Boockmann, Bernhard; Steiner, Viktor

Working Paper

Cohort effects and the returns to education in West Germany

ZEW Discussion Papers, No. 00-05

Provided in cooperation with:
Zentrum für Europäische Wirtschaftsforschung (ZEW)

Cohort effects and the returns to education in West Germany

Bernhard Boockmann and Viktor Steiner
Centre for European Economic Research (ZEW)
P.O. Box 10 34 43
D-68034 Mannheim

Abstract: Using a Mincer-type wage function, we estimate cohort effects in the returns to education for West German workers born between 1925 and 1974. The main problem to be tackled in the specification is to separately identify cohort, experience, and possibly also age effects in the returns. For women, we find a large and robust decline in schooling premia: in the private sector, the returns to a further year of post-compulsory education fell from twelve per cent for the 1945-49 cohort to about seven per cent for those born in the early 1970s. Cohort effects in men's returns to education are less obvious, but we do find evidence that they, too, have declined. We conclude by identifying possible reasons for the decline.

JEL-classification: J 31

Key Words: returns to education, cohort effects, population ageing.

Acknowledgement: We should like to thank Charlotte Lauer for useful suggestions, as well as for supplying us with the data.

1 Introduction

In contrast to the experience in the United States, estimated rates of return for human capital have shown to be relatively constant over time in West Germany (Fitzenberger and Franz, 1998; Lauer and Steiner, 2000; Steiner and Wagner, 1998a). However, the relative stability of educational premia over time could disguise a much larger change over cohorts. While the first studies which quantified cohort effects in educational premia in the U.S. appeared some twenty years ago, to date there has been only little research on cohort effects in the returns to education in Germany, [1] although there are strong a priori reasons why they might be present. First, the number of births changed dramatically over the post-war period, with a peak of the baby-boom around 1965 and a steep subsequent decline in fertility. Second, enrolment in higher education rose to levels previously unknown during the decades after 1945. Third, female labour force participation increased widely. For example, among women in their thirties it almost doubled between 1970 and 1995 to about 75 per cent. A factor that could limit the extent of cohort effects on wages, however, could be the relatively centralised German system of wage determination. If wages are not allowed to vary over age groups, cohort effects should mainly show up in unemployment figures.

In this paper, we concentrate on the question whether cohort effects in the returns to education can be observed. Our answer is that there are, indeed, significant changes over cohorts, but they are much more marked for women than for men. In the private sector, women born in the 1960s suffer a decline in the returns of three percentage points compared to women born in the 1940s. This development is part of a secular decline affecting all cohorts which we observe, i.e. individuals born between 1925 and 1975. For men, we observe a weak but significant decline between individuals born in the early 1950s and the mid-1960s.

In the following section, we give a brief account of the literature. We distinguish several reasons for cohort effects. We also discuss the empirical evidence available thus far. The third section introduces the dataset and discusses problems of estimation. The results are discussed in section four. Section five concludes.

---

[1] Some indirect evidence is contained in Fitzenberger et al. (1995). Lauer and Steiner (2000) present results similar to those discussed in this paper.

2 Cohort effects in educational premia: theory and empirical evidence

Put very broadly, there will be cohort effects if older and younger workers are imperfect substitutes in production. In that case, the relative scarcity within each birth cohort will result in different wage levels. If the relative scarcity varies not only between birth cohorts but also across educational groups within each cohort, there will be differences not only in wage levels but also in the wage premia paid for post-compulsory education.

Why should the relative scarcity of workers with different education change over cohorts? A first reason may be that there are exogenous shifts in educational attainment between cohorts. If more workers receive higher education, the relative scarcity of college- or university-educated workers falls, and so should the wage premium paid to them.

Apart from the proportions of workers choosing certain educational levels, the overall number of workers may matter, too. Thus an exogenous decline in the number of workers – due to changes in fertility, wars or epidemics – may lead to a change in the relative rewards of different skill groups. There are two reasons for this. First, the own-price elasticities of demand may differ across skill groups. One would assume that they are lower for high-skilled workers because these workers can less easily be substituted against capital; indeed, empirical findings for Germany support this hypothesis (Falk and Koebel, 1998; Steiner and Wagner, 1998b). If a smaller cohort enters the labour market (and substitution between educational groups is limited), wages will be driven up in the market for qualified workers more than in other labour market segments. Second, the elasticity of substitution between younger and older workers may differ across qualification levels. If it is higher at low education levels, the effect of a demographic change will be subdued for these workers and more pronounced among highly skilled workers. Indeed, substitution elasticities typically show this pattern in empirical studies (Stapleton and Young, 1988, section II).

How large these effects are depends on the precise form of the production function. Consider the effect of technological change. Following the adoption of new technology, firms may depend on the services of young workers who have received their education (e.g., computer literacy skills) relatively recently. Thus labour demand for young, well-educated workers becomes very inelastic with respect to their relative wages, while demand elasticities for young workers with little education are unaffected (or even decline if the technological change also entails easier substitution of unskilled labour by capital). The faster technological change is, the lower is the substitution elasticity among younger and older skilled workers and the larger are changes in the returns to education over cohorts.

Still another source of cohort effects may be downward rigidity of incumbents‘ wages, caused, for instance, by „social norms“. If young workers with high levels of education enter the labour market in large numbers, one would expect wages of older highly skilled workers to decline. However, older workers will have become accustomed to a certain wage level, and a reduction below the previous level may, by an efficiency wages argument, lead to less effort by these workers. In that case, employers may be unwilling to adjust incumbents‘ wages, thus concentrating the effect of the supply change on new entrants.[2]

In all our arguments so far, we have assumed that the number of workers in each age-education category is determined by exogenous factors (such as demography). In reality, individuals make choices on education and labour force participation, a problem which will be discussed below.

There are a number of empirical studies on cohort effects in educational premia, starting with Welch's (1979) investigation into the wage implications of the US baby boom of the 1950s. Welch concentrated on cohort size as a determinant of wages. His estimations of separate wage functions by educational levels imply that wages of college graduates suffered more than others from the expansion of supply caused by the baby boom's entry into the labour market. Freeman (1979) decomposed weekly earnings into age and education components. He found that between 1969 and 1976, the difference between wages of younger and older workers grew significantly (implying a steeper wage-age profile), and that this increase was particularly significant for higher education groups. Freeman attributes this change to the baby boom. From this finding, however, an unambiguous conclusion to the returns to education cannot be drawn, since wages might have gone up for college graduates at all ages during this period.

[2] Of course, the pay differential between old and young workers would imply that older workers are paid above their productivity, while younger workers are paid below. Employers would then try to sack older and hire younger workers, creating unemployment among the first. This may explain why, in the empirical results reported by Zimmermann (1991), an increase in the size of young cohorts has a stronger impact on unemployment among older than among younger workers. At the same time, however, older workers are usually better protected against redundancy than younger workers.

Recent analyses of education, age and cohort effects on wages have mostly used Mincer's (1974) approach. This procedure yields coefficients for on-the-job experience and the number of years spent in education, with the latter having (under certain assumptions) a structural interpretation as the implied rates of return for an additional year of schooling. Several authors have included cohort variables into the Mincer equation. For instance, Berger (1985) and recently Macunovich (1999) have entered cohort size as an additional explanatory variable. The latter study also tries to separately identify demand and supply effects of demographic change.

In contrast to studies which use cohort size as an independent variable, our focus is not exclusively on demographics because this is certainly not the only cohort-specific impact on wages. Other characteristics pertaining to particular cohorts, such as skills in particular technologies, may also be of importance. Using only size as a cohort-specific explanatory variable may lead to biased estimates if the variables left out from the estimation are correlated with cohort size. Therefore, we estimate Mincer equations with cohorts (consisting of adjacent birth years) entering as dummy variables, both directly on wages as well as in interaction with education.

This approach has been taken in a number of other studies. For German data, Fitzenberger et al. (1995) investigate whether the same age-earnings profiles can be observed across cohorts once macroeconomic time effects are accounted for. In one of their estimations, they look at wages as a function of age, distinguishing between (a) two different birth groups (being five years apart from each other), and (b) four educational categories. Their finding is that there has been, between 1978 and 1983, a reduction in the entry-level wages for workers with completed apprenticeship and/or Abitur (A-levels), while there is no such effect for workers without these qualifications. This points to a decline in educational premia. They do not find a similar effect for university-educated workers, but this may be due to their data source: since they use data from social security files, in which income is recorded only up to an upper threshold, their results for highly educated workers suffer from a severe problem of right-censoring.

Apart from the identification of cohort effects, there is an abundance of empirical studies using the Mincer equation approach which are geared towards other problems. A large strand of the literature concentrates on the potential endogeneity of schooling, as well as on measurement error in this variable. Little consensus seems to have emerged regarding the best way to proceed. For example, family background variables (in particular, parents‘ education) are often used to instrument schooling.[3] However, this may even further bias upwards the estimated returns (Card, 1999).

From an institutional point of view, the distinction between returns to education in private and in public employment is also very interesting. In many countries, wage structures in the public sector appear to be more compressed, thus diminishing the returns which can be achieved. This may, of course, lead to systematic selection of individuals into the two sectors. Dustmann and van Soest (1998) provide a systematic empirical treatment of this kind of endogeneity, as well as endogeneity in schooling, hours worked and years of work experience.

In our paper, we do differentiate between sectors but neglect the endogeneity problem. The results should thus be seen as a first empirical assessment of cohort effects, while leaving the endogeneity issue to further research. There seems to be no a priori expectation, however, in which direction our estimates for the cohort effects should be biased in the presence of endogeneity problems.

3 Methodology and Data

Our basic specification is

$$ln\, w_{it} = \alpha_0 + \alpha_1 X_{it} + \alpha_2 X_{it}^2 + \alpha_3 S_i + \sum_{k=2}^{K} \beta_k D_{ik} + \sum_{k=2}^{K} \gamma_k D_{ik} S_i + \ldots + \varepsilon_{it}, \qquad (1)$$

$i = 1, \ldots, n$

$t = 1984, \ldots, 1997$

where w stands for wages, X is work experience in years, S is years spent in education, and D is a dummy variable which takes the value one if the individual observed is a member of birth cohort k (defined on a number of years) and zero otherwise. As the subscripts indicate, equation (1) is estimated on an unbalanced panel of individuals i observed over a number of years t.

Our data base is the German Socio-Economic Panel (GSOEP). We use all 14 waves (from 1984 to 1997) currently available but confine ourselves to men and

---

3 Using the same data source as in this paper, Lauer and Steiner (2000) instrument years of schooling by variables such as father's education, parents' employment status, etc. This procedure has only minor effects on the estimated returns to schooling.

women living in West Germany. In order to have cohorts with a sufficient number of observations, our estimations are restricted to individuals born between 1925 and 1974. Several groups of people are excluded from the dataset: this concerns students, military personnel, pensioners, and civil servants („Beamte“). In some of the estimations, we exclude all public employees from the sample. We estimate separate wage equations for men and women. Only employees with German nationality are represented in the sample used here.

The dependent variable is net nominal earnings per hour worked. The hours measure includes paid overtime. Among the independent variables, we use two different concepts to measure education. The first is the number of years spent in education. The other is the highest degree reached in education. We distinguish seven broad education/qualification categories: (1) secondary schooling without apprenticeship, (2) secondary schooling and completed apprenticeship, (3) master craftsman, (4) Abitur (A-levels), (5) Abitur and completed apprenticeship, (6) polytechnic degree, (7) university degree.

Experience is a crucial variable in our estimations. Rather than using potential experience, which is typically defined as age minus years of schooling minus six years, we construct a variable for actual labour market experience from retrospective data contained in the GSOEP. Individuals are asked about past spells in full-time or part-time employment since the age of 15. The durations of these spells can then be added to obtain measures for total length of experience. The decisive assumption that we make is that spells out of employment (e.g., due to child raising or unemployment) do not contribute towards the accumulation of human capital rewarded by the labour market, and hence to higher wages. In the German case, this assumption is more problematic for the public than for the private sector because public employees‘ salaries rise automatically with age.

We also differentiate between full-time and part-time experience because these may affect productivity and wages differently. However, since the number of men in part-time employment is very small, we use variables measuring part-time experience only in the estimations for women.

The experience variable is important because the identification of year, cohort and experience effects hinges on it. Suppose we measured experience as potential experience, e.g. the number of years beyond age 15 minus the years of post-compulsory education. If, in addition, we included dummy variables for each birth year and the current year in our estimations, there would be a linear dependency because birth year plus 15 years plus years of post-compulsory schooling plus years of potential experience always equals the current year.[4] By contrast, our experience coefficients are identified (a) because actual experience differs from potential experience, and (b) because our cohort measures are dummy variables for a number of birth years (five or ten), such that there is variation in the length of experience *within* each cohort thus defined.

The Mincer equation implicitly assumes that the returns to education are constant throughout the working life. One may object to this assumption because it takes time to realise the full impact of education on productivity (and thus on wages). Put loosely, education and experience may be complements. Moreover, in the presence of seniority wages a young worker is not compensated according to the full productivity effect of his or her education; instead, wages are held back until later ages. Hence, the returns to education may increase with experience. On the other hand, the effect of initial education on productivity will typically decline at long levels of experience because knowledge depreciates. These effects may bias our estimations of cohort effects because the younger cohorts are observed at earlier ages when they do not reap the full benefits from their educational attainment. Similarly, an observed increase in wages for men over the pre-war cohorts could be due to the greater impact of schooling on productivity for the middle-aged than for older workers.

To correct for this possible distortion, we estimate an alternative specification to equation (1) which includes an interaction between (full time-) experience and education:

$$\begin{aligned} \ln w_{it} &= \alpha_0 + \alpha_1 X_{it} + \alpha_2 X_{it}^2 + \alpha_3 S_i \\ &\qquad + \alpha_4 X_{it} S_i + \alpha_5 X_{it}^2 S_i + \sum_{k=2}^{K} \beta_k D_{ik} + \sum_{k=2}^{K} \gamma_k D_{ik} S_i + \end{aligned} \qquad (1')$$

The interaction is allowed to be non-monotonic to take into account both seniority wages and knowledge depreciation. While educational premia are now allowed to vary parametrically both over cohorts and over the life cycle, we are implicitly assuming that the shape of the interaction stays constant over cohorts. A drawback of using interactions is that the schooling coefficient does not any longer

---

4 For a discussion of the identification problem, see Heckman and Robb (1985) who also address identification of interaction and higher order terms.

have a structural interpretation as the returns to education, as in the original Mincer equation (1).[5]

Apart from schooling and experience, there are two other variables that control for human capital accumulation on the job. The first is tenure, the number of years an individual stayed with his or her current employer. The other is the number of months spent in full-time or part-time employment during the current year to control for the continuity of employment during the observation period. We also control for a number of other factors, like the region (German Laender, i.e. federal states), regional unemployment rates, industry, and firm size. Year dummies are also included to control for macroeconomic effects on individual wages.

4 Results

Estimation results for specification (1) are displayed in the left columns of tables 1(a) and 1(b). The coefficients for the cohort dummies are given in the upper part of the table, followed by the number of years spent in education and cohort-education interactions (Schooling 30-34, etc.). The birth years 1925 to 1929 form the base category for the cohort variables. Employment status is the number of months spent in full-time or part-time employment during the current year, as just defined.

Since there are interactions, the cohort effects on wages and the returns to education cannot be inferred from a single coefficient alone but must be calculated from the estimated parameters.[6] Figure 1 presents the returns to education by birth cohorts for both men and women. We observe a decline over cohorts in women's returns to education which is monotonic apart from a spike for the 1940-44 cohort.

---

5 Another decomposition of the education coefficient in a Mincer equation could be made by using quantile regression techniques as in Hartog et a. (1999).

6 The same is true for the standard errors. In the graphs, standard errors are calculated using the delta method. Suppose the total effect of education on wages is $h(\alpha_1, \alpha_2) = \alpha_1 + \alpha_2 X$, where the α are parameters and X is some independent variable. The variance of the total effect is $Var(\hat{\alpha}) = AV'A$, with A a matrix of partial derivatives $[\partial h / \partial \alpha_1 \quad \partial h / \partial \alpha_2]$ and V the covariance matrix of the estimated parameters α_1 and α_2.

Table 1: Returns to years spent in education

a) Men

	(1)		(1‘)	
Variable	**Coeff.**	**t–stat.**	**Coeff.**	**t–stat.**
Cohort 1930-34	–0.078	–1.031	–0.064	–0.837
Cohort 1935-39	–0.119	–1.845	–0.112	–1.632
Cohort 1940-44	–0.137	–2.084	–0.166	–2.245
Cohort 1945-49	–0.123	–1.813	–0.210	–2.644
Cohort 1950-54	–0.088	–1.332	–0.240	–3.012
Cohort 1955-59	–0.030	–0.466	–0.234	–2.897
Cohort 1960-64	0.060	0.890	–0.212	–2.540
Cohort 1965-69	0.230	3.296	–0.073	–0.844
Cohort 1970-74	0.317	3.201	–0.026	–0.230
Schooling	0.078	15.138	0.044	6.155
Schooling x Experience			0.002	6.777
Schooling x Experience²			–0.003	–3.734
Schooling x 30-34	0.006	0.950	0.005	0.752
Schooling x 35-39	0.007	1.157	0.006	1.018
Schooling x 40-44	0.009	1.571	0.012	1.858
Schooling x 45-49	0.005	0.880	0.013	1.936
Schooling x 50-54	–0.001	–0.179	0.013	1.896
Schooling x 55-59	–0.009	–1.578	0.010	1.379
Schooling x 60-64	–0.016	–2.941	0.007	1.009
Schooling x 65-69	–0.034	–5.798	–0.008	–1.036
Schooling x 70-74	–0.047	–5.498	–0.018	–1.884
Tenure	0.002	3.292	0.002	3.002
Tenure²/100	0.005	2.642	0.005	2.832
Employment status	0.019	19.018	0.018	18.314
Experience	0.022	21.178	–0.005	–1.284
Experience²/100	–0.050	–22.002	–0.009	–0.899
Unemployment rate	–0.007	–1.278	–0.007	–1.328
Unemployment rate²/100	0.070	1.112	0.066	1.049
Year Dummies	YES		YES	
Industry Dummies	YES		YES	
Laender Dummies	YES		YES	
Firm size dummies	YES		YES	
Number of observations	19,004		19,004	
R^2	0.500		0.502	
adjusted R^2	0.499		0.501	

F–test of specification 1‘ against 1: $F(2, 18937) = 37.32$

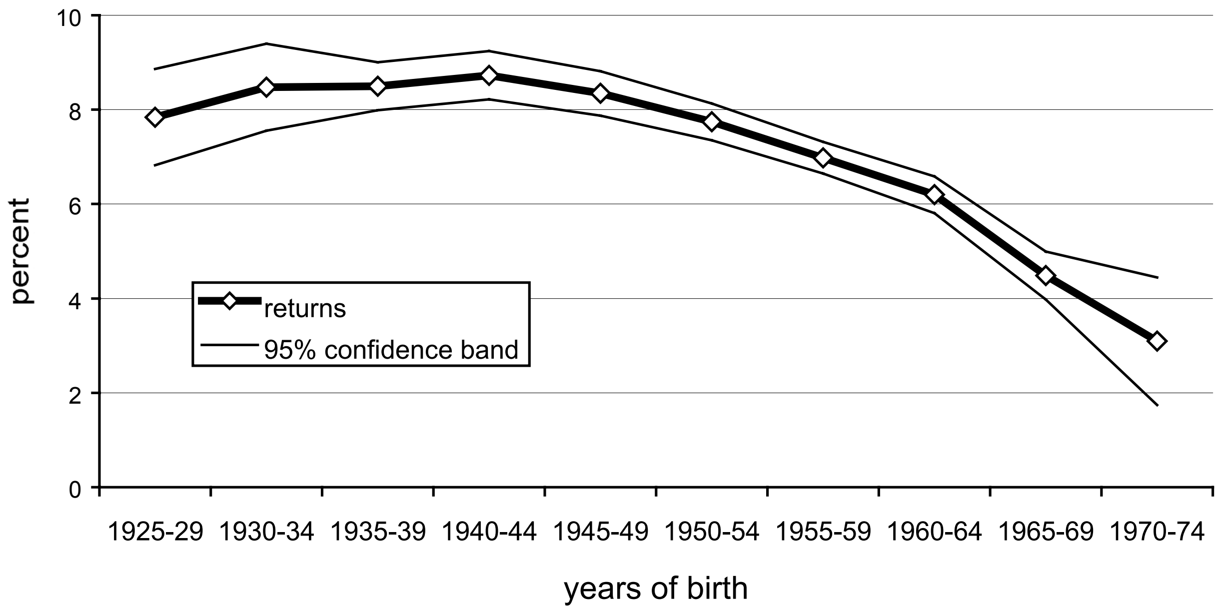
b) Women

	(1)		(1‘)	
Variable	**Coeff.**	**t–stat.**	**Coeff.**	**t–stat.**
Cohort 1930-34	0.139	1.194	0.095	0.791
Cohort 1935-39	0.368	3.319	0.437	3.919
Cohort 1940-44	0.368	3.300	0.387	3.408
Cohort 1945-49	0.478	4.383	0.566	5.035
Cohort 1950-54	0.560	5.407	0.604	5.572
Cohort 1955-59	0.613	6.057	0.635	5.960
Cohort 1960-64	0.645	6.326	0.631	5.839
Cohort 1965-69	0.862	8.238	0.820	7.387
Cohort 1970-74	0.941	7.582	0.881	6.789
Schooling	0.115	13.020	0.103	10.504
Schooling x Experience			0.003	6.681
Schooling x Experience²			–0.009	–6.101
Schooling x 30-34	–0.014	–1.294	–0.009	–0.797
Schooling x 35-39	–0.036	–3.525	–0.043	–4.128
Schooling x 40-44	–0.029	–2.861	–0.031	–2.944
Schooling x 45-49	–0.037	–3.747	–0.046	–4.410
Schooling x 50-54	–0.044	–4.606	–0.048	–4.789
Schooling x 55-59	–0.050	–5.399	–0.052	–5.288
Schooling x 60-64	–0.053	–5.782	–0.052	–5.289
Schooling x 65-69	–0.073	–7.692	–0.069	–6.817
Schooling x 70-74	–0.086	–7.687	–0.080	–6.823
Tenure	0.007	7.169	0.007	6.931
Tenure²/100	–0.014	–4.034	–0.012	–3.695
Employment status: FT	0.019	14.696	0.019	14.404
Full–time Experience	0.017	15.659	–0.017	–3.269
FT Experience²/100	–0.027	–9.381	0.071	4.415
Part–time dummy	0.055	6.043	0.054	5.893
Employment status: PT	0.010	7.176	0.010	7.034
Part–time Experience	0.003	1.806	0.003	2.092
PT Experience²/100	0.013	2.233	0.011	1.781
Unemployment rate	0.014	1.989	0.014	2.067
Unemployment rate²/100	–0.159	–2.061	–0.162	–2.104
Married	–0.010	–1.485	–0.008	–1.244
Year dummies	YES		YES	
Industry dummies	YES		YES	
Laender dummies	YES		YES	
Firm size dummies	YES		YES	
Number of observations	13,492		13,492	
R^2	0.428		0.430	
adjusted R^2	0.425		0.427	

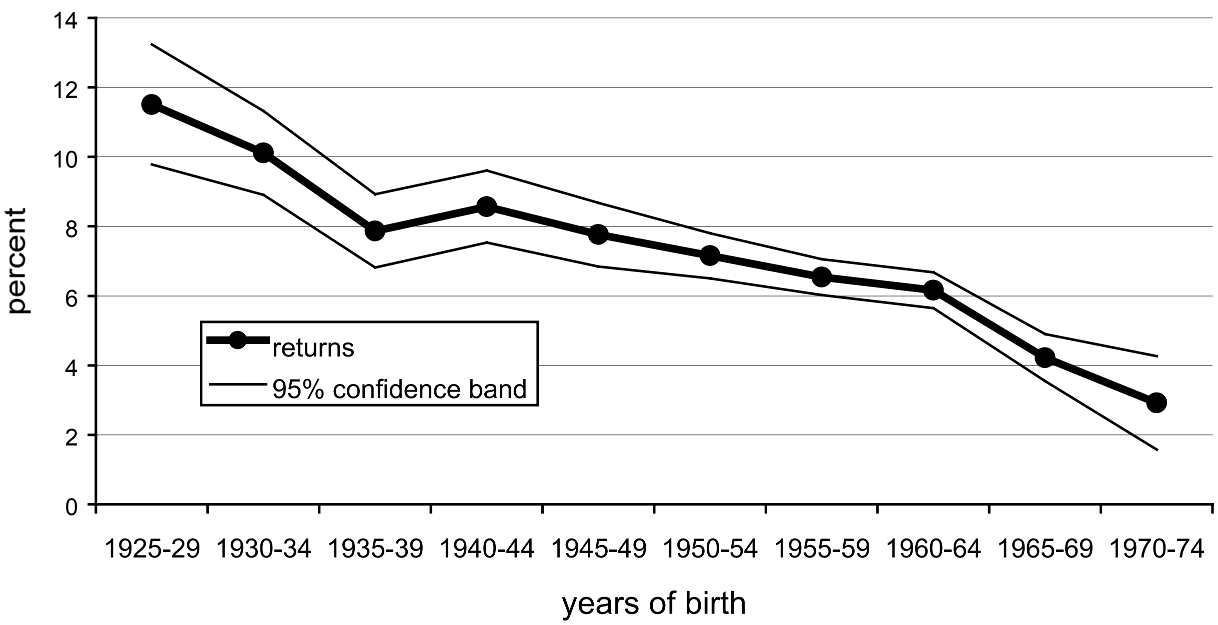
Note: F–test of specification 1‘ against 1: F(2, 18937) = 37.32

Figure 1: Cohort-specific returns to education

a) men



b) women



By contrast, the returns to education for men increase slightly up to the 1940-44 cohort and decline afterwards. Among the cohorts born after the Second World War, there is no difference in the returns to education between men and women. Overall, the decrease in the returns for later cohorts appears to be quite dramatic.

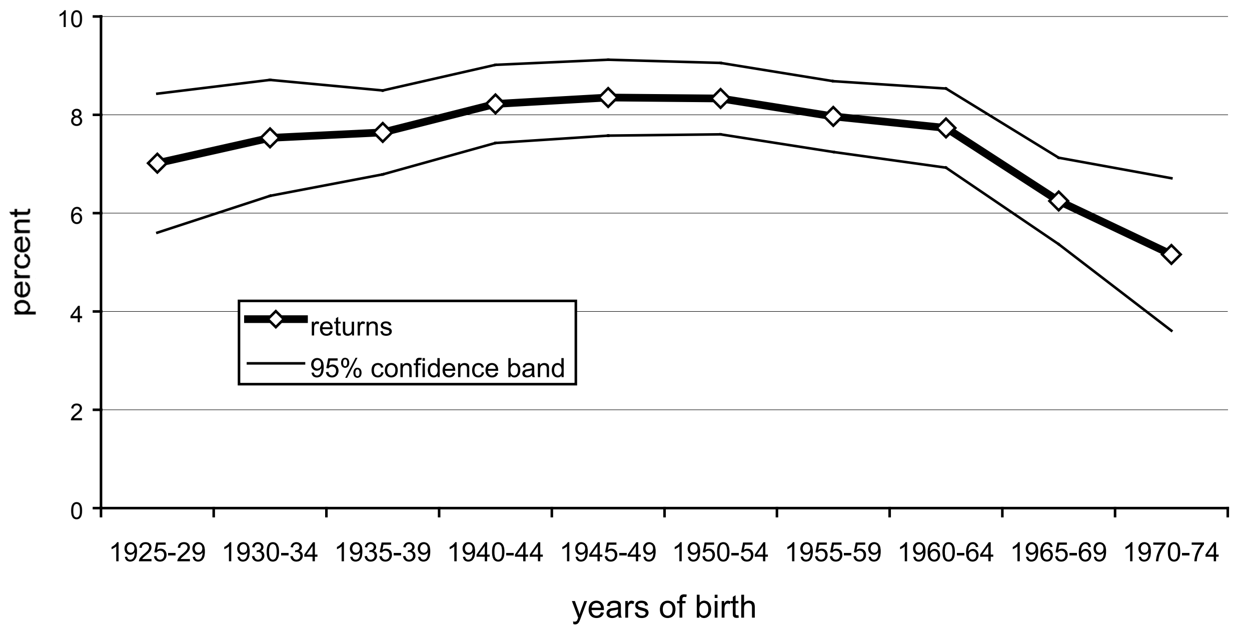
Our first robustness check concerns the size of the education premium over the life cycle. Results from the estimation of (1') are given in the second column of table 1 and in figure 2. In order to obtain a similar cohort average over the returns to education as in figure 1, we fix experience at 15 years of full-time employment. In interpreting the graph, it has to be kept in mind that this inevitably produces out-of sample predictions. For example, since the last year of the observation period is 1997, there are no individuals with 15 years of work experience in the 1970-74 cohort. Hence, the *absolute magnitudes* of the returns to education should not be interpreted for cohorts which have, on average, much more or much less labour market experience in the observation period. The purpose of the figure is to show the difference in the returns to education across cohorts, not their level.

We observe that allowing the educational premia to vary over the working life takes out some of the decline found in the results for specification (1). In particular, the difference in schooling premia between the youngest cohort (1970-74) and the cohort born 20 years earlier now amounts to roughly three per cent for men, as compared to the five per cent found earlier. A similar change can be observed for women. Overall, there is still a clear negative trend over birth years in the returns to education for women. Concerning men, however, the decline is confined to the two youngest birth cohorts. Using an F-test, we find that the interaction between schooling and experience is statistically significant, and we will therefore continue to work with specification (1') instead of (1).

The results from (1') can also be displayed with respect to the cohort effects in wage levels. Figure 3 shows cohort wage differentials (the differences in log wages) for different durations spent in education (nine years for a worker without apprenticeship; twelve years for a person who has finished apprenticeship; and 18 years for a university graduate). To a certain extent, this picture is just another way of showing the results seen from the previous graph. On average, wages for younger cohorts are lower than for older cohorts. For women, this decline is the larger the longer an individual stayed in education. Indeed, at a very low level of education, female individuals from younger cohorts actually earn more than women born earlier. For men, the three curves are not that clearly ordered. Up to the 1960-64 cohort, the negative cohort effects are greater for individuals with shorter pe-

Figure 2: Returns to education, accounting for schooling-experience interaction

a) men



b) women

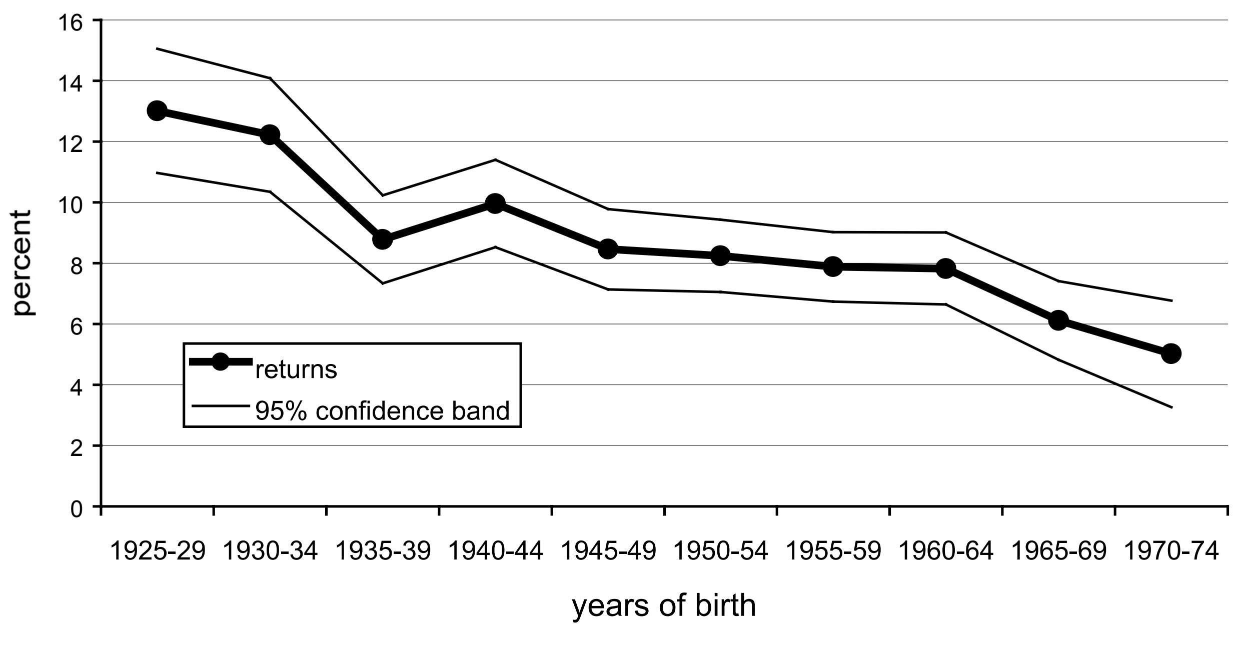
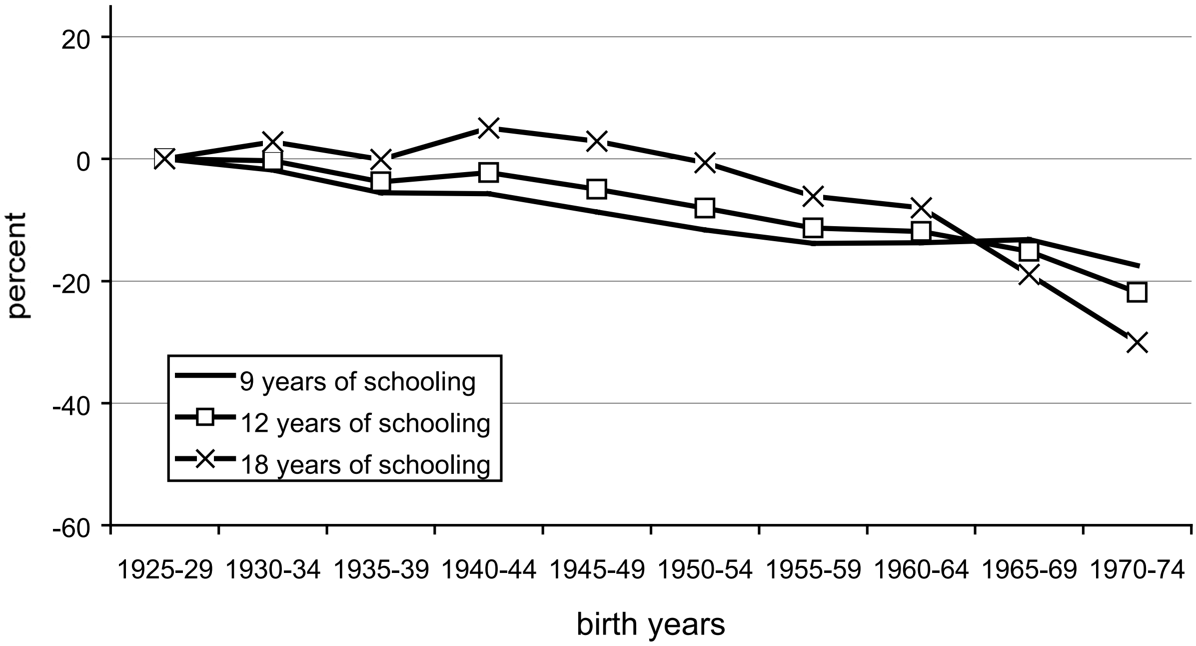
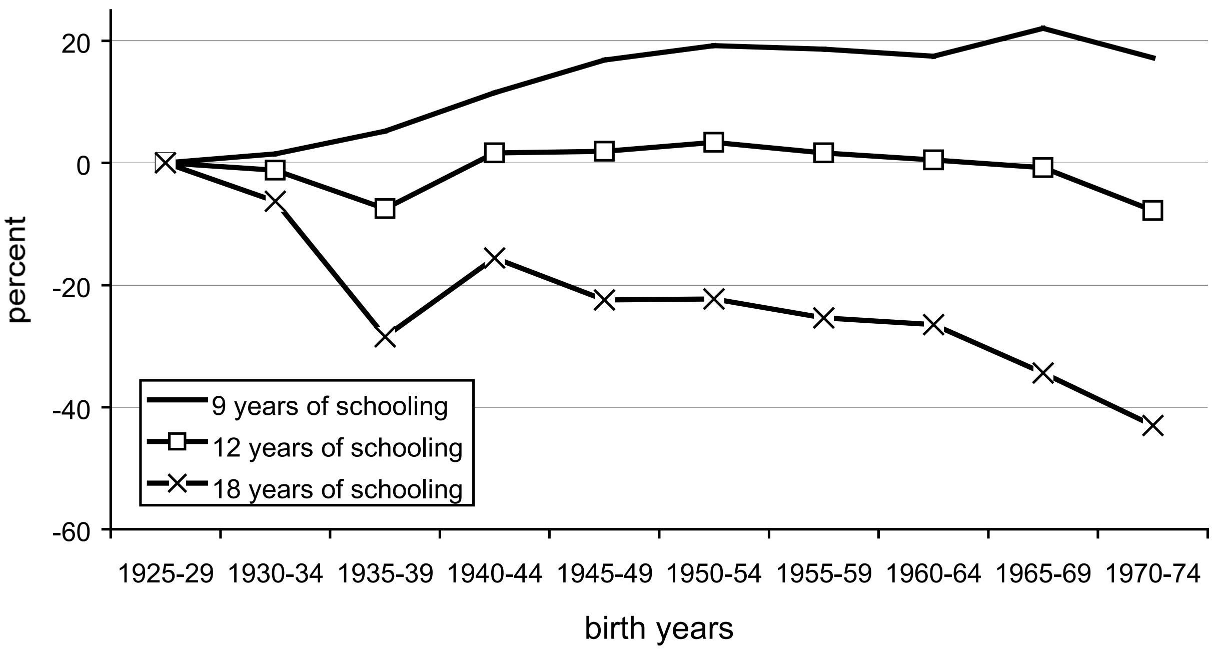


Figure 3: Cohort effects on wage levels

a) men



b) women



periods of education; only for the youngest two cohorts is the pattern similar to the cohort effects for women.

In specification (1‘), we have accounted for experience effects in the returns to education, but we clearly could not account for age effects because of the linear dependency between cohorts, calendar year and age. If, apart from the experience effect, there is also an effect of age on educational premia, this effect is entirely interpreted as a cohort effect in our results. This seems to be a problem mainly in the public sector. In Germany, the salaries of public employees are raised according to age every two years, independently of the work history of the individual. For the public sector, therefore, pure cohort effects are not properly identified. One would also believe that pure cohort effects are more pronounced in the private than in the public sector because wage determination in the public sector is more rigid. Since there are fixed pay scales, the only remaining instrument to adjust wages in the public sector is to grade individuals of different cohorts into different pay categories.

We account for these problems by estimating specification (1‘) on the basis of a subsample of private employees only. The results, displayed in table 2 and figure 4, show that there remains a significant decline in educational returns for women. However, the youngest cohort now still earns returns to education of about seven percent, compared to the five percent found earlier. This means that the decline in the returns ceases to be significant at the five per cent level for men, but it is still very marked for women.

Another way, apart from using interaction terms between years of schooling and labour market experience, of checking whether the estimated cohort effects in the returns to education are due to misspecification of the experience (or age) part of the equation is to estimate the returns to education for different cohorts *observed at the same age*. We thus compare wages of an early cohort observed at an early year to those of a subsequent cohort observed a corresponding number of years later. In order to obtain a larger set of observations, we slightly change the definition of birth cohorts to include individuals born within an interval of seven years. Since there are, by construction, no age effects in the differences between the two cohorts, we do not have to distinguish between public and private sector wages.

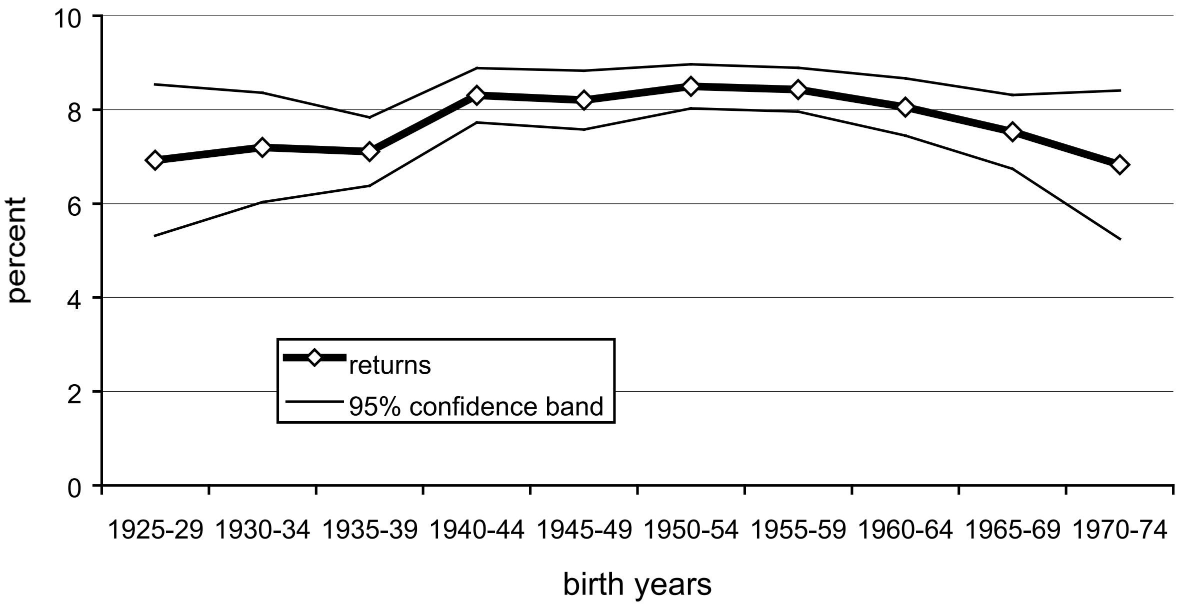
In a first step, we compare individuals born in the years 1962 to 1968 with the cohort born 13 years earlier, i.e. the 1949-55 cohort. The distance in time between the two cohorts is chosen because the first and the last obtainable waves of the GSOEP are just 13 years apart. The earlier cohort is observed in 1984, the second

Table 2: Estimation results for equation (1‘), private sector employees only

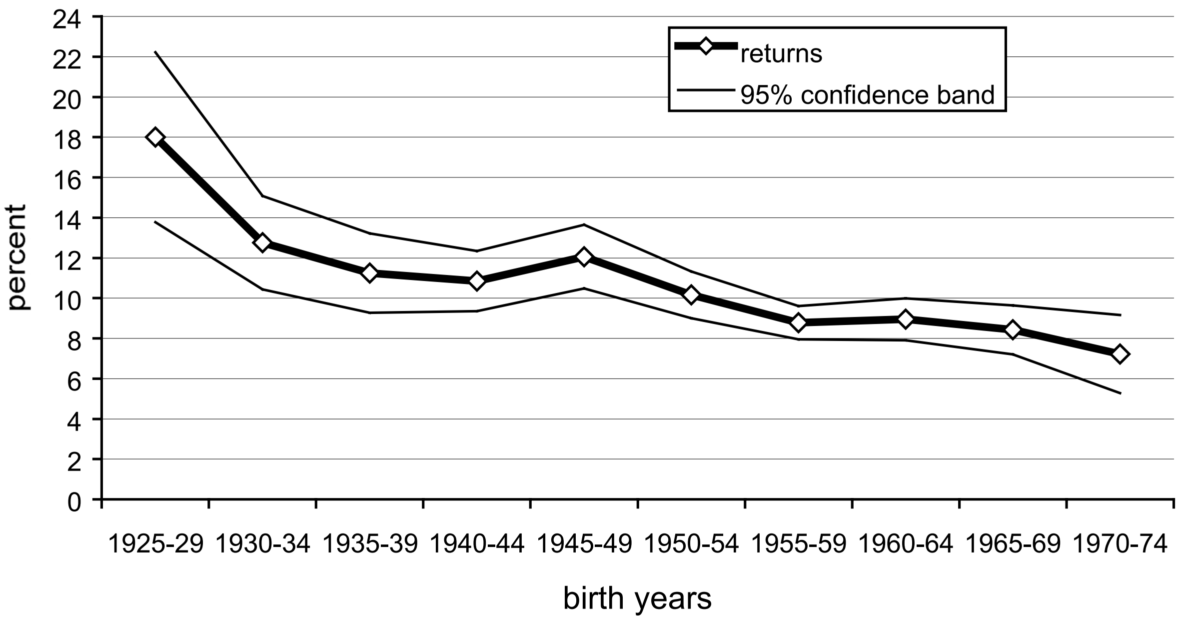
	Men		Women	
Variable	**Coeff.**	**t–stat.**	**Coeff.**	**t–stat.**
Cohort 1930-34	–0.034	–0.388	0.546	2.404
Cohort 1935-39	–0.058	–0.706	0.704	3.010
Cohort 1940-44	–0.187	–2.087	0.816	3.562
Cohort 1945-49	–0.213	–2.180	0.730	3.127
Cohort 1950-54	–0.271	–2.783	0.937	4.112
Cohort 1955-59	–0.300	–3.035	1.085	4.833
Cohort 1960-64	–0.264	–2.598	1.047	4.629
Cohort 1965-69	–0.223	–2.113	1.118	4.906
Cohort 1970-74	–0.207	–1.595	1.199	4.945
Schooling	0.031	3.582	0.129	5.852
Schooling x Experience	0.003	8.238	0.005	6.924
Schooling x Experience²	–0.004	–4.276	–0.011	–4.625
Schooling x 30-34	0.003	0.344	–0.052	–2.360
Schooling x 35-39	0.002	0.249	–0.068	–2.933
Schooling x 40-44	0.014	1.747	–0.072	–3.189
Schooling x 45-49	0.013	1.500	–0.059	–2.597
Schooling x 50-54	0.016	1.848	–0.078	–3.508
Schooling x 55-59	0.015	1.754	–0.092	–4.188
Schooling x 60-64	0.011	1.284	–0.090	–4.081
Schooling x 65-69	0.006	0.662	–0.096	–4.298
Schooling x 70-74	–0.001	–0.088	–0.108	–4.587
Tenure	0.002	2.580	0.008	6.197
Tenure²/100	0.004	2.116	–0.020	–4.346
Employment status:FT	0.019	16.768	0.019	11.883
Full–time Experience	–0.015	–3.373	–0.036	–4.614
FT Experience²/100	0.004	0.317	0.091	3.743
Part–time dummy			0.040	3.570
Employment status:PT			0.009	5.464
Part–time Experience			0.005	2.614
PT Experience²/100			0.004	0.477
Unemployment rate	–0.007	–1.149	0.020	2.447
Unemployment rate²/100	0.082	1.169	–0.251	–2.770
Married			–0.028	–3.433
Year dummies	YES		YES	
Industry dummies	YES		YES	
Laender dummies	YES		YES	
Firm size dummies	YES		YES	
Number of observations	15661		9344	
R^2	0.513		0.425	
adjusted R^2	0.510		0.421	

Figure 4: Cohort-specific returns to education, private sector only

a) men



b) women



cohort in 1997. This means that in both cohorts, the individuals included in the sample are between 29 and 35 years of age, i.e. they are individuals a couple of years into their careers. We make a similar comparison for cohorts born 20 years earlier, i.e. between individuals in the 1942-48 and the 1929-35 cohorts who are between 49 and 55 years of age in 1984 and 1997, respectively.

There does remain an untestable identifying assumption in these estimations. We attribute all of the difference in the returns to education between the two sub-samples to the cohort effects, while they may also be produced by year effects, since both sub-groups are observed at different calendar years. This identifying assumption does not strike us as overly restrictive, however, because macroeconomic conditions did not differ much between the two years.

Estimation results in table 3 confirm our findings for the 29 to 35 year olds: the cohort-schooling interaction is significant at the one percent level for men, although it is only significant at the ten percent level for women. For men, the estimated returns are 8.25 per cent for the 1949-55 cohort, and 5.88 per cent for the 1962-68 period; for women, the estimations yield rates of return of 6.68 and 4.29, respectively. This order of magnitude is in-between of the one estimated on the whole sample in specifications (1) and (1‘).

Concerning older workers we do not find significant changes in educational premia across cohorts. This was to be expected for men (cf. figure 2), but not for women. The insignificance of the decline for earlier periods probably stems from the fact that the number of women with long post-compulsory education spells is very low in the sample.

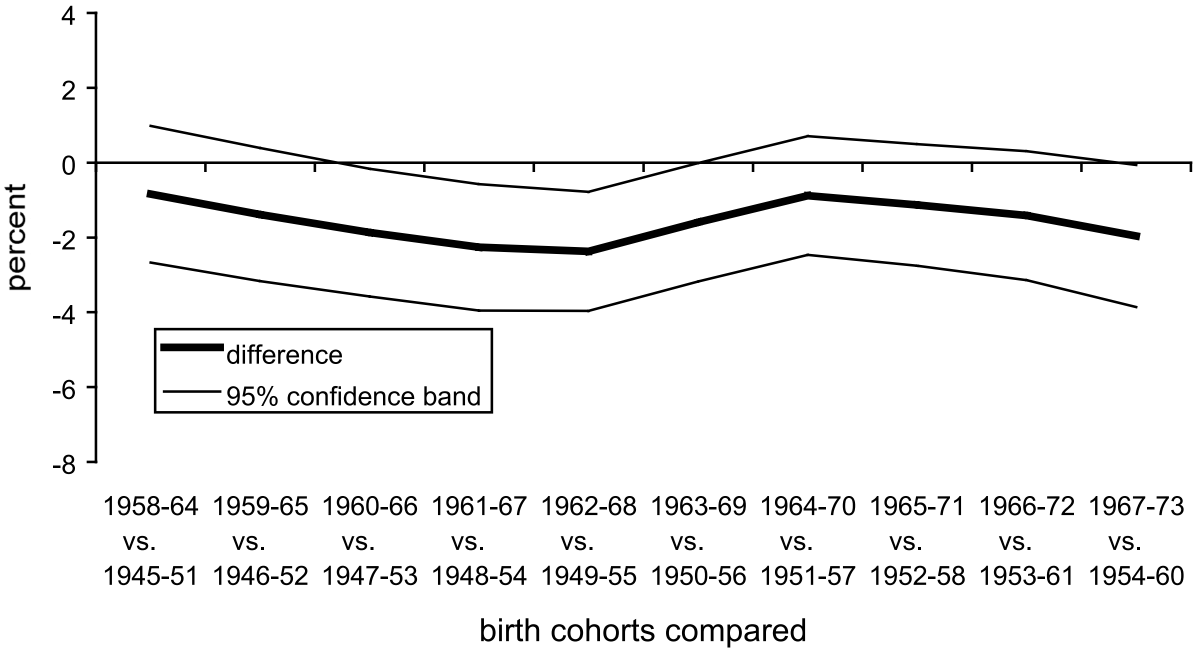
Comparing the returns of particular birth cohorts is open to the criticism that the start and end years of the cohorts are chosen arbitrarily. To give a more complete picture, we performed separate estimations for particular cohorts at a particular age for different cohort definitions. As can be seen from figure 4 (a), the decline in the returns for men is significant at the five per cent level for four cohorts, the oldest of which consists of individuals born 1960 to 1966 and the youngest of which contains individuals born from 1963 to 1969. When we turn to female workers, the picture is different in that returns do not recover for cohorts younger than the 1963-69 birth group. However, for only two cohorts are the estimated returns for women outside the 95 percent confidence interval for men. While at first glance the cohort effects appear to be smaller than in other graphs, it has to be stressed that

Table 3: Comparison of different cohorts observed at the same age

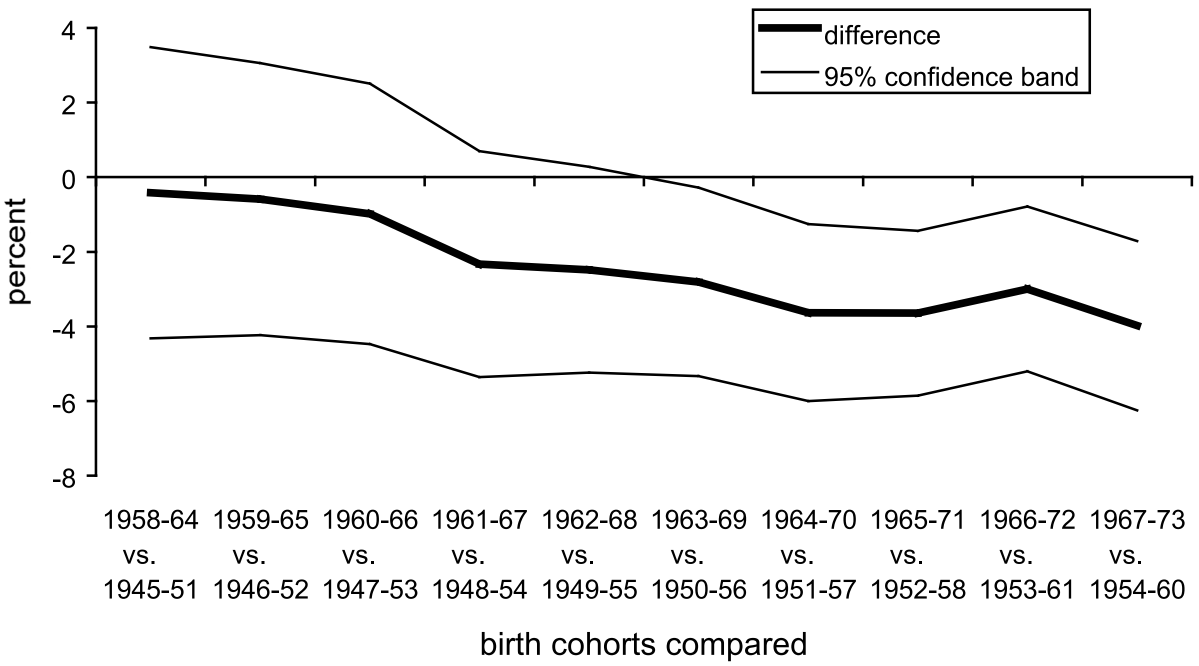
	men				*women*			
	29–37 yrs old		*49–57 yrs old*		*29–37 yrs old*		*49–57 yrs old*	
Variable	**Coeff.**	**t–stat.**	**Coeff.**	**t–stat.**	**Coeff.**	**t–stat.**	**Coeff.**	**t–stat.**
Second cohort / 1997	0.497	4.930	0.258	1.768	0.458	2.787	0.410	1.564
Schooling	0.082	11.543	0.087	9.577	0.067	5.445	0.097	6.104
Schooling x 2^{nd} cohort	–0.024	2.924	–0.004	0.317	–0.025	1.769	–0.021	0.871
Tenure	0.643	0.856	1.359	3.101	0.010	0.964	0.006	0.862
$Tenure^2/100$	–0.230	0.059	2.192	2.011	–0.028	0.482	–0.004	0.198
Employment status: FT	0.631	0.942	4.129	2.468	0.019	1.952	0.014	0.660
FT Experience	0.043	3.519	0.005	0.628	0.020	1.561	0.006	0.800
FT $Experience^2/100$	–0.151	2.668	0.018	1.042	–0.080	1.085	–0.009	0.454
Parttime Dummy					0.011	0.147	–0.023	0.214
Employment status: PT					0.013	1.450	0.000	0.009
PT experience					–0.028	1.443	0.003	0.300
PT $experience^2$					0.235	1.386	0.010	0.238
Unemployment rate	–0.013	1.221	0.042	1.794	–0.030	2.258	–0.021	0.441
Married					–0.009	0.244	–0.033	0.560
Industry dummies	YES		YES		YES		YES	
Laender dummies	YES		YES		YES		YES	
Firm size dummies	YES		YES		YES		YES	
Number of observations	668		404		452		267	
R^2	0.394		0.560		0.451		0.539	
adjusted R^2	0.361		0.521		0.401		0.465	

Figure 5:Differences in the returns to education for specific cohorts observed at the same age

a) men



b) women



we are looking at an interval of 13 years only, while figures 1 and 3 show the differences for a much larger set of cohorts.

Finally, we turn to the estimations using educational *levels* rather than durations. It may be argued that in Germany, due to institutional reasons such as entry requirements to particular occupations, degrees reached are more important for individuals‘ earnings than years of schooling spent to achieve a given degree. In order to retain reasonable sample sizes at each educational level, we define cohorts as ten-year intervals rather than five-year intervals as before. Again, we allow for the interaction between the highest educational degree, on one hand, and labour market experience and its square, on the other. Table 4 gives the estimated coefficients while figure 6 displays the effects of education on wages for the most relevant of the educational categories. As in figure 2, experience is fixed at 15 years in the graphs.[7] For the same reason as before, the level of the returns to education should not be interpreted for cohorts observed at ages where 15 years of labour market exerience are uncommon.

In absolute terms, the steepest decline in educational premia between the 1945-54 and the 1965-74 cohorts is in the group of workers with university education. For men, the premium falls from 115 percent to 80 percent between these cohorts, while the decline for women is from 100 percent to 60 percent. In relative terms, we also find a marked decline in the premium for apprenticeship. For both men and women, the premium halved between the 1945-54 and the 1965-74 cohorts. By contrast, there is a positive cohort effect for male graduates from polytechnics. The university premium for the oldest female cohort appears to be extremely high (about 250 per cent), which is clearly due to the very small number of older university-educated women in the sample. The relative decline in the apprenticeship premium is particularly visible if we calculate the implied rates of return on human capital investment (table 5), taking account of the fact that a higher education level means less time in employment and, therefore, lower lifetime earnings.

[7] Figures are percentage differences to the lowest category (no apprenticeship, no higher education). They are calculated as:

$$exp(\beta_{level} + \beta_{level*cohort} + \beta_{level*experience} * 15 + \beta_{level*experience^2} * 225) - 1 .$$

Table 4: Returns to education levels, men and women

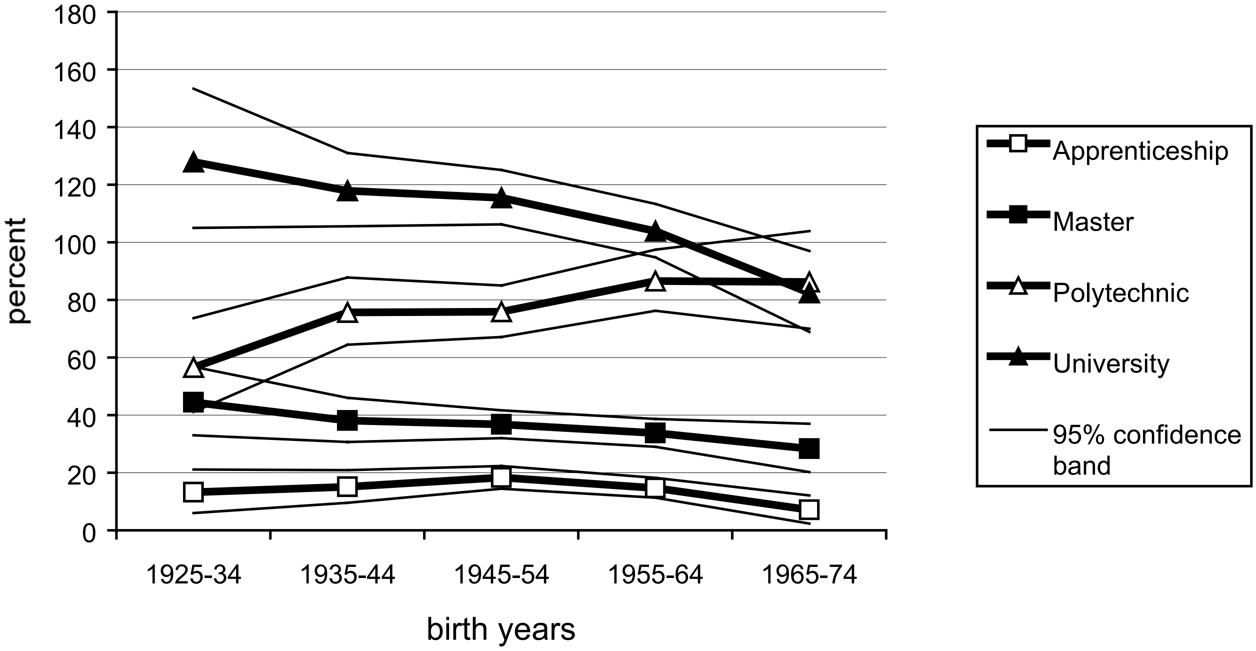
	Men		*Women*	
Variable	**Coeff.**	**t–stat.**	**Coeff.**	**t–stat.**
Cohort 1935-44	–0.048	–2.288	0.049	2.777
Cohort 1945-54	–0.099	–3.220	0.100	4.718
Cohort 1955-65	–0.116	–3.195	0.157	7.162
Cohort 1965-74	–0.092	–2.204	0.181	6.994
Apprenticeship	0.096	1.975	0.211	7.247
Master	0.382	5.759	0.443	7.074
Abitur	0.747	3.385	–0.424	–1.757
Abitur+Apprenticeship	0.240	2.613	0.361	4.212
Polytechnic	0.254	3.671	0.386	4.326
University	0.598	8.843	1.146	9.794
Apprenticeship x 35-44	0.016	0.684	–0.038	–1.483
Apprenticeship x 45-54	0.044	1.310	–0.031	–1.105
Apprenticeship x 55-64	0.013	0.324	–0.109	–3.843
Apprenticeship x 65-74	–0.056	–1.270	–0.154	–5.022
Master x 35-44	–0.044	–1.448	–0.175	–3.026
Master x 45-54	–0.054	–1.297	–0.122	–2.077
Master x 55-64	–0.076	–1.527	–0.207	–3.449
Master x 65-74	–0.118	–2.008	–0.269	–4.283
Abitur x 35-44			–0.261	–1.248
Abitur x 45-54	–0.820	–3.700	0.568	2.480
Abitur x 55-64	–0.656	–2.863	0.266	1.128
Abitur x 65-74	–0.749	–3.273	0.308	1.283
Abitur+Appr. x 35-44	0.184	2.912	–0.215	–2.239
Abitur+Appr. x 45-54	0.113	1.446	0.007	0.09
Abitur+Appr. x 55-64	–0.008	–0.088	–0.102	–1.201
Abitur+Appr. x 65-74	–0.176	–1.928	–0.189	–2.143
Polytechnic x 35-44	0.115	2.801	–0.165	–1.946
Polytechnic x 45-54	0.116	2.192	–0.302	–3.715
Polytechnic x 55-64	0.175	2.980	–0.037	–0.443
Polytechnic x 65-74	0.173	2.505	–0.155	–1.554
University x 35-44	–0.045	–0.971	–0.611	–5.195
University x 45-54	–0.056	–1.014	–0.564	–4.667
University x 55-64	–0.112	–1.867	–0.632	–5.274
University x 65-74	–0.223	–3.264	–0.793	–6.381
Apprent. x Experience	0.003	1.118	0.004	1.619
Master x Experience	0.000	0.103	0.001	0.246
Abitur x Experience	0.047	4.220	0.067	5.904
Abit.+App. x Experience	0.009	1.954	–0.001	–0.129
Polytech. x Experience	0.017	3.632	0.029	3.663
Univ. x Experience	0.028	7.031	0.017	2.91
Apprent. x Experience²	–0.008	–1.397	–0.012	–1.949

Table 4 (continued)

	Men		*Women*	
Variable	**Coeff.**	**t–stat.**	**Coeff.**	**t–stat.**
Polytech. × Experience²	–0.029	–2.493	–0.085	–4.201
Univ. × Experience²	–0.086	–7.847	–0.064	–2.855
Tenure	0.002	3.593	0.008	7.312
Tenure²/100	0.005	2.468	–0.014	–3.986
Employment status: FT	0.017	16.354	0.019	14.285
FT Experience	0.015	5.921	0.013	6.671
FT Experience²/100	–0.036	–6.707	–0.014	–2.916
Part-time dummy			0.053	5.734
Employment status: PT			0.010	7.074
PT Experience			0.002	1.577
PT Experience²/100			0.010	1.595
Unemployment rate	–0.009	–1.685	0.014	1.989
Unemployment rate²/100	0.073	1.167	–0.156	–2.003
Married			–0.023	–3.479
Year dummies	YES		YES	
Industry dummies	YES		YES	
Laender dummies	YES		YES	
Firm size dummies	YES		YES	
Number of observations	19004		13492	
R^2	0.505		0.415	
adjusted R^2	0.503		0.411	

Figure 6: Premia for education levels

a) men



b) women

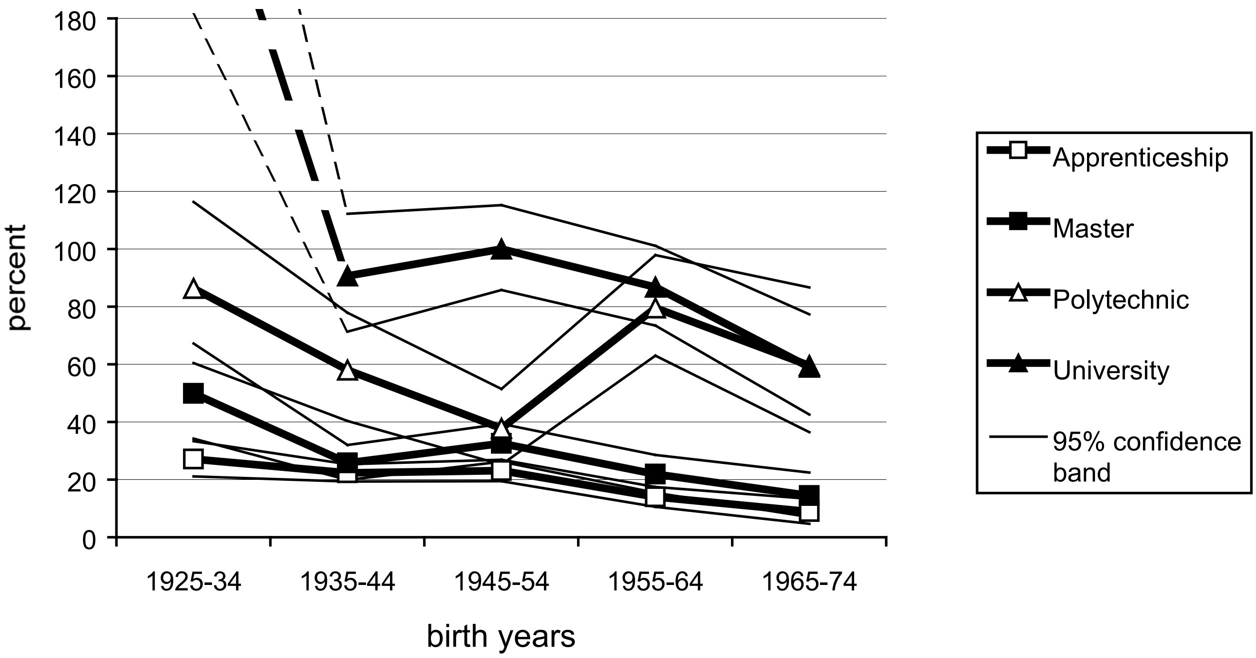


Table 5: Calculated rates of return of different educational levels

a) men

Education levels compared		*Rates of return*	
		Cohort 1945-54	Cohort 1965-74
secondary education only	apprenticeship (+ 3 yrs)	5.15	1.71
secondary education only	university degree (+8 yrs)	9.24	6.99

b) women

Education levels compared		*Rates of return*	
		Cohort 1945-54	Cohort 1965-74
secondary education only	apprenticeship (+ 3 yrs)	6.23	1.96
secondary education only	university degree (+8 yrs)	8.00	4.95

Note: Results are based on table 4. Implied rates of return are calculated for 15 years of potential working experience:

$$\exp\left\{\frac{1}{E_j}\Big(\beta_{jk} + \beta_{level*EXP} * 15 + \beta_{level*EXP^2} * 225 + \beta_{EXP} * (15 - E_j) - \beta_{EXP^2} * (15 - E_j)^2 - \beta_{EXP} * 15 + \beta_{EXP^2} * 225\Big)\right\}$$

where E_j are the years typically spent in education in order to reach a certain educational level.

5 Conclusion

We have found evidence for a decline in the returns to education in West Germany for cohorts born after the Second World War. The decline appears to be much stronger and affects a larger number of cohorts for women than for men. It seems to have taken place at high as well as low educational levels. Our findings also appear to be reasonably robust against different specifications of the experience part of the equation. In particular, we have relaxed the assumption that the returns of education are independent of labour market experience, allowing for years of schooling and years of labour market experience and it square to interact. We also used sub-samples of workers observed at the same age to get rid of possible age or work experience effects in the cohort-specific returns to education.

A question which we have not tackled in this paper is how these cohort effects can be explained. It appears that there are several main candidates for an explanation. First, there was a strong increase in female labour force participation. Second, educational attainment increased over the post-war cohorts, and third, West Germany experienced a baby boom which peaked in the mid-1960s. The expansion of female labour supply is certainly the prime source of the gender differences found in our estimations, although shifts in the demand for specific qualifications probably have mattered, too. The increase in educational attainment which enhanced the supply of qualified relative to unskilled workers could also have led to a reduction in educational premia. However, this explanation does not sit well with the fact that the decline is not limited to higher education but is also visible in the apprenticeship premia. This suggests that the development of educational premia, insofar as they affect both men and women, may have more to do with demography.

An objection to this explanation could be that, if it were correct, we should observe an increase in educational premia for individuals born after 1968 because cohort size declined hugely after that year. According to our estimations, however, returns to education continued to shrink. We do not find this objection compelling because it is still too early for a precise estimate of the educational premium for these cohorts, many of its members not having entered the labour market by the time they were observed. However, the question of what caused the drop in educational premia clearly needs to be addressed more carefully in future research.

References

Berger, M.C. (1985): The effect of cohort size on earnings growth: A re-examination of the evidence, *Journal of Political Economy*, 93, 561-573.

Card, D. (1999): The causal effect of education on earnings, in: Ashenfelter, O., D. Card (eds.): *Handbook of Labor Economics*, vol. 3, Amsterdam.

Dustmann, C., A. van Soest (1998): Public and private sector wages of male workers in Germany, *European Economic Review*, 42, 1417-1441.

Falk, M.; B. Koebel (1998): Determinanten der qualifikatorischen Arbeitsnachfrage in der westdeutschen Industrie 1978-90: FuE-intensive versus nicht FuE-intensive Industrien, in: Pfeiffer, F.; W. Pohlmeier (eds.): *Qualifikation, Weiterbildung und Arbeitsmarkterfolg*, Schriftenreihe des ZEW, 31, Baden-Baden.

Fitzenberger, B., R. Hujer, T.E. MaCurdy, R. Schnabel (1995): The dynamic structure of wages in Germany 1976 - 1984: A cohort analysis, University of Mannheim, Discussion Paper no. 533.

Fitzenberger, B., W. Franz (1998): Flexibilität der qualifikatorischen Lohnstruktur und Lastverteilung der Arbeitslosigkeit: Eine ökonometrische Analyse für Westdeutschland. In: Gahlen, B., H. Hesse, H.J. Ramser (Hrsg.): *Zunehmende Ungleichheit – Erklärungen und Konsequenzen*, Schriftenreihe des Wirtschaftswissenschaftlichen Seminars Ottobeuren, 27, Tübingen.

Freeman, R.B. (1979): The effect of demographic factors on age-earnings profiles, *Journal of Human Resources*, 14, 289-318.

Hartog, J., P.T.Pereira, J.A.C. Vieira (1999): Changing returns to education in Portugal during the 1980s and early 1990s: OLS and quantile regression estimators. Ms., University of Amsterdam.

Heckman, J., R. Robb (1985): Using longitudinal data to estimate age, period and cohort effects in earnings equations. In: Mason, W.M., S.F. Fienberg (eds.): *Cohort analysis in social research*, Beyond the identification problem, New York, Heidelberg.

Lauer, C., V. Steiner (2000): Returns to education in West Germany: An empirical assessment, discussion paper no. 00–04, Centre for European Economic Research (ZEW), Mannheim.

Macunovich, D.J. (1999), The fortunes of one's birth: Relative cohort size and the youth labor market in the United States, *Journal of Population Economics*, 12, 215-272.

Mincer, J. (1974): *Schooling, Experience and Earnings*, New York.

Stapleton, D.C., D.J. Young (1988): Educational attainment and cohort size, *Journal of Labour Economics*, 6, 330-361.

Steiner, V., K. Wagner (1998a): Has earnings inequality in Germany changed in the 1980s? *Zeitschrift für Wirtschafts- und Sozialwissenschaften*, 118, 29-60.

Steiner, V., K. Wagner (1998b): Relative earnings and the demand for unskilled labour in West German manufacturing, in: Black, S.W. (ed.): *Globalization, technological change, and labor markets*, Boston.

Welch, F. (1979): Effects of cohort size on earnings: The baby boom babies' financial bust, *Journal of Political Economy*, 87, S65-S97.

Zimmermann, K.F. (1991): Ageing and the labour market, Age structure, cohort size and unemployment. *Journal of Population Economics*, 4, 177-200.