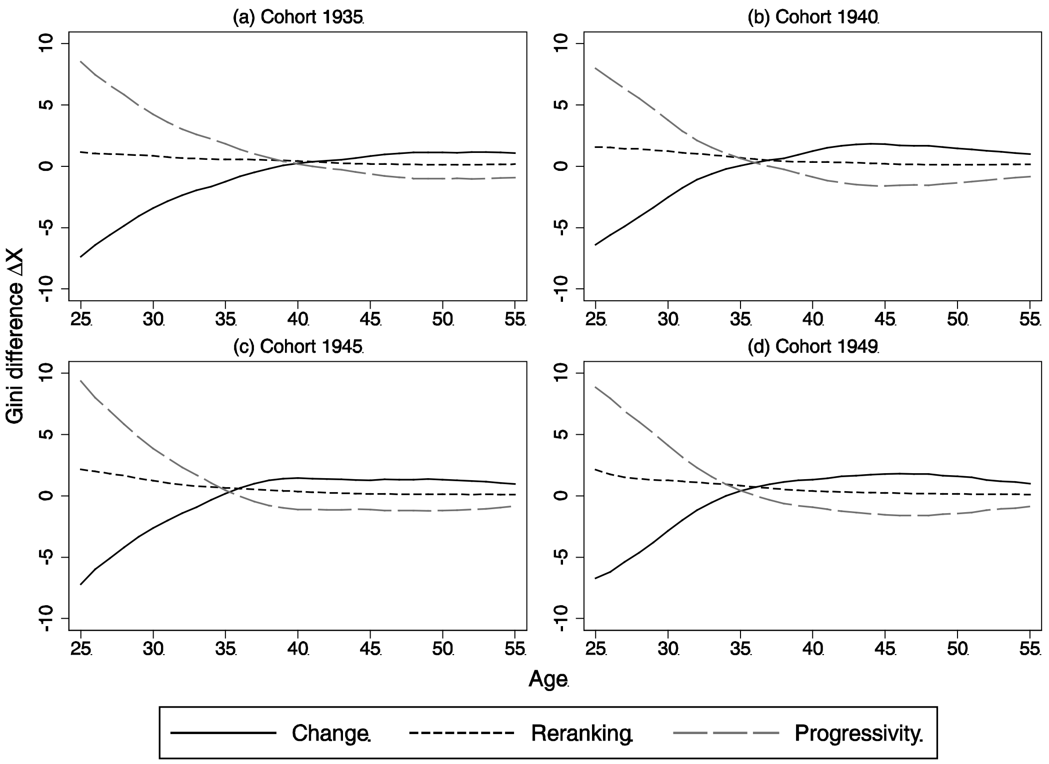


Fig. III.45.– Decomposition of changes in inequality as of Eq. (3) in the paper for selected cohorts, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

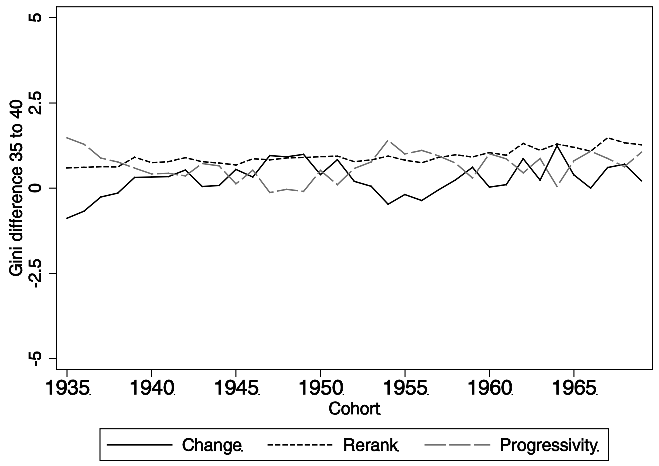


Fig. III.46.– Decomposition of changes in inequality as of Eq. (3) in the paper for changes of lifetime earnings from age 35 to age 40 for cohorts 1935-1969, men

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

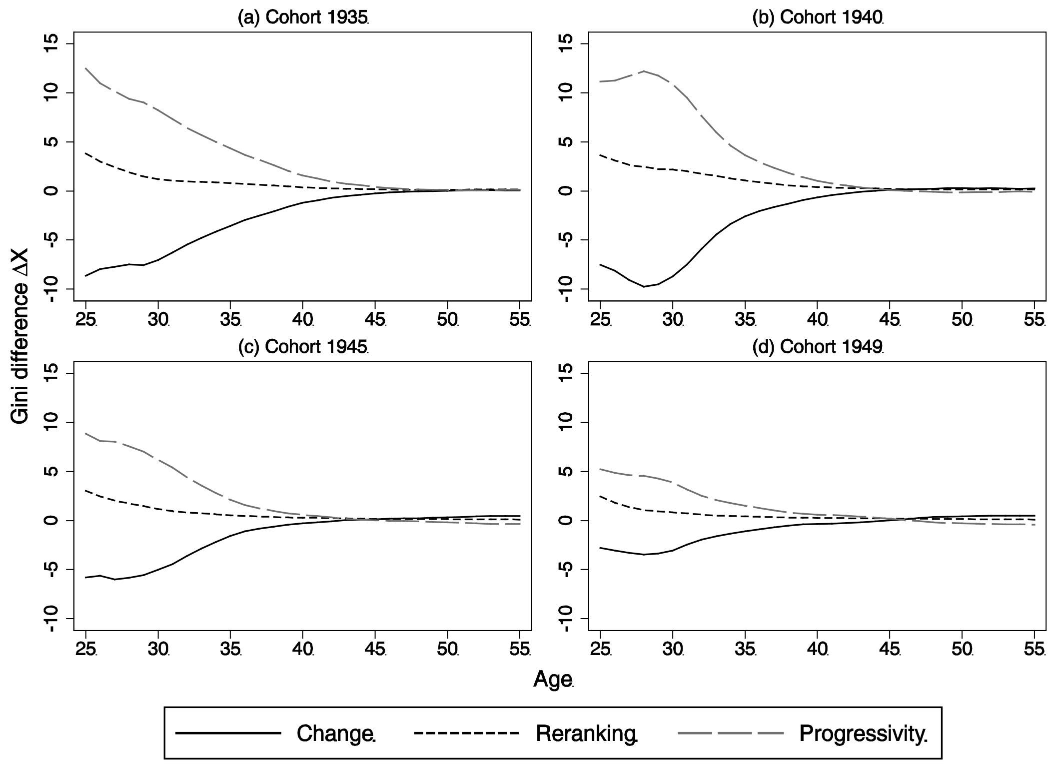


Fig. III.47.– Decomposition of changes in inequality as of Eq. (3) in the paper for selected cohorts, women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

III.6 NPV after 25

All graphs in this section are based on federal bond discounting.

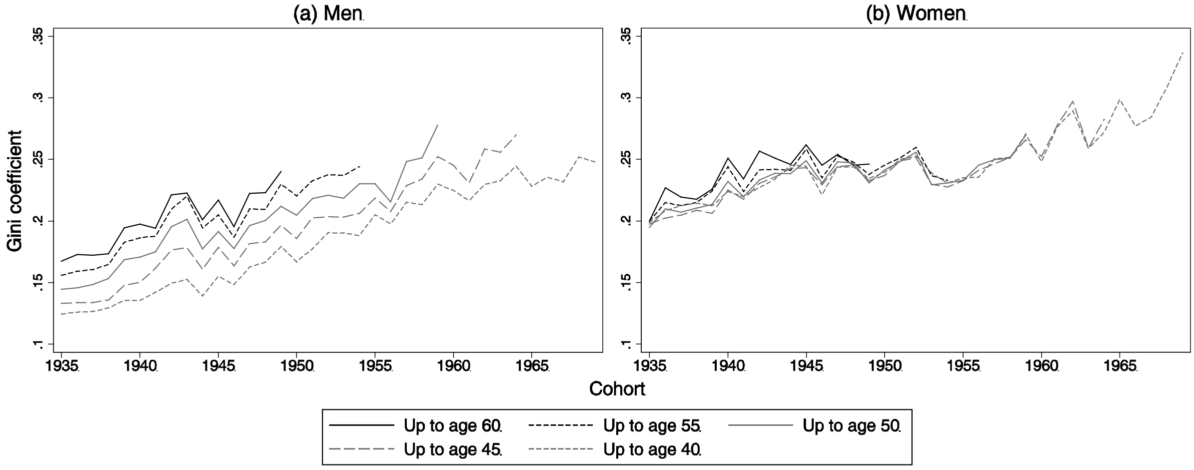


Fig. III.48.– Gini coefficients of UAX for cohorts 1935-1969, men and women

Note.– The NPV is based on annual earnings from age 25 to age *X* instead of age 17 to age *X*.
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

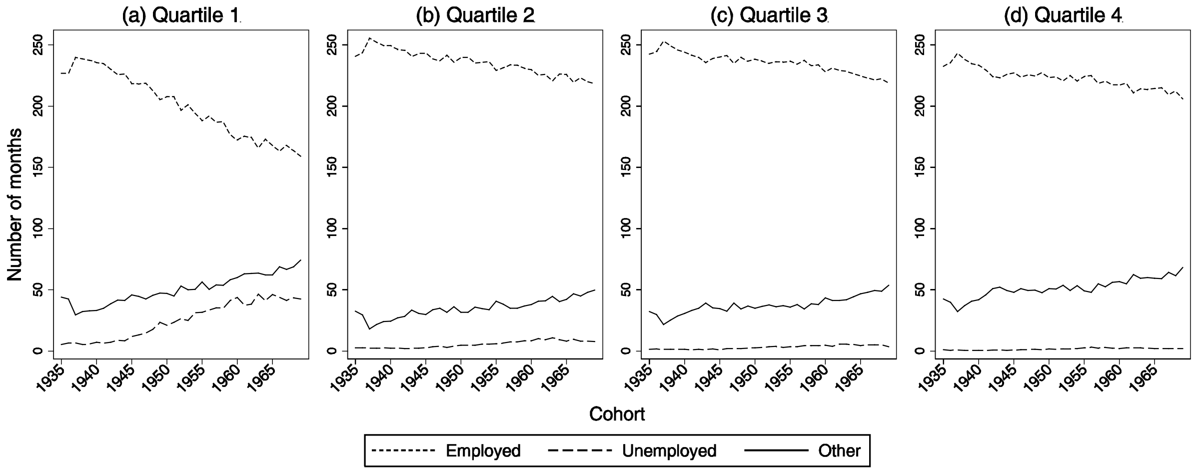


Fig. III.49.– Months of employment status up to age forty by quartile of UA-40 for cohorts 1935-1969, men

Note.– Earnings quartiles based on UA-40 with federal bond discounting. The NPV is based on annual earnings from age 25 to age *X* instead of age 17 to age *X.*
Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

III.7 Marginal employment "minijobs"

All graphs in this section are based on federal bond discounting.

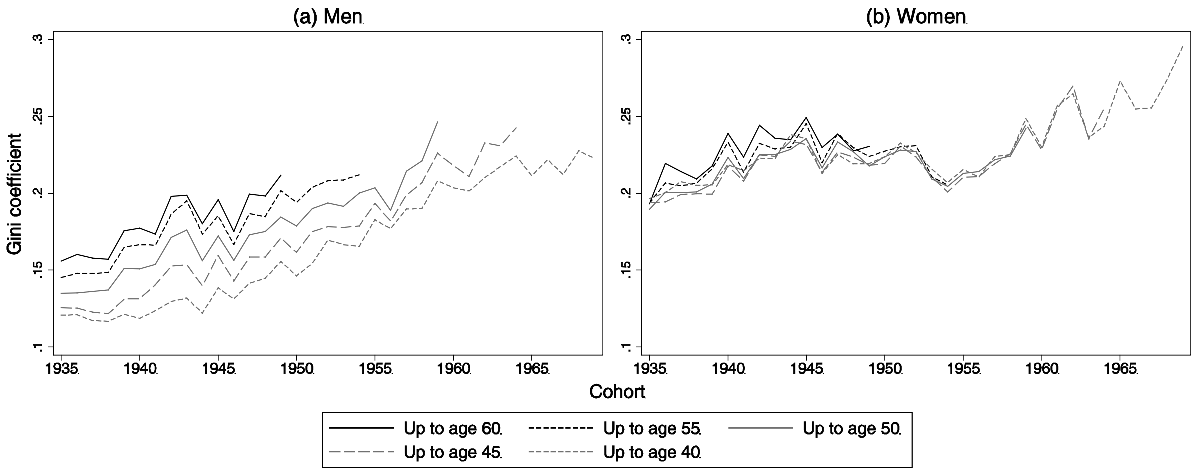


Fig. III.50.– Gini coefficients of UAX with earnings from marginal employment set to zero for cohorts 1935-1969, men and women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

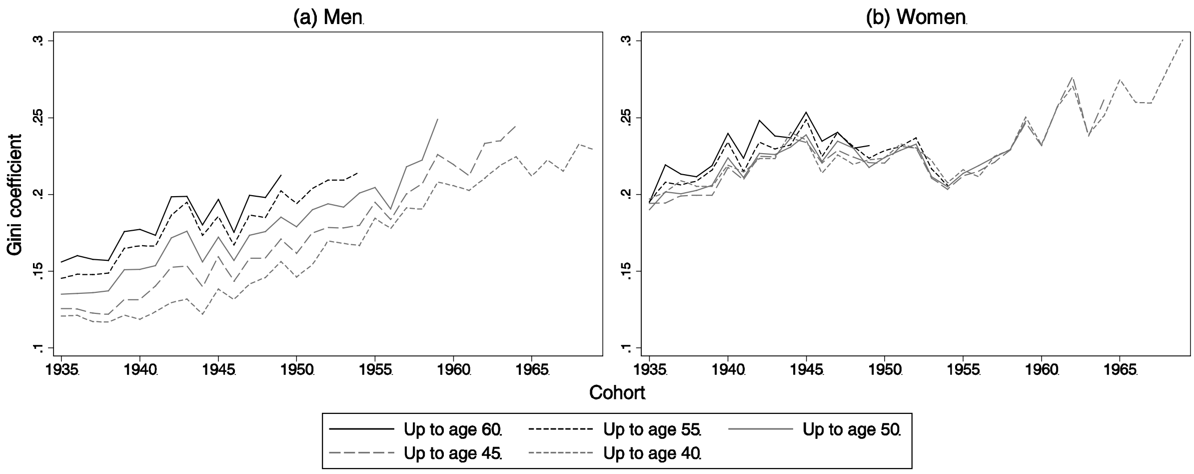


Fig. III.51.– Gini coefficients of UAX with earnings from marginal employment set to missing for cohorts 1935-1969, men and women

Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

Appendix IV: Decomposing the rise of inequality

We consider two cohorts, "parents" and "children", and the distribution of their UAX at the same $\boldsymbol{X}$, e.g. when the cohorts are forty-five. Let $\boldsymbol{a}$ denote the actually measured Gini coefficient of the UAX-distribution of parents and let $\boldsymbol{A}$ denote that coefficient for the children. The difference

$$\boldsymbol{A}-\boldsymbol{a} \qquad \text{(IV.1)}$$

is to be decomposed into two parts: the increase of inequality due to the rise of unemployment and the increase due to changes in the wage structure.

Let's denote by $\boldsymbol{h}$ the hypothetical Gini coefficient of the UAX-distribution of parents had they lived under full employment. Empirically, we obtain $\boldsymbol{h}$ by imputing earnings to parents for the few months when they were unemployed. The ratio of $\boldsymbol{a}$ to $\boldsymbol{h}$ captures the inequality increase due to "quasi-full-employment" instead of full employment.

In order to gauge the effect on the inequality increase $\boldsymbol{A}$ – $\boldsymbol{a}$ due to the rise of unemployment, we have to estimate the hypothetical Gini coefficient of the UAX-distribution of children in case they had lived under the same situation of "quasi-full-employment" as their parents. We do this by computing the hypothetical Gini coefficient of the UAX-distribution of children had they lived under full employment and by assuming that the inequality-increasing effect of having "quasi-full-employment" rather than full employment is symmetric to the effect we found for the parents' cohort.

Let's denote by $\boldsymbol{H}$ the hypothetical Gini coefficient of the UAX-distribution of children had they lived under full employment. Empirically, we obtain it by imputing earnings to children for the months they were unemployed. The hypothetical Gini coefficient of the UAX-distribution had they lived under the same "quasi-full-employment" as their parents is:

$$(\boldsymbol{a}/\boldsymbol{h})\cdot\boldsymbol{H}. \qquad \text{(IV.2)}$$

The difference $(\boldsymbol{a}/\boldsymbol{h})\boldsymbol{H}$ – $\boldsymbol{a}$ is thus the increase in UAX inequality between the two cohorts that we attribute to changes in the wage structure while $\boldsymbol{A}$ – $(\boldsymbol{a}/\boldsymbol{h})\boldsymbol{H}$ is the increase we attribute to the rise of unemployment.

In the main text of the paper we refer to the shares of the inequality increase that can be attributed to the two factors. The share due to changes in the wage structure is:

$$\frac{(\boldsymbol{a}/\boldsymbol{h})\boldsymbol{H}-\boldsymbol{a}}{\boldsymbol{A}-\boldsymbol{a}}=\frac{(\boldsymbol{H}-\boldsymbol{h})/\boldsymbol{h}}{(\boldsymbol{A}-\boldsymbol{a})/\boldsymbol{a}}. \qquad \text{(IV.3)}$$

This is the formula we use in the main text of our paper: growth rate of the Gini in the hypothetical situation of full employment divided by the growth rate of the actually measured Gini.

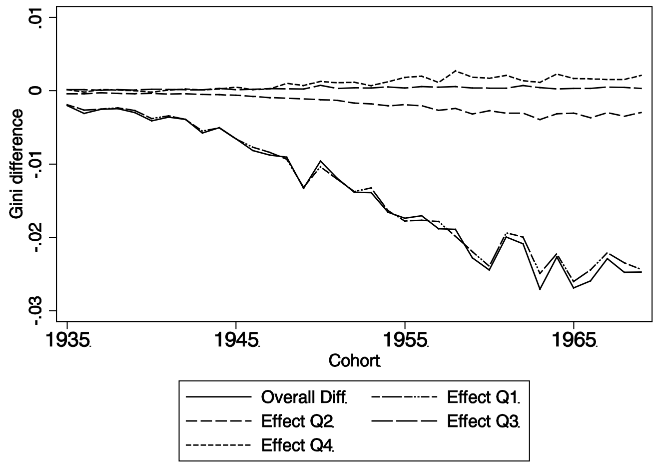


Fig. IV.1.– Effect of the imputation of earnings for times of unemployment on total inequality for cohorts 1935-1969, men

Note.– Earnings quartiles based on up-to-age 40 earnings with federal bond discounting without imputation of earnings for times of unemployment. Gini difference denotes the difference compared to the Gini coefficient of UA-40 without imputation of times of unemployment if times of unemployment are imputed across all quartiles ("Overall Diff") or for each quartile separately while leaving the other quartiles unchanged ("Effect Q1", "Effect Q2", "Effect Q3", "Effect Q4"). Source.– FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations using weighted data.

Figure VI.1 decomposes the effect on the Gini of UA-40 if earnings are imputed for times of unemployment. To measure the influence of the imputation on total inequality reduction by quartile, first the total effect of the imputation is calculated (also see Figure 13 in Section 5) and the difference to the actually observed UA-40 Gini (solid black line labelled "Overall Diff") is computed. Then, earnings are imputed for unemployment spells for each earnings quartile separately while leaving the UA-40 in the other three quartiles unaltered. Thereby, earnings quartiles are based on the original UA-40 distribution. The effect on overall inequality for imputing in the first quartile only is labelled "Effect Q1" and so forth for quartile two, three and four in the above figure. This exercise reveals that almost the whole inequality reduction stems from the imputation of unemployment spells in the first quartile.

Chapter 2:

The Dynamics of Earnings in Germany: Evidence from Social Security Records*

Abstract: This paper uncovers ongoing trends in idiosyncratic earnings volatility across generations by decomposing residual earnings auto-covariances into a permanent and a transitory component. We employ data on complete earnings life cycles for prime age men born 1935 through 1974 that covers earnings between 1960 and 2009. Over this period, the German labor market undergoes a heavy transformation and experiences strong deregulation, deunionization and a shift in employment from the industrial to the service sector. Our findings of increases in both components reflect the distinct phases of this transformation process. In magnitude, the transitory component increases most strongly in the early 1970s and the 1990s for young workers, whereas the permanent component displays the strongest increases for older workers in the early 1980 and the 2000s. Thus, the changes complicate the labor market entry for young workers while widening wage differences for established workers.

Keywords: Earnings dynamics, Life cycle, Earnings distribution, Inequality, Earnings volatility

JEL Classification: D31, D33, H24.

*This chapter is joined work with Timm Bönke and Matthias Giesecke.

1 Introduction

Labor markets and their earnings structure are continuously subject to profound changes. Examples are globalization, skill biased technological change, demographic trends, booms and recessions – frequently followed by adjustments of (labor market related) institutions. All of these are discussed extensively in the literature, impacting labor market earnings and their volatility over the life cycle, altering idiosyncratic earnings risks and earnings levels associated with labor market experience, age, cohort or skill set. In an economic environment characterized by incomplete insurance, a thorough analysis of these earnings dynamics and earnings risks over the life cycle is linked not only to individual financial decisions like wealth accumulation (Hugget, 1996; Castaneda et al., 2003), but also to lifetime earnings inequality (Bönke et al., 2015) and consumption capabilities (Gourinchas and Parker, 2002; Guvenen, 2007). It is also connected to the welfare costs related to earnings fluctuations (Storesletten et al., 2001; Blundell and Preston, 2008), and how insurance through welfare states is able to enhance overall welfare by mitigating these earnings risks efficiently (e.g. Blundell et al., 2014). For these issues, a deep understanding of the (changing) nature of labor market outcomes and of the persistence and variance of labor market shocks is needed.

This paper sheds light on the age related patterns of idiosyncratic earnings volatility over complete life cycles for West German males born between 1935 and 1974 from 1960 through 2009. Considering consistency and comparability, we focus on the main employment phase between 25 and 59. The period extends from the German "Wirtschaftswunder"-era up to the post-unification downturns that coined Germany the *sick man of Europe* (e. g. Economist, 2004). The long time frame offers unique possibilities to analyze cohorts' earnings dynamics against the background of varying economic circumstances and institutional changes like labor market deregulation, deunionization and a shift in employment from the industrial to the service sector.

To analyze earnings dynamics over extended periods, we employ a model that distinguishes between long- and short-term shocks to individual earnings trajectories. This allows disentangling earnings inequality and instability. The model relies on decomposing the auto-covariances of residual earnings into a permanent and a transitory component. Essentially, our model extends the model of Baker and Solon (2003). Our extension explicitly enables us to model the two sources of variation in earnings data (MaCurdy, 2007; Bowlus and Robin, 2012): (1) Macroeconomic dynamics relate to business cycle fluctuations, institutional changes or growth that cause changes to cross-sectional distributions over time. (2) Microeconomic dynamics define the changes of individuals' relative positions within cross-sectional distributions of successive periods. Microeconomic dynamics are modeled as follows: The permanent component considers permanent shocks to, as well as differences in, individual earnings trajectories by the inclusion of both a random walk and random growth. This captures differences in earnings levels and growth patterns due to education, effort, tenure, as well as permanent up- or downward shifts of earnings paths due to, e.g., health shocks. The transitory component is modeled as an AR(1) process with additional flexibility through a quartic age term that allows diverging shock levels by age. To correctly identify these life cycle parameters, macroeconomic dynamics are explicitly modeled as calendar time shifters for both permanent and transitory component. For an accurate identification of generational differences, the model also includes cohort shifters for both components.

As a general pattern across life cycles, we find that the permanent component steadily increases as the individual ages. The transitory component is almost u-shaped over the life cycle. In the early stages, the predominant share of earnings volatility is explained by short-term fluctuations, which typically vanish after about two years. Long-term divergences then become more relevant, surpassing the transitory component in its relative importance around age 35. This mirrors the structure of earnings trajectories, which are typically settled after age 35 in Germany (e.g. Bönke et al., 2015), and implies that shocks endured thereafter are more likely to be permanent. At the end of the life cycle, the transitory component again increases in relevance. Thus, shocks to the cohorts' earnings paths in close distance to retirement are not likely to be permanent but rather reflect an opting out of the labor market.

Comparing earnings dynamics from 1960 to 2009, our results indeed suggest a rising overall variance through an increase in both permanent and transitory component. For the transitory component, we identify that the increase started in the mid-1970s and intensified in the mid-1990s. The increase is especially pronounced for younger workers. Thus, establishing themselves on the labor market became increasingly more demanding for labor market entrants because earnings paths were interrupted more often. For the permanent component, we find strong increases since the 1980s that amplify in the early 2000s. In terms of magnitude, the permanent component increases more strongly for workers well established in their careers. Hence, persistent differences, such as education, entail a lower earnings path for low skilled workers and a higher one for highly skilled workers (Blundell et al., 2014). Further, the increasing importance of permanent shocks indicates that it becomes more difficult for individuals to reestablish themselves on the labor market after large shocks like health shocks or involuntary job loss both over the life cycle and across generations. The findings relate well to the overall developments on the German labor market.

Our paper relates particularly to three strands of the literature. First, we relate to studies on inequality in Germany, which typically document increasing cross-sectional and lifetime earnings. Reasons are, e. g., overall wage dispersion, increasing plant level heterogeneity (Card et al., 2013), deunionization, deregulation, job polarization (Dustmann et al., 2009) or a steep decrease of employment spells (Bönke et al., 2015). Our findings are consistent with these explanations, as they imply more divergent earnings paths and decreasing job stability. We complement by uncovering what part of inequality is transitory and what part is permanent at various points of the life cycle and how these patterns evolve across generations. This gives a deeper understanding of how past and current and inequality trends are composed.

Second, our study relates to papers similarly decomposing the development of earnings inequality and instability in a specific country over time, e.g. Shin and Solon (2011) and Gottschalk and Moffitt (1994; 2002; 2012) for the United States, Baker and Solon (2003) for Canada or Cappellari (2004) for Italy.[33] Similar to our results, most studies find increasing earnings volatility over time, which is to a larger extent driven by permanent inequality. Since our data allows the analysis of entire earnings life-cycles, we complement by showing how to fit variance decompositions over extensive time periods.[34] Further, we provide comprehensive results for Germany. We confirm many previously documented results and therefore validate the decomposition approach for the shorter panels used in previous studies.

Finally, we look at complete life cycles. Therefore, this paper relates to studies that contribute to the microeconomic dynamics of life cycle earnings risk[35] with the purpose of providing evidence for an improved calibration of macroeconomic models, stressing the importance of heterogeneous age-specific innovations (e.g. Guvenen, 2009; Karahan and Ozkan, 2013). While the parameters of our model can be also used for calibration, our results foremost emphasize the inclusion of cohort differences. Microeconomic dynamics of the life cycle are also analyzed with regard to education (Meghir and Pistaferri, 2004), family context (Blundell et al., 2014; Bingley et al., 2014), and shocks of higher moments across the distribution (Guvenen et al. 2014; Guvenen et al., 2015). We contribute by modeling complete life-cycles with the inclusion of macroeconomic dynamics and generational differences. While still identifying common microeconomic dynamics, we show that permanent and transitory shocks vary substantially across generations.

The remainder of this study is structured as follows: Section 2 describes key facts on the evolution of the German labor market. Section 3 provides the theoretical model on earnings dynamics, while Section 4 presents the underlying dataset, related issues and sample descriptives. Section 5 covers

[33] There are also studies on e.g. Great Britain by Dickens (2000), Luxembourg by Sologon and Van Kerm (2014), Sweden by Gustavsson (2008) or Denmark (Bingley et al., 2013). Oftentimes, subgroup developments are compared (e. g. blue vs. white collar workers, education groups, immigrants vs. natives).

[34] Most studies focus on shorter 15 to 25 year periods and none of the underlying datasets used in other studies include enough data to cover complete life cycles.

[35] These papers disregard macroeconomic dynamics and abstract from cohort and calendar time effects.

the main estimation results, discusses the implications and relates the findings to the developments on the German labor market. Section 6 concludes.

2 Macroeconomic trends and institutional changes in Germany since 1950

For classification and interpretation of empirical long run trends, this section gives a concise overview on major changes affecting the West German labor market since 1950 – supported by key indicators in Figures 1 and 2. In addition to standard indicators for overall economic performance, such as annual GDP growth and unemployment, Figure 1 provides an indicator for openness and the shares of employees by sector.[36] Openness relates to international connectedness and increasing connectedness likely threatens wages of low-skilled workers and potentially increases inequality (e. g. Krugman and Venables, 1995; Wood, 1995). The shares of employees reveal which of the three sectors employs most: industry, services or agriculture. Each sector entails distinct properties regarding e.g. remuneration rules or type of employment contracts. Therefore, shifts in sectoral importance can translate into changes in wage dispersion and job security. Figure 2 provides the ratio of union members and employees to the percentage of employees covered by sectoral contracting agreements. Sectoral contracting implies that contracts for these employees are negotiated between employer associations and trade unions on national or federal state level. Both indicators describe union power, which in turn relates to wage compression and inequality (e.g. Acemoglu et al., 2001). Figure 2 also covers indicators of labor market deregulation and shows the shares of subcontracted employees and of those with fixed term contracts.[37]

The developments on the West German labor market following World War II can be divided into three distinct phases. The first phase, the German *Wirtschaftswunder,* lasted from after World War II in the late 1940s throughout the early 1970s. After regaining some political independence from Allied Powers, the West German economy transformed rapidly and began producing consumer goods and equipment. Labor demand increased immensely through a combination of ongoing reconstruction of war damages, increasing consumer demand, as well as the relocation of firms and manufacturing bases from East Germany to the West.[38] Until around 1950, large inflows of about 8 million displaced German workers from the former eastern territories of the German Reich satisfied this demand (Bauer et al., 2013). Labor demand was then met by the westward migration of East

[36] We define openness as the combined share of imports and exports over GDP. Alternative measures of openness like foreign direct investment show similar trends.
[37] In addition, Table C.1 provides an overview on the chronology of laws regarding labor market (de)regulation since 1972.
[38] For example, Buenstorf and Guenther (2010) find that 23% of the East German machine tools industry reallocated to West Germany shortly after World War II.

Germans (until 1961) and the recruitment of guest workers (late 1950s to early 1970s).[39] Naturally, the strong labor demand and high GDP growth rates coincided with extremely low unemployment rates (Figure 1). More than half of the employees worked in the industrial sector, characterized by strong unions, high job security and a rather compressed wage distribution due to sectoral agreements (Figures 1 and 2). Figure 2 also reveals that there was one union member for every three employees[40] and that sectoral contracting covered more than four-fifths of all employees. During this period, legislators expanded the welfare state and enhanced labor contract protection (Bartels, 2014).

Between the mid-1970s and German reunification in 1990, this successful system started to dissolve. Global developments gained influence and increasingly affected the interconnected German economy (Figure 1), while competitiveness became a growing issue. The first oil price shock in 1973 caused a recession with unemployment rates tripling, reaching 5%. The share of employment in the manufacturing sector started declining steadily while that of the service sector grew continuously; employment trends that continue to this day (Figure 1).[41] While unions remained strong, legislators slightly deregulated the labor market and introduced subcontracted work in 1972 to increase flexibility (Table C.1 and Figure 2). After the second oil price shock in 1979/80, another major recession hit Germany, causing unemployment to rise to more than 9% (Figure 1). Legislators considered labor market rigidity to be a key problem and lowered employment protection, expanded possibilities for subcontracted work, and introduced fixed term contracts (Table C.1). At the same time, the ratio of union members to employees declined, while sectoral coverage remained about constant (Figure 2).

After a short lived boom following reunification in the early 1990s, a subsequent recession marked another turning point for the German labor market. Already experiencing mass unemployment, growing competition from the former socialist European countries put additional pressure especially on low skilled individuals. Further, the West German labor market was the target of migration for

[39] Until 1950, the labor force grew mainly due to forced migration of Germans from Eastern Europe following the conclusion of World War II. The bulk of the displaced originated from the former eastern territories of the German Reich (Pomerania, Prussia, Silesia). The inflow of migrants from the German Democratic Republic numbers about 2.6 million and stopped with the closing of the inner German border, best symbolized with the Berlin Wall in 1961. In the late 1950s, the West German government started a large scale recruitment of guest workers due to a shortage in low-skilled labor (Bauer et al., 2005). This active manpower recruitment included treaties with several countries, most notably Italy (1955), Spain and Greece (1961), Turkey (1961) and former Yugoslavia (1968). For a detailed description of the recruitment procedure, see Bauer et al. (2005).

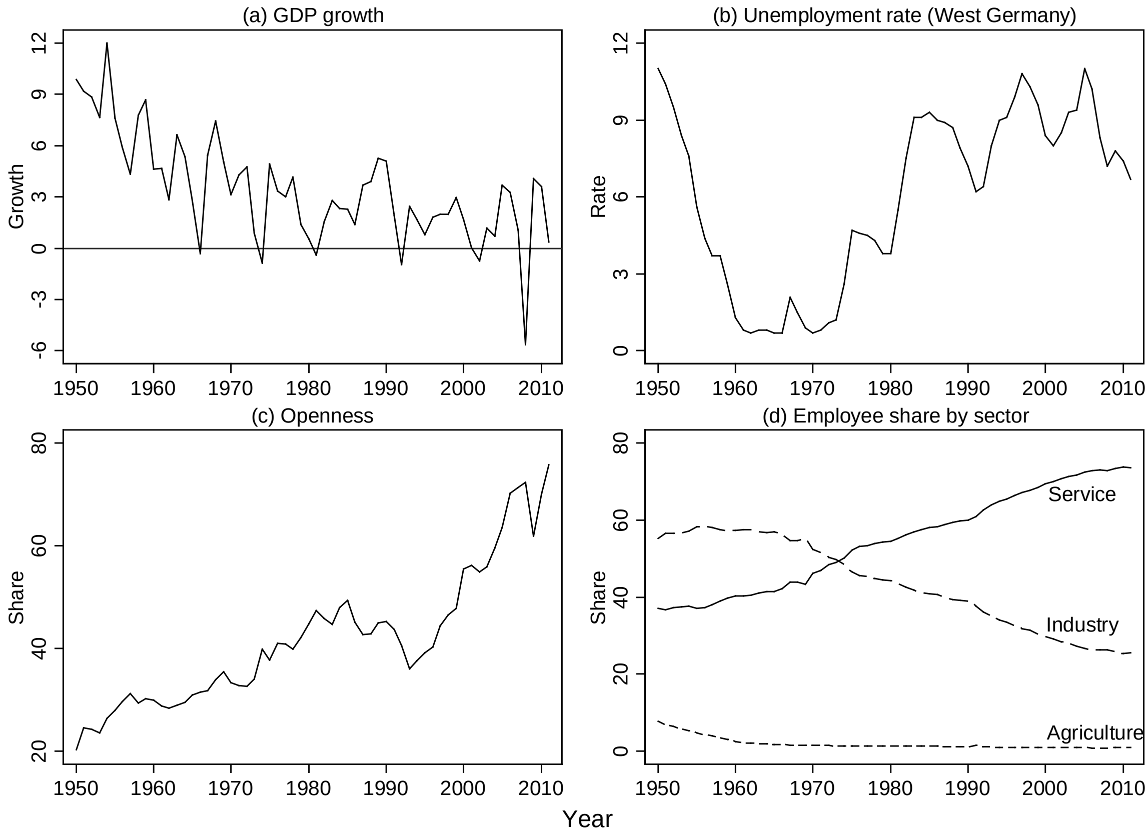
[40] The large migration inflow reduced the ratio of union members to employees until 1960, but this does not qualify as a trend.

[41] For example, the number of West German firms in textile industry dropped from 14,400 in 1960 to 4,000 in 2000, a trend common in industrialized countries (Bartels, 2014).

about 5 million people between 1989 and 1995, amplifying this pressure.[42] Influx and availability of new labor directly affected unemployment, reaching 10.8% in 1997. In addition, East Germans started to leave unions after reunification. Overall union membership dwindled even more rapidly than before, falling below 25% in 1997 (Figure 2). At the same time, sectorial contracting covered only about two-thirds of employees. Especially after 1996, newly established plants are no longer part of the classical sectoral contracting system (Card et al., 2013). The manufacturing sector employed less than one-third of the work force, with the remainder finding employment in the service sector (Figure 1). The decline in both union coverage and the industrial employment reflected continuing trends that started in the 1970s. Simultaneously, the fiscal imbalance grew: in particular social expenditures steadily rose due to costs related to unifying Germany's labor market and social security system (Bartels, 2014). By the mid-1990s, a high public deficit, low growths rates and peaking unemployment made Germany the *sick man of Europe* (e. g. Economist, 2004). Again, legislators saw labor market rigidity and high per unit labor costs as the key labor market problem and strongly expanded the possibilities of fixed term and subcontracted work (see Table C.1.). In what followed, economic openness strongly increased as Germany became a more integrated economy. Germany eventually recovered from being *the sick man of Europe,* but its labor market radically changed in the process- with effects on the evolution of earnings dynamics.

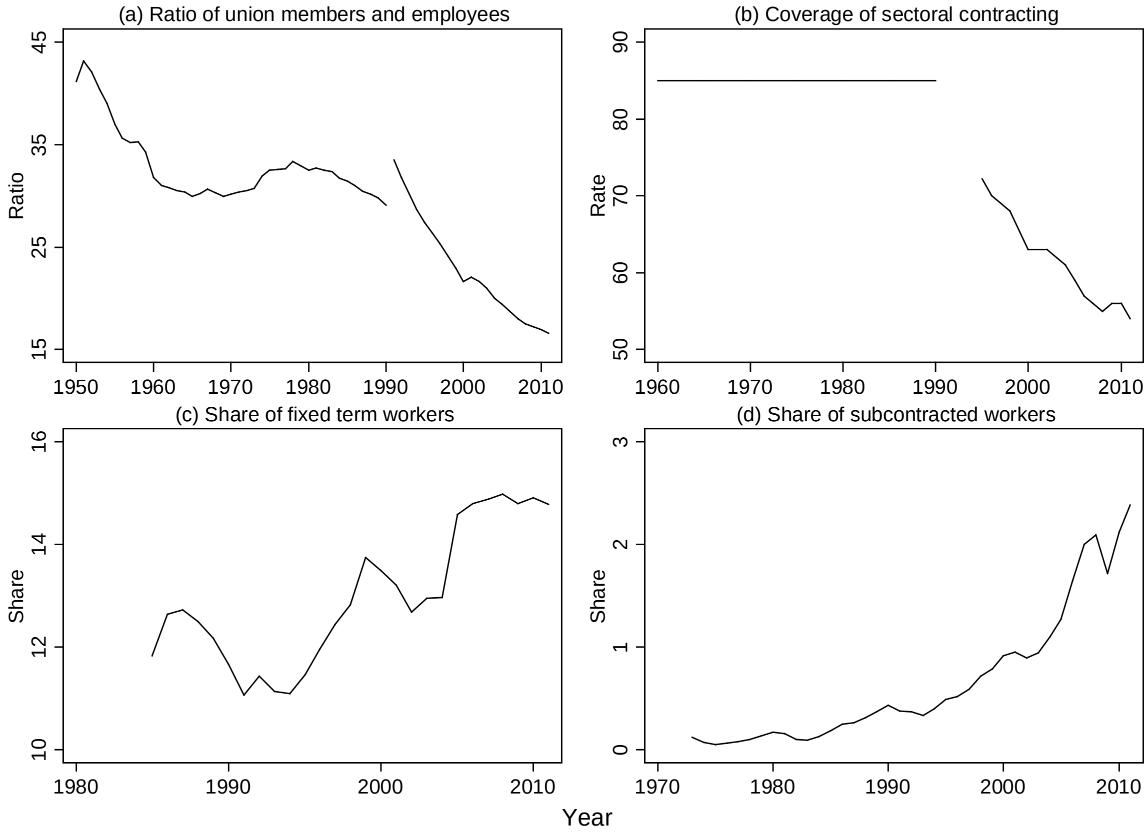
[42] The majority of the immigrants to West Germany originated from former socialist Eastern Germany. However, starting with the fall of the Irion Curtain in 1989 until 1995, each year several hundred thousand native German immigrants (Spätaussiedler) and foreign workers from former socialist Eastern European territories immigrated to Germany (Bauer et al., 2005). Bauer et al. (2005) further report that asylum-seekers and refugees led to the historical peak of 782,000 net immigrants in 1992.

Figure 1: Macroeconomic development in Germany: GDP, unemployment, openness and sectoral employment



Note: Panels (a), (c) and (d) display West Germany before 1990 and reunited Germany thereafter. Panel (b) shows West Germany only, as East German unemployment rates are substantially higher. Openess is defined as as the combined share of imports and exports over GDP.
Source: Federal Statistical Office (2015), own calculations.

Figure 2: Union membership, sectoral coverage and shares of fixed term and subcontracted workers



Note: Panels (a), (c) and (d) display West Germany before 1990 and reunited Germany thereafter. Panel (b) displays results for West Germany only.

Source: Panel (a): Deutscher Gewerkschaftsbund (2015); Panel (b) until 1990: Armingeon et al. (2014): Panel (b) after 1995: Kohaut and Schnabel (2002), Ellgut and Kohaut (2005, 2008, 2013); Panels (c) and (d): Federal Statistical Office (2015)

3 Model and estimation

Our aim is to model earnings dynamics over entire life cycles, while explicitly modeling micro- and macroeconomic dynamics. Further, we distinguish between permanent (or long-term) and transitory (or short-term) earnings path deviations.[43] The microeconomic dynamics of the permanent income component should mirror the most important, well documented, features of labor markets. Therefore, we rely on two processes- a random growth and a random walk. The random growth process relates to a Mincerian approach and captures earnings growth due to labor market experience or on the job training/tenure. This allows individuals to have permanently higher or lower growth paths than other individuals (or the cohort average). Different paths are caused by e. g. different levels of innate abilities, effort levels or education. The random walk captures permanent shifts from the individual's expected earnings path. It models a random permanent shock to the

[43] Comparable models date back to Lillard and Weiss (1979) and Hause (1980). Our model is essentially an extension of the model developed by Baker and Solon (2003). See Meghir and Pistaferri (2011) for an overview on the evolution of related models.

expected path that does not fade over time, e. g. through job displacements, negative health shocks, or additional qualifications achieved after entering the labor market.[44] Transitory shocks on the other hand describe temporary deviations from individual permanent earnings paths that fade as the individual ages. The shock persistence is modeled by assuming an AR(1) process. We now build the model step by step.

Decomposing individual i's log earnings into period t and cohort c specific mean earnings $\bar{y}_{ct}$ and the deviations from it, we get:

(1) $$earnings_{ict} = \bar{y}_{ct} + y_{ict},$$

where y_{ict} is the individual deviation from the cohort mean. In the present case, individuals range from $i = 1, \dots N$, periods and cohorts covered are $t = 1960, \dots, 2009$ and $c = 1935, \dots, 1974$. An important feature of modelling individual deviations from cohort and period specific mean earnings is its equivalency to including cohort specific age dummies. This is crucial as we investigate individual life cycles of up to 35 years (from age 25 to 59) and cover a 50 year period (from 1960 to 2009). Therefore, individual profiles are likely to be subject to cohort and age specific wage growth. By subtracting the mean (de-meaning), this growth is controlled for.[45] The individual specific deviation is now assumed to be additively decomposable into a permanent (y^P_{ict}) and a transitory component (y^T_{ict}):

(2) $$y_{ict} = y^P_{ict} + y^T_{ict}$$

Further, we define $E\left[y^P_{ict}\right] = E\left[y^T_{ict}\right] = 0$ and $E\left[y^P_{ict} y^T_{ict}\right] = 0$. Thus, expected values of both components are zero and orthogonal. Considering the aforementioned specification of the permanent earnings as a combination of a random walk and a random growth, the assumed process has the following form:

(3) $$y^P_{ict} = \pi_t \kappa_c [\mu_i + \gamma_i (t - c - 25) + r_{ict}]$$

The permanent component differs by period through shifters π_t and by cohort through shifters κ_c. These factor loadings capture macroeconomic dynamics and time trends in the permanent earnings component and ensure correct identification of the microeconomic dynamics in core model parameters. They allow institutional changes like the introduction of temporary employment to

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[44] This captures the idea that an additional degree obtained parallel to working from e.g. evening classes or weekend seminars permanently shifts the individual earnings path.

[45] This idea is introduced by Baker and Solon (2003) and is used by e. g. Bingley et al. (2013). Alternatives are regression approaches that include individual characteristics (e.g. Gottschalk and Moffitt, 2012; Meghir and Pistaferri, 2004). Since our dataset lacks most of the commonly used socio-economic characteristics, de-meaning seems the superior strategy. Further, Bingley et al. (2013) find that de-meaning gives similar results to first-stage regressions that include information on industry, education or local unemployment.

affect cohorts in a different way. Still, all cohorts share the same core process of initial earnings μ_i, random growth $\gamma_i(t-c-25)$ and random walk r_{ict}. The random growth process $\gamma_i(t-c-25)$ reads as follows. Starting at age 25, the initial earnings of an individual μ_i grow with the individual specific growth rate γ_i over time. This specification ensures that earnings levels vary both in absolute terms and by the individual's ability to accumulate skills or exert effort over the life cycle.[46] Initial earnings as well as the growth rate are assumed to stem from zero mean distributions:

$$(\mu_i, \gamma_i) \sim \left[(0,0); \left(\sigma_\mu^2, \sigma_\gamma^2, \sigma_{\mu\gamma}\right)\right],$$

where σ_μ^2 captures the variance of the starting level and σ_γ^2 the variance of subsequent earnings growth. Then, $\sigma_{\mu\gamma}$ denotes the covariance between the two components. A positive $\sigma_{\mu\gamma}$ means that those with initially high earnings also experience higher subsequent earnings growth. If the covariance is negative, this suggests the existence of Mincerian cross-overs (e. g. Mincer 1974; Lillard and Weiss 1979; Hause, 1980; Baker and Solon, 2003; Bingley et al., 2013). Then, individuals with initially high earnings upon entering the labor market experience lower subsequent earnings growth. If so, within cohort earnings inequality will decrease in the beginning and then increase at later stages of the life cycle.

The random walk component r_{ict} is defined as:

(4) $$r_{ict} = r_{ic(t-1)} + u_{ict}.$$

As mentioned above, the permanent component includes shocks with permanent effects like job changes, job displacements or disabling injuries (e.g. MaCurdy 1982; Moffitt and Gottschalk, 1995; 2012; Baker and Solon, 2003). The random walk component is assumed to be i.i.d. with $u_{ict} \sim (0; \sigma_u^2)$. Then, the (independent) variance of permanent re-orderings is captured by σ_u^2, which allows a linear 'white noise' innovation in the permanent component (Baker and Solon, 2003). Note that these innovations do not vanish over the life course. In sum, the auto-covariance structure of permanent earnings for period t and period s can be written as:

(5) $$Cov\left(y_{ict}^P y_{ics}^P\right) = \pi_t \pi_s \kappa_c^2 \left[\sigma_\mu^2 + \sigma_\gamma^2 ts + \sigma_{\mu\gamma}(t+s) + \sigma_u^2 s\right]$$

In an exemplary case for cohort $c = 1935$ in period $t = 1970$ the variance (hence for $t = s$) of the permanent component according to (5) is displayed in equation (5a). Note that the process builds on earnings of individuals who are at least 25 years of age. Therefore, cohort 1935 entered in 1960 and is 10 years past its entry in 1970:

[46] See e.g. Baker (1997), Baker and Solon (2003), Cappellari (2004), Bingley et al. (2013).

(5a) $$Var\left(y^P_{i,1935,1970}\right) = \pi^2_{1970}\kappa^2_{1935}\left(\sigma^2_\mu + 100\sigma^2_\gamma + 20\sigma_{\mu\gamma} + 10\sigma^2_u\right)$$

For the microeconomic dynamics of the transitory component, several studies establish that a low order ARMA-process is sufficient. [47] Specifically, we follow e. g. Baker and Solon (2003) and model an AR(1) process for the transitory earnings component. Similarly, we adopt period (τ_t) and cohort (λ_c) specific shifters to explicitly model the influence of institutional changes or macroeconomic trends on specific cohorts and to correctly identify the microeconomic dynamics of earnings insecurity. For the transitory component we obtain:

(6) $$y^T_{ict} = \tau_t\lambda_c\nu_{it} \text{ and } \nu_{it} = \rho\nu_{i,t-1} + \epsilon_{it}$$

where ϵ_{it} is a random shock with $\epsilon_{it}{\sim}(0;\sigma_\epsilon)$ and $0 < \rho < 1$ the persistence of the transitory shock. The initial transitory variation at the first period of observation, $\nu_{i0}{\sim}(0;\sigma^2_0)$, is observed at $t - c - 25 = 0$ (thus at age 25). Subsequent earnings instability is captured by the variance of innovations σ^2_ϵ. Typically, earnings or wage instability is associated with a u-shaped pattern in age with higher instability for young (labor market entry) and old (labor market exit) workers. To allow earnings instability to vary with age, we follow Baker and Solon (2003) and incorporate a quartic age function (polynomial of the fourth degree) of the variance σ_ϵ.[48] In sum, the auto-covariance structure of transitory earnings can be written as:

(7) $$\begin{aligned} Var\left(y^T_{ict}y^T_{ict}\right) = \lambda^2_c\big[&\rho^2 Var\left(\nu_{i,t-1}\right) \\ &+ \tau^2_t\left(\sigma^2_{\epsilon,0} + (t-c)\sigma^2_{\epsilon,1} + (t-c)^2\sigma^2_{\epsilon,2} + (t-c)^3\sigma^2_{\epsilon,3}\right. \\ &\left.+ (t-c)^4\sigma^2_{\epsilon,4}\right)\big] \end{aligned}$$

And for $t \neq s$ we obtain:

(7a) $$Cov\left(y^T_{ict}y^T_{ics}\right) = \lambda^2_c\left(\rho Cov(\nu_{i,t-1}\nu_{is})\right)$$

Returning to our example from (5a), the variance of the transitory component for cohort 1935 in period 1960 amounts to:

(7b) $$Var\left(y^T_{i1935,1970}\right) =$$

[47] E. g. Moffitt and Gottschalk (2012) find that higher order ARMA-parameters are not significant.

[48] However, we deviate from Baker and Solon (2003) not only in the incorporation of permanent cohort shifters, but also in the incorporation of a transitory cohort shifter, λ_c , in addition to the usual transitory period shifter, τ_t. Further, Baker and Solon (2003) model cohort specific initial variances. However, more recent literature shows that those are subject to a potential bias due to left-censoring (e.g. Moffitt and Gottschalk, 2012). Since we observe cohorts from the beginning (here age 25), this bias does not apply to our setting. In order to ensure a comparison with other recent models and in order to be able to shorten the timeframe for a robustness test, we moved away from the cohort specific initial variances to cohort specific transitory shifters. The results for both specifications are not qualitatively different. Still, the latter specification gives a slightly better fit. Additional results are displayed in Appendix A.

$$\lambda_{1935}^2\left[\rho^2 Var\left(\nu_{i1935,1969}\right) + \tau_{1970}^2(\sigma_{\epsilon,0}^2 + 10\sigma_{\epsilon,1}^2 + 100\sigma_{\epsilon,2}^2 + 1{,}000\sigma_{\epsilon,3}^2 + 10{,}000\sigma_{\epsilon,4}^2)\right]$$

Due to the orthogonality assumption, the total auto-covariance structure results from the sum of the permanent component (5) and the transitory component (7) or (7a) respectively. For the estimation procedure, we apply equally weighted minimum distance. See Appendix B for details.

4 Data and descriptives

4.1 Sample selection

We use *Versicherungskontenstichprobe* (VSKT), German social security data, as provided by Deutsche Rentenversicherung.[49] A stratified random sample, the VSKT provides the records of mandatorily insured employees in Germany. The requirements are at least one (pension relevant) entry in the employment biography and 30 to 67 years of age in the reference year. We use the waves of the reference years 2002 and 2004-2009. The VSKT contains the employment biographies after 14 years of age until the age in the reference year (up to a maximum of 67 years of age). These biographies include monthly information on (un-)employment, sickness and pension contributions. The latter are used to calculate the individual earnings. In line with most of the literature on earnings component models using administrative data, we only consider earnings covered by social security. Earnings from self-employment and government transfers are not included in our wage measure.[50] In addition, civil servants are not covered. Still, the VSKT represents about 80% of the total male work force in West Germany (Bönke et al., 2015).

We consider men only to ensure comparability to related studies and to avoid sample selection issues due to changing labor market participation rates of women (Bönke et al., 2015). Further, we focus on men between 25 and 59 years of age. This excludes from our analysis both the unstable years of very young workers (including military and civil service) and the retirement transition period. This enables comparisons to other studies, which exclusively focus on comparable populations. Then, we focus on native Germans who have always worked in West Germany to avoid the problem of fractured biographies.[51] Individuals who have worked in East Germany are excluded because their earnings information and earnings level over time is not comparable to that of West Germans. This

[49] Our dataset, FDZ-RV—VSKT2002, 2004–9_Bönke, is accessible through controlled remote computing and provided by the Data Research Centre of Deutsche Rentenversicherung (the German statutory pension scheme). Cohorts and the underlying sample are constructed in the same way as in Bönke et al. (2015).

[50] Jenson and Shore (2015) find that earnings volatility and its evolution differ between self-employed and employed.

[51] This excludes immigrants as well as native German immigrants ("Spätaussiedler") who worked in their country of origin. Further, West-East migration is negligible before reunification and extremely small thereafter (Fuchs-Schündeln and Schündeln, 2009).

especially holds for the older sample cohorts. Younger East German cohorts are then excluded to ensure sample consistency for the investigation of long run trends.

The oldest cohort we observe is born in 1935. For this cohort and all others up to cohort 1950, we observe complete life cycles from age 25 to age 59. For those born after 1950, we observe biographies that are right censored at the cohort's age in 2009. We include 40 cohorts up to the one born in 1974 to ensure a sufficiently long period of observation.[52]

Although the VSKT is virtually free from measurement errors, we perform three adjustments in order to ensure time consistency in the earnings data. First, since one-time payments are only subject to social security since 1984, earnings prior to 1984 are adjusted according to their spurious growth between 1983 and 1984.[53] Second, we deal with the problem of different levels of social security contributions over time and subgroups. Therefore, we add the employers' social security contributions to the individual gross wages. These contributions can be seen as an approximation of the value of insurance that employees would have bought if the insurance had not been supplied by governmental institutions (Bönke et al., 2015). In this sense, the earnings we analyze represent the market value of labor.[54] Our third adjustment is an imputation of top-coded earnings. In Germany earnings are only subject to social security up to a contribution ceiling. This causes our earnings data to be right-censored at this ceiling. Our imputation method is extensively documented in Bönke et al. (2015) and assumes a Pareto-distribution for the upper tail. The imputation is done separately by year and cohort. Since we do not want to artificially impute variance into the sample, we follow Bönke et al. (2015) in the assignment of wage above the contribution ceiling and preserve the individual ranks prior to the censored wages. This is an assumption of minimal mobility for individuals who consistently earn wages above the ceiling.[55]

[52] This subsequent entry of younger cohorts might be a problem for the identification of time and cohort effects of early calendar years since in these years only few cohorts are observed at the same time. Therefore we include a robustness test and estimate the model starting in 1979, discarding all prior years and adjusting the sample selection. We observe no qualitative difference in the results (see Appendix A). It seems that the auto-covariances ensure consistent estimates even for periods when only few cohorts are present.

[53] The method is documented in Bönke et al. (2015). It is an extension of Fitzenberger's (1999) cross sectional adjustment of administrative data to spurious growth. It exploits the panel structure of the VSKT and adjusts the wage according to the individual age and rank in the earnings distribution.

[54] Since, e. g., miners have higher levels of social security contributions and a changing relative weight over the cohorts, subgroup consistency can only be assured when using the market value concept. This also solves the problem of changing levels of social security contributions (to pension, unemployment, health, and nursing care insurance) over time. For instance, contributions were lower in the 1960s than in the 2000s. All parameters of the social security system used for constructing the market values are provided in Bönke et al. (2015).

[55] The opposite would be an assumption of maximal mobility, which would introduce artificial variance into the sample. Further, to limit the influence of outliers, we censor the highest wage at 5 times the average social security wage. Very few observations are affected by this censoring. Limiting the influence of outliers is common in the literature; see e.g. Bingley et al. (2013). Further details of the imputation method are provided

Finally, the sample is restricted to those with consistent earnings biographies. We are left with at least 1,000 observations per cohort and about 50,000 in total, amounting to about 1.2 million person years. Details are provided in Appendix A. Our results are robust to conditioning on at least 5 years of consecutive earnings. [56] All earnings are real earnings with the base year 2000.

4.2 Sample descriptives

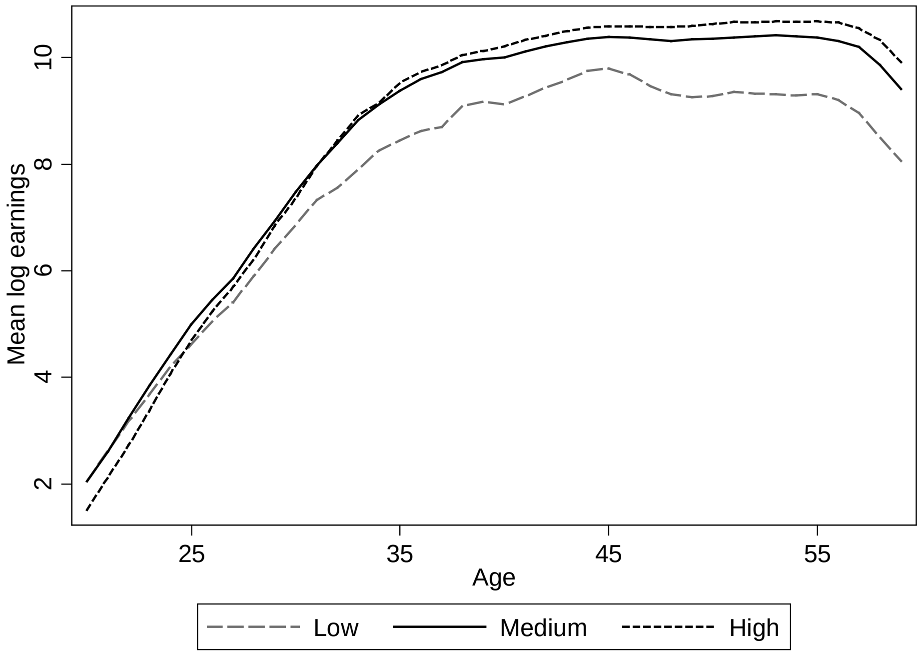
To provide some empirical motivation to our model, this section presents important attributes of the evolution of earnings and their dispersion in Germany. Figure 3 displays age-earnings profiles for three groups sorted according to their lifetime earnings into low (1st quartile), medium (2nd and 3rd quartile) and high (4th quartile) earners. Lifetime earnings are calculated as real (CPI-deflated) net present values from ages 17 to 59. The means show the expected inverted u-shape over the life cycle, are closest at young ages and fan out at later ages. In line with theoretical predictions and empirical findings, Figure 3 reveals Mincerian cross-overs,[57] e. g. when the high earners' mean passes the low earners' mean at 26 and does not fall below again.

by Bönke et al. (2015), who also validate the imputation procedure with survey data and find no difference between the cross-sectional earnings distributions of the VSKT and the survey data. There is no robustness-test with completely censored data on purpose. Since the ceiling changes by calendar year and is, in general, increasing over time, it must be imputed. For a thorough representation of the ceiling's evolution see e.g. Lüthen (2015).

[56] We consider biographies to be consistent if the sample provides a nearly gapless record of individual labor market activities after age 30 (equal to Bönke et al., 2015). The idea of consecutive information follows Bingley et al. (2013), who sought a criterion that neither constructs a fully unbalanced panel nor one that excludes too many observations. Conditioning on consecutive earnings yields a slightly worse fit but no qualitative differences. The results are displayed in Appendix B.

[57] This implies that those with high earnings at young ages are not those with the steepest permanent earnings paths (see Section 2.1 for more details).

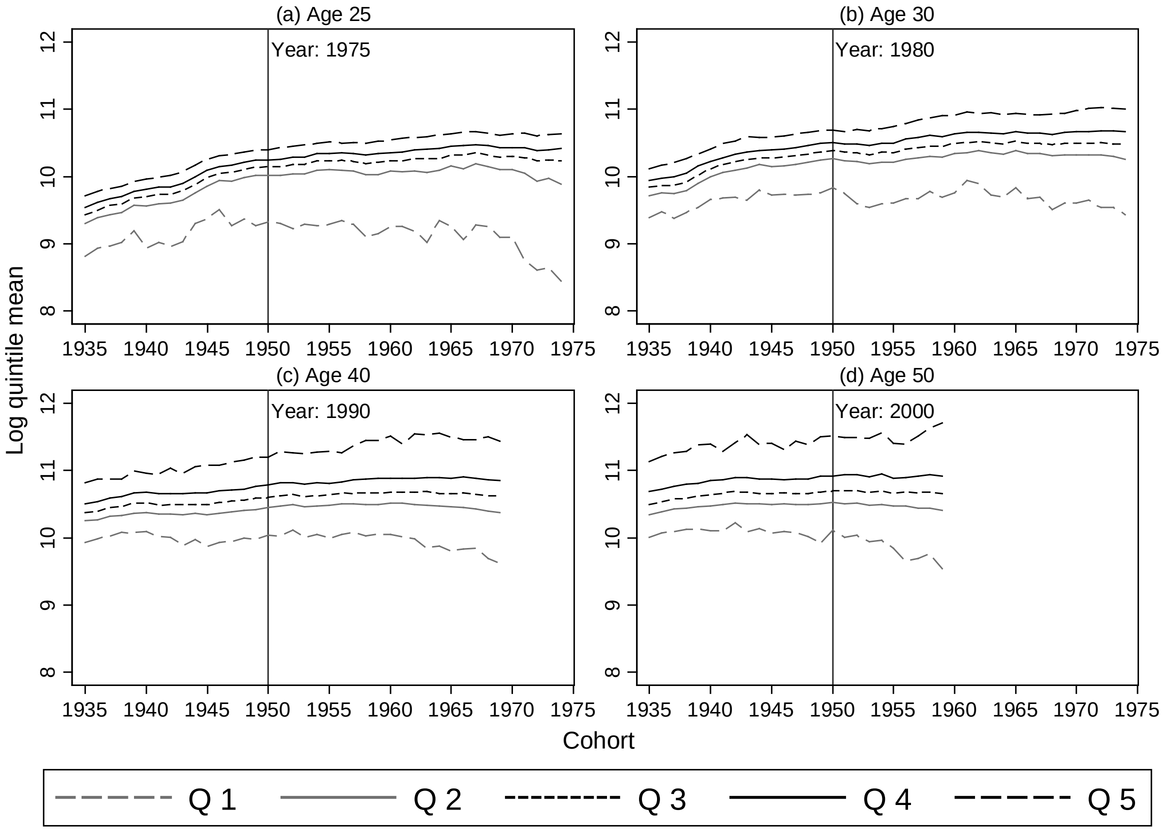
Figure 3: Means of logarithmic earnings by lifetime earnings, pooled cohorts 1935-1950



Source: FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations.
Note: "Low" depicts mean earnings of individuals in the lowest quarter of lifetime earnings, "Medium" of those between the 25th and the 75th percentile and "High" of those in the highest quarter.

Figure 4 concentrates on developments across cohorts and shows quintile means of logarithmic earnings for selected ages by cohort. For ages 25 and 30, Figure 4 displays stable earnings growth as well as stable quintile distances across cohorts except for the lowest quintile. The lowest quintile fluctuates strongly and its distance to the other quintiles increases. For later ages, Figure 4 reveals moderate cohort specific earnings growth for quintiles 2 to 4. The highest quintile gains more and the lowest quintile declines across cohorts. However, distances between the lowest quintile and other quintiles decrease for later ages and its evolution stabilizes. This indicates more earnings instability in the early stages of the life cycle, which increases for younger cohorts and decreases after age 30 for all cohorts. Widening distances between the earnings quintiles on the other hand suggest increasing permanent divergences for younger cohorts and later ages. These findings are in line with Dustmann et al. (2014), who find decreasing wages for the 15th percentile, a rather stable median and increasing wages for the 85th percentile since 1990. This first impression underlines the importance of certain key aspects of our model: An age and cohort specific modeling of permanent and transitory components is needed to uncover underlying trends across life-cycles and generations.

Figure 4: Means of logarithmic earnings for selected ages



Source: FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations.
Note: Q1 to Q5 relate to the respective quintile means of logarithmic earnings at various ages.

5 Results

5.1 Microeconomic dynamics: Core model estimates

Table 1 presents our core model estimates based on Equations (5) and (7). It shows that the assumed flexibility of the theoretical structure of the permanent and the transitory component are key to fitting the model to life-time earnings data. The model identifies heterogeneity both in starting levels (σ_{μ}^{2}) and in subsequent earnings growth (σ_{γ}^{2}). The estimates suggest that individuals whose earnings grow one standard deviation above the mean accumulate an average income advantage of about 1.6% per year ($100 \cdot \sqrt{\sigma_{\gamma}^{2}} \ = 1.63$). The result lies between the findings of Baker and Solon (2003) for Canada (1%) and those of Baker (1997) for the USA and Bingley et al. (2013) for Denmark (both about 2.8%). Since we estimate the average annual growth rate to be 0.24%, our model outcome indicates considerable growth rate heterogeneity.[58] Like most studies on earnings dynamics (e. g. Baker and Solon, 2003; Moffitt and Gottschalk, 2012; Bingley et al., 2013), we estimate a negative covariance between initial earnings and subsequent earnings growth, ($\sigma_{\mu\gamma}$<0). This is typically interpreted as a

[58] We follow Bingley et al. (2013) and estimate the comparison estimate of average annual growth as a regression of low-wages on a linear age trend. Bingley et al. (2013) find a larger estimate of 0.9% for Denmark.

trade-off between initially relative high earnings and subsequent earnings growth as predicted by the Mincer-earnings-model. Following Hause (1980) and Bingley et al. (2013), $t^* = -\sigma_{\mu\gamma}/\sigma_{\gamma}^2$ gives the point of lowest (permanent) inequality due to diverging earnings paths (abstracting from permanent shocks of the random walk). Here, we find an estimate of 1.2 years, implying a low shortly after 26 years of age. The estimate confirms the impressions of Figure 3 and can be related to Bönke et al. (2015), who find that earnings paths of highly educated individuals start below the earnings paths of the lesser educated, rise steeper, cross in the late 20s and exceed thereafter.

Table 1: core model estimates

Permanent component			Transitory component		
	Coeff.	SE		Coeff.	SE
σ_{μ}^2	0.05	0.02	σ_0^2	0.1	0.057
σ_{γ}^2	0.0003	0.0001	$\sigma_{\epsilon,0}^2$	0.0955	0.0547
σ_{u}^2	0.01	0.004	$\sigma_{\epsilon,1}^2$	-0.018	0.01
$\sigma_{\mu\gamma}$	-0.0003	0.00017	$\sigma_{\epsilon,2}^2$	0.002	0.001
			$\sigma_{\epsilon,3}^2$	-7.3E-05	4.2E-05
			$\sigma_{\epsilon,4}^2$	1.17E-06	6.81E-07
			ρ	0.28	0.005

Note: Remaining model estimates are provided in Appendix B.
Source: FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations.

The estimation of transitory innovations, $\sigma_{\epsilon,i}^2$, suggests a u-shape over the life cycle and is in line with Baker and Solon's (2003) most comparable estimates. The innovations fall more than 60% from the mid-twenties to the early forties, flatten out over the forties and rise again in the early fifties and then reach the levels observed in the early stages of the life cycle again. Our estimate of 0.28 for ρ is relatively low compared to other studies and suggests low shock persistence for transitory innovations in Germany (Baker and Solon (2003) find a value of 0.54 for Canada). The results are robust to a left-censoring like a later start of the analysis in 1979. This robustness test shows that the low value of ρ is not driven by the higher weight of older cohorts.[59] Further robustness tests include an estimation of the baseline model with a different sample selection criterion (5 years of consecutive employment) and an estimation of a model that exchanges the transitory cohort shifters for cohort-specific initial variances (following Baker and Solon, 2003). All robustness checks and the remaining parameters for period and cohort shifters of both the transitory and the permanent component are presented in Appendix B.

[59] The robustness-test further reveals no qualitative difference apart from a strong increase in significance. In line with Gottschalk and Moffitt (2012), the robustness test includes an additional parameter to deal with left-censoring. Details are provided in Appendix A.

Our findings for the underlying microeconomic dynamics indicate more earnings instability during very early and very late stages of the working career and increasing permanent divergences over the life cycle. This finding is similar even regarding country-specific aspects and differences in data quality. Studies that model age-dependent innovations typically find a considerable decline of earnings instability after age 25, reaching a trough between ages 35 – 45 and rising again thereafter (e .g. Baker and Solon (2003) for Canada, Karahan and Ozkan (2013) for the U.S. and Blundell et al. (2014) for Norway). Guvenen et al. (2015), while accounting for variation in higher order moments, conclude with similar results for the U.S.

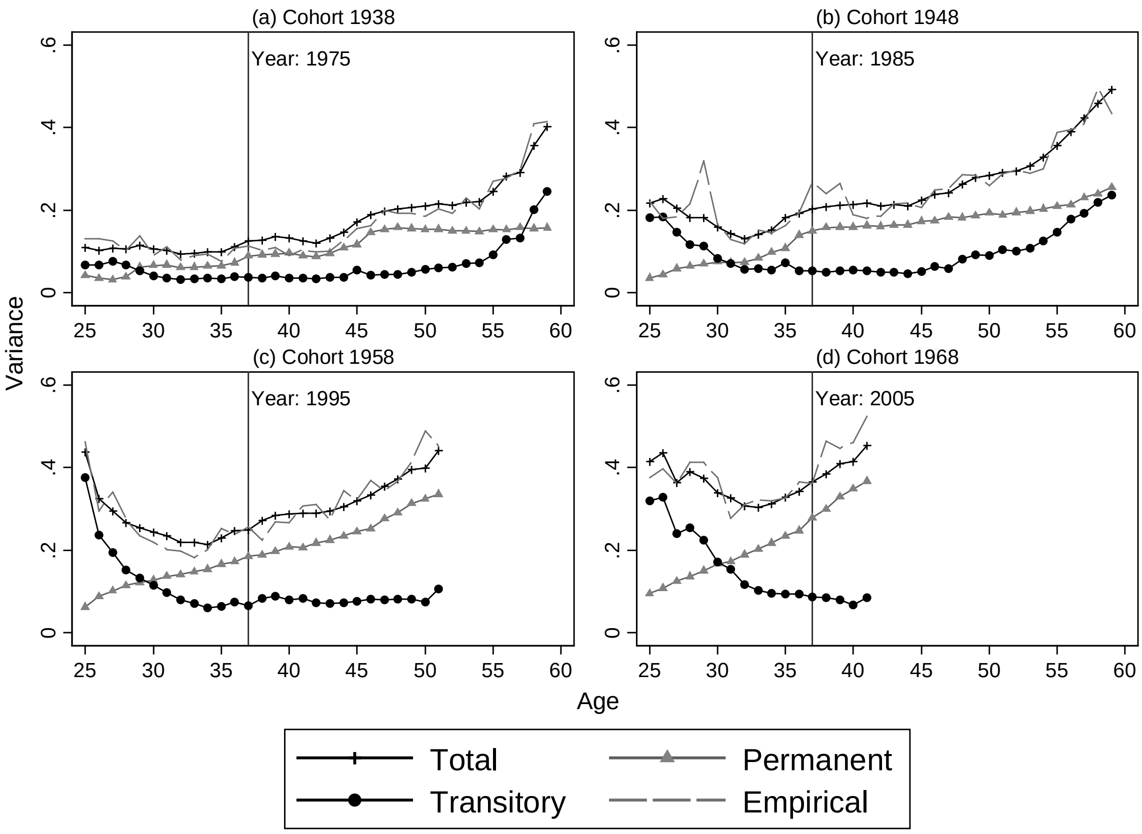
5.2 Earnings dynamics over the life cycle

Figure 5 outlines the empirical and predicted variances over the life cycle for selected cohorts.[60] The empirical variance evolutions (line: dash) are well matched by the predictions of the total variances (line: +). The total variance decreases until the early 30s and increases afterwards. This is in line with studies examining inequality over the life-cycle (e. g. Björklund, 1993; Kopczuk et al., 2010; Bönke et al., 2015), which estimate the lowest point of overall cross-sectional inequality to be around this age. The evolution of the transitory component (line: ●), which is about u-shaped over the life cycle also after the inclusion of period and cohort shifters. The permanent component (line: Δ) usually rises over life cycle.

Figure 5 also reveals two other important findings. First, younger cohorts face higher total earnings variance and both higher transitory and permanent variances. Second, the results suggest a different composition of variance components across generations. For younger cohorts we find a more pronounced u-shape of the transitory component and a steeper rising permanent component. Thus, younger cohorts face higher earnings instability at the beginning of their life cycle and a steeper rising permanent component, hence more divergence between earnings paths over their life cycle.

[60] The interpretation given for the selected cohorts is in line with the results for all cohorts. Figures for all cohorts are available from the authors upon request.

Figure 5: Empirical and predicted variance for selected cohorts



Source: FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations.

5.3 Evolution of earnings dynamics across generations

The structural shift in the variance components across cohorts becomes more apparent when comparing variances at various ages. Figure 6 displays the actual observed empirical variance (line: dash) and the according estimation for the permanent (line: Δ) and transitory (line: ●) component for each cohort at the beginning (age 25 and 30), in the midst (age 40), and toward the end (age 50) of the earnings career. In addition, the total estimated variance as the sum of both components is displayed (line: +) to give an impression of the model fit. Comparing empirical and estimated variances reveals a satisfying fit across all cohorts and age groups. Figure 7 complements Figure 6 and displays the respective growth of the permanent and transitory components at these respective ages, normalized by estimates for cohort 1935.

We comment on the transitory component first. Confirming the upward shift pictured in Figure 5, Figure 6 and 7 reveal a marked increase at younger ages (upper panels), a considerable less pronounced trend for age 40 and almost no clear trend for age 50 (lower panels). The differences by age pertain to the importance of the transitory component at the beginning of the working career (e. g. Baker and Solon, 2003). The strong trend at young ages relates to the three phases of institutional changes and macroeconomic trends outlined in Section 2, which particularly affect labor market entrants. Cohorts born until the mid-1940s entered the labor market before 1973 during favorable economic circumstances of the *Wirtschaftswunder*-period. For these cohorts, the transitory component remains comparably low and rather stable, mirroring low unemployment, high job security and no fixed term or temporary employment. Cohorts entering between the early 1970s and the mid-1990s encountered less favorable economic conditions and ongoing labor market deregulation, including the introduction of fixed term employments for first time employees (see Figure 4 and Table C.1.). This reflects in a tripling transitory component between pre-1946 cohorts and 1960s cohorts at age 25 and a doubling at age 30. Finally, those born in the early 1970s experienced another strong increase in the transitory innovations upon labor market entry. They joined the labor force around and after the mid-1990s during Germany's period as the *sick man of Europe*, a time of economic hardship characterized by mass immigration, high unemployment, sectoral shifts, deunionization, and competition with the former socialist East. This conjoins further labor market deregulation regarding dismissal protection, fixed term contracts and temporary employment (Figures 3 and 4; Table C.1). All this contributes to the steep surge of the transitory variance observed at ages 25 and 30 for cohorts born in 1970s. In comparison to cohorts born before the mid-1940s, 1970s cohorts face an earnings insecurity that is about five to seven (age 25) and three to five (age 30) times higher. The earnings risk still doubles for 40 year olds in course of the mid-1990s events. For 50-year-olds, the earnings risks increases only slightly after 1995, mirroring long and stable earnings careers and favorable employment contracts.

For the permanent variance, Figures 6 and 7 display an increase at all ages. The largest relative increase occurs for the young at age 30. Still, since the permanent component is more pronounced at later ages, its absolute gain is largest at ages 40 and 50 (see also Figure 5). In contrast to the transitory component, the increase initially starts after the second oil crises in 1980. Between 1980 and 1990, the permanent component doubles for ages 25 and 30; the 40 year olds are slightly less and the 50 years olds are not affected. Thus, our results suggest that the favorable conditions of the *Wirtschaftswunder*-period only diminish after the more fierce recession following the second oil crises. In the 1980s, mass unemployment and deregulation put permanent pressure on the wage structure. In addition, the increasing number of workers in the service sector was mainly recruited from younger cohorts. These contracts do not offer the same security and wage compression as the long-term industry contracts most prominent for older cohorts. This could also explain why older, well-established workers (aged 50) are not affected by the 1980s recession.

After reunification, we observe a second surge in the permanent component, coinciding with several severe global and local changes. While the increase at age 25 is rather small, the increase at age 30 is already distinct. At ages 40 and 50, a steady rise begins around 1995 and surges in the early 2000s. These increases coincide with four important developments starting in the early 1990s. First, the ongoing globalization puts pressure on low skilled labor, e. g. through offshoring and growing international competition. This also reflects in the strong increases in openness since 1990 (Figure 3). Second, changing job requirements cause job polarization on German labor markets. Job polarization describes a shift in demand toward very highly skilled, non-routine labor at the expense of workers tasked with routine operations, e.g. due to the effect of computerization on clerical work. For Germany since the mid-1990s, Dustman et al. (2009) find job polarization a driving force for wage inequality. Third, deunionization and the opting out of sectoral agreements have also had long term consequences (e. g. Acemoglu et al., 2001).[61] Antonczyk et al. (2010) attribute a considerable share of rising wage inequality to de-unionization and the decline in collective wage bargaining coverage, especially at the lower end of the wage distribution. Card et al. (2013) find that this joint decline of traditional German wage bargaining institutions increases employer-specific compensations and widens wage differences among employees in the same industry. Since the 1990s, wage negotiations have shifted from collective bargaining to the individual level. In particular establishments founded after 1996 are more likely to pay lower wages, to exhibit larger wage heterogeneity and to not participate in the sectoral contracting system. Since younger workers are more likely to work at these establishments, our result of rising permanent dispersion and earnings instability for younger cohorts are in line with a declining coverage by collective sectoral wage bargaining. Forth, the probability of job changes increased since the 1990s. Voluntary turnover grew since the pecuniary gains of job changes increase, which provides an incentive to change employers more often (Card et al., 2013). Involuntary job changes increased due to lowered dismissal protection and the enhancement of fixed term contracts and subcontracted work (Figure 4; Table C.1; Figure C.1). The rising permanent variance since the 1990s therefore reflects more diverse permanent earnings paths as well as larger shifts of the paths. The increasing importance of permanent shocks also implies greater difficulty in returning to the previous path after a negative shock like a health shock or involuntary job loss.

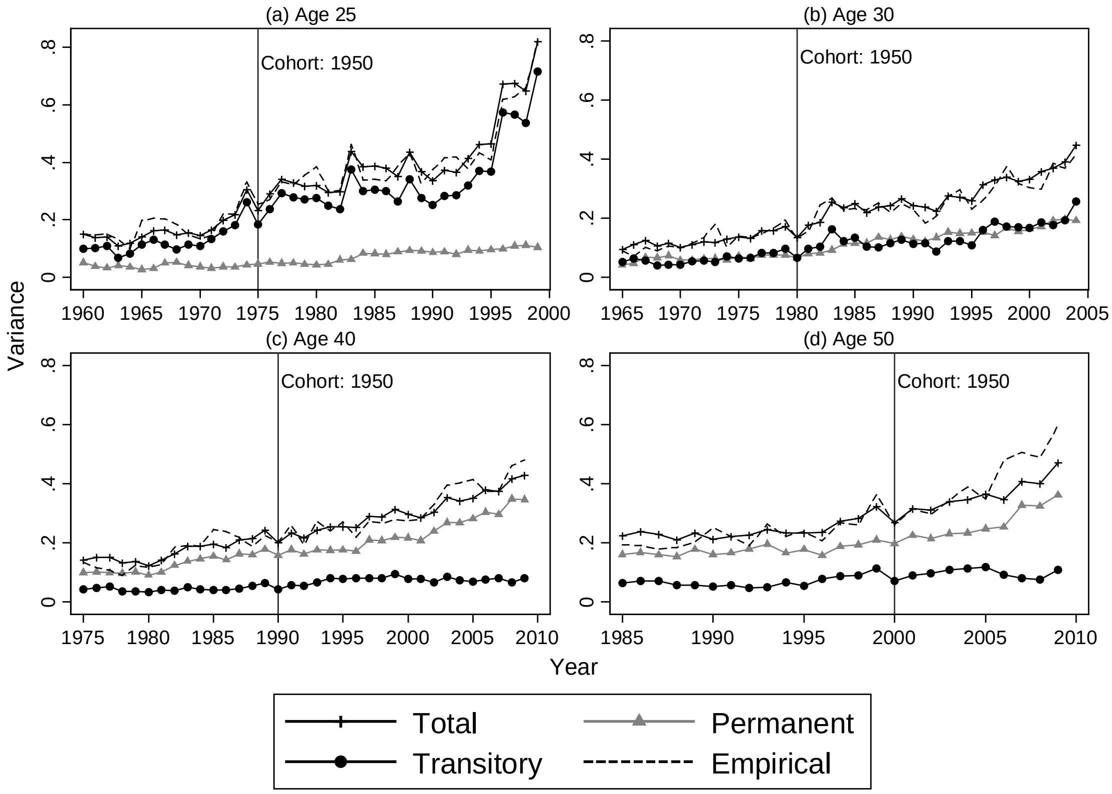
The observed developments mirror macroeconomic trends and institutional changes that affect income distributions in the long run. Therefore, the permanent variance and its growth follow a smoother trend than the transitory variance. In contrast to the transitory component, the increase in the permanent components starts after the second oil crisis in 1979 at ages 25, 30 and 40 and after the mid-1990s at age 50. Existing contracts seem to dampen immediate effects of large scale events on the permanent component, causing a slowed response. On the other hand, these events immediately hit the most vulnerable- young workers without a strong labor market attachment and

[61] These developments are not independent; e.g. skill-biased technological change is a likely driver of both job polarization (Dustmann et al., 2009) and deunionization (e. g. Acemoglu et al., 2001).

job seekers. This causes immediate effects in insecurity for young workers after macro-shocks and delayed effects on permanent divergences at later ages.

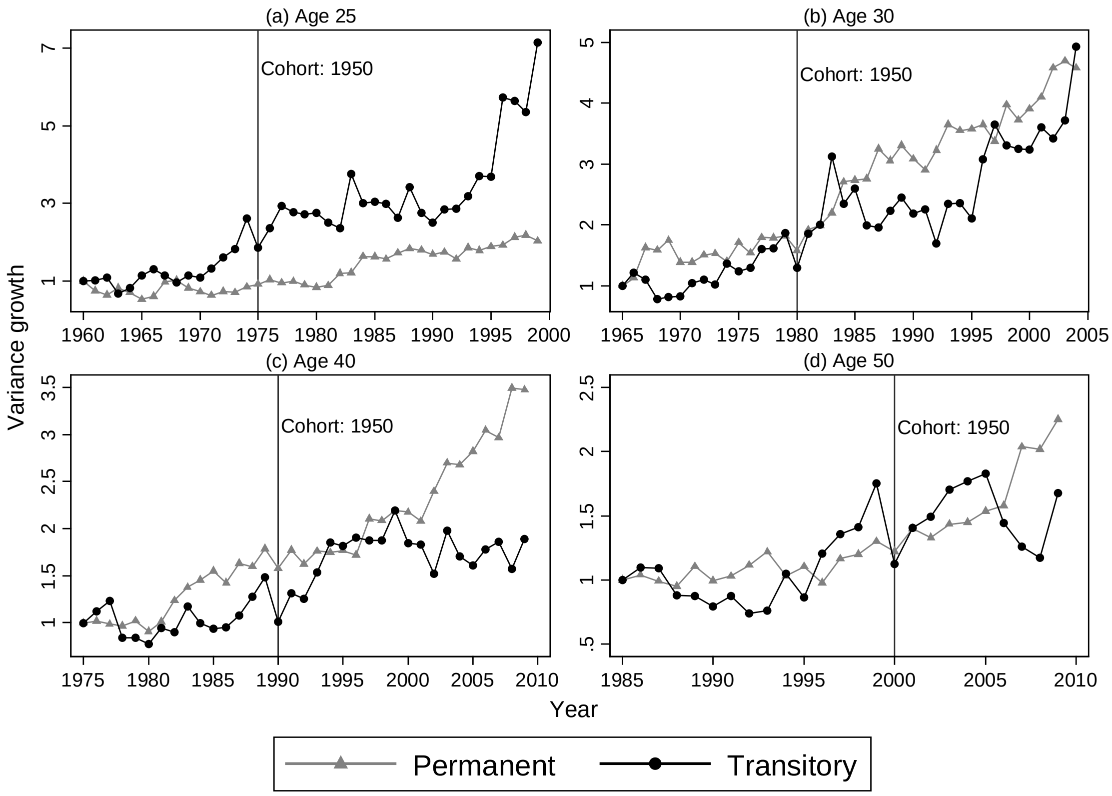
Studies on other countries report comparable results for men. For the U.S., Moffitt and Gottschalk (2012) report a substantial increase in earnings instability throughout the 1970s and 1980s and major more immediate shifts during recessions. They further identify a considerable rise in the permanent component since the mid-1990s. Although their data does not allow for controlling cohort differences, these results roughly align with ours. For Canada, Baker and Solon (2003) and Ostrovsky (2010) document a rise in both components after the second oil crises in 1980 and another steady increase since the early 1990s. For Italy, Cappellari (2004) finds similar trends with a stronger increasing permanent component. Apart from recessions, he ascribes the rise to higher demand for skilled labor and the decline of the strongly regulated pay-system in Italy. In this process wages become more often determined at the firm level, which can be compared to what happened in Germany. Further, Sologon and Van Kerm (2014) provide a visual summary of existing studies on European countries and the US, confirming the upward trends in both components (except for the transitory component in Luxembourg).

Figure 6: Empirical and predicted variance at selected ages



Source: FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations.

Figure 7: Variance growth rates at selected ages



Source: FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations.

5.4 Implications and discussion

Albeit all the presented findings relate to gross earnings, they have several implications for net disposable income and consumption. The extent that an individual or society and its welfare is affected depends on the welfare state's ability to insure against earnings risk and to compensate for permanent income differences through redistribution. Younger cohorts and lower skilled individuals experience higher transitory fluctuations of gross earnings and more pronounced inequality in terms of a more dispersed permanent income component. Without insurance or the adjustment of existing social security schemes, those transitory fluctuations translate directly into additional welfare costs (Storesletten et al., 2001; Blundell and Preston, 2008). So far the German welfare state seems to cope well in smoothing transitory earnings shocks, even for younger cohorts (Bartels and Bönke, 2013). Mitigating increasing long-term disparities, on the other hand, would require increasing redistributive capabilities, e.g. by more progressive income tax schedule. However, recent modifications to German income taxes show an opposite trend and it is unlikely that this trend will change.[62]

As discussed above, many forces that drive rising permanent disparities are global in nature. However, developments in Germany tend to amplify this trend. The formerly strong equalizing influence of trade unions is diminishing, reflected declines in coverage of sectoral contracting agreements and union membership. Further, to strengthen international competitiveness, the adopted labor market deregulation aimed at cutting employment costs and increasing flexibility. In terms of employment and economic recovery, the deregulation is successful, however at the cost of higher inequality (Dustmann et al., 2014). This flexibility, along with decentralization of wage determination from the industry level to single firms or even workers, coincides with a decrease of real wages at the lower end of the wage distribution. In sum, the changing German labor market institutions further fostered the dispersion of wages and earnings careers. At the same time, adjustments in the tax and transfer system reduced the redistributive impact of the German welfare state (Bartels and Bönke, 2013). Hence, gross earnings inequality translates into net earnings and disposable income inequality.

Although Germany's economy is recently performing exceptionally well, it is unlikely that Germany will ever regulate the labor market in the way that it was in the 1960s, due to the fear of losing its competitive advantage. Increased flexibility might also attract more volatility industries and therefore amplify the trend of increasing earnings instability and inequality (Cunat and Melitz, 2012).

[62] Figure C.2 displays the evolution of marginal and average income tax rates from 1958 through 2013 for varying earnings levels. Figure C.2 shows a reduction in progressivity of the German income tax since the mid-1980s.

Hence, employees starting their earnings career after the 2000s will likely experience a continuing trend in rising levels of inequality and uncertainty. Still, excesses at the lower end of the wage distribution are of concern and in 2015 Germany introduced a nationwide minimum wage. Its impact on earnings inequality and volatility remains to be investigated. On the one hand, it might increase the unemployment risk for low-skilled workers, but on the other hand, it might decrease the pressure on low wages (see e.g. Lee and Saez (2012) for a discussion).

6 Conclusion

We scrutinize the effects of historic and recent event which transformed the German labor market on earnings dynamics in Germany by decomposing earnings' variances into a permanent and a transitory component. Using administrative data covering complete earnings life cycles of West German males born between 1935 and 1974, we can show how the profound changes of the German labor market affected inequality and stability over a period of 50 years. To model the evolution of earnings within individual life-cycles, we specify both a random walk and a random growth for the permanent and an AR(1)-process and a quartic age term for the transitory component. Next to these microeconomic dynamics, both components include period and cohort specific shifters to explicitly model macroeconomic dynamics and generational differences. In this regard the model leaves us with greater detail compared to approaches utilized in comparable studies. For the development of microeconomic dynamics across life cycles, we find an increasing permanent and about u-shaped transitory component. We also identify a trade-off between initially high earnings and subsequent earnings growth. While our results validate most of the findings from studies on shorter panels, we find that modeling extensive time frames requires explicit accounting for cohort specific differences.

Although we identify common life-cycle features, our main results stem from comparing volatility across generations. Looking at the evolution at different stages of the life-cycle, we find an upward trend for both transitory and permanent component. This finding is also commonly identified in studies on other countries despite differences regarding institutions, periods investigated, data used or methodology applied. The results mirror how some global long-term trends like declining manufacturing sectors, deunionization, increasing international economic integration and job polarization affect many Western societies. Still, first the unique situation following World War II and second the reunification with its both its financial obligations and the massive inflow of migrants make Germany a special case. The order of magnitude and explanatory power of the two components differs substantially across countries and time. For Germany, we find a strongly increasing transitory component at young ages with a trend starting in the early 1970s and intensification in the mid-1990s. For older workers, we find moderate increases. The permanent component starts increasing in the early 1980s and strongly increases in the early 2000s for all ages. Our results suggest that structural labor market changes affect both components and immediately translate into increasing short-term earnings risks especially for young workers. With delay, these structural changes also translate into increasing permanent divergences, especially at later ages.

The described trends of earnings dynamics are likely to continue and have several implications. Earnings risks upon labor market entry will remain high and it will become increasingly difficult to obtain stable employment. At later stages of the life cycle after labor market entry, permanent divergences will become more important. This implies increasing lifetime earnings inequality. Thus, although the flexibility gained through deregulation is deemed an important source of Germany's recent economic success, the downsides are rising insecurity and inequality. This burden, is carried mainly by the younger generations.

In general, short-term earnings risks are rather successfully mitigated by welfare state insurance (Bartels and Bönke, 2013). Mitigating increasing long-term disparities on the contrary would require increasing redistributive capabilities, which is currently unlikely. Therefore, permanent disparities are likely to gain even more importance in the future, reflecting in a continuing trend of rising lifetime earnings inequality. By the nature of our study, the most recent developments cannot be captured. The most interesting event is probably the introduction of the German minimum wage in 2015, which is left for future research.

Appendix A: Data

Bönke, Corneo and Lüthen (2015) provide further information on sample selection, dataset and indicators of the social security system; see also their online Appendix.

Table A.1: Sample descriptives

Cohort	Number of observations	Person years	Start year	End year	Last age observed	Years included	Moments	VSKT wave
1935	1005	33745	1960	1994	59	35	630	2002
1936	962	32168	1961	1995	59	35	630	2002
1937	993	33094	1962	1996	59	35	630	2004
1938	1026	34235	1963	1997	59	35	630	2005
1939	1054	34979	1964	1998	59	35	630	2006
1940	1025	33382	1965	1999	59	35	630	2007
1941	1072	34889	1966	2000	59	35	630	2008
1942	1035	33822	1967	2001	59	35	630	2009
1943	1029	33639	1968	2002	59	35	630	2009
1944	987	32267	1969	2003	59	35	630	2009
1945	1091	35327	1970	2004	59	35	630	2009
1946	1063	34377	1971	2005	59	35	630	2009
1947	1058	34124	1972	2006	59	35	630	2009
1948	1066	33971	1973	2007	59	35	630	2009
1949	1027	32483	1974	2008	59	35	630	2009
1950	1069	34434	1975	2009	59	35	630	2009
1951	1096	34083	1976	2009	58	34	595	2009
1952	1097	33218	1977	2009	57	33	561	2009
1953	1118	32942	1978	2009	56	32	528	2009
1954	1150	32930	1979	2009	55	31	496	2009
1955	1178	32803	1980	2009	54	30	465	2009
1956	1232	33209	1981	2009	53	29	435	2009
1957	1231	31854	1982	2009	52	28	406	2009
1958	1258	31696	1983	2009	51	27	378	2009
1959	1290	31012	1984	2009	50	26	351	2009
1960	1315	30498	1985	2009	49	25	325	2009
1961	1379	30855	1986	2009	48	24	300	2009
1962	1432	30706	1987	2009	47	23	276	2009
1963	1443	29352	1988	2009	46	22	253	2009
1964	1426	27798	1989	2009	45	21	231	2009
1965	1480	27487	1990	2009	44	20	210	2009
1966	1505	26387	1991	2009	43	19	190	2009
1967	1519	25195	1992	2009	42	18	171	2009
1968	1554	24304	1993	2009	41	17	153	2009
1969	1635	23791	1994	2009	40	16	136	2009
1970	1619	22138	1995	2009	39	15	120	2009
1971	1470	18764	1996	2009	38	14	105	2009
1972	1464	17354	1997	2009	37	13	91	2009
1973	1510	16576	1998	2009	36	12	78	2009
1974	1469	14797	1999	2009	35	11	66	2009
Total	49,432	1,200,685					17000	

Source: FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations.

Appendix B: Model estimation and robustness

(1) Estimation

After de-meaning the earnings, cohort specific variances and covariances are calculated and then stacked upon each other. This provides the vector of sample moments, $C = f(\theta)$. Then, we employ GMM to minimize the distance between this vector and the theoretical vector provided by the model parameters:

(B.1) $$Q = [C - f(\hat{\theta})]'W[C - f(\hat{\theta})]$$

As shown by Altonji and Segal (1996) and Clark (1996), the asymptotically optimal weighting matrix induces a bias in finite sample. Therefore, following e. g. Haider (2001) and Altonji and Segal (1996), we use the identity matrix as weighting matrix W. The estimation, often called equally weighted minimum distance, effectively becomes a nonlinear least squares estimation (Chamberlain, 1984). Standard errors are derived using the delta-method employing the fourth moments matrix. Standard errors are calculated with the delta-method, $V(\theta) = (G'G)^{-1}G'VG(G'G)^{-1}$, with V being the fourth moment matrix and G the gradient matrix derived from the estimation (e. g. Cappellari, 2004). Our dataset provides 17,000 sample moments used in the estimation procedure.

(2) Robustness

At first, we give a brief overview about the models shown in this section. Although the estimation of some parameters varies, our results of a shift in the variance components as well as our other results are qualitatively alike. Figures for all scenarios can be obtained from the authors upon request.

Model 1: This is the baseline model. See Section 3 in the main text for description.

Model 2: Here we estimate the baseline model on a different sample selection. We follow Bingley et al. (2013) and condition on 5 years of consecutive earnings.

Model 3: This model deviates from our baseline model in equation (7). We follow Baker and Solon (2003) and, instead of including cohort shifters for the transitory component, we estimate cohort specific initial variances $\nu_{i0} \sim (0; \sigma^2_{0,c})$. Equation (7) now becomes:

(B.2) $$\begin{aligned} Var(y^T_{ict} y^T_{ict}) = \rho^2 Var(\nu_{i,t-1}) \\ + \tau_t^2\big(\sigma^2_{\epsilon,0} + (t-c)\sigma^2_{\epsilon,1} + (t-c)^2\sigma^2_{\epsilon,2} + (t-c)^3\sigma^2_{\epsilon,3} \\ + (t-c)^4\sigma^2_{\epsilon,4}\big) \end{aligned}$$

Model 4: Because the oldest cohorts are included over their entire life cycle, we might face a bias in our estimation results due to their "overrepresentation." Therefore, we start estimating our model in 1979 instead of starting in the estimation in 1960. This leads to a decreased weight of the older cohorts in the model estimation. Still, we estimate similar shock persistence ρ. A comparison of the permanent and the transitory component does not show qualitative differences. Still, the core model parameter estimates are expected to differ because the cohort shifters are normalized to 1979 and not to 1960, as in the baseline model. Since our observation period is left-censored, the estimation of the initial transitory variance σ_0^2 might be biased. Therefore, we follow Moffitt and Gottschalk (2012) and estimate an additional parameter α for all left-censored cohorts. For left-censored cohorts, σ_0^2 is now included as follows: $(1 + \alpha \cdot age79)\sigma_0^2$. At this, $age79$ is the distance of the cohorts' age in 1979 and age 25. By way of an example, this bias-correction obtains 19 for cohort 1935 in the year 1979 and yields a transitory variance of $Var\left(y_{i1935,1970}^{T}\right) = (1 + 19\alpha)\sigma_0^2$.

Table B.1: Core model estimates

Coefficient	**Model 1**		**Model 2**		**Model 3**		**Model 4**	
	Coeff	SE	Coeff	SE	Coeff	SE	Coeff	SE
σ_μ^2	0.051	0.022	0.084	0.026	0.050	0.022	0.016	0.001
σ_u^2	2.65E-04	1.18E-04	9.02E-05	3.86E-05	2.70E-04	1.22E-04	8.32E-05	9.34E-06
σ_γ^2	0.0101	0.0043	0.0154	0.0048	0.0097	0.0042	0.0033	0.0003
$\sigma_{\mu\gamma}^2$	-3.28E-04	1.75E-04	-9.83E-04	3.23E-04	-3.17E-04	1.72E-04	-1.76E-04	3.79E-05
σ_0^2	0.100	0.057	0.123	0.067			0.089	0.015
$\sigma_{\epsilon,0}^2$	0.095	0.055	0.114	0.062	0.090	0.004	0.090	0.015
$\sigma_{\epsilon,1}^2$	-0.018	0.010	-0.013	0.007	-0.016	0.001	-0.018	0.003
$\sigma_{\epsilon,2}^2$	1.73E-03	9.97E-04	1.59E-03	8.88E-04	1.27E-03	5.78E-05	1.91E-03	2.82E-04
$\sigma_{\epsilon,3}^2$	-7.27E-05	4.20E-05	-7.96E-05	4.52E-05	-4.33E-05	2.09E-06	-8.54E-05	1.23E-05
$\sigma_{\epsilon,4}^2$	1.17E-06	6.81E-07	1.60E-06	9.27E-07	5.54E-07	2.73E-08	1.44E-06	2.05E-07
ρ	0.278	0.005	0.258	0.005	0.277	0.005	0.269	0.005
α							-0.055	0.012

Source: FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations.

Table B.2: Permanent cohort shifter

Cohort	Model 1: Base		Model 2: Adj. sample selection		Model 3: Cohort inital variances		Model 4: Shorted time frame	
	Coeff	SE	Coeff	SE	Coeff	SE	Coeff	SE
1935	1	-	1	-	1	-	1	-
1936	1.03	0.03	1.04	0.04	1.04	0.03	1.03	0.05
1937	1.04	0.03	1.04	0.04	1.04	0.03	1.03	0.04
1938	1.06	0.03	1.04	0.04	1.05	0.03	1.05	0.04
1939	1.16	0.04	1.10	0.04	1.15	0.03	1.17	0.05
1940	1.15	0.04	1.14	0.04	1.14	0.04	1.14	0.05
1941	1.20	0.04	1.17	0.04	1.19	0.04	1.18	0.05
1942	1.28	0.04	1.20	0.04	1.27	0.04	1.27	0.05
1943	1.34	0.04	1.26	0.04	1.33	0.04	1.33	0.05
1944	1.27	0.04	1.20	0.04	1.25	0.04	1.25	0.05
1945	1.31	0.04	1.23	0.04	1.30	0.04	1.30	0.05
1946	1.27	0.04	1.24	0.04	1.26	0.04	1.26	0.05
1947	1.41	0.04	1.29	0.04	1.40	0.04	1.40	0.06
1948	1.44	0.05	1.34	0.04	1.43	0.05	1.43	0.06
1949	1.56	0.05	1.43	0.04	1.56	0.05	1.55	0.06
1950	1.52	0.05	1.42	0.04	1.51	0.05	1.51	0.06
1951	1.65	0.05	1.49	0.04	1.65	0.05	1.64	0.07
1952	1.63	0.05	1.47	0.05	1.62	0.05	1.62	0.06
1953	1.70	0.05	1.56	0.05	1.70	0.05	1.69	0.06
1954	1.74	0.06	1.58	0.05	1.74	0.05	1.73	0.07
1955	1.75	0.06	1.52	0.05	1.76	0.06	1.74	0.07
1956	1.78	0.06	1.66	0.05	1.77	0.06	1.77	0.07
1957	2.00	0.07	1.77	0.06	1.98	0.07	1.99	0.08
1958	2.00	0.06	1.77	0.06	1.99	0.06	1.99	0.08
1959	2.13	0.06	1.87	0.06	2.13	0.06	2.12	0.08
1960	2.14	0.07	1.89	0.06	2.13	0.07	2.13	0.09
1961	2.13	0.07	1.85	0.06	2.12	0.07	2.11	0.08
1962	2.30	0.07	1.94	0.06	2.28	0.07	2.29	0.09
1963	2.46	0.08	2.08	0.07	2.46	0.08	2.45	0.10
1964	2.49	0.08	2.06	0.07	2.47	0.08	2.47	0.10
1965	2.51	0.08	2.06	0.07	2.48	0.08	2.49	0.10
1966	2.61	0.08	2.12	0.07	2.59	0.08	2.59	0.11
1967	2.54	0.08	2.07	0.07	2.54	0.08	2.53	0.11
1968	2.78	0.09	2.21	0.08	2.77	0.09	2.76	0.12
1969	2.80	0.09	2.26	0.08	2.77	0.09	2.78	0.12
1970	2.88	0.10	2.23	0.08	2.88	0.10	2.87	0.13
1971	3.00	0.11	2.32	0.09	3.03	0.11	2.98	0.14
1972	3.20	0.12	2.44	0.09	3.21	0.12	3.18	0.15
1973	3.26	0.12	2.43	0.09	3.27	0.12	3.25	0.15
1974	3.27	0.13	2.44	0.10	3.37	0.13	3.26	0.16

Source: FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations.

Table B.3: Transitory cohort shifter/Cohort specific transitory initial variances

Cohort	Model 1: Base Coeff	SE	Model 2: Adj. sample selection Coeff	SE	Model 3: Cohort inital variances Coeff	SE	Model 4: Shorted time frame Coeff	SE
1935	1	-	1	-	0.10	0.06	1	-
1936	1.08	0.02	1.08	0.02	0.14	0.08	1.11	0.02
1937	1.06	0.02	1.05	0.03	0.11	0.06	1.10	0.02
1938	0.95	0.02	0.97	0.03	0.13	0.08	0.94	0.02
1939	0.97	0.02	0.99	0.03	0.06	0.04	0.99	0.02
1940	0.97	0.02	1.05	0.03	0.17	0.06	0.95	0.03
1941	1.03	0.03	1.10	0.03	0.15	0.05	1.04	0.02
1942	1.01	0.03	1.08	0.03	0.14	0.05	0.99	0.03
1943	0.98	0.03	1.15	0.04	0.14	0.05	0.96	0.03
1944	1.05	0.04	1.18	0.04	0.09	0.07	1.07	0.03
1945	1.03	0.03	1.21	0.05	0.08	0.03	1.00	0.03
1946	1.07	0.04	1.27	0.05	0.08	0.03	1.07	0.03
1947	1.12	0.04	1.25	0.06	0.12	0.03	1.16	0.04
1948	1.21	0.05	1.38	0.07	0.11	0.04	1.24	0.04
1949	1.34	0.05	1.50	0.08	0.15	0.04	1.37	0.05
1950	1.16	0.05	1.47	0.08	0.09	0.02	1.22	0.05
1951	1.34	0.06	1.61	0.09	0.10	0.03	1.45	0.06
1952	1.39	0.07	1.66	0.10	0.12	0.04	1.46	0.07
1953	1.48	0.07	1.74	0.11	0.12	0.03	1.59	0.08
1954	1.49	0.07	1.79	0.12	0.14	0.02	1.61	0.09
1955	1.58	0.08	1.81	0.12	0.18	0.03	1.70	0.09
1956	1.43	0.08	1.91	0.14	0.09	0.02	1.55	0.09
1957	1.40	0.09	1.82	0.14	0.09	0.02	1.51	0.10
1958	1.48	0.10	1.96	0.16	0.11	0.01	1.62	0.11
1959	1.58	0.10	2.04	0.17	0.09	0.02	1.74	0.12
1960	1.58	0.10	2.08	0.17	0.09	0.01	1.75	0.12
1961	1.62	0.10	2.21	0.19	0.08	0.02	1.80	0.12
1962	1.49	0.11	2.13	0.20	0.10	0.01	1.66	0.12
1963	1.69	0.12	2.31	0.22	0.11	0.01	1.90	0.14
1964	1.55	0.11	2.29	0.22	0.07	0.01	1.75	0.13
1965	1.57	0.12	2.37	0.23	0.10	0.02	1.78	0.14
1966	1.69	0.12	2.40	0.24	0.12	0.02	1.92	0.15
1967	1.80	0.13	2.50	0.26	0.13	0.02	2.07	0.16
1968	1.82	0.13	2.51	0.27	0.09	0.01	2.09	0.16
1969	1.78	0.13	2.66	0.29	0.08	0.01	2.06	0.16
1970	1.93	0.14	2.83	0.32	0.09	0.01	2.24	0.17
1971	2.10	0.16	2.96	0.34	0.11	0.01	2.44	0.19
1972	2.06	0.16	3.11	0.37	0.10	0.01	2.40	0.20
1973	2.15	0.17	3.27	0.40	0.12	0.01	2.51	0.21
1974	2.44	0.20	3.54	0.44	0.15	0.01	2.86	0.24

Source: FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations.

Table B.4: Permanent period shifter

	Model 1: Base		Model 2: Adj. sample selection		Model 3: Cohort inital variances		Model 4: Shorted time frame	
Period	Coeff	SE	Coeff	SE	Coeff	SE	Coeff	SE
1960	1		1		1			
1961	0.84	0.13	0.81	0.09	0.84	0.14		
1962	0.77	0.13	0.82	0.11	0.77	0.14		
1963	0.86	0.15	0.83	0.11	0.87	0.16		
1964	0.72	0.13	0.73	0.10	0.72	0.13		
1965	0.63	0.12	0.68	0.09	0.65	0.12		
1966	0.65	0.12	0.69	0.09	0.66	0.13		
1967	0.77	0.15	0.77	0.11	0.79	0.16		
1968	0.75	0.14	0.77	0.10	0.77	0.15		
1969	0.72	0.14	0.70	0.09	0.73	0.14		
1970	0.65	0.13	0.65	0.09	0.66	0.13		
1971	0.62	0.12	0.64	0.09	0.63	0.13		
1972	0.61	0.12	0.61	0.08	0.62	0.12		
1973	0.58	0.11	0.64	0.09	0.59	0.12		
1974	0.59	0.12	0.64	0.09	0.60	0.12		
1975	0.63	0.12	0.68	0.10	0.64	0.13		
1976	0.61	0.12	0.66	0.09	0.62	0.13		
1977	0.60	0.12	0.65	0.09	0.61	0.12		
1978	0.58	0.12	0.63	0.09	0.59	0.12		
1979	0.55	0.11	0.61	0.08	0.55	0.11	1	
1980	0.52	0.10	0.58	0.08	0.53	0.11	0.95	0.01
1981	0.53	0.11	0.59	0.08	0.54	0.11	0.96	0.01
1982	0.55	0.11	0.60	0.08	0.55	0.11	0.99	0.01
1983	0.55	0.11	0.60	0.08	0.56	0.11	1.00	0.01
1984	0.60	0.12	0.63	0.09	0.61	0.12	1.09	0.01
1985	0.60	0.12	0.64	0.09	0.60	0.12	1.08	0.01
1986	0.59	0.12	0.63	0.09	0.60	0.12	1.07	0.01
1987	0.57	0.11	0.62	0.09	0.58	0.12	1.03	0.01
1988	0.55	0.11	0.61	0.09	0.56	0.11	1.00	0.01
1989	0.54	0.11	0.60	0.08	0.55	0.11	0.97	0.01
1990	0.52	0.10	0.58	0.08	0.53	0.11	0.94	0.01
1991	0.51	0.10	0.57	0.08	0.52	0.11	0.91	0.01
1992	0.49	0.10	0.55	0.08	0.50	0.10	0.89	0.01
1993	0.49	0.10	0.55	0.08	0.50	0.10	0.89	0.01
1994	0.48	0.10	0.54	0.08	0.49	0.10	0.86	0.01
1995	0.48	0.10	0.54	0.08	0.48	0.10	0.86	0.01
1996	0.46	0.09	0.53	0.08	0.47	0.10	0.84	0.01
1997	0.46	0.09	0.53	0.08	0.46	0.09	0.82	0.01
1998	0.45	0.09	0.54	0.08	0.46	0.09	0.82	0.02
1999	0.44	0.09	0.52	0.07	0.44	0.09	0.79	0.01
2000	0.43	0.09	0.52	0.07	0.44	0.09	0.78	0.02
2001	0.43	0.09	0.51	0.07	0.43	0.09	0.77	0.02
2002	0.42	0.09	0.51	0.07	0.43	0.09	0.76	0.02
2003	0.42	0.08	0.51	0.07	0.43	0.09	0.76	0.02
2004	0.41	0.08	0.51	0.07	0.42	0.09	0.75	0.02
2005	0.42	0.08	0.54	0.08	0.43	0.09	0.76	0.02
2006	0.42	0.08	0.53	0.08	0.42	0.09	0.76	0.02
2007	0.43	0.09	0.54	0.08	0.43	0.09	0.77	0.02
2008	0.42	0.09	0.54	0.08	0.42	0.09	0.76	0.02
2009	0.42	0.08	0.54	0.08	0.42	0.09	0.76	0.02

Source: FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations.

Table B.5: Transitory period shifter

Period	Model 1: Base Coeff	SE	Model 2: Adj. sample selection Coeff	SE	Model 3: Cohort inital variances Coeff	SE	Model 4: Shorted time frame Coeff	SE
1960	1		1		1			
1961	0.93	0.32	1.03	0.30	0.89	0.12		
1962	0.98	0.31	0.96	0.28	1.03	0.09		
1963	0.86	0.28	0.91	0.26	0.84	0.10		
1964	0.94	0.28	1.00	0.28	1.08	0.06		
1965	1.10	0.32	1.04	0.29	0.99	0.09		
1966	1.11	0.33	1.04	0.28	1.05	0.06		
1967	1.06	0.31	1.08	0.30	1.04	0.07		
1968	1.00	0.29	0.93	0.26	0.98	0.06		
1969	1.01	0.32	0.93	0.26	1.08	0.06		
1970	1.01	0.30	0.92	0.26	1.12	0.06		
1971	1.07	0.32	0.99	0.27	1.22	0.05		
1972	1.13	0.33	0.95	0.26	1.21	0.07		
1973	1.11	0.33	0.99	0.27	1.29	0.06		
1974	1.21	0.35	1.02	0.28	1.38	0.06		
1975	1.17	0.34	0.99	0.27	1.47	0.06		
1976	1.15	0.33	0.96	0.27	1.46	0.05		
1977	1.23	0.36	1.03	0.29	1.61	0.06		
1978	1.13	0.33	0.88	0.25	1.52	0.06		
1979	1.11	0.32	0.88	0.25	1.49	0.06	1	
1980	1.05	0.31	0.87	0.24	1.38	0.05	1.02	0.07
1981	1.10	0.32	0.85	0.24	1.67	0.05	1.07	0.07
1982	1.10	0.32	0.88	0.25	1.66	0.06	1.06	0.07
1983	1.31	0.38	0.92	0.26	1.93	0.05	1.25	0.09
1984	1.09	0.32	0.82	0.23	1.73	0.05	1.04	0.07
1985	1.10	0.32	0.83	0.23	1.76	0.05	1.04	0.07
1986	1.06	0.31	0.78	0.22	1.78	0.05	1.00	0.07
1987	1.09	0.32	0.73	0.21	1.73	0.04	1.02	0.07
1988	1.09	0.32	0.75	0.21	1.76	0.04	1.02	0.07
1989	1.07	0.31	0.71	0.20	1.85	0.05	0.99	0.07
1990	1.01	0.30	0.67	0.19	1.68	0.04	0.93	0.07
1991	1.00	0.29	0.67	0.20	1.64	0.04	0.92	0.07
1992	0.94	0.28	0.63	0.18	1.57	0.04	0.86	0.06
1993	0.98	0.29	0.63	0.18	1.80	0.04	0.89	0.07
1994	1.08	0.32	0.68	0.20	2.03	0.04	0.98	0.07
1995	0.99	0.29	0.64	0.19	1.91	0.04	0.90	0.07
1996	1.14	0.34	0.72	0.21	2.20	0.04	1.03	0.08
1997	1.15	0.34	0.72	0.21	2.30	0.05	1.04	0.08
1998	1.08	0.32	0.68	0.20	2.18	0.04	0.97	0.08
1999	1.10	0.33	0.68	0.20	2.24	0.05	0.98	0.08
2000	1.00	0.30	0.61	0.18	2.19	0.05	0.89	0.07
2001	0.97	0.29	0.59	0.18	2.14	0.04	0.87	0.07
2002	0.97	0.29	0.57	0.17	2.16	0.04	0.86	0.07
2003	0.97	0.29	0.59	0.18	2.21	0.05	0.86	0.07
2004	0.99	0.30	0.55	0.17	2.29	0.05	0.87	0.07
2005	0.94	0.28	0.56	0.17	2.29	0.05	0.82	0.07
2006	0.93	0.28	0.52	0.16	2.34	0.05	0.80	0.07
2007	0.88	0.27	0.51	0.16	2.31	0.05	0.76	0.07
2008	0.80	0.24	0.46	0.14	2.17	0.05	0.68	0.06
2009	0.91	0.28	0.49	0.15	2.45	0.00	0.76	0.07

Source: FDZ-RV – VSKT2002, 2004-2009_Bönke, own calculations.

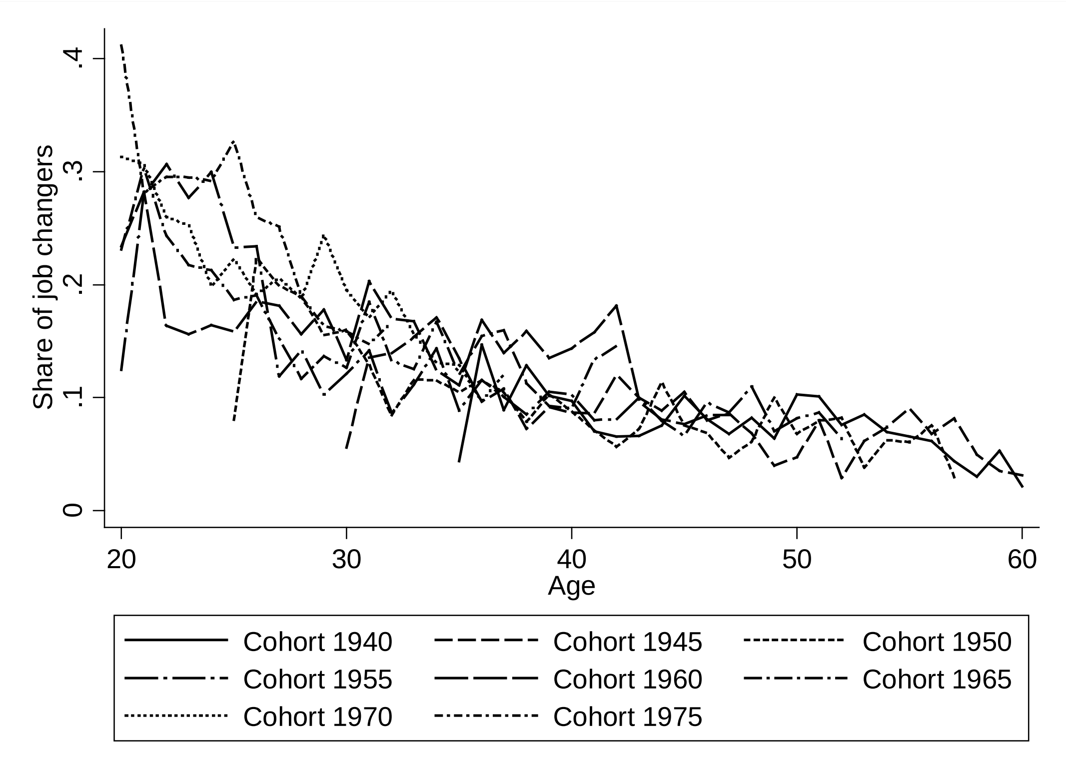
Appendix C: Supplements

Table C.1: Changes in labor market regulations since 1972

Year	Law (German Abbreviation)	Summary of Content
1972	AÜG	Permission of subcontracted work for up to 3 month
1985	BeschFG	1. Reduction of dismissal protection and weakening of standard employment contracts
		2. Introduction of fixed term contracts for first time employees (up to 18 month)
		3. Extension of maximum time for subcontracted work from 3 to 6 month
1990	BeschFG	Relaxation of justification requirements for fixed term contracts
1993	KündFG	Harmonization of employment protection (abolishment of special arrangements)
	1. SKWGP	Extension of maximum time for subcontracted work from 6 to 9 month
1996	BeschFG	1. Fixed term contracts can be applied multiple times
		2. Fixed term contracts enhanced to 24 month
		3. Further reduction of employment protection through the introduction of severance pay rules and for employees in small businesses
1997	ARFG	Extension of maximum time for subcontracted work from 9 to 12 month
1998	Gesetz zur Sicherung der Arbeitnehmerrechte	Rollback of employment protection legislation to the regulations in place prior to BeschFG 1996
2002	Job-AQTIV_Gesetz	Extension of maximum time for subcontracted work from 12 to 24 month
2003	Hartz 1	Abolishment of time limit for subcontracted work
		Reintroduction of employment protection legislation according to BeschFG 1996

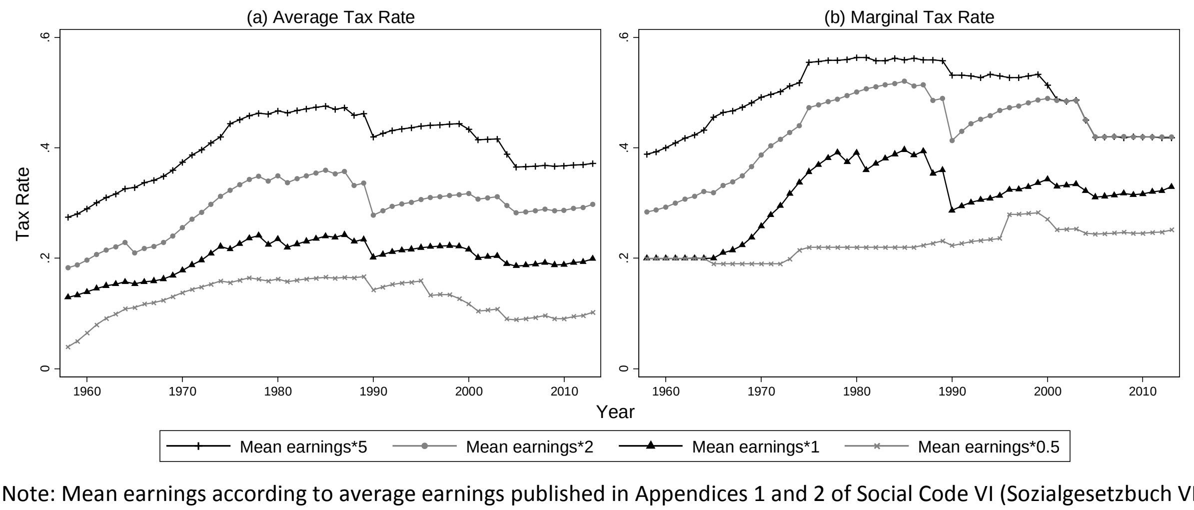
Source: Bundesgesetzblätter, various issues (available on request).
Note: Year is the year of parliamentary decision on passing the law, the entry of the law into force can deviate. For a more detailed overview on labor market regulation see Bartels (2014).

Figure C.1: Share of job changers by age and cohort



Source: Bartels et al. (2015)

Figure C.2: Average and marginal tax rates, 1958 - 2013



Note: Mean earnings according to average earnings published in Appendices 1 and 2 of Social Code VI (Sozialgesetzbuch VI), Federal Ministry of Labour and Social Affairs. Marginal and average tax rates on yearly wage income of unmarried employees without children.

References

Acemoglu, D., Aghion, P. and Violante, Giovanni L., (2001). Deunionization, technical change and inequality, Carnegie-Rochester Conference Series on Public Policy, 55(1), 229-264.

Altonji, Joseph G., and Segal, Lewis M. (1996). Small-Sample Bias in GMM Estimation of Covariance Structures. Journal of Business and Economic Statistics 14, 353–66.

Antonczyk, D., Fitzenberger, B. and K. Sommerfeld (2010). Rising wage inequality, the decline of collective bargaining, and the gender wage gap. Labour Economics 17 (5), 835 – 847.

Armingeon, K., L. Knöpfel, D. Weisstanner and S. Engler (2014). *Comparative Political Data Set I 1960-2012.* Bern: Institute of Political Science, University of Berne.

Baker, M. (1997). "Growth-Rate Heterogeneity and the Covariance Structure of Life-Cycle Earnings," Journal of Labor Economics 15(2), 338-75.

Baker, M. and G. Solon (2003). Earnings Dynamics and Inequality among Canadian Men, 1976-1992: Evidence from Longitudinal Income Tax Records. Journal of Labor Economics 21 (2), 267–288.

Bartels, C. (2014). Versicherung und Anreize im deutschen Wohlfahrtsstaat, Springer Gabler: Wiesbaden.

Bartels, C. and T. Bönke (2013). Can households and welfare states mitigate rising earnings instability?, Review of Income and Wealth 59(2), 250-282.

Bartels, C., T. Bönke and H. Lüthen (2015). Bildungs- und Jobmobilität in Deutschland. mimeo.

Bauer, T. K., Braun, S. and Michael Kvasnicka (2013). The Economic Integration of Forced Migrants: Evidence for Post-War Germany. The Economic Journal 123 (571), 998–1024.

Bauer, T. K., Dietz, B., Zimmermann, K. F. and Zwintz, Eric (2005): German Migration: Development, Assimilation, and Labor Market Effects. In: European Migration - What do We Know?, edited by Klaus F. Zimmermann, Oxford University Press, Oxford.

Bingley, P., Cappellari, L. and Westergård-Nielsen, N. (2013), Unemployment Insurance, Wage Dynamics and Inequality Over the Life Cycle. The Economic Journal 123 (568): 341–372.

Björklund, A. (1993). A comparison between actual distributions of annual and lifetime income: Sweden, 1951–89. Review of Income and Wealth 39, 377–86.

Blundell, R. and I. Preston (1998). Consumption Inequality And Income Uncertainty. The Quarterly Journal of Economics 113 (2), 603–640.

Blundell, R., M. Graber and M. Mogstad (2014). Labor income dynamics and the insurance from taxes, transfers and the family. Journal of Public Economics (forthcoming).

Bönke, T., G. Corneo, and H. Lüthen (2015). Lifetime Earnings Inequality in Germany. Journal of Labor Economics 33 (1), pp. 171–208.

Bowlus, A. J. and Robin, J.-M. (2012). An International Comparison of Lifetime Inequality: How Continental Europe Resembles North America. Journal of the European Economic Association

(10), 1236–1262.

Card, D., Heining, J. and P. Kline (2013). Workplace Heterogeneity and the Rise of West German Wage Inequality. Quarterly Journal of Economics 128 (3), 967 – 1015.

Castaneda, A., Diaz-Jimenez, J., Rios-Rull, J.V., 2003. Accounting for earnings and wealth inequality. Journal of Political Economy 111 (4), 818–857.

Cappellari, L. (2004). The Dynamics and Inequality of Italian Men's Earnings: Long-term Changes or Transitory Fluctuations? Journal of Human Resources 39 (2), 475–499.

Chamberlain, G. (1984). 'Panel data', in (Z. Griliches and M. Intriligator, eds.), Handbook of Econometrics, vol. 2, pp. 1247–318, Amsterdam: North-Holland.

Clark, Todd. "Small-Sample Properties of Estimators of Nonlinear Models of Covariance Structure." Journal of Business and Economic Statistics 14 (July 1996): 367–73.

Cuñat, A. and Melitz, M. (2012). Volatility, Labor Market Flexibility, And The Pattern Of Comparative Advantage. Journal of the European Economic Association 10(2), 225–254.

Dickens, R. (2000). The Evolution of Individual Male Earnings in Great Britain: 1975-95. Economic Journal 110 (460), 27–49.

Dustmann, C., J. Ludsteck and U. Schönberg (2009): Revisiting the German wage structure, Quarterly Journal of Economics 124, 843-881.

Dustmann, C., Fitzenberger, B., Schönberg, U. and A. Spitz-Oener (2014): From Sick Man of Europe to Economic Superstar: Germany's Resurgent Economy. Journal of Economic Perspectives 28 (1), 167-188.

Economist (2004). "Germany on the Mend." November 17. http://www.economist.com/node/3352024.

Ellgut, P. and Kohaut, S. (2005). Tarifbindung und betriebliche Interessenvertretung: Aktuelle Ergebnisse aus dem IAB-Betriebspanel. WSI Mitteilungen 7/2005, 398-406.

Ellgut, P. and Kohaut, S. (2008). Tarifbindung und betriebliche Interessenvertretung: Aktuelle Ergebnisse aus dem IAB-Betriebspanel 2007. WSI Mitteilungen 9/2008, 1-5.

Ellgut, P. and Kohaut, S. (2013). Tarifbindung und betriebliche Interessenvertretung: Ergebnisse aus dem IAB-Betriebspanel 2012. WSI Mitteilungen 4/2013, 281-288.

Fuchs-Schündeln, N. and M. Schündeln (2009): Who stays, who goes, who returns? Economics of Transition 17, 703–38.

Gottschalk, P. and R. Moffitt (1994). The Growth of Earnings Instability in the U.S. Labor Market. Brookings Papers on Economic Activity 25 (2), 217–272.

Gourinchas, P.-O., Parker, J.A., 2002. Consumption over the life cycle. Econometrica 70, 47–89.

Guvenen, F., 2007. Learning your earning: Are labor income shocks really very persistent? American

Economic Review 97 (3), 687–712.

Guvenen, F., 2009. An Empirical Investigation of Labor Income Processes. Review of Economic Dynamics 12(1), 58-79.

Gustavsson, M. (2008). A New Picture Of Swedish Earnings Inequality: Persistent And Transitory Components, 1960-1990. Review of Income and Wealth 54 (3), 324–349.

Guvenen, F., S. Ozkan, and J. Song (2014). The Nature of Countercyclical Income Risk. Journal of Political Economy 122 (3), 621 – 660.

Guvenen, F., F. Karahan, S. Ozkan, and J. Song (2015). What Do Data on Millions of U.S. Workers Reveal about Life-Cycle Earnings Risk? Federal Reserve Bank of New York Staff Reports, no. 710.

Fitzenberger, B. (1999): Wages and employment across skill groups: An analysis for West Germany, ZEW Economic Studies, Physika-Verlag: Heidelberg, New York.

Haider, S. J. (2001). Earnings Instability and Earnings Inequality of Males in the United States: 1967-1991. Journal of Labor Economics 19 (4), 799–836.

Hause, J.C. (1980). 'The fine structure of earnings and the on-the-job training hypothesis', Econometrica, vol. 48(4), pp. 1013–29.

Huggett, M., 1996. Wealth distribution in life-cycle economies. Journal of Monetary Economics 38, 469–494.

Jenson, S. and S. Shore (2015): Changes in the Distribution of Earnings Volatility. The Journal of Human Resources 50(3), 811-836.

Karahan, F. and S. Ozkan (2013). On the Persistence of Income Shocks over the Life Cycle: Evidence, Theory, and Implications. Review of Economic Dynamics 16 (3), 452–476.

Kohaut, S. and Schnabel, S. (2002). Tarifverträge – nein danke? IAB Materialien 1/2002, 4-5.

Kopczuk, W., E. Saez and J. Song (2010). Earnings inequality and mobility in the United States: Evidence from social security data since 1937. Quarterly Journal of Economics 125, 91–128.

Krugman, P. and A.J. Venables (1995). Globalization and the inequality of nations. The Quarterly Journal of Economics 110, 857–880.

Lee, D., & Saez, E. (2012). Optimal minimum wage policy in competitive labor markets. Journal of Public Economics 96(9), 739-749.

Lillard, L. and Weiss, Y. (1979). Components of variation in panel earnings data: American scientists, 1960–70, Econometrica, vol. 47(2), pp. 437–59.

MaCurdy, Thomas (2007). A Practitioner's Approach to Estimating Intertemporal Relationships using Longitudinal Data: Lessons from Applications in Wage Dynamics. In: Handbook of Econometrics, Vol. 6, J. Heckman and E. Leamer (eds), Elsevier.

Meghir, C. and Pistaferri, L. (2011). Earnings, consumption, and life cycle choices. In: Handbook of

Labor Economics', Vol. 4B, O. Ashenfelter and D. Card (eds), Elsevier.

Mincer, J. (1974). Schooling, Experience and Earnings, New York: Columbia University Press.

Moffitt, R. A. and P. Gottschalk (2002). Trends in the Transitory Variance of Earnings in the United States. The Economic Journal 112 (478), C68–C73.

Moffitt, R. and Gottschalk, P. (2012). Trends in the transitory variance of male earnings: methods, and evidence, Journal of Human Resources, vol. 47(1), pp. 204–36.

Ostrovsky, Yuri. 2010. Long-Run Earnings Inequality and Earnings Instability among Canadian Men Revisited, 1985-2005. The B.E. Journal of Economic Analysis & Policy. 10(1).

Shin, D. and Solon, G. (2011). Trends in men's earnings volatility: what does the panel study of income dynamics show?, Journal of Public Economics, vol. 95(7–8), pp. 973–82.

Sologon, D. and Van Kerm, P. (2014). Earnings dynamics, foreign workers and the stability of inequality trends in Luxembourg 1988-2009, CEPS/INSTEAD Working Paper Series 2014-03.

Storesletten, K., Telmer, C., Yaron, A., (2001). The welfare costs of business cycles revisited: Finite lives and cyclical variation in idiosyncratic risk. European Economic Review 45, 1311–1339.

Wood, A. (1995). How Trade Hurt Unskilled Workers. The Journal of Economic Perspectives 9 (3), 57-80.

Chapter 3:

Effectiveness of early retirement disincentives: individual welfare, distributional and fiscal implications

Abstract: In aging societies, information on how to reform pension systems is essential to policy makers. This study scrutinizes effects of early retirement disincentives on retirement behavior, individual welfare, pensions and public budget. We employ administrative pension data and a detailed model of the German tax and social security system to estimate a structural dynamic retirement model. We find that labor market participation and retirement behavior in general are strongly influenced by the level of disincentives. Further, disincentives come at the cost of increasing inequality and individual welfare losses. Still, net public returns are more than five times as high as monetarized individual welfare losses. Our estimates also suggest that similar levels of net public returns achieved by indiscriminating pension cuts are associated with individual welfare losses that are at least twice as high.

Keywords: dynamic discrete choice, retirement, tax and pension system, pension reform.

JEL Classification: C61, H55, J26

*This Chapter is joined work with Timm Bönke and Daniel Kemptner.

1 Introduction

Aging populations exert increasing financial pressure on pension systems around the globe. Therefore, this central feature of modern welfare states is, and has been, subject to many fundamental reforms. Typical examples include increasing eligibility ages (Mastrobuoni, 2009; Staubli and Zweimüller, 2013; Ataly and Barret, 2015), pension level adjustments (Haan and Prowse, 2014), and pension system restructuring (Laun and Wallenius, 2013).[63] Apart from debates fueled by the Great Recession, the imminent retirement of baby-boomer cohorts calls for fundamental reforms of old age security in most welfare states in the near future. Thus, evaluations on different pension reforms are highly relevant when discussing future pension policy design.

The German case is an excellent example. Until the late 1970s, the German pay-as-you-go (PAYG) system was expanded, becoming one of the world's most generous programs, both in terms of replacement rates and early retirement provisions. Population aging, German reunification, and high unemployment rates since the late 1970s, however, caused a rising fiscal imbalance. Since the early 1990s, the eligibility age has been increased, replacement rates have been lowered, and subsidies stimulating private old-age provisions have been introduced (e.g. Bönke et al., 2010). These reforms have direct implications for the financial situation of Germany's current and future pensioners. They alter the legal framework under which individual labor supply, retirement, savings, and fertility decisions are made (e.g. Börsch-Supan, 2000; Blundell, 2002). The effects are vast as statutory pensions account for about 85% of the average household disposable income for the elderly population (Börsch-Supan and Reil-Held, 2001). While many of these reforms have been undertaken in other countries in a similar fashion, some reforms are not yet fully investigated and deserve further attention.

This study scrutinizes disincentives that lead to permanent pension deductions and increase with the distance between the actual/early and normal retirement age. Since individuals still have a (limited) choice, disincentives differ from indiscriminating pension cuts or raising the legal eligibility age for early retirement. Further, from a theoretical perspective, Diamond and Mirrless (1978) find similar reforms to reduce moral hazard problems in the pension scheme. We contribute to the existing literature with an analysis of both actual and potential behavior to provide a detailed overview on effects of retirement disincentives. At this, we contrast positive effects on public finances to negative effects on affected individuals. To provide comprehensive evidence on disincentives in general, we model a broad range of disincentive levels. This range includes pension deductions of 0.3% per

[63] A broad overview of select reforms is provided by Gruber and Wise (2007).

month of early retirement, which were actually introduced through a major pension reform in Germany in 1992 (Hanel, 2010; Lüthen, 2015). We analyze to what extent disincentives are able to steer retirement behavior and provide evidence on distributional, individual welfare and fiscal implications of reducing pensions for early retirees. Typically for pension reforms, the institutional changes were phased in, impacting birth cohorts to different degrees. Thus, evaluation is not trivial due to the lack of intra-cohort variation. We incorporate comprehensive dynamic incentives of labor market participation and retirement behavior by estimating a structural dynamic retirement model (e. g. Rust and Phelan, 1997; Gustman and Steinmeier, 2015). Then, we model forward-looking agents who consider option values of possible retirement decisions and, thus, recognize the impact of their choices on the accumulation of pension wealth and future consumption possibilities.

Retirement disincentives give individuals the choice to retire within a certain time period at the cost of actuarial adjustments. When modeling retirement behavior, dynamic incentives are particularly relevant because individuals account for the entire future stream of pension benefits (Coile and Gruber, 2007). For an accurate estimation, we model the German tax and social security system in great detail and utilize high quality German administrative pension data. This enables us to disentangle other changes in the tax and pension system from the introduction of the disincentives, which induce cohort specific dynamic incentives. The inter-cohort variation in dynamic incentives helps identifying the structural parameters of our retirement model (e.g. Manoli et al., 2014). Then, based on the estimated parameters, we simulate a variety of economic outcomes for a number of counterfactual scenarios with changing levels of retirement disincentives.

For working males and the disincentive level of the 1992 reform, we find a retirement entry delay of 5.5 months. Increasing the disincentives causes further delay; a tripling of the 1992 disincentive level encourages most individuals to completely abandon early retirement. We also find disincentives to increase inequality in expected consumption, to cause individual welfare losses, and to lead to positive net public returns. All three outcomes increase with the disincentive level, although with diminishing marginal returns. The welfare losses are heterogeneously spread across the earnings distribution and greatest for medium income earners. Still, at each disincentive level, the net public returns are more than five times as high as monetarized individual welfare losses. Further, depending on disincentive level, net public returns can correspond to up to 16% of total pension expenditure per individual. It follows that early retirement disincentives are able to substantially increase the pension system's financial stability. Comparing disincentives to indiscriminating pension cuts, we find that at similar levels of net public returns, pension cuts result in individual welfare losses that are more than twice as high.

The remainder of the article is structured as follows. The next section describes the institutional setting in Germany and the data. Section 3 illustrates the conceptual framework. The core of the paper is Section 4, where we present our estimation results and conduct a policy analysis. Section 5 concludes.

2 Institutional setting and data

2.1 German pension scheme

The German statutory pension system is a pay-as-you-go system of Bismarckian variety. The great majority of employees is mandatorily insured, contributing a percentage of their income up to a contribution ceiling based on their gross wage. For their contributions, the insurants acquire pension entitlements in form of earnings (or remuneration) points. Earnings points are calculated as ratio of employee's wage to average wage. Hence, the number of earnings points corresponds to one (per year) if the employee's yearly wage corresponds to the average yearly wage. Over their working life employees accumulate earnings points until retirement. At retirement the individual pension level is calculated on the basis of these accumulated earnings points (EP). Thus, the pension level mirrors the length of the working life and the average position in the earnings distribution. The *pension formula* (§ 64, Sozialgesetzbuch VI) provides the details on how to calculate the monthly pension $p_{n,t}$ for individual n: [64]

$$p_{n,t} = A_t \cdot RA_n \cdot Z_n \cdot EP_n$$

where A_t corresponds to the *pension value*. Basically, the pension value is the amount of money that is multiplied with the sum of earnings points EP to calculate the monthly pension. The value is adjusted every calendar year (for an overview see Table 2 below). RA represents the pension type, which is 1 for old-age pensions. The factor Z is introduced by the 1992 reform to reflect the retirement age and the deductions due to early retirement: $Z = (1 - deduction)$.

The pension scheme offers various retirement possibilities depending on the retiree's individual situation. We focus on agents who have a choice between continuing to work and retirement, therefore abstracting from previously unemployed or disabled individuals. The individuals considered are able to claim the *normal old-age pension* at age 65 or the *pension for long-term insured* after age 63, which is conditioned on having spent at least 35 years in the pension system.[65] Retiring before

[64] Appendix A1 provides an overview on key institutional figures. For further details on the calculation of pensions in Germany see Lüthen (2015).

[65] We disregard individuals claiming old-age pensions for previously unemployed or disabled persons. These can be claimed at age 60 under different eligibility criteria like time spend in the pension system. These

age 65 is viewed as early retirement. Women are excluded due to their diverging pension prospects and the low number of cases when conditioning on similar early retirement eligibility. In sum, we concentrate on men with a strong labor market attachment who are eligible to retire at age 63, even if they choose to work longer.

2.2 Introduction of early retirement disincentives

In 1992, Germany introduced a major pension reform to equalize different retirement ages monetarily. The aim was to balance the pension wealth of early retirees and normal retirees. However, the budget relief was also needed to ensure stable contribution rates (e.g. Schmähl, 2011). Since early retirees have a prolonged benefit period, one possibility was to reduce their pension wealth. Therefore, the reform implemented permanent pension deductions of 0.3% per month of early retirement. The deduction level results from the distance (in month times 0.3%) between the actual retirement age and normal retirement age of 65.[66] Still, all cohorts were allowed to retire at 63. The deductions were gradually phased in for the 1937 and 1938 cohorts, then fully affecting those born thereafter. At this, the maximum deduction starts at 0.3% for those born in January 1937 and increases by 0.3 % points per month of birth up to 7.2% for cohorts born after 1938. Thus, the individuals born during the phase-in are only partially affected by the reform. Table 1 provides an overview and exemplary date of birth examples.

Table 1: Phase-in of disincentives

Date of birth	Retirement age without deductions	Maximal deduction (month)	Maximal deduction (share)
Before 1937	63	0	0%
January 1937	63+1 month	1	0.3%
June 1937	63+6 month	6	1.8%
January 1938	64+1 month	13	3.9%
June 1938	64+6 month	19	5.7%
After 1938	65	24	7.2%

Note: The maximal deduction (share) determines the age factor Z in the pension formula. Source: SUFVSKT2002, 2004-12.

2.3 Data

To calculate pension entitlements as described above, the pension insurance collects information on all contributors' earnings biographies. The dataset we use, the Insurance Account Sample

"waiting periods" consist of periods of contributions, wage replacement benefits (unemployment, sick-pay, invalidity), child-raising and times of education. A detailed overview on eligibility and pension types is provided in Lüthen (2015).

[66] See Lüthen (2015) for further details. The reform also introduces a pension bonus of 0.5 % per month retiring after 65, but this affects only a negligible amount of individuals. Due to dominance of collective bargaining for cohorts considered, most contracts force workers to retire at 65.

(*Versicherungskontenstichprobe*, VSKT), is a stratified random sample of these records. Each wave contains information on individuals aged between 30 and 67 in the reference year.[67] From age 14 through age 65, the VSKT provides a monthly history of employment, unemployment, sickness, and earnings points. The latter are used to compute monthly gross earnings. The total sum of earnings points provides the foundation for calculating gross pensions. To obtain net incomes, we subtract taxes and social security contributions. We account for all regulations and changes affecting monthly disposable income and pensions. An additional scenario that implements the regulations of the first year considered in this study, 1998, for all later years is also estimated. This allows the disentanglement of other changes in the tax and pension system from changes induced by the disincentives. For further details, see Figure 2 and Appendix A.

To ensure early retirement eligibility, we restrict the sample to those who have spent at least 35 years in the pension system before turning 63. This also ensures that the sample does not include individuals with substantial labor market earnings unnoticed by the Federal Pension Insurance (i.e. self-employed, civil servants, or long-term emigrants). Further, we exclude individuals who have worked in the German Democratic Republic (GDR; the former East Germany). For the cohorts considered, neither the labor market situation nor working life is comparable to the West German context.[68] The final sample contains 945 individuals (Table 2).

While German social security data records earnings very accurately, one major drawback is the top coding of earnings information at the contribution ceiling. For a better approximation of true distribution of earnings above the ceiling, we impute earnings of all individuals affected by top coding. The imputation method is based on the assumption of Pareto-distributed earnings in the upper tail of the distribution.[69] Further, the VSKT lacks information on other income sources, wealth or household context. A comparison with survey data reveals that these limitations are not harmful. The considered group of individuals receives income almost exclusively from statutory pensions and wages.[70] This income also accounts on average for more than 90% of their total household income (Table A4). Nearly 80% are married. Accordingly, we assume a married single-earner household and

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[67] We use the scientific use files for on-site-use (waves SUFVSKT2002 and SUFVSKT2004 to SUFVSKT2012), provided to researchers by the Data Research Center of the German Federal Pension Insurance. We use all 10 waves in our analysis (see Appendix B for further information).

[68] West-East migration only affects an empirically negligible share of the population (Schündeln and Schündeln, 2009).

[69] Bönke et al. (2015) provide a detailed description of the imputation procedure in Online Appendix III.3. They find that, on average, top coding affects 7 % of all West German men in the VSKT.

[70] Tables A3 and A4 provide information on the relevance of different income sources and marital status. For the overall population, Bönke et al. (2015) document that the VSKT represents about 80% of the total male labor force in West Germany and that its cross-sectional earnings distributions are similar to those found in survey data.

joint taxation. For robustness, we also calculate a scenario assuming only single households (Tables B1 and B2). Our results are robust to this assumption.

Table 2 provides key descriptives of the sample. Column 1 shows that the observed average retirement entry age increases by about 8 months across cohorts. Column 2 reveals declining average pensions in real terms, although pension entitlements remain stable across cohorts (column 3). Columns 4 and 5 add further insights to this development by showing the pension value and the average amount of disincentives. At age 65, the pension value slightly increases up to cohort 1937 and then decreases for later cohorts (calendar years 2000-2010).[71] The column "disincentives" gives the average deduction on the monthly pension realized by each cohort. For fully affected cohorts, the average deduction fluctuates between 2.6% and 4.3%. All changes are accounted for when modeling the institutional background.

Table 2: Sample descriptives

Cohort	Entry age	Monthly pension	Earnings points	Pension value at age 65	Disincentives in %	Number of observations
1935	63.55	1680.97	57.73	28.91	0.00	53
1936	63.67	1660.76	55.98	28.71	0.00	43
1937	63.61	1636.40	55.72	29.18	1.06	50
1938	63.75	1565.26	54.42	29.03	3.70	72
1939	63.89	1607.84	56.33	28.70	4.28	84
1940	64.03	1558.46	54.84	28.27	3.77	93
1941	64.06	1564.30	55.85	27.83	3.69	77
1942	64.32	1580.62	56.08	27.28	2.67	95
1943	64.36	1574.17	55.01	26.81	2.56	122
1944	64.31	1555.39	54.57	27.18	2.73	115
1945	64.23	1562.70	55.82	27.20	3.08	141

Note: The average pensions and the pension values are in 2010 Euro values. The numbers of observations represent the final sample. Source: SUFVSKT2002, 2004-12, Deutsche Rentenversicherung (2014), (own calculations).

[71] Figure 1 and Appendix A provide estimations on the effect of these changes.

3 Model and estimation

3.1 Dynamic retirement model

In the following we introduce our theoretical framework step by step. The model aims to explain retirement behavior in a time frame of 24 months, namely between age 63 and age 65. Still, implications for future years spent in retirement are also accounted for and correspond to individuals' particular choices. We rely on a theoretical framework where agents' utility in period t depends on consumption c and disutility of labor l. The total number N of individuals is indexed by n. Discrete time is measured in months t, running up to age 100 (period T). Then, t also expresses individual age, where $t = 0$ corresponds to the month an agent turns 63. Consumption in t for individual n equals net income flow from earnings, pensions or social security transfers, and disutility of labor depends on working or not in t. Agent n's utility in month t is then:

(1) $$u(c_{nt}, l_{nt})$$

Further, we assume risk averse agents. Current and future consumption possibilities in month t depend on earnings biography and choices until the current period, whereas disutility of labor is allowed to vary in age t. Equation (1) becomes:

(2) $$u(c(\boldsymbol{s}_{nt}, d_{nt}), l(d_{nt}, t_n))$$

where $\boldsymbol{s}_{nt}$ denotes a vector of state variables (age, birth cohort, accumulated pension points, gross wage, and previous period's choice) and $d_{nt} \in \{0,1\}$ is a dummy variable indicating the retirement choice. Hence, $c(\boldsymbol{s}_{nt}, d_{nt})$ denotes the level of consumption associated with state $\boldsymbol{s}_{nt}$ and choice d_{nt}. Due to our short time frame of two years, we abstract from private savings. Thus, individual disposable income corresponds to consumption in the respective period.[72] Disutility of labor $l(d_{nt}, t_n)$ is both a function of d_{nt} (since there is no more disutility of labor after retirement) and age t_n. For the explicit form, we assume a time separable random utility model representing individual preferences that satisfy our assumptions on consumption and disutility of labor:

[72] Before retirement, individuals earn a gross wage. In case of early retirement, some monthly wages between ages 63 and 65 are unobserved. We impute the counterfactual wage relying on the last real wage observed in the respective month of the previous year. This corresponds to the wage observed 12 or 24 month before the imputation. This accounts for monthly wage volatility.

$$u(c_{nt}, l_{nt}) = \alpha \frac{c(\boldsymbol{s}_{nt}, d_{nt})^{(1-\rho)} - 1}{(1-\rho)} + l(d_{nt}, t_n) + \epsilon_{nt}(d_{nt}) \tag{3}$$

We assume the random component $\epsilon_{nt}(d_{nt})$ to be type 1 extreme value distributed. The random component represents individual utility shocks not observed by the researcher. ρ depicts the coefficient of relative risk aversion and α a consumption weight. To allow the disutility of labor to vary in age, age enters $l(d_{nt}, t_n)$ as a linear spline function. The function is allowed to change the slope every three months of age to ensure a flexible specification:

$$l(d_{nt}, t_n) = \delta_1[1 - d_{nt}] + [1 - d_{nt}] \begin{Bmatrix} & \delta_2\, min(t\ \ ,3) & if\ t > 0 \\ + & \delta_3\, min(t - \ 3,3) & if\ t > 3 \\ + & \delta_4\, min(t - \ 6,3) & if\ t > 6 \\ \vdots & \vdots & \vdots \\ + & \delta_9\, min(t - \ 21,3) & if\ t > 21 \end{Bmatrix} \tag{4}$$

When timing retirement decisions, agents are forward looking and maximize their expected lifetime utility according to their preferences constrained by the institutional setting. Here, for each month between ages 63 and 65, agents decide between continuing to work or retirement. When continuing to work, utility stems from consumption only, but individuals experience disutility of labor. After retirement, agents receive utility from consumption only. In line with the rules and regulations of the pension system, working individuals accumulate pension claims proportional to real wages. This creates dynamic incentives for individuals taken into account by the dynamic choice framework. Retirement is an absorbing state and agents are not allowed to return to work, making utility maximization an optimal stopping problem. Earliest possible retirement choice is at $t = 1$ in the month following the 63rd birthday, latest possible early retirement decision is at age 64, month 12 ($t = 24$). Each month t, individual n observes state variables $\boldsymbol{s}_{nt}$ and makes retirement choice d_{nt} to maximize expected lifetime utility E. We define $D(\boldsymbol{s}_{nt})$ to be the choice set available to individual n in period t and to contain the choice between employment and retirement:

$$\max_{d_{nt} \in D(\boldsymbol{s}_{nt})} E\left[\sum_{j=0}^{T-t} \theta_{bt+j}\, \beta^j u\big(c_{nt+j}, l_{nt+j}\big)\right] \tag{5}$$

with β denoting a monthly subjective time discount factor, which we derive from a yearly discount factor of 0.96 (Gourinchas and Parker, 2002). To accommodate our monthly setting, we implement $\beta = \sqrt[12]{0.96}$. θ_{bt+j} indicates individual probabilities of being alive in period $t + j$, conditional on survival until period t and belonging to cohort b. Cohort specific mortality rates ensure a realistic setup and also help identifying parameters in the estimation procedure by inducing cohort-specific heterogeneity in dynamic incentives.[73]

We further define a Markov transition function $q(\boldsymbol{s}_{nt+1}|\boldsymbol{s}_{nt}, d_{nt})$ to capture individual beliefs about future states. Since $\boldsymbol{s}_{nt+1}$ evolves from state variables and agents are assumed to have perfect foresight about future states, $q(\boldsymbol{s}_{nt+1}|\boldsymbol{s}_{nt}, d_{nt})$ is a deterministic function. The only function not evolving deterministically is the utility shock $\epsilon_{nt}(d_{nt})$, which is not regarded as a state variable. Therefore, agents' maximization problem corresponds to the following value function $v(\boldsymbol{s}_{nt})$:

(6)
$$v_t(\boldsymbol{s}_{nt}) = \max_{d_{nt} \in D(\boldsymbol{s}_{nt})} \Bigg\{ u(c_{nt}, l_{nt}) + \theta_{bt+1}\beta \int_{\epsilon} \left[\sum_{S(\boldsymbol{s}_{nt})} v(\boldsymbol{s}_{nt+1}) q(\boldsymbol{s}_{nt+1}|\boldsymbol{s}_{nt}, d_{nt}) \right] g(\boldsymbol{\epsilon}_{nt+1}) \Bigg\}$$

where $g(\cdot)$ represents a multivariate probability density function of the random components. $S(\boldsymbol{s}_{nt})$ contains all possible different states in $t+1$ given state $\boldsymbol{s}_{nt}$. The difference in the expected discounted future utility between working and not working reflects option values of respective choices.

3.2 Choice probabilities and estimation

This section features the model estimation. Given the finite horizon of the individuals' optimization problem, it can be solved recursively. Starting point is the expected value function $V(\cdot)$ for particular choice options in the last period T. $V(\cdot)$ needs to be computed for all possible choices. In the last period T, it corresponds to

(7)
$$V(\boldsymbol{s}_{nT}, d_{nT}) = E[u(c_{nT}, l_{nT})]$$

By Bellman's principle of optimality, the individual's optimization problem can be written as a two-period problem for all other time periods t, which take into account the optimal decision for $t+1$. Due to the type 1 extreme value distribution of utility shock $\epsilon_{nt}(d_{nt})$, the expected value function has a closed form solution (Rust, 1987):

(8)
$$V(\boldsymbol{s}_{nt}, d_{nt}) = E[u(c_{nt}, l_{nt})]$$

[73] To account for increasing life expectancy, we use official mortality tables supplying cohort-specific projections (Statistisches Bundesamt, 2006).

$$+\theta_{bt+1}\beta \sum_{s_{nt+1}} \log\left\{ \sum_{d_{nt+1} \in D(s_{nt+1})} exp\big(V(\boldsymbol{s}_{nt+1}, d_{nt+1})\big) \right\} q(\boldsymbol{s}_{nt+1} | \boldsymbol{s}_{nt}, d_{nt})$$

Computation of expected value functions between mandatory retirement (age 65) and T is comparatively simple as individual choices are limited until age 65. Thereafter, real net income streams remain constant. Rust (1987) shows that when assuming additive separability and conditional independence of utility shocks, conditional choice probabilities have a closed form solution (here mixed logit probabilities):

(9) $$Prob(d_{nt} | \boldsymbol{s}_{nt}) = \frac{exp\big(V(\boldsymbol{s}_{nt}, d_{nt})\big)}{\sum_{j \in D(s_{nt})} exp\big(V(\boldsymbol{s}_{nt}, j)\big)}$$

The model is estimated by maximum likelihood. The log-likelihood function of the sample is given by

(10) $$\sum_{n=1}^{N} \sum_{t=1}^{T} \log\left\{ \sum_{d_{nt}} Prob(d_{nt} | \boldsymbol{s}_{nt}, \boldsymbol{\lambda}) \times I(d_{nt}) \right\}$$

with $I(d_{nt})$ indicating the individual choice observed in period t and the vector $\boldsymbol{\lambda} = (\alpha, \rho, \delta_1, .., \delta_9)$ containing all parameters of the utility function. The likelihood contributions then correspond to the respective conditional choice probabilities, abstracting from random transitions of state variables. For robustness, we also estimate a model specification allowing for unobserved heterogeneity in δ_1 (Heckman and Singer, 1984). We further include a robustness test where unobserved types are modeled as a function of lifetime earnings until age 63 to account for a possible correlation between leisure preferences and employment history (Wooldridge, 2005). Still, neither the central preference parameter ρ nor any of our postestimation outcomes are sensitive to these extensions.[74]

3.3 Parameter estimates and model fit

An overview on parameter estimates is displayed in Table 4. Our estimate of the relative risk aversion, $\rho = 1.5$, is in line with previous studies (see e.g. Chetty, 2006), although our identification is based on retirement choices only. The estimates of the spline function $l(\cdot)$ indicate the results typically found in the literate: a spiking retirement hazard at early eligibility and normal retirement age that cannot be explained entirely by incentives but rather mirrors institutional constraints (e.g. Coile and Gruber, 2007). Here, this is reflected by the high negative estimate for δ_1 (mitigated by δ_2 when continuing to work) and the high positive estimate for δ_9. All estimates are independent of

[74]Although we can identify two types, the second type is estimated to make up only a small fraction of the population and precision of the respective type-specific parameter is low. Initial conditions exert no significant effect on type probabilities. See Appendix B for results and details on the estimation.

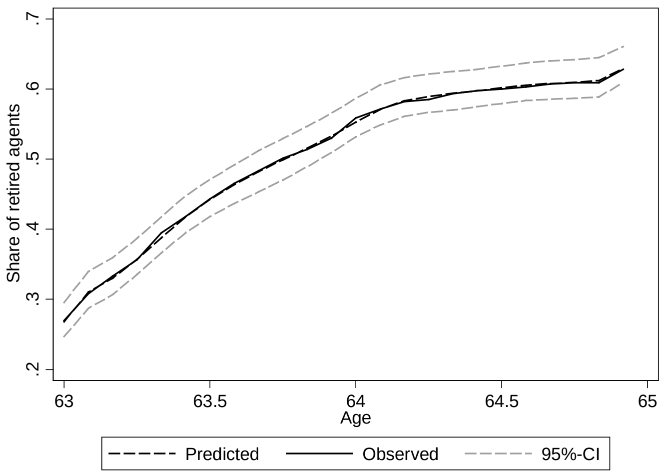
their starting values and the small standard errors indicate precise estimation. In the following, confidence intervals of postestimates are computed by applying a parametric bootstrapping method. Based on the inverse of the Hessian of the log-likelihood function, the procedure relies on 200 draws from the asymptotic sampling distribution of the estimated model parameters. Figure 1 compares predicted and observed shares of retirees by age and shows a very good internal validity.

Table 3: Parameter estimates

α	0.215	δ_3	-0.164	δ_7	0.255
	(0.0241)		(0.0665)		(0.1281)
ρ	1.506	δ_4	0.032	δ_8	-0.232
	(0.0716)		(0.0647)		(0.1411)
δ_1	-3.015	δ_5	0.080	δ_9	1.212
	(0.2476)		(0.0733)		(0.1343)
δ_2	1.057	δ_6	-0.261		
	(0.1182)		(0.0916)		
Log-likelihood	-1942.31				

Note: Standard errors in parenthesis.

Figure 1: Comparison of predicted and observed shares of retirees



Source: SUFVSKT2002, 2004-12

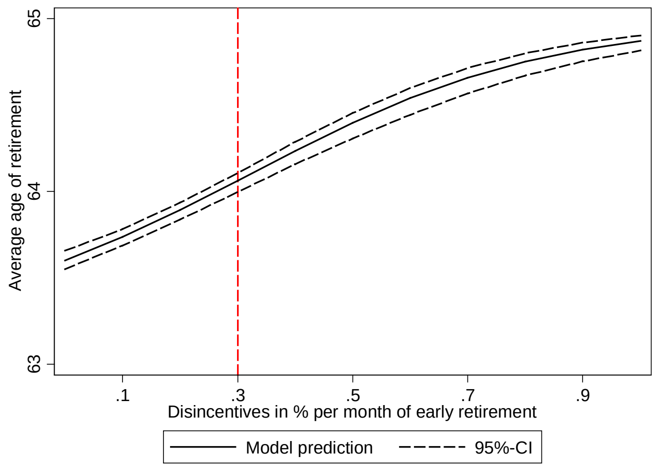
4 Results and policy analysis

To analyze the economic effects of retirement disincentives in general, we simulate scenarios for different disincentive levels. Those levels range from 0% to 1% per month of early retirement. We set the distance between each disincentive level to 0.1%, resulting in 10 counterfactual scenarios. Unless stated otherwise, the results are based on cohorts 1939 to 1945 which are fully affected by the 1992 reform. This ensures meaningful comparisons among counterfactual scenarios. We present predominantly graphical results; the actually implemented disincentive level of 0.3% per month is marked with a vertical dashed line. In sum, this section sheds light on the “dose-response” relationship between disincentive level and outcome measure and still includes a full analysis of the 1992 reform.

A. Labor market effects

Here we look at the effects of disincentives on labor market exit timing. Figure 2 displays a concave relationship between average retirement age and disincentive level. This suggests that disincentives can be used to steer retirement behavior. While low disincentive levels lead to small postponements in retirement, high levels induce most individuals to retire at age 65 such that hardly any penalties are actually realized (“prohibitive effect”). The actually implemented level causes a postponement of about 5.5 months, whereas the highest disincentive level would have delayed retirement by about 15 months.

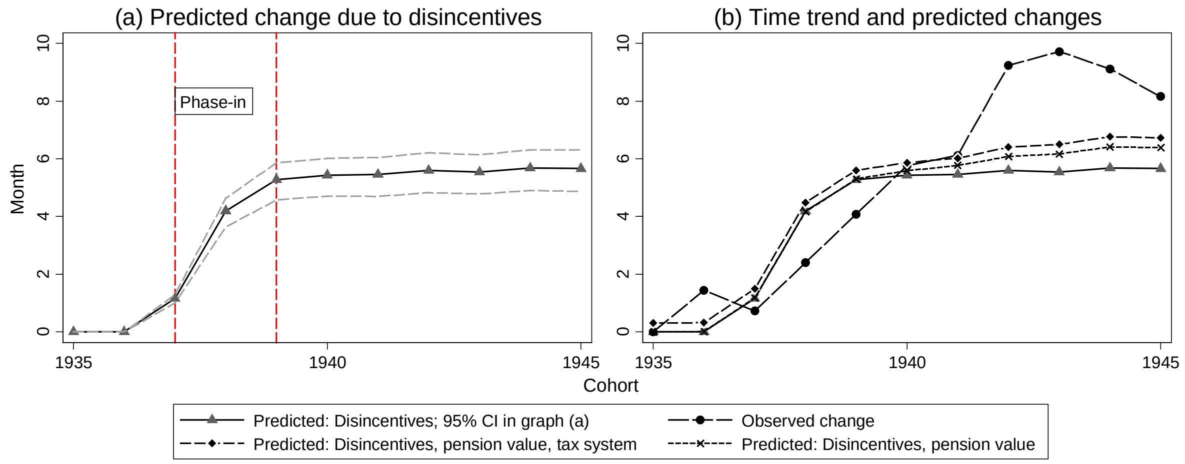
Figure 2: Expected retirement age by disincentive level



Source: SUFVSKT2002, 2004-12

To look more closely at the labor market effect, Figure 3 provides changes in expected retirement age by birth cohort. We focus on the 0.3%-disincentive level introduced by the 1992 reform and show how much of the cohorts’ retirement postponement can be attributed to the reform. Panel (a) displays predicted changes due to the introduction of disincentives. Panel (b) additionally shows observed changes from cohorts 1935 to 1945 as well as predicted changes attributed to disincentives, the pension value and the tax system. Panel (a) indicates that the disincentives delay average retirement entries by 5.5 months. This finding is stable across the fully affected 1939 to 1945 cohorts. We identify smaller effects for cohorts 1937 and 1938, which were affected by the reform’s phase-in. Comparing observed entries and predictions across cohorts, the introduction of disincentives explains 68% of the observed change in retirement patterns. Panel (b) demonstrates that changes in tax system and pension value delay retirement by an additional month. In total, the model explains about 80% of the average increase in retirement age between the 1935 and 1945 cohorts. This total predicted change relies on a counterfactual where tax and pension legislation from 1998 hold for all agents. This mirrors the institutional setting of agents born in 1935 at age 63, which is the point of their first decision about early retirement.

Figure 3: Reform effects on expected retirement age by birth cohort



Source: SUFVSKT2002, 2004-12

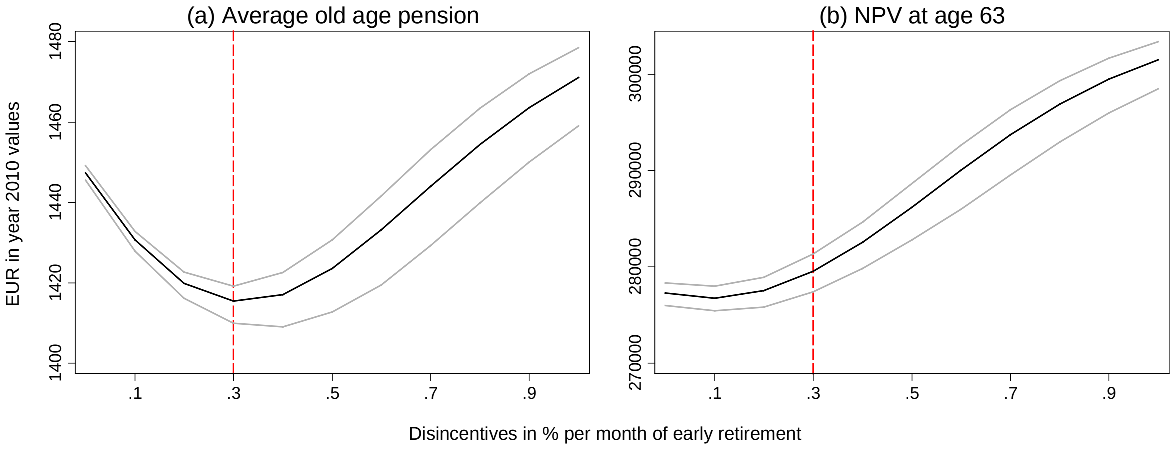
B. Financial implications

Here we analyze disincentive effects on pension level and NPVs of remaining lifetime consumption.[75] The individual pension level is affected by two countervailing effects induced by the reform. First, early retirement entails a penalty on pension benefits. Second, individuals delay exiting the labor market and receive labor earnings for a longer period of time (notice that wages exceed pension benefits in most cases). More contributions then translate into higher pension claims. Thus, the behavioral effect of delayed retirement is able to counteract the disincentive effect at some point. With that in mind, it is not surprising that Figure 4 shows a u-shaped relationship between pension and disincentive level. The actually implemented disincentive level of 0.3% per month yields the lowest average pension — both reducing and enhancing the disincentive level increases average retirement income. When decreasing the disincentive level, the behavioral reactions are small but pensions still rise. When increasing, the behavioral effect outweighs the penalty effect and pensions increase. For the actually implemented level of the 1992 reform, we find that pensions decrease by €32 per month. Put another way, the average individual loses a bit more than the equivalent of one year of average pension entitlements (i.e. one earnings point). Figure 6 also shows a similar relationship between disincentive level and remaining lifetime consumption. Interestingly, the lowest

[75] The NPV constitutes the sum of the discounted expected consumption stream at age 63.

NPV realized is associated with a disincentive level of 0.1%. At a level of 0.2%, the increases in labor market earnings start to outweigh the decreases in pension level.

Figure 4: NPVs of expected consumption and retirement income by disincentive level



Note: Euro in 2010 real values. Source: SUFVSKT2002, 2004-12

C. Individual welfare effects

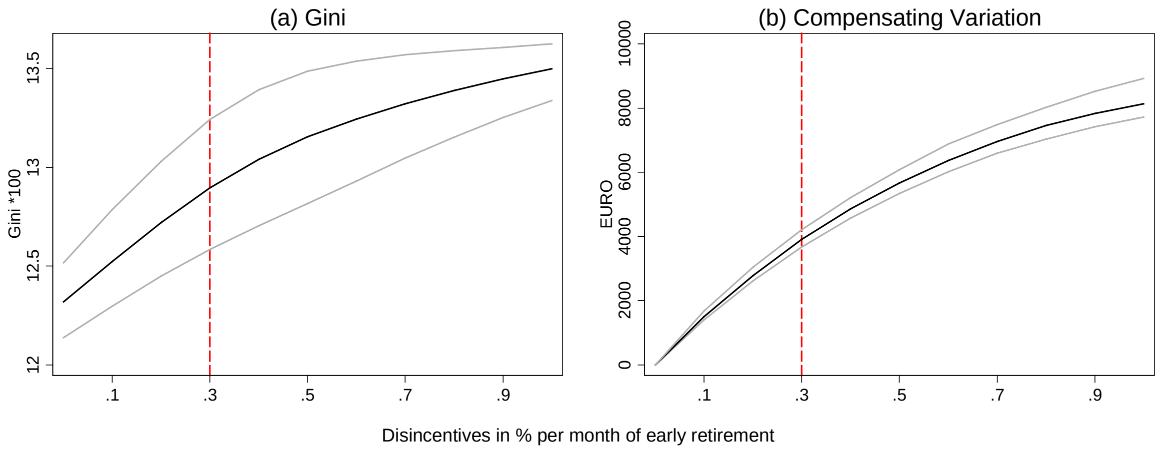
Figures 5 displays potential costs associated with the considered range of disincentive levels. Panel (a) provides estimates on increasing inequality in remaining lifetime consumption (Gini coefficient). Without disincentives retirement behavior is more heterogeneous, which offsets some initial inequalities in pension claims at age 63. Panel (b) assesses expected individual welfare losses (compensating variations, CV). [76] The variations refer to NPVs at age 63 that are annuitized over the remaining lifetime. Obviously, individuals who would have worked until age 65 even without disincentives are unaffected. The estimates provide quantifications for the average decline in individual welfare and further allow a disaggregated analysis of individual welfare losses along the income distribution.

Both Gini and CV show a concave relationship to the level of disincentives. Increases at low disincentives levels cause large increases in Gini and CV. Increases at higher disincentive levels have smaller effects since at high disincentive levels, the average retirement age is close to 65 already (see Figure 2). We find that the relationship flattens out around a disincentive level of 0.7%. For the highest considered disincentive level, the overall effects amount to twice the effects attributed to

[76] A compensating variation (CV) indicates the amount of money that an individual would have to receive at age 63 to be fully compensated for a particular reform. Here, to compute CVs, we employ an iterative algorithm targeting the expected remaining lifetime utilities at age 63 without retirement disincentives. The algorithm converges when the differences in individuals' expected utilities under both scenarios (disincentives and no disincentives) are very small. The payment is then annuitized over the remaining lifespan.

the actually implemented reform (vertical dashed line) – a 10% increase in the Gini and a CV of about €8000.

Figure 5: Gini and CVs by disincetive level

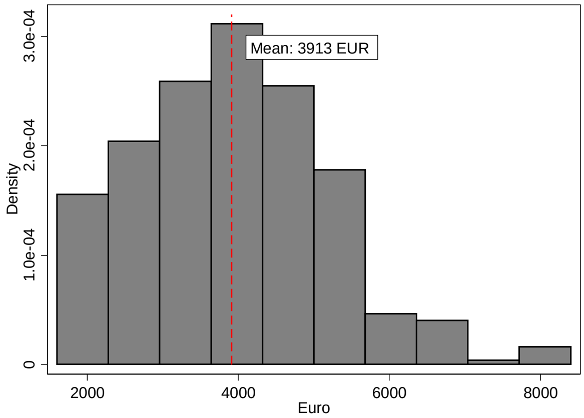


Note: Euro in 2010 real values. Source: SUFVSKT2002, 2004-12

For a detailed analysis, we again focus on the 0.3% disincentive level implemented by the 1992 reform. Figure 6 concentrates on the CV and reveals that individual welfare losses are heterogeneously distributed in the sample population, ranging from negligible amounts up to almost €9,000. This complicates compensation through e.g. saving subsidies because such a scheme may not allow for the targeting of individuals according to their specific losses.[77] Figure 7 shows a non-parametric regression of estimated compensating variations on NPVs of expected consumption. The results suggest that medium income earners lose most through the introduction of retirement disincentives. This is driven by earnings-level heterogeneity in the expected retirement age. Low and high income individuals tend to retire closer to age 65 regardless, which is due to low pension claims and high opportunity costs of retirement, respectively.

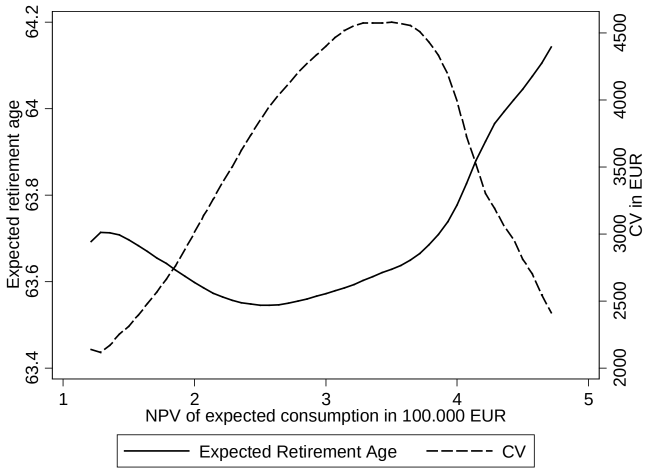
[77] Indeed, in 2002 Germany introduced subsidies for private pension plans to compensate employees for lower levels of expected PAYG-pensions due to various reforms. For a distributional analysis and further details see Corneo et al. (2015).

Figure 6: Distribution of compensating variations



Note: Euro in 2010 real values. Source: SUFVSKT2002, 2004-12

Figure 7: Predicted CVs and expected retirement age by NPVs of consumption



Note: Euro in 2010 real values. Source: SUFVSKT2002, 2004-12

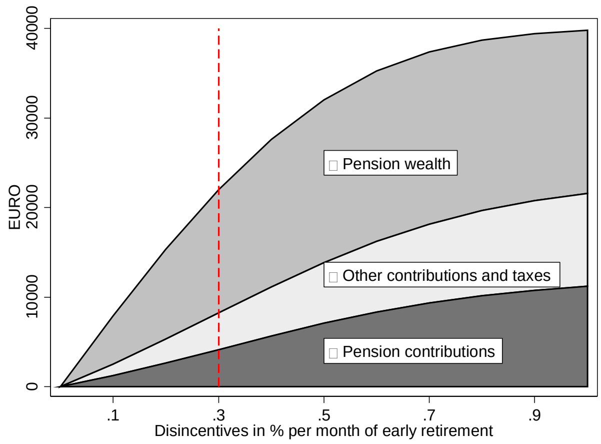
D. Fiscal implications

Figure 8 displays the (fiscal) benefits of introducing disincentives − the net public returns at varying levels. We again find the relationship to be concave. Although the returns are diminishing, increasing disincentives beyond the implemented level (vertical dashed line) would further foster the pension system's financial sustainability. At each disincentive level, about half of the net public returns are generated by reduced pension wealth, while the remainder is divided into increases in pension contributions and increases in tax payments and other contributions.

The net public returns can be linked to pension expenditures under a no disincentive scenario. This reveals that net public returns correspond to about 9% of average pension wealth under the actually implemented 0.3%-scenario (€ 21,994; vertical dashed line) and to 16% at the 1%-disincentive level. These fiscal implications are substantial. Resorting to aggregate data of the German pension insurance, we find that our sample population corresponds to 424,286 individuals for the 1939 to 1945 cohorts affected by the 1992 reform (Deutsche Rentenversicherung, 2014). We assess that the simulated public returns per capita at the 0.3% disincentive level translate into overall public gains of $424{,}286 \times € 21{,}994 \approx € 9.33$ billion for these cohorts.

Relating costs (Figure 5) and benefits (Figure 8) demonstrates that this increase in financial stability comes at the cost of increasing inequality and non-negligible individual welfare losses within the population of retirees. Still, at each disincentive level, net public returns are about five times as high average individual welfare losses.

Figure 8: Net public returns by disincetive level

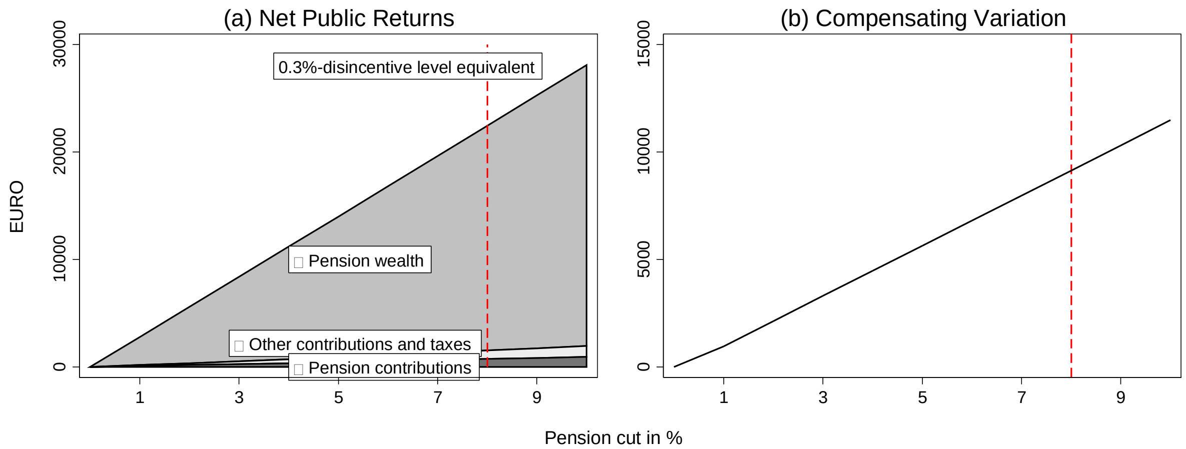


Note: Euro in 2010 real values. Source: SUFVSKT2002, 2004-12

E. Alternative reforms

To set the welfare losses into perspective, we simulate scenarios where we indiscriminately cut all pensions by a certain amount, ranging from 1% to 10% (Figure 9). It turns out that to yield equal net public returns to introducing a 0.3%-disincentive level, all pensions would have to be lowered by about 8%. However, pension cuts more than double the individual welfare losses. This holds also true for higher levels of net public returns. Since individuals barely adjust their retirement behavior when confronted with a pension cut, nearly all the net public returns stem from decreased pensions: €113 per month instead of €32 under the 0.3% disincentive level. These findings suggest that disincentives realize financial gains at lower individual costs than pension cuts.

Figure 9: Pension cuts – costs and benefits



Note: Euro in 2010 real values. The red line indicates the level of pension cuts that correspond to the net public returns under a 0.3%-disincentive level. Source: SUFVSKT2002, 2004-12, own calculations

5 Conclusion

This study evaluates the effectiveness of early retirement disincentives and its distributional, individual welfare and fiscal implications. We focus on disincentives leading to permanent pension deductions that increase with the distance between actual/early and normal retirement age. We model different disincentive levels and analyze to what extent disincentives are able to steer retirement behavior. Our range of disincentive levels includes the level actually implemented by the 1992 pension reform in Germany, which introduced permanent pension deductions of 0.3% per month of early retirement. For the actually implemented level, we estimate an increase in retirement age of 5.5 months. This implies that the reform is responsible for 68% of the observed change in retirement patterns across cohorts. Further simulations demonstrate that tripling the actually implemented level would essentially prevent individuals from opting for early retirement at all. Dose response reveals that disincentives increase inequality in expected consumption, cause individual welfare losses, and lead to positive net public returns. All three show a concave relationship with the disincentive level. The individual welfare losses are largest for medium income earners and difficult to compensate due to their heterogeneous distribution. However, at each disincentive level, the net public returns are more than five times as high as the individual welfare losses. Overall welfare in the economy may increase regardless, given that longer life expectancies and demographic change requires a reform of either the contribution scheme or the level of pension benefits.

Contrary to many public claims, disincentives do not correspond to an indiscriminating pension cut. In fact, at equal levels of net public returns, disincentives cause individual welfare losses that are less than half as large as those under a pension cut. Concerning future implications, Germany introduced various major pension reforms, two of which can be directly related to our results. The first reform increases the normal retirement age to 67 while the early retirement age remains at 63. This increases the disincentives for early retirees. Here, our results suggest that average retirement age increases and average pensions adjust slightly for individuals still employed at 62. The second reform introduces an exception to the rule by abolishing disincentives for pensioners with very long employment histories. According to our results, this will cause a substantial decline in average retirement age for eligible individuals. A more detailed analysis is left for future research.

Appendix

Appendix A: Taxation, social security contributions and sample selection

The income from PAYG-pensions and employment constructed from the available information provided in the VSKT is gross. To obtain net incomes, we subtract social security contributions and personal income taxes from gross earnings and pensions. Because the burden of taxes and social security contributions heavily depends on whether being an employee or retiree, a concise overview of the procedure and underlying assumptions to obtain net incomes is provided in subsections A.1 and A.2.

A.1 Social security contributions

The calculation of social security contributions is straightforward. Regular employees considered in our sample must contribute to the pension, unemployment, health and long term care insurance. Pensioners only have to contribute to the health and long term care insurance. Note that rates for pensioners and regular employees differ. Assessment basis is insurable income up to the respective contribution ceiling. Tables A1 and A2 list the key determinants used for calculating statutory social security contributions for the 1998 to 2011 assessment years. Displayed contribution rates are annual averages. In case of the statutory health insurance, actual contribution rates differ between insurance providers. Our calculation assumes the average contribution rates published by Deutsche Rentenversicherung (2014)**.** Further, employees with earnings above the compulsory insurance exemption limit may opt for a private health insurance instead of the statutory. We disregard this possibility.

Between 1998 and 2011, employees face a joint burden on gross earnings from contributions of roughly 23%, not including the employer's share. Social security contributions are usually almost evenly split between employee and employer. Gross earnings are net of employer's contribution and therefore only the employee's contributions need to be deducted. The burden differs with total remuneration. Low income earners and those receiving incomes above the respective contribution ceilings of the various branches of the social security system are subject to a lower relative burden. Social security contributions are calculated on hypothetical gross annual earnings and then deducted from gross monthly earnings. In contrast to employees, pensioners are subject to a combined average burden of 8 - 10%, which is deducted from the monthly pension.

Table A1: Pension and unemployment insurance

Year	Average social security income	Contribution ceiling	Contribution rate	
			Pension insurance	Unemployment insurance
1998	DM 52925	DM 100800	10.15	3.25
1999	DM 53507	DM 102000	9.85	3.25
2000	DM 54256	DM 103200	9.65	3.25
2001	DM 55216	DM 104400	9.55	3.25
2002	€ 28626	€ 54000	9.55	3.25
2003	€ 28938	€ 61200	9.75	3.25
2004	€ 29060	€ 61800	9.75	3.25
2005	€ 29202	€ 62400	9.75	3.25
2006	€ 29494	€ 63000	9.75	3.25
2007	€ 29951	€ 63000	9.95	2.1
2008	€ 30625	€ 63600	9.95	1.65
2009	€ 30506	€ 64800	9.95	1.4
2010	€ 31144	€ 66000	9.95	1.4
2011	€ 32100	€ 66000	9.95	1.5

Note: Values until 2001 in DM and in Euro thereafter. One Euro corresponds to 1.95583 DM. Contribution rates are annual averages for employees, contributions for employers differ slightly. Pensioners are not subject to pension or unemployment insurance contributions. Source: Deutsche Rentenversicherung (2014) (own calculations).

Table A2: Health and long-term care insurance

Year	Contribution ceiling	Contribution rate Health insurance – employees	Long-term care insurance – employees	Health and long term care insurance – pensioners
1998	DM 75600	6.8	7.575	0.85
1999	DM 76500	6.8	7.6253	0.85
2000	DM 77400	6.8	7.6	0.85
2001	DM 78300	6.8	7.6	0.85
2002	€ 40500	7	7.725	0.85
2003	€ 41400	7.2	7.925	0.85
2004	€ 41856	7.2	8.27505	0.85
2005	€ 42300	8	9.05	1.1
2006	€ 42756	7.4	9.25	1.1
2007	€ 42756	7.7	9.4	1.1
2008	€ 43200	7.8	9.7	1.1
2009	€ 44100	7.9	10	1.225
2010	€ 45000	7.9	9.85	1.225
2011	€ 44550	8.2	10.15	1.225

Note: Values until 2001 in DM and in Euro thereafter. One Euro corresponds to 1.95583 DM. Contribution rates are annual averages for employees/pensioners, contribution rates for employers/pensions insurance differ slightly. Source: Deutsche Rentenversicherung (2014) (own calculations).

A.2 Personal income tax

In Germany, personal income tax depends on several characteristics of the tax unit not available in our data. For our calculation we assume that all taxable income solely stems either from employment and/or PAYG pensions. Other sources of income are not recorded in our data. In Table A3 and A4 we provide an overview of the actual composition of household and individual incomes for the considered population according to the SOEP. The population depicted in Tables A3 and A4 mirrors our sample regarding age, region, employment status, earnings biography and gender. For our sample, household and individual incomes are predominantly comprised of earnings from employment and PAYG-pensions, which can be observed in our data. Other pensions, transfers or asset income are negligible small.

Table A3: Composition of individual income and marital status

Age	Employment				Pensions				Unempl. benefit		Married	
	Employed		Self		PAYG		Other					
	Share	(Sd)	Share	(Sd)	Share	(Sd)	Share	(Sd)	Share	(Sd)	Share	(Sd)
62	96	(16)	0	(2)	3	(14)	0	(0)	1	(7)	78	(48)
63	92	(24)	1	(10)	7	(21)	0	(0)	1	(7)	76	(48)
64	83	(32)	1	(9)	16	(31)	0	(3)	0	(4)	70	(47)
65	58	(46)	0	(3)	41	(46)	1	(7)	0	(2)	73	(48)
66	36	(41)	0	(6)	64	(41)	0	(1)	0	(1)	72	(47)
67	10	(22)	1	(8)	89	(23)	0	(0)	0	(0)	72	(44)
68	8	(21)	1	(7)	92	(22)	0	(0)	0	(0)	76	(44)

Note: Income shares of total individual income in percent. Standard deviation (Sd) in parentheses. Sample comprised of West German males born between 1935 and 1945 in regular insurable employment at age 62. Source: SOEP waves 1984-2012.

Table A4: Composition of household income

Age	Labor income		PAYG Pensions		Asset income	
	Share	(Sd)	Share	(Sd)	Share	(Sd)
62	87.81	(18.69)	6.10	(15.83)	4.18	(8.96)
63	84.65	(22.48)	8.77	(19.01)	3.53	(6.82)
64	75.71	(28.65)	17.23	(25.31)	3.90	(8.40)
65	56.69	(38.86)	32.22	(34.73)	5.12	(8.72)
66	39.22	(35.40)	47.93	(32.81)	4.58	(7.59)
67	20.24	(25.77)	63.67	(28.14)	6.37	(10.38)
68	17.54	(26.69)	67.73	(28.65)	5.57	(9.05)

Note: Income shares of total household income in percent. Standard deviation (Sd) in parentheses. Sample comprised of households with a West German male born between 1935 and 1945 in regular insurable employment at age 62. Source: SOEP waves 1984-2012.

Table A3 shows that roughly three-quarters of the individuals are married. Because the martial status is not recorded in the data, we assume all tax units to be married and eligible for joint assessment.[78] For robustness, we also calculate a scenario where the tax units are assumed to be single. Due to the ages considered, we do not regard the case of tax relevant children.

In general, after deductions of e. g. social security, the income tax schedule is applied. The income tax is calculated on yearly taxable income (earnings and pensions). To obtain the monthly income tax, the yearly tax burden is distributed according to the monthly share of taxable income on yearly

[78] Married couples profit from a splitting rule (Bönke and Eichfelder, 2010). We assume joint assessment and a single earner/pensioner without spousal income.

taxable income. From 1998 to 2011, the code was subject to several changes, e.g: top marginal tax rates were reduced from 53% to 45%; taxation of pensions was reformed by the introduction of deferred taxation and changes in the deductibility of social security contributions. In addition, there were some minor alterations like changes in lump sum deductions. All these changes occur regularly between 1998 and 2011, impacting the birth cohorts accordingly and influencing their retirement decisions. To disentangle the impact from changes in the income tax law from changes in the pension system, we simulate a counterfactual assuming the governing law of 1998 (see Appendix C). Concerning the taxation of income from employment and PAYG-pensions, our tax model in particular includes the following regulations:[79]

- Income from employment: In order to obtain the taxable portion of income, gross earnings are reduced by a lump sum deduction for work related expenses (*Werbungskostenpauschale*).
- Income form PAYG-pension: In case of pensions, the return portion (*Ertragsanteil*) is taxable only if the pensioner retired before 2005. For our sample, the return portion varies between 27% and 29%, depending on retirement age. Beginning with 2005, the taxable portion (*Besteuerungsanteil*) depends on the year of retirement and ranges from 50% in 2005 to 62% in 2011. Further, the lump sum deduction for pensions is subtracted.
- Special expenses (*Sonderausgaben*): The modelling concerning the deduction of social security contribution from taxable income (*Vorsorgeaufwendungen*) accounts for all changes between 1998 and 2011. Further, the lump sum deduction for special expenses (*Sonderausgabenpauschbetrag*) is subtracted.

A.3 Data

The dataset consists of the waves of SUFVSKT of calendar years 2002 and 2004-2012. Each SUF is a 25% stratified random sample of the VSTK of the respective year and includes the same information. Since we need completed biographies to clearly identify the timing of old-age retirement, we focus on cohorts aged 66 or 67 in the respective year only. This means that usable observations for cohorts 1938-1945 appear in two different waves, once aged 66 and once aged 67. Due to the sampling structure it is possible to match those two waves for each of these cohorts and enhance the number of observations. Since there is no unique identifier across all waves, we identify duplicates (whom appear in both waves) on the basis of their employment biographies. For the selected cohorts, those

[79] For a detailed description of work related deduction and special expenses see Bönke und Eichfelder (2010).

biographies consist of monthly earnings points observations included from age 14 onwards up to the age of 66. Therefore, we draw on a large number of data points for the matching procedure and do not have to make any assumptions. For identification we use all of the at least 420 month (35 years) history as well as the year and month of birth. Verification checks further confirm the correctness of our procedure. Certainly, the matching procedure might be problematic for individuals without a strong labor market attachment - but those are not the persons we focus on.

Appendix B: Robustness

B.1 Inclusion of type-specific preference heterogeneity

We implement preference heterogeneity in disutility of work by assuming two unobserved types[80] $m \in \{1,2\}$ that comprise a fixed proportion of the population (Heckman and Singer, 1984). We assume that the constant in the spline function of the disutility of work δ_1 is heterogeneous for the unobserved types. Hence, equation (4) becomes:

$$l(d_{nt}, t_n) = \delta_{1m}[1 - d_{nt}] + [1 - d_{nt}] \begin{Bmatrix} & \delta_2\, min(t, 3) & if\ t > 0 \\ + & \delta_3\, min(t - \ 3,3) & if\ t > 3 \\ + & \delta_4\, min(t - \ 6,3) & if\ t > 6 \\ \vdots & \vdots & \vdots \\ + & \delta_9\, min(t - \ 21,3) & if\ t > 21 \end{Bmatrix}$$

The results suggest that the second type with lower disutility comprises only about 9% of the population. The specification yields unprecise estimates for the parameter related to this type (δ_{12}). Still, central parameters remain stable, the model fit improves only slightly, and postestimation outcomes are almost unaffected. Therefore, we conclude that our results are insensitive to the inclusion of preference heterogeneity and do not add it to the baseline specification.

B.2 Type probabilities (unconditional and conditional on initial conditions)

The probability that individual n is of type m is given by π_{nm}, where π_{nm} is assumed to be logistic and can be modeled conditional on initial conditions at age 63. For the unconditional specification we assume:

$$\pi_{nm} = \frac{\exp(\gamma_1)}{1 + \exp(\gamma_1)}$$

For the conditional specification we assume:

[80]For more than two unobserved types, the optimization algorithm did not converge.

$$\pi_{nm} = \frac{\exp(\gamma_1 + \gamma_2 Lifetime\ earnings_n^{63}/10)}{1 + \exp(\gamma_1 + \gamma_2 Lifetime\ earnings_n^{63}/10)}$$

In the conditional specification, the probability that individual n is of type m is modeled as a function of the employment history at age 63. Thereby, we use real accumulated lifetime earnings (Bönke et. al., 2015) as a summary measure because it reflects both wage history and employment pattern over the working life cycle.

By making the type probabilities a function of the employment and wage history at age 63, we account for non-random initial conditions at age 63. This approach follows Wooldridge (2005) and only requires the assumption that the initial condition is random conditional on real accumulated lifetime earnings at age 63. The log-likelihood function of the sample is then given by

$$\sum_{n=1}^{N} \log\left\{\sum_{m=1}^{2} \pi_{nm} \prod_{t=1}^{T}\left[\sum_{d_{nt}} Prob_m(d_{nt}|\boldsymbol{s}_{nt}, \boldsymbol{\lambda}) \times I(d_{nt})\right]\right\}$$

with $I(d_{nt})$ indicating the individual choice observed in period t.

B.3 Results

Table B1: Parameter estimates and robustness

	Baseline	Single	Unobserved Heterogeneity w/o init. cond.	Unobserved Heterogeneity with init. cond.
α (consumption)	0.21516	0.2024	0.3203	0.32503
	(0.024108)	(0.026465)	(0.035013)	(0.040995)
ρ (CRRA)	1.5065	1.5	1.5661	1.5784
	(0.071666)	(0.087501)	(0.065832)	(0.079646)
δ_1, δ_{11} (disutility, constant)	-3.0154	-2.9948	-3.0742	-3.0743
	(0.24757)	(0.23875)	(0.25671)	(0.23648)
δ_{12} (disutility, constant)			-1.5747	-1.5228
			(2.3507)	(2.384)
δ_2 (disutility, spline)	1.0568	1.0577	1.0561	1.0561
	(0.11822)	(0.11174)	(0.12249)	(0.11377)
δ_3 (disutility, spline)	-0.16413	-0.16571	-0.16444	-0.16443
	(0.066455)	(0.062373)	(0.071532)	(0.070221)
δ_4 (disutility, spline)	0.032382	0.038399	0.032745	0.032774
	(0.064722)	(0.064915)	(0.072272)	(0.070505)
δ_5 (disutility, spline)	0.080143	0.077559	0.078887	0.078822
	(0.073298)	(0.072578)	(0.078636)	(0.07622)
δ_6 (disutility, spline)	-0.26074	-0.26343	-0.25952	-0.2594
	(0.091609)	(0.093278)	(0.099041)	(0.091424)
δ_7 (disutility, spline)	0.25537	0.2551	0.24973	0.2493
	(0.12815)	(0.13244)	(0.14189)	(0.11686)
δ_8 (disutility, spline)	-0.23207	-0.22357	-0.21365	-0.21193
	(0.14114)	(0.14541)	(0.15405)	(0.13698)
δ_9 (disutility, spline)	1.2119	1.2055	1.1328	1.1277
	(0.13425)	(0.13741)	(0.15709)	(0.15359)
γ_1 (constant, type)			2.2783	2.4235
			(0.36287)	(0.71499)
γ_2 (initial condition, type)				-0.0088864
				(0.031546)
Log-likelihood	-1942	-1947	-1930	-1930

Source: SUFVSKT2002, 2004-12, own calculations.

Table B2: Exemplary robustness results for a disincentive level of 0.3%

	Baseline	Single	Unobserved Heterogeneity w/o init. cond.	Unobserved Heterogeneity with init. cond.
ΔE[retirement age] (months)	5.54	5.07	5.80	5.77
ΔE[NPV of consumption]	2,271	€ -783	€ 3,369	€ 3,249
ΔE[NPV of consumption] (%)	0.65	-0.40	1.07	1.04
ΔGini coefficient (%)	4.68	3.6	5.09	4.81
ΔMonthly retirement income	€ -31.92	€ -35.13	€ -30.60	€ -30.74
Average compensating variation	€ 3,913	€ 3,561	€ 4,615	€ 4,604
Average equivalent variation	€ 3,754	€ 3,453	€ 4,296	€ 4,297
NPV of net public returns	€ 21,994	€ 22,825	€ 22,784	€ 22,650
ΔE[NPV of pension benefits]	€ 13,754	€ 13,450	€ 13,851	€ 13812
ΔE[NPV of pension contributions]	€ 4,137	€ 3,762	€ 4,381	€ 4345
ΔE[NPV of other contr. & taxes]	€ 4,103	€ 5,613	€ 4,552	€ 4493

Note: Euro in 2010 real values. Source: SUFVSKT2002, 2004-12, own calculations

Appendix C: Effects of different institutional changes

Here we document the effects of various alternative reforms apart from the introduction of disincentives. Tables C1 shows the effects of certain tax/pension parameters on retirement age. These effects are measured by assuming the values from 1998 (the first year of this study) for all other years. Figure C1 presents an alternative reform-scenario, which implements a pension cut of a various levels. Panel (a) shows how much the pensions would have to decrease to yield a certain amount of net public returns. Panel (b) provides estimates on the corresponding individual welfare losses. Table C2 shows further results of this counterfactual scenario.

Table C1: Predicted effects of various reforms

	Predicted effect on retirement age in month			
Cohort	Disincentives	Pension value constant	Tax system constant	Total predicted change
1935	0	-0.0037798	0.3086298	0.30485
1936	0	0.010305	0.306345	0.31665
1937	1.1605	0.0136	0.3205	1.4946
1938	4.1914	-0.0299	0.3171	4.4786
1939	5.2811	0.0228	0.2984	5.6023
1940	5.4339	0.1542	0.2788	5.8669
1941	5.4546	0.3112	0.2548	6.0206
1942	5.597	0.4855	0.3239	6.4064
1943	5.5353	0.6321	0.3298	6.4972
1944	5.6771	0.7385	0.3522	6.7678
1945	5.6608	0.7227	0.3398	6.7233

Note: Euro in 2010 real values. Pension value and tax system constant refer to scenarios where both are constant at the 1998 level, the first year of this study. Source: SUFVSKT2002, 2004-12, own calculations.

Table C2: Effects of pension cuts

Pension cut	Net public returns	CV	Δ Retirement age (month)	Δ Gini	Δ NPV of expected lifetime consumption	Δ Monthly pension benefits
1%	2792	959	0.11033	0.21055	-2279	-14
2%	5589	2132	0.224	0.42715	-4551	-28
3%	8389	3303	0.34041	0.64724	-6815	-42
4%	11193	4476	0.45975	0.87447	-9070	-56
5%	14001	5642	0.58111	1.1089	-11318	-71
6%	16813	6813	0.70536	1.3496	13557	-85
7%	19629	7979	0.83244	1.5997	-15788	-99
8%	22451	9147	0.96449	1.848	-18007	-113
9%	25278	10313	1.1001	2.1045	-20218	-127
10%	28111	11483	1.239	2.3701	-22419	-141

Note: Euro in 2010 real values. Source: SUFVSKT2002, 2004-12, own calculations.

References

Atalay, K., and G. F. Barrett (2015): "The Impact of Age Pension Eligibility Age on Retirement and Program Dependence: Evidence from an Australian Experiment," The Review of Economics and Statistics, 97(1), 71–87.

Blundell, R., C. Meghir and S. Smith, (2002): "Pension Incentives and the Pattern of Early Retirement," Economic Journal 112(478), 153-170.

Bönke, T., G. Corneo, and H. Lüthen (2015): "Lifetime earnings inequality in Germany," Journal of Labour Economics 33(1), 171-208.

Bönke, T., and S. Eichfelder (2010): "Horizontal Equity in the German Tax- Benefit System: A Simulation Approach for Employees," FinanzArchiv: Public Finance Analysis, 3(37), 295–331.

Bönke, T., C. Schröder, and K. Schulte (2010): "Incomes and Inequality in the Long Run: the Case of German Elderly," German Economic Review, 11, 487–510.

Börsch-Supan, A. (2000): "A Model under Siege: A Case Study of the Germany Retirement Insurance System," The Economic Journal, 110(461), 24-45.

Börsch-Supan, and A. Reil-Held (2001): "How Much is Transfer and How Much is Insurance in a Pay-as-you-go System? The German Case," The Scandinavian Journal of Economics, 3(103), 505–524.

Börsch-Supan, A. (2002): "Incentive Effects of Social Security on Labor Force Participation: Evidence in Germany and Across Europe," Journal of Public Economics, 78, 25–49.

Coile, C. and J. Gruber (2007): "Future social security entitlements and the retirement decision," The Review of Economics and Statistics 89(2), 234-246.

Corneo, G., C. Schröder and J. König (2015): "Distributional Effects of Subsidizing Retirement Savings Accounts: Evidence from Germany," Economics Discussion Paper, School of Business & Economics, FU Berlin, Nr. 2015/18.

Diamond, P. and Mirrless, J. (1978): "A Model of Social Insurance with variable Retirement", Journal of Public Economics 10, 295-336.

Deutsche Rentenversicherung (2014): Rentenversicherung in Zeitreihen, vol. 22 of DRV-Schriften.

French, E., and J. B. Jones (2011): "The Effects of Health Insurance and Self- Insurance on Retirement Behavior," Econometrica, 79(3), 693–732.

Fuchs-Schündeln, N., and M. Schündeln (2009): "Who stays, who goes, who returns?" Economics of Transition, 17, 703–738.

Gourinchas, P.-O., and J. A. Parker (2002): "Consumption Over the Life Cycle," Econometrica, 70(1), 47–89.

Gruber, j. and D. Wise (2007): "Social Security Programs and Retirement around the World," NBER

Publications, University of Chicago Press, Chicago.

Gustman, A. and Steinmeier, T. (2015): "Effects of social security policies on benefit claiming, retirement and saving," Journal of Public Economics 129(C), 51-62.

Haan, P., and V. Prowse (2014): "Longevity, life-cycle behavior and pension reform," Journal of Econometrics, 178(P3), 582–601.

Hanel, B. (2010): "Financial incentives to postpone retirement and further effects on employment - Evidence from a natural experiment," Labour Economics, 17, 474–486.

Heckman, J., and B. Singer (1984): "A Method for Minimizing the Impact of Distributional Assumptions in Econometric Models for Duration Data," Econometrica, 52(2), 271–320.

Laun, T., and J. Wallenius (2015): "A Life Cycle Model of Health and Retirement: The Case of the Swedish Pension Reform," Journal of Public Economics, forthcoming.

Lüthen, H. (2015): "Rates of Return and Early Retirement Disincentives: Evidence from a German Pension Reform," German Economic Review, forthcoming.

Manoli, D., Mullen, K. and Wagner, M. (2014): "Policy variation, labor supply elasticities, and a structural model of retirement," IZA discussion paper 8659.

Mastrobuoni, G. (2009): "Labor supply effects of the recent social security benefit cuts: Empirical estimates using cohort discontinuities," Journal of Public Economics, 93(11-12), 1224–1233.

R., B., C. Meghir, and S. Smith (2002): "Pension Incentives and the Pattern of Early Retirement," The Economic Journal, 112, 153–170.

Rust, J. (1987): "Optimal Replacement of GMC Bus Engines: An Empirical Model of Harold Zurcher," Econometrica, 55(5), 999–1033.

Rust, J., and C. Phelan (1997): "How Social Security and Medicare Affect Retirement Behavior in a World of Incomplete Markets," Econometrica, 65(4), 781–832.

Schmähl, W. (2011), 'Von der Ergänzung der gesetzlichen Rentenversicherung zu deren partiellem Ersatz: Ziele, Entscheidungen sowie sozial- und verteilungspolitische Wirkungen – Zur Entwicklung von der Mitte der 1990er Jahre bis 2009', in: Eichendorfer, E., H. Rische and W. Schmähl (eds.), Handbuch der Gesetzlichen Rentenversicherung SGB IV, Luchterhand, Köln, 169-249.

Statistisches Bundesamt (2006): Generationen-Sterbetafeln Für Deutschland, Modellrechnungen für die Geburtsjahrgänge 1871-2004. Wiesbaden.

Staubli, S., and J. Zweimüller (2013): "Does raising the early retirement age increase employment of older workers?" Journal of Public Economics, 108, 17–32.

Wooldridge, J. (2005): "Simple solutions to the initial conditions problem in dynamic, nonlinear panel-data models with unobserved heterogeneity," Journal of Applied Econometrics 20(1), 39-54.

Chapter 4:

Rates of Return and Early Retirement Disincentives: Evidence from a German Pension Reform*

Abstract: To counteract the financial pressure emerging in aging societies, statutory pension schemes are undergoing fundamental reforms in many Western countries. Starting with cohort 1937, Germany introduced permanent pension deductions for early retirement. This study examines the profitability of pension contributions against the background of this reform for cohorts 1935-1945. Internal rates of return (IRR) are used to measure the profitability. For men, the IRR declines from 2.4% to 1.2% and for women from 5.2% to 3.7%. The results suggest that the majority of the trend, about 75%-80%, is caused by increased pension contributions and not by the reform.

JEL CODES: D02, D14, H55

Keywords: Pensions, reform, early retirement, disincentives, pay-as-you-go, rates of return, Germany

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Zusammenfassung

Diese kumulative Dissertation besteht aus vier eigenständigen Studien, die thematisch miteinander verknüpft sind. Insgesamt geht es um die Lebensverläufe westdeutscher Arbeitnehmer seit dem zweiten Weltkrieg. Die ersten beiden Aufsätze vergleichen die Erwerbsverläufe verschiedener Generationen im Zeitablauf, wobei sich unterschiedliche Phasen auf dem bzw. diverse Transformationsprozesse des Arbeitsmarkt(s) in den unterschiedlichen Einkommenspfaden widerspiegeln. Die beiden hinteren Kapitel widmen sich dem Übergang vom Erwerbsleben in die Rente. Da die alternde Gesellschaft zunehmend die Finanzierung der deutschen Rentenversicherung erschwert, sind bei dieser wichtigen Säule des deutschen Wohlfahrtsstaats verschiedene Reformen durchgeführt worden. Eine dieser Reformen sowie diverse Zeittrends werden in den letzten beiden Kapiteln genauer untersucht.

Das erste Kapitel, *Lifetime earnings inequality in Germany*, ist eine gemeinsame Arbeit mit Timm Bönke und Giacomo Corneo, wobei jeder Autor einen Beitrag von einem Drittel geleistet hat. Die Arbeit ist im *Journal of Labor Economics* erschienen (siehe Seite II dieser Dissertation). In der Studie untersuchen wir auf Grundlage der Versicherungskontenstichprobe (VSKT) die Ungleichheit und Mobilität von Erwerbseinkommen, welche über den gesamten Lebensverlauf erzielt wurden. Im Zentrum der Arbeit steht ein intragenerationaler Vergleich zwischen 1935 und 1969 geborenen westdeutschen männlichen Arbeitnehmern. Unsere Ergebnisse zeigen, dass sowohl kurz- als auch langfristige Einkommensmobilität unserer Untersuchungspopulation konstant geblieben ist. Dahingegen finden wir über die Kohorten einen starken Anstieg der Ungleichheit der Lebenseinkommen um 85%. Hierbei ist eine erhöhte Arbeitslosigkeit von Arbeitnehmern im unteren Viertel der Lohnverteilung zu 20-40% für diesen Anstieg verantwortlich. Eine erhöhte Lohnspreizung verursacht die verbleibenden 60-80% des Anstiegs.

Der zweite Teil, *The Dynamics of Earnings in Germany: Evidence from Social Security Records*, ist eine Gemeinschaftsarbeit mit Timm Bönke und Matthias Giesecke, die jeweils ein Drittel zu diesem Projekt beisteuerten. Wiederum auf Grundlage der VSKT untersuchen wir laufende Trends idiosynkratischer Einkommensvolatilität, indem wir die Autokovarianzen von Einkommensresiduen in eine permanente (langfristige) und eine transitorische (kurzfristige) Komponente zerlegen. Hierbei verwenden wir komplette Erwerbslebensverläufe westdeutscher Männer der Kohorten 1935 bis 1974 für die Jahre 1960 bis 2009. Auf dem deutschen Arbeitsmarkt fand in diesem Zeitraum ein deutlicher Transformationsprozess statt, gekennzeichnet durch starke Deregulierung, einer Schwächung der Gewerkschaften und einer Verlagerung von Arbeitsplätzen vom Industrie- in den

Servicesektor. Wir finden deutliche Anstiege beider Komponenten, was gleichzeitig die verschiedenen Phasen dieses Transformationsprozesses widerspiegelt. Die größten Zuwächse der transitorischen Komponente zeigen sich in den frühen 1970er und den 1990er Jahren bei jungen Arbeitnehmern. Bei der permanenten Komponente hingegen finden wir die deutlichsten Steigerungen bei älteren Arbeitnehmern in den frühen 1980er und den 2000er Jahren. Insgesamt ergeben sich damit ein erschwerter Arbeitsmarkteintritt für jüngere Arbeitnehmer und ein starker Anstieg von Lohndifferenzen für etablierte Arbeitnehmer.

Das dritte Kapitel, *Rates of Return and Early Retirement Disincentives: Evidence from a German Pension Reform*, ist ein alleiniges Projekt und im German Economic Review veröffentlicht (siehe Seite II dieser Dissertation). Der Beitrag untersucht die reale Verzinsung von Rentenversicherungsbeiträgen für die Kohorten 1935-1945 in Deutschland. Als Datenbasis dienen die Scientific-Use-Files (SUFs) der VSKT, wobei individuelle Erwerbsbiografien westdeutscher Frauen und Männer die Untersuchungsgrundlage bilden. Die Analyse findet vor dem Hintergrund der Rentenreform 1992 statt, welche Abschläge auf den vorzeitigen Renteneintritt für die Jahrgänge ab 1937 einführte. Die Verzinsung wird mit dem internen Zinsfuß gemessen. Dies ermöglicht sowohl Vergleiche zwischen Untergruppen innerhalb einer Kohorte als auch Vergleiche zwischen Kohorten. Die Ergebnisse zeigen, dass der Zinssatz von Beiträgen für Altersrenten über die betrachteten Kohorten bei Männern von 2,4% auf 1,2% und bei Frauen von 5,2% auf 3,7% absinkt. Für die Empfänger von Erwerbsminderungsrenten ergibt sich kein eindeutiger Trend, wobei der Zinssatz hier bei ca. 5% für Frauen und ca. 3% für Männer liegt. Mit Hilfe einer kontrafaktischen Analyse wird des Weiteren deutlich, dass der Großteil der Abnahme des Zinsfußes durch gestiegene Beiträge zur Rentenversicherung und nicht durch die Einführung von Abschlägen verursacht wird.

Der vierte Artikel, *Effectiveness of early retirement disincentives: individual welfare, distributional and fiscal implications*, untersucht den Übergang vom Erwerbsleben in die Rente und ist ein gemeinsames Werk mit Timm Bönke und Daniel Kemptner, wobei der Anteil eines jeden Autors bei einem Drittel liegt. Dadurch, dass in alternden Gesellschaften zunehmend finanzieller Druck auf Alterssicherungssysteme ausgeübt wird, sind Erkenntnisse über das Renteneintrittsverhalten äußerst wertvoll für politische Entscheidungsträger. Der Fokus dieser Studie liegt auf der Beantwortung der Frage, welche Effekte Abschläge auf vorzeitigen Renteneintritt auf das Renteneintrittsverhalten haben und wie sich Abschläge auf die individuelle Wohlfahrt, das Staatsbudget oder die Ungleichheit auswirken. Hierbei fokussieren wir uns auf Westdeutsche Männer mit langen Erwerbsbiografien. Wiederrum auf Basis der Scientific-Use-Files (SUFs) der VSKT und einem detaillierten Model des deutschen Steuer- und Transfersystems schätzen wir ein dynamisches strukturelles Rentenzugangsmodel. Unsere Ergebnisse zeigen, dass Abschläge ein effektives Mittel sind, um

Individuen dazu zu bewegen, länger im Arbeitsmarkt zu verbleiben. Auf der anderen Seite erfahren die Betroffenen individuelle Wohlfahrtsverluste und die Ungleichheit im Restlebenskonsum steigt. Die Zuwächse der öffentlichen Einnahmen sind allerdings deutlich höher als die individuellen Wohlfahrtsverluste. Im Vergleich mit einer kontrafaktischen Reform, die pauschale Rentenkürzungen einführt und dasselbe öffentliche Aufkommen erzielt, zeigt sich zudem, dass pauschale Rentenkürzungen doppelt so große individuelle Wohlfahrtsverluste hervorrufen.

Erklärung

Erklärung gem. § 9 Abs. 4 der Promotionsordnung zum Dr. rer. pol. des Fachbereichs Wirtschaftswissenschaft der Freien Universität Berlin vom 14. August 2008.

Hiermit erkläre ich, dass ich für die Dissertation folgende Hilfsmittel und Hilfen verwendet habe:

Stata 10, 11, 12 und 13 in den Versionen SE und MP

Zitierte Literatur

Auf dieser Grundlage habe ich die Arbeit selbständig verfasst.

Holger Lüthen