Three Essays on Labor Markets

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Yao Yao

Contents

Chapter 1

General Introduction

This thesis consists of three essays on labor markets with focus of idiosyncratic labor income risk, its link to trade openness and unemployment inflows and durations. The second chapter, the first essay, is joint work with Tom Krebs. We decompose labor income uncertainty into permanent and transitory income risk using micro data in Germany, where the study on idiosyncratic labor income risk is inadequately documented so far. Permanent income risk is defined as variance of shocks to income that do not fade out over time and are assumed to be not self-insurable. The permanent income risk can be accumulated during the whole working life of individuals, and consists of the stochastic trend of individual's future income.

Much of the literature finds that the nature of income risk, in particular, the nature of permanent income risk, has profound effects on the consumption-savings decision of individuals, which in turn plays a central role in individual welfare. The pioneering study of Aiyagari (1994) shows that income risk has influence on the determination of wealth inequality and income taxation, when individuals cannot fully insure away their idiosyncratic risks and can only self-insure by borrowing and lending money. Deaton (1991) provides an analysis that the ability of individuals to self-insure and the precautionary saving motives are sensitive to the feature of microeconomic labor income process and income uncertainty. Constantinides and Duffie (1996) argue that idiosyncratic income risk has impact on the asset pricing and could explains risk premium in a specification using power utility. Moreover, Krusell and Smith (1999) and Krebs (2003) claim that idiosyncratic income risk causes the welfare cost of business cycles.

We follow the estimation method in Krebs, Krishna, and Maloney (2010) employing complete set of moment conditions for permanent income risk, which leads us deviating from existing literature estimating idiosyncratic income risk in Germany. The income measure used is individual annual earning. We provide several stylized facts about idiosyncratic income risk in Germany over the recent three decades. Firstly, the risk to individual earning is more transitory rather than persistent. Both permanent and transitory income risk demonstrate an upward trend after the early 1990, the German reunification time, and this uptrend seems to cease around the year 2000. It implies that the main driving force of increasing trend in permanent income risk is due to the German reunification, instead of continuously growing trade activities. Besides, we also discover that East German residents and females are two groups with considerably higher income risks, compared to West Germans and males, respectively. Moreover, education might bring more income uncertainty.

The third chapter, the second essay, analyzes how increased offshoring impacts on idiosyncratic labor income risk. This is a joint work with Jan Hogrefe. It provides an assessment that directly connects labor income risk and offshoring trends in a panel setting at the industry level. We are also the first to conduct such an investigation for Germany; a large trading economy with exceptional reliance on international integration.

Our analysis proceeds in three steps. First, the unpredictable part of individual income is estimated as the residual term from wage regressions based on individual-level data. Second, values for permanent income risk at the sector level are estimated as the variance of persistent shocks to this unpredictable component. We link these to offshoring in a panel framework in step three. Offshoring at the sector level is understood as the amount of intermediate inputs imported from the same industry abroad divided by industry output. We do not distinguish between within-firm versus arm's length transactions as we consider our offshoring variable to approximate the outcome of any make-or-buy decision.

We do, however, treat offshoring to non-OECD countries with special attention since this is usually the type of production relocation which stirs up the most anxiety among the public. It is also the type of cost-savings driven offshoring modeled in most of the theoretical literature on the topic.

We find an increase in offshoring in a given sector correlates with a decrease in permanent income risk in that sector. Offshoring to non-OECD countries has a particularly strong impact. We attribute this offshoring related decline in income risk to the fact that, in countries with relatively rigid labor markets, firms tend to offshore volatile production intensive parts of their undertakings which leaves the remaining tasks less volatile on average. Overall, a decrease in income risk provides a channel through which offshoring can be welfare enhancing.

In the last chapter, the third essay, is a joint work with Esther Perez. Here we read search theory's unemployment equilibrium condition as an Iso-Unemployment Curve (IUC). A country's position along the curve reveals its preferences over the destruction-duration mix. Using a panel of 20 OECD countries over 1985-2008, we distinguish two types of labor markets, dynamic labor market displaying higher turnover rate of workers, and stagnant labor market with lower turnover rate.

One fundamental question raised is whether dynamic labor markets significantly outperform more stagnant ones. Furthermore, is high unemployment inflow rates with short unemployment durations a better combination (in terms of unemployment rate), compared to limited unemployment inflows and longer unemployment spells? We find that labor market institutions plays different roles on unemployment inflows and durations. In particularly, the employment protection legislation to have opposing effects on destructions and durations, while the effects of the remaining key institutional factors on both variables tend to reinforce each other. Implementing the right reforms could reduce job destruction rates by about 0.05 to 0.25 percentage points and shorten unemployment spells by around 10 to 60 days. Correspondingly, unemployment rates declines as well.

Chapter 2

Labor Income Risk in Germany

2.1 Introduction

This chapter focuses on the estimation of labor income risk and its time trend in Germany over the last three decades. We address two questions in this chapter. Firstly, how volatile has innovative labor income been in Germany since the early 1980s? We distinguish between two types of idiosyncratic labor income risk: permanent and transitory. The permanent component must be considered when agents make decisions of future consumption and asset holdings and therefore has substantial effects on individual welfare. In contrast, transitory component can be self-insured by saving and dis-saving and thus has less effect on consumption smoothing and individual welfare. We model the stochastic individual labor income process and specify and estimate the two components in order to quantitatively analyze permanent and transitory idiosyncratic income risk in Germany over the last three decades. Secondly, apart from most existing literature, we estimate *time-dependent* idiosyncratic labor income risk, which allows us to examine how estimated labor income risk evolves over the last thirty years. German reunification in 1990 is the first special feature of this time period. Alongside this was the rapid expansion of trade volume between Germany and the rest of the world. The reunification and trade globalization would be potential driving forces of any time pattern of individual labor income risk that emerges over this thirty year period.

Much of the literature on heterogeneous-agent models has shown the importance of distinguishing between permanent and transitory income risk and estimating idiosyncratic labor market income risk. In this type of model, individuals maximize expected life time utility by choosing consumption and asset holdings for each period, given idiosyncratic labor income risk and the lack of perfect insurance to insure the risk away. Hence, the patterns of idiosyncratic labor income risk have profound effects on the consumption-savings decision of individuals, as well as on individual welfare. Aiyagari (1994) examines the influence of individual labor income risk on wealth inequality. Moreover, Krebs (2003, 2007)shows the welfare relevance of idiosyncratic income risk. A reduction in uninsurable idiosyncratic labor income risk increases human capital investment and welfare by a significant amount. A quantitative analysis has been conducted by Krebs, Krishna, and Maloney (2010) and Kuhn (2013). Individuals are willing to give up a large amount of consumption (and therefore lose individual welfare) in return for avoiding an increase in income risk. Furthermore, the amount of consumption given up by agents becomes stronger if they are more risk averse. In all, when permanent income risk is present, welfare losses under incomplete insurance market are substantial and the welfare gain is non-negligible when persistent income risk is removed.

Besides the quantitative investigation of idiosyncratic income risk, we are also interested in its time pattern in recent decades. German reunification was a unique and huge shock to the labor market and the German economy in general. After the fall of the Berlin Wall, there was substantial turbulence, labor force flows and migration between the two regions of Germany. Big changes happened before and after the event. As Fuchs-Schuendeln, Krueger, and Sommer (2010) pointed out, there exists large differences in labor income and income inequality prior and subsequent to the reunification. Before the reunification, income grew at healthy rate, and there was no rising income inequality. In contrast, income has grown much more slowly and income inequality started to go up after the reunification. In this chapter, we examine the remaining and lasting effect of reunification on income uncertainty in the German labor market using our measure of labor income risk.

On the other hand, the globalization has been proceeded at a rapid rate as the trade volume has expanded. To illustrate the globalization process in Germany, we measure trade openness using the exports and imports of goods and services as proportion of GDP as plotted for the last three decades in figure 2.1. The exports and imports share of GDP in Germany declined slightly in the late 1980s and early 1990s, and has been continuously rising from the early 1990s to 2011, except for one drop during the recent crisis in around 2009. Notably, the exports share has increased from about 20 percent at the time of reunification to above 40 percent in 2011, and imports share, which is at lower level than exports share at all times, has grown from less than 20 percent to about 35 percent. The observation in figure 2.1 motivates our second research question as how idiosyncratic income risk evolves in a more open economy. The trade literature has conducted both theoretical and empirical analysis on the correlation between trade openness and income volatility. For example, Grossman and Rossi-Hansberg (2008) present a theoretical framework in which trade can reallocate tasks with different income volatilities among different countries. Furthermore, Krebs, Krishna, and Maloney (2010) and Hogrefe and Yao (2012) provide empirical evidence that trade openness can either positively or negatively related to idiosyncratic income risk.

A large body of the literature has analyzed the difference in labor market performance pre- and post-reunification, how trade openness and financial globalization affect wages and employment, and provided theoretical and empirical findings on both aspects. The value added in this chapter is that we focus on a less mentioned but nevertheless important aspect of the labor market: that is, the stochastic individual labor income process and in particular the evolution of idiosyncratic labor income risk over the last three decades.

Idiosyncratic labor income risk in the US has been broadly studied in the existing literature. Nevertheless, the stylized facts concerning labor income risk in the German labor market are less well-documented so far. This chapter aims to provide more empirical findings on idiosyncratic labor income risk in Germany, employing a relatively complete set of moment conditions to generate plausible estimates, particularly of permanent risk, using German Socio-Economic Panel Study (henceforth GSOEP) data.

Figure 2.1: *Trade globalization in Germany*

The reminder of the chapter is organized as follows. In section 2.2, we review the existing literature on idiosyncratic income risk. We then describe our data selection in section 2.3. The empirical model is specified in section 4.3, and section 2.5 summarizes our empirical findings. We finally conclude in section 4.5.

2.2 Literature Review

Working with Panel Study of Income Dynamics (henceforth PSID) data, a bulk of research estimates the idiosyncratic income risk in the US. For instance, Carroll and Samwick (1997) provide a fundamental method for estimating time-independent individual income risk. They define income risk as the variance of (unpredictable) changes of income, and disentangle permanent and transitory components. The estimated permanent income risk is 14.5% and the transitory risk is approximately 21% in the US over a 1981-1987 sample period.

Another study by Storesletten, Telmer, and Yaron (2004) asks how idiosyncratic income risk varies over the business cycle. They conclude that permanent income risk is strongly countercyclical, and find that estimates of persistent income risk are roughly 21% in a recession and 12% in an expansion. On top of permanent and transitory income risk, Meghir and Pistaferri (2004) add a third component, measurement error in the income process, which has partial persistent power. They treat this measurement error as an MA(q) process, and model labor income as a parsimonious ARCH process with both observable and unobserved heterogeneity.

Besides the study of labor income risk in the US, Krebs, Krishna, and Maloney (2010) develop a GMM estimation of income risk, using longitudinal quarterly data of Mexican workers in 21 different industries. They find that the annualized permanent income risk in Mexico is around 18% but it varies among different manufacturing sectors.

Meanwhile, there has been much less study of income risk based on micro data for the German labor market. Fuchs-Schuendeln, Krueger, and Sommer (2010) estimate the transitory and permanent income risk to the innovative wage rate and household labor income. They extract transitory risk from the covariance of growth rate of individual labor incomes, and specify a permanent component by eliminating the transitory part from the variance of the growth rate. They report the average permanent income risk in Germany from early 1980s to late 1990s as 12.7%, and higher transitory income risk than permanent income risk and a slight upward trend of permanent income risk during the 1990s.

Gernandt and Pfeiffer (2007) investigate the evolution of wage inequality in East and West Germany, especially after German reunification. Their findings are that rising inequality emerges in the East and West, but with different characteristics. There is higher degree of wage inequality in lower part of the wage distribution in West Germany, while in the upper part of the wage distribution in East Germany.

In this chapter, we follow and extend the estimation method in Krebs, Krishna, and Maloney (2010). That is, we adopt a more complete set of moment conditions including the growth rate(, which is used in Fuchs-Schuendeln, Krueger, and Sommer (2010)), as well as moment conditions of higher order differences, such that we specify the component of innovative individual income with a high degree of persistence, that is, with welfare relevance. More details on the estimation procedure are presented in section 4.3.

2.3 Data

The data used in this study are drawn from the 1983-2010 GSOEP. The GSOEP is a longitudinal survey of private households and individuals in Germany. It has a similar structure to PSID, which is widely used in various topics of economic research. We include all subsamples in GSOEP except for the 'Oversampling of high income' subsample which is eliminated to minimize potential selection bias. We concentrate on analyzing household heads, the main earner in the household between 25 and years of age 60, which is assumed to be the working period of individuals in the data. The labor income measure employed is individual annual earnings, which is 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. Moreover, we eliminate individuals with too low hourly wage rate (less than 3 euros per hour) in order to minimize measurement error by ruling out implausibly low income. Lastly, self-employed people are not included. In GSOEP, self-employed people report capital and labor income together as individual earnings. We exclude self-employment, as we focus on the labor income of individuals.

In GSOEP, East Germany (the previous German Democratic Republic, GDR) joins the West (officially, the Federal Republic of Germany) in the year 1991. Therefore, our 1983 to 2011 sample allows us to investigate if there is any remaining impact of reunification on idiosyncratic labor income risk in East and West 20 years after this unique event.

2.4 Estimation of the Labor Income Risk

2.4.1 The Model

In this chapter, labor income risk is defined as the unpredictability of income and refers to income variability from an ex ante perspective. Therefore, labor income risk arises whenever a person's future income deviates from their expectation and it is idiosyncratic to them. In general, this could happen during the whole working period of an individual. In this sense, labor income risk is estimated as the variance of the unpredictable part of individual income. It can be divided into two components: permanent income risk and transitory income risk. Permanent income risk consists of the stochastic trend of innovative income and has persistent power over the remaining length of working life, while its transitory counterpart follows a random white noise process and is defined as the deviation of income from a stochastic trend. Hence, individuals consume a certain proportion of a permanent shock rather than a transitory one. Permanent risk has a considerable effect on future consumption and asset holdings, as well as individual welfare.

The innovative individual income is derived from a standard Mincerian type regression model, where each individual generates her expectations about future income from a projection based upon her observables and predictable characteristics.
Consider the following equation of individual labor income i in time t:

$$y_{it} = \alpha_t + X_{it}\beta_t + u_{it} \tag{2.1}$$

y_{it} is logarithm of our measure of labor income, individual annual earnings. The constant α_t captures the effect of aggregate shocks on individual earnings, e.g. oil price shocks and business cycle effects. X_{it} is a vector of all observables and demographic characteristics of individuals: gender, age, experience, region, family size, education and experience. β_t are year-specific coefficients. The expected return to individual characteristics is identical for all individuals in the same period but could change over time. For instance, education is found to have a higher return on individual labor income in recent years because of global technological progress. u_{it} is the stochastic part of individual labor income, which is idiosyncratic and unpredictable ex ante. The first-stage Mincerian regression (2.1) provides the estimated parameters $\hat{\alpha_t}$, $\hat{\beta_t}$ and predicts the residuals $\hat{u_{it}}$.

We make further assumptions on the residuals u_{it} so that it can be decomposed into the permanent component ω_{it} and a transitory innovation ϵ_{it}. The permanent component ω_{it} is assumed to be a random walk without drift.

$$u_{it} = \omega_{it} + \epsilon_{it} \tag{2.2}$$

$$\omega_{it} = \omega_{it-1} + \eta_{it} \tag{2.3}$$

Moreover, we suppose η_{it} and ϵ_{it} are time dependently distributed, that is,

$$\eta_{it} \sim N(0, \sigma^2_{\eta,t}) \tag{2.4}$$

$$\epsilon_{it} \sim N(0, \sigma^2_{\epsilon,t}) \tag{2.5}$$

We assume that both the permanent and transitory shock η_{it} and ϵ_{it} are i.i.d. across agents, and that η_{it} and ϵ_{it} are independent with each other at all leads and lags, i.e. $cov(\eta_{it}, \eta_{is}) = 0, \forall t \neq s, cov(\epsilon_{it}, \epsilon_{is}) = 0, \forall t \neq s, cov(\eta_{it}, \epsilon_{is}) = 0, \forall t, s$. The standard deviations $\sigma_{\eta,t}$ and $\sigma_{\epsilon,t}$ of the permanent shock η_{it} and the transitory shock ϵ_{it} are the measures of permanent and transitory income risks. They are time-varying, such that individuals face different income risks in each year.

2.4.2 Estimation Method

We follow closely the estimation procedure in Krebs, Krishna, and Maloney (2010), and extend it to panel data with a longer time dimension. From equations (2.2) and (2.3) above, the n-year difference of u_{it} can be expressed as

$$\triangle_n u_{it} = u_{it+n} - u_{it} = \eta_{it+1} + ... + \eta_{it+n} + \epsilon_{it+n} - \epsilon_{it} \tag{2.6}$$

Then the variance of $\triangle_n u_{it}$ is given by

$$V[\triangle_n u_{it}] = \sigma^2_{\eta,t-1} + ... + \sigma^2_{\eta,t-n} + \sigma^2_{\epsilon,t} + \sigma^2_{\epsilon,t-n} \tag{2.7}$$

We estimate (2.7) by minimizing the sum of squared distance between the moment conditions in the model and the corresponding moments in the data. Put differently, we run an OLS regression of (2.7), where $V[\triangle_n u_{it}]$ is the dependent variable with observations $\hat{V}[\triangle_n \hat{u}_{it}]$, while the right hand side is a combination of $\sigma^2_{\eta,t}$ and $\sigma^2_{\epsilon,t}$ with varying n and t. If $\sigma^2_{\eta,t}$ or $\sigma^2_{\epsilon,t}$ is included in a particular $V[\triangle_n u_{it}]$, we put 1 at the corresponding position of independent variable. If not, we put 0 instead. This makes our independent variables with observation 1 or 0 and with coefficients $\sigma^2_{\eta,t}$ or $\sigma^2_{\epsilon,t}$. We have a sufficiently number of observations of $V[\triangle_n u_{it}]$ because the time period in our data is long enough. Then the OLS estimators $\hat{\sigma}^2_{\eta,t}$ and $\hat{\sigma}^2_{\epsilon,t}$ are unbiased estimates of $\sigma^2_{\eta,t}$ and $\sigma^2_{\epsilon,t}$.

2.5 Empirical Results

2.5.1 Labor Income Risk Estimates

Table 2.1: *Average of estimated labor income risk*

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
Male in West	High School dropouts	0.2867	0.1241
	High School graduates	0.2445	0.0785
	College graduates	0.1827	0.0975

We summarize the labor income risk estimates in this section. Table 2.1 presents the average over time of estimates of transitory and permanent income risk to individual labor income using GSOEP data according to different subgroups. For all individuals in our sample, as shown in the upper part of table 2.1 the permanent income risk to individual labor income has a mean value of 10.78% between 1985 and 2008[1], while the transitory labor income risk is approximately 30% on average over the same period. To interpret, an individual who earns the average labor income, i.e. 30,000 euros per year, faces an average of permanent shock of 10.78% or 3000 euros, between 1985 and 2008.

Furthermore, we classify our sample into several subgroups, that is, males in the West, females in the West, males in the East and females in the East, and look at the labor income risk in those subgroups separately in order to investigate the region and gender difference in idiosyncratic labor income risk. In the middle part of table 2.1, we display the average estimated permanent and transitory income risk in these subgroups over time.

As we see in table 2.1, estimated labor income risk differs substantially between East and West Germany. The permanent and transitory labor income risks in East Germany are higher for both females and males than in the West. The average transitory income risk for females in the East is almost 42% and average permanent income risk is 16%, while transitory income risk is a bit higher than 36% for females in the West, and permanent risk is approximately 12% for the West females. Males have transitory income risk of nearly 31% and permanent income risk of about 10.48% in the East, while males in the West have transitory risk of 24% and permanent risk of 10.31% to individual earning in the West.

We are not the first to observe the different patterns of the micro labor income process in the East and West. Fuchs-Schuendeln, Krueger, and Sommer (2010) report that the increase in income inequality is much stronger in the East than in the West, and that there exists a convergence of income inequality in males between East and West. Interestingly, we find that the gap of labor income risk is smaller for males than for females when comparing East and West residents.

On the other hand, females have more income uncertainty than males in both regions.

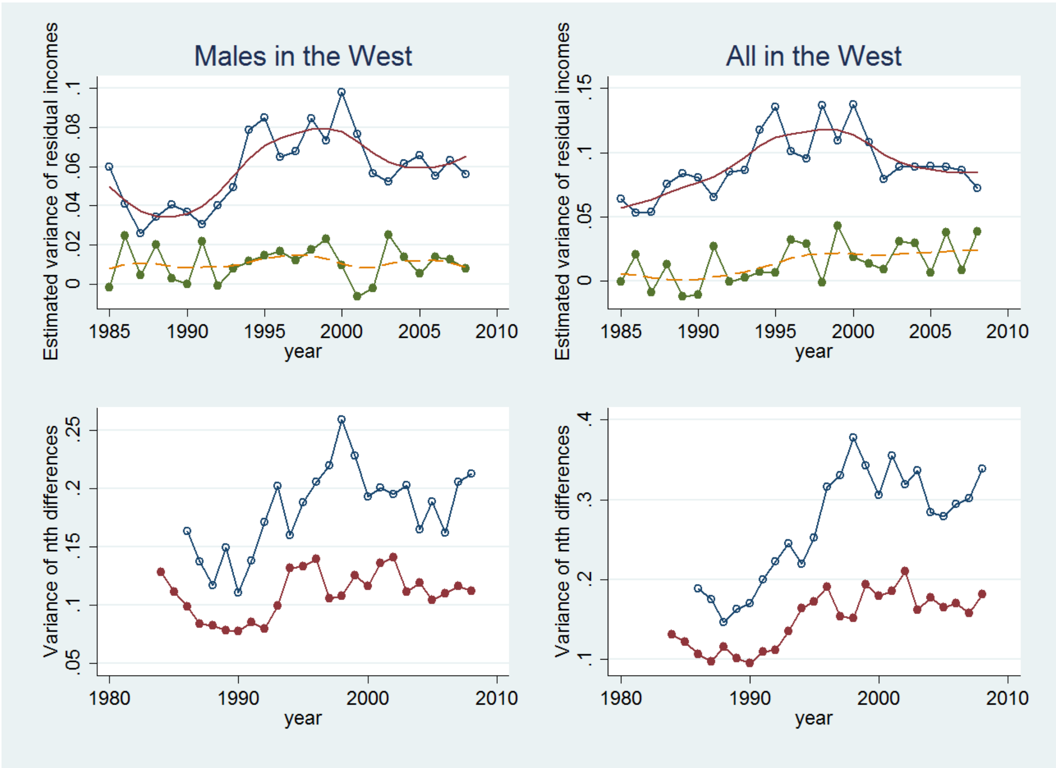
[1]We drop estimates in the first two and the last two years, as we need further assumptions to specify the yearly estimates in those years.

Precisely, females have about 12% more transitory income risk, and about 2% more permanent income risk than males in West Germany, while females have around 11% higher of transitory income risk and about 6% higher of permanent income risk than males in East Germany. Much of the literature has found that females have quite different labor participation decisions compared to males. Their labor supply is more sensitive to their marital status, the presence of young children, home productivity, tax incentives etc. As observed (e.g. in U.S. Census Bureau, International Statistics (2012)), the marriage rate has declined and the divorce rate has increased over the last several decades in Germany. Not surprisingly, there has been notable increase in single-parent households over this time. All these facts can be associated with the higher income volatility of females in our sample.

In addition, we allow for education-specific differences when estimating idiosyncratic income risk. In order to disentangle the effect of education on labor income risk from the effect of region and gender, we classify males in the West into three groups according to their education level with respect to high school, namely, high school dropouts, high school graduates including other equivalent diplomas, and college graduates which includes people with degrees more advanced than high school.

High school graduates not receiving more education or training have the lowest permanent risk, compared to high school dropouts and college graduates. Consider a typical high school graduate who receives the necessary training and stays in the same job role for most of her working life; the income risk for her is relatively low. However for a college graduate who has more skill and is more adaptive and flexible to new jobs and positions, the excess volatility of labor income possibly originates from switching jobs or obtaining promotion. In the meantime, people with limited training, say, without high school graduation or enough professional training, mostly work part-time or short term. Therefore, the income process of high school dropouts exhibits greater fluctuations. This observation implies that an increase of human capital investment does not necessarily reduce the uninsurable income risk.

Figure 2.2: *Time trend of estimated labor income risk*

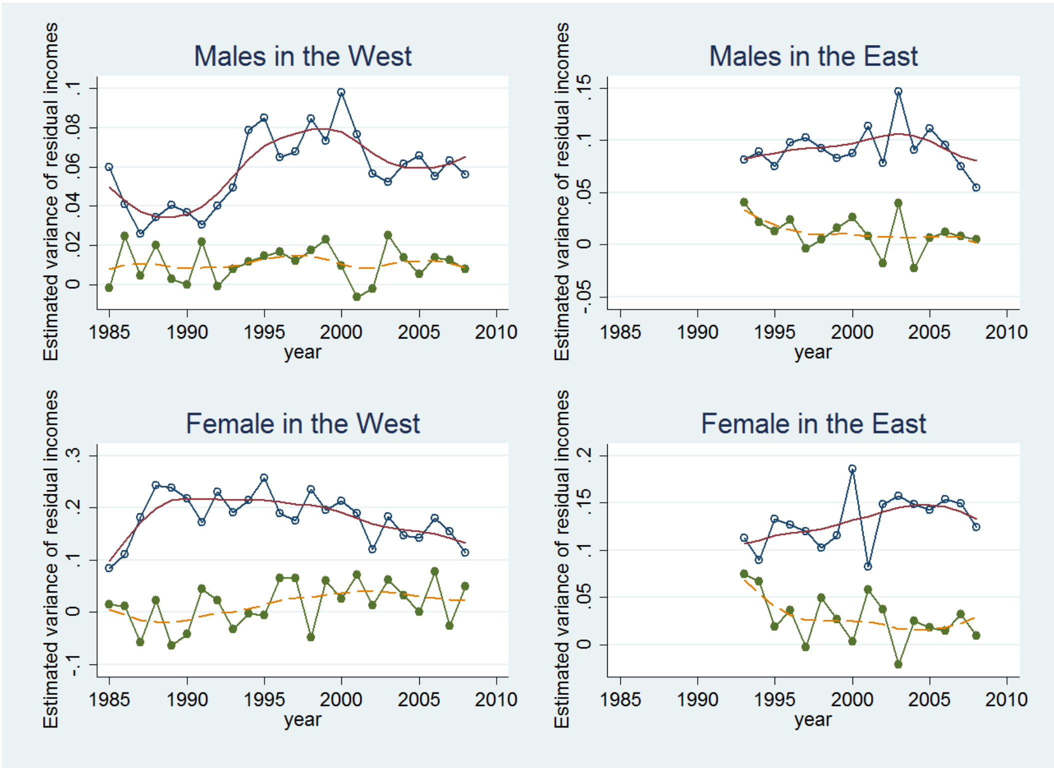


Note: In the upper diagrams, the hollow dots represent the estimated variance of transitory components, solid dots are the estimated yearly variance of permanent components, and the solid and dashed lines are the corresponding HP filtered trend components of transitory and permanent estimates. In the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

2.5.2 Main Results

Figure 2.2 illustrates the time trend of estimated idiosyncratic labor income risk between 1985 and 2008. In the upper panel of figure 2.2 we plot the estimates of permanent and transitory labor income risk for males in the West and for all agents in the West. 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 components of transitory and permanent estimates. The lower panel in figure 2.2 presents the estimated variance of one-year and 5-year differences of residual incomes generated from the first stage regression, $\hat{V}[\triangle_1 \hat{u}_{it}]$ and $\hat{V}[\triangle_5 \hat{u}_{it}]$, for males in the West and for all agents in the West. Details on the yearly estimates and

Figure 2.3: *Time trend of estimated labor income risk*



Note: In the upper diagrams, 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 components of transitory and permanent estimates. In the lower diagrams, 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.

corresponding standard errors can be found in appendix 2.B.

Due to the limited number of years in the sample prior to the 1990s, there is no clear pattern shown in the estimates of transitory and permanent components, as well as in the estimated variances of the first and fifth differences of residual income. However, clearer time patterns emerge after the early 1990s in our estimates. For males in the West, transitory income risk, variance of the 5-year and 1-year difference of residual incomes have started to rise since early 1990, the post reunification period for Germany, until the year 2000. In the meanwhile, yearly estimated variances of permanent components exhibit a moderate yet increasing trend which terminates around the year 2000. Increases in permanent income risk are accumulated year after year and eventually lead to a substantial effect on individual consumption-saving decisions as well as individual welfare, given the persistence of permanent income risk.

The observation that income risk increases after reunification, and that there is an end in the uptrend of permanent income risk after the year 2000, implies that the adjustment process after reunification is wearing off over time, and that the increasing trend (which is similarly observed by Fuchs-Schuendeln, Krueger, and Sommer (2010) is no longer there. This seems to be the leftover turbulence after reunification. Once the process is over, the trend changes. The unique event of reunification in Germany tends to increase both transitory and permanent income risks for residents in the West, but the effect of reunification on our measure of income risk reecks one decade after the event.

In the mean time, export and import shares of GDP have been continuously rising, except for one drop during the recently crisis. We can conclude that the earlier uptrend is due to turmoil after the reunification, not because of globalization. Hence, for our measure of labor income risk, no effect is observed due to globalization or the increasing activities of exports or imports.

As we can see in the plot for all agents in the West (the right diagrams in figure 2.2), transitory income risk, variance of 5-year difference of residuals and variance of 1-year difference of residual incomes have similar trends, compared to males in the West. These three yearly estimates started to rise from the early 1990 until around the year 2000. However, there are more fluctuations in the estimated variance of permanent components, which might be because of the female counterpart in our sample.

We further compare males and females in the West in order to estimate the gender differences in the time trend. The result is shown in figure 2.3. Compared to males in the West, females in the West seem to have no clear trend of permanent income risk, the welfare-relevant income risk. The time trend of the estimated variance of fifth differences of residual incomes reinforces the finding for males in the West that variance has increased after reunification until around the year 2000, and a decreasing trend seems to emerge afterwards. The impact of reunification works in a different way for males in the East, as we observe that permanent income risk declines over time and that there is no clear trend in transitory income risk for males in East Germany. Furthermore, there seems to be no uptrend in permanent income risk in the West before reunification. Moreover, reunification seems to increase transitory risk to labor income as well.

2.5.3 Robustness

We further test the robustness of our results by adjusting the estimation sample, and we focus on the time trend of estimates of permanent income risk. We firstly eliminate the minimum hourly wage rate threshold of agents' earnings, that is, we include all household heads who are not self-employed and between age 25 and 60 regardless of their labor income. The estimation results without the minimum wage threshold are shown in figure 2.4. Similar estimates of permanent and transitory labor income risk and the variance of 1-year and 5-year differences of residual incomes emerge. Firstly, there is an increasing trend of transitory and permanent income risks during the 1990s for males and all agents in the West when excluding the minimum wage threshold. Secondly, the increasing trend seems to be discontinuous after around the year 2000.

We then take self-employment into the sample, such that we include all household heads who are between the age of 25 and 60 and earn a credible hourly wage rate, regardless of whether they are self-employed. Estimates of labor income risk and variance of differences are plotted in figure 2.5. For males in the West, the uptrend of permanent income risk seems to be less clear compared to other data selection criteria, however, it still shows that there is a halt to increasing labor income risk after the year 2000, and even the beginning of a decline. For all agents in the West, the right panel of figure 2.5, there are more fluctuations of estimated permanent income risk, as the income of self-employed includes their capital income which incorporates more volatile sources of income.

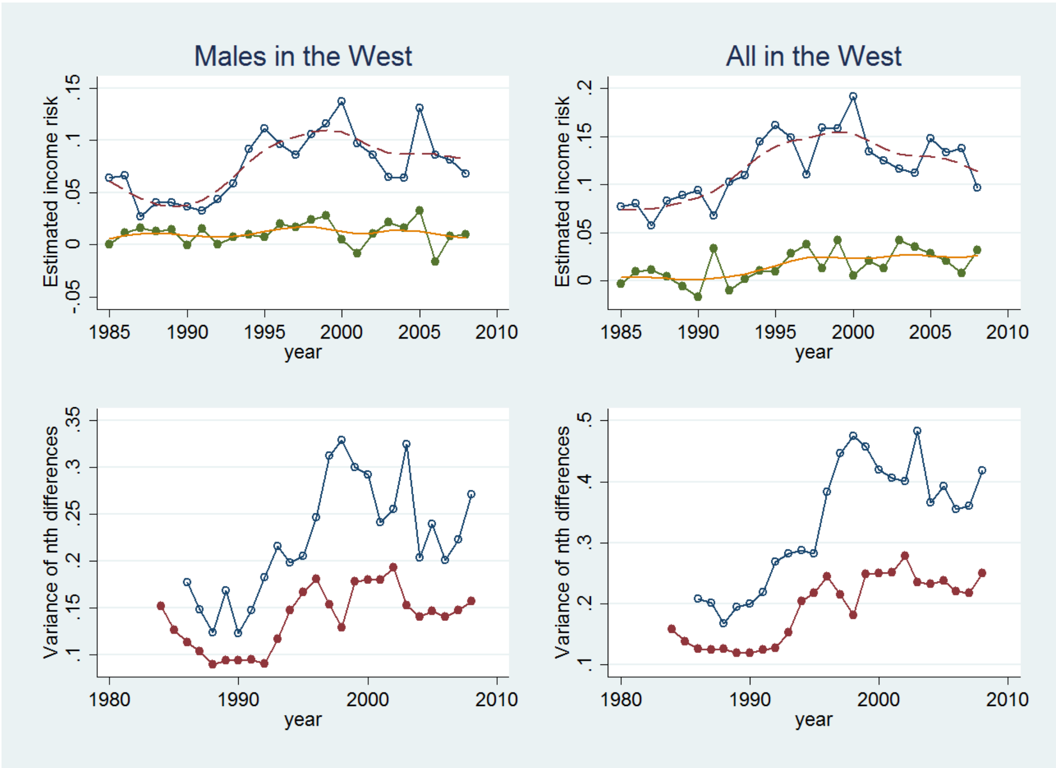
Finally, we drop one of the regressors, years of education, in our first stage regression, in order that we could examine if the time trend of estimated labor income risk relies on the choice of the independent variables in our first stage regression (2.1). For males in the West, the increasing trend in 1990s and its subsequent expiration are still observed in figure 2.6.

2.6 Conclusion

This chapter decomposes idiosyncratic income risk into permanent and transitory components and estimates the annual transitory and permanent income risks in Germany over the last three decades. We assume that there are two shocks to individual earnings, transitory and permanent shocks, where the permanent shock is the stochastic trend of income and has persistence over the remaining working period of individuals. The standard deviation of permanent and transitory shocks serve as our measures of permanent and transitory income risk. This chapter focuses on estimation of permanent income risk and its recent time trend.

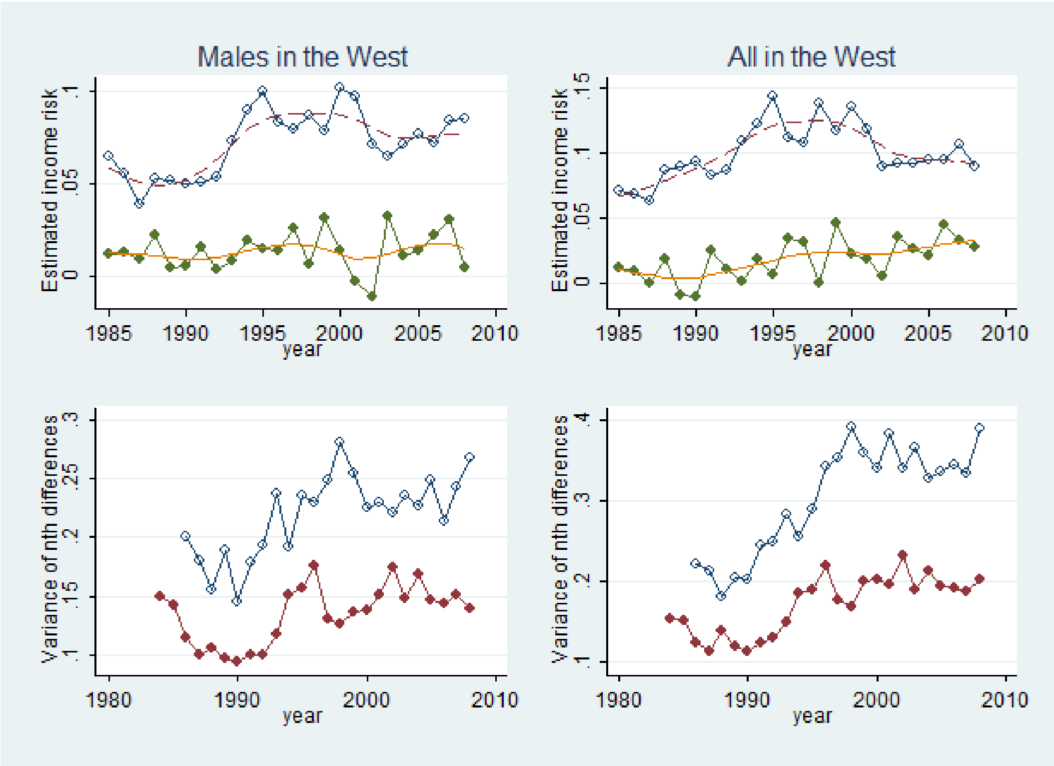
We find that the average estimated permanent labor income risk in Germany between 1985 and 2008 is about 11%, and that average of transitory income risk is approximately 30% over the same time period. Individuals in East Germany face a more volatile income process compared to those in West Germany, and male household heads have lower

Figure 2.4: *Time trend of labor income risk without minimum wage threshold*



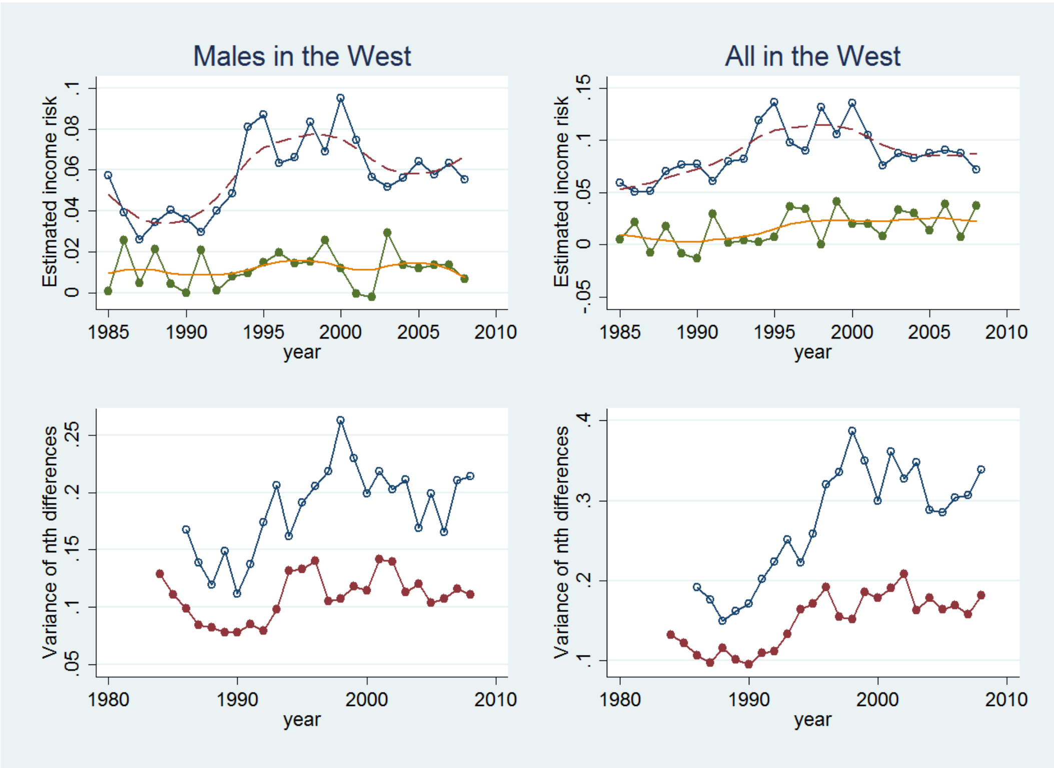
Note: In the upper diagrams, the hollow dots represent the estimated variance of transitory components, solid dots are the estimated yearly variance of permanent components, and the solid and dashed lines are the corresponding HP filtered trend components of transitory and permanent estimates. In the lower diagrams, 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 2.5: *Time trend of labor income risk including self-employment*



Note: In the upper diagrams, 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 components of transitory and permanent estimates. In the lower diagrams, 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 2.6: *Labor income risk without education in the first stage regression*



Note: In the upper diagrams, 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 components of transitory and permanent estimates. In the lower diagrams, 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.

income risks than females. Interestingly, education could bring more risk to individual earnings. We also observe that permanent labor income risk in Germany has an upward trend after the early 1990s, the post German reunification period, and the uptrend seems to cease around the year 2000. This implies that the growing permanent income risk during the 1990s could be mainly driven by turbulence in labor market in the wake of reunification rather than the process of globalization. It implies that the German labor market remained stable during rapidly growing trading activities in world markets.

2.A Descriptive Statistics and First Stage Regression

Table 2.A.1 displays the descriptive statistics of variables in our sample, and figure 2.A.1 and 2.A.2 show the time trend of means and medians of real and nominal earnings in our sample. We also compare our means and medians to the real and nominal earnings reported by the Federal Statistical Office in Germany. In figure 2.A.1 and 2.A.2, the solid line represents connected means of real and nominal monthly labor income, the dashed line shows connected 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 one can see, our sample demonstrate similar time trend to what German Federal Statistical Office reports, that is, there was a significant drop after reunification for real earning and a relative moderate but still considerable drop for nominal earnings in Germany. Average real earnings hadn't increasing during 1990s, and declined further after the year 2000. The average nominal earnings seem to grow smoothly during 1990s, and have slowed down after 2000.

Table 3.A.1 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. Put differently, individual labor income grows with age, education and family size.

Table 2.A.1: *Descriptive statistics of our sample (1983-2010)*

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

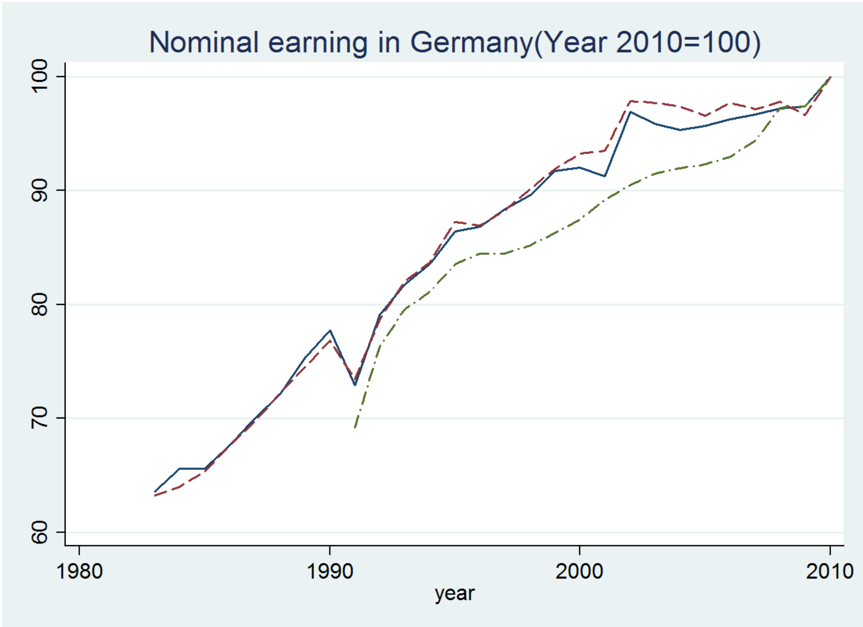
Figure 2.A.1: *Mean and median of real income*

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.

Table 2.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

Figure 2.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.

2.B Estimates of Income Risk and Standard Errors

In table 2.B.1 and 2.B.2, we report our yearly estimated variances of transitory and permanent components of residual incomes, and the corresponding standard errors, for all agents in the West and only males in the West. Table 2.B.3 shows the estimated variance of one-year and 5-year differences of residual incomes for all agents in the West and males in the West.

Table 2.B.1: *Estimation results : All agents in the West*

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)

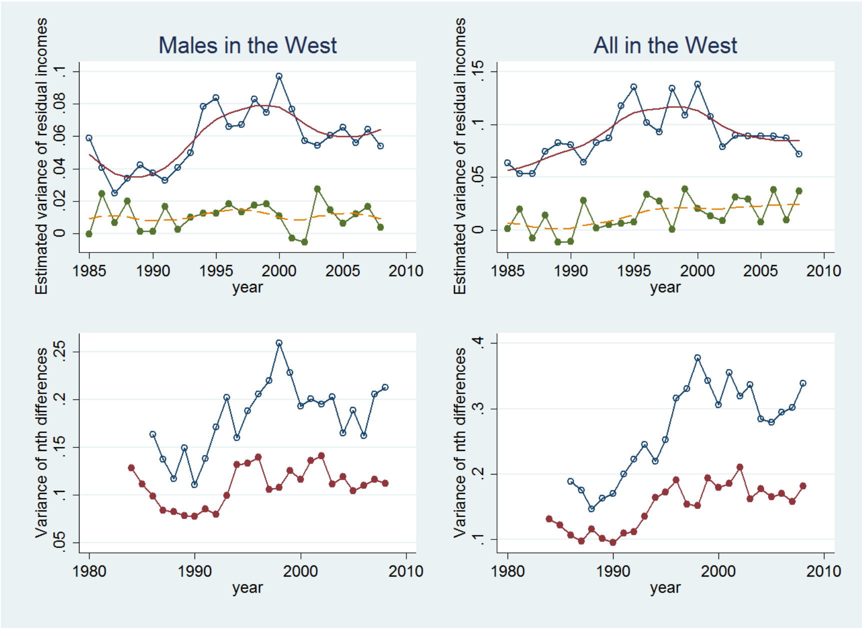
Table 2.B.2: *Estimation results : Males in the West*

Year	Estimates of transitory components	(standard errors)	Estimates of permanent components	(standard errors)
1985	0.060	(0.010)	-0.002	(0.017)
1986	0.041	(0.008)	0.024	(0.013)
1987	0.026	(0.007)	0.005	(0.011)
1988	0.034	(0.007)	0.020	(0.010)
1989	0.041	(0.006)	0.003	(0.009)
1990	0.037	(0.006)	0.000	(0.009)
1991	0.031	(0.006)	0.022	(0.008)
1992	0.040	(0.006)	-0.001	(0.008)
1993	0.049	(0.006)	0.008	(0.008)
1994	0.079	(0.005)	0.012	(0.007)
1995	0.085	(0.005)	0.015	(0.007)
1996	0.065	(0.005)	0.017	(0.007)
1997	0.068	(0.005)	0.012	(0.007)
1998	0.085	(0.005)	0.017	(0.007)
1999	0.073	(0.005)	0.023	(0.007)
2000	0.098	(0.006)	0.010	(0.007)
2001	0.077	(0.006)	-0.006	(0.008)
2002	0.056	(0.006)	-0.002	(0.008)
2003	0.052	(0.006)	0.025	(0.008)
2004	0.061	(0.006)	0.014	(0.009)
2005	0.066	(0.007)	0.005	(0.009)
2006	0.055	(0.007)	0.014	(0.010)
2007	0.063	(0.008)	0.013	(0.011)
2008	0.056	(0.010)	0.008	(0.013)

Table 2.B.3: *Estimated variance of first and fifth differences of* $\hat{u}_{it}$

Year	All agents in the West		Males in the West	
	$\hat{V}[\triangle_1 \hat{u}_{it}]$	$\hat{V}[\triangle_5 \hat{u}_{it}]$	$\hat{V}[\triangle_1 \hat{u}_{it}]$	$\hat{V}[\triangle_5 \hat{u}_{it}]$
1986	0.106	0.188	0.098	0.163
1987	0.096	0.174	0.083	0.137
1988	0.115	0.146	0.082	0.117
1989	0.100	0.162	0.077	0.149
1990	0.094	0.169	0.077	0.110
1991	0.109	0.200	0.085	0.137
1992	0.111	0.222	0.079	0.171
1993	0.134	0.244	0.098	0.202
1994	0.163	0.218	0.131	0.159
1995	0.171	0.252	0.133	0.187
1996	0.190	0.316	0.139	0.205
1997	0.153	0.330	0.105	0.220
1998	0.151	0.378	0.107	0.258
1999	0.193	0.343	0.125	0.228
2000	0.178	0.305	0.116	0.193
2001	0.185	0.355	0.136	0.200
2002	0.209	0.319	0.140	0.194
2003	0.161	0.336	0.111	0.202
2004	0.176	0.283	0.119	0.164
2005	0.164	0.278	0.104	0.188
2006	0.169	0.294	0.109	0.161
2007	0.157	0.301	0.116	0.205
2008	0.181	0.338	0.111	0.212

Figure 2.C.1: *Weighted estimation of income risk*



2.C Weighted and Unweighted Estimation

Individual weights are usually applied when using micro data in order to match the whole population. In order to see how sensitive our estimation is to inclusion of weights, we construct the weighted estimates for income measure as individual labor earnings. The results are shown in figure 2.C.1.

Chapter 3

Offshoring and Labor Income Risk: an Empirical Investigation

3.1 Introduction

Globalization is often perceived as creating a more volatile working environment on the labor market. In particular, trends such as the relocation of parts of production abroad (offshoring) induce fears of job loss and higher fluctuations in individual income. While the long-run *level* effects of different types of offshoring on income and employment have been documented by a large literature, see e.g. Feenstra (2010), a lot less academic attention has been paid to the analysis of effects on the *variability* of incomes. Our paper further completes the picture of how offshoring has an impact on characteristics of labor income by estimating its relationship to income risk at the industry level using data from German manufacturing. To the best of our knowledge, our paper is the first to put the link between offshoring and income risk at the heart of an empirical analysis.

Income risk is defined as the variance of changes in the unexplained component of individual income. As such, it describes changes in income that are not a result of observable and predictable characteristics like age or education. Crucially, and in line with the literature, we econometrically distinguish between transitory and permanent risks to income. Transitory shocks to income are more likely to be smoothed out by self-insurance mechanisms such as saving and borrowing. However, this does not hold for permanent shocks, i.e. shocks that permanently shift an individual's income trajectory. Following the literature, we assume permanent income risk to be uninsurable from an individual perspective. Then, unexpected permanent variation in income affects the present value of lifetime earnings, which impacts on individual welfare (Aiyagari, 1994). It is thus the permanent component of income shocks we are interested in. Linking offshoring to changes in the variance of permanent income shocks yields evidence on the effect of offshoring on labor income which allows for a discussion of welfare consequences.

Our analysis proceeds in two steps. First, we derive and estimate measures of income risk at the industry level. Second, we link it to offshoring. Income risk is estimated at the industry level from individual income data as the average variance of changes in the unexplained component of individual income. The latter is retrieved from standard Mincerian wage regressions. For this first stage we use variation in individual level data from a sample of official German social security records to estimate industry level income risk. In the second step, we rely on panel methods, helping us to answer the question of whether an increase in offshoring over time is correlated with a change in the permanent component of income risk. The offshoring measures are calculated at the industry level in a way similar to Feenstra and Hanson (1999). We use detailed yearly import matrices from input-output tables in combination with output and trade data.

From the outset, it is not clear whether offshoring increases or decreases income risk – especially with respect to the permanent component. On the one hand, there is empirical evidence at the industry level that offshoring tends to raise labor demand elasticities, which could lead to higher income risk, e.g. (Senses, 2010). On the other hand, this evidence is in part contradicted at finer levels of aggregation. Sascha and Marc-Andreas (2008) find offshoring to actually lower separation rates in employment at the firm level and Buch and Lipponer (2010) directly cast doubt on the claim that offshoring is responsible for changes in labor demand elasticities within multinational firms. It is important to note, however, that most studies within the rather inconclusive empirical literature are only indirectly related to the concept of income risk, and its permanent component in particular. As mentioned above, our analysis specifically tries to address a measure of "insecurity" that has clear and well-documented welfare implications – a characteristic generally attributed to the permanent component of income risk.

In addition to the mixed empirical results, theory recently suggested offshoring to be much less of a specter to workers than what is reflected in public anxiety and job loss fears. For example, Bergin, Feenstra, and Hanson (2009) show that offshoring has the potential to exert a dampening effect on economic volatility in the offshoring country if demand shocks are buffered by excess production activity in offshore plants. In other words, fluctuations are "exported" and firms face a less volatile domestic economic environment; and potentially their workers do as well.[1] It is also possible that offshoring induces what may be called a "composition effect". If offshoring is understood as trade in tasks, as in Grossman and Rossi-Hansberg (2008), and the tasks as such differ in their specific income volatilities, the relocation of certain tasks abroad might lead to aggregate changes in industry level income risk. If the offshored tasks are at the same time more volatile with respect to income, the average income risk of the tasks remaining onshore falls. A similar mechanism could arise from different income risk and trade patterns across skill groups. One could think of this effect as arising from firms effectively insuring themselves against fluctuations in economic activity. If institutional rigidities in the home market make adjustment costly, firms would be expected to relocate the activities most affected in places where adjustment is less costly. Such considerations seem particularly plausible in light of the European Union's enlargement to the East and Germany's location close to the new EU member states.[2]

The particular focus on offshoring also sets this paper apart from the recent literature studying effects on income risk arising from other forms of globalization such as import competition and tariff reductions. Krebs, Krishna, and Maloney (2010) analyse how tariff

[1] Yet, the opposite holds true for the receiving country. Volatility abroad (e.g. in Mexico for the case of US offshoring) is amplified.

[2] Note that this does not necessarily lead to an aggregate employment loss with less volatile, yet lower, overall employment at home since offshoring also triggers productivity effects possibly leading to net job creation (Kohler and Wrona, 2011). Additionally, Wright (2012) shows empirical evidence for output increases from offshoring.

reductions and the ensuing integration of the Mexican economy into the world market (in particular the North American part of it) affected income risk. They show income risk to increase as a response to trade liberalization, inducing the emergence of negative welfare effects. Yet, the Mexican economy may be considered a rather special case, in particular with regard to its proximity to the US and the existence of the "maquiladora" sector near the northern border.[3] Krishna and Senses (2009) set out to find the roots of income risk in the US labor market. Their prime candidate is import competition, which they show to raise the permanent component of income risk. As a robustness check these authors also employ an offshoring-type variable, which shows a negative coefficient in their estimations. We build on these intriguing insights and additionally put a clear focus on offshoring to explicitly contrast the concerns about increased risks often raised in the public debate with the potentially risk smoothing effects of a more efficient international allocation of production tasks.

Our findings contradict the general impression of offshoring as a major factor in raising long-run income volatility. They suggest an increase in offshoring is correlated with a *decrease* in the permanent component of income risk at the industry level. The effect is particularly visible for offshoring to non-OECD countries. For instance, a one percentage point rise in the overall offshoring intensity implies, on average, at least a 40% fall in permanent income risk compared to its mean value. Looking at offshoring as a particular type of international trade, we thus find the *opposite* effect in comparison with other studies relating more general measures of globalization to permanent income risk.

This chapter is structured as follows. The next section details the approach for estimating income risk, presents the data we use, and gives further insight into measuring the offshoring intensity at the industry level. In sections 3.3 and 3.4, we describe in detail the econometric specification and provide results on how income risk and offshoring are linked, respectively. A final section discusses potential welfare effects and concludes.

[3] In fact, this "maquiladora" sector has been shown in Bergin, Feenstra, and Hanson (2009) to have a particularly high volatility due to its role in the production sharing with the US economy.

3.2 Estimation and Calculation of Variables

3.2.1 Estimating Labor Income Risk

The approach taken in this chapter involves a two-stage procedure to first estimate the permanent component of individual income risk (stage one), and subsequently relating it to carefully constructed offshoring indices at the industry level (stage two). The goal of this section is to motivate our measure of income risk and to derive the corresponding estimation procedure. We follow the bulk of the literature and define income risk as the unpredictability of individual income, while referring to this variability from an ex-ante perspective (Carroll and Samwick, 1997; Meghir and Pistaferri, 2004). As such, income risk accompanies people whenever their future income is stochastic. In this sense, income risk is conceptualized as a deviation of the future income stream from its expectation, and is estimated as the variance of changes in the unexpected component of individual income.

In this chapter, as in most of the related literature, the estimated income risk has two components: a transitory one and a permanent one. This distinction is important since the two components have vastly different welfare effects. Transitory risk refers to the standard deviation of stochastic income changes without persistence. It describes the standard deviation of the purely random part of residual labor income and has no effect on the stochastic trend of residual individual earnings. Therefore, it could be effectively "self-insured" by individuals through saving and borrowing. Such unexpected transitory variation could be introduced by windfall labor income, extra or unusual (un)employment in the short-run, or changes in hours worked, which do not persist until the end of an individual's working life. Thus, following common theoretical considerations, there are no reasons for individuals to change their consumption and savings pattern, and therefore there are hardly any welfare effects (Levine and Zame, 2002). Moreover, in the standard estimation procedure, measurement error in the data is captured in transitory risk.

For the permanent component of income shocks, however, a different picture emerges. Permanent income risk, the standard deviation of permanent shocks, has profound effects on the consumption and savings decision of individuals in environments with imperfect insurance possibilities. Permanent income shocks reflect the stochastic trend of income. These shocks have persistent power over the remaining working period of individuals. This affects the present value of lifetime earnings and thus individuals "consume out" a certain amount of permanent shocks. Therefore, and in contrast to transitory risk, permanent income risk has an effect of future consumption and asset holding and, hence, affects individual welfare (Constantinides and Duffie, 1996; Krebs, 2003). Permanent shocks are observed as permanent events during workers' working life – for example, promotion beyond expectation or changes in employment resulting in a different matching quality of an individual's abilities and the job's requirements. A quantitative welfare analysis is conducted in Krebs (2003), where a baseline consumption-saving model with heterogeneous agents is considered. Individuals have log utility function with degree of risk aversion 1, and time discount factor is 0.96. In this case, the individual welfare, measured in equivalent consumption, declines by 1%, if permanent labor income risk increases by 2 percentage points, from 9% to 11%. Furthermore, if the degree of risk aversion rises to 2, that is, individual are more risk averse, then the cost of consumption-equivalent welfare is 2%, given the same increase in permanent labor income risk. Kuhn (2013) consider a more extreme case where half of the persistent income risk is removed. The consumption-equivalent welfare gain in such an environment with incomplete insurance markets rises up to 20%. Given this documented welfare effect, we focus our analysis on the connection between offshoring and the permanent component of labor income risk and disregard the transitory component. This is the same approach as in Krebs, Krishna, and Maloney (2010) and Krishna and Senses (2009).

The procedure for estimating the components of income risk starts with the identification of the unexplained component of individual income. This component is retrieved as the residual from standard Mincerian wage regressions of the following form:

$$y_{it} = \alpha_{jt} + \beta_t X_{ijt} + u_{ijt} \tag{3.2.1}$$

Note that the regressions are run year-by-year and include fixed effects for industries j. The control vector X_{ijt} includes the commonly used wage determinants such as age, age squared, education, marital status, nationality and firm-size as well as 2-digit occupation fixed effects. Notice that the estimation allows for changes in the returns to observable characteristics over time. An increase in the skill premium, for instance, is not regarded as contributing to income risk. The regressions are run on a restricted sample, which includes male individuals fully employed in manufacturing industries in West Germany. y_{it} is the natural logarithm of our income variable for individual i in year t, specified in more detail in the database descriptions below. The assumptions underlying the above model imply that individuals develop expectations about their future income from a projection based upon individually observable and predictable characteristics. Thus, the panel of all u_{ijt}'s represents the unexpected and stochastic component of individual earnings, which is idiosyncratic and unpredictable to individuals. We show exemplary results from this first stage regression in the appendix.

For the estimation of income risk and its components, we make the following assumptions. Suppose u_{ijt} has two components: a permanent one ω_{ijt} and a transitory one ϵ_{ijt}. Furthermore, assume ω_{it} to follow a random walk process.[4]

$$u_{ijt} = \omega_{ijt} + \epsilon_{ijt} \tag{3.2.2}$$

$$\omega_{ijt} = \omega_{ijt-1} + \eta_{ijt} \tag{3.2.3}$$

η_{ijt} and ϵ_{ijt} are the corresponding permanent and transitory shocks to individual earnings, respectively. ϵ_{ijt} is white noise, which has a temporary effect on labor income only and would vanish in the next time period. η_{ijt}, however, has persistence because ω_{ijt} follows a random walk process.

ϵ_{ijt} and η_{ijt} in each period are white noise and i.i.d distributed.

$$\epsilon_{ijt} \sim N(0, \sigma^2_{\epsilon,j,t}) \tag{3.2.4}$$

$$\eta_{ijt} \sim N(0, \sigma^2_{\eta,j,t}) \tag{3.2.5}$$

[4]The random walk assumption is not the only possible structure underlying the income process. For instance, other papers have suggested including a third, MA(1), component. Yet, as Krebs and Yao (2009) show, the permanent component of income risk is hardly affected by different assumptions on the income process. We therefore stick to the random walk assumption.

ϵ_{ijt} and η_{ijt} are independent for all leads and lags, that is, $cov(\epsilon_{ijt}, \epsilon_{ijs}) = 0, \forall t \neq s, cov(\eta_{ijt}, \eta_{ijs}) = 0, \forall t \neq s, cov(\epsilon_{ijt}, \eta_{ijs}) = 0, \forall t, s$. $\sigma^2_{\epsilon,j,t}$ and $\sigma^2_{\eta,j,t}$ are the variances of the transitory and permanent shocks to income, while $\sigma_{\epsilon,j,t}$ and $\sigma_{\eta,j,t}$ are the standard deviation of transitory and permanent labor income shocks, respectively.

Based on this assumed structure of the unexplained part of income, we can single out the permanent component of income risk. There are two different strategies usually employed in the literature. They differ in their assumptions on whether income risk can be assumed to be time-independent.[5] While assuming time-independence substantially simplifies the estimation, we will nevertheless calculate time-varying income risk. Recall that we are interested in measuring changes in income risk in order to link them to offshoring developments.

Estimation of Time-specific Income Risk

The assumption of time-independence taken by many related studies may seem to be a strong one. Shocks to permanent labor income in reality could differ across time periods due to, e.g. macroeconomic factors such as business cycle movements or trade related influences.[6] In fact, this is exactly what our paper is aiming to identify: How *changes* in permanent income risk can be explained. We therefore briefly describe the procedure for estimation of yearly values of permanent income risk. This subsection closely follows and extends the GMM estimation method employed in Krebs, Krishna, and Maloney (2010) for a short run panel data of 5 quarters to long run panel data of 10 years.[7]

For the changes in the unexplained income over time, we can generally write the n-year difference of u_{it} as

$$\triangle_n u_{it} = u_{it+n} - u_{it} = \eta_{it+1} + ... + \eta_{it+n} + \epsilon_{it+n} - \epsilon_{it} \tag{3.2.6}$$

[5] For instance, the assumption of time-constant income risk is taken by (see Carroll and Samwick, 1997; Gottschalk and Moffitt, 1994; Krishna and Senses, 2009).

[6] Storesletten, Telmer, and Yaron (2004) argue that the conditional variance of these permanent income shocks is counter-cyclical, increasing during contractions and decreasing during expansions. Krebs, Krishna, and Maloney (2010) find that trade policy has a significant effect on income risk.

[7] We drop the industry subscript j for expositional reasons in this subsection.

The variance of changes in the unexplained component of individual income between period t and $t+n$ thus is given by:

$$V[\triangle_n u_{it+n}] = \sigma^2_{\eta,t+1} + ... + \sigma^2_{\eta,t+n} + \sigma^2_{\epsilon,t} + \sigma^2_{\epsilon,t+n}. \quad (3.2.7)$$

Deviating from **?**, (3.2.7) is a larger and more complicated system of moment conditions due to more time periods included in the data. The estimation furthermore relies on additional moment conditions for the transitory component. In particular, it is assumed that this component of income risk is identical for the first and last two periods. Naturally, this also restricts the permanent component to being the same for those periods. According to **?**, the permanent component of income risk can be estimated from (3.2.7). $V[\triangle_n u_{it}]$ can be estimated as $\hat{V}[\triangle_n \hat{u}_{it}]$, where $\hat{u}_{it}$ are the predicted residual income terms from the first stage regressions (3.2.1) for individual i in time t. $\triangle_n \hat{u}_{it}$ is the n order difference of $\hat{u}_{it}$. In the estimation $V[\triangle_n u_{it}]$ is the dependent variable with observations $\hat{V}[\triangle_n \hat{u}_{it}]$. In (3.2.7), the left hand side observations are thus different combinations of $\sigma^2_{\eta,t}$ and $\sigma^2_{\epsilon,t}$, when n and t on the right hand side vary. If $\sigma^2_{\eta,t}$ and\or $\sigma^2_{\epsilon,t}$ are included in a certain $V[\triangle_n u_{it}]$, we put a 1 at the corresponding position in the matrix as the observation of the independent variable. If it is not included, we put a 0 instead. This renders our independent variables different combinations of entries with 1 or 0, with corresponding coefficients $\sigma^2_{\eta,t}$ or $\sigma^2_{\epsilon,t}$. We will have enough observations of $V[\triangle_n u_{it}]$, if there are sufficiently many time periods in our data. Then, we can run OLS regression as the last step of our GMM estimation, and the OLS estimators $\hat{\sigma}^2_{\eta,t}$ and $\hat{\sigma}^2_{\epsilon,t}$ are the unbiased estimates of $\sigma^2_{\eta,t}$ and $\sigma^2_{\epsilon,t}$. Hence, the result are industry level estimates for the two components of income risk.

3.2.2 Data and Implementation

Our main data source is a sample from official social security records from the German Employment Agency ("BA-Employment panel").[8]

[8]This study uses the factually anonymous BA-Employment Panel (Years 1998 - 2007). Data access was provided via a Scientific Use File supplied by the Research Data Centre (FDZ) of the German Federal Employment Agency (BA) at the Institute for Employment Research (IAB). For detailed information on the database, see Schmucker and Seth (2009).

We use information on income for individuals that stay within the same 2-digit manufacturing industry for at least two consecutive periods. On the one hand, we thus observe income variation for people who remain employed, yet face income changes due to wage changes and changes in other payments such as bonuses. On the other hand, we do not exclude individuals that lose their job once or several times as long as they are re-employed in the same industry at some point in our sample, irrespective of how long the unemployment spell is. In fact, temporary job loss is likely an important source of variation in income as job transition is often accompanied by a loss of occupation or employer-specific human capital leading to persistent changes in income. If an individual is employed in several industries over the sample period, we treat it as if it were two different individuals. That is, we use variation occurring during employment within an industry, but not between industries. We thus do not include variation based on individuals switching between industries or out of manufacturing in general. We admit that switching industries can be a source of income risk, yet one that is difficult to assign to any particular industry's aggregate income risk.[9] In terms of variation coming from within industry income changes, our approach is the same as in Krebs, Krishna, and Maloney (2010). Even within these limits, we will show that there is considerable variation in individual income and that a substantial part of this is reflected in permanent income risk. With the above in mind, the results for industry level income risk are best interpreted as representing an average expression for income risk akin to a representative industry characteristic.

The BA-Panel has a 10 year time period. More importantly, it has a large number of individual observations per industry and year. Thus, it allows us to estimate yearly income risk. It represents a 2 percent random sample drawn from official German employment records based on social security contributions for the years from 1998 to 2007. Income information in this case is log-monthly income and includes non-wage payments such as bonuses to the employees. We further restrict the panel to full-time employed, working age, male, West German residents. This still leaves us with a total of more than

[9] Krishna and Senses (2009) estimate income risk to be higher for individuals experiencing a transition from one industry to another when compared to individuals staying in one industry. We thus regard our estimations of income risk as representing a lower bound. If one were to link individual income risk arising from switching industries to offshoring one would have to be able to clearly and causally identify offshoring as the reason for the switch - a task that we cannot accomplish with our data at hand.

800,000 individual observations. We then proceed by applying the estimation approach for time-varying income risk. Note that our focus on permanent income implies that we only obtain estimates for the years up to 2005, since persistence of shocks is hard to observe when approaching the last years of the sample. Furthermore our estimation approach technically relies on the additional assumption that in the first two periods, income risk values are identical. This leaves us with observations for the years 1999 to 2005.

A common concern about such official German employment records is that the accompanying income information is censored at the legal threshold for social security contributions. This is the case in the present data as well. It is of potential concern that some income variation might be precluded from the analysis. Note that even an approach as ours, which relies on variation over time within industries when linking offshoring and income risk, is affected since the share of individuals at the income threshold is non-constant. On the other hand, the problem is less severe in manufacturing and particularly among low-skilled workers. These individuals simply rarely reach the threshold income.

Finally, the BA-panel data is quarterly in its original style. Yet, most of the income information is based on one entry per year only (so called "Jahresmeldung"). Thus, only yearly information can be used for our purposes. We do so by using monthly income data points from the December waves only.[10]

3.2.3 Income Risk: Results

In table 3.2.1, we present results derived from the more detailed BA-Employment panel. There is considerable heterogeneity across industries. The employment weighted average risk to residual monthly income stands at around 6 %. The estimates in table 3.2.1 are somewhat lower than those found in other studies, e.g. in Krebs, Krishna, and Maloney (2010). Note, however, that this latter study, as well as others, overestimates permanent income risk since it assumes all remaining income variation after 4 quarters to be permanent, whereas we treat changes from one year to the next as transitory still. Furthermore,

[10]The few remaining assumed imputed incomes, which are reported in the absence of a report by the employer, so-called "Fortschreibefälle", are deleted.

some studies rely on total household income which inherently has higher risk since it includes the outcomes of labor-leisure choice and substitution effects between household members. Additionally, it is plausible that by international standards the German labor market features lower income risk due to stronger institutions such as employment protection and wage bargaining coordination. In terms of identification of the connection to offshoring, we will not use levels information but either use industry fixed effects estimation or first-difference the data. Hence, we use information on the variation within industries rather than the heterogeneity in the levels between them.

Table 3.2.1: *Descriptives: estimated variances of permanent components*

Industry	Code	$\hat{\sigma}^2_{\eta,j}$	$\hat{\sigma}_{\eta,j}$	Change ($\hat{\sigma}^2_{\eta,j}$)
Food Products and Beverages	15	0.0043	0.0657	-0.0009
Textiles	17	0.0039	0.0621	0.0041
Wearing Apparel; Dressing	18	0.0055	0.0740	-0.0003
Wood Products, Except Furniture	20	0.0032	0.0564	0.0045
Pulp, Paper and Paper Products	21	0.0028	0.0528	-0.0028
Publishing, Printing	22	0.0040	0.0634	0.0018
Chemicals and Chemical Products	24	0.0022	0.0472	0.0043
Rubber and Plastic Products	25	0.0035	0.0593	0.0000
Other Non-metallic Mineral Products	26	0.0031	0.0558	0.0040
Basic Metals	27	0.0030	0.0543	0.0004
Fabricated Metal Prod., ex. Machinery	28	0.0038	0.0619	0.0006
Machinery and Equipment NEC	29	0.0034	0.0587	0.0011
Office Machinery and Computers	30	0.0052	0.0723	-0.0056
Electrical Machinery	31	0.0030	0.0546	-0.0018
Radio, Television and Communication	32	0.0039	0.0621	0.0034
Medical, Precision and Optical	33	0.0031	0.0554	0.0046
Motor Vehicles, Trailers	34	0.0026	0.0506	0.0005
Other Transport Equipment	35	0.0038	0.0614	-0.0039
Furniture; Manufacturing NEC	36	0.0040	0.0632	0.0000

Notes: Values for income risk are averages over time. Changes are first-to-last period differences of absolute values. The employment weighted industry average is 5.9 % ($\bar{\hat{\sigma}}_{\eta,j} = 0.0588$). Industry names may be incomplete.

3.2.4 Measuring the Offshoring Intensity

Offshoring is measured using input-output tables and trade data following a method introduced by Feenstra and Hanson (1999) and extended by Geishecker (2006). The offshoring intensities are calculated to represent the amount of an industry's intermediate inputs purchased from the same industry abroad in total industry output. This emphasizes the fact that the product could have likely been produced at home as well, and precludes situations in which traditionally imported goods count as offshoring. The offshoring intensity therefore is assumed to describe the outcome of multiple firm's make-or-buy decisions aggregated to the industry level. Note that it captures offshoring that occurs within as well as outside of a firm. In terms of the original notation introduced by Feenstra and Hanson (1999) our measure is the offshoring intensity in a "narrow" sense. Technically it looks as follows:

$$OFF_{jt} = \frac{IMP_{j^*t} \times \Omega_{j^*jt}}{Y_{jt}}. \tag{3.2.8}$$

Y_{jt} is output of j at time t. Ω_{j^*jt} describes the share of all imports from a specific 2-digit NACE 1.1 industry (j^*) abroad, used in the respective industry (j) at home. These shares are derived from yearly import matrices that are part of the input-output tables provided by the Statistical Office in Germany. IMP_{j^*t} are all imports from the foreign industry j^*, taken from the OECD STAN database, just as the output values. The data on imports and industry output are deflated using an aggregate manufacturing import price deflator and industry specific producer price indices, respectively.[11]

We furthermore differentiate between worldwide offshoring and offshoring to non-OECD countries. Here we again draw on the OECD STAN database and multiply the imports in (3.2.8) by the share of imports coming from non-OECD countries.[12] Note that this region-specific calculation of offshoring entails the common assumption of identical Ω_{j^*jt} for the two groups of countries, since the input-output tables do not hold any region

[11]The import price index is taken from Destatis and the industry specific indices for producer prices come from the EU KLEMS (March 2008) release.

[12]When calculating import shares for non-OECD countries, we had to rely on aggregates for industries 15-16; 17-19 and 21-22 and thus use the same non-OECD share for each industry within the respective group. Note, however, that this only applies to the non-OECD trade share and neither to total imports IMP_{j^*t} nor Ω_{j^*jt}.

specific information. The special distinction of non-OECD offshoring is meant to reflect the cost savings motive inherent in offshoring – a concept at the core of most theoretical approaches as well as the common public worries.

Table 3.2.2 shows offshoring intensities for the different manufacturing industries. Overall, worldwide offshoring has reached significantly higher levels than offshoring to low-income countries. Yet, starting from low values, growth rates are higher for offshoring to non-OECD countries. Additionally, we observe positive growth in all industries but wood products for non-OECD offshoring, while two more industries had a lower worldwide offshoring intensity in 2005 compared to 1999. Interestingly, for both measures the industries show quite some heterogeneity with respect to variations over time. This variation will be important in identifying the effect of offshoring on income risk later on.

3.3 Econometric Specification

We now turn to developing a suitable estimation strategy for an evaluation of the impact of offshoring on income risk. The general idea for the identification of the effect of offshoring on income risk is that exogeneous improvements in communication technology and the integration of many large Emerging Market Economies into the world trading system led to the observed surge in offshoring. Against this background it seems plausible to us to assume exogeneity of offshoring. Yet, some doubt remain. In particular, industry level shocks related to institutions or technology that affect both offshoring and income risk could pose problems to our identification strategy. We thus specify a model that tries to capture most of these possible disturbing influences. The data at hand permits a panel approach controlling for unobserved heterogeneity in two dimensions: industry and time. Industry-specific effects may well matter for the relationship between offshoring and income risk. Some industries are probably more inherently risky than others. This may be due to different demand elasticities for their products or unique employment structures in terms of jobs and tasks that can differ in their idiosyncratic risk. These industry fixed effects (FE) also play an important role in ameliorating endogeneity concerns. It could be argued for instance that labor market institutions are correlated with both income risk

Table 3.2.2: *Offshoring - descriptives*

		worldwide			non-OECD		
Industry	Code	1999	2005	change	1999	2005	change
Food Products and Beverages	15	3.04	3.97	0.93	0.44	0.63	0.20
Textiles	17	10.87	8.79	-2.08	4.27	4.42	0.15
Wearing Apparel; Dressing	18	10.81	12.94	2.13	4.24	6.51	2.27
Wood Products, Except Furniture	20	4.71	3.49	-1.22	1.13	1.11	-0.01
Pulp, Paper and Paper Products	21	7.95	8.87	0.92	0.40	0.54	0.14
Publishing, Printing	22	0.12	0.92	0.80	0.01	0.06	0.05
Chemicals and Chemical Products	24	13.28	13.73	0.45	0.87	0.98	0.12
Rubber and Plastic Products	25	0.95	1.48	0.53	0.09	0.19	0.10
Other Non-metallic Mineral Produc	26	1.98	2.08	0.10	0.22	0.34	0.12
Basic Metals	27	10.55	16.35	5.79	1.70	3.36	1.67
Fabricated Metal Prod., ex. Mach.	28	1.45	1.81	0.36	0.24	0.37	0.13
Machinery and Equipment NEC	29	6.36	7.35	0.99	1.12	1.94	0.82
Office Machinery and Computers	30	3.67	13.85	10.18	0.89	6.06	5.18
Electrical Machinery	31	6.50	6.57	0.07	1.17	1.52	0.35
Radio, Television, Communication	32	7.06	19.75	12.68	1.67	6.99	5.32
Medical, Precision and Optical	33	4.12	4.52	0.40	0.62	0.71	0.09
Motor Vehicles, Trailers	34	7.93	10.21	2.29	0.30	0.55	0.24
Other Transport Equipment	35	21.09	13.03	-8.06	1.02	1.18	0.16
Furniture; Manufacturing NEC	36	7.17	9.42	2.25	2.29	3.81	1.51

Notes: Values are calculated according to $OFF_{jt} = \frac{IMP_{j^*t} \times \Omega_{j^*jt}}{Y_{jt}}$ (see text) and represent percentage values. Changes are absolute changes.

and offshoring. Strong unions could be responsible for less fluctuations in income and dampened offshoring activities at the same time, introducing bias in our results. Yet, variables capturing labor market institutions usually have very limited time series variation. Hence, as long as these characteristics are specific to an industry and do not vary over time, a fixed effects setup will capture this type of unobserved heterogeneity. The same logic holds true for using an approach based on first-differencing (FD) the data. In our case, we show results from both FE and FD estimation. In addition to industry fixed effects, we can employ time fixed effects for time-varying factors that are unobservable to us, such as business cycle effects at the country or world level, which capture this variation as long as it is uniform across industries. Some further remaining variation that could possibly bias our results will be picked up by the control variables included in the model. In particular, these will be the industry specific and time-varying research and development spending share in value added (to capture industry specific technology shocks) and the labor share of income (as an approximation to time-varying union influences). Furthermore, we include measures for the export share in production (capturing another dimension of dependency on international output fluctuations) and the import penetration ratio (to provide for a comparison with the literature – Krishna and Senses (2009), in particular).[13] A further point deserves attention. Given the structure of our data set, we have to be careful when calculating standard errors (Krebs, Krishna, and Maloney, 2010). Our dependent variable $\hat{\sigma}^2_{\eta jt}$ is by itself the outcome of an estimation at the industry level. While, according to Krebs, Krishna, and Maloney (2010), this does not bias the regression coefficients in the second stage, with different standard errors across industries from the first-stage estimations, we are facing heteroscedasticity. We therefore follow Krebs, Krishna, and Maloney (2010) and Krishna and Senses (2009) and report robust standard errors based on a Huber-White-correction.

With the above considerations in mind, we arrive at the following empirical model. The model also allows for the inclusion of lagged effects of offshoring on income risk if we chose $N > 0$:

$$\sigma^2_{\eta jt} = \Sigma^N_{i=0}\beta_i OFF_{jt-i} + \gamma X_{jt} + \phi_j + \varphi_t + \nu_{jt} \tag{3.3.1}$$

[13] All these data are retrieved from the OECD STAN database.

In this model, X_{jt} is the control vector. ϕ_j represents the industry fixed effects and φ_t are indicators for years (t), respectively. ν_{jt} represents the model's independent error term.

3.4 Results: Offshoring and Income Risk

3.4.1 Main Results

In this subsection we present the results based on the above model - estimated once as the FE model in equation (3.3.1) and once as an alternative FD specification. Table 3.4.1 reveals the main result: an increase in offshoring correlates with lower permanent income risk. We present results for one-year lagged values of offshoring as explanatory variables, because we do not find any significant contemporaneous correlation. This points to the impact offshoring has on income risk as the outcome of a change in how employment and production are organized internationally. Recall, that income risk measures shocks from an ex-ante perspective, i.e. it describes how shocks at a given time play out over future periods. We can therefore state that, on average, workers in an industry that shifts more tasks abroad will subsequently face less severe shocks to permanent income.[14]

We find negative coefficients on the different offshoring variables throughout columns (1) to (4) in table 3.4.1. The results are statistically significant at conventional levels for the preferred specification including all controls. The coefficient value in column (3) implies that, on average, an increase in the offshoring intensity by one percentage point decreases the variance of permanent component by -0.00035. Compared to its (employment weighted) mean across industries and over time of 0.00345, this represents a decrease of a little more than 10% for every percentage point increase in the overall offshoring intensity. The effect is stronger and more precisely estimated for offshoring to non-OECD countries. The results in column (4) show the effect to be roughly four times the size of the corresponding value for worldwide offshoring. This type of offshoring, however, shows a smaller absolute increase over time: roughly half a percentage point.
The coefficients of the control variables, which were mainly included to control for possible

[14]This does not say anything about the possible effects of displacements at the margin of offshoring. Yet, according to some recent studies, offshoring does not seem to be a major cause of overall job-loss at the industry level (Harrison and McMillan, 2011; OECD, 2007).

sources of bias, only hold a small bit of additional information. The coefficient for the import penetration ratio is found to be positive, yet not within the conventional significance bounds. We take this finding to weakly confirm the result found in Krishna and Senses (2009) for the US also for Germany. The R&D share does not have any statistically significant effect. It seems as if this variable is unable to approximate technological change above and beyond common aggregate-level trends captured by the time effects. The export share in production does not emerge as a main influence on industry level income risk either.

Turning to the results based on a first-difference estimation, the overall picture is confirmed. We find negative coefficients of offshoring throughout, albeit with stronger significance for offshoring to non-OECD countries. The coefficients are mostly of a similar magnitude as well, with the one on non-OECD offshoring in the fully specified model being higher than before. In summary, our results from both models show a negative effect of offshoring on the permanent component of income risk. With respect to worldwide offshoring the picture regarding statistical significance is a bit mixed. Offshoring to non-OECD countries on the other hand has a particularly strong and significant negative effect.

3.4.2 Robustness

In addition to the various specifications above, we further test the robustness of our results by manipulation the estimation sample. We focus on the negative link between income risk and offshoring to non-OECD economies since this connection was stronger and statistically significant in the main tables. Results for worldwide offshoring are as above, negative but with mixed statistical significance. Table 3.4.3 shows the results of our robustness analysis. In the first column, we drop the industry with the largest change in the permanent component of income risk (30). In the following column, we take the industry with the highest increase in non-OECD offshoring (32) out of the sample. Subsequently, we use an alternative data source for the offshoring measures - the data

Table 3.4.1: *Results from **fixed effects** regression, 1999-2005*

variance of permanent component of income risk	1	2	3	4
1-year lagged offshoring intensity (world)	-0.0333		-0.0350*	
	(0.0201)		(0.0186)	
1-year lagged offshoring intensity (non-OECD)		-0.148*		-0.139**
		(0.0760)		(0.0620)
export-share in production			-0.0291	-0.0242
			(0.0181)	(0.0181)
import penetration			0.0260	0.0241
			(0.0192)	(0.0200)
R&D share in value added			0.0185	0.0042
			(0.0114)	(0.0095)
labor share			0.00031	0.0021
			(0.0052)	(0.0051)
year fixed effects	yes	yes	yes	yes
industry fixed effects	yes	yes	yes	yes
Observations	114	114	108	108
R-squared	0.307	0.428	0.465	0.518
Number of sectors	19	19	18	18

Notes: Estimation is by fixed effects. The coefficient values on the offshoring measures are to be understood as follows: a one unit change in offshoring (= percentage point change) corresponds to a $\hat{\beta}/100$ change in the variance of persistent changes in the unexplained component of income (= permanent income risk). Industry 36 has incomplete data coverage which leads to a slightly reduced number of observations in some cases. Cluster-robust standard errors are shown in parentheses; $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 3.4.2: *Results from* ***first-difference*** *regression, 2000-2005*

variance of permanent component of income risk	1	2	3	4
1-year lagged offshoring intensity (world)	-0.0308*		-0.0376	
	(0.0172)		(0.0272)	
1-year lagged offshoring intensity (non-OECD)		-0.148**		-0.202*
		(0.0558)		(0.0979)
export-share in production			-0.0213	-0.0289
			(0.0301)	(0.0370)
import penetration			0.0185	0.0270
			(0.0253)	(0.0309)
R&D share in value added			0.0004	0.0011
			(0.0143)	(0.0162)
labor share			0.0076	0.0106
			(0.0068)	(0.0078)
year fixed effects	yes	yes	yes	yes
Observations	95	95	90	90
R-squared	0.229	0.260	0.299	0.367
Number of sectors	19	19	18	18

Notes: Estimation is by OLS after first-differencing the data. The coefficient values on the offshoring measures are to be understood as follows: a one unit change in offshoring (= percentage point change) corresponds to a $\hat{\beta}/100$ change in the variance of persistent changes in the unexplained component of income (= permanent income risk). Industry 36 has incomplete data coverage which leads to a slightly reduced number of observations in some cases. Cluster-robust standard errors are shown in parentheses; $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

made available online by Ingo Geishecker.[15] Finally, we also constructed income risk measures from a different sample wave (June) of our income data. We find the results to be highly robust to these data alterations. All offshoring coefficients are negative again. Statistical significance is even stronger for all altered samples.

Table 3.4.3: *Results from robustness analyses*

variance of permanent component of income risk	2	4	6	8
	without sec. 32	without sec. 30	Geishecker offsh.	June wave
1-year lagged offshoring intensity (non-OECD)	-0.0755***	-0.1945***	-0.3834***	-0.0845***
	(0.0150)	(0.0379)	(0.0474)	(0.0104)
export-share in production	-0.0179	-0.0524***	-0.0556***	-0.0590***
	(0.0149)	(0.0100)	(0.0180)	(0.0161)
import penetration	0.0206	0.0479***	0.0531**	0.0572***
	(0.0125)	(0.0104)	(0.0201)	(0.0188)
R&D share in value added	-0.0001	-0.0001	-0.0001	0.0001
	(0.0001)	(0.0002)	(0.0002)	(0.0001)
labor share	0.0057	0.0064	0.0051	-0.0021
	(0.0048)	(0.0056)	(0.0080)	(0.0036)
year fixed effects	yes	yes	yes	yes
industry fixed effects	yes	yes	yes	yes
Observations	75	75	80	80
R-squared	0.4323	0.8175	0.6811	0.5109
Number of sector	15	15	16	16

Notes: Estimation is by fixed effects. The coefficient values on the offshoring measures are to be understood as follows: a one unit change in offshoring (= percentage point change) corresponds to a $\hat{\beta}/100$ change in the variance of persistent changes in the unexplained component of income (= permanent income risk). Industry 36 has incomplete data coverage which leads to a slightly reduced number of observations in some cases. In column (6) offshoring is to Asian countries instead of all non-OECD countries since this is the closest definition available in the data provided by Ingo Geishecker (`http://www.uni-goettingen.de/en/99958.html`) available until 2004 and for 17 of the industries used here. Cluster-robust standard errors are shown in parentheses; $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

3.5 Conclusion and Welfare Considerations

The analysis in this paper presents offshoring as a source of changes in permanent income risk at the industry level. Income risk is an important factor in determining the consumption, savings and thus welfare patterns in an economy. We single out offshoring as a potential influence, given the anxiety it regularly stirs up in the public debate as

[15]The data are available at (`http://www.uni-goettingen.de/en/99958.html`). Further information on their exact construction can be found in Geishecker (2006).

well as its large role in international trade transactions. Within the limits of our available data, we seeks to answer whether the fears regarding income insecurity often associated with it are justified. We find that they are not. On the contrary, within manufacturing industries, increased offshoring is associated with a *decrease* in the permanent component of income risk.

In our empirical analysis, we first estimate industry level income risk from individual level data, isolating the welfare-relevant permanent component. We then link it to offshoring at the industry level in a panel framework. We find offshoring to have a negative and (mostly) statistically significant effect on income risk for employees within industries in manufacturing. Furthermore, there is strong evidence for a differentiated impact across destination regions, with a stronger than average effect for offshoring to non-OECD countries. This is expected as offshoring facilitates the allocations of different tasks. Put differently, home producers tend to offshore the more volatile part of their production to foreign countries. In our case, German manufacturing sectors have offshored the production or occupations with more uncertainty and fluctuations to non-OECD countries while the occupations with less volatility have been kept.

However, with respect to welfare implications the results are less straight forward. Clearly, taken by itself, a reduction in income risk brought about by a higher offshoring intensity would imply a positive welfare effect. Yet, this effect might not be the only welfare-affecting change. Two points deserve particular attention. First, the wage level still matters as well. Individuals may have a smaller benefit if risk decreases but this comes as a trade-off with lower average wages. Yet, on an aggregate level, this is not necessarily to be expected. **?** theoretically show that the wage effects of offshoring are ambiguous, and empirical evidence often documents relative wages for different skill groups to change while *overall* wages are hardly affected. Leaving considerations with respect to a skill-specific effect to further research, we are therefore leaning towards the conclusion that lower income risk does not come at the cost of lower average wages in manufacturing.

The second possible concern is related to employment levels. A shift of more volatile occupations (or tasks) abroad may change average income risk in the home country at the expense of lower overall employment levels. The volatile jobs would move offshore and – as a consequence – the remaining ones show a lower average income risk. Yet, it is hard to argue that this situation is desirable from an aggregate perspective if overall employment falls due to replacement of foreign for domestic jobs. Ideally, if composition effects are at work, one would want the home employment to stay constant or to grow due to productivity effects from offshoring and the workers whose tasks are moved offshore would find re-employment in less volatile jobs. There are some hints that offshoring is not responsible for falling employment levels in manufacturing. For instance, the OECD states that "*(...) the industrial sectors that have most downsized their workforce are not the ones that have most engaged in offshoring. Offshoring does not therefore emerge as a major cause of job losses.*" (OECD, 2007). This finding has recently been confirmed by Harrison and McMillan (2011) for the United States, who find most of the manufacturing employment decline to be a result of capital-labor substitution rather than international labor reallocation. We therefore conclude on a slightly optimistic tone. If offshoring lowers the permanent component of income risk, while average wages do not fall and overall employment levels stays widely unaffected, there may be positive effects on welfare.

Finally, considering our sample includes only full time working males in manufacturing sectors the effect of offshoring on labor income risk and welfare implication on other groups is not fully clear. For example, female workers and part time workers. The extension of our research is to investigate further on the link of offshoring and labor income risk for other groups of workers.

3.A Appendix

This additional table (3.A.1) shows results from the first stage regression generating the income residuals. Coefficients for fixed effects are not shown. The results are based on BA data with imputed wages from the cross section for the year 2005. Results for any other year look nearly the same. All coefficients have the expected sign and significance. That is, income grows with age, skill, firm size, etc.

Table 3.A.1: *First stage wage regression for 2005*

ln wage	1
Age	0.0374***
	(0.0001)
Age-squared	-0.0004***
	(0.0000)
Foreign nationality	0.0104***
	(0.0034)
Firm size	0.0607***
	(0.0006)
Medium-skilled	0.0503***
	(0.0050)
High-skilled	0.1640***
	(0.0032)
Constant	7.1712***
	(0.0040)
Industry fixed effects	yes
2-digit occupation fixed effects	yes
Observations	76,830
R-squared	0.5730

Chapter 4

Can Institutional Reform Reduce Job Destruction and Unemployment Duration? Yes, It Can

4.1 Introduction

This paper investigates how labor market policies affect the unemployment rate through its two defining factors, the duration of unemployment spells and job destruction rates. To this aim, we look at search theory's unemployment equilibrium condition as an Iso-Unemployment Curve (IUC). The IUC represents the locus of job destruction rates and expected unemployment durations rendering the same unemployment level. A country's position along the curve reveals its preferences over the destruction-duration mix, while its distance from the origin indicates the unemployment level at which such preferences are satisfied. We next provide micro-foundations for the link between destructions, durations and policy variables. This allows us to explore the relevance of institutional features using a sample of 20 OECD countries over the period 1985-2008.

The empirical literature investigating the influence of labor market institutions on overall unemployment rate is sizable (Blanchard and Wolfers, 1999; Nickell, Nunziata, Ochel, and Quintini, 2001). Equally numerous are the studies splitting unemployment into job creation and job destruction flows (Blanchard and Portugal, 1998; Elsby, Hobijn, and Sahin, 2013; Shimer, 2012). This work connects these two strands of the literature by investigating how labor market policies shape both job separations and unemployment spells, which together determine the overall unemployment rate in the economy. The IUC schedule used in our analysis is novel and is motivated by the need to understand the nature of unemployment, as essentially coming from destructions, durations or a combination of both these factors. This can help clarify whether policy makers should focus primarily on speeding up workers' reallocation across job positions rather than protecting them in the workplace.

One fundamental question raised in this context is whether countries with dynamic labor markets significantly outperform countries with more stagnant markets. By dynamic (stagnant) we mean labor markets displaying high (low) levels of workers' turnover in and out of unemployment. Is it the case that countries featuring high job destruction rates but brief unemployment spells tend to display lower unemployment rates than labor markets characterized by limited job destruction but longer unemployment durations? And how do institutional features shape destructions and durations?

The remainder of the chapter is organized as follows. The next section looks at the empirical evidence on unemployment inflows and durations. Section 4.3 introduces the concept of IUCs, which is the backbone of our theoretical construct. Section 4.4lays out the empirical strategy, discusses the empirical results and uses the estimated model for simulation purposes. Section 4.5 concludes the paper.

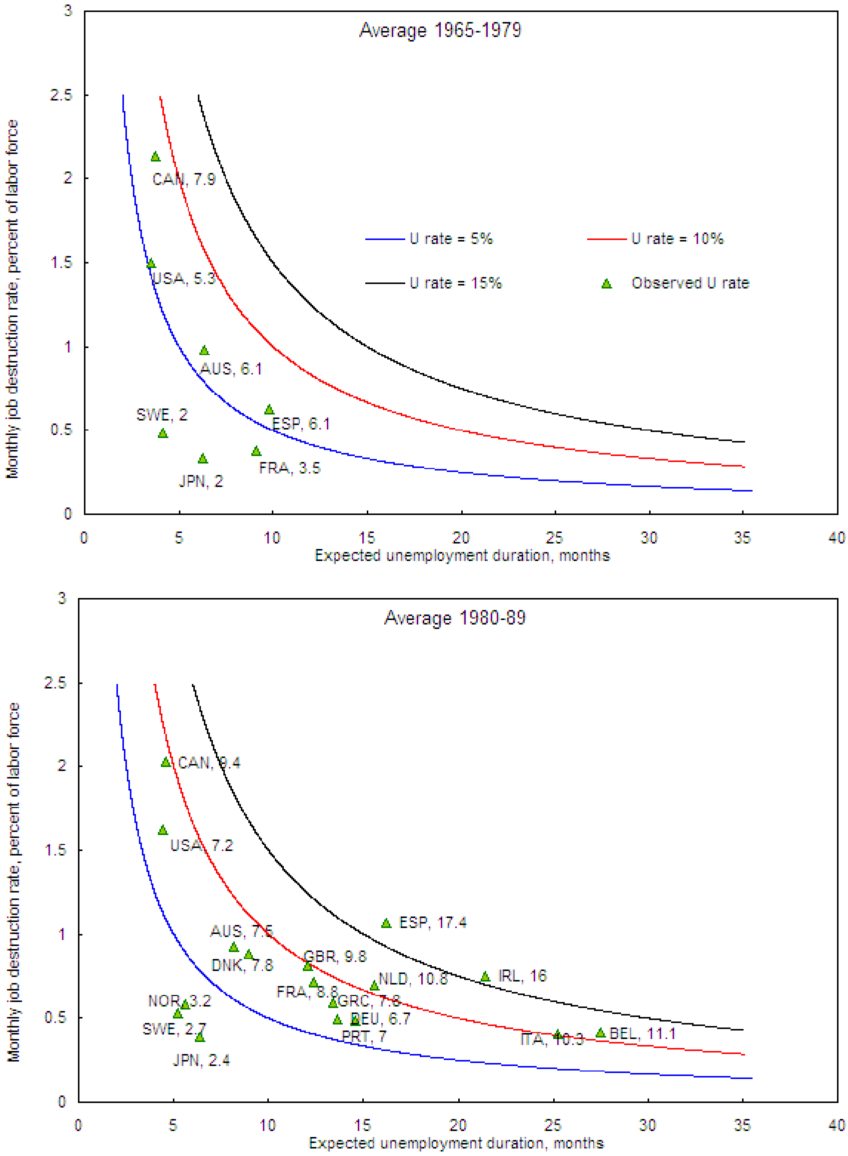
4.2 Inflows and Durations: The Facts

Search theory allows for a decomposition of the steady-state unemployment rate into the product of the number of workers that leave employment as percent of the labor force and the expected duration of unemployment. As discussed in Section 4.3, it is possible to interpret this equilibrium relationship as an "Iso-Unemployment Curve" (or IUC), reflecting the different combinations of inflows and durations compatible with the same unemployment level. It is precisely the reasons underlying the position of countries in the space of IUCs that is investigated in this paper. Before we do this, it is interesting to see where countries actually stand in the IUC schedule.

To this aim, Figures 4.2.1 and 4.2.2 decompose annual unemployment rates into the yearly destruction rates and the expected duration of unemployment for 20 OECD countries over the four decades ranging from 1965 to 2009. The sample comprises eleven euro-area countries (Austria, France, Germany, Belgium, Italy, Spain, Greece, Portugal, Ireland, the Netherlands, and Finland) alongside nine other OECD countries (US, Canada, Australia, Japan, UK, Denmark, Norway, Sweden, and Switzerland). The empirical job destruction rates and unemployment durations have been extracted from available data on unemployment rates following the methodology developed by Shimer (2012). The detailed procedure is described in the appendix of this chapter.

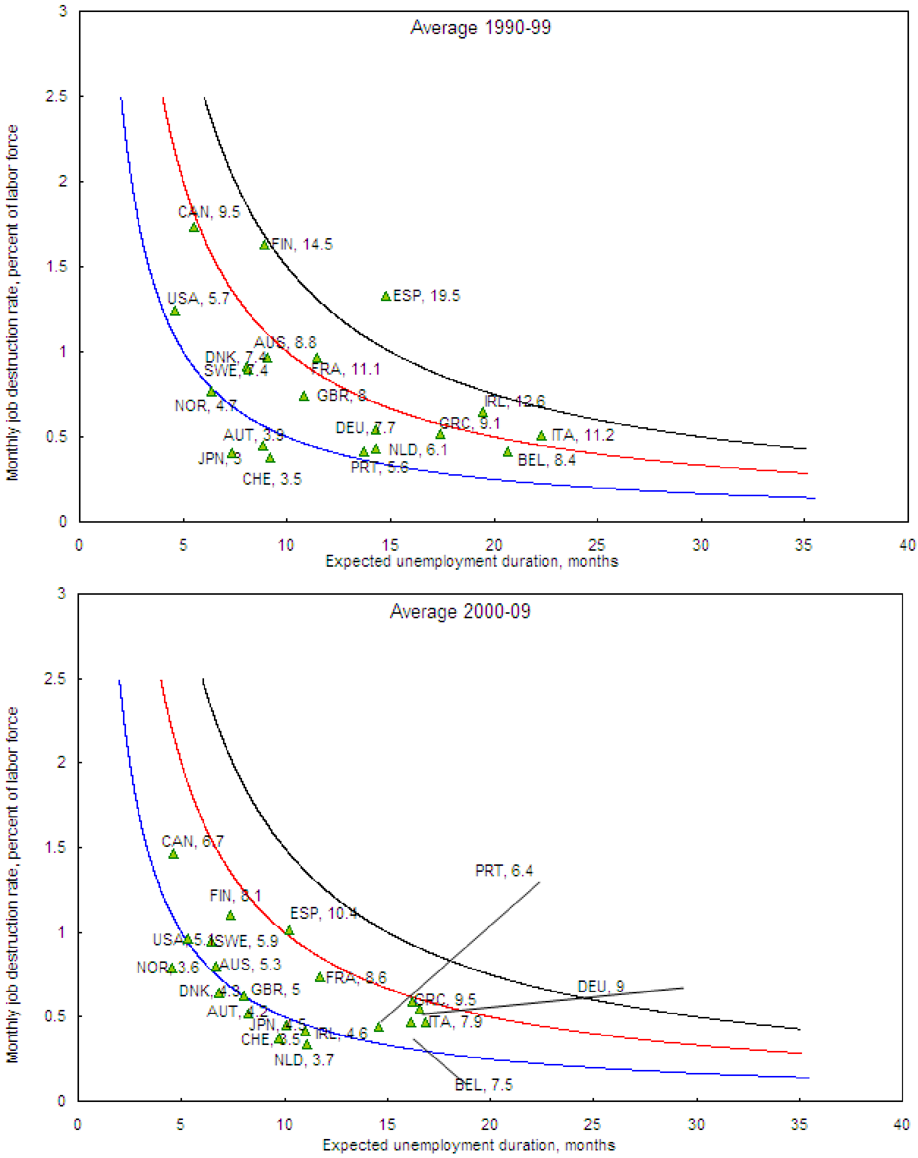
Figures 4.2.1 and 4.2.2 reveal a huge variation in the unemployment inflow rates and durations across countries. With both low inflows and durations, Norway, Sweden, Japan, Switzerland, and to a lesser extent Austria, display the lowest unemployment levels. Reflecting a combination of brief unemployment spells and high inflow rates, North America is situated on IUCs with low-to-moderate unemployment levels. Belgium, Italy, Ireland and Greece have consistently featured low job destruction rates, coupled with relatively long durations and high unemployment rates. Australia, France, Germany, the Netherlands, Denmark and the UK, characterized by intermediate levels of both inflows and durations, tend to post higher unemployment rates than the first group, while their position relative to North America is ambiguous. Reflecting elevated inflows and durations, Spain tends to be located on IUCs with the highest unemployment levels (bar the seventies). Significant variation in the relative importance of the two variables of interest makes it difficult to characterize unemployment's behavior in Finland.

Figure 4.2.1: *Unemployment duration and job destruction rates in 20 OECD Countries, 1965-1989*



Source: Authors' estimates of job destruction rates and expected unemployment duration based on OECD data and the methodology described in the appendix of this chapter.

Figure 4.2.2: *Unemployment duration and job destruction rates in 20 OECD Countries, 1990-2009*



Source: Authors' estimates of job destruction rates and expected unemployment duration based on OECD data and the methodology described in the appendix of this chapter.

4.3 The Model

4.3.1 Equations

Theories of unemployment may be divided into two broad groups, depending on whether they are based on stocks or flows. The specification of the labor market adopted in this paper focuses on flows. It builds on the standard matching model developed by Pissarides (2000). Suppose that the size of the labor force is constant[1]. We let u_t and v_t at time t denote the unemployment and vacancy rates, both expressed as percent of the labor force. The number of job matches taking place per unit of time is assumed to follow a Cobb-Douglas specification of the form

$$m_t = \mu v_t^{1-\rho} u_t^{\rho} \tag{4.3.1}$$

where μ denotes the degree of efficiency in the matching technology and $\rho \in (0,1)$ represents the unemployment elasticity of job matches. The job vacancies and unemployed workers that are matched at any time t are randomly selected from the sets u_t and v_t. Hence the hazard rate out of unemployment follows a Poisson process with rate

$$f_t = \frac{m_t}{u_t} = \mu \theta_t^{1-\rho} \tag{4.3.2}$$

where $\theta_t = v_t/u_t$ is the number of vacancies per unemployed workers or labor market tightness. This generates an exogenous hazard rate into employment equal to $f_t u_t$.

The flow into unemployment results from job-specific productivity shocks arriving to firms at Poisson rate s_t. When the shock arrives, the net product of the job changes to some new value drawn from a general probability distribution $G(x_t)$,where $0 \leq x_t \leq 1$ and $w(x_t)$ stand for the job's productivity and its corresponding wage. Let us denote by $J(x_t)$ the

[1]This assumption is needed to apply the methodology developed by Shimer (2012) to measure job creation and job destruction flows. This is consistent with the view that, compared with workers' flows within the labor force, flows in and out of the labor force play a lesser role in explaining unemployment movements. Once this assumption is relaxed, one cannot use publicly available data to construct all possible flows between employment, unemployment and inactivity. Micro-data on individuals' employment status should be used instead. Data availability issues prevent us from taking this route.

firm's expected profit from an occupied job. The firm chooses a reservation productivity R_t (with $J(R_t) = 0$) and destroys jobs whose productivity fall below it because of the shock. In words, job destruction operates when the reservation productivity is such that the expected profit from an occupied job is 0. This generates an endogenous hazard rate into unemployment equal to the product of the fraction of firms hit by a shock s_t, and the probability that a shock is below reservation $G(R_t\ )$, with associated job destruction of $s_t G(R_t\ )(1-u_t)$.The out-of-steady-state unemployment dynamics is therefore given by

$$u_{t+1} - u_t = s_t G(R_t\ )(1-u_t) - f_t u_t \tag{4.3.3}$$

At equilibrium the unemployment rate is constant, thus

$$sG(R)(1-u) = fu \tag{4.3.4}$$

with the interpretation that the mean number of workers who enter unemployment is equal to the mean number of workers who leave unemployment. Put another way,

$$\frac{1}{f}(1-u) = \frac{1}{sG(R\ )}u \tag{4.3.5}$$

Substituting (4.3.2) into (4.3.4) and solving for the u yields

$$u = sG(R)(1-u)\frac{1}{\mu\theta^{1-\rho}} \tag{4.3.6}$$

which tells us that the equilibrium unemployment rate is the product of the hazard rate into unemployment $sG(R)$ times the expected duration of unemployment $(1-u)\frac{1}{\mu\theta^{1-\rho}}$.

Solving the model requires obtaining the equilibrium conditions for the two endogenous R and θ, to which we now turn. Suppose that job creation satisfies the zero-profit condition[2], that job destruction operates when $J(R) = 0$, and that the wage sharing rule is set in a Nash bargaining fashion. Let $\beta \in (0,1)$ be the labor's relative strength in wage bargaining, z the unemployment income, c the hiring and firing costs of firms, κ an indicator of the

[2]As the model allows for regulatory costs, the zero-profit condition implies that firms' profits are dissipated by such costs.

regulatory stance in the market, tw the tax wedge, and r the rental cost of capital. One can then express labor market tightness and reservation productivity (characterizing job creation and job destruction respectively) as a function of the exogenous variables in the model, that is

$$\theta = \theta(\beta, c, \kappa, tw, s, r, \mu, R) \tag{4.3.7}$$

$$R = R(\beta, c, z, \kappa, tw, s, r, \mu, \theta) \tag{4.3.8}$$

It is convenient to represent this system in the space (R, θ) while viewing the remaining parameters as constant (Figure 4.3.1). The model behaves according to the following logic:

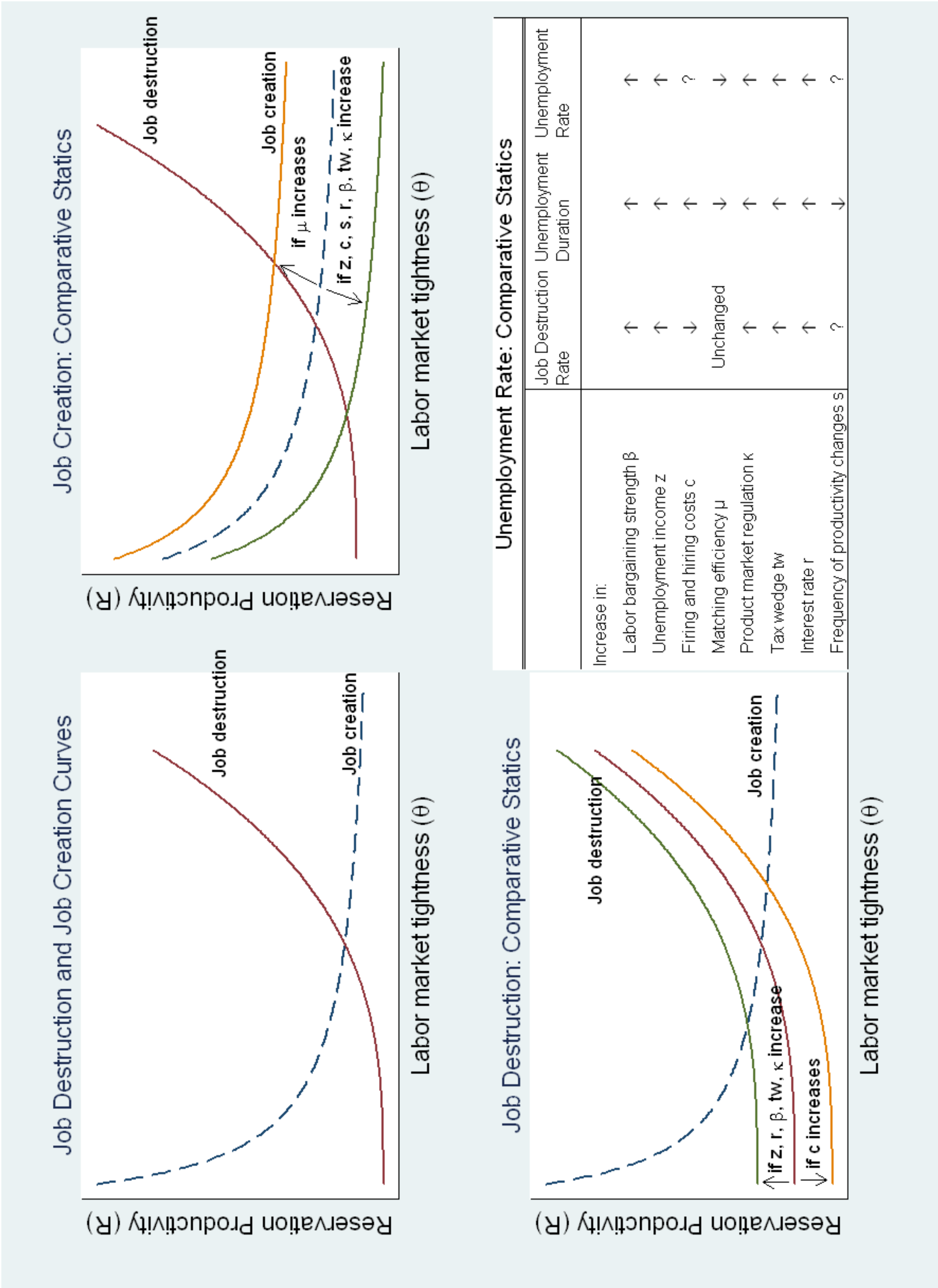
- Expression (4.3.7) on job creation corresponds to the free-entry condition postulated by search models and reflects that firms post vacancies as long as the marginal recruiting cost per vacancy is equal to the expected value of holding it. At equilibrium, the expected gain from a new job to the firm must be equal to the expected hiring cost paid by the firm. It slopes down in the space (R, θ) because at higher reservation productivity , the expected life of (and gain from) a job is reduced. Firms create fewer jobs as a result, leading to a fall in market tightness θ.

- Expression (4.3.8) on job destruction slopes up in the space (R, θ) because at higher θ the worker's outside opportunities are better and wages are higher. Higher wages translate into a lower net product of labor. Thus more marginal jobs are destroyed. A noteworthy property implicit in the equilibrium condition (4.3.8) is that the reservation productivity R is less than the reservation wage $w(R)$. This implies some labor hoarding in equilibrium: because job productivity might change, the firm keeps some currently unprofitable jobs occupied.

4.3.2 Model's predictions

Comparative statics implied by equations (4.3.7) and (4.3.8) is as follows (Figure 4.3.1):

Figure 4.3.1: *Model's predictions*



Increase in:	Job Destruction Rate	Unemployment Duration	Unemployment Rate
Labor bargaining strength β	↑	↑	↑
Unemployment income z	↑	↑	↑
Firing and hiring costs c	↓	↑	?
Matching efficiency μ	Unchanged	↓	↓
Product market regulation κ	↑	↑	↑
Tax wedge tw	↑	↑	↑
Interest rate r	↑	↑	↑
Frequency of productivity changes s	?	↓	?

Source: The authors.

- At given θ,

$$\frac{dR}{d\beta} > 0, \frac{dR}{dz} > 0, \frac{dR}{dtw} > 0, \frac{dR}{d\kappa} > 0, \frac{dR}{dc} < 0, \frac{dR}{ds} \lesseqgtr 0, \frac{dR}{dr} > 0 \tag{4.3.9}$$

- At given R,

$$\frac{d\theta}{d\beta} < 0, \frac{d\theta}{dz} < 0, \frac{d\theta}{dc} < 0, \frac{d\theta}{dtw} < 0, \frac{d\theta}{d\kappa} < 0, \frac{d\theta}{ds} < 0, \frac{d\theta}{dr} < 0, \frac{d\theta}{d\mu} < 0 \tag{4.3.10}$$

Consider now the influence of parameter changes on the job destruction and job creation equilibrium values, the latter being inversely related to our second variable of interest, the expected duration of unemployment. Bearing in mind that unemployment will change only if the new flows implied by the change in R and θ are not equal, the model's predictions can be summarized as follows:

- Both a higher unemployment income z and a higher labor share in bargaining β raise reservation wages. This reduces the expected gain from a job, shrinking job creation (increasing duration) and increasing job destruction. As a result, the unemployment rate will be higher in the new equilibrium.

- An increase in the tax wedge tw makes leisure more attractive relative to work, putting upward pressure on bargained wages. This renders marginal workers unprofitable and, all else being equal, more jobs are destroyed. At the same time, higher wages reduce the expected value from job matches to the firm, thereby inhibiting their recruiting activities. Higher job destruction and lower job creation therefore lead higher unemployment rates.

- Stricter regulation κ tends to lower labor productivity in equilibrium, hence dampening job creation. A deterioration in competition-restraining regulations conveys higher mark-ups to producers, who will now seek to sell their goods at higher prices. As there is excess supply at the new prices, some unprofitable firms will exit the market and jobs will be destroyed until a new equilibrium is reached characterized by lower production and employment levels. Overall, the equilibrium unemployment rate is hypothesized to increase following a deterioration in the regulatory conditions.

- Job creation increases (duration falls) if productivity changes more frequently (higher s). Job destruction is subject to two opposing forces. On the one hand, job destruction increases because there are more shocks on average, but on the other hand, it decreases because firms are now more willing to hold on to labor if they expect a quick arrival of better conditions. Thus the effect of increased s on unemployment is ambiguous.

- Higher interest rate r reduces job creation (increases duration) as future profits on new jobs are discounted more heavily. At the same time, firms respond to higher interest rates by destroying more jobs. This is so because the option value of keeping unprofitable jobs is lower, given that the returns from a productivity change only accrue in the future. Less job creation and more job destruction combine into higher unemployment rates.

- By reducing the expected duration of vacancies, higher matching efficiency μ reduces hiring costs for firms, increasing job creation (reducing duration) and bringing down the unemployment rate.

- Higher hiring and firing costs c lower the expected gain from a job and reduce job creation. Likewise, larger firing costs make firms more conservative in their firing decisions and limit job destruction. This makes for a more stagnant labor market, but the effect on lower inflows and higher durations on equilibrium is itself ambiguous.

4.4 Estimating the Model

4.4.1 Data and Empirical Strategy

In order to assess the impact of labor policies on job destruction rates and the duration of unemployment spells, we estimate empirical versions of (4.3.7) and (4.3.8) of the form:

$$Job\ destruction\ rate_{it} = fe_i + d_t + \sum_j b_j x_{ijt} + OG_{it} + \varepsilon_{it} \tag{4.4.1}$$

$$Expected\ duration\ of\ unemployment_{it} = fe_i + d_t + \sum_j b_j x_{ijt} + OG_{it} + \varepsilon_{it} \qquad (4.4.2)$$

Equations (4.4.1) and (4.4.2) control for out-of-steady-state dynamics by entering the output gap as an indicator of cyclical position in the regressions. In our notation t , i and j are indexes for time, countries and institutions, and x represents a vector of institutions. The specification allows for country fixed and common time effects. The two dependent variables are extracted from available data on unemployment rates following the methodology described in the appendix of this chapter.

All institutional variables in the model are normalized to have zero mean and unit standard deviation. They are mapped into empirical concepts in the way we turn to describe. Reservation wages z are proxied by benefit replacement rates during the first year of unemployment, as averaged over various family types and earning levels (Nickell and Nunziata, 2001). Firms' firing costs c are approximated by the OECD indicators of Employment Protection Legislation for regular workers (EPL) and the incidence of temporary employment.

The matching efficiency is supposed to be influenced by three categories of Active Labor Market Policies (ALMPs). Given that heterogeneity in the synthetic indicator of ALMPs is one reason for non-significance of these policies in previous studies (Bassanini and Duval, 2006), we proceed with a disaggregated analysis whereby ALMP spending is decomposed into three main categories: job seeker support by Public Employment Services (PES), training policies and financial support to labor-demand. Data on ALMPs programs is also taken from the OECD labor market policies database. As suggested by Estevao (2003), we express ALMP expenditures per unemployed person as a percentage of GDP per capita to ensure cross-country comparability.

We use union centralization and coordination measures (Visser, 2007) to calibrate workers' strength (β) in wage bargaining. Tax wedges comprise social security contributions, personal income and indirect tax rates (Nickell and Nunziata, 2001).

Estimating the model further requires identifying measures for productivity shocks and overall financial conditions, which we take to be the standard deviation of labor productivity[3] and the real interest rate. As a measure of the regulatory stance in product markets, we use the OECD indicator of regulation in Energy, Transport and Communications (ETRC). This is narrower in terms of sectoral coverage than the OECD aggregate indicator for product market regulation but has the advantage of being available in long-time series. Moreover, as the EU Single Market Program has been more effective in liberalizing traditionally monopolistic sectors than in opening up services in general, the potential for reform presented in the following section can be interpreted as a lower bound to overall reform gains.

4.4.2 Results

We estimate the two-equation system on job destruction and unemployment durations given by (4.4.1) and (4.4.2) using data from the 20 OECD countries listed in Section 4.2 over the period 1985-2008. The Simultaneous Equations Model (SEM) estimator we use assumes that the two unemployment components are determined simultaneously by the model postulated in section 4.3. SEM further allows for cross-correlation in the error terms of the two variables of interest. Results are summarized in table 4.4.1.

As expected, unemployment inflows and durations are strongly affected by cyclical conditions. At times where the economy operates above potential output and labor markets are tighter, both job losses and the length of unemployment spells are reduced. The opposite holds during downturns, characterized by negative output gaps, slack labor markets, and higher unemployment inflows and durations.

[3] Conceptually one may distinguish between *aggregate* shocks, interpreted as a general increase or decrease in job productivity, and *reallocation* shocks, interpreted as shocks that increase or decrease the variance of productivity. Consistent with the theoretical approach adopted in this paper, our empirical analysis focuses on reallocation shocks.

For a right interpretation of the link between the set of institutional factors and the two endogenous variables, note that, by construction, the coefficients presented in table Shimer (2012) capture the estimated effect of each policy on unemployment inflows and durations when the economy is producing at full capacity (i.e. zero output gap).

Our estimates for job destruction suggest strong correlations with the tax wedge, EPL (alone and interacted with the incidence of temporary employment) and product market regulation, which we turn to rationalize.

By increasing the price of leisure relative to working activities, a higher tax wedge pushes up workers' reservation wage. Workers at the margin are rendered unprofitable to firms and more jobs are destroyed. There is some evidence that this impact depends on the economic cycle, as indicated by the significance of the coefficient of the interaction term between the tax wedge and the output gap. A positive sign implies that the effect of the tax wedge on job destruction is augmented in periods when labor markets are tight (when the output gap is positive) and less marked in periods with considerable labor market slack (when the output gap is negative). This is consistent with workers attaching greater value to their leisure time during upturns as compared to downturns.

In keeping with previous findings in the literature, we find that stricter job protection rules tend to reduce job losses. However, as indicated by the positive sign of the interaction term between EPL and temporary employment, the dampening effect of EPL on job destruction is reduced in economies characterized by a high share of temporary jobs–where labor legislation tends to differ across permanent and temporary workers.

Hurdles to competition in the services sectors result into higher equilibrium levels of job destruction. This is because regulations that inhibit competition grant more market power to producers, each one charging now higher prices for the goods they sell. But, as aggregate demand is less than supply at the new prices, a number of unprofitable firms will exit the market. And more jobs will be destroyed. Thus stricter regulations lead to market configurations characterized by higher job destruction and lower employment levels.

Consistent with our theoretical priors, ALMP programs (our proxy to matching efficiency) do not appear to have a significant impact on job destruction. Neither could we find evidence of a significant relationship between bargaining coordination and the inflow rate. Using bargaining centralization to measure the capacity of the social dialogue to internalize the impact of wage demands on employment destruction preserved this result. And the coefficients of non-linear specifications testing the Calmfors-Driffill hypothesis turned out to be also non-significant.

Unlike direct indicators of wage bargaining, more generous replacement rates do seem to push up reservation wages and aggravate job destruction. Our estimates could not shed light on the theoretical ambiguity surrounding the influence of productivity changes on job destruction.

Turning to the duration of unemployment, high tax wedges appear to discourage job creation and prolong unemployment spells. By increasing equilibrium wages, higher tax wedges tend to depress labor demand and increase the flow of workers into unemployment. Likewise, the unemployment benefit replacement rate is found to have a positive impact on unemployment duration. This is likely to capture the negative impact of the replacement rate on search incentives, as stressed in earlier works (Bassanini and Duval, 2006; Blanchard and Wolfers, 1999). EPL also has a positive, albeit weakly significant, influence on unemployment duration.

Labor-demand support does not appear to influence unemployment durations while job-seeker support is only borderline significant. The former result is consistent with previous studies pointing to large dead-weight losses and substitution effects associated with employment subsidies. These studies have often been disappointing in terms of bringing the unemployed back into unsubsidized work (**?**).

In contrast with both theoretical priors and economic intuition, more spending on formation policies tends to result in longer unemployment spells. This could partly reflect

Table 4.4.1: *Regression results*

SEM Regression: Job Destruction Rate and Unemployment Duration

	Job Destruction Rate (% of Labor Force)	Unemployment Duration (Months)
Output Gap	-0.0522***	-0.0201***
	[-6.036]	[-2.829]
EPL	-0.1363**	0.0589*
	[-2.449]	[1.712]
EPL*Temporary Employment	0.0630**	0.0132
	[2.493]	[0.640]
ALMP		
-Job Seeker Support	-0.0241	-0.1052*
	[-0.335]	[-1.789]
-Training	-0.0421	0.2043***
	[-0.436]	[2.581]
-Labor Demand Support	0.0791	-0.0573
	[1.064]	[-0.940]
Tax Wedge	0.1338***	0.1028***
	[3.664]	[3.437]
Output Gap*Tax Wedge	0.0055*	0.0017
	[1.796]	[0.294]
Product Market Regulation	0.1191**	0.0358
	[2.403]	[-0.881]
Replacement Rate	0.0624*	0.0705*
	[1.716]	[1.813]
Bargaining Coordination	-0.0085	-0.0396
	[-0.244]	[-1.379]
Real Interest Rate	0.0082	0.0897***
	[0.295]	[3.924]
Productivity Volatility	0.0042	-0.0274*
	[0.234]	[-1.862]
Total Observations	324	324
Adjusted R-squared	0.851	0.9

Note: The sample includes data for 20 OECD countries listed in Section 2. The estimation includes cross-country fixed effects (not displayed) and year dummies as common time effects. Dependent variables are in logs. The t-statistics are reported in parenthesis. Superscripts *, ** and *** indicate that the estimated coefficient is significantly different from zero at 10, 5 and 1 percent levels, respectively.

reverse causality, as those programs can be adjusted more rapidly in response to an economic downturn and longer unemployment spells than, say, EPL or tax wedges. Moreover, training programs may reduce search efforts if not properly designed, and in the case where participation in such programs represents a more attractive alternative to workers than open unemployment, they could even augment wage demands. In any event, any beneficial effects of ALMPs need to be weighed against the costs of taxes required to finance them, which may in turn increase unemployment.

Our estimates confirm the theoretical prediction that job creation increases (duration falls) in response to more frequent changes in productivity while the time spent in looking for a job appears to be inversely related with real interest rates.

One interesting finding from our analysis is that, with the notable exception of EPL and ALMPs, all remaining labor market institutions seem to impact the job destruction rate and the unemployment duration in the same direction. Countries with both high inflows into unemployment and long durations may therefore find efficient to use the policies that affect both dimensions at the same time.

4.4.3 Policy Implications from Some Illustrative Simulations

To get an idea of the model's projection capacity of both inflows and durations, we apply the coefficients estimated intable 4.4.1 to output gap forecasts over the period 2010-12[4]. These estimates are based on the most updated (i.e. 2008) institutional configurations in each country. The projection exercise is meant to illustrate the link between the model's institutional variables and the two unemployment components, rather than to provide accurate forecasts over the referred period. The projections presented in this section should therefore be taken with caution.

Figure 4.4.1 ranks countries according to their expected behavior in terms of inflows and

[4]The cut-off date for the forecasts is July 2011.

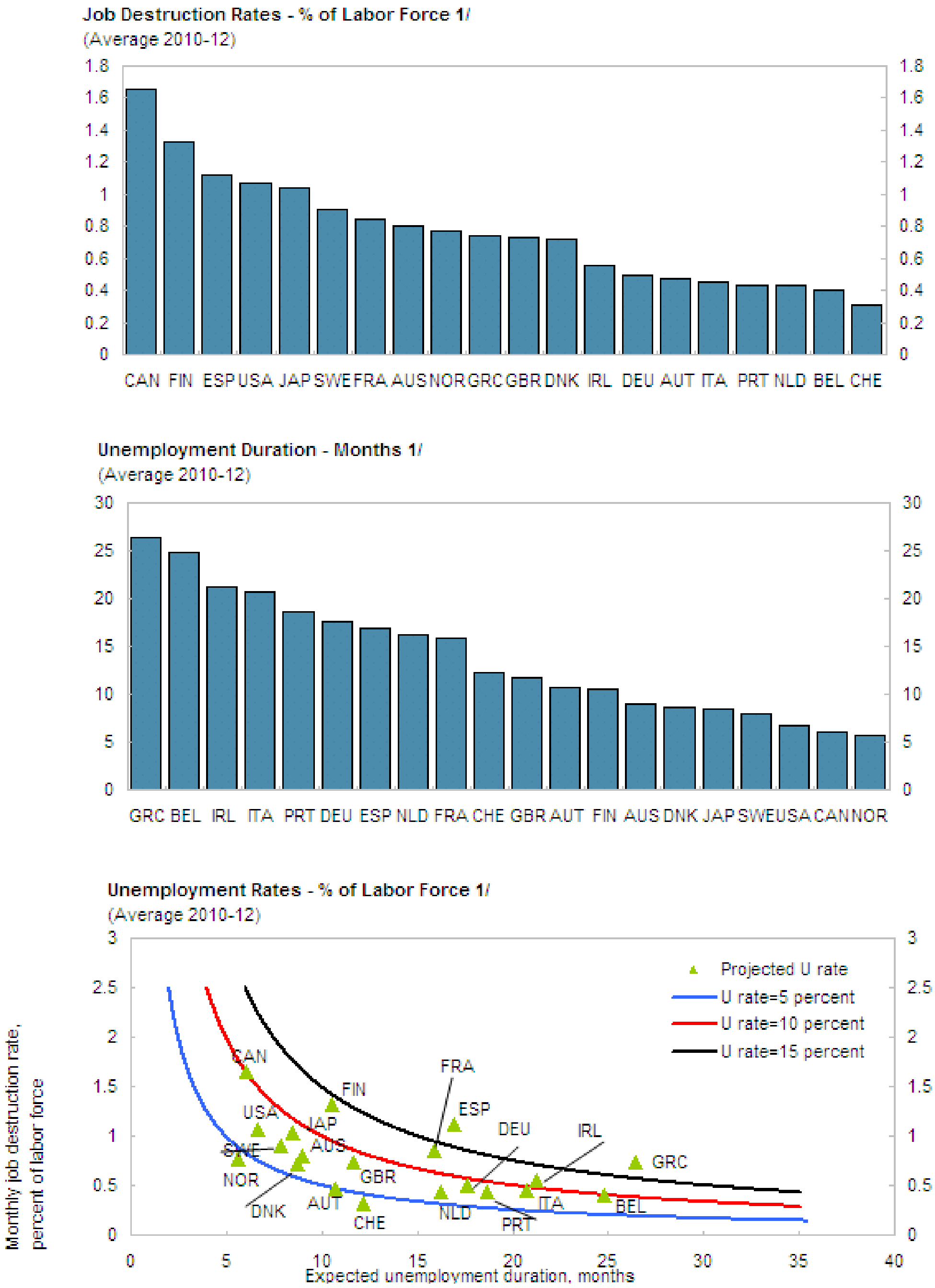
durations. Projections reflect a wide variation in countries' experiences. The duration of unemployment spells ranges from around 26 months in Greece to 5.6 in Norway. Belgium, Ireland, Italy and Portugal also feature comparatively long durations. The shortest unemployment spells are projected for the Nordic and Anglo-Saxon countries. Job destruction rates are envisaged to be the highest in countries as heterogeneous as Canada, Finland, Spain, and the US; and the lowest in Portugal, Netherlands, Belgium, Belgium and Switzerland.

We now turn to calibrate the impact of changes in institutional variables on unemployment inflows and durations. To this aim, we quantify the (percentage point) difference from baseline projections induced by i) a move to best sample practices in services regulation; ii) a 5-percentage-point reduction in tax wedges; iii) a 5-percentage-point reduction in replacement rates; iv) a simulation combining all the above changes. Rather than as a literal description of what every country ought to do, the simulations presented here are meant to illustrate the scope for less employment destruction and lower unemployment duration, as well as to generate a benchmark against which to make cross-country comparisons. Clearly, countries need not to adopt such reform package in full, but may target specific policy areas depending on the nature of the unemployment problem and decide on the pace of reform that best suits their needs.

The results are indicative of substantial gains from reform (figure 4.4.2). Our simulations imply that removing the hurdles to entering key services and alleviating the tax burden on labor can each reduce job destruction rates permanently by 0.05 percentage points on average for the whole sample. And by stimulating job search, lower replacement rates and tax wedges can limit the duration of unemployment spells up to 20 and 50 days, respectively. The combined impact of those changes would yield sizable unemployment reductions in all countries, with Greece, Spain, France, Finland, Ireland, Belgium, and Italy benefiting the most from the implementation of the combined package (figure 4.4.3). Of course, actual unemployment impact will vary with the ambition of the reform agenda, the speed of its implementation, and the time needed for these reforms to take hold.

These reforms could go hand in hand with reducing inequalities. In countries where labor market duality remains unacceptably high, measures to harmonize employment protec-

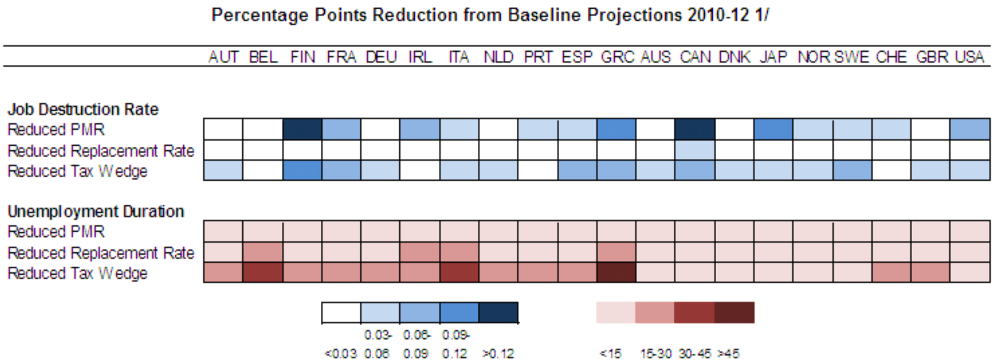
Figure 4.4.1: *Job destruction rates and unemployment duration: baseline Projections*



Source: Authors' Calculations.

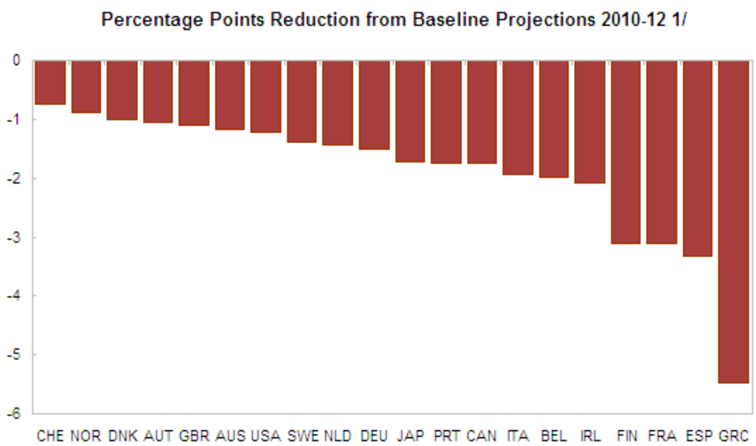
1/ Projections are based on the estimated coefficients reflected in Table 1, 2010-2012 output gap data and 2008 institutional configurations in each country.

Figure 4.4.2: *Reform impact on job destruction rates and unemployment duration*



Source: Authors' Calculations.
1/ The heatmap illustrates changes in job destruction rates and unemployment durations assuming that (i) every country adjusts PMR to best sample practices; (ii) replacement rates are cut down by 5 percentage points across-the-board.
2/ Percentage point reduction for job destruction rates and reduction in days for unemployment duration.

Figure 4.4.3: *Combined reform impact on unemployment rates*



Source: Authors' Calculations.
1/ The simulation illustrates the scope for reform by assuming that tax wedges and benefit replacement rates are each reduced by 5 percentage points in every country; and PMR is adjusted to match best sample practices.

tion benefits between types of job contracts should reduce the disproportionate burden on temporary workers—those last hired and first fired. Our simulations suggest that a reduction in job protection rules for regular workers to cover only $\frac{1}{4}$ of the distance with the least regulated economies would have a neutral impact on unemployment, provided it is accompanied by a reduction in temporary unemployment to match best practices.

4.5 Conclusions

This paper reads the basic unemployment equilibrium condition postulated by search theory as an Iso-Unemployment Curve (IUC). The IUC is the locus of job destruction rates and expected unemployment durations that render the same unemployment level. We use this schedule to classify countries according to their preferences over the job destruction-unemployment duration trade-off. The upshot of this analysis is that labor markets characterized by high levels of job destruction but brief unemployment spells do not necessarily outperform countries characterized by the opposite behavior. But, the IUC construct makes it clear that high unemployment rates result from extreme values in either durations or destructions, or intermediate-to-high levels in both.

Looking at unemployment through the lenses of the IUC schedule focuses the attention on each economy's revealed social preferences over the destruction-duration mix. Policy packages fighting unemployment should take into consideration such preferences. Some countries seem to tolerate relatively high destruction rates as long as unemployment duration is short. Others are biased towards job security and do not mind financing longer job search spells. A few unfortunate countries are trapped in a high inflow-high duration combination, seemingly condemned for long periods of high unemployment.

An optimistic message arising from this study, especially for countries located on higher IUCs, is that an ambitious structural reform program tackling high labor tax wedges, activating unemployment benefits and removing barriers to competition in key services can effectively reduce job losses, limit the duration of unemployment spells and yield substantial reduction in unemployment.

4.A Appendix

Measuring Job Destruction and Unemployment Durations

This appendix presents the methodology used to estimate annual time series of flow hazard rates into and out of unemployment. It also discusses how to infer the average duration of unemployment spells consistent with such flows. The procedure builds on Shimer (2007). However, this approach cannot be directly applied to European countries as unemployment duration is not available at monthly frequencies in the European Labor Force Survey. To overcome this limitation, we follow the methodology proposed by Elsby and others (2008). This methodology exploits annual and quarterly data to measure annual averages of monthly unemployment flows.

Let us denote by $F_t^{<12}$ the probability that an unemployed worker exits unemployment within one year. The annual change in the unemployment stock can be expressed as

$$u_{t+12} - u_t = u_{t+12}^{<12} - F_t^{<12} u_t \tag{4.A.1}$$

Here $u_{t+1}^{<12}$ represents the stock of unemployed workers with duration less than one year (i.e. the yearly flow into unemployment), and $F_t^{<12} u_t$ represents the flows out of unemployment. Solving for the annual outflow probability $F_t^{<12}$, one obtains

$$1 - F_t^{<12} = \frac{u_{t+12} - u_{t+12}^{<12}}{u_t} = \underbrace{\frac{u_{t+12} - u_{t+12}^{<12}}{u_{t+12}}}_{A} \underbrace{\frac{u_{t+12}}{u_t}}_{B} \tag{4.A.2}$$

where the factor A represents the fraction of unemployment with duration longer than one year and the ratio B is the annual gross growth rate of unemployment. Assuming that the monthly outflow hazard rate for workers unemployed less than one year $f_{out,t}^{<12}$ is constant within years, the annual outflow probability $F_t^{<12}$ is related to $f_{out,t}^{<12}$ through

$$e^{-12 f_{out,t}^{<12}} = 1 - F_t^{<12} \tag{4.A.3}$$

so that $F_t^{<12}$ can be mapped into $f_{out,t}^{<12}$ in the following manner

$$f_{out,t}^{<12} = -\ln\left(1 - F_t^{<12}\right)/12 \tag{4.A.4}$$

where $f_{out,t}^{<12}$ is the hazard rate for unemployed workers of duration lower than one year, which is related to the probability that an unemployed worker at time t completes her spell within the subsequent twelve months.

In order to obtain estimates of the corresponding inflow hazard rates $f_{in,t}^{<12}$, let us reformulate the evolution of the monthly unemployment rate over time as

$$\frac{du_t}{dt} = f_{in,t}^{<12}(1-u_t) - f_{out,t}^{<12}u_t \tag{4.A.5}$$

Assuming that the flow hazards are constant within years and solving equation (4.A.5) forward one year, we can relate the variation in the unemployment stock u_t over the course of the year to the variation in the underlying hazard rates $f_{in,t}^{<12}$ and $f_{out,t}^{<12}$:

$$u_t = \lambda_t u_t^* + (1-\lambda_t)u_{t-12} \tag{4.A.6}$$

where the steady-state unemployment rate u_t^* is given by:

$$u_t^* = \frac{f_{in,t}^{<12}}{f_{in,t}^{<12} + f_{out,t}^{<12}} \tag{4.A.7}$$

and the annual rate of convergence to the steady state λ_t is found to be:

$$\lambda_t = 1 - e^{-12(f_{in,t}^{<12} + f_{out,t}^{<12})} \tag{4.A.8}$$

To operationalize the methodology described above, we use OECD annual data on unemployment rates and unemployment rates by duration to compute equation (4.A.2). Given $F_t^{<12}$,we use equation (4.A.4) to estimate $f_{in,t}^{<12}$, which together with u_t allows us to obtain

$f_{out,t}^{<12}$ through equations (4.A.6), (4.A.7) and (4.A.8).

The inflow rates estimated above are combined with annual data for the unemployment rates to estimate through equation (4.A.4) the average duration of unemployment spells for the four decades ranging from the 1970s throughout the 2000s.

Cross Correlation in Error Terms of SEM Regression

Table 4.A.1 reports the correlation coefficients of residual unemployment inflows and durations in the 20 countries of our sample over the four decades ranging from 1965 to 2009.

Table 4.A.1: *Cross correlation in error terms of SEM regression of (4.4.1) and (4.4.2)*

Country	Correlation coefficients of residual job destruction rates and unemployment durations
Australia	-0.4378
Austria	-0.9482***
Belgium	0.1165
Canada	-0.835***
Denmark	0.2212
Finland	-0.7299***
France	-0.9915***
Germany	-0.8207***
Greece	-0.6761***
Ireland	-0.4226*
Italy	-0.4271
Japan	0.3236
Netherlands	-0.235
Norway	-0.6157**
Portugal	-0.4726*
Spain	-0.1038
Sweden	-0.6499*
Switzerland	-0.7417***
United Kingdom	-0.4621***
United States	-0.71***

Note: Superscripts *, ** and *** indicate that the estimated coefficient is significantly different from zero at 10, 5 and 1 percent levels, respectively.

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Eidesstattliche Erklärung

Hiermit erkläre ich, die vorliegende Dissertation selbständig angefertigt und mich keiner anderen als der in ihr angegebenen Hilfsmittel bedient zu haben. Insbesondere sind sämtliche Zitate aus anderen Quellen als solche gekennzeichnet und mit Quellenangaben versehen.

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