Assessing Intergenerational Earnings Persistence Among German Workers

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In this study we assess the relationship between father and son earnings among (West) German Workers. To reduce the lifecycle and attenuation bias a novel sampling procedure is developed and applied to the German Socio-Economic Panel (SOEP) 1984-2006. Our preferred point estimate indicates an intergenerational earnings elasticity of $\frac{1}{3}$.

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1. Introduction and Motivation

> The vitality and stability of our democracy - as well as the economy -
> eventually depend on the social permeability of our society[1].

This statement draws attention to the strong meritocratic believes concerning the
equality of opportunity that govern public debates. This is especially true for the
education system. According to public rhetoric it is aimed to guarantee social mobil-
ity in Germany. 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 aid for students from low-income families. But, does this general
concern translate into a society in which one's economic success is independent of
the family born into? And if so, to what degree?

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[1] Horst Köhler, German Federal President, in an interview with the Frankfurter Allgemeine
Zeitung, 29.12.2007, Berlin (own translation).

To empirically analyze the intergenerational relationship, the following econometric model

$$(1) \qquad y_1^i \quad = \quad \alpha + \beta_y y_0^i + \epsilon_1^i$$

is used as a starting point (Corak 2004). A linear relationship between long-run economic status y_0^i and y_1^i of family i in generation 0 and 1 is assumed, allowing for shifts in mean economic status independent of parental status via the parameter α. Deviations from predicted status due to market luck or other random elements in the intergenerational transmission of skills and personal traits are summarized in the idiosyncratic error term ϵ_1^i. Ideally, permanent earnings are chosen as the measure of economic status (Friedman 1957). We use both terms to describe the long-run economic success of an individual. In the case all status variables are measured in their natural logarithm, β_y in equation (1) is the intergenerational elasticity of permanent earnings. It measures the (expected) percentage change in offspring's economic status associated with a one percent change in parental success. In principle, β_y can take any value but most studies find a value between zero and one[2]. A positive value does indicate generational persistence of permanent earnings in which higher parental long-run status favors economic success of one's offspring; a negative number indicates generational reversal of economic status. A value of zero for the intergenerational elasticity β_y (child's and parental economic success are unrelated) 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. $(1 - \beta_y)$ provides a measure of the degree to which economic status regresses to the mean (Becker & Tomes 1986, Goldberger 1989). If it takes value one ($\beta_y = 0$), a child from parents who attain below average long-run status can expect average status just as the offspring of high status parents.

Although there is agreement about the existence of an intergenerational link in economic status, a number of recent studies debate its varying magnitude across countries (Solon 2002, Grawe 2006, Jäntti, Røed, Naylor, Björklund, Bratsberg, Raaum, Österbacka & Eriksson 2006, Vogel 2007). While many features of the human skill formation process are universal, there may however be unique features in German data. In an international perspective, low tuition fees and federal student

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[2]See Solon (2002) for a recent survey.

aid might ease the impact of borrowing constraints and thus enhance mobility in Germany compared to other countries.

The contribution of our paper to the literature on intergenerational persistence is twofold. First, based upon recent improvements in the understanding of the association between short- and long-run economic status we asses the potential biases in previous studies. Deviations of current from permanent economic status arise due to transitory fluctuations (Bowles 1972, Solon 1992) and a time-varying association between the two (Haider & Solon 2006, Grawe 2006). We introduce a novel sampling procedure that accounts for both and allows to observe father-son pairs at a rather similar stage of their lifecycle. Second, the relationship is assessed for Germany with samples drawn from the German Socio-Economic Panel (SOEP) 1984-2006.

Our results suggest that the best conservative point estimate of intergenerational earnings persistence among West German workers is $\frac{1}{3}$. This indicates a lower degree of mobility (and a higher degree of persistence) in Germany compared to Couch & Dunn (1997) and Wiegand (1997) but is in line with Vogel (2007), who compares intergenerational mobility between Germany and the United States. In an international perspective, the intergenerational earnings persistence seems to be lower compared to the United States $\beta_y^{US} = 0.4$ (Solon 1992), and higher compared to Sweden $\beta_y^{S} = 0.2$ (Björklund & Jäntti 1997). There still seems to be substantial intergenerational earnings mobility among West German workers, but more persistence than previous research suggested.

The remainder of this paper is organized as follows: Section 2 provides an introduction to the econometric methods applied to estimate intergenerational persistence with incomplete data. Section 3 presents our novel sampling procedure with the SOEP. Section 4 discusses the econometric findings, whereas section 5 concludes.

2. Econometric Problems and Findings from the Literature

In this section the econometric problems associated with measuring intergenerational persistence and the consequences we draw regarding its estimation among German workers are pointed out.

2.1. Measurement Error Problems

The deduction of an individual's permanent earnings requires a lifelong earnings history. Since researchers usually lack direct measures of long-run status y_0^i and y_1^i

for two generations in order to investigate intergenerational mobility, they rely on proxies (y_{0h}^i, y_{1t}^i) of permanent earnings for each generation $(0, 1)$ observed at age h and t. Sometimes only single-year measures of earnings[3] have been used. Usually, however, a short-run measure of economic status is an imperfect proxy of long-run status. It is subject to measurement error due to transitory fluctuations and lifecycle variation in the association between current and lifetime earnings[4].

2.1.1. Transitory Fluctuations

Current earnings of fathers y_{0h}^i and sons y_{1t}^i can be decomposed as follows (Friedman 1957).

$$
\begin{align}
(2) \qquad y_{1t}^i &= y_1^i + v_{1t}^i \\
(3) \qquad y_{0h}^i &= y_0^i + v_{0h}^i
\end{align}
$$

y_0^i and y_1^i describe time-invariant permanent earnings, while (v_{0h}^i, v_{1t}^i) indicates time-varying transitory fluctuations. The latter might arise from job mobility, business cycle effects or variable compensation schemes. If current earnings deviate from permanent status, using them as a proxy for long-run status introduces attenuation bias in the estimation of equation (1). Assuming that v_{1t}^i and v_{0h}^i are uncorrelated with each other and permanent earnings y_0^i and y_1^i, a deviation implies a downward inconsistency of the estimated slope coefficient $\hat{\beta}_y^{OLS}$ in an OLS estimation by the factor θ_h (Solon 1992).

$$
\begin{align}
(4) \qquad \text{plim } \hat{\beta}_y^{OLS} &= \theta_h \beta_y < \beta_y \\
(5) \qquad \theta_h &= \left(\frac{Var[y_0]}{Var[y_0] + Var[v_{0h}]} \right)
\end{align}
$$

The attenuation factor θ_h captures how much signal $Var[y_0]$ is provided by the measure y_{0h} relative to its total noise, $Var[y_{0h}] = Var[y_0] + Var[v_{0h}]$.

Based on single-year snapshots, empirical findings by Corcoran, Laren, Gordon & Solon (1991), Card (1994) and Hyslop (2001) suggest an attenuation factor around $\theta_h = 0.5$. This implies a (considerable) signal-to-noise ratio of observed parental earnings and an attenuation bias of $(1 - \theta_h) = 0.5$. Note also, that transitory

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[3] See Behrman & Taubman (1985) as an example.
[4] For a further errors-in-reporting problem see Bound & Krueger (1991) and Duncan & Hill (1985).

fluctuations in offspring's earnings v_{1t}^i do not bias the OLS estimation in equation (1) as long as they are uncorrelated with v_{0h}^i. However, the higher their variance, the larger the confidence interval of $\hat{\beta}_y^{OLS}$ will be.

Averaging Parental Earnings

To decrease the magnitude of the inconsistency, Solon (1992) suggests to average parental status over T years which 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).

$$(6) \qquad \theta_h \;=\; \left(\frac{Var[y_0]}{Var[y_0] + \frac{1}{T}Var[v_{0h}]} \right)$$

As more years of data are used, the attenuation factor θ_h rises and the attenuation bias $(1 - \theta_h)$ declines. According to Mazumder (2005), the attenuation factor θ_h rises to $\theta_h = 0.7$ (from $\theta_h = 0.5$) when relying on a 5-year average of earnings. The attenuation bias is reduced to $[(1 - \theta_h) = 0.3]$. Solon (1992) and Wiegand (1997) estimated an intergenerational elasticity of father's and son's earnings 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.

Instrumenting Parental Earnings

In a second approach to estimate β_y, the direct projection of y_1^i on y_0^i, Solon (1992) proposes an IV estimation. Acknowledging the difficulties of finding an instrument, that is correlated with parental long-run status but not a structural determinant of offspring's permanent earnings, the basic model in equation (1) is amended to include an additional factor I_0^i.

$$(7) \qquad y_1^i \;=\; \beta_1 y_0^i + \beta_I I_0^i + \omega_1^i$$

In this case, performing an IV estimation of β_y using I_0^i as the instrument, yields

the following probability limit, Solon (1992).

$$\text{(8)} \qquad \text{plim } \hat{\beta}_y^{IV} \;=\; \beta_y + \beta_I \left(\frac{1 - \kappa^2}{\kappa} \right) \left(\frac{Sd[I_0]}{Sd[y_0]} \right)$$

$$\text{(9)} \qquad \kappa \;=\; \frac{Cov[I_0, y_0]}{Sd[y_0]Sd[I_0]}$$

$\hat{\beta}_y^{IV}$ is an unbiased estimator for β_y only if the instrument does not influence off-spring's status ($\beta_I = 0$) or the instrument and parental status are perfectly corre-lated, $|\kappa| = 1$. The closer $|\kappa|$ is to one, the smaller the bias as there is less variation in earnings that is not captured by the instrument. Assuming a positive but imper-fect correlation between the instrument and parental long-run status, the direction of the inconsistency is determined by β_I. If the instrument I_0^i has a positive impact on offspring' s status ($\beta_I > 0$), the estimator will be biased upward. If the opposite is true, the estimated coefficient is downward biased like the OLS estimate.

In empirical research, parental years of education (Solon 1992, Dearden, Machin & Reed 1997) or indicators of occupational prestige (Zimmerman 1992, Wiegand 1997) have been used to instrument long-run parental status. Since years of education enhance labor market earnings, it may capture an important part of parental per-manent earnings, although not necessarily to a 100%[5]. 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 there-fore suggests to bracket the coefficient (Solon 1992). The OLS estimate is downward inconsistent due to error-in-variable bias, whereas the IV estimate is presumably up-ward biased. Accounting for the associated standard errors, β_y is located between the two estimates.

$$\boxed{\hat{\beta}_y^{OLS} < \beta_y < \hat{\beta}_y^{IV}}$$

2.1.2. Lifecycle Variations

Empirical research as well as theoretical reasoning suggest that wage workers differ with respect to their age-earnings profiles[6]. This may occur due to age-specific het-erogeneity in human capital investment or variations in the wage structure across

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[5]See Card (1999) for a recent survey.
[6]See Mincer (1975) and Baker (1997) among others and Vogel (2007) for an application to inter-generational mobility.

jobs erected by firms for the purpose of effort regulation and incentive compatibility. For estimation purposes, the projection of current on permanent earnings is generalized to include a time-varying parameter $\lambda_{t,h}$ to capture age-specific aspects in the association between current and permanent earnings over the lifecycle (Haider & Solon 2006).

$$(10) \qquad y_{1t}^i \;=\; \lambda_t y_1^i + v_{1t}^i$$

$$(11) \qquad y_{0h}^i \;=\; \lambda_h y_0^i + v_{0h}^i$$

Averaging parental earnings y_{0h}^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 earnings, equation 12 determines the potential bias.

$$(12) \qquad \text{plim } \hat{\beta}_y^{OLS} \;=\; \lambda_t \theta_h \beta_y$$

$$(13) \qquad \theta_h \;=\; \frac{\lambda_h Var[y_0]}{\lambda_h^2 Var[y_0] + \frac{1}{T} Var[v_{0h}]}$$

Assuming $\theta_h = 1$, the probability limit of the estimated coefficient $\hat{\beta}_y^{OLS}$ is $\lambda_t \beta_y$ instead of β_y. In the case of $\lambda_t = 1$ (as implicitly assumed in the discussion of 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 earnings are observed. Focusing on the impact of θ_h (setting $\lambda_t = 1$), it is not obvious whether the combination of transitory fluctuations and lifecycle variation leads to an amplification bias instead of an attenuation bias. For $\lambda_h > 1$ the estimation is downward biased, but for values smaller than one and minor transitory variance the opposite is true. θ_h is a summary measure of the attenuation bias resulting from transitory fluctuations as well as lifecycle variation. Therefore the age composition of the sample matters (Jenkins 1987, Grawe 2006). In summary, measurement error in offspring's status is *not* innocuous for consistency as well as measurement error in parental long-run status. Both induce either amplification or attenuation bias of the OLS estimation.

Using U.S. Social Security Administration earnings 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 earnings separately. Their dataset ranges from 1951 to 1991 and provides nearly career-long earnings histories for a broadly representative sample of the U.S. population. This allows to derive a
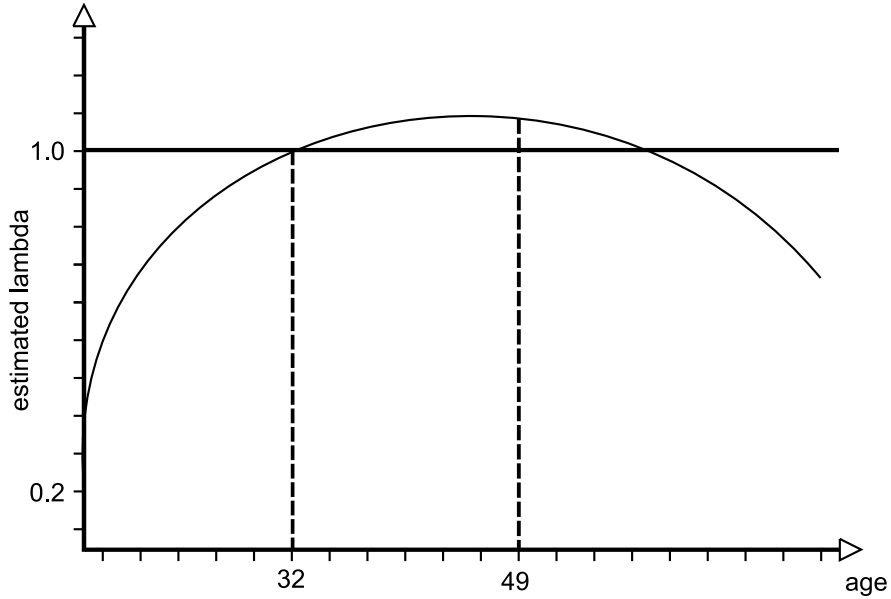
Figure 1: Estimated Correlation Between Current and Permanent Earnings

more precise estimate of the (logarithmized) present value of lifetime earnings $\ln V^i$. Starting with the impact of measurement error in offspring's (permanent) earnings level, the forward regression of $\ln V^i$ on $y^i_{t,h}$ leads to the estimated slope coefficient $\hat{\lambda}_{t,h}$ depicted in figure 1. Starting at a value around $\hat{\lambda}_{t,h} = 0.2$ it increases steadily. At age 32, the textbook assumption of $\lambda_{t,h} = 1$ seems reasonable. Thenceforward, $\hat{\lambda}_{t,h}$ declines some in the late forties. Turning to the case of measurement error in parental permanent earnings, the estimated reliability ratio $\hat{\theta}_h$ is depicted in figure 2. It is the result of a backward regression of $\ln V^i$ on a 5-year average of $y^i_{t,h}$. A significant increase till age 30 is followed by a quite robust factor between 0.6 and 0.8, but after the age of 50, $\hat{\theta}_h$ declines and the bias rises. Unfortunately, we are not aware of any comparable work for the case of Germany.

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.
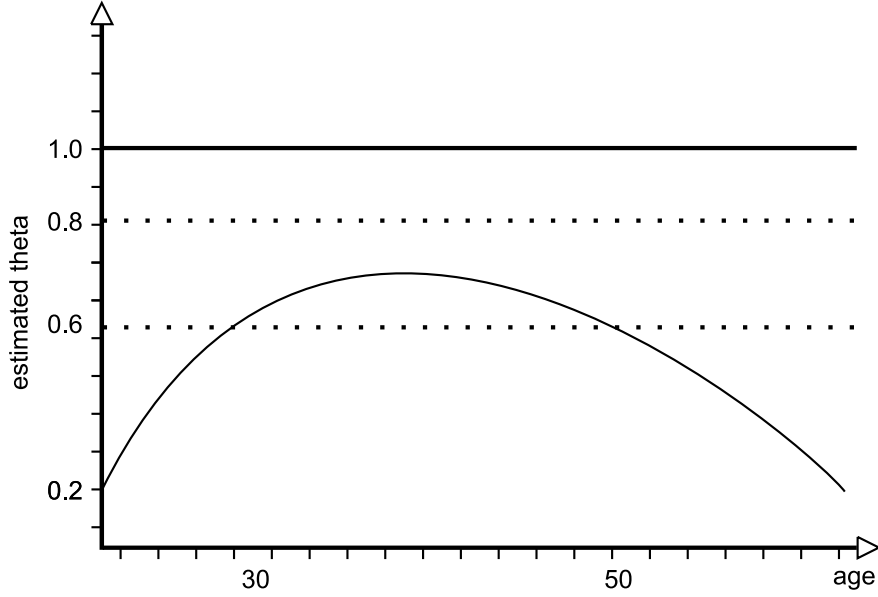
Figure 2: Estimated Reliability Ratio

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 may 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 earnings $Var[y_{j=0,1}^*]$ is only a fraction τ of the population variance $Var[y_{j=0,1}]$.

$$(14) \qquad\qquad Var[y_{j=0,1}^*] = \tau Var[y_{j=0,1}]$$

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).

$$(15) \qquad\qquad \text{plim } \hat{\beta}_{y*}^{OLS} \;=\; \phi\beta_y < \beta_y$$
$$(16) \qquad\qquad\qquad \phi \;=\; \frac{\tau}{1 - R^2(1-\tau)}$$

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

A sample exhibiting homogeneity in parental earnings 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 (Solon 1992, Wiegand 1997), see equation (17).

$$(17) \quad \left(\frac{Var[y_{0*}]}{Var[y_{0*}] + Var[\upsilon_{0h}]} \right) \beta_y = \operatorname{plim} \hat{\beta}_{y*}^{OLS} <$$

$$\operatorname{plim} \hat{\beta}_y^{OLS} = \left(\frac{Var[y_0]}{Var[y_0] + Var[\upsilon_{0h}]} \right) \beta_y$$

In applied empirical research, inclusion into an intergenerational dataset requires for father and son to both report positive labor market earnings in the periods of interest. Presumably, in such samples β_y is underestimated, but the use of larger representative samples eases this problem. To the best of our knowledge, however, there is no research on the magnitude of this bias available.

3. Econometric Approach and Sampling Procedure

We estimate the econometric model presented in equation (18). Son's observed status y_{1t}^i in year t is expressed as a regression function of father's observed status y_{0h}^i in year h, including age-controls for both (Solon 1992, Zimmerman 1992, Wiegand 1997, Vogel 2007). It is derived by the incorporation of age-earnings profiles into equations (2) and (3) and substitution into the basic equation (1).

$$(18) \qquad y_{1t}^i \;\; = \;\; \beta_0 + \beta_y y_{0h}^i + \beta_1 A_{0h}^i + \beta_2 A_{0h}^{2i} + \beta_3 A_{1t}^i + \beta_4 A_{1t}^{2i} + \omega_{1t}^i$$

An individual's current earnings are determined by the level of permanent earnings (y_1^i, y_0^i), the stage in the lifecycle $[(A_t^i; A_t^{2i}), (A_h^i; A_h^{2i})]$, a general level of economic well-being in the corresponding generation (α_1, α_0), and an idiosyncratic error term $(\upsilon_{1t}^i, \upsilon_{0h}^i)$.

$$(19) \qquad\qquad y_{1t}^i \;\; = \;\; y_1^i + \alpha_1 + \gamma_1 \, A_{1t}^i + \delta_1 \, A_{1t}^{2i} + \upsilon_{1t}^i$$
$$(20) \qquad\qquad y_{0h}^i \;\; = \;\; y_0^i + \alpha_0 + \gamma_0 \, A_{0h}^i + \delta_0 \, A_{0h}^{2i} + \upsilon_{0h}^i$$

The empirical part builds on samples from the German Socio-Economic Panel[7] (SOEP) from 1984 to 2006. To assure comparability of real earnings observed in different years, they are adjusted by the real GDP-Growth Rate. Our measure of long-run economic status are real[8] monthly earnings before tax and social security deductions as reported in each cross-section of the SOEP[9]. This allows for international comparison[10]. Measuring all earnings variables in their natural logarithm, we choose the intergenerational earnings elasticity as our indicator for intergenerational persistence (or mobility) and use both terms interchangeably.

Thus, our indicator is a summary measure of personal characteristics shared by parent and offspring that are valued in the German labor market. This includes similarities in educational attainment, cognitive and noncognitive skills and personal traits (Bowles, Gintis & Osborne-Groves 2001, Bowles & Gintis 2002, Heckman 2007, Pfeiffer & Reuß 2007).

Table 1: Final Sample Overview

Groups Excluded from Sample	Measures of Economic Status	Age - Restrictions
self-employed	<u>Son</u>	
part-time employed	monthly earnings (1984 - 2006)	between 30 - 50
East Germans	<u>Father</u>	
migrants	monthly earnings (1984 - 2006)	
younger brothers	years of education	

A novel feature of our study is the sampling procedure. We select pairs of fathers and sons in a way that their earnings are observed as close in their lifecycle as possible. Furthermore, the bias due to transitory fluctuations and lifecycle variation is minimized, see table 1. As a start, the self-employed, who have more volatile earnings (Baker & Solon 2003, Albarrán, Carrasco & Martínez-Granado 2007, Pfeiffer 1994)

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[7]Consult Haisken-DeNew & Frick (2005) for further information on the dataset.

[8]Deflated by the consumer price index (base year 2000) supplied by the German Federal Statistical Office.

[9]This approach is similar to Wiegand (1997), but different from Vogel (2007), who calculates a measure of yearly earnings from monthly earnings records.

[10]See Solon (1999) for a survey on intergenerational earnings mobility. We concentrate on persistence of labor market earnings. For research using a more inclusive measure of total economic status made up by a variety of differing types of income, earnings and monetary inheritance see Piketty (2000) and Mulligan (1997) among others. It is left for future research to construct a more inclusive measure with the SOEP, since sample size is reduced and the problem of measurement error increases.

are excluded. 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 dynamic wage growth[11] may have changed the reliability of current earnings to reflect permanent status. To avoid sample homogeneity, only the oldest sibling is included in our baseline specification (Solon 1992). Migrants, identified by their country of origin, are dropped for our basic analysis. Migration might distort the long-run relationship between labor market earnings of father and son due to the change of the labor market (Borjas 2006, Friedberg 2000). However, we perform a separate analysis for migrants and discuss the results. For the group of fathers, moving 5-year averages of earnings and age are calculated to reduce the attenuation bias. Thus, if for a given observation earnings are not observable in each of the four following years, it is dropped.

Furthermore, the following age restrictions are imposed (and relaxed again for further discussion) to account for the time-varying association between short- and long-run economic status. This procedure accounts for the pattern of the variance of the transitory component over the lifecycle (which flattens out at mid age in the U.S. (Baker & Solon 2003)). Since the association between monthly and lifetime earnings is still low for workers below the age of 30, we select workers above that age. For younger workers job mobility is high and earnings are more volatile, partly because of lower tenure (Haider & Solon 2006, Björklund 1993). Workers aged above 50 years are excluded as well. Labor market status and hours worked may become more volatile again which might depress the estimated level of persistence (Grawe 2006). However, this line of reasoning may differ between countries, for instance as a result of different industrial structures or different degrees of employment protection laws (Blau & Kahn 1996, OECD 1999, Pries & Rogerson 2005).

Finally, father (obs_k and obs_l) and son (obs_j and obs_i) observations (of family n) satisfying the sampling rule are matched in all possible combinations, see figure 3. This procedure leads to numerous matched observations for each father-son pair. To identify a unique pair, intended to lead to the most reliable estimate of the intergenerational elasticity, a decision rule is implemented. For each observation we select the one with the smallest absolute age difference between father and son. This is to ensure that father and son are observed at as similar stages in their lifecycle as

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[11]See Hunt (2002) among others.
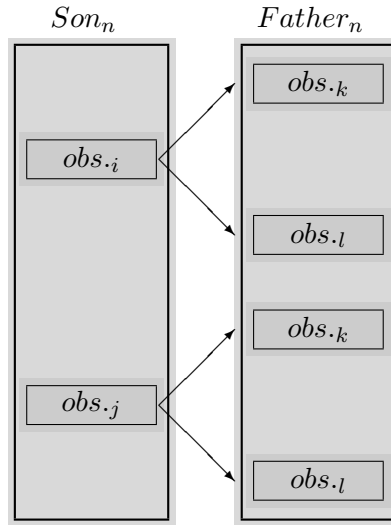
Figure 3: Sampling Strategy

possible. 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. For comparison and discussion, other samples with less restrictive selection rules are utilized in the next section.

The sample contains 180 father-son pairs compared to Wiegand's (1997) 130 and Vogel's (2007) 300. Table 2 depicts the basic statistics. The age difference between father and son amounts to 8.68 years. Sons in the sample report less earnings than their matched fathers which is mainly explained by their early stage in the lifecycle. While most information on father's economic status is obtained within the early SOEP waves, the collection of offsprings' information is not confined to the most recent wave. The age-composition of our sample differs substantially from previous works. Sons are 35 years old which is an increase of 4 years compared to Wiegand (1997) and 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. 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).

Table 2: Final Sample Statistics

Statistic	Fathers	Sons
Gross Earnings in €[1]	2,307.03	1,936.87
Sd. of Gross Earnings	716.82	640.40
Year of Observation	1,987	2,004
Age in Years	44.40	35.73
Age - Difference in Years	8.68	
Number of Observations	180	

[1] reported 5/1 - year average of adjusted real gross monthly earnings

Selection could rise from the blind eye on individuals not meeting the selection rules, table 2. The final sample is compared to all workers living up to the sample requirements except for the need to report positive earnings 5 years in a row and being matched with their offspring. The father-sample is contrasted in 1984, while the son-sample is compared in 2004. Earnings in the father-sample are nearly identical to the one reported by all workers in 1984 (2,331.01€). However, the standard deviation is higher in the comparison group (782.43€) in 1984. Using 5-year averages of earnings in the father-sample, therefore, as intended, reduces transitory fluctuations. Comparing the son-sample, earnings are higher (1,917.13€ in the comparison group) and show a higher standard deviation (574.15€). In our son-sample the average age is lower which induces higher wage dispersion.

4. Econometric Findings

4.1. Basic Results

Table 3: Basic Results

	5 - Year Avg. Earnings	*Single - Year Earnings*
<u>OLS Estimate</u>		
Intergenerational Elasticity	0.282	0.205
95% Confidence Interval	(0.09 - 0.44)	(0.08 - 0.32)
Standard Error	0.087	0.061
Observations	180	249
<u>IV Estimate</u>[1]		
Intergenerational Elasticity	0.374	
95% Confidence Interval	(0.09 - 0.65)	
Standard Error	0.144	
Observations	180	

[1] using years of education

The OLS estimate based on a 5-year average of earnings $\hat{\beta}_y^{OLS} = 0.282$ is higher compared to Wiegand (1997), whereas the one-year snapshot is about the same, table 3. Compared to Vogel (2007), the result based on the 5-year average of earnings is similar. We use years of education[12] as an instrument to bracket the intergenerational elasticity. According to the IV estimate the intergenerational elasticity is higher, $\hat{\beta}_y^{IV} = 0.374$. Following Solon's (1992) approach, the intergenerational elasticity of German Workers should lie between the two estimates and we suggest a reasonable value of $\frac{1}{3}$.

$$\hat{\beta}_y^{OLS} = 0.282 < \beta_y < 0.374 = \hat{\beta}_y^{IV}$$

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[12]This variable includes both, school and occupational education. The German school system introduces differentiated educational tracks 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.

The 95% confidence interval of the IV estimate $[0.09 \leq \hat{\beta}_y^{IV} \leq 0.66]$ includes the OLS estimate. Although the two point estimates contain some useful information, the degree of precision seems to be rather low. We come back to this issue in the conclusion.

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	0.2822***	0.2841***	0.2751***	0.2441***	0.1984**
95% Confidence Interval	(0.09 - 0.44)	(0.11 - 0.45)	(0.11 - 0.44)	(0.08 - 0.41)	(0.04 - 0.35)
Standard Error	0.0870	0.0866	0.0854	0.0841	0.0798
Observations	180	180	180	180	180

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 earnings

Table 4 and 5 report the general pattern that $\hat{\beta}_y^{OLS}$ 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, father earnings need 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 earnings 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 earnings 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 earnings measure utilized. This result is 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	0.2822***	0.2867***	0.2596***	0.2076***	0.2045***
95% Confidence Interval	(0.11 - 0.45)	(0.13 - 0.45)	(0.10 - 0.41)	(0.07 - 0.34)	(0.08 - 0.32)
Standard Error	0.0870	0.0815	0.0790	0.0695	0.0614
Observations	180	190	217	227	249

Basic Specification

Source: own calculations

Level of Significance: *** 1% **5% *10%

[1] see Table 11 in the Appendix for the detailed results

[2] average of father's logarithmized adjusted real gross monthly earnings

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 earnings in the following year likely due to un- or part-time employment. This implies that father-son pairs with larger transitory fluctuations are consecutively added to the panel.

4.3. Investigating the Bias from Lifecycle Variation

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 persistence. However, table 6 reveals an increase in the estimate when continuing to soften the age restriction. This seems to be in line with Vogel (2007), whose estimate of intergenerational persistence in Germany based including individuals aged above 50 is slightly higher. We offer two explanations. First, the increase could point at sample selection with only pairs added that exhibit a particular strong persistence of earnings. But a comparison of the descriptive statistics (years of education, monthly earnings) did not offer any evidence on the type of selection. Second, the increase in the estimated level of mobility could be explained by an increase in the reliability ratio θ_h in our sample rather than a decrease as documented for the United States, figure 2. This is presumably the result of the comparatively high degree of centralization governing wage determination in Germany and employment protection laws which, together

with the accumulation of specific human capital, favor incumbent workers[13]. This could decrease the transitory fluctuations among older German workers.

Table 6: Summary Results[1]: Relaxing Age - Restrictions for Fathers

Father's Maximal Age	50	55	60	65
Intergenerational Elasticity[2]	0.2822***	0.2509***	0.3538***	0.3584***
95% Confidence Interval	(0.11 - 0.45)	(0.10 - 0.41)	(0.22 - 0.49)	(0.22 - 0.49)
Standard Error	0.0870	0.0794	0.0696	0.0686
Observations	180	240	281	285

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 earnings; 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 for sons is consecutively lowered to 20 years. This seems to be in line with Haider & Solon (2006). The parameter λ_t (see equation (12) in section 2) is lowered as younger and younger workers are added to the sample and the lifecycle bias rises.

Table 7: Summary Results[1]: Relaxing Age - Restrictions for Sons

Son's Minimum Age	30	25	20
Intergenerational Elasticity[2]	0.2822***	0.2553***	0.2402***
95% Confidence Interval	(0.11 - 0.45)	(0.12 - 0.39)	(0.13 - 0.30)
Standard Error	0.0870	0.0666	0.0558
Observations	180	282	385

Basic Specification

Source: own calculations

Level of Significance: *** 1% **5% *10%

[1] see Table 14 in the Appendix for the detailed results

[2] 5-year average of father's logarithmized adjusted real gross monthly earnings; father at most 50 years of age

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[13]See Botero, Djankov, Porta & Lopez-De-Silanes (2004) and Franz & Pfeiffer (2006) among others.

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. An Analysis of Migrants and Further Sensitivity Checks

Analysis of Migrants

The analysis of the migrant population[14] is based on 93 father-son pairs when relying on a 5-year average of earnings, see table 15 in the Appendix for the detailed results. It turns out, that the point estimate of the intergenerational earnings elasticity ($\hat{\beta}_y = 0.41$) is higher compared to our sample of German workers. While the age structure of the samples are rather identical, there are substantial differences in average earnings. For migrants, son earnings are lower (1,617€ compared to 1,936€). The same is true for migrant fathers, who earn 1,789€ on average compared to 2,307€ for a German father. This hints at possible non-linearities in the intergenerational link. For example Corak & Heisz (1999) and Hertz (2005) present evidence for stronger persistence among low-income families.

Including Younger Siblings

The inclusion of younger siblings raises the sample size from 180 to 224 when relying on a 5-year average of father's earnings. The point estimate is slightly reduced to $\hat{\beta}_y^{OLS} = 0.276$. 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 slightly, see table 16 in the Appendix for the detailed results.

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[14]Identified by the fact, that at least the father is not born in Germany.

Adjustment of Monthly Earnings

To ensure robustness with respect to the measure of comparability (GDP-Growth in the baseline estimation), earnings are deflated by the growth rate of average real gross monthly earnings in Germany's industry sector (as reported by the German Federal Statistical Office). The estimated intergenerational elasticity is not affected, see table 17 in the Appendix for the detailed results.

Instrumenting Parental Status

To compare our findings with Wiegand (1997), the IV estimation is repeated instrumenting parental status using the Wegener-Index, a standard index for occupational prestige. The baseline estimate ($\hat{\beta}_y^{IV} = 0.372$) remains unchanged. The finding that both instruments lead to rather identical results is robust to changes in the sampling rule, see table 18 in the Appendix for the detailed results.

5. Discussion and Concluding Remarks

Figure 8 compares our result to the international evidence. Although the studies differ with respect to data and methods, the comparison suggests higher mobility (that is less persistence) in Germany compared to the United States and the United Kingdom, but lower mobility compared to Sweden.

Our preferred point estimate of the elasticity in Germany is $\beta_y^{GER} = \frac{1}{3}$, compared with $\beta_y^{US} = 0.4$ for the United States and $\beta_y^{S} = 0.2$ for Sweden. In comparison to former studies by Couch & Dunn (1997) and Wiegand (1997) on intergenerational persistence in Germany, Vogel (2007) and our results suggest higher persistence. This is the result of our special attention on the sources of potential lifecycle bias. However, common to all studies presented in table 8 are the considerable confidence intervals, which currently forbid any strong comparative statements on the level of intergenerational persistence.

Table 8: International Perspective

Country	OLS Result	IV Result
United States		
Mazumder (2005)	0.613 (0.09)[a]	—
Solon (1992)	0.413 (0.09)	0.526 (0.14)
Zimmerman (1992)	0.400 (0.06)	0.330 (0.27)
United Kingdom		
Dearden, Machin & Reed (1997)	0.240 (0.03)	0.443 (0.03)
Sweden		
Björklund & Jäntti (1997)	0.216 (0.04)	—
Germany		
Couch & Dunn (1997)	0.124 (0.07)	—
Wiegand (1997)	0.238 (0.06)	0.402 (0.13)
Vogel (2007)	0.266 (0.06)	—
this volume	0.282 (0.09)	0.374 (0.14)

[a] Standard errors in parentheses.

For illustrative reasons, we conclude by figuring out some consequences of the value of $\beta_y^{GER} = \frac{1}{3}$ for Germany. The intergenerational elasticity β_y translates intragenerational inequality in parental long-run labor market status into the economic advantage, which a child from parents with higher economic status can hope for compared to one from lower status parents. Table 9 depicts the advantage of a child with parents in the top permanent earnings decile compared to offspring born to parents in the bottom decile as determined by equation (21) (Corak 2004).

$$(21) \qquad \frac{y_1^{90th}}{y_1^{10th}} = \left(\frac{y_0^{90th}}{y_0^{10th}} \right)^{\beta_y}$$

Table 9: Inequality and the Expected Permanent Earnings Advantage

90/10 - Ratio	Intergenerational Elasticity			
	0.2	$\frac{1}{3}$	0.4	0.5
2.0	15%	25%	32%	41%
2.5	20%	35%	44%	58%
3.0	25%	44%	55%	73%
3.5	28%	51%	65%	87%
4.0	32%	59%	74%	100%

For Germany, Gernandt & Pfeiffer (2007) calculate a 90/10-percentile earnings ratio of 2.5 for a cross-section sample of prime age dependent male workers in 2005, which is rather close to our sample. Then, taking our advocated value for an intergenerational elasticity in Germany of $\beta_y^{GER} = \frac{1}{3}$, the expected earnings advantage amounts to 35%. If β_y^{GER} would be 0.5, the advantage increases to 59%. We would like to add that a value of $0 < \beta_y < 1$ does not necessarily imply a compression of the earnings distribution, because the variance of the error term ϵ_1^i in equation (1) matters, too.

Summarizing our findings, we find intergenerational earnings persistence among West German workers is higher than previously suggested. A value of $\beta_y^{GER} = \frac{1}{3}$ still indicates that there is substantial intergenerational mobility, which presumably is a result of the massive expansion of publicly funded education in Germany from the seventies and the openness of the German labor market.

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A. Estimation Appendix

Table 10: Detailed Results: Balanced Panel

	5 - Year Average		4 - Year Average		3 - Year Average		2 - Year Average		1 - Year Average	
	Coeff.[1]	SE.[1]	Coeff.	SE.	Coeff.	SE.	Coeff.	SE.	Coeff.	SE.
Father										
Measure[2]	0.2822***	0.0870	0.2841***	0.0866	0.2751***	0.0854	0.2441***	0.0841	0.1984**	0.0798
Age	-0.1100	0.0735	-0.1106	0.0734	-0.1072	0.0733	-0.1040	0.0738	-0.0968	0.0741
Age (squared)	0.0015*	0.0009	0.0016*	0.0009	0.0015*	0.0009	0.0015	0.0009	0.0014	0.0009
Son										
Age	0.2245**	0.1050	0.2249**	0.1049	0.2249**	0.1050	0.2160**	0.1054	0.2150**	0.1061
Age (squared)	-0.0029**	0.0014	-0.0029**	0.0014	-0.0029**	0.0014	-0.0028*	0.0014	-0.0028*	0.0014
F - Test	5.6816***		5.7364***		5.6521***		5.2225***		4.7309***	
Adjusted R^2	0.1156		0.1168		0.1150		0.1055		0.0944	
Observations	180		180		180		180		180	

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 earnings

Table 11: Detailed Results: Unbalanced Panel

	5 - Year Average		4 - Year Average		3 - Year Average		2 - Year Average		1 - Year Average	
	Coeff.[1]	SE.[1]	Coeff.	SE.	Coeff.	SE.	Coeff.	SE.	Coeff.	SE.
Father										
Measure[2]	0.2822***	0.0870	0.2867***	0.0815	0.2596***	0.0790	0.2076***	0.0695	0.2045***	0.0614
Age	-0.1100	0.0735	-0.0983	0.0754	-0.0725	0.0730	-0.0553	0.0744	-0.1027	0.0687
Age (squared)	0.0015*	0.0009	0.0014	0.0009	0.0010	0.0009	0.0008	0.0009	0.0014*	0.0008
Son										
Age	0.2245**	0.1050	0.2333**	0.1033	0.2433**	0.1011	0.2384**	0.0990	0.2647***	0.0904
Age (squared)	-0.0029**	0.0014	-0.0030**	0.0014	-0.0031**	0.0014	-0.0031**	0.0013	-0.0034***	0.0012
F - Test	5.6816***		6.5795***		6.1001***		5.9149***		7.1282***	
Adjusted R^2	0.1156		0.1286		0.1056		0.0982		0.1100	
Observations	180		190		217		227		249	

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 earnings

Table 12: 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						
Measure[2]	0.3241***	0.1139	0.3735**	0.1435	0.3365***	0.1269
Son						
Age	0.2149***	0.0478	0.1966*	0.1049	0.2290***	0.0828
Age (squared)	-0.0027***	0.0007	-0.0025*	0.0014	-0.0029***	0.0011
F - Test	34.6628***		4.6729***		5.7988***	
Adjusted R^2	0.2693		0.0717		0.0752	
Observations	282		180		240	

Basic Specification

Source: own calculations

Level of Significance: *** 1% **5% *10%

[1] Coeff. = Coefficient ; SE. = Standard Error

[2] years of education

IV

Table 13: Detailed Results: Relaxing Age - Restrictions for Fathers

	Max. Age 50		Max. Age 55		Max. Age 60		Max. Age 65	
	Coeff.[1]	SE.[1]	Coeff.	SE.	Coeff.	SE.	Coeff.	SE.
Father								
Measure[2]	0.2822***	0.0870	0.2509***	0.0794	0.3538***	0.0696	0.3584***	0.0686
Age	-0.1100	0.0735	0.0909*	0.0524	0.0488	0.0378	0.0536	0.0341
Age (squared)	0.0015*	0.0009	-0.0010*	0.0006	-0.0005	0.0004	-0.0006	0.0004
Son[3]								
Age	0.2245**	0.1050	0.2008**	0.0850	0.1377**	0.0635	0.1258**	0.0618
Age (squared)	-0.0029**	0.0014	-0.0026**	0.0011	-0.0017**	0.0008	-0.0016*	0.0008
F - Test	5.6816***		5.3447***		7.5867***		7.8214***	
Adjusted R^2	0.1156		0.0833		0.1052		0.1072	
Observations	180		240		281		285	

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 earnings

[3] son at least 30 years of age

Table 14: Detailed Results: Relaxing Age - Restrictions for Sons

	Min. Age 30		Min. Age 25		Min. Age 20	
	Coeff.[1]	SE.[1]	Coeff.	SE.	Coeff.	SE.
Father[3]						
Measure[2]	0.2822***	0.0870	0.2553***	0.0666	0.2402***	0.0558
Age	-0.1100	0.0735	-0.1139**	0.0477	-0.0806**	0.0373
Age (squared)	0.0015*	0.0009	0.0015**	0.0006	0.0011**	0.0005
Son						
Age	0.2245**	0.1050	0.2275***	0.0465	0.1227***	0.0248
Age (squared)	-0.0029**	0.0014	-0.0029***	0.0007	-0.0015***	0.0004
F - Test	5.6816***		24.2022***		48.0827***	
Adjusted R^2	0.1156		0.2922		0.3801	
Observations	180		282		385	

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 earnings

[3] father at most 50 years of age

Table 15: Robustness Check I: Analysis of Migrant Sample

	5 - Year Average		4 - Year Average		3 - Year Average		2 - Year Average		1 - Year Average	
	Coeff.[1]	SE.[1]	Coeff.	SE.	Coeff.	SE.	Coeff.	SE.	Coeff.	SE.
Father										
Measure[2]	0.4056***	0.1362	0.3598***	0.1326	0.3183**	0.1334	0.2808**	0.1241	0.2732**	0.1073
Age	0.2708	0.1518	0.2134	0.1466	0.1507	0.1233	0.1300	0.1192	0.0026*	0.0953
Age (squared)	-0.0033	0.0019	-0.0026	0.0018	-0.0018	0.0015	-0.0030	0.0014	-0.0033*	0.0011
Son										
Age	0.2556*	0.1411	0.2437*	0.1377	0.1949*	0.1128	0.2267**	0.1100	0.2442**	0.1076
Age (squared)	-0.0032*	0.0019	-0.0030	0.0019	-0.0024	0.0015	-0.0028*	0.0015	-0.0031**	0.0015
F - Test	4.26***		4.08***		3.27***		4.00***		4.20***	
Adjusted R^2	0.1504		0.1372		0.1275		0.1042		0.0981	
Observations	93		98		118		130		148	

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 earnings

Table 16: Robustness Check II: Including Younger Siblings

	5 - Year Average		4 - Year Average		3 - Year Average		2 - Year Average		1 - Year Average	
	Coeff.[1]	SE.[1]	Coeff.	SE.	Coeff.	SE.	Coeff.	SE.	Coeff.	SE.
Father										
Measure[2]	0.2767***	0.0833	0.2726***	0.0831	0.2639***	0.0827	0.2300***	0.0820	0.1780**	0.0784
Age	-0.1315*	0.0722	-0.1308*	0.0722	-0.1289*	0.0722	-0.1255*	0.0726	-0.1185	0.0730
Age (squared)	0.0018**	0.0009	0.0018*	0.0009	0.0018*	0.0009	0.0017*	0.0009	0.0016*	0.0009
Son										
Age	0.2518**	0.0992	0.2508**	0.0993	0.2509**	0.0994	0.2452**	0.0999	0.2440**	0.1005
Age (squared)	-0.0032**	0.0014	-0.0032**	0.0014	-0.0032**	0.0014	-0.0032**	0.0014	-0.0031**	0.0014
F - Test	6.7625***		6.7032***		6.5778***		6.0692***		5.4723***	
Adjusted R^2	0.1144		0.1134		0.1112		0.1021		0.0911	
Observations	224		224		224		224		224	

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 earnings

Table 17: Robustness Check III: Earnings Adjustment Using Growth Rate of Industry Wages

	5 - Year Average		4 - Year Average		3 - Year Average		2 - Year Average		1 - Year Average	
	Coeff.[1]	SE.[1]	Coeff.	SE.	Coeff.	SE.	Coeff.	SE.	Coeff.	SE.
Father										
Measure[2]	0.2806***	0.0873	0.2816***	0.0869	0.2720***	0.0857	0.2414***	0.0845	0.1976**	0.0803
Age	-0.1072	0.0734	-0.1078	0.0733	-0.1048	0.0733	-0.1021	0.0737	-0.0955	0.0741
Age (squared)	0.0015	0.0009	0.0015	0.0009	0.0015	0.0009	0.0014	0.0009	0.0014	0.0009
Son										
Age	0.2396**	0.1050	0.2399**	0.1050	0.2395**	0.1051	0.2308**	0.1055	0.2298**	0.1062
Age (squared)	-0.0031**	0.0014	-0.0031**	0.0014	-0.0031**	0.0014	-0.0030**	0.0014	-0.0030**	0.0014
F - Test	5.7383***		5.7774***		5.6818***		5.2625***		4.7995***	
Adjusted R^2	0.1169		0.1177		0.1157		0.1064		0.0959	
Observations	180		180		180		180		180	

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 earnings

Table 18: Robustness Check IV: 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						
Measure[2]	0.3809***	0.1116	0.3725***	0.1421	0.3223***	0.1219
Son						
Age	0.2029***	0.0480	0.2121*	0.1077	0.2421***	0.0851
Age (squared)	-0.0026***	0.0007	-0.0027*	0.0015	-0.0031***	0.0011
F - Test	34.3379***		4.7035***		5.9378***	
Adjusted R^2	0.2607		0.0717		0.0752	
Observations	273		174		231	

Source: own calculations

Level of Significance: *** 1% **5% *10%

[1] Coeff. = Coefficient ; SE. = Standard Error

[2] Wegener - Index of occupational prestige