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ABSTRACT

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This paper uses annual data drawn from the GSOEP to estimate individual earnings risk (labor market risk) in Germany for the period 1983-2012. The econometric specification of the earnings process allows for transitory shocks and permanent shocks to individual earnings. We find that both the transitory component and the permanent component of earnings risk have been rising in West Germany in the 1990s and have remained at elevated levels in the 2000s. In contrast, labor market risk in East Germany did not rise. These findings are robust to different sample selection criteria and changes in the specification of the earnings process. We provide a simple welfare calculation that suggests that the negative welfare consequences of the observed rise in the permanent component of earnings risk in West Germany are substantial. We argue that the time series evidence is not consistent with the view that international trade integration has been a main driver of the observed rise in labor market risk in West Germany.

JEL Classification: J31, D52

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1. Introduction

Labor market risk is a prevalent phenomenon with important economic implications. Specifically, labor market risk affects consumption and welfare of risk-averse workers. In addition, labor market risk has important macroeconomic implications through its effect on physical capital accumulation (Aiyagari, 1994), human capital accumulation (Krebs, 2003), and labor supply (Heathcote, Storesletten, and Violante, 2010). Finally, labor market risk interacts in interesting ways with labor market institutions (Ljungqvist and Sargent, 1998). A large literature has estimated individual wage and earnings risk in the US using panel data drawn from the PSID.[1] A common approach in the literature is to assume that wage/earnings risk has a transitory component as well as a persistent/permanent component. A common finding in the literature is that the variances of both transitory and permanent shocks to wages and earnings are large, and that both have risen in the 1980s in the US (Gottschalk and Moffitt, 1994, and Meghir and Pistaferri, 2004).

There is little work in the literature estimating individual wage or earnings risk for Germany that is comparable to the US studies.[2] This lack of comparable evidence is somewhat surprising given that Germany is the largest economy in Europe and has experienced a number of interesting labor market developments over the last 30 years. Specifically, German reunification occurred in 1990, international trade strongly increased from the middle of the 1990s until the onset of the Great Recession in 2008, and deunionization started in the middle of the 1990s and continued until the mid 2000s. Estimates of individual wage or earnings risk for Germany can potentially shed some light on the important question to what extent these developments have affected labor market risk. In this paper, we fill this gap in the literature. Specifically, we use data drawn for the GSOEP, a data set comparable in scope and design to the PSID, to estimate the transitory and permanent components of earnings risk in Germany over the period 1983-2012 thereby filling an important gap in the

[1] See, for example, Meghir and Pistaferri (2011) for a recent survey.

[2] Two noteworthy exceptions are Bayer and Juessen (2012) and Fuchs-Schuendeln, Krueger, and Sommer (2009), which we discuss in more detail below.

literature. In addition, we assess the welfare significance of our empirical findings using a simple consumption-saving model along the lines of Constantinides and Duffie (1994) and Krebs (2007).

We follow the bulk of the empirical literature on the topic and assume that individual log-earnings depend on a number of observable characteristics and a stochastic variable. Further, the stochastic variable is normally distributed and the sum of two components, a transitory component reflecting transitory earnings shocks and measurement error, and a persistent component reflecting persistent shocks to earnings. In line with several contributions to the literature, we further assume that the transitory component is an i.i.d. process and that the persistent component is fully permanent (random walk).[3] The respective variance parameters may depend on time and are interpreted as earnings risk. Finally, as in most of the literature on the topic, we estimate these variance parameters using selected moment conditions and the equally weighted minimum distance estimator.

Our main results are summarized in Figures 1 and 2. For West Germany, both the transitory component and the permanent component of individual earnings risk started to move up at the beginning of the 1990s reaching its peak around the end of the 1990s and followed by a slight decline of the transitory component. Further, the long-run levels of earnings risk in West Germany in the 2000s, transitory as well as permanent, are substantially higher than their respective levels at the beginning of the 1990s. The developments in East Germany are very different from the West German experience. Transitory as well as permanent earnings risk initially declined in the early 1990s and then remained roughly constant until the end of our sample period. These findings are robust to different sample selection criteria and changes in the specification of the earnings process.

From an economic point of view, the distinction between transitory and persistent income shocks is crucial, and a decomposition of observed income changes into individual components

[3]See section 3.1 for a detailed discussion of the various assumptions used in the literature. Note that in our robustness analysis we also consider an extension with a third component that follows an MA(1) process.

is therefore essential if the estimates are used as an input into economic models. However, one drawback of this approach is that the estimates often have high standard errors and heavily depend on the moment conditions used, which makes the precise detection of time trends somewhat difficult.[4] To confirm that the time trends in earnings risk depicted in Figures 1 and 2 are not spurious, we follow Guvenen, Ozkan, and Song (2014) and estimate the variance of the one-year changes and the variance of the five-year changes in (log)-earnings. This approach has the advantage that it delivers very precise estimates of earnings risk, but has the drawback that it does not allow for an unambiguous decomposition into transitory and persistent/permanent income shocks. Figures 3 and 4 show our estimates of the variances of one-year and five-year changes in log-earnings and confirm the main time trends depicted in Figure 1 and 2. Specifically, in West Germany the variances of both the one-year changes and the five-year changes of log-earnings started to move up at the beginning of the 1990s and then somewhat declined around the end of the 1990s. Further, these variances settled down at substantially elevated long-run levels in the 2000s. In contrast, the development in East Germany have been very different from the West German experience: the variances of both one-year changes and five-year changes initially declined and then remained roughly constant.

In addition to the empirical analysis, we provide a quantitative evaluation of the welfare consequences of the rise in labor market risk in West Germany in the 1990s based on a dynamic general equilibrium model with incomplete insurance markets along the lines of Constantinides and Duffie (1996) and Krebs (2007). The model allows for a closed-form solution for welfare as a function of the underlying parameters governing preferences and the earnings process. We find that the welfare cost of the rise in the permanent component of earnings risk in West Germany are substantial: 8.1 percent of lifetime consumption for moderate degree of risk aversion (log-utility). In comparison, the welfare cost of business cycles in the US are around two orders of magnitude smaller (Lucas, 2003). We note,

[4] This problem is well-known in the literature. See, for example, Meghir and Pistaferri (2011) for a thorough discussion.

however, that our welfare calculations do not take into account any endogenous change in insurance (Krueger and Perri, 2006) or aggregate productivity (Heathcote, Storesletten, Violante, 2010).

Our findings can in principle shed some light on the question to what extent international trade affects labor market risk. Specifically, the German economy experienced a strong and steady increase in international trade from the middle of the 1990s until the onset of the Great Recession in 2008 – see Figure 5. Popular discussions often suggest that this process of "globalization" might have led to heightened levels of risk in the labor market.[5] Our finding that individual earnings risk in West Germany has been declining or steady since the end of the 1990s (see Figure 1), a time when openness to international trade strongly increased, is difficult to reconcile with the idea that international trade has a significant positive effect on labor market risk.[6] Clearly, the timing of the increase in labor market risk in West Germany suggests that German reunification played an important role,[7] but the time series evidence presented here can only provide suggestive evidence. We leave for future research a deeper analysis of the link between labor market risk in Germany and various fundamental factors (international trade, technological progress, labor market institutions).

Bayer and Juessen (2012) and Fuchs-Schuendeln, Krueger, and Sommer (2009) are two important contributions that estimate earnings/wage risk in Germany following an approach similar to ours. Bayer and Juessen (2012) do not estimate time-dependent variance parameters and therefore cannot discuss the time trend in labor market risk, which is the main issue

[5] See Krebs, Krishna, and Maloney (2010) for a discussion of the theoretical literature on the link between trade and income risk.

[6] Our finding is, however, in line with the results reported in Krebs, Krishna, and Maloney (2010), who used Mexican data and found no evidence that trade liberalization has long-lasting effects on labor income risk. Thus, our paper provides additional evidence that there is no simple long-run relationship between globalization (openness to international trade) and individual labor market risk.

[7] Note that German reunification affected the West German labor market in various ways. For example, large net migration from East Germany to West Germany occured (Fuchs-Schuendeln, Krueger, and Sommer, 2009). Further, eastward expansion of the European Union allowed German firms to move production plans to close-by low-wage countries in Eastern Europe.

we address in this paper. Our paper is closely related to Fuchs-Schuendeln, Krueger, and Sommer (2009), who use an econometric methodology similar to ours to generate year-by-year estimates of transitory and persistent earnings risk. In line with the estimates reported in Fuchs-Schuendeln, Krueger, and Sommer (2009), we find that both the transitory and permanent component of earnings risk increased in the 1990s. Our work goes beyond the work by Fuchs-Schuendeln, Krueger, and Sommer (2009) in a number of ways. First, we provide separate estimates for West Germany and East Germany, a distinction that turns out to be crucial since only earnings risk in West Germany increased in the 1990s. Second, we have eight more years of observation available and therefore provide estimates of the various components of earnings risk up to 2008, which is important to infer that most of the increase of labor market risk that occurred in the 1990s was not reversed in the 2000s.[8] Third, we provide a number of checks that show that our results are robust to different sample selection criteria and different specifications of the earnings process. Finally, we provide a simple calculation that suggests that the estimated increase in labor market risk in West Germany is quantitatively important from a welfare point of view.

2. Data

Our data are drawn from the German Socio-Economic Panel (GSOEP). The GSOEP is an annual longitudinal survey of private households and individuals in Germany with a structure and design similar to the PSID, which is the data set mainly used in the literature to estimate individual income risk in the US. It was first conducted in 1984 in West Germany with about 4500 households and after German re-unification 2170 households from East Germany were included in 1991. In 1998 and again in 2000 refreshment samples were added increasing the total sample size substantially. We use data drawn from the waves from 1984 to 2013. Since the variables of interest (earnings) refer to the year prior to the survey this means that we have observations for the years 1983 to 2012.

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[8]Fuchs-Schuendeln, Krueger, and Sommer (2009) use a sample of annual earnings that ends in 2004 and their estimates of earnings risk go only up to 2002.

We include all subsamples in GSOEP except for the 'Oversampling of high income' subsample, which is eliminated to minimize potential selection bias. We focus on individuals who are head of households of age 25 to 60. Our measure of labor income is annual earnings of individuals, defined as the summation of wages and salaries from all employment including training, primary and secondary jobs and self-employment, plus income from bonuses, over-time, and profit sharing. We eliminate individuals whose hourly wage rate is below a certain threshold (less than 3 euros per hour) in order to minimize measurement error by ruling out implausibly low income. Lastly, self-employed head of households are not included since reported earnings of this group has a capital income component. In our robustness analysis we return to the issue of minimal wages and self-employment. We often split the sample into two subsamples for East Germany (the previous German Democratic Republic, GDR) and West Germany (the Federal Republic of Germany). Note that the first year of observation for East Germany is 1991.

3. Econometric Approach

3.1 Specification

The GSOEP provides us with annual earnings (labor income) of individuals. As in previous empirical work, we assume that the log of labor income of individual i in year t, $\ln y_{it}$, is given by:

$$\ln y_{it} = \alpha_t + X_{it}\beta_t + u_{it} \tag{1}$$

In equation (1), α_t and β_t denote time-varying coefficients and X_{it} is a vector of observable characteristics: gender, age, experience, education, and region. Note that speciation (1) restricts the return to observable characteristics to be the same across individuals, but allows these returns to vary over time. The variable u_{it} is the stochastic component of earnings and represents individual income changes that are unpredictable. We assume that this stochastic term is the sum of two components, a permanent component ω_{it} and a transitory component η_{it}

$$u_{it} = \omega_{it} + \eta_{it} \tag{2}$$

Shocks to ω_{it} are fully persistent in the sense that this component follows a random walk:

$$\omega_{i,t+1} = \omega_{it} + \epsilon_{i,t+1} \tag{3}$$

where the innovation terms, ϵ_{it}, are independently distributed over time and i.i.d. across individuals with $\epsilon_{i,t+1} \sim N(0, \sigma^2_{\epsilon,t+1})$. The transitory shocks have no persistence, that is, they are independently distributed over time. We further assume that they are i.i.d across individuals with $\eta_{it} \sim N(0, \sigma^2_{\eta,t})$. Clearly, η_{it} captures both temporary income shocks and measurement error.

Our specification for the labor income process is in accordance with the empirical work on US labor income risk. For example, Moffitt and Gottschalk (1994), Carroll and Samwick (1997), Gourinchas and Parker (2002), and Heathcote, Perri, and Violante (2010) estimate earnings risk based on PSID data and use exactly our specification. Further, Fuchs-Schuendeln, Krueger, and Sommer (2009) use the GSOEP data for Germany and Krebs, Krishna, and Maloney (2010) use Mexican data to estimate earnings risk based on the specification (1)-(3). Storesletten, Telmer and Yaron (2004) assume that the permanent component is an AR(1) process, but estimate an autocorrelation coefficient close to one (the random walk case). Finally, some papers have allowed for a third, MA(1), component (Meghir and Pistaferri, 2004). See Meghir and Pistaferri (2011) for a survey of the literature. In our robustness analysis we returns to this issue and introduce an additional MA(1) component as in Meghir and Pistaferri (2004).

The random walk assumption for ω is particularly attractive for understanding the fanning out of consumption over the life-cycle as first documented in Deaton and Paxson (1994). However, it has been questioned by some authors, most recently by Guvenen (2009), who shows that the estimation of the parameters of an earnings process with heterogeneity in income profiles suggest an autocorrelation coefficient for the persistent component substantially less than one (the random walk). In contrast, Baker and Solon (2003) using Canadian data find evidence in favor of a substantial random walk component and heterogeneity in growth rates. Further, Hryshko (2012) uses Monte Carlo simulations to show that the random walk hypothesis cannot be rejected in the PSID data, and finds little evidence for heterogeneous growth rates. Meghir and Pistaferri (2011) survey the literature and conclude that a (near) random walk component in earnings is not rejected by the existing evidence, though some conflicting results remain in the literature.

3.2 Estimation

Consider the change in the residual of earnings of individual i between year t and $t+n$:

$$\begin{aligned} \Delta_n u_{it} &= u_{i,t+n} - u_{it} \\ &= \epsilon_{i,t+1} + \ldots + \epsilon_{i,t+n} + \eta_{i,t} + \eta_{i,t+n} \end{aligned} \tag{4}$$

The variances of these n-year changes are given by

$$var\left[\Delta_n u_{it}\right] = \sigma^2_{\epsilon,t+1} + \ldots + \sigma^2_{\epsilon,t+n} + \sigma^2_{\eta,t} + \sigma^2_{\eta,t+n} \tag{5}$$

Following a number of contributions in the literature (see our discussion below), we us the moment restrictions (5) to estimate the risk parameters $\sigma^2_{\epsilon,t}$ and $\sigma^2_{\eta,t}$ using an equally weighted minimum distance estimator. Specifically, we choose the risk parameters $\sigma^2_{\epsilon,t}$ and $\sigma^2_{\eta,t}$ to minimize the distance function

$$\sum_{t,n} \left(v\hat{a}r[\Delta_n u_{it}] - (\sigma^2_{\epsilon,t+1} + \ldots + \sigma^2_{\epsilon,t+n} + \sigma^2_{\eta,t} + \sigma^2_{\eta,t+n}) \right)^2 \tag{6}$$

where $v\hat{a}r[\Delta_n u_{it}]$ are the corresponding sample variances. Note that in general we have more moment conditions than parameters and the system is therefore overidentified. The first-order conditions associated with the minimum-distance problem give rise to a linear equation system for the estimates $\hat{\sigma}^2_{\epsilon,t}$ and $\hat{\sigma}^2_{\eta,t}$.

Some intuition for the way our approach separates permanent from transitory shocks can be obtained by looking at the case of time-independent risk parameters $\sigma^2_{\epsilon,t}$ and $\sigma^2_{\eta,t}$. In this case, the moment restrictions (5) become

$$var[\Delta_n u_{it}] = 2\sigma^2_\eta + \sigma^2_\epsilon \, n \tag{7}$$

Thus, the variance of observed n-year changes in earnings is a linear function of n, where the slope coefficient is equal to σ_ϵ^2. Further, our minimum-distance estimator is obtained by OLS regression of the sample variances $v\hat{a}r[\Delta_n u_{it}]$ on n.

The moment conditions we select in (5) to estimate the parameters of the earnings process are in line with several contributions in the literature. Specifically, Carroll and Samwick (1997), Deaton and Paxson (1994), Gottschalk and Moffitt (1994), and Krebs, Krishna, and Maloney (2010) all use, as we do here, the variances of n-year earnings changes as moment restrictions. In contrast, a number of papers have worked with the auto-covariances of earnings levels or one-year earnings changes as moments restrictions (Meghir and Pistaferri, 2004, and the contributions to the special issue in the Review of Economic Dynamics surveyed in Krueger et al., 2009). In particular, Fuchs-Schuendeln, Krueger and Sommers (2009) present two estimates of the earnings process (1)-(3) based on GSOEP data and two different sets of moment conditions: i) the auto-covariances of in earnings levels and ii) the auto-covariances of earnings changes. They find that the estimates of the variances of permanent shocks is unreasonably large when using the first-difference auto-covariance restrictions. In contrast, their estimates of permanent earnings risk based on auto-covariance restriction of earnings levels are in line with the estimates we obtain.

4. Results

4.1 Earnings Risk: Time-Average

We begin with a brief discussion of the estimation results averaged over time. Table 1 summarizes the time-averages of the estimates $\sigma_{\epsilon,t}^2$ and $\sigma_{\eta,t}^2$. We emphasize three results.

First, the average value for permanent labor income risk, σ_ϵ, is about 11% and the average value for transitory labor income risk (plus measurement error), σ_η^2, is 30%. These findings are in line with the results reported by Fuchs-Schuendeln et al. (2009) when they estimate earnings risk using covariance restrictions based on level variables. For the US, estimates of σ_ϵ cluster around 15%. Thus, permanent labor market risk in Germany is somewhat lower than in the US. This finding seems intuitive given that wage volatility in Germany can be expected to be smaller than in the US and keeping in mind that our sample selection excludes most long-term unemployed.

Our second result is that labor market risk in West Germany is smaller than in East Germany, though the difference with respect to permanent income risk is small. Finally, we find that females face significantly more labor market risk than males. The literature has found that females have different labor participation decisions compared to males and that their labor supply decision is more sensitive to their marital status, the presence of young children, home productivity, tax incentives etc. All these facts point towards higher income volatility for females.

4.2 Earnings Risk: Time Series Evidence

Figure 1 shows the time series of our estimates of the transitory component, $\sigma^2_{\eta,t}$, and the permanent component, $\sigma^2_{\epsilon,t}$, of earnings risk in West Germany. The hollow dots represent the estimated annual variances of the transitory component, the solid dots the estimated variances of the permanent component, and the dashed and solid lines are the corresponding HP-filtered trend components. For the transitory component we observe a gradual increase that accelerates around 1990, the time of re-unification. Transitory earnings risk reaches its peak around the end of the 1990s and then declines to a long-run level that is significantly higher than the value at the beginning of our sample period. The permanent component of earnings risk is constant at the beginning of our sample period, begins to increase around 1990 until the end of the 1990s, then declines slightly and finally settles down at an elevated level until the end of our sample period.

Figure 2 shows the time series of our estimates of the transitory component, $\sigma^2_{\eta,t}$, and the permanent component, $\sigma^2_{\epsilon,t}$, of earnings risk in East Germany. The developments in East Germany are very different from the West German experience. Specifically, in East Germany transitory as well as permanent earnings risk initially declined in the early 1990s and then remained roughly constant until the end of our sample period.

In the line with previous results in the literature, we note that individual variances of the permanent and transitory component are estimated with substantial stand errors (see Table A3 in the Appendix). In Figures 3 and 4 we therefore also show the variances of the one-year change, $var[\Delta_1 u_{it}]$, and five-year changes, $var[\Delta_5 u_{i,t}]$, which are estimated with negligible standard error given the large number of individual observations i for any t. Figure 3 provides evidence in support of the general pattern depicted in Figure 1. Specifically, in West Germany both the variance of the one-year change in earnings and the variance of the five-year change in earnings start to rise around 1990 until reaching their respective peaks at the end of the 1990s. Both variances then decline slightly, but stay at elevated levels until the end of our sample period. Note that the increase in the variance of the five-year difference is substantially larger than the increase in the variance of the one-year difference (in absolute terms), which suggests that a substantial part of the increase in the variance of the five-year difference is due to an increase in the variance of persistence shocks. Figure 4 corroborates the trend for East Germany depicted in Figure 2: transitory as well as permanent earnings risk initially declined in the early 1990s and then remained roughly constant until the end of our sample period.

Figure 6 shows our estimates of the transitory and permanent component of earnings risk for the subsample of all males in West Germany. The results in Figure 6 are very similar to the results shown in Figure 1. Figure 7 shows the estimates of earnings risk for the subsample of all females in West Germany. In this case the general pattern is also broadly consistent with the pattern depicted in Figure 1, but the risk parameters are estimated with much less precision because of the small sample size and the pattern is therefore less clear. Overall, Figures 6 and 7 support the view that for both men and women in West Germany labor market risk has been strongly rising between the beginning of the 1990s and the end of the 1990s, after which it declined for a few years until settling down at a new long-run level that is significantly higher than its level in 1990.

4.3 International Trade in Germany

The time series evidence depicted in Figure 1 is difficult to reconcile with the view that openness to international trade has permanent effects on labor market risk. Specifically, the German economy has experienced a steady and strong increase in international trade starting in the middle of the 1990s until the onset of the Great Recession in 2008 – see Figure 5. The fact that labor market risk in Germany has been declining at a time when international trade was growing suggests that gradual changes in international trade have no substantial effects on labor market risk. This finding is in line with the results reported in Krebs, Krishna, and Maloney (2010) using Mexican data, who found no evidence that trade liberalization has long-lasting effects on labor income risk.[9] In other words, neither σ_η^2 nor σ_ϵ^2 were permanently affected by the dramatic trade liberalization that occured in Mexico at the end of the 1980s, and the results presented here suggest that the rapid trade expansion that occured in Germany in the period 1995-2005 had also no significant impact on these risk parameters.

In addition to the rapid trade expansion, in the mid 1990s a process of deunionization began in Germany that continued until the mid 2000s. See Dustmann, Ludsteck, and Schoenberg (2009) for the evidence of on deunionization in Germany.[10] If a decline in union coverage makes wages more flexible, then this could lead to an increase in labor market risk. Thus, deunionization could in principle be one factor driving the rise in labor market risk in West Germany in the 1990s, but the fact that earnings risk in West Germany has been declining or steady since the end of the 1990s, a time when union coverage was still decreasing, suggests that union coverage is at best part of the explanation of the pattern

[9] Krebs, Krishna, and Maloney (2010) find, however, that the dramatic trade liberalization in Mexico at the end of the 1980s led to a temporary increase in σ_η^2 and σ_ϵ^2. This result suggests that episodes of abrupt structural change require an adjustment of the labor market that generates a short-run rise in labor market risk.

[10] Dustmann, Ludsteck, and Schoenberg (2009) also show that this decline in union coverage has increased wage inequality. The concept of wage inequality is related to our concept of wage/earnings risk, but they are distinct concepts and wage inequality can increase without a corresponding increase in wage risk.

depicted in Figures 1 and 3.

Clearly, the timing of the increase in labor market risk in West Germany suggests that German reunification played an important role,[11] but the time series evidence presented here can only provide suggestive evidence. We leave for future research a deeper analysis of the issue of the link between labor market risk in Germany and various fundamental factors (international trade, technological progress, labor market institutions).

4.4 Robustness

We now turn to the discussion of the robustness of our main results to a number of changes in data selection. First, we drop the minimum threshold for the hourly wage rate, that is, we include all head of households who are not self-employed and between the age of 25 and 60 regardless of their level of earnings. The estimation results for this case are shown in Figures 8 and 9 for the West German sample, respectively East German sample. A comparison of Figures 8 and 9 with the corresponding Figures 1 and 2 shows that this change in sample selection has almost no effect on the estimates of the transitory and permanent component of labor income risk. We found similar results when we considered the variances of the one-year changes and 5-year changes. Thus, we conclude that the choice of a lower bound on the wage rate has negligible effects on our estimation results.

So far, we used a sample selection that excluded all self-employed. Clearly, the income of the self-employed includes an component that is earnings, and in this sense the self-employed also face earnings risk. We therefore re-estimate the transitory and permanent component of earnings risk with a sample that also includes the self-employed. The estimation results for this case are shown in Figures 10 and 11 for the West sample, respectively East sample. A comparison of Figures 10 and 11 with the corresponding Figures 1 and 2 shows that this change in sample selection has only small effects on the estimates of the transitory and

[11] Note that German reunification affected the West German labor market in various ways. For example, large net migration from East Germany to West Germany occured (Fuchs-Schuendeln, Krueger, and Sommer, 2009). Further, eastward expansion of the European Union allowed German firms to move production plans to close-by low-wage countries in Eastern Europe.

permanent component of labor income risk. We found similar results for the variances of the one-year changes and 5-year changes. Thus, we conclude that the inclusion of income from self-employment into our analysis has negligible effects on our estimation results.

Finally, we consider a generalization of the stochastic process governing the earnings dynamics that allows for a third MA-component. The details of the estimation approach with an additional MA-component are discussed in the Appendix. Figure 12 depicts the estimation results for West Germany. The results shown in Figure 12 confirm the general pattern shown in Figure 1: both transitory and permanent component of earnings risk rise between the beginning of the 1990s and the end of the 1990s, then decline slightly to settle down at elevated levels. The main difference between Figure 1 and Figure 12 is that the average values of transitory risk (i.i.d. component) and permanent risk (random walk component) are somewhat higher in Figure 1 than in Figure 12. This is not surprising since part of the observed earnings changes are picked up by the MA-component leaving less for the i.i.d. component and the random walk component.

5. Earnings Risk and Welfare

In this section, we provide use a simple consumption-saving model that allows us to link changes in earnings risk to changes in welfare. The voluminous literature on consumption and saving with individual income risk and incomplete insurance markets has generated a number of insights.[12] One important insight is that workers can effectively self-insure against transitory income shocks through borrowing or own saving, and that the effect of these shocks on equilibrium prices and quantities are relatively small.[13] A second important insight of this literature is that very persistent or fully permanent income shocks have substantial effects on consumption and welfare even if individual households have own savings, but no or only limited access to insurance markets. Indeed, when labor income is the main source of income

[12]See, for example, Heathcote, Storesletten, and Violante (2009) for a recent survey.

[13]See, for example, Aiyagari (1994) for quantitative work and Levine and Zame (2002) for a theoretical argument.

and labor income shocks are highly persistent, we would expect that consumption responds (almost) one-for-one to labor income shocks. This point has been made more formally by Constantinides and Duffie (1996) using dynamic general equilibrium exchange models with incomplete markets and used in Krebs (2007) to analyze the link between earnings risk and welfare.[14] Section 5.1 briefly discusses the model and its welfare implications and section 5.2 uses the resulting welfare formula to provide a quantitative assessment of the welfare impact of the rise in earnings risk that occured in West Germany in the 1990s.

5.1 Model Outline

The model features long-lived, risk-averse workers with homothetic preferences who make consumption/saving choices in the face of uninsurable income shocks. Workers' preferences over consumption plans, $\{c_{it}\}$, allow for a time-additive expected utility representation with one-period utility function of the CRRA-type, where in this paper we confine attention to the log-utility case (degree of relative risk aversion of one):

$$U(\{c_{it}\}|\omega_{i0}) = E\left[\sum_{t=0}^{\infty} \beta^t ln c_{it}|\omega_{i0}\right] . \tag{8}$$

Workers maximize expected lifetime utility subject to a sequential budget constraint that allows them to transfer wealth across periods through saving (or borrowing). The model is an exchange economy with endogenous interest rate (general equilibrium).

Suppose that workers face a labor income process that only contains the permanent component ω_{it}, that is, the ex-ante heterogeneity X_{it} in (1) and the transitory component η_{it} in (2) have been removed.[15] Assume further that the distribution of the innovation term, ϵ, and the distribution of initial income, ω_0, include a mean-adjustment: $\epsilon \sim N(-\sigma_\epsilon^2/2, \sigma_\epsilon^2)$ and $\omega_0 \sim N(-\sigma_{\omega_0}^2/2, \sigma_{\omega_0}^2)$. This adjustment is common in the literature and necessary to

[14] See Blundell, Pistaferri, and Preston (2008) for empirical evidence that, in line with the model outlined here, permanent income shocks have large effects on individual consumption.

[15] Clearly, the part of η that captures measurement error should not enter into the worker's budget constraint and this part of η should therefore be omitted. Further, previous work in the literature suggests that the effect of transitory income shocks on consumption and welfare is small and can be neglected as a first approximation – see also footnote 13.

ensure that σ_ϵ^2 and $\sigma_{\omega_0}^2$ can be interpreted as uncertainty parameters, which ensures that our welfare results are sensible (see below).

We assume that all workers have the opportunity to borrow or save at the common interest rate r. If the risk-free asset is in zero net supply, then the equilibrium interest rate will adjust so that individual workers will optimally decide to set consumption equal to labor income: $c_{it} = \omega_{it}$. See Constantinides and Duffie (1996) and Krebs (2007) for details of the formal arguments. In other words, in the equilibrium of this stylized model there is no self-insurance against permanent income shocks – consumption and income move one-to-one.

5.2 Welfare

Using $c_{it} = \omega_{it}$ and the law of motion for ω_{it} given in (3) we can evaluate the expected lifetime utility (8) of an individual with initial income ω_{i0}. Taking the expectation over ω_{i0} yields social welfare, W, where we assume that each individual household is assigned equal weight in the social welfare function. In other words, social welfare is the expected lifetime utility from an *ex ante* point of view when the initial condition, ω_0, is not yet known (veil of ignorance). More formally, we have

$$\begin{aligned} W &= E\left[\sum_{t=0}^{\infty}\beta^t ln c_{it}\right] \qquad (9) \\ &= E\left[E\left[\sum_{t=0}^{\infty}\beta^t ln c_{it}|\omega_0\right]\right] \\ &= E\left[-\frac{\beta}{(1-\beta)(1-\beta)}\frac{\sigma_\epsilon^2}{2}+\frac{1}{1-\beta}\omega_0\right] \\ &= -\frac{\beta}{(1-\beta)(1-\beta)}\frac{\sigma_\epsilon^2}{2}-\frac{1}{1-\beta}\frac{\sigma_{\omega_0}^2}{2} \end{aligned}$$

The formula (9) shows how social welfare depends on the income parameters σ_ϵ^2 and $\sigma_{\omega_0}^2$ and the preference parameter β. In particular, (9) shows that an increase in labor market risk, σ_ϵ^2, reduces welfare. In order to express welfare changes in economically meaningful units, we calculate the corresponding change in consumption in each period and possible future state that is necessary to compensate the worker for the change in uncertainty. For example, suppose we compare two economies, one with risk parameter σ_ϵ^2 and one with risk parameter $\hat{\sigma}_\epsilon^2$. We then define the consumption-equivalent welfare change, Δ, of moving from an economy with σ_ϵ^2 to an economy with $\hat{\sigma}_\epsilon^2$ as

$$E\left[\sum_{t=0}^{\infty} \beta^t ln\left(c_{it}(1+\Delta)\right)\right] = E\left[\sum_{t=0}^{\infty} \beta^t ln\hat{c}_{it}\right] , \tag{10}$$

where c is consumption in the first economy and $\hat{c}$ is consumption in the second economy. Using the definition (10) and the welfare formula (9), we find:[16]

$$\ln(1+\Delta) \quad = \quad \frac{\beta}{1-\beta}\left(\frac{\hat{\sigma}_\epsilon^2}{2} - \frac{\sigma_\epsilon^2}{2}\right) + \left(\frac{\hat{\sigma}_{\omega_0}^2}{2} - \frac{\sigma_{\omega_0}^2}{2}\right) \tag{11}$$

Equation (11) provides a convenient formula to translate changes in labor market risk, $\hat{\sigma}_\epsilon^2 - \sigma_\epsilon^2$, into changes in welfare, Δ. Note that these welfare changes depend on only two factors: the change in income risk and the discount factor β (the degree of risk aversion is fixed to one).

5.3 Application

We now apply the formula (11) to the West German experience in the 1990s. In order to do so, we first infer from Figure 1 a long-run change in permanent labor market risk, σ_ϵ^2, and then compute the corresponding welfare change using formula (11). The long-run change in labor market risk is computed as follows. We first take the average of the estimates of σ_ϵ^2 in the 1990s and the average of the estimates of σ_ϵ^2 in the 2000s, and then compute the difference between these two averages. This yields $\bar{\sigma}_{\epsilon,2000s}^2 - \sigma_{\epsilon,1990s}^2 = 0.018$. Note that we find a very similar increase in permanent earnings risk if we compare the HP-trend component at the beginning of the 1990s with the value of the HP-trend component in the 2000s.

To use the welfare formula (11), we need to specify a value for the discount factor β. The quantitative macroeconomic literature often uses a value of $\beta = 0.96$ for the annual discount factor. However, one could argue that many workers are less patient, specifically those worker with little or no financial wealth. In addition, introducing finite lifes of household

[16]For general CRRA utility function with degree of relative risk aversion γ the corresponding formula can be found in Krebs (2007).

(OLG model) would reduce the effective discount factor. Thus, a value of $\beta = 0.90$ seems more reasonable for our welfare analysis based on the model outlined above.

Using $\beta = 0.9$, we find that the welfare cost of the increase in labor market risk observed in West Germany, Δ, is equal to 8.1 percent of lifetime consumption. For $\beta = 0.96$, we find a welfare cost of 21.6 percent of lifetime consumption. Thus, for a wide range of values of household impatience, the welfare costs associated with the observed rise in labor market risk in West Germany are very large. In comparison, Lucas (2003) suggests that the welfare cost of business cycles is less than 0.1 percent of lifetime consumption, which is two orders of magnitude smaller than the values we find here. Note that our computations assume that the degree of relative risk aversion is one (log-utility), which is quite moderate. In addition, we find that the welfare cost of the rise in earnings risk strongly depends the value of β, which is not surprising given that the earnings shocks considered here are fully permanent.

Appendix

A. Descriptive Statistics and First-Stage Regression

Table A1 displays the descriptive statistics of variables in our sample, and Figures A1 and A2 show the time trend of means and medians of real and nominal earnings in our sample. We also compare these means and medians from our sample to the real and nominal earnings reported by the Federal Statistical Office in Germany. In Figures A1 and A2, the solid line represents means of real and nominal monthly labor income, the dashed line shows medians of real and nominal monthly labor income, the dash-dot line is the plot of average real and nominal monthly labor income reported by the Federal Statistical Office starting from early 1990s. As can be seen from the plots, our sample shows time trends in mean and median similar to the trends shown by the aggregate data from the German Federal Statistical Office. Specifically, average real earnings dropped after reunification and there was a further decline after the year 2000. Average nominal earnings also show a drop after reunification, grow smoothly during 1990s, and remained roughly constant in the first years of 2000.

Table A2 shows the regression results from the first stage Mincerian wage equation regression in the year 2005, from which the corresponding residuals are generated. The regression results in the other years look virtually the same, in the sense that all coefficients have the expected signs and significance. We see from table A2 that individual labor income grows with age, education and family size.

In Table A3 we report our y estimated variances of transitory and permanent components of residual incomes, and the corresponding standard errors, for all individuals in West Germany. .

B. Earnings process with MA(1) Component

There is empirical evidences that suggests the presence of a third component in labor income process. For example, Meghir and Pistaferri (2004) use a specification of the earnings process with an i.i.d. component, a random walk component, and an MA component. In line with this approach, we consider an earnings process for which u_{it} is equation (2) is replaced by

$$u_{it} = \omega_{it} + m_{it} + \eta_{it} \tag{12}$$

where as in the previous analysis η_{it} is a normally distributed i.i.d. component and ω_{it} is a random walk component with normally distributed innovation term. In addition, there is a MA component, m_{it}, given by

$$m_{it} = r_{it} + \theta r_{i,t-1} \tag{13}$$

where the shocks r_{it} are normally distributed and i.i.d.

To derive the relevant moment conditions, note that

$$u_{it} = \omega_{it} + r_{it} + \theta r_{i,t-1} + \epsilon_{it} \tag{14}$$

$$u_{i,t+1} = \omega_{it} + \epsilon_{it+1} + r_{i,t+1} + \theta r_{it} + \eta_{i,t+1}$$

Thus, we have

$$\triangle_1 u_{it} = u_{i,t+1} - u_{it} = \epsilon_{i,t+1} + r_{i,t+1} + (\theta - 1) r_{it} - \theta r_{i,t-1} + \eta_{i,t+1} - \eta_{it} \tag{15}$$

and

$$\triangle_n u_{it} = u_{it+n} - u_{it} = \epsilon_{i,t+1} + ... + \epsilon_{i,t+n} + r_{i,t+n} + \theta r_{i,t+n-1} - r_{it} - \theta r_{i,t-1} + \eta_{i,t+n} - \eta_{it}.$$

Therefore, the moment conditions read

$$V(\triangle_1 u_{it}) = \sigma^2_{\epsilon,t+1} + (\theta - 1)^2 \sigma^2_{r,t} + \theta^2 \sigma^2_{r,t-1} + \sigma^2_{\eta,t+1} + \sigma^2_{\eta,t} \tag{16}$$

and

$$V(\triangle_n u_{it}) = \sigma^2_{\epsilon,t+1} + ... + \sigma^2_{\epsilon,t+n} + \sigma^2_{r,t+n} + \theta^2 \sigma^2_{r,t+n-1} + \sigma^2_{r,t} + \theta^2 \sigma^2_{r,t-1} + \sigma^2_{\eta,t+n} + \sigma^2_{\eta,t}.$$

The moment conditions (16) define an over-identified equation system with parameters $\{\sigma^2_{\eta,t}, \sigma^2_{\epsilon,t}, \sigma^2_{r,t}, \theta\}$. Therefore, we apply nonlinear least square to generate the consistent estimates of parameters $\{\sigma^2_{\eta,t}, \sigma^2_{\epsilon,t}, \sigma^2_{r,t}\}$. In order to simplify our estimation approach, we take the value of θ as given. Biewen (2005) shows that θ is in the range from -0.5 to -0.3 and Myck et al. (2008) estimates that θ is with value in the range from -0.5 to -0.4. Based on these results, we use $\theta = -0.4$.

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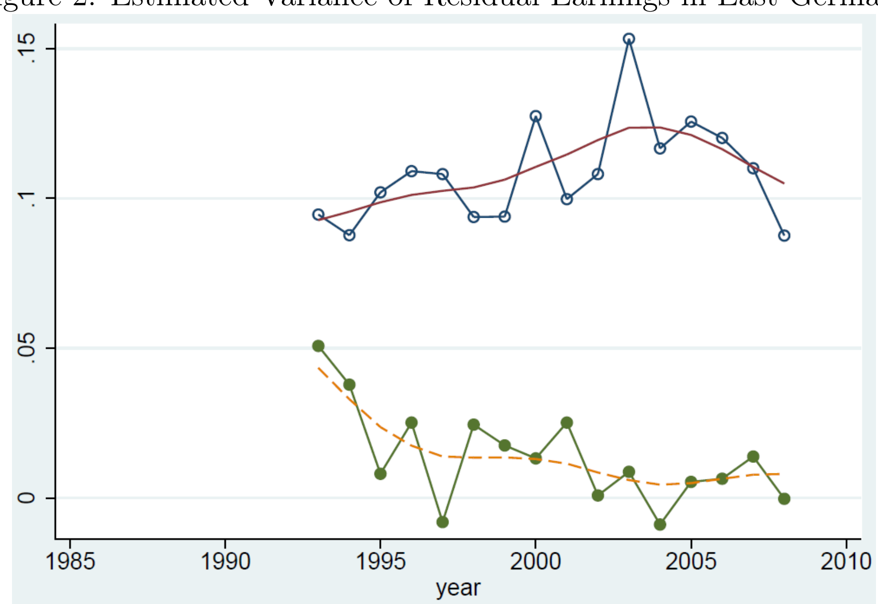
Figure 1: Estimated Variance of Residual Earnings in West Germany

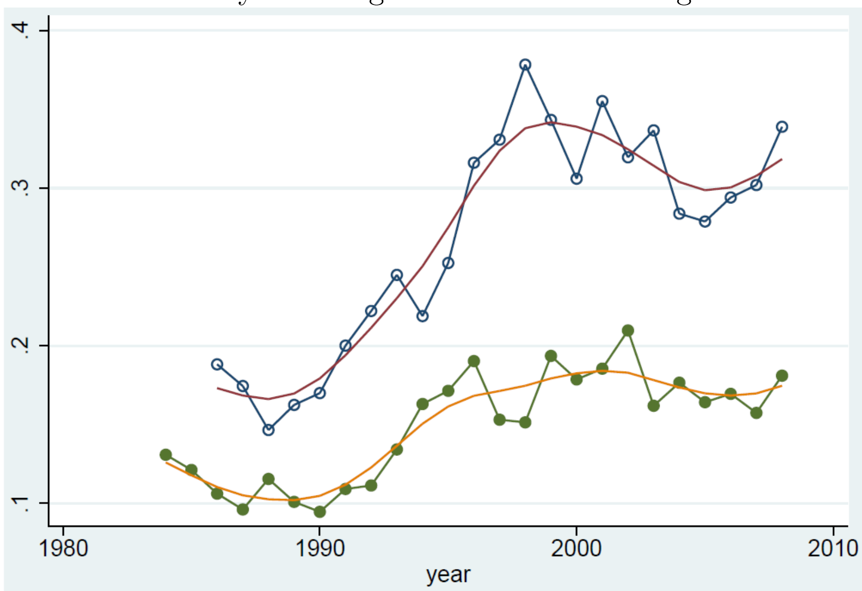
Note: The hollow dots represent the estimated variance of transitory components, solid dots are the estimated yearly variance of permanent components, and the solid and dashed lines are the corresponding HP filtered trend of transitory and permanent components estimates.

Figure 2: Estimated Variance of Residual Earnings in East Germany



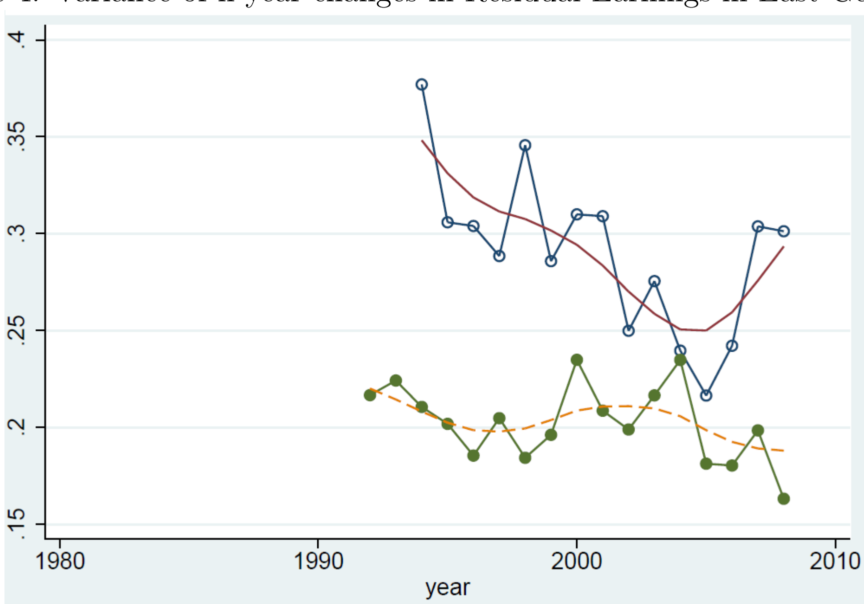
Note: The hollow dots represent the estimated variance of transitory components, solid dots are the estimated yearly variance of permanent components, and the solid and dashed lines are the corresponding HP filtered trend of transitory and permanent components estimates.

Figure 3: Variance of n-year changes in Residual Earnings in West Germany



Note: The solid and hollow dots show correspondingly the estimated variance of one-year and 5-year differences of residual incomes generated from the first stage regression.

Figure 4: Variance of n-year changes in Residual Earnings in East Germany



Note: The solid and hollow dots show correspondingly the estimated variance of one-year and 5-year differences of residual incomes generated from the first stage regression.

Figure 5: Export-to-GDP and Import-to-GDP Ratio in Germany

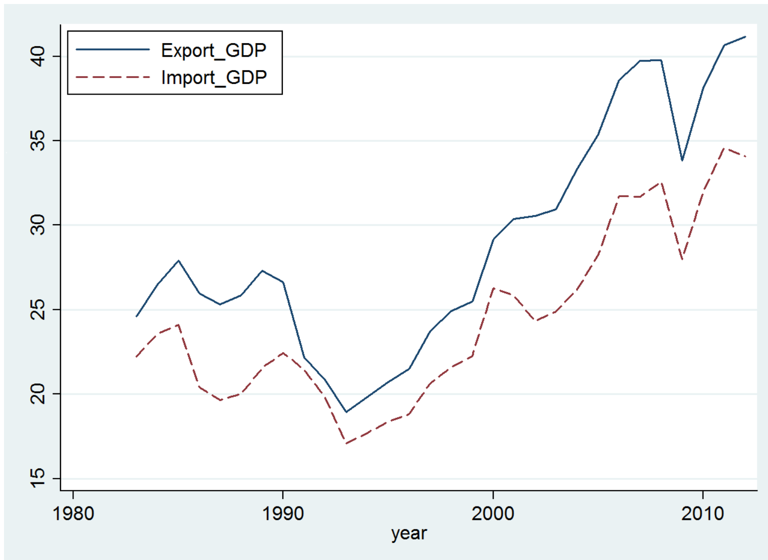
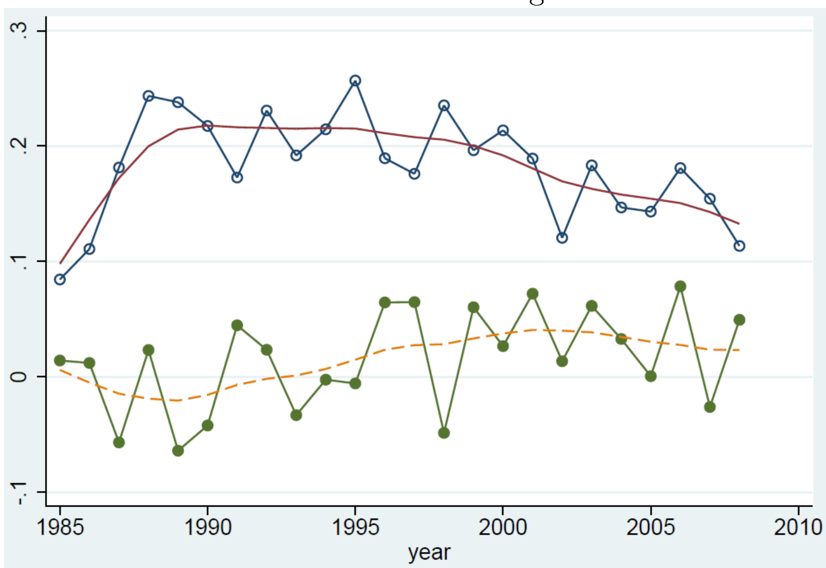


Figure 6: Estimated Variance of Residual Earnings for Males in West Germany

Note: The hollow dots represent the estimated variance of transitory components, solid dots are the estimated yearly variance of permanent components, and the solid and dashed lines are the corresponding HP filtered trend of transitory and permanent components estimates.

Figure 7: Estimated Variance of Residual Earnings for Females in West Germany

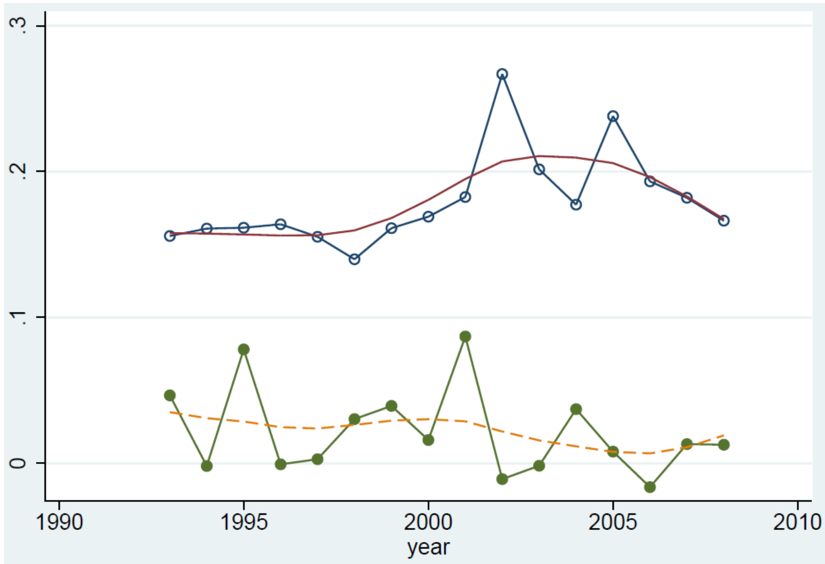


Note: The hollow dots represent the estimated variance of transitory components, solid dots are the estimated yearly variance of permanent components, and the solid and dashed lines are the corresponding HP filtered trend of transitory and permanent components estimates.

Figure 8: Estimated Variance of Residual Earnings in West Germany - no lower bound on wages

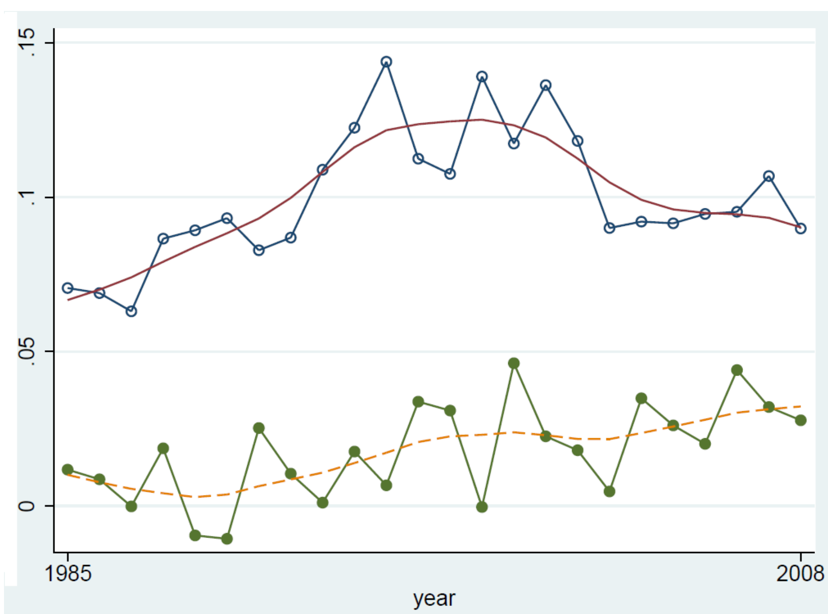
Note: The hollow dots represent the estimated variance of transitory components, solid dots are the estimated yearly variance of permanent components, and the solid and dashed lines are the corresponding HP filtered trend of transitory and permanent components estimates.

Figure 9: Estimated Variance of Residual Earnings East Germany - no lower bound on wages



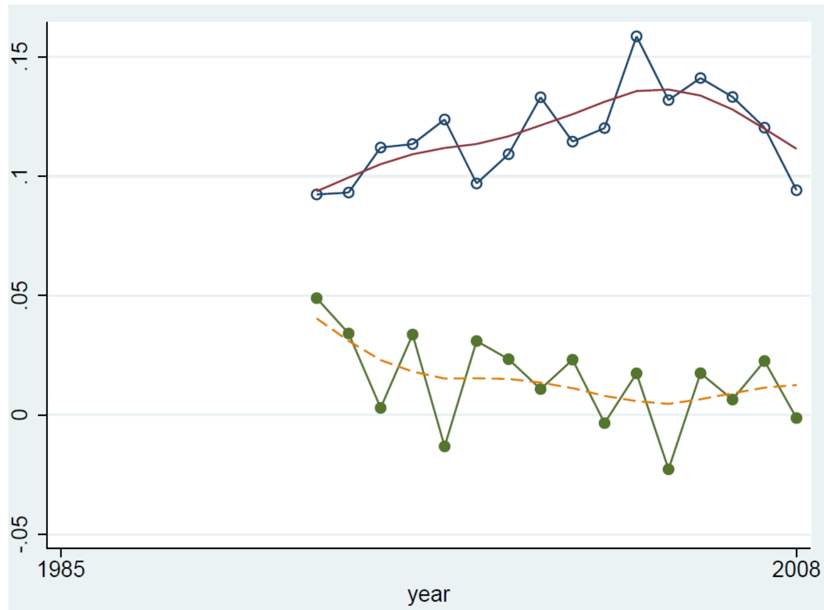
Note: The hollow dots represent the estimated variance of transitory components, solid dots are the estimated yearly variance of permanent components, and the solid and dashed lines are the corresponding HP filtered trend of transitory and permanent components estimates.

Figure 10: Estimated Variance of Residual Earnings in West Germany including self-employment



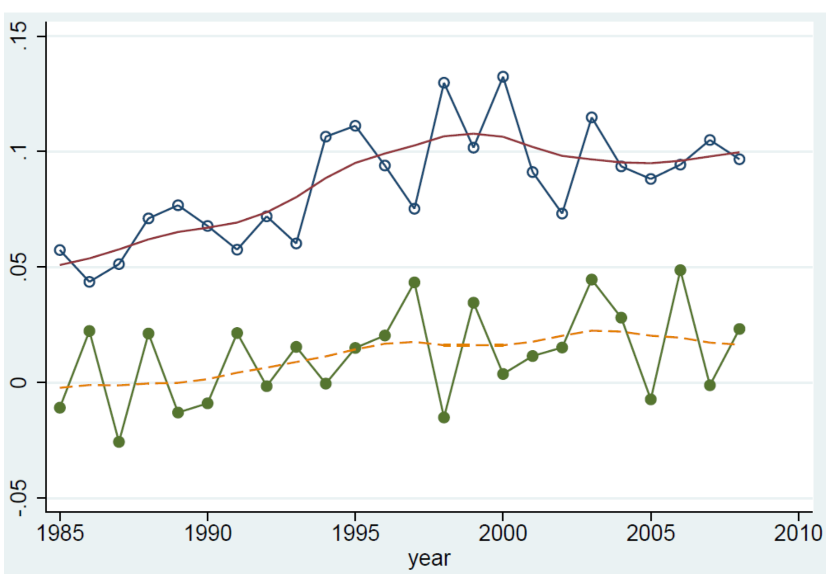
Note: The hollow dots represent the estimated variance of transitory components, solid dots are the estimated yearly variance of permanent components, and the solid and dashed lines are the corresponding HP filtered trend of transitory and permanent components estimates.

Figure 11: Estimated Variance of Residual Earnings in East Germany including self-employment



Note: The hollow dots represent the estimated variance of transitory components, solid dots are the estimated yearly variance of permanent components, and the solid and dashed lines are the corresponding HP filtered trend of transitory and permanent components estimates.

Figure 12: Estimated Variance of Residual Earnings with MA-Component in West Germany

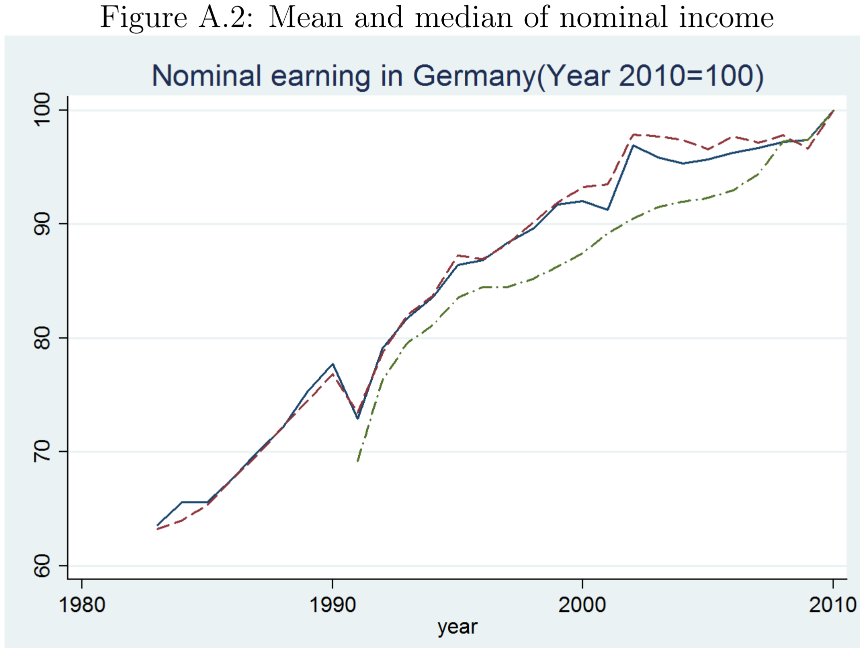


Note: The hollow dots represent the estimated variance of transitory components, solid dots are the estimated yearly variance of permanent components, and the solid and dashed lines are the corresponding HP filtered trend of transitory and permanent components estimates.

Figure A.1: Mean and median of real income

Note: the solid line represents connected means of real monthly labor income, the dashed line shows connected medians of real monthly labor income, the dash-dot line is the plot of average real monthly labor income reported by the German Federal Statistical Office.

Figure A.2: Mean and median of nominal income



Note: the solid line represents connected means of nominal monthly labor income, the dashed line shows connected medians of nominal monthly labor income, the dash-dot line is the plot of average nominal monthly labor income reported by the German Federal Statistical Office.

Table 1: Average of Estimated Variances of Residual Earnings

Region	Subgroups	Average estimates of transitory risk	Average estimates of permanent risk
All	all	0.3038	0.1078
West	Female	0.3620	0.1208
	Male	0.2416	0.1031
East	Female	0.4185	0.1625
	Male	0.3064	0.1048

Appendix

A Descriptive Statistics and First Stage Regression

Table A.1: Descriptive Statistics for Household Sample

Variable	Mean	(Std. Dev.)	Min	Max
Family size	2.831017	1.338483	1	14
Years of education	12.09746	2.709664	7	18
Ind. working hrs	2043.418	642.8531	13	5965
Ind. earning	30069.63	20386.59	56.5213	1774492

Table A.2: First-Stage Wage Regression for Year 2005

ln individual labor income	coefficients
Age	0.1076 ***
	(0.0095)
Age-squared	-0.1069***
	(0.0110)
Family size	0.0311***
	(0.0079)
Years of education	0.0987***
	(0.0034)
Constant	5.8062***
	(0.1996)
Observations	5623
R-squared	0.3015

Table A.3: Estimated Variance of Residual Earnings in West Germany

Year	Estimates of transitory components	(standard errors)	Estimates of permanent components	(standard errors)
1985	0.064	(0.012)	0.000	(0.021)
1986	0.053	(0.010)	0.021	(0.016)
1987	0.054	(0.009)	-0.009	(0.013)
1988	0.075	(0.008)	0.013	(0.012)
1989	0.084	(0.008)	-0.012	(0.011)
1990	0.081	(0.007)	-0.011	(0.010)
1991	0.065	(0.007)	0.027	(0.010)
1992	0.085	(0.007)	0.000	(0.010)
1993	0.086	(0.007)	0.003	(0.009)
1994	0.117	(0.007)	0.007	(0.009)
1995	0.136	(0.007)	0.007	(0.009)
1996	0.101	(0.006)	0.032	(0.009)
1997	0.095	(0.006)	0.029	(0.009)
1998	0.136	(0.007)	-0.001	(0.009)
1999	0.110	(0.007)	0.043	(0.009)
2000	0.137	(0.007)	0.018	(0.009)
2001	0.014	(0.009)	0.108	(0.007)
2002	0.079	(0.007)	0.009	(0.010)
2003	0.089	(0.007)	0.031	(0.010)
2004	0.089	(0.008)	0.029	(0.010)
2005	0.090	(0.008)	0.006	(0.011)
2006	0.089	(0.009)	0.038	(0.012)
2007	0.086	(0.010)	0.009	(0.013)
2008	0.072	(0.012)	0.038	(0.016)