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Educational Expansion and Its Heterogeneous Returns for Wage Workers

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Abstract

This paper examines the evolution of returns to education in the West German labour market over the last two decades. During this period, graduates from the period of educational expansion entered the labour market and an upgrading of the skill structure took place. In order to tackle the issues of endogeneity of schooling and its heterogeneous returns, we apply two estimation methods: Wooldridge's (2004) approach that relies on conditional mean independence and Garen's (1984) control function approach that requires an exclusion restriction. For the population of wage workers from the SOEP, we find that both approaches produce estimates of average returns to education that decrease until the late 1990s and increase afterwards. The gender gap in returns to education seems to vanish. Furthermore, we find that the so-called "baby boomer" cohort has the lowest average return to education in early working life. However, this effect disappears when the "baby-boomer" cohort grows older.

Zusammenfassung

Dieser Artikel untersucht die Entwicklung der Bildungsrenditen auf dem westdeutschen Arbeitsmarkt während der letzten beiden Jahrzehnte. In dieser Periode betraten Schulabgänger der Generation der Bildungsexpansion den Arbeitsmarkt und die Qualifikationsstruktur der Beschäftigten verbesserte sich deutlich. Um der Endogenität des Bildungserwerbs und der Heterogenität der Bildungsrenditen gerecht zu werden, vergleichen wir zwei Schätzmethoden: die Methode von Wooldridge (2004), die auf der Annahme der konditionalen 40 Mittelwertunabhängigkeit basiert, und den Kontrollfunktionsansatz von Garen (1984), der eine Ausschlussrestriktion voraussetzt. Beide Schätzmethoden kommen zu dem gleichen Ergebnis: Die durchschnittliche Bildungsrendite für

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unsere Stichprobe abhängig Beschäftigter aus dem Sozioökonomischen Panel (SOEP) ist von 6,6% im Jahre 1984 auf 4,9% bis Ende der 1990er Jahre gesunken, jedoch danach wieder auf 6,6% angestiegen. Die Bildungsrenditen zwischen Frauen und Männern haben sich in den vergangenen Jahren angeglichen. Es zeigt sich auch, dass die Babyboom-Generation die niedrigste Bildungsrendite am Anfang ihrer Karriere verzeichnet. Dieser anfängliche Nachteil scheint sich jedoch im Laufe der Erwerbsleben der Babyboom-Generation wieder zu verringern.

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1. Introduction

In West Germany, a major expansion of higher secondary and tertiary education occurred during the 1960s and 70s. Even though this is a moderate expansion from an international perspective (Müller / Wolbers, 2003), it significantly raised the qualification structure of the workforce when the students from the period of educational upgrading entered the labour market later on. For instance, in the year 2005, 40 percent of West Germans aged 25 to 30 held an upper secondary degree, as compared to 20 percent of West Germans aged 55 to 60 (Statistisches Bundesamt, 2006). A similar picture appears at the tertiary level, where 17 percent of West Germans aged 30 to 35 held a higher technical college or university degree, compared to 13 percent of West Germans aged 55 to 60. Our main research question is whether the upgrading of schooling "devaluated" average long-run labour market returns, taking the changing cohort and gender composition of the labour force into account.

The contribution of our study will be threefold. First, we investigate the evolution of heterogeneous returns to education in the twenty-two-year period from 1984 to 2006 to study the longer run labour market impacts of educational expansion. The empirical assessment is based on data from the German Socio-Economic Panel (SOEP). Second, we take the endogeneity and selectivity of school choice (Card, 1999; Willis / Rosen, 1979) into account, which results in models of heterogeneous rates of returns (see Blundell / Dearden / Sianesi, 2005; Flossmann / Pohlmeier, 2006, among others, and, Heckman / Lochner / Todd, 2008 for a discussion on the economics of rates of returns studies). A correlated random coefficient model is employed, where the explanatory variable "years of schooling" is measured as a continuous treatment variable, which can be correlated with unobserved heterogeneity. Identification is based on different assumptions. Following Wooldridge (2004), we identify the average return to education via conditional mean independence assumptions, and, following Garen (1984), a control function approach is employed that uses exclusion restrictions to control for selection on unobservable heterogeneity. Under the validity of its assumption, the latter approach provides addi-

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tional insights into the direction and size of selection on unobservables. Third, we differentiate the returns to education in West Germany by gender and birth cohort in order to learn more about trends for specific groups of workers. In this way, we investigate the relationship induced by female labour force participation and the rise of newborns until the 1960s and its decline afterwards.

Our findings indicate that the average returns to education decreased until the late 1990s and increased afterwards. Using Wooldridge's approach, our results vary between 4.9 and 6.6 percent for the average partial effect of an additional year of schooling, which are at the lower end of previous findings for Germany (Boockmann / Steiner, 2006; Flossmann / Pohlmeier, 2006; Lauer / Steiner, 2001, among others). Regarding the gender aspects, the average returns to education seem to have been larger for women during the 1980s and early 1990s. However, the gap decreases over time, which may be a consequence of increased participation of women. Furthermore, we find that the so-called "baby boomer" cohort has the lowest average return to education compared to the cohort before and the one thereafter (the former is characterised by lower and the latter by sharply decreasing cohort sizes). While this finding seems to be in line with the literature on wages and cohort size (Macunovich, 1999), in our data the effect vanishes when the "baby-boomer" cohort grows older.

This paper is organized as follows: Section 2 discusses factors that influence returns to education over time. Section 3 develops the idea of heterogeneous returns to education in a correlated random coefficient model, and compares estimates from the conventional, as well as Wooldridge's (2004) and Garen's (1984), approaches. Section 4 describes the data set and variables used. In section 5, we discuss estimation results differentiated by estimation techniques, gender, and cohorts over time. Section 6 concludes the paper.

2. Educational Expansion, Wages, and the Labour Market in West Germany

Educational attainment started to increase in the 1960s in West Germany, eventually leading to the upgrading of educational qualifications in the labour market. In our sample of workers from West Germany, extracted from the SOEP, average years of schooling increased for women (men) from 11.3 (11.9) years in 1984 to 12.8 (12.9) years in 2006. In recent years, male and female average educational attainment has become similar. In a standard economic supply and demand labour market framework, a rising supply of (high) skilled workers induces, ceteris paribus, a decline in the returns to education.

Another argument focuses on the "quality" of different educational groups in terms of unobserved characteristics that may have changed in the process of educational expansion (Taber, 2001). High-educated workers may have on average higher levels of unobserved ability or motivation, which results both

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in higher educational attainment and higher wages. They also may select jobs where they expect the relatively highest monetary returns (Willis / Rosen, 1979). However, educational expansion may have resulted in institutions starting to accept students increasingly from the lower end of the distribution of student abilities, such that weaker students might have been admitted to higher education, leading to a decrease in the average productivity level of higher educated workers (Hægeland, 2001, Walker / Zhu, 2005). As a consequence, the degree of positive selection into the higher educational groups might have decreased and lowered the observed returns to education in the course of educational expansion.

Besides the educational expansion, there exist other factors that have influenced demand and supply conditions on German labour markets over the last two decades. Some important factors have been, for instance, increasing female labour market participation, birth cohort sizes, wage determination processes, and skill-biased technological change. In West Germany, the female participation rate has been rising during the last decades, growing closer to the male participation rate and resulting in increased competition for college slots and labour market positions. Based on the decreasing gender-gap in educational attainment and labour market participation, a convergence of gender-specific returns to education can be expected.

West Germany, as well as many other western countries, experienced a demographic change due both to a baby boom that peaked during the mid-1960s and sharply decreasing cohort sizes afterwards. Changes in the number of births alter the supply of workers entering the market about twenty years later, i.e., during our period of investigation. If larger birth cohorts enter the labour market and substitution in production is limited between younger and older workers, ceteris paribus, a downward pressure on returns to education for labour market entrants arises (Freeman, 1979, among others, Macunovich, 1999). Therefore, one may expect decreasing returns to education for the baby boom cohorts and increasing returns to education for individuals born after 1964 (when cohort sizes started to decline sharply). In addition, there was fierce wage competition for entrants due to unemployment rates as high as ten percent in Germany. Compared to workforce entrants, incumbent workers in Germany enjoy some protection against wage competition due to strong unions and / or efficiency wage considerations (Franz / Pfeiffer, 2006). Because large cohorts are absorbed gradually by the labour market when experience increases, we expect lower returns to education for labour market entrants.

The computer revolution that started around 1970 changed the organisation of labour away from routine manual tasks to non-routine analytical and creative tasks (Autor / Katz / Kearney, 2006; Spitz-Oener, 2006). The demand shift towards analytical skills presumably favoured the high skilled and may even have increased returns to education, despite increasing supply (Acemoglu, 2002).

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To sum up, we expect supply side factors like educational expansion and the increase in female participation to lower the returns to education (in a ceteris paribus sense). Similarly, supply side factors such as a decreasing cohort size and demand side factors such as skill-biased technological change and workplace innovations are likely to increase the returns to education. In Germany, the impact of educational expansion on wages is also likely to be formed by the process of wage determination, the regulation of labour, as well as by the rate of unemployment and active labour market policies. We would like to analyze the empirical evolution of the returns to education in West Germany from 1984 to 2006 that resulted from the factors outlined above.

3. Econometric Approach

Our empirical framework is the correlated random coefficient model (Blundell / Dearden / Sianesi, 2005; Heckman / Vytlacil, 1998; Wooldridge, 2004):

$$\ln Y_i = a_i + b_i S_i \quad \text{with} \quad a_i = a' X_i + \varepsilon_{ai} \quad \text{and} \quad b_i = b' X_i + \varepsilon_{bi} \tag{1}$$

where the outcome variable, $\ln Y_i$, is the natural log of wages and the explanatory variable, S_i, is the years of schooling of individual *i*. This equation has been derived from optimal schooling choice where education is determined by a respective individual's observed and unobserved marginal benefits, and the costs of schooling (Card, 1999). The model has an individual-specific intercept a_i and slope b_i that may depend on observable variables X_i and unobservable heterogeneity ε_{ai} and ε_{bi}. The heterogeneity components capture influences from gender, family background, age, preferences, ability, etc., such that a_i and b_i represent random coefficients. We do not assume that b_i and S_i are independent. a_i and S_i, as well as b_i and S_i, can be correlated (Wooldridge, 2004). Since individuals with higher expected benefits from education are more likely to participate longer in education, the returns to education b_i may in general be correlated with S_i if variation in unobserved (to the econometrician) benefits implies positive self-selection. In this case, the schooling variable is influenced by its own coefficient, yielding an endogeneity problem.

Our research interest is the effect of S_i on $\ln Y_i$, represented by b_i in equation (1). In this model, the return to education varies across individuals in both the observable heterogeneity in returns X_i and the unobserved individual-specific returns to schooling, ε_{bi}. The resulting distribution of returns will be summarized with the average partial effect (APE) (Flossmann / Pohlmeier, 2006; Wooldridge, 2004). APE measures the average return per additional year of education for a randomly chosen individual from our population:

$$E(\partial \ln Y / \partial S) = E(b_i) = \beta\,. \tag{2}$$

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The earnings equation (1) nests more specific models. If returns to education are homogenous, the outcome equation can be re-written as the classical Mincer-type of earnings function (Blundell / Costa Dias, 2000):

$$\ln Y_i = a'X_i + \bar{b}S_i + \varepsilon_{ai} \tag{3}$$

where $\bar{b}$ is the common rate of return. Unobserved heterogeneity may exclusively enter the intercept of the wage equation but not the slope coefficient. In that case, there might still be endogeneity problems if the unobserved general individual earnings capacity ε_{ai} is correlated with S_i. One appealing feature of the general approach (1) is that variation in unobserved heterogeneity affects the slope as well, i.e., unobserved heterogeneity influences the wage effect of education.

3.1 Conventional Methods

When estimating (1) by OLS, there are three potential sources of bias. First, if individuals with high absolute earnings capacity acquire more education and earn higher wages, schooling S_i will be positively correlated with ε_{ai} (Griliches, 1977). This ability bias induces an upward bias in the estimated average return (Behrman / Rosenzweig, 1999). Second, classical measurement error in the schooling variable S_i induces a downward bias (Griliches, 1977). Third, a bias can exist if individuals differ in their relative earnings capacity and act upon their comparative advantage when choosing their level of education (Willis / Rosen, 1979). If returns to education are homogeneous, the latter bias is absent.

In the literature on the return to schooling in Germany, instrumental variables (IV) methods are commonly used to handle the endogeneity problems. For instance, Lauer / Steiner (2001) estimate homogeneous returns to education using different family background variables as instruments. The results depend on the instruments used and vary between 6.6 and 14.8 percent. However, when schooling is also correlated with unobserved individual heterogeneity, standard IV methods may fail to identify APE (Heckman / Li, 2004). Ichino / Winter-Ebmer (1998) are among those who instead estimate the local average treatment effect (LATE) of schooling for Germany using different instruments. Since each instrument implies its own LATE and the group of compliers cannot be identified without further assumptions, this may be regarded as a drawback. However, LATE is especially interesting when school reforms are used as instruments since LATE measures the returns to schooling for those who changed their level of schooling because of the reform. With this approach, Pischke / van Wachter (2008) find rather low marginal returns to education in Germany. In our empirical analysis, we employ methods that reduce the potential bias from OLS and IV techniques. Furthermore, in contrast to the LATE interpretation of IV techniques, we are interested in assessing the APE.

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3.2 Wooldridge's (2004) Conditional Mean Independence (CMI) Approach

The methods we employ rely on different identifying assumptions: Wooldridge's (2004) conditional mean independence (CMI) approach and Garen's (1984) control function (CF) approach. According to Wooldridge (2004), APE is identified by the following two assumptions if the linear outcome equation (1) holds:

$$(4) \qquad E(\ln Y_i | a_i, b_i, S_i, X_i) = E(\ln Y_i | a_i, b_i, S_i) = a_i + b_i S_i$$

$$(5) \qquad E(S_i | a_i, b_i, X_i) = E(S_i | X_i) \quad \text{and} \quad Var(S_i | a_i, b_i, X_i) = Var(S_i | X_i)$$

where the elements of X_i are suitable proxy variables for the observed and unobserved heterogeneity, i.e., the X_i elements should be good enough predictors of S_i. The first assumption postulates that the vector X_i is redundant given S_i and (a_i, b_i) in the structural conditional expectation (4). This identification assumption obviously holds since the control variables X_i enter the earnings function through a_i, b_i, and S_i only. The second assumption is a redundancy condition of the form that both heterogeneity terms a_i and b_i are redundant in the first two conditional moments of the schooling variable S_i, when conditioning on a set of covariates X_i. The latter is the strongest assumption as it requires a differentiated set of variables that control sufficiently for observable and unobservable heterogeneity. These conditional moment independence (CMI) conditions are a weaker form of conditional independence assumptions (CIA) (Wooldridge, 2002: 607). Based on assumptions (4) and (5), Wooldridge (2004) derives the following estimator for APE:

$$(6) \qquad \hat{\beta} = \frac{1}{N} \sum_{i=1}^{N} \left(\left(S_i - \hat{E}(S_i | X_i) \right) \ln Y_i \Big/ \hat{Var}(S_i | X_i) \right).$$

Because $\ln Y_i$, S_i, and X_i are observable, one needs to estimate the conditional mean and variance, $E(S_i / X_i)$ and $Var(S_i / X_i)$. Since S_i is nonnegative, simple linear models have shortcomings. Therefore, we employ a generalized linear model (GLM) with a Poisson distributional assumption for years of schooling S_i:

$$(7) \qquad E(S_i | X_i) = e^{\gamma' X_i} \quad \text{and} \quad Var(S_i | X_i) = \sigma^2 e^{\gamma' X_i}.$$

This specification guarantees positive estimates of both conditional mean and variance. Contrary to the standard variance assumption of equality between the conditional variance and the mean equation, (7) relies on the weaker Poisson GLM variance assumption that allows the variance-mean ratio to be any positive constant (Wooldridge, 2002). A consistent estimator of σ^2 is ob-

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tained as the mean of squared Pearson residuals. Since analytical standard errors have not been developed so far, standard errors of the APE are bootstrapped.

3.3 Garen's (1984) Control Function (CF) Approach

Garen (1984) proposed a possible alternative solution to the random coefficient estimation problem – called the control function (CF) approach – that is similar to Heckman's (1978) two-step estimator. While the standard IV approach does generally not identify APE in the heterogeneous returns context, the CF approach does. The CF approach is implemented by a simultaneous modelling of both schooling and wages. Hence, an explicit model of the schooling selection process, which relates the rule for assigning individuals to treatment to the potential treatment outcomes, is required:

$$S_i = c'X_i + dZ_i + v_i \quad \text{with} \quad E(v_i|Z_i, X_i) = 0 \tag{8}$$

where both X_i and Z_i influence the educational decision and v_i represents the usual error, incorporating unobserved components that determine the choice of education. Z_i is an exclusion restriction, i.e., it should have no correlations with unobserved heterogeneity in the wage equation. The error terms v_i, ε_{ai}, and ε_{bi} are normally distributed with zero means and positive variances that are possibly correlated with each other.[1] Following Garen (1984), one can formulate an augmented wage equation based on the linear outcome specification (1) of the form:[2]

$$\ln Y_i = a_i + \beta S_i + \gamma_1 \nu_i + \gamma_2 \nu_i S_i + w_i \tag{9}$$

where $\gamma_1 \nu_i$ and $\gamma_2 \nu_i S_i$ re the control functions with $\gamma_1 = \mathrm{cov}(\varepsilon_{ai}, \nu_i)/\mathrm{var}(\nu_i)$ and $\gamma_2 = \mathrm{cov}(\varepsilon_{bi}, \nu_i)/\mathrm{var}(\nu_i)$. Once these terms are included in the outcome equation (and implicitly subtracted from its error term), the error term w_i has all the desirable properties, i.e., it is orthogonal to all of the regressors in the new equation: $E(w_i|X_i, S_i, v_i) = 0$ (Heckman / Robb, 1985a).

This model can be estimated using a generalization of the two-step approach. In the first step, an estimation of the schooling choice is used to construct the control functions that are included as additional regressors in the augmented wage equation. The estimated coefficients of v_i and $v_i S_i$ provide information about the selection on the unobserved absolute earnings capacity term and about selection on the comparative earnings capacity, respectively. If

[1] This trivariate normality assumption can be weakened to the condition that conditional expectations of the unobserved earnings components ε_{ai} and ε_{bi} are linear in the residual of the selection equation (Blundell et al., 2005).

[2] Thus, both Garen's CF approach and Wooldridge's CMI approach rely on the assumption of the linearity of the wage equation.

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an individual attains a higher (lower) level of education than according to our expectations, the value of v_i is positive (negative). For example, if the coefficient γ_1 of the first control function is positive, this implies that the unobserved factors that lead to educational "over-achievement" (positive v_i) have a positive impact on earnings. The sign of the coefficient γ_2 of the second control function describes how this effect changes with increasing levels of education. Following the comparative advantage hypothesis (Willis / Rosen, 1979), we expect that γ_2 is positive, i.e., those with unexpectedly large amounts of schooling (positive v_i) tend to earn more than the others with higher education. Based on their higher unobserved marginal returns, they are likely to enrol into higher education according to their comparative advantage.

4. Data and Variables

The empirical analysis is based upon samples from 23 waves of the German Socio-Economic Panel Study (SOEP, see Haisken-DeNew / Frick, 2006). SOEP contains information on education, employment, earnings, and retrospective information about family background. We include SOEP refreshment samples from 1998 and 2000. Due to lack of comparability, foreign-born individuals were excluded from the sample. Furthermore, the analysis is restricted to West German citizens for comparison reasons. Self-employed workers are excluded from the sample since they are exposed to different earnings-generating mechanisms. The resulting sample is composed of full-time dependent workers between the ages of 25 and 60 who work 30 hours or more per week. After eliminating observations with missing values, we obtain a final sample size that ranges from 1,545 observations in 1984 to 3,965 in 2000.[3]

The dependent variable is the natural logarithm of real earnings per hour worked. The measure of years of schooling is derived by attaching a standard number of years to the highest educational level (cf. Table 1). As control variables, gender and individual age (in linear and quadratic terms) are included. We use age variables instead of labour market experience because the latter might be endogenous with respect to schooling. We do not include variables such as labour market experience, tenure, or occupation that are not predetermined to education and, thus, not exogenous. This should guarantee that the "gross" effect of schooling on wages is assessed, notwithstanding any later intervening mechanisms.

To justify the conditional moment independence assumptions, a set of family background information is utilized that is covered by recall questions and that is available for a sufficient number of persons in each wave considered.

[3] Although sample size almost doubles due to refreshment samples, the following estimation results do not change substantially if the new samples are excluded.

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Table 1

Description of variables and summary statistics

Variable Name	*Description*	*Mean (std. dev.)* 1984	1995	2006
log wage	Log gross hourly wage	2.36 (0.42)	2.56 (0.41)	2.58 (0.50)
years of education	Years of education: constructed with standard times for highest educational and vocational degree obtained: no degree (7 years), lower secondary (9 years), intermediate secondary (10 years), technical secondary (12 years), higher secondary (13 years), vocational training (+1.5 years), vocational school (+2 years), higher technical college (+3 years), university (+5 years)	11.68 (2.48)	12.20 (2.66)	12.84 (2.76)
age	Age in years	40.21 (9.81)	40.18 (9.66)	43.10 (8.87)
age squared	Age squared			
female	Dummy for sex (1 = female; 0 = male)	0.35	0.39	0.43
Father's Education				
elementary	Reference category: compulsory education or less	0.17	0.12	0.12
compulsory + voc	Compulsory education plus vocational training	0.65	0.64	0.60
secondary intermediate	Secondary intermediate education, with / without vocational training	0.09	0.11	0.14
full secondary	Full secondary education (Abitur), with / without vocational training	0.03	0.04	0.03
university	University / university of applied sciences	0.07	0.09	0.11
Mother's Education				
high education	Dummy (1= intermediate secondary, full secondary, university; 0 rest)	0.12	0.17	0.23
Father's Occupational Position				
blue collar	Reference category: father blue collar	0.49	0.47	0.44
white collar	Father white collar worker	0.18	0.14	0.13
self-employed	Father self-employed	0.16	0.21	0.25
civil servant	Father civil servant	0.11	0.12	0.11
Place of Socialisation				
rural socialisation	Dummy (1 = rural socialisation; 0 = urban socialisation, i.e. city, big town, small town)	0.40	0.37	0.38
Family Composition				
number siblings	Number of siblings	1.70 (1.78)	1.68 (1.68)	1.66 (1.64)
Number of observations		1,545	2,075	3,477

Note: Standard deviations for continuous variables in parentheses.

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Family background, as well as parental educational and occupational attainment proxies the parental influence on educational attainment and later employment carriers (Erikson / Jonsson, 1996, among others). The measure for parents' education follows the CASMIN educational classification, which has the advantage of combining information on the highest school degree and the highest vocational degree of the parents (Erikson / Goldthorpe, 1992). The CASMIN categories have been summarized in five categories for fathers and in a dummy-variable for mothers' higher education (cf. Table 1). There are four categories of parents' occupational position. These categories should proxy for the economic circumstances of the family, which affect educational choice by influencing costs of schooling. A further proxy for costs of schooling is the dummy variable "rural socialisation" (Card, 1995).

For the control function approach, the number of siblings is used as an exclusion restriction. We assume that it satisfies the two conditions for valid instrumental variables (Wooldridge, 2002). First, Becker / Tomes (1976), and Hanushek (1992), among others, hypothesize a positive correlation between the number of siblings and individual educational attainment even after controlling for other family background characteristics. Parents try to optimally allocate financial and non-financial resources to their children, who compete for the attention and resources of their parents. Therefore, educational achievement and total family size might be negatively related given limited educational resources.

Second, the instrumental variable should be uncorrelated with unobserved individual's earnings capacities, i.e., the number of siblings should not have an effect on income other than the indirect effect transmitted over educational attainment. Because we control for a set of other family background variables like parents' education, occupation, and the place of socialisation, we do not expect a non-negligible, systematic, and independent effect of the number of siblings on earnings. In the case of Wooldridge's (2004) CMI approach, we do not include the number of siblings because we assume that it has the character of an exclusion restriction and, thus, it should not serve as a simple control variable. Table 1 gives an overview of the variable definitions and summary statistics in selected years.

5. Estimation Results

5.1 Evolution over Time: Comparison of Different Estimation Techniques

Figure 1 compares the evolution of our three different estimates of individual returns to education in West Germany during the period 1984 to 2006 (for detailed estimation results, see Table A.1). Other than the results from a Mincerian OLS regression with years of schooling and controlling for age in

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linear and squared functional form on log wage, the APE from the conditional mean independence (CMI) approach and from the control function (CF) approach are graphed. With OLS, we find a slight downward trend in the evolution of returns to schooling until the late 1990s. The returns to one additional year of education fell from 7.4 percent in 1984 to 5.4 percent in 1998. From 1998 onwards, we find increasing returns to education reaching a new local maximum of 6.8 percent in 2002. The estimates until 1998 are in line with the findings of Lauer / Steiner (2001), among others. We are not aware that the increase in returns starting around 1998 has been documented so far.

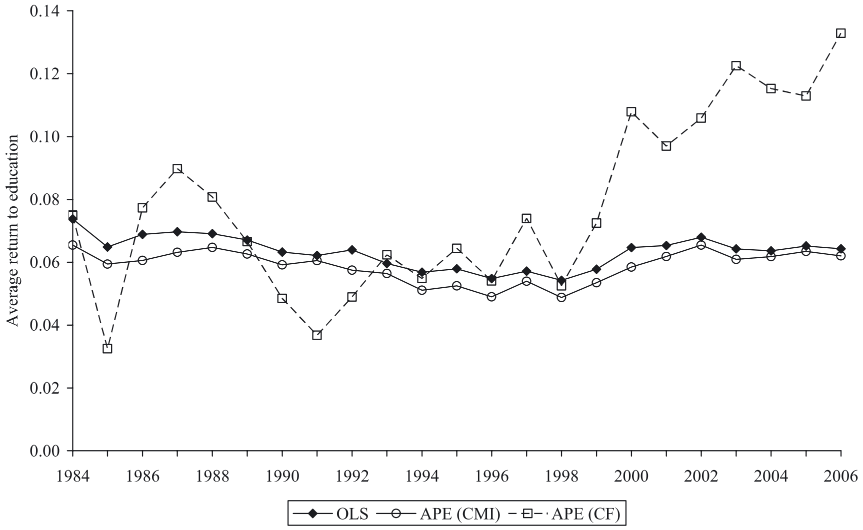

Figure 1: Average return to education, 1984 – 2006:
OLS, CMI, and CF approach compared

Interestingly, the APE estimated under conditional moment independence (CMI) follows a fairly similar evolution pattern over time. Both approaches also produce similar small standard errors varying between 0.003 and 0.004 in the observation period. However, there are differences. First, the estimated APE is always lower than standard OLS, between 0.2 and 1 percentage points. According to our interpretation, this difference reflects the potential ability bias from OLS estimates. Taking into account the heterogeneity of returns to education and controlling for family background variables, the CMI approach controls to a certain degree for positive ability. Second, although OLS and APE estimates are comparable over time, their content varies. APE measures the average of the distribution of heterogeneous returns, whereas OLS measures the average return to education that is homogenous for all individuals.

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Compared to the literature, our estimate of the APE seems to be rather low. Maier et al. (2004), for instance report an estimated APE of 8.7 percent for the year 1999 for German male workers.

The CF approach has been implemented in a two-stage estimation procedure (for detailed estimation results, see Table 2). The first stage, the educational attainment selection equation, has been used for testing the validity of the instrumental variables. A regression that includes the number of siblings in a simple OLS log-wage equation, together with other family background variables, was insignificant, suggesting that the number of siblings seems to be a reasonable exclusion restriction. According to our findings, the number of siblings has a strongly negative influence on educational attainment, holding constant other family background characteristics. This seems to be in line with the literature mentioned in the previous section above.

The evolution pattern of the estimated APE under CF approach deviates substantially from the CMI results. First, the yearly estimates derived from the CF approach are more volatile and less precise. Hence, the standard errors are higher (cf. Table A.1).[4] Detailed tests show that the deviations are usually not significant during the 1980s and the initial fluctuations (see Figure 3) might be a consequence of fewer observations for the 1980s. Second, there is a stronger and significant increase of the APE after 1998 compared to the CMI approach. In 2006, the APE is 13 percent, which is in line with some recent IV studies for Germany (Flossmann / Pohlmeier, 2006). From a methodological point of view, the deviations reflect differences in the identification strategies. From a substantive point of view, the discrepancy during the last years could be a hint for changing selection on unobservables, which the CF approach controls for.

Detailed inspections of the CF results show that the control function coefficient for the selection on unobserved absolute earnings capacity is positive during the 1980s and 90s, decreases over time, and becomes negative in the last years (cf. Table 2). However, the coefficients are insignificant, which is probably related to the fact that the educational and occupational background of the family is a reasonable proxy for the absolute earnings capacity. The coefficient of the control function for the selection on comparative earnings capacity is always negative, which contradicts the comparative advantage hypothesis. In our data, we find that those with unexpectedly high amounts of schooling ("educational over-achievement") have lower marginal returns to education, which seems to be similar to findings from Maier et al. (2004), and Pischke / van Wachter (2008). There are individuals who have done worse after more schooling. The effect is significant in the period from 1984 to 1989, and especially from 2000 onwards.

4 They vary around 0.04 in the 1980s and between 0.02 and 0.03 afterwards, i.e., they are about ten times larger than the CMI and OLS standard errors.

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Table 2

Returns to education, 1984 – 2006: CF approach

	1. Stage		2. Stage				N
	IV: Number of siblings		Selection on unobserved absolute earnings capacity		Selection on unobserved comparative earnings capacity		
	coeff.	(s.e.)	coeff. γ_1	(s.e.)	coeff. γ_2	(s.e.)	
1984	–0.091***	(0.032)	0.034	(0.081)	–0.003***	(0.002)	1,545
1985	–0.129***	(0.032)	0.081	(0.132)	–0.004**	(0.002)	1,600
1986	–0.094***	(0.031)	0.053	(0.096)	–0.005***	(0.002)	1,682
1987	–0.122***	(0.031)	0.062	(0.054)	–0.006***	(0.002)	1,775
1988	–0.132***	(0.031)	0.032	(0.047)	–0.003***	(0.002)	1,798
1989	–0.128***	(0.030)	0.059	(0.047)	–0.004***	(0.002)	1,922
1990	–0.151***	(0.029)	0.036	(0.035)	–0.002	(0.001)	2,007
1991	–0.148***	(0.028)	0.050	(0.033)	–0.002	(0.001)	2,122
1992	–0.164***	(0.028)	0.030	(0.031)	–0.002	(0.001)	2,107
1993	–0.180***	(0.028)	0.032	(0.036)	–0.003*	(0.002)	2,124
1994	–0.185***	(0.029)	–0.004	(0.031)	0.000	(0.002)	2,082
1995	–0.188***	(0.031)	0.023	(0.032)	–0.002	(0.002)	2,075
1996	–0.182***	(0.031)	0.024	(0.033)	–0.002	(0.002)	2,057
1997	–0.190***	(0.031)	0.018	(0.032)	–0.003*	(0.002)	2,011
1998	–0.220***	(0.031)	0.009	(0.026)	–0.001	(0.001)	2,145
1999	–0.211***	(0.031)	0.001	(0.029)	–0.001	(0.001)	2,163
2000	–0.158***	(0.023)	0.011	(0.031)	–0.004***	(0.001)	3,965
2001	–0.154***	(0.023)	–0.001	(0.028)	–0.002***	(0.001)	3,961
2002	–0.137***	(0.024)	0.009	(0.035)	–0.003***	(0.001)	3,668
2003	–0.158***	(0.025)	–0.010	(0.035)	–0.004***	(0.001)	3,476
2004	–0.150***	(0.025)	–0.011	(0.036)	–0.003***	(0.001)	3,366
2005	–0.153***	(0.026)	–0.010	(0.040)	–0.003**	(0.002)	3,220
2006	–0.157***	(0.027)	–0.001	(0.038)	–0.005***	(0.001)	3,477

Notes: (1) The first stage includes additional regressors such as gender, age, rural socialisation, educational and occupational background of the parents. (2) The second stage includes additional regressors such as years of education, gender, age, rural socialisation, educational and occupational background of the parents. The IV number of siblings is excluded. (3) Standard errors on the second stage are bootstrapped each with 500 repetitions. (4) Significant: *** at the 1% level; ** at the 5% level; * at the 10% level.

The existence of negative selection on unobservables in both components of the control function in the new millennium partially explains the increasing gap between the OLS and CF estimates. This evidence for a negative selection on unobserved factors supports the idea that educational expansion resulted in institutions digging deeper into the distribution of student abilities so that

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weaker students might have been admitted to higher education. Whereas this supply-sided process works more slowly, an explanation for the acceleration of this process in the last years might be demand-induced change in returns to unobserved skills. Technological change such as the IT revolution might have changed the skill requirements of jobs (Ludwig / Pfeiffer, 2006; Spitz-Oener, 2006) and / or recently devaluated the labour market value of some unobserved skills that have driven children into higher education in previous times (Taber, 2001). The recent increase in returns to education should however not be over-interpreted due to the higher impression of CF compared to CMI estimates and its different identification strategy.

To summarize our findings so far: Independently of the method used, the returns to education (APE) were fairly constant during the 1980s and 1990s in (West) Germany. Despite a continuous upgrading of educational qualification however, they started to increase from 1998 onwards, although only moderately. This finding seems to be line with rising wage inequality in Germany that set in around 1994 (see Gernandt / Pfeiffer, 2007).

5.2 Gender and Cohort Effects

The following analysis differentiates gender and birth cohorts to take a second look at the recent increase in the APE. The comparison rests on the CMI approach because it produces lower standard errors than the CF approach. The average return to an additional year of education in our population of women from the SOEP declined from 8.6 percent in 1984 to about 4.9 percent in 1996 (cf. Figure 2). As the return to education decreased more for women than for men, the gender gap diminished, which is in line with findings from Lauer / Steiner (2001), among others. While women registered significantly higher relative returns, especially during the 1980s, statistical tests indicate that this difference is no longer significant after 1995. According to our interpretation, the convergence in gender-specific returns to education should be mainly the result of the female educational expansion and rising labour force participation, perhaps also intensified by non-neutral technical progress (Spitz-Oener, 2006). Since 1999, the APE is increasing again for women as well as for men. The gap does not become significant again.

In a further step, the average returns to education are compared for four different birth cohorts. We differentiate the cohorts based on relative birth cohort sizes to avoid interpretation problems stemming from self-selection into the labour market. Cohort size in the labour market might be endogenous because individuals change their educational attainment and labour market entry with respect to cohort size (Berger, 1989; Macunovich, 1999). Each birth cohort is composed of eight years: people born between 1942 and 1949, 1950 and 1957, 1958 and 1965, and 1966 and 1973. The cohort boundaries are

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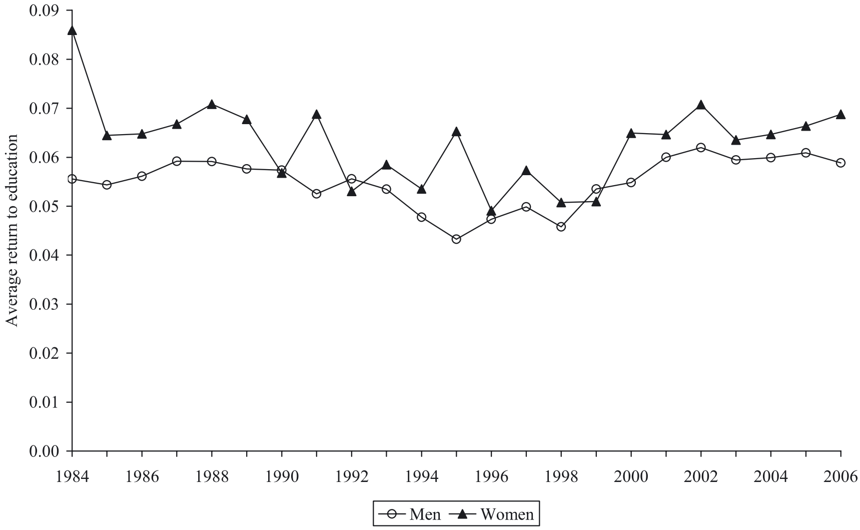

Figure 2: Average partial effect (CMI approach) by gender, 1984 – 2006

geared to cohort sizes: The first cohort (1942 – 49) has low birth rates due to the 2nd World War and the post-war period. The second cohort (1950 – 57) is of relatively constant size. The third cohort (1958 – 65) is the "baby boom" cohort, with strongly increasing cohort sizes peaking in 1964. Finally, the fourth cohort (1966 – 73) is characterized by a sharply declining cohort size. In order to have cohorts with a sufficient number of observations, our estimations are restricted to birth cohorts that are older than 27 to 34 years, e.g., we estimate APE for cohorts born 1958 – 65 starting at the year 1992.

Time, cohort, and life cycle effects cannot be disentangled empirically because it is impossible to observe two different birth cohorts at the same age and in the same year (Heckman / Robb, 1985b). Nevertheless, we try to shed some light on these effects. To empirically assess birth cohort effects in average returns to education in Germany, different birth cohorts are compared at the same age during their early working life. Figures 3 and 4 display estimation results for the four birth cohorts at given age intervals. We form age intervals of seven years in accordance with the birth cohort intervals and in order to have enough observations.[5] For example, all cohorts at the first observation point in Figure 3 are observed at ages 27 – 34. However, we do this for the

[5] The cohort comparison in figure 3 relies for each age group on about 398 – 502 cases for the cohort 1950 – 57, 599 – 641 cases for the cohort 1966 – 73, and 749 – 871 cases for the cohort 1966 – 73. The cohort comparison in Figure 4 relies for each age group on about 356 – 415 cases for the cohort 1942 – 49, 478 – 532 cases for the cohort 1950 – 57, and 995 – 1235 cases for the cohort 1958 – 65.

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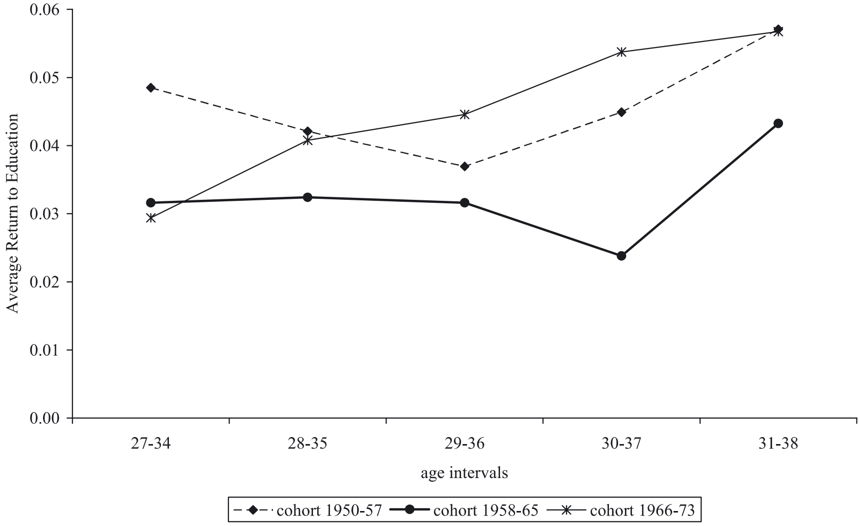
cohort 1950 – 57 in 1984, the cohort 1958 – 65 in 1992, and the cohort 1966 – 73 in 2000, such that the birth cohort effects at given age might be confounded with time effects. We repeat this exercise for several age intervals in order to test how robust the results are with regard to the chosen age interval and in order to see whether the birth cohort differences change when the young workers get established in the labour market.



Remark: Birth cohorts are observed at the same age but at different points in time. The cohort 1950 – 57 is observed during the period 1984 – 1988, the cohort 1958 – 65 is observed during the period 1992 – 1996 and the birth cohort 1966 – 73 during the period 2000 – 2004.

Figure 3: Average partial effect (CMI approach) by birth cohorts at same age intervals (27 – 38)

Figure 3 reveals that the "baby-boomer" cohort has the lowest average return to education compared to the cohort before (1950 – 57) and the cohort thereafter (1966 – 73). A larger cohort size seems to reduce the average return to education at young ages (27 – 38 years old).[6] As previously mentioned, however, the cohort effects might be confounded by time effects because the cohorts are observed at different time points, although we keep constant the age of the birth cohorts. If the effect is caused by the different years of observation rather than by the birth cohort, we should observe similar patterns of returns to education for all birth cohorts in the respective period. As can be seen from figure A.1, birth cohorts just entering the labour market display different pat-

6 From a statistical point of view, the cohort differences are weakly significant for most of the age comparison groups.

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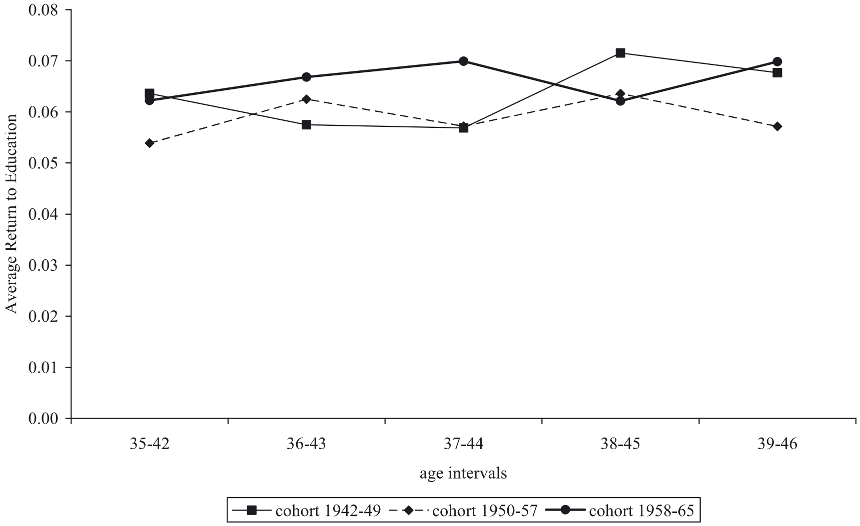
terns in returns to education than older workers who are observed in the same year. For example, the bad performance of the birth cohort 1958 – 65 during the years 1992 – 99 is probably not related to time effects because all older cohorts show high and stable returns during these years that are similar to earlier and later time periods. The additional insights from figure A.1 support our view that the observed pattern is related to cohort effects and not to general time effects.

Thus, our results weakly support that higher supply of labour market entrants increases wage competition, which seems to be also in line with findings from Boockmann / Steiner (2006) and Lauer / Steiner (2001), among others. As a new result, we find that the quantitative differences in the APE disappear when the "baby-boomer" cohort is compared at older ages (35 – 46 years old) with other cohorts at the same age when the young workers have got established in the labour market (see Figure 4). If so, cohort effects seem to exist only for the young when they enter the labour market and seem to vanish with increasing age. This might result from wage rigidity for incumbent workers and a higher degree of wage flexibility for entrants to the labour market (Gernandt / Pfeiffer, 2006; Pfeiffer, 2003).



Remark: Birth cohorts are observed at the same age interval but at different years. The cohort 1942 – 49 is observed during the period 1984 – 1988, the cohort 1950 – 57 is observed during the period 1992 – 1996 and the birth cohort 1958 – 65 during the period 2000 – 2004.

Figure 4: Average partial effect (CMI approach) by birth cohorts at same age intervals (35 – 46)

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6. Conclusions

In West Germany, graduates from the period of educational expansion in the 1960s and 70s entered the labour market during the period of observation from 1984 to 2006 and contributed to a skill upgrade of the labour force. While educational expansion might have devaluated the returns to education, other factors, such as skill-biased technological change, changing cohort sizes, or increased female labour supply might have altered the evolution of returns to education as well. In our study, we estimate the evolution of heterogeneous returns to education during the period from 1984 to 2006 to assess potential long-run consequences of educational expansion for wage workers. We employ two different estimation techniques that deal with the endogeneity of school choice and the heterogeneity of returns: Wooldridge's (2004) CMI approach and Garen's (1984) CF approach. The former method relies crucially on the conditional moment independence assumption, which requires sufficient observable control variables. The latter method employs distributional assumptions and needs an exclusion restriction such that it can control for selection on unobservables.

Our empirical findings indicate that both approaches produce estimates of average returns to education that decrease until the late 1990s and increase afterwards. During the period from 1984 to 2006, the estimated APE follows a roughly similar evolution pattern over time, although standard errors from Garen's approach are larger. According to the Wooldridge approach, returns to one additional year of education fell from 6.6 percent in 1984 to 4.9 percent in 1998 and increased to 6.6 percent in 2002 again. Interestingly, the CF approach shows a stronger and significant increase of the APE after 1998 compared to the CMI approach, although the increase should not be over-interpreted due to the higher degree of imprecision of CF estimates and its different identification strategy. Nevertheless, the increase in the CF estimates is a hint for increasing negative selection on unobservables in the sense that educational overachievers with unexpectedly high amounts of schooling have lower marginal returns to education. This supports the idea that educational expansion has resulted in institutions digging deeper into the distribution of student abilities and / or a devaluation of unobserved skills that have driven children into higher education in previous times due to changing employers' demand.

Furthermore, we estimated heterogeneous returns to education differentiated by gender and birth cohort in order to learn more about trends in central dimensions of observed heterogeneity in returns to education. We find that during the 1980s and early 1990s, returns to education have been higher for women than for men, but the gender gap in returns vanish after 1995, which is probably the result of increasing female education and labour supply, intensified by non-neutral technical progress. Furthermore, we find that the cohorts of "baby boomers" (workers born between 1958 and 1965) who, as a conse-

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quence, were exposed to stronger competition at labour market entry, had the lowest average return to education. However, the effect exists only at young ages and disappears when employees become older, which might result from wage rigidities for incumbent workers.

In this study, education is measured as years of schooling. Future research could be directed to specific characteristics of the German educational system, like early ability tracking and dual vocational educational qualifications. In addition, while our results point at the changing role of unobserved heterogeneity, future research could focus on the determinants of cognitive and non-cognitive ability formation in childhood (see Blomeyer / Coneus / Laucht / Pfeiffer, 2009), and its long run consequences for returns to education for wage workers.

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Appendix

Table A.1

Average partial effect, 1984 – 2006: OLS, CMI, and CF compared

	OLS coeff.	OLS (s.e.)	APE (CMI) coeff.	APE (CMI) (s.e.)	APE (CF) coeff.	APE (CF) (s.e.)	*N*
1984	0.074	(0.004)	0.066	(0.004)	0.075	(0.079)	1,545
1985	0.065	(0.004)	0.059	(0.004)	0.032	(0.131)	1,600
1986	0.069	(0.004)	0.061	(0.004)	0.077	(0.091)	1,682
1987	0.070	(0.004)	0.063	(0.004)	0.090	(0.048)	1,775
1988	0.069	(0.004)	0.065	(0.004)	0.081	(0.041)	1,798
1989	0.067	(0.003)	0.063	(0.004)	0.067	(0.038)	1,922
1990	0.063	(0.003)	0.059	(0.004)	0.048	(0.031)	2,007
1991	0.062	(0.003)	0.060	(0.004)	0.037	(0.030)	2,122
1992	0.064	(0.003)	0.057	(0.004)	0.049	(0.027)	2,107
1993	0.060	(0.003)	0.057	(0.004)	0.062	(0.026)	2,124
1994	0.057	(0.003)	0.051	(0.004)	0.055	(0.022)	2,082
1995	0.058	(0.003)	0.053	(0.004)	0.064	(0.024)	2,075
1996	0.055	(0.003)	0.049	(0.004)	0.054	(0.025)	2,057
1997	0.057	(0.003)	0.054	(0.003)	0.074	(0.025)	2,011
1998	0.054	(0.003)	0.049	(0.003)	0.053	(0.021)	2,145
1999	0.058	(0.003)	0.054	(0.003)	0.072	(0.023)	2,163
2000	0.065	(0.002)	0.059	(0.003)	0.108	(0.024)	3,965
2001	0.065	(0.002)	0.062	(0.003)	0.097	(0.022)	3,961
2002	0.068	(0.003)	0.066	(0.003)	0.106	(0.030)	3,668
2003	0.064	(0.003)	0.062	(0.003)	0.123	(0.028)	3,476
2004	0.064	(0.003)	0.062	(0.003)	0.115	(0.030)	3,366
2005	0.065	(0.003)	0.064	(0.003)	0.113	(0.032)	3,220
2006	0.064	(0.003)	0.063	(0.003)	0.133	(0.033)	3,477

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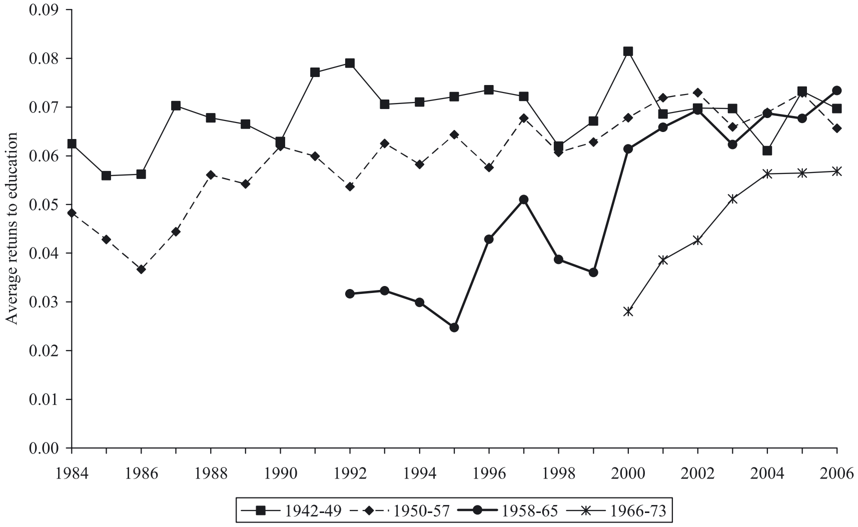

Remarks: (1) The different starting points of the lines are a result of different cohort ages at a given time point and due to low sample sizes at some points in time. The lines for the three youngest birth cohorts begin when the birth cohort is 27 – 32 years old. (2) Reading the graph horizontally, we can compare the returns to education for older (left) and younger (right) birth cohorts at a given age, but at different time periods. Figures 3 and 4 are produced based on this procedure.

Figure A.1: Average partial effect (CMI approach) by birth cohorts at same years

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