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Wage Dispersion, Trade Unions, and Heterogeneous Labor Demand – Microeconometric Analyses for Germany

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To My Parents

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Abstract

The dissertation consists of four self-contained essays. The first paper provides a detailed description of the evolution of wage inequality in East and West Germany in the late years of the twentieth century. In contrast to previous decades, wage inequality has been rising in several dimensions during that period. The second paper identifies cohort effects in the evolution of both wages and employment. Observed structures are consistent with a labor demand framework that incorporates steady skill-biased technical change. Substitutability between skill and age groups in the German labor market is found to be relatively high. Simulations based on estimated elasticities of substitution illustrate that higher wage dispersion between skill groups would have contributed to a reduction in unemployment. The third paper estimates determinants of individual union membership decisions and studies the erosion of union density in East and West Germany. Using corresponding predictions of net union density, the fourth paper analyzes the link between union strength and the structure of wages. A higher union density is associated with lower residual wage dispersion, reduced skill wage differentials, and a lower wage level. This finding is in line with an insurance motive for union action.

List of Papers

The thesis comprises the following articles:

(1) "Rising Wage Dispersion, After All! The German Wage Structure at the Turn of the Century," IZA Discussion Paper 2098, April 2006.

(2) "Skill Wage Premia, Employment, and Cohort Effects: Are Workers in Germany All of the Same Type?", IZA Discussion Paper 2185, June 2006, joint with Bernd Fitzenberger.

(3) "The Erosion of Union Membership in Germany: Determinants, Densities, Decompositions," IZA Discussion Paper 2193, July 2006, joint with Bernd Fitzenberger and Qingwei Wang.

(4) "Equal Pay for Equal Work? On Union Power and the Structure of Wages in West Germany, 1985–1997," translation of "Gleicher Lohn für gleiche Arbeit? Zum Zusammenhang zwischen Gewerkschaftsmitgliedschaft und Lohnstruktur in Westdeutschland 1985–1997," *Zeitschrift für ArbeitsmarktForschung*, 38 (2/3), 125-146, joint with Bernd Fitzenberger, 2005.

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1 Introduction

Over the past decades, the German economy has been plagued with high unemployment, especially among persons with low educational attainment. In the year 2004, unemployment rates in West Germany were 21.7% among individuals without a vocational degree, 7.3% among those with a completed apprenticeship, and 3.5% among university graduates, and the respective numbers for East Germany were as high as 51.2%, 19.4%, and 6.0% (Reinberg and Hummel, 2005). Different types of rigidities hampering flexible adjustments in the labor market have been hypothesized to have contributed to this situation.

Among the most prominent labor market features, the structure of wages is crucial for the evolution of employment and different degrees of unemployment incidence. With the growing availability of large micro data sets not only the wage level, but also the degree of wage dispersion or compression has received increasing attention. Numerous studies employ microeconometric approaches to analyze various facets of heterogeneity in the wage structure (Katz and Autor, 1999). Particular emphasis is put on returns to human capital and the evolution of skill wage premia.

The evolution of the West German wage structure between the mid-1970s and the mid-1990s has been extensively studied. By and large, the structure has been found to be relatively compressed in international comparison and rather stable over time. Returns to human capital components as well as residual wage inequality showed fairly little variation. However, as elaborated in the discussion about skill-biased technical change, relative demand for low-skilled labor decreases faster over time than does relative supply. In line with neoclassical demand theory, market clearing would in this case require an increase of qualificational wage differentials. Therefore, the "unbearable stability" (Prasad, 2004) of the German wage structure is considered a key determinant for the asymmetric increase in unemployment.

As a related key issue, the influence of labor market institutions on economic performance in general, and on wage setting in particular, is currently under debate (OECD, 2006). In times of increasingly heterogeneous economic conditions the catchword is eurosclerosis, stating that institutional inflexibilities restrain labor market performance and the dynamics of economic development. A major focus is on the impact of trade unions.

Trade union membership has declined remarkably in many developed countries over the last decades. In Germany, gross union density, i. e., the ratio of the number of union members and the number of employees in the labor market, was down to a historically low level of 27% by the year 2004 (Fitzenberger, Kohn, and Wang, 2006). A large international literature investigates into causes of this pervasive phenomenon as well as its economic consequences (Addison and Schnabel, 2003). Lacking a direct influence on firms' hiring and firing decisions, trade unions influence economic outcomes mainly through the channel of wage bargaining or, put differently, through their impact on the wage structure. For example, as wage moderation may lead to a higher level of employment, a higher union-bargained wage level may be detrimental to employment. A strong union taste for wage equality, resulting in wage compression, may again worsen the employment prospects of the low-skilled. In any case, the observed trends towards deunionization are likely to weaken unions' bargaining power and therefore their impact on the labor market.

This dissertation collects four self-contained articles which contribute to the active literature on wage structures, heterogeneous labor demand, and the impact of trade unions. The following paragraphs put the essays into perspective; they provide an overview of central results and point to scope for future research.

Chapter 2 (Kohn, 2006) presents a detailed picture of the evolution of wage inequality in East and West Germany in the late years of the twentieth century.

Wage dispersion in West Germany has been rather low in international comparison and remarkably stable between the mid-1970s and the mid-1990s. Yet studies of the East German wage structure report an even higher degree of wage compression in the late years of the GDR, reflecting the egalitarian doctrine of the socialist system. Newly available register data from the IAB employment sample (IABS) 1975–2001 allow to reinvestigate the empirical evidence for more recent years. My paper scrutinizes the evolution of wage levels and wage inequality within and between different labor market groups for the years 1992–2001. I find that wage inequality has in fact been rising in many dimensions throughout this period.

First, the inspection of year-specific wage distributions reveals that wage dispersion has been rising both in East and West Germany. The increase was more pronounced in East Germany such that, starting out from a lower degree in the year 1992, wage dispersion in the East largely caught up with the West German level until 2001. The increase in wage inequality among women was more pronounced compared to the increase among men. Moreover, the larger part of the increase in dispersion among women happened in the lower parts of the distribution, but dispersion among men increased disproportionately in the upper parts. Convergence in wage levels between East and West Germany took place only until the year 1996, but substantial differences persist since then.

Second, the estimation of censored quantile wage regressions provides insights into the determinants of the observed wage distributions. The bottom line of the regression results meets a-priori expectations: Age-earnings wage profiles not only are the steeper the higher the skill level, but they are also relatively flat in East Germany in 1992. The unification shock clearly led to a depreciation of human capital in the East. However, this effect wears out with the aging of post-unification labor market cohorts, and differences between East and West Germany have lessened by the year 2001. The quantile regression approach further reveals remarkable differences in the effects across the wage distribution.

Third, and finally, Machado-Mata-type analyses decompose differences in the wage distributions between East and West Germany as well as corresponding changes over time into "characteristics effects" capturing differences in the composition of the workforce, and "coefficients effects" capturing differences in the returns associated with observed characteristics. Building on the flexible quantile regressions, this approach depicts heterogeneous effects. In East-West comparison, differences in the composition of the male work force turn out largely negligible. However, characteristics of full-time working women are mostly in favor of higher wages in the East. Yet this effect ceases to apply at the lower end of the distribution by the year 2001. With respect to the evolution of wages over time, changes in the composition of the workforce capture major parts of the respective wage increases in the upper halves of the wage distributions for West Germany. This finding reflects a skill upgrading in the work force. Shifts in the industry structure and skill upgrading yet played only a minor role for explaining the wage increases in East Germany. For women in the lower parts of the Eastern distribution, the characteristics effect even worked towards real wage cuts, substantiating the particular increase in wage dispersion among this group.

The focus of the paper does not aim at providing an in-depth analysis of the economic causes and consequences of the revealed trends. In face of alternative explanatory hypotheses—such as accelerating non-neutral technical or organizational change, increased exposure to international competition, reduced union power, or more flexible labor market institutions—estimates of structural models as well as reduced-form approaches using richer data sets may be expected to complement the descriptive evidence. The papers in the following chapters expand along these lines.

Chapter 3 (Fitzenberger and Kohn, 2006) focusses on the link between wage structures and employment in a labor demand framework. Using data from the IABS 1975–1997 for West Germany, it incorporates age and skill as important dimensions of labor heterogeneity.

On the one hand, the evolution of age-specific skill wage premia shows that the age profiles of skill wage differentials have not moved in parallel fashion over time, but rather experienced a twist. Accordingly, it is unlikely that these developments are associated merely with age and time effects applying uniformly to all cohorts. On the other hand, we observe a break in the inter-cohort trend of skill- and age-specific relative employment such that young birth cohorts do not follow the path of the older ones towards further skill upgrading. The empirical evidence thus suggests the existence of cohort effects affecting the evolution of both skill wage premia and relative employment. Following the testing approach suggested in MaCurdy and Mroz (1995), we find such cohort effects for both relative employment and wage premia.

A coherent operationalization of wages and employment in a labor demand framework is generally difficult due to the heterogeneous nature of the input factor labor. We extend the structural approach of Card and Lemieux (2001). In this set-up based on a nested CES model, the simultaneous inclusion of skill and age as dimensions of heterogeneity not only enables the separation of age, time, and cohort effects, but also facilitates the estimation of a specification with a relatively large number of different input factors. Moreover, the model incorporates steady skill-biased technical change. We estimate the model with and without instruments taking account of the endogeneity of both wages and employment. Our preferred specifications estimate the elasticity of substitution between skill groups to range between 4.9 and 6.9, and the elasticity of substitution between age groups between 5.2 and 20.1. Compared to the literature, these numbers are rather high. In international comparison, this finding reflects the fairly small amount of over-all wage dispersion in Germany as well as the relatively compressed distribution of skills.

Based on the estimated parameters, we conduct some simulation experiments. In particular, we simulate the magnitude of wage changes for the different skill groups that would have been necessary to reduce skill-specific unemployment rates in 1997 by one half. With wage changes equal for all age groups within the respective skill classes, this would have left the wage structure within skill groups unaffected. The necessary nominal wage reductions range between 8.8 and 12.2% and are the higher the lower the employees' qualification. This finding provides evidence for the existence of wage compression—relative to a situation with reduced unemployment, there is in fact too little wage dispersion across the different skill groups.

Our analysis shows the necessity to integrate different dimensions of heterogeneity into empirically meaningful models of labor demand. The nested CES approach allows to do this in a parsimonious way and the results give rise to a number of interesting questions for future research: What are the reasons for the high degree of substitutability among German workers? What led to the slowdown of skill upgrading? And does it continue in more recent years? The literature on technical and organizational change (Autor, Levy, and Murnane, 2003; Spitz-Oener, 2006) suggests that skill contents and job tasks changed significantly in recent years. Is there a polarization of the labor market into "lousy and lovely jobs" as observed by Goos and Manning (2003) for Great Britain? And how do labor market institutions contribute to any of the developments?

Chapter 4 (Fitzenberger, Kohn, and Wang, 2006) turns towards trade unions as one important labor market institution. Trade unions bargain for higher wages, equal pay, reduced working hours, fair working conditions, or employment protection. However, in Germany—as in a number of other countries—the results of union activity apply to most of the workers irrespective of membership. Membership is not compulsory and closed shop regulations are illegal. The public good character of core services offered by trade unions may give rise to free-rider behavior.

Why do people join a union, then? Our study draws on micro-data from the German Socio-Economic Panel (GSOEP) to estimate determinants of individual membership decisions for West (1985–2003) and for East Germany (1993–2003). The merits of this approach are threefold. First, our findings quantify the influence of socio-demographic personal characteristics, such as age or marital status; the influence of workplace characteristics, i. e., match, firm, or industry-specific effects; and the influence of attitudinal factors for the individual choice to be or not to be a union member. Determinants are allowed to differ between East and West Germany as well as over time, and the application of a Chamberlain-Mundlack-type correlated random effects probit model controls for unobserved heterogeneity.

Second, we use our estimates to project net union density (NUD), defined as the share of employed union members in the number of employees. The projections for East and West Germany, which consistently trace the trends towards deunionization in both parts of the country, are then analyzed by means of decomposition techniques in order to shed light on the differences in NUD between East and West Germany and on the corresponding changes over time. We find that changes in the composition of the work force only played a minor role for the observed decline in both East and West Germany. In East-West comparison, the West German work force exhibits more attributes supporting union membership. The higher union density in East Germany in the year 1993 and the stronger subsequent decline therefore reflect a lower quality of membership matches resulting from a widespread, arbitrary membership recruitment directly after unification.

Third of all, while the erosion of union membership in Germany is likely to weaken unions' bargaining power and therefore their impacts on the labor market, the availability of adequate data from union records to study these impacts is limited. From 1991 onwards, only aggregate numbers for unified Germany are available, and unions' publications do not distinguish between employed members on the one hand and unemployed, retired, or student members on the other. Yet this distinction is important from an economic point of view. Net union density is a better measure of union power than gross union density because it is more closely related to the union's financial resources and to the potential to mobilize workers within firms. Union power differs significantly between different labor market segments. For example, unions are traditionally strong in manufacturing industries, but they are of minor importance in personal service sectors. Official membership information does not distinguish between sufficiently homogenous segments.

Therefore, our microeconometric membership estimations can be used to predict union densities for homogeneously defined labor market segments, such as industries and/or regions. These predictions can then be employed to study the impact of unionization on economic performance, and on employment and the structure of wages in particular.

Chapter 5 uses predictions of net union densities as derived in the preceding chapter in order to study the impact of union power on the structure of wages in West Germany.[1] In this study we apply a cell-level approach, focussing on labor market segments which correspond to the structure of the German wage bargaining system. This is important since in contrast to Anglo-Saxon countries, it is not meaningful to estimate a wage effect of individual union membership in Germany, where collective agreements on individual membership premia are forbidden by constitutional law. Given the high rate of collective bargaining coverage in the German labor market, union-bargained wages apply to the major part of all employees and unions influence the wage structure of members as well as of non-members.

How does union power affect wage levels and the degree of wage dispersion against this background? Do unions aim to realize "equal pay for equal work"? That is, does the impact of unions reduce residual wage inequality between employees with similar observable characteristics? Moreover, does union power also reduce wage dispersion between employees with different characteristics? Are the effects asymmetric across the conditional wage distribution, corresponding to a minimum wage argument, for instance? Is there a trade-off between reduced inequality and higher wage levels, such that a reduction in wage inequality comes along with a lower wage level? And did any of the effects change over time?

In order to answer these questions we combine data from two different data sets. First, we use GSOEP-based membership estimations to calculate net union densities for labor market segments defined by the dimensions industry, skill, age, and time. The density estimates are then imputed to the IABS 1975–1997, whose larger sample size allows for detailed investigations into the distribution of wages. Using the combined data set, we analyze the correlation of NUD on the one hand and different measures of the wage structure within and between labor market segments on the other.

Our results show that a higher net union density is ceteris paribus not only associated with reduced residual wage dispersion, but also with a lower wage level. This—possibly surprising—result is in line with an insurance motive for union membership. As discussed by Agell and Lommerud (1992) and Burda (1995), a higher net union density can be accompanied by a lower wage level if unions have a strong preference for wage equality and also want to prevent negative employment effects. Skill wage differentials also turn out smaller in segments with strong

[1] This chapter is a translation of our German paper Fitzenberger and Kohn (2005).

unions—the union's impact even goes beyond "equal pay for equal work". In line with a minimum wage interpretation of union-bargained wages, the wage distribution is compressed disproportionately from below. Union effects further vary with the age of the workforce and over time, but there are no clear trends against the background of the perpetual decline in union membership since the 1980's.

The analysis so far can not take account of a possible endogeneity of union density, as the implementation of structural models as well as an appropriate instrumentation of labor market institutions prove to be intricate. In future research, the use of richer linked employer-employee data such as the German Structure of Earnings Survey or the linked IAB data might provide valid instruments. It further makes sense to disentangle the effects of union density and collective bargaining coverage (Fitzenberger, Kohn, and Lembcke, 2006). Finally, union effects on the wage structure and on employment should be analyzed simultaneously.

2 Rising Wage Dispersion, After All! The German Wage Structure at the Turn of the Century

2.1 Introduction

The structure of wages is crucial for economic performance and the evolution of employment in particular; see the handbook article of Katz and Autor (1999) and the more recent survey of Autor, Katz, and Kearney (2005b). With the growing availability of large micro data sets not only the wage level, but also the degree of wage dispersion or compression has received increasing attention. The evolution of the West German wage structure between the mid-1970s and the mid-1990s has been extensively studied. By and large, the wage structure has been found to be relatively compressed in international comparison and rather stable over time; see Fitzenberger (1999) and Prasad (2004) and the literature cited therein. Returns to human capital components as well as residual wage inequality showed fairly little variation. In face of an ongoing skill-biased technical change (Acemoglu, 2002), this "unbearable stability" (Prasad, 2004) is considered a key determinant for the growing unemployment among low-skilled workers and it is frequently attributed to institutional rigidities.

Studies of the East German wage structure report an even higher degree of wage compression in the late years of the GDR, reflecting the egalitarian doctrine of the socialist system; see Krueger and Pischke (1995). This finding of strong wage compression still holds for the early years after the German unification. Exceptionally flat age-earnings or experience-earnings profiles suggest that experience accumulated under the old system is poorly remunerated afterwards. The unification shock led to a massive depreciation of human capital. However, as post-unification labor market cohorts started to age, wage dispersion increased, catching up to the West German level; see Franz and Steiner (2000) and Burda and Hunt (2001).

More up-to-date data lately allow us to trace the evolution of the wage structure towards the turn of the century. Recent evidence from survey data in Gernandt and Pfeiffer (2006) and from administrative data in Möller (2005) suggests that inequality has in fact been rising in both

East and West Germany. In this paper, I employ the recently available regional file of the IAB employment sample (IABS) 1975–2001 for a comprehensive description of the structure of wages for different labor market groups in the first decade after the German unification.

An inspection of year-specific unconditional wage distributions for the different groups generally supports the notion of rising wage inequality. As measured by the interquintile range $QD8020$, in the year 1992 dispersion was lower in East Germany than in West Germany, but it was higher by the year 2001. The increase was highest for full-time working women in East Germany, for whom $QD8020$ went up by remarkable 25 log points. Moreover, the larger part of the increase in dispersion among women happened in the lower parts of the respective distributions. Dispersion among men increased disproportionately in the upper parts, though. Convergence in wage levels between East and West Germany has not been achieved.

The subsequent analysis contributes to the literature by means of two approaches. First, I estimate wage equations in order to shed light on the determinants of observed wages. The large sample size of the IABS allows the application of quantile regression techniques, which are more flexible than the least squares estimations employed by most existing studies. Due to censoring of the wage data at the social security taxation threshold, I use censored quantile regressions (CQR). In sum, the regression results meet a-priori expectations. Age-earnings profiles not only are the steeper the higher the skill level, but they are also relatively flat in East Germany in 1992. The effect of the unification shock in fact wears out with the aging of post-unification labor market cohorts, and differences in the profiles have lessened by the year 2001. The quantile regression approach reveals significant differences in the effects across the wage distribution. The result that low-skilled women working full-time in East Germany are left particularly worse-off at the lower end of the distribution substantiates the high and asymmetric increase in dispersion for this group.

Second, I employ the decomposition technique introduced by Machado and Mata (MM, 2005), which builds on the estimation of quantile regressions, in order to shed light on (1) differences of the wage distributions between East and West Germany and (2) changes of the wage structure over time. The MM decomposition is well-suited to depict heterogeneous characteristics and coefficients effects across the wage distribution. In East-West comparison, differences in the composition of the work force turn out to be largely negligible for men. However, characteristics of full-time working women are mostly in favor of higher wages in the East. Yet this effect ceases to apply at the lower end of the distribution in 2001. With respect to the evolution of wages over time, characteristics effects capture major parts of the respective wage increases in the upper half of the wage distribution for West Germany. This finding reflects a skill upgrading in the work force. Restructuring and skill upgrading yet played only a minor role in explaining the wage increases in East Germany. For women in the lower parts of the distribution the characteristics effect even worked towards real wage cuts, substantiating also the particular increase in wage dispersion among this group.

With these two approaches, the paper goes beyond the recent studies of Möller (2005) and Gernandt and Pfeiffer (2006) which also report rising wage dispersion in Germany. Using the IABS 2001, Möller compares raw decile ratios of wage distributions for some selective labor market groups, but he does not investigate into the nature of observed differences by means of regression or decomposition techniques. Gernandt and Pfeiffer do employ wage regressions and decompositions, but their analysis is restricted by the small sample size of the GSOEP survey data such that they do not run separate analyses for women and have to rely on OLS regressions and the decomposition technique introduced by Juhn, Murphy, and Pierce (1993). As it turns out in this paper below, the more flexible MM decompositions unveil important differences across the respective distributions.

The course of the paper is organized as follows. Section 2.2 starts out from related analyses of the German wage structure in the literature. It introduces the data at use and offers a snapshot of raw wage distributions for different labor market groups. Section 2.3 introduces the estimation approach and discusses estimation results. The particular focus is on differences in estimated coefficients for age and skill and on the shape of age-earnings profiles. Decomposition techniques for the setting at hand are introduced in section 2.4. The subsequent discussion of results scrutinizes patterns in the respective wage distributions and discusses the effects underlying the wage differentials between East and West Germany as well as the changes of the wage structure over time. Section 2.5 concludes.

2.2 Approaching the German Wage Structure

The evolution of the West German wage structure between the mid-1970s and the mid-1990s has been extensively studied since large micro data sets have become available. Studies used the survey data provided by the German Socio-Economic Panel (GSOEP) or the administrative IAB employment samples (*IAB-Beschäftigtenstichproben, IABS*). By and large, the wage structure has been found to be relatively compressed in international comparison and rather stable over time; see Fitzenberger (1999) and Prasad (2004) and the literature cited therein. Returns to human capital components such as skill and experience as well as residual wage inequality showed fairly little variation. In face of an ongoing skill-biased technical change (Acemoglu, 2002), this “unbearable stability” (Prasad, 2000) is considered a key aspect for the growing unemployment among low-skilled workers. Compression and stability is often attributed to institutional factors. However, some degree of variability is found by a few studies with more in-depth focus. For example, Riphahn (2003) reports higher income inequality among foreign workers, Fitzenberger

(1999) reports some changes in the upper and in the lower parts of the wage structure, and when looking separately at different age groups Fitzenberger and Kohn (2006) find that there was quite some variation in the skill premia across age: cohort effects differently affected the different skill groups. Observed trends prove consistent with steady skill-biased technical change.

In their early study of the East German wage structure, Krueger and Pischke (1995) use the 1988 Survey on Income of Blue and White Collar Households in the GDR (*Einkommensstichprobe in Arbeiter- und Angestelltenhaushalten*) and the retrospective 1989 information of the 1990 GSOEP-East to find an even more compressed wage structure in the late years of the GDR, expressing the egalitarian doctrine of the socialist system. Follow-up comparative studies using different GSOEP waves[1] confirm this effect for the first years after the German unification. In particular, they report flat age-earnings or experience-earnings profiles in the East. The findings suggest that experience accumulated under the old system is poorly remunerated afterwards. The unification shock led to a massive depreciation of human capital. However, as post-unification labor market cohorts start to age, increasing wage dispersion is observed in East Germany during the 1990s.

More up-to-date administrative data for both parts of the country have recently been made available with the regional file of the IAB employment sample (IABS) 1975–2001. This version of the IABS is a 2% random sample of German social security accounts; see Hamann et al. (2004) for a description of the data set.[2] While excluding mainly self-employed workers and civil servants, the IABS covers about 80% of all employed persons. Employment in East Germany is included from 1992 onwards. The IABS offers a large sample size and—due to its administrative character—a reliable quality of data. In particular, the wage data are very accurate compared to survey data. On the downside, the data set provides relatively few covariates and no information on working time except from a distinction between full-timers and part-timers. Besides, the wage data are top-coded at the social security taxation threshold (SSTT).

Möller (2005) uses the years 1992–2001 of the IABS to compare raw decile ratios of log wage distributions for some selective labor market groups in 1992 to the respective ratios in 2001. His main findings are that wage inequality has generally been rising between 1992 and 2001 and that the rise in inequality has been more pronounced for low-skilled compared to medium-skilled

[1] Schwarze and Wagner (1992), Schwarze (1993), and Bird, Schwarze, and Wagner (1994) also use the retrospective information for 1989 in addition to waves up to 1991. Burda and Schmidt (1997) employ the waves 1990–1993. Steiner and Wagner (1997a), Franz and Steiner (2000), as well as Steiner and Hölzle (2000) estimate wage regressions based on the waves 1990–1995 or 1990–1997, respectively. Burda and Hunt (2001) compare the waves 1990–1999 and Hunt (2001) studies wage growth and job mobility in East Germany based on the waves 1990–1996. She concludes that the observed wage growth patterns provided insufficient incentives for worker mobility, which impeded efficient restructuring and employment recovery.

[2] For further information (on antecedent versions of the IABS) see also Bender, Hilzendegen, Rohwer, and Rudolph (1996) and Bender, Haas, and Klose (2000).

workers and for women compared to men. Starting out at a lower level in 1992, wage inequality in East Germany has largely caught up with the level of inequality in West Germany by 2001. GSOEP survey data employed by Gernandt and Pfeiffer (2006) suggest that the trends towards increasing inequality continued at least until the year 2004.

In this paper, I also employ the years 1992–2001 of the IABS. In order to give a comprehensive description for different groups in the labor market I take advantage of the large sample size and consider separate distributions for men working full-time, women working full-time, and women working part-time in East and West Germany in each of the years 1992–2001. For each of these subsamples, I select individuals aged between 25 and 55 years who are not currently in education. Marginal part-time workers (*geringfügig Beschäftigte*) are not included in the analysis in order to avoid spurious effects through changes in the employment of this group.

Figure 2.1 depicts the evolution of log nominal daily wages for the different labor market groups.[3] The deciles changed rather smoothly over the period 1992–2001 so that it makes sense to focus on the two boundary years in the following. Table 2.1 depicts median wages $LNW50$ and percentile differences $QD8050$ and $QD5020$ for the different groups in 1992 and 2001. As expected, the wage level is generally higher for men compared to women and for workers in West Germany compared to East Germany in 1992. Until 2001, the gender wage gap narrowed especially in East Germany. The East-West wage gap in the wage level also went down, but persisted to some degree for full-time employees. Figure 2.2 reveals that convergence in—nominal as well as real—wage levels took place until the year 1996, but then basically stopped: Starting out at 58%, 34%, and 17% in 1992, the respective nominal differences for men, full-time working women, and part-time working women all shrunk by 7 to 10 real log points (pp). However, only little variation is observed from 1996 on. Nominal differences of 38–40% and 18–20% remain for full-time working men and women, but there is virtually no more difference for part-time working women.[4]

Table 2.1 further shows that wage dispersion as measured by the percentile differences generally increased for all groups between 1992 and 2001. With the only exception of part-time working women the increase was considerably stronger in East Germany than in West Germany. By the year 2001, the level of wage dispersion in the East even exceeded the level in the West. Moreover, there are remarkable differences across groups. For Men in West Germany, $QD8050$

[3] At this point I examine nominal wages in order to facilitate East-West comparisons because it is not clear a priori which price deflator and which base year to choose when comparing East and West Germany in real terms; see the discussion in Franz and Steiner (2000). When comparing East and West German wage levels in figure 2.2 below I also present alternative price normalizations. All comparisons across time in section 2.4 are based on real wages, deflated by consumer price indices.

[4] This effect has already been extensively discussed in the literature; see, e. g., Burda and Schmidt (1997) and Burda and Hunt (2001).

increased by about 6 pp and $QD5020$ by 3 pp, adding up to an increase in the interquintile range $QD8020$ of 9 pp. The larger part of this increase therefore is due to changes in the upper part of the distribution.[5] Since the 80th percentile for men in East Germany is censored in 1992, an analogous statement for this group cannot be inferred directly from table 2.1. Yet the results in section 2.4 show that wages for men in East Germany also went up disproportionately in the upper part of the distribution.[6] Having said that, wage inequality among full-time working women increased disproportionately in the lower half of the distribution, and most strikingly so in East Germany: whereas $QD8050$ and $QD5020$ went up by 3 and 4 pp in the West, the respective numbers for East Germany are 8 and 17 pp, adding up to a remarkable increase of the interquintile range $QD8020$ of 25 pp.

In what follows, the observed distributions are investigated by means of wage regressions for the years 1992 and 2001 to capture the changes over time. The application of (censored) quantile regressions allows to look at between and within inequality, and it sets the stage for the decomposition analyses in section 2.4. Considering the years 1992 and 2001 is warranted for the following two reasons. First, both years are similar with respect to their location in the West German business cycle: Whereas the unification boom faded out in 1992, the year 2001 marked the end of the new economy boom. Second, the labor force in East Germany dropped sharply from about 10 to below 7 million in the course of the German unification and most of the immediate downturn took place in 1990 and 1991; see Kommission (1996). Net emigration from East Germany was highest between 1989 and 1991; see Hunt (2006). 1992 was the first year with positive GDP growth in East Germany after the unification shock (Burda and Hunt, 2001) and thus is the first year not heavily exposed to distortions resulting from the unification.

2.3 Wage Regressions

Let $Y_{s,i} \equiv \ln W_{s,i}$ denote log wages for individuals i, drawn from a distribution $F_s(Y_s)$ in an adequately defined labor market segment s. Given the focus of this paper one might think of segments as regions (East and West Germany) or different points in time (years).

Since the wage data used use are censored from above at the social security taxation threshold c_s, one observes only $\tilde{Y}_{s,i} = \min\{Y_{s,i}, c_s\}$. One thus might apply Tobit regression (after Tobin,

[5] This finding is similar to the trends observed by Fitzenberger (1999) for the period 1975–1990.

[6] The conclusion is also corroborated by Möller's (2005) result for the core group of medium-skilled men.

1956) to estimate the conditional expected value $E(Y_s|X_s)$ based on covariates X_s, assuming normality of the error term u_s in

$$Y_s = E(Y_s|X_s) + u_s = X_s\beta_s + u_s. \tag{2.1}$$

A more informative approach is to employ quantile regressions, which do not only capture the expected value, but the entire distribution. As introduced by Koenker and Bassett (1978) and generalized by Powell (1984, 1986), conditional quantiles

$$Q_\theta(Y_s|X_s) = X_s\beta_s(\theta) \tag{2.2}$$

in the case of censoring from above can be estimated for a given quantile $\theta \in (0, 1)$ by minimizing over β_s the objective function

$$N_s^{-1}\sum_{i=1}^{N_s}\rho_\theta(\tilde{Y}_{s,i} - \min\{X_{s,i}\beta_s, c_s\}), \tag{2.3}$$

where the residuals $u_{s,i}$ for individuals $i = 1, ..., N_s$ are weighted in an asymmetric way by the check function

$$\rho_\theta(u_{s,i}) = \begin{cases} \theta u_{s,i} & \text{for} \quad u_{s,i} \geq 0 \\ (\theta - 1)u_{s,i} & \text{for} \quad u_{s,i} < 0 \end{cases}. \tag{2.4}$$

There are different algorithms to solve this non-convex optimization problem in the literature; see, e. g., Buchinsky (1994), Fitzenberger (1997a, 1997b), or Koenker and Park (1996). In the following applications, I apply the Buchinsky algorithm as well as the Fitzenberger algorithm for different starting values and choose the respective best estimator in terms of the objective function (2.3). Heteroscedasticity consistent standard errors are obtained by means of design matrix bootstraps. Here, it asymptotically suffices to draw on observations for which predicted values are not censored; see Bilias, Chen, and Ying (2000).

Quantile regressions are particularly suited for the purpose of this paper because they do not only reveal differences between, say, different skill or age groups, but also allow these differences to differ across the wage distribution.

2.3.1 Coefficients Across the Distribution

The estimated log wage equations include a set of formal skill dummies (low-skilled d_l: workers without vocational training and without university degree, medium-skilled (base category): those with vocational training and no university degree, and high-skilled d_h: employees with university or technical college degree)[7], (normalized) age, and age squared (agesq). In order to allow for different age-earnings profiles across skill groups I include interaction terms of skill and age as well as skill and agesq, yielding the following specification which is estimated separately for all segments s:

$$\begin{aligned} Y_{si} &= \beta_{1s} + d_{l,si}\beta_{2s} + d_{h,si}\beta_{3s} + \text{age}_{si}\beta_{4s} + \text{agesq}_{si}\beta_{5s} \\ &\quad + d_{l,si}\text{age}_{si}\beta_{6s} + d_{l,si}\text{agesq}_{si}\beta_{7s} + d_{h,si}\text{age}_{si}\beta_{8s} + d_{h,si}\text{agesq}_{si}\beta_{9s} + u_{si}. \end{aligned} \tag{2.5}$$

All regressions further include a set of industry dummies (16 industries as provided with the IABS 1975–2001) and a dummy for individuals working in Berlin. Observations are weighted by the length of the respective employment spells. Summary statistics of the covariates are displayed in tables 2.2 and 2.3.

Figures 2.3 to 2.8 show coefficient estimates for censored quantile regressions (CQR) at different deciles of the distributions as well as Tobit coefficients for comparison. The results are grouped by labor market groups (full-time working men, women working full-time, and women working part-time) and years (1992 and 2001), and each of the figures shows coefficients for West (left panel) and East Germany (middle panel) as well as differences between the two parts of the country (right panel).

In general, the estimated effects are significantly different from zero. Merely some age×skill interactions in East Germany prove insignificant in some parts of the distributions. Moreover, CQR coefficients generally vary significantly across the distribution and differ from the more restrictive Tobit estimates, with the only exception of part-time working women, for whom the confidence bands are relatively wide. The censoring problem is most severe for older high-skilled employees. The interaction terms of age and high skill thus are somewhat sensitive. For example, the median coefficient of age×high skill for full-time working men in West Germany 2001 is extraordinarily low, whereas the median effect of agesq×high skill jumps up. At the 60% quantile, things are reversed. This effect might affect the shape of single age-earnings profiles (see this section below), but its impact on predictions (as used for the decomposition analyses in the next section) can be expected to be small.

Due to the inclusion of the interaction effects, the interpretation of some of the coefficients is not apparent, and I resort to looking at age-earnings profiles in the next subsection. Nevertheless, there are some notable differences of coefficients across quantiles. For example, the effect of age is found to become steeper and more concave at higher quantiles for full-timers. The (negative)

[7] In order to deal with measurement error in the education variable when defining skill groups, I correct the skill information such that formal degrees an individual has once obtained are not lost later on; see also Fitzenberger (1999).

base effect of low skill tends to be smallest at low quantiles, and so does the (positive) base effect of high skill. These results are well in line with the predictions of human capital theory; see Becker (1964) and Card (1999).

Looking at West-East differences in the coefficients for the year 1992, differences in the base effects of skill turn out to be relatively small. The base trajectory of age is steeper and slightly more concave for men in West Germany, but the picture is reversed for full-time working women, most strikingly in the lower half of the distribution. Differences in the returns to skill among part-time working women are relatively large in the lower half of the distribution. In the year 2001, the differences in the age effects are basically the same as in 1992, but now low-skilled men are particularly worse off in East Germany in the lower half of the distribution. On the other hand, the base return to high skill in East Germany has increased disproportionately at the upper end of the distribution so that one finds a negative difference there.

Changes of the coefficients between 1992 and 2001 can be inferred from figures 2.9 to 2.14, which rearrange the estimation results in the left two panels and show the changes between 1992 and 2001 explicitly in the right panel. In West Germany, the base wage has increased, and for full-timers this effect was stronger at higher quantiles. Base skill differentials for both men and women (except for high-skilled part-timers at the top of the distribution) have increased, hinting at an increasing inequality between skill groups. The base returns to age only changed little, though. The changes in East Germany are qualitatively comparable to those in the West. Yet the baseline increased even more distinctly over time, and more pronounced differences across quantiles hint at a higher degree of within dispersion. The negative base wage premium for low skill has grown most strikingly at the lower end of the distribution, whereas the base premium for high-skilled men has grown most at the top of the distribution.

2.3.2 Age-Earnings Profiles

Figures 2.15 to 2.18 present age-earnings profiles used to judge differences in the remunerations of formal skill and age. The two panels display results for West and East Germany, respectively. In most cases, the profiles have the familiar concave form. However, some profiles for high-skilled employees, for whom the censoring problem is most severe, should be interpreted with caution; compare the discussion above.

Figure 2.15 displays results of the median regressions for different skill groups. Trajectories are generally the steeper the higher the skill level. The only exception is the group of low-skilled women working part-time in East Germany in 2001 which exhibits an exceptionally steep profile. In West Germany, the profiles for women are usually flatter than those for men, but men and women do not differ as much in East Germany.

In East-West comparison, the profiles in the East are flatter and decrease more pronouncedly for older workers in the year 1992. This finding mirrors the low returns to age or experience as human capital components in East Germany in the aftermath of the unification. Yet the difference has lessened by the year 2001, indicating some recovery of returns. Whereas the profiles for West Germany are rather similar between 1992 and 2001, changes occurred in the East, where the profile for high-skilled men became particularly concave—returns recovered most distinctly for high-skilled post-unification cohorts. On the other hand, the profiles of low-skilled women improved for part-timers, but deteriorated for full-timers. Given the position of the low-skilled at the lower end of the unconditional wage distribution, this effect contributes substantially to the asymmetric rise in dispersion among women working full-time in East Germany.

The general picture is also reflected in figures 2.16 to 2.18 which display skill-specific profiles at different quantiles (20%, 50%, and 80%). Standard profiles with steeper trajectories in higher regions of the distribution are primarily observed for the core labor market group of male full-timers with an apprenticeship degree. When looking at (full-time as well as part-time working) women in West Germany in the year 2001, one finds an analogous standard ordering of the profiles for the high-skilled, but a reversed ordering for the low-skilled: Women with low formal qualification gain most from accumulating experience at the lower end of the pay scale. In East Germany, the profiles decrease for older workers across all quantiles. Yet high-skilled men at younger age gained most in the upper part of the distribution and the position of older low-skilled women deteriorated particularly at the lower end. Again, the findings underline the depreciation of human capital and the asymmetric recovery in the aftermath of the unification.

2.4 Decomposing Differences Across Wage Distributions

The above regression analyses provided detailed insights into the remuneration of observed worker characteristics in different labor market segments and in different parts of the wage distribution. Decomposition analyses are well-suited to complement the regression evidence by answering the question whether differences in observed distributions result from differences in estimated coefficients or from differences in the composition of the workforce. I focus on differences between East and West Germany and on changes of the respective wage structures over time.

A Blinder (1973)-Oaxaca (1973)-type decomposition for the difference between the expected wages in two segments s and $\tilde{s}$ is:

$$E(Y_s|X_s) - E(Y_{\tilde{s}}|X_{\tilde{s}}) = (X_s - X_{\tilde{s}})\beta_s + X_{\tilde{s}}(\beta_s - \beta_{\tilde{s}}). \tag{2.6}$$

To apply the Blinder-Oaxaca (B-O) decomposition in case of censored data, I evaluate equation (2.6) at mean values of the characteristics and use the coefficients estimated by means of Tobit regressions.[8]

The first summand on the right hand side of equation (2.6), traditionally labeled "characteristics effect", captures the part of the difference that is attributable to differences in the covariates across the two segments. The second summand known as "returns" or "coefficients effect" captures the part of the difference that is attributed to differences in the returns to the covariates. When decomposing West-East wage gaps in the next section, I choose the counterfactual $X_{\text{East}}\beta_{\text{West}}$ to answer the question what the expected log wage would have been, had a population with the same distribution of characteristics as East Germany faced returns to characteristics as in the West.[9] The approach assumes that the West German returns are the relevant benchmark for the distribution in the absence of any "discrimination". In case of the comparison across time in section 2.4.2 the counterfactual $X_{1992}\beta_{2001}$ hypothesizes what the expected wage would have been in face of returns in the year 2001, had the distribution of characteristics not changed since 1992.[10]

A further method introduced by Juhn, Murphy, and Pierce (1991) and applied in a series of papers by Blau and Kahn (1992, 1994, 1997) also decomposes the change of a wage gap over time. This approach has got the additional merit that it decomposes also residual effects into a quantity and a price effect. However, it suffers from the shortcoming that it assumes unique coefficients across segments s and $\tilde{s}$. What is more, the decomposition of the residual terms is inapplicable in the case of censored data, in which residuals can only be used for uncensored observations.

The main disadvantage of all techniques discussed so far is that all of them consider only mean effects. In contrast, Machado and Mata (2005) build on quantile regressions to decompose differ-

[8] In contrast to the traditional OLS case, however, the predicted conditional difference does not necessarily coincide with the observed mean difference. "Observed" mean wages in the censoring case have to be estimated by means of Tobit regressions on a constant.

[9] It is well known that the partition depends on the ordering of the effects and that the decomposition results may not be invariant with respect to the choice of the involved counterfactual $X_{\tilde{s}}\beta_s$; see the surveys of Oaxaca and Ransom (1994) and Silber and Weber (1999). Therefore, the choice of a counterfactual should be guided by the question of economic interest.

[10] There are alternative methodologies to the standard B-O decompositions in the literature. In light of the present focus on differences in two dimensions, techniques to decompose changes of wage gaps over time in one single exercise—as proposed by Smith and Welch (1989) or Wellington (1993)—would be of particular interest. However, I opt to consider both decompositions separately for two reasons. First, any combination of involved counterfactuals—be it with or without interaction terms between the differences in characteristics and differences in coefficients—bears an even higher degree of arbitrariness; see Le and Miller (2004). Second, and most importantly, each of the two comparisons, the differences between East and West Germany as well as the changes of the wage distributions within the two regions over time, is interesting on its own.

ences across entire distributions. They propose an estimator $F_s^*(Y_s)$ of the marginal distribution of wages which conforms to the linear conditional model (2.2) as follows:

1. Draw M numbers $\theta^1, ..., \theta^M$ at random from a uniform distribution $U(0,1)$.
2. For each θ^m, estimate the conditional quantile (2.2), using the sample $\{Y_{s,i}, X_{s,i}\}_{i=1}^{N_s}$. This yields coefficient estimates $\hat{\beta}_s(\theta^1), ..., \hat{\beta}_s(\theta^M)$.
3. Draw M random draws $X_s^1, ..., X_s^M$ from the sample $\{X_{s,i}\}_{i=1}^{N_s}$.
4. Then, the data set $\{Y_s^{*m} \equiv X_s^m \hat{\beta}_s(\theta^m)\}_{m=1}^M$ constitutes a random sample from $F_s^*(Y_s)$.

An estimator $F_s^*(Y_s(X_{\tilde{s}}))$ of the counterfactual marginal distribution, which relies on the coefficients of segment s but on the characteristics of segment $\tilde{s}$, can be obtained in an analogous way by drawing resamples from $X_{\tilde{s}}$ rather than from X_s in the third step. The Machado/Mata (MM) decomposition based on the estimated distributions therefore writes

$$\begin{aligned} \hat{F}_s(Y_s) - \hat{F}_{\tilde{s}}(Y_{\tilde{s}}) &= F_s^*(Y_s) - F_{\tilde{s}}^*(Y_{\tilde{s}}) + \epsilon \\ &= [F_s^*(Y_s) - F_s^*(Y_s(X_{\tilde{s}}))] + [F_s^*(Y_s(X_{\tilde{s}})) - F_{\tilde{s}}^*(Y_{\tilde{s}})] + \epsilon, \end{aligned} \tag{2.7}$$

where $\hat{F}_s(\cdot)$ denotes an estimator of the distribution based on the observed sample. Similar to the B-O decomposition, the term in the first brackets on the right hand side of (2.7) is a characteristics effect, and the one in the second brackets a returns effect. Provided that the linear specification (2.2) is appropriate, the residual term ϵ is negligible for large samples. With respect to the choice of a counterfactual distribution the same caveat as in the B-O case applies.

I employ the MM technique, resorting to quantile measures for the involved distributions in order to gauge the elements of the decompositions. However, a couple of adaptations are undertaken. First, I estimate CQR as explained above. Second, I follow Albrecht, Björklund, and Vroman (2001) to save computation time: Rather than drawing M random numbers for θ^m and then estimating M (censored) quantile regressions, I estimate one regression for each single percentile and then draw $M = 1000$ random draws from the distributions of the covariates for each $\hat{\beta}_s(\cdot)$. Third, and finally, predictions above the SSTT are censored to this value in order to replicate the censoring of the wage data. As a consequence, all comparisons of the simulated distributions $F_s^*(\cdot)$ consider only the respective uncensored parts.

There are also alternative approaches in the literature for decomposing differences across entire distributions. The decomposition introduced by Juhn, Murphy, and Pierce (JMP, 1993), which is also used by Blau and Kahn (1996) for cross-country comparisons and by Steiner and Wagner (1997, 1998) and Gernandt and Pfeiffer (2006) for German data, employs the distribution of residuals resulting from wage regressions to rank observations. This approach gives a structural interpretation to the regression residual. Yet it faces a couple of shortcomings. First, its focus on the distribution of residuals renders the approach as inapplicable in the case of censored data as the related (1991) approach. Second, even without censoring of the data, the JMP (1993) decomposition is valid only in the case of homoscedasticity, which is usually rejected for empirical wage regressions. Third, and most importantly, it is more restrictive than the MM technique because it assumes a single linear model to hold for the entire wage distribution, whereas the latter approach based on quantile regressions allows for flexibility across the distribution.

Autor, Katz, and Kearney (2005a) also build on the MM approach, while DiNardo, Fortin, and Lemieux (1996) exploit kernel density estimations to decompose differences in a nonparametric setting. Compared to this approach, the semiparametric MM framework is restrictive by nature. Yet by quantifying differences in the coefficients it sheds light on that part of a difference which would be left unexplained in the nonparametric framework.

2.4.1 Differences between East and West Germany

Table 2.4 reports observed and predicted West-East differences in log wages across quantiles for the years 1992 and 2001.[11] Observed and predicted quantiles of the unconditional wage distributions show a close resemblance, therefore suggesting that estimation and specification error is of minor importance. The predicted gaps thus broaden the snapshot discussion of section 2.2. Decile differences which cannot be interpreted due to the censoring problem are marked by a dot. The Tobit results reported in the last column are usually close to the values at the median.

For the group of full-time working men the gap varies between 55% at the first decile and 61% at the eighth decile in 1992. The observation that the gap at the upper end of the distribution exceeds the gap at the lower end by 6 pp indicates a higher wage dispersion in the West as compared to the East. In 2001 the East-West differential varies less between quantiles (38% at the first decile and 40% at the eighth): Wage dispersion in East Germany has caught up to a large degree. Except for the difference in the level, the picture for women working full-time in 1992 is very similar to that for males in the upper two thirds of the distribution: The gaps at the third and at the eighth decile differ by 4 pp. However, the gap of 22% at the first decile falls below the gap at the third decile by remarkable 14 pp—at the very low end of the distribution the West-East gap is less severe. This finding still holds for the year 2001, but now the differential at the third decile also exceeds the differential at the eighth by 9 pp: The upper half of this group's distribution participated most strikingly in the closing of the West-East wage gap. Women working part-time in East Germany in 1992 were relatively well off at the low and at the high end of the distribution, and the West-East differential was highest around the median.

[11] The analysis in this section is based on nominal numbers; see the discussion in section 2.2.

The differential for this group had basically vanished by 2001, though. At the first decile wages were even slightly higher in the East.

When decomposing West-East wage differentials in order to judge whether the differentials stem from different decompositions of the work force or whether employees' characteristics are remunerated differently in East and West Germany, one generally finds relatively small impacts of the characteristics. The better parts of the differentials are in most cases captured by differences in the coefficients.

For full-time working men the characteristics effect is largely negligible in both years 1992 and 2001. If anything, different characteristics explain 2 pp of the West-East differential in the upper part of the distribution in 2001. In the group of women working full-time in 1992, the characteristics effect ranges between –9 pp at the first decile and –6 pp at the eighth. It therefore is in favor of higher earnings in East Germany and most pronounced in the lower half of the distribution. In relative terms, women selecting into full-time jobs in East Germany had preferable characteristics in 1992. This tendency still holds for 2001, but to a lesser degree and mainly in the upper half of the distribution. In the lower part of the distribution the relative deterioration of characteristics contributed substantially to the worsened position in the pay scale. A similar reasoning also applies for women working part-time in 1992. However, there are only little offsetting characteristics and coefficients effects in the year 2001, by which convergence of wages has been achieved for this group.

The conclusion that differences in employees' characteristics only play a minor role in explaining East-West wage differentials is supported by the summary statistics of the covariates in tables 2.2 and 2.3. By and large, differences are very small. In both years 1992 and 2001 and for all labor market groups, the level of formal education in East Germany is higher than in the West. Only the proportion of male employees with a university degree is higher in West Germany in 2001.

The latter finding is in line with the results of the B-O decompositions in Burda and Schmidt (1997) and the JMP decompositions in Steiner and Wagner (1997a), both of which use GSOEP data for the early 1990s and report a minor importance of differences in the characteristics of the work forces. Görzig, Gornig, and Werwatz (2004), using a decomposition based on establishment-level data, compare wages in East and West Germany for the years 1994 and 1998. They stress the importance of differences in establishment types and conclude that the catching-up in the East was in part offset by an increasing share of low-wage-type establishments in East Germany. The analysis of East-West migrants in Kirbach and Smolny (2004) also concludes that only a small part of observed East-West wage gaps can be attributed to observed socioeconomic characteristics of the workers.

2.4.2 Changes in the Wage Structure Over Time

In order to analyze changes in the wage structure over time, I use real wages (normalized by consumer prices of 1992, differentiated by regions). In a setup analogous to that of table 2.4 in the previous section, the panels in table 2.5 display the observed and predicted log wage changes between 1992 and 2001. Differences of the numbers across quantiles give account of the evolution of wage inequality.

Among the group of men working full-time in West Germany, inequality as measured by percentile differences $QD8020$ has increased by 9 pp and this increase was slightly more pronounced in the upper half of the distribution. The eighth decile gained 5% while the second decile lost by 4%. Due to the censoring problem, changes at the very high end of the distribution cannot be assessed, but wages at the very low end exhibited a remarkable real loss of almost 8%. The (predicted) interquintile range $QD8020$ of 14 pp for men in East Germany shows that wage dispersion went up even more remarkably. Moreover, most of this increase (10 pp) took place in the upper half of the distribution. Yet even at the lower end real wage growth was positive for this group.

Real wages of women working full-time in West Germany hardly changed in the lower third of the distribution. Only the first decile exhibited a decline of 2%. Negative real wage growth of up to –4% is found in the lower third of the distribution for this group in East Germany. The gender wage gap in East Germany thus did not close, but rather widened in this part of the distribution. Wage growth further differed substantially at higher quantiles: Whereas Western wages increased by up to 8%, wages in the East went up by remarkable 23% at the eighth decile. The corresponding interquintile range $QD8020$ of 25 pp shows that the increase in inequality was most striking among this group.

The group of part-time working women in West Germany experienced real wage growth between 5 and 11%, with highest increases at the extreme deciles. In East Germany, the range of differences across quantiles is 9 pp. However, the biggest increase is observed in the middle part of the distribution and—well in line with the observed closing of the East-West gap for this group—the level of changes exceeds that in the West by about 10 pp.

The decomposition of the wage changes reveals characteristics effects in the range between 1 pp (in favor of higher earnings in 2001) at the first decile and 5 pp at the eighth decile for all three labor market groups working in West Germany. With shares of about one half for women and virtually full coverage for men, changes in the characteristics therefore capture the better part of the respective wage increases in the upper halves of the distributions. The finding likely reflects some skill upgrading in the prime-age work force. In fact, reconsidering the summary statistics of the covariates in tables 2.2 and 2.3, one finds that skill upgrading took place in both East and West Germany between 1992 and 2001. As the proportion of low-skilled workers decreased in all labor market groups, the proportion of high-skilled went up. This increase was more pronounced in West Germany than in the East. With respect to changes in the industry structure of the work force, employment in public and social security system services (sector 16) decreased most remarkably in East Germany.

Restructuring and skill upgrading yet played only a minor role in explaining the striking wage increase (especially in the upper half of the distribution) for men working in East Germany: The characteristics effect does not exceed 2 pp. A similar result holds for the majority of women working full-time in East Germany, but for this group the characteristics effect goes down up to –7 pp in the lower middle of the distribution. The characteristics in that part of the distribution working towards real wage cuts, the increasing inequality was driven by a more advantageous development of characteristics at the upper end. Finally, the contribution of changes in the characteristics is largely negligible across the entire distribution of wage changes for women working part-time in East Germany.

A bottom line of this exercise is that the diverse patterns of changing wage levels and increasing inequality are due to changes in the composition of the respective work forces *and* changing remunerations of relevant characteristics. This result differs from that of related studies in the literature[12], all of which use the more restrictive B-O or JMP decompositions for different periods of time and find basically no composition effects among prime-age employees.

2.5 Conclusions

The German wage structure has been rather compressed in international comparison and "unbearabl[y] stable" (Prasad, 2004) between the mid-1970s and the mid-1990s. Newly available register data from the IAB employment sample 1975–2001 now allow researchers to reinvestigate the empirical evidence for more recent years. This paper studies the evolution of wage levels and wage inequality within and between different labor market groups for the years 1992–2001. I find that wage inequality has in fact been rising in many dimensions throughout this period.

A comparison of mean wage differences reveals that convergence in wage levels between West and East Germany took place until the year 1996, but nominal differences of about 40% for men and 20% for full-time working women persisted until 2001. No more difference is observed in the wages of part-time working women.

[12] Steiner and Wagner (1998a) analyze the evolution of wage inequality among West German males by means of JMP decompositions applied to GSOEP and IABS data for the years 1984–1990. Note that their analysis for the IABS bears some problems because it only considers uncensored wages. Burda and Hunt (2001) apply B-O decompositions to the GSOEP East 1990–1999. Gernandt and Pfeiffer (2006) also use GSOEP data for 1984–2004 and apply JMP decompositions.

The inspection of year-specific wage distributions unveils rising wage dispersion. As measured by interquintile ranges $QD8020$, dispersion was generally lower in East Germany than in West Germany in the year 1992, but it caught up until 2001: Whereas $QD8020$ increased by 8 to 9 log points (pp) for men and full-time working women in West Germany, the corresponding numbers are 14 to 25 pp in the East. Moreover, the larger part of the increase in dispersion among women happened in the lower parts of the respective distributions, but dispersion among men increased disproportionately in the upper parts.

The estimation of censored quantile wage regressions provides insights into the determinants of the observed differences and changes. The bottom line of the regression results meets a-priori expectations. Age-earnings profiles not only are the steeper the higher the skill level, but they are also relatively flat in East Germany in 1992. The unification shock clearly led to a depreciation of human capital in the East. However, this effect wears out with the aging of post-unification labor market cohorts, and differences in the profiles between East and West Germany have lessened by the year 2001. The quantile regression approach further reveals significant differences in the effects across the wage distribution. The result that low-skilled women working full-time in East Germany are left particularly worse-off at the lower end of the distribution substantiates the high and asymmetric increase in dispersion for this group.

Drawing on the flexible quantile regression approach, the decomposition technique introduced by Machado and Mata (2005) is well-suited to depict heterogeneous characteristics and coefficients effects across the respective wage distributions. In East-West comparison, differences in the composition of the work force turn out largely negligible for men. However, characteristics of full-time working women are mostly in favor of higher wages in the East. Yet this effect ceases to apply at the lower end of the distribution by the year 2001.

With respect to the evolution of wages over time, characteristics effects can account for major parts of the respective wage increases in the upper halves of the wage distributions for West Germany. This finding reflects a skill upgrading in the work force. Restructuring and skill upgrading yet played only a minor role in explaining the wage increases in East Germany. For women in the lower parts of the Eastern distribution the characteristics effect even worked towards real wage cuts, substantiating again the particular increase in wage dispersion among this group.

The finding of rising wage inequality is broadly in line with the evidence in Möller (2005), who compares decile ratios for selective labor market groups and also stresses the importance to distinguish between men and women when assessing asymmetries in the evolution of wage inequality. Gernandt and Pfeiffer (2006), also reporting increasing wage inequality, do not distinguish between sexes and therefore do not give account of the striking asymmetries between the groups in East Germany. As a consequence, their JMP decompositions do not detect this effect, either.

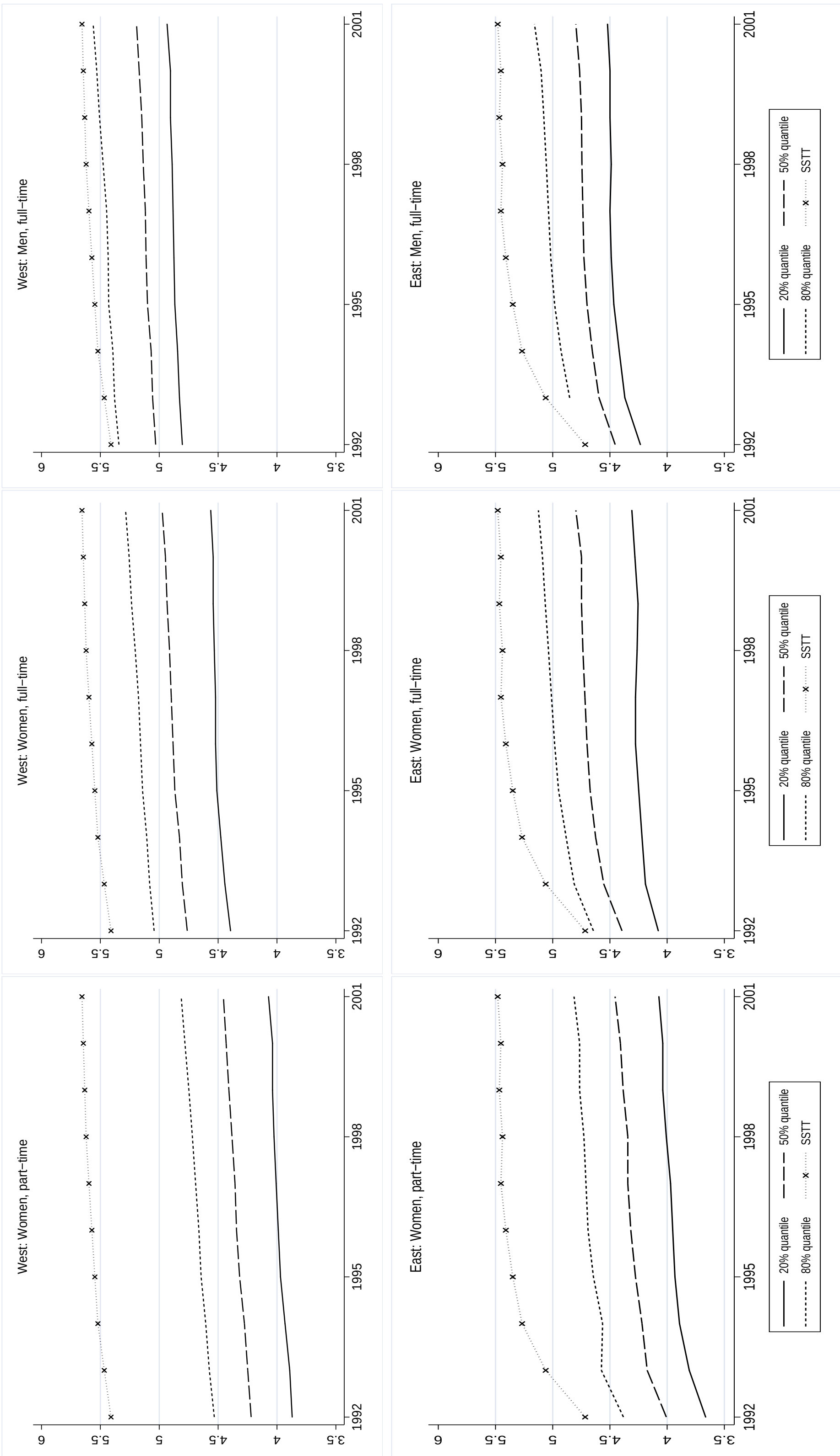
All of the results discussed in this paper are descriptive by nature. Unfortunately, the IABS provides only relatively few covariates, such that it is impossible to venture upon instrumental variable estimation or a control function approach in order to account for a possible endogeneity of educational attainment or differences in the selection into the labor market. The analysis focusses on core labor market groups and leaves aside marginal part-time workers (*geringfügig Beschäftigte*), among others. This is important to note because it renders the finding of increasing inequality even more meaningful.

An analogous argument applies with respect to migration, which is not modeled explicitly. East-West migration in the aftermath of the unification had already come down to stable numbers by the year 1992 and the evidence for the existence of a brain drain is mixed; see Arntz (2006), Büchel, Frick, and Witte (2002), and Hunt (2006). However, if emigration from East Germany during the observation period is skill- or age-biased, i. e., if migrants are in fact either better educated workers or low-skilled who have been laid-off (Hunt, 2006), the observation that wage inequality increases faster in East Germany is even more remarkable.

Finally, it is not the aim of this paper to speculate about the economic causes and consequences of the unveiled trends. In face of alternative explanatory hypotheses—such as accelerating non-neutral technical change, increasingly heterogeneous work environments, more flexible labor market institutions, or a decline in union power—estimates of structural models may be expected to complement the descriptive evidence in future research.

2.A Tables and Figures

Figure 2.1: Nominal Wage Distributions, 1992–2001



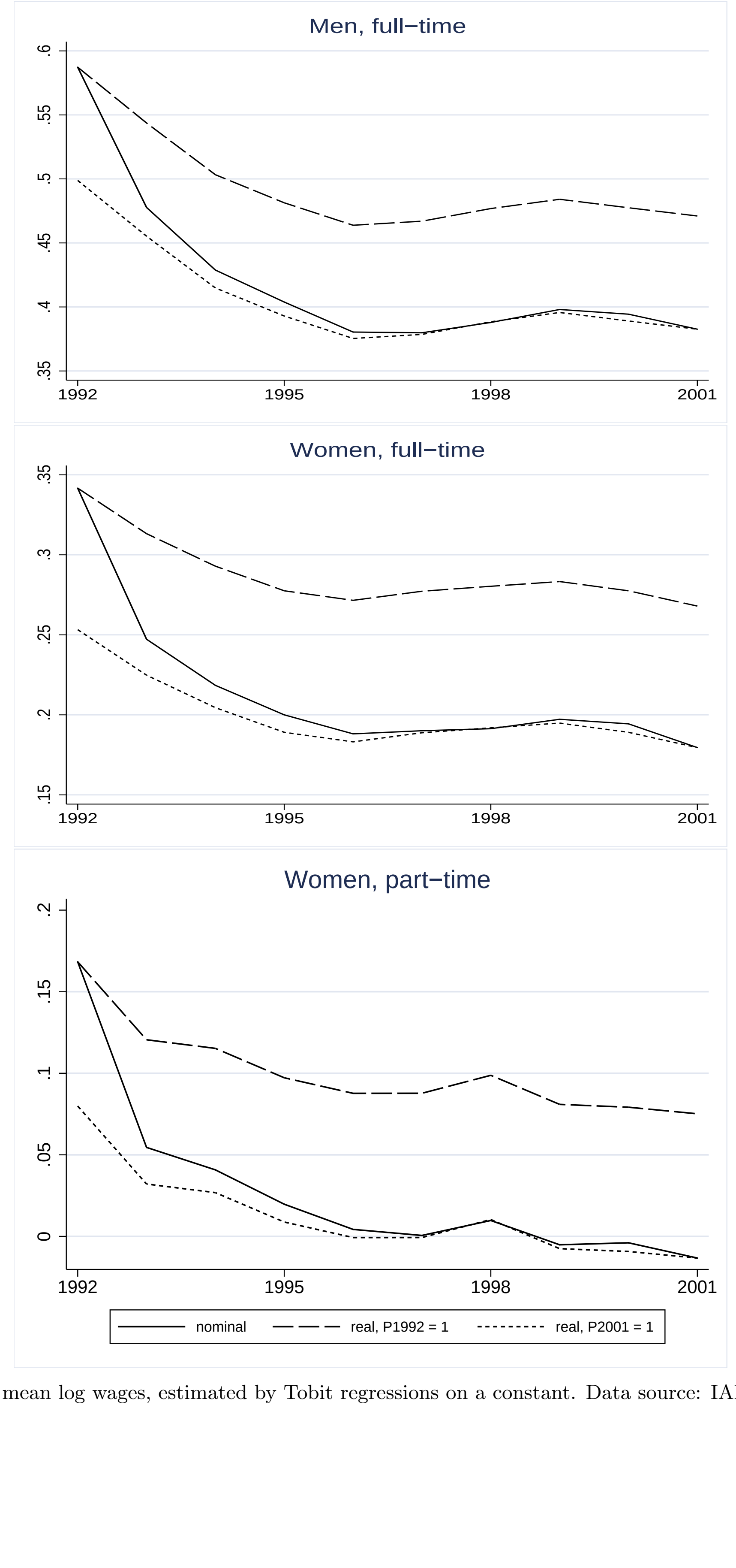
Raw quantiles of log nominal daily wage distributions. SSTT: social security taxation threshold. Data source: IABS 1975–2001.

Table 2.1: Wage Levels and Wage Dispersion, 1992 and 2001

1992	Men f.-t., West	Men f.-t., East	Wom. f.-t., West	Wom. f.-t., East	Wom. p.-t., West	Wom. p.-t., East
QD8050	0.312	·	0.281	0.250	0.313	0.375
LNW50	5.030	4.454	4.762	4.394	4.220	4.007
QD5020	0.226	0.220	0.368	0.317	0.348	0.344
2001	**Men f.-t., West**	**Men f.-t., East**	**Wom. f.-t., West**	**Wom. f.-t., East**	**Wom. p.-t., West**	**Wom. p.-t., East**
QD8050	0.369	0.362	0.311	0.327	0.359	0.359
LNW50	5.193	4.798	4.975	4.798	4.455	4.455
QD5020	0.259	0.277	0.412	0.490	0.383	0.383

Nominal log daily wages. · indicates a censored quantile. Data source: IABS 1975–2001.

Figure 2.2: West-East Wage Gaps, 1992–2001



Differences of mean log wages, estimated by Tobit regressions on a constant. Data source: IABS 1975–2001.

Table 2.2: Description and Summary Statistics of Covariates, 1992

Covariate	Description	Men f.-t., West		Men f.-t., East		Wom. f.-t., West		Wom. f.-t., East		Wom. p.-t., West		Wom. p.-t., East	
DL	dummy for low-skilled	0.111	(0.31)	0.054	(0.22)	0.160	(0.36)	0.059	(0.23)	0.181	(0.38)	0.082	(0.27)
DM*	dummy for medium-skilled	0.769	(0.42)	0.810	(0.39)	0.769	(0.42)	0.811	(0.39)	0.759	(0.42)	0.800	(0.39)
DH	dummy for high-skilled	0.119	(0.32)	0.135	(0.34)	0.069	(0.25)	0.128	(0.33)	0.058	(0.23)	0.117	(0.32)
AGE	age (25–55 years)	38.78	(8.60)	39.21	(8.64)	37.39	(8.83)	39.17	(8.50)	40.99	(7.97)	40.12	(8.73)
DSEC1	agriculture & mining	0.035	(0.18)	0.043	(0.20)	0.010	(0.10)	0.024	(0.15)	0.006	(0.08)	0.010	(0.10)
DSEC2	production of basic materials	0.097	(0.29)	0.066	(0.24)	0.042	(0.20)	0.032	(0.17)	0.021	(0.14)	0.015	(0.12)
DSEC3*	metal industry, machinery	0.145	(0.35)	0.095	(0.29)	0.042	(0.20)	0.030	(0.17)	0.021	(0.14)	0.018	(0.13)
DSEC4	vehicles & technical appliances	0.111	(0.31)	0.065	(0.24)	0.085	(0.27)	0.036	(0.18)	0.035	(0.18)	0.020	(0.14)
DSEC5	consumer goods	0.068	(0.25)	0.040	(0.19)	0.071	(0.25)	0.044	(0.20)	0.038	(0.19)	0.027	(0.16)
DSEC6	food, beverages, tobacco	0.027	(0.16)	0.024	(0.15)	0.035	(0.18)	0.030	(0.17)	0.018	(0.13)	0.016	(0.12)
DSEC7	main construction	0.067	(0.25)	0.146	(0.35)	0.009	(0.09)	0.019	(0.13)	0.008	(0.09)	0.011	(0.10)
DSEC8	subconstruction work	0.036	(0.18)	0.061	(0.23)	0.010	(0.10)	0.010	(0.10)	0.010	(0.10)	0.006	(0.08)
DSEC9	wholesale trade	0.067	(0.25)	0.041	(0.20)	0.059	(0.23)	0.032	(0.17)	0.039	(0.19)	0.030	(0.17)
DSEC10	retail trade	0.043	(0.20)	0.035	(0.18)	0.102	(0.30)	0.070	(0.25)	0.184	(0.38)	0.194	(0.39)
DSEC11	transport & communication	0.062	(0.24)	0.106	(0.30)	0.031	(0.17)	0.056	(0.23)	0.047	(0.21)	0.084	(0.27)
DSEC12	business-related services	0.095	(0.29)	0.062	(0.24)	0.139	(0.34)	0.085	(0.28)	0.119	(0.32)	0.060	(0.23)
DSEC13	household-oriented services	0.024	(0.15)	0.020	(0.14)	0.063	(0.24)	0.044	(0.20)	0.036	(0.18)	0.045	(0.20)
DSEC14	medical services	0.043	(0.20)	0.054	(0.22)	0.152	(0.35)	0.153	(0.36)	0.213	(0.40)	0.177	(0.38)
DSEC15	associations & organizations	0.023	(0.15)	0.028	(0.16)	0.069	(0.25)	0.057	(0.23)	0.087	(0.28)	0.060	(0.23)
DSEC16	public services, social security	0.052	(0.22)	0.106	(0.30)	0.072	(0.25)	0.271	(0.44)	0.109	(0.31)	0.220	(0.41)
DBERLIN	dummy for Berlin	0.035	(0.18)	0.078	(0.26)	0.045	(0.20)	0.092	(0.28)	0.041	(0.19)	0.043	(0.20)
N	number of observations	227243		57435		114324		44024		50338		7257	

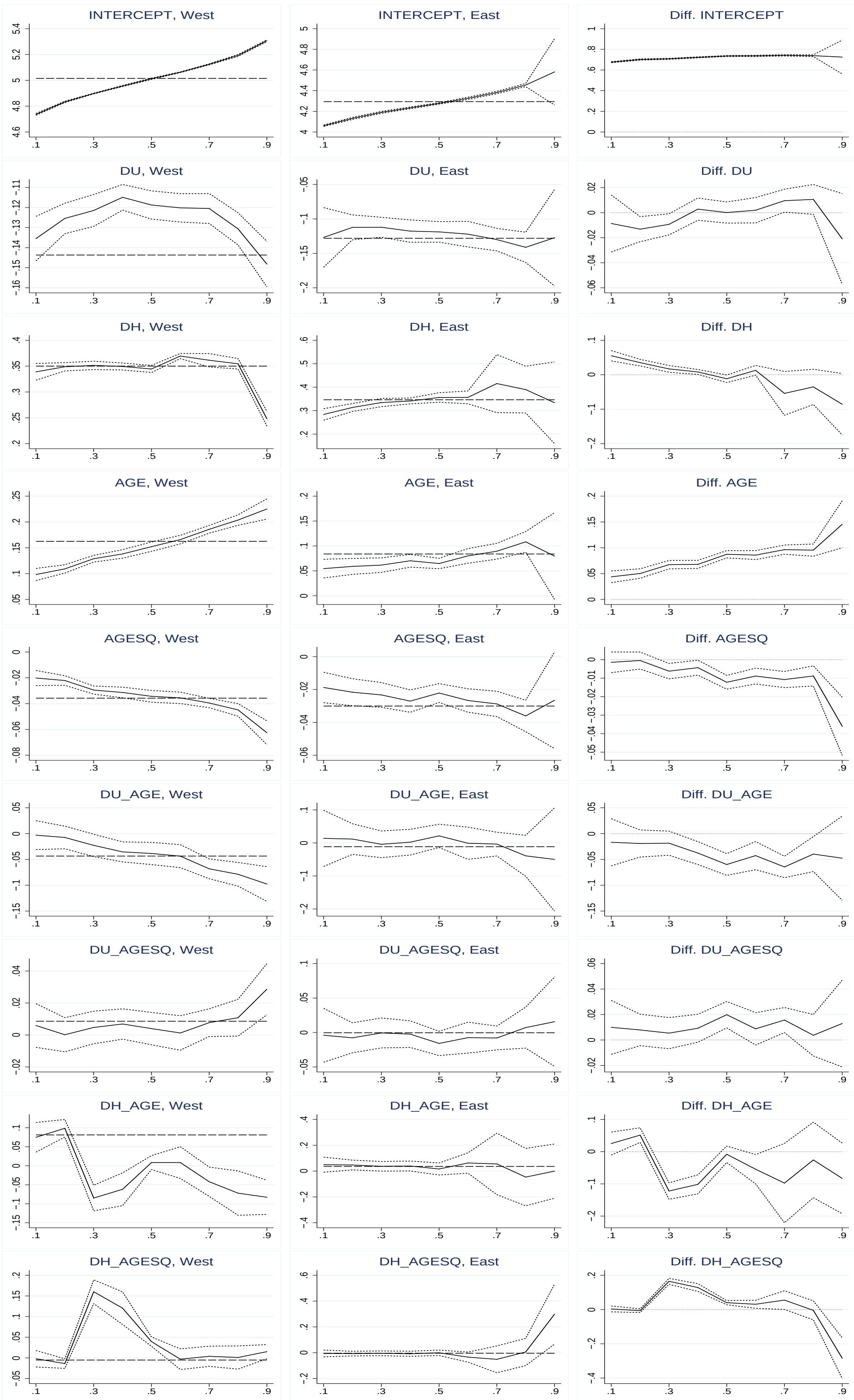
Mean values; standard deviations in parentheses. * indicates base categories. Observations weighted with length of resp. employment spells. Data source: IABS 1975–2001.

Table 2.3: Description and Summary Statistics of Covariates, 2001

Covariate	Description	Men f.-t., West		Men f.-t., East		Wom. f.-t., West		Wom. f.-t., East		Wom. p.-t., West		Wom. p.-t., East	
DL	dummy for low-skilled	0.087	(0.28)	0.019	(0.13)	0.103	(0.30)	0.019	(0.13)	0.118	(0.32)	0.035	(0.18)
DM*	dummy for medium-skilled	0.753	(0.43)	0.832	(0.37)	0.786	(0.41)	0.818	(0.38)	0.789	(0.40)	0.838	(0.36)
DH	dummy for high-skilled	0.158	(0.36)	0.148	(0.35)	0.110	(0.31)	0.162	(0.36)	0.091	(0.28)	0.125	(0.33)
AGE	age (25–55 years)	39.47	(7.89)	40.10	(7.83)	38.82	(8.34)	40.77	(7.68)	41.57	(7.25)	40.65	(7.51)
DSEC1	agriculture & mining	0.029	(0.17)	0.058	(0.23)	0.011	(0.10)	0.036	(0.18)	0.006	(0.08)	0.016	(0.12)
DSEC2	production of basic materials	0.082	(0.27)	0.054	(0.22)	0.033	(0.18)	0.021	(0.14)	0.015	(0.12)	0.003	(0.06)
DSEC3*	metal industry, machinery	0.134	(0.34)	0.091	(0.28)	0.037	(0.19)	0.021	(0.14)	0.017	(0.13)	0.005	(0.07)
DSEC4	vehicles & technical appliances	0.106	(0.30)	0.059	(0.23)	0.070	(0.25)	0.036	(0.18)	0.025	(0.15)	0.009	(0.09)
DSEC5	consumer goods	0.060	(0.23)	0.044	(0.20)	0.048	(0.21)	0.036	(0.18)	0.021	(0.14)	0.011	(0.10)
DSEC6	food, beverages, tobacco	0.023	(0.15)	0.020	(0.14)	0.031	(0.17)	0.035	(0.18)	0.015	(0.12)	0.017	(0.12)
DSEC7	main construction	0.051	(0.22)	0.119	(0.32)	0.008	(0.09)	0.015	(0.12)	0.005	(0.07)	0.010	(0.10)
DSEC8	subconstruction work	0.035	(0.18)	0.068	(0.25)	0.010	(0.10)	0.012	(0.10)	0.007	(0.08)	0.005	(0.07)
DSEC9	wholesale trade	0.069	(0.25)	0.048	(0.21)	0.059	(0.23)	0.033	(0.17)	0.031	(0.17)	0.018	(0.13)
DSEC10	retail trade	0.046	(0.21)	0.043	(0.20)	0.096	(0.29)	0.078	(0.26)	0.161	(0.36)	0.203	(0.40)
DSEC11	transport & communication	0.067	(0.25)	0.103	(0.30)	0.036	(0.18)	0.048	(0.21)	0.032	(0.17)	0.028	(0.16)
DSEC12	business-related services	0.144	(0.35)	0.113	(0.31)	0.186	(0.38)	0.141	(0.34)	0.141	(0.34)	0.103	(0.30)
DSEC13	household-oriented services	0.027	(0.16)	0.028	(0.16)	0.065	(0.24)	0.067	(0.25)	0.040	(0.19)	0.047	(0.21)
DSEC14	medical services	0.048	(0.21)	0.060	(0.23)	0.157	(0.36)	0.188	(0.39)	0.261	(0.43)	0.245	(0.43)
DSEC15	associations & organizations	0.028	(0.16)	0.034	(0.18)	0.082	(0.27)	0.095	(0.29)	0.108	(0.31)	0.129	(0.33)
DSEC16	public services, social security	0.044	(0.20)	0.049	(0.21)	0.064	(0.24)	0.132	(0.33)	0.107	(0.30)	0.143	(0.35)
DBERLIN	dummy for Berlin	0.027	(0.16)	0.070	(0.25)	0.039	(0.19)	0.077	(0.26)	0.032	(0.17)	0.068	(0.25)
N	number of observations	240974		46845		121960		32593		65083		12982	

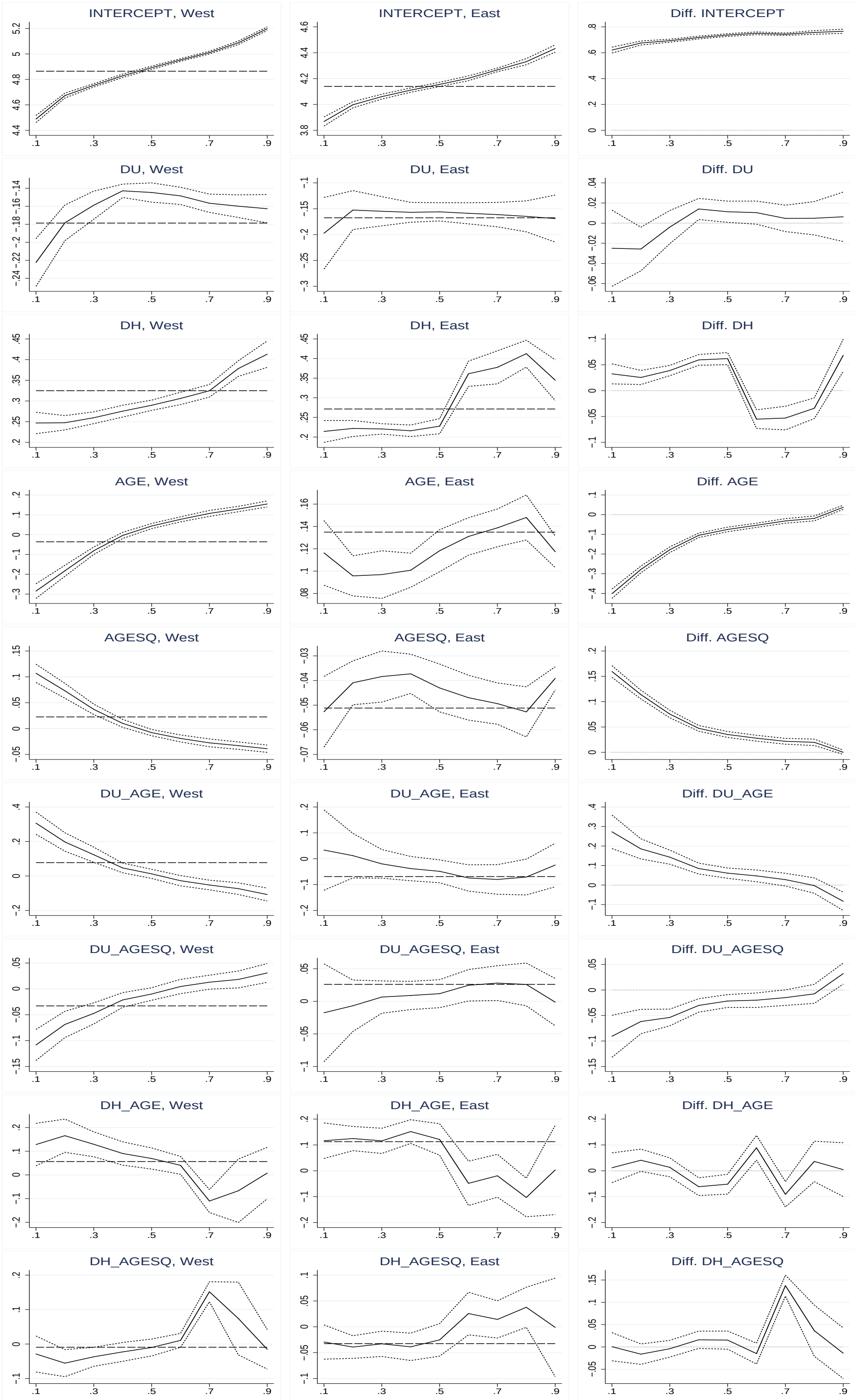
Mean values; standard deviations in parentheses. * indicates base categories. Observations weighted with length of resp. employment spells. Data source: IABS 1975–2001.

Figure 2.3: Regression Coefficients by Deciles in East-West Comparison: Men Working Full-Time, 1992



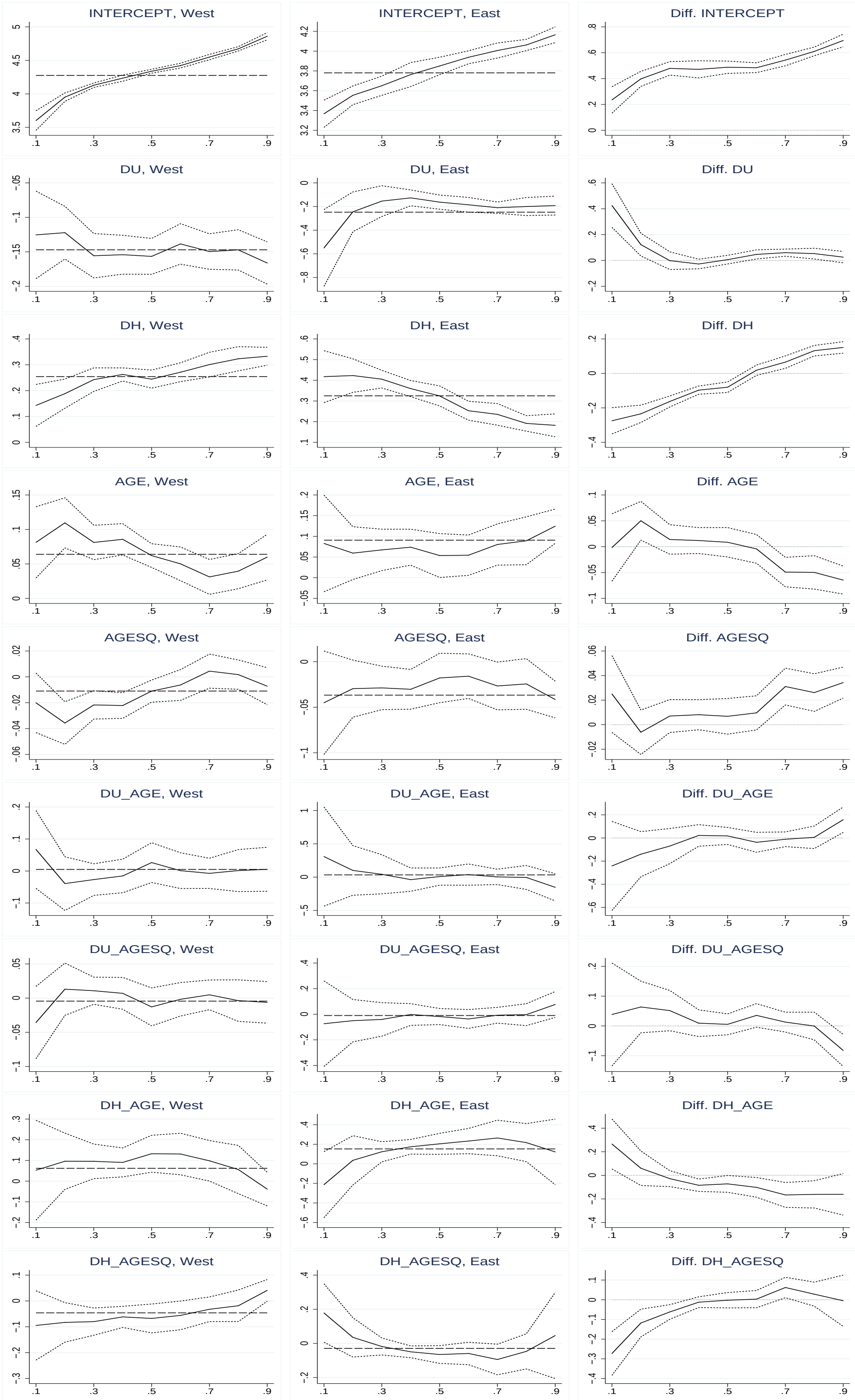
Coefficients from censored quantile regressions. Left panel: West Germany; middle panel: East Germany; right panel: West-East difference. Dashed lines: 95% confidence bands based on 50 bootstrap resamples. Long dashed lines: Tobit regression coefficients. Data source: IABS 1975–2001.

Figure 2.4: Regression Coefficients by Deciles in East-West Comparison: Women Working Full-Time, 1992



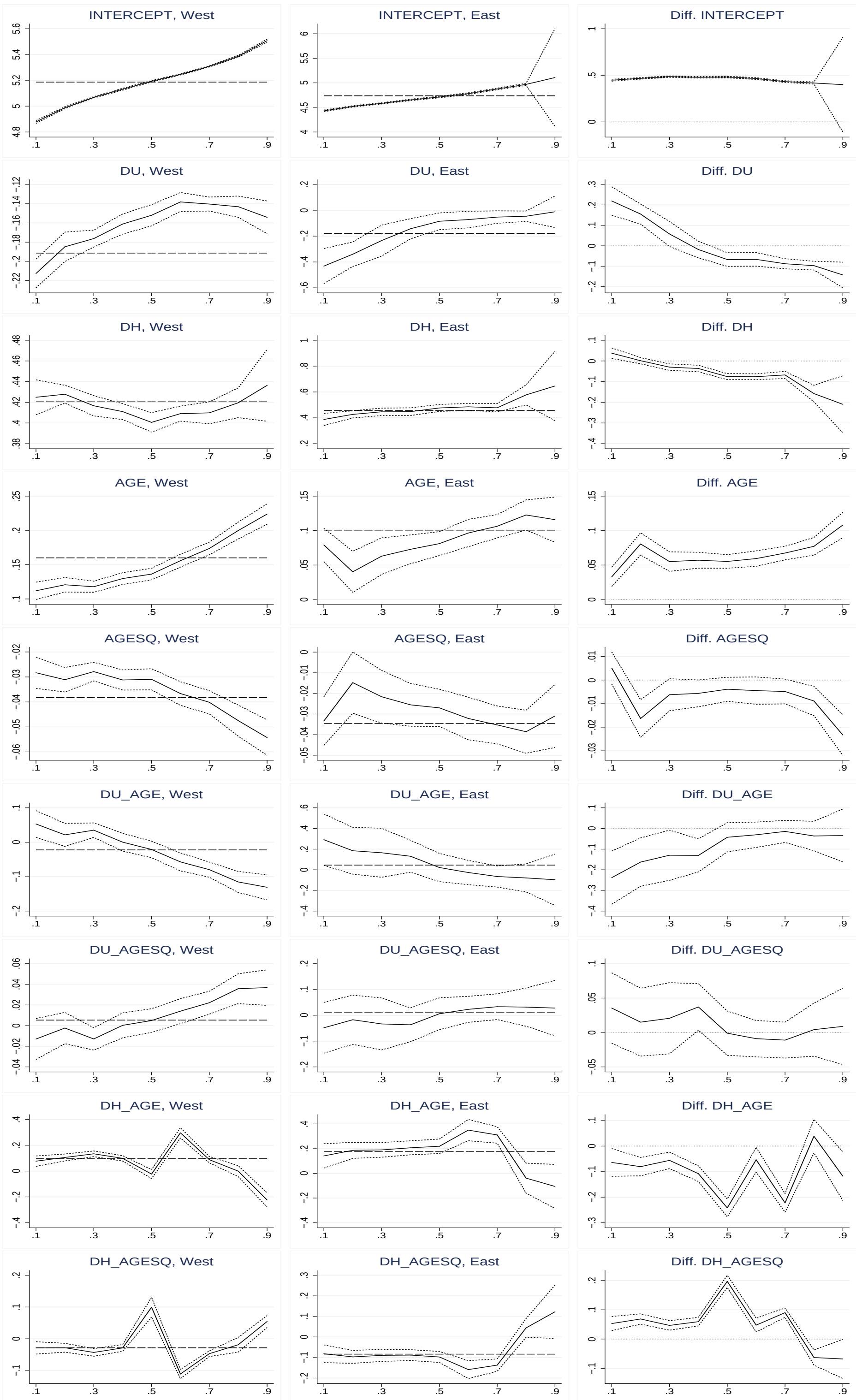
Coefficients from censored quantile regressions. Left panel: West Germany; middle panel: East Germany; right panel: West-East difference. Dashed lines: 95% confidence bands based on 50 bootstrap resamples. Long dashed lines: Tobit regression coefficients. Data source: IABS 1975–2001.

Figure 2.5: Regression Coefficients by Deciles in East-West Comparison: Women Working Part-Time, 1992



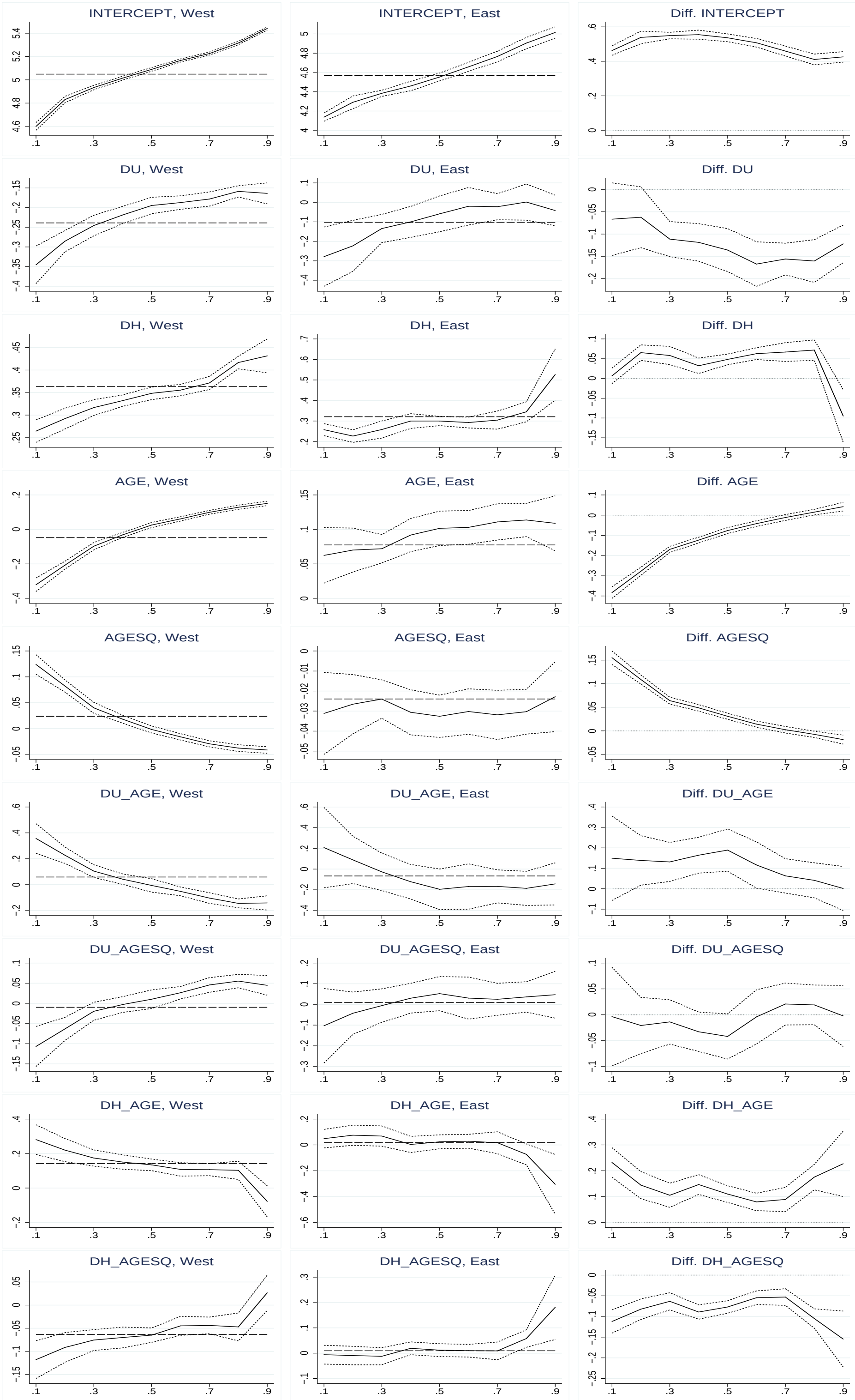
Coefficients from censored quantile regressions. Left panel: West Germany; middle panel: East Germany; right panel: West-East difference. Dashed lines: 95% confidence bands based on 50 bootstrap resamples. Long dashed lines: Tobit regression coefficients. Data source: IABS 1975–2001.

Figure 2.6: Regression Coefficients by Deciles in East-West Comparison: Men Working Full-Time, 2001



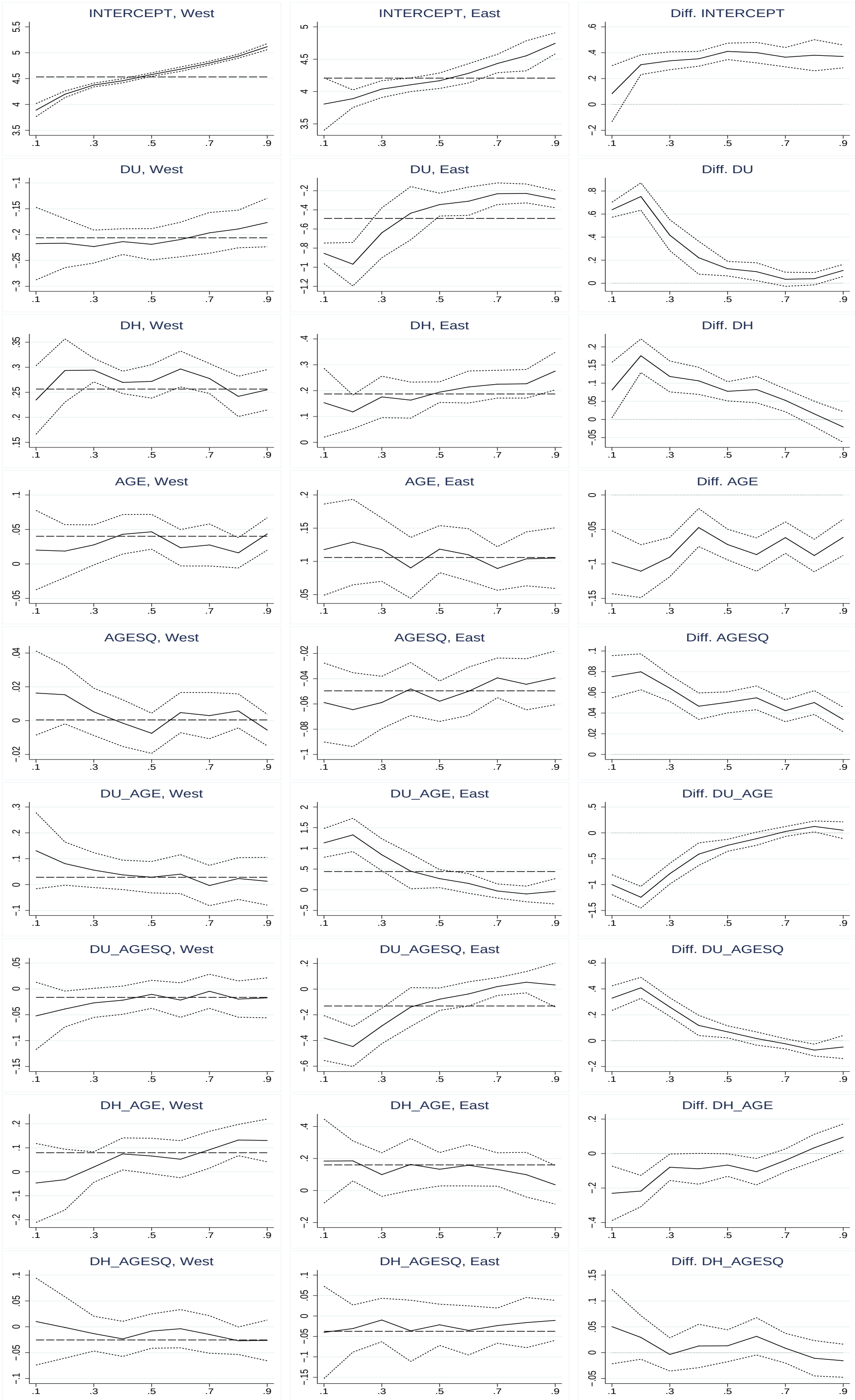
Coefficients from censored quantile regressions. Left panel: West Germany; middle panel: East Germany; right panel: West-East difference. Dashed lines: 95% confidence bands based on 50 bootstrap resamples. Long dashed lines: Tobit regression coefficients. Data source: IABS 1975–2001.

Figure 2.7: Regression Coefficients by Deciles in East-West Comparison: Women Working Full-Time, 2001



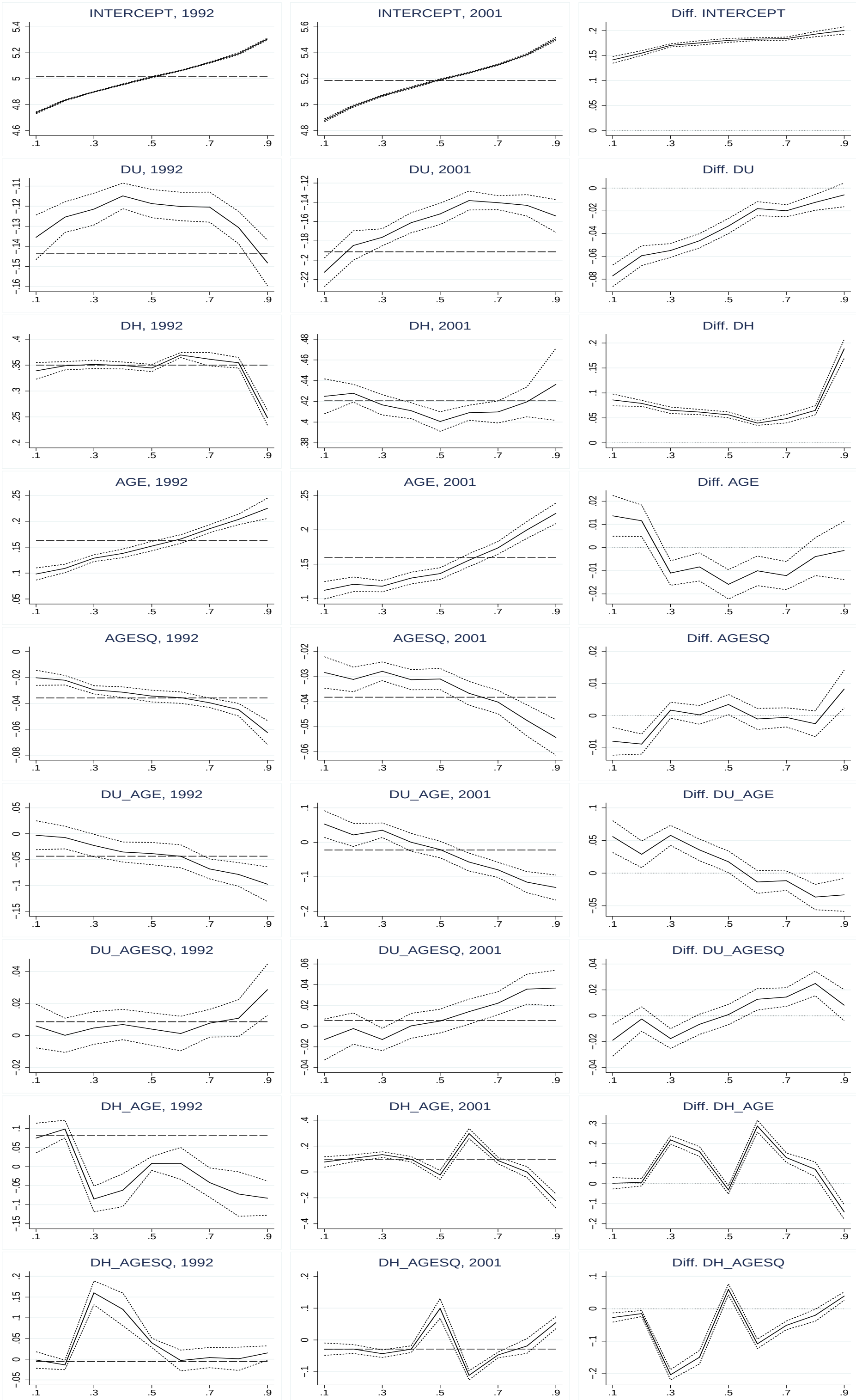
Coefficients from censored quantile regressions. Left panel: West Germany; middle panel: East Germany; right panel: West-East difference. Dashed lines: 95% confidence bands based on 50 bootstrap resamples. Long dashed lines: Tobit regression coefficients. Data source: IABS 1975–2001.

Figure 2.8: Regression Coefficients by Deciles in East-West Comparison: Women Working Part-Time, 2001



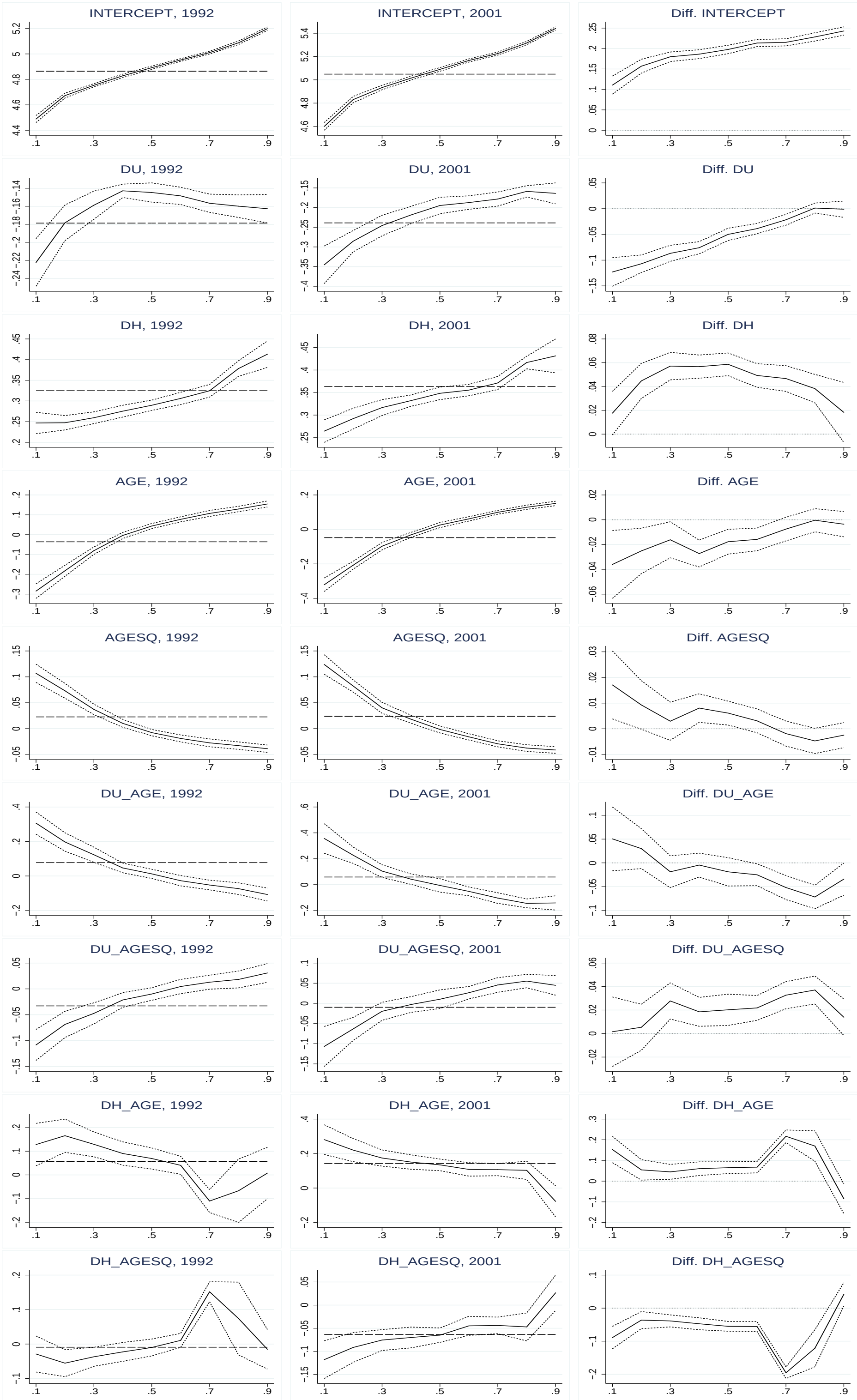
Coefficients from censored quantile regressions. Left panel: West Germany; middle panel: East Germany; right panel: West-East difference. Dashed lines: 95% confidence bands based on 50 bootstrap resamples. Long dashed lines: Tobit regression coefficients. Data source: IABS 1975–2001.

Figure 2.9: Regression Coefficients by Deciles in Comparison over Time: Men Working Full-Time, West



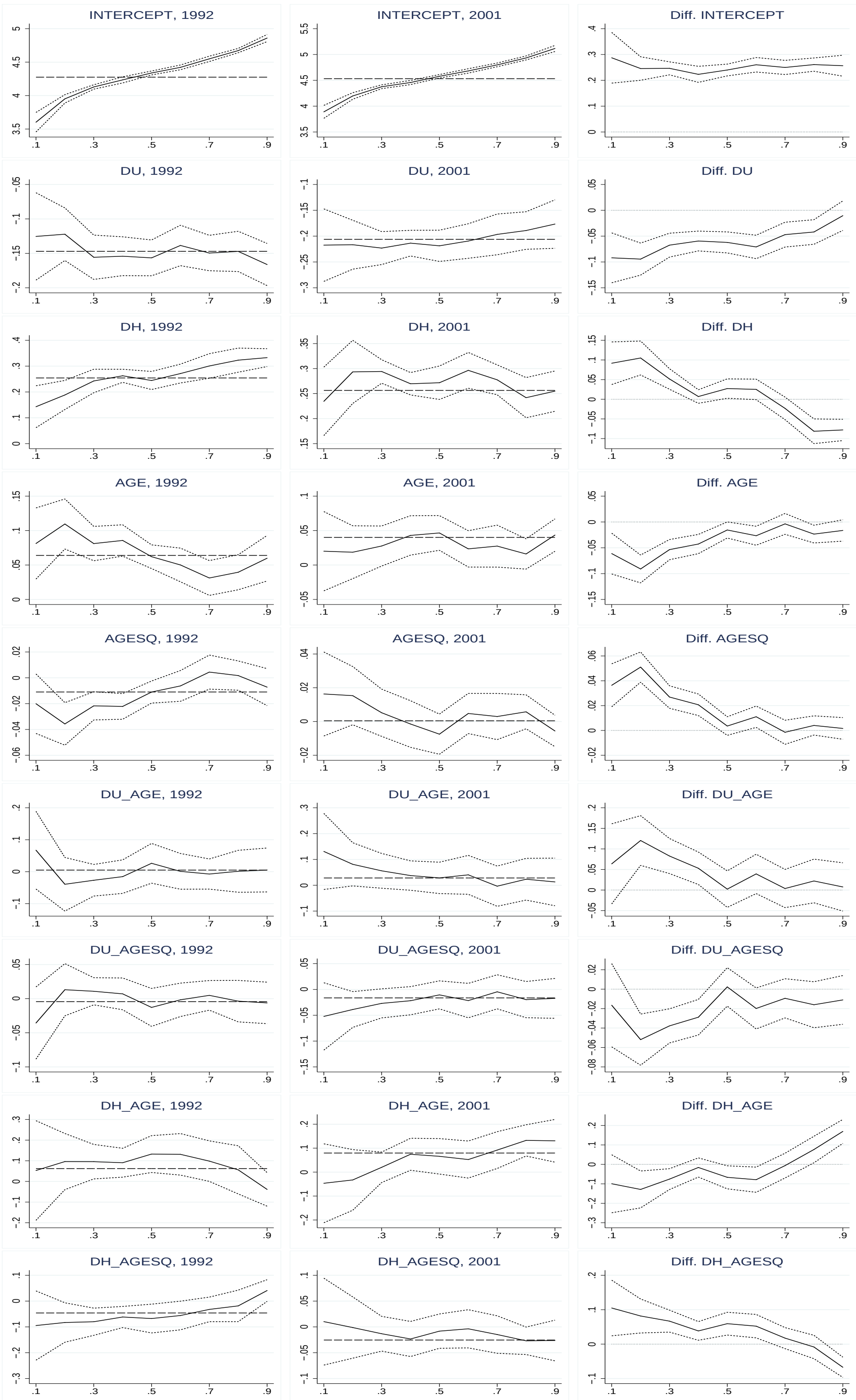
Coefficients from censored quantile regressions. Left panel: 1992; middle panel: 2001; right panel: difference 2001–1992. Dashed lines: 95% confidence bands based on 50 bootstrap resamples. Long dashed lines: Tobit regression coefficients. Data source: IABS 1975–2001.

Figure 2.10: Regression Coefficients by Deciles in Comparison over Time: Women Working Full-Time, West



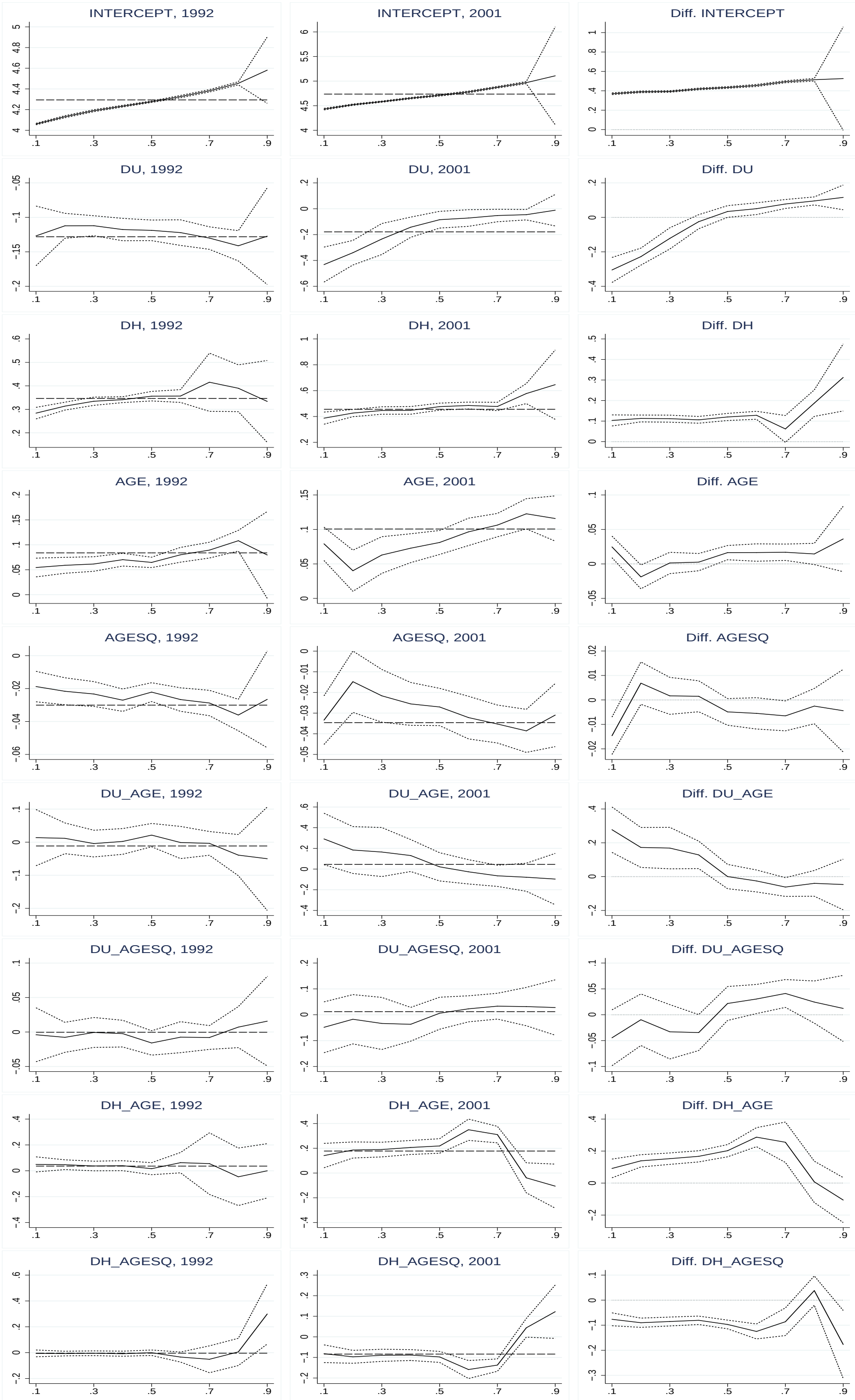
Coefficients from censored quantile regressions. Left panel: 1992; middle panel: 2001; right panel: difference 2001–1992. Dashed lines: 95% confidence bands based on 50 bootstrap resamples. Long dashed lines: Tobit regression coefficients. Data source: IABS 1975–2001.

Figure 2.11: Regression Coefficients by Deciles in Comparison over Time: Women Working Part-Time, West



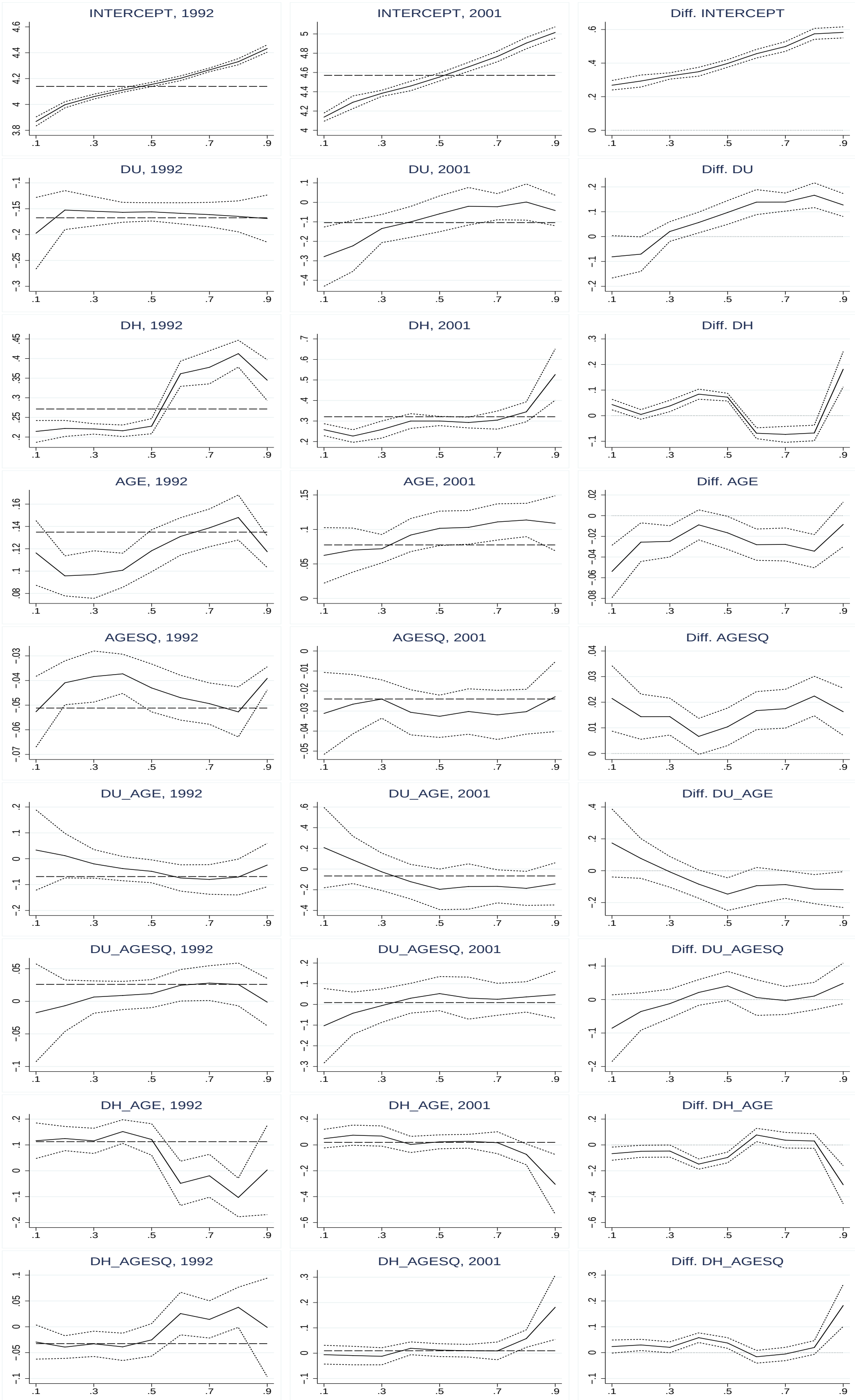
Coefficients from censored quantile regressions. Left panel: 1992; middle panel: 2001; right panel: difference 2001–1992. Dashed lines: 95% confidence bands based on 50 bootstrap resamples. Long dashed lines: Tobit regression coefficients. Data source: IABS 1975–2001.

Figure 2.12: Regression Coefficients by Deciles in Comparison over Time: Men Working Full-Time, East



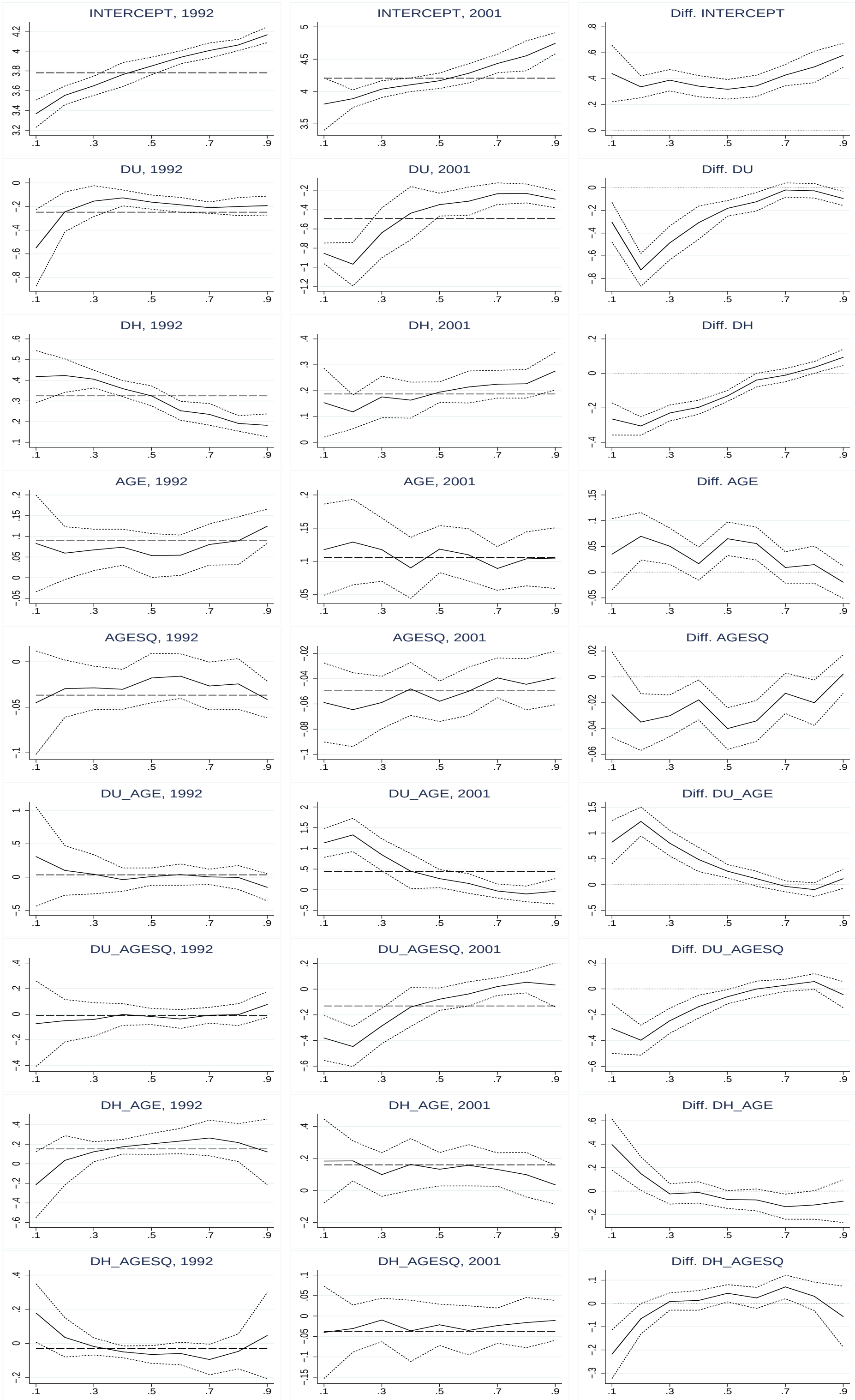
Coefficients from censored quantile regressions. Left panel: 1992; middle panel: 2001; right panel: difference 2001–1992. Dashed lines: 95% confidence bands based on 50 bootstrap resamples. Long dashed lines: Tobit regression coefficients. Data source: IABS 1975–2001.

Figure 2.13: Regression Coefficients by Deciles in Comparison over Time: Women Working Full-Time, East



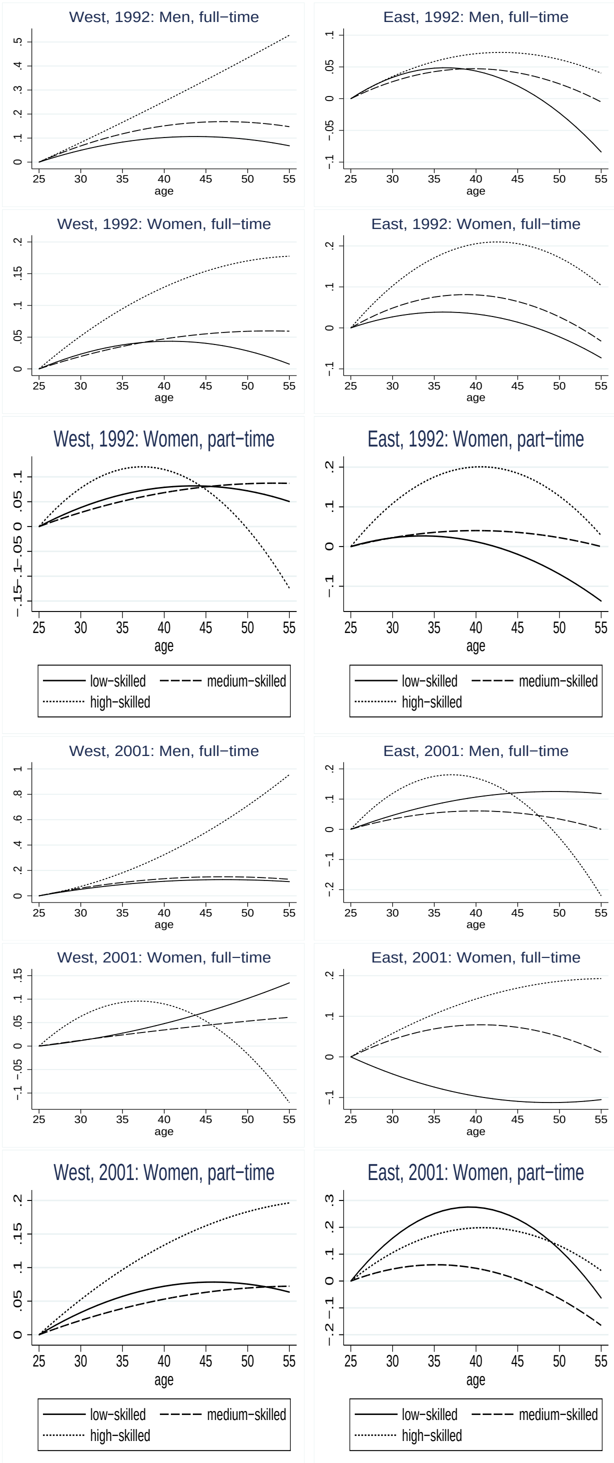
Coefficients from censored quantile regressions. Left panel: 1992; middle panel: 2001; right panel: difference 2001–1992. Dashed lines: 95% confidence bands based on 50 bootstrap resamples. Long dashed lines: Tobit regression coefficients. Data source: IABS 1975–2001.

Figure 2.14: Regression Coefficients by Deciles in Comparison over Time: Women Working Part-Time, East



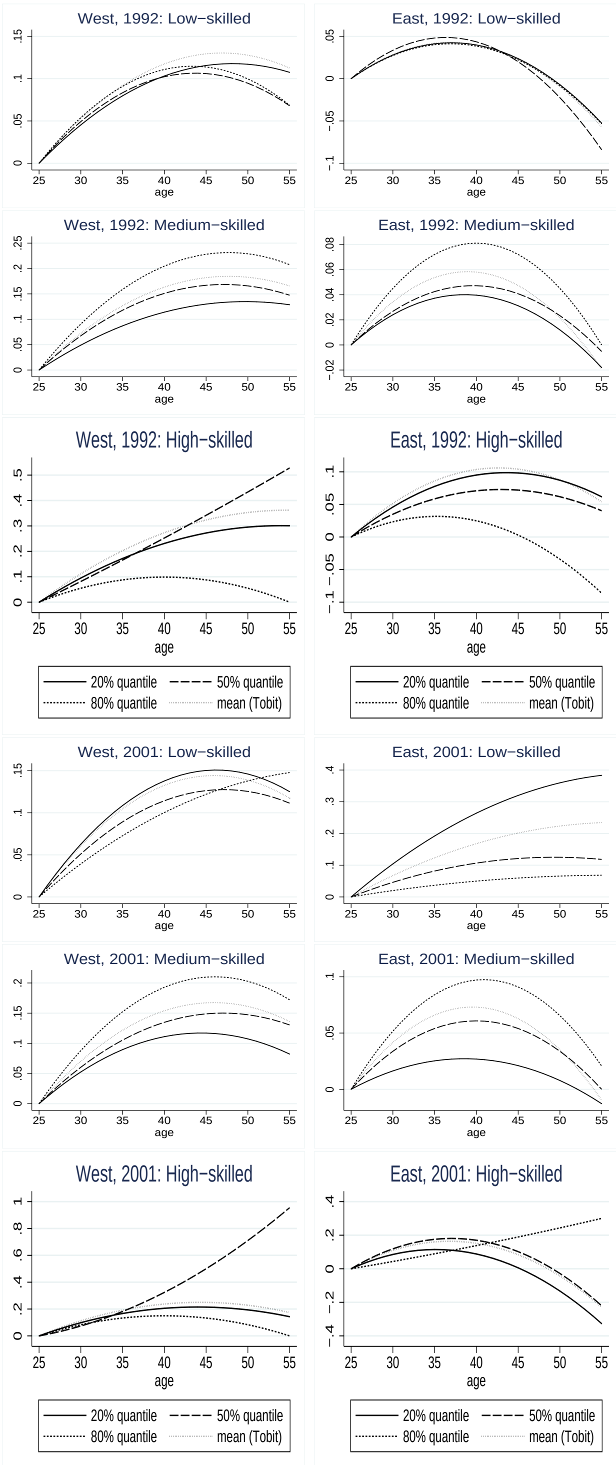
Coefficients from censored quantile regressions. Left panel: 1992; middle panel: 2001; right panel: difference 2001–1992. Dashed lines: 95% confidence bands based on 50 bootstrap resamples. Long dashed lines: Tobit regression coefficients. Data source: IABS 1975–2001.

Figure 2.15: Median Age-Earnings Profiles for Different Skill Groups



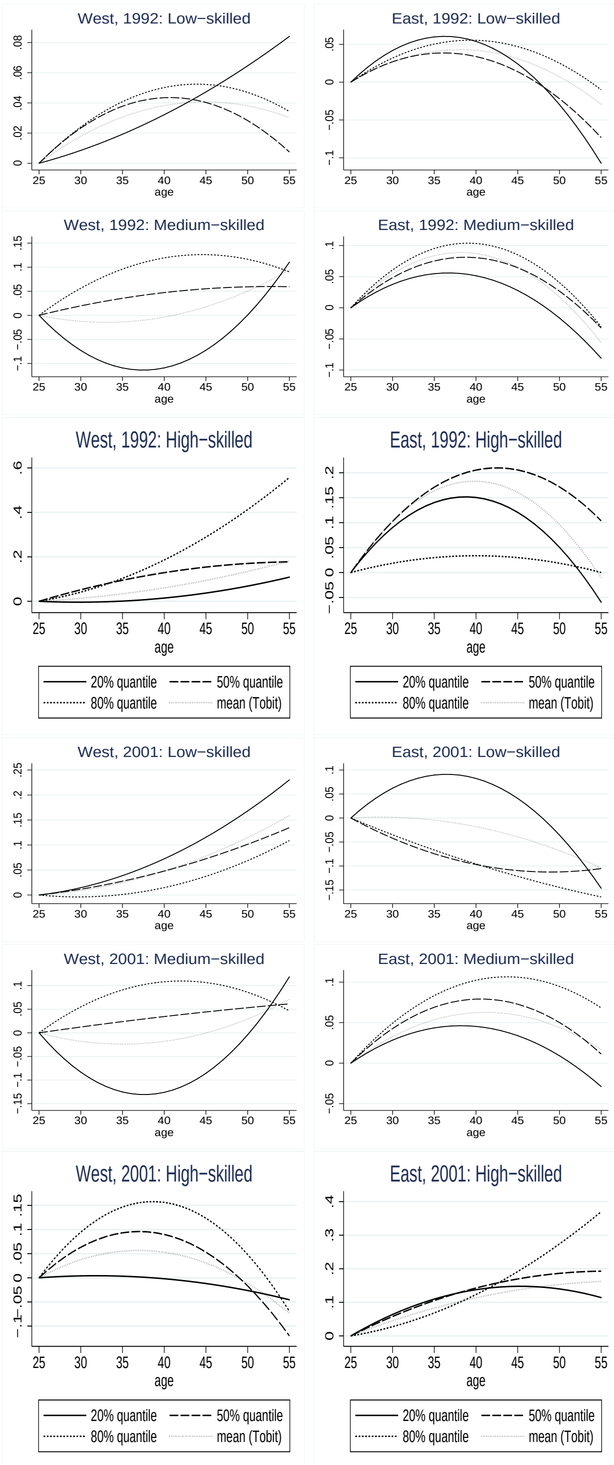
Results of censored median regressions. Solid lines: low-skilled; long dashed lines: medium-skilled; short dashed lines: high-skilled. Data source: IABS 1975–2001.

Figure 2.16: Age-Earnings Profiles across the Wage Distribution, by Skill Groups: Men Working Full-Time



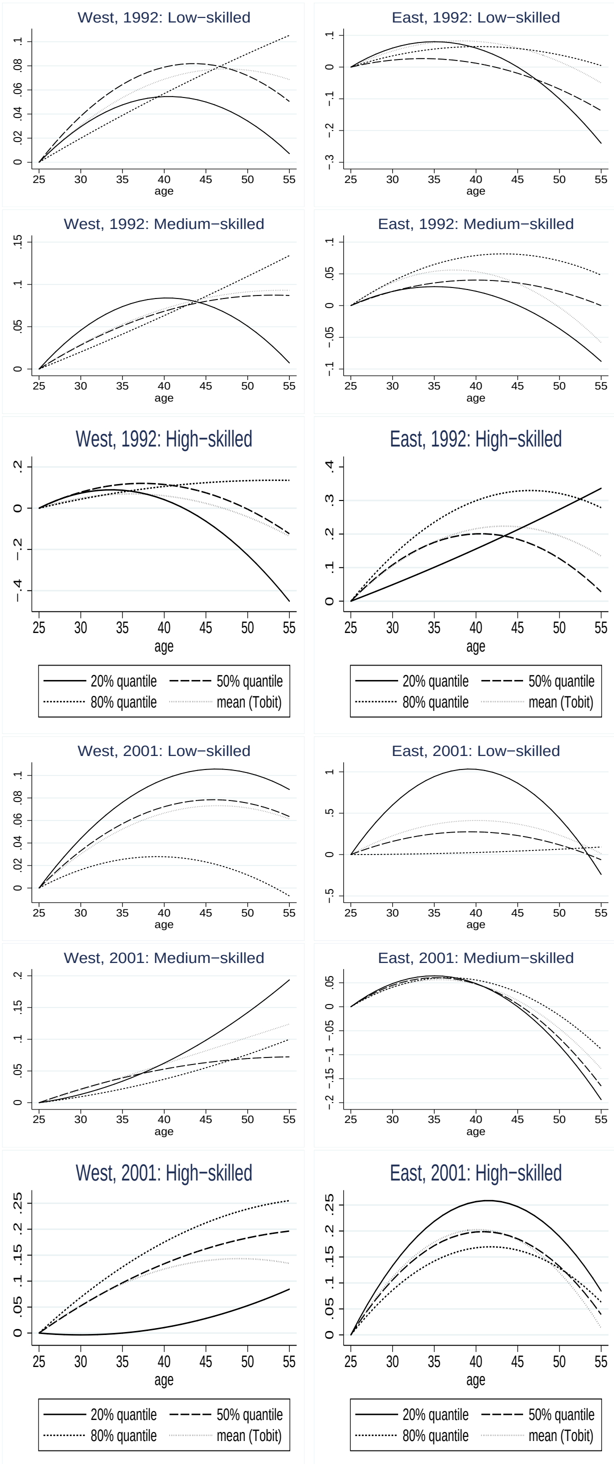
Results of censored quantile regressions. Solid lines: 20% quantile; long dashed lines: 50% quantile; short dashed lines: 80% quantile; dotted lines: mean (Tobit). Data source: IABS 1975–2001.

Figure 2.17: Age Earnings Profiles across the Wage Distribution, by Skill Groups: Women Working Full-Time



Results of censored quantile regressions. Solid lines: 20% quantile; long dashed lines: 50% quantile; short dashed lines: 80% quantile; dotted lines: mean (Tobit). Data source: IABS 1975–2001.

Figure 2.18: Age Earnings Profiles across the Wage Distribution, by Skill Groups: Women Working Part-Time



Results of censored quantile regressions. Solid lines: 20% quantile; long dashed lines: 50% quantile; short dashed lines: 80% quantile; dotted lines: mean (Tobit). Data source: IABS 1975–2001.

Table 2.4: Decomposition I: West-East Wage Differences Across the Distribution

	Men Working Full-Time									
1992	10th	20th	30th	40th	50th	60th	70th	80th	90th	Tobit
Observed gap	0.546	0.570	0.579	0.574	0.576	0.590	0.604	·	·	0.587
Predicted gap	0.555	0.569	0.573	0.578	0.579	0.595	0.607	0.614	·	0.578
Char. effect	-0.003	0.000	0.003	0.006	0.008	0.009	0.004	0.000	·	-0.018
Coeff. effect	0.557	0.570	0.570	0.572	0.571	0.585	0.603	0.614	·	0.596
2001										
Observed gap	0.381	0.413	0.405	0.417	0.395	0.405	0.405	0.402	·	0.382
Predicted gap	0.370	0.398	0.405	0.413	0.412	0.412	0.412	0.398	·	0.386
Char. effect	0.010	0.007	0.010	0.012	0.021	0.024	0.024	0.022	·	0.008
Coeff. effect	0.360	0.390	0.396	0.401	0.391	0.388	0.388	0.376	·	0.378
	Women Working Full-Time									
1992	10th	20th	30th	40th	50th	60th	70th	80th	90th	Tobit
Observed gap	0.219	0.317	0.360	0.369	0.368	0.378	0.381	0.399	·	0.342
Predicted gap	0.238	0.327	0.358	0.369	0.373	0.376	0.378	0.388	·	0.345
Char. effect	-0.086	-0.086	-0.079	-0.076	-0.066	-0.056	-0.056	-0.059	·	-0.073
Coeff. effect	0.324	0.414	0.437	0.445	0.439	0.433	0.434	0.447	·	0.417
2001										
Observed gap	0.150	0.254	0.276	0.234	0.177	0.146	0.143	0.161	0.171	0.179
Predicted gap	0.171	0.249	0.263	0.227	0.192	0.164	0.157	0.158	0.156	0.179
Char. effect	-0.015	-0.030	-0.034	-0.039	-0.041	-0.036	-0.042	-0.043	-0.055	-0.047
Coeff. effect	0.186	0.279	0.297	0.266	0.233	0.200	0.199	0.201	0.211	0.226
	Women Working Part-Time									
1992	10th	20th	30th	40th	50th	60th	70th	80th	90th	Tobit
Observed gap	0.061	0.187	0.201	0.215	0.212	0.210	0.187	0.151	0.125	0.168
Predicted gap	0.096	0.197	0.213	0.215	0.208	0.199	0.181	0.161	0.147	0.186
Char. effect	-0.049	-0.038	-0.028	-0.016	-0.015	-0.013	-0.017	-0.018	-0.033	-0.016
Coeff. effect	0.145	0.235	0.241	0.231	0.223	0.212	0.198	0.179	0.180	0.202
2001										
Observed gap	-0.043	0.000	0.000	0.000	0.000	0.000	-0.018	0.000	0.013	-0.013
Predicted gap	-0.063	-0.020	0.000	-0.010	-0.004	-0.009	-0.017	0.003	0.008	0.003
Char. effect	-0.023	-0.021	-0.028	-0.015	-0.009	-0.013	-0.017	-0.017	-0.015	-0.004
Coeff. effect	-0.040	0.001	0.028	0.005	0.006	0.004	0.000	0.020	0.024	0.007

Nominal differences, evaluated at various percentiles. Tobit "observed" gaps estimated by Tobit regressions on a constant. · indicates censored deciles. Data source: IABS 1975–2001.

Table 2.5: Decomposition II: Changes of the Wage Structure, 1992–2001

	Men Working Full-Time									
West Germany	10th	20th	30th	40th	50th	60th	70th	80th	90th	Tobit
Observed change	-0.075	-0.038	-0.023	-0.010	-0.006	0.016	0.030	0.051	·	-0.004
Predicted change	-0.075	-0.039	-0.027	-0.015	0.001	0.012	0.030	0.051	·	0.001
Char. effect	0.010	0.013	0.013	0.023	0.032	0.037	0.045	0.052	·	0.036
Coeff. effect	-0.084	-0.052	-0.040	-0.038	-0.030	-0.025	-0.015	-0.001	·	-0.036
East Germany										
Observed change	0.001	0.031	0.062	0.058	0.087	0.112	0.140	·	·	0.112
Predicted change	0.022	0.044	0.052	0.062	0.080	0.106	0.136	0.179	·	0.104
Char. effect	-0.012	-0.013	-0.015	-0.013	-0.007	-0.003	0.006	0.013	·	-0.011
Coeff. effect	0.034	0.057	0.067	0.076	0.087	0.109	0.130	0.166	·	0.115
	Women Working Full-Time									
West Germany	10th	20th	30th	40th	50th	60th	70th	80th	90th	Tobit
Observed change	-0.024	0.000	-0.001	0.035	0.045	0.053	0.068	0.075	0.106	0.040
Predicted change	-0.017	-0.009	0.008	0.027	0.043	0.053	0.065	0.082	0.103	0.032
Char. effect	0.006	0.004	0.012	0.021	0.027	0.034	0.036	0.047	0.054	0.017
Coeff. effect	-0.023	-0.013	-0.004	0.007	0.016	0.019	0.029	0.035	0.049	0.015
East Germany										
Observed change	-0.043	-0.025	-0.006	0.081	0.147	0.197	0.217	0.225	·	0.113
Predicted change	-0.038	-0.019	0.015	0.081	0.137	0.177	0.197	0.224	·	0.110
Char. effect	-0.044	-0.044	-0.067	-0.064	-0.038	-0.012	0.003	0.013	·	-0.026
Coeff. effect	0.006	0.025	0.082	0.144	0.175	0.190	0.195	0.212	·	0.136
	Women Working Part-Time									
West Germany	10th	20th	30th	40th	50th	60th	70th	80th	90th	Tobit
Observed change	0.112	0.054	0.051	0.065	0.068	0.091	0.104	0.114	0.111	0.079
Predicted change	0.076	0.037	0.039	0.054	0.072	0.087	0.097	0.108	0.106	0.077
Char. effect	0.011	0.025	0.021	0.024	0.026	0.031	0.026	0.034	0.030	0.024
Coeff. effect	0.066	0.012	0.017	0.030	0.046	0.056	0.071	0.074	0.076	0.053
East Germany										
Observed change	0.127	0.152	0.163	0.191	0.191	0.212	0.220	0.176	0.134	0.173
Predicted change	0.147	0.166	0.163	0.190	0.195	0.206	0.207	0.178	0.156	0.172
Char. effect	0.020	0.020	0.000	0.017	0.005	0.007	0.017	0.013	0.008	0.008
Coeff. effect	0.127	0.146	0.163	0.173	0.190	0.198	0.190	0.165	0.149	0.164

Real differences, evaluated at various percentiles. Tobit "observed" gaps estimated by Tobit regressions on a constant. · indicates censored deciles. Data source: IABS 1975–2001.

3 Skill Wage Premia, Employment, and Cohort Effects: Are Workers in Germany All of the Same Type?

3.1 Introduction

Numerous empirical studies record descriptive micro-data evidence on the evolution of wages and employment measures; see the survey article of Katz and Autor (1999). To capture the heterogeneity of labor, authors usually undertake a grouping into different classes based on observed covariates like age and sex of employees or on the basis of job characteristics. Available studies typically report considerable wage dispersion both between and within adequately defined classes. Variation over time yet generates another important dimension of heterogeneity.

Particular attention is given to skill wage premia and the evolution of skill-specific employment. As a stylized fact, the unemployment rate is the higher the lower the (formal) qualificational level of the employees. In West Germany, for example, the respective rates for employees without a vocational degree, for those with, and for those with a university degree were 21.7%, 7.3%, and 3.5% in the year 2004 (Reinberg and Hummel, 2005).

Rigidity of the wage structure is often referred to as a major cause for the different degrees of incidence of unemployment; compare, e. g., Fitzenberger and Franz (2001). As elaborated in the discussion about employment impacts of skill-biased technical change (SBTC; see Katz and Autor, 1999; Acemoglu, 2002), relative demand for low-skilled labor decreases faster over time than does relative supply. In line with neoclassical demand theory (Hamermesh, 1993), market clearing would in this case require an increase of qualificational wage differentials.

Despite the popularity and plausibility of this hypothesis an empirical operationalization of the interrelation between wage structures and employment that goes beyond mere descriptive evidence proves difficult due to the heterogeneity of labor, among other things. Conventional empirical analyses of qualificational labor demand typically take into account only a small number of homogeneous skill groups—mostly not more than three; cf. the surveys in Hamermesh (1993) and Katz and Autor (1999) and for Germany, e. g., the studies of Fitzenberger (1999), Steiner and Wagner (1998b), or Falk and Koebel (1999, 2002). These approaches are often justified in light of the fact that satisfactory solutions to the problem of aggregation do not exist.[1] Also, standard approaches based on cost-minimizing behavior like flexible translog systems, which allow for a larger number of factors, quickly become impracticable.

Based on US data, Katz and Murphy (1992) analyze wage differentials between high school and college graduates in the context of supply and demand effects. A CES model proves compatible with the developments of wage premia and employment over time. These are consistent with the labor market entry of young and the exit of older birth cohorts on the one hand and an increase in average educational attainment on the other. The literature interprets these trends as a race between changes in the skill structure of labor supply and that of labor demand; cf., for example, Johnson (1997), Topel (1997), and Machin (2002). However, in addition to the variation of skills between different cohorts, human capital endowments also change with age. Whereas increasing labor market experience and job tenure augment human capital stocks with age, skill-biased and accelerating structural change might invalidate individual endowments of older workers. Freeman (1979) and Welch (1979) thus account for imperfect substitutability between workers of different age by means of CES technologies for workers from discrete age or "career phase" groups.

Card and Lemieux's (2001) – henceforth CL – investigation using US, UK, and Canadian data reconciles the analysis of Katz and Murphy (1992) with those of Freeman (1979) and Welch (1979). In a set-up which uses the nested CES model developed by Sato (1967) the simultaneous inclusion of skill and age as dimensions of heterogeneity not only enables the separation of age, time, and cohort effects, but also facilitates the estimation of a specification with a relatively large number of different input factors. The estimation strategy undertaken in particular yields elasticities of substitution both between high school and college graduates and between workers belonging to different age classes.

The starting point of the study by CL is the observation that the college-high school gap in wages has increased strongly for younger US men whereas the gap for older men has remained nearly constant. The driving force for these observed cohort-specific changes is the slowdown in the growth of college-educated labor which did not keep up with the steady skill bias in labor demand; see also Autor, Katz, and Kearney (2005b) for a recent reassessment.

Wage trends in West Germany differed from what happened in countries like the US, UK, or Canada over the last decades. In particular, wage dispersion for male workers did not increase

[1] For discussions of the problem of aggregation in the context of labor demand estimations see, e. g., Koebel (2005) and Katz and Autor (1999).

to the same extent (Prasad, 2004). In fact, skill wage differentials decreased between workers with and those without a vocational training degree; see Fitzenberger (1999) and Fitzenberger and Wunderlich (2002). Regarding the differential between workers with a vocational degree and workers with university-type education there is conflicting evidence; see Steiner and Wagner (1998a), Möller (1999), and Fitzenberger (1999). Little evidence exists regarding age-related wage differentials. Fitzenberger (1999) and Fitzenberger and Wunderlich (2002) find that cross-sectional age profiles became somewhat steeper for male workers without and for those with a vocational training degree. According to the SBTC hypothesis, skill upgrading in employment should thus have occurred at a faster rate in Germany compared to countries like the US, the UK, or Canada and, in the spirit of CL, cohort effects are likely to be of importance in West Germany as well. There is recent concern that the necessary skill upgrading of the labor force in Germany is too slow to combat the high unemployment of the low-skilled; see the stylized facts reported in OECD (2004).

This paper broadens the scope of the nested CES framework and provides estimates based on the IAB employment sample (IABS) for Germany. While consistently reconciling the developments of relative wages and employment, our treatment extends upon the existing literature in several directions. First, we let three skill groups account for heterogeneity within the qualification dimension. This extension is necessary in light of the coexistence of vocational training and university education in Germany. Second, we treat the identification of cohort effects more rigorously. Tests for the existence of cohort effects and their separability from age and time effects (as suggested by MaCurdy and Mroz, 1995) are applied to check the validity of the specification. Third, rather than merely running regressions for skill wage differentials, we estimate a full system of skill *and* age premia implied by the nested CES model. Fourth, we take a closer look at the notions of observed employment and let instrumental variable techniques account for the endogeneity of both wages and employment. Finally, we draw on the estimated substitution parameters in order to conduct two simulation experiments: We calculate the magnitude of wage changes in the three skill groups that would have been necessary to halve skill-specific unemployment rates in 1997 (the latest period available). While allowing for relative changes between skill groups, this would have left the wage structure within skill groups unaffected. Due to the particularly high unemployment rate among low-skilled employees in Germany, the design imposes a disproportionately prominent increase in employment of this group, and thus is of high policy relevance. Alternatively, one might be interested in changes of the wage structure within skill groups, holding the structure across the respective groups constant. Here, the model set-up may provide an answer to the question how wages for employees of different age would have had to change to reduce all age-specific unemployment rates by one half.

The remainder of the paper is organized as follows: Section 3.2 outlines the trends in skill wage premia and skill-specific employment in the IABS between 1975 and 1997. Following an investigation into the nature of cohort effects in section 3.3, section 3.4 discusses different facets of the nested CES model which allow for the reconciliation of the stylized empirical facts, and section 3.5 estimates elasticities of substitution across and within skill groups. Based on the resulting parameters, the simulation experiments are presented in section 3.6. Section 3.7 concludes.

3.2 Descriptive Evidence

A number of recent empirical studies provides descriptive evidence for skill wage differentials in the German labor market. Among the analyses—comprising, e. g., Christensen (2003), Christensen and Schimmelpfennig (1998), Fitzenberger (1999), Fitzenberger and Wunderlich (2002), Möller (1999), Prasad (2004), Riphahn (2003), Steiner and Mohr (2000), and Steiner and Wagner (1998a)—there is some consensus that, by and large, the earnings distribution across skill groups stayed relatively stable during the 1980's and 1990's.

A closer look calls for detailed investigations which take into consideration further aspects of heterogeneity. In the tradition of Mincer (1974) work experience is an important additional determinant of individual earnings, and the effects of age—often used as a proxy for experience—are of interest themselves. Prasad (2004) and Riphahn (2003), for example, estimate year-specific Mincer equations and depict the evolution of returns to potential experience. Studies explicitly accounting for the age dimension of wage distributions examine single cross-sectional age profiles, like Fitzenberger and Reize (2003), or focus specifically on cohort analyses, as Boockmann and Steiner (2000) or Fitzenberger, Hujer, MaCurdy, and Schnabel (2001), for example. Beißinger and Möller (1998) account for the age dimension in the distribution of (un)employment for discrete years between 1980 and 1990.

Our study scrutinizes both wages and employment across the two dimensions skill and age for the time span 1975–1997. It is based on the IAB employment subsample (IABS), a 1% random draw of German employment spells subject to social insurance contributions. The IABS covers about 80% of all employed persons, and it provides detailed information on daily wages for blue and white collar workers as well as the exact timing of employment spells. We classify employees into three skill groups and consider six age classes. An extensive description of the data and classifications used is given in appendix 3.A.

3.2.1 Stylized Facts I: The Evolution of Wage Differentials

Age-specific skill wage premia or skill wage differentials $r_{sm,a,t}$ among workers of age a at time t are defined as the difference in mean *log* wage of high-skilled ($s = h$, employees with a university degree) or low-skilled workers ($s = l$, employees with neither university nor vocational training degree) and that of medium-skilled workers ($s = m$, employees with a vocational training degree). Using dummy variables $d_{s,a,t}$ for the different skill groups and possibly controlling for further influences,[2] they can be derived from regressions

$$\ln(w_{a,t}) = \text{constant}_{a,t} + r_{l,a,t} \cdot d_{l,a,t} + r_{h,a,t} \cdot d_{h,a,t} + \text{controls}_{a,t} + \epsilon_{a,t} \tag{3.1}$$

in the respective age-time cells. Due to the social security taxation threshold, wage data in the IABS are censored from above. Thus (3.1) is estimated by means of Tobit regressions. Observations are weighted by the length of the respective employment spells. Results are provided in tables 3.4 and 3.5 in appendix 3.B.

Figure 3.2 illustrates the evolution of age-specific wage differentials for males over time. Skill premia generally grow with age. Taking age as a proxy for experience, this corresponds to classical human capital theory (Becker, 1964). The estimated premia have evolved quite differently, though.

The education premium for high-skilled employees compared to the medium-skilled stayed roughly constant for the oldest age class until 1987 and declined by about 9 percentage points (ppoints) thereafter. The relative position of 30- to 35-year-old high-skilled, on the other hand, deteriorated by about 9 ppoints during the late 1970's, partly rose again in the first half of the 80's, and stayed constant from 1986 on.

The differential between older medium- and low-skilled workers exhibited a decline of about 5 ppoints during the eighties and recovered to an overall decline of about 2 ppoints during the nineties. In the youngest age class this wage premium exhibited an even higher volatility: Between 1975 and 1986, low-skilled workers on average gained around 6 ppoints compared to the medium-skilled. Later on, the differential increased again and even exceeded the 1975-level in 1997.

To infer the evolution of age profiles across time, we plot the wage differentials for three years against the age dimension in figure 3.3. Average wage differentials between high- and medium-skilled generally increase rather steeply with age: The premium grows by up to 29 ppoints. However, the shape of the profiles changes over time.

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[2] Cf. appendix 3.A for details on implemented specifications.

In 1975 the profile is considerably curved, showing especially a pronounced rise for young individuals. In transition to the mid-1980's, the curvature declines whilst the profile still shows a similarly high increase over the entire age span: In particular for middle-aged workers the premium for higher education declines compared to 1975. Starting in the second half of the eighties, one observes a twist of the profile. Whereas the increase in the premium for higher education for workers up to their mid-thirties is much the same in 1997 as in 1986, the profile has become flatter for older employees: The relative position of older high-skilled workers has deteriorated in comparison to the situation in 1986.

In comparison, the profile of the wage differential between low- and medium-skilled workers is typically much flatter, especially for older workers. The differential declines strongly for younger workers between 1975 and 1986 and it increases again strongly between 1986 and 1997. But even though the maximum decline—roughly 8 ppoints in 1986—is found to be small relative to the one experienced by the high-to-medium-skilled differential, the picture of the developments over time is still striking. In 1975 the average education premium moderately rose with age, showing increments declining with age. Up to 1986, the profile shifted downward by about 2–6 ppoints, becoming steeper for younger age classes. In 1997, however, the profile shows a twisted shape: Whilst the differential for older workers partly recovered in a parallel kind of manner, the youngest workers now face a premium increased by 6 ppoints that renders the entire profile nearly flat.

Taking the above results together, we assert a first stylized fact:

> Between the mid-1970's and the mid-1990's, age profiles of skill wage premia have not moved in parallel fashion over time. Skill wage premia declined over time (especially between the 1970's and 1980's) in a non-uniform fashion across age groups.

Thus, the developments are not likely to be the result of pure age and time effects alone. Cohort effects, i. e., systematic differences across birth cohorts, supposedly play an additional important role. Our subsequent theoretical and empirical investigation into the development of skill wage premia hence takes account of age, time, and cohort effects.

3.2.2 Stylized Facts II: Trends in Relative Employment

Based on the individual spell data, we use a weighted headcount as our measure of employment: In each age-time cell, the number of skill-specific employed is summed up, weighted by the duration of the respective employment spells.

Figure 3.4 presents relative employment trends for the different age classes. These are the employment counts of the high- and the low-skilled relative to the employment in the mediumskill group, respectively. The measures show the skill upgrading over the past decades: For most of the sample period, both the ratio of high-skilled to medium-skilled and that of medium-skilled to low-skilled employment were the higher the younger the respective age class. Furthermore, the skill-intensity of employment has increased over time. Starting from a situation of uniform skill upgrading in all age classes, however, the increase of relative employment of the skilled slows down considerably or even comes to an end at some point in time. Beginning in the mid 1980's, this break occurs first for the youngest age group. It then works through the older classes during the following years until it affects the oldest employees in the second half of the 1990's.[3]

We record a second stylized fact:

> There is a break in the inter-cohort trend of relative employment such that younger birth cohorts do not follow the older ones towards further skill upgrading.

The empirical evidence thus suggests the existence of cohort effects in the employment dimension, too.

3.3 Testing for Cohort Effects

To distinguish age, cohort, and time effects in wage premia $r_{a,t}$, CL undertake a decomposition of wage premia by the following regression:

$$r_{a,t} = b_a + c_{t-a} + d_t + \epsilon_{a,t} \tag{3.2}$$

where b_a, c_{t-a}, and d_t denote age, cohort, and time dummies, respectively. However, one should be cautious with respect to the identification of wage premia. When separating cohort effects from pure time and age effects an identification issue arises because the cohort (defined by the individual's year of birth) is calendar time minus age.

As a first identification approach, we follow CL by estimating equations (3.2), setting the effects for the oldest birth-cohorts (up to 1928) equal to zero. The model is formally "identified" based on annual data by using five-year age intervals and implicitly assuming age and cohort effects to be constant within each interval.[4] A test for the existence of cohort effects is then conducted

[3] Note that the approximate zero-growth of the relative employment of high-skilled in the first age class should not be over-interpreted in our context, because it likely reflects the extension of education durations and the corresponding deferments of labor market entries during the last decades; cf., for example, Reinberg and Hummel (1999).

[4] Boockmann and Steiner (2000) follow a similar identification strategy by defining their cohorts to span periods of five or ten years. In addition, the study considers actual experience rather than age.

by testing for joint significance of all other cohort terms. This approach is suggestive from an economic point of view. However, it resolves the identification problem in a rather ad hoc way; see Heckman and Robb (1985) for a detailed discussion of the identification issue. We employ an alternative approach introduced by MaCurdy and Mroz (1995) and also used in Fitzenberger, Hujer, MaCurdy, and Schnabel (2001) which deals with the identification issue both more explicitly and more rigorously.

Following this approach, we formalize cohort effects as the outcome of interaction between age and time by allowing for interaction terms of different order. For identification, the linear cohort effect is explicitly set to zero.[5] To test for the existence of cohort effects, we estimate the following specification:

$$r_{sm,a,t} = b_{sm,a} + d_{sm,t} + \sum_{i=1}^{4} \gamma_{i,sm} R_{i,a,t} + \xi_{a,t} K_{sm,\text{after}}(c_{a,t}) \tag{3.3}$$

$$+(1-\xi_{a,t}) K_{sm,\text{before}}(c_{a,t}) + \epsilon_{sm,a,t}, \quad s \in \{l,h\}, \quad \xi_{a,t} = \begin{cases} 1 & : \quad c_{a,t} \geq 0 \\ 0 & : \quad \text{else} \end{cases},$$

using age and time dummy variables as well as year of birth $c_{a,t}$ as cohort variable, normalized to zero for those aged 25 in the year 1975. The pure, separable cohort effects for those entering the labor market after and before 1975, respectively, are given by

$$K_{sm,k}(c_{a,t}) = \delta_{k,1,sm} c_{a,t}^2 + \delta_{k,2,sm} c_{a,t}^3 + \delta_{k,3,sm} c_{a,t}^4, \quad k \in \{\text{after,before}\}, \quad s \in \{l,h\}. \tag{3.4}$$

The terms $R_{i,a,t}$ capture polynomial interaction terms between age and cohorts in the time derivative of $r_{sm,a,t}$ as defined in MaCurdy and Mroz (1995).[6]

As a second specification, we use polynomials of order four in time instead of time dummies. In both specifications separability of age and time effects on the wage differentials holds if $\gamma_{i,sm} = 0$ for all i. Under this assumption, additive models can be valid representations. Uniform growth in wage ratios holds if additionally the pure effects for the cohorts after 1975 are equal to zero: $\gamma_{i,sm} = \delta_{\text{after},j,sm} = 0$ for all i, j. In this case, the existence of cohort effects is denied for those whose entire working life cycle falls into the observation period. Finally, one may test whether even older cohorts do not face any cohort effects: $\gamma_{i,sm} = \delta_{\text{after},j,sm} = \delta_{\text{before},h,sm} = 0$ for all h, i, j; see MaCurdy and Mroz (1995) for further details.

[5] It is natural to set the linear cohort effect to zero because in a model with separable age and time effects and only a linear cohort effect, one only observes parallel shifts of the cross-sectional age profiles over time; see Fitzenberger, Hujer, MaCurdy, and Schnabel (2001).

[6] Adapted to our notation, the integrals of interaction terms up to second order are given by $R_{1,a,t} = c_{a,t} a_{a,t}^2/2 + a_{a,t}^3/3$, $R_{2,a,t} = c_{a,t}^2 a_{a,t}^2/2 + 2a_{a,t}^3 c_{a,t}/3 + a_{a,t}^4/4$, $R_{3,a,t} = c_{a,t} a_{a,t}^3/3 + a_{a,t}^4/4$, and $R_{4,a,t} = c_{a,t}^2 a_{a,t}^3/3 + a_{a,t}^4 c_{a,t}/2 + a_{a,t}^5/5$.

The approach is also applied to test for the existence of cohort effects in the employment dimension suggested by the graphical inspection in section 3.2.2. In this case, $r_{sm,a,t}$ in equation (3.3) is replaced by $\ln(L_{s,a,t}/L_{m,a,t})$.

The detailed estimation results for the cohort effects and the associated tests can be found in tables 3.6 and 3.7 in appendix 3.B. Our major findings are that there is evidence for cohort effects in skill wage differentials as well as in relative employment measures.[7] Yet additive separability of age, time, and cohort effects in the evolution of wage differentials does not have to be rejected. Based on these results, the estimation of the structural model introduced and discussed in the subsequent section is in fact justified.

3.4 Estimation Framework

Building on the stylized facts, we follow CL in applying a model based on the two-level CES production function developed by Sato (1967). The model treats not only workers with different educational attainment, but—well in line with the conjectures of Freeman (1979) and Welch (1979)—also similarly educated workers of different age as imperfect substitutes. Given factor remunerations according to their respective marginal products, it can be transformed into relative wage equations which permit to separate age, time, and cohort effects on the wage gaps—and therefore provides an analytical framework to link the stylized facts outlined above.

Our study extends the analysis of CL in several directions. First, we consider the three skill groups introduced above. Second, we do not only look at skill premia, but also take account of age premia implied by the model. When estimating the model, these two points require system estimation techniques. Third, we additionally allow for the possibility of cohort effects in age-specific productivity terms. Fourth, we are concerned with possible endogeneity of employment and estimate the model with and without instrumental variables.

3.4.1 The Two-Level CES Model

The Sato (1967) framework suggests a CES model of aggregate production y_t:

$$y_t = \left(\theta_{l,t} L_{l,t}^{\rho} + \theta_{m,t} L_{m,t}^{\rho} + \theta_{h,t} L_{h,t}^{\rho}\right)^{\frac{1}{\rho}}, \tag{3.5}$$

[7] Since the restrictive decomposition of cohort and age effects in equation (3.2) following CL is rejected, we do not discuss the associated results. Though, if accepted as such, this approach suggests the existence of cohort effects as well.

where $L_{s,t}$, the measures of employment in skill group s and period t, themselves are CES subaggregates of the skill- and time-specific employment quantities $L_{s,a,t}$ of individuals in age groups a:

$$L_{s,t} = \left[\sum_a \phi_{s,a} L^{\pi}_{s,a,t}\right]^{\frac{1}{\pi}}, \quad s \in \{l, m, h\}. \tag{3.6}$$

The productivity parameters $\theta_{s,t}$ covering the usual CES distribution parameters as well as the (relative) efficiency terms of the different skill groups are allowed to vary over time to capture (skill-biased) technical change, and $\phi_{s,a}$ map the productivities of the different age classes within the skill classes.[8] $\sigma_S = 1/(1-\rho)$ and $\sigma_A = 1/(1-\pi)$ denote the elasticity of substitution between two skill groups and the elasticity of substitution between different age groups within the same skill group, respectively.

Let wages be determined by the respective marginal products:

$$\frac{w_{s,a,t}}{w_{\tilde{s},\tilde{a},t}} = \frac{\frac{\partial y_t}{\partial L_{s,a,t}}}{\frac{\partial y_t}{\partial L_{\tilde{s},\tilde{a},t}}} = \frac{\theta_{s,t} \cdot L^{\rho-\pi}_{s,t} \cdot y^{1-\rho}_t \cdot \phi_{s,a} \cdot L^{\pi-1}_{s,a,t}}{\theta_{\tilde{s},t} \cdot L^{\rho-\pi}_{\tilde{s},t} \cdot y^{1-\rho}_t \cdot \phi_{\tilde{s},\tilde{a}} \cdot L^{\pi-1}_{\tilde{s},\tilde{a},t}} \tag{3.7}$$

for all $s, \tilde{s} \in \{l, m, h\}$ and $a, \tilde{a} \in \{27, ..., 52\}$. Then age-specific skill premia $r_{s\tilde{s},a,t} = \ln(w_{s,a,t}/w_{\tilde{s},a,t})$ result as

$$r_{s\tilde{s},a,t} = \ln\left(\frac{\theta_{s,t}}{\theta_{\tilde{s},t}}\right) + \ln\left(\frac{\phi_{s,a}}{\phi_{\tilde{s},a}}\right) - \frac{1}{\sigma_A}\ln\left(\frac{L_{s,a,t}}{L_{\tilde{s},a,t}}\right) + \left[\frac{1}{\sigma_A} - \frac{1}{\sigma_S}\right]\ln\left(\frac{L_{s,t}}{L_{\tilde{s},t}}\right), \quad s \neq \tilde{s}. \tag{3.8}$$

Moreover, the production technology specifies the skill-specific wage premia across age $r_{s,a\tilde{a},t} = \ln(w_{s,a,t}/w_{s,\tilde{a},t})$ as

$$r_{s,a\tilde{a},t} = \ln\left(\frac{\phi_{s,a}}{\phi_{s,\tilde{a}}}\right) - \frac{1}{\sigma_A}\ln\left(\frac{L_{s,a,t}}{L_{s,\tilde{a},t}}\right), \quad a \neq \tilde{a}. \tag{3.9}$$

Note that CL base their empirical analysis just on (3.8) whereas our study considers all information implied by both (3.8) and (3.9).

The occurrence of perfect substitutability between different age groups, i. e., $\sigma_A \to \infty$, nests the standard case of a CES with skill groups being homogeneous in the age dimension.[9] One would expect substitutability to be higher within skill groups than across, i. e., $\sigma_A > \sigma_S$. In this case both age group-specific relative employment $\ln(L_{s,a,t}/L_{\tilde{s},a,t})$ and aggregate relative employment $\ln(L_{s,t}/L_{\tilde{s},t})$ exert a negative impact on the skill premia in (3.8).

[8] Note that the evolution of the relative efficiency terms over time captured by trends in $\theta_{s,t}$ also includes drifts in the overall efficiency of age-specific labor that are common across all age groups but may vary across skill classes. Any changes, e. g., in capital endowments affecting the productivity of the different skills implicitly enter this way.

[9] Note that this occurrence does not preclude differences in the productivity parameters $\phi_{s,a}$ across age groups.

Moreover, rewriting equation (3.8) as

$$r_{s\tilde{s},a,t} = \ln\left(\frac{\theta_{s,t}}{\theta_{\tilde{s},t}}\right) + \ln\left(\frac{\phi_{s,a}}{\phi_{\tilde{s},a}}\right) - \frac{1}{\sigma_S}\ln\left(\frac{L_{s,t}}{L_{\tilde{s},t}}\right) - \frac{1}{\sigma_A}\left[\ln\left(\frac{L_{s,a,t}}{L_{\tilde{s},a,t}}\right) - \ln\left(\frac{L_{s,t}}{L_{\tilde{s},t}}\right)\right] \tag{3.10}$$

unveils the nature of incorporated cohort effects. If $\ln(L_{s,a,t}/L_{\tilde{s},a,t}) - \ln(L_{s,t}/L_{\tilde{s},t})$ varies over time, there are cohort effects in relative employment in the sense that age-specific relative employment evolves differently from the aggregate measure.[10] If, in addition, σ_A is finite, then differences in cohort size affect $r_{s\tilde{s},a,t}$ through the term in brackets.

To show that cohort effects identify σ_A, consider the general decomposition

$$\ln\left(\frac{L_{s,a,t}}{L_{\tilde{s},a,t}}\right) = \tilde{\psi}_{s\tilde{s},a} + \tilde{\mu}_{s\tilde{s},t-a} + \tilde{\nu}_{s\tilde{s},(a,t-a)} \tag{3.11}$$

such that the model involves year- (index t), age- (index a), and cohort-specific (index $t-a$) effects. Then,

$$\begin{aligned} r_{s\tilde{s},a,t} &= \ln\left(\frac{\theta_{s,t}}{\theta_{\tilde{s},t}}\right) + \left[\frac{1}{\sigma_A} - \frac{1}{\sigma_S}\right]\ln\left(\frac{L_{s,t}}{L_{\tilde{s},t}}\right) + \ln\left(\frac{\phi_{s,a}}{\phi_{\tilde{s},a}}\right) - \frac{1}{\sigma_A}\tilde{\psi}_{s\tilde{s},a} \\ &\quad -\frac{1}{\sigma_A}\left(\tilde{\mu}_{s\tilde{s},t-a} + \tilde{\nu}_{s\tilde{s},(a,t-a)}\right) \\ &\equiv \lambda_{s\tilde{s},t} + \psi_{s\tilde{s},a} + \mu_{s\tilde{s},t-a} + \nu_{s\tilde{s},(a,t-a)}, \quad s \neq \tilde{s}. \end{aligned} \tag{3.12}$$

Observe that equations (3.3) and (3.4) used to test for cohort effects in section 3.3 are flexible parameterizations of (3.12). It is clear that the identification of σ_A depends on the existence of cohort size effects.

By disregarding variations of age-specific productivity $\phi_{s,a}$ over time, any cohort effects found in the skill wage premia are implicitly attributed to changes in labor quantities. This assumption is suited in light of our main focus to operationalize the relationship between relative wages and employment, and it is not contradicted by the test results of section 3.3; compare also Juhn, Murphy, and Pierce (1993) and Welch (1979). Yet we also show how it can be relaxed in section 3.4.3 below.

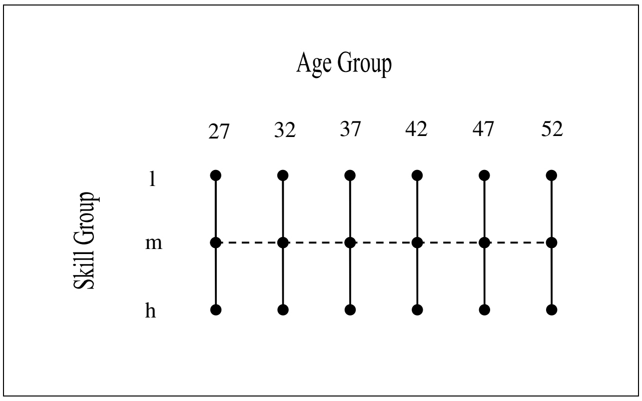
Apart from being operational for a large number of input factors, being consistent with a neo-classical production function is a great merit of the CES framework. The two-level CES offers the additional advantage that it accounts for an important aspect of heterogeneity within the skill groups: Workers of different age are allowed to be imperfect substitutes.[11]

[10] This is what has been tested explicitly in section 3.3 using the MaCurdy and Mroz (1995) approach.

3.4.2 Empirical Implementation

Equations (3.8) and (3.9) specify all wage ratios $r_{s\tilde{s},a\tilde{a},t}$ across skill and age groups for given t; see figure 3.1.[12]

Figure 3.1: System Structure of Wage Ratios



Knots: input factors. Solid lines: age-specific skill wage premia; dashed lines: skill-specific age premia.

The solid lines correspond to the age-specific skill wage premia in equation (3.8), the dashed lines to the skill-specific age premia in equation (3.9). With a total of $3 \times 6 = 18$ input factors (knots), the system implies $6 \times 3 = 18$ age-specific skill premia plus $3 \times 5(5-1)/2 = 45$ skill-specific age premia, adding up to 63 possible equations. However, there are only 17 independent wage ratios (lines). The additional 46 ratios are redundant in the sense that they can be expressed by means of the 17 independent ones.[13] For example, the wage differential between high-skilled and medium-skilled workers on the one hand and the differential between medium-skilled and low-skilled on the other hand add up to the differential between the high-skilled and the low-skilled, i. e., $r_{hl,a,t} = r_{hm,a,t} + r_{ml,a,t}$. Analogously, age premia add up; for example, $r_{s,3727,t} = r_{s,3732,t} + r_{s,3227,t}$.

The adding-up constraints translate to cross-equation restrictions in the estimation of the system. Our basic approach is to estimate the full-rank system (17 ratios) by means of Feasible GLS

[11] Prima facie, one might judge the model's functional form restrictive. In particular, the elasticities of substitution between (identically skilled) workers of different age are restricted to be all equal, so that, say, a 55-year-old executive can be replaced by an experienced 50-year-old as well as by a 25-year-young entrant. However, the model is well-suited to tell apart the effects of the two dimensions age and time. In contrast to feasible translog systems, for example, its age×time dimensioning allows to incorporate a relatively large number of input factors. For discussions on functional specification and aggregation see, e. g., Koebel (2005).

[12] One could also write down wage ratios across time based on the CES production function. However, the output in different time periods y_t would not cancel.

[13] There exist even more redundant wage ratios across both age and skill ($r_{s\tilde{s},a\tilde{a},t}$, $s \neq \tilde{s}$, $a \neq \tilde{a}$).

(FGLS).[14] In order to achieve invariance with respect to the choice of the excluded equations, we estimate the 63-equation system by System OLS (SOLS) at the first stage. As compared to the approaches pursued in the literature so far, the inclusion of equations for age premia (3.9) promises more accurate estimation of, in particular, σ_A. The following paragraphs describe the estimation strategy in more detail.

The two-level CES basically entails nonlinear system equations. However, estimation can be achieved by estimating linear models in three steps:[15]

(1) Estimate the equation system

$$\begin{aligned} r_{s\tilde{s},a,t} &= b_{s\tilde{s},a} + d_{s\tilde{s},t} - \frac{1}{\sigma_A} \ln\left(\frac{L_{s,a,t}}{L_{\tilde{s},a,t}}\right) + \epsilon_{s\tilde{s},a,t}, \quad s \neq \tilde{s} \qquad (3.13) \\ r_{s,a\tilde{a},t} &= b_{s,a\tilde{a}} - \frac{1}{\sigma_A} \ln\left(\frac{L_{s,a,t}}{L_{s,\tilde{a},t}}\right) + \epsilon_{s,a\tilde{a},t}, \quad a \neq \tilde{a} \qquad (3.14) \end{aligned}$$

to obtain an estimate for $\frac{1}{\sigma_A}$, which is equal across equations. At this step, the first-stage SOLS contains the full set of 63 equations, but the system rank, i. e., the number of equations estimated in the second stage of FGLS, is reduced to 15: in comparison with the above number of 17 independent ratios, one additional degree is lost for each of the two estimated skill ratios due to the inclusion of the skill-specific time dummies $d_{s\tilde{s},t}$.

(2) SOLS estimation of the 18-equation system

$$\ln(w_{s,a,t}) + \frac{1}{\hat{\sigma_A}} \ln(L_{s,a,t}) = d_{s,t} + \ln(\phi_{s,a}) + \epsilon_{s,a,t} \qquad (3.15)$$

then provides estimates of $\phi_{s,a}$ and allows to calculate the skill group aggregates $L_{s,t}$ defined in (3.6).

(3) Finally, the entire model—equations (3.9) and (3.10) extended by additive error terms—can be estimated, using the generated aggregates and taking account of the cross-equations restrictions concerning $1/\sigma_A$ and $1/\sigma_S$. Again, first-stage SOLS employs all 63 equations, but now FGLS uses 17 equations as explained above.

FGLS makes use of the covariance of the error terms across equations within a year. The relative productivity of workers over time, $\ln(\theta_{s,t}/\theta_{\tilde{s},t})$, is assumed to follow a linear time trend. This

[14] As will become evident below, there are also cross-equation restrictions within the non-redundant part.

[15] In general, the model can be estimated in one step using nonlinear techniques. Following CL, we proceed in three steps to avoid numerical difficulties. This is the only viable alternative because we apply bootstrapping to obtain standard errors.

approach captures the steady demand hypothesis (Acemoglu, 2002)—the steady shift towards a higher relative demand for more highly skilled labor reflects a constant rate of SBTC.[16]

Concerning the age-productivity within skill groups, $\phi_{s,a}$, two specifications are possible: First, $\ln(\phi_{s,a}/\phi_{\tilde{s},a})$ and $\ln(\phi_{s,a}/\phi_{s,\tilde{a}})$ can be estimated freely at the third step, using age dummies in analogy to the first step (model version (a)). Alternatively, $\phi_{s,a}$ may be treated as predetermined by the estimate from the second step (model version (b)). Furthermore, a version (c) would treat σ_A at the third step as predetermined by first-step estimate, and finally, both $\phi_{s,a}$ and σ_A can be taken as predetermined from previous steps (version (d)). We compare the four versions (a) to (d) in a Monte Carlo study in appendix 3.C. By and large, version (a) performs best in terms of closest point estimates and minimum root mean squared error. All of our estimations thus use this specification.

To account for estimation error in all steps of the estimation approach, we obtain bootstrap standard errors (details can be found in appendix 3.D). This is crucial because the third step estimates are based on the generated regressor $L_{s,t}$.

3.4.3 Model Relaxations and Extensions

We consider two types of model relaxations (specification tests) of the tight specification of the production technology introduced in section 3.4.1. First, we allow for elasticities of substitution between age groups being different across skill groups by replacing (3.6) with

$$L_{s,t} = \left[\sum_a \phi_{s,a} L_{s,a,t}^{\pi_s} \right]^{\frac{1}{\pi_s}}, \quad s \in \{l, m, h\}. \tag{3.16}$$

The third step now estimates

$$r_{s\tilde{s},a,t} = b_{s\tilde{s},a} + \beta_{s\tilde{s}} t - \frac{1}{\sigma_S} \ln\left(\frac{L_{s,t}}{L_{\tilde{s},t}}\right) - \frac{1}{\sigma_{As}} \ln\left(\frac{L_{s,a,t}}{L_{s,t}}\right) + \frac{1}{\sigma_{A\tilde{s}}} \ln\left(\frac{L_{\tilde{s},a,t}}{L_{\tilde{s},t}}\right) + \epsilon_{s\tilde{s},a,t} \tag{3.17}$$

$$r_{s,a\tilde{a},t} = b_{s,a\tilde{a}} - \frac{1}{\sigma_{As}} \ln\left(\frac{L_{s,a,t}}{L_{s,\tilde{a},t}}\right) + \epsilon_{s,a\tilde{a},t}. \tag{3.18}$$

This relaxation is quite plausible and the hypothesis $\sigma_{As} = \sigma_A$ for all $s \in \{l, m, h\}$ is easily tested.

[16] See also Murphy and Welch (1992, 2001) and Card and DiNardo (2002) regarding the pros and cons of this hypothesis. We also experimented with modeling breaks in SBTC which might have resulted from German unification; compare the discussion of results in section 3.5.

A second relaxation, regarding additionally the uniform elasticity of substitution between skill groups, can be implemented at the third step replacing (3.9), (3.10) by

$$r_{s\tilde{s},a,t} = b_{s\tilde{s},a} + \beta_{s\tilde{s}} t - \frac{1}{\sigma_{Ss}} \ln(L_{s,t}) + \frac{1}{\sigma_{S\tilde{s}}} \ln(L_{\tilde{s},t}) - \frac{1}{\sigma_{As}} \ln\left(\frac{L_{s,a,t}}{L_{s,t}}\right) + \frac{1}{\sigma_{A\tilde{s}}} \ln\left(\frac{L_{\tilde{s},a,t}}{L_{\tilde{s},t}}\right) + \epsilon_{s\tilde{s},a,t} \quad (3.19)$$

$$r_{s,a\tilde{a},t} = b_{s,a\tilde{a}} - \frac{1}{\sigma_{As}} \ln\left(\frac{L_{s,at}}{L_{s,\tilde{a}t}}\right) + \epsilon_{s,a\tilde{a},t} \quad (3.20)$$

and testing whether $\sigma_{Ss} = \sigma_S$ for all $s \in \{l, m, h\}$. Note however that this ad hoc relaxation abandons the theoretical consistency of the model and has to be viewed as a specification test. The parameters σ_{Ss} are no longer elasticities of substitution.

The model specifications discussed so far are referred to as “benchmark version (i)”. Next, we describe the model versions (ii) to (ix) which are estimated as extensions of (i) as part of our extensive sensitivity analysis. Table 3.8 in the appendix conveniently summarizes all estimated versions.

As discussed in section 3.4.1, a further type of cohort effects could arise if age-specific productivity $\phi_{s,a}$ were allowed to vary with time. This case would match an age bias in the evolution of the returns to (i. e., the price of) experience over time: There might be differential trends in the relative productivity of different age groups thus implying an age/cohort bias in technological progress (see footnote 8). The separability of these particular productivity components from the time effects captured by the educational skill measure $\theta_{s,t}$ has to rely on further functional form assumptions; cf. the discussion about the identification of cohort effects in section 3.3. A simple and convenient form is to assume multiplicative interaction of age or cohort and time. Then, equation (3.6) is replaced by

$$L_{s,t} = \left[\sum_a \phi_{s,a} \exp(\delta_s z t) L_{s,a,t}^{\pi}\right]^{\frac{1}{\pi}}, \quad s \in \{l, m, h\}, \quad (3.21)$$

where $z = a$ (version (ii)) or $z = t - a$ (version (iii)), respectively. Then, equations (3.9), (3.10) become

$$r_{s\tilde{s},a,t} = b_{s\tilde{s},a} + \beta_{s\tilde{s}} t + (\delta_s - \delta_{\tilde{s}}) z t - \frac{1}{\sigma_S} \ln\left(\frac{L_{s,t}}{L_{\tilde{s},t}}\right) - \frac{1}{\sigma_A}\left[\ln\left(\frac{L_{s,a,t}}{L_{\tilde{s},a,t}}\right) - \ln\left(\frac{L_{s,t}}{L_{\tilde{s},t}}\right)\right] + \epsilon_{s\tilde{s},a,t} \quad (3.22)$$

$$r_{s,a\tilde{a},t} = b_{s,a\tilde{a}} + \delta_s (z - \tilde{z}) t - \frac{1}{\sigma_A} \ln\left(\frac{L_{s,a,t}}{L_{s,\tilde{a},t}}\right) + \epsilon_{s,a\tilde{a},t}, \quad (3.23)$$

where $\tilde{z} = \tilde{a}$ (version (ii)) or $\tilde{z} = t - \tilde{a}$ (version (iii)), respectively. We test for significance of the additional parameters δ_s.

Further sensitivity tests are straightforward. Following CL, we estimate the limited information version (iv) which involves estimating solely equation (3.10) for skill wage differentials but disregards the skill-specific age premia (3.9). Version (v) excludes the youngest group of university graduates (25 to 29 years old) because the descriptive results in section 3.2.2 showed quite different trends for this group. Version (vi) examines possible breaks in the SBTC trends caused by German unification in 1990.

As comparison with a traditional CES model we estimate model version (vii) with σ_A being restricted to infinity. Focusing on the estimation of σ_S, this version still allows for productivity differences across age. Moreover, we estimate a traditional CES model (version (viii))

$$r_{s\tilde{s},t} = \text{constant}_{s\tilde{s}} + \beta_{s\tilde{s}} t - \frac{1}{\sigma_S} \ln\left(\frac{L_{s,t}}{L_{\tilde{s},t}}\right) + \epsilon_{s\tilde{s},t}, \quad s \neq \tilde{s}, \tag{3.24}$$

again testing for uniqueness of σ_S. Here, time-specific mean wage differences $r_{s\tilde{s},t} = \ln(w_{s,t}/w_{\tilde{s},t})$ are calculated as a weighted average

$$r_{s\tilde{s},t} = \frac{1}{L_{s,t} + L_{\tilde{s},t}} \sum_a (L_{s,a,t} + L_{\tilde{s},a,t})(\omega_{s,a,t} - \omega_{\tilde{s},a,t}), \quad s \neq \tilde{s} \tag{3.25}$$

of time- and age-specific differences $\omega_{s,a,t}$ pre-estimated by period specific Tobit regressions

$$\ln(w_t) = \sum_s \sum_a \omega_{s,a,t} \cdot d_{s,a,t} + \text{controls}_t + \epsilon_t\ , \tag{3.26}$$

where $d_{s,a,t}$ indicate dummies for the different skill$\times$age groups. Alternatively, version (ix) obtains skill differentials $r_{sm,t}$ directly from standard regressions which include separate skill and age dummies:

$$\ln(w_t) = r_{lm,t} \cdot d_{l,t} + r_{hm,t} \cdot d_{h,t} + \sum_a \varpi_{a,t} \cdot d_{a,t} + \text{controls}_t + \epsilon_t. \tag{3.27}$$

In contrast to the previous versions, (viii) and (ix) average out the age dimension before estimating the elasticity of substitution and $L_{s,t}$ measures aggregate employment as a headcount rather than in efficiency units. Hence, the resulting elasticities should be comparable to those found in the literature for employment in persons.

3.4.4 Endogeneity of Employment

The different measures of employment (headcounts for $L_{s,a,t}$ and efficiency units for $L_{s,t}$) are crucial here. The estimation approach builds on inverting labor demand and the literature (e. g., CL for the US, UK, and Canada) assumes equality of demand and (effective) supply and supply being inelastic in the short run. Market clearing is highly questionable in the case of Germany, since it disregards unemployment driving a wedge between demand and supply of labor.

Moreover, both observed wage premia, i. e., the relative price of skilled labor, as well as observed relative employment generally result as outcomes of all labor market processes—and should therefore be treated as endogenous in the empirical implementation. Endogeneity of employment follows, for instance, in wage-setting models with right-to-manage (RTM) assumption or efficient bargaining, in which wage bargaining takes account of the firms' employment decisions (McDonald and Solow, 1981, Nickell and Andrews, 1983,, or Arnsperger and de la Croix, 1990).

Under the RTM assumption—in contrast to efficient bargaining—observations on wages and employment lie on the demand curve. Then, the coefficient on relative employment $-1/\sigma$ in any of the models above represents the (negative) relationship between wage premia $r_{s\tilde{s}} = \ln(w_s/w_{\tilde{s}})$ and relative employment $\ln(L_s/L_{\tilde{s}})$ on the demand schedule. Unobserved shocks in output demand affect wages and employment in the same direction. Such shocks render relative employment endogenous and dilute the negative labor demand relation. Least squares estimation then yields (in absolute terms) downward-biased estimates of the true relationship or, put differently, upward-biased estimates for the elasticity of substitution σ.

As a remedy, we implement an instrumental variable (IV) approach. As instruments, we use measures of labor supply, which is assumed inelastic in the short run, possibly due to past human capital investment (Katz and Autor, 1999 and CL).[17] We compile measures of skill×age-specific labor force numbers $L^{\text{supply}}_{s,a,t}$ and aggregate numbers $L^{\text{supply}}_{s,t} = \sum_a L^{\text{supply}}_{s,a,t}$ from German Microcensus data available at the Federal Statistical Office and estimate a system of IV equations with equation-specific instruments as follows.

To instrument age-specific employment $L_{s,a,t}$ at the first and the third step of the estimation approach, we have the $3 \times 6 = 18$-equation system

$$\ln\left(L_{s,a,t}\right) = \alpha_{s,a} + d_{s,t} + \sum_{\tilde{s}} \alpha_{\tilde{s}s} \ln\left(L^{\text{supply}}_{\tilde{s},a,t}\right) + \epsilon_{s,a,t} \tag{3.28}$$

with skill×age-specific intercepts $\alpha_{s,a}$, skill-specific time dummies $d_{s,t}$, and the impact $\alpha_{\tilde{s}s}$ of the excluded instruments $L^{\text{supply}}_{\tilde{s},a,t}$.[18] The system (3.28) is also estimated first by SOLS and subsequently by FGLS in order to increase efficiency.[19] We then use predicted employment values $\hat{L}_{s,a,t}$ from (3.28) at the first and at the third step of the estimation approach; see appendix 3.D for estimation of the covariance of this sequential estimator.

[17] Accounting for the endogeneity of relative employment is also crucial as traditional demand analysis treats prices/wages as exogenous (Hamermesh, 1993).

[18] For versions (ii) and (iii), the equations of (3.28) also involve the terms $\alpha_s at$ and $\alpha_s(t-a)t$, respectively.

At the third step, we additionally instrument aggregate employment $L_{s,t}$ based on

$$\ln\left(L_{s,t}\right) = \alpha_s + \beta_s t + \sum_{\tilde{s}} \alpha_{\tilde{s}s} \ln\left(L_{\tilde{s},t}^{\text{supply}}\right) + \epsilon_{s,t} \tag{3.29}$$

with skill-specific intercepts α_s, skill-specific linear time trends $\beta_s t$, and $\alpha_{\tilde{s}s}$ as coefficients of the excluded instruments $L_{\tilde{s},t}^{\text{supply}}$.[20] In this case, there are no cross-equation restrictions and the estimation by SOLS and FGLS is straightforward. Predicted values $\hat{L}_{s,t}$ based on (3.29) are used at the third step of the estimation approach.

Results for the IV equations and tests for significance of the excluded instruments are displayed in tables 3.12 and 3.13 in appendix 3.B. The excluded instruments are jointly significant both for age-specific employment and aggregate employment, which we take as sufficient evidence in favor of our IV estimation approach. Note, however, that lack of individual significance is often the case, especially for aggregate employment.

3.5 Estimation Results

We estimate the model versions (i) to (ix) described in section 3.4.3 (see table 3.8 for a short description). Tables 3.9 to 3.11 comprise the estimates of the substitution parameters estimated at the first and at the third step and report the results of specification tests.

3.5.1 Preferred Specifications

Table 3.1 shows estimates of our preferred specifications (i), (ii), and (iii). Recall that version (i) is the benchmark specification and versions (ii) and (iii) allow for cohort effects in age-specific productivity.

In effect, the estimated elasticity σ_A usually proves finite, meaning that the estimated $1/\sigma_A$ is significantly positive: Employees of different age are imperfect substitutes. The structural model consistently mirrors the dimensions of cohort effects uncovered by the descriptive inspection in section 3.2. The preferred specifications in table 3.1 let the elasticity between age groups σ_{As} vary across skill classes (relaxation 1), but stick to a single elasticity of substitution across skill classes σ_S. The assumption of identical σ_A turns out overly restrictive in almost all cases; compare the

[19] Due to the cross-equation restrictions for the time dummies $d_{s,t}$, the rank of the system covariance matrix is reduced to 15, which does not allow us to estimate freely all age specific intercepts $\alpha_{s,a}$ with FGLS. Therefore, we replace the skill×age dummies $\alpha_{s,a}$ with skill-specific age polynomials of order three. The SOLS estimates for $\alpha_{s,a}$ do not differ significantly from polynomials of order three in age.

[20] In case of the model version (iv), the equations of (3.29) further include skill-specific breaks in the linear time trends.

Table 3.1: Elasticities of Substitution, Preferred Specifications of the Nested CES

model version		(i)		(ii)		(iii)	
Estimation		FGLS	FGLS-IV	FGLS	FGLS-IV	FGLS	FGLS-IV
	l	**8.58**	**10.31**	**9.18**	**9.44**	**9.20**	**8.68**
		(0.67)	(1.64)	(0.74)	(1.62)	(0.74)	(1.40)
σ_A	m	**4.81**	**5.27**	**5.23**	**6.01**	**5.22**	**5.38**
		(0.32)	(0.66)	(0.38)	(0.86)	(0.37)	(0.73)
	h	**19.52**	**20.13**	**10.36**	**8.50**	**10.15**	**8.59**
		(5.87)	(11.11)	(1.98)	(2.57)	(1.83)	(2.63)
σ_S		**9.36**	**6.15**	**9.49**	**6.97**	**6.32**	**4.91**
		(0.91)	(2.87)	(0.93)	(2.85)	(0.55)	(1.54)

Model versions: (i) benchmark model; (ii) with age×time interaction in age-specific productivity; (iii) with cohort×time interaction in age-specific productivity. Standard errors in parentheses based on 500 bootstrap repetitions. Bold numbers: Elasticities finite (inverses significant at 0.95 level). Data sources: IABS 1975–1997, German Microcensus.

test results in tables 3.9 and 3.10. However, differences in σ_{Ss} across skill groups are significant in most cases for FGLS, but no so for FGLS-IV. Moreover, the restricted FGLS-estimate for σ_S provides reasonable estimates of the average σ_{Ss}. Imposing the restriction of a uniform σ_{Ss} does not affect σ_{As} in a noticeable way. For these reasons, we will focus on the restricted estimates.

Regarding instrumentation, we find only little differences in the point estimates for σ_{As}. Yet the IV estimates for σ_S are considerably reduced. Along the reasoning of the previous section, unobserved shocks affect particularly aggregate relative employment, rendering this measure endogenous and estimation of σ_S without instruments inconsistent. Not surprisingly, though, the estimated standard errors are higher in case of IV estimation such that in some cases σ_S is not statistically different from 1.[21]

Our estimates of σ_S, ranging from 4.9 to 6.9 in case of IV, imply a rather high degree of substitutability compared to findings in the related literature; cf. the surveys in Hamermesh (1993) and Katz and Autor (1999). CL report elasticities of substitution between college graduates and high school graduates for Canada, the UK, and the US between 2 and 2.5.[22] In international comparison, our high elasticities are likely to reflect the small amount of over-all wage dispersion

[21] Large standard errors of our estimates may result for various reasons: First, the aggregate employment measures included at the third step are pre-generated regressors, the variation of which the bootstrap procedure takes account of. Second, FGLS instrumentation has to account for the estimation error in earlier estimation steps. Third, we lose precision by instrumentation. The labor force numbers taken to instrument employment do not differ strongly from linear time trends such that especially σ_S, the coefficient of predicted aggregate employment, is difficult to estimate precisely.

as well as the more compressed distribution of skills in Germany; cf. Nickell and Bell (1996) and Freeman and Schettkat (2001). Comparable studies for Germany also take account of three skill types, but they find elasticities not higher than 3.6.[23] Differences in the estimates may be attributed to the selection of data or the model's functional form. We address this issue when discussing additional model specifications in the next section.

Employees with different skill levels are more difficult to substitute than those with identical skill levels. The substitutability across different age groups with values of σ_{As} between 5.3 and 8.6 (version (iii), IV) is lowest among the medium-skilled.[24] This finding supports the view that low-skilled employees, mainly in positions which do not require intensive training, can be substituted relatively easily. Contrary to the hypothesis that substitutability between young and old workers diminishes (monotonically) with educational attainment (Welch, 1979),[25] an analogous reasoning applies to university graduates of different age, whose education is often said to provide them with a high competence in general problem solving. Workers with a vocational degree, however, qualify for specific tasks such that, say, younger colleagues can substitute older workers less easily.

3.5.2 Sensitivity Analysis

Tables 3.9 to 3.11 report the results for all different model versions. First of all, table 3.11 reports the outcomes of models which assume perfect substitution across age classes. Estimates of the nested CES with σ_A restricted to infinity at the third step (version (vii)) as well as from the CES model (viii), which still incorporates age×skill specific intercepts, are very close to the

[22] Other studies quantifying elasticities for the US present σ-estimates within a similar range: Autor, Katz, and Kearney (2005b), also applying a nested CES model, report elasticities around 1.6. Bound and Johnson (1992), Katz and Murphy (1992), and Krusell, Ohanian, Rios-Rull, and Violante (2000) report 1.8, 1.4, and 1.7, respectively. Ciccone and Peri (2005) prefer a span between 1.2 and 2.2, and Stapleton and Young (1988) note a value of 3.0.

[23] Fitzenberger and Franz (2001) estimate elasticities of substitution between medium- and low-skilled of 0.6–1.4 for manufacturing and of 3.0–3.6 for non-manufacturing industries, while Steiner and Wagner (1998b) and Steiner and Mohr (2000) report values for all three classes of merely 0.3–0.5 for manufacturing and 1.4 for construction and transportation. Falk and Koebel (1999, 2002) find at most substitutability between medium- and low-skilled employees, whereas Koebel, Falk, and Laisney (2003) bilaterally classify high- and medium-skilled as well as medium- and low-skilled as substitutes, but they find complementarity between low- and high-skilled employees. Entorf (1996) finds elasticities between 0.5 and 1.5 for blue and white collar workers and Beißinger and Möller (1998) of 1.8 for males and 3.3 for females.

[24] σ_{Am} even turns out lower than σ_S in several specifications reported in tables 3.9 and 3.10.

[25] Studies for the US report a much higher degree of substitutability between age classes within the group of high school graduates than among those with a college degree: Freeman (1979) finds elasticities of 14 and 2, respectively (even if the estimated reciprocals of both values show insignificant). Stapleton and Young (1988) note amounts of 73.6 (reciprocal insignificant) and 2.5. CL do not find any significant differences, though. They report significantly finite values of σ_A in the range of 4–6. Our higher estimates then reflect a higher degree of homogeneity within the skill groups defined for Germany, compared to that within the college and high school groups pertinently classifying Anglo-Saxon education systems.

results above. However, we get lower estimates of σ_S between 3.7 (IV) and 4.9 (no IV) from a traditional CES model (ix), which does not allow for age×skill interaction. On the one hand, the fact that the latter estimate is still relatively high in comparison with the literature suggests that prime age males are indeed a relatively homogeneous group. On the other hand, the finding that all specifications (i) to (viii) yield higher elasticities than the more restrictive version (ix) warrants the conclusion that models (including those in the literature) disregarding differences in the relative productivity of the different age classes incorrectly attribute too much variation in (relative) wages to changes in (relative) employment.

Second, disregarding the equations for age premia as in version (iv) in table 3.10 basically leaves σ_S unchanged, while estimates for σ_A increase. Since this approach does not use the full information content of the system, we consider these results to be less reliable.

Third, the additional interaction terms in versions (ii) and (iii) in table 3.9 are significant in most cases. Thus, there is evidence for cohort effects in age-specific relative productivities in addition to the cohort effects in relative employment. Yet the resulting elasticities are comparable to those obtained from the benchmark model (i). If anything, σ_{Ah} is lower in the versions with interactions (ii) and (iii).

Fourth, while we have so far assumed a constant rate of SBTC, we also estimated breaks in the linear time trends to capture a possible slowdown or an increase in SBTC resulting from German unification. The corresponding breaks generally turned out insignificant; see the results for version (vi) in table 3.10. Finally, the results do not change notably either when we exclude the particular group of 25–29-year-old university graduates[26] from the analysis; see version (v) in table 3.10. We are thus somewhat confident about the results of our preferred specifications in table 3.1.

3.6 Simulation Experiments

In light of the ongoing policy debate about cures for unemployment and the creation of employment, estimates from the above structural model can be used to assess the effect of wage changes on employment by means of simulation experiments.

First, and similar to the experiment conducted by Fitzenberger and Franz (2001), we estimate the magnitude of wage changes in the three skill groups that would be necessary to induce, say, a reduction of unemployment rates by one half in all three skill groups. Due to the particularly high unemployment rate among the low-skilled,[27] this design enquires about a disproportionately

[26] Compare footnote 3.

sizeable increase in employment of that deprived group, and thus is of high policy relevance. The relative wage changes are assumed to be equal for all age groups within the respective skill groups: $\Delta\ln(w_{s,a}) = \Delta\ln(\bar{w}_s)$ for all a. While allowing for relative changes between skill groups, this leaves the wage structure within skill groups unaffected.

The calculations are done for the latest available year 1997. The time index t is omitted for notational simplicity. We use a first order Taylor approximation of overall employment in each skill group s as the sum of employment in the respective age groups a:

$$L_s^* = \sum_a L_{s,a}^* = \sum_a \left(L_{s,a} + \sum_{\tilde{s}} \sum_{\tilde{a}} \frac{\partial L_{s,a}}{\partial \ln(w_{\tilde{s},\tilde{a}})} \Delta \ln(w_{\tilde{s},\tilde{a}}) \right), \quad s \in \{l, m, h\}, \tag{3.30}$$

where L_s^*, $L_{s,a}^*$ are the employment targets consistent with the goal to reduce unemployment rates by one half. Drawing on the wage elasticity of labor demand

$$\eta_{s\tilde{s},a\tilde{a}} = \frac{\partial L_{s,a}}{\partial w_{\tilde{s},\tilde{a}}} \frac{w_{\tilde{s},\tilde{a}}}{L_{s,a}} = \frac{\partial \ln(L_{s,a})}{\partial \ln(w_{\tilde{s},\tilde{a}})} = \frac{\partial L_{s,a}}{\partial \ln(w_{\tilde{s},\tilde{a}})} \frac{1}{L_{s,a}}, \tag{3.31}$$

equation (3.30) can be written in terms of relative changes:

$$\frac{\Delta L_s}{L_s} = \frac{L_s^* - L_s}{L_s} = \sum_a \frac{L_{s,a}}{L_s} \sum_{\tilde{s}} \sum_{\tilde{a}} \eta_{s\tilde{s},a\tilde{a}} \Delta \ln(w_{\tilde{s},\tilde{a}}). \tag{3.32}$$

The relationship between wage elasticities $\eta_{s\tilde{s},a\tilde{a}}$, Allen-Uzawa elasticities of substitution $\sigma_{s\tilde{s},a\tilde{a}}$, and cost shares $S_{s,a}$ implied by cost minimizing behavior of employers is given by

$$\eta_{s\tilde{s},a\tilde{a}} = S_{\tilde{s},\tilde{a}} \sigma_{s\tilde{s},a\tilde{a}} + S_{\tilde{s},\tilde{a}} \eta \quad \text{for} \quad a \neq \tilde{a} \quad \vee \quad s \neq \tilde{s}, \tag{3.33}$$

where η denotes the price elasticity of product demand and

$$\eta_{ss,aa} = \eta - \sum_{\tilde{s}} \sum_{\tilde{a} \neq a} \eta_{s\tilde{s},a\tilde{a}} - \sum_{\tilde{s} \neq s} \eta_{s\tilde{s},aa} = S_{s,a}\eta - \sum_{\tilde{s}} \sum_{\tilde{a} \neq a} S_{\tilde{s},\tilde{a}} \sigma_{s\tilde{s},a\tilde{a}} - \sum_{\tilde{s} \neq s} S_{\tilde{s},a} \sigma_{s\tilde{s},aa}; \tag{3.34}$$

see, e. g., Hamermesh (1993). Based on the nested CES production function, inter-class Allen-Uzawa partial elasticities of substitution and intra-class elasticities,[28] write

$$\sigma_{s\tilde{s},a\tilde{a}} = \sigma_S \quad \text{for} \quad s \neq \tilde{s}, \quad \text{and} \quad \sigma_{ss,a\tilde{a}} = \sigma_S + \frac{1}{S_s}(\sigma_A - \sigma_S) \quad \text{for} \quad a \neq \tilde{a}. \tag{3.35}$$

[27] The skill-specific and age-specific rates of unemployment in West Germany our simulations make use of are displayed in appendix 3.A.

[28] For model relaxation (3.16), σ_A in equation (3.35) has to be replaced by σ_{As}.

On principle, cost shares for the nested CES model can be derived directly from the model via Shepard's Lemma as functions of the productivity parameters θ_s and $\phi_{s,a}$ and wages $w_{s,a}$; cf., for example, Chung (1994). Yet the actual calculation fails this way due to the underidentification of the productivity parameters. Hence, we employ observed cost shares

$$S_{s,a} = \frac{w_{s,a} L_{s,a}}{\sum_{\tilde{s}} \sum_{\tilde{a}} w_{\tilde{s},\tilde{a}} L_{\tilde{s},\tilde{a}}} \quad \text{and} \quad S_s = \sum_a S_{s,a}. \tag{3.36}$$

The targeted relative change of employment can be inferred from the unemployment rates $ur_s = U_s/WF_s = 1 - L_s/WF_s$, where U_s and WF_s denote unemployment and work force in skill group s, respectively:

$$\frac{\Delta L_s}{L_s} = \frac{L_s^* - L_s}{L_s} = \frac{(0.5WF_s + 0.5L_s) - L_s}{L_s} = 0.5\frac{ur_s}{1 - ur_s}. \tag{3.37}$$

As η we take a weighted average of the elasticities estimated by Fitzenberger and Franz (2001) separately for the manufacturing and the non-manufacturing sector, with employment ratios in the respective sectors as weights.

Since we set $\Delta \ln(w_{s,a}) = \Delta \ln(\bar{w}_s)$ for all a, the system (3.32) yields unique solutions for the necessary wage changes based on our estimation results. The calculation of standard errors is based on the errors of the estimated parameters.

Alternatively, one might be interested in changes of the wage structure within the skill groups, holding the structure across the respective groups constant. In this context, the model set-up allows us to answer the question how the wages for employees of different age would have to change—identically in all skill groups—to reduce all age-specific unemployment rates $ur_a = U_a/WF_a = 1 - L_a/WF_a$ by one half.

In analogy to (3.30), we write

$$L_a^* = \sum_s L_{s,a}^* = \sum_s \left(L_{s,a} + \sum_{\tilde{s}} \sum_{\tilde{a}} \frac{\partial L_{s,a}}{\partial \ln(w_{\tilde{s},\tilde{a}})} \Delta \ln(w_{\tilde{s},\tilde{a}}) \right), \quad a \in \{27, ..., 52\}. \tag{3.38}$$

Now assuming $\Delta \ln(w_{s,a}) = \Delta \ln(\bar{w}_a)$ for all s, the system

$$\frac{\Delta L_a}{L_a} = \frac{L_a^* - L_a}{L_a} = \sum_s \frac{L_{s,a}}{L_a} \sum_{\tilde{s}} \sum_{\tilde{a}} \eta_{s\tilde{s},a\tilde{a}} \Delta \ln(\bar{w}_{\tilde{a}}) \tag{3.39}$$

can be solved for the necessary wage changes within the skill groups.

To evaluate the respective real magnitudes of the wage changes, we calculate the price adjustments induced by the nominal wage reductions. Here, the assumption of profit maximizing behavior under monopolistic competition takes account of endogenous output effects. We consider the Amoroso-Robinson relation for the output price level p and a constant elasticity of product demand η,

$$\left(1+\frac{1}{\eta}\right)p = MC, \quad \text{such that} \quad d\ln(p) = d\ln(MC), \tag{3.40}$$

with marginal costs

$$MC = \sum_s \sum_a w_{s,a} \frac{\partial L_{s,a}}{\partial y} = \sum_s \sum_a w_{s,a} \frac{L_{s,a}}{y} \frac{\partial L_{s,a}}{\partial y} \frac{y}{L_{s,a}} = \sum_s \sum_a \frac{w_{s,a} L_{s,a}}{y}. \tag{3.41}$$

The last equality in (3.41) follows from the constant returns to scale assumption. Relative price changes then arise from (3.40) as

$$d\ln(p) = \frac{\sum_s \sum_a \frac{L_{s,a} w_{s,a}}{y} d\ln(w_{s,a})}{\sum_{\tilde{s}} \sum_{\tilde{a}} \frac{L_{\tilde{s},\tilde{a}} w_{\tilde{s},\tilde{a}}}{y}} = \sum_s \sum_a \frac{L_{s,a} w_{s,a}}{\sum_{\tilde{s}} \sum_{\tilde{a}} L_{\tilde{s},\tilde{a}} w_{\tilde{s},\tilde{a}}} d\ln(w_{s,a}). \tag{3.42}$$

Now let $\Delta \ln(w_{s,a}) = \Delta \ln(\bar{w}_s)$ for all a in the first experiment. Then,

$$\Delta \ln(p) = \sum_s \Delta \ln(\bar{w}_s) \sum_a \frac{L_{s,a} w_{s,a}}{\sum_{\tilde{s}} \sum_{\tilde{a}} L_{\tilde{s},\tilde{a}} w_{\tilde{s},\tilde{a}}}. \tag{3.43}$$

In the second experiment, $\Delta \ln(w_{s,a}) = \Delta \ln(\bar{w}_a)$ for all s, and so

$$\Delta \ln(p) = \sum_a \Delta \ln(\bar{w}_a) \sum_s \frac{L_{s,a} w_{s,a}}{\sum_{\tilde{s}} \sum_{\tilde{a}} L_{\tilde{s},\tilde{a}} w_{\tilde{s},\tilde{a}}}. \tag{3.44}$$

Table 3.2 displays the outcome of the first simulation experiment and compares it to results obtained in Fitzenberger and Franz (2001).

Considering the employment target of reducing skill-specific unemployment rates, wages paid are too high in all skill groups, and the necessary wage reductions—ranging from 8.8 to 12.2%—are the higher the lower the skill level. This result provides evidence for wage compression across skill groups. The fact that estimated wage reductions appear rather modest may be ascribed to at least two reasons: on the one hand to the high wage elasticities resulting from the substantial elasticities of substitution, and to the assumption of constant returns to scale on the other. The latter point becomes evident by the comparison of our results to those of Fitzenberger and Franz (2001): Their specification 4, which likewise postulates constant returns to scale, yields estimates very similar to ours, whilst their unrestricted specification 3 indicates higher (nominal) reductions. The range of dispersion, however, turns out rather similar in all models.

Table 3.2: Wage Changes for Different Skill Groups Necessary to Halve Skill-Specific Unemployment Rates in 1997 and Induced Price Change

Model	$\Delta \ln(w_l)$	$\Delta \ln(w_m)$	$\Delta \ln(w_h)$	$\Delta \ln(p)$
(i) FGLS[a]	-0.109	-.093	-0.091	-0.094
	(0.0139)	(0.0139)	(0.0139)	(0.0139)
(i) FGLS-IV[a]	-0.117	-0.092	-0.089	-0.094
	(0.0175)	(0.0135)	(0.0137)	(0.0135)
(ii) FGLS[a]	-0.109	-0.093	-0.091	-0.094
	(0.0139)	(0.0139)	(0.0139)	(0.0139)
(ii) FGLS-IV[a]	-0.114	-0.093	-0.090	-0.094
	(0.0176)	(0.0150)	(0.0151)	(0.0151)
(iii) FGLS[a]	-0.116	-0.092	-0.089	-0.094
	(0.0140)	(0.0139)	(0.0139)	(0.0139)
(iii) FGLS-IV[a]	-0.122	-0.092	-0.088	-0.094
	(0.0173)	(0.0151)	(0.0152)	(0.0151)
F/F (2001)[b]	-0.141	-0.103	-	-0.105
	(0.019)	(0.020)	(-)	(0.020)
F/F (2001)[c]	-0.342	-0.313	-	-0.314
	(0.099)	(0.020)	(-)	(0.020)

[a] Calculations based on the results displayed in table 3.1. Standard errors in parentheses based on 500 bootstrap repetitions.
[b] Fitzenberger and Franz (2001), specification 4; assumption of constant returns to scale; elasticities of substitution between the high-skilled on the one hand and medium- and low-skilled on the other restricted to equal 1; no changes in wages and employment for the high-skilled; results for 1995.
[c] Fitzenberger and Franz (2001), specification 3; elasticities of substitution between the high-skilled on the one hand and medium- and low-skilled on the other restricted to equal 1; no changes in wages and employment for the high-skilled; results for 1995.

The induced relative price changes are a weighted average of the wage reductions; compare equation (3.43). Thus, given our estimates of nominal wage reductions, the high-skilled experience a real wage increase, whereas the low-skilled face real losses ex constructione.

To put this result into perspective, some remarks are in order: First, the experiment models a shock to the employment decision of the firm—we do not attempt to account for supply-side reactions to the wage changes. Second, capital and other inputs are assumed to be constant, as well.[29] Third, and similarly, the simulation does not consider substitutability with respect to

[29] None of the simulations reported in table 3.2 takes into consideration substitution effects with respect to intermediate inputs or capital stocks. Given capital-skill complementarities, for example, the reported numbers might overstate actual necessary wage changes. For the importance of capital issues in labor demand cf. Krusell, Ohanian, Rios-Rull, and Violante (2000).

participants in different labor market segments, like women or employees in mini jobs not subject to social security contributions.

The results of the second experiment, regarding a reduction of age specific unemployment rates, are displayed in table 3.3.

Table 3.3: Wage Changes for Different Age Groups Necessary to Halve Age-Specific Unemployment Rates in 1997 and Induced Price Change

Model	$\Delta\ln(w_{27})$	$\Delta\ln(w_{32})$	$\Delta\ln(w_{37})$	$\Delta\ln(w_{42})$	$\Delta\ln(w_{47})$	$\Delta\ln(w_{52})$	$\Delta\ln(p)$
(i) FGLS[a]	-0.087	-0.087	-0.086	-0.086	-0.086	-0.087	-0.087
	(0.0128)	(0.0128)	(0.0128)	(0.0128)	(0.0128)	(0.0128)	(0.0128)
(i) FGLS-IV[a]	-0.087	-0.087	-0.087	-0.086	-0.086	-0.087	-0.087
	(0.0126)	(0.0125)	(0.0125)	(0.0125)	(0.0125)	(0.0125)	(0.0125)
(ii) FGLS[a]	-0.088	-0.087	-0.086	-0.086	-0.086	-0.087	-0.087
	(0.0128)	(0.0128)	(0.0128)	(0.0128)	(0.0128)	(0.0128)	(0.0128)
(ii) FGLS-IV[a]	-0.088	-0.087	-0.087	-0.086	-0.086	-0.087	-0.087
	(0.0140)	(0.0140)	(0.0140)	(0.0140)	(0.0140)	(0.0140)	(0.0140)
(iii) FGLS[a]	-0.088	-0.087	-0.087	-0.086	-0.086	-0.087	-0.087
	(0.0128)	(0.0128)	(0.0128)	(0.0128)	(0.0128)	(0.0128)	(0.0128)
(iii) FGLS-IV[a]	-0.088	-0.087	-0.08	-0.086	-0.086	-0.087	-0.087
	(0.0140)	(0.0140)	(0.0140)	(0.0140)	(0.0140)	(0.0140)	(0.0140)

Calculations based on the results displayed in table 3.1. Standard errors in parentheses based on 500 bootstrap repetitions.

The calculated wage reductions in the different age groups are very similar. Yet the small degree of variation comes as no surprise because the differences in unemployment rates across the age classes are rather small. In particular, it has to be recalled that very young as well as older participants close to (early) retirement age, who can be expected to face deviant labor market conditions that result in differing unemployment rates, have been excluded from the analysis. For our sample of prime age males there is no evidence of wage compression across the age distribution. As to the underlying high elasticities of substitution and concerning the interpretation of the induced price changes, the same caveats as for the first experiment apply.

3.7 Conclusions

The evolution of age-specific skill wage premia in the German labor market between 1975 and 1997 shows that the age profiles of skill wage differentials have not moved in parallel fashion over time, but rather experienced a twist. Accordingly, it is unlikely that these developments are associated merely with age and time effects which apply uniformly to all cohorts. Furthermore, we observe a break in the inter-cohort trend of skill- and age-specific relative employment such that young birth cohorts do not follow the path of the older ones towards further skill upgrading. The empirical evidence thus suggests the existence of cohort effects affecting the evolution of both skill wage premia and relative employment. Following the approach suggested in MaCurdy and Mroz (1995), we find such cohort effects for both relative employment and wage premia.

A coherent operationalization of wages and employment in a labor demand framework is generally difficult due to the heterogeneous nature of the input factor labor. We extend the structural approach of Card and Lemieux (2001) based on the nested CES model of Sato (1967), giving rise to a complex picture of German labor demand. On the one hand, the model consistently maps rational behavior within the framework of neoclassical production theory. On the other hand, its age$\times$time dimensioning allows to incorporate a relatively large number of input factors. This way, we analyze wage differences between 18 types of labor.

The results are compatible with the steady demand hypothesis of a constant rate of SBTC as in Acemoglu (2002). Moreover, employees of different age are found to be imperfect substitutes—the model indeed takes account of age, time, and cohort effects. Our preferred specifications estimate the elasticity of substitution between skill groups to range between 4.9 and 6.9, and the elasticity of substitution between age groups between 5.2 and 20.1. Compared to the literature, these numbers are rather high. In international comparison, this finding reflects the fairly small amount of over-all wage dispersion in Germany as well as the relatively compressed distribution of skills. In comparison with alternative studies using different functional forms to model labor demand in Germany, we reckon that, on the one hand, our focus on prime age male employees in the IABS in fact results in a considerably homogeneous sample. On the other hand, approaches in the literature which disregard the interaction of skill and age are likely to report spuriously small elasticities.

On the basis of the estimated parameters, simulation experiments allow for policy-relevant implications. We simulate the magnitude of wage changes in the different skill groups that would have been necessary to reduce skill-specific unemployment rates in 1997 by one half. With wage changes equal for all age groups within the respective skill classes, this would have left the wage structure within skill groups unaffected. The necessary nominal wage changes range between 8.8 and 12.2% and are the higher the lower the employees' qualification. This finding provides evidence for the existence of wage compression—relative to a situation with reduced unemployment, there is too little wage dispersion across the different skill groups.

Our analysis shows the necessity to integrate different dimensions of heterogeneity into empirically meaningful models of labor demand. The nested CES approach allows to do this in a parsimonious way. However, it comes at the price of strong functional form assumptions.

As a final caveat, it should be mentioned that our neoclassical production function framework fails to incorporate residual wage inequality that remains within cells defined by skill, age, and year. Residual wage inequality is a major part of total wage dispersion (Juhn, Murphy, and Pierce (1993) and Fitzenberger, Garloff, and Kohn (2003)). This should be taken account of in future research on the link between wage differences and labor demand. Yet no conceptual framework exists so far to do so.

3.A Data

Throughout the empirical investigation, we make use of the IAB employment subsample (IABS) 1975–1997, a representative 1% random draw of German employees with employment spells subject to social insurance contributions. Excluding civil servants, self-employed, and freelancers, the IABS covers about 80% of all employed persons. For an extensive description of these register-based data see Bender, Hilzendegen, Rohwer, and Rudolph (1996) and Bender, Haas, and Klose (2000). Selected data at first comprise spells of both men and women employed full-time in West-Germany, excluding parallel employment and training spells.

We restrict attention to prime-age employees between 25 and 55 years to circumvent a number of sample selection problems. Since the IABS contains no information on hours worked, we undertake a headcount to derive an employment measure, weighting each observation with the length of the respective employment spell. This procedure assumes that the number of, say, monthly hours does not change over time nor does it differ by individual, justifying the concentration on full-time employees only.

Concerning the wage data, Steiner and Wagner (1997b) report a structural break between 1983 and 1984. In order not to deceivingly interpret this as increasing wage inequality across skill groups, we apply the correction procedure suggested by Fitzenberger (1999).

Observations are classified into three skill groups according to the individuals' educational attainment. The group of the low-skilled consists of employees without any vocational training. Those with a vocational training degree are considered medium-skilled, and individuals with a university or technical college degree form the group of the high-skilled. To deal with measurement error in the education information when defining the skill groups, we correct the skill information such that formal degrees an individual has once obtained are not lost later.

Stage zero of the estimation approach estimates wage differentials by means of Tobit regressions due to the censoring of wage data induced by the social security taxation threshold (*Beitragsbemessungsgrenze*). Observations are weighted by the length of the respective employment spells. As a first approach, equation (3.1) includes dummies for foreigners and women as control variables and further allows for possible interactions of these with the skill variables. Besides, a linear age term captures variation within the age classes. Cross terms of female and skill dummies prove significant in nearly all cells. Consequently, we base our analysis on males only. Period-specific wage differentials for the traditional CES are similarly estimated by pre-step Tobit estimations (3.26), using age-specific skill dummies and a dummy for foreigners.

Estimation equations at the first and at the second step include a full set of age dummies and time dummies for 1976–1997. The latter are replaced by a linear time trend at the third step.

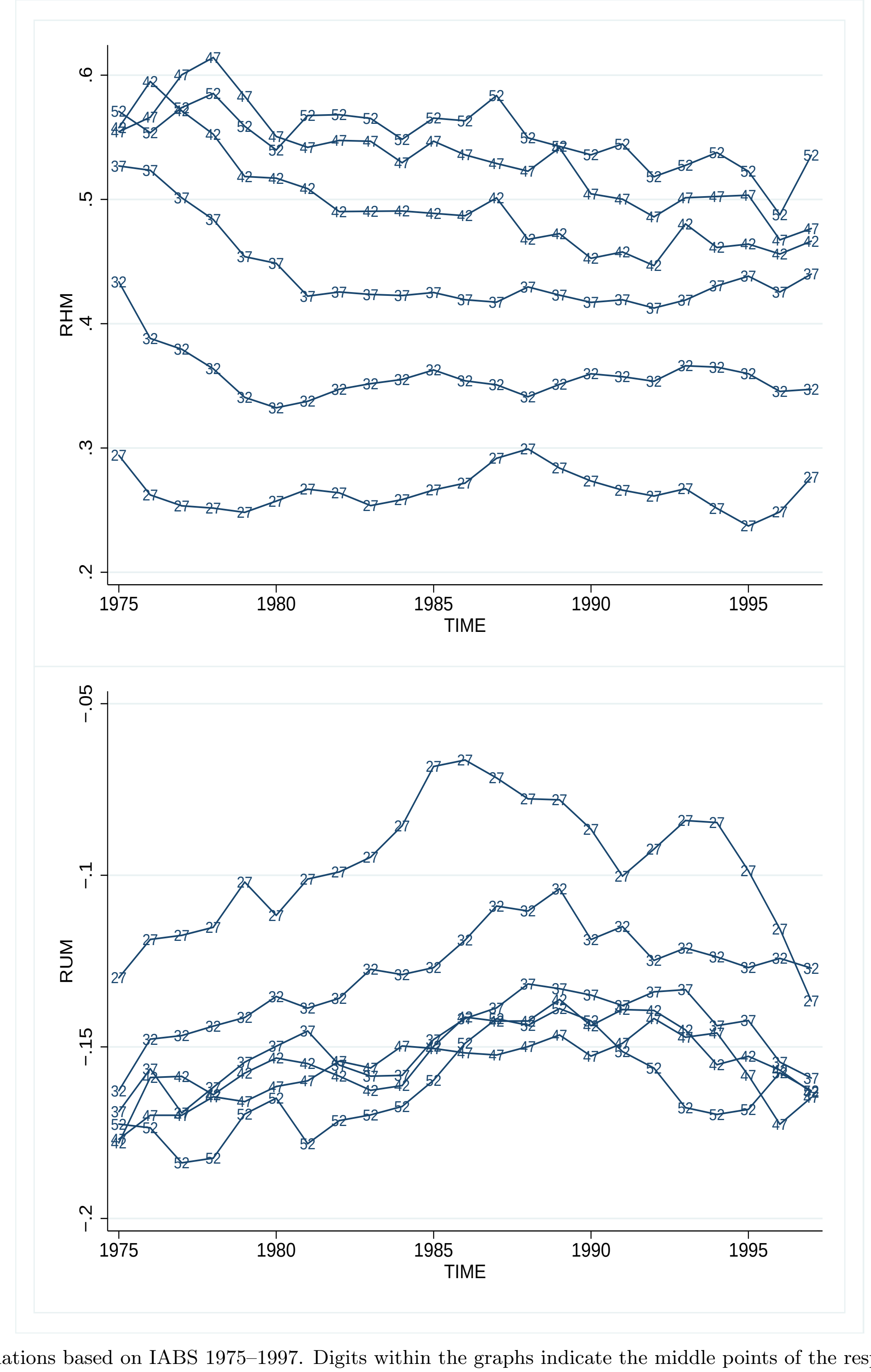
At steps one and three we instrument observed employment measures by means of the size of the labor force obtained from the German Microcensus, a representative 1% population sample collected annually, typically via face-to-face interviews. We use representative subsamples available through the Federal Statistical Office (*Statistisches Bundesamt*). The cell-specific labor force is imputed as the sum of (male) employed and unemployed workers within the skill×age groups. For several years within our sampling period, however, individual records of educational attainment were voluntary, leading to sizable shares of missing values. We apply the procedure developed in Fitzenberger, Schnabel, and Wunderlich (2004) to assign the shares of missings to the three skill groups in each cell. For the years without any skill information in the German Microcensus, we interpolate; see also Fitzenberger (1999).

For the first simulation experiment, skill-specific unemployment rates are taken from Reinberg and Hummel (2002). Rates for low-, medium-, and high-skilled males in 1997 read 27.1%, 6.8%, and 3.0%, respectively. Age group-specific unemployment rates for the second experiment are calculated based on Statistisches Bundesamt (1998). For the six age groups (from young to old) the rates are 8.5%, 7.5%, 7.4%, 7.1%, 7.0%, and 8.1%.

In order to obtain employment weights for the manufacturing and the non-manufacturing sector, we assign the IABS sector codes to the two categories as done in Fitzenberger (1999). Using the 1997-weights (0.4412 for manufacturing and 0.4746 for non-manufacturing), we calculate the price elasticity of demand, η, as a weighted average of the elasticities $\eta_{\text{man}} = -0.7994$ and $\eta_{\text{non-man}} = -0.1762$ estimated by Fitzenberger and Franz (2001).

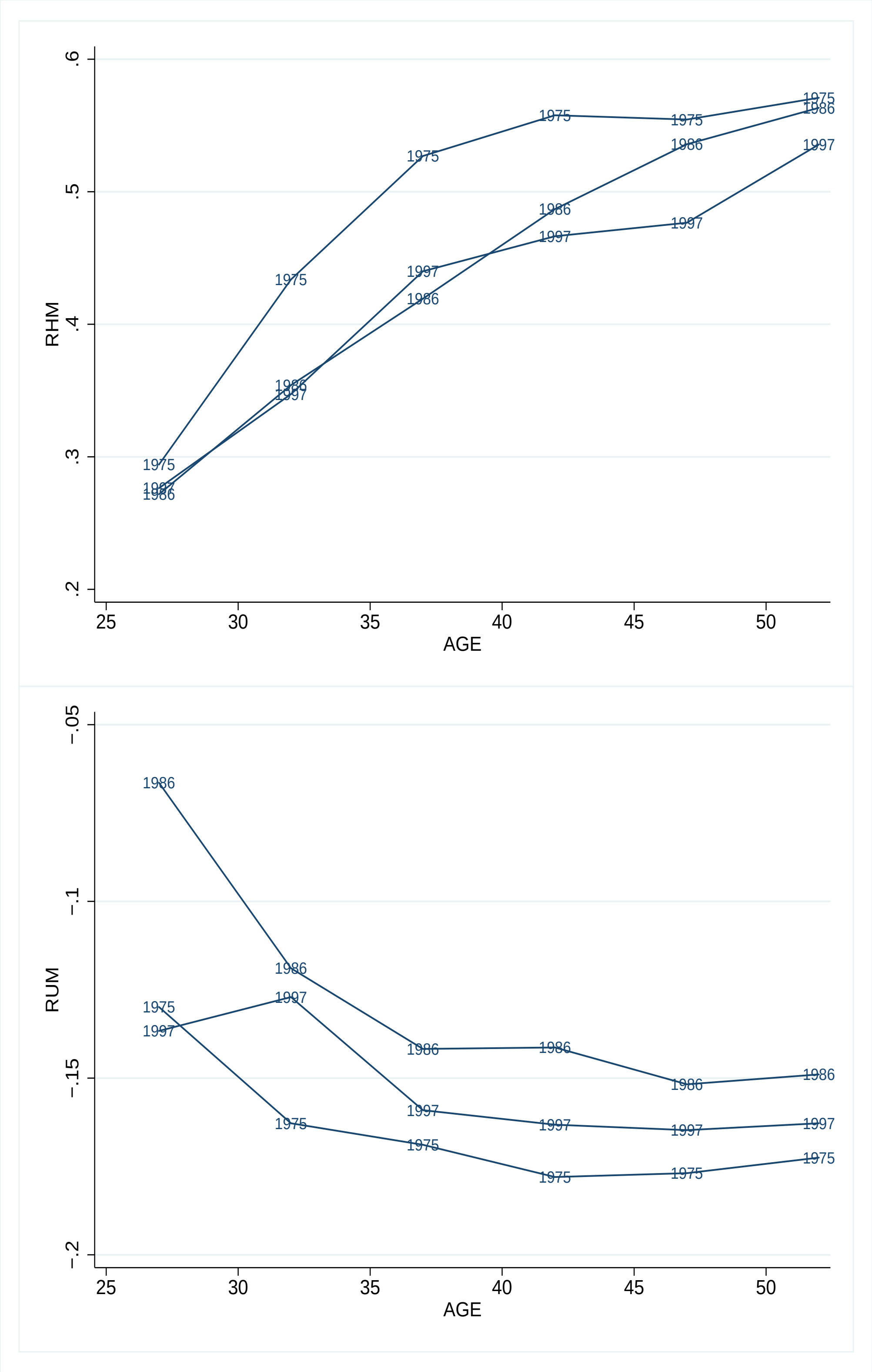
3.B Additional Tables and Figures

Figure 3.2: Evolution of Wage Differentials over Time



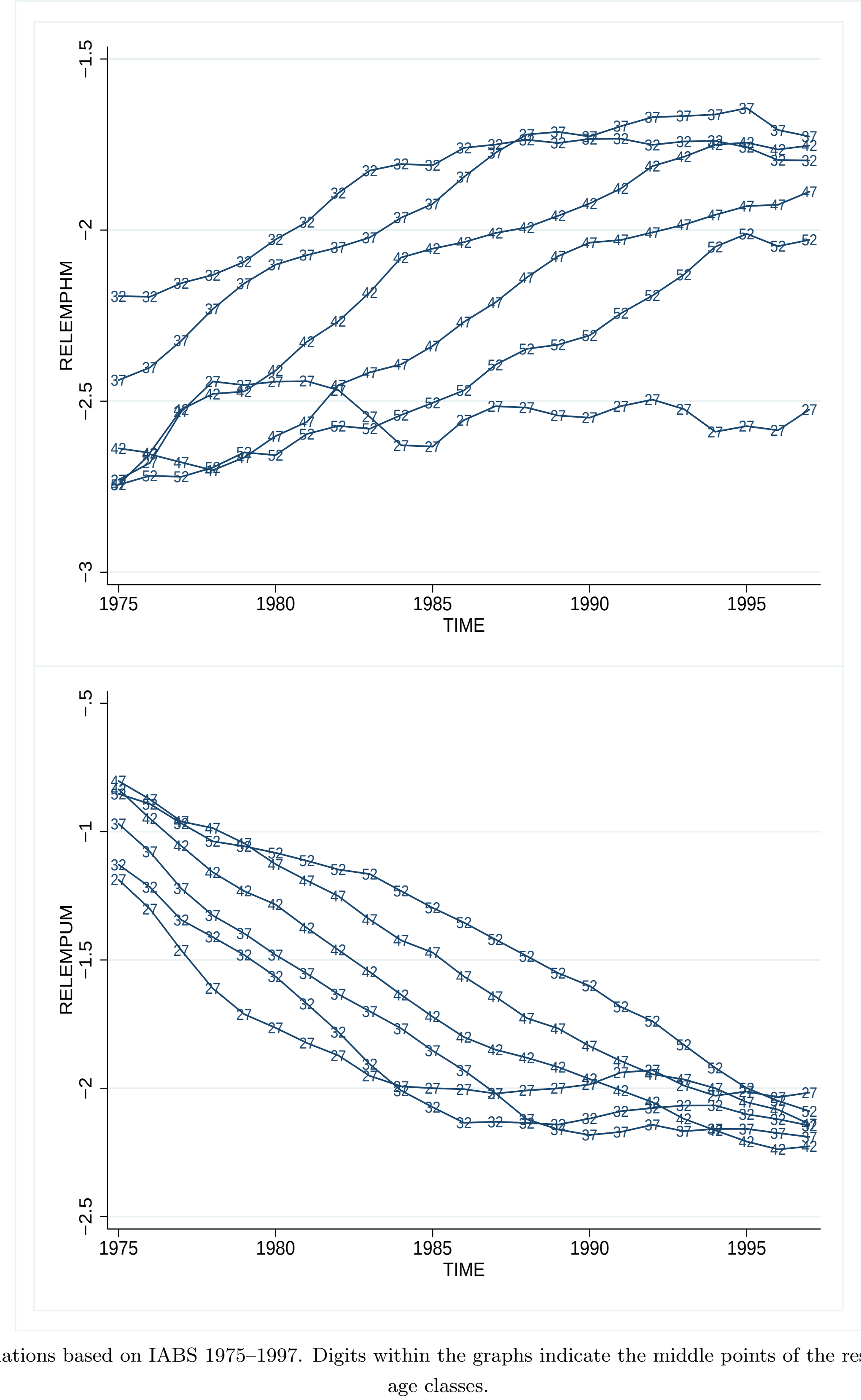
Calculations based on IABS 1975–1997. Digits within the graphs indicate the middle points of the respective age classes.

Figure 3.3: Age Profiles of Wage Differentials



Calculations based on IABS 1975–1997. Digits within the graphs indicate the calendar years of the respective age-time cells.

Figure 3.4: Trends in Relative Employment



Calculations based on IABS 1975–1997. Digits within the graphs indicate the middle points of the respective age classes.

Table 3.4: Estimated Wage Differentials by Age and Time I

Age / Time	25–29 $r_{l,a,t}$	25–29 $r_{h,a,t}$	30–34 $r_{l,a,t}$	30–34 $r_{h,a,t}$	35–39 $r_{l,a,t}$	35–39 $r_{h,a,t}$
1975	-0.1299	0.2942	-0.1628	0.4338	-0.1689	0.5270
	(0.0057)	(0.0111)	(0.0058)	(0.0113)	(0.0048)	(0.0140)
1976	-0.1187	0.2622	-0.1477	0.3882	-0.1564	0.5235
	(0.0057)	(0.0102)	(0.0062)	(0.0102)	(0.0050)	(0.0123)
1977	-0.1175	0.2535	-0.1467	0.3795	-0.1692	0.5013
	(0.0060)	(0.0093)	(0.0065)	(0.0097)	(0.0055)	(0.0113)
1978	-0.1151	0.2517	-0.1439	0.3638	-0.1618	0.4839
	(0.0064)	(0.0089)	(0.0067)	(0.0093)	(0.0059)	(0.0104)
1979	-0.1020	0.2482	-0.1414	0.3407	-0.1544	0.4540
	(0.0067)	(0.0090)	(0.0066)	(0.0085)	(0.0062)	(0.0098)
1980	-0.1117	0.2573	-0.1353	0.3324	-0.1498	0.4486
	(0.0068)	(0.0088)	(0.0066)	(0.0081)	(0.0068)	(0.0102)
1981	-0.1011	0.2669	-0.1387	0.3376	-0.1454	0.4222
	(0.0070)	(0.0088)	(0.0068)	(0.0078)	(0.0072)	(0.0100)
1982	-0.0990	0.2640	-0.1358	0.3473	-0.1553	0.4256
	(0.0071)	(0.0089)	(0.0071)	(0.0075)	(0.0075)	(0.0098)
1983	-0.0947	0.2537	-0.1274	0.3518	-0.1585	0.4236
	(0.0072)	(0.0090)	(0.0078)	(0.0075)	(0.0078)	(0.0096)
1984	-0.0856	0.2585	-0.1289	0.3551	-0.1583	0.4227
	(0.0073)	(0.0094)	(0.0084)	(0.0077)	(0.0081)	(0.0094)
1985	-0.0683	0.2663	-0.1269	0.3628	-0.1479	0.4251
	(0.0074)	(0.0096)	(0.0087)	(0.0078)	(0.0084)	(0.0092)
1986	-0.0664	0.2718	-0.1189	0.3541	-0.1417	0.4193
	(0.0073)	(0.0090)	(0.0089)	(0.0076)	(0.0087)	(0.0089)
1987	-0.0716	0.2917	-0.1090	0.3509	-0.1386	0.4174
	(0.0072)	(0.0087)	(0.0088)	(0.0074)	(0.0090)	(0.0086)
1988	-0.0777	0.2992	-0.1104	0.3413	-0.1317	0.4296
	(0.0072)	(0.0088)	(0.0084)	(0.0069)	(0.0094)	(0.0083)
1989	-0.0780	0.2839	-0.1039	0.3512	-0.1330	0.4230
	(0.0070)	(0.0086)	(0.0084)	(0.0070)	(0.0096)	(0.0084)
1990	-0.0866	0.2735	-0.1187	0.3597	-0.1349	0.4172
	(0.0069)	(0.0086)	(0.0081)	(0.0069)	(0.0094)	(0.0082)
1991	-0.1002	0.2661	-0.1149	0.3574	-0.1379	0.4192
	(0.0068)	(0.0084)	(0.0078)	(0.0067)	(0.0093)	(0.0081)
1992	-0.0924	0.2614	-0.1248	0.3537	-0.1340	0.4126
	(0.0066)	(0.0081)	(0.0075)	(0.0065)	(0.0089)	(0.0077)
1993	-0.0840	0.2672	-0.1212	0.3662	-0.1333	0.4190
	(0.0069)	(0.0083)	(0.0075)	(0.0065)	(0.0089)	(0.0076)
1994	-0.0846	0.2514	-0.1238	0.3651	-0.1438	0.4305
	(0.0073)	(0.0089)	(0.0075)	(0.0063)	(0.0090)	(0.0075)
1995	-0.0987	0.2373	-0.1269	0.3598	-0.1423	0.4383
	(0.0075)	(0.0091)	(0.0077)	(0.0065)	(0.0090)	(0.0075)
1996	-0.1156	0.2486	-0.1242	0.3456	-0.1544	0.4255
	(0.0079)	(0.0097)	(0.0078)	(0.0066)	(0.0091)	(0.0076)
1997	-0.1366	0.2764	-0.1271	0.3473	-0.1591	0.4400
	(0.0089)	(0.0106)	(0.0083)	(0.0069)	(0.0092)	(0.0077)

Tobit estimations, observations weighted with the length of the respective employment spells. Standard errors in parentheses. Data source: IABS 1975–1997.

Table 3.5: Estimated Wage Differentials by Age and Time II

Age	40–44		45–49		50–54	
Time	$r_{l,a,t}$	$r_{h,a,t}$	$r_{l,a,t}$	$r_{h,a,t}$	$r_{l,a,t}$	$r_{h,a,t}$
1975	-0.1780	0.5577	-0.1769	0.5445	-0.1725	0.5709
	(0.0052)	(0.0189)	(0.0055)	(0.0213)	(0.0069)	(0.0248)
1976	-0.1586	0.5713	-0.1699	0.5662	-0.1736	0.5537
	(0.0053)	(0.0147)	(0.0056)	(0.0183)	(0.0065)	(0.0208)
1977	-0.1586	0.5713	-0.1699	0.6003	-0.1838	0.5739
	(0.0053)	(0.0147)	(0.0056)	(0.0185)	(0.0066)	(0.0205)
1978	-0.1640	0.5525	-0.1645	0.6142	-0.1824	0.5853
	(0.0055)	(0.0132)	(0.0059)	(0.0182)	(0.0065)	(0.0191)
1979	-0.1577	0.5185	-0.1659	0.5829	-0.1696	0.5589
	(0.0054)	(0.0118)	(0.0057)	(0.0161)	(0.0063)	(0.0169)
1980	-0.1532	0.5171	-0.1615	0.5507	-0.1649	0.5398
	(0.0054)	(0.0115)	(0.0057)	(0.0148)	(0.0060)	(0.0163)
1981	-0.1549	0.5087	-0.1599	0.5420	-0.1782	0.5675
	(0.0057)	(0.0111)	(0.0057)	(0.0143)	(0.0060)	(0.0164)
1982	-0.1585	0.4902	-0.1542	0.5475	-0.1715	0.5682
	(0.0060)	(0.0105)	(0.0057)	(0.0128)	(0.0062)	(0.0155)
1983	-0.1626	0.4905	-0.1561	0.5468	-0.1698	0.5652
	(0.0066)	(0.0101)	(0.0060)	(0.0122)	(0.0062)	(0.0149)
1984	-0.1613	0.4907	-0.1497	0.5296	-0.1674	0.5484
	(0.0073)	(0.0102)	(0.0062)	(0.0115)	(0.0066)	(0.0145)
1985	-0.1498	0.4888	-0.1504	0.5469	-0.1597	0.5655
	(0.0082)	(0.0109)	(0.0065)	(0.0116)	(0.0069)	(0.0145)
1986	-0.1413	0.4869	-0.1517	0.5359	-0.1490	0.5633
	(0.0090)	(0.0112)	(0.0067)	(0.0111)	(0.0069)	(0.0140)
1987	-0.1425	0.5013	-0.1524	0.5289	-0.1418	0.5836
	(0.0095)	(0.0117)	(0.0073)	(0.0113)	(0.0070)	(0.0138)
1988	-0.1425	0.4679	-0.1499	0.5228	-0.1437	0.5495
	(0.0096)	(0.0110)	(0.0077)	(0.0108)	(0.0071)	(0.0124)
1989	-0.1362	0.4725	-0.1465	0.5416	-0.1388	0.5427
	(0.0098)	(0.0111)	(0.0083)	(0.0115)	(0.0073)	(0.0125)
1990	-0.1438	0.4527	-0.1527	0.5045	-0.1424	0.5359
	(0.0096)	(0.0104)	(0.0092)	(0.0118)	(0.0073)	(0.0122)
1991	-0.1391	0.4578	-0.1490	0.5002	-0.1513	0.5446
	(0.0095)	(0.0102)	(0.0101)	(0.0126)	(0.0077)	(0.0124)
1992	-0.1394	0.4468	-0.1417	0.4860	-0.1562	0.5183
	(0.0095)	(0.0095)	(0.0102)	(0.0121)	(0.0080)	(0.0120)
1993	-0.1450	0.4803	-0.1472	0.5014	-0.1677	0.5274
	(0.0098)	(0.0095)	(0.0104)	(0.0121)	(0.0088)	(0.0121)
1994	-0.1552	0.4614	-0.1459	0.5024	-0.1697	0.5378
	(0.0102)	(0.0092)	(0.0106)	(0.0106)	(0.0099)	(0.0123)
1995	-0.1527	0.4641	-0.1583	0.5034	-0.1682	0.5226
	(0.0102)	(0.0091)	(0.0106)	(0.0114)	(0.0112)	(0.0130)
1996	-0.1568	0.4561	-0.1725	0.4675	-0.1575	0.4877
	(0.0105)	(0.0091)	(0.0105)	(0.0107)	(0.0120)	(0.0132)
1997	-0.1632	0.4664	-0.1647	0.4766	-0.1628	0.5356
	(0.0105)	(0.0092)	(0.0110)	(0.0106)	(0.0129)	(0.0141)

Tobit estimations, observations weighted with the length of the respective employment spells. Standard errors in parentheses. Data source: IABS 1975–1997.

Table 3.6: Cohort Effects in Wage Differentials?

Coefficients	Wage Differential l/m				Wage Differential h/m			
DJ32	-0.03277	(-10.72)	-0.03278	(-9.51)	0.10880	(12.00)	0.10887	(15.74)
DJ37	-0.05143	(-9.00)	-0.05145	(-8.17)	0.19881	(13.88)	0.19895	(15.78)
DJ42	-0.05798	(-5.37)	-0.05801	(-4.93)	0.25360	(9.79)	0.25384	(10.69)
DJ47	-0.06539	(-3.37)	-0.0655	(-3.10)	0.29732	(6.75)	0.29769	(7.29)
DJ52	-0.06898	(-2.19)	-0.06907	(-2.03)	0.32937	(4.87)	0.32994	(5.46)
DT76	0.01123	(3.58)			-0.00966	(-0.80)		
DT77	0.00835	(2.45)			-0.01215	(-1.09)		
DT78	0.01117	(3.50)			-0.01760	(-1.32)		
DT79	0.01841	(5.94)			-0.0428	(-3.41)		
DT80	0.02124	(5.34)			-0.05278	(-3.93)		
DT81	0.02133	(5.24)			-0.05340	(-3.73)		
DT82	0.02226	(5.37)			-0.05424	(-3.71)		
DT83	0.02336	(4.55)			-0.05669	(-3.64)		
DT84	0.02661	(4.39)			-0.06185	(-3.54)		
DT85	0.03486	(5.06)			-0.05445	(-2.89)		
DT86	0.04075	(5.36)			-0.05967	(-2.89)		
DT87	0.04317	(4.89)			-0.05404	(-2.26)		
DT88	0.04341	(4.27)			-0.0663	(-2.53)		
DT89	0.04695	(4.04)			-0.06747	(-2.30)		
DT90	0.04024	(3.01)			-0.08155	(-2.49)		
DT91	0.03848	(2.66)			-0.08363	(-2.30)		
DT92	0.03976	(2.22)			-0.09726	(-2.38)		
DT93	0.03882	(1.96)			-0.08609	(-1.89)		
DT94	0.03565	(1.62)			-0.09096	(-1.81)		
DT95	0.03279	(1.33)			-0.09749	(-1.74)		
DT96	0.02853	(1.02)			-0.11586	(-1.86)		
DT97	0.02480	(0.79)			-0.09967	(-1.40)		
TIME			0.00174	(1.08)			-0.01827	(-4.84)
TIME2			0.00032	(1.09)			0.00217	(3.16)
TIME3			0.00002	(-1.06)			-0.00012	(-2.67)
TIME4			2.16e-07	(0.47)			2.38e-06	(2.26)
R1	-0.00054	(-1.05)	-0.00054	(-0.96)	0.00059	(0.35)	0.00059	(0.32)
R2	-6.09e-06	(-0.23)	-6.14e-06	(-0.22)	-0.00009	(-0.80)	-0.00009	(-0.78)
R3	0.00017	(1.87)	0.00017	(1.68)	-3.65e-06	(-0.01)	-3.26e-06	(-0.01)
R4	2.89e-06	(0.44)	2.90e-06	(0.42)	0.00003	(1.30)	0.00003	(1.19)
COHORTA2	0.00018	(1.86)	0.00018	(1.78)	0.00071	(3.22)	0.00071	(4.60)
COHORTB2					0.00093	(3.52)	0.0009	(2.89)
COHORTA3	-0.00001	(-3.67)	-0.00001	(-3.84)	-0.00003	(-3.32)	-0.00003	(-3.62)
COHORTB3					0.00003	(1.75)	0.00003	(1.74)
CONSTANT	-0.12597	(-41.59)	-0.12178	(-35.41)	0.28023	(23.91)	0.28895	(43.45)
Tests[a]								
Separability[b]	7.83^{*}		6.27		8.11^{*}		9.13^{**}	
Cohorts after 1975[c]	36.57^{***}		34.79^{***}		34.14^{***}		36.00^{***}	
Any cohort effects[d]					233.12^{***}		232.72^{***}	

Data source: IABS 1975–1997. White robust t-values in parentheses. Specification of equation (3.4): Inclusion of additional polynomial cohort terms as long as neither the respective coefficient nor those of lower orders turn insignificant.

[a] Wald tests, χ^2-values. $^{*}(^{**},^{***})$ Hypothesis rejected at 0.90 (0.95, 0.99) level.

[b] $H_0 : \mathrm{R}_i = 0$ for all i.

[c] $H_0 : \mathrm{R}_i = \mathrm{COHORTA}_j = 0$ for all i, j.

[d] $H_0 : \mathrm{R}_i = \mathrm{COHORTA}_j = \mathrm{COHORTB}_h = 0$ for all h, i, j.

Table 3.7: Cohort Effects in Relative Employment?

Coefficients	Log. Relative Employment l/m				Log. Relative Employment h/m			
DJ32	0.20288	(6.78)	0.20273	(7.04)	0.40904	(10.08)	0.40907	(10.61)
DJ37	0.43241	(9.80)	0.43213	(9.96)	0.19945	(3.48)	0.19952	(3.67)
DJ42	0.74647	(9.19)	0.74594	(9.46)	-0.19470	(-2.02)	-0.19458	(-2.13)
DJ47	1.0750	(7.67)	1.07414	(7.90)	-0.6767	(-4.19)	-0.67656	(-4.39)
DJ52	1.271	(5.76)	1.26952	(5.89)	-1.164	(-4.61)	-1.16373	(-4.84)
DT76	-0.12099	(-5.19)			0.02528	(0.69)		
DT77	-0.26984	(-12.28)			0.08666	(2.69)		
DT78	-0.38845	(-15.32)			0.13468	(3.18)		
DT79	-0.48621	(-17.13)			0.17800	(3.95)		
DT80	-0.58195	(-19.86)			0.23592	(5.19)		
DT81	-0.6860	(-22.36)			0.30340	(6.54)		
DT82	-0.79031	(-24.24)			0.3775	(8.62)		
DT83	-0.90355	(-23.41)			0.43387	(9.11)		
DT84	-1.0147	(-22.55)			0.50096	(9.02)		
DT85	-1.1114	(-21.79)			0.57136	(10.11)		
DT86	-1.2122	(-20.47)			0.67825	(11.37)		
DT87	-1.3021	(-20.57)			0.78194	(11.87)		
DT88	-1.3900	(-19.29)			0.87782	(11.65)		
DT89	-1.4639	(-17.60)			0.95953	(11.15)		
DT90	-1.5333	(-16.10)			1.0478	(10.62)		
DT91	-1.5957	(-15.04)			1.1553	(10.39)		
DT92	-1.6613	(-13.83)			1.2654	(10.11)		
DT93	-1.7531	(-12.75)			1.3681	(9.60)		
DT94	-1.8375	(-11.92)			1.473	(9.18)		
DT95	-1.9217	(-11.27)			1.5817	(8.76)		
DT96	-2.0027	(-10.59)			1.6518	(8.23)		
DT97	-2.0752	(-9.73)			1.771	(7.92)		
TIME			-0.11679	(-12.16)			0.04054	(2.86)
TIME2			-0.00076	(-0.42)			0.00105	(0.49)
TIME3			0.00020	(1.75)			0.00013	(0.99)
TIME4			-5.38e-06	(-2.04)			-4.31e-06	(-1.55)
R1	0.01396	(3.96)	0.01394	(4.10)	-0.02248	(-4.80)	-0.02248	(-5.10)
R2	0.00011	(0.62)	0.00011	(0.64)	0.00007	(0.32)	0.00007	(0.35)
R3	-0.00443	(-6.99)	-0.00443	(-7.34)	0.00455	(5.28)	0.00455	(5.63)
R4	-0.00009	(-2.05)	-0.00009	(-2.12)	0.00003	(0.66)	0.00003	(0.71)
COHORTA2	0.00484	(7.97)	0.00484	(8.05)	-0.00600	(-8.44)	-0.00600	(-8.66)
COHORTB2					-0.00326	(-4.16)	-0.00326	(-4.48)
COHORTA3	-0.00010	(-7.00)	-0.00010	(-6.91)	0.00011	(7.32)	0.00011	(7.20)
CONSTANT	-1.25039	(-40.19)	-1.26223	(-42.44)	-2.62419	(-47.43)	-2.62877	(-55.56)
Tests[a]								
Separability[b]	377.59^{***}		355.74^{***}		35.76^{***}		43.66^{***}	
Cohorts after 1975[c]	1000.3^{***}		1035.6^{***}		1004.0^{***}		1034.1^{***}	
Any cohort effects[d]					1011.1^{***}		1043.8^{***}	

Data source: IABS 1975–1997. White robust t-values in parentheses. Specification of equation (3.4): Inclusion of additional polynomial cohort terms as long as neither the respective coefficient nor those of lower orders turn insignificant.

[a] Wald tests, χ^2-values. *(**,***) Hypothesis rejected at 0.90 (0.95, 0.99) level.

[b] $H_0 : \mathrm{R}_i = 0$ for all i.

[c] $H_0 : \mathrm{R}_i = \mathrm{COHORTA}_j = 0$ for all i, j.

[d] $H_0 : \mathrm{R}_i = \mathrm{COHORTA}_j = \mathrm{COHORTB}_h = 0$ for all h, i, j.

Table 3.8: Summary of Model Versions (Specifications)

Label	Specification	Description
(i)	benchmark model	two-level CES
(ii)	extended benchmark	two-level CES with age×time interaction in age-specific relative productivity
(iii)	extended benchmark	two-level CES with cohort×time interaction in age-specific relative productivity
(iv)	sensitivity check	two-level CES disregarding age-premia
(v)	sensitivity check	two-level CES excluding university graduates aged 25–29 years
(vi)	sensitivity check	two-level CES with break in SBTC
(vii)	restricted benchmark	two-level CES with $1/\sigma_A = 0$
(viii)	CES with interaction	CES with age×skill interaction
(ix)	traditional CES	traditional CES

Table 3.9: Elasticities of Substitution, Specifications of the Nested CES

model version		(i)	(i)	(i)	(ii)	(ii)	(ii)	(iii)	(iii)	(iii)
	l		**7.10**	**7.10**		**7.24**	**7.24**		**7.24**	**7.24**
			(0.49)	(0.49)		(0.50)	(0.50)		(0.50)	(0.50)
$\sigma_{A,\text{1st step}}^{\text{FGLS}}$	m	**8.28**	**7.07**	**7.07**	**7.53**	**7.20**	**7.20**	**7.53**	**7.20**	**7.20**
		(0.55)	(0.48)	(0.48)	(0.54)	(0.50)	(0.50)	(0.54)	(0.50)	(0.50)
	h		**18.55**	**18.55**		**8.98**	**8.98**		**8.98**	**8.98**
			(5.59)	(5.59)		(1.51)	(1.51)		(1.51)	(1.51)
	l		**8.58**	**8.59**		**9.18**	**9.21**		**9.20**	**9.26**
			(0.67)	(0.69)		(0.74)	(0.75)		(0.74)	(0.76)
$\sigma_{A,\text{3rd step}}^{\text{FGLS}}$	m	**8.71**	**4.81**	**4.81**	**7.85**	**5.23**	**5.25**	**7.92**	**5.22**	**5.26**
		(0.63)	(0.32)	(0.32)	(0.60)	(0.38)	(0.39)	(0.61)	(0.37)	(0.38)
	h		**19.52**	**19.69**		**10.36**	**10.47**		**10.15**	**10.49**
			(5.87)	(5.98)		(1.98)	(2.02)		(1.83)	(1.99)
	l			**12.56**			**12.72**			**8.18**
				(1.58)			(1.75)			(0.88)
σ_{S}^{FGLS}	m	**8.97**	**9.36**	**7.15**	**9.04**	**9.49**	**6.98**	**5.65**	**6.32**	**4.74**
		(0.81)	(0.91)	(1.06)	(0.84)	(0.93)	(1.01)	(0.46)	(0.55)	(0.52)
	h			**6.81**			**6.78**			**5.36**
				(1.00)			(0.97)			(0.74)
	l		**6.87**	**6.87**		**7.23**	**7.23**		**7.23**	**7.23**
			(0.53)	(0.53)		(0.54)	(0.54)		(0.54)	(0.54)
$\sigma_{A,\text{1st step}}^{\text{FGLS-IV}}$	m	**8.11**	**6.86**	**6.86**	**7.44**	**7.20**	**7.20**	**7.44**	**7.20**	**7.20**
		(0.60)	(0.53)	(0.53)	(0.60)	(0.54)	(0.54)	(0.59)	(0.54)	(0.54)
	h		**28.95**	**28.95**		**9.85**	**9.85**		**9.86**	**9.86**
			(14.57)	(14.57)		(2.25)	(2.25)		(2.25)	(2.25)
	l		**10.31**	**11.53**		**9.44**	**10.20**		**8.68**	**9.22**
			(1.64)	(2.12)		(1.62)	(1.98)		(1.40)	(1.62)
$\sigma_{A,\text{3rd step}}^{\text{FGLS-IV}}$	m	**10.25**	**5.27**	**5.67**	**9.22**	**6.01**	**6.55**	**8.23**	**5.38**	**5.79**
		(1.52)	(0.66)	(0.82)	(1.34)	(0.86)	(1.14)	(1.19)	(0.73)	(0.91)
	h		**20.13**	21.26		**8.50**	**9.10**		**8.59**	**9.85**
			(11.11)	(12.10)		(2.57)	(3.21)		(2.63)	(3.65)
	l			$9.67^\sharp$			$11.54^\sharp$			$8.03^\sharp$
				(13.37)			(24.50)			(8.44)
$\sigma_{S}^{\text{FGLS-IV}}$	m	**8.14**	**6.15**	$5.65^\sharp$	7.94	**6.97**	$6.69^\sharp$	**5.92**	**4.91**	**4.48**$^\sharp$
		(3.11)	(2.87)	(5.72)	(4.07)	(2.85)	(7.16)	(1.65)	(1.54)	(2.16)
	h			$6.8^\sharp$			$7.19^\sharp$			$6.38^\sharp$
				(7.76)			(8.13)			(4.56)

Model versions: (i) benchmark model; (ii) with age×time interaction in age-specific productivity; (iii) with cohort×time interaction in age-specific productivity. Standard errors in parentheses based on 500 bootstrap repetitions. Bold numbers: Elasticities finite (inverses significant at 0.95 level). $^\sharp$ Respective parameters not significantly different at 0.95 level. Data sources: IABS 1975–1997, German Microcensus.

Table 3.10: Elasticities of Substitution, Further Specifications of the Nested CES

model version		(iv)	(iv)	(iv)	(v)	(v)	(v)	(vi)	(vi)	(vi)
	l		**15.19**	**15.19**		**6.85**	**6.85**		**7.10**	**7.10**
			(1.27)	(1.27)		(0.47)	(0.47)		(0.49)	(0.49)
$\sigma_{A,\text{1st step}}^{\text{FGLS}}$	m	**16.64**	**15.13**	**15.13**	**7.23**	**6.83**	**6.83**	**8.28**	**7.07**	**7.07**
		(1.37)	(1.26)	(1.26)	(0.50)	(0.46)	(0.46)	(0.55)	(0.48)	(0.48)
	h		**26.76**	**26.76**		**9.86**	**9.86**		**18.55**	**18.55**
			(6.37)	(6.37)		(1.74)	(1.74)		(5.59)	(5.59)
	l		**17.09**	**17.20**		**8.54**	**8.53**		**8.62**†	**8.64**
			(1.77)	(1.82)		(0.66)	(0.66)		(0.69)	(0.69)
$\sigma_{A,\text{3rd step}}^{\text{FGLS}}$	m	**19.01**	**9.21**	**9.26**	**7.65**	**4.79**	**4.79**	**8.86**†	**4.84**†	**4.84**
		(2.04)	(0.76)	(0.79)	(0.57)	(0.32)	(0.33)	(0.66)	(0.32)	(0.33)
	h		**24.53**	**25.54**		**11.19**	**11.23**		**19.19**†	**18.93**
			(6.41)	(7.00)		(2.39)	(2.33)		(5.66)	(5.46)
	l			**12.31**			**12.54**			24.55
				(1.69)			(1.70)			(14.19)
σ_S^{FGLS}	m	**8.94**	**8.92**	**6.83**	**8.59**	**9.24**	**6.84**	**14.42**†	**16.69**†	**13.02**
		(0.83)	(0.85)	(1.09)	(0.77)	(0.88)	(0.96)	(4.98)	(6.77)	(5.64)
	h			**5.95**			**6.46**			**7.13**
				(0.79)			(1.08)			(2.03)
	l		**14.83**	**14.83**		**6.71**	**6.71**		**6.87**	**6.87**
			(1.33)	(1.33)		(0.56)	(0.56)		(0.56)	(0.56)
$\sigma_{A,\text{1st step}}^{\text{FGLS-IV}}$	m	**15.57**	**14.78**	**14.78**	**7.02**	**6.70**	**6.70**	**8.11**	**6.86**	**6.86**
		(1.36)	(1.32)	(1.32)	(0.58)	(0.56)	(0.56)	(0.65)	(0.56)	(0.56)
	h		**21.71**	**21.71**		**9.13**	**9.13**		28.95	28.95
			(4.99)	(4.99)		(2.00)	(2.00)		(14.94)	(14.94)
	l		**15.10**	**15.03**		**8.70**	**9.83**		**10.48**†	**10.23**†
			(1.33)	(1.27)		(1.15)	(1.57)		(1.70)	(1.65)
$\sigma_{A,\text{3rd step}}^{\text{FGLS-IV}}$	m	**14.73**	**8.27**	**8.26**	**8.76**	**4.67**	**5.16**	**10.94**†	**5.31**†	**5.16**†
		(1.22)	(0.65)	(0.63)	(1.01)	(0.48)	(0.68)	(1.73)	(0.65)	(0.61)
	h		**12.15**	**11.95**		**8.40**	**11.02**		26.43†	**20.52**†
			(1.64)	(1.60)		(2.56)	(4.80)		(16.49)	(10.03)
	l			12.48$^\sharp$			13.87$^\sharp$			-26.87$^{\sharp\dagger}$
				(7.71)			(25.22)			(72.13)
$\sigma_S^{\text{FGLS-IV}}$	m	**8.09**	**8.72**	**6.00**$^\sharp$	**8.01**	**8.75**	6.10$^\sharp$	15.87†	-172.7†	-20.97$^{\sharp\dagger}$
		(2.98)	(3.30)	(1.64)	(1.90)	(3.43)	(3.51)	(12.29)	(2749.2)	(72.13)
	h			**9.10**$^\sharp$			6.11$^\sharp$			6.70$^{\sharp\dagger}$
				(3.32)			(7.11)			(16.96)

Model versions: (iv) excluding equations for age premia; (v) excluding high-skilled of age 25–29; (vi) with break in SBTC. Standard errors in parentheses based on 500 bootstrap repetitions. Bold numbers: Elasticities finite (inverses significant at 0.95 level). $^\sharp$ Respective parameters not significantly different at 0.95 level. † Time break in SBTC insignificant at 0.95 level. Data sources: IABS 1975–1997, German Microcensus.

Table 3.11: Estimates of σ_S, Assuming Perfect Substitution Between Age Classes

model version		(vii)	(vii)	(viii)	(viii)	(ix)	(ix)
	l		**12.45**		**11.95**		**10.57**
			(1.70)		(2.00)		(2.28)
σ_S^{FGLS}	m	**8.82**	**6.79**	**8.25**	**5.86**	**4.93**	**3.81**
		(0.78)	(0.97)	(0.97)	(0.94)	(0.67)	(0.45)
	h		**6.15**		**4.67**		**3.36**
			(0.80)		(0.54)		(0.24)
	l		$15.95^\sharp$		$13.80^\sharp$		198.3
			(47.95)		(8.26)		(3132.6)
$\sigma_S^{\text{FGLS-IV}}$	m	9.14	$6.93^\sharp$	**6.26**	**6.21**$^\sharp$	**3.76**	**5.15**
		(8.03)	(7.26)	(1.16)	(1.79)	(0.95)	(1.99)
	h		$8.32^\sharp$		**5.65**$^\sharp$		**4.53**
			(14.30)		(1.18)		(1.08)

Model versions: (vii) nested CES (benchmark) with $1/\sigma_A$ restricted to zero at the third step; (viii) CES model with skill differentials as employment-weighted average of age-specific premia; (ix) CES model without skill×age-interaction. Standard errors in parentheses. Bold numbers: Elasticities finite (inverses significant at 0.95 level). $^\sharp$ Respective parameters not significantly different at 0.95 level. Data sources: IABS 1975–1997, German Microcensus.

Table 3.12: Instrumental Variables: First Stage Results for Age-Specific Employment

Model	(i)/(iv)		(ii)		(iii)	
	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.
α_{uu}	0.7411^{***}	(0.0716)	0.8091^{***}	(0.0762)	0.8091^{***}	(0.0762)
α_{mu}	0.5254^{***}	(0.1204)	0.5237^{***}	(0.1296)	0.5237^{***}	(0.1296)
α_{hu}	-0.3894^{***}	(0.0677)	-0.5690^{***}	(0.0832)	-0.5690^{***}	(0.0832)
α_{um}	0.0620	(0.0458)	0.1158^{**}	(0.0490)	0.1158^{**}	(0.0490)
α_{mm}	0.9232^{***}	(0.0766)	0.9106^{***}	(0.0814)	0.9106^{***}	(0.0815)
α_{hm}	-0.1008^{**}	(0.0450)	-0.2272^{***}	(0.0572)	-0.2272^{***}	(0.0572)
α_{uh}	-0.1603^{*}	(0.0899)	-0.1044	(0.0947)	-0.1044	(0.0947)
α_{mh}	0.6185^{***}	(0.1504)	0.6178^{***}	(0.1592)	0.6178^{***}	(0.1592)
α_{hh}	0.6484^{***}	(0.0848)	0.4987^{***}	(0.1081)	0.4987^{***}	(0.1081)
χ^2	12722.7^{***}		5150.5^{***}		5150.5^{***}	

Coefficients of additional instruments. See the text for a description of the instrumentation strategy. See the text and tables 3.9 and 3.10 for descriptions of the model versions (i) to (iv). χ^2: Test for joint significance of additional instruments. $^{*}(^{**},^{***})$ parameter significant at 0.90 (0.95, 0.99) level.
Data sources: IABS 1975–1997, German Microcensus.

Table 3.13: Instrumental Variables: First Stage Results for Aggregate Employment

Model	(i)		(ii)		(iii)		(iv)	
	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.	Coeff.	Std.Err.
α_{uu}	0.3466	(0.4031)	0.3409	(0.4028)	0.3187	(0.3906)	-0.9379***	(0.2001)
α_{mu}	0.3597	(0.9445)	0.3595	(0.9436)	0.4125	(0.9151)	1.8051***	(0.4019)
α_{hu}	1.4425	(1.3686)	1.4435	(1.3673)	1.3901	(1.3260)	0.3551	(0.5588)
α_{um}	-0.2373**	(0.1074)	-0.2435**	(0.1068)	-0.3352**	(0.1424)	-0.0797	(0.1203)
α_{mm}	1.4841***	(0.2517)	1.4763***	(0.2501)	1.6945***	(0.3336)	1.3068***	(0.2417)
α_{hm}	-0.0854	(0.3647)	-0.0778	(0.3624)	-0.2980	(0.4834)	0.0480	(0.3360)
α_{uh}	-0.5004**	(0.2492)	-0.4982**	(0.2465)	-0.5909**	(0.2954)	0.1934	(0.1846)
α_{mh}	2.5600***	(0.5838)	2.5478***	(0.5774)	2.768***	(0.6921)	1.7792***	(0.3707)
α_{hh}	-1.5078*	(0.8459)	-1.4936*	(0.8367)	-1.7162*	(1.0028)	-0.9205*	(0.5154)
χ^2	235.04***		240.74***		233.86***		250.51***	

Coefficients of additional instruments. See the text for a description of the instrumentation strategy. See the text and tables 3.9 and 3.10 for descriptions of the model versions (i) to (iv). χ^2: Test for joint significance of additional instruments. *(**,***) parameter significant at 0.90 (0.95, 0.99) level.
Data sources: IABS 1975–1997, German Microcensus.

3.C Monte Carlo Study

We conduct a Monte Carlo Study in order to compare the following estimation approaches:

(a) Age-specific relative productivities $\phi_{s,a}/\phi_{\tilde{s},\tilde{a}}$ and elasticities σ_{As} estimated freely at the third step.

(b) $\phi_{s,a}$ at the third step taken as predetermined from the second step.

(c) σ_{As} at the third step taken as predetermined from the first step.

(d) $\phi_{s,a}$ as well as σ_{As} at the third step taken as predetermined from previous steps.

We assume the following parameter values for the benchmark model:

- elasticities $\sigma_{A_u} = 15$, $\sigma_{A_m} = 10$, $\sigma_{A_h} = 20$, and $\sigma_S = 2$.
- skill-specific linear time trends of 1% per year for $\ln(\theta_{h,t}/\theta_{m,t})$ and of 2% for $\ln(\theta_{m,t}/\theta_{l,t})$.
- age×skill-specific productivities $\phi_{s,a}$ set to appropriate values between $\exp(9.3)$ and $\exp(9.8)$.

When simulating log wages, we assume a normally distributed additive error term with standard deviation $STDDEV$, which captures residual wage dispersion. The chosen values for $STDDEV$ between 0.001 and 0.2 correspond to 90–10-percentile wage differences between 0.2% and 50%.

We then estimate the benchmark version (i) of the model. Results for the different approaches are displayed in table 3.14. None of the approaches strictly dominates the others in terms of minimum bias or minimum root mean squared error for all parameters. However, approach (a), which (re)estimates all parameters freely at the third step, performs best in most of the cases and, what is more, its performance is also fairly good when coming off second-best. We therefore decide to use approach (a) for the estimations throughout the paper.

Table 3.14: Monte Carlo Study: Average Third-Step Estimates for $1/\sigma$ and Root Mean Squared Errors

STDDEV	Parameter	Value	(a)	(b)	(c)	(d)
0.001	$1/\sigma_{Au}$	0.0667	**0.0666**	0.0763	0.0739	0.0739
			(**0.00046**)	(0.00966)	(0.00720)	(0.00720)
	$1/\sigma_{Am}$	0.1	**0.1000**	0.0976	0.0742	0.0742
			(**0.00085**)	(0.00244)	(0.02581)	(0.02581)
	$1/\sigma_{Ah}$	0.05	**0.0500**	0.0477	0.0488	0.0488
			(**0.00082**)	(0.00239)	(0.00144)	(0.00144)
	$1/\sigma_S$	0.5	**0.5002**	0.4730	0.4993	0.3208
			(**0.00139**)	(0.02704)	(0.00145)	(0.17936)
0.005	$1/\sigma_{Au}$	0.0667	**0.0666**	0.0767	0.0746	0.0746
			(**0.00230**)	(0.01023)	(0.00810)	(0.00810)
	$1/\sigma_{Am}$	0.1	**0.0999**	0.0981	0.0748	0.0748
			(0.00423)	(**0.00262**)	(0.02529)	(0.02529)
	$1/\sigma_{Ah}$	0.05	**0.0499**	0.0479	0.0496	0.0496
			(0.00409)	(**0.00357**)	(0.00385)	(0.00385)
	$1/\sigma_S$	0.5	**0.5004**	0.4758	0.4992	0.3180
			(0.00691)	(0.02530)	(**0.00659**)	(0.18496)
0.010	$1/\sigma_{Au}$	0.0667	**0.0665**	0.0771	0.0747	0.0747
			(**0.00459**)	(0.01116)	(0.00879)	(0.00879)
	$1/\sigma_{Am}$	0.1	**0.0998**	0.0986	0.0748	0.0748
			(0.00847)	(**0.00401**)	(0.02542)	(0.02542)
	$1/\sigma_{Ah}$	0.05	**0.0498**	0.0482	**0.0498**	**0.0498**
			(0.00818)	(**0.00607**)	(0.00767)	(0.00767)
	$1/\sigma_S$	0.5	**0.5006**	0.4765	0.4992	0.3130
			(0.01381)	(0.02757)	(**0.01347**)	(0.19854)
0.050	$1/\sigma_{Au}$	0.0667	**0.0658**	0.0790	0.0741	0.0741
			(0.02296)	(0.02341)	(**0.01983**)	(**0.01983**)
	$1/\sigma_{Am}$	0.1	**0.0989**	0.1020	0.0743	0.0743
			(0.04234)	(**0.01985**)	(0.03162)	(0.03162)
	$1/\sigma_{Ah}$	0.05	0.0492	0.0463	**0.0495**	**0.0495**
			(0.04088)	(**0.03095**)	(0.03842)	(0.03842)
	$1/\sigma_S$	0.5	0.5026	0.4989	**0.5009**	0.2976
			(0.06903)	(0.07197)	(**0.06884**)	(0.34478)
0.100	$1/\sigma_{Au}$	0.0667	**0.0649**	0.0795	0.0733	0.0733
			(0.04592)	(0.03973)	(**0.03737**)	(**0.03737**)
	$1/\sigma_{Am}$	0.1	**0.0978**	0.1064	0.0735	0.0735
			(0.08469)	(**0.03821**)	(0.04534)	(0.04534)
	$1/\sigma_{Ah}$	0.05	0.0484	0.0432	**0.0491**	**0.0491**
			(0.08177)	(**0.06613**)	(0.07685)	(0.07685)
	$1/\sigma_S$	0.5	0.5049	0.5578	**0.5031**	0.4805
			(0.13808)	(0.16231)	(**0.13778**)	(0.44487)
0.200	$1/\sigma_{Au}$	0.0667	**0.0632**	0.0746	0.0717	0.0717
			(0.09185)	(0.07540)	(**0.07372**)	(**0.07372**)
	$1/\sigma_{Am}$	0.1	**0.0955**	0.1087	0.0718	0.0718
			(0.16939)	(**0.07247**)	(0.07874)	(0.07874)
	$1/\sigma_{Ah}$	0.05	0.0469	0.0405	**0.0481**	**0.0481**
			(0.16356)	(**0.14403**)	(0.15371)	(0.15371)
	$1/\sigma_S$	0.5	0.5094	0.6651	**0.5073**	0.7133
			(0.27625)	(0.30859)	(**0.27571**)	(0.51902)

Simulation of the benchmark model based on 1000 resamples. RMSE in parentheses. Bold numbers: minimum bias and minimum RMSE, respectively. See the text of appendix 3.C for a description of the estimation approaches (a) to (d).

3.D Calculating Standard Errors

The calculation of standard errors and test statistics for the parameters obtained from the multi-step estimation approach and in the simulation experiment has to take account of pre-step estimation variability. We therefore use bootstrapping techniques.

We resample from the distribution of $\ln(w_{s,a,t})$ estimated by the Tobit regressions at stage zero (section 3.2.1, equation (3.1)). All three subsequent estimation steps as well as the calculations for the simulation experiments are put into a single bootstrap loop. We use 500 repetitions to obtain the variance-covariance matrix of the estimated parameters from the empirical bootstrap distribution.

Standard errors for the reported elasticities σ can then be calculated by means of the Delta method relying on the estimated bootstrap distribution of $1/\sigma$. Direct bootstrapping of σ is not possible because of the discontinuity of the inverse function at the argument zero. With negative estimates of the inverse elasticities for single extreme resamples, direct calculation of the variance of σ would not be not well-defined.

In case of IV estimation the three-step approach is extended at steps one and three by the estimation of IV equations. In each iteration of the bootstrap loop we draw from the estimated distribution of IV parameters, calculate predicted values for employment $L_{s,a,t}$ and $L_{s,t}$, and estimate the three-step model. Note that while predicted employment values are used in the SOLS and FGLS estimation, the calculation of the respective FGLS weighting matrices from SOLS residuals relies on actual employment.

When drawing inference on the estimated wage changes in the simulation experiment, we assume η_{man} and $\eta_{\text{non-man}}$, the price elasticities of product demand taken from Fitzenberger and Franz (2001), to be independently normally distributed.

4 The Erosion of Union Membership in Germany: Determinants, Densities, Decompositions

4.1 Introduction

Trade unions bargain for higher wages, equal pay, reduced working hours, fair working conditions, or employment protection (Freeman and Medoff, 1984). However, in Germany—and in a number of other countries—the results of union activity apply to most of the workers irrespective of membership. Membership is not compulsory and closed shop regulations are illegal. The public good character of core services offered by trade unions may give rise to free-rider behavior.

Thence, why do people join a union? Given the economic importance of union activity (Addison and Schnabel, 2003), interest in the determinants of union membership is of its own right. It is essential to disentangle and quantify the determinants of union membership in order to understand the recent decline in union membership in developed countries; see OECD (2004b) and Ebbinghaus (2003) for Germany. Moreover, facing the lack of information on union membership in various data sets, microeconometric membership estimations can be used to predict union density for homogeneously defined labor market segments, such as industries and/or regions. These predictions can then be employed to study the impact of unionization on economic performance, and on employment and the structure of wages in particular; see Fitzenberger and Kohn (2005). This is of importance since in contrast to the Anglo-Saxon literature (see the survey by Card, Lemieux, and Riddell, 2003), it is not meaningful to estimate a wage effect of individual union membership in Germany, where the public good nature of union activity results in union coverage being much higher than union density.

For Germany, a couple of microeconometric analyses of union membership using survey data are available. For example, Windolf and Haas (1989), Lorenz and Wagner (1991), and Schnabel and Wagner (2005) use different sets of cross sectional survey data to estimate binary choice models of union membership in West Germany. Schnabel and Wagner (2003) also estimate determinants of union membership in East Germany. However, none of the above studies employs panel data methods to control for unobserved heterogeneity. This was first established by Fitzenberger, Haggeney, and Ernst (1999) and Beck and Fitzenberger (2004), whose analyses for West Germany are based on three and four waves of the German Socio-Economic Panel (GSOEP), respectively.

Our study extends upon this literature in two main directions. First, we estimate determinants of union membership in East as well as in West Germany, using the panel structure of the GSOEP and applying a Chamberlain (1980)-Mundlack (1978)-type correlated random effects probit model. Our estimations are based on six waves during the period 1985–2003 providing union membership status for individuals in West Germany, and on four waves between 1993 and 2003 for East Germany. In fact, it proves important to control for individual-specific effects in the membership decision. Our findings show the influence of socio-demographic personal characteristics, such as age or marital status; the influence of workplace characteristics, i. e., match, firm, or industry specific effects; as well as the influence of attitudinal factors for the individual choice to be or not to be a union member, and we analyze differences of these factors between East and West Germany and across time.

Second, we use our estimates to predict net union density (NUD) as a measure for union strength in East and West Germany. The predictions, which consistently trace the trends towards deunionization in both parts of the country, are then analyzed by means of decomposition techniques in order to shed light on (1) the changes in unionization over time and (2) the differences in NUD between East and West Germany. We find that changes in the composition of the work force only played a minor role for the deunionization trends in both East and West Germany. In East-West comparison, differences in the characteristics of the work force are in favor of higher NUD in the West. The stronger decline in union membership in East Germany reflects a stronger change in the impact of these characteristics.

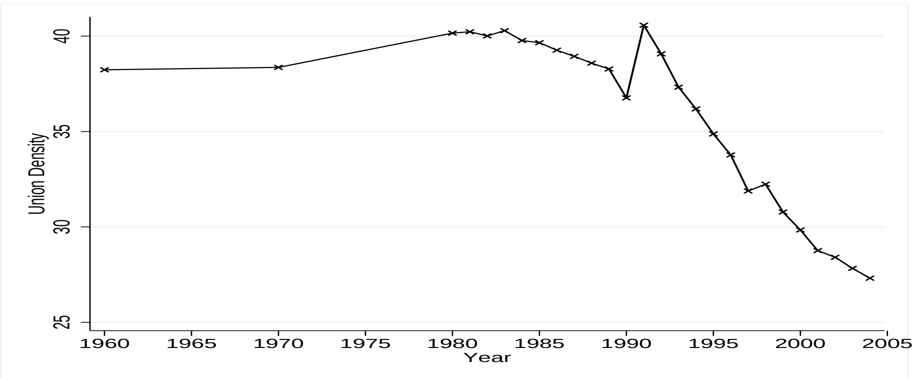
The remainder of the paper is organized as follows: Section 4.2 reviews related literature on union membership. Section 4.3 discusses potential determinants of membership decisions. Our econometric investigation is presented in section 4.4. Corresponding projections of net union densities and the decomposition analyses are discussed in section 4.5. Section 4.6 concludes. The appendix includes further information on the data and empirical results.

4.2 Related Literature

Studies of union membership in Germany face three challenges. First, collective bargaining is an open shop system. Negotiation outcomes apply not only to union members, but to the vast majority of all employees; see, e. g., Bosch (2004). Membership is not compulsory and closed shop regulations are forbidden by constitutional law. By nature, the core services trade unions offer have public good character, which gives rise to the possibility of free-riding behavior. Thus, why would people want to join a union at all? Who joins the union? And how much do different determinants such as personal or workplace characteristics contribute to people's membership decision?

Second, union membership rates have been steadily declining in recent decades. Figure 4.1 depicts gross union density (GUD), defined as the ratio of the number of union members and the number of employees in the German labor market. After a period of slight increases in the 1970's, the early 1980's mark the beginning of a pronounced trend towards deunionization, which started out at a level of about 40%. By 2004, GUD was down to a historically low level of 27%. Deunionization was merely interrupted by a unification effect in 1990, when West German institutions were transferred to the East, and unions were initially very successful in recruiting members in the East. However, the upsurge in aggregate GUD of about five percentage points (pp)—which was built on the grounds of the GDR labor organization *Freier Deutscher Gewerkschaftsbund*, whose members had comprised the largest part of the GDR work force—was not sustainable, and deunionization went on even more rapidly in the 1990's and 2000's. Some trade unions have responded to the decline in size by merging; see, e. g., Keller (2005). To date, however, unions have not been able to reverse the trend; see also Ebbinghaus (2003) and Fichter (1997). Against this background, micro-level studies which unveil how individual membership decisions have been changing over time, can give insight into the nature of observed trends.

Figure 4.1: Evolution of Gross Union Density



Gross union density in percent; 1960–1990: West Germany; 1991–2004: Unified Germany.
Union membership in CGB (*Christlicher Gewerkschaftsbund*, data until 1999: German Statistical Office (*Statistisches Bundesamt, Statistische Jahrbücher*), union information thenceforward), DAG (*Deutsche Angestelltengewerkschaft*, until 2000; data: German Statistical Office), DGB (*Deutscher Gewerkschaftsbund*, data: www.dgb.de), DBB (*Deutscher Beamtenbund*, data: www.dgb.de), and DPolG (*Deutsche Polizeigewerkschaft*, until 1970, data: German Statistical Office).

Employment (*abhängig Beschäftigte ohne mithelfende Familienangehörige*) from German Microcensus: www.destatis.de.

Third of all, the availability of adequate data from union records is limited. From 1991 onwards, only aggregate numbers for unified Germany are available, and unions' publications do not distinguish between employed members on the one hand and unemployed, retired, or student members on the other. Yet this distinction is important from an economic point of view. Net union density (NUD), defined as the share of employed union members in the number of employees, is a better measure of union power than GUD because it is more closely related to the union's financial resources and to the potential to mobilize workers within firms. Net union density is lower than gross union density by definition. Estimates of aggregate NUD usually fall short of GUD by about 10 pp and this difference also varies with the business cycle; see Ebbinghaus (2003). Union power further differs significantly between different labor market segments. For example, unions are traditionally strong in manufacturing industries, but they are of minor importance in personal service sectors. Official membership information does not distinguish between sufficiently homogenous segments. Detailed NUD estimates obtained from survey data should thus help providing meaningful measures of union strength.

The following paragraphs review contributions from the theoretical literature and existing empirical studies on these issues. The purpose is to motivate the determinants under study and the adequacy of the econometric approach pursued later on; see Naylor (2003), Riley (1997), and Schnabel (2003) for more extensive surveys.

4.2.1 Theoretical Analyses

In a traditional cost-benefit framework, potential union members balance the utility derived from being member with the costs associated to it. Membership is costly due to membership fees, and in an open shop system the key benefits arising from union bargaining—like higher wages and equal pay, reduced working hours, fair working conditions, or employment protection—are basically public goods. Therefore, a rational individual would not join a union, but rather free ride in this setting.

Several ways to accommodate the free-rider problem have been considered. First, membership could be compulsory. Second, membership would be voluntary in a dichotomous labor market with union and non-union sectors (Grossman, 1983), but all employees in unionized firms choose to be member. Both of these arrangements are essentially closed shop solutions, where only members are eligible for benefits. However, the German as well as most European empirical evidence dismisses closed shop solutions.

Third, unions offer selective incentives (Olson, 1965) in addition to the collective goods. On the one hand, these can be actual private goods (Booth and Chatterji, 1995), such as legal aid and grievance procedures, accident insurances, or even education and further training. Moreton (1998), for instance, considers greater job security for union members. On the other hand, members may comply with a social custom to support the union. The notion of social custom, as introduced to labor economics by Akerlof (1980), captures the idea that individuals abide by internal rules or norms set within society because non-conformance would result in a loss of reputation, which would be costly for the individual (Booth, 1985). Naylor (1989) considers the case where individuals' believes about a social custom are heterogeneous, thereby explaining stable equilibria at intermediate union densities. Alternatively, the incorporation of management opposition in Corneo (1995) uses the union's interaction with other institutional agents to explain stable intermediate levels of unionization. All of the economic models above rationalize a minimum level of unionization below which a union loses its ability to effectively provide services. Reduced services would in turn induce more and more members to quit and, at the end of the day, the union would cease to exist.

Complementary approaches in social and political sciences (Wallerstein and Western, 2000, for example) discuss internal rules, class consciousness, social values, political attitudes, etc. Though difficult to measure, these are considered to be highly relevant for union membership. Most of these factors are likely to induce unobserved heterogeneity in the empirical analysis.

4.2.2 Empirical Studies

There are three important strands of literature; see Riley (1997) and Schnabel (2003) for overviews. A first class of studies uses aggregate time series data. In the tradition of Ashenfelter and Pencavel (1969) the analyses focus on long-run trends and business cycle effects. For Germany, Armingeon (1989) analyzes changes in union density in the period 1950–1985. He finds that the stability of gross union density was caused by membership gains in shrinking segments of the labor market on the one hand and stagnation in growing industries on the other. Similarly, Schnabel (1989) studies trade union growth in the period 1955–1986 and links it to changes in price levels and wages, employment, and unemployment. The aggregate evidence stresses the importance of environmental factors.

The second strand of the literature analyzes the impact of institutional regulations and interactions in social environments. Centralization and coordination of collective bargaining, coverage rates, and codetermination are the pillars of an industrial relation system (Hassel, 1999). Interaction between these constituents is closely related to union membership. For example, union membership of a works council increases union density at the firm level since it facilitates union access to the work force (Windolf and Haas, 1989). Similar effects can be expected from a Ghent system of union-managed unemployment insurance. The contact of union officials to insured employees supports recruitment efforts. Frege (1996) emphasizes that the membership decision and the question whether people actually participate in collective action are two separate issues.

She finds that there is no difference in actual behavior of union members between East and West Germany.

The third class of studies uses micro data to model individual membership decisions. At this level, determinants can be grouped into three categories.

- Personal characteristics: Observed socio-demographic variables such as age or marital status, but also attitudes determine an individual's decision to be union member.
- Workplace characteristics: Match-specific, firm-specific, as well as industry-specific effects can facilitate or impede unionization.
- Social environment: The influence of reference groups frames the individual decision.

Studies based on micro-level data for Germany analyze membership determinants along these lines. Windolf and Haas (1989) provide logit estimates based on cross-sectional survey data for the period 1976–1984 and Lorenz and Wagner (1991) use the 1985 wave of the German Socio-Economic Panel (GSOEP). Fitzenberger, Haggeney, and Ernst (1999) and Beck and Fitzenberger (2004) use various GSOEP waves with union membership information in 1985, 1989, 1993, and 1998 (only Beck and Fitzenberger) to estimate West German union membership. Both studies use panel probit estimators and conclude that the propensity for union membership has not changed considerably over time. Hence, the observed aggregate decline in union membership would mainly be driven by composition effects.

Goerke and Pannenberg (2004) also use GSOEP data. They employ fixed effects estimations to back up a theoretical social custom model. Schnabel and Wagner (2005) conduct an analysis based on West German data from the general social survey (ALLBUS), a collection of independent biannual cross sections. Their probit estimates for years between 1980 and 2000 yield no consistent picture of the influence of most variables over time. Applying the same method, Schnabel and Wagner (2003) use the years 1992, 1996, and 2000 of the ALLBUS data to compare determinants in East and West Germany. They conclude that the factors influencing an individual's propensity to be union member have converged between East and West Germany between 1992 and 2000.

4.3 Determinants of Union Membership

Our empirical analysis investigates the following theoretical hypotheses about determinants of union membership; see Schnabel (1993) or Beck and Fitzenberger (2004) for more detailed discussions.

- Age: Mobility tends to decrease with the age of a worker. Family ties and specific human capital increase with age. Thus, older workers are more interested in job security and therefore in union membership as an implicit insurance. Yet the interest in union representation may fade out once people know that they are successful in the labor market. The link between age and union membership may also mirror cohort effects. Differences between generations in value orientation or social custom may result in different attitudes towards unions; see Blanchflower (2006) for an extensive discussion.

- Gender: Compared to males, women are less attached to the labor market and tend to accumulate less specific human capital. Besides, trade union services have traditionally been directed to the needs of male members. Therefore, women are less likely to be union members. The higher rate of female labor force participation in East Germany and its increase in the West should reduce the gender differential.

- Education: Higher education generally implies a higher participation and workplace-related involvement, hence increasing the propensity for unionization. However, higher education is usually associated with higher professional status involving a closer relationship to management, which reduces the desire for a union voice. The latter effect may outweigh the former in particular for employees with a university degree, whereas the former may dominate among workers with vocational training in comparison to less educated workers.

- Marital Status: If both partners of a couple are working, the risk of a job loss is diversified to some extent, which reduces the need for job protection. However, married workers are responsible not only for themselves but also for their family. This would increase the propensity to be a union member, especially with children or when being the only earner.

- Citizenship: Foreigners can be expected to have a weaker attachment to the German labor market and cultural differences might be an obstacle to unionization. Thus, a lower unionization rate among foreigners is likely.

- Political Preference: The historically close relationship between the Social Democratic Party (SPD) and unions suggests that individuals who share values of the Social Democrats are more likely to be union members. There also exists a strong workers' wing within the Christian Democratic Party (CDU), suggesting a somewhat smaller positive effect for the CDU (relative to the omitted category).

- Wage: Membership fees increase with wages and a higher wage tends to be associated with a higher professional status, both of which reduce the propensity to join a union. However, higher wages may indicate higher firm-specific human capital, thus increasing the demand for stability. Similar to education, a hump-shaped relationship may arise with a positive influence for low wage levels and a negative one for higher wages.

In closed shop systems as, e. g., in the United States, union membership itself may result in higher wages; see the broad literature on union wage gaps surveyed in Card, Lemieux, and Riddell (2003). In Germany, however, there are no wage effects of union membership per se at the individual level; compare the discussion in Goerke and Pannenberg (2004).

- Employee Status: Trade unions historically evolved as organizations of blue-collar workers, whose relatively homogenous preferences accommodate unionization efforts. A similar argument applies to civil servants, who share a preference for a stable work environment. Both of these groups are thus more likely to be union members compared to white-collar workers. The latter have moved into the focus of union action only recently with the relative decline of blue-collar employment. With respect to working time, the weaker labor market adherence of part-time workers renders them less likely to be union members than full-time workers.

- Unemployment History: Employees who experienced unemployment in the past might join a union to increase job protection. However, unemployment spells might be associated with a lower attachment to the labor market, thus reducing membership. The overall effect is ambiguous.

- Job Satisfaction: In cooperation with works councils, unions provide a platform of voice and support for dissatisfied workers. They can offer legal advice and financial support in case of a lawsuit between employer and employee. Therefore, membership may be more attractive for dissatisfied workers. However, union intermediation also facilitates communication and understanding between employer and employees which will result in a higher degree of job satisfaction. The overall effect is ambiguous.

- Tenure: With increasing tenure, the worker accumulates more firm specific human capital, which would call for protection. At the same time, an increasing job duration builds up identification with the job, trust, and loyalty towards the employer, thereby decreasing the propensity to unionize. The overall effect is ambiguous.

- Firm Size: The existence of fixed set-up and organizational costs favors union recruitment in larger firms. Larger firms also provide larger subsets of homogeneous workers which accommodate recruitment efforts. Works councils and supervisory boards in large firms support union access to the firm. Large firms show more scope for rent sharing and, therefore, the higher is the relevance of wage bargaining. All of these arguments imply a positive effect.

- Industry: Unions are traditionally pervasive in manufacturing and they are also strong in the public sector, where competition is generally low and high rents exist which can be shared between employees and employers. Private services, however, have less of a union tradition, feature more heterogeneous work forces, and face fierce competition in goods and factor markets as well as rapid structural changes. All of these factors make union recruitment efforts more difficult in private services.

Each of the above factors may influence union membership differently in East and West Germany, and its impact may change over time. In addition, further unobserved individual factors (e. g., social customs) are likely to be of importance.

4.4 Empirical Analysis

4.4.1 Correlated Random Effects Probit Model

We employ a Chamberlain (1980)-Mundlack (1978)-type correlated random effects probit model, which allows us to control for unobserved heterogeneity and to take account of possible correlation of individual- specific effects with observed characteristics. This is central because it is likely that people's attitudes towards unions differ considerably and these attitudes are correlated with observed characteristics.

Let union membership y_{it} of individuals $i = 1, 2, ..., N$ in periods $t = 1, 2, ..., T$ be captured by a binary choice model

$$y_{it} = \begin{cases} 1 & \text{if} \quad y_{it}^* \geq 0 \\ 0 & \text{else} \end{cases} , \tag{4.1}$$

where the latent variable y_{it}^* driving the membership decision of individual i in period t is a linear function of observable characteristics x_{it} and an unobservable individual-specific, time-invariant effect c_i:

$$y_{it}^* = x_{it}\beta + c_i + u_{it}. \tag{4.2}$$

The error term u_{it} is assumed to be normally distributed with unit variance in all periods. The individual-specific effect c_i controls for unobserved heterogeneity in the membership decision. What is more, we consider c_i as a random effect which can be correlated to some variables in x_{it}. In the tradition of Chamberlain (1980) and Mundlack (1978) we assume that c_i is related to the time averages $\bar{x}_{ji}$ of some variables x_{jit}, and that it follows a conditional normal distribution

$$c_i | x_{i1}, ..., x_{iT} \sim N(\mu + \bar{x}_i \xi, \sigma_\epsilon^2), \tag{4.3}$$

where σ_ϵ^2 is the variance of ϵ_i in the regression $c_i = \mu + \bar{x}_i\xi + \epsilon_i$, therefore constituting the conditional variance of c_i. A detailed discussion of this model can be found in Wooldridge (2002).

Given this specification, the model can be written as

$$P(y_{it} = 1|x_{i1}, ..., x_{iT}, c_i) = \Phi\left(\theta(x_{it}\beta + \mu + \bar{x}_i\xi)\right), \tag{4.4}$$

where $\theta = (1 + \sigma_\epsilon^2)^{-1/2}$. As in a standard random effects probit model, the estimation of (4.4) is straightforward. Adding $\bar{x}_i$ is quite intuitive: β estimates the effect of x_{it} on the union participation decision at time t, holding the time average $\bar{x}_i$ fixed. $\bar{x}_i$ contributes to the decision through its effect on the time-invariant individual- specific effect. Note that c_i can only be correlated to averages of time-varying variables, because the effect of the average $\bar{x}_{ji}$ of a time-invariant characteristic x_{jit} could not be discriminated from the direct effect of x_{jit} itself. Furthermore, a constant in x_{it} cannot be distinguished from μ. Details on the empirical model selection are presented in the next section.

4.4.2 Data and Model Selection

We use data of the German Socio-Economic Panel (GSOEP), a longitudinal survey of individuals in private households in the Federal Republic of Germany. The GSOEP started in (West) Germany in the year 1984, and it was extended to East Germany in 1990; see Haisken-DeNew and Frick (2003) for detailed information on the GSOEP. Among others, questions related to the labor market are at the heart of the yearly survey. The question of membership in a trade union, however, is not included in every wave. Six waves contain accordant information for West Germany: 1985, 1989, 1993, 1998, 2001, and 2003. For East Germany, we can use four waves: 1993, 1998, 2001, and 2003. To analyze the determinants of employees' union membership decisions, we focus on individuals in gainful dependent employment who are aged between 16 and 65 years and who earn not more than DM 15,000 per month.[1] Definitions of variables considered in the analysis can be found in tables 4.2 and 4.3 in the appendix. Tables 4.4 and 4.5 report summary statistics for our subsamples of West and East Germany, respectively.

In order to avoid the loss of a large number of observations due to missing values, we add dummy variables for missings in single regressors into the regression equations. In particular, we include dummy variables for missing values in ABITUR, FIRM-SIZE, and SECTOR since these variables

[1] We consider the earnings threshold in order to measure the impact of EARNINGS in the main part of the distribution, which is skewed to the right. Median earnings lie between DM 2,000 (East, 1993) and DM 3,000 (West, 2003) per month, and the 99% quantile varies between DM 5,100 (East, 1993) and DM 10,00 (West, 2003). However, there are outliers with earnings as high as DM 31,400. Applying the earnings threshold, we lose only 22 observations in West Germany and none in East Germany.

contribute most to the problem of missing values. At the same time, some individuals appear in several, but not in all sample periods—due to unemployment spells, for example. We control for this by introducing missing-period dummies. For instance, a vector (1, 1, 0, 1, 0, 1) is assigned to West German individuals observed only in 1993 and 2001—that is, the third and the fifth of the six waves. Furthermore, time dummies and interactions of these with other regressors are included to allow each of the effects to vary between different years.

We estimate several specifications of model (4.4), separately for both West and East Germany. These specifications are as follows:

(A) Selected Model: The estimation of a random effects probit model is computationally involved due to the numerical integration needed. Therefore, we start by applying pooled probit estimations, which are consistent and need significantly less computation time, to select variables for a preferred specification. The resulting specification is then estimated and tested by means of a random effects probit.

More specifically, we first apply a backward selection procedure to select those time-average regressors in $\bar{x}_i$ which are correlated to the individual- specific effect. Starting from a model which includes all x_{jit} as well as averages of all time-varying regressors, we stepwise drop the $\bar{x}_{ji}$ which is least significant, until all remaining averages are significant at the 5% level. At the end of this stage, the list of variables related to the individual- specific effect comprises for West Germany: CHRISTIAN-DEMOCRAT, SOCIAL-DEMOCRAT, WHITE-COLLAR, TRAINEE, UNEMPLOYMENT HISTORY, EARNINGS, FIRM-SIZE, and SECTOR. For East Germany, EARNINGS, TENURE, FIRM-SIZE, and SECTOR turn out to be correlated to the individual- specific effect.

Given the above choice of $\bar{x}_i$, we estimate specifications which include interactions of the regressors x_{it} with year dummies in order to allow for the possibility of time-varying coefficients. Again, effects significant at the 5% level are kept as time-varying. At this stage, the variations of AGE and AGE SQUARED are tested jointly, and so are the variations of EARNINGS and EARNINGS SQUARED as well as those of the SECTOR and the FIRM-SIZE categories.

At the end, we estimate a correlated random effects probit model using the selected variables and test it against a model which again includes averages of all time-varying regressors. Joint significance of the excluded variables is rejected for both West and East Germany.

(B) Reduced Selected Model: Some regressors x_{jit} which are generally time-varying show only limited variation within individuals. For example, an individual's educational attainment rarely changes during his or her working life, and civil servants seldom change back to a private employer. Nevertheless, the averages of these variables might turn out significant in the selected model (A). This could be due to problems of multicollinearity, with the direct effects of x_{jit} becoming insignificant. For this reason, we also estimate a model without averages of educational attainment and vocational status variables.

(C) Benchmark Model: We further estimate a standard random effects probit as a benchmark model. Here we use the same procedure as described above to consider time-varying coefficients, but we do not include any averages $\bar{x}_i$.

(D) IABS Model: Estimates of individual union membership status can be used to predict union densities; compare section 4.5. Accordant predictions can be based on different data sets with larger sample sizes in order to achieve more detailed and more precise predictions.[2] In order to facilitate predictions for individuals in the IAB employment sample (*IAB Beschäftigtenstichprobe*, IABS) 1975–2001, we estimate an additional specification, using only those explanatory variables which are available in the IABS. This specification excludes the variables CHRISTIAN-DEMOCRAT, SOCIAL-DEMOCRAT, SATISFACTION, TENURE, and MARRIED. The estimation for East Germany further excludes UNEMPLOYMENT HISTORY because the IABS offers accounts for individuals in East Germany from 1992 onward only. We also incorporate the fact that the IABS distinguishes between different vocational statuses only among full-timers. In all other respects the selection process is the same as in (A).

(E) GSES Model: For the same purpose as in (D), we also estimate a specification which includes only variables available in the German Structure of Earnings Survey (GSES, *Gehalts- und Lohnstrukturerhebung*) 2001. We thus exclude CHRISTIAN-DEMOCRAT, SOCIAL-DEMOCRAT, SATISFACTION, UNEMPLOYMENT HISTORY, FOREIGNER, and MARRIED. Since the GSES is a cross-sectional data set, this specification is estimated without averages $\bar{x}_i$.

4.4.3 Estimation Results

Estimated coefficients for West and East Germany are reported in tables 4.6 and 4.7 in the appendix. Note that there are only four specifications for East Germany because the selected and the reduced selected specification coincide for the East. In the following, we compare the different models and then turn to our preferred models in more detail.

[2] Most large micro-level data sets for Germany—such as the administrative IAB employment sample or the Structure of Earnings Survey carried out by the German Statistical Office—provide no information on union membership. In order to take advantage of the big sample sizes of these data sets membership propensities thus have to be imputed.

Comparing first the correlated random effects models (A) and (B) to the benchmark models (C), we find significance of several elements of ξ for both parts of the country. Individual-specific effects are in fact correlated to averages $\bar{x}_i$ of some observed characteristics, for which the effects in the benchmark model are quite similar to the joint impact of the direct effect and the indirect effect through c_i in the models (A) and (B). Yet the economic reasoning behind the estimated determinants is more subtle in the correlated models. The latter do not only take account of direct impacts on the membership decision, but also allow for the possibility that some determinants are correlated with unobserved individual-specific attitudes towards unions, which again drive the membership decision. Given this richer interpretation and the statistical significance of ξ, we prefer the models (A) and (B) to the benchmark ones (C). There are only small differences of coefficients β_j for those variables x_{jit} whose time-averages are not included.

When comparing the selected specification (A) to the reduced selected model (B), we also find very similar effects. In fact, the direct effect of being a TRAINEE is even insignificant if the corresponding average is included. This finding suggests multicollinearity between the regressor and its average. Predictions based on either specification should not differ fundamentally, though. Both specifications yield very similar percentages of correct predictions. Since, above all, the true channel through which the determinants work is not clear a priori, we prefer the statistically validated specification (A).

The IABS models (D) and the GSES models (E) include only regressors that are also available in the respective target data sets. Estimated coefficients for the included variables are generally very similar to the results obtained from the full models, and predictive power is also comparable. More specifically, recalling that the GSES is a cross-sectional data set, the coefficients of specifications (E) match those from the benchmark models (C). The coefficients of the IABS models (D) match the results of specifications (A).

We now describe the preferred specifications (A) in more detail. Results for East and for West Germany are remarkably similar despite some notable exceptions. First, the baseline (TIME dummies) and the impacts of EARNINGS and UNEMPLOYMENT HISTORY are the only effects which vary significantly across time in the East, whereas some more effects vary in the West. On the one hand this is to be expected given the longer sample period for West Germany. On the other hand, East-West convergence (Schnabel and Wagner, 2003) is likely to be driven by changes in the East. Second, while MARRIED individuals ceteris paribus have a lower propensity to be a union member in the West, the respective effect is significantly positive in the East. This finding likely reflects East-West differences in labor force participation. Third, working PART-TIME has the expected negative effect in West Germany, but it has a positive but insignificant effect in East Germany. Fourth, differences between sectors are stronger in East Germany, and most direct SECTOR effects are insignificant, possibly due to the relatively small number of observations in some sectors (compare table 4.5) and to less within-individual variation.

The coefficients are generally allowed to vary over time. However, most of the effects do not change significantly. Those which do change mainly show a consistent, monotonic pattern. For instance, both the linear effect of AGE and the curvature effect of AGE SQUARED decrease in West Germany over some time, rendering the total impact less concave. In East Germany, the impact of EARNINGS becomes also less concave. Thus, in contrast to Beck and Fitzenberger (2004) and Schnabel and Wagner (2005), we find some clear patterns of changes. For East Germany, we find a significant positive time effect only for 1993, whereas for West Germany there is a negative time trend throughout the entire sample period. Therefore, the estimated time trend contributes to the continuous deunionization in West Germany but not in East Germany.

Turning to the effects of the covariates, we can confirm most of the hypotheses in section 4.3. Women are less likely to be union members. The effect of FOREIGNER is positive but not significant in East Germany. We further find that a positive, concave impact of AGE. As expected, supporters of the Social Democrats (but not those of the Christian Democrats) have a higher propensity to join a union. Regarding education, ABITUR and UNIVERSITY have a sizeable negative impact, but the influence of APPRENTICESHIP is not as clear. In contrast to SEMISKILLED and—even more substantially—SKILLED BLUE-COLLAR workers, CIVIL SERVANTS and WHITE-COLLAR workers show a significantly lower propensity. The effect for individuals working PART-TIME has the expected negative sign in West Germany, but it is positive and insignificant in the East. The effect of UNEMPLOYMENT HISTORY is negative. However, the effect of average UNEMPLOYMENT HISTORY is strongly positive. Employees who have recently been unemployed are less likely to be union member due to their lower labor market attachment, whereas employees who are generally at a higher risk of unemployment have a higher need for protection. Job SATISFACTION shows virtually no effect in the West and only a limited one in the East.

The concave effect of EARNINGS generally meets our expectations. As discussed above, the impact becomes flatter over time in East Germany. The EARNINGS effect is more sizeable in East Germany, being attenuated by converse effects through average EARNINGS and average EARNINGS SQUARED. The positive but small TENURE coefficient supports the human capital argument. In contrast, FIRM-SIZE shows a substantial positive impact. However, the differences between firms with more than 200 employees and ones with more than 2000 are negligible. Finally, considerable differences in unionization exist between industries. Compared to our reference SECTOR "Miscellaneous Manufacturing (7)", the large positive effects of "Chemical Products (5)" and the formerly public industries "Transport and Communication (11)" are most striking.

In contrast, “Hotels and Restaurants (10)”, “Financial Intermediation (12)”, or “Other Services (16)” show significantly lower union membership.

4.5 Prediction and Decomposition Analysis

Based on our preferred models, we predict propensities to be a union member for each of the individuals in our samples. These propensities can be averaged to an estimator for net union density. More specifically, we estimate net union density NUD_{rt} separately for regions $r \in \{\text{East, West}\}$ in each year t by

$$\widehat{NUD}_{rt} = N_{rt}^{-1} \sum_{i=1}^{N_{rt}} \Phi\left(\hat{\theta}^{rt}(x_{irt}\hat{\beta}^{rt} + \hat{\mu}^{rt} + \bar{x}_{ir}\hat{\xi}^{rt})\right). \tag{4.5}$$

The observed as well as the predicted net union densities are depicted in figure 4.2. In general, the predicted densities match the observed frequencies fairly well. Compared to the West, membership in East Germany started out at a higher level in the year 1993, but exhibited a stronger decline afterwards. NUD for 1993 and 2003 were 38% and 19% in East and 27% and 21% in West Germany.[3] Aggregate NUD is about 10 percentage points (pp) lower than gross union density (compare section 4.2).

We now investigate (1) the changes of NUD over time and (2) the differences in NUD between East and West Germany by means of Blinder (1973) and Oaxaca (1973) decomposition techniques, which we adapt to the nonlinear probit framework. To decompose the changes of NUD within the two regions between 1993 and 2003,[4] we write

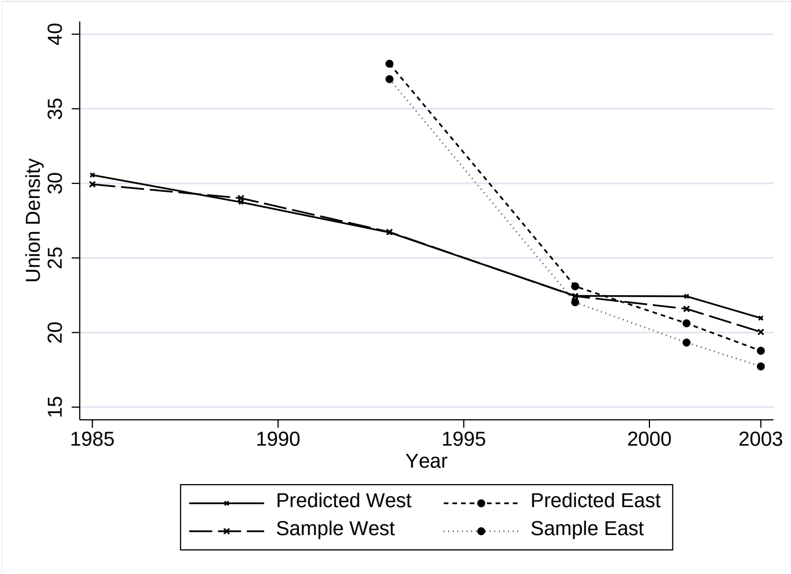
$$\begin{aligned}
\widehat{NUD}_{2003} - \widehat{NUD}_{1993} &= \underbrace{(\widehat{NUD}_{2003} - \widehat{NUD}^{1993}_{2003})}_{\text{coefficients effect}} + \underbrace{(\widehat{NUD}^{1993}_{2003} - \widehat{NUD}_{1993})}_{\text{characteristics effect}} &(4.6)\\
&= \underbrace{(\widehat{NUD}_{2003} - \widehat{NUD}^{2003}_{1993})}_{\text{characteristics effect}} + \underbrace{(\widehat{NUD}^{2003}_{1993} - \widehat{NUD}_{1993})}_{\text{coefficients effect}}, &(4.7)
\end{aligned}$$

where $\widehat{NUD}_t$ are estimated as described in equation (4.5). The decompositions (4.6) and (4.7) differ with respect to the chosen counterfactual densities $\widehat{NUD}^{\tilde{t}}_t$. In equation (4.6), $\widehat{NUD}^{1993}_{2003}$

[3] Estimates of individual union membership can generally be used to predict densities for more narrowly defined labor market segments. Lorenz and Wagner (1991), Fitzenberger, Haggeney, and Ernst (1999), and Beck and Fitzenberger (2004), for instance, predict union densities for two-digit industries. In particular, the latter studies impute membership propensities for individuals in the IAB employment sample, taking advantage of the bigger sample size of this data set to achieve more detailed and more precise NUD predictions.

[4] It would also have been possible to analyze the change for West Germany over the even longer period 1985–2003. However, we opt for 1993–2003 in order to facilitate East-West comparisons in table 4.1.

Figure 4.2: Net Union Density in East and West Germany



Sample frequencies and predicted densities in percent. Data source: GSOEP.

denotes the prediction for individuals in the year 2003, assuming that the coefficients stayed as in 1993. In equation (4.7), $\widehat{NUD}_{1993}^{2003}$ uses predictions for individuals in 1993 based on the coefficients of 2003. To investigate the sensitivity of the decomposition result, we compute both versions (4.6) and (4.7).[5] The characteristics effect involves the part of the overall change between 1993 and 2003 which can be attributed to changes in personal, workplace, and social characteristics of the individuals in the sample at given coefficients. The coefficients effect captures the part which is due to changes in the coefficients at given characteristics. The necessary counterfactuals can be estimated as averages analogous to equation (4.5).

For the differences between East (E) and West (W) Germany in any given year, we use

$$\widehat{NUD}_W - \widehat{NUD}_E = \underbrace{(\widehat{NUD}_W - \widehat{NUD}_W^E)}_{\text{coefficients effect}} + \underbrace{(\widehat{NUD}_W^E - \widehat{NUD}_E)}_{\text{characteristics effect}} \quad (4.8)$$

$$= \underbrace{(\widehat{NUD}_W - \widehat{NUD}_E^W)}_{\text{characteristics effect}} + \underbrace{(\widehat{NUD}_E^W - \widehat{NUD}_E)}_{\text{coefficients effect}}, \quad (4.9)$$

where the involved counterfactual densities $\widehat{NUD}_r^{\tilde{r}}$ are defined as above.

[5] It is well-known that the decompositions resulting from the different counterfactuals do not necessarily yield identical results. Different approaches to the issue of non-uniqueness have been proposed in the literature; see Oaxaca and Ransom (1994) and Silber and Weber (1999) for surveys. Yet each of the approaches relies on ad-hoc assumptions of some type, so we choose to report the two most prominent cases.

The results of the different decompositions are reported in table 4.1. Standard errors to assess the accuracy of the decompositions are obtained by means of a parametric bootstrap by resampling from the estimated distribution of the parameters $\left(\beta^{rt\,\prime}, \mu^{rt}, \xi^{rt\,\prime}, \sigma_{\epsilon}^{rt}\right)^{\prime}$.

Table 4.1: Differences in Net Union Density: Decomposition Analyses

	Net Union Density [%] 1993	Net Union Density [%] 2003	Change	Char. Effect[a]		Coeff. Effect[a]	
West	26.71	20.97	-5.74	-1.81	(0.58)	-3.93	(0.80)
Germany	(0.57)	(0.61)	(0.71)	*-0.4*	*(0.55)*	*-5.34*	*(0.82)*
East	38.03	18.78	-19.25	-3.56	(1.84)	-15.69	(1.70)
Germany	(1.67)	(1.22)	(1.12)	*-0.11*	*(1.64)*	*-19.14*	*(2.18)*

	1993		2003	
Difference	-11.32		2.19	
	(1.75)		(1.35)	
Char. Effect[b]	4.7	*7.94*	6.03	*6.37*
	(4.27)	*(0.66)*	(2.77)	*(0.56)*
Coeff. Effect[b]	-16.02	*-19.26*	-3.84	*-4.18*
	(4.13)	*(1.80)*	(2.80)	*(1.44)*

[a] Counterfactual with characteristics of 2003/*1993* in normal/*italic* font.
[b] Counterfactual with West/*East* characteristics in normal/*italic* font.
Standard errors in parentheses estimated by 1000 bootstrap resamples. Data source: GSOEP.

The results are not very sensitive to the choice of counterfactuals in (4.6) and (4.7), nor in (4.8) and (4.9). Interpreting first the (horizontal) decompositions of the changes in NUD over time, both characteristics and coefficients effects contribute to the observed deunionization. However, the coefficients effect dominates in almost all cases. The characteristics effect does not explain more than a third of the 6 pp decline in West Germany and does not even account for a fifth of the 19 pp decline in the East. This result is in contrast to the finding in Beck and Fitzenberger (2004) that the decline in union density in West Germany between the 1980's and 1990's was mainly driven by changes in the composition of the work force.[6] The small impact of the characteristics effect in East Germany is quite remarkable in light of the structural change during the 1990's; compare the summary statistics in tables 4.4 and 4.5. The strong coefficients effect involves the negative time trend as well as changing impacts of particular characteristics.

Regarding the (vertical) East-West comparison, the characteristics effects and the coefficients effects generally work in opposite directions. The characteristics effect is in favor of a higher density in West Germany by 5 to 7 pp: The composition of the West German work force exhibits more attributes supporting union membership. Thus, the higher NUD in East Germany in 1993 resulted from differences in the coefficients in the order of 16 to 19 pp: For given characteristics,

[6] Note that Beck and Fitzenberger (2004) do not apply decomposition techniques.

East Germans were more strongly unionized than West Germans. This finding suggests a lower quality of union membership matches in East Germany resulting from the widespread, arbitrary recruitment after unification. A stronger decline in union membership thus comes as no surprise. Ten years later, in 2003, union density in East Germany is already 2 pp smaller than in West Germany. The composition of the work force still being more in favor of union membership in the West, the coefficients in the two parts of the country have become more similar such that the—still negative—coefficients effect has lost its bite.

4.6 Conclusions

The importance of unions in the German labor market is undisputed. However, the question why people join a union is anything but beyond dispute. This study uses detailed micro-panel data to provide insights into the determinants of individual union membership. We use the German Socio-Economic Panel (GSOEP) to estimate membership equations for West (1985–2003) and for East Germany (1993–2003). The application of a Chamberlain (1980)-Mundlack (1978)-type correlated random effects probit model controls for unobserved heterogeneity and allows for a correlation between individual- specific effects and observed characteristics.

Our findings quantify the influence of socio-demographic personal characteristics, such as age or marital status; the influence of workplace characteristics, i. e., match, firm, or industry specific effects; and the influence of attitudinal factors for the individual choice to be or not to be a union member. The membership equations are allowed to differ between East and West Germany and over time.

Projections of net union densities (NUD) based on our estimates consistently trace the trends towards deunionization in both parts of the country. Compared to the West, membership in East Germany started out from a higher level at the beginning of the 1990's, but exhibited a stronger decline afterwards. By the year 2003, NUD was even lower in East Germany than in the West.

Decomposition analyses shed light on (1) the changes in unionization over time and (2) on the differences in NUD between the two regions. Changes in the composition of the work force do in no case explain more than one third of the observed decline in NUD over time. In East-West comparison, the West German work force exhibits more attributes supporting union membership. The higher union density in East Germany in the year 1993 and the stronger subsequent decline reflect a lower quality of membership matches resulting from the widespread, arbitrary membership recruitment after unification.

The erosion of union membership in Germany is likely to weaken the bargaining power of unions and therefore the unions' impact on the labor market (Fitzenberger and Kohn, 2005). Despite the still high coverage of collective agreements (especially in West Germany), the results of wage bargaining are likely to deteriorate from the perspective of union members—but possibly result in higher employment. We plan to explore the link between union membership, wages, and employment in future research, for which the results of this study provide a necessary input.

4.A Additional Tables

Table 4.2: Description of Variables

Dummy Variables	= 1 if true
MEMBER	being a union member
FEMALE	being female
MARRIED	being married
FOREIGNER	being a foreigner
Education:	
ABITUR	*Abitur* is the highest educational attainment
APPRENTICESHIP	apprenticeship or a similar vocational training is the highest professional degree
UNIVERSITY	person has obtained a technical college or a university degree
Political Orientation:	
CHRISTIAN-DEMOCRAT	person feels close to the Christian Democratic Party
SOCIAL-DEMOCRAT	person feels close to the Social Democratic Party
Vocational Status:	
PART-TIME	working part-time
SEMISKILL-BLUE	being an unskilled or a semi-skilled blue-collar worker
SKILL-BLUE	being a skilled blue-collar worker
WHITE-COLLAR	being a white-collar worker
CIVIL SERVICE	being employed in the civil service
TRAINEE	being currently in professional training
UNEMPLOYMENT HISTORY	person has been unemployed at least once during past 5 years (10 years for 1985 wave)
Firm Size:	
FIRM-SIZE19	firm has less than 20 employees
FIRM-SIZE199	firm has 20–199 employees
FIRM-SIZE1999	firm has 200–1999 employees
FIRM-SIZE_MORE	firm has more than 1999 employees
SECTORj:	working in sector j[a]
MISSINGt:	person is not observed in year t
TIMEt:	observation is in year t
Other Variables	
AGE	age of person divided by 10
EARNINGS	total earnings last month in thousands of DM, at constant prices (1985 = 100)
TENURE	duration of employment in the current firm, in years
SATISFACTION	satisfaction of the worker with her/his job, scaled from 0 (not satisfied) to 10 (very satisfied)

[a] See table 4.3 for the industry classification and grouping of sectors.

Table 4.3: NACE Industry Classification in the GSOEP and Grouping Used in our Empirical Analysis

No.[a]	Industry	NACE[b]
01	Agriculture, Forestry, and Fishing; Mining; Energy and Water Supply	01–14, 40–41
02	Manufacture of Food, Beverages, and Tobacco	15–16
03	Textiles	17–19
04	Woodwork, Paper, Printing, Publishing	20–22
05	Chemical Products	23–26
06	Manufacture of Iron, Steel, Metal; Machinery; Vehicles	27–29, 34–35
07	Other Manufacturing; Recycling	30–33, 36–37, 96–97, 100
08	Construction	45
09	Trade	50–52
10	Hotels and Restaurants	55
11	Transport and Communication	60–64
12	Financial Intermediation	65–67
13	Education; Research	73, 80
14	Health Care System and Social Work	85
15	Public Administration and Defence, Social Security	75
16	Other Services	70–74, 90–95, 98–99

[a] Sector classification used in the empirical analysis.
[b] GSOEP industry classification based on 2-digit NACE.

Table 4.4: Summary Statistics, West Germany

Variable	1985	1989	1993	1998	2001	2003
MEMBER	29.94	29.01	26.74	22.44	21.59	20.04
FEMALE	38.04	38.86	41.67	42.52	44.27	45.32
MARRIED	66.88	63.62	63.36	61.12	62.04	61.19
FOREIGNER	29.45	28.84	28.03	21.25	19.36	17.62
APPRENTICESHIP	59.09	60.42	59.51	64.57	63.51	64.15
ABITUR	8.90	9.73	11.45	15.48	17.36	18.29
MISSING_ABITUR	0.20	0.32	0.70	0.93	1.74	2.45
UNIVERSITY	8.80	8.84	9.73	13.23	15.39	15.88
CHR-DEM	11.50	8.62	10.06	9.54	10.01	12.83
SOC-DEM	30.15	29.75	21.09	24.20	24.10	21.18
PART-TIME	11.05	10.85	13.40	14.93	16.52	17.53
SEMISKILL-BLUE	29.52	30.03	26.60	21.57	20.46	18.51
SKILL-BLUE	20.68	18.29	18.10	16.20	16.41	15.24
WHITE-COLLAR	34.94	37.64	41.48	48.87	50.78	53.89
CIVIL SERVICE	8.26	7.67	7.67	7.86	7.29	7.59
TRAINEE	6.59	6.38	6.15	5.50	5.03	4.76
UNEMPL_HIST	18.47	9.71	7.84	9.91	9.84	7.24
FIRM-SIZE19	18.55	20.11	21.05	23.05	21.85	21.63
FIRM-SIZE199	26.96	26.95	26.01	26.03	27.11	25.63
FIRM-SIZE1999	22.30	25.20	25.46	25.76	23.32	22.99
FIRM-SIZE_MORE	27.88	27.48	27.17	24.99	23.55	24.04
MISSING_FIRM-SIZE	4.30	0.25	0.31	0.16	4.17	5.72
SECTOR01	2.11	2.25	2.46	2.34	2.00	1.97
SECTOR02	3.01	3.12	2.44	2.31	2.08	1.87
SECTOR03	3.05	3.09	2.31	1.22	0.95	0.83
SECTOR04	2.64	2.78	2.79	2.66	2.52	2.41
SECTOR05	5.44	6.25	6.61	6.14	5.15	4.92
SECTOR06	16.90	19.22	16.56	13.86	14.38	12.73
SECTOR07	6.16	7.46	7.03	8.79	6.68	6.38
SECTOR08	8.28	7.65	7.64	5.50	5.87	5.34
SECTOR09	7.75	9.15	11.18	12.88	12.67	11.34
SECTOR10	1.68	1.82	2.13	1.54	2.05	1.84
SECTOR11	4.81	4.53	5.03	4.62	4.75	4.48
SECTOR12	2.76	3.58	3.98	4.67	4.22	4.73
SECTOR13	4.05	4.05	3.93	4.38	4.57	4.99
SECTOR14	5.67	6.55	8.15	10.12	9.40	10.00
SECTOR15	8.16	8.39	8.24	8.02	8.04	8.45
SECTOR16	6.40	7.63	7.67	9.91	11.20	11.34
MISSING_SECTOR	11.11	2.48	1.85	1.04	3.44	6.38
AGE	3.76	3.74	3.78	3.84	3.92	3.98
EARNINGS	2.58	2.82	2.95	3.11	3.12	3.25
SATISFACTION	7.40	7.23	7.26	7.24	7.25	7.07
TENURE	9.76	10.16	9.84	10.17	10.15	10.65
N. of Obs.	5111	4719	4552	3765	3456	3149

Mean values of variables.
See text for details on the selected sample.
Data source: GSOEP.

Table 4.5: Summary Statistics, East Germany

Variable	1993	1998	2001	2003
MEMBER	36.99	22.03	19.33	17.73
FEMALE	47.03	47.13	49.19	49.56
MARRIED	70.47	63.13	60.36	57.56
FOREIGNER	0.20	0.17	0.17	0.19
APPRENTICESHIP	75.37	74.79	71.40	72.11
ABITUR	16.17	17.96	21.35	23.00
MISSING_ABITUR	0.30	1.06	1.32	1.78
UNIVERSITY	28.14	27.22	27.79	27.89
CHR-DEM	7.91	7.75	10.64	14.74
SOC-DEM	10.48	9.98	11.05	9.97
PART-TIME	7.22	9.48	12.14	14.55
SEMISKILL-BLUE	12.76	12.10	12.43	11.25
SKILL-BLUE	30.02	26.77	25.43	24.78
WHITE-COLLAR	48.22	49.25	49.71	52.16
CIVIL SERVICE	1.68	3.51	4.32	4.45
TRAINEE	7.32	8.31	8.06	7.37
UNEMPL_HIST	12.17	19.91	19.85	15.44
FIRM-SIZE19	24.48	29.50	27.10	24.71
FIRM-SIZE199	33.68	34.30	33.03	31.39
FIRM-SIZE1999	22.45	19.07	18.12	19.63
FIRM-SIZE_MORE	18.74	16.68	16.46	16.96
MISSING_FIRM-SIZE	0.64	0.45	5.29	7.31
SECTOR01	7.42	4.91	4.09	3.56
SECTOR02	1.53	1.56	1.61	1.65
SECTOR03	0.89	0.89	1.09	1.21
SECTOR04	1.73	1.73	1.61	1.72
SECTOR05	3.51	3.07	2.47	2.67
SECTOR06	8.51	7.53	8.29	7.56
SECTOR07	3.96	6.41	4.55	4.76
SECTOR08	12.41	10.93	9.72	7.12
SECTOR09	11.67	12.16	13.23	11.94
SECTOR10	1.68	1.84	2.59	2.48
SECTOR11	7.67	5.41	5.06	4.83
SECTOR12	2.52	2.96	2.99	3.30
SECTOR13	7.17	5.19	6.67	6.16
SECTOR14	8.85	11.82	11.45	10.67
SECTOR15	11.18	10.82	9.32	9.97
SECTOR16	7.52	11.43	10.99	10.86
MISSING_SECTOR	1.78	1.34	4.26	9.53
AGE	3.72	3.83	3.89	3.94
EARNINGS	2.07	2.44	2.45	2.59
SATISFACTION	6.42	6.72	6.76	6.55
TENURE	6.91	7.27	7.54	8.30
N. of Obs.	2022	1793	1738	1574

Mean values of variables.
See text for details on the selected sample.
Data source: GSOEP.

Table 4.6: Determinants of Union Membership, West Germany

	Specification									
	(A)		(B)		(C)		(D)		(E)	
	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)
FEMALE	-0.263 ***	(0.077)	-0.405 ***	(0.074)	-0.435 ***	(0.064)	-0.248 ***	(0.076)	-0.443 ***	(0.064)
FOREIGNER	-0.007	(0.070)	0.051	(0.068)	0.063	(0.064)	-0.028	(0.071)		
AGE_1985	-0.032	(0.249)	-0.095	(0.246)			0.219	(0.243)		
AGE_1989	0.380	(0.262)	0.301	(0.259)			0.482 *	(0.259)		
AGE_1993	0.725 ***	(0.278)	0.573 **	(0.274)	0.501 ***	(0.134)	0.964 ***	(0.279)	0.591 ***	(0.136)
AGE_1998	1.039 ***	(0.325)	0.790 **	(0.320)			1.254 ***	(0.330)		
AGE_2001	1.187 ***	(0.342)	0.893 ***	(0.335)			1.359 ***	(0.349)		
AGE_2003	1.368 ***	(0.386)	1.032 ***	(0.379)			1.521 ***	(0.391)		
AGE_SQU_1985	0.013	(0.032)	0.021	(0.032)			-0.004	(0.031)		
AGE_SQU_1989	-0.034	(0.033)	-0.025	(0.033)			-0.029	(0.033)		
AGE_SQU_1993	-0.085 **	(0.034)	-0.069 **	(0.034)	-0.054 ***	(0.016)	-0.095 ***	(0.034)	-0.065 ***	(0.016)
AGE_SQU_1998	-0.118 ***	(0.039)	-0.091 **	(0.038)			-0.125 ***	(0.039)		
AGE_SQU_2001	-0.125 ***	(0.040)	-0.095 **	(0.039)			-0.126 ***	(0.040)		
AGE_SQU_2003	-0.138 ***	(0.044)	-0.105 **	(0.044)			-0.137 ***	(0.045)		
MARRIED	-0.060	(0.052)	-0.051	(0.051)	-0.095 *	(0.049)				
CHR-DEM	-0.048	(0.092)	-0.045	(0.092)	-0.260 ***	(0.073)				
SOC-DEM	0.110 **	(0.055)	0.114 **	(0.055)	0.392 ***	(0.044)				
ABITUR	-0.387 ***	(0.119)	-0.399 ***	(0.114)	-0.432 ***	(0.102)	-0.409 ***	(0.112)	-0.417 ***	(0.104)
APPRENTICESHIP	0.041	(0.057)	0.030	(0.057)	0.038	(0.054)	0.041	(0.057)	0.025	(0.052)
UNIVERSITY	-0.265 **	(0.110)	-0.261 **	(0.108)	-0.311 ***	(0.103)	-0.348 ***	(0.104)	-0.307 ***	(0.103)
SKILL-BLUE	0.100	(0.063)	0.102	(0.063)	0.094	(0.061)	0.098	(0.064)	0.102 *	(0.060)
WHITE-COLLAR	-0.178 **	(0.086)	-0.490 ***	(0.070)	-0.526 ***	(0.066)	-0.249 ***	(0.083)	-0.538 ***	(0.064)
CIVIL SERVICE	-0.276 **	(0.139)	-0.142	(0.134)	0.015	(0.122)	-0.344 ***	(0.129)	-0.006	(0.122)
TRAINEE	-0.083	(0.138)	0.058	(0.120)	0.234 **	(0.115)	-0.105	(0.139)	0.227 **	(0.113)
PART-TIME_1985						(0.080)	-0.593 ***	(0.162)		
PART-TIME_1989						(0.080)	-0.118	(0.156)		
PART-TIME_1993	-0.042	(0.084)	-0.070	(0.083)	-0.101	(0.080)	-0.576 ***	(0.159)	-0.115	(0.080)
PART-TIME_1998						(0.080)	-0.078	(0.158)		
PART-TIME_2001						(0.080)	0.031	(0.156)		
PART-TIME_2003						(0.080)	0.133	(0.158)		
UNEMPL_HIST_1985							-0.328 ***	(0.097)		
UNEMPL_HIST_1989							0.108	(0.126)		

Continued on next page...

... table 4.6 continued

	Specification									
	(A)		(B)		(C)		(D)		(E)	
	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)
UNEMPL_HIST_1993	-0.116 *	(0.062)	-0.107 *	(0.062)	-0.025	(0.057)	0.054	(0.139)		
UNEMPL_HIST_1998							-0.330 **	(0.156)		
UNEMPL_HIST_2001							-0.241	(0.165)		
UNEMPL_HIST_2003							-0.300	(0.211)		
SATISFACTION	-0.007	(0.011)	-0.007	(0.011)	-0.019 *	(0.010)				
EARNINGS_1985					0.741 ***	(0.101)			0.767 ***	(0.102)
EARNINGS_1989					0.621 ***	(0.100)			0.622 ***	(0.102)
EARNINGS_1993	0.320 ***	(0.068)	0.387 ***	(0.067)	0.562 ***	(0.095)	0.327 ***	(0.067)	0.578 ***	(0.096)
EARNINGS_1998					0.482 ***	(0.099)			0.486 ***	(0.100)
EARNINGS_2001					0.469 ***	(0.094)			0.432 ***	(0.094)
EARNINGS_2003					0.464 ***	(0.097)			0.415 ***	(0.099)
EARNINGS_SQU_1985					-0.102 ***	(0.014)			-0.107 ***	(0.014)
EARNINGS_SQU_1989					-0.085 ***	(0.014)			-0.086 ***	(0.014)
EARNINGS_SQU_1993	-0.039 ***	(0.008)	-0.044 ***	(0.007)	-0.067 ***	(0.012)	-0.037 ***	(0.007)	-0.068 ***	(0.011)
EARNINGS_SQU_1998					-0.055 ***	(0.012)			-0.055 ***	(0.012)
EARNINGS_SQU_2001					-0.058 ***	(0.011)			-0.054 ***	(0.011)
EARNINGS_SQU_2003					-0.056 ***	(0.011)			-0.052 ***	(0.011)
TENURE_1985									0.023 ***	(0.005)
TENURE_1989									0.022 ***	(0.005)
TENURE_1993	0.025 ***	(0.003)	0.025 ***	(0.003)	0.028 ***	(0.003)			0.018 ***	(0.005)
TENURE_1998									0.026 ***	(0.005)
TENURE_2001									0.035 ***	(0.005)
TENURE_2003									0.038 ***	(0.005)
FIRM-SIZE199_1985	0.716 ***	(0.133)	0.695 ***	(0.132)	0.846 ***	(0.124)	0.659 ***	(0.134)	0.814 ***	(0.123)
FIRM-SIZE199_1989	0.705 ***	(0.140)	0.692 ***	(0.139)	0.903 ***	(0.130)	0.716 ***	(0.140)	0.937 ***	(0.130)
FIRM-SIZE199_1993	0.618 ***	(0.142)	0.602 ***	(0.142)	0.801 ***	(0.132)	0.630 ***	(0.143)	0.813 ***	(0.132)
FIRM-SIZE199_1998	0.731 ***	(0.157)	0.718 ***	(0.156)	0.956 ***	(0.146)	0.762 ***	(0.157)	0.940 ***	(0.146)
FIRM-SIZE199_2001	0.469 ***	(0.157)	0.473 ***	(0.156)	0.734 ***	(0.148)	0.530 ***	(0.158)	0.725 ***	(0.148)
FIRM-SIZE199_2003	0.302 *	(0.169)	0.303 *	(0.169)	0.599 ***	(0.158)	0.347 **	(0.171)	0.587 ***	(0.159)
FIRM-SIZE1999_1985	1.374 ***	(0.137)	1.347 ***	(0.136)	1.595 ***	(0.127)	1.334 ***	(0.138)	1.584 ***	(0.126)
FIRM-SIZE1999_1989	0.978 ***	(0.141)	0.963 ***	(0.140)	1.291 ***	(0.130)	1.016 ***	(0.141)	1.330 ***	(0.130)
FIRM-SIZE1999_1993	0.903 ***	(0.142)	0.891 ***	(0.141)	1.249 ***	(0.131)	0.947 ***	(0.142)	1.287 ***	(0.130)
FIRM-SIZE1999_1998	0.800 ***	(0.160)	0.792 ***	(0.158)	1.192 ***	(0.148)	0.873 ***	(0.160)	1.182 ***	(0.149)

Continued on next page...

... table 4.6 continued

	Specification									
	(A)		(B)		(C)		(D)		(E)	
	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)
FIRM-SIZE1999_2001	0.710 ***	(0.160)	0.702 ***	(0.159)	1.170 ***	(0.151)	0.827 ***	(0.161)	1.136 ***	(0.152)
FIRM-SIZE1999_2003	0.735 ***	(0.170)	0.712 ***	(0.170)	1.221 ***	(0.161)	0.857 ***	(0.173)	1.185 ***	(0.162)
FIRM-SIZE_MORE_1985	1.378 ***	(0.135)	1.356 ***	(0.134)	1.729 ***	(0.125)	1.332 ***	(0.136)	1.724 ***	(0.125)
FIRM-SIZE_MORE_1989	1.138 ***	(0.142)	1.125 ***	(0.141)	1.616 ***	(0.131)	1.172 ***	(0.142)	1.646 ***	(0.132)
FIRM-SIZE_MORE_1993	0.937 ***	(0.143)	0.925 ***	(0.143)	1.423 ***	(0.133)	1.003 ***	(0.144)	1.472 ***	(0.133)
FIRM-SIZE_MORE_1998	0.873 ***	(0.160)	0.847 ***	(0.159)	1.383 ***	(0.150)	0.962 ***	(0.161)	1.395 ***	(0.151)
FIRM-SIZE_MORE_2001	0.819 ***	(0.162)	0.807 ***	(0.161)	1.404 ***	(0.154)	0.942 ***	(0.164)	1.391 ***	(0.155)
FIRM-SIZE_MORE_2003	0.705 ***	(0.172)	0.691 ***	(0.171)	1.327 ***	(0.162)	0.855 ***	(0.174)	1.313 ***	(0.164)
SECTOR1	0.215	(0.191)	0.216	(0.191)	0.497 ***	(0.143)	0.257	(0.189)	0.484 ***	(0.145)
SECTOR2	0.110	(0.176)	0.108	(0.177)	0.032	(0.139)	0.077	(0.180)	0.027	(0.139)
SECTOR3	0.212	(0.188)	0.202	(0.188)	0.315 **	(0.151)	0.247	(0.191)	0.281 *	(0.159)
SECTOR4	-0.189	(0.174)	-0.186	(0.174)	0.061	(0.142)	-0.175	(0.174)	0.025	(0.142)
SECTOR5	0.302 **	(0.129)	0.289 **	(0.129)	0.331 ***	(0.102)	0.303 **	(0.129)	0.330 ***	(0.100)
SECTOR6	0.223 **	(0.095)	0.227 **	(0.095)	0.476 ***	(0.080)	0.209 **	(0.095)	0.469 ***	(0.079)
SECTOR8	-0.342 **	(0.134)	-0.354 ***	(0.133)	-0.426 ***	(0.106)	-0.346 **	(0.137)	-0.419 ***	(0.105)
SECTOR9	-0.246 **	(0.120)	-0.225 *	(0.120)	-0.200 **	(0.100)	-0.211 *	(0.120)	-0.210 **	(0.099)
SECTOR10	-0.527 **	(0.263)	-0.487 *	(0.263)	-0.684 ***	(0.220)	-0.489 *	(0.260)	-0.718 ***	(0.222)
SECTOR11	0.553 ***	(0.153)	0.538 ***	(0.152)	0.693 ***	(0.117)	0.627 ***	(0.152)	0.715 ***	(0.117)
SECTOR12	-0.540 ***	(0.208)	-0.514 **	(0.211)	-0.598 ***	(0.157)	-0.606 ***	(0.208)	-0.657 ***	(0.154)
SECTOR13	-0.191	(0.191)	-0.211	(0.189)	-0.068	(0.147)	-0.217	(0.188)	-0.089	(0.149)
SECTOR14	0.014	(0.154)	0.018	(0.151)	-0.083	(0.115)	0.052	(0.152)	-0.091	(0.116)
SECTOR15	-0.165	(0.142)	-0.175	(0.141)	-0.037	(0.112)	-0.136	(0.142)	-0.081	(0.112)
SECTOR16	-0.437 ***	(0.130)	-0.414 ***	(0.130)	-0.390 ***	(0.107)	-0.438 ***	(0.130)	-0.417 ***	(0.107)
MISSING_SECTOR	-0.097	(0.143)	-0.098	(0.143)	-0.072	(0.122)	-0.103	(0.142)	-0.080	(0.121)
MISSING_FIRMSIZE	0.393 **	(0.181)	0.352 *	(0.180)	0.702 ***	(0.164)	0.420 **	(0.184)	0.669 ***	(0.164)
MISSING_ABITUR	0.110	(0.214)	0.110	(0.213)	0.063	(0.206)	0.131	(0.217)	0.033	(0.207)
CHR-DEM_av	-0.340 **	(0.156)	-0.384 **	(0.155)						
SOC-DEM_av	0.793 ***	(0.098)	0.767 ***	(0.097)						
WHITE-COLLAR_av	-0.794 ***	(0.125)					-0.816 ***	(0.115)	(0.125)	
TRAINEE_av	0.779 ***	(0.249)					0.871 ***	(0.251)	(0.250)	
UNEMPL-HIST_av	0.752 ***	(0.198)	0.676 ***	(0.187)			0.789 ***	(0.185)		
EARNINGS_av	0.626 ***	(0.104)	0.420 ***	(0.097)			0.676 ***	(0.105)		
EARNINGS_SQU_av	-0.087 ***	(0.013)	-0.073 ***	(0.013)			-0.091 ***	(0.013)		

Continued on next page...

... table 4.6 continued

	Specification									
	(A)		(B)		(C)		(D)		(E)	
	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)
FIRM-SIZE199_av	1.254 ***	(0.256)	1.296 ***	(0.248)			1.202 ***	(0.265)		
FIRM-SIZE1999_av	1.788 ***	(0.246)	1.822 ***	(0.245)			1.664 ***	(0.235)		
FIRM-SIZE_MORE_av	2.714 ***	(0.255)	2.710 ***	(0.246)			2.853 ***	(0.246)		
SECTOR1_av	1.968 ***	(0.480)	1.831 ***	(0.503)			1.664 ***	(0.500)		
SECTOR2_av	0.455	(0.450)	0.443	(0.505)			0.681	(0.600)		
SECTOR3_av	1.212 ***	(0.444)	1.171 **	(0.477)			1.132 **	(0.557)		
SECTOR4_av	2.073 ***	(0.445)	2.058 ***	(0.458)			2.001 ***	(0.457)		
SECTOR5_av	0.984 ***	(0.325)	0.993 ***	(0.342)			0.809 **	(0.375)		
SECTOR6_av	1.731 ***	(0.288)	1.720 ***	(0.303)			1.708 ***	(0.363)		
SECTOR8_av	0.177	(0.328)	0.283	(0.342)			0.128	(0.487)		
SECTOR9_av	1.034 ***	(0.335)	0.701 **	(0.352)			0.709 *	(0.384)		
SECTOR10_av	-0.199	(0.870)	-0.662	(0.880)			-0.584	(0.768)		
SECTOR11_av	1.348 ***	(0.447)	1.244 ***	(0.455)			1.129 **	(0.447)		
SECTOR12_av	0.949 **	(0.424)	0.539	(0.480)			0.947 **	(0.462)		
SECTOR13_av	1.379 ***	(0.482)	1.373 ***	(0.424)			1.432 ***	(0.488)		
SECTOR14_av	0.595	(0.396)	0.450	(0.372)			0.220	(0.409)		
SECTOR15_av	1.206 ***	(0.316)	1.080 ***	(0.331)			0.954 **	(0.413)		
SECTOR16_av	0.888 **	(0.404)	0.577	(0.409)			0.670	(0.422)		
MISSING_SECTOR_av	0.553	(0.562)	0.381	(0.570)			0.834	(0.568)		
MISSING_FIRM-SIZE_av	1.431 **	(0.587)	1.830 ***	(0.580)			0.865	(0.531)		
MISSING1985	0.308 ***	(0.096)	0.251 ***	(0.097)	0.055	(0.081)	0.331 ***	(0.098)	0.054	(0.084)
MISSING1989	0.062	(0.091)	-0.005	(0.091)	-0.099	(0.082)	0.070	(0.090)	-0.107	(0.083)
MISSING1993	0.142	(0.093)	0.073	(0.091)	-0.022	(0.083)	0.196 **	(0.094)	0.003	(0.084)
MISSING1998	0.278 ***	(0.104)	0.239 **	(0.101)	0.127	(0.091)	0.229 **	(0.102)	0.130	(0.092)
MISSING2001	-0.169	(0.119)	-0.174	(0.118)	-0.203 *	(0.113)	-0.064	(0.119)	-0.196 *	(0.114)
MISSING2003	0.114	(0.099)	0.077	(0.100)	0.076	(0.099)	0.079	(0.105)	0.056	(0.099)
TIME1989	-0.817	(0.537)	-0.774	(0.535)	0.122	(0.226)	-0.749	(0.539)	0.146	(0.226)
TIME1993	-1.281 **	(0.610)	-1.071 *	(0.606)	0.179	(0.234)	-1.505 **	(0.612)	0.162	(0.234)
TIME1998	-2.206 ***	(0.739)	-1.749 **	(0.731)	0.045	(0.257)	-2.428 ***	(0.747)	0.021	(0.257)
TIME2001	-2.609 ***	(0.800)	-2.009 **	(0.789)	0.191	(0.246)	-2.816 ***	(0.813)	0.146	(0.247)
TIME2003	-3.167 ***	(0.902)	-2.442 ***	(0.889)	0.151	(0.258)	-3.378 ***	(0.911)	0.076	(0.261)
Intercept	-6.221 ***	(0.572)	-5.830 ***	(0.568)	-4.296 ***	(0.319)	-6.388 ***	(0.572)	-4.529 ***	(0.315)

Continued on next page...

... table 4.6 continued

	Specification									
	(A)		(B)		(C)		(D)		(E)	
	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)
N. of Obs.	24752		24752		24752		24910		24910	
Log-likelihood	-9035.035		-9064.15		-9535.12		-9185.97		-9371.74	
sigma	1.88	(0.05)	1.86	(0.05)	1.93	(0.05)	1.94	(0.05)	1.77	(0.04)
rho	0.78	(0.009)	0.78	(0.009)	0.79	(0.009)	0.79	(0.009)	0.76	(0.008)
Correct Predictions [%]:										
$y = 0$	81.99		81.03		79.23		80.63		78.79	
$y = 1$	65.46		65.71		68.78		64.81		67.58	

(Correlated) random effects probit models of union membership.
See section 4.4.2 for a description of the model specifications.
* / ** / *** indicate significance at 10% / 5% / 1% level.
Data source: GSOEP.

Table 4.7: Determinants of Union Membership, East Germany

	Specification							
	(A)/(B)		(C)		(D)		(E)	
	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)
FEMALE	-0.273 **	(0.132)	-0.217 *	(0.126)	-0.330 **	(0.131)	-0.203	(0.124)
FOREIGNER	0.791	(1.072)	0.672	(1.017)	0.995	(1.081)		
AGE	0.812 **	(0.341)	0.741 **	(0.325)	1.270 ***	(0.323)	1.038 ***	(0.313)
AGE_SQU	-0.061	(0.041)	-0.049	(0.039)	-0.091 **	(0.038)	-0.078 **	(0.038)
MARRIED	0.387 ***	(0.122)	0.343 ***	(0.116)				
CHR-DEM	-0.249 *	(0.144)	-0.182	(0.141)				
SOC-DEM	0.392 ***	(0.134)	0.415 ***	(0.126)				
ABITUR	-0.741 ***	(0.196)	-0.655 ***	(0.169)	-0.818 ***	(0.177)	-0.678 ***	(0.164)
APPRENTICESHIP	-0.083	(0.120)	-0.072	(0.116)	-0.116	(0.117)	-0.074	(0.114)
UNIVERSITY	-0.096	(0.179)	-0.213	(0.166)	-0.171	(0.169)	-0.205	(0.164)
SKILL-BLUE	0.110	(0.153)	0.062	(0.147)	0.252	(0.154)	0.107	(0.144)
WHITE-COLLAR	-0.082	(0.159)	-0.118	(0.154)	0.155	(0.161)	-0.083	(0.150)
CIVIL SERVICE	-0.131	(0.360)	-0.104	(0.312)	0.215	(0.341)	-0.116	(0.312)
TRAINEE	0.238	(0.268)	0.164	(0.259)	0.428	(0.264)	0.228	(0.255)
PART-TIME	0.106	(0.185)	0.147	(0.173)	0.210	(0.197)	0.163	(0.172)
UNEMPL_HIST_1993	-0.441 **	(0.204)	-0.468 **	(0.200)				
UNEMPL_HIST_1998	0.013	(0.193)	-0.022	(0.185)				
UNEMPL_HIST_2001	-0.114	(0.208)	-0.139	(0.200)				
UNEMPL_HIST_2003	-0.234	(0.251)	-0.244	(0.247)				
SATISFACTION	-0.055 **	(0.027)	-0.049 *	(0.025)				
EARNINGS_1993	1.427 ***	(0.335)	1.205 ***	(0.296)	1.578 ***	(0.332)	1.276 ***	(0.292)
EARNINGS_1998	1.004 ***	(0.266)	0.826 ***	(0.229)	1.073 ***	(0.255)	0.841 ***	(0.224)
EARNINGS_2001	0.905 ***	(0.262)	0.657 ***	(0.219)	0.923 ***	(0.253)	0.701 ***	(0.218)
EARNINGS_2003	0.450 **	(0.205)	0.231	(0.178)	0.554 ***	(0.213)	0.279	(0.177)
EARN_SQU_1993	-0.268 ***	(0.063)	-0.251 ***	(0.056)	-0.295 ***	(0.064)	-0.260 ***	(0.055)
EARN_SQU_1998	-0.137 ***	(0.040)	-0.128 ***	(0.034)	-0.143 ***	(0.039)	-0.129 ***	(0.034)
EARN_SQU_2001	-0.123 ***	(0.039)	-0.103 ***	(0.033)	-0.123 ***	(0.038)	-0.107 ***	(0.033)
EARN_SQU_2003	-0.047 *	(0.024)	-0.033	(0.021)	-0.060 **	(0.027)	-0.039 *	(0.022)
TENURE	0.019 **	(0.009)	0.041 ***	(0.006)			0.043 ***	(0.006)
FIRM-SIZE199	0.349 **	(0.142)	0.384 ***	(0.112)	0.359 ***	(0.137)	0.381 ***	(0.111)
FIRM-SIZE1999	0.715 ***	(0.169)	0.970 ***	(0.138)	0.700 ***	(0.162)	0.945 ***	(0.136)
FIRM-SIZE_MORE	1.023 ***	(0.187)	1.296 ***	(0.146)	1.069 ***	(0.181)	1.302 ***	(0.145)

Continued on next page...

... table 4.7 continued

	Specification							
	(A)/(B)		(C)		(D)		(E)	
	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)
SECTOR1	-0.181	(0.342)	0.009	(0.284)	-0.092	(0.339)	-0.055	(0.284)
SECTOR2	0.397	(0.459)	0.354	(0.385)	0.369	(0.452)	0.329	(0.379)
SECTOR3	0.054	(0.627)	0.556	(0.407)	0.208	(0.609)	0.561	(0.416)
SECTOR4	-0.256	(0.428)	0.490	(0.333)	-0.317	(0.419)	0.477	(0.331)
SECTOR5	0.448	(0.392)	0.631 **	(0.308)	0.393	(0.382)	0.600 **	(0.302)
SECTOR6	-0.308	(0.294)	0.291	(0.251)	-0.242	(0.289)	0.273	(0.249)
SECTOR8	-0.369	(0.287)	-0.220	(0.240)	-0.410	(0.283)	-0.242	(0.239)
SECTOR9	-0.389	(0.299)	-0.079	(0.245)	-0.463	(0.292)	-0.085	(0.244)
SECTOR10	-0.774	(0.530)	-0.597	(0.424)	-0.781	(0.515)	-0.603	(0.415)
SECTOR11	0.105	(0.348)	0.751 ***	(0.265)	0.017	(0.339)	0.719 ***	(0.260)
SECTOR12	-1.082 *	(0.617)	-1.022 **	(0.442)	-1.085 *	(0.612)	-1.034 **	(0.416)
SECTOR13	-0.149	(0.367)	0.426	(0.281)	-0.291	(0.356)	0.399	(0.279)
SECTOR14	-0.598 *	(0.341)	-0.108	(0.262)	-0.671 **	(0.335)	-0.136	(0.260)
SECTOR15	-0.207	(0.328)	0.159	(0.260)	-0.296	(0.322)	0.171	(0.260)
SECTOR16	-0.637 **	(0.297)	-0.200	(0.244)	-0.708 **	(0.293)	-0.251	(0.244)
MISSING_SECTOR	-0.439	(0.401)	-0.140	(0.356)	-0.530	(0.393)	-0.197	(0.348)
MISSING_FIRM-SIZE	0.551	(0.365)	0.774 **	(0.321)	0.523	(0.359)	0.725 **	(0.316)
MISSING_ABITUR	0.048	(0.504)	0.200	(0.487)	0.070	(0.471)	0.233	(0.488)
EARNINGS_av	-0.538 **	(0.241)			-0.431 *	(0.236)		
EARNINGS_SQU_av	0.038	(0.036)			0.024	(0.035)		
TENURE_av	0.045 ***	(0.012)						
FIRM-SIZE199_av	0.161	(0.338)			0.429	(0.309)		
FIRM-SIZE1999_av	1.158 ***	(0.385)			1.755 ***	(0.391)		
FIRM-SIZE_MORE_av	1.053 ***	(0.387)			1.415 ***	(0.379)		
SECTOR1_av	1.719 **	(0.718)			0.944	(0.741)		
SECTOR2_av	0.343	(0.997)			0.618	(0.982)		
SECTOR3_av	2.586 **	(1.100)			2.114 *	(1.138)		
SECTOR4_av	5.235 ***	(0.932)			5.084 ***	(0.870)		
SECTOR5_av	1.260	(0.900)			1.201	(0.834)		
SECTOR6_av	3.182 ***	(0.643)			2.715 ***	(0.654)		
SECTOR8_av	1.596 **	(0.671)			1.228 *	(0.656)		
SECTOR9_av	1.886 ***	(0.672)			1.567 **	(0.617)		
SECTOR10_av	1.688	(1.124)			1.251	(1.139)		

Continued on next page...

... table 4.7 continued

	Specification							
	(A)/(B)		(C)		(D)		(E)	
	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)	Coeff.	(Std.Err.)
SECTOR11_av	2.571 ***	(0.680)			2.844 ***	(0.744)		
SECTOR12_av	0.897	(1.250)			0.251	(1.094)		
SECTOR13_av	2.515 ***	(0.683)			2.491 ***	(0.663)		
SECTOR14_av	2.357 ***	(0.695)			2.052 ***	(0.671)		
SECTOR15_av	1.976 ***	(0.652)			1.592 **	(0.664)		
SECTOR16_av	2.129 ***	(0.648)			1.650 **	(0.646)		
MISSING_SECTOR_av	1.815	(1.177)			1.182	(1.131)		
MISSING_FIRM-SIZE_av	0.932	(1.064)			1.163	(1.032)		
MISSING1993	-0.574 **	(0.242)	-0.410 **	(0.160)	-0.438 **	(0.214)	-0.396 **	(0.161)
MISSING1998	-0.191	(0.195)	-0.075	(0.135)	-0.123	(0.191)	-0.050	(0.132)
MISSING2001	-0.169	(0.196)	-0.093	(0.147)	0.013	(0.220)	-0.110	(0.144)
MISSING2003	-0.410 **	(0.197)	-0.356 **	(0.140)	-0.363 *	(0.189)	-0.362 ***	(0.138)
TIME1998	-1.025 **	(0.440)	-1.029 **	(0.426)	-0.996 **	(0.419)	-0.951 **	(0.411)
TIME2001	-1.127 **	(0.446)	-1.056 **	(0.428)	-1.062 **	(0.432)	-1.063 **	(0.418)
TIME2003	-0.839 *	(0.442)	-0.826 *	(0.427)	-0.915 **	(0.433)	-0.851 **	(0.415)
Intercept	-6.054 ***	(0.940)	-4.316 ***	(0.772)	-7.337 ***	(0.907)	-5.212 ***	(0.736)
N. of Obs.	7127		7127		7152		7152	
Log likelihood	-2641.50		-2739.62		-2702.79		-2767.36	
sigma	2.42	(0.14)	2.24	(0.12)	2.33	(0.12)	2.22	(0.11)
rho	0.85	(0.014)	0.83	(0.014)	0.84	(0.014)	0.83	(0.014)
Correct Predictions [%]:								
$y = 0$	80.48		79.32		79.35		79.18	
$y = 1$	62.15		63.52		61.54		63.02	

(Correlated) random effects probit models of union membership.
See section 4.4.2 for a description of the model specifications.
* / ** / *** indicate significance at 10% / 5% / 1% level.
Data source: GSOEP.

5 Equal Pay for Equal Work? On Union Power and the Structure of Wages in West Germany, 1985–1997

5.1 Introduction

The influence of institutions on economic performance in general, and on wage setting in the labor market in particular, is currently under debate (OECD 2004, 2006). In times of increasingly heterogeneous economic conditions the catchword is eurosclerosis, stating that institutional rigidities restrain labor market performance and the dynamics of economic development. A major focus is on the impact of trade unions; see, e. g., the handbook of Addison and Schnabel (2003).

The main channel for unions to influence the wage structure is via collective bargaining. In Germany, this influence goes beyond mere negotiating of wage premia for union members since collective agreements on individual membership premia are forbidden by constitutional law. Given the high rate of collective bargaining coverage in the German labor market, union-bargained wages apply to the major part of all employees and unions influence the wage structure of members as well as of non-members.[1]

How does union power affect wage levels and the degree of wage dispersion against this background? Do unions aim to realize "equal pay for equal work"? That is, does the impact of unions reduce residual wage inequality between employees with similar observable characteristics? Moreover, does union power also reduce wage dispersion between employees with different characteristics? Are the effects asymmetric across the conditional wage distribution, corresponding to a minimum wage argument, for instance? Is there a trade-off between reduced inequality

[1] It is generally possible to distinguish between union members and non-members when applying opening clauses. For example, the metalworking industry union *IG Metall* negotiated bonuses exclusively for their members in some firm-level agreements in the collective bargaining area of North Rhine-Westphalia; see Berliner Zeitung (2004). However, such discriminatory agreements constitute rare exceptions. Firms do not want to create incentives for their employees to become union members.

and higher wage levels, such that a reduction in wage inequality comes along with a lower wage level? And did any of the effects change over time?

In order to answer these questions, our paper draws a detailed descriptive picture of the correlation between union membership and the wage structure in the German labor market. We argue that net union density is a good proxy for union power because of the associated financial power and the personal representation within firms, both governing the union's threat point in the collective bargaining process and therefore *determining* the bargaining outcome. This approach is complementary to studies based on collective bargaining coverage which focus on the *application of* bargaining agreements.

Since there is no detailed information on net union densities available for Germany, we employ estimations by Beck and Fitzenberger (2004) to impute propensities to be a union member for individuals in the IAB employment sample (*IAB Beschäftigtenstichprobe*, IABS) 1975–1997. Taking averages of these propensities, we project net union densities (NUD) for adequately defined labor market segments in the period 1985–1997 and analyze the correlation of NUD on the one hand and measures of the wage structure within and between the segments on the other.

Our descriptive results show that a higher net union density is ceteris paribus associated with a lower wage level and reduced residual wage dispersion. This—possibly surprising—result is in line with an insurance motive for union membership as discussed by Agell and Lommerud (1992) and Burda (1995). Furthermore, skill wage differentials are smaller in segments with strong unions, too—the union's impact even goes beyond "equal pay for equal work". Finally, wage compression is in fact asymmetric and occurs notably in the lower parts of the distribution. This finding corresponds to a minimum wage interpretation of union-negotiated wages.

The course of the paper is organized as follows. Section 5.2 motivates the use of net union density as a proxy for union power and sketches theoretical considerations about the relationship between union power and the distribution of wages. Section 5.3 provides a brief overview of complementary microeconometric studies of the German collective bargaining system. Our empirical analysis is conducted in section 5.4. Section 5.5 concludes.

5.2 Union Power, Union Density, and the Distribution of Wages

In this section we discuss the use of union density as a measure of trade union power and we outline considerations in the theoretical literature in order to derive testable hypotheses on the relationship between union membership on the one hand and wage levels and wage dispersion on the other.

5.2.1 Measuring Labor Market Institutions

In the tradition of Calmfors and Driffill (1988), cross-country comparative studies usually use single indices to reduce dimensionality when operationalizing labor market institutions; see the surveys in Kenworthy (2001, 2003), OECD (1997, 2004), or Schettkat (2003). Corresponding empirical evidence on the relationship between the institutional design of the labor market on the one hand and measures of economic performance such as GDP, unemployment rates, and the level of wages on the other hand, is generally ambiguous. The only exception is a stable correlation of institutional settings and the wage structure: Higher degrees of centralization and coordination of wage bargaining are ceteris paribus associated with lower wage dispersion; see the synopses in Aidt and Tzannatos (2002), Blanchflower (2006), Flanagan (1999), Gerlach and Meyer (1995), OECD (1997, 2004), and the handbook articles of Blau and Kahn (1999) and Nickell and Layard (1999). Blau and Kahn (1996) find an asymmetric effect such that a higher degree of centralization compresses the wage distribution mainly from below.

Schettkat (2003) offers a two-step interpretation of institutional impacts and points out that in contrast to the negative correlation between centralization or coordination and wage dispersion in the first step, there is a lack of unambiguous evidence on the relationship between wage dispersion and (un)employment in the second step.

In addition to the sensitivity with respect to the chosen measure, comparative studies based on single measures exhibit several shortcomings. For example, Soskice (1990) and Rowthorn (1992) emphasize the concomitant importance of coordination *and* centralization. Changes of institutional settings over time (Wallerstein and Western, 2000) should be taken into account as well as the endogeneity of institutions (Flanagan, 1999). According to Acemolgu, Aghion, and Violante (2001), skill-biased technological change may be responsible for the deunionization trends in recent decades. Moreover, the effect of particular institutional elements is likely to interfere with further country-specific institutions; see, e. g., Hübler and Jirjahn (2003), Jirjahn (2003), and Klikauer (2004) on the interaction of collective bargaining, union representation, and firm-level co-determination in Germany. Fitzenberger and Franz (1994) conclude that detailed institutional knowledge is necessary in order to judge the actual design of the German wage-setting system. Country-specific studies based on micro data can meet the requirement to interpret institutional effects with reference to social norms within society (Flanagan, 1999). Against this background, our study focuses on union power and the structure of wages in the German labor market.

5.2.2 Union Density as a Measure of Union Power

The game theoretical literature puts emphasis on the “bargaining power” of negotiating parties, which empirical studies need to operationalize. In the Anglo-Saxon literature union power is defined as the product of union density and the union wage gap; see Addison, Bailey, and Siebert (2004). This concept is inappropriate for the German case because collective agreements constituting discriminatory wage policies with disadvantages for non-members are forbidden by constitutional law (negative freedom of association, *negative Koalitionsfreiheit, Grundgesetz Art. 9*). As wage gains from union membership are not internalized, there exists a free-rider problem of missing individual incentives to join a union.[2] In fact, the scope of collective agreements goes beyond the organized parties. Wages set at the firm level as well as individually bargained wages are geared to collective bargaining agreements, be it in order to reduce transaction costs or to avoid incentives for employees to join a union. In West Germany collective agreements are binding for the majority of workers; see Kohaut and Schnabel (2003a, 2003b).[3] Prevalent wage-setting models in the literature therefore even assume that collective bargaining agreements apply to all employees; see, e. g., Fitzenberger (1999), chapter 6.

Firms' decision whether to apply collective bargaining agreements is pivotal to the degree of bargaining coverage. Having said that, net union density governs the union's threat point in the collective bargaining process and therefore is pivotal to the outcome of the bargaining. The more union members pay membership fees, the better is the union's funding. Yet financial obligations also increase with the size of the union. Relative financial power thus is best represented by net union density defined as the share of contributors among all potentially represented workers. In case of industrial conflicts, higher financial power enables the union to pay strike benefits for a longer period of time. This increases individual support for union action, the probability and the length of a strike, and therefore the expected damage inflicted upon employers.[4] Furthermore, financially powerful unions can invest more in public relations in order to sanitize their public image. The more homogenous the public appearance of union representatives is, the higher is the sustainability of press coverage. Again, the focus is on union density instead of absolute size since union growth is associated with increased heterogeneity within the union. Conflicting interests and contradictory statements then undermine the union's representative role; see also Ebbinghaus (2003) and Keller (2005).

In sum, a high net union density strengthens the union's position in the wage bargaining process. The succeeding analysis of the effect of union power on the wage structure thus builds on union

[2] However, there are additional motives for union membership. The literature discusses selective incentives provided in addition to public goods (Olson, 1965), collective-voice mechanisms (Hirschmann, 1970), or the existence of social norms (Akerlof, 1980; Booth, 1985).

[3] Collective agreements can also be declared generally binding by the Minister for Labor and Social Affaires. The direct impact of this provision may be of minor relevance—only 0.8% of all employees subject to social security contributions are covered by agreements which are binding by declaration (BMWA, 2004). Yet the mere possibility of such a declaration constitutes incentives per se; see OECD (1994).

[4] Skeels and McGrath (1997) confirm this aspect for the US. They find that more liquid funds per union member ceteris paribus increase the probability of a strike.

membership data. This approach is complementary to studies of bargaining coverage and the application of collective wage agreements, which are discussed in section 5.3 below.

5.2.3 Union Power and the Distribution of Wages

Bargaining models regard the negotiation of wages as a rent-sharing problem, the solution to which depends on the bargaining power of the negotiating parties. Classical models predict a monotonic positive (negative) relationship between union power and the level of bargained wages (employment); see the surveys of Farber (1986) and Oswald (1985).

Some more recent studies incorporate effects on higher moments of the wage distribution in addition to effects on the wage level. Agell and Lommerud (1992) and Burda (1995) focus on wage dispersion and discuss an insurance motive for union membership. Faced with uncertainty of future productivity or wages, risk aversive employees a priori have a taste for wage compression. Unions act as agents of the work force and bargain for a compression of the wage distribution relative to the productivity distribution. The compression comes at the price of an insurance premium in the form of a reduced wage level.[5] Ceteris paribus, the degree of wage compression is higher, the higher the bargaining power of the union.

Ex ante, individuals prefer wage compression along unobservable as well as observable dimensions which may influence future productivity. For instance, they might demand insurance against negative shocks to their skill level or secure smoothed earnings over the life cycle. Ex post, union membership is irrational for individuals who have reached a position in the upper part of the (conditional) wage distribution and who may expect to keep this position because, say, match-specific information about their productivity has been revealed in the course of their working life. In this case the insurance motive would be time inconsistent. Unions could counteract this problem by approving an increase of wage inequality with increasing age or experience of the employees.

In the reasoning above, wage differences result from differences in workers' marginal productivity. Besides, search and matching theories (Mortensen and Pissarides, 1999) stress the importance of imperfect information and the existence of search frictions for match-specific rents, which can cause wage dispersion even among individuals with identical characteristics. In case of wage-posting by the firms, workers also accept wage offers which are below their marginal productivity. Unions, equipped with higher bargaining power than individual workers, counteract the monopsony power of the firms and retard downward deviations, realizing a lower wage dispersion.

[5] If the wage distribution is skewed to the right the redistribution motive is consistent with median voter models of the union; see Burda (1990), Farber (1986), and Freeman (1982). The reallocation implied by a compressed wage structure can be understood as a substitute for explicit means of redistribution such as taxation; see Agell (1999, 2002).

By enforcing "equal pay for equal work" unions additionally seek to limit favoritism and discrimination by superiors and colleagues, and to encourage solidarity among the work force; see Freeman (1982).

On principle, union action can also work towards higher wage dispersion. For example, some search and matching models predict that, in the absence of unions, firms pay the reservation wage to all workers, regardless of their actual productivity. This would lead to wage compression if the distribution of reservation wages was more compressed than the productivity distribution. Again, powerful unions would claim a share of the match specific rents, now enforcing a higher wage dispersion. It is also conceivable that powerful unions generally enforce a high level of wages when bargaining with successful firms, but offer opening clauses or renegotiations to firms flourishing less; see, e. g., Fitzenberger and Franz (2000).

Due to the extreme assumptions underlying the arguments for a positive correlation, empirical evidence is unlikely to match these arguments. Instead, we expect a negative correlation between union power and wage dispersion, and we anticipate a trade-off between reduced inequality and a higher wage level; see also Calmfors (1993).

At any rate, the impact of unions on the wage structure likely varies across the wage distribution. If collective wages serve as wage floors, the (conditional) wage distribution is compressed from below. In the wage bargaining model of Büttner and Fitzenberger (2003) efficiency wages are paid in the upper part of a productivity distribution, whereas union-bargained wages above marginal productivity are binding for less productive matches. This is in line with the perception of unions which represent mainly less productive employees, striving for higher wages particularly at the lower end of the distribution. Then, compression of the wage distribution from below is the higher, the stronger the influence of the union.

Referring to the interaction of industry-level collective agreements and local wage formation, Büttner and Fitzenberger (1998) consider two kinds of wage flexibility. Wages can react to the overall economic conditions (operationalized by the national rate of unemployment) as well as to local conditions (regional unemployment rates): Baseline wages are set at the industry level with respect to the economy-wide environment, but deviations are negotiated under consideration of the particular conditions at the local level. In this case, a high impact of the national unemployment rate on the remuneration of low-skilled workers at the lower end of the wage distribution reflects a minimum wage interpretation of industry-level contract wages. Since union influence is greater at the industry level than at the local level, this finding also alludes to the asymmetry of the union impact across the wage distribution.

In sum, we expect to find a positive correlation of union power and the degree of wage compression from below.

5.3 Complementary Empirical Studies for Germany

Bargaining coverage, as measured by the share of employment contracts following collective agreements, was relatively stable in West Germany until the end of the 1990's. Yet union density, which had also shown mere variation with the business cycle in former decades, has been declining distinctly since the end of the 1980's; see Franz (2003) and OECD (2004) for overviews of the German collective bargaining system and Ebbinghaus (2003) and Bosch (2004) for comprehensive inquiries into recent institutional developments. Detailed insights can be gained on the basis of micro data. This section surveys related studies in the literature which scrutinize the design of the collective bargaining system or analyze institutional wage effects.

5.3.1 The Collective Wage Bargaining System

Individual determinants of union membership are estimated by Lorenz and Wagner (1991), Fitzenberger, Haggeney, and Ernst (1999), and Beck and Fitzenberger (2004) based on different waves of the German Socio-Economic Panel (GSOEP) and by Schnabel and Wagner (2003, 2005) based on data from the German General Social Survey (ALLBUS). The effects of socio-demographic personal and job-related characteristics[6] on individual decisions to join a union are found to be rather similar in West Germany across the different waves, and effects have also converged between East and West Germany during the first decade after unification. Therefore, the decline in union membership during the 1990's is to a large extent attributed to composition effects among the labor force. In light of the free-rider problem of union membership, the study of Goerke and Pannenberg (2004), which is also based on the GSOEP, explicitly addresses the impact of social norms on individual propensities to join a union.

Collective bargaining coverage is markedly lower in East Germany than in the West, and it has been declining in both parts of the country in recent years. This "erosion" is examined by several studies using firm data. Based on the IAB establishment panel, Kohaut and Bellmann (1997), Bellmann, Kohaut, and Schnabel (1999), and Kohaut and Schnabel (2003b) estimate determinants of firms' decision to apply collective wage agreements. Kohaut and Schnabel (2003a) differentiate between a direct commitment to collective agreements and firm-level or individual agreements which are geared to the collective ones. Age and skill level of the work force and firm size positively affect the probability of bargaining coverage apart from industry-specific effects. In addition, the existence of a works council significantly reduces the propensity of a firm to abolish collective coverage. Survey evidence from works councils discussed in Bispinck and

[6] See Schnabel (1993) for an extensive discussion of various personal and job-related determinants of union membership.

Schulten (2003) reconfirms tendencies to more decentralized wage setting in recent years; see also Bosch (2004).

5.3.2 Effects of the Wage Bargaining System

The design of the collective wage bargaining system affects the level and the dispersion of wages. There is a vast literature on union wage effects in Anglo-Saxon systems; see Card, Lemieux, and Riddell (2003) for an overview. However, this literature is of limited relevance for the case of Germany as collective agreements on membership premia are forbidden by constitutional law (negative freedom of association); see also Schmidt (1994).

Using different cross sections (1990, 1995, 2001) of the Structure of Earnings Survey (SES, *Gehalts- und Lohnstrukturerhebung*) for Lower-Saxony, Gerlach and Stephan (2002, 2005a, 2005b) estimate wage distributions for labor market regimes with and without collective and firm-level wage agreements. In the manufacturing sector, average hourly wages paid in accordance with a collective or a firm-level agreement are higher than the average of individually negotiated wages. Yet unconditional as well as conditional wage dispersion is highest among individual contracts. Similar results are obtained in Statistisches Landesamt (2004) based on the SES subsample for Baden-Württemberg. Multi-level regression models in Stephan and Gerlach (2003, 2005) reveal that differences in the wage level are consistent with a higher base wage in case of collective coverage. Returns to human capital—skill, experience, and tenure—as well as residual wage dispersion are lower under collective coverage. Gerlach and Stephan (2006) find that collective agreements compress within-firm compensation schemes across occupations.

Dustmann and Schönberg (2004) focus on the relationship between collective coverage and professional training. Firms applying collective contracts as a commitment device ceteris paribus employ a higher share of workers with an apprenticeship degree. Moreover, the employed linked data of the IAB employment statistics and the IAB establishment panel suggest that under collective coverage, employee turnover is higher, wage cuts happen more often, and (conditional) wages have a lower variance.

Büttner and Fitzenberger (1998) analyze the joint impacts of industry-level collective bargaining and local agreements on the wage distribution; compare section 5.2.3. Using the IAB employment sample (IABS), they find that overall economic conditions—as measured by the national rate of unemployment—are taken into account at the centralized level of wage bargaining. Resulting contract wages work as minimum wages and affect the wage distribution mainly in the lower part. On the other hand, local specifics—captured by regional unemployment rates—result in incentive wages which cause higher flexibility at the upper end of the wage distribution. Pooled cell-data regressions for the period 1976–1990 further indicate that union influence reduces wage dispersion: A higher net union density ceteris paribus comes along with an (albeit insignificant) increase in wages at low quantiles of the distribution and a (significant) decrease at higher quantiles.

Also drawing on the IABS 1975–1990, Fitzenberger (1999, chapter 6) estimates a structural model of industry-level wage bargaining. In this study, a union maximizes a Stone-Geary utility function with specific weights for employment, average wages, and—in some of the specifications—wage dispersion within two skill classes of the work force. In line with a right-to-manage assumption as in Pencavel and Holmlund (1988), employment is determined by the firms. There are effects of habit formation in the function weights for employment and average wages, and unions put specific emphasis on the employment goal. In specifications which include wage dispersion in the objective, unions put a positive weight on the reduction of dispersion and make concessions in particular with respect to the employment goal. In manufacturing, an increase in net union density is associated with a significantly stronger preference for high employment relative to the wage levels and the reduction of wage dispersion.

5.4 Empirical Analysis

This section draws a detailed descriptive picture of the relationship between union power and the wage structure within and between segments of the German labor market. The analysis is based on the IAB employment sample (*IAB Beschäftigtenstichprobe*, IABS) 1975–1997, which is appreciated for its big sample size and exact administrative wage information; see Bender, Hilzendegen, Rohwer, and Rudolph (1996) and Bender, Haas, and Klose (2000) for detailed descriptions of the data set. In order to circumvent a number of selection problems, we focus on prime-age males employed full-time in West Germany; see the appendix for a review of our data selection. Since the IABS does not provide data on union membership, Beck and Fitzenberger (2004) employ survey information provided with the waves 1985, 1989, 1993, and 1998 of the GSOEP to estimate determinants of individual union membership decisions. Estimated determinants are rather stable over time and are thus used to predict probabilities to be a union member for individuals in the IABS over the period 1985–1997.

For our analysis we additionally allow the individual propensities—i. e., the predicted probabilities—to vary by industry sector in case of workers who were employed in different sectors within a year. Net union densities for different labor market segments can then be estimated by means of aggregation. The following sections motivate our segmentation, illustrate developments of union membership, and analyze the correlation between union density and the structure of wages within and between the chosen segments.

In light of the literature on bargaining coverage and the wage structure summarized in the previous section, it would be desirable to simultaneously analyze the effects of collective bargaining coverage *and* union density on the wage structure. Unfortunately, this is not possible as the IABS does not provide information on bargaining coverage.[7]

A further alternative to our approach would be to use the GSOEP for the entire analysis. On principle, labor market segments could be constructed (as explained in the next section) with the GSOEP in the same manner as with the IABS data. However, the much smaller sample size of the GSOEP would result in a set of cells with very few or even no observations. Corresponding estimates of the conditional wage structure, which are central to our analysis, would be rather inaccurate. For this reason we rely on the IABS to generate measures of the wage structure and only impute union membership from GSOEP results. At the chosen level of aggregation the prediction error should be largely negligible such that our estimates are not significantly biased by measurement error.

The evolution of the German wage structure over the last decades has been extensively documented in the literature; see Fitzenberger and Kohn (2006), Fitzenberger and Kurz (2003), Kohn (2006) and the literature cited therein. We thus refrain from an expatiated discussion at this point.

5.4.1 Labor Market Segments and the Operationalization of Union Power and Wage Inequality

In order to partition the labor market, we first define cells spanned by the dimensions time (13 years) × age (seven five-year classes for individuals aged between 25 and 60 years) × industry (17 sectors).

The industry dimension is central to analyses of the wage structure; see Krueger and Summers (1988) and Bellmann and Gartner (2003), Bellmann and Möller (1995), or Fitzenberger and Kurz (2003) for the case of Germany. Not only do these studies find inter-industry wage differentials, but they also detect asymmetries such as greater dispersion for higher wage quantiles. In addition, union strength traditionally varies between industries; see Beck and Fitzenberger (2004) who highlight significant inter-industry differences in union density. Table 5.5 in the appendix contains our grouping of industry sectors.

Employees age as they advance on the life cycle path. The interaction of age and time then allows to study cohort effects on the wage structure (Fitzenberger and Kohn, 2006). In the

[7] For this reason, the use of linked employer-employee data—such as the newly available LIAB data—is warranted in future research.

context of union membership, the age dimension further promises information about the ex-post rationality of membership. Besides, age serves as an—albeit imperfect—proxy for work experience. Educational attainment likewise measures the formal qualification of an employee. We distinguish between three skill classes: high-skilled employees with a university or technical college degree, medium-skilled employees who have completed a vocational training, and low-skilled employees with neither university nor apprenticeship degree.

The definition of cells is advantageous for two reasons. First, it reflects the structure of the German wage bargaining system. The sector classification accounts for the fact that collective negotiations take place at the industry level. The observation that collective agreements further differentiate between various wage groups is captured by the dimensions age and skill. Unfortunately, an additional segmentation by regions is not possible because the number of observations in the respective cells would get too small. In each cell z we calculate net union density NUD_z, our measure of union power, as the mean of individual membership propensities, weighted by the durations of the respective employment spells. The cell level aggregation enables us to analyze the effect of union power independently of individual membership. As pointed out in the discussion in section 5.3.2 above, it would not make sense to estimate individual membership premia.

Second, the cell-level approach allows us to study both wage dispersion *between* the defined segments and residual wage dispersion, which remains *within* the homogenous segments. Due to the censoring of wages at the social security taxation threshold, we estimate cell-specific wage levels (LNW_z) using Tobit regressions of individual log real wages on a constant:

$$\ln w_{iz} = LNW_z + \epsilon_{iz}. \tag{5.1}$$

The regressions are run separately for each industry×age×time-cells z, weighting individual observations i by the length of the respective employment spells. Measures of wage dispersion between skill groups, i. e., skill premia BH_z for the high-skilled and (negative) premia BL_z for the low-skilled, are obtained from analogous Tobit regressions

$$\ln w_{iz} = LNWM_z + BL_z \cdot DL_{iz} + BH_z \cdot DH_{iz} + \epsilon_{iz} \tag{5.2}$$

which contain dummies for low (DL) and for high-skilled workers (DH) as additional regressors. The level effect $LNWM_z$ now comprises the average wage of a medium-skilled worker.

Quantile differences QD serve as measures of cell-specific unconditional wage dispersion. We focus on the difference between the fourth and the first quintile, $QD8020$, and on the differences $QD5020$ and $QD8050$ in order to examine asymmetries across the distribution. In addition, the Tobit-sigma ($\sigma TOBIT$) estimated from (5.2) represents a first measure of residual dispersion.

However, $\sigma TOBIT$ depends on the functional form of the Tobit specification, and so we intend to additionally employ quantile differences from conditional log wage distributions. For this purpose, we take skill as an additional cell dimension and estimate equation (5.1) separately for industry×age×skill×time-cells $\tilde{z}$. Compared to the first approach above, this approach has the disadvantage that the number of observations gets rather small in several cells and that—especially for the high-skilled—the censoring problem is more severe such that larger parts of the data cannot be used for the analysis. However, it offers the advantage that quantile differences QD are additionally conditioned on the skill level of the workers. We therefore obtain a measure for residual wage dispersion which is independent of any underlying distribution, and we can examine asymmetries as well.

5.4.2 Evolution of Net Union Density

Net union density varies between 4.4 and 85.1% in the different segments. In order to capture the variation along different dimensions, table 5.6 in the appendix shows the results of weighted regressions of NUD on the components industry, age, and time (column 2) as well as additionally on skill (column 4).

Differences by industry are most prominent. Union density is highest in the formerly state-owned postal and railway services, and it is lowest in financial intermediation. Moreover, union density ceteris paribus increases with age. Insofar as this finding reflects an actual age effect, membership does not decline over the life cycle and there is no evidence for a lack of ex-post rationality of membership. Yet an interpretation as a cohort effect is also intuitive: Younger cohorts do not identify themselves with unions as much as older cohorts. With respect to differences by skill level, net union density among the low-skilled is only marginally higher than NUD among the medium-skilled, whereas NUD is lower by 12 percentage points among the high-skilled, ceteris paribus. The time dummies further indicate that NUD remained relatively stable during the 1980's but it has declined significantly since 1990.

Figure 5.2 in the appendix illustrates these trends by plotting NUD by industries against age for the two cross sections 1985 and 1997.[8] Again, the significant differences across industries are striking. In the year 1985, NUD increases uniformly with age in almost all industries. The only exception is the oldest age group, from which union members disproportionately select into early retirement. The 1997 cross section reveals that the decline in NUD is heavily borne by decreasing membership rates among younger cohorts. Moreover, the decline is the more distinct

[8] The evolution of NUD is rather similar across skill levels. So it is unlikely that our results merely reflect the skill upgrading in the work force.

the higher the base level of industry-specific NUD. This leads to a continuing compression of the membership structure across industries.

5.4.3 Net Union Density and the Structure of Wages

We regress the measures of wage level and wage dispersion on NUD in order to quantify the effect of union power on the wage structure. Dummy variables for the cell dimensions control for additional effects in the determination of the wage bargaining outcome.[9] Tables 5.1 to 5.4 report the ceteris paribus effects of a transition from a situation without union representation ($NUD = 0$) to a situation with full representation ($NUD = 1$). In the following, we discuss marginal effects resulting from an increase in net union density by one percentage point.

Table 5.1: Net Union Density and the Structure of Wages I

Level	LNW	$LNWM$	$LNW50$			
NUD	-0.249	-0.255	-0.355			
	(0.049)	(0.045)	(0.051)			
# Cells	1545	1545	1545			
Dispersion	$QD8020$	$QD5020$	$QD8050$	BU	BH	$\sigma TOBIT$
NUD	-0.530	-0.427	-0.144	0.565	-0.239	-0.248
	(0.089)	(0.054)	(0.044)	(0.076)	(0.095)	(0.036)
# Cells	1235	1545	1235	1545	1513	1545

Estimations by weighted least squares, cells weighted by cell-specific employment. Exclusion of cells censored at the 50% quantile (80% quantile for $QD8020$ and $QD8050$). Covariates: dummies for time, age, and industry. White robust standard errors in parentheses. Data source: extended IABS 1985–1997.

Table 5.1 displays evidence for industry×age×time-cells z. An increase in NUD by one percentage point ceteris paribus comes along with a decrease of the level of mean wages (LNW or $LNWM$, respectively) by 0.25 log percentage points (pp). This result is surprising at first glance as studies in the literature on the effects of collective bargaining coverage find higher wages in firms applying collective contracts (compare section 5.3.2), and bargaining coverage and union density within firms are likely to be positively correlated. Unfortunately, the IABS data do not allow us to analyze the impact of bargaining coverage. Possible explanations for the different results should be investigated in future research using appropriate data: On the one hand, it is not clear a priori whether a positive correlation between collective bargaining coverage and

[9] For example, industry dummies take account of differences in the intensity of competition in different product markets as well as of different rates of employer unionization. Employees' fall-back options are determined by their age and the level of skills.

union representation persists if additional covariates are controlled for. On the other hand, union density affects the structure of wages in a labor market segment irrespective of actual bargaining coverage. In light of the argument put forward in section 5.2, we consider this the most important cause of the difference. It is plausible to assume that only a strong union with high union density has got the power to achieve equality objectives for a whole labor market segment independently of the degree of collective coverage. Collective contracts are most likely accepted in firms capturing high rents, and on average these firms pay higher wages to their workers.

The median $LNW50$ is even decreased by 0.36 pp. Moreover, higher union representation involves a reduction of the unconditional quantile difference $QD8020$ by about 0.53 pp, most of which is observed in the lower half of the distribution ($QD5020$: -0.43 pp, $QD8050$: -0.14 pp). The union effect on wage compression is in fact asymmetrical: While total dispersion decreases, the skewness of the log wage distribution rises.

The skill differentials BL and BH are significantly reduced by respective 0.57 and 0.24 pp: Higher union power comes along with a lower wage dispersion between skill groups. This finding is in line with the result of Kahn (2000), who observes a negative correlation of union density and skill wage differentials in a cross-country study. Provided that an employee's productivity is adequately reflected by her/his skill level, the finding suggests that the union impact exceeds the objective "equal pay for equal work", compressing wages even between groups of workers with observationally different productivity.

The Tobit-Sigma ($\sigma TOBIT$) estimated from equation (5.2) is lower in case of a stronger union representation. This finding indicates a reduction of residual wage dispersion in addition to the compression of skill differentials BL and BH. Further insights about within-group dispersion are gained from table 5.2, which reports the ceteris paribus effects for industry×age×skill×time-dimensioned cells $\tilde{z}$.

The first panel of table 5.2 displays the results of regressions for a pooled sample of all skill groups. We observe a decline in the level of wages similar to that in table 5.1 above. What is more, the finding of reduced quantile differences ($QD8020$: -0.65 pp) now reflects a decrease in residual wage dispersion: In segments with stronger unions, within wage dispersion as well as the average wage are lower, ceteris paribus. This result corresponds to the insurance motive for union action discussed by Agell and Lommerud (1992) and Burda (1995), and it corroborates the results of Büttner and Fitzenberger (1998) on the effects of net union density as well as those of Dustmann and Schönberg (2004) on the application of collective contracts.[10]

The dispersion effects obtained from the pooled estimations are rather symmetrical. However, separate estimations for the three skill groups in the panels 2 to 4 of table 5.2 draw a more

Table 5.2: Net Union Density and the Structure of Wages II: Differences by Skill Level

Total	LNW	$LNW50$	$QD8020$	$QD5020$	$QD8050$	$\sigma TOBIT$
NUD	-0.282	-0.327	-0.654	-0.319	-0.320	-0.306
	(0.036)	(0.037)	(0.056)	(0.029)	(0.031)	(0.025)
# Cells	3442	3442	3032	3442	3032	3442
Low-skilled						
NUD	0.459	0.496	-0.598	-0.229	-0.369	-0.330
	(0.049)	(0.057)	(0.066)	(0.034)	(0.044)	(0.029)
# Cells	1474	1474	1441	1474	1441	1474
Medium-skilled						
NUD	-0.270	-0.407	-0.584	-0.452	-0.166	-0.316
	(0.048)	(0.049)	(0.072)	(0.050)	(0.036)	(0.038)
# Cells	1545	1545	1400	1545	1400	1545
High-skilled						
NUD	-0.658	-0.394	0.805	0.350	0.169	-0.026
	(0.129)	(0.118)	(0.421)	(0.158)	(0.169)	(0.146)
# Cells	423	423	191	423	191	423

Estimations by weighted least squares, cells weighted by cell-specific employment. Exclusion of cells censored at the 50% quantile (80% quantile for $QD8020$ and $QD8050$). Covariates: dummies for time, age, industry, and skill. White robust standard errors in parentheses. Data source: extended IABS 1985–1997.

heterogeneous picture. In the core group of medium-skilled workers, wage level and dispersion are reduced to a similar extent as in the pooled regressions, but the wage distribution is compressed asymmetrically from below ($QD5020$: -0.45 pp, $QD8050$: -0.17 pp). At the lower end of the (conditional) wage distribution, union-bargained wages serve as minimum wages.

In the group of low-skilled workers, a more powerful union is associated with a higher average wage (LNW: +0.46 pp). This finding reflects the equalizing effect with regard to the skill differentials. Again, residual wage dispersion declines with rising labor union power, but now a

[10] A referee pointed out two alternative explanations for the negative correlation between union density and the wage level. First, higher wages could be paid in specific segments in order to prevent workers from unionization. Second, a higher proportion of women in a labor market segment could be accompanied by lower wages for men as well as a higher union density among men. In a strict sense, we can not reject these alternatives on the basis of our analysis. However, we consider them as little plausible for the following reasons: First, it is not clear how to explain our findings with respect to wage dispersion in these settings. Second, our estimations control for industry and other effects which take up the alternative explanations. Third, the positive correlation between collective bargaining coverage and the wage level found by other studies (compare section 5.3.2) is in contrast to the deterrence effect. Fourth, the partial effect of being female is negative in the union membership estimations of Beck and Fitzenberger (2004). Hence, it is unlikely that the union density of men is particularly high in segments with a high proportion of female workers.

larger part of the compression occurs in the upper half of the conditional distribution ($QD5020$: -0.23 pp, $QD8050$: -0.37 pp). Wages in the mid-parts of the distribution are most pronouncedly moved up by union action.

High-skilled workers on average earn significantly less in sectors with strong unions (LNW: -0.66 pp). However, the number of utilizable cells for this group is severely restricted due to the high degree of wage censoring, such that we obtain no significant effects on wage dispersion.

We inferred from section 5.4.2 that net union density is higher among workers of older age or older cohorts. In order to scrutinize the ex-post rationality of union membership over the life cycle, we further examine whether the union wage effects vary systematically with age. For that purpose, the regressions in table 5.3 include interaction terms of NUD and age dummies. The analyses are implemented separately for low-skilled and for medium-skilled workers. Due to the severe problem of wage censoring among the high-skilled, we do not expect reliable results for this group.

Table 5.3: Net Union Density and the Structure of Wages III: Differences across Age

Low-skilled	LNW	$LNW50$	$QD8020$	$QD5020$	$QD8050$	$\sigma TOBIT$
$NUD \times D(25 \leq \text{age} \leq 29)$	0.592	0.649	-0.620	-0.224	-0.394 *	-0.326
$NUD \times D(30 \leq \text{age} \leq 34)$	0.498	0.539	-0.602	-0.222	-0.377 *	-0.334
$NUD \times D(35 \leq \text{age} \leq 39)$	0.448	0.481	-0.558	-0.202	-0.354 *	-0.335
$NUD \times D(40 \leq \text{age} \leq 44)$	0.388	0.425	-0.597	-0.234	-0.364 *	-0.336
$NUD \times D(45 \leq \text{age} \leq 49)$	0.331	0.363	-0.571	-0.222	-0.351 *	-0.305
$NUD \times D(50 \leq \text{age} \leq 54)$	0.335	0.367	-0.540	-0.204	-0.339 *	-0.293
$NUD \times D(55 \leq \text{age} \leq 59)$	0.351	0.417	-0.483	-0.140	-0.343 *	-0.259
# Cells	1474	1474	1441	1474	1441	1474
Medium-skilled						
$NUD \times D(25 \leq \text{age} \leq 29)$	-0.184	-0.302	-0.496	-0.423 *	-0.110	-0.233
$NUD \times D(30 \leq \text{age} \leq 34)$	-0.237	-0.369	-0.551	-0.451 *	-0.134	-0.275
$NUD \times D(35 \leq \text{age} \leq 39)$	-0.265	-0.401	-0.603	-0.450 *	-0.186	-0.311
$NUD \times D(40 \leq \text{age} \leq 44)$	-0.288	-0.430	-0.593	-0.459 *	-0.166	-0.328
$NUD \times D(45 \leq \text{age} \leq 49)$	-0.278	-0.412	-0.585	-0.446 *	-0.192	-0.328
$NUD \times D(50 \leq \text{age} \leq 54)$	-0.268	-0.402	-0.534	-0.444 *	-0.143	-0.321
$NUD \times D(55 \leq \text{age} \leq 59)$	-0.264	-0.404	-0.527	-0.445 *	-0.146	-0.302
# Cells	1545	1545	1400	1545	1400	1545

Estimations by weighted least squares, cells weighted by cell-specific employment. Exclusion of cells censored at the 50% quantile (80% quantile for $QD8020$ and $QD8050$). Covariates: dummies for time, age, industry, and skill. * coefficients not significantly different ($\alpha = 0.05$, variances White robust). Data source: extended IABS 1985–1997.

With growing age, the extent of the reductions in wage levels among the medium-skilled increases and the union-inflicted mark-up among the low-skilled decreases. In both skill groups younger workers benefit more considerably from stronger unions. There is no evidence for seniority bonuses paid to older workers or for premia to older cohorts which are more closely related to trade unions. On the contrary, the result is consistent with an insurance motive securing wage smoothing over the life cycle.

The effect towards residual wage compression among the low-skilled also decreases slightly with age. With the exception of the oldest age group, however, the differences are generally small or even insignificant. A similar finding applies to the wage compression in the lower half of the distribution ($QD5020$) among the medium-skilled. In this group, the wage-compressing effect is slightly lower for younger as well as for older workers compared to those of age 35–39, but most of the differences stem from differences in the effects on $QD8050$. Summing up, the union impact concerning the equality objective declines slightly with age, mostly as a result of reduced wage compression in the upper half of the distribution. This finding is consistent with the notion that unions do in fact take account of the ex-post rationality of membership to some extent.

In order to analyze whether the union effects on the wage structure have changed over time, we additionally run regressions which include interactions of NUD and year dummies. Figure 5.1 depicts the changes of the NUD coefficients over time, relative to the base period 1985.

Most of the individual effects are not significant. In light of the joint significance of the differences across years (especially with respect to $LNW50$, $QD5020$, and $\sigma TOBIT$) it is nevertheless worthwhile to look at the developments in more detail. To a large extent, the effects on the wage level, LNW and $LNW50$, evolve parallely to the effects on wage dispersion, $QD8020$ and $\sigma TOBIT$. There is a general trade-off between the two union objectives regarding higher wages and the reduction of wage inequality. Until 1989 the unions put increasing weight on wage levels, but the equality objective got an increasing weight during the subsequent years until 1994. Starting from 1990 this development was accompanied by a decrease in union membership (compare section 5.4.2). However, while the decline in union membership even continued to grow in strength after 1994, the effect on the wage level increased again starting from 1995. Unlike the effect on $QD8020$, the influence of NUD on $\sigma TOBIT$ continued to change towards decreasing dispersion over the years 1994 to 1997. Generally, larger parts of the changes over time happened in the lower half of the conditional distribution ($QD5020$), while only small changes are observed in the upper half ($QD8050$) after 1994.

Finally, we analyze whether the correlations estimated on the basis of both longitudinal and cross-sectional variation are compatible with observed long differences. Does the decline in union membership comply with the results of tables 5.1 and 5.2 over the entire period 1985–1997—i. e., does it come along with an increasing wage inequality and a rise in the wage level? In order to illustrate this point, we regress changes of our wage measures between 1985 and 1997, ΔLNW etc., on a constant and the change in net union density over the same period, ΔNUD. The approach filters out industry-, age-, and, if applicable, skill-specific effects. Table 5.4 displays the results.

Table 5.4: Changes of Net Union Density and the Structure of Wages, 1985–1997

Segmentation z	ΔLNW	$\Delta LNWM$	$\Delta LNW50$	ΔBL	ΔBH	$\Delta\sigma TOBIT$
ΔNUD	0.182	0.030	0.096	0.496	-0.101	-0.079
	(0.100)	(0.100)	(0.116)	(0.131)	(0.195)	(0.068)
# Differences	119	119	119	119	113	119
Segmentation $\tilde{z}$	ΔLNW	$\Delta LNW50$	$\Delta QD8020$	$\Delta QD5020$	$\Delta QD8050$	$\Delta\sigma TOBIT$
ΔNUD	0.105	-0.036	-0.453	-0.273	-0.173	-0.150
	(0.079)	(0.092)	(0.123)	(0.081)	(0.062)	(0.054)
# Differences	249	249	219	249	219	249

Estimations by weighted least squares, differences weighted by employment average of 1985 and 1997. Exclusion of cells censored at the 50% quantile (80% quantile for $QD8020$ and $QD8050$). Segmentation z: industry×age×time cells; segmentation $\tilde{z}$: industry×age×skill×time cells. White robust standard errors in parentheses. Data source: extended IABS 1985–1997.

By and large, the results for the measures of wage dispersion meet the implications of tables 5.1 and 5.2. In particular, the decline in union density over the observation period came along with significant reductions of both residual wage dispersion and the wage differential between low- and medium-skilled workers. However, there is no significant effect with respect to the wage level—the negative effect estimated from the combination of longitudinal and cross-sectional data is not maintained when relying on long differences only.

5.5 Conclusions

Net union density is an appropriate proxy for union power because the associated financial power and the personal representation within firms govern the union's threat point in the collective wage bargaining process and therefore determine the bargaining outcome. In order to analyze empirically the correlation between union power and the structure of wages in the German labor market, we thus impute union membership propensities for individuals in the IAB employment sample (IABS) 1985–1997 from the estimations of Beck and Fitzenberger (2004) and compute net union densities for labor market segments corresponding to the structure of the German wage bargaining system.

According to an insurance motive for union membership as discussed by Agell and Lommerud (1992) and Burda (1995), a higher net union density should ceteris paribus come along with reduced residual wage dispersion. If unions have a strong preference for wage equality and also want to prevent negative employment effects, a higher net union density can even be accompanied by a lower wage level. Our empirical results corroborate the insurance argument with respect to negative effects on the wage level and on wage dispersion within and between labor market segments. Even beyond "equal pay for equal work", skill wage differentials are also lower in segments with powerful unions. In line with a minimum wage interpretation of union-bargained wages, the wage distribution is compressed disproportionately from below—powerful unions ceteris paribus increase the skewness of the distribution. When looking at the age dimension we also find a wage smoothing effect or, put differently, an insurance motive over the life cycle. Neither is there evidence for the conjecture that unions counteract ex post rational membership quits of older workers by bargaining for seniority bonuses, nor do unions privilege older cohorts which stand more closely by the unions. However, there is some evidence that the compression of wages in the upper parts of the distribution decreases with age. Union effects further vary over time, but there are no clear trends against the background of the perpetual decline in union membership since 1990.

The literature also discusses employment effects of union representation which our study does not consider explicitly. Yet our results imply that higher union power can have a repressive effect on employment if the inflexibility associated with excessive wage compression leads to higher unemployment. According to Schettkat (2003), empirical cross-country evidence on the relationship between wage inequality and employment performance is mixed. However, in Fitzenberger and Kohn (2006) we find that a higher degree of wage dispersion between skill groups would contribute to a reduction of skill-specific unemployment rates in Germany.

Although our results are surprising at first glance, they are not in contrast to the positive correlation between collective bargaining coverage and the wage level as discussed in the literature. Union power and collective coverage do not necessarily coincide. Wages paid by firms without collective coverage are often geared to collective agreements in the corresponding labor market segment, and the wage policy implemented in collective agreements crucially depends on the power of the union.

Unfortunately, our regressions can not take account of the apparent endogeneity of union density, and so the results should not be interpreted as causal effects. The implementation of structural models proves to be intricate (compare Fitzenberger (1999), chapter 6) and it is awkward to find appropriate instruments. While we consider the insurance motive an intuitively plausible explanation for the descriptive evidence, further research to investigate the robustness of the results is certainly warranted.

Valuable insights are to be expected from an analysis on the basis of linked employer-employee data which would capture the effects of both union density *and* collective bargaining coverage.

Finally, union effects on the wage structure *and* on employment should be analyzed simultaneously.

5.A Data

Our empirical analysis is based on the administrative IAB employment sample (IABS) 1975–1997, a representative 1% random draw of German employees with employment spells subject to social insurance contributions. Excluding civil servants, self-employed, and freelancers, the IABS covers about 80% of all employed persons. For an extensive description of these register-based data see Bender, Hilzendegen, Rohwer, and Rudolph (1996) and Bender, Haas, and Klose (2000). Selected data comprise spells of men who are employed full-time in West Germany (excluding Berlin), not in training, and without parallel employment.

Individuals aged between 25 and 60 years are classified into five-year age groups (25–29 years, 30–34 years, ..., 55–59 years). With respect to the individuals' educational attainment, we construct three skill groups: The group of the low-skilled consists of employees without any vocational training. Those with a vocational training degree are considered medium-skilled, and individuals with a university or technical college degree form the group of the high-skilled. To deal with measurement error in the education information when defining the skill groups, we correct the skill information such that formal degrees an individual has once obtained are not lost later.

Since the IABS contains no information on hours worked, we undertake a headcount to derive a measure of cell-specific employment, weighting each observation with the length of the respective employment spell. This procedure assumes that the number of (weekly) working hours does not change over time and does not differ between individuals. It therefore justifies our concentration on full-time employees only.

We extend the IABS by imputing individual propensities for union membership from Beck and Fitzenberger (2004), who estimate determinants of union membership using survey data from the German Socio-Economic Panel (GSOEP) for different years between 1985 and 1998. We can thus use the years 1985–1997 of the extended data set. Net union densities are obtained by means of aggregation at the cell level.

At any rate, our cell-data analysis is restricted to cells with a minimum of 20 observations. Table 5.7 provides information about the cell size for the different segmentations.

5.B Additional Tables and Figures

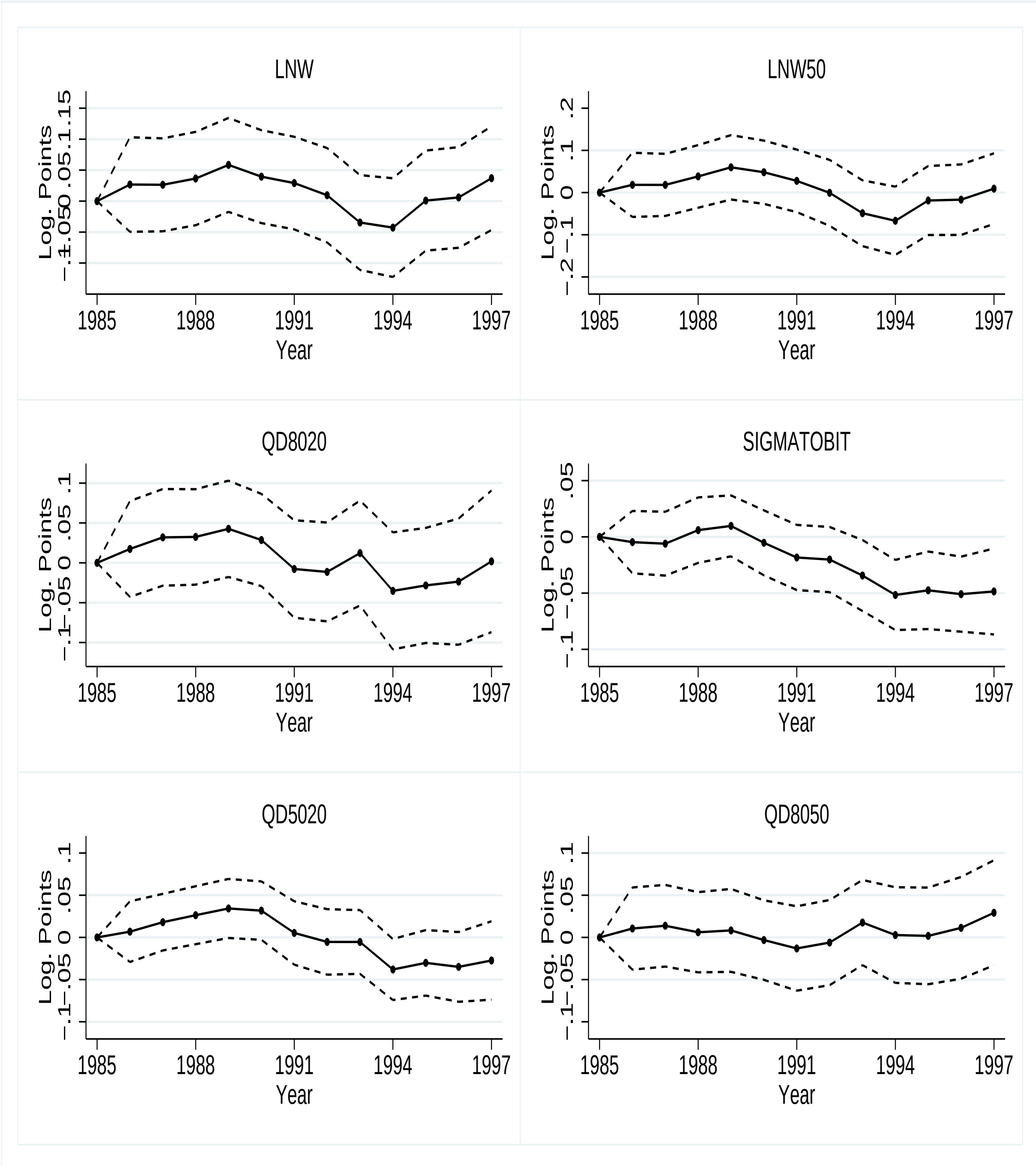
Table 5.5: Industry Classification and Grouping of Sectors Used in the Empirical Analysis

No.[a]	Industry	WZW[b]
01	Agriculture, Forestry, Fishing, Mining, Energy and Water Supply	01–08
02	Chemical Products	09–11
03	Synthetic Materials	12–13
04	Stone and Earth Products	14–16
05	Iron and Steel, Machinery, Vehicle Construction	17–32
06	Electric Appliances, Precision Instruments and Optics	33–39
07	Woodwork, Paper, Printing, Publishing	40–44
08	Textiles and Apparel	45–53
09	Food, Beverages, and Tobacco	54–58
10	Construction, Building Expenses, and Conversion	59–61
11	Wholesale Trade, Commercial Mediation, Retail Trade	62
12	Railways, Postal Services	63–64
13	Other Transportation	65–68
14	Financial Intermediation	69
15	Hotels and Restaurants, Personal Services, Private Non-profit Organizations, Private Households, Government, Social Insurance, Other Services	70–73, 79–99
16	Education and Science	74–77
17	Health Care System	78

[a] Sector classification used in the empirical analysis.

[b] Industry classification provided with the IABS 1975–1997.

Figure 5.1: Net Union Density and the Structure of Wages IV: Variation over Time



Estimations by weighted least squares, cells weighted by cell-specific employment. Deviations of NUD-coefficients relative to effects in the base period 1985; 95%-confidence intervals (White robust). Exclusion of cells censored at the 50%-quantile (80%-quantile for $QD8020$ and $QD8050$); no high-skilled. Covariates: dummies for time, age, industry, and skill; interactions $NUD\times$skill (base: medium-skilled) and $NUD\times$age. Data source: extended IABS 1985–1997.

Table 5.6: Decomposition of Net Union Density

	NUD		NUD	
D(agriculture, mining, energy)	-0.035	(0.005)	-0.035	(0.004)
D(chemical products)	-0.039	(0.002)	-0.030	(0.003)
D(synthetic materials)	-0.163	(0.002)	-0.166	(0.002)
D(stone and earth)	-0.170	(0.002)	-0.174	(0.002)
D(iron and steel, machinery, vehicles)	base		base	
D(electric appliances)	-0.208	(0.004)	-0.198	(0.004)
D(paper, printing)	-0.105	(0.002)	-0.110	(0.002)
D(textiles)	-0.087	(0.002)	-0.093	(0.002)
D(food, beverages)	-0.193	(0.002)	-0.198	(0.002)
D(construction)	-0.284	(0.002)	-0.288	(0.002)
D(trade)	-0.311	(0.003)	-0.311	(0.003)
D(railways, postal services)	0.289	(0.005)	0.283	(0.005)
D(other transportation)	-0.240	(0.002)	-0.245	(0.002)
D(financial intermediation)	-0.366	(0.003)	-0.357	(0.003)
D(misc. services)	-0.244	(0.003)	-0.230	(0.002)
D(education, science)	-0.291	(0.002)	-0.254	(0.003)
D(health care)	-0.325	(0.002)	-0.297	(0.003)
$D(25 \leq \text{age} \leq 29)$	base		base	
$D(30 \leq \text{age} \leq 34)$	0.023	(0.003)	0.030	(0.003)
$D(35 \leq \text{age} \leq 39)$	0.044	(0.003)	0.051	(0.003)
$D(40 \leq \text{age} \leq 44)$	0.062	(0.003)	0.066	(0.003)
$D(45 \leq \text{age} \leq 49)$	0.078	(0.003)	0.080	(0.003)
$D(50 \leq \text{age} \leq 54)$	0.091	(0.003)	0.091	(0.003)
$D(55 \leq \text{age} \leq 59)$	0.090	(0.003)	0.088	(0.003)
$D(\text{time} = 1985)$	base		base	
$D(\text{time} = 1986)$	-0.004	(0.003)	-0.004	(0.003)
$D(\text{time} = 1987)$	-0.007	(0.003)	-0.005	(0.003)
$D(\text{time} = 1988)$	-0.008	(0.003)	-0.006	(0.003)
$D(\text{time} = 1989)$	-0.009	(0.003)	-0.006	(0.003)
$D(\text{time} = 1990)$	-0.017	(0.003)	-0.014	(0.003)
$D(\text{time} = 1991)$	-0.024	(0.003)	-0.021	(0.002)
$D(\text{time} = 1992)$	-0.030	(0.003)	-0.025	(0.002)
$D(\text{time} = 1993)$	-0.034	(0.003)	-0.029	(0.003)
$D(\text{time} = 1994)$	-0.040	(0.003)	-0.035	(0.003)
$D(\text{time} = 1995)$	-0.047	(0.004)	-0.042	(0.003)
$D(\text{time} = 1996)$	-0.054	(0.004)	-0.049	(0.004)
$D(\text{time} = 1997)$	-0.061	(0.005)	-0.055	(0.004)
D(low-skilled)			0.020	(0.002)
D(medium-skilled)			base	
D(high-skilled)			-0.124	(0.002)
Intercept	0.462	(0.004)	0.460	(0.004)
N	1547		4119	
R^2	0.978		0.960	

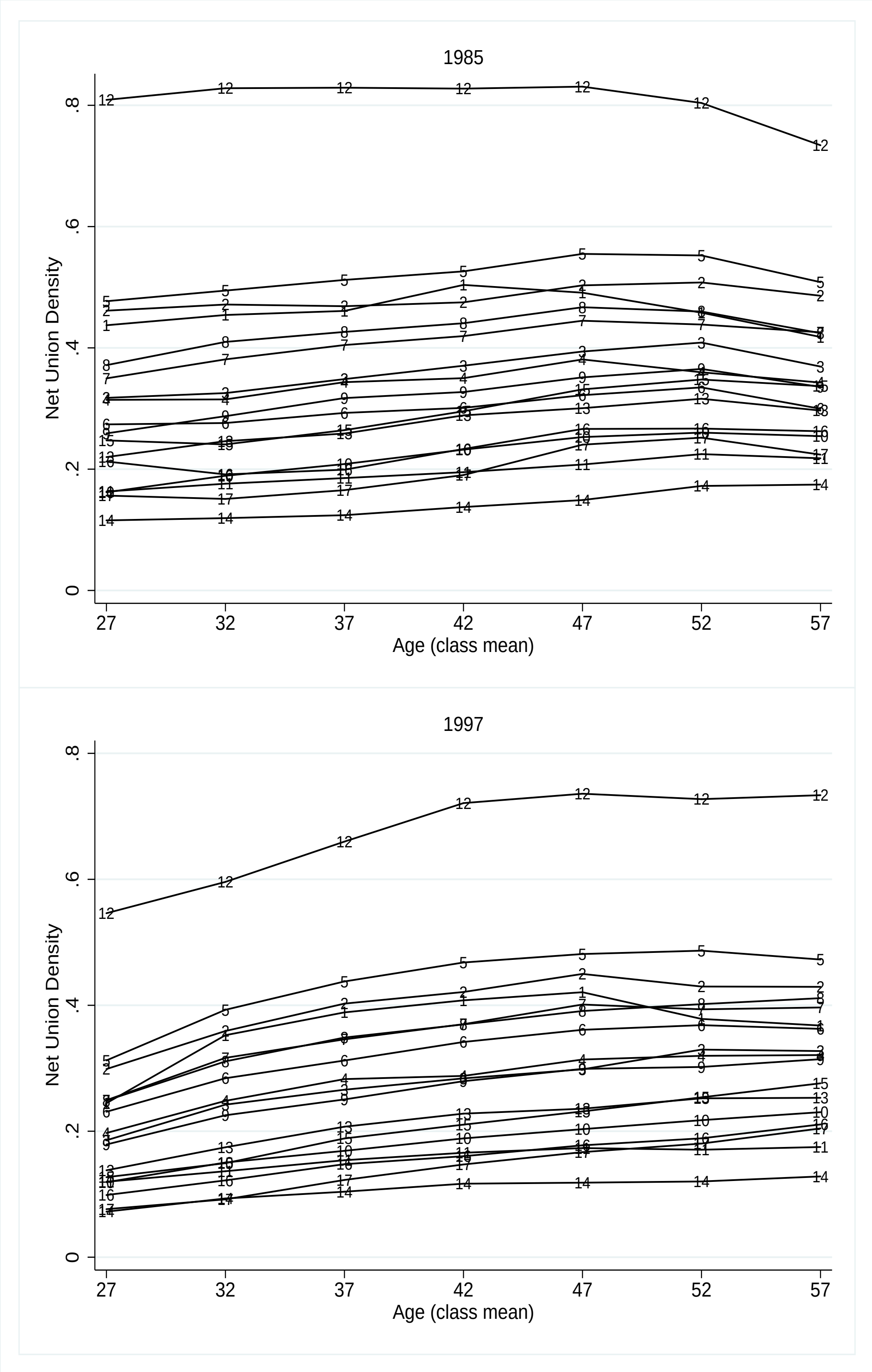
Estimations by weighted least squares, cells weighted by cell-specific employment. White robust standard errors in parentheses. See table 5.5 for a complete definition of the industry sectors. Data source: extended IABS 1985–1997.

Table 5.7: Cell Size for the Different Segmentations

Segmentation industry×age×time	# Cells	# Observations in Cells Mean	Std.Dev.	Min.	Max.
Total	1547	1040.6	932.7	113	5024
$LNW50$ not censored	1545	1040.6	933.3	113	5024
$LNW80$ not censored	1235	1092.5	1000.7	113	5024
Segmentation industry×age×skill×time	**# Cells**	**# Observations in Cells Mean**	**Std.Dev.**	**Min.**	**Max.**
Total	4119	389.3	549.9	20	4091
$LNW50$ not censored	3442	444.7	584.3	20	4091
$LNW80$ not censored	3032	455.5	605.5	20	4091
Low-Skilled	1474	168.2	148.4	20	952
$LNW50$ not censored	1474	168.2	148.4	20	952
$LNW80$ not censored	1441	170.9	149.0	20	952
Medium-Skilled	1547	783.1	722.6	75	4091
$LNW50$ not censored	1545	783.0	723.1	75	4091
$LNW80$ not censored	1400	787.7	749.2	75	4091
High-Skilled	1098	131.4	136.3	20	976
$LNW50$ not censored	423	172.1	183.9	20	976
$LNW80$ not censored	191	167.9	191.2	20	976

Data Source: extended IABS 1985–1997.

Figure 5.2: Evolution of Net Union Density



Calculations based on the extended IABS 1985–1997. Numbers in the diagrams indicate industry sectors; see table 5.5 for reference.

6 Deutschsprachige Zusammenfassung

Der deutsche Arbeitsmarkt ist seit einigen Jahrzehnten durch steigende Arbeitslosigkeit gekennzeichnet, die insbesondere gering qualifizierte Arbeitnehmer betrifft. So betrugen im Jahr 2004 die qualifikationsspezifischen Arbeitslosenquoten in Westdeutschland für Personen ohne Berufsabschluss 21,7%, für jene mit Ausbildungsabschluss 7,3% und für Hochschulabsolventen 3,5%. Zudem lagen die entsprechenden Quoten in Ostdeutschland mit 51,2%, 19,4% und 6,0% weit über dem westdeutschen Niveau (Reinberg und Hummel, 2005). Eine Mitschuld an dieser Situation wird verschiedenen Arten von Rigiditäten zugeschrieben, die flexiblen Anpassungen auf dem Arbeitsmarkt entgegen stehen.

Eine zentrale Rolle für die Anpassung der Beschäftigung und die Entwicklung der Arbeitslosigkeit kommt der Lohnstruktur oder den Lohnverteilungen für verschiedene Arbeitnehmergruppen zu. Seit einiger Zeit stehen neben dem Lohnniveau auch Lohndispersion oder -kompression zwischen und innerhalb von homogenen Arbeitnehmergruppen im Zentrum arbeitsmarktökonomischen Interesses. Mit der zunehmenden Verfügbarkeit von großen Mikrodatensätzen widmet sich eine schnell wachsende Anzahl empirischer Untersuchungen mit mikroökonometrischen Methoden unterschiedlichen Facetten heterogener Lohnstrukturen. Besonderes Augenmerk gilt dabei den Erträgen aus Humankapitalinvestitionen und der Entwicklung von Qualifikationsprämien (Katz und Autor, 1999).

Die Entwicklung von Lohnstrukturen in Westdeutschland seit den 1970er Jahren bis zur Mitte der 1990er ist in dieser Hinsicht intensiv untersucht worden, wobei sich die Struktur im großen und ganzen als im internationalen Vergleich wenig gespreizt und über die Zeit hinweg recht stabil erwiesen hat. Sowohl Qualifikationsprämien als auch residuale Lohnungleichheiten zeigten eine bemerkenswert geringe Variation. Andererseits sank über den gleichen Zeitraum die relative Nachfrage nach gering qualifizierten Arbeitnehmern – etwa durch qualifikationsverzerrten technischen Fortschritt – schneller als das relative Angebot, so dass für eine Markträumung steigende Qualifikationsprämien erforderlich gewesen wären. Insofern wird der "unerträglichen Stabilität" (Prasad, 2004) eine Schlüsselrolle für den asymmetrischen Anstieg der Arbeitslosigkeit zugeschrieben.

In engem Zusammenhang damit wird der Einfluss von institutionellen und organisatorischen Rahmenbedingungen auf die wirtschaftliche Entwicklung im allgemeinen und auf die Lohnsetzung auf dem Arbeitsmarkt im speziellen diskutiert. In Zeiten zunehmend heterogener wirtschaftlicher Bedingungen lautet der Tenor in der arbeitsmarktökonomischen Debatte, dass institutionelle Inflexibilitäten das Funktionieren der Arbeitsmärkte und damit die Dynamik der Beschäftigungsentwicklung hemmen (OECD, 2006). Im Mittelpunkt des Interesses steht nicht zuletzt die Bedeutung der Gewerkschaften.

Gewerkschaftliche Organisationsgrade haben im Laufe der letzten Jahrzehnte in vielen Industrieländern einschließlich Deutschlands beträchtlich abgenommen. Im Jahr 2004 etwa befand sich der Bruttoorganisationsgrad, d. h. der Anteil aller Gewerkschaftsmitglieder an der Gesamtzahl der Beschäftigten, für Deutschland auf dem historischen Tiefstand von 27% (Fitzenberger, Kohn und Wang, 2006). Zu den Determinanten dieses Phänomens wie auch den daraus resultierenden Konsequenzen existiert eine breite Literatur (Addison und Schnabel, 2003). Wenngleich die Gewerkschaften – von auf der Betriebsebene abgeschlossenen Beschäftigungspakten vielleicht abgesehen – keinen direkten Einfluss auf die Personalpolitik von Arbeitgebern besitzen, beeinflussen sie die Beschäftigung auf dem Arbeitsmarkt im Zuge von Tariflohnverhandlungen, also mittels ihres Einflusses auf die Lohnstruktur. So mag etwa Lohnzurückhaltung zu mehr Beschäftigung führen, während hohe Tariflohnsteigerungen beschäftigungsschädlich wirken können; und ein starkes gewerkschaftliches Gleichheitsziel, welches sich in einer Stauchung der Lohnverteilung zwischen gering und höher qualifizierten Arbeitnehmern niederschlägt, kann wiederum die Beschäftigungsaussichten der gering Qualifizierten schmälern. Grundsätzlich ist angesichts des beobachteten Trends über die Zeit abnehmender Organisationsgrade und der damit einhergehenden Einbuße gewerkschaftlicher Verhandlungsmacht eine Verringerung der Gewerkschaftseinflüsse zu erwarten.

Die vorliegende Arbeit basiert auf vier eigenständigen Aufsätzen, die sich dem Spannungsfeld von Lohnstrukturen, heterogener Arbeitsnachfrage und Gewerkschaftsmacht auf dem deutschen Arbeitsmarkt mit Hilfe mikroökonometrischer Methoden nähern. Die folgenden Abschnitte fassen die Ansätze und die Kernergebnisse dieser Arbeiten zusammen und geben Ausblicke auf weiterführende Forschungsfragen.

Zur Entwicklung der Lohnungleichheit

Der erste Aufsatz (Kohn, 2006) zeichnet ein detailliertes Bild der Entwicklung der Lohnungleichheit in Ost- und Westdeutschland in den Jahren des ausgehenden 20. und des beginnenden 21. Jahrhunderts.

Bis Anfang der 1990er Jahre zeichnete sich der westdeutsche Arbeitsmarkt durch eine im internationalen Vergleich niedrige und lediglich geringen Veränderungen über die Zeit unterworfene Lohnstreuung aus, und in Ostdeutschland war in den späten Jahren der DDR ein noch höheres Maß an Lohnkompression zu beobachten. Neuere Daten der IAB-Beschäftigtenstichprobe (IABS) 1975–2001 lassen nun Analysen für jüngere Jahre zu. Der vorliegende Aufsatz untersucht die Entwicklung von Lohnniveau und Lohndispersion zwischen und innerhalb von verschiedenen Gruppen von Arbeitnehmern für den Zeitraum zwischen 1992 und 2001. Als zentrales Resultat lässt sich festhalten, dass über diesen Zeitraum die Lohnungleichheit entlang verschiedener Dimensionen zugenommen hat.

Zunächst offenbaren unbedingte jahresspezifische Lohnverteilungen, dass die Lohnstreuung sowohl in Ost- als auch in Westdeutschland insgesamt angestiegen ist. Ausgehend von einem geringeren Niveau war der Anstieg in Ostdeutschland stärker als in Westdeutschland, so dass im Jahr 2001 beide Landesteile kaum mehr Unterschiede im Ausmaß der Lohnstreuung aufwiesen. Grundsätzlich war der Anstieg bei Frauen ausgeprägter als bei Männern, wobei der Großteil der Zunahme bei Frauen im unteren Bereich der Lohnverteilung, bei Männern jedoch im oberen Bereich stattgefunden hat. Während in den ersten Jahren nach der Wiedervereinigung eine Konvergenz der Lohnniveaus in Ost- und Westdeutschland zu verzeichnen war, findet seit 1996 keine weitere Angleichung mehr statt, so dass beachtliche Ost-West-Gefälle im Lohnniveau fortbestehen.

Die Schätzung zensierter Quantilsregressionen liefert daraufhin Erkenntnisse hinsichtlich zentraler Determinanten der beobachteten Lohnverteilungen, wie etwa Alter und Qualifikationsniveau der Beschäftigten. Der Humankapitaltheorie entsprechend, zeigen sich Alters-Lohn-Profile umso steiler, je höher das Qualifikationsniveau der Beschäftigten ist. Zudem waren die Profile für Ostdeutschland im Jahr 1992 besonders flach. Offenbar führte der Wiedervereinigungsschock im Osten zu einer Abschreibung von Humankapital. Mit dem Eintreten jüngerer Kohorten in den Arbeitsmarkt verliert dieser Effekt jedoch an Bedeutung, so dass sich die Profile in Ost- und Westdeutschland für 2001 in stärkerem Maße ähneln.

Unterschiede in den Lohnverteilungen zwischen Ost- und Westdeutschland sowie entsprechende Veränderungen über die Zeit lassen sich mit Hilfe von Zerlegungstechniken näher untersuchen. Ich verwende Machado-Mata-Dekompositionen (Machado und Mata, 2005), die auf der Schätzung der flexiblen Quantilsregressionen aufbauen und über die Lohnverteilung heterogene Charakteristika- und Koeffizienten-Effekte offenbaren, welche Unterschiede in der Zusammensetzung der Beschäftigten (Charakteristika-Effekt) beziehungsweise Unterschiede im geschätzten Einfluss der Beschäftigtenmerkmale (Koeffizienten-Effekt) auffangen. Im Ost-West-Vergleich leisten unterschiedliche Charakteristika bei Männern praktisch keinen Beitrag zur Erklärung des Lohngefälles. Bei vollzeitbeschäftigten Frauen jedoch spricht der

Charakteristika-Effekt grundsätzlich für höhere Löhne in Ostdeutschland; nur am unteren Ende der Lohnverteilung ist dieser Effekt im Jahr 2001 nicht mehr zu beobachten. Hinsichtlich der Veränderung der Löhne über die Zeit tragen Änderungen in der Zusammensetzung der Beschäftigten einem Großteil der Lohnzuwächse in den oberen Teilen der Verteilungen für Westdeutschland Rechnung. Dieses Ergebnis spiegelt insbesondere die Zunahme des durchschnittlichen Qualifikationsniveaus der Beschäftigten wider. Veränderungen des Qualifikationsniveaus und Verschiebungen in der sektoralen Beschäftigungsstruktur erklären allerdings nur einen geringen Teil der Lohnzuwächse in Ostdeutschland. Am unteren Ende der Lohnverteilung für Frauen in Ostdeutschland wirkte der Charakteristika-Effekt sogar in Richtung realer Lohnkürzungen, was den in dieser Gruppe besonders ausgeprägten Anstieg der Lohnungleichheit erhärtete.

In dem ersten Aufsatz steht die detaillierte empirische Bestandsaufnahme, nicht jedoch eine tiefer greifende ökonomische Analyse möglicher Gründe und Konsequenzen der aufgezeigten Entwicklungen in der Lohnstruktur im Vordergrund. Angesichts alternativer Erklärungsansätze, wie etwa sich beschleunigenden, nicht-neutralen technischen oder organisatorischen Wandels, zunehmender internationaler Verflechtung, abnehmender Gewerkschaftsmacht oder flexiblerer Arbeitsmarktinstitutionen, gilt es die deskriptive Evidenz durch die Schätzung struktureller Modelle oder die Verwendung noch aussagekräftigerer Datensätze zu komplementieren. Die nachfolgenden Aufsätze unternehmen Schritte in dieser Richtung.

Qualifikationsprämien, Kohorteneffekte und heterogene Arbeitsnachfrage

Der zweite Aufsatz (Fitzenberger und Kohn, 2006) untersucht den Zusammenhang zwischen Lohnstrukturen und Beschäftigung im Rahmen eines Arbeitsnachfragemodells für Westdeutschland, welches Qualifikationsniveau und Alter der Beschäftigten als wichtige Dimensionen der Heterogenität des Faktors Arbeit inkorporiert.

Zunächst zeigt die Schätzung altersspezifischer Qualifikationsprämien anhand der IABS 1975–1997, dass die Altersprofile von Lohndifferenzialen sich im Zeitverlauf nicht parallel entwickelt, sondern vielmehr eine Verdrehung erfahren haben. Dementsprechend ist es unwahrscheinlich, dass die Entwicklungen in der Lohnstruktur allein Alters- und Zeiteffekten zuzuschreiben sind, welche alle Arbeitnehmerkohorten gleichermaßen beeinflusst haben. Zudem finden wir einen Bruch in altersspezifischen Qualifizierungstrends dergestalt, dass jüngere Geburtsjahrgänge dem Pfad der älteren in Richtung eines zunehmenden durchschnittlichen Qualifikationsniveaus nicht mehr folgen. Insofern suggeriert die empirische Evidenz, dass Kohorteneffekte sowohl Qualifikationsprämien im Lohn als auch die relative Beschäftigung von Qualifikationsgruppen beeinflusst haben. Wir weisen die Existenz solcher Effekte anhand des Testverfahrens von MaCurdy und Mroz (1995) nach.

Die entsprechende Heterogenität des Faktors Arbeit erschwert die Operationalisierung von Löhnen und Beschäftigung in einem Arbeitsnachfragemodell. Wir erweitern deshalb den strukturellen Ansatz von Card und Lemieux (2001), in deren zweistufigem CES Modell die simultane Berücksichtigung von Qualifikationsniveau und Alter nicht nur die Separation von Alters-, Zeit- und Kohorteneffekten erlaubt, sondern auch die Schätzung von nachfragetheoretisch konsistenten Modellspezifikationen mit einer relativ großen Anzahl verschiedener Einsatzfaktoren ermöglicht. Darüber hinaus berücksichtigt das Modell qualifikationsverzerrten technischen Wandel. Wir schätzen Modellvarianten mit und ohne Instrumentierung, welche der Endogenität von Löhnen und Beschäftigung Rechnung trägt, und erhalten Substitutionselastizitäten von 4,9 bis 6,9 für die Substitution zwischen den Qualifikationsklassen sowie von 5,2 bis 20,1 für die Substitution zwischen den Altersklassen. Im Vergleich zu den Ergebnissen anderer Studien in der Literatur sind diese Werte recht hoch und spiegeln die in Deutschland insgesamt geringe Lohndispersion sowie ein relativ geringes Maß an Qualifikationsunterschieden zwischen den jeweiligen Ausbildungs- und Erfahrungskategorien wider.

Auf Basis der geschätzten Substitutionsparameter lassen sich Simulationsexperimente zum Zusammenhang von Lohnstruktur und Beschäftigung durchführen. Im besonderen simulieren wir Lohnänderungen, die in den einzelnen Qualifikationsklassen notwendig gewesen wären, um die qualifikationsspezifischen Arbeitslosenquoten im Jahr 1997 um die Hälfte zu reduzieren. Dabei nehmen wir identische Änderungen für alle Altersklassen an, halten also die relativen Löhne innerhalb der Qualifikationsklassen unverändert. Die notwendigen Nominallohnsenkungen liegen zwischen 8,8 und 12,2% und fallen umso größer aus, je geringer das Qualifikationsniveau der Beschäftigten ist. Gemessen an einer Situation mit niedrigerer Arbeitslosigkeit war die Lohndispersion zwischen den Qualifikationsklassen also zu gering.

Unsere Analyse unterstreicht die Notwendigkeit, verschiedene Dimensionen von Heterogenität in aussagekräftige empirische Arbeitsnachfragemodelle zu integrieren. Die aus dem zweistufigen CES Ansatz resultierenden Ergebnisse führen zu einer Anzahl von interessanten Ansatzpunkten für zukünftige Forschungsfragen: Worin liegt beispielsweise der hohe Grad an Substituierbarkeit der Beschäftigten auf dem deutschen Arbeitsmarkt begründet? Woher rührt der beobachtete Bruch in der Zunahme des durchschnittlichen Qualifikationsniveaus? Und hat sich dieser Trend in jüngeren Jahren fortgesetzt? Die Literatur zu technischem und organisatorischem Wandel etwa weist darauf hin, dass jenseits der formalen Qualifikation der Beschäftigten die Tätigkeitsmerkmale per se in den letzten Jahren starken Veränderungen unterworfen waren (Autor, Levy und Murnane, 2003; Spitz-Oener, 2006). Findet auch in Deutschland eine zunehmende Polarisierung auf dem Arbeitsmarkt in gute und schlechte Jobs statt, wie von Goos und Manning (2003) für

Großbritannien beobachtet? Und welche Rolle kommt den Arbeitsmarktinstitutionen bei den jeweiligen Entwicklungen zu?

Zur Entwicklung der Gewerkschaftsmitgliedschaft

Einer zentralen Arbeitsmarktinstitution wendet sich der dritte Aufsatz (Fitzenberger, Kohn und Wang, 2006) zu, der die Entwicklung der Gewerkschaftsmitgliedschaft in West- und Ostdeutschland untersucht. Gewerkschaften setzen sich für höhere Löhne, gerechte Bezahlung, Verringerung der Wochenarbeitszeit, faire Arbeitsbedingungen oder etwa verstärkten Kündigungsschutz ein. Von den Errungenschaften gewerkschaftlichen Engagements profitieren in Deutschland allerdings nicht nur Gewerkschaftsmitglieder, sondern der Großteil aller Beschäftigten, da die im Grundgesetz verankerte negative Koalitionsfreiheit expliziten und impliziten Mitgliedschaftszwang verbietet. Die zentralen Gewerkschaftsleistungen haben daher den Charakter eines öffentlichen Gutes und sind Trittbrettfahrerverhalten seitens der Arbeitnehmerschaft ausgesetzt.

Warum treten also Arbeitnehmer überhaupt einer Gewerkschaft bei? Unsere Studie nutzt Daten des Deutschen Sozio-oekonomischen Panels (SOEP), um Determinanten individueller Mitgliedschaftsentscheidungen im Zeitraum 1985–2003 (Westdeutschland) beziehungsweise 1993–2003 (Ostdeutschland) zu schätzen. Dieser Ansatz besitzt dreierlei Vorzüge. Zum ersten quantifizieren unsere Ergebnisse die Einflüsse soziodemographischer individueller Eigenschaften, wie etwa Alter oder Familienstand; die Einflüsse von Arbeitsplatzcharakteristika, d. h. match-, firmen- oder sektorspezifische Effekte; sowie den Einfluss persönlicher Einstellungen auf die Entscheidung für oder gegen eine Gewerkschaftsmitgliedschaft. Dabei lassen die verwendeten Probit-Modelle mit korrelierten zufälligen Effekten in der Tradition von Chamberlain (1980) und Mundlack (1978) nicht nur Unterschiede in den Einflüssen zwischen Ost- und Westdeutschland und zwischen den einzelnen Jahren zu, sondern berücksichtigen auch unbeobachtete Heterogenität.

Zum zweiten verwenden wir die Schätzergebnisse, um gewerkschaftliche Nettoorganisationsgrade für Ost- und Westdeutschland zu projizieren, welche jeweils den Anteil beschäftigter Gewerkschaftsmitglieder an der Gesamtzahl der Beschäftigten angeben. Die Projektionen, die den Rückgang der Gewerkschaftsmitgliedschaft in beiden Landesteilen konsistent nachzeichnen, lassen sich dann mit Hilfe von Zerlegungstechniken näher untersuchen. Wir passen Blinder-Oaxaca-Dekompositionen (Blinder, 1973; Oaxaca, 1973) für unseren nichtlinearen Schätzansatz an und analysieren die Unterschiede im Organisationsgrad zwischen Ost- und Westdeutschland sowie die entsprechenden Veränderungen über die Zeit. Zum einen kommt Veränderungen in der Zusammensetzung der Arbeitnehmerschaft nur eine untergeordnete Rolle bei der Erklärung der Trends zu abnehmender Gewerkschaftsmitgliedschaft zu. Zum anderen sprechen im Ost-West-Vergleich Arbeitnehmer- und Arbeitsplatzeigenschaften zu jedem Zeitpunkt für einen höheren Organisationsgrad in Westdeutschland. Der tatsächlich höhere Organisationsgrad in Ostdeutschland im Jahr 1993 und der sich anschließende stärkere Rückgang in diesem Landesteil reflektieren daher eine schwächere Bindung der Mitglieder, welche aus der breit gestreuten, umgehenden Rekrutierung direkt nach der Wiedervereinigung resultiert.

Der dritte Vorzug unseres Ansatzes erschließt sich vor dem Hintergrund, dass Gewerkschaften selbst keine detaillierten Informationen über die Gruppenzusammensetzung ihrer Mitglieder veröffentlichen. Seit 1991 stehen nur Zahlen für Gesamtdeutschland zur Verfügung, und es wird nicht zwischen beschäftigten Mitgliedern und arbeitslosen, pensionierten oder studentischen Mitgliedern unterschieden. Eine entsprechende Unterscheidung ist jedoch aus ökonomischer Sicht wichtig, da die Finanzkraft einer Gewerkschaft und ihre Fähigkeit, Arbeitskräfte in den Firmen zu mobilisieren, maßgeblich vom Anteil der beschäftigten Mitglieder abhängt. Insofern ist der Nettoorganisationsgrad ein besserer Indikator für die Macht einer Gewerkschaft als der Bruttoorganisationsgrad. Gleichzeitig differiert die Macht von Gewerkschaften beachtlich zwischen verschiedenen Arbeitsmarktsegmenten. Beispielsweise sind Gewerkschaften im produzierenden Gewerbe traditionell stark, während sie in privaten Dienstleistungssektoren schwerer Fuß fassen. Die veröffentlichten Mitgliederzahlen der Gewerkschaften lassen sich jedoch nicht auf spezifische Arbeitsmarktsegmente herunterbrechen. Stattdessen sind mikroökonometrische Mitgliedschaftsschätzungen zur Bestimmung gewerkschaftlicher Nettoorganisationsgrade in homogenen Segmenten – wie etwa Regionen und/oder Sektoren – heranzuziehen. Mit Hilfe der resultierenden Machtindikatoren lässt sich dann der Gewerkschaftseinfluss auf die wirtschaftliche Entwicklung, die Beschäftigung oder die Lohnstruktur abschätzen.

Gewerkschaftsmacht und Lohnstruktur

Das letzte Kapitel[1] der Dissertation greift diesen Gedanken auf und verwendet entsprechend geschätzte Organisationsgrade zur Untersuchung des Gewerkschaftseinflusses auf die Lohnstruktur in Westdeutschland. Die Analyse erfolgt auf der Ebene aggregierter Arbeitsmarktsegmente, deren Definition sich an der Struktur des deutschen Tarifverhandlungssystems orientiert. Dies ist insofern von zentraler Bedeutung, als es in Deutschland im Gegensatz zu angelsächsischen Ländern nicht sinnvoll ist, Lohneffekte individueller Gewerkschaftsmitgliedschaft zu schätzen, da die negative Koalitionsfreiheit tarifvertragliche Mitgliedschaftsprämien verbietet und sich die Anwendung von Tarifverträgen auf den Großteil aller Beschäftigten erstreckt.

Wie wirkt sich vor diesem Hintergrund die Gewerkschaftsmacht auf das Lohnniveau und den Grad der Lohndispersion auf dem deutschen Arbeitsmarkt aus? Verfolgen Gewerkschaften das

[1] Hierbei handelt es sich um eine Übersetzung des deutschen Aufsatzes Fitzenberger und Kohn (2005).

Ziel "Gleicher Lohn für gleiche Arbeit", reduziert der Gewerkschaftseinfluss mithin die residuale Lohnungleichheit zwischen Arbeitnehmern mit gleichen beobachtbaren Eigenschaften? Oder geht der Einfluss gar darüber hinaus, so dass es auch zwischen Arbeitnehmern mit unterschiedlichen Eigenschaften zu verringerter Lohndispersion kommt? Zeigt sich der Zusammenhang asymmetrisch über die Lohnverteilung, etwa einem Mindestlohnargument entsprechend? Werden für eine Ungleichheitsreduktion Zugeständnisse in Bezug auf die Lohnhöhe gemacht? Und hat sich der Einfluss im Laufe der Zeit verändert?

Zur Beantwortung dieser Fragen imputieren wir die Gewerkschaftsinformation aus dem SOEP für Individuen in der IABS 1975–1997, deren umfangreichere Stichprobengröße verlässliche Untersuchungen der Lohnverteilung erlaubt, projizieren Nettoorganisationsgrade (NOG) für entlang der Dimensionen Wirtschaftszweig, Qualifikation, Alter und Zeit definierte Arbeitsmarktsegmente und analysieren den Zusammenhang zwischen NOG auf der einen und Lohnstrukturmaßen auf der anderen Seite.

Unsere Ergebnisse zeigen, dass ein höherer Nettoorganisationsgrad ceteris paribus mit einer reduzierten residualen Lohndispersion und einem geringeren Lohnniveau einhergeht. Dieses auf den ersten Blick überraschende Resultat entspricht einem von Agell und Lommerud (1992) sowie Burda (1995) theoretisch diskutierten Versicherungsmotiv für die Aktivität von Gewerkschaften, welche ein starkes Gleichheitsziel verfolgen und, um negative Beschäftigungseffekte zu vermeiden, dabei Einbußen im Hinblick auf das Lohnniveau hinnehmen.

Überdies fallen auch qualifikatorische Lohndifferenziale in Segmenten mit starken Gewerkschaften geringer aus, so dass der Einfluss in der Tat über "Gleichen Lohn für gleiche Arbeit" hinausgeht. In Übereinstimmung mit einem Mindestlohncharakter gewerkschaftlich ausgehandelter Löhne erfolgt die Stauchung der Lohnverteilung ferner asymmetrisch insbesondere in der unteren Hälfte – ceteris paribus erhöhen starke Gewerkschaften die Schiefe der Verteilung. Mit Blick auf die Altersdimension finden wir auch über den Lebenszyklus einen Lohnglättungseffekt. Es liegt damit keine Evidenz dafür vor, dass Gewerkschaften ex post rationale Mitgliederaustritte mittels der Durchsetzung von die Ungleichheit erhöhenden Senioritätszuschlägen verhindern oder Mitglieder älterer Kohorten, die den Gewerkschaften näher stehen als jüngere, besonders bevorzugen. Schließlich variieren die Gewerkschaftseffekte über die Zeit. Klare, signifikante Trends vor dem Hintergrund des Rückgangs in der Gewerkschaftsmitgliedschaft sind jedoch nicht zu erkennen.

Bezüglich einer kausalen Interpretation der vorliegenden Ergebnisse ist Vorsicht geboten, da der Analyserahmen einer möglichen Endogenität des Organisationsgrades nicht Rechnung trägt. Weder stellt die theoretische Literatur bisher Modelle für eine strukturelle Modellierung bereit, noch stehen valide Instrumente im Datensatz zur Verfügung. Im Rahmen zukünftiger Forschung lässt die Verwendung aussagekräftigerer verknüpfter Arbeitgeber-Arbeitnehmer-Daten auf Abhilfe hoffen. Arbeitgeber-Arbeitnehmer-Daten, wie etwa die Verdienststrukturerhebungen der

Statistischen Ämter oder die verknüpften Daten des IAB-Betriebspanels und der Beschäftigtenstatistik (LIAB), ermöglichen zudem die gemeinsame Berücksichtigung von Tarifbindung und gewerkschaftlichem Organisationsgrad (Fitzenberger, Kohn und Lembcke, 2006). Schließlich sollte eine simultane Untersuchung des Gewerkschaftseinflusses auf die Lohnstruktur und auf die Beschäftigung erfolgen.

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Ehrenwörtliche Erklärung

Ich habe die vorgelegte Dissertation eigenständig verfasst und dabei nur die von mir angegebenen Quellen und Hilfsmittel benutzt. Alle Textstellen, die wörtlich oder sinngemäß aus veröffentlichten oder nicht veröffentlichten Schriften entnommen sind, sowie Angaben, die auf mündlichen Auskünften beruhen, sind als solche kenntlich gemacht.

Mannheim, 21. November 2006

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