# Assessing Intergenerational Income Mobility Among German Workers 

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Studies on intergenerational income mobility intend to assess the relationship between parental and offspring's human capital. This empirical relationship is utilized as a measure of openness and equality of opportunity in an economy. Due to data restrictions human capital is proxied by current income measures in empirical studies. Such a proxy introduces attenuation bias due to transitory fluctuations as well as life - cycle bias. This paper introduces an improved sampling procedure and assesses a variety of father - son incomes from samples of West German workers drawn from the GSOEP 1984-2005. Our results indicate that the best point estimate of intergenerational income mobility among West German workers is $\frac{1}{3}$. Although the estimates still lack precision, our results nevertheless suggest that there is substantial intergenerational income mobility among West German workers.


#### Abstract

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[^0]
## 1. How Close Does the Apple Fall to the Tree in Germany

> ...the real explanation why the poor are where they are is that they made the mistake of being born to the wrong parents... Once that mistake is being made, they could have been paragons of will and morality, but most of them would never even have had a chance to get out (Harrington 1962)...

This statement, aimed at the United States, contrasts the meritocratic believes concerning the equality of opportunity that govern public debates. This is also true for the German education and welfare system. According to public rhetoric it is aimed to guarantee social permeability. No apparent legal measure ensure the economic success of an advantaged child or hinder the attempts of disadvantaged adolescents to escape poverty. Families receive a child benefit transfer, schooling for up to 13 years is free of charge and, if education is continued at a university, the cost of living is covered by federal student aid for low-income families.
Does this general concern translate into a society in which ones economic success is independent of the family born into? Are the results of the German type of welfare state effective in ensuring equality of opportunity and if so to what degree? Or does the position of the father in the distribution of income already determine the position of the son just as the citation suggests?

To empirically analyze the intergenerational relationships the following econometric model...

$$
\begin{equation*}
y_{1}^{i}=\alpha+\beta_{y} y_{0}^{i}+\epsilon_{1}^{i} \tag{1}
\end{equation*}
$$

is used as a starting point, see Corak (2004). A linear relationship between the human capital of family $i$ in generation 0 and 1 is assumed, allowing for shifts in mean economic status independent of parental status via the parameter $\alpha$. Deviations from predicted status due to market luck or other random elements in the intergenerational transmission of skills and personal traits, which constitute the human capital of family $i$, are summarized in the idiosyncratic error term $\epsilon_{1}^{i}$. Ideally, permanent income, as suggested by Friedman (1957), is chosen as the measure of economic status or human capital. In our study it is defined as the annuitized present discounted value of lifetime income generated in the German labor market. In the
empirical analysis we take into account the lag between income of fathers and sons and the general economic progress that occurred in West Germany in the observation period between 1984 and 2005. The terms (permanent) income and long - run (economic) status are used interchangeably, both describing the economic success of an individual.

All income variables will be measured in their natural logarithms. In that case $\beta_{y}$ in equation (1) is the intergenerational elasticity of income. This relationship is utilized as a measure of openness and equality of opportunity among West German workers of different generations. It measures the percentage change in offspring's human capital associated with a one percent change in parental human capital. The elasticity $\beta_{y}$ is determined by a multitude of factors forwarded from parents to children and influenced by investment in education as well as the economic order that affects a generation's economic possibilities. In Germany these factors include public education and social transfers, the labor market regulations and collective bargaining, among others.

The (expected) deviation of offspring's economic status from the mean is $\beta_{y}$ times the deviation of parental status from mean status. A value of zero for the intergenerational elasticity $\beta_{y}$ (child's and parental economic success are uncorrelated) corresponds to complete intergenerational mobility, while a value of unity (the child's economic success is completely determined by parental achievement) is associated with complete immobility. A positive value does indicate generational persistence of permanent income in which higher parental long - run status is associated with increased permanent income for the child; a negative number indicates generational reversal of economic success. $\left(1-\beta_{y}\right)$ provides a measure of the degree to which economic status regresses to the mean. If it takes value one $\left(\beta_{y}=0\right)$, a child from parents who attain below average long - run status can expect average status just as the offspring of high - status parents.

While many features of the human skill formation process are universal, there may however be unique features in German data. For instance, credit constraints may be less binding in the German system of social transfers to families and publicly provided education. Huggett, Ventura \& Yaron (2007) identify initial conditions at age 20 as the major source of lifetime inequality. Findings from neurobiology and child development research suggest that childhood is a critical period for skill formation, see Cunha, Heckman, Lochner \& Masterov (2006) for summarizing the evidence. Families therefore seem to be more important for the intergenerational
transmission of human capital than schools. Empirical research on the relationship between the human capital of fathers and sons may be helpful for exploring the facts. Due to data restrictions human capital is proxied by current income measures in empirical studies. Such a proxy introduces attenuation bias due to transitory fluctuations as well as a life - cycle bias, whose impact depends on an individual's stage in the life - cycle (Haider \& Solon 2006).

The contribution of our paper to the literature on intergenerational mobility is twofold. First, recent improvements in the understanding of the association between short - and long - run economic status allow for a new assessment of potential biases. Second, the income relationships between fathers and sons is estimated with a more suitable dataset and an improved sampling procedure. The data are drawn from the GSOEP 1984-2005.

Our results indicate that the best point estimate of intergenerational income mobility among West German workers is $\frac{1}{3}$. This indicates lower mobility in Germany compared to Couch \& Dunn (1997) and Wiegand (1997) and is in line with Vogel (2007) who compares intergenerational mobility between Germany and the United States. In an international perspective, the intergenerational income connection seems to be lower compared to the United States 0.4 (Solon 1992), and higher compared to Sweden 0.2 (Björklund \& Jäntti 1997). Even though the estimates still lack precision, our results suggest that there is substantial intergenerational income mobility among West German workers, which however is lower than previous work suggested.

The remainder of this paper is organized as follows: Section 2 introduces to the econometric methods dealing with the estimation of intergenerational income mobility based on incomplete data. Section 3 discusses our sampling procedure with the GSOEP. Following up, section 4 discusses the econometric findings. Section 5 concludes.

## 2. Econometric Problems and Findings from the Literature

The deduction of an individual's permanent income requires a life - long income history. It comes as no surprise that researchers usually lack direct measures of long - run status $\left(y_{0,1}^{i}\right)$ which are required for two generations in order to investigate intergenerational mobility. That is why researchers rely on proxies $\left(y_{0 s}^{i}, y_{1 t}^{i}\right)$ of permanent income for each generation $(0,1)$ observed at age $s$ and $t$. Sometimes only single - year measures of income ${ }^{1}$ are used. In this section the econometric problems associated with this approach and the consequences we draw regarding the estimation of intergenerational income mobility are pointed out.

### 2.1. Measurement Error Problems

Usually, a current measure of economic status $\left(y_{0 s}^{i}, y_{1 t}^{i}\right)$ is an imperfect proxy of long - run status. It is subject to measurement error due to transitory fluctuations and life - cycle variation in the association between current and lifetime income ${ }^{2}$

### 2.1.1. Transitory Fluctuations

A standard permanent - transitory decomposition of current income $\left(y_{0 s}^{i}, y_{1 t}^{i}\right)$ can be written as follows (Friedman 1957)...

$$
\begin{align*}
y_{1 t}^{i} & =y_{1}^{i}+v_{1 t}^{i}  \tag{2}\\
y_{0 s}^{i} & =y_{0}^{i}+v_{0 s}^{i} . \tag{3}
\end{align*}
$$

$y_{0,1}^{i}$ describes time - invariant permanent income, while $\left(v_{0 s}^{i}, v_{1 t}^{i}\right)$ describe time varying transitory fluctuations. The latter might arise from job mobility, business cycle effects, and variable compensation schemes. If current income deviates from permanent - status, using it to proxy for long - run status introduces attenuation bias in estimating equation (1). Assuming that $v_{1 t}^{i}$ and $v_{0 s}^{i}$ are uncorrelated with each other and permanent income $y_{0,1}^{i}$, the deviation of current from permanent income implies a downward inconsistency of the estimated slope coefficient $\hat{\beta}_{y}^{O L S}$ in

[^1]an OLS estimation by the factor $\theta_{s}$ (Solon 1992).
\[

$$
\begin{align*}
\operatorname{plim} \hat{\beta}_{y}^{O L S} & =\theta_{s} \beta_{y}<\beta_{y}  \tag{4}\\
\theta_{s} & =\left(\frac{\operatorname{Var}\left[y_{0}\right]}{\operatorname{Var}\left[y_{0}\right]+\operatorname{Var}\left[v_{0 s}\right]}\right) \tag{5}
\end{align*}
$$
\]

The attenuation factor $\theta_{s}$ captures how much signal $\operatorname{Var}\left[y_{0}\right]$ is provided by the measure $y_{0 s}$ relative to its total noise, $\operatorname{Var}\left[y_{0 s}\right]=\operatorname{Var}\left[y_{0}\right]+\operatorname{Var}\left[v_{0 s}\right]$.

Based on single - year snapshot, empirical findings by Corcoran, Laren, Gordon \& Solon (1991), Mazumder (2001), Card (1994) and Hyslop (2001) suggest an attenuation factor around $\theta_{s}=0.5$. This implies a considerable signal - to - noise ratio of observed parental income and an attenuation bias of $\left(1-\theta_{s}\right)=0.5$. Note also, that transitory fluctuations in offspring's income, $v_{1 t}^{i}$, do not bias the OLS estimation in equation (1) as long as they are uncorrelated with $v_{0 s}^{i}$. However, the higher their variance, the higher the standard errors of $\hat{\beta}_{y}^{O L S}$ will be.

## Averaging Parental Income

To reduce the magnitude of the inconsistency, Solon (1992) suggests to average parental status over $T$ years. The averaging reduces the variance of the noise relative to the signal. Transitory shocks are averaged away, as long as the process is stationary, see Mazumder (2005).

$$
\begin{equation*}
\theta_{s}=\left(\frac{\operatorname{Var}\left[y_{0}\right]}{\operatorname{Var}\left[y_{0}\right]+\frac{1}{T} \operatorname{Var}\left[v_{0 s}\right]}\right) \tag{6}
\end{equation*}
$$

As more years of data are used, the attenuation factor $\theta_{s}$ rises and the attenuation bias $\left(1-\theta_{s}\right)$ declines.

## Instrumenting Parental Income

It is natural to assume, that offspring's permanent income $y_{1}^{i}$ is not solely determined by parental long - run status $y_{0}^{i}$ as in equation (11), but an additional factor $I_{0}^{i}$ does play a role, see equation (7) (Solon 1992).

$$
\begin{equation*}
y_{1}^{i}=\beta_{1} y_{0}^{i}+\beta_{I} I_{0}^{i}+\omega_{1}^{i} \tag{7}
\end{equation*}
$$

Then the direct projection of offspring's permanent income on parental long - run status introduces an omitted - variable bias.

$$
\begin{align*}
y_{1}^{i} & =\beta_{y} y_{0}^{i}+\epsilon_{1}^{i}  \tag{8}\\
\beta_{y} & =\beta_{1}+\beta_{I}\left(\frac{\operatorname{Cov}\left[I_{0}, y_{0}\right]}{\operatorname{Var}\left[y_{0}\right]}\right) \tag{9}
\end{align*}
$$

An instrumental variable estimation, with an adequate instrument, $I_{0}^{i}$, will have the following probability limit (Solon 1992).

$$
\begin{align*}
\operatorname{plim} \hat{\beta}_{y}^{I V} & =\beta_{y}+\beta_{I}\left(\frac{1-\kappa^{2}}{\kappa}\right)\left(\frac{S d\left[I_{0}\right]}{S d\left[y_{0}\right]}\right)  \tag{10}\\
\kappa & =\frac{\operatorname{Cov}\left[I_{0}, y_{0}\right]}{S d\left[y_{0}\right] S d\left[I_{0}\right]} . \tag{11}
\end{align*}
$$

$\hat{\beta}_{y}^{I V}$ is an unbiased estimator for $\beta_{y}$ only if the instrument does not influence offspring's status $\left(\beta_{I}=0\right)$ or the instrument and parental status are perfectly correlated $|\kappa|=1$. The closer $|\kappa|$ is to one, the smaller the bias as there is less variation in income that is not captured by the instrument. Assuming a positive but imperfect correlation between the instrument and parental long - run status, the direction of the inconsistency is determined by $\beta_{I}$. If the instrument $I_{0}^{i}$ has a positive impact on offspring's status $\left(\beta_{I}>0\right)$, the estimator will be biased upward. If the opposite is true, the estimated coefficient is downward biased just as the OLS estimate.

In empirical research, parental years of education (for instance Solon (1992) and Dearden, Machin \& Reed (1997)) or indicators of occupational prestige (see Zimmerman (1992) and Wiegand (1997)) are used to instrument long - run parental status. Since years of education enhance labor market income, it may capture an important part of parental permanent income, although not necessarily to a $100 \%$ (see Card (1999) for a recent survey). Years of education may result from higher unobserved ability or less credit constraints at the time when people invest in their human capital (see Cunha \& Heckman (2007)). In this case an IV - estimate using years of education will be upward biased.

Estimating the intergenerational elasticity $\hat{\beta}_{y}$ using OLS and IV techniques suggests to bracket the coefficient. The OLS estimate is downward inconsistent due to error - in - variable bias, whereas the IV - Estimate is presumably upward biased. Accounting for the associated standard errors, the true $\beta_{y}$ may be located between the two estimates (Solon 1992).

$$
\hat{\beta}_{y}^{O L S}<\beta_{y}<\hat{\beta}_{y}^{I V}
$$

### 2.1.2. Life - Cycle Variations

Empirical research as well as theoretical reasoning suggest that wage workers differ with respect to their age - income profiles (see Mincer (1975) and Baker (1997) for general research and Vogel (2007) for an application to intergenerational mobility). This may occur due to age - specific heterogeneity in human capital investment or variations in the wage structure across jobs erected by firms for the purpose of effort regulation and incentive compatibility. For estimation purposes, the projection of current on permanent income is generalized to include a time - varying parameter to capture age - specific aspects in the association between current and permanent income over the lifecycle (Haider \& Solon 2006).

$$
\begin{align*}
y_{1 t}^{i} & =\lambda_{t} y_{1}^{i}+v_{1 t}^{i}  \tag{12}\\
y_{0 s}^{i} & =\lambda_{s} y_{0}^{i}+v_{0 s}^{i} \tag{13}
\end{align*}
$$

Averaging parental income $y_{0 s}^{i}$ across $T$ years, the interaction of both types of measurement error is considered. If parental and offspring's long - run status is proxied by short - run income, the resulting attenuation factor is calculated as...

$$
\begin{align*}
\operatorname{plim} \hat{\beta}_{y}^{O L S} & =\lambda_{t} \theta_{s} \beta_{y}  \tag{14}\\
\theta_{s} & =\frac{\lambda_{s} \operatorname{Var}\left[y_{0}\right]}{\lambda_{s}^{2} \operatorname{Var}\left[y_{0}\right]+\frac{1}{T} \operatorname{Var}\left[v_{0 s}\right]} \tag{15}
\end{align*}
$$

Assuming $\theta_{s}=1$, the probability limit of the estimated coefficient $\hat{\beta}_{y}^{O L S}$ is $\lambda_{t} \beta_{y}$ instead of $\beta_{y}$. In the case of $\lambda_{t}=1$ (as implicitly assumed in the discussion of


Figure 1: Estimated Correlation Between Current and Permanent Income
transitory fluctuations) this does no harm, but in general, the estimator will be inconsistent and the inconsistency varies as a function of age $t$ at which income is observed. Focusing on the impact of $\theta_{s}$ (setting $\lambda_{t}=1$ ), it is not obvious whether the combination of transitory fluctuations and life - cycle variation leads to an amplification bias instead of a attenuation bias. For $\lambda_{s}>1$ the estimation is downward biased, but for values smaller than one and minor transitory variance the opposite is true. $\theta_{s}$ can be interpreted as the slope coefficient in a backward regression of lifetime income on (averaged) income at age $s$. In fact, it is a summary measure of the attenuation bias resulting from transitory fluctuations as well as life - cycle variation. This indicates that results from studies of intergenerational mobility may be sensitive to the age - composition of the sample (see Jenkins (1987) and Grawe (2006)). In summary, measurement error in offspring's status is not innocuous for consistency as well as measurement error in parental long - runs status. Taken together, both induce either amplification or attenuation bias of the OLS estimation.
Using U.S. Social Security Administration income histories of members of the Health and Retirement Study sample, Haider \& Solon (2006) asses the magnitude of measurement error in offspring's and parental permanent income separately. Their dataset ranges from 1951 to 1991 and provides nearly career - long income histories


Figure 2: Estimated Reliability Ratio
for a broadly representative sample of the U.S. population. This allows to derive a more precise estimate of the (logarithmized) present value of lifetime income $\ln V^{i}$.

Starting with the impact of measurement error in offspring's (permanent) income level, the forward regression of $\ln V^{i}$ on $y_{t, s}^{i}$ leads to the estimated slope coefficient $\hat{\lambda}_{t, s}$ depicted in Figure 1. Starting at a value around $\hat{\lambda}_{t, s}=0.2$ it increases steadily. At age 32, the textbook assumption of $\lambda_{t, s}=1$ seems reasonable. Thenceforward, $\hat{\lambda}_{t, s}$ declines some in the late forties. Turning to the case of measurement error in parental permanent income, the estimated reliability ratio $\hat{\theta}_{s}$ is depicted in Figure 2. It is the result of a backward regression of $\ln V^{i}$ on a 5 - year average of $y_{t, s}^{i}$. A significant increase till age 30 is followed by a quite robust factor around 0.6 and 0.8 , but after the age of $50 \hat{\theta}_{s}$ declines and the bias rises.

### 2.2. Sample Homogeneity

In selected sub - populations with respect to location, socioeconomic status, or occupation, the sample variance in long - run economic status is possibly less than in the whole population. For example, a study by Sewell \& Hauser (1975) was based on a selective son - sample from Wisconsin, who graduated in 1957 and thus excluded high - school dropouts, leaving only rather successful sons in the sample.

Similarly, Behrman \& Taubman (1985) are confined to parental data on white male twins born between 1927 and 1929, who both served in the Army. Presumably, this father - sample is rather homogeneous. Both types of selectivity introduce a third source of inconsistency as Solon (1989) points out. To concentrate on the effect of sample homogeneity, long - run status is assumed to be measured correctly until indicated otherwise. Formally speaking, the parental/offspring - sample is more homogeneous in long - run status, if the variance in permanent income $\operatorname{Var}\left[y_{j=0,1}^{*}\right]$ is only a fraction $\tau$ of the population variance $\operatorname{Var}\left[y_{j=0,1}\right]$.

$$
\begin{equation*}
\operatorname{Var}\left[y_{j=0,1}^{*}\right]=\tau \operatorname{Var}\left[y_{j=0,1}\right] \tag{16}
\end{equation*}
$$

Under normality of parental economic status, selection on the dependent variable leads to a proportional change in the estimated intergenerational elasticity, where $R^{2}$ is the coefficient of determination of the population - based regression model (Goldberger 1981).

$$
\begin{align*}
\operatorname{plim} \hat{\beta}_{y *}^{O L S} & =\phi \beta_{y}<\beta_{y}  \tag{17}\\
\phi & =\frac{\tau}{1-R^{2}(1-\tau)} \tag{18}
\end{align*}
$$

If $\tau<1$ (implying $\phi<1$ ) the estimated intergenerational elasticity $\hat{\beta}_{y_{*}}^{O L S}$ is downward inconsistent even though long - run status is measured correctly.

A sample exhibiting homogeneity in parental income does not affect the consistency of intergenerational elasticity estimates. This is true as long as economic status is measured correctly. If this is not the case, the downward bias is worsened, as evident from equation (Solon 1992, Wiegand 1997).
$\begin{aligned}\left.(19)^{\left(\frac{\operatorname{Var}\left[y_{0 *}\right]}{\operatorname{Var}\left[y_{0 *}\right]+\operatorname{Var}\left[v_{0 s}\right]}\right)}\right) \beta_{y}=\operatorname{plim} \hat{\beta}_{y *}^{O L S}< & \\ \operatorname{plim} \hat{\beta}_{y}^{O L S} & =\left(\frac{\operatorname{Var}\left[y_{0}\right]}{\operatorname{Var}\left[y_{0}\right]+\operatorname{Var}\left[v_{0 s}\right]}\right) \beta_{y}\end{aligned}$
In applied empirical research, inclusion into an intergenerational dataset requires for father and son to both report positive labor market income in the periods of interest. Presumably, in such samples $\beta_{y}$ is underestimated. Unfortunately, there is
no research on the magnitude of this bias available.

### 2.3. An Econometric Model

The following econometric model is estimated.

$$
\begin{equation*}
y_{1 t}^{i}=\beta_{0}+\beta_{y} y_{0 s}^{i}+\beta_{1} A_{0 s}^{i}+\beta_{2} A_{0 s}^{2 i}+\beta_{3} A_{1 t}^{i}+\beta_{4} A_{1 t}^{2 i}+\omega_{1 t}^{i} \tag{20}
\end{equation*}
$$

Son's observed status in year $t$ is expressed as a regression function of father's observed status in year $s$, including age - controls for both (Solon 1992). Equation (20) is derived by the incorporation of age - income profiles into equations (2) and (3) and substitution into the basic equation (1)...

$$
\begin{align*}
y_{1 t}^{i} & =y_{1}^{i}+\alpha_{1}+\gamma_{1} A_{1 t}^{i}+\delta_{1} A_{1 t}^{2 i}+v_{1 t}^{i}  \tag{21}\\
y_{0 s}^{i} & =y_{0}^{i}+\alpha_{0}+\gamma_{0} A_{0 s}^{i}+\delta_{0} A_{0 s}^{2 i}+v_{0 s}^{i} \tag{22}
\end{align*}
$$

,where an individual's current income is determined by the level of permanent income , $\left(y_{1}^{i}, y_{0}^{i}\right)$, the period in life $\left[\left(A_{t}^{i} ; A_{t}^{2 i}\right),\left(A_{s}^{i} ; A_{s}^{2 i}\right)\right]$, a general level of economic well - being in the corresponding generation $\left(\alpha_{1}, \alpha_{0}\right)$, and an idiosyncratic error term $\left(v_{1 t}^{i}, v_{0 s}^{i}\right)$.
For investigating the intergenerational income mobility, we would like to address the problems of transitory fluctuations and life - cycle variation, when using short run proxies for long - run economic status. The averaging over several years reduces the transitory variance and we employ a 5 - year average of parental income for the baseline estimation. According to Mazumder (2005) the attenuation factor $\theta_{s}$ rises to $\theta_{s}=0.7$ (from $\theta_{s}=0.5$ ) when relying on a 5 - year average of income. The attenuation bias is reduced to $\left[\left(1-\theta_{s}\right)=0.3\right]$. Solon (1992) and Wiegand (1997) estimated an intergenerational elasticity of father's and son's income from wage work based on 5 - year averages of 0.4 for the United States and 0.2 for Germany. Given the attenuation factor mentioned above the "true" elasticities would come closer to 0.6 for the United States and 0.3 for Germany. Furthermore, empirical research as well as theoretical reasoning suggests that the variance of the transitory component $\operatorname{Var}\left[v_{0 s}\right]$ exhibits a U - shaped pattern over the lifecycle and flattens out a mid age as depicted in Figure3 (see Baker \& Solon (2003) and Mazumder (2001) among


Figure 3: Transitory Variance over the Life - Cycle
others). Before the age of thirty, job mobility is higher resulting in a higher transitory income component. Productivity shocks, e.g. through technological progress, could lead to the observed increase in the variance at older ages. Therefore, we restrict the sample to individuals within that range. Those measures in place, an OLS estimate is still suspicious of attenuation bias. This necessitates an supplementary IV - Estimation to bracket the intergenerational coefficient.

Equation (20) is estimated with OLS and IV techniques which, presumably, leads to an upper and lower bound for the intergenerational income elasticity. Furthermore, samples with varying requirements concerning the number of years averaged and with regard to the age structure are utilized to assess the remaining bias in the estimation.

## 3. Data

The empirical part builds on samples form the German Socio - Economic Panel (GSOEP)from 1984 to 2005 (see Haisken DeNew \& Frick (2005)). One innovative feature of our study on intergenerational income mobility in Germany is the sampling procedure. From the GSOEP, we select pairs of fathers and sons in a way that their wages are observed as close in their life - cycle as possible and the bias due to transitory fluctuations and life - cycle variations is minimized. The measure of long - run economic status is rea ${ }^{3}$ monthly income before tax and social security deductions as reported in each cross - section of the GSOEP ${ }^{4}$.
Initially, separate samples, one for potential sons and one for potential fathers, are selected. Perspective sons, for which the father is at any point a member of the GSOEP, are recognized by a special code identifying the father. Hence, father and son observations can be sorted in two separate groups.

To reduce the bias introduced by possible measurement error in long - run economic status, several restrictions are imposed on the sample, see Table (11). Groups suspected of high measurement error are dropped. This is the case for the self - employed, who have more volatile income (Baker \& Solon (2003), Albarrán, Carrasco \& Martínez-Granado (2007), and Pfeiffer (1994)). Only full - time employed are retained in the sample, that is individuals reporting to work more than 35 hours the last week. Workers from East Germany are excluded as well since the possibility for mobility increased dramatically after the fall of the Berlin Wall and drastic wage growth $h^{5}$ may have changed the reliability of current income to reflect permanent status.
Since the association between monthly and lifetime income is low for workers below the age of 30 , we select only workers above that age. For younger workers job mobility is high and at the same time labor income is lower and more volatile because of lower tenure (see Björklund (1993) and Haider \& Solon (2006)). Workers aged above 50 years are excluded as well. Labor market income and hours worked may become more volatile again which might increase the bias of the estimated intergenerational elasticity (see Grawe (2006)). However, this line of reasoning may differ

[^2]between countries, for instance as a result of different industrial structures or different degrees of employment protection laws. Therefore, we will perform robustness tests relaxing the imposed age restrictions. Migrants are dropped from the analysis for two reasons. First, migration might change the long - run relationship between labor market income of father and son, and, second, the transitory component is presumably high (see also Borjas (2006) and Friedberg (2000)).

Table 1: Sample Overview

| Groups Excluded from Sample | Measures of Economic Status | Age - Restrictions |
| :---: | :---: | :---: |
| younger brothers | $\underline{\text { Son }}$ |  |
| East Germans | monthly income $(1984-2005)$ | between $30-50$ |
| migrants | $\underline{\text { Father }}$ |  |
| self - employed | monthly income (1984-2005) |  |
| part - time employed | years of education |  |



Figure 4: Sampling Strategy I
Observations from the group of sons who do not offer the needed information are dropped. For the group of fathers, moving 5 - year averages of the required information are calculated. Thus, if for a given observation income is not observable in each of the four following years, it is dropped. The remaining observations include information on single - year income and the age that same year for sons (obs ${ }_{j}$ and $o b s_{i}$ ), while for fathers ( $o b s_{k}$ and $o b s_{l}$ ) averaged income and the associated age as well as years of education are contained (see Figure 4).
Finally, father - son observations (of family $h$ ) satisfying the sampling rule are matched in all possible combinations (see Figure 5).


Figure 5: Sampling Strategy II

This procedure leads to numerous matched observations for each pair. To identify a unique pair, expected to lead to the most reliable estimate of the intergenerational elasticity, a decision rule is implemented. For each observation the age - difference between father and son is calculated and only the one with the smallest absolute value retained. This ensures that father and son are observed at similar stages in their life - cycle. If still more than one observation for a particular father - son pair fulfills the requirement, the one associated with the lowest father age is used. Furthermore, only the oldest son is retained in the case two siblings comply with the sampling rule. To assure comparability of real income observed in different years, they are adjusted by the real GDP - Growth Rate. Other measures like the growth rate of average real gross monthly income in Germany's industry sector $\sqrt{6}$ do not change the result. The final sample exhibits the descriptive statistics depicted in Table 2 .

[^3]Table 2: Sample Statistics

| Statistic | Fathers | Sons |
| :--- | ---: | ---: |
| Gross Income in $€^{1}$ | $2,272.03$ | $1,937.35$ |
| Sd. of Gross Income | 665.47 | 627.85 |
| Year of Observation | $1,984.64$ | $2,003.30$ |
| Age in Years | 44.34 | 35.24 |
| Age - Difference in Years | 9.12 |  |
| Number of Observations | 173 |  |
| ${ }^{1}$ reported 5/1 - year average of adjusted real |  |  |
| gross monthly income |  |  |

While most information on father's economic status is obtained within the very first GSOEP - Wave, the collection of offsprings' information is not confined to the most recent wave. The age - composition in our sample differs from other studies. An average age of 44 years for fathers is slightly lower than the one reported by Wiegand (1997) with 46 years, while Couch \& Dunn's (1997) fathers are 51 years old. Solon's (1992) fathers are reported to be 42 years of age on average, nearly identical to an average father in Björklund \& Jäntti's (1997) sample (43 years). But in contrast, sons are almost 35 years old which is an increase of 4 years compared to Wiegand (1997) and even 13 years compared to Couch \& Dunn (1997). Solon (1992) reports an average age of 29 for sons, while Björklund \& Jäntti (1997) rely on sons at the age of 34 on average. The age - difference between father and son in this paper amounts to 9,65 years. This comes at a price. The resulting sample contains only 173 father - son pairs compared to Wiegand's (1997) 130 and Vogel's (2007) 300. The rather small increase is a concession to the strict sampling rule imposed. Selection could rise from the blind eye on individuals not reporting positive income 5 years in a row. This ignores all short - term unemployed, potentially leaving only successful labor market participants to comply with the sampling rule. Individuals without any offspring reporting in the GSOEP are ignored as well.

Meeting these concerns, the final sample is compared to all individuals living up to the sample requirements except for the need to report positive income 5 years in a row and being matched with their offspring. Essentially, it is selection on unobserved permanent status that matters, but the comparison of current (averaged) income might at least give some indication on the issue. The father - sample is contrasted in 1984, while the son - sample is compared in 2005. Income in the father - sample
is nearly identical to the one reported by all individuals in 1984 ( $2,340.27$ €). The standard deviation of income is rather high in the comparison group (776.83€) in 1984. Using 5 - year averages of income in the father - sample, therefore reduces transitory fluctuations (see equation 233 )

$$
\begin{equation*}
\operatorname{Var}\left[y_{0 s}\right]=\lambda^{2} \operatorname{Var}\left[y_{0}\right]+\frac{1}{T} \operatorname{Var}\left[v_{0 s}\right] \tag{23}
\end{equation*}
$$

Comparing the son - sample, income is slightly higher than in the comparison group ( $1,872.04 €$ ) and the same is true for its standard deviation ( 542.23 €). Sons in the sample report less income than their matched fathers which is mainly explained by their early stage in the life - cycle. For comparison, other samples with less restrictive selection rules are utilized in the econometric part below.

## 4. Econometric Findings

### 4.1. Basic Results

Table 3: Basic Results

|  | 5 - Year Avg. Income | Single - Year Income |
| :---: | :---: | :---: |
| OLS - Estimate <br> Intergenerational Elasticity 95\% Confidence Interval Standard Error | $\begin{gathered} 0.262 \\ (0.09-0.43)^{1} \\ 0.087 \end{gathered}$ | $\begin{gathered} 0.193 \\ (0.07-0.32) \\ 0.063 \end{gathered}$ |
| Observations | 173 | 241 |
| IV - Estimate ${ }^{2}$ <br> Intergenerational Elasticity 95\% Confidence Interval Standard Error | $\begin{gathered} 0.353 \\ (0.06-0.65) \\ 0.150 \end{gathered}$ |  |
| Observations | 173 |  |

${ }^{1} 95 \%$ confidence intervals in parenthesis
${ }^{2}$ using years of education

Our basic results, depicted in Table 3, differ to the estimates by Wiegand (1997). The OLS estimate based on a 5 - year average of income $\hat{\beta}_{y}^{O L S}=0.262$ is higher, whereas the one - year snapshot is lower and similar to Wiegand (1997). Compared to Vogel (2007), the result based on a 5 - year average of income is similar.

Years of education $7^{7}$ and the Wegener - Index, a standard indicator of occupational prestige, are used as an instrument to bracket the intergenerational elasticity. According to the IV - estimate the intergenerational elasticity is higher, 0.35 , for both instruments. Following Solon (1992), the intergenerational elasticity for West German Workers should lie between 0.26 and 0.35 .

[^4]$$
\hat{\beta}_{y}^{O L S}=0.26<\beta_{y}<0.35=\hat{\beta}_{y}^{I V}
$$

The lower bound, determined by the downward inconsistent OLS estimate, amounts to $\hat{\beta}_{y}^{O L S}=0.26$, while the upward biased IV - estimation indicates $\hat{\beta}_{y}^{I V}=0.35$ as the upper bound. However, both values are estimated with considerable standard errors. The $95 \%$ confidence interval of the IV - estimate $\left[0.06 \leq \hat{\beta}_{y}^{I V} \leq 0.65\right]$ even includes the OLS - estimate. Although the different estimation method contains some useful information the degree of precision seems to be rather low.

### 4.2. Investigating the Bias from Transitory Fluctuations

Table 4: Summary Results ${ }^{1}$ : Balanced Panel

| Father Measure $^{2}$ | $5-$ Year | $4-$ Year | $3-$ Year | $2-$ Year | $1-$ Year |
| :--- | :---: | :---: | :---: | :---: | :---: |
| Intergenerational Elasticity | $0.2615^{* * *}$ | $0.2685^{* * *}$ | $0.2621^{* * *}$ | $0.2441^{* * *}$ | $0.1997^{* *}$ |
| $95 \%$ Confidence Interval | $(0.09-0.43)$ | $(0.10-0.44)$ | $(0.09-0.43)$ | $(0.08-0.41)$ | $(0.04-0.35)$ |
| Standard Error | 0.0871 | 0.0867 | 0.0855 | 0.0842 | 0.0789 |
| Observations | 173 | 173 | 173 | 173 | 173 |

Basic Specification
Source: own calculations
Level of Significance: *** $1 \%{ }^{* *} 5 \% * 10 \%$
${ }^{1}$ see Table 9 in the Appendix for the detailed results
${ }^{2}$ average of father's logarithmized adjusted real gross monthly income

Table 4 and 5 report the general pattern that $\hat{\beta}_{y}^{O L S}$ increases with the number of years averaged as the attenuation bias declines. This is in line with equation (6).

For inclusion in the balanced panel, parental income needs to be observed for 5 years in a row even though only lower averages are used for the supplementary estimations. The changing estimate is due to the reduced number of years averaged and not to a change in the sample composition. For this reason, the number of observations remains constant. The unbalanced panel, however, includes all pairs with the necessary number of successive income observations for the father that is needed for the respective estimation.

A comparison of the OLS results in the balanced and unbalanced panel reveals that the difference between a 5 - and 4 - year average of father's income is negligible.

However, it makes a difference in our sample whether the estimate is based on a $1 / 2$ - year average compared to an $4 / 5$ - year average. Averaging only a small number of years amplifies the attenuation bias due to a high volatility of the income measure utilized. This result seems to be in line with the literature as reported in section 2 .

Table 5: Summary Results ${ }^{1}$ : Unbalanced Panel

| Father Measure $^{2}$ | $5-$ Year | $4-$ Year | $3-$ Year | $2-$ Year | $1-$ Year |
| :--- | :--- | :--- | :--- | :--- | :--- |
| Intergenerational Elasticity | $0.2615^{* * *}$ | $0.2730^{* * *}$ | $0.2382^{* * *}$ | $0.1956^{* * *}$ | $0.1931^{* * *}$ |
| $95 \%$ Confidence Interval | $(0.09-0.43)$ | $(0.11-0.43)$ | $(0.08-0.39)$ | $(0.06-0.33)$ | $(0.07-0.31)$ |
| Standard Error | 0.0871 | 0.0815 | 0.0792 | 0.0695 | 0.0626 |
| Observations | 173 | 183 | 210 | 220 | 241 |

Basic Specification
Source: own calculations
Level of Significance: ${ }^{* * *} 1 \%{ }^{* *} 5 \%{ }^{*} 10 \%$
${ }^{1}$ see Table 10 in the Appendix for the detailed results
${ }^{2}$ average of father's logarithmized adjusted real gross monthly income

The rather early decrease of the estimated coefficient in the unbalanced panel might be attributable to the construction of the panel. When lowering the number of years averaged, the added individuals do not report income in the following year likely due to un - or part - time employment. This implies that individuals with rather large transitory fluctuations are consecutively added to the panel.

### 4.3. Investigating the Bias from Life - Cycle Variation

Following the empirical evidence in Haider \& Solon (2006), inclusion of older fathers decreases the reliability ratio $\theta_{s}$ due to an increase in the transitory variance $\operatorname{Var}\left[v_{0 s}\right]$ and a rather unchanged association between current and permanent income $\lambda_{s}$, see equation 14 in section 2. Raising the upper age - limit from 50 to 55 years results in a rather sharp increase in sample size and a slight decrease in estimated intergenerational mobility. However, Table 6 reveals an increase in the estimate when continuing to soften the age - restriction, which seems to be in line with Vogel (2007). This could point at sample selection with only pairs added that exhibit a particular strong persistence of income. A comparison of the descriptive statistics (years of education, monthly income) does not offer any evidence on the type of selection. The added individuals do not differ distinctively with respect to these
characteristics. Another explanation is that income of workers above 50 years and older in Germany who report income for five consecutive years are less volatile compared to the U.S.. If this interpretation is valid, $\beta_{y}$ may lie near $\frac{1}{3}$ according to the estimates in Table 4 and 5 .

Table 6: Summary Results ${ }^{1}$ : Relaxing Age - Restrictions for Fathers

| Father's Maximal Age | 50 | 55 | 60 | 65 |
| :--- | :---: | :---: | :---: | :---: |
| Intergenerational Elasticity $^{2}$ | $0.2615^{* * *}$ | $0.2315^{* * *}$ | $0.3331^{* * *}$ | $0.3387^{* * *}$ |
| $95 \%$ Confidence Interval | $(0.09-0.43)$ | $(0.07-0.39)$ | $(0.19-0.47)$ | $(0.20-0.47)$ |
| Standard Error | 0.0871 | 0.0804 | 0.0707 | 0.0696 |
| Observations | 173 | 233 | 274 | 278 |

Basic Specification
Source: own calculations
Level of Significance: ${ }^{* * *} 1 \%{ }^{* *} 5 \%$ * $10 \%$
${ }^{1}$ see Table 12 in the Appendix for the detailed results
${ }^{2} 5$ - year average of father's logarithmized adjusted real gross monthly income; son at least 30 years of age

Table 7 documents a significant rise in the number of observations and a sharp decline in the estimated intergenerational elasticity when the age - requirement is consecutively lowered to 20 years. This seems to be in line with Haider \& Solon (2006). The parameter $\lambda_{t}$ (see equation $(14)$ in section 2) is lowered as younger and younger individuals are added to the sample and the life - cycle bias rises.

Table 7: Summary Results ${ }^{1}$ : Relaxing Age - Restrictions for Sons

| Son's Minimum Age | 30 | 25 | 20 |
| :--- | :--- | :--- | :--- |
| Intergenerational Elasticity $^{2}$ | $0.2615^{* * *}$ | $0.1955^{* * *}$ | $0.1885^{* * *}$ |
| $95 \%$ Confidence Interval | $(0.09-0.43)$ | $(0.06-0.33)$ | $(0.07-0.30)$ |
| Standard Error | 0.0871 | 0.0688 | 0.0588 |
| Observations | 173 | 273 | 370 |

## Basic Specification

Source: own calculations
Level of Significance: *** $1 \% * * 5 \% * 10 \%$
${ }^{1}$ see Table 13 in the Appendix for the detailed results
${ }^{2} 5$ - year average of father's logarithmized adjusted real gross monthly income; father at most 50 years of age

The analysis above gave the impression that the age - composition of either sample is changed without affecting the other. Obviously, this is not true since father son pairs are added. However, negligible changes in the age - composition of the unchanged (with respect to the age - restrictions imposed) sample support this approach.

### 4.4. Further Sensitivity Checks

## Including Younger Siblings

The inclusion of younger siblings raises the sample size from 173 to 216 when relying on a 5 - average of income. The coefficient is slightly reduced to $\hat{\beta}_{y}^{O L S}=0.24$. Siblings share the same family and community background which makes similar long - run economic status more likely and increases homogeneity within the sample. This depresses the estimated coefficient, see Table 14 in the Appendix for the detailed results.

## Adjustment of Monthly Income

To ensure robustness with respect to the measure of comparability (GDP - Growth in the baseline estimation), income is deflated by the growth rate of average real gross monthly income in Germany's industry sector (as reported by the German Federal Statistical Office). The estimated intergenerational elasticity is not affected, see Table 15 in the Appendix for the detailed results.

## Instrumenting Parental Status

To provide complete comparison with Wiegand (1997), the IV - Estimation is repeated instrumenting parental status using the Wegener - Index, a standard index for occupational prestige. The baseline estimate $\left(\hat{\beta}_{y}^{I V}=0.38\right)$ is marginally higher. The finding that both instruments lead to rather identical results is robust to changes in the sampling rule, see Table 16 in the Appendix for the detailed results.


Figure 6: Comparing the Results for Bracketed $\beta_{y}$

## 5. Discussion and Concluding Remarks

Figure 6 illustrates the bracketed level of intergenerational mobility for the U.S., Germany, and the United Kingdom.

From our estimates and from related findings of the literature, the intergenerational income mobility between father - son pairs in the period from 1984 to 2005 seems to be higher compared to the United States and the United Kingdom but lower compared to Sweden. According to our interpretation, the point estimate of the elasticity in Germany is around $\beta_{y}^{G E R}=\frac{1}{3}$, while it is around $\beta_{y}^{U S}=0.4$ (or higher) in the United States and $\beta_{y}^{S}=0.2$ in Sweden (see Solon (1992), Mazumder (2005) for the U.S. and Björklund \& Jäntti (1997) for Sweden as examples). However, despite the considerable number of studies, the point estimates are not really precise. Common to all studies are the rather large confidence intervals, which currently forbid any comparative statements on the level of intergenerational mobility. The findings from the studies by Couch \& Dunn (1997) and Wiegand (1997) for Germany seem to have a larger attenuation bias. Ours and Vogel's (2007) work suggest that mobility in Germany is lower. Be that as it may, we try to figure out some consequences of the value of $\beta_{y}^{G E R}=\frac{1}{3}$ for Germany.

The intergenerational elasticity $\beta_{y}$ translates intragenerational inequality in parental
long - run status into the economic advantage, which a child from parents with higher economic status can hope for in the next generation compared to one from lower (permanent) income parents. Assuming inequality in permanent income to be reflected in cross - section inequality in annual gross income, Table 8 depicts the expected (permanent) income advantage in percentage terms of a child with parents in the top income decile compared to offspring born to parents in the bottom decile (Corak 2004).

$$
\begin{equation*}
\frac{y_{1}^{90 t h}}{y_{1}^{10 t h}}=\left(\frac{y_{0}^{90 t h}}{y_{0}^{10 t h}}\right)^{\beta} \tag{24}
\end{equation*}
$$

Table 8: Inequality and the Expected Permanent Income Advantage

| Intergenerational Elasticity |  |  |  |  |
| :--- | :---: | :---: | :---: | :---: |
| $90 / 10$ - Ratio | 0.2 | $\frac{1}{3}$ | 0.4 | 0.5 |
| 2.0 | $15 \%$ | $25 \%$ | $32 \%$ | $41 \%$ |
| 2.5 |  | $20 \%$ | $35 \%$ | $44 \%$ |
| 380 |  | $25 \%$ | $44 \%$ | $55 \%$ |
| 3.5 |  | $28 \%$ | $51 \%$ | $65 \%$ |
| 4.0 |  | $32 \%$ | $59 \%$ | $74 \%$ |

For Germany, Gernandt \& Pfeiffer (2006) calculate a $90 / 10$ - percentile income ratio of 2.5 for a sample of prime age dependent male workers, which is rather close to our one. Taking the middling value for an intergenerational elasticity in Germany of $\beta_{y}=\frac{1}{3}$, the (expected) income advantage amounts to $35 \%$. If $\beta_{y}$ would be 0.5 , the (expected) income advantage increases to $58 \%$.

For countries with a lower degree of intragenerational inequality, like Sweden, the advantage will be lower. As opposed to countries with higher levels of inequality, where the opposite is true. Summarizing our findings, intergenerational mobility among West German workers is lower than previously suggested. However, a value of $\beta_{y}=\frac{1}{3}$ still indicates that there is substantial income mobility, which presumably is one result of publicly funded education and the welfare system.
A. Estimation Appendix
Table 9: Detailed Results: Balanced Panel

|  | 5 - Year | verage | 4 - Year A | verage | 3 - Year A | verage | 2 - Year | verage | 1 - Year | verage |
| :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: |
|  | Coeff. ${ }^{1}$ | SE. ${ }^{1}$ | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. |
| Father |  |  |  |  |  |  |  |  |  |  |
| Measure ${ }^{2}$ | $0.2615{ }^{* * *}$ | 0.0871 | $0.2685^{* * *}$ | 0.0867 | $0.2621^{* * *}$ | 0.0855 | $0.2441^{* * *}$ | 0.0842 | 0.1997** | 0.0789 |
| Age | -0.0740 | 0.0708 | -0.0749 | 0.0706 | -0.0717 | 0.0705 | -0.0703 | 0.0707 | -0.0639 | 0.0710 |
| Age (squared) | 0.0011 | 0.0009 | 0.0011 | 0.0009 | 0.0011 | 0.0009 | 0.0011 | 0.0009 | 0.0010 | 0.0009 |
| Son |  |  |  |  |  |  |  |  |  |  |
| Age | 0.2681** | 0.1163 | 0.2668** | 0.1161 | 0.2658** | 0.1162 | $0.2573 * *$ | 0.1165 | $0.2634^{* *}$ | 0.1172 |
| Age (squared) | $-0.0036^{* *}$ | 0.0016 | -0.0036 ** | 0.0016 | -0.0036 ** | 0.0016 | $-0.0035^{* *}$ | 0.0016 | -0.0036 ** | 0.0016 |
| F - Test | $5.3096{ }^{* * *}$ |  | $5.4356^{* * *}$ |  | $5.3967^{* * *}$ |  | $5.1756^{* * *}$ |  | $4.7363^{* * *}$ |  |
| Adjusted $R^{2}$ | 0.1113 |  | 0.1142 |  | 0.1133 |  | 0.1082 |  | 0.0980 |  |
| Observations | 173 |  | 173 |  | 173 |  | 173 |  | 173 |  |

Basic Specification
Source: own calculations
Level of Significance: ${ }^{* * *} 1 \%{ }^{* * 5} \%{ }^{*} 10 \%$
${ }^{1}$ Coeff. $=$ Coefficient ; SE. $=$ Standard Error
${ }^{2}$ average of father's logarithmized adjusted real gross monthly income
Table 10: Detailed Results: Unbalanced Panel

|  | 5 - Year | verage | 4 - Year A | verage | 3 - Year A | verage | 2 - Year | verage | 1 - Year | verage |
| :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: |
|  | Coeff. ${ }^{1}$ | SE. ${ }^{1}$ | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. |
| Father |  |  |  |  |  |  |  |  |  |  |
| Measure ${ }^{2}$ | $0.2615^{* * *}$ | 0.0871 | $0.2730^{* * *}$ | 0.0815 | $0.2382^{* * *}$ | 0.0792 | $0.1956^{* * *}$ | 0.0695 | $0.1931^{* * *}$ | 0.0626 |
| Age | -0.0740 | 0.0708 | -0.0677 | 0.0729 | -0.0393 | 0.0711 | -0.0352 | 0.0728 | -0.0881 | 0.0685 |
| Age (squared) | 0.0011 | 0.0009 | 0.0010 | 0.0009 | 0.0006 | 0.0009 | 0.0006 | 0.0009 | 0.0012 | 0.0008 |
| Son |  |  |  |  |  |  |  |  |  |  |
| Age | 0.2681** | 0.1163 | 0.2728** | 0.1135 | 0.2836** | 0.1129 | $0.2946 * *$ | 0.1111 | 0.2667** | 0.1060 |
| Age (squared) | $-0.0036^{* *}$ | 0.0016 | $-0.0037^{* *}$ | 0.0016 | -0.0037** | 0.0016 | -0.0039** | 0.0015 | -0.0035** | 0.0015 |
| F - Test | $5.3096^{* * *}$ |  | 6.2493 *** |  | $5.6532^{* * *}$ |  | $5.8107^{* * *}$ |  | $6.2945^{* * *}$ |  |
| Adjusted $R^{2}$ | 0.1113 |  | 0.1260 |  | 0.1002 |  | 0.0990 |  | 0.0993 |  |
| Observations | 173 |  | 183 |  | 210 |  | 220 |  | 241 |  |

Basic Specification
Source: own calculations
Level of Significance: *** $1 \%{ }^{* * 5} \%{ }^{*} 10 \%$
${ }^{1}$ Coeff. $=$ Coefficient ; SE. $=$ Standard Error
${ }^{2}$ average of father's logarithmized adjusted real gross monthly income
Table 11: Detailed Results: IV - Estimation Using Years of Education

|  | Lower Age Limit -5 |  | Basic Specification |  | Upper Age Limit +5 |  |
| :---: | :---: | :---: | :---: | :---: | :---: | :---: |
|  | Coeff. ${ }^{1}$ | SE. ${ }^{1}$ | Coeff. | SE. | Coeff. | SE. |
| Father <br> Measure ${ }^{2}$ | 0.2796** | 0.1173 | 0.3531** | 0.1502 | 0.3049** | 0.1335 |
| Son <br> Age <br> Age (squared) | $\begin{gathered} 0.2653^{* * *} \\ -0.0035^{* * *} \end{gathered}$ | $\begin{aligned} & 0.0556 \\ & 0.0008 \end{aligned}$ | $\begin{gathered} 0.2521^{* *} \\ -0.0033^{* *} \end{gathered}$ | $\begin{aligned} & 0.1162 \\ & 0.0016 \end{aligned}$ | $\begin{gathered} 0.2371^{* * *} \\ -0.0031^{* *} \end{gathered}$ | $\begin{aligned} & 0.0898 \\ & 0.0012 \end{aligned}$ |
| F - Test <br> Adjusted $R^{2}$ <br> Observations | 32.671 0.25 273 |  | 4.435 0.07 17 |  | 5.517 0.0 233 |  |

Basic Specification
Source: own calculations
Level of Significance: ${ }^{* * *} 1 \%{ }^{* * 5} \% * 10 \%$
${ }^{1}$ Coeff. $=$ Coefficient ; SE. $=$ Standard Error
${ }^{2}$ years of education
Table 12: Detailed Results: Relaxing Age - Restrictions for Fathers

|  | Max. A | e 50 | Max. A | e 55 | Max. A | e 60 | Max. A | e 65 |
| :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: |
|  | Coeff. ${ }^{1}$ | SE. ${ }^{1}$ | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. |
| Father |  |  |  |  |  |  |  |  |
| Measure ${ }^{2}$ | $0.2615^{* * *}$ | 0.0871 | $0.2315^{* * *}$ | 0.0804 | $0.3331^{* * *}$ | 0.0707 | $0.3387^{* * *}$ | 0.0696 |
| Age | -0.0740 | 0.0708 | 0.0884* | 0.0515 | 0.0609 | 0.0378 | 0.0629* | 0.0340 |
| Age (squared) | 0.0011 | 0.0009 | -0.0010 | 0.0006 | -0.0007 | 0.0004 | -0.0007 | 0.0004 |
| $\underline{S o n}^{3}$ |  |  |  |  |  |  |  |  |
| Age | 0.2681** | 0.1163 | 0.2084** | 0.0915 | 0.1241* | 0.0667 | $0.1144^{*}$ | 0.0648 |
| Age (squared) | $-0.0036^{* *}$ | 0.0016 | $-0.0027^{* *}$ | 0.0012 | -0.0016* | 0.0009 | -0.0014 | 0.0009 |
| F - Test | $5.3096{ }^{* * *}$ |  | $5.0960^{* * *}$ |  | $7.0635^{* * *}$ |  | $7.3204^{* * *}$ |  |
| Adjusted $R^{2}$ | 0.1113 |  | 0.0811 |  | 0.1000 |  | 0.1024 |  |
| Observations | 173 |  | 233 |  | 274 |  | 278 |  |

[^5]Table 13: Detailed Results: Relaxing Age - Restrictions for Sons

|  | Min. A | e 30 | Min. A | e 25 | Min. | e 20 |
| :---: | :---: | :---: | :---: | :---: | :---: | :---: |
|  | Coeff. ${ }^{1}$ | SE. ${ }^{1}$ | Coeff. | SE. | Coeff. | SE. |
| Father ${ }^{3}$ |  |  |  |  |  |  |
| Measure ${ }^{2}$ | $0.2615^{* * *}$ | 0.0871 | $0.1955^{* * *}$ | 0.0688 | $0.1885^{* * *}$ | 0.0588 |
| Age | -0.0740 | 0.0708 | -0.0696 | 0.0497 | -0.0569 | 0.0411 |
| Age (squared) | 0.0011 | 0.0009 | 0.0010 | 0.0006 | 0.0009 | 0.0005 |
| Son |  |  |  |  |  |  |
| Age | 0.2681** | 0.1163 | $0.2744^{* * *}$ | 0.0541 | $0.1300^{* * *}$ | 0.0288 |
| Age (squared) | $-0.0036^{* *}$ | 0.0016 | $-0.0037^{* * *}$ | 0.0008 | $-0.0016^{* * *}$ | 0.0005 |
| F - Test | $5.3096{ }^{* * *}$ |  | $22.7786^{* * *}$ |  | $40.7848^{* * *}$ |  |
| Adjusted $R^{2}$ | 0.1113 |  | 0.2859 |  | 0.3503 |  |
| Observations | 173 |  | 273 |  | 370 |  |

Note: The drastic increase in the Adjusted $R^{2}$ - Statistic, interpreted as the fraction of the sample variation in $y_{1}$ that is explained by $y_{0}$, remains an issue to be solved. Basic Specification
Source: own calculations
Level of Significance: ${ }^{* * *} 1 \%{ }^{* *} 5 \%{ }^{*} 10 \%$
${ }^{1}$ Coeff. $=$ Coefficient; SE. $=$ Standard Error
${ }^{2} 5$ - year average of father's logarithmized adjusted real gross monthly income
${ }^{3}$ father at most 50 years of age
Table 14: Robustness Check I: Including Younger Siblings

|  | 5 - Year | verage | 4 - Year | verage | 3 - Year | Average | 2 - Year | verage | 1 - Year | verage |
| :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: |
|  | Coeff. ${ }^{1}$ | SE. ${ }^{1}$ | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. |
| Father |  |  |  |  |  |  |  |  |  |  |
| Measure ${ }^{2}$ | $0.2492^{* * *}$ | 0.0843 | $0.2485^{* * *}$ | 0.0840 | $0.2450^{* * *}$ | 0.0835 | $0.2258^{* * *}$ | 0.0827 | $0.1752^{* * *}$ | 0.0781 |
| Age | -0.1043 | 0.0707 | -0.1041 | 0.0707 | -0.1028 | 0.0707 | -0.1014 | 0.0709 | -0.0950 | 0.0712 |
| Age (squared) | 0.0014 | 0.0009 | 0.0014 | 0.0009 | 0.0014 | 0.0009 | 0.0014 | 0.0009 | 0.0013 | 0.0009 |
| Son |  |  |  |  |  |  |  |  |  |  |
| Age | $0.3204^{* * *}$ | 0.1127 | $0.3191^{* * *}$ | -0.1127 | $0.3185^{* * *}$ | -0.1127 | $0.3143^{* * *}$ | -0.1131 | $0.3194^{* * *}$ | -0.1137 |
| Age (squared) | $-0.0043^{* * *}$ | 0.0016 | $-0.0042^{* * *}$ | 0.0016 | $-0.0042^{* * *}$ | 0.0016 | $-0.0042^{* * *}$ | 0.0016 | $-0.0043^{* * *}$ | 0.0016 |
| F - Test | $6.5570^{* * *}$ |  | $6.5590^{* * *}$ |  | $6.5282^{* * *}$ |  | $6.2706^{* * *}$ |  | $5.7326^{* * *}$ |  |
| Adjusted $R^{2}$ | 0.1144 |  | 0.1145 |  | 0.1139 |  | 0.1092 |  | 0.0991 |  |
| Observations | 216 |  | 216 |  | 216 |  | 216 |  | 216 |  |

Source: own calculations
Level of Significance: *** $1 \% * * 5 \% * 10 \%$
${ }^{1}$ Coeff. $=$ Coefficient ; SE. $=$ Standard Error
${ }^{2}$ average of father's logarithmized adjusted real gross monthly income
Table 15: Robustness Check II: Income Adjustment Using Growth Rate of Industry Wages

|  | 5 - Year A | verage | 4 - Year A | verage | 3 - Year A | verage | 2 - Year | verage | 1 - Year | verage |
| :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: | :---: |
|  | Coeff. ${ }^{1}$ | SE. ${ }^{1}$ | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. | Coeff. | SE. |
| Father |  |  |  |  |  |  |  |  |  |  |
| Measure ${ }^{2}$ | $0.2548^{* * *}$ | 0.0874 | $0.2610^{* * *}$ | 0.0871 | $0.2542^{* * *}$ | 0.0859 | $0.2373^{* * *}$ | 0.0847 | 0.1970** | 0.0794 |
| Age | -0.0692 | 0.0707 | -0.0701 | 0.0706 | -0.0673 | 0.0705 | -0.0663 | 0.0707 | -0.0608 | 0.0709 |
| Age (squared) | 0.0010 | 0.0009 | 0.0010 | 0.0009 | 0.0010 | 0.0009 | 0.0010 | 0.0009 | 0.0009 | 0.0009 |
| Son |  |  |  |  |  |  |  |  |  |  |
| Age | 0.2801** | 0.1164 | 0.2789** | 0.1162 | $0.2776 * *$ | 0.1163 | 0.2693** | 0.1165 | 0.2750** | 0.1171 |
| Age (squared) | $-0.0038^{* *}$ | 0.0016 | $-0.0037^{* *}$ | 0.0016 | $-0.0037^{* *}$ | 0.0016 | $-0.0036^{* *}$ | 0.0016 | $-0.0037^{* *}$ | 0.0016 |
| F - Test | $5.2340^{* * *}$ |  | $5.3396{ }^{* * *}$ |  | $5.2925^{* * *}$ |  | $5.0919^{* * *}$ |  | $4.7188^{* * *}$ |  |
| Adjusted $R^{2}$ | 0.1096 |  | 0.1120 |  | 0.1109 |  | 0.1109 |  | 0.0976 |  |
| Observations | 173 |  | 173 |  | 173 |  | 173 |  | 173 |  |

[^6]Table 16: Robustness Check III: IV - Estimation Using Wegener - Index

|  | Lower Age Limit -5 |  | 30-50 Age Limit |  | Upper Age Limit +5 |  |
| :---: | :---: | :---: | :---: | :---: | :---: | :---: |
|  | Coeff. ${ }^{1}$ | SE. ${ }^{1}$ | Coeff. | SE. | Coeff. | SE. |
| Father <br> Measure ${ }^{2}$ | 0.2890** | 0.1148 | $0.3800^{* * *}$ | 0.1394 | $0.3182^{* *}$ | 0.1229 |
| Son <br> Age <br> Age (squared) | $\begin{array}{r} 0.2528^{* * *} \\ -0.0033^{* * *} \end{array}$ | $\begin{aligned} & 0.0556 \\ & 0.0008 \end{aligned}$ | $\begin{gathered} 0.2608^{* *} \\ -0.0034^{* *} \end{gathered}$ | $\begin{aligned} & 0.1187 \\ & 0.0016 \end{aligned}$ | $\begin{gathered} 0.2408^{* * *} \\ -0.0031^{* *} \end{gathered}$ | $\begin{aligned} & 0.2408 \\ & 0.0012 \end{aligned}$ |
| F - Test <br> Adjusted $R^{2}$ <br> Observations | 31.361 0.25 26 |  | 5.0566 0.07 167 |  | 6.00 0.0 2 | 5*** 20 4 |

Source: own calculations
Level of Significance: ${ }^{* * *} 1 \% * * 5 \% * 10 \%$
Coeff. = Coefficient ; SE. $=$ Standard Error
${ }^{2}$ Wegener - Index of occupational prestige

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[^1]:    ${ }^{1}$ See Behrman \& Taubman (1985) as an example.
    ${ }^{2}$ For a further errors - in - reporting problem see Bound \& Krueger (1991) and Duncan \& Hill (1985).

[^2]:    ${ }^{3}$ Deflated by the appropriate price index supplied by the German Federal Statistical Office.
    ${ }^{4}$ This approach is similar to Wiegand (1997), but different from Vogel (2007), who calculates a measure of monthly yearly income form monthly income records.
    ${ }^{5}$ See Hunt (2001) for an empirical analysis.

[^3]:    ${ }^{6}$ As reported by the German Federal Statistical Office

[^4]:    ${ }^{7}$ This variable includes both, school and occupational education. The German school system introduces differentiated educational tracks already after four grades of primary education. The basic school (Hauptschule) graduates individuals after five years of secondary education and is traditionally a preparation for blue collar occupations. The middle school (Realschule) lasts six years and trains for white collar employment. The highest track (Gymnasium) offers nine years of schooling and a degree (Abitur), which is a precondition for academic studies. Completion of an apprenticeship adds another 1.5 years, a technical college 3 years, and graduation form university increases years of education by 5 years.

[^5]:    Basic Specification
    Source: own calculations
    Level of Significance: ${ }^{* * *} 1 \% * * 5 \% * 10 \%$
    ${ }^{1}$ Coeff. = Coefficient ; SE. $=$ Standard Error
    ${ }^{2} 5$ - year average of father's logarithmized adjusted real gross monthly income
    ${ }^{3}$ son at least 30 years of age

[^6]:    Source: own calculations
    Level of Significance: *** $1 \% * * 5 \% * 10 \%$
    ${ }^{1}$ Coeff. $=$ Coefficient ; SE. $=$ Standard Error
    ${ }^{2}$ average of father's logarithmized adjusted real gross monthly income

